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# Association between exposure to health information and mortality: Reduced mortality among women exposed to information via TV programs

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ABSTRACT

Health communication inequality is one of the potential mechanisms linking socioeconomic status (SES) to health disparities. To our knowledge, no previous study has examined the association between exposure to health information and mortality. We analyzed 3-year follow-up cohort data from the Japan Gerontological Evaluation Study (JAGES), involving 8544 males and 9698 females aged 65 years or older, to examine associations between exposure to health information via different types of media and mortality. The baseline survey was conducted from October to December 2013 in 21 municipalities in Japan. Adjusted for health conditions, health behaviors, and other potential confounders, Cox proportional hazards models were used to estimate hazard ratios (HRs) of all-cause mortality. Over a mean of 3.2 years of follow-up, 956 deaths occurred. Among females, receiving health information from TV programs was associated with lower mortality (HR = 0.90; 95% confidence interval [CI] = 0.83, 0.98). By contrast, there were no significant associations among males. Our findings suggest that improving the accuracy of health information delivered via television might be beneficial.

## 1. Introduction

Health communication inequality is one of the potential mechanisms linking socioeconomic status (SES) to health disparities. According to Viswanath, health communication inequalities are defined as: "differences in access to and use of information resources; attention to and processing of health information; and differential capacity to act on the information between different social groups at the individual level and differential capacity in generating, processing, and disseminating information at the group level" (Viswanath, 2006). The model highlights the concern that disparities exist not only in access but also in the capacity and capability to take advantage of new developments in communication science and technology which may ultimately result in increased health disparities (Viswanath et al., 2007).

In general, females, whites (compared to racial/ethnic minorities), and people with more education and higher income seek more health information than their counterparts (Berry et al., 2011; Hay et al., 2009; Hogue et al., 2012; Ishikawa et al., 2016; Ishikawa et al., 2012; Viswanath, 2006; Weaver et al., 2010). Social capital (for example, levels of civic participation, trust in the community, and mutual social supports) promotes health media usage (Ishikawa et al., 2016; Viswanath et al., 2007). It is also known that there are distinct patterns of health communication by gender and that females are more likely to be involved in the exchange of health information with others (Kontos et al., 2011; Mertens et al., 2017; Rodgers et al., 2007; Saint-Charles et al., 2012; Triana et al., 2016). In turn, exposure to health information improves people's knowledge and awareness of health risk (Jung et al., 2013; Viswanath et al., 2006; Wong and Sam, 2010), ultimately promoting changes in risk behaviors such as diet, smoking, disease prevention, and use of medical treatment (Beaudoin and Hong, 2011;

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Finney Rutten et al., 2009; Freisling et al., 2010; Huston et al., 2009; Kim and Jung, 2017; Redmond et al., 2010). Also, knowledge and awareness can aid patients with time-sensitive conditions such as acute coronary and cerebral symptoms in presenting hospital early (Bray et al., 2015; Choudhary et al., 2016; De Silva et al., 2007; Khan et al., 2007), and this can affect survival directly. However, to our knowledge, no previous study has examined the association between exposure to different types of health information and mortality.

Understanding the differences between reliance on alternative forms of media is crucial for the effective delivery of health information to different population groups. Generally speaking, people from higher SES backgrounds tend to prefer printed media (i.e., newspapers and magazines) or the Internet rather than television, while those from lower SES backgrounds tend to prefer television (Beaudoin and Hong, 2011; Viswanath and Ackerson, 2011; Wong and Sam, 2010). Nonetheless, how health outcomes vary according to the use of different media remains unclear. Some studies found that printed media and reliance on information obtained through interpersonal relationships had more favorable associations with health-related knowledge and behaviors than television (Beaudoin and Hong, 2011; Freisling et al., 2010; Kim and Jung, 2017; Redmond et al., 2010; Viswanath et al., 2006). In contrast, other studies pointed out some positive aspects of obtaining health information from television (Jung et al., 2013; Mertens et al., 2017; Paek et al., 2017). Also, there is a growing literature on the impact of the Internet on health (Viswanath, 2005), but the evidence is still mixed as well. Some studies concluded that the Internet exerts a positive influence (Freisling et al., 2010; Hesse et al., 2005; Ishikawa et al., 2012; Viswanath et al., 2006), while others failed to reach the same conclusion (Beaudoin and Hong, 2011; Mertens et al., 2017; Redmond et al., 2010). For example, some studies reported a negative impact of the Internet on uptake of flu vaccination (Jung et al., 2013; Kim and Jung, 2017).

Trust in the information source can affect the awareness of health risk and the adoption of behavior changes (Bleich et al., 2007; Finney Rutten et al., 2009; Huston et al., 2009; Lindström and Janzon, 2007; Liu et al., 2014). According to the World Values Survey (2010–2014) (Inglehart et al., 2014), about 70% of Japanese people had confidence in the press and television either "a great deal" or "quite a lot," whereas only about 20% of American people answered that they had confidence in their press and television. In the survey, the Japanese indicated the highest level of confidence in mass media among 60 surveyed countries.

In this study, we used large cohort data of older Japanese and examined differential associations between exposure to health information and mortality according to the medium.

## 2. Methods

We used cohort data from the Japan Gerontological Evaluation Study (JAGES). The JAGES is an on-going longitudinal study of Japanese people aged 65 or older who were physically and cognitively independent. From October to December in 2013, self-reported questionnaires were mailed to eligible residents in 21 municipalities in 9 (out of 47) prefectures. Random sampling methods were used in 11 large municipalities, while a Census of all eligible residents (65 years or older) were conducted in 10 smaller municipalities. Of 28,448 people invited to participate, 19,909 people returned the questionnaires, corresponding to a response rate of 70.0%. 1085 subjects for which sex and age could not be confirmed or were reported in error were excluded.

### 2.1. Mortality outcome

To ascertain vital status, all participants were linked by the cooperating municipalities to the administrative databases of the public long-term care insurance registers. Of the eligible sample of 18,824 individuals, 18,242 subjects were successfully linked to the administrative records through 2016, corresponding to a follow-up rate of 96.9%. The mean follow-up period was 3.2 years, and we observed 956 deaths during the period.

### 2.2. Exposure to health information

The JAGES asked the subjects to look back over the past one month and answer the following six questions: "How often did you see a news program about health on TV?," "How often did you see an information program about health, doctors or hospitals on TV?," "How often did you read an article about health in a newspaper or magazine for the general?," "How often did you read a magazine or newsletter with a special column on health or medical care?," "How often did you read an article about health in a government-issued announcement or newsletter?," and "How often did you see information about health on the Internet?" Options for each of the questions were "(1) Twice a week or more, (2) About once a week, (3) Less than once a week, and (4) Not at all." We dichotomized answers into from (1) to (3) or (4) and created binary variables for each of the items indicating whether the subject was exposed to health information via the medium at least once in a month. The cutoff point of frequency was determined by referring to previous literature (Hay et al., 2009; Ishikawa et al., 2012; Redmond et al., 2010; Rodgers et al., 2007).

### 2.3. Covariates

Demographic and SES variables included age (65–69, 70–74, 75–79, 80–84, ≥85 years), years of education (low (≤9), middle (10–12), high (≥13)), marital status (married and spouse is alive, other), and annual equivalized household income (low (≤1.9), middle (2.0–3.9), high (≥4.0) million JPY). Variable related to health conditions included self-rated health (very good, good, poor, very poor), depressive symptoms assessed with the short version of the Geriatric Depression Scale (GDS) (non-depressed (< 5), depressed (≥5)) (Burke et al., 1991), instrumental activities of daily living (IADL) measured with the Tokyo Metropolitan Institute of Gerontology Index of Competence (fully capable (5), less capable (< 5)) (Koyano et al., 1991). We also included self-reported disease diagnoses as follows: cancer, heart disease, stroke, diabetes, respiratory disease and others (hypertension, hyperlipidemia, gastrointestinal, liver, or gallbladder disease, kidney or prostate gland disease, musculoskeletal disease, traumatic injury, blood or immune system disease, dementia, Parkinson's disease, eye disease, ear disease, and other). We used three scales to measure social capital, previously validated, consisting of civic participation, social cohesion, and reciprocity (Saito et al., 2016). Civic participation is measured as the number of the following groups engaged in a month: volunteer groups, sports groups, hobby activities, study or cultural groups, and activities for teaching skills. Social cohesion is measured as the number of the subject answering "strongly/moderately agree" in three questions about community trust, norms of reciprocity, and community attachment. Reciprocity is measured as the number of the subject answering "any one or more" in three questions about receiving and providing emotional support and receiving instrumental support. Each variable scores from 0 to 3. Variables related to health behaviors included the frequency of meat or fish intake (≥1/day, < 1/day) and vegetable or fruit intake (≥1/day, < 1/day), walking time (≥30min/day, < 30 min/day), alcohol intake (drinker, ex-drinker, non-drinker), smoking status (smoker, ex-smoker, non-smoker), and body mass index (BMI) (underweight (< 18.5 kg/$m^2$), normal (18.5–24.9 kg/$m^2$), overweight (25.0–29.9 kg/$m^2$), obesity (≥30.0 kg/$m^2$)).

### 2.4. Statistical analyses

First, we carried out an exploratory factor analysis to find patterns of media exposure. The factor analysis enabled us to find similar media in users and cluster them into a factor. It was conducted using maximum likelihood estimation with Promax rotation. The number of factors was restricted to two at a maximum because the estimation appeared to be Heywood case if three or more factors were chosen under the Kaiser-Guttman rule (i.e., a factor should be retained if its eigenvalue is larger than 1). Then we used Cox proportional hazards models to evaluate the associations of all-cause mortality with factor scores (i.e., clustered measures of multiple media exposure) calculated by the factor analysis with use of Bartlett's method. We statistically tested the proportional hazards assumption and confirmed that the assumption was satisfied (see Table A in a supplement). Data were separately analyzed for males and females considering distinct patterns of health communication by gender (Kontos et al., 2011; Mertens et al., 2017; Rodgers et al., 2007; Saint-Charles et al., 2012; Triana et al., 2016). To assure the validity of the gender stratification, we also tested the significance of interaction terms between type of media and gender using the full sample (see Table B in a supplement). Model 1 was adjusted for age, years of education, marital status, and equivalized household income as potential confounders. Model 2 was additionally adjusted for health conditions (self-rated health, depressive symptoms, instrumental activities of daily living (IADL), and self-reported disease diagnoses) as potential confounders. Model 3 was further adjusted for scales of social capital (civic participation, social cohesion, and reciprocity) as potential confounders. Lastly, Model 4 was additionally adjusted for health behaviors (meat/fish intake, vegetable/fruit intake, walking time, alcohol intake, smoking status, and BMI) as potential mediators linking exposure to health information to survival.

In our data, 3682 out of 8544 subjects (43.1%) for males and 5633 out of 9698 subjects (58.1%) for females included some missing values. To mitigate potential biases caused by the missing values, we adopted multiple imputation under the missing at random (MAR) assumption. Incomplete variables were imputed by multivariate normal model using all the covariates as explanatory variables: sex, age, years of education, marital status, equivalized household income, self-rated health, depressive symptoms, instrumental activities of daily living (IADL), self-reported disease diagnoses, civic participation, social cohesion, reciprocity, alcohol intake, smoking status, meat/fish intake, vegetable/fruit intake, walking time, BMI, and exposure to health information (TV news program, TV information program, newspaper or magazine for the general public, magazine or newsletter with a special column, government-issued newsletter, and the Internet). We created 20 imputed datasets, and the estimates were combined. For sensitivity analysis, we performed the same analyses on the subset of subjects without missing values and confirmed that the significance of associations did not change (see Table C in a supplement). In addition, to check robustness of imputation, we carried out sensitivity analysis using an alternative modeling strategy, i.e., multiple imputation by chained equations (MICE). The results from the corresponding model were similar to results from the multivariate normal model (not shown). All analyses were performed with use of Stata 14.2 (Stata Corp, College Station, TX).

## 3. Results

The sample comprised 8544 males and 9698 females. Females were more likely to seek health information through all types of media (except for the Internet) compared to males. For both males (87.4%) and females (89.6%), TV news program was the most prevalent source of health information, followed by TV information program, newspaper or magazine for the general public (Table 1).

Table 2 displays the percentages of those who were exposed to health information more than once in the past one month according to

**Table 1**
Characteristics of older Japanese male and female participants: Japan, 2013.

| | Male (n = 8544) | | Female (n = 9698) | | $\chi^2$ test |
|---|---|---|---|---|---|
| | n | % | n | % | P-value |
| Age (years) | | | | | 0.04 |
| 65-69 | 2405 | 28.1 | 2606 | 26.9 | |
| 70-74 | 2589 | 30.3 | 2922 | 30.1 | |
| 75-79 | 1901 | 22.2 | 2215 | 22.8 | |
| 80-84 | 1134 | 13.3 | 1274 | 13.1 | |
| ≥85 | 515 | 6.0 | 681 | 7.0 | |
| Education (years) | | | | | < 0.001 |
| Low (≤9) | 3251 | 38.1 | 4458 | 46.0 | |
| Middle (10–12) | 2964 | 34.7 | 3519 | 36.3 | |
| High (≥13) | 2147 | 25.1 | 1470 | 15.2 | |
| Missing and other | 182 | 2.1 | 251 | 2.6 | |
| Marital status | | | | | < 0.001 |
| Married and spouse is alive | 7187 | 84.1 | 5690 | 58.7 | |
| Other | 1191 | 13.9 | 3668 | 37.8 | |
| Missing | 166 | 1.9 | 340 | 3.5 | |
| Annual equivalized household income (million JPY) | | | | | < 0.001 |
| Low (≤1.9) | 3490 | 40.8 | 3829 | 39.5 | |
| Middle (2.0–3.9) | 3024 | 35.4 | 2706 | 27.9 | |
| High (≥4.0) | 843 | 9.9 | 804 | 8.3 | |
| Missing | 1187 | 13.9 | 2359 | 24.3 | |
| Self-rated health | | | | | 0.006 |
| Very good | 1006 | 11.8 | 1114 | 11.5 | |
| Good | 5732 | 67.1 | 6621 | 68.3 | |
| Poor | 1379 | 16.1 | 1391 | 14.3 | |
| Very poor | 214 | 2.5 | 211 | 2.2 | |
| Missing | 213 | 2.5 | 361 | 3.7 | |
| Depressive symptoms (GDS) | | | | | 0.03 |
| Not depressed (< 5) | 5453 | 63.8 | 5683 | 58.6 | |
| Depressed (≥5) | 1980 | 23.2 | 1903 | 19.6 | |
| Missing | 1111 | 13.0 | 2112 | 21.8 | |
| Instrumental activities of daily living (IADL) | | | | | < 0.001 |
| Fully capable (5) | 6046 | 70.8 | 8203 | 84.6 | |
| Less capable (< 5) | 2263 | 26.5 | 1179 | 12.2 | |
| Missing | 235 | 2.8 | 316 | 3.3 | |
| Self-reported disease diagnoses | | | | | |
| Cancer (yes) | 392 | 4.6 | 235 | 2.4 | < 0.001 |
| Heart disease (yes) | 1062 | 12.4 | 720 | 7.4 | < 0.001 |
| Stroke (yes) | 376 | 4.4 | 200 | 2.1 | < 0.001 |
| Diabetes (yes) | 1384 | 16.2 | 981 | 10.1 | < 0.001 |
| Respiratory disease (yes) | 549 | 6.4 | 412 | 4.2 | < 0.001 |
| Others (yes) | 5888 | 68.9 | 7094 | 73.1 | < 0.001 |
| Missing | 481 | 5.6 | 669 | 6.9 | |
| Civic participation (number of participating groups in five indicators) | | | | | < 0.001 |
| None | 4273 | 50.0 | 3901 | 40.2 | |
| One | 1396 | 16.3 | 1535 | 15.8 | |
| Two | 850 | 9.9 | 1034 | 10.7 | |
| Over three | 545 | 6.4 | 779 | 8.0 | |
| Missing | 1480 | 17.3 | 2449 | 25.3 | |
| Social cohesion (number of "strongly/moderately agree" in three indicators) | | | | | 0.003 |
| None | 1060 | 12.4 | 1279 | 13.2 | |
| One | 1389 | 16.3 | 1682 | 17.3 | |
| Two | 1821 | 21.3 | 1942 | 20.0 | |
| Three | 3990 | 46.7 | 4269 | 44.0 | |
| Missing | 284 | 3.3 | 526 | 5.4 | |
| Reciprocity (number of "any one or more" in three indicators) | | | | | < 0.001 |
| None | 175 | 2.0 | 81 | 0.8 | |
| One | 411 | 4.8 | 170 | 1.8 | |
| Two | 597 | 7.0 | 591 | 6.1 | |
| Three | 6945 | 81.3 | 8229 | 84.9 | |
| Missing | 416 | 4.9 | 627 | 6.5 | |
| Frequency of meat or fish intake over the past month | | | | | < 0.001 |
| ≥1/day | 3528 | 41.3 | 4605 | 47.5 | |
| < 1/day | 4852 | 56.8 | 4907 | 50.6 | |
| Missing | 164 | 1.9 | 186 | 1.9 | |
| Frequency of vegetable or fruit intake over the past month | | | | | < 0.001 |

(*continued on next page*)

**Table 1** (*continued*)

| | Male (n = 8544) | | Female (n = 9698) | | $\chi^2$ test |
|---|---|---|---|---|---|
| | n | % | n | % | P-value |
| ≥1/day | 6127 | 71.7 | 8080 | 83.3 | |
| < 1/day | 2294 | 26.8 | 1469 | 15.1 | |
| Missing | 123 | 1.4 | 149 | 1.5 | |
| Walking time | | | | | 0.93 |
| ≥30min/day | 6263 | 73.3 | 7028 | 72.5 | |
| < 30min/day | 2152 | 25.2 | 2408 | 24.8 | |
| Missing | 129 | 1.5 | 262 | 2.7 | |
| Alcohol intake | | | | | < 0.001 |
| Drinker | 4757 | 55.7 | 1390 | 14.3 | |
| Ex-drinker | 788 | 9.2 | 164 | 1.7 | |
| Non-drinker | 2912 | 34.1 | 7981 | 82.3 | |
| Missing | 87 | 1.0 | 163 | 1.7 | |
| Smoking | | | | | < 0.001 |
| Smoker | 1566 | 18.3 | 303 | 3.1 | |
| Ex-smoker | 2585 | 30.3 | 225 | 2.3 | |
| Non-smoker | 4286 | 50.2 | 8982 | 92.6 | |
| Missing | 107 | 1.3 | 188 | 1.9 | |
| BMI (kg/m$^2$) | | | | | < 0.001 |
| Underweight (< 18.5) | 443 | 5.2 | 871 | 9.0 | |
| Normal (18.5–24.9) | 5878 | 68.8 | 6283 | 64.8 | |
| Overweight (25.0–29.9) | 1767 | 20.7 | 1713 | 17.7 | |
| Obesity (≥30.0) | 126 | 1.5 | 227 | 2.3 | |
| Missing | 330 | 3.9 | 604 | 6.2 | |
| Exposure to health information in the past one month | | | | | |
| TV news program | | | | | < 0.001 |
| Once or more | 7465 | 87.4 | 8690 | 89.6 | |
| None | 744 | 8.7 | 508 | 5.2 | |
| Missing | 335 | 3.9 | 500 | 5.2 | |
| TV information program | | | | | < 0.001 |
| Once or more | 6841 | 80.1 | 7941 | 81.9 | |
| None | 1330 | 15.6 | 1072 | 11.1 | |
| Missing | 373 | 4.4 | 685 | 7.1 | |
| Newspaper or magazine for the general public | | | | | < 0.001 |
| Once or more | 6698 | 78.4 | 7752 | 79.9 | |
| None | 1495 | 17.5 | 1353 | 14.0 | |
| Missing | 351 | 4.1 | 593 | 6.1 | |
| Magazine or newsletter with a special column | | | | | < 0.001 |
| Once or more | 4209 | 49.3 | 5360 | 55.3 | |
| None | 3949 | 46.2 | 3617 | 37.3 | |
| Missing | 386 | 4.5 | 721 | 7.4 | |
| Government-issued newsletter | | | | | < 0.001 |
| Once or more | 5400 | 63.2 | 6270 | 64.7 | |
| None | 2697 | 31.6 | 2519 | 26.0 | |
| Missing | 447 | 5.2 | 909 | 9.4 | |
| Internet | | | | | < 0.001 |
| Once or more | 1587 | 18.6 | 987 | 10.2 | |
| None | 6471 | 75.7 | 7792 | 80.3 | |
| Missing | 486 | 5.7 | 919 | 9.5 | |

the medium. For each category of characteristics, we compared the percentages using a chi-squared test. People with a high level of education were more likely to receive health information via media. People who rated health conditions as very good or good, those who were fully capable in IADL, and those who were not depressed were more likely to receive health information. Also, variables related to social capital were positively correlated with exposure to health information for all types of the media. Those who practiced healthier behaviors (i.e., non-smoker, taking meat/fish and vegetable/fruit more than once per day, walking more than 30 min per day) were also more likely to report exposure to health information.

Table 3 shows results of an exploratory factor analysis. The factor analysis suggested that two factors existed. For the first factor, variables of TV news program and TV information program had high factor loadings, and thus we labeled it "TV programs." For the second factor, variables of newspaper, magazine, and government-issued newsletter had high factor loadings, and thus we labeled it "printed media." The Internet had low factor loadings for both of the factors. Hence, we treated the Internet independently, and then estimated scoring coefficients of the two factors and factor scores for each of the subjects.

Table 4 shows hazard ratios (HRs) for the association of all-cause mortality with exposure to health information by gender and according to different types of media calculated with use of Cox hazards proportional models. Predictors of TV programs and printed media were the factor scores calculated by the factor analysis, while a predictor of the Internet was a standardized z-score for ease of comparison with factor scores. For males, there were no significant associations. In contrast, for females, we found significant associations with factors of TV programs and printed media in Model 1. For printed media, the significance disappeared after adjusting for health conditions at the baseline as confounders (Model 2). Meanwhile, an association with TV programs was significant in both Model 2 and Model 3 (adjusting for social capital). Also, after adjusting for health behaviors as mediators (Model 4), its beneficial association was remained significant. In Model 4, we found a reduced mortality among females who were exposed to health information in a month through TV programs (HR = 0.90; 95% confidence interval [CI] = 0.83, 0.98). We also tested associations of mortality with individual sources of health information. As well as the result above, exposure to health information via TV news program and TV information program was independently associated with a reduced mortality among females (see Table D in a supplement).

## 4. Discussion

To our knowledge, this is the first study which explored associations between exposure to health information via different media and survival. We found that older Japanese women who were exposed to health information through television programs more than once a month had reduced mortality. In contrast to women, mortality for men was not sensitive to exposure to health information.

The observed differences by gender and by type of media may be explained by differential trust in health information. Our results are consistent with a previous study, which showed that Japanese women were more likely to trust media than men and that television was more trusted as a source of health information than newspapers, magazines, and the Internet (Ishikawa et al., 2012). Surveys suggest that the Japanese report high levels of trust in the public broadcasts from the NHK, Nippon Hoso Kyokai (Japan Broadcasting Corporation) out of all media channels (Japan Press Research Institute, 2017). NHK (like the BBC in Britain) is publicly funded and delivers many health-specific programs about nutrition, exercise, disease prevention and medical treatment without commercial advertisements. Also, it often broadcasts special programs related to health with thorough coverage and compliance with international medical guidelines for an hour. The more credible is a medium, the more it can affect health behaviors. A study found a significant and larger effect of national television news on taking a flu shot than local television news, local newspapers, and the Internet (Jung et al., 2013).

Unlike other media, television conveys wellness information with vivid images and plausible stories. It helps one imagine living a healthy life and shape positive self-perceptions. A previous study showed that positive self-perceptions of aging affected survival (Levy et al., 2002).

**Table 2**
Percentages of those who were exposed to health information more than once a month: Japan, 2013.

| | TV news program | TV information program | Newspaper or magazine for the general public | Magazine or newsletter with a special column | Government-issued newsletter | Internet |
|---|---|---|---|---|---|---|
| Age (years) | | | | | | |
| 65-69 | 91.8 | 84.8 | 83.0 | 53.9 | 66.3 | 17.9 |
| 70-74 | 93.6 | 86.4 | 85.1 | 56.7 | 70.2 | 15.8 |
| 75-79 | 94.2 | 88.4 | 84.6 | 59.3 | 71.8 | 13.6 |
| 80-84 | 93.0 | 86.8 | 82.3 | 56.3 | 70.9 | 12.4 |
| ≥85 | 88.1 | 79.3 | 76.9 | 47.1 | 63.2 | 12.7 |
| P-value for χ2 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| Education (years) | | | | | | |
| Low (≤9) | 91.8 | 84.1 | 78.1 | 50.5 | 64.4 | 8.9 |
| Middle (10–12) | 94.1 | 87.3 | 87.2 | 58.7 | 72.9 | 15.8 |
| High (≥13) | 93.1 | 87.9 | 89.1 | 62.0 | 72.8 | 27.3 |
| P-value for χ2 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| Marital status | | | | | | |
| Married and spouse is alive | 93.1 | 86.4 | 85.1 | 56.3 | 70.5 | 16.9 |
| Other | 92.6 | 85.4 | 80.0 | 54.8 | 66.1 | 11.1 |
| P-value for χ2 | 0.32 | 0.09 | < 0.001 | 0.08 | < 0.001 | < 0.001 |
| Annual equivalized household income (million JPY) | | | | | | |
| Low (≤1.9) | 92.7 | 85.3 | 81.2 | 54.0 | 67.9 | 12.1 |
| Middle (2.0–3.9) | 94.0 | 88.0 | 87.6 | 58.4 | 73.4 | 19.6 |
| High (≥4.0) | 92.1 | 87.0 | 86.9 | 59.6 | 68.3 | 21.8 |
| P-value for χ2 | 0.003 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| Self-rated health | | | | | | |
| Very good | 91.6 | 85.0 | 85.0 | 59.1 | 69.9 | 18.4 |
| Good | 93.4 | 86.5 | 85.1 | 56.6 | 70.5 | 15.6 |
| Poor | 91.7 | 85.0 | 78.0 | 51.9 | 63.8 | 11.9 |
| Very poor | 88.3 | 81.2 | 67.3 | 40.6 | 52.4 | 11.1 |
| P-value for χ2 | < 0.001 | 0.003 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| Depressive symptoms (GDS) | | | | | | |
| Not depressed (< 5) | 93.8 | 87.5 | 86.7 | 59.0 | 72.7 | 16.9 |
| Depressed (≥5) | 90.0 | 82.2 | 75.5 | 47.2 | 59.9 | 12.1 |
| P-value for χ2 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| Instrumental activities of daily living (IADL) | | | | | | |
| Fully capable (5) | 94.2 | 87.7 | 86.5 | 59.2 | 72.5 | 16.6 |
| Less capable (< 5) | 87.4 | 79.3 | 71.8 | 42.5 | 55.9 | 10.5 |
| P-value for χ2 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| Self-reported disease diagnoses | | | | | | |
| Cancer | | | | | | |
| Yes | 92.4 | 86.0 | 86.3 | 56.9 | 66.9 | 19.1 |
| No | 92.9 | 86.1 | 83.5 | 55.8 | 69.3 | 15.1 |
| P-value for χ2 | 0.63 | 0.96 | 0.07 | 0.59 | 0.22 | 0.01 |
| Heart disease | | | | | | |
| Yes | 92.7 | 85.2 | 81.7 | 54.2 | 69.2 | 15.9 |
| No | 92.9 | 86.2 | 83.8 | 56.0 | 69.2 | 15.2 |
| P-value for χ2 | 0.78 | 0.26 | 0.03 | 0.15 | 0.98 | 0.49 |
| Stroke | | | | | | |
| Yes | 90.5 | 83.7 | 78.5 | 52.5 | 66.2 | 12.9 |
| No | 92.9 | 86.2 | 83.7 | 55.9 | 69.3 | 15.4 |
| P-value for χ2 | 0.03 | 0.11 | 0.001 | 0.12 | 0.13 | 0.13 |
| Diabetes | | | | | | |
| Yes | 91.9 | 84.8 | 81.2 | 54.7 | 65.7 | 15.9 |
| No | 93.0 | 86.3 | 84.0 | 56.0 | 69.7 | 15.2 |
| P-value for χ2 | 0.07 | 0.07 | 0.001 | 0.26 | < 0.001 | 0.37 |
| Respiratory disease | | | | | | |
| Yes | 92.1 | 85.2 | 79.8 | 52.1 | 64.6 | 13.7 |
| No | 92.9 | 86.2 | 83.8 | 56.0 | 69.5 | 15.4 |
| P-value for χ2 | 0.35 | 0.43 | 0.002 | 0.02 | 0.003 | 0.19 |
| Others | | | | | | |
| Yes | 93.4 | 86.9 | 84.1 | 56.9 | 70.1 | 15.2 |
| No | 91.1 | 83.7 | 82.0 | 52.5 | 66.2 | 15.5 |
| P-value for χ2 | < 0.001 | < 0.001 | 0.002 | < 0.001 | < 0.001 | 0.70 |
| Civic participation (number of participating groups in five indicators) | | | | | | |
| None | 90.9 | 83.1 | 78.8 | 50.4 | 63.1 | 11.7 |
| One | 94.4 | 87.6 | 87.3 | 57.4 | 72.0 | 17.8 |
| Two | 95.0 | 89.2 | 90.3 | 62.0 | 76.8 | 20.6 |
| Over three | 96.7 | 92.4 | 94.0 | 72.0 | 83.4 | 28.5 |
| P-value for χ2 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| Social cohesion (number of "strongly/moderately agree" in three indicators) | | | | | | |
| None | 89.0 | 80.4 | 75.6 | 48.4 | 55.9 | 12.5 |
| One | 91.2 | 83.8 | 80.3 | 52.2 | 65.0 | 13.3 |
| Two | 92.8 | 85.8 | 84.5 | 53.3 | 68.8 | 15.3 |
| Three | 94.7 | 88.9 | 87.2 | 60.7 | 75.0 | 16.9 |
| P-value for χ2 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| Reciprocity (number of "any one or more" in three indicators) | | | | | | |

(*continued on next page*)

**Table 2** (*continued*)

| | TV news program | TV information program | Newspaper or magazine for the general public | Magazine or newsletter with a special column | Government-issued newsletter | Internet |
|---|---|---|---|---|---|---|
| None | 78.4 | 70.5 | 64.7 | 41.3 | 41.7 | 8.1 |
| One | 83.3 | 71.3 | 65.8 | 37.1 | 48.3 | 12.8 |
| Two | 89.2 | 80.3 | 74.8 | 44.4 | 57.2 | 13.1 |
| Three | 93.8 | 87.6 | 85.6 | 57.9 | 71.6 | 15.8 |
| P-value for $\chi2$ | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| Frequency of meat or fish intake over the past month | | | | | | |
| ≥1/day | 93.8 | 87.5 | 86.8 | 59.6 | 72.1 | 17.3 |
| < 1/day | 92.0 | 84.8 | 81.0 | 52.8 | 66.8 | 13.6 |
| P-value for $\chi2$ | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| Frequency of vegetable or fruit intake over the past month | | | | | | |
| ≥1/day | 93.7 | 87.1 | 85.6 | 57.6 | 71.1 | 15.9 |
| < 1/day | 89.7 | 81.9 | 76.3 | 49.1 | 61.3 | 13.0 |
| P-value for $\chi2$ | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| Walking time | | | | | | |
| ≥30min/day | 93.4 | 86.9 | 85.5 | 57.8 | 71.3 | 16.2 |
| < 30min/day | 91.2 | 83.7 | 78.2 | 50.6 | 63.1 | 12.8 |
| P-value for $\chi2$ | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| Alcohol intake | | | | | | |
| Drinker | 92.0 | 85.1 | 84.3 | 54.3 | 69.1 | 19.2 |
| Ex-drinker | 91.5 | 84.9 | 80.0 | 54.4 | 65.4 | 19.0 |
| Non-drinker | 93.4 | 86.7 | 83.5 | 56.9 | 69.5 | 12.7 |
| P-value for $\chi2$ | 0.001 | 0.01 | 0.004 | 0.01 | 0.04 | < 0.001 |
| Smoking | | | | | | |
| Smoker | 87.6 | 78.3 | 74.0 | 43.7 | 59.4 | 12.6 |
| Ex-smoker | 91.6 | 84.1 | 80.9 | 49.8 | 65.4 | 19.4 |
| Non-smoker | 93.8 | 87.6 | 85.5 | 58.9 | 71.4 | 14.7 |
| P-value for $\chi2$ | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 | < 0.001 |
| BMI ($kg/m^2$) | | | | | | |
| Underweight (< 18.5) | 92.3 | 85.2 | 82.4 | 52.3 | 68.1 | 13.7 |
| Normal (18.5–24.9) | 93.1 | 86.4 | 84.8 | 56.8 | 70.3 | 15.8 |
| Overweight (25.0–29.9) | 93.0 | 86.4 | 81.6 | 55.8 | 68.2 | 15.1 |
| Obesity (≥30.0) | 92.6 | 84.4 | 81.4 | 52.7 | 60.9 | 14.0 |
| P-value for $\chi2$ | 0.71 | 0.51 | < 0.001 | 0.01 | < 0.001 | 0.20 |

Also, people may watch health programs on TV with their families in a living room. Such a situation would facilitate discussions on health issues and the adoption of health behaviors (Mertens et al., 2012).

One may question how exposure to health information could impact survival in such a short period. We propose that health information seeking behavior is a lifelong habit and that it did not necessarily bring about effects just within the follow-up period. The influence of health information could have been already manifest as favorable health behaviors at the baseline. As we presented in Table 2, there were strong correlations between exposure to health information and healthier behaviors though we cannot rule out reverse causation due to the study design.

This study has several limitations. First, we cannot infer causality from the results. Although the subjects were physically and cognitively independent at the baseline and we controlled for various potential confounders including health conditions, we cannot eliminate the possibility viewing TV programs is endogenous, i.e., health-conscious people watch these programs. Second, we considered health conditions and social capital as confounders and health behaviors as mediators. However, we could not distinguish between the confounding versus mediation model because we did not follow up those variables in the study period. Also, we were not free from model misspecification due to complex pathways of health communication. Third, exposure to health information and other covariates were self-reported, and thus they could be biased (e.g., health-conscious people recalled viewing more programs). Fourth, we did not have detailed information such as the length of time watching TV, the quality of the contents, health literacy, and actual changes in health behaviors after exposure to health information. Thus, further studies are needed to reveal the mechanisms for the reduction in mortality among those who watched health programs on TV. Finally, some variables were missed, and we multiply imputed them under the MAR assumption. However, the missing mechanism may depend on the value of the missing data, i.e., missing not at random (MNAR). In this case, our estimation can be biased.

In summary, we found significant associations of TV programs with survival among different forms of media. Health information via television can reach out to people with low SES (Beaudoin and Hong, 2011; Hay et al., 2009; Wong and Sam, 2010). Therefore, improving the

**Table 3**
Exploratory factor analysis using Promax rotation.

| | Rotated factor loading | | Scoring coefficient | |
|---|---|---|---|---|
| | TV programs (Factor 1) | Printed media (Factor 2) | TV programs (Factor 1) | Printed media (Factor 2) |
| TV news program | **0.87** | −0.06 | 0.88 | −0.13 |
| TV information program | **0.66** | 0.14 | 0.32 | 0.15 |
| Newspaper or magazine for the general public | 0.16 | **0.59** | 0.05 | 0.54 |
| Magazine or newsletter with a special column | −0.06 | **0.72** | −0.04 | 0.57 |
| Government-issued newsletter | 0.02 | **0.58** | −0.003 | 0.38 |
| Internet | −0.04 | 0.29 | | |

Note. Factor loadings greater than 0.40 were indicated by boldface.

**Table 4**
Adjusted hazard ratios (HRs) for associations of all-cause mortality with exposure to health information by gender and by types of media: Japan, 2013 through 2016.

| | Model 1 | | Model 2 | | Model 3 | | Model 4 | |
|---|---|---|---|---|---|---|---|---|
| | HR | 95% CI | HR | 95% CI | HR | 95% CI | HR | 95% CI |
| **Males** | | | | | | | | |
| TV programs | 1.00 | (0.94–1.07) | 1.03 | (0.96–1.10) | 1.04 | (0.97–1.11) | 1.04 | (0.97–1.11) |
| Printed media | 0.96 | (0.90–1.02) | 1.01 | (0.94–1.08) | 1.02 | (0.95–1.09) | 1.03 | (0.96–1.10) |
| Internet | 1.00 | (0.92–1.08) | 1.02 | (0.94–1.10) | 1.02 | (0.94–1.10) | 1.02 | (0.94–1.11) |
| **Females** | | | | | | | | |
| TV programs | 0.86*** | (0.80–0.93) | 0.90** | (0.83–0.97) | 0.90** | (0.82–0.97) | 0.90* | (0.83–0.98) |
| Printed media | 0.86** | (0.79–0.94) | 0.92 | (0.84–1.01) | 0.92 | (0.84–1.01) | 0.94 | (0.86–1.04) |
| Internet | 0.95 | (0.83–1.09) | 0.98 | (0.85–1.13) | 0.99 | (0.86–1.14) | 0.99 | (0.86–1.15) |

Note. HRs were calculated with use of Cox proportional hazards models. Predictors of TV programs and printed media were factor scores calculated by an exploratory factor analysis, while a predictor of the Internet was a standardized z-score. 95% confidence intervals (CI) are in parentheses. Model 1 (demographic, SES): adjusted for age, years of education, marital status, and annual equivalized household income. Model 2 (health conditions): model 1 + self-rated health, depressive symptoms, instrumental activities of daily living (IADL), and self-reported disease diagnoses. Model 3 (social capital): model 1 + civic participation, social cohesion, and reciprocity. Model 4 (health behaviors): model 3 + meat/fish intake, vegetable/fruit intake, walking time, alcohol intake, smoking status, and BMI. *P < 0.05; **P < 0.01; ***P < 0.001.

accuracy of health information delivered via television might be a worthwhile investment to shrink health disparities.

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## Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.socscimed.2018.12.019.

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# Relative Deprivation, Poverty, and Mortality in Japanese Older Adults: A Six-Year Follow-Up of the JAGES Cohort Survey

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**Abstract:** Most studies have evaluated poverty in terms of income status, but this approach cannot capture the diverse and complex aspects of poverty. To develop commodity-based relative deprivation indicators and evaluate their associations with mortality, we conducted a 6-year follow-up of participants in the Japan Gerontological Evaluation Study (JAGES), a population-based cohort of Japanese adults aged 65 and older. We analyzed mortality for 7614 respondents from 2010 to 2016. Cox regression models with multiple imputation were used to estimate hazard ratios (HRs) for mortality. Seven indicators were significantly associated with mortality: no refrigerator, no air conditioner, cut-off of essential services in the past year for economic reasons, and so on. Among participants, 12.0% met one item, and 3.3% met two items or more. The HRs after adjusting for relative poverty and some confounders were 1.71 (95%CI: 1.18–2.48) for relative deprivation, and 1.87 (95%CI: 1.14–3.09) for a combination of relative poverty and deprivation. Relative deprivation was attributable to around 27,000 premature deaths (2.3%) annually for the older Japanese. Measurement of relative deprivation among older adults might be worthwhile in public health as an important factor to address for healthy aging.

**Keywords:** relative deprivation; material poverty; relative poverty; mortality; older people

## 1. Introduction

Several studies have shown that relative poverty based on low income is significantly associated with poor health [1]. However, a relative income approach has limitations when attempting to capture the diverse and complex aspects of poverty. In order to assess poverty line, researchers of poverty have proposed the concept of relative deprivation to measure the lack of a living standard that most people in society enjoy. Townsend reported that people experience relative deprivation when they lack the resources to follow a proper diet, cannot participate in activities, and do not have living conditions and amenities that are customary, or are at least widely encouraged, in the societies in which they belong [2]. It has also been suggested that people who live in relative deprivation have different characteristics from those living in relative poverty [3–5].

The association between relative deprivation and mortality was examined by some ecological studies. The use of social indicator approaches such as the Townsend deprivation index or Carstairs deprivation score has found that people in relatively deprived areas have a higher risk for standardized mortality rates [6,7], cancer mortality rates [8], and suicide rates [9]. However, considering the possibility of ecological fallacy, analysis at an individual level is needed. Although some Japanese cross-sectional studies have shown an association with individual poor subjective health [10,11], to our knowledge, no study has examined the association between relative deprivation and premature death at the individual level.

Therefore, we examined the association between relative deprivation as one aspect of poverty and mortality among older Japanese adults after controlling for relative income poverty and other confounding factors.

## 2. Materials and Methods

### 2.1. Study Design and Participants

We used prospective cohort data from the Japan Gerontological Evaluation Study (JAGES), a large-scale population-based study of Japanese people aged 65 or older who were physically and cognitively independent. Baseline data were collected from August 2010 to January 2012, with a response rate of 66.3%. Of these, the present analyses used data on 7614 participants who answered a relative deprivation questionnaire, after excluding participants with missing information on sex and age. The average age of the respondents was 73.5 years (standard deviation (SD) = 5.6), and 53.0% were women. This study was performed on the basis of a collaborative research agreement with the municipality. Ethical permission (No. 13–14) was provided by the Ethics Board at Nihon Fukushi University.

### 2.2. Mortality Outcome

We retrieved information on death records from 2010 to 2016 from the government database of public long-term care insurance. This government database covers all respondents. Among these records, there were 514 (6.75%) deaths identified in the analysis sample.

### 2.3. Relative Deprivation and Relative Poverty

Commodity-based relative deprivation indicators have been developed using a consensual approach based on public opinion. These indicators were drawn from a review of daily necessities and basic needs in society [4,10,12–16]. Based on these previous papers, we evaluated 13 indices that equated with “lack of daily necessities,” “lack of living environment,” and “lack of social life” due to economic reasons. These factors are associated with a low standard of living in current Japanese society [5,11]. Lack of daily necessities indicators included having no television, no refrigerator, no air conditioner, no microwave oven, or no water heater due to economic reasons. Lack of living environment indicators included having no private toilet, kitchen, or bathroom in the house, and having a dining room that was not separate from the bedroom. Lack of social life indicators included having no telephone or ceremonial dress, being absent from family celebrations and events during the previous year due to economic reasons, and having essential services such as water, electricity, or gas cut-off in the previous year (except in cases of forgetting to make a payment). The relative deprivation index was assessed by counting the number of these items that the respondent experienced.

Relative poverty was defined as an income of less than half of the median annual equivalent income in the government statistics [17]; the threshold was 1.49 million Japanese yen in 2009. This definition is accepted by the Organisation for Economic Co-operation and Development (OECD), and is conceptually based on the relative approach of the Luxembourg Income Study [18]. We used annual pre-tax household income. For each response, we calculated the equivalent household income by dividing the income by the square root of the number of household members. This square root scale implies that, for instance, a household of four persons has needs twice as large as one composed of a single person (it is not quadrupled).

### *2.4. Covariates*

Demographic variables included sex, age, years of education, and marital status at the baseline survey. In order to account for the health status at the baseline, presence of medical treatment, self-recognition of forgetfulness, and depressive symptoms were also considered. Medical treatment was determined by asking, "Are you currently receiving any medical treatment?". Self-recognition of forgetfulness was measured by asking, "Do people around you notice your forgetfulness, for example, by telling you that you often ask the same thing?". Depressive symptoms were assessed using the short version of the Geriatric Depression Scale (GDS-15), which was developed for self-administration in the community, using a simple yes/no format [19].

### *2.5. Statistical Analysis*

First, we extracted the relative deprivation index that was associated with premature death among older Japanese adults by calculating crude hazard ratios (HRs) for mortality using Cox regression analysis. Second, we applied Cox regression analysis, starting with assessing the relationship between relative poverty (monetary poverty) and mortality, adjusting for the above covariates (Model 1). In Model 2, we assessed the association between relative deprivation and mortality, adjusting for relative poverty and covariates. The relative deprivation index was divided to ternary (none, only one, two and over) in Model 2. In addition, we examined the binary category (none or anyone) in Model 2a, and the quaternion category (none, one, two, three and over) in Model 2b, respectively. We also examined the combination of relative deprivation and poverty in Model 3.

To mitigate potential biases caused by missing information in predictors and covariates, we adopted the multiple imputation approach, under the missing at random assumption. We generated 20 imputed data sets using the multiple imputation by chained equations procedure. Finally, we calculated population attributable risk percentage (PAR%) in an older Japanese population. This estimation assumed that the adjusted HRs truly reflected causal impact, and that our results represented the entire older Japanese population. Data on annual mortality were obtained from governmental reports [20]. We used STATA 15.1 for all analyses.

## 3. Results

The crude HR showed that the seven relative deprivation indicators were significantly associated with a higher risk for death, respectively (Table 1 & Table S1). In particular, the experience of cut-off of essential services in the past year had a higher risk for premature death compared with other deprivation factors. A total of 15.3%, 3.3%, and 1.5% of respondents reported experiencing a single deprivation item, two or more deprivation items, and three or more deprivation items, respectively. A significantly greater HR for higher mortality was seen for these subjects, compared with non-deprived people: 1.53 (95% confidence interval (CI): 1.24–1.90), 2.10 (95%CI: 1.46–3.02), and 2.36 (95%CI: 1.41–3.95), respectively.

**Table 1.** Relative deprivation index—its prevalence and association with mortality.

| Item | Category | % | Mortality | |
|---|---|---|---|---|
| | | | **Crude HR** | **(95%CI)** |
| No television | No | 98.1 | ref. | |
| | Yes (+) | 1.9 | 1.83 * | (1.12–2.96) |
| No refrigerator | No | 98.9 | ref. | |
| | Yes (+) | 1.1 | 2.01 * | (1.11–3.65) |
| No air conditioner | No | 95.5 | ref. | |
| | Yes (+) | 4.5 | 1.51 * | (1.07–2.13) |
| No private bathroom | No | 93.2 | ref. | |
| | Yes (+) | 6.8 | 1.45 * | (1.08–1.96) |
| No ceremonial dress | No | 98.6 | ref. | |
| | Yes (+) | 1.4 | 1.84 * | (1.04–3.27) |
| Absence from family ceremonial occasions | No | 94.6 | ref. | |
| | Yes (+) | 5.4 | 1.65 ** | (1.21–2.26) |
| Cut-off of essential services in the past year | No | 98.9 | ref. | |
| | Yes (+) | 1.1 | 2.12 * | (1.17–3.85) |
| Relative deprivation index [a] | None | 84.7 | ref. | |
| | 1+ | 15.3 | 1.53 *** | (1.24–1.90) |
| | None | 84.7 | ref. | |
| | 1 | 12.0 | 1.38 * | (1.07–1.77) |
| | 2+ | 3.3 | 2.10 *** | (1.46–3.02) |
| | None | 84.7 | ref. | |
| | 1 | 12.0 | 1.38 * | (1.07–1.77) |
| | 2 | 1.9 | 1.90 * | (1.16–3.14) |
| | 3+ | 1.5 | 2.36 ** | (1.41–3.95) |

*** $p < 0.001$, ** $p < 0.01$, * $p < 0.05$. HR: hazard ratio, 95%CI: 95% confidence interval. (+) is related to relative deprivation. [a] This index was assessed by counting the number of items.

After adjustment for individual attributes and relative poverty (Table 2), among respondents with two or more deprivation items, mortality risk was 1.71 (95%CI: 1.18–2.48) times higher than that in non-deprived subjects (Model 2). Respondents with only one deprivation item did not have a significantly higher mortality risk. In addition, the adjusted HRs were 1.62 (95%CI: 0.97–2.69) for respondents with two items, and 1.82 (95%CI: 1.10–3.01) for those with three or more items (Appendix A, Table A1). Relative poverty had a marginally significant association with mortality. In addition, respondents who fell under both relative deprivation and relative poverty had 1.87 (95%CI: 1.14–3.09) times higher mortality risk compared with those who fell under non-deprivation and poverty. The HR for relative deprivation was comparatively higher than that for relative poverty (Model 3). When we analyzed raw data that did not impute missing values, the major results and trends were similar to those reported above (Appendix A, Table A2).

The estimation of PAR% showed that about 27,000 premature deaths (2.3% of all deaths) or about 15,000 premature deaths (1.2% of all deaths) could be avoided annually if there was less severe relative deprivation in Japan (Table 3).

**Table 2.** Hazard ratios (HRs) for association of mortality with relative deprivation in multiple-imputed dataset [a,b].

| | Model 1 | Model 2 | Model 3 |
|---|---|---|---|
| | HR (95%CI) | HR (95%CI) | HR (95%CI) |
| Relative poverty | | | |
| Non-poverty | ref. | ref. | |
| Poverty | 1.26 * (1.02–1.56) | 1.22 † (0.98–1.53) | |
| Relative deprivation [c] | | | |
| Non-deprivation | | ref. | |
| 1 | | 1.14 (0.87–1.49) | |
| 2 + | | 1.71 ** (1.18–2.48) | |
| Combination [d] | | | |
| No dep. & pov. | | | ref. |
| Poverty only | | | 1.22 † (0.98–1.52) |
| Deprivation only | | | 1.86 † (0.92–3.76) |
| Pov. & dep. | | | 1.87 * (1.14–3.09) |

** $p < 0.01$, * $p < 0.05$, † $p < 0.10$. HR: hazard ratio, 95%CI: 95% confidence interval. [a] Multiple imputation by chained equations was performed using relative deprivation index, relative poverty, sex, age, years of education, marital status, disease and/or impairment, self-recognition of forgetfulness, depressive symptoms (m = 20). [b] Sex, age, years of education, marital status, disease and/or impairment, self-recognition of forgetfulness, and depressive symptoms were controlled. [c] This index was assessed by counting the number of items. [d] Relative deprivation in combination variable was defined as respondents who fell under two and over deprivation index. Proportions of each category were as follows: No dep. & pov.: 70.1%; poverty only: 27.0%; deprivation only: 1.1%; and pov. & dep.: 1.7%. In addition, the proportion is not coincident with other tables, because it was confined to the respondents which answered the relative deprivation and poverty index.

**Table 3.** Estimated population attributable risks (PARs) in Japan.

| | Mortality | | | |
|---|---|---|---|---|
| | % Exposed [a] | HR | PAR | |
| | | | % [b] | n [c] |
| Relative deprivation (1+) | 15.3 | 1.25 | 3.7 | 44,197 |
| (2+) | 3.3 | 1.71 | 2.3 | 27,465 |
| (3+) | 1.5 | 1.82 | 1.2 | 14,577 |
| Relative poverty | 18.0 | 1.22 | 3.8 | 45,698 |

[a] The % exposed of relative deprivation is in our study participants. That of relative poverty is from Japanese official statistics (comprehensive survey of living conditions). [b] PAR(%) = $Pe(HR - 1)/(Pe(HR - 1) + 1)$; $Pe$: the proportion of exposure in the target population; $HR$: hazard ratio. [c] The denominator is the annual number of mortality among people 65 years and older in 2015 (N = 1,199,686) which was obtained from governmental reports.

## 4. Discussion

Relative deprivation is an important element in poverty, although it might be unsuitable for an international comparative study because the standard decent life varies by nation, culture, and period. To the best of our knowledge, this is the first study to examine the effect and impact of relative deprivation and relative poverty on mortality among older Japanese adults. Our results suggest the effectiveness of our seven relative deprivation indicators as social determinants of healthy ageing. Our findings are consistent with previous findings that analyzed poor social support [5,21] and subjective health [10,11]. Our results also showed that the association between the relative deprivation index and premature death remained even after adjusting for monetary poverty.

At the same time, our study added new evidence that relative deprivation has a stronger association with mortality than relative poverty if subjects experience relative deprivation in two or more items. There are several possible reasons for this finding. First, a relative deprivation index might capture severe poverty conditions better than a relative income approach. Whelan et al. revealed that people living in relatively deprived conditions experienced long-term and severe poverty throughout their life course [13]. Some studies also reported that people who have overlapping multidimensional disadvantages are more likely to be socially excluded, have poor self-rated health, and experience psychological distress [3,22]. Second, unlike monetary poverty, a poor standard of living (which relative deprivation measures) might be closely related to unhealthy lifestyles, including poor eating habits and nutrition, and lack of access to healthcare and welfare services. Third, relative deprivation might increase psychosomatic stresses and anxieties related to complaints or dissatisfaction with life [23,24]. Among Japanese older adults, a lower relative income compared with their reference group was associated with the onset of functional disability and death from cardiovascular diseases, regardless of the amount of the objective income [25,26].

Although the Europe 2020 strategy has adopted a strong material deprivation index as a goal for social inclusion in the next decade, there are few discussions regarding specific policy in Japan. Our results suggest that relative deprivation indicators could more accurately represent severe or absolute poverty in society that relative poverty indicators cannot address. The income approach has some limitations in the discussion of the poverty line, because the relationship between income and consumption behavior is complex. Relative deprivation indicators, which are composed of specific primary goods and resources, might be relevant to capture their part of "capabilities" [27]. Our results showed that subjects that fulfilled two or more deprivation indicators had a higher mortality risk, although there are some discussions about the cut-off point of relative deprivation indicators. The proportion of relative deprivation indicators was low, but PAR% was not. It is important to assess and discuss relative deprivation in addition to the conventional relative income approach.

This study has several limitations. First, although we included indicators used in previous studies, our relative deprivation indicators did not cover the full range of daily resources among older people in Japan. Second, our analysis was limited to all-cause mortality. Third, our findings may be underestimated because people living in serious poverty and deprivation may have been less likely to participate in our survey. Finally, our data were not representative of the whole country. On the other hand, it is important to note that we did perform a large-scale survey concerning non-monetary poverty among older people in more than one municipality. Only a few studies have focused on relative deprivation among Asian older adults [5,11,28].

## 5. Conclusions

It is well-known that poverty is one of the social determinants of health. One important implication of our findings is that measurement of relative deprivation, along with relative poverty (monetary poverty), might be worthwhile in public health as an important factor for healthy aging. From a life course perspective, the impact of relative deprivation on health should be evaluated in older people as one of the cumulative disadvantages.

**Supplementary Materials:** The following are available online at http://www.mdpi.com/1660-4601/16/2/182/s1, Table S1: Original thirteen relative deprivation index; its prevalence and association with mortality.

**Author Contributions:** All authors contributed to the conception and design of this study. Data collection was primarily conducted by M.S., N.K., and K.K. Analyses were performed by M.S. and supported by N.K. and T.O. M.S. prepared the initial manuscript and N.K., T.O., T.T., and K.K. significantly contributed to revising it. All authors read and approved the final manuscript.

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**Conflicts of Interest:** The authors declare no conflict of interest.

## Appendix A

**Table A1.** Hazard ratios (HRs) for the association of mortality with other relative deprivation criteria in multiple-imputed dataset [a,b].

| | Model 2-a | Model 2-b |
|---|---|---|
| | HR (95%CI) | HR (95%CI) |
| Relative poverty | | |
| Non-poverty | ref. | ref. |
| Poverty | 1.23† (0.99–1.54) | 1.22† (0.98–1.53) |
| Relative deprivation [c] | | |
| Non-deprivation | ref. | ref. |
| 1 + | 1.25 † (0.99–1.59) | |
| 1 | | 1.14 (0.87–1.49) |
| 2 | | 1.62 † (0.97–2.69) |
| 3 + | | 1.82 * (1.10–3.01) |

** $p < 0.01$, * $p < 0.05$, † $p < 0.10$. HR: Hazard ratio; 95%CI: 95% confidence interval. [a] Multiple imputation by chained equations was performed using relative deprivation index, relative poverty, sex, age, years of education, marital status, disease and/or impairment, self-recognition of forgetfulness, depressive symptoms (m = 20). [b] Sex, age, years of education, marital status, disease and/or impairment, self-recognition of forgetfulness, and depressive symptoms were controlled. [c] This index was assessed by counting the number of items.

**Table A2.** Hazard ratios (HRs) for association of mortality with relative deprivation in raw data [a,b].

| | Model 1 | Model 2 | Model 3 | Model 4 | Model 5 |
|---|---|---|---|---|---|
| | HR (95%CI) | HR (95%CI) | HR (95%CI) | HR (95%CI) | HR (95%CI) |
| Relative poverty | | | | | |
| Non-poverty | ref. | ref. | ref. | | |
| Poverty | 1.24 † (1.00–1.53) | 1.25 * (1.00–1.56) | 1.24 † (0.99–1.54) | 1.23 † (0.99–1.54) | |
| Relative deprivation [c] | | | | | |
| Non-deprivation | | ref. | ref. | ref. | |
| 1+ | | 1.21 (0.94–1.57) | | | |
| 1 | | | 1.09 (0.81–1.47) | 1.09 (0.81–1.47) | |
| 2+ | | | 1.67 * (1.10–2.53) | | |
| 2 | | | | 1.68 † (0.97–2.93) | |
| 3+ | | | | 1.64 † (0.92–2.96) | |
| Combination [d] | | | | | |
| No dep. & pov. | | | | | ref. |
| Poverty only | | | | | 1.28 * (1.03–1.60) |
| Deprivation only | | | | | 1.98 † (1.00–3.97) |
| Pov. & dep. | | | | | 2.07 ** (1.26–3.39) |

** $p < 0.01$, * $p < 0.05$, † $p < 0.10$. HR: hazard ratio; 95%CI: 95% confidence interval. [a] Missing values in control variables were included as a dummy variable. [b] Sex, age, years of education, marital status, disease and/or impairment, self-recognition of forgetfulness, depressive symptoms were controlled. [c] This index was assessed by counting the number of items. [d] Relative deprivation in combination variable was defined as respondents who fell under two and over deprivation index.

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RESEARCH **Open Access**

# Socioeconomic inequalities in low back pain among older people: the JAGES cross-sectional study

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## Abstract

**Background:** Low back pain is an important public health issue across the world. However, it is unclear whether socioeconomic status (SES) is associated with low back pain. This study determines an association between SES and low back pain among older people.

**Methods:** We used cross-sectional data derived from the year 2013 across 30 Japanese municipalities. The survey was conducted between October 2013 to December 2013. Functionally independent community-dwelling older adults aged 65 and above ($n = 26{,}037$) were eligible for the study. Multilevel Poisson regression analysis with a robust variance estimator was used to examine the association between SES and low back pain. Self-reported low back pain in the past year was used as a dependent variable. Educational attainment, past occupation, equivalized household income, wealth, and subjective economic situation represented SES and were separately analyzed as independent variables, adjusted for covariates including age and sex.

**Results:** The prevalence of low back pain was 63.4%. Overall, lower SES were more likely to suffer from low back pain compared with that for the highest. First, as for the educational attainment, the prevalence ratio (PR) (95% credible interval (CI)) for the lowest level was 1.07 (1.02–1.12). Second, as for the past occupation, the PR (95% CI) for the blue-collared workers compared with professionals was 1.06 (1.01–1.11). Third, as for the equalized household income, the PRs (95% CI) for lower middle and the lowest income levels were 1.08 (1.02–1.13) and 1.16 (1.10–1.23), respectively. Fourth, as for the wealth, the PRs (95% CI) for lower middle and the lowest wealth levels were 1.11 (1.04–1.19) and 1.18 (1.11–1.27), respectively. Fifth, as for the subjective economic situation, the PRs (95% CI) for lower middle and the lowest financial conditions were 1.18 (1.10–1.26) and 1.32 (1.22–1.44), respectively.

**Conclusions:** Significant socioeconomic inequalities were observed in low back pain among older individuals in Japan. Policymakers and clinicians must understand the nature of these inequalities.

**Keywords:** Income, Educational attainment, Subjective economic situation, Occupation, Low back pain, Socioeconomic status, Health inequalities

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## Introduction

Low back pain is the number one cause of disability [1]. It is also most commonly experienced among musculoskeletal pains for all age groups [2, 3]. As a whole, musculoskeletal pains impact individuals' diseases and functional status such as depression [4], dementia [5], falls [6], and disability [5]. Based on the systematic review of the prevalence of low back pain in the adult populations, the estimated one-year prevalence was 38.0% ± 19.4% and more likely to be higher in the older populations [2].

Socioeconomic inequalities in health among older populations have emerged as a global concern [7, 8]. Recent studies have reported that such inequalities were observed not only in diseases but also in symptoms, including musculoskeletal pains [9–12]. Various studies reported socioeconomic inequalities in the risk factors of low back pain [13–16] such as depression [17], obesity [18], and smoking [18]. However, the results of previous studies on socioeconomic status (SES) and low back pain have been inconsistent. A recent large-scale cross-sectional study from the United States reported that the lowest income levels are significantly associated with low back pain compared with the highest income levels [12]. On the other hand, another cross-sectional study from France reported that there was no association between educational attainment and low back pain [19]. The difference in results might be explained by the different aspects of SES indicators; income is a proxy of the present SES and education is a proxy of the past SES. Seldom studies have investigated the associations between various SES factors and low back pain. Here, we conducted a cross-sectional study to determine the association of past and present SES with low back pain among older Japanese people.

## Methods

### Study population

We used data from the Japan Gerontological Evaluation Study (JAGES) project, which was cross-sectional data derived from the year 2013. Self-reported questionnaires were mailed to 112,123 people aged ≥65 years, who were not part of the long-term care insurance system [20]. Based on official residential registers obtained from respective municipal governments, the questionnaires were randomly mailed to residents selected from the 17-city areas, all of which have larger populations. In the other 13 municipalities, all of which have a smaller population, questionnaires were mailed to all eligible residents. The survey was conducted between October 2013 and December 2013. The questionnaires were divided into five subsets because many items were inquired as the whole questionnaire. The 112,123 eligible individuals were each distributed one of the five questionnaire subsets. Therefore, 38,724 individuals were mailed the questionnaire that included questions on low back pain. Of them, 27,684 individuals responded, with a response rate of 71.5%. Consequently, we used the data from 24,285 individuals in the analysis (see Fig. 1).

### Dependent variables: Low back pain in the past year

To measure chronic pain rather than acute one, the most widely used period is one-year prevalence of low back pain in previous studies [2]. Although a previous cohort study targeted at older people examined the association between pain intensity within one month and incident disability [5], we consider that such relatively acute pain might not have been enough when considering the long-term mechanism of disability. Therefore, we used one-year prevalence of low back pain. Information

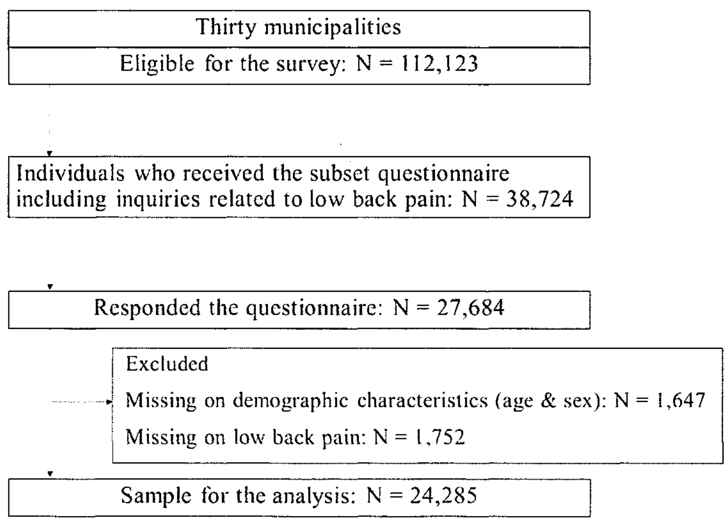


**Fig. 1** Recruitment diagram of the complete data

on low back pain was obtained by asking the following question to the participants: "Have you felt pain in or around your lower back in the past year?" Responses of "yes" indicated the presence of low back pain. If individuals responded "yes" in the previous question, we further obtained information on the severity of low back pain by asking the following question: "Have you felt physically limited in your daily life because of low back pain?" Responses of "yes" indicated the presence of intense low back pain in our study. Furthermore, we obtained information on medical access for low back pain by asking the following question: "Have you consulted a doctor for low back pain?" with possible answers of "yes" or "no."

### Independent variables: Socioeconomic status

We assessed five SES indicators as independent variables: Educational attainment, past occupation, equivalized household income, subjective economic situation, and wealth. For older adults, SES can be divided into past or present SES. Thus, we firstly examined the maximum likelihood method with Promax rotations for factor analysis to detect the type of SES indicators. Using factor analysis, educational attainment and past occupation were categorized into past SES. On the other hand, equivalized household income, wealth, as well as subjective economic situation were categorized into present SES (Additional file 1: Table S1). Educational attainment, representing past SES, was categorized into three groups: < 10 years (junior high school; lower secondary education), 10–12 years (high school; upper secondary education), and ≥ 13 years (college or university) [21]. Past occupation, representing past SES, was ascertained by the type of occupation in which participants had been engaged for the longest period. Occupations were categorized as follows: professionals, white-collared workers except for professionals, blue-collared workers, and those who had never worked before. These categories referred to a previous study [22]. We classified equivalized household incomes per year, as a present SES, into four groups: < 1,000,000 yen, 1,000,000–1,999,999 yen, 2,000,000–2,999,999 yen, and ≥ 3,000,000 yen. Wealth, as a present SES, was ascertained as household assets including savings, real estate (e.g. house, land, condominium), stocks, golf membership and was classified into five groups: < 1,000,000 yen, 1,000,000–4,999,999 yen, 5,000,000–9,999,999 yen, 10,000,000–49,999,999 yen, and ≥ 50,000,000 yen. Subjective economic situation, as a present SES, was ascertained by asking the following question: "Which of the following best describes your feelings against your current financial living conditions as a whole?" with possible answers of "very difficult," "difficult," "comfortable," and "very comfortable."

### Covariates

We used several covariates on the basis of previous works: age (65–69, 70–74, 75–79, 80–85, and ≥ 85 years), sex, number of people living together (living alone, living with others), marital status [23] (married, widowed, divorced, and never married), musculoskeletal disease, body mass index (BMI) [15] (< 18.5 kg/m$^2$, 18.5–24.9 kg/m$^2$, 25.0–29.9 kg/m$^2$, ≥30 kg/m$^2$), drinking habit [24, 25] (current, former, never), smoking [16] (current, former, never), physical activity [26, 27] (≥4 times a week, 2–3 times a week, once a week, 1–3 times a month, a few times a year and rare), and depression [13, 14] (none, mild, severe). Following a previous study [26], physical activity was measured as comprising the frequency of moderately intensive activities such as walking (at a brisk pace), dancing, gymnastics, golf, yard work, and car washing. Drinking habit was ascertained by asking the following question: "Do you drink alcohol?" with possible answers of "current", "former," and "never." The Japanese short version of Geriatric Depression Scale (GDS), consisting of 15 questions, which has been previously reported to be validated as a screening index for major depression, was used to assess the prevalence of depressive symptoms [28, 29]. We classified the participants into three groups: those with non-depressive symptoms (GDS < 5), those with mild depression (GDS of 5–9), and those with severe depression (GDS ≥10) [29].

### Statistical analysis

Considering the hierarchical structures of municipalities compared to individuals, multilevel Poisson regression analysis with a robust variance estimator was used to examine the association between SES and low back pain [30]. The data was structured as two levels: individual as level 1 and municipality as level 2. Estimates were obtained from Bayesian estimation using Markov Chain Monte Carlo (MCMC) methods. To avoid multicollinearity, the variables of educational attainment, past occupation, equivalized household income, subjective economic situation, and wealth were separately analyzed in different models, adjusting for the covariates. Dummy variables for all covariates were appropriately added to the models.

We built four regression models: Model 1, a crude model; Model 2, with age and sex adjusted to Model 1; Model 3, with number of persons living together, marital status, musculoskeletal disease, BMI, drinking habit, smoking, physical activity added to Model 2; Model 4, with depression added to Model 3. Depression was considered to be a possible intermediate factor in our analyses based on previous studies [31, 32]. To evaluate the social gradient of pain, we also examined P for trends for each model. Stratified analyses were also performed regarding sex and age (< 75 years old, ≥75 years old), with reference to previous studies [2, 12]. This stratification of age was made for the following reason: copayment for medical services differs between older adults aged < 75 years and ≥ 75 years in Japan; thus, medical access to low back pain might have been affected by economic reasons [33]. Moreover, we performed the same regression analysis in which SES indicators concurrently analyzed in the same model: Model A, past SES indicators (educational attainment and past occupation) concurrently added; Model B, present SES indicators (income, subjective economic situation and wealth) concurrently added; Model C, all SES indicators concurrently added. All of the previously mentioned models were adjusted for age, sex, number of persons living together, marital status, musculoskeletal disease, BMI, drinking habit, smoking and physical activity.

Before performing regression analyses, we employed multiple imputation under the missing at random (MAR) assumption to handle the problem of missing values. Missing variables were imputed by multivariate imputation chained equations (MICE) using following variables; sex, age, equivalized household income, educational attainment, past occupation, wealth, subjective economic situation, number of people living together, marital status, presence of knee pain, presence of low back pain, smoking, drinking habit, BMI, GDS, physical activity, and residential municipality. On the basis of a previous work, we imputed not only independent variables and covariates, but also dependent variable [34]. Rubin's rule was used to combine the results across 10 imputed datasets [35]. For the complete case analysis, we used listwise deletion methods.

For sensitivity analysis, we performed the same analysis among participants who subjectively reported physical limitation due to low back pain in daily life ($n$ = 7878) [36], as intense low back pain might shorten healthy life expectancy [37]. Multilevel analyses were performed with MLwiN, version 3.02 (Centre for Multilevel Modelling, University of Bristol) via Stata, version 15.1 (Stata Corp, College Station, TX). All other analyses were conducted using Stata.

## Results

### Demographic characteristics

Table 1 and Additional file 1: Table S2 summarize the demographic characteristics, health status, and health behaviors of all eligible participants, respectively. The prevalence of low back pain in the past year was 63.4% in the complete data. (Additional file 1: Table S2) Those who were older, female, living alone, less educated, lower income level and with lower wealth were more likely to suffer from low back pain in the past year (Table 1).

### Back pain and socioeconomic status

Table 2 summarizes the results of multilevel Poisson regression analyses for the five SES independent variables after imputation for missing data. The municipality level variances were small in all models (Table 2).

First, as for educational attainment, after adjusting for covariates and risk factors (Model 3), those of the lowest educational level were more likely to experience low back pain compared with the highest educational level–a prevalence ratio (PR) (95% credible interval (CI)) of 1.07 (1.02, 1.12). This association was attenuated after additional adjustment for depression–a PR (95% CI) of 1.05 (1.002, 1.10) (Model 4). Second, as with past occupation, the PR (95% CI) of experiencing low back pain for blue-collared workers compared with professionals was 1.06 (1.01, 1.11). This association was attenuated after additional adjustment for depression–a PR (95% CI) of 1.04 (1.001, 1.10) (Model 4). Third, with regard to equivalized household income, after adjusting for covariates and risk factors (Model 3), the PRs (95% CI) for lower middle and the lowest income levels were 1.08 (1.02, 1.13) and 1.16 (1.10, 1.23), respectively. Significant associations persisted after controlling for additional adjustment for depression (Model 4). Fourth, with regard to subjective economic situation, Model 3 showed that the PRs (95% CIs) for the "difficult" and the "very difficult" situations were 1.18 (1.10, 1.26) and 1.32 (1.22, 1.44), respectively. The associations similarly persisted in Model 4. Finally, with regard to wealth, Model 3 showed that PRs (95% CIs) for the lower middle and the lowest wealth levels were 1.11 (1.04, 1.19) and 1.18 (1.11, 1.27), respectively. The associations similarly persisted in Model 4. P for trends in education, past occupation, income, subjective economic situation, and wealth were significant (Table 2). All results using complete data were similar to those from multiple imputation pooled data (Additional file 1: Table S3).

Additional file 1: Table S3 also shows the associations of covariates and risk factors with low back pain. Overall, being older, female, the presence of musculoskeletal pain, obesity, and depression were associated with low back pain in several models. As for BMI, both overweight and obesity were associated with low back pain compared with normal weight (Model 3). For example, when SES was determined by income level, PRs (95% CI) for overweight and obesity were 1.07 (1.02, 1.12) and 1.16 (1.02, 1.32), respectively. The associations persisted after additional adjustment for depression (Model 4). Similarly, as for depression, both mild and severe depression were associated with low back pain compared with non-depression (Model 3). When SES was

**Table 1** The presence of low back pain in participants by characteristic ($n = 26,037$)

| Characteristics | Total<br>N | Having low back pain<br>N (%) | P-value |
|---|---|---|---|
| Sex | | | |
| Male | 12,088 | 6892 (61.0) | <.01 |
| Female | 13,949 | 8509 (65.5) | |
| Age, years | 74.02 (6.25) [a] | | |
| 65–69 | 7220 | 4155 (60.5) | <.01 |
| 70–74 | 7796 | 4479 (61.3) | |
| 75–79 | 5754 | 3439 (64.8) | |
| 80–84 | 3519 | 2229 (69.0) | |
| ≥ 85 | 1748 | 1099 (69.8) | |
| Educational attainment, years | | | |
| < 10 (junior high school; lower secondary education) | 10,847 | 6602 (66.0) | <.01 |
| 10–12 (high school; upper secondary education) | 9509 | 5597 (62.4) | |
| ≥ 13 (college or university degree) | 5072 | 2866 (59.3) | |
| Past occupation | | | |
| Professionals | 3742 | 2179 (59.4) | <.01 |
| White-collared workers | 5305 | 3150 (60.4) | |
| Blue-collared workers | 9627 | 6010 (64.3) | |
| Never worked before | 1302 | 826 (65.5) | |
| Equivalized household income, yen | | | |
| < 1 million | 3119 | 2107 (70.3) | <.01 |
| 1 million–1.99 million | 7735 | 4795 (63.7) | |
| 2 million–2.99 million | 4819 | 2839 (60.3) | |
| ≥ 3 million | 4974 | 2863 (58.6) | |
| Subjective economic situation | | | |
| Very difficult | 1963 | 1378 (76.0) | <.01 |
| Difficult | 8853 | 5556 (67.2) | |
| Comfortable | 12,247 | 6871 (59.8) | |
| Very comfortable | 2286 | 1225 (56.5) | |
| Wealth, yen | | | |
| < 1 million | 2101 | 1428 (70.3) | <.01 |
| 1 million–4.99 million | 2815 | 1800 (65.8) | |
| 5 million–9.99 million | 3255 | 1971 (62.0) | |
| 10 million–49.99 million | 7842 | 4693 (61.0) | |
| ≥ 50 million | 2828 | 1586 (57.2) | |
| Depression | | | |
| Non (GDS < 5) | 15,592 | 8578 (58.2) | <.01 |
| Mild (GDS of 5–9) | 4178 | 2834 (71.9) | |
| Severe (GDS ≥ 10) | 1485 | 1101 (79.6) | |

[a] mean age (SD). Chi-squared test was performed
1 US dollar is approximately 100 yen and 1 EURO is approximately 130 yen

determined by income level, PRs (95% CI) for mild and severe were 1.19 (1.13, 1.26) and 1.29 (1.19, 1.40), respectively. Additional file 1: Table S4 shows the results of stratified analyses. When considering sex stratification, sex differences were observed regarding education, past occupation, and equivalized household income–the associations were observed among males. When considering age stratification, no

**Table 2** The association of each socioeconomic status with low back pain after multiple data imputations (n = 26,037. Multilevel Poisson regression analysis)

| Socioeconomic status | Model 1 | | Model 2 | | Model 3 | | Model 4 | |
|---|---|---|---|---|---|---|---|---|
| | PR | 95% CI | PR | 95% CI | PR | 95% CI | PR | 95% CI |
| Fixed parameter | | | | | | | | |
| Education, years (ref, ≥13) | | | | | | | | |
| 10–12 | 1.05 | 1.01, 1.10 | 1.04 | 0.99, 1.09 | 1.03 | 0.99, 1.09 | 1.03 | 0.98, 1.08 |
| < 10 | 1.12 | 1.07, 1.17 | 1.09 | 1.04, 1.14 | 1.07 | 1.02, 1.12 | 1.05 | 1.002, 1.10 |
| *P* for trend | < 0.01 | | < 0.01 | | < 0.01 | | 0.01 | |
| Random part | | | | | | | | |
| Municipality level variance (standard error) | 0.001 (< 0.001) | | 0.001 (< 0.001) | | 0.001 (< 0.001) | | 0.001 (< 0.001) | |
| Past occupation (ref, professionals) | | | | | | | | |
| Fixed parameter | | | | | | | | |
| White-collared workers | 1.02 | 0.97, 1.07 | 1.00 | 0.95, 1.06 | 1.01 | 0.95, 1.06 | 1.01 | 0.95, 1.06 |
| Blue-collared workers | 1.08 | 1.03, 1.14 | 1.06 | 1.02, 1.11 | 1.06 | 1.01, 1.11 | 1.04 | 1.001, 1.10 |
| Never worked before | 1.11 | 1.04, 1.19 | 1.04 | 0.97, 1.12 | 1.03 | 0.95, 1.12 | 1.01 | 0.93, 1.09 |
| *P* for trend | < 0.01 | | < 0.01 | | < 0.01 | | 0.02 | |
| Random part | | | | | | | | |
| Municipality level variance (standard error) | 0.001 (< 0.001) | | 0.001 (< 0.001) | | 0.001 (< 0.001) | | 0.001 (< 0.001) | |
| Income, yen (ref, ≥3 million) | | | | | | | | |
| 2 million–2.99 million | 1.03 | 0.98, 1.09 | 1.03 | 0.98, 1.08 | 1.03 | 0.97, 1.09 | 1.02 | 0.97, 1.08 |
| 1 million–1.99 million | 1.09 | 1.04, 1.14 | 1.09 | 1.04, 1.14 | 1.08 | 1.02, 1.13 | 1.05 | 1.002, 1.11 |
| < 1 million | 1.20 | 1.14, 1.26 | 1.17 | 1.11, 1.23 | 1.16 | 1.10, 1.23 | 1.12 | 1.06, 1.19 |
| *P* for trend | < 0.01 | | < 0.01 | | < 0.01 | | < 0.01 | |
| Random part | | | | | | | | |
| Municipality level variance (standard error) | 0.001 (< 0.001) | | 0.001 (< 0.001) | | 0.001 (< 0.001) | | 0.001 (< 0.001) | |
| Subjective economic situation (ref, very comfortable) | | | | | | | | |
| Fixed parameter | | | | | | | | |
| Comfortable | 1.06 | 1.00, 1.12 | 1.07 | 1.004, 1.13 | 1.05 | 0.99, 1.12 | 1.04 | 0.98, 1.11 |
| Difficult | 1.19 | 1.12, 1.26 | 1.21 | 1.14, 1.29 | 1.18 | 1.10, 1.26 | 1.14 | 1.07, 1.22 |
| Very difficult | 1.34 | 1.25, 1.44 | 1.36 | 1.26, 1.47 | 1.32 | 1.22, 1.44 | 1.22 | 1.11, 1.33 |
| *P* for trend | < 0.01 | | < 0.01 | | < 0.01 | | < 0.01 | |
| Random part | | | | | | | | |
| Municipality level variance (standard error) | 0.001 (< 0.001) | | 0.001 (< 0.001) | | 0.001 (< 0.001) | | 0.001 (< 0.001) | |
| Wealth, yen (ref, ≥50 million) | | | | | | | | |
| Fixed parameter | | | | | | | | |
| 10 million–49.99 million | 1.07 | 1.01, 1.12 | 1.07 | 1.01, 1.12 | 1.05 | 0.99, 1.11 | 1.04 | 0.98, 1.10 |
| 5 million–9.99 million | 1.09 | 1.03, 1.16 | 1.09 | 1.02, 1.15 | 1.07 | 1.004, 1.14 | 1.05 | 0.98, 1.12 |
| 1 million–4.99 million | 1.15 | 1.09, 1.23 | 1.14 | 1.08, 1.22 | 1.11 | 1.04, 1.19 | 1.08 | 1.01, 1.16 |
| < 1 million | 1.22 | 1.15, 1.30 | 1.21 | 1.14, 1.29 | 1.18 | 1.11, 1.27 | 1.13 | 1.06, 1.21 |
| *P* for trend | < 0.01 | | < 0.01 | | < 0.01 | | < 0.01 | |
| Random part | | | | | | | | |
| Municipality level variance (standard error) | 0.001 (< 0.001) | | 0.001 (< 0.001) | | 0.001 (< 0.001) | | 0.001 (< 0.001) | |

Abbreviations: *PR* prevalence ratio, *95% CI* 95% credible interval
Socioeconomic status was separately added to each model. Model 1, a crude model; Model 2, with age and sex adjusted to Model 1; Model 3, with number of persons living together, marital status, musculoskeletal disease, BMI, drinking habit, smoking and physical activity added to Model 2; Model 4, with depression added to Model 3

**Table 3** The association of combined socioeconomic status with low back pain after multiple data imputations (*n* = 26,037. Poisson regression analysis)

| Socioeconomic status | Model A | | Model B | | Model C | |
|---|---|---|---|---|---|---|
| | PR | 95% CI | PR | 95% CI | PR | 95% CI |
| Fixed effect parameters | | | | | | |
| Educational attainment, years (ref, ≥13) | | | | | | |
| 10–12 | 1.03 | 0.98, 1.08 | – | – | 1.01 | 0.97, 1.06 |
| < 10 | 1.05 | 1.002, 1.11 | – | – | 1.02 | 0.97, 1.08 |
| Past occupation (ref, professionals) | | | | | | |
| White-collared workers | 1.01 | 0.95, 1.06 | – | – | 1.01 | 0.96, 1.07 |
| Blue-collared workers | 1.05 | 0.99, 1.10 | – | – | 1.03 | 0.97, 1.08 |
| Never worked before | 1.02 | 0.94, 1.11 | – | – | 1.00 | 0.92, 1.08 |
| Income, yen (ref, ≥3 million) | | | | | | |
| 2 million–2.99 million | – | – | 1.00 | 0.95, 1.06 | 1.00 | 0.95, 1.06 |
| 1 million–1.99 million | – | – | 1.02 | 0.96, 1.07 | 1.02 | 0.96, 1.08 |
| < 1 million | – | – | 1.06 | 0.99, 1.13 | 1.06 | 0.99, 1.14 |
| Subjective economic situation (ref, very comfortable) | | | | | | |
| Comfortable | – | – | 1.04 | 0.97, 1.11 | 1.04 | 0.97, 1.11 |
| Difficult | – | – | 1.12 | 1.04, 1.20 | 1.15 | 1.07, 1.24 |
| Very difficult | – | – | 1.17 | 1.07, 1.29 | 1.27 | 1.15, 1.39 |
| Wealth, yen (ref, ≥50 million) | | | | | | |
| 10 million–49.99 million | – | – | 1.01 | 0.95, 1.07 | 1.00 | 0.94, 1.06 |
| 5 million–9.99 million | – | – | 0.99 | 0.93, 1.07 | 0.99 | 0.92, 1.06 |
| 1 million–4.99 million | – | – | 1.01 | 0.94, 1.09 | 1.00 | 0.93, 1.08 |
| < 1 million | – | – | 1.03 | 0.95, 1.12 | 0.99 | 0.95, 1.11 |
| Random parameter | | | | | | |
| Municipality level variance (standard error) | 0.001 (< 0.001) | | 0.001 (< 0.001) | | 0.001 (< 0.001) | |

Abbreviations: *PR* prevalence ratio, *95% CI* 95% credible interval
Model A, educational attainment and past occupation concurrently added to the model adjusting for age, sex, number of persons living together, marital status, musculoskeletal disease, BMI, drinking habit, smoking and physical activity
Model B, income, subjective economic situation, wealth concurrently added to the model adjusting for age, sex, number of persons living together, marital status, musculoskeletal disease, BMI, drinking habit, smoking and physical activity
Model C, all socioeconomic status concurrently added to the model adjusting for age, sex, number of persons living together, marital status, musculoskeletal disease, BMI, drinking habit, smoking and physical activity

clear differences were observed between ≥75 years old and < 75 years old.

Table 3 shows the associations when SES indicators concurrently added to the regression models. For past SES indicators, the significant association of educational attainment persisted in the lowest group. Meanwhile, for past occupation, the association was no longer statistically significant for blue-collared (Model A). For present SES indicators, the significant associations of subjective economic situation persisted among difficult and very difficult, while the other associations were attenuated (Model B). When all SES indicators were included in the model, only the significant association of subjective economic situation being difficult and very difficult persisted, while associations were no longer statistically significant for other indicators.

For sensitivity analysis, associations were emphasized for all models when performing the same analysis among participants who experienced low back pain with limitations in daily life (*n* = 7878). (see Additional file 1: Table S5).

## Discussion

To the best of our knowledge, our study was the first to reveal the association of past and present SES with low back pain in the older population. We found that participants with low SES, as measured by education, past occupation, income, subjective economic situation, and wealth, were more prone to experience low back pain compared with those with high SES. Moreover, these results showed that there was a socioeconomic gradient in low back pain; people with lower socioeconomic background were more likely to suffer from pain. Therefore, low back pain is a problem for not only of the deprived people, but also a problem for the whole society. Expectedly, the associations of SES with low back pain dramatically attenuated when depression was adjusted for.

Regarding present SES, a cross-sectional study from the United States found lower-income levels to be associated with low back pain in the general population [12]. This study also indicated that associations between income and low back pain were stronger among males than among females [12]. The findings of our study are also in line with those of this cross-sectional study. We found that older individuals with a lower income level were more likely to suffer from low back pain. This association was strongly observed among older males.

We also newly elucidated the association between other present SES, as represented by wealth or subjective economic situation, and low back pain. Accordingly, we found that participants with a lower level of both wealth and subjective economic situation were more likely to experience low back pain, when separately analyzed.

Our further analyses which included all SES factors showed that the impact of more difficult subjective economic situation remained significant while the effects of other SES indicators were attenuated (see Table 3). Recently, subjective economic situation has been focused upon as a new SES indicator representing the perceived relative deprivation of individuals [38, 39]. A cross-sectional study from Germany showed that subjective economic situation mediates associations between objective SES indicators (education, occupation, and income) and depressive symptoms in adults [39]. Moreover, the study reported that the association of subjective economic situation with poor mental health was stronger than that of other SES indicators [39]. Our findings have the same context with these results to show that the subjective economic situation had the largest impact. Furthermore, we revealed that present SES was found to be associated with low back pain among participants aged < 75 years as well as ≥75 years. This indicates that present SES-related inequalities persist throughout the life.

According to our understanding, this study is among the first to reveal the associations of past SES, as measured by educational attainment and past occupation, with low back pain among older individuals. We found that participants with the lowest educational level and blue-collared workers were more likely to suffer from low back pain. Furthermore, the association between education/occupation and low back pain was stronger among males than among females. For educational attainment, in contrast to our study, a cross-sectional study from France that interviewed labor population reported that the association of educational attainment with low back pain was no longer statistically significant when adjusting for several lifestyle indicators, including BMI and smoking [19]. The difference in educational inequalities between studies might be explained as follows: educational inequalities affect health via health literacy [40], and health literacy is significantly higher in labor generations compared with that in older generations [41, 42]. Therefore, such differences between studies emerged due to demographic differences. No previous study has investigated the association of occupational inequalities with low back pain among older populations. However, numerous previous studies have indicated that heavy labor—a common issue faced by many blue-collared workers—is a risk factor of low back pain [43–46]. Our study is in accordance with the results of these prior studies. Similar to present SES, associations of past SES attenuated when all status indicators were mutually adjusted (see Table 2, Model 4). Furthermore, the association of educational attainment with low back pain was also observed among participants aged < 75 years as well as ≥75 years, indicating that educational inequalities persist throughout the life.

When considering the mechanism of low back pain, the role of risk factors must be determined. Previous studies have indicated that depression [13, 14], obesity [15], smoking [16], and lower-income level [12] are risk factors of low back pain, which is partially in accordance with our findings. Consistent with the results of a previous study [12], present SES as represented by income, subjective economic situation, and wealth were found to be statistically associated with low back pain among older adults. Two possible pathways exist for present SES-related inequalities in health: psychosocial stress and material poverty [47]. Subjective economic situation is considered to be a result of income level and is considered to represent psychosocial stress rather than material poverty [47, 48]. Moreover, individuals with lower income levels are more likely to face barriers in accessing medical care [49]. In our study, among participants with low back pain, medical access to low back pain was significantly different by SES (see Additional file 1: Table S6). This indicated that barriers in accessing medical care would be a proxy for material poverty to account for socioeconomic inequalities in low back pain. A previous study indicated a mutual effect between depression and low back pain [14]. Additionally, a causal relation between low SES and depression has been previously reported [31, 32], which supports our idea of depression as an intermediary factor. In addition to depression, numerous earlier studies have reported obesity [15] to be risk factors of low back pain. In our study, overweight and obesity were associated with low back pain. The associations of obesity somewhat attenuated when depression was additionally adjusted for. Previous studies have reported that such adverse health-related factors were strongly related to psychosocial stress [38, 39], derived from relative deprivation. Therefore, in addition to depression, obesity might contribute to low back pain through psychosocial stress that is affected by SES. Furthermore, the association of drinking habit with low back pain was not statistically significant in our study. However, previous studies have indicated that alcohol abuse might be associated with low back pain [24, 25]. We could not identify participants with alcohol abuse; however, alcohol abuse is associated with low SES [50].

There are several strengths and limitations of our study. First, we examined the association of past and present SES with low back pain. Second, we analyzed a large sample size ($n = 26{,}037$), which is higher than that analyzed in previous studies [12, 19]. The first limitation of our study is that we were unable to distinguish between acute and chronic pain, which leads to regression dilution bias. In contrast to chronic pain, a previous study has shown that individuals with a higher income level were more likely to experience acute low back pain [12]. Hence, we believe that our results are under-estimating the associations when considering such biases. Second, the pain questionnaire we used lacked information on degree of pain. There is a possibility that inequalities in low back pain might differ in degree of pain. In fact, in our sensitivity analysis, the associations were emphasized for all models when performing the same regression analysis among participants who experienced low back pain with limitations in daily life (see Additional file 1: Table S5). Future studies should include question about degree of pain. Third, we could not clarify the causal pathway because this is a cross-sectional study. Thus, the probable mediation by depressive conditions is not always consistent. However, we revealed that past SES and present SES were associated with low back pain. Longitudinal or cohort studies are necessary for future studies. Fourth, our study participants were not disabled and were not eligible for the Japanese long-term care insurance system. Future study is expected to investigate association between SES and low back pain among population including those of highly physically limited older people. Fifth, the generalizability of the present results to the entire Japanese population remains unclear. This is because the 30 municipalities investigated in this study were not randomly selected, and the sampling method for residents differed according to the population of the municipality. It was difficult to compare our study population with the entire older population due to lack of demographic characteristics in national survey.

## Conclusion

We analyzed data from a cross-sectional study, revealing that socioeconomic inequalities were significantly associated with low back pain among the older Japanese population. Policymakers and clinicians must understand the nature of these inequalities.

## Additional file

**Additional file 1: Table S1.** Factor loadings of each socioeconomic status. **Table S2.** Health status and health behaviors of all eligible participants ($n = 26{,}037$). **Table S3.** The associations of each parameter with low back pain in complete data ($n = 24{,}285$. Multilevel Poisson regression analysis). **Table S4.** The association of socioeconomic status with low back pain, stratified by sex or age after multiple data imputations ($n = 26{,}037$. Multilevel Poisson regression analysis). **Table S5.** The association of socioeconomic status with *severe* low back pain, stratified by sex or age in the complete dataset ($n = 16{,}762$. Separately Multilevel Poisson regression analysis). **Table S6.** Differences in medical access for low back pain among participants having low back pain by socioeconomic status ($n = 15{,}401$). (DOCX 119 kb)

#### Abbreviations

BMI: Body mass index; CI: Credible interval; GDS: Geriatric Depression Scale; JAGES: Japan Gerontological Evaluation Study; PR: Prevalence ratio; SES: Socioeconomic status

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#### Availability of data and materials

The data that support the findings of this study are available from the JAGES project but restrictions apply to the availability of these data, which were used under license for the current study, and so are not publicly available. Data are however available from the authors upon reasonable request and with permission of the JAGES project.

#### Authors' contributions

TI designed the study and wrote the initial draft of the manuscript. KS contributed to the study design and directed its implementation, including quality assurance and control. JA contributed to analysis and interpretation of data and assisted in the preparation of the manuscript. TT and NW conducted the literature review and prepare the Methods and the Discussion sections of the text. KK and KO contributed to data collection. All authors contributed to data interpretation, and critically reviewed the manuscript. All authors have approved the final version of the manuscript and agree to be accountable for all aspects of the work in ensuring that questions related to the accuracy or integrity of any part of the work are appropriately investigated and resolved.

#### Ethics approval and consent to participate
The JAGES protocols were approved by the ethics committee of Tohoku University and Nihon Fukushi University. We considered that individuals who responded to our survey had consented to participate.

#### Consent for publication
Not applicable.

#### Competing interests
The authors declare that they have no competing interests.

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# Social and Behavioural Determinants of the Difference in Survival among Older Adults in Japan and England

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**Keywords**

Cross-country comparative study · Laplace regression · Cohort study · Mortality · Social relationships

**Abstract**

***Background:*** A rapidly ageing population presents major challenges to health and social care services. Cross-country comparative studies on survival among older adults are limited. In addition, Japan, the country with the longest life expectancy, is rarely included in these cross-country comparisons. ***Objective:*** We examined the relative contributions of social and behavioural factors on the differences in survival among older people in Japan and England. ***Methods:*** We used data from the Japan Gerontological Evaluation Study (JAGES; *n* = 13,176) and the English Longitudinal Study of Ageing (ELSA; *n* = 5,551) to analyse all-cause mortality up to 9.4 years from the baseline. Applying Laplace regression models, the 15th survival percentile difference was estimated. ***Results:*** During the follow-up, 31.3% of women and 38.6% of men in the ELSA died, whereas 19.3% of women and 31.3% of men in the JAGES died. After adjusting for age and baseline health status, JAGES participants had longer survival than ELSA participants by 318.8 days for women and by 131.6 days for men. Family-based social relationships contributed to 105.4 days longer survival in JAGES than ELSA men. Fewer friendship-based social relationships shortened the JAGES men's survival by 45.4 days compared to ELSA men. Currently not being a smoker contributed to longer survival for JAGES women (197.7 days) and ELSA men (46.6 days), and having lower BMI reduced the survival of JAGES participants by 129.0 days for women and by 212.2 days for men. ***Conclusion:*** Compared to participants in England, Japanese older people lived longer mainly because of non-smoking for women and family-based social relationships for men. In contrast, a lower rate of underweight, men's better friendship-based social relationships, and a lower smoking rate contributed to survival among participants in England.

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## Introduction

Overall, Japanese people have the longest life expectancy in the world, especially Japanese women [1]. Research exploring the reasons for this greater longevity among the Japanese population has identified a range of potential explanations [2–5]. A milestone review by Marmot and Smith [5] identified economic growth, a low-fat diet, the public health system, and job security in Japan as the most important factors for greater longevity in this population. More recently, "a culture of hygiene, high levels of educational attainment, an egalitarian society, and strong government that led public health programmes, particularly for tuberculosis control" [3] have been proposed as key factors explaining the longevity of the Japanese population. On the other hand, high blood pressure, higher smoking rates in men, increasing social inequalities, and increasing obesity have been reported as recent threats to Japanese longevity [2, 3].

The health of the older population has become more important in ageing societies, and life expectancy at birth does not necessarily correspond to life expectancy at older age [6]. To date, there is limited evidence from cross-country comparative studies that have directly examined the respective roles of health, behavioural, and social factors in relation to longevity among older adults. One study examined the association between behavioural factors and life expectancy among older adults though it focused on European countries only [7]. However, most of the cross-country comparison studies on mortality only included Western countries. Here we compared population data from Japan, the country with the highest life expectancy, with population data from England. The aim of this prospective cohort study was to determine the relative contributions of social and behavioural factors to differences in survival between older adults in England and Japan.

## Methods

*Data Source*

Cohort data from two on-going prospective cohort studies, the English Longitudinal Study of Ageing (ELSA) [8, 9] and the Japan Gerontological Evaluation Study (JAGES) Project [10, 11], were used in this analysis. ELSA targets independent-living older adults in England aged 50 years or older and the first wave was conducted between 2002 and 2003. The first survey of the JAGES Project was conducted in 2003, drawing participants from 6 municipalities in the Aichi prefecture (at the time, the project was named the Aichi Gerontological Evaluation Study [AGES]). Although Japan has Okinawa, known in the past years as the Blue Zone [12], which is characterised as longer life in the world, the baseline survey did not cover this area. JAGES participants were community-dwelling individuals aged 65 years or older that were randomly selected from each municipality. Further details of ELSA and JAGES can be found in each cohort profile [8, 10]. In our study, we analysed the data from individuals aged 65 years or older at the first wave in both studies.

*Ethical Considerations*

Ethical approval for the ELSA was granted from the Multicentre Research and Ethics Committee (MREC/01/2/91). The AGES and JAGES study protocol and informed consent procedure were approved by the Ethics Committee on the Research of Human Subjects at Nihon Fukushi University (04-05 and 10-05).

*Outcome Measure*

The outcome was all-cause mortality occurring until 9.4 years (3,436 days) from baseline. The end of follow-up was March 15, 2012 for the ELSA and March 28, 2013 for the JAGES. For both studies, mortality data were obtained by linking with administrative records. For the ELSA, all-cause mortality was ascertained for the consenting study members (95% of the eligible participants) by linking to the National Health Service mortality register. For the JAGES, the mortality records of the national long-term care insurance database (which contains each death reported by physicians to their local municipal government) were linked with 99% of the cohort participants.

*Behavioural and Social Factors*

For health-related behaviours, we used physical activity, smoking, alcohol drinking, and body mass index (BMI). A level of physical activity was indicated by self-report on participants' active involvement in a sports club. The response was dichotomised. Based on the self-reported smoking status, participants were categorised into never smokers, past smokers, or current smokers. For alcohol drinking, participants were categorised into: non-drinking, not drinking every day, or drinking every day. A previous study reported that a BMI of 22.5–25.0 had the lowest risk of mortality [13]; we set a BMI of 22.5–24.9 as the reference category of BMI variable. In our data, a small number of Japanese participants had a BMI of 30 or over, participants with a BMI of 25 or over were put into a category. Then, BMI was divided into the following groups: less than 18.5, between 18.5 and 22.4, between 22.5 and 24.9, or 25 or over.

For social factors, socioeconomic status and social relationships were used. As an indicator for socioeconomic status, participants' age when they completed their formal education (15 years or younger, 16–18 years, 19 years or older) was used for both countries. In our study, social relationships were assessed via family (marital status) and friendship-based social relationships (social network). For family-based social relationships, marital status was categorised into married, divorced/widowed, or non-married. For friendship-based social relationships, we assessed the frequency of meeting up with friends, grouped into: once or more a week, once or twice a month, or less than once a month.

*Covariates*

Age at baseline was categorised as 65–69, 70–74, 75–79, or 85 years or older. In JAGES, older adults receiving the public disability insurance benefit were not invited to the study even if they were

**Table 1.** Baseline characteristics and mortality on the English Longitudinal Study of Ageing (ELSA) and the Japan Gerontological Evaluation Study (JAGES) Project

| | ELSA (women) | | | JAGES (women) | | | $p$ value[a] | ELSA (men) | | | JAGES (men) | | | $p$ value[a] |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $n$ | (%) | mortality, % | $n$ | (%) | mortality, % | | $n$ | (%) | mortality, % | $n$ | (%) | mortality, % | |
| *Age* | | | | | | | | | | | | | | |
| 65–69 years | 912 | (29.6) | 11.4 | 2,273 | (33.0) | 7.9 | <0.001 | 805 | (32.6) | 19.1 | 2,341 | (37.2) | 16.4 | <0.001 |
| 70–74 years | 797 | (25.9) | 22.8 | 1,951 | (28.3) | 13.9 | | 680 | (27.6) | 33.2 | 1,936 | (30.8) | 27.4 | |
| 75–79 years | 596 | (19.3) | 36.1 | 1,492 | (21.7) | 22.1 | | 498 | (20.2) | 45.2 | 1,242 | (19.7) | 44.4 | |
| 80–84 years | 498 | (16.2) | 52.2 | 769 | (11.2) | 35.9 | | 308 | (12.5) | 65.3 | 548 | (8.7) | 58.4 | |
| 85– years | 280 | (9.1) | 72.5 | 397 | (5.8) | 69.0 | | 177 | (7.2) | 83.1 | 227 | (3.6) | 79.3 | |
| *ADL limitation* | | | | | | | | | | | | | | |
| No | 2,333 | (77.4) | 26.0 | 6,194 | (96.6) | 18.2 | <0.001 | 2,029 | (83.7) | 34.2 | 5,891 | (97.3) | 30.3 | <0.001 |
| Partially disabled | 682 | (22.6) | 47.1 | 216 | (3.4) | 45.8 | | 394 | (16.3) | 60.4 | 162 | (2.7) | 59.3 | |
| *Comorbidity* | | | | | | | | | | | | | | |
| Cancer | | | | | | | | | | | | | | |
| No | 2,835 | (92.1) | 30.3 | 6,248 | (97.4) | 19.0 | <0.001 | 2,275 | (92.3) | 37.3 | 5,732 | (95.9) | 30.3 | <0.001 |
| Yes | 243 | (7.9) | 42.4 | 166 | (2.6) | 33.1 | | 191 | (7.7) | 54.5 | 247 | (4.1) | 51.0 | |
| Heart disease | | | | | | | | | | | | | | |
| No | 2,283 | (74.1) | 27.9 | 5,646 | (87.7) | 17.7 | <0.001 | 1,692 | (68.6) | 34.8 | 4,997 | (83.4) | 29.5 | <0.001 |
| Yes | 796 | (25.9) | 40.7 | 789 | (12.3) | 31.1 | | 774 | (31.4) | 47.0 | 998 | (16.6) | 39.5 | |
| Stroke | | | | | | | | | | | | | | |
| No | 2,880 | (93.5) | 29.8 | 6,333 | (98.8) | 19.2 | <0.001 | 2,265 | (91.8) | 37.0 | 5,827 | (97.5) | 30.7 | <0.001 |
| Yes | 199 | (6.5) | 52.3 | 79 | (1.2) | 32.9 | | 201 | (8.2) | 57.2 | 152 | (2.5) | 50.0 | |
| Hypertension | | | | | | | | | | | | | | |
| No | 1,618 | (52.5) | 29.5 | 4,078 | (62.6) | 19.1 | <0.001 | 1,431 | (58.0) | 38.3 | 4,062 | (67.4) | 31.3 | <0.001 |
| Yes | 1,461 | (47.5) | 33.1 | 2,436 | (37.4) | 19.7 | | 1,035 | (42.0) | 39.1 | 1,968 | (32.6) | 30.8 | |
| Diabetes | | | | | | | | | | | | | | |
| No | 2,835 | (92.0) | 30.5 | 5,786 | (90.0) | 18.8 | 0.003 | 2,187 | (88.6) | 37.6 | 5,167 | (86.2) | 29.9 | 0.002 |
| Yes | 248 | (8.0) | 40.3 | 644 | (10.0) | 23.9 | | 281 | (11.4) | 46.6 | 826 | (13.8) | 39.0 | |
| Psychiatric disorders | | | | | | | | | | | | | | |
| No | 2,906 | (94.3) | 31.3 | 6,309 | (98.4) | 19.3 | <0.001 | 2,372 | (96.1) | 38.5 | 5,915 | (98.9) | 31.0 | <0.001 |
| Yes | 177 | (5.7) | 30.5 | 105 | (1.6) | 21.0 | | 95 | (3.9) | 41.1 | 63 | (1.1) | 41.3 | |
| Education | | | | | | | | | | | | | | |
| 15 years or under | 2,022 | (67.1) | 33.9 | 4,365 | (64.6) | 20.5 | <0.001 | 1,698 | (70.4) | 42.0 | 3,594 | (57.7) | 33.9 | <0.001 |
| 16–18 years | 786 | (26.1) | 26.7 | 2,045 | (30.2) | 17.0 | | 501 | (20.8) | 31.9 | 1,752 | (28.1) | 28.1 | |
| 19 years or more | 204 | (6.8) | 23.5 | 351 | (5.2) | 15.4 | | 213 | (8.8) | 28.2 | 882 | (14.2) | 26.3 | |
| Marital status | | | | | | | | | | | | | | |
| Married | 1,424 | (46.2) | 22.5 | 3,730 | (55.9) | 14.4 | <0.001 | 1,803 | (73.1) | 34.6 | 5,520 | (89.4) | 29.2 | <0.001 |
| Divorced and widowed | 1,490 | (48.3) | 38.8 | 2,753 | (41.2) | 25.8 | | 551 | (22.3) | 50.6 | 616 | (10.0) | 46.8 | |
| Non-married | 169 | (5.5) | 39.1 | 192 | (2.9) | 19.3 | | 114 | (4.6) | 43.9 | 39 | (0.6) | 35.9 | |
| Social network | | | | | | | | | | | | | | |
| Meet up with friend once or more a week | 1,518 | (64.8) | 28.1 | 4,486 | (69.1) | 17.0 | <0.001 | 1,072 | (58.2) | 35.0 | 3,144 | (52.6) | 27.1 | <0.001 |
| Once or twice a month | 479 | (20.4) | 22.5 | 830 | (12.8) | 20.2 | | 439 | (23.8) | 30.5 | 1,014 | (17.0) | 30.3 | |
| Less than once a month | 346 | (14.8) | 28.3 | 1,177 | (18.1) | 24.0 | | 330 | (17.9) | 37.3 | 1,815 | (30.4) | 37.5 | |

**Table 1** (continued)

| | ELSA (women) | | | JAGES (women) | | | *p* value[a] | ELSA (men) | | | JAGES (men) | | | *p* value[a] |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | *n* | (%) | mortality, % | *n* | (%) | mortality, % | | *n* | (%) | mortality, % | *n* | (%) | mortality, % | |
| Exercise | | | | | | | | | | | | | | |
| No sports club participation | 2,166 | (85.8) | 30.5 | 4,725 | (80.6) | 21.0 | <0.001 | 1,779 | (84.7) | 39.0 | 4,352 | (78.1) | 32.8 | <0.001 |
| Yes | 359 | (14.2) | 14.5 | 1,140 | (19.4) | 9.2 | | 321 | (15.3) | 23.4 | 1,220 | (21.9) | 20.4 | |
| Alcohol drinking | | | | | | | | | | | | | | |
| Non | 545 | (17.9) | 45.5 | 5,881 | (88.0) | 19.8 | | 225 | (9.3) | 44.9 | 2,678 | (43.3) | 36.3 | |
| Not every day | 2,060 | (67.6) | 27.9 | 588 | (8.8) | 13.8 | | 1,641 | (67.7) | 39.0 | 1,131 | (18.3) | 27.1 | |
| Every day | 443 | (14.5) | 32.1 | 214 | (3.2) | 14.5 | | 559 | (23.1) | 37.9 | 2,379 | (38.4) | 27.2 | |
| Smoking | | | | | | | | | | | | | | |
| Current smoker | 395 | (13.1) | 36.5 | 160 | (2.5) | 30.6 | <0.001 | 290 | (12.0) | 46.9 | 1,449 | (24.0) | 35.7 | <0.001 |
| Past smoker | 1,270 | (42.1) | 33.1 | 234 | (3.6) | 24.8 | | 1,575 | (65.0) | 39.0 | 2,964 | (49.0) | 30.3 | |
| Never smoker | 1,350 | (44.8) | 27.0 | 6,093 | (93.9) | 18.4 | | 558 | (23.0) | 32.6 | 1,636 | (27.0) | 28.0 | |
| BMI | | | | | | | | | | | | | | |
| <18.5 | 48 | (1.7) | 54.2 | 566 | (8.8) | 33.9 | <0.001 | 17 | (0.8) | 70.6 | 456 | (7.5) | 57.7 | <0.001 |
| 18.5–22.4 | 332 | (11.9) | 36.7 | 2,505 | (38.8) | 19.1 | | 178 | (7.9) | 50.0 | 2,387 | (39.5) | 33.6 | |
| 22.5–24.9 | 503 | (18.1) | 33.6 | 1,855 | (28.7) | 14.9 | | 402 | (17.8) | 39.6 | 1,960 | (32.4) | 24.5 | |
| ≥25 | 1,898 | (68.2) | 25.9 | 1,532 | (23.7) | 15.3 | | 1,657 | (73.5) | 34.8 | 1,244 | (20.6) | 24.5 | |

ADL, activities of daily living; BMI, body mass index. [a] $\chi^2$ *p* value for the distribution of the participants between ELSA and JAGES.

living at their home [10], whereas ELSA targets independently living older adults. Due to this research design, it is likely that systematic differences exist in participants' disability status between two studies. To account for this study design factor, activities of daily living (ADL) were adjusted in all models as covariates. ADL were determined from the participants' response to three questions addressing the presence of difficulties in walking, bathing or showering, or using the toilet in both studies. Participants who reported one or more difficulties were regarded as partially disabled. We also used comorbidities to indicate the baseline health status of the participants as covariates. The ELSA data contained self-reports of doctor-diagnosed cancer, heart disease, stroke, hypertension, diabetes, and psychiatric disorders, whereas we used the receipt of medical treatments for cancer, heart disease, stroke, hypertension, diabetes, and psychiatric disorders for the JAGES sample.

*Analysis*

To estimate the country differences in survival, the Laplace regression analysis was applied. The Laplace regression analysis enables the estimation of the percentiles of survival time [14–17]. In our study, the lowest mortality was observed among the female participants of JAGES at 18.5%. Therefore, we used the 15th survival percentile difference (15th PD), i.e., the difference in the time when the first 15% of the participants had died, to estimate the difference in survival between the groups [17]. We also estimated the country difference in survival in relation to the explanatory factors. For sensitivity analysis, we also ran the Cox proportional hazards regression analyses [16, 17].

At first, to examine the contributions of each factor separately among the ELSA and JAGES participants, stratified analyses in each cohort were conducted. Then, the datasets for each cohort were combined, and we conducted pooled analyses. In the pooled analyses, we included the country dummy variable as a fixed effect to calculate the difference in survival between the ELSA and JAGES participants, and then examined the factors contributing to the difference in survival between the two cohorts, by calculating the number of days attributable to the different factors. For this purpose, several models were constructed. At first, before examining the contribution of each factor, age, ADL, and baseline health status (i.e., comorbidities) were included to adjust (Model 1). In Models 2–4, socioeconomic status (i.e., age at which the participants left full-time education) and social relationships variables (family-based and friendship-based) were separately added to Model 1. Then, these 3 variables were simultaneously included (Model 5). Modifiable health risks were separately added for: exercise (Model 6), alcohol drinking status (Model 7), smoking status (Model 8), and BMI (Model 9). To evaluate the degree of explained difference between the countries in survival time (days) by each factor, the change in the differences in survival between the models was calculated and displayed in a graph.

All analyses were stratified by gender. We did not apply weights on our analyses, given that ELSA and JAGES used the different sample weighting strategy. To support our decision, we conducted sensitivity analysis applying weights on both datasets. Because the results showed very close findings (see online suppl. Tables 1 and 2; see www.karger.com/doi/10.1159/000485797 for all online suppl. material), we report non-weighted results. Although the mortality outcome did not have any missing information, missing data because of item non-response were dealt with multiple imputation using chained equations by the STATA programme [18]. The numbers of missing responses are shown in online supplementary Table 3. Imputation was conducted for each dataset (country) and gender separately and 10 multiply imputed datasets were produced. Models were independently applied for each of the 10 imputed datasets. The imputation models included auxiliary variables known to be predictive of missingness. Then, single mean estimates and adjusted standard errors were calculated using Rubin's rule [19]. For sensitivity analysis, complete case analysis was also performed (in online suppl. Tables 4 and 5), which revealed similar results to those from the imputed data. As other sensitivity analysis, we excluded data from participants who died or dropped out during the first 2 years from the baseline to reduce the possibility of reverse causation (in online suppl. Tables 6 and 7). STATA SE version 14.1 (Stata Corp., College Station, TX, USA) was utilised for all analyses.

## Results

We analysed the data obtained from 3,083 women and 2,468 men in ELSA (mean age: 74.7 [SD = 7.3] and 73.7 [SD = 6.7] years, respectively) and 6,882 women and 6,294 men in JAGES (mean age: 73.4 [SD = 6.3] and 72.5 [SD = 5.7] years, respectively). The maximum follow-up period was 3,435 days (9.4 years), and the mean follow-up period was 3,047 days for women and 2,871 days for men, respectively.

Table 1 shows the baseline characteristics and mortality in both cohorts, and reveals significant differences between the countries in these characteristics. The number of deaths that occurred between the baseline and the end of the follow-up period was 964 (31.3%) women and 953 (38.6%) men in ELSA and 1,330 (19.3%) women and 1,968 (31.3%) men in JAGES. Mortality rates per person-year were 3.9 and 5.1% for ELSA women and men and 2.3 and 3.9% for JAGES women and men, respectively.

The results of the univariate and multivariable Laplace regression analyses on the differences in the 15th percentile of survival according to the explanatory variables are presented separately for women (Table 2) and men (Table 3). Older age, partial disability, and comorbid status were associated with a smaller percentile of survival time in both ELSA and JAGES in the univariate and fully adjusted multivariate models. Participants with the single status had a smaller percentile of survival especially for ELSA men. Poorer friendship-based social relationships were associated with a smaller percentile of survival especially in JAGES men. More years of education, any drinking, and not being a current smoker tended to be associated with longer survival in both the ELSA and JAGES participants. Lower (<18.5) BMI was associated with shorter survival especially for JAGES participants. The associations between each variable and 15th PD were similar in both men and women. These associations are similar to the sensitivity analyses through Laplace regression models that excluded those who died or dropped out during 1 year or 2 years from the baseline (in online suppl. Tables 6 and 7) and Cox proportional hazards models (in online suppl. Table 8).

There were differences in survival between the JAGES and ELSA participants. After adjusting for age, ADL, and baseline health status and comparing the samples when the first 15% of the participants had died, JAGES women survived for 318.8 more days than ELSA women (in online suppl. Table 9, Model 1), and JAGES men survived for 131.6 more days than ELSA men (in online suppl. Table 10, Model 1).

Figure 1 shows how much each of the aforementioned factors explains these differences in 15th percentile of survival between the two cohorts. For example, smoking contributed to 197.7 days (obtained from online suppl. Table 9, Models 5 and 8; 320.8 minus 123.1) of longer survival for JAGES women compared to ELSA women. This is explained as follows. There were more smokers among ELSA women compared to JAGES women (Table 1), and smokers had shorter survival than non-smokers (Table 2, online suppl. Table 9). In contrast, smoking was more prevalent among JAGES men compared to ELSA men. As a result, although smoking was associated with increased mortality in both JAGES and ELSA men (Table 3), the lower smoking rate among ELSA men contributed to 46.6 days longer survival compared to JAGES men.

Among women, not smoking were the most important factor for survival among JAGES compared to ELSA participants, whereas being a non-daily drinker was the most important contributing factor for longer survival among ELSA compared to JAGES women. Not being underweight also contributed to longer survival among ELSA compared to JAGES women.

Among men, rich family-based social relationships in JAGES was the most important contributor to longer survival compared to ELSA. Although the association of family-based social relationships seemed to be stronger among ELSA men (Table 3), the prevalence of being married was higher among JAGES men (Table 1). Therefore, a higher prevalence of being married seemed to contribute to 105.4 days longer survival in JAGES compared to ELSA.

For ELSA men, not being underweight, non-daily drinking, being a non-smoker, and better friendship-based social relationships contributed to increased survival compared to JAGES men. Friendship-based social

**Table 2.** Results of Laplace regression with multiple imputation among women in the English Longitudinal Study of Ageing (ELSA) and Japan Gerontological Evaluation Study (JAGES)

| | ELSA (women, *n* = 3,083) differences in 15th percentile of survival, days (95% CI) | | JAGES (women, *n* = 6,882) differences in 15th percentile of survival, days (95% CI) | |
|---|---|---|---|---|
| | univariate | multivariate | univariate | multivariate |
| *Age (ref. = 65–69)* | | | | |
| 70–74 years | –945.0 (–1,290.4; –599.7) | –522.3 (–853.8; –190.7) | –576.1 (–768.6; –383.5) | –545.3 (–753.2; –337.3) |
| 75–79 years | –2,071.8 (–2,532.9; –1,610.8) | –1,429.6 (–1,821.3; –1,037.9) | –1,380.5 (–1,591.2; –1,169.8) | –1,070.8 (–1,324.9; –816.8) |
| 80–84 years | –2,348.4 (–2,568.1; –2,128.6) | –1,845.3 (–2,254.2; –1,436.5) | –2,035.5 (–2,301.7; –1,769.3) | –1,735.8 (–2,046.9; –1,424.6) |
| 85– years | –3,076.0 (–3,276.3; –2,875.7) | –2,288.9 (–2,686.4; –1,891.3) | –3,158.5 (–3,368.5; –2,948.6) | –2,633.8 (–2,947.3; –2,320.3) |
| *ADL (ref. = no ADL limitation)* | | | | |
| Partially disabled | –1,080.8 (–1,356.2; –805.3) | –560.1 (–855.1; –265.0) | –1,823.2 (–2,039.1; –1,607.3) | –876.1 (–1,271.8; –480.5) |
| Comorbidity | | | | |
| Cancer | –1,051.9 (–1,284.9; –818.8) | –691.2 (–1,101.6; –280.7) | –1,692.2 (–2,172.3; –1,212.1) | –1,713.1 (–2,505.0; –921.2) |
| Heart disease | –816.5 (–1,095.2; –537.8) | –369.4 (–619.8; –119.0) | –1,109.7 (–1,563.9; –655.5) | –384.2 (–634.9; –133.4) |
| Stroke | –1,048.5 (–1,312.9; –784.1) | –305.4 (–725.3; 114.6) | –957.0 (–1,638.7; –275.3) | –497.3 (–1,334.0; 339.3) |
| Hypertension | –142.4 (–424.4; 139.6) | –280.2 (–496.9; –63.4) | –55.0 (–305.1; 195.1) | 55.8 (–119.3; 230.9) |
| Diabetes | –671.7 (–1,309.1; –34.3) | –604.5 (–895.7; –313.3) | –335.4 (–593.1; –77.7) | –455.9 (–751.0; –160.7) |
| Psychiatric disorders | –239.5 (–954.5; 475.4) | 3.6 (–322.4; 329.6) | –648.0 (–1,326.3; 30.4) | –78.9 (–960.0; 802.1) |
| *Education (ref. = 15 years or under)* | | | | |
| 16–18 years | 314.2 (–35.7; 664.1) | 134.6 (–140.6; 409.8) | 377.5 (177.1; 577.9) | 138.2 (–71.2; 347.7) |
| 19 years or more | 980.3 (376.1; 1,584.4) | 380.4 (–155.2; 915.9) | 539.3 (57.2; 1,021.4) | 51.6 (–348.5; 451.7) |
| *Family-based social relationships (marital status, ref. = married)* | | | | |
| Divorced and widowed | –1,072.3 (–1,387.4; –757.2) | –87.0 (–368.9; 194.9) | –1,146.0 (–1,318.3; –973.7) | –93.9 (–278.0; 90.1) |
| Non-married | –1,246.1 (–1,698.3; –794.0) | –436.5 (–875.2; 2.2) | –707.4 (–1,262.1; –152.6) | –128.2 (–685.6; 429.2) |
| *Friendships-based social relationships (social network, ref. = meet up with friend once or more a week)* | | | | |
| Once or twice a month | 104.3 (–417.9; 626.5) | –1.2 (–342.5; 340.0) | –484.4 (–757.0; –211.8) | –87.0 (–328.9; 154.9) |
| Less than once a month | –259.4 (–766.1; 247.3) | –34.2 (–385.6; 317.3) | –759.4 (–1,090.9; –427.9) | –145.6 (–358.9; 67.8) |
| *Physical activity (ref. = no sports club participation)* | | | | |
| Yes | 1,487.6 (1,035.0; 1,940.2) | 256.2 (–133.8; 646.3) | 1,305.9 (1,056.9; 1,554.9) | 389.6 (141.8; 637.4) |
| *Alcohol drinking (ref. = non)* | | | | |
| Not every day | 882.9 (629.2; 1,136.6) | 447.4 (222.9; 671.8) | 717.4 (389.3; 1,045.5) | 21.3 (–313.3; 355.9) |
| Every day | 653.9 (177.4; 1,130.4) | 268.0 (–185.3; 721.2) | 595.4 (–241.3; 1,432.1) | 400.1 (16.4; 783.8) |
| *Smoking (ref. = current smoker)* | | | | |
| Past smoker | 131.2 (–293.8; 556.2) | 496.0 (13.7; 978.3) | 274.5 (–977.9; 1,526.8) | 718.7 (74.0; 1,363.3) |
| Never smoker | 597.5 (192.6; 1,002.3) | 767.7 (302.1; 1,233.3) | 1,167.8 (596.7; 1,738.9) | 1,146.5 (554.3; 1,738.8) |
| *BMI (ref. = 22.5–24.9)* | | | | |
| <18.5 | –557.2 (–1,237.2; 122.8) | –140.3 (–595.1; 314.4) | –1,541.5 (–1,927.0; –1,156.1) | –564.0 (–837.4; –290.7) |
| 18.5–22.4 | –66.8 (–507.4; 373.8) | –161.5 (–538.8; 215.9) | –467.3 (–780.1; –154.6) | –151.7 (–354.7; 51.4) |
| ≥25 | 454.2 (77.1; 831.4) | 242.7 (–21.5; 506.8) | 63.6 (–262.6; 389.8) | –136.7 (–392.3; 118.8) |

ADL, activities of daily living; BMI, body mass index; CI, confidence interval.


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**Table 3.** Results of Laplace regression with multiple imputations among men in the English Longitudinal Study of Ageing (ELSA) and the Japan Gerontological Evaluation Study (JAGES)

| | ELSA (men, *n* = 2,468) differences in 15th percentile of survival, days (95% CI) | | JAGES (men, n = 6,294) differences in 15th percentile of survival, days (95% CI) | |
|---|---|---|---|---|
| | univariate | multivariate | univariate | multivariate |
| *Age (ref. = 65–69)* | | | | |
| 70–74 | –949.6 (–1,495.8; –403.3) | –818.6 (–1,207.6; –429.6) | –829.3 (–1,115.5; –543.2) | –686.2 (–906.4; –466.0) |
| 75–79 | –1,499.2 (–2,072.7; –925.8) | –1,215.7 (–1,692.3; –739.2) | –1,522.3 (–1,795.6; –1,249.0) | –1,234.8 (–1,434.3; –1,035.2) |
| 80–84 | –1,952.6 (–2,482.1; –1,423.0) | –1,736.6 (–2,156.2; –1,316.9) | –2,080.2 (–2,424.9; –1,735.4) | –1,851.9 (–2,098.7; –1,605.2) |
| 85– | –2,256.6 (–2,760.1; –1,753.0) | –1,974.7 (–2,390.6; –1,558.8) | –2,509.8 (–2,792.2; –2,227.5) | –2,169.5 (–2,505.6; –1,833.4) |
| *ADL (ref. = no ADL limitation)* | | | | |
| Partially disabled | –935.6 (–1,147.5; –723.7) | –591.5 (–891.8; –291.1) | –1,498.5 (–1,761.5; –1,235.5) | –816.9 (–1,277.9; –356.0) |
| *Comorbidity* | | | | |
| Cancer | –937.6 (–1,246.9; –628.3) | –1,104.6 (–1,475.5; –733.7) | –1,449.5 (–1,625.9; –1,273.0) | –1,524.3 (–1,822.0; –1,226.5) |
| Heart disease | –669.8 (–875.2; –464.5) | –474.1 (–719.5; –228.7) | –492.5 (–717.6; –267.4) | –338.9 (–514.1; –163.7) |
| Stroke | –548.0 (–752.0; –343.9) | –451.8 (–756.2; –147.3) | –445.1 (–766.1; –124.1) | –241.0 (–644.3; 162.2) |
| Hypertension | 92.6 (–126.4; 311.7) | –13.4 (–266.0; 239.2) | 143.5 (–47.2; 334.2) | –20.8 (–173.9; 132.3) |
| Diabetes | –304.0 (–711.2; 103.2) | –225.7 (–560.0; 108.5) | –508.7 (–774.3; –243.1) | –671.0 (–840.8; –501.2) |
| Psychiatric disorders | 241.5 (–219.7; 702.7) | 227.0 (–290.4; 744.4) | –501.4 (–1,270.6; 267.8) | –669.2 (–1,273.9; –64.5) |
| *Education (ref. = 15 years or under)* | | | | |
| 16–18 years | 127.6 (752.5; 0.0) | 160.8 (–139.2; 460.7) | 441.8 (209.4; 674.2) | –4.9 (–185.4; 175.6) |
| 19 years or more | 736.5 (158.3; 1,314.7) | 458.6 (93.2; 824.1) | 476.7 (230.2; 723.1) | 172.7 (–43.5; 389.0) |
| *Family-based social relationships (marital status, ref = married)* | | | | |
| Divorced and widowed | –729.2 (–913.3; –545.1) | –237.0 (–509.3; 35.3) | –841.9 (–1,097.9; –585.9) | –190.9 (–383.5; 1.7) |
| Non–married | –640.2 (–1,152.1; –128.4) | –765.6 (–1,343.4; –187.7) | 86.8 (–1,155.2; 1,328.8) | –525.3 (–1,372.2; 321.6) |
| *Friendships-based social relationships (social network, ref. = meet up with friend once or more a week)* | | | | |
| Once or twice a month | 130.8 (–206.9; 468.5) | 133.0 (–213.0; 479.0) | –76.3 (–287.4; 134.8) | –60.2 (–239.3; 118.9) |
| Less than once a month | 105.6 (–244.5; 455.7) | 165.6 (–170.6; 501.9) | –700.4 (–884.2; –516.5) | –297.9 (–472.9; –122.9) |
| *Physical activity (ref. = no sports club participation)* | | | | |
| Yes | 830.4 (94.7; 1,566.2) | 285.1 (–314.1; 884.3) | 839.5 (578.2; 1,100.8) | 245.7 (43.0; 448.5) |
| *Alcohol drinking (ref. = non)* | | | | |
| Not every day | 190.4 (–349.6; 730.4) | –60.9 (–517.7; 396.0) | 670.6 (444.8; 896.4) | 386.6 (160.8; 612.4) |
| Every day | 190.7 (–363.5; 745.0) | 44.9 (–461.7; 551.6) | 777.0 (611.4; 942.7) | 330.2 (158.9; 501.6) |
| *Smoking (ref. = current smoker)* | | | | |
| Past smoker | 350.1 (–4.3; 704.5) | 592.9 (194.2; 991.6) | 306.1 (79.3; 532.9) | 433.6 (271.9; 595.4) |
| Never smoker | 562.7 (135.9; 989.6) | 830.5 (417.7; 1,243.3) | 494.9 (224.1; 765.6) | 579.1 (370.5; 787.7) |
| *BMI (ref. = 22.5–24.9)* | | | | |
| <18.5 | –745.7 (–1,418.3; –73.2) | –415.2 (–1,537.7; 707.2) | –1,536.0 (–1,724.8; –1,347.1) | –678.9 (–875.0; –482.7) |
| 18.5–22.4 | –252.0 (–704.0; 200.0) | –292.8 (–889.9; 304.2) | –623.5 (–821.1; –425.8) | –297.3 (–487.1; –107.5) |
| ≥25 | 137.7 (–226.2; 501.6) | 78.8 (–189.6; 347.1) | –70.9 (–296.1; 154.3) | –120.5 (–329.6; 88.6) |

ADL, activities of daily living; BMI, body mass index; CI, confidence interval.

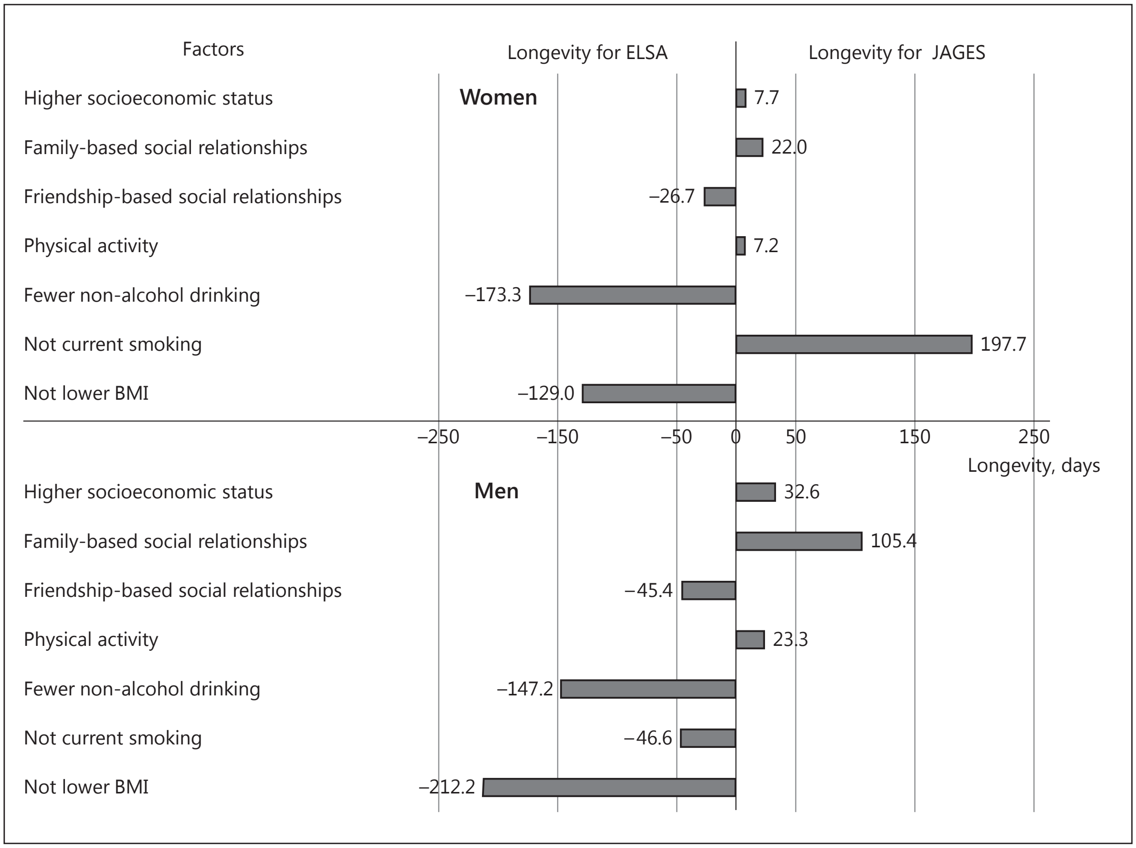


**Fig. 1.** Reasons for the difference in longevity (days) by each factor between the English Longitudinal Study of Ageing (ELSA) and the Japan Gerontological Evaluation Study (JAGES) presented separately by gender. Explained difference in longevity (days) was calculated from the results in online Supplementary Tables 9 and 10: socioeconomic status (model 1 minus model 2), family-based social relationships (model 1 minus model 3), friendships-based social relationships (model 1 minus model 4), physical activity (model 5 minus model 6), alcohol drinking (model 5 minus model 7), smoking (model 5 minus model 8), and BMI (model 5 minus model 9).

relationships predicted survival among JAGES participants only (Table 3), and the prevalence of poorer friendship-based social relationships was higher among JAGES men (Table 1). Therefore, a lack of friends contributed to 45.4 fewer days of survival in JAGES compared to ELSA men.

## Discussion

Research evidence on survival among older adults based on cross-country comparative studies is limited, especially with Japan, which has the longest life expectancy. In this study, we examined the differences in 15th percentile of survival between older individuals living in Japan and England from two prospective cohort studies, and identified the factors contributing to these differences. The participants in Japan lived longer than those in England at the 15th percentile of all-cause mortality after adjusting for age, ADL, and baseline health status. Major contributing factors for the higher survival among the JAGES compared to ELSA participants were lower smoking rate in women, and better family social relationships in men. In contrast, several factors contributed to longer survival in ELSA compared to JAGES participants: a lower rate of underweight, moderate alcohol drinking for both women and men, better friendship-based social relationships in men, and a lower smoking rate in men.

### *Strengths and Limitations of this Study*

To the best of our knowledge, this is the first cross-country comparative population-level study to show the absolute difference in survival (days) between the two countries and the size of the effect of the factors contrib-



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uting to that difference. Using absolute measure is more appropriate in public health decision making because they are easier to explain to the general public than relative risks [20, 21]. Several previous studies directly compared country differences using pooled data, but they provided relative measures rather than absolute measures [22, 23]. Traditional life table methods enable the direct assessment of absolute longevity, yet a limited number of factors can be tested with these methods [6, 24, 25]. Although recently applied advanced methods enable the consideration of multiple factors with absolute longevity, direct comparison between countries using pooled data has not been conducted [7].

In this study, because we could use pooled cohort data, the country in which the participant resided was treated as a fixed effect, using a dummy variable. As there is lower possibility of reverse temporal association between the country of residence among older people and individual sociodemographic characteristics and behaviours, the possibility of bias when estimating the mediation effect is small.

In both ELSA and JAGES, the mortality information was reliable and obtained via credible official sources. Additionally, most of the explanatory variables used in this analysis were obtained through a similar approach across the two studies. On the other hand, while we had relevant information for when participants were older adults, we could not consider various demographic, psychosocial, and environmental factors in early life due to a lack of available and comparable data.

There are notable differences between the two countries in those factors and they could have played a role in explaining our results; for example, accumulation of stress through divergent historical experiences could have affected the difference in health between the two cohorts. Inclusion of objective measures to indicate physical health status or environmental factors that are comparable in two cohorts could elaborate explaining the differences in the mortality between the two cohorts. Although both England and Japan have universally accessible health care systems, there are differences in health care provision [26]. The effect of medical care access on the outcome should be considered in future studies.

High blood pressure is reportedly the greatest risk factor for mortality among Japanese women [2]. Although we used comorbidities including hypertension in our models, having an objective blood pressure measure could have allowed further explanation of the differences in longevity between the two countries. In addition, because this study used cohorts of older adults who survived at least at the baseline of the survey, selection bias could overestimate the association of factors such as moderate alcohol drinking [27]. Our results suggest that moderate drinkers have a longer life than non-drinkers.

Although a J-shaped relationship between alcohol consumption and all-cause mortality has been previously reported [28], a recent systematic review has challenged the merits of moderate drinking after considering several biases [29]. In addition, a study on the British cohort of people born in 1958 suggested that non-drinking would be likely to be associated with long-term illnesses in early life [30]. Although we accounted for comorbidities at baseline, a precise measure to indicate quantity of alcohol intake that was comparable to both countries was impossible to obtain; instead, we used frequency of alcohol use. Results from alcohol use need to be carefully interpreted.

Another limitation relates to differences in the study designs between the two cohorts. The sampling strategies were different and the JAGES 2003 did not contain (or aim for) a national representative sample. In addition, older adults with disability were excluded from the JAGES but not from the ELSA sample. We accommodated the potential selection bias by accounting for the baseline ADL and we only included those aged 65 years and over in both studies. Additionally, in the sensitivity analyses, we excluded participants who died within 1 year or 2 years from the baseline, which likely reduced the possibility of health selection bias in our study.

Our study is based on the use of self-reported questionnaires from two cohorts. It is possible that comparability in these cohorts is limited to some degree, such as ADL. Given the cultural differences in these two cohorts, their response to the ADL questions may be different. However, the questions were phrased similarly in the two cohorts with regard to their ability to perform certain daily tasks. Moreover, this is treated as a covariate in our study. We are therefore confident that potential response bias is minimal in our study. Questions were phrased differently to the cohort members to obtain baseline health status for this study, which raises the question for comparability in this study. However, baseline health status is treated as a covariate in the model; we think that the effect on the longitudinal associations between social determinants and longevity is minimal.

In relation to the analyses, we estimated the 15th survival PD, the difference in the time when the first 15% of the participants had died. Future studies with longer follow-up would better inform cross-country differences in social determinants of longevity.

Although the effect of the long-term change in the baseline characteristics on longevity is not in our study focus, this is likely to affect the course of longevity. Future studies including time-varying factors would inform us how the policies relevant to healthy ageing to be formulated.

*Comparison with Previous Studies*

The reasons for the longevity observed in Japanese people have been examined in previous studies. The life expectancy at birth for Japanese women was the longest internationally with 86.8 years in 2015 [1]. Japanese men also have a comparatively long life expectancy at 80.5 years. Since the milestone studies related to longevity among Japanese individuals conducted by Marmot and his colleagues [4, 5], lower mortality from coronary heart disease, which is partly attributable to lower serum cholesterol levels, has been considered one of the reasons for the longer life expectancy found among the Japanese [3]. However, present analyses did not show any significant association of overweight, a possible indication of higher serum low density cholesterol, and therefore shorter survival. On the contrary, we found a significant association between underweight and shorter survival among Japanese participants, which can offer additional support to a J-shaped or U-shaped association between BMI and mortality among Japanese [12, 30]. Previous studies have reported that BMI has a J-shaped or U-shaped relationship with mortality [13, 31]. A recent study that pooled cohort data from 10 million participants reported that those with a BMI of 22.5–25.0 had the lowest risk of death compared to those who were categorised as underweight or overweight [13]. Findings from our previous study on older Japanese individuals also support the same trend [32].

In addition, the association of BMI with all-cause mortality varies according to age; a relatively higher BMI is associated with lower mortality among older populations compared to younger populations [33, 34]. In a meta-analysis of older adults, higher all-cause mortality was observed among people with a BMI of 35 or more and a significant increment of hazard ratio was not observed among people with a BMI of 30–34 [34]. This may explain why we did not observe shorter survival among the overweight participants in ELSA. Being underweight among older adults is associated with frailty [35], and frailty is linked to shorter longevity [36]. In the present study, a relatively large number of older Japanese had a BMI of less than 18.5, which was associated with increased mortality. Therefore, being underweight in older populations should be considered an important public health issue, especially in Japan.

Our results also showed that smoking reduced longevity by about 2 years even when making comparisons only for the period that the first 15% of participants died. Smoking is the greatest risk factor for death among Japanese men [2, 37]. We also confirmed the harmful effects of smoking, as the high smoking rates among Japanese men reduced their survival and decreased the difference in survival between the Japanese and English cohorts by 46.6 days. In contrast, the lower rates of smoking among women in Japan partially explained why their survival was 197.7 days longer than women in England. Therefore, among older populations, smoking is still an important public health issue that predominantly affects different genders in these two cohorts (Japanese men and English women).

Previous studies have shown the considerable health benefits of social relationships, such as marriage and contact with friends, with both factors associated with longer longevity [38, 39]. Our study added the knowledge about the different distribution of social relationships regarding health in two cohorts. The longevity of the participants in Japan could partly be explained by social relationships with families, especially for men. In contrast, longevity for men in England compared to men in Japan was partly derived from their social relationships with friends.

Similar to previous studies, the present study also confirmed that a higher educational level was associated with longer survival [40, 41]. However, education did not substantially explain the differences in longevity between the two cohorts of older adults, since their education levels were similar. This similar distribution of educational attainment in this generation in Japan and UK has already mentioned in a previous study [5].

*Public Health Implications*

Although the mean longevity was longer among the participants in Japan, we found some factors that contributed to survival among the participants in England. Tackling the factors that contribute to reduced survival, such as smoking for JAGES men and ELSA women, could effectively improve survival in both countries. In relation to social relationships, public health interventions to promote social relationships with friends have been advocated [42]. In Taketoyo in Japan, a resident-centred community intervention programme was developed to provide opportunities to promote social interactions with community members. The programme improved social interactions and reduced the incidence of functional disability among older adults. Improving social relationships with family members, skills in handling worries and demands, and conflict management within families are all potentially important components of community interventions [43].

## Conclusions

The participants in Japan lived significantly longer than those in England. Lower smoking rates in women, and better family social relationships in men were major contributors to the longevity among the Japanese participants. In contrast, a lower rate of underweight, better friendship-based social relationships in men, and a lower smoking rate in men contributed to survival among participants in England. Modifying health-related behaviours and enriching social relationships could make a difference in survival among elderly at the population level.

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## Disclosure Statement

The authors have no conflicts of interest to declare.

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Research Article

# Adverse Childhood Experiences and Higher-Level Functional Limitations Among Older Japanese People: Results From the JAGES Study

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## Abstract

**Background:** A life-course perspective is essential in understanding the determinants of higher-level functional limitations. We examine the impact of adverse childhood experiences (ACEs) on higher-level functional limitations in older people.
**Methods:** Data were from the Japan Gerontological Evaluation Study 2013, a population-based cohort of independent people aged 65 years or older across Japan ($n$ = 19,220). ACEs before the age of 18 were assessed in terms of seven adversities: parental death, parental divorce, parental mental illness, family violence, physical abuse, psychological neglect, and psychological abuse. Associations between the cumulative number of ACEs and higher-level functional limitations were investigated by multivariate Poisson regression with robust error variances, adjusted for age, gender, childhood disadvantage, adult sociodemographics, adult health behaviors, and health status.
**Results:** Of the older people, 36.3% reported at least one ACE. Older people who had experienced two or more ACEs showed significantly greater higher-level functional limitations than those with no ACE in a crude model (prevalence ratio, PR = 1.61, 95% confidence interval, CI = 1.51–1.71). After adjusting the covariates, this association remained (PR = 1.19, 95% CI = 1.12–1.27).
**Conclusions:** ACEs showed robust independent effects on higher-level functional limitations among older Japanese without disabilities, even after adjusting for potential covariates in childhood and adulthood. The current findings may help in understanding the impact of the latent effects of ACEs on functional limitations in older people.

**Keywords:** Childhood disadvantage—Latent effect—Life-course epidemiology

Recently, a life-course perspective, including exposure to major events at an early life stage, has been proposed as an important determinant of disability or higher-level functional limitations, such as socializing and activities that require higher cognitive functions (1–3). Previous studies suggested that low childhood socioeconomic status (SES), assessed by parental education or parental occupation (2), was also a risk factor for disability later in middle life. However, the mechanism(s) by which low childhood SES induces higher-level functional limitations later in life remain(s) unknown.

Adverse childhood experiences (ACEs), including interpersonal loss, family psychopathology, abuse, and neglect (4), are more likely to be observed among children in poverty. Several studies have reported that ACEs are linked with health risk behaviors and diseases, such as cardiovascular disease (5), cancer (6), chronic obstructive pulmonary disease (7), depression (8), obesity (9), and alcohol abuse (8), which are associated with disabilities. ACEs have been suggested as having enduring emotional, immune, and metabolic effects, which may ultimately increase the risks of associated diseases

(10). ACEs have also been associated with poorer verbal retrieval and visual space memory among people aged more than 65 years (11). In terms of disability, Rose et al. reported a dose–response relationship between ACEs and disability (self-reported activity limitations and/or assistive device use) in middle-age adulthood (mean age, 48 years) with adjustment for potentially mediating health conditions (12). These results indicated that ACEs may lead directly to disabilities in later life (a latent effect model (13)). In addition, more recent research has pointed out the importance of investigating factors in young adults that could mitigate the influence of ACEs on older people (14). Thus, increasing awareness of the potential long-term impact of ACEs could be an essential and potentially cost-effective target for avoiding at least some disabilities in older people. However, to our knowledge, the impact of ACEs on higher-level functional limitations in older people has not been reported before.

To explore any associations between ACEs and higher-level functional limitations in older people, we used data from the Japan Gerontological Evaluation Study (JAGES), which evaluated both ACEs and higher-level functional limitations in community-dwelling independent older people, aged 65 years or older, in Japan, some of whom survived World War II (1937–1945) during their early childhoods. Because the period of WWII was associated with ACEs, including severe poverty and parental death (15,16), these data provide a unique opportunity to investigate the impact of ACEs on higher-level functional limitations in older people. Thus, the purpose of this study was to investigate the independent effects of ACEs and higher-level functional limitations; and to explore potential mediation paths, including sociodemographics, adult health behavior, and health status as mediators and childhood disadvantage as a confounder of ACE effects in Japanese older people (Figure 1), referring to the recent life-course model on aging (14).

## Methods

### Sample

We used data from the JAGES 2013 Project, which was designed to investigate the social determinants of health among noninstitutionalized, functionally independent individuals aged 65 or older. The survey was initially conducted to elucidate the determinants of disability in older people; therefore, older adults with disabilities were not included in the survey. The survey data used were sampled in 2013 from 30 municipalities in 15 prefectures in Japan. Among the 30 municipalities, we sampled the entire population of residents in 14 municipalities, whereas in the remaining 16 municipalities, we mailed questionnaires to a random sample based on public long-term care insurance data acquired from municipal authorities. In total, 193,694 residents were targeted for the questionnaire. Of the eligible participants, 137,736 people responded (response rate = 71.1%). ACEs were investigated in one-fifth of the sample ($n$ = 25,928). We excluded those who did not respond to questions on gender or age. We also excluded those with limitations in activities of daily living, meaning those who could not walk, take a bath, and use the toilet without assistance. Furthermore, to focus on associations between ACEs and higher-level functional limitations, we excluded those who did not respond fully to the 13 questions on functional limitations and ACEs ($n$ = 4,606); thus, our analytic sample consisted of 19,220 individuals.

The JAGES protocol was approved by the Ethics Committee on Research of Human Subjects at Nihon Fukushi University.

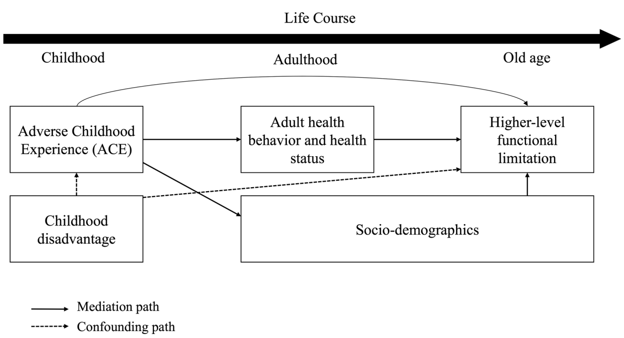


**Figure 1.** Conceptual model on the association between adverse childhood experience and higher-level functional limitation in life-course model.

### Assessing Higher-Level Functional Limitations

Higher-level functional limitations were evaluated using the Tokyo Metropolitan Institute of Gerontology Index of Competence (TMIG-IC) (17), based on the Lawton instrumental activities of daily living scale (18). This is a 13-item index consisting of three sublevels of competence with yes/no responses (1): instrumental self-maintenance (five items: ability to use public transport, shop for daily necessities, prepare meals, pay bills, and handle a bank account) (2); intellectual activities (four items: ability to fill out forms, read newspapers, read books or magazines, and interest in television programs or news articles dealing with health); and (3) social role (four items: ability to visit the home of friends, give advice to relatives and friends, visit someone in hospital, and initiate conversation with younger people). This scale is used widely to assess higher-level functional capacity in older people in Japan (19). Although a higher TMIG-IC score denotes a higher capacity, we reversed the score so that a higher score denotes a higher functional limitation in our analysis.

### Assessing ACEs

The ACE questionnaire was developed based on previous ACE studies (4), modified to suit older Japanese people (20). The respondents were asked whether they had experienced the following adversities before the age of 18 (yes/no responses): interpersonal loss (parental loss, parental divorce), family psychopathology (parental mental illness, family violence), and abuse and neglect (physical abuse, psychological neglect, psychological abuse). The sum of the number of ACEs was categorized as 0, 1, and 2 or more, because the percentage of people who experienced three or more ACE was only 2.0%, which needed to be collapsed.

### Covariates

Potential covariates were also assessed with the questionnaire, including age, gender, childhood disadvantage, sociodemographics, and adult health behaviors and health status. Childhood disadvantages were childhood economic hardship (a proxy measure of childhood SES) and adulthood height (a marker of cumulative nutritional and unhygienic exposure during early childhood (21)). Childhood economic hardship was assessed with a question about whether the financial condition during childhood provided difficulties to thriving (yes/no response). Self-reported height was obtained from the questionnaire. People reporting heights above 3 *SD*s from the mean height (from the 2010 Japanese National Health and Nutrition Survey (22)) were eliminated because they were assumed to be outliers. Furthermore, height was categorized into five groups in 5-cm intervals for men (<155, 155–159.9, 160–164.9, 165–169.9, and ≥170 cm) and women (<145, 145–149.9, 150–154.9, 155–159.9, and ≥160 cm) (3). In terms of adult sociodemographics, SES was assessed by educational attainment (<10 years, 10–12 years, ≥13 years), annual household income (<1.5, 1.5–2.9, 3.0–4.9, ≥5 million yen: 1 million yen is equivalent to 10,000 US$), and employment status (currently working, retired, or never worked). Marital status (married, widowed, divorced/never married/other) and living status (living alone, living with wife/husband, living with other) were also obtained. Health behaviors included smoking (current, quit, never) and drinking (current, quit, never) were investigated. Body mass index was determined from the self-reported height and weight (kg/m$^2$), then categorized as underweight (<18.5), normal weight (18.5–24.9), and overweight (≥25). Depression was assessed with the Geriatric Depression Scale (GDS) (23), then categorized as no depression (score 0–4), mild depression (score 5–9), and severe depression (score 10–15). The respondents were also questioned as to whether they were currently under medical care. Missing values were treated as dummy variables (24).

### Analysis

Multiple Poisson regression analysis with robust error variances was used to examine any associations between ACEs and higher-level functional limitations, treated as count data. In addition to a crude model, Model 1 was adjusted for age and gender and Model 2 was additionally adjusted for childhood disadvantage. Model 3 was adjusted for adult sociodemographics beyond Model 2. Model 4, the final model, was also adjusted for adulthood health behaviors and health status. Further, we conducted mediation analysis to convert higher-level functional limitations into dichotomous (0–1 vs 2 or more) limitations, using a Stata program called "binary_mediation", which allows for multiple mediator variables (25). All statistical analyses were conducted using Stata (ver. 13.1; StataCorp, College Station, TX). The level of statistical significance was set at $p$ <0.05 (two-tailed).

## Results

The mean age was 73.2 (standard deviation [*SD*] = 5.9) years old, and 52.9% were women (Table 1). Approximately three quarters (73.5%) were married, and 13.0% were living alone. Regarding adult SES, 36.4% had less than 10 years of education, 12.4% earned less than 1.5 million yen, and 60.8% were retired. Approximately one-third were currently drinking and one-tenth were currently smoking. Most of the participants were receiving medical treatment (78.5%). Regarding the prevalence of each ACE, parental death was the highest (22.1%), followed by psychological neglect (11.6%), and psychological abuse (5.1%). In total, 36.3% reported at least one ACE and 7.4% reported two or more ACEs. Prevalence of ACEs by age cohort showed that the rates of parental death were higher in the age group of 70+ years, who were 2+ years old when World War II ended, compared to the age group of 65–69 years (Supplementary Table S1). On the other hand, the rates of parental divorce and family violence were higher in the 65–69 age group than in the 70+ age group.

Comorbidity of ACEs was also confirmed. For example, 21.2% of those who experienced physical abuse reported any one more additional ACE and 64.4% experienced any two or more additional ACEs (Supplementary Table S2).

Table 2 shows the prevalence ratio (PR) of higher-level functional limitations, according to cumulative numbers of ACEs, using multiple Poisson regression analysis with robust error variance. Regarding higher-level functional limitations, the mean reversed TMIG-IC score was 1.38 (*SD* = 1.82) and the median was 1, similar to other studies (26–28). Overall, the cumulative number of ACEs showed a positive association with higher-level functional limitations; that is, the "dose–response" relationship was statistically significant. For example, the PR of one ACE for higher-level functional limitation was 1.21 (95% confidence interval [CI] = 1.16–1.26) in comparison with no ACE, and the PR of two or more ACEs was 1.61 (95% CI = 1.51–1.71), showing a significant dose–response association ($p$ for trend < 0.001). For two or more ACEs, the significant association remained even after correction for possible confounders. Association of other covariates and higher-level functional limitations can be found in Supplementary Table S3. Childhood economic hardship was also associated with higher-level functional limitations, after adjusting for age and gender (PR = 1.23, 95% CI = 1.18–1.27), and the point estimate was attenuated after adjusting for ACE and adulthood height (PR = 1.16, 95% CI = 1.11–1.20), suggesting that ACE and adulthood height (as a marker of cumulative nutritional and unhygienic exposure during early childhood) mediated the association between childhood economic hardship and higher-level functional limitations in older people.

Mediation analysis revealed that adult sociodemographics (marital status, living status, education, annual income, employment status), adult health behaviors (drinking, smoking), and health status (body mass index, depression, and disease status) mediated the association between ACEs and higher-level functional limitations by 66.1% (95% CI: 54.8%–77.5%), after adjustment of confounders (i.e., age, gender, childhood economic hardship, adult height). Thus, a significant portion of the association between ACEs and higher-level functional limitations was mediated by adult sociodemographics, health behaviors, and health status; however, independent effects exist, as hypothesized in Figure 1.

Regarding the three sublevels of competence (instrumental-self maintenance, intellectual activities, social role), the "dose–response" relationship between the cumulative number of ACEs was also statistically significant with all three sublevels of competence of the TMIG-IC (Supplementary Table S4).

The PR for higher-level functional limitations according to cumulative numbers of ACE stratified by age group and stratified by sex are shown in Supplementary Tables S5 and S6, respectively. There was no statistically significant interaction between the cumulative number of ACEs and age or between ACEs and gender.

## Discussion

In this large population-based study, the cumulative number of ACEs was significantly associated with higher-level functional limitations in older people without ADL limitations. Childhood disadvantage, adult sociodemographics, and adult health behaviors and health status attenuated the impact of ACEs on higher-level functional limitations. Generally, significant associations remained after adjustment for confounders and mediators, suggesting an independent effect of ACEs in older age (13).

Possible pathways include the following. First, ACE may harmfully affect the development of cognitive function during childhood, known to be a sensitive period. Danese and McEwen reported that those who with a history of ACEs showed smaller volumes of the prefrontal cortex and hippocampus in adulthood, suggesting that ACEs caused enduring changes in the nervous system (29). The prefrontal cortex plays important roles in cognitive control and the hippocampus has a specific role in the storage and recall of memory

**Table 1.** Characteristics of Sample (*n* = 19,220)

| Characteristics | Total | |
|---|---|---|
| | *N* | % |
| Demographics | | |
| Age | | |
| 65–69 years old | 6,138 | 31.9 |
| 70–74 years old | 6,063 | 31.6 |
| 75–79 years old | 3,959 | 20.6 |
| 80+ years old | 3,060 | 15.9 |
| Gender | | |
| Men | 9,060 | 47.1 |
| Childhood disadvantage | | |
| Economic hardship | 8,617 | 44.8 |
| Adulthood height | | |
| Short | 1,536 | 8.0 |
| Middle-short | 4,009 | 20.9 |
| Middle | 6,508 | 33.9 |
| Middle-tall | 4,760 | 24.8 |
| Tall | 2,407 | 12.5 |
| Adult sociodemographics | | |
| Education | | |
| <10 years | 7,000 | 36.4 |
| 10–12 years | 7,638 | 39.7 |
| 13+ years | 4,348 | 22.6 |
| Missing | 234 | 1.2 |
| Annual income | | |
| <1.5 million | 2,382 | 12.4 |
| 1.5–2.9 million | 6,078 | 31.6 |
| 3.0–4.9 million | 4,705 | 24.5 |
| 5.0+ million | 3,934 | 20.5 |
| Missing | 2,121 | 11.0 |
| Employment status | | |
| Current | 4,624 | 24.1 |
| Retirement | 11,684 | 60.8 |
| Never | 2,037 | 10.6 |
| Missing | 875 | 4.6 |
| Marital status | | |
| Married | 14,135 | 73.5 |
| Widow | 3,683 | 19.2 |
| Divorce/never married/other | 1,191 | 6.2 |
| Missing | 211 | 1.1 |
| Living status | | |
| Living alone | 2,497 | 13.0 |
| Living with wife/husband | 15,805 | 82.2 |
| Living with other | 238 | 1.2 |
| Missing | 680 | 3.5 |
| Health behaviors | | |
| Drinking | | |
| Current | 6,969 | 36.3 |
| Quit | 896 | 4.7 |
| Never | 11,261 | 58.6 |
| Missing | 94 | 0.5 |
| Smoking | | |
| Current | 1,973 | 10.3 |
| Quit | 3,077 | 16.0 |
| Never | 14,055 | 73.1 |
| Missing | 115 | 0.6 |
| Health status | | |
| BMI | | |
| <18.5 | 1,338 | 7.0 |
| 18.5–24.9 | 13,703 | 71.3 |
| 25.0+ | 4,179 | 21.7 |
| GDS | | |
| 0–4 | 12,918 | 67.2 |
| 5–9 | 3,073 | 16.0 |
| 10+ | 852 | 4.4 |
| Missing | 2,377 | 12.4 |
| Disease status, currently in treatment | 15,086 | 78.5 |
| Retrospectively reported adverse childhood experience (ACE) | | |
| Interpersonal loss | | |
| Parental death | 4,245 | 22.1 |
| Parental divorce | 378 | 2.0 |
| Family psychopathology | | |
| Parental mental illness | 138 | 0.7 |
| Family violence | 717 | 3.7 |
| Abuse and neglect | | |
| Physical abuse | 222 | 1.2 |
| Psychological neglect | 2,233 | 11.6 |
| Psychological abuse | 980 | 5.1 |
| ACE total score | | |
| 0 | 12,248 | 63.7 |
| 1 | 5,556 | 28.9 |
| 2+ | 1,416 | 7.4 |

*Note:* ACE = adverse childhood experience; BMI = body mass index; GDS = Geriatric Depression Scale.

sequences in context. These cognitive functions are important for higher-level functional limitations; thus, ACEs may lead to higher-level functional limitations in older people.

Alternatively, a pathway effect may explain the association (13,30). That is, ACEs may increase the risk of a low adult SES, which is associated with poor health behavior (such as smoking and heavy drinking) and disease status (such as cardiovascular disease), which may all be associated with higher-level functional limitations (a pathway model). In our study, the effects of ACEs on higher-level functional limitations were attenuated by adding adult sociodemographics, health behaviors, and health status, suggesting that low adult SES, poor health behaviors, and health status are possible leading causes of higher-level functional limitations in older people.

Consistent with results of previous studies investigating associations between childhood SES and disabilities in older people (1–3), we found that childhood economic hardship and adulthood height (as a marker of cumulative social and environmental exposures during early childhood) were important factors for higher-level functional limitations. A reason may be that poverty exposes children to the risk of low birthweight (31) and poor nutrition (32), and more harmful health risks, such as environmental pollutants and second hand smoke (33), which are all associated with cognitive function (34–37), which are important for higher-level functional capacity in older people. We add to the literature that the association between childhood SES and high-level functional limitations in older people may be explained by ACEs, because we observed that the point estimate of PR of childhood economic hardship for high-level functional limitation was attenuated after adjusting for ACEs (from 1.23 to 1.16). These results were consistent with a previous study showing that the impact of low childhood SES on mental disorders can be explained by ACEs (38). Moreover, to our knowledge, this is the first study to investigate the association between ACEs and higher-level functional capacity in older people.

The present study has several limitations. First, the retrospective assessment of ACEs may induce recall bias, which can be expected to occur within an older age cohort. However, the prevalence of psychological neglect, for example, was similar across each age cohort, and the

**Table 2.** Prevalence Ratio (95% Confidential Interval) of ACE for Higher-Level Functional Limitation

| Number of ACE | Reversed TIMG-IC (mean ± *SD*) | Crude | | Model 1 (adjusted age, sex) | | Model 2 (model 1 + childhood disadvantage) | | Model 3 (Model 2 + adult sociodemographics) | | Model 4 (Model 3 + adult health behavior and health status) | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | | PR | 95% CI | PR | 95% CI | PR | 95% CI | PR | 95% CI | PR | 95% CI |
| 0 | (1.25 ± 1.70) | Ref. | | Ref. | | Ref. | | Ref. | | Ref. | |
| 1 | (2.50 ± 1.92) | 1.21 | 1.16–1.26 | 1.17 | 1.12–1.22 | 1.13 | 1.09–1.18 | 1.09 | 1.05–1.14 | 1.06 | 1.02–1.10 |
| 2+ | (2.00 ± 2.22) | 1.61 | 1.51–1.71 | 1.55 | 1.46–1.65 | 1.46 | 1.37–1.55 | 1.34 | 1.26–1.43 | 1.19 | 1.12–1.27 |

*Note:* ACE = adverse childhood experience; BMI = body mass index; CI = confidence interval; GDS = Geriatric Depression Scale; PR = prevalence ratio; Ref = reference; *SD* = standard deviation; TIMG-IC = Tokyo Metropolitan Institute of Gerontology Index of Competence.

prevalence of parental death was higher among the older age cohort. This is reasonable to expect, as parents of the people in this cohort were more likely to have participated in World War II, suggesting that recall bias may have no significant fault. In addition, there is evidence that severe abuse is often remembered well and that false-positive reports are probably rare, whereas less severe adversities are more likely to be under reported (39). Thus, the association may be underestimated, but is less likely to be overestimated. Second, we assessed seven categories of ACE, each via a single item. Thus, other ACEs, such as sexual abuse and parental substance abuse, were not included, because of the feasibility according to the municipalities. However, childhood sexual abuse and parental substance abuse are rare in Japan (40). Third, there is a common method bias because all measures were assessed by a single self-reported questionnaire. Reuben et al. reported that as compared to prospective ACE studies, retrospective ACE studies showed stronger associations with subjectively assessed outcomes, that is, our results might have overestimated the association (41). Fourth, poor childhood health may also affect a variety of health outcomes, but these were not assessed, and we were unable to examine any effects on functional limitations. Fifth, because we used data from the JAGES study, in which the participants were healthy enough to participate in the survey, surviving bias could result in an underestimation of the association between ACEs and higher-level functional limitations among older people. Finally, our analytical sample excluded people who have a disability, so the generalizability of the findings may not be extended to those with disabilities.

In conclusion, our findings provide evidence for an independent effect of ACEs on higher-level functional limitations among older people in Japan. In addition to interventions aimed at improving health behaviors in adulthood, policies aimed at reducing ACEs may help decrease demand for medical and social care services associated with functional limitations in old age.

## Supplementary Material

Supplementary data is available at *Journals of Gerontology, Series A: Biological Sciences and Medical Sciences* online.

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## Conflict of Interest

The authors have no conflict of interest directly relevant to the content of this article.

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# Long-term Impact of Childhood Disadvantage on Late-Life Functional Decline Among Older Japanese: Results From the JAGES Prospective Cohort Study

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Abstract

Background

Increasing evidence suggests an impact of childhood disadvantage on late-life functional impairment in Western countries. However, the processes by which childhood disadvantage affects functional capacity are influenced by several factors unique to particular societies. We examined the impact of childhood disadvantage on functional decline among older Japanese, using a large-scale prospective cohort study.

Methods

Data came from surveys conducted in 2010 and 2013 as part of the Japan Gerontological Evaluation Study (JAGES), a nationwide cohort study targeting community-dwelling people aged 65 years and over. Childhood disadvantage included subjective childhood socioeconomic status (SES), body height, and educational level. The sample was stratified by age at baseline (65–69, 70–74, 75–79, and ≥ 80 years).

Results

A total of 11,601 respondents were analyzed. In the 65–69-year group, lower childhood SES was associated with functional decline but this association was mediated by adult SES. In contrast, childhood SES was independently associated with functional decline in the older cohort. In the 75–79-year group, lower childhood SES was associated with functional decline. However, in the ≥ 80-year group, people with higher childhood SES were more likely to experience functional decline. Shorter height was associated with functional decline in the 70–74-year group. Higher education was related to functional decline in all age groups except the ≥ 80-year group.

Conclusions

These findings suggest that childhood disadvantage affects functional decline but its effect varies by age cohort. The mechanisms underlying the association between childhood disadvantage and functional decline may be influenced by social and historical context.

Functional decline, Childhood disadvantage, Socioeconomic status, Body height, Education, Japan
Topic: adult body height child prospective studies socioeconomic factors japanese
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# Community social capital and inequality in depressive symptoms among older Japanese adults: A multilevel study

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ABSTRACT

Although studies have suggested that community social capital contributes to narrow income-based inequality in depression, the impacts may depend on its components. Our multilevel cross-sectional analysis of data from 42,208 men and 45,448 women aged 65 years or older living in 565 school districts in Japan found that higher community-level civic participation (i.e., average levels of group participation in the community) was positively associated with the prevalence of depressive symptoms among the low-income groups, independent of individual levels of group participation. Two other social capital components (cohesion and reciprocity) did not significantly alter the association between income and depressive symptoms.

## 1. Introduction

Ageing is a major risk factor for many chronic diseases and physical and mental illnesses. Specifically, in Japan, Wada demonstrated that in the years 2000/2001, 33.5% of older people in four rural towns had depressive symptoms, defined as scoring more than 5 points on the Geriatric Depression Scale (GDS-15) (Wada et al., 2004). Depression can cause other critical health issues including suicide, frailty, functional disability, and mortality (Waern et al., 2003; Kondo et al., 2008; Wada et al., 2004). Therefore, depression is a key target of the public health actions targeted to older adults in Japan and worldwide (World Health Organization, 2010; Ministry of Health Labour and Welfare, 2012).

Socioeconomic inequalities in the prevalence of depressive symptoms among older adults are also major targets of public health actions, as social epidemiology studies have identified depression as strongly concentrated among the socially disadvantaged older population (Murata et al., 2008; Cole and Dendukuri, 2003).

Depression can be affected by psychosocial conditions in the community. Recent studies have suggested that high community social capital—defined as “resources that are accessed by individuals as a result of their membership in a network or a group”—is associated with fewer individual risks for depressive symptoms in older adults (Ivey et al., 2015; Kawachi et al., 1999). However, the evidence is scarce on whether or not high community social capital is associated with the smaller socioeconomic disparity in depressive symptoms. Poverty puts one at risk for social isolation, which, in turn, puts one at risk for depression. Therefore, we could hypothesize that communities rich in social capital may contribute more to those who are worse off, according to their contextual characteristics as positive externalities (Berkman et al., 2014). Our previous ecological study in Japan showed an inverse relationship between community-level social capital and income-based inequality in depressive symptoms among older adults (Haseda et al., 2018). Although this ecological study supported the hypothesis, it could not distinguish the compositional and contextual effects of community social capital in reducing income-based inequality in depressive symptoms; thus, a more sophisticated study is necessary: that is, a multilevel analysis to assess whether or not community-level social capital modifies the association between individual income and the prevalence of depressive symptoms. Moreover, we also hypothesize that the association between community social capital and income-based inequality in depressive symptoms may differ by the dimensions of the former. Specifically, based on recent discussions, we focused on the structural and cognitive aspects of community social capital, which when evaluated as cognitive social cohesion may contribute to individuals who comprise the community universally, regardless of socioeconomic background. This is because people inhabiting a community can reap the benefits of social cohesion equally, given that cohesion is nonexcludable (Berkman et al., 2014). In theory and practice, cognitive social capital should and has been evaluated in terms of

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the levels of social cohesion (trust and reciprocity) (Berkman et al., 2014). However, community social capital evaluated as a structural aspect (levels of civic participation or the opportunities for participation (Berkman et al., 2014)) may not necessarily benefit all people. This may be specifically so in a highly segregated community, because the organizations for the rich and the poor may be completely different, making positive spillover less likely to occur. Hence, the average level of structural social capital cannot reflect equality in the opportunities for social interaction. To differentiate the structural and cognitive aspects of community social capital, we used a recently developed validated social capital scale.

Thus, the purpose of this study was to evaluate the variable associations between individual income and depressive symptoms across communities with different levels of structural and cognitive social capital, using large-scale multilevel data of older Japanese adults.

## 2. Methods

### 2.1. Data

We used cross-sectional data of the 2013 wave of the Japan Gerontological Evaluation Study (JAGES). The JAGES 2013 wave was an anonymous self-administered mail-in survey across 30 municipalities in 13 out of 47 prefectures in Japan. Although the participating municipalities are not nationally representative, they vary significantly and include small rural towns to large metropolitan cities from areas in the North (Hokkaido) and South (Kumamoto) ends. Their population ranged between 1246 and 3.7 million people, and the share of the older population was between 18.0% and 47.6% (Statistics Bureau; Ministry of Internal Affairs and Communications, 2013). Participants who were aged 65 years or older were functionally independent in their daily living; that is, they did not receive benefits from public long-term care insurance at that time. Participants living in 16 large municipalities were randomly selected by multistage sampling, while all eligible individuals living in 14 small municipalities were selected. We mailed 193,694 questionnaires, of which 137,736 were returned (response rate = 71.1%). We excluded the responses without valid values for key variables (age, gender, the area of residence, and depressive symptoms) from the analyses.

### 2.2. Measures

#### 2.2.1. Depressive symptoms

We used the Japanese short version of the Geriatric Depression Scale (GDS-15) developed for self-administered surveys to assess depressive symptoms (Yesavage and Sheikh, 1986; Niino et al., 1991). We set a cut-off score of 4/5, which has been universally adopted as indicating a depressive tendency based on validation studies (Nyunt et al., 2009).

#### 2.2.2. Income

We gathered information on annual income by asking, "What was your pretax annual household income for 2012 (including pension)?" in 15 predetermined categories (in thousands of yen). We calculated household income equivalized by dividing each response by the square root of the number of household members and further dividing them into tertiles.

#### 2.2.3. Community social capital

We used the Health-Related Social Capital Scale developed and validated among older Japanese people by Saito et al. (2017). The scale has three subscales that assess civic participation, social cohesion, and reciprocity. Scores for each subscale are derived from the summation of the percentage that responded to multiple questions and are standardized. Items for civic participation concern participation in five types of groups in the community: volunteering, sports, hobby, culture, and skill teaching. Social cohesion items concern trust ("Do you think that people living in your area can be trusted, in general?"), others' perceived intention to help ("Do you think that people living in your area try to help others in most situations?"), and attachment to the residential area ("How attached are you to the area in which you live?"). The sum of the percentage of those who answered "very" or "moderately" to the items formed the score. Reciprocity items concern having someone to provide or receive emotional support or to receive instrumental support: "Do you have someone who listens to your concerns and complaints?" "Do you listen to someone's concerns and complaints?" and "Do you have someone who looks after you when you are sick and confined to a bed for a few days?" The sum of the percentage of those who designated anyone to the questions formed the score.

### 2.3. Covariates

Referring to recent social epidemiology studies, we considered the following variables as potential confounding factors in the association between income and depression among older adults: age, years of education (less than nine years or not), marital status (having a spouse or not), living alone or not, comorbidities (having past medical history of stroke, heart disease, diabetes mellitus, cancer, dementia, or Parkinson's disease), and frequency of going out (Chang-Quan et al., 2010; Yan et al., 2011; Stahl et al., 2016; Evans et al., 2005; Cohen-Mansfield et al., 2010). We also considered the fixed effect of each municipality.

### 2.4. Statistical analysis

We conducted a multilevel Poisson regression analysis incorporating both individuals and school districts. As recent studies report that community-level factors had different impacts with regard to gender, we modeled men and women separately (Eriksson et al., 2011; Pattyn et al., 2011; Haseda et al., 2018). School districts are the smallest areal units identifiable using JAGES data. Within 30 municipalities, there are 565 school districts. The district originally represents the district unit determining the catchment area of each public school. We chose to use this areal unit because a school district is likely to represent Japanese "communities" developed in its local history, such as "*kyu-son*" (former village areas). We believed that the school district could represent the adequate areal size to reflect community social capital that could work as the informal resource contributing to community autonomy.

We statistically investigated the effect modifications of community social capital in terms of civic participation, social cohesion, and reciprocity on the association between income and depressive symptoms. To do so, in the Poisson regression, we modeled cross-level interaction terms between each standardized community social capital component and income besides those variables' main effects. In addition to the covariates explained above, we modeled individual responses to each social capital component as potential confounders. Each individual-level social capital component was binarized in our analyses; that is, those who participated in any of the five kinds of groups in their community were considered to engage in individual civic participation; those who showed trust, others' intention to help, or attachment to the residential area were regarded as individually socially cohesive; and those who answered that they have someone to provide or receive emotional or instrumental support were considered to have individual social support. We took into account missing values assigning dummy variables for the missing category. We used Stata version 14.1 for these statistical analyses. (Stata Corp. Texas, USA)

## 3. Results

We analyzed 87,656 individuals (42,208 men and 45,448 women) after excluding those with missing responses to key variables (n = 7996

**Table 1**
Descriptive characteristics of participants with depressive symptoms.

| | Participants (n = 87,656) | Have depressive symptoms | | | |
|---|---|---|---|---|---|
| | Total (Men/Women) | Men (n = 12,704) | | Women (n = 13,861) | |
| | n or mean [SD] | n (%) or mean [SD] | | n (%) or mean [SD] | |
| Age | | | | | |
| 65–74 | 52,294 (25,531/ 26,763) | 6996 | (25.3) | 7220 | (23.4) |
| 75–84 | 29,963 (14,333/ 15,630) | 4808 | (29.5) | 5317 | (27.4) |
| 85 and older | 5399 (2344/ 3055) | 900 | (32.0) | 1324 | (34.2) |
| Income | | | | | |
| T1 (Low) | 22,816 (10,349/ 12,467) | 4369 | (38.3) | 4896 | (34.0) |
| T2 (Middle) | 20,047 (10,880/ 9367) | 3268 | (28.2) | 2683 | (25.0) |
| T3 (High) | 28,928 (15,366/ 13,562) | 2847 | (17.1) | 2639 | (17.3) |
| Missing | 15,865 (5813/ 10,052) | 2220 | (30.9) | 3643 | (26.6) |
| Education | | | | | |
| < 9 years | 37,669 (16,690/ 20,979) | 6132 | (32.6) | 7344 | (29.1) |
| > =9 years | 48,481 (24,904/ 23,577) | 6309 | (23.2) | 6102 | (22.3) |
| Missing | 1506 (614/ 892) | 263 | (31.1) | 415 | (28.1) |
| Comorbidities | | | | | |
| none | 58,160 (25,514/ 32,646) | 6942 | (24.5) | 9127 | (23.6) |
| 1 or more | 24,004 (14,185/ 9819) | 5112 | (32.8) | 3908 | (33.8) |
| Missing | 5492 (2509/ 2983) | 650 | (22.3) | 826 | (21.4) |
| Living alone | | | | | |
| Yes | 72,576 (3604/ 7581) | 1844 | (46.3) | 2982 | (32.5) |
| No | 11,185 (37,005/ 35,571) | 10,205 | (25.0) | 9973 | (23.9) |
| Missing | 3895 (1599/ 2296) | 655 | (31.8) | 906 | (28.5) |
| Having spouse | | | | | |
| Yes | 65,850 (38,815/ 27,035) | 9741 | (24.7) | 7131 | (22.6) |
| No | 20,027 (2732/ 17,295) | 2620 | (41.0) | 6208 | (29.9) |
| Missing | 1779 (661/ 1118) | 343 | (35.0) | 522 | (28.6) |
| Frequency of going out | | | | | |
| > = 1/week | 82,904 (39,978/ 42,926) | 11,495 | (26.0) | 12,422 | (24.4) |
| < 1/week | 3521 (1650/ 1871) | 983 | (52.9) | 1159 | (51.4) |
| Missing | 1231 (580/ 651) | 226 | (30.3) | 280 | (27.8) |
| Participation in social groups | | | | | |
| Any participation | 35,105 (15,300/ 19,805) | 2906 | (17.3) | 4043 | (17.4) |
| No participation | 41,247 (21,982/ 19,265) | 8011 | (33.4) | 7289 | (33.1) |
| Missing | 11,304 (4926/ 6378) | 1787 | (29.9) | 2529 | (28.8) |
| Individual social cohesion | | | | | |
| Cohesive | 74,739 (36,331/ 38,408) | 9459 | (50.4) | 10,065 | (46.4) |
| Not cohesive | 11,814 (5454/ 6360) | 2982 | (23.7) | 3392 | (22.2) |
| Missing | 1103 (423/ 680) | 263 | (28.2) | 404 | (27.2) |
| Individual social support | | | | | |
| Any support | 85,783 (41,044/ 44,739) | 11,929 | (26.4) | 13,379 | (25.3) |
| No support | 1177 (818/ 359) | 566 | (65.2) | 266 | (66.8) |
| Missing | 696 (346/ 350) | 209 | (27.1) | 216 | (26.3) |
| Community-level social capital (unstandardized value[a]) | | | | | |
| Community civic participation | 0.84 [0.17] (0.85 [0.17]/ 0.84 [0.18]) | | | | |
| Community social cohesion | 2.01 [0.15] (2.01 [0.15]/ 2.01 [0.15]) | | | | |
| Community reciprocity | 2.82 [0.05] (2.82 [0.05]/ 2.82 [0.05]) | | | | |

[a] These values were standardized in the later analysis.

for age and gender, n = 28,786 for the area of residence, and n = 2506 for depressive symptoms) among eligible subjects. Descriptive statistics showed that, similar to previous studies, the percentage of people who showed depressive symptoms was more prevalent in lower-income groups. The prevalence of depressive symptoms in lowest-, middle-, and highest-income groups were 42.2%, 30.6%, and 18.5% among men and 39.3%, 28.6%, and 19.5% among women, respectively (Table 1).

Results of the Poisson regression showed that even after adjusting for age, educational attainment, marital status, living arrangement, the presence of comorbidities, frequency of going out, and the dummy variables representing the municipality or residence, low income was associated with the high prevalence of depressive symptoms. The adjusted prevalence ratio (PR) of depressive symptoms in the low-income group among men compared to the high-income group was 1.89 (95% confidence intervals [CI]: 1.80, 1.98), and the adjusted PR was 1.49 (95% CI: 1.42, 1.57) for the middle-income group (Model 2 in Table 2). Among women, the PR of depressive symptoms in the low-income group was 1.65 (95% CI: 1.57, 1.73) and 1.35 (95% CI: 1.27, 1.42) in the middle-income group compared to the high-income group. All community social capital components tended to be inversely associated with depressive symptoms. The adjusted PR per 1 standard deviation (SD) unit increase in community civic participation was 0.99 (95% CI: 0.97, 1.01) among men and 1.00 (95% CI: 0.97, 1.02) among women (Table 2). The adjusted PR per 1 SD unit increase in community social cohesion was 0.98 (95% CI: 0.96, 1.00) among men and 0.97 (95% CI: 0.95, 1.00) among women (Table 3). The adjusted PR per 1 SD unit increase in community reciprocity was 0.98 (95% CI: 0.96, 1.00) among men and 0.98 (95%CI: 0.96, 1.00) among women (Table 4).

When we adjusted for covariates and cross-level interaction terms (Model 3 in Table 2), the association between low income and the high prevalence of depressive symptoms was stronger among those residing in the areas with high community-level civic participation ($P = 0.016$ in men, $P = 0.080$ in women, Fig. 1, Supplementary 1). The difference in predicted prevalence of depressive symptoms between the lowest- and highest-income groups was 17.4% points among men and 14.4% points among women where community-level civic participation was the mean level, while the difference was 18.8% points among men and 15.4% points among women where community-level civic participation

**Table 2**
Prevalence ratio (95% confidence intervals [CI]) for depressive symptoms: results of multilevel Poisson regression analysis modeling community civic participation.

| | Men (n = 42,208) | | | | Women (n = 45,448) | | | |
|---|---|---|---|---|---|---|---|---|
| | Null model PR [95% CI] | Model 1 Age-adjusted PR [95% CI] | Model 2 Multilevel PR [95% CI] | Model 3 Interaction added PR [95% CI] | Null model PR [95% CI] | Model 1 Age-adjusted PR [95% CI] | Model 2 Multilevel PR [95% CI] | Model 3 Interaction added PR [95% CI] |
| **Individual factors** | | | | | | | | |
| Intercept | 0.30 [0.29, 0.31] | – | 0.08 [0.06, 0.11] | 0.08 [0.06, 0.11] | 0.30 [0.29, 0.30] | – | 0.12 [0.09, 0.16] | 0.13 [0.10, 0.17] |
| Income (ref: T3 [High]) | | | | | | | | |
| T1 (Low) | | 2.24 [2.14, 2.35] | 1.89 [1.80, 1.98] | 1.81 [1.71, 1.93] | | 1.95 [1.86, 2.05] | 1.65 [1.57, 1.73] | 1.50 [1.41, 1.61] |
| T2 (Middle) | | 1.65 [1.57, 1.73] | 1.49 [1.42, 1.57] | 1.43 [1.34, 1.52] | | 1.46 [1.38, 1.54] | 1.35 [1.27, 1.42] | 1.25 [1.16, 1.34] |
| Education ≤ 9 years (ref: > 9 years) | | 1.40 [1.35, 1.45] | 1.13 [1.09, 1.17] | 1.13 [1.09, 1.17] | | 1.29 [1.24, 1.33] | 1.08 [1.04, 1.12] | 1.08 [1.04, 1.12] |
| Comorbidities: yes (ref: none) | | 1.30 [1.26, 1.35] | 1.25 [1.20, 1.29] | 1.25 [1.20, 1.29] | | 1.37 [1.32, 1.42] | 1.29 [1.24, 1.34] | 1.29 [1.24, 1.34] |
| Living alone (ref: no) | | 1.84 [1.75, 1.93] | 1.42 [1.33, 1.53] | 1.42 [1.33, 1.53] | | 1.34 [1.28, 1.39] | 1.23 [1.17, 1.29] | 1.23 [1.17, 1.29] |
| No spouse (ref: having spouse) | | 1.63 [1.56, 1.70] | 1.20 [1.13, 1.28] | 1.20 [1.13, 1.28] | | 1.24 [1.20, 1.29] | 1.07 [1.03, 1.12] | 1.07 [1.02, 1.11] |
| Freq. of going out < 1/week (ref: ≥1/week) | | 1.94 [1.82, 2.08] | 1.52 [1.42, 1.63] | 1.53 [1.43, 1.63] | | 1.88 [1.76, 2.00] | 1.56 [1.47, 1.67] | 1.57 [1.47, 1.67] |
| Participation in social groups (ref: no participation)[a] | | 0.52 [0.50, 0.55] | 0.61 [0.58, 0.63] | 0.55 [0.50, 0.60] | | 0.56 [0.54, 0.58] | 0.62 [0.59, 0.64] | 0.54 [0.50, 0.59] |
| **Community-level factors** | | | | | | | | |
| Community civic participation (CP) (per 1SD increment) | | 0.93 [0.91, 0.94] | 0.99 [0.97, 1.01] | 0.96 [0.92, 0.99] | | 0.92 [0.90, 0.94] | 1.00 [0.97, 1.02] | 0.97 [0.93, 1.01] |
| **Interaction terms** | | | | | | | | |
| Individual SC × income level (ref: T3 [High]) | | | | | | | | |
| T1 (Low) | | | | 1.14 [1.01, 1.27] | | | | 1.26 [1.13, 1.40] |
| T2 (Middle) | | | | 1.14 [1.01, 1.28] | | | | 1.16 [1.03, 1.31] |
| Community CP x income level (ref: T3 [High]) | | | | | | | | |
| T1 (Low) | | | | 1.06 [1.01, 1.11] | | | | 1.04 [0.99, 1.10] |
| T2 (Middle) | | | | 1.02 [0.97, 1.08] | | | | 1.05 [0.99, 1.11] |
| **Variance components** | | | | | | | | |
| Residual district-level variance | 0.004 [0.002] | N/A | $2.43e^{-34}$ [$4.31e^{-19}$] | $1.99e^{-34}$ [$3.99e^{-19}$] | 0.014 [0.003] | N/A | $1.12e^{-34}$ [$2.85e^{-19}$] | $2.71e^{-34}$ [$5.07e^{-19}$] |

All models except for the null model are age-adjusted.
Missing values were taken into account in all variables including interaction terms.
[a] Evaluated using the same questions for calculating the community civic participation score.

**Table 3**
Prevalence ratio (95% confidence intervals [CI]) for depressive symptoms: results of multilevel Poisson regression analysis modeling community social cohesion.

| | Men (n = 42,208) | | | | Women (n = 45,448) | | | |
|---|---|---|---|---|---|---|---|---|
| | Null model | Model 1 Age-adjusted | Model 2 Multilevel | Model 3 Interaction added | Null model | Model 1 Age-adjusted | Model 2 Multilevel | Model 3 Interaction added |
| | PR [95% CI] | PR [95% CI] | PR [95% CI] | PR [95% CI] | PR [95% CI] | PR [95% CI] | PR [95% CI] | PR [95% CI] |
| **Individual factors** | | | | | | | | |
| Intercept | 0.30 [0.29, 0.31] | | 0.09 [0.07, 0.12] | 0.11 [0.08, 0.15] | 0.30 [0.29, 0.30] | | 0.11 [0.09, 0.14] | 0.12 [0.09, 0.16] |
| Income (ref: T3 [High]) | | | | | | | | |
| T1 (Low) | | 2.24 [2.14, 2.35] | 1.88 [1.78, 1.97] | 1.44 [1.30, 1.60] | | 1.95 [1.86, 2.05] | 1.65 [1.57, 1.73] | 1.44 [1.30, 1.60] |
| T2 (Middle) | | 1.65 [1.57, 1.73] | 1.49 [1.42, 1.57] | 1.29 [1.15, 1.44] | | 1.46 [1.38, 1.54] | 1.34 [1.27, 1.42] | 1.27 [1.13, 1.43] |
| Education < =9 year (ref: > 9 year) | | 1.40 [1.35, 1.45] | 1.17 [1.12, 1.21] | 1.17 [1.12, 1.21] | | 1.29 [1.24, 1.33] | 1.12 [1.08, 1.16] | 1.12 [1.08, 1.16] |
| Comorbidities: yes (ref: none) | | 1.30 [1.26, 1.35] | 1.25 [1.20, 1.29] | 1.25 [1.20, 1.29] | | 1.37 [1.32, 1.42] | 1.30 [1.25, 1.35] | 1.30 [1.25, 1.35] |
| Living alone (ref: no) | | 1.84 [1.75, 1.93] | 1.35 [1.25, 1.44] | 1.35 [1.25, 1.44] | | 1.34 [1.28, 1.39] | 1.17 [1.11, 1.23] | 1.17 [1.12, 1.23] |
| No spouse (ref: having spouse) | | 1.63 [1.56, 1.70] | 1.19 [1.12, 1.27] | 1.19 [1.12, 1.27] | | 1.24 [1.20, 1.29] | 1.06 [1.02, 1.11] | 1.06 [1.02, 1.11] |
| Freq. of going out < 1/week (ref: > =1/week) | | 1.94 [1.82, 2.08] | 1.54 [1.45, 1.65] | 1.55 [1.45, 1.66] | | 1.88 [1.76, 2.00] | 1.59 [1.49, 1.70] | 1.60 [1.50, 1.70] |
| Individual social cohesion (SC)[a] (ref: not cohesive) | | 0.47 [0.45, 0.49] | 0.56 [0.53, 0.58] | 0.45 [0.41, 0.49] | | 0.48 [0.46, 0.50] | 0.54 [0.52, 0.57] | 0.49 [0.44, 0.54] |
| **Community-level factors** | | | | | | | | |
| Community-level social cohesion (SC) (per 1SD increment) | | 0.95 [0.93, 0.96] | 0.98 [0.96, 1.00] | 0.99 [0.95, 1.03] | | 0.95 [0.94, 0.97] | 0.97 [0.95, 1.00] | 0.98 [0.94, 1.02] |
| **Interaction term** | | | | | | | | |
| Individual SC × Income level (ref: T3 [High]) | | | | | | | | |
| T1 (Low) | | | | 1.40 [1.25, 1.57] | | | | 1.19 [1.06, 1.34] |
| T2 (Middle) | | | | 1.19 [1.05, 1.35] | | | | 1.06 [0.93, 1.22] |
| Community SC x Income level (ref: T3 (High)) | | | | | | | | |
| T1 (Low) | | | | 0.98 [0.94, 1.03] | | | | 0.99 [0.94, 1.04] |
| T2 (Middle) | | | | 0.98 [0.93, 1.03] | | | | 0.98 [0.93, 1.04] |
| **Variance components** | | | | | | | | |
| Residual district-level variance | 0.004 [0.002] | N/A | $4.79e^{-35}$ [$1.53e^{-19}$] | $4.11e^{-35}$ [$1.71e^{-19}$] | 0.014 [0.003] | N/A | $6.24e^{-35}$ [$2.02e^{-19}$] | $5.53e^{-35}$ [$1.74e^{-19}$] |

All models except for the null model are age-adjusted.
Missing values were taken into account in all variables including interaction terms.
[a] Evaluated using the same questions for calculating the community social cohesion score.

**Table 4**
Prevalence ratio (95% confidence intervals [CI]) for depressive symptoms: results of multilevel Poisson regression analysis modeling community reciprocity.

| | Men (n = 42,208) | | | | Women (n = 45,448) | | | |
|---|---|---|---|---|---|---|---|---|
| | Null model PR [95% CI] | Model 1 Age-adjusted PR [95% CI] | Model 2 Multilevel PR [95% CI] | Model 3 Interaction added PR [95% CI] | Null model PR [95% CI] | Model 1 Age-adjusted PR [95% CI] | Model 2 Multilevel PR [95% CI] | Model 3 Interaction added PR [95% CI] |
| **Individual factors** | | | | | | | | |
| Intercept | 0.30 [0.29, 0.31] | | 0.10 [0.07, 0.13] | 0.13 [0.09, 0.19] | 0.30 [0.29, 0.30] | | 0.14 [0.11, 0.19] | 0.21 [0.12, 0.35] |
| Income (ref: T3 [High]) | | | | | | | | |
| T1 (Low) | | 2.24 [2.14, 2.35] | 1.96 [1.86, 2.06] | 1.39 [1.03, 1.88] | | 1.95 [1.86, 2.05] | 1.73 [1.64, 1.81] | 1.22 [0.75, 1.98] |
| T2 (Middle) | | 1.65 [1.57, 1.73] | 1.52 [1.45, 1.60] | 1.20 [0.87, 1.65] | | 1.46 [1.38, 1.54] | 1.37 [1.29, 1.44] | 1.00 [0.59, 1.71] |
| Education < =9 year (ref: > 9 year) | | 1.40 [1.35, 1.45] | 1.19 [1.14, 1.23] | 1.19 [1.14, 1.23] | | 1.29 [1.24, 1.33] | 1.15 [1.11, 1.19] | 1.15 [1.11, 1.19] |
| Comorbidities: yes (ref: none) | | 1.30 [1.26, 1.35] | 1.27 [1.22, 1.31] | 1.27 [1.22, 1.31] | | 1.37 [1.32, 1.42] | 1.31 [1.26, 1.36] | 1.31 [1.26, 1.36] |
| Living alone (ref: no) | | 1.84 [1.75, 1.93] | 1.35 [1.25, 1.45] | 1.35 [1.25, 1.45] | | 1.34 [1.28, 1.39] | 1.18 [1.12, 1.24] | 1.18 [1.12, 1.24] |
| No spouse (ref: having spouse) | | 1.63 [1.56, 1.70] | 1.21 [1.14, 1.28] | 1.21 [1.13, 1.28] | | 1.24 [1.20, 1.29] | 1.08 [1.03, 1.12] | 1.08 [1.03, 1.13] |
| Freq. of going out < 1/week (ref: > =1/week) | | 1.94 [1.82, 2.08] | 1.62 [1.51, 1.73] | 1.62 [1.52, 1.73] | | 1.88 [1.76, 2.00] | 1.68 [1.58, 1.79] | 1.69 [1.58, 1.80] |
| Individual social support (ref: no support)[a] | | 0.41 [0.38, 0.45] | 0.65 [0.60, 0.72] | 0.49 [0.37, 0.64] | | 0.42 [0.37, 0.47] | 0.55 [0.49, 0.63] | 0.39 [0.25, 0.61] |
| **Community-level factors** | | | | | | | | |
| Community reciprocity (per 1SD increment) | | 0.95 [0.94, 0.97] | 0.98 [0.96, 1.00] | 0.99 [0.95, 1.03] | | 0.96 [0.94, 0.97] | 0.98 [0.96, 1.00] | 0.99 [0.95, 1.03] |
| **Interaction terms** | | | | | | | | |
| Individual social support × income level (ref: T3 [High]) | | | | | | | | |
| T1 (Low) | | | | 1.42 [1.05, 1.93] | | | | 1.42 [0.87, 2.30] |
| T2 (Middle) | | | | 1.27 [0.92, 1.76] | | | | 1.37 [0.80, 2.34] |
| Community reciprocity x income level (ref: T3 [High]) | | | | | | | | |
| T1 (Low) | | | | 0.98 [0.94, 1.03] | | | | 1.00 [0.95, 1.04] |
| T2 (Middle) | | | | 0.99 [0.94, 1.04] | | | | 0.99 [0.94, 1.05] |
| **Variance components** | | | | | | | | |
| Residual district-level variance | 0.004 [0.002] | N/A | $6.91e^{-34}$ [$2.06e^{-19}$] | $1.17e^{-34}$ [$2.72e^{-19}$] | 0.014 [0.003] | N/A | $4.49e^{-34}$ [$1.39e^{-19}$] | $9.34e^{-35}$ [$2.29e^{-19}$] |

All models except for the null model are age-adjusted.
Missing values were taken into account in all variables including interaction terms.
[a] Evaluated using the same questions for calculating the community reciprocity score.

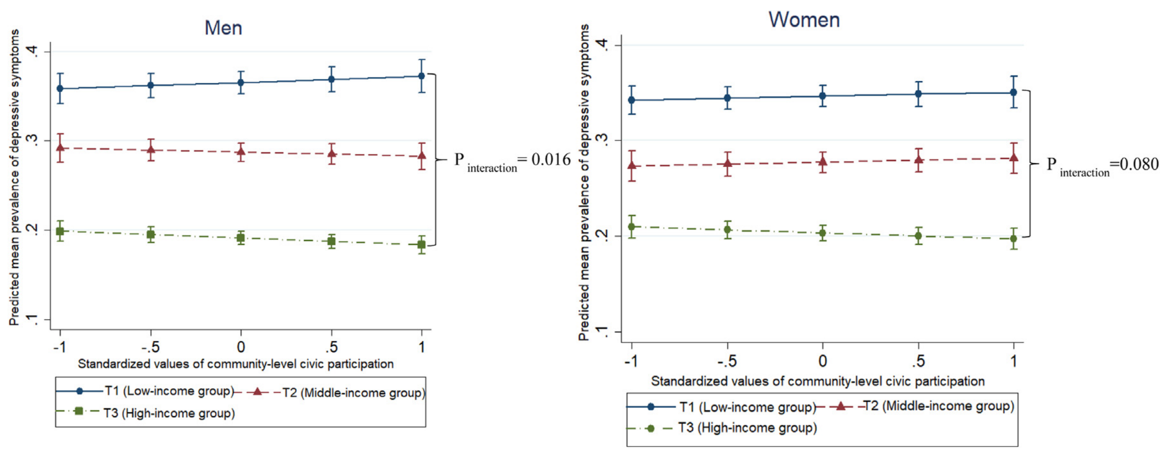


**Fig. 1.** Effect modification by community civic participation levels on the association between income and depressive symptoms: predicted mean values of depressive symptoms (with 95% confidence intervals) by community civic participation across income tertiles. Error bars are 95% confidence intervals of predicted mean values of depressive symptoms. $P_{interaction}$ is the probability that the slope of low- and high-income groups in the prevalence of depressive symptoms across community civic participation levels would be the same as or more extreme than the actual observed results.

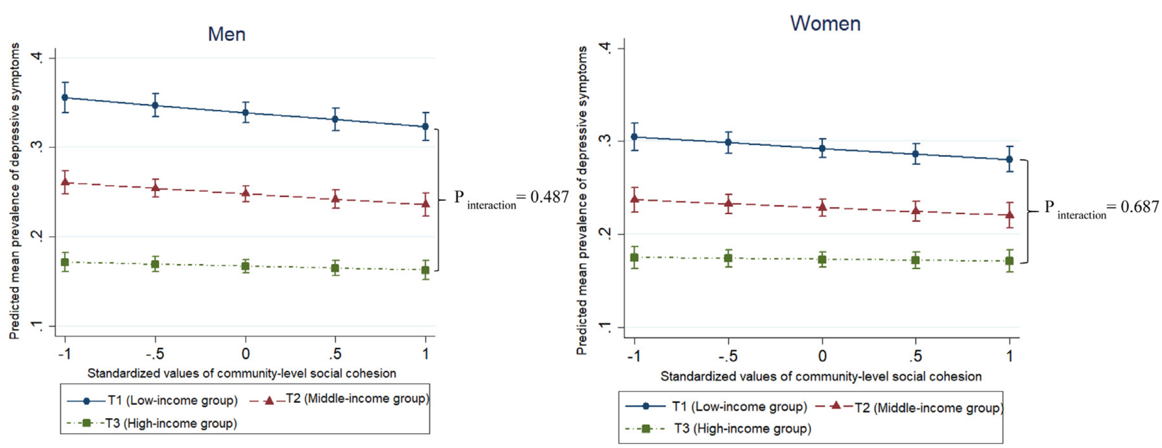


**Fig. 2.** Effect modification by community social cohesion levels on the association between income and depressive symptoms: predicted mean values of depressive symptoms (with 95% confidence intervals) by community social cohesion across income tertiles. Error bars are 95% confidence intervals of predicted mean values of depressive symptoms. $P_{interaction}$ is the probability that the slope of low and high-income groups in the prevalence of depressive symptoms across community social cohesion levels would be the same as or more extreme than the actual observed results.

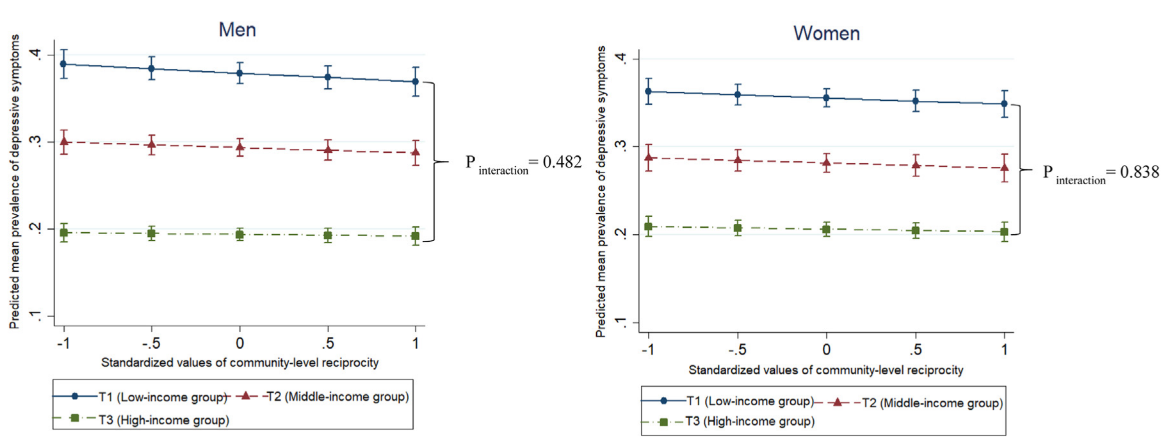


**Fig. 3.** Effect modification by community social cohesion levels on the association between income and depressive symptoms: predicted mean values of depressive symptoms (with 95% confidence intervals) by community reciprocity across income tertiles. Error bars are 95% confidence intervals of predicted mean values of depressive symptoms. $P_{interaction}$ is the probability that the slope of low and high-income groups in the prevalence of depressive symptoms across community reciprocity levels would be the same as or more extreme than the actual observed results.

was the + 1 SD level. We did not find statistically significant effect modification by the levels of community social cohesion or community reciprocity for the association between income and depressive symptoms (Tables 3 and 4). The decrement of difference in the predicted prevalence of depressive symptoms between the lowest- and highest-income groups was 0.8% points among men (P = 0.487) and

0.6% points among women (P = 0.687) per 1 SD unit increase in community-level social cohesion and 0.8% points among men (P = 0.482) and 0.4% points among women (P = 0.838) per 1 SD unit increase in community-level reciprocity (Figs. 2 and 3, Supplementary 1).

## 4. Discussion

The results of this study showed that the association between income and depressive symptoms varied across communities in relation to the levels of community social capital components, independent of community- and individual-level potential confounders and individual responses to social capital questions. Although the results suggest the overall preventive effect of community social capital on depressive symptoms, the effect size was small. The direction of the effect modification for the association between low income and high prevalence of depression differed by components; civic participation potentially strengthened the association, whereas social cohesion and community reciprocity did not modify the association.

Recent literature reviews showed that high community-level cognitive social capital is associated with a lower risk of developing common mental disorders (Ehsan and De Silva, 2015). The results of our study corresponded to this finding, showing the tendency of a smaller prevalence of depressive symptoms among residents of communities with richer social capital than among those living in communities with poorer social capital. However, in our study, the effect size was very small. Moreover, in contrast to the expectation, we could not detect a clear buffering effect of community-level social cohesion and reciprocity. Thus, our previous ecological observation might reflect not the contextual effect but the compositional effect of community-level social cohesion and reciprocity toward income-based inequality in depressive symptoms (Haseda et al., 2018). The results of the present study were not in line with another study based on JAGES, which suggests that neighborhood social cohesion could alleviate the negative effect of living alone (Honjo et al., 2018). This might happen because the gap of prevalence in depressive symptoms would be wider among different living arrangements than among different income groups, which could indicate the clear protective effect of community-level social cohesion.

Another literature review also suggests that a cohesive community environment might enforce a sense of solidarity among residents and serve as a barrier to discrimination and stigma toward socially vulnerable community members (i.e., ethnic minorities) (Pickett and Wilkinson, 2008). While the mental burden caused by loneliness (as a consequence of living alone) and discrimination could be directly buffered by community-level cognitive social capital (namely social cohesion and reciprocity), the financial burden (as reflected by a low-income status) might be affected more by the structural characteristics of the community, that is, the stronger physical and mental barriers to have new personal interactions and participation. To date, many studies have consistently shown that poorer individuals are less likely to participate in local activities (Mather, 1941; Marmot, 2002). As such, in our study, the multilevel models controlling for individual levels of community group participation suggest that community civic participation might actually increase the risk of depressive symptoms among the poor. That is, the benefit of community environments that promote group participation might not be effective for financially disadvantaged people in general. The possible mechanism of the deteriorating effect of community-level civic participation toward income-based inequality in depressive symptoms could be explained by both theoretical and empirical evidence. Social capital theorist Bourdieu conceptualizes that privileged people can easily exploit their capital of social connections (i.e., networks, by becoming members of groups) to reinforce their status (Bourdieu, 1984). Arneil and Offer also state that social pressure could exclude people of lower socioeconomic status, reflecting that the community context expressed by the high overall civic participation only benefits people in mainstream society, while the remaining people have few resources to develop opportunities (Arneil, 2006; Offer, 2012). In addition, Harpham et al. claim that group membership could damage mental health by imposing an extra burden or introducing stigma and peer pressure among vulnerable women (Harpham et al., 2006). Such a situation could produce negative feelings such as an inferiority complex in lower-income people, which could lead to the spread of depressive symptoms among them. The systematic review by Uphoff and Pickett, which reported that certain types of social capital might be beneficial only to those who had access to better health through their affluent assets, also supports our findings (Uphoff et al., 2013).

Although the overall trends in the findings were similar between men and women, as is indicated in another study (Vyncke et al., 2014), men showed potentially clearer results on the positive association between community civic participation and individual depressive symptoms and on the strengthening effects toward inequality in depressive symptoms. This might be because of a larger gap in the prevalence of depressive symptoms among income groups in men than women (Gero et al., 2017). Alternatively, mental stresses due to poverty may be stronger among older Japanese men. Saito et al. have reported that relative income deprivation increases the mortality risks more for men than for women among JAGES participants. Relative deprivation compared to others in their community might become a harmful psychosocial stressor for men (Saito et al., 2012).

This study has important policy implications. The Japanese government has officially recommended that local governments empower communities and community social capital to promote community health (Ministry of Health Labor and Welfare, 2015). Given the findings of this study, such efforts might decrease the prevalence of depressive symptoms in middle- or higher-income people. However, a conventional population approach that simply promotes opportunities for local group participation could widen the gap (Benach et al., 2013). Thus, continuous assessments of community intervention policies in terms of not only their overall effects but also their differential impacts on subpopulations should be conducted.

There are several limitations to this study. First, although this observational study suggests a higher prevalence of depressive symptoms among the poor in a community with rich civic participation, it cannot be inferred that actual interventions to promote community networking expand the socioeconomic gaps with regard to mental risks. For example, in the community intervention program in Taketoyo town, the municipal government strategically developed multiple voluntary social gathering programs called "*ikoi-no-saron* (recreation salon)," which halved the incidence of functional disabilities among its participants compared to non-participants. Importantly, the participants of the salon were primarily those whose incomes were low (Hikichi et al., 2015). Thus, a greater number of activities involving low-income people could alleviate inequality in depressive symptoms. Second, this cross-sectional study cannot account for temporal associations between community-level social capital and inequality in depressive symptoms. There can be two causal pathways between community social capital and depression. The former can affect not only individual mental health but also the depressed residents' decisions on moving into or out of the community, potentially causing sample selection issues. Further studies using longitudinal data are required. Third, although we used a small area (i.e., school district) as the unit of community, because it was the smallest identifiable unit, we do not have strong support for the validity of utilizing this unit in evaluating community social capital. Hence, the risk of the modifiable area unit problem (MAUP) should be considered (Mobley et al., 2007). However, using the school district is correct in terms of the scale that we used to assess social capital, as the unit was exactly the same for developing the Health-Related Social Capital Scale (Saito et al., 2017). Fourth, there is a substantial amount of missing data. Given the general tendency of people with low-income levels and mental illnesses to avoid responding, the prevalence of depressive symptoms may be underestimated, potentially causing an underestimation of the relationships between income and depressive symptoms, which could in turn result in biasing the modification effects of community factors toward null. Fifth, the generalizability of this study to Japan as whole or to other countries may be limited as the participating municipalities are not nationally representative, though JAGES covers a wide variety of municipalities.

## 5. Conclusion

This cross-sectional research suggests that despite the potential overall benefits to decrease depressive symptoms, high levels of community civic participation may increase the income-based inequality in depressive symptoms. Policies aiming to build healthy communities should consider the potential positive and negative effects of community social capital (Portes, 1998). Universal activities promoting opportunities for civic participation might be insufficient in terms of building an equitable community, and it potentially expands the disparity in mental health. Such universal promotion should be coupled with additional targeted interventions toward populations with high isolation risks (Marmot et al., 2010).

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## Appendix A. Supplementary material

Supplementary data associated with this article can be found in the online version at https://doi.org/10.1016/j.healthplace.2018.04.010.

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# Association between childhood socioeconomic status and fruit and vegetable intake among older Japanese: The JAGES 2010 study.

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## Abstract

Fruit and vegetable intake (FVI) contributes to the prevention of non-communicable diseases. Although food preference is considered to be determined early in life, few studies have investigated the association between childhood socioeconomic status (SES) and FVI in older age. Because a school lunch program was initiated in Japan after World War II, we were able in this study to examine this association in an older Japanese population. We used data from a population of physically and cognitively independent adults aged 65years or older who were living independently in the community and were recruited from 27 municipalities in the Japan Gerontological Evaluation Study 2010 project (August 2010-January 2012). Three categories of childhood SES (low, middle, and high) and current FVI were evaluated via a self-reported questionnaire. Poisson regression was used to investigate the association between childhood SES and FVI in 19,920 individuals. After adjustment for age and sex, older people with low childhood SES were 1.36 times more likely (95% CI 1.23-1.52) to have poor FVI than those with high childhood SES. In the fully adjusted model, the significant association disappeared. Further age-stratified analysis revealed a positive association between childhood SES and FVI among people aged 70-76years who were partially exposed to the school lunch program, but not among people aged 65-69years old who were fully exposed to the program. In conclusion, social policy such as school lunches targeting children with low SES could help improve FVI in old age.



**KEYWORDS:**

Childhood socioeconomic status; Fruit and vegetable intake; Life course; Older people; School lunch

OPEN

# Community-level Sports Group Participation and Older Individuals' Depressive Symptoms

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*[1]Center for Preventive Medical Sciences, Chiba University, Chiba, Chiba, JAPAN; [2]Department of Preventive Medicine and Public Health, Tokyo Medical University, Shinjuku, Tokyo, JAPAN; [3]Human Resource Management Department, ITOCHU Techno-Solutions Corporation, Shinagawa, Tokyo, JAPAN; and [4]Center for Gerontology and Social Science, National Center for Geriatrics and Gerontology, Obu, Aichi, JAPAN*

**ABSTRACT**

TSUJI, T., Y. MIYAGUNI, S. KANAMORI, M. HANAZATO, and K. KONDO. Community-level Sports Group Participation and Older Individuals' Depressive Symptoms. *Med. Sci. Sports Exerc.*, Vol. 50, No. 6, pp. 1199–1205, 2018. **Purpose**: Community-level group participation is a structural aspect of social capital that may have a contextual influence on an individual's health. Herein, we sought to investigate a contextual relationship between community-level prevalence of sports group participation and depressive symptoms in older individuals. **Methods**: We used data from the 2010 Japan Gerontological Evaluation Study, a population-based, cross-sectional study of individuals 65 yr or older without long-term care needs in Japan. Overall, 74,681 participants in 516 communities were analyzed. Depressive symptoms were diagnosed as a 15-item Geriatric Depression Scale score of ≥5. Participation in a sports group 1 d·month$^{-1}$ or more often was defined as "participation." For this study, we applied two-level multilevel Poisson regression analysis stratified by sex, calculated prevalence ratios (PR), and 95% confidence intervals (CI). **Results**: Overall, 17,420 individuals (23.3%) had depressive symptoms, and 16,915 (22.6%) participated in a sports group. Higher prevalence of community-level sports group participation had a statistically significant relationship with a lower likelihood of depressive symptoms (male: PR, 0.89 (95% CI, 0.85–0.92); female: PR, 0.96 (95% CI, 0.92–0.99), estimated by 10% of participation proportion) after adjusting for individual-level sports group participation, age, diseases, family form, alcohol, smoking, education, equivalent income, and population density. We found statistically significant cross-level interaction terms in male participants only (PR, 0.86; 95% CI, 0.77–0.95). **Conclusions**: We found a contextual preventive relationship between community-level sports group participation and depressive symptoms in older individuals. Therefore, promoting sports groups in a community may be effective as a population-based strategy for the prevention of depression in older individuals. Furthermore, the benefit may favor male sports group participants. **Key Words:** DEPRESSION, ELDERLY, MULTILEVEL ANALYSIS, SOCIAL CAPITAL, EXERCISE EPIDEMIOLOGY, JAPAN GERONTOLOGICAL EVALUATION STUDY



Depression in older adults is strongly associated with being house bound and isolated (1), which may lead to a decline in physical and cognitive function and eventually to premature death (2). Physical activity, which is a modifiable behavior, can prevent or alleviate depressive symptoms in older individuals (3–6). In particular, group participation in sports may have positive effects of physical activity and social participation on mental health (7–11), leading to enjoyment, enhanced self-esteem, and decreased stress (8). Increasing the frequency of sports group participation may alleviate the worsening depressive symptoms among older individuals who walk at an even rate compared with those who increase their daily walking time (11). Participation in a sports group may also lower the risk of functional disability compared with participation in other kinds of social activities (e.g., local community, volunteer, industry, or politics) (12). Therefore, growing evidence suggests that sports group participation may have preventive effects on psychological and physical functional decline in older individuals.

The definition of social participation mostly focuses on a "person's involvement in activities that provide interaction with others in society or the community" and is an index of

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social capital (13). There has been an increasing application of social capital in public health (14). According to Putnam (15), social capital refers to "features of social organization, such as trust, norms, and networks, that can improve the efficacy of society by facilitating coordinated actions" (p. 167). It is also described as "resources that are accessed by individuals as a result of their membership of a network or a group" (p. 291) (16). Social capital has two levels: individual level, which refers to resources accessed by the individual through their ego-centered networks, and group level, which pertains to a characteristic of the whole social network (16). Findings among studies investigating the relationship between group (community)-level social capital or social participation and older individuals' mental health were inconsistent. Although community-level bonding social capital (perceived homogeneous network) and bridging social capital (perceived heterogeneous network) do not necessarily improve mental health in older individuals (17), Saito and colleagues (18) revealed that a composite score of community-level social participation had a protective relationship for individual-level poor self-rated health and depressive symptoms in older individuals after adjusting for individual-level social participation. Inconsistent with this study, community-level formal group participation did not correlate with older depressed individuals' symptoms in another study with small samples (19). There is little evidence to support an association between community-level sports group participation and depressive symptoms in older individuals.

The present study investigated whether older individuals living in community areas with a higher prevalence of sports group participation among older individuals are less likely to have depressive symptoms compared with those living in community areas with a lower prevalence of such participation after controlling for individual-level sports group participation. We performed community- and individual-level multilevel analyses to clarify the contextual relationship between community-level sports group participation and depressive symptoms in older individuals.



## METHODS

**Study sample.** We used cross-sectional data from the Japan Gerontological Evaluation Study (JAGES). The JAGES project is an ongoing cohort study investigating social and behavioral factors related to the loss of health with respect to functional decline or cognitive impairment among individuals 65 yr or older (20). Between August 2010 and January 2012, a total of 169,215 community-dwelling people 65 yr or older were randomly selected from 31 municipalities including metropolitan, urban/semiurban, and rural communities in 12 prefectures from as far north as Hokkaido (i.e., the northernmost prefecture) and as far south as Okinawa (i.e., the southernmost prefecture) in Japan and were mailed a set of questionnaires. Overall, 112,123 people participated (response rate, 66.3%). We used data from 74,681 respondents (valid response rate, 44.1%) in 516 community areas, after excluding (i) 46 community areas with ≤30 respondents each (a total of 980 respondents), (ii) 4099 respondents whose areas of residence were unknown, and (iii) 32,363 respondents whose status of sex, age, depressive symptoms, or extent of sports group participation were unknown. One community area was essentially equivalent to one elementary or junior high school district. Generally, a school district represents a geographical scale in which the Japanese elderly can travel easily by foot or bicycle (21). Therefore, the present study used the community area as a proxy for the neighborhoods of people in our sample. JAGES participants were informed that participation in the study was voluntary and that completing and returning the questionnaire via mail indicated their consent to participate in the study. Ethical approval for the study was obtained from the Ethics Committee at Nihon Fukushi University, Japan (Approval No.: 10-05).

**Dependent variable.** We assessed depressive symptoms using the 15-item Geriatric Depression Scale (GDS) (22,23). Following previous research (22–24), mild or severe depressive symptoms (GDS ≥ 5) were set as an outcome of the present study. Cronbach's $\alpha$ for internal consistency of the scale was 0.80 (24), and the cutoff point was previously validated as a screening instrument for major depressive disorder with 96% sensitivity and 95% specificity (24).

**Community- and individual-level independent variables.** Participants were queried on their frequency of sports group participation. Possible answers were ≥4 $d{\cdot}wk^{-1}$, 2–3 $d{\cdot}wk^{-1}$, 1 $d{\cdot}wk^{-1}$, 1–3 $d{\cdot}month^{-1}$, a few times a year, or zero. We defined participating 1 $d{\cdot}month^{-1}$ or more often as "participation" in a sports group (18,25) and aggregated individual-level sports group participation by community area as a community-level independent variable. Previous research (18) indicates that the correlation of the proportion of older individuals with depressive symptoms (GDS ≥ 5) in areas with community-level sports group participation 1 $d{\cdot}month^{-1}$ or more often ($r = -0.355$) tended to be strong compared with that with community-level sports group participation 1 $d{\cdot}wk^{-1}$ or more often ($r = -0.314$).

**Covariates.** Sex was controlled by conducting a stratified analysis. Age groups were categorized as 65–69, 70–74, 75–79, 80–84, and ≥85 yr. To gather information on disease status in treatment, participants were asked if they were currently receiving any medical treatment; answer choices were yes and no. Participants were asked about household members living with and were categorized as living with others or living alone. Drinking status (none, past, or current), smoking status (none, past, or current), and education (<10, 10–12, or ≥13 yr) were classified by each answer choice. Annual equivalent income was calculated by dividing household income by the square root of the number of household members and categorized into three groups: <2,000,000; 2,000,000–3,999,999; or ≥4,000,000 yen. If participants did not respond to the individual-level covariates, corresponding observations were assigned to "missing" categories. As a community-level covariate, we calculated population density per square kilometer of inhabitable area for each community area and categorized into quartile categories (<1800, 1800–6999, 7000–9999, and ≥10,000 persons per kilometer squared).

**Statistical analysis.** Descriptive statistics were used to characterize the participants and community areas. In the data set, 35,975 men and 38,706 women were nested in 516 community areas (community level). Having depressive symptoms (GDS ≥ 5) was the outcome variable, and after adjusting for individual- and community-level covariates, the effect of community-level sports groups was inferred. To examine the contextual relationship of community-level prevalence of sports group participation to individual-level depressive symptoms, we applied two-level multilevel Poisson regression analysis (the individual as level 1 and the community as level 2) with random intercepts and fixed slopes and calculated the multilevel prevalence ratio (PR) and 95% confidence interval (CI). Because percentages of individuals with depressive symptoms (23.3%) were >10%, adjusted odds ratio derived from logistic regression can no longer approximate the PR (26). Two models of analysis were used. Both community- and individual-level sports group participation and cross-level interaction terms were included in model 1. In model 2, all individual- and community-level covariates were added. The PR and 95% CI of community-level sports group participation were estimated by 10% of participation proportion. We used STATA 13/SE (StataCorp, College Station, TX) for all statistical analyses.

## RESULTS

Of 74,681 analytic samples (mean age ± SD, 73.4 ± 6.0 yr in male and 73.8 ± 6.3 yr in female participants), 17,420 (23.3%) had depressive symptoms and 16,915 (22.6%) participated in a sports group 1 d·month$^{-1}$ or more often. When the proportions of depressive symptoms and sports group participation were calculated for each community area, the ranges were 0.0%–60.6% and 0.0%–56.5%, respectively. Tables 1 and 2 show the descriptive statistics and crude PR (95% CI) for having depressive symptoms for the individual- and community-level variables by sex, respectively. Participants of both sexes with older age who had disease(s) under treatment, lived alone, were past drinkers, or were past/current smokers were more likely to have depressive symptoms. Participants of both sexes who were current drinkers, with higher educational levels, and higher equivalent incomes were less likely to have depressive symptoms. Compared with male and female participants who lived in community areas with first (lowest) quartile of population density, those who lived in areas with the second and fourth (highest) quartiles as well as female participants who lived in areas with the third quartile were less likely to have depressive symptoms. Mean proportions of sports group participants by population density categories were 18.9%, 25.6%, 26.2%, and 26.0% from the first, second, third, and fourth quartiles of population density, respectively.

Table 3 shows the results of the multilevel Poisson regression analyses. Regardless of whether the model included covariates and sex, community-level higher prevalence of sports group participation had a statistically significant relationship with lower likelihood of depressive symptoms (male: PR, 0.89 (95% CI, 0.85–0.92); female, PR, 0.96 (95% CI, 0.92–0.99)) in the fully adjusted model estimated by 10% of participation proportion. Individual-level sports group participation also had a significant relationship with low likelihood of depressive symptoms (male: PR, 0.56 (95% CI, 0.52–0.60); female: PR, 0.58 (95% CI, 0.55–0.62), in the fully adjusted model). We found statistically significant cross-level interaction terms in male participants only (PR, 0.86 (95% CI, 0.77–0.95), in the fully adjusted model).

## DISCUSSION

To the best of our knowledge, this is the first study to find the contextual relationship between community-level prevalence of sports group participation in older individuals and depressive symptoms in older individuals. A 10% increase in community-level sports group participation was associated with an 11% and 4% reduction in the prevalence of depressive symptoms after adjusting for individual-level sports group participation and covariates. Not all older individuals are able to participate in a sports group because of factors such as work, lower socioeconomic status (i.e., lower income and lower educational attainment), or less social support (27). Furthermore, not all of those individuals are physically active because of demographic and biological, psychosocial, behavioral, social and cultural, and environmental factors (28). The results of the present study suggested that promoting sports groups in a community may be effective for the prevention of depression not only in sports group participants but also in nonparticipants having those barriers.

The individual-level relationship between sports group participation and mental health in adults has been frequently discussed in previous reports (7–10). A systematic review indicated consistent evidence that club- and team-based sports participation, when compared with individual forms of physical activity, was associated with improved psychological and social health (7). On the basis of a 6-yr longitudinal study using nationally representative data in Japan (10), adults 50–59 yr old who participated in exercise or sports activities were less likely to have worse mental health compared with those who did not participate. In the current study, we also found an individual-level preventive relationship of sports group participation with depressive symptoms in older individuals. The reciprocals of the PR for participating in a sports group were 1.79 (1/0.56) in male and 1.72 (1/0.58) in women. Male sports group participation may mitigate the adverse effects of living alone on depressive symptoms (PR, 1.70). In women, sports group participation may help in overcoming the adverse effects of disease status

TABLE 1. Characteristics of respondents and their associated PR for having depressive symptoms.

| | Male | | | | | Female | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | | Prevalence of Depressive Symptoms | | | | | Prevalence of Depressive Symptoms | | | |
| | Total *n* | *n* | Pct. | PR | 95% CI | Total *n* | *n* | Pct. | PR | 95% CI |
| Total | 35,975 | 8327 | 23.1 | | | 38,706 | 9093 | 23.5 | | |
| Participating in a sports group | | | | | | | | | | |
| Zero | 25,790 | 7017 | 27.2 | 1.00 | | 28,567 | 7694 | 26.9 | 1.00 | |
| A few times a year | 2419 | 359 | 14.8 | 0.55 | (0.49–0.61) | 990 | 155 | 15.7 | 0.58 | (0.50–0.68) |
| 1–3 d·month$^{-1}$ | 2078 | 258 | 12.4 | 0.46 | (0.40–0.52) | 1392 | 216 | 15.5 | 0.58 | (0.50–0.66) |
| 1 d·wk$^{-1}$ | 1947 | 256 | 13.1 | 0.48 | (0.43–0.55) | 3320 | 466 | 14.0 | 0.52 | (0.47–0.57) |
| 2–3 d·wk$^{-1}$ | 2815 | 344 | 12.2 | 0.45 | (0.40–0.50) | 3622 | 470 | 13.0 | 0.48 | (0.44–0.53) |
| ≥4 d·wk$^{-1}$ | 926 | 93 | 10.0 | 0.37 | (0.30–0.45) | 815 | 92 | 11.3 | 0.42 | (0.34–0.51) |
| Participating in a sports group, 1 d·month$^{-1}$ or more often | | | | | | | | | | |
| No | 28,209 | 7376 | 26.1 | 1.00 | | 29,557 | 7849 | 26.6 | 1.00 | |
| Yes | 7766 | 951 | 12.2 | 0.47 | (0.44–0.50) | 9149 | 1244 | 13.6 | 0.51 | (0.48–0.54) |
| Age, yr | | | | | | | | | | |
| 65–69 | 11,409 | 2354 | 20.6 | 1.00 | | 11,923 | 2427 | 20.4 | 1.00 | |
| 70–74 | 10,658 | 2304 | 21.6 | 1.05 | (0.99–1.11) | 11,271 | 2399 | 21.3 | 1.05 | (0.99–1.11) |
| 75–79 | 7771 | 1919 | 24.7 | 1.20 | (1.13–1.27) | 8170 | 2032 | 24.9 | 1.22 | (1.15–1.30) |
| 80–84 | 4297 | 1190 | 27.7 | 1.34 | (1.25–1.44) | 4719 | 1323 | 28.0 | 1.38 | (1.29–1.47) |
| ≥85 | 1840 | 560 | 30.4 | 1.48 | (1.35–1.62) | 2623 | 912 | 34.8 | 1.71 | (1.58–1.84) |
| Disease status in treatment | | | | | | | | | | |
| No | 8326 | 1349 | 16.2 | 1.00 | | 8094 | 1247 | 15.4 | 1.00 | |
| Yes | 27,134 | 6865 | 25.3 | 1.56 | (1.47–1.66) | 29,920 | 7678 | 25.7 | 1.67 | (1.57–1.77) |
| Missing | 515 | 113 | 21.9 | 1.35 | (1.12–1.64) | 692 | 168 | 24.3 | 1.58 | (1.34–1.85) |
| Living with others | | | | | | | | | | |
| Yes | 32,914 | 7140 | 21.7 | 1.00 | | 31,821 | 7101 | 22.3 | 1.00 | |
| No (living alone) | 2461 | 1008 | 41.0 | 1.89 | (1.77–2.02) | 6157 | 1754 | 28.5 | 1.28 | (1.21–1.35) |
| Missing | 600 | 179 | 29.8 | 1.38 | (1.19–1.60) | 728 | 238 | 32.7 | 1.47 | (1.29–1.67) |
| Drinking status | | | | | | | | | | |
| None | 12,199 | 3213 | 26.3 | 1.00 | | 30,445 | 7360 | 24.2 | 1.00 | |
| Past | 2111 | 692 | 32.8 | 1.24 | (1.15–1.35) | 370 | 144 | 38.9 | 1.61 | (1.37–1.90) |
| Current | 19,806 | 3961 | 20.0 | 0.76 | (0.72–0.80) | 5991 | 1110 | 18.5 | 0.77 | (0.72–0.82) |
| Missing | 1859 | 461 | 24.8 | 0.94 | (0.85–1.04) | 1900 | 479 | 25.2 | 1.04 | (0.95–1.14) |
| Smoking status | | | | | | | | | | |
| None | 8625 | 1730 | 20.1 | 1.00 | | 31,856 | 7060 | 22.2 | 1.00 | |
| Past | 18,405 | 4182 | 22.7 | 1.13 | (1.07–1.20) | 1777 | 551 | 31.0 | 1.40 | (1.28–1.53) |
| Current | 6664 | 1840 | 27.6 | 1.38 | (1.29–1.47) | 1124 | 399 | 35.5 | 1.60 | (1.45–1.77) |
| Missing | 2281 | 575 | 25.2 | 1.26 | (1.14–1.38) | 3949 | 1083 | 27.4 | 1.24 | (1.16–1.32) |
| Education, yr | | | | | | | | | | |
| <10 | 14,713 | 4157 | 28.3 | 1.00 | | 18,158 | 4925 | 27.1 | 1.00 | |
| 10–12 | 12,135 | 2481 | 20.4 | 0.72 | (0.69–0.76) | 14,294 | 2951 | 20.6 | 0.76 | (0.73–0.80) |
| ≥13 | 8533 | 1491 | 17.5 | 0.62 | (0.58–0.66) | 5353 | 930 | 17.4 | 0.64 | (0.60–0.69) |
| Missing | 594 | 198 | 33.3 | 1.18 | (1.02–1.36) | 901 | 287 | 31.9 | 1.17 | (1.04–1.32) |
| Annual equivalent income, yen | | | | | | | | | | |
| <2,000,000 | 14,633 | 4348 | 29.7 | 1.00 | | 15,758 | 4482 | 28.4 | 1.00 | |
| 2,000,000–3,999,999 | 13,589 | 2322 | 17.1 | 0.58 | (0.55–0.60) | 11,845 | 1997 | 16.9 | 0.59 | (0.56–0.62) |
| ≥4,000,000 | 3854 | 430 | 11.2 | 0.38 | (0.34–0.41) | 3535 | 457 | 12.9 | 0.45 | (0.41–0.50) |
| Missing | 3899 | 1227 | 31.5 | 1.06 | (0.99–1.13) | 7568 | 2157 | 28.5 | 1.00 | (0.95–1.05) |

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in treatment (PR, 1.52) and smoking (PR, 1.53). Furthermore, it is worth noting that the mitigational relationship of 10% increases in community-level sports group participation, which were estimated by the reciprocals of the PR (1.12 in men and 1.04 in women), was comparable with the age-related differences by 15 yr in men (PR, 1.12) and 10 yr in women (PR, 1.04).

Among fundamental dimensions of social capital, the previous report suggested that community-level social participation

TABLE 2. Characteristics of community areas and their associated prevalence.

| *n* = 516 | | Male PR | Male 95% CI | Female PR | Female 95% CI |
|---|---|---|---|---|---|
| Proportion of depressive symptoms | | | | | |
| Mean (SD) | 22.6% (6.9%) | — | | — | |
| (Min–Max) | (0.0%–60.6%) | | | | |
| Proportion of sports group participants[a] | | | | | |
| Mean (SD) | 24.3% (8.2%) | 0.85 | (0.82–0.88) | 0.89 | (0.86–0.92) |
| (Min–Max) | (0.0%–56.5%) | | | | |
| Population density, persons per kilometer squared of inhabitable area | | | | | |
| <1800, *n* | 129 | 1.00 | | 1.00 | |
| 1800–6999, *n* | 129 | 0.92 | (0.86–0.98) | 0.91 | (0.86–0.96) |
| 7000–9999, *n* | 129 | 0.93 | (0.85–1.01) | 0.84 | (0.77–0.91) |
| ≥10,000, *n* | 129 | 0.91 | (0.84–0.99) | 0.86 | (0.80–0.93) |

[a]PR and 95% CI were calculated by 10% estimation.

TABLE 3. Association of depressive symptoms with community- and individual-level variables determined by multilevel Poisson regression.

| | Male ($n$ = 35,975) | | | | Female ($n$ = 38,706) | | | |
|---|---|---|---|---|---|---|---|---|
| | Model 1 | | Model 2 | | Model 1 | | Model 2 | |
| | PR | 95% CI | PR | 95% CI | PR | 95% CI | PR | 95% CI |
| Fixed effects | | | | | | | | |
| Community-level independent variable | | | | | | | | |
| Proportion of sports group participants | | | | | | | | |
| 10% estimation | 0.82 | (0.79–0.85) | 0.89 | (0.85–0.92) | 0.89 | (0.86–0.92) | 0.96 | (0.92–0.99) |
| Individual-level independent variable | | | | | | | | |
| Participating in a sports group (1 d·month$^{-1}$ or more often) | | | | | | | | |
| No | 1.00 | | 1.00 | | 1.00 | | 1.00 | |
| Yes | 0.49 | (0.46–0.52) | 0.56 | (0.52–0.60) | 0.52 | (0.49–0.55) | 0.58 | (0.55–0.62) |
| Cross-level interaction | | | | | | | | |
| Community-level exposure × individual-level exposure | | | | | | | | |
| 10% estimation | 0.86 | (0.77–0.96) | 0.86 | (0.77–0.95) | 1.03 | (0.93–1.13) | 1.04 | (0.94–1.14) |
| Community-level covariate | | | | | | | | |
| Population density, persons per kilometer squared of inhabitable area | | | | | | | | |
| <1800 | | | 1.00 | | | | 1.00 | |
| 1800–6999 | | | 1.01 | (0.96–1.06) | | | 1.00 | (0.95–1.05) |
| 7000–9999 | | | 1.04 | (0.95–1.12) | | | 0.94 | (0.87–1.03) |
| ≥10,000 | | | 1.00 | (0.92–1.08) | | | 0.95 | (0.88–1.03) |
| Individual-level covariates | | | | | | | | |
| Age, yr | | | | | | | | |
| 65–69 | | | 1.00 | | | | 1.00 | |
| 70–74 | | | 0.99 | (0.94–1.05) | | | 0.96 | (0.91–1.02) |
| 75–79 | | | 1.06 | (0.99–1.12) | | | 1.04 | (0.98–1.11) |
| 80–84 | | | 1.12 | (1.04–1.20) | | | 1.11 | (1.04–1.19) |
| ≥85 | | | 1.17 | (1.06–1.28) | | | 1.34 | (1.24–1.45) |
| Disease status in treatment | | | | | | | | |
| No | | | 1.00 | | | | 1.00 | |
| Yes | | | 1.51 | (1.42–1.60) | | | 1.52 | (1.43–1.61) |
| Missing | | | 1.18 | (0.98–1.43) | | | 1.35 | (1.15–1.59) |
| Living with others | | | | | | | | |
| Yes | | | 1.00 | | | | 1.00 | |
| No (living alone) | | | 1.70 | (1.59–1.82) | | | 1.14 | (1.08–1.21) |
| Missing | | | 1.14 | (0.98–1.33) | | | 1.21 | (1.06–1.38) |
| Drinking status | | | | | | | | |
| None | | | 1.00 | | | | 1.00 | |
| Past | | | 1.19 | (1.09–1.29) | | | 1.41 | (1.20–1.67) |
| Current | | | 0.85 | (0.81–0.89) | | | 0.86 | (0.81–0.92) |
| Missing | | | 1.01 | (0.89–1.16) | | | 0.98 | (0.88–1.09) |
| Smoking status | | | | | | | | |
| None | | | 1.00 | | | | 1.00 | |
| Past | | | 1.16 | (1.09–1.22) | | | 1.35 | (1.23–1.47) |
| Current | | | 1.37 | (1.29–1.47) | | | 1.53 | (1.38–1.70) |
| Missing | | | 1.12 | (0.99–1.27) | | | 1.13 | (1.05–1.22) |
| Education, yr | | | | | | | | |
| <10 | | | 1.00 | | | | 1.00 | |
| 10–12 | | | 0.86 | (0.82–0.90) | | | 0.89 | (0.85–0.93) |
| ≥13 | | | 0.80 | (0.75–0.85) | | | 0.80 | (0.74–0.86) |
| Missing | | | 1.13 | (0.98–1.30) | | | 1.07 | (0.95–1.21) |
| Annual equivalent income, yen | | | | | | | | |
| <2,000,000 | | | 1.00 | | | | 1.00 | |
| 2,000,000–3,999,999 | | | 0.64 | (0.60–0.67) | | | 0.66 | (0.63–0.70) |
| ≥4,000,000 | | | 0.44 | (0.40–0.48) | | | 0.51 | (0.46–0.56) |
| Missing | | | 0.96 | (0.90–1.02) | | | 0.95 | (0.90–1.00) |
| Intercept (SE) | 0.220 | 0.003 | 0.185 | 0.008 | 0.227 | 0.003 | 0.189 | 0.007 |
| Random effects | | | | | | | | |
| Community-level variance | | | | | | | | |
| $\Omega\mu$ (SE) | 0.0040 | 0.0027 | 0.0003 | 0.0020 | 0.0022 | 0.0025 | 0.0014 | 0.0022 |

was more closely related to older individuals' depressive symptoms than was social cohesion and reciprocity (18). In that study, community-level social participation was evaluated by merging sports group, volunteer group, and hobby activity participation. Those groups with egalitarian relationships could be categorized as horizontal social capital and had a stronger contextual association with individuals' health status than were groups with hierarchical relationships (i.e., vertical social capital) (29,30). Even when extracting sports group participation in the present study, we found that the protective relationship with mental health remained.

Social contagion, informal social control, and collective efficacy are regarded as pathways from group-level social capital to individual-level health outcomes (16). Social contagion references the notion that behaviors spread more quickly through a tightly knit social network (16). Sometimes the behavior that spreads via the network can promote healthy lifestyle changes, that is, the spread of smoking cessation (31). Informal social control refers to the ability of adults in a community to maintain social order, that is, to step in and intervene when they witness deviant behavior (16). In relation to the present study, some older individuals may be encouraged by sports group participants to get more exercise or take up a sport irrespective of whether it is an individual or a team-based sport. Collective efficacy is the group-level analog of the concept of self-efficacy and refers to the ability of the collective to mobilize to undertake collective action (16,32). Facilities, systems, bylaws for health promotion, or sports business and management may develop to reflect the opinions and actions of communities with many sports groups and participants. Group-level mechanisms of widespread sports group participation may result in positive spillover effects.

In the present study, a statistically significant cross-level interaction term was observed in men only, suggesting that male sports group participants living in community areas with a higher prevalence of sports group participation might be less likely to have depressive symptoms compared with those living in community areas with a lower prevalence of sports group participation. One of the possible benefits in group sports is the opportunity to play key roles in those groups. Takagi and colleagues (21) reported that Japanese older men who held group leadership positions reported lower depressive symptoms than those who participated but did not have leadership roles. However, this interaction effect did not apply to women. Because Japanese society is characterized by strong patriarchal values, men seek meaning and identity by being valued in the workplace (21). Male retirees may feel rewarded by seeking positions of authority or responsibility within the social organizations in which they participate. It is assumed that a community with many sports group participants naturally also has many sports groups. Accordingly, leadership positions for managing such groups are necessary. Therefore, older male participants who live in an active area with sports groups may have opportunities to fill these key roles.

The strength of the present study is its large, nationwide, population-based sample enabling sex-stratified community- and individual-level multilevel analysis for clarifying the contextual relationship of sports group participation. However, the study has several limitations. First, reverse causality could occur because of the nature of the cross-sectional design, and further longitudinal studies are needed to resolve this limitation. Second, the response rate and the valid response rate were 66.3% and 44.1%, respectively; therefore, selection bias may have affected the results. Several missing values may also bring systematic bias. In our previous study, we reported that both response rates and percentage of sports group participants with lower incomes were significantly lower than those in participants with higher incomes (33). Furthermore, respondents with unknown status of depressive symptoms or extent of sports group participation reported lower equivalent income compared with valid respondents in our survey (data not shown). The present study showed that respondents with lower income tended to have depressive symptoms. Therefore, nonrespondents or respondents who were excluded from the analysis might have worse depressive symptoms and be less likely to participate in sports groups compared with valid respondents. Although the results showed a significant relationship of sports group participation with depressive symptoms, this slightly low valid response rate might attenuate the relationship. Third, we could not discuss the preferable frequency of sports group participation because we set the cutoff point for participation/nonparticipation to only one time point: 1 d·month$^{-1}$ or more often/less than 1 d·month$^{-1}$. Although we attempted an analysis in which the cutoff was set to 1 d·wk$^{-1}$ or more often/less than 1 d·wk$^{-1}$, the models did not converge. However, we believe "at least 1 d·month$^{-1}$" to be an adequate and feasible frequency for both individuals who are not familiar with sports and those who want to establish a leadership position and manage a sports group in their community.

## CONCLUSIONS

Older individuals living in community areas with a higher prevalence of sports group participation in older individuals are less likely to have depressive symptoms compared with those living in a community area with lower prevalence of participation after adjusting for individual-level sports group participation and other covariates; that is, we found a contextual preventive relationship between community-level sports group participation and depressive symptoms in older individuals. Furthermore, the benefit may favor male sports group participants. Promoting sports groups in a community may be effective as a population-based strategy for the preventing depression in older individuals regardless of each individual's participation status.

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None of the authors have a conflict of interest in relation to this manuscript.

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# Socioeconomic Disparity in the Prevalence of Objectively Evaluated Diabetes Among Older Japanese Adults: JAGES Cross-Sectional Data in 2010

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## ABSTRACT

**Background:** Studies on sex-specific socioeconomic gradients in objectively evaluated diabetes among older adults are scarce.

**Methods:** We used cross-sectional data of 9,893 adults aged 65 years and older in Aichi Prefecture without long-term care insurance from the Japan Gerontological Evaluation Study (JAGES) in 2010 (Response rate: 66.3%). We collected demographic, socioeconomic (income, years of education, and longest occupation) and behavioral information using a mail-in self-reported survey. Blood samples for the objectively evaluated diabetes and self-reported medical history were collected at annual municipal health checkups. Poisson regression analysis stratified by sex with multiple imputations was conducted to calculate prevalence ratio and 95% confidence interval.

**Results:** A clear income gradient in diabetes prevalence was observed among women, from 11.7% in the lowest income quartile (Q1) to 7.8% in the highest (Q4). Among men, the findings were 17.6% in Q1 to 15.1% in Q4. The prevalence ratios for diabetes with incomes Q1 to Q4 were 1.43 (95% confidence interval [CI], 1.07–1.90) for women and 1.16 (95% CI, 0.90–1.50) for men after adjusting for age and other socioeconomic factors. Even after adjusting for marital status, body mass index, other metabolic risk factors, and lifestyle factors, the income-based gradient remained among women. Education and occupation were not significantly associated with diabetes in the study population.

**Conclusions:** Only women showed an income-based gradient in diabetes. Monitoring income gradient in diabetes is important in public health actions, even in older populations. Future longitudinal and intervention studies should evaluate the causal link of income to diabetes onset, determine the mechanisms of the potential sex differences in the income/diabetes association, and identify ways to mitigate the income-based inequality.

**Key words:** socioeconomic status; diabetes mellitus; sex differences; elderly adults; Japan



## INTRODUCTION

According to global reports on diabetes 2016, about 422 million adults were living with diabetes worldwide in 2014. During the past 4 decades, the global prevalence of adult diabetes has nearly doubled, increasing from 4.7% to 8.5%.[1] In Japan, the National Health and Nutrition Survey reported that 16.2% of men and 9.2% of women aged over 20 years were suspected to have diabetes in 2015.[1] About 3.2 million people (1.8 million men and 1.4 million women) received treatment for diabetes in 2014,[2] and 70% were over 65 years old.

Socioeconomic disparities in diabetes prevalence and incidence have been well documented in Western countries[3–11] and some Asian countries, including South Korea,[12] China,[13,14] Taiwan,[15] and Japan.[16,17] Except for one study in China,[14] inverse relationships between socioeconomic status (SES) and diabetes prevalence or incidence have been observed according to occupational class,[3,18,19] income level,[10,18,20] and educational attainment.[3,10,18–22] However, few studies have investigated the social gradient in diabetes among older adults, and the findings of studies using data among older populations have been inconsistent with respect to the association between SES and

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diabetes. For example, a Finnish study has shown that the low-income group had a higher diabetes prevalence compared with the high-income group among women, whereas among men the high-income group was more likely to have diabetes.[6] Studies in Germany[23] and the United Kingdom[5] have shown that neither income, education, nor occupation was associated with the incidence of diabetes.

To date, many studies have suggested sex differences in the association between SES and diabetes. Most studies have reported more obvious social gradients in diabetes among women than men in Western countries.[3,6,18–20,22] In Asian countries, only Lee et al in South Korea investigated sex differences in the association between SES and diabetes, similarly showing a stronger SES-diabetes association among women than men.[21] No evidence of sex differences in the association between SES and diabetes has been reported from other parts of Asia, including Japan.

These studies, other than the study in the United Kingdom ($N = 7{,}432$), are not large (Finland: $N = 379$, Germany $N = 1{,}223$), which may be a limitation in detecting the between-group gaps. Moreover, in the recent studies, the definitions of diabetes vary. For example, diabetes was defined using self-reports,[6] first diabetes medication prescribed,[5] and oral glucose tolerance tests.[23] Specifically, self-report of having diabetes could induce reporting bias. The validity study by Goto et al found that positive predictive value of self-report diabetes was 75.7%, whereas negative predictive value was 96.5% in the Japanese population.[24] The bias may go toward null on the association between SES and diabetes prevalence, given that health-conscious people with high health literacy recognize their health status more accurately. According to a Japanese nationally representative survey, health-conscious people are likely to be more educated.[25] To our knowledge, no studies have investigated the association between objectively diagnosed diabetes and SES among the older population.

Therefore, the purpose of this study was to investigate (1) whether there is an association between SES and diabetes prevalence among Japanese older adults, and (2) whether there is a sex difference in this association, using large-scale cross-sectional data with objectively measured biomarkers of diabetes.

## METHODS

### Study participants

We used cross-sectional data of the 2010 wave of the Japan Gerontological Evaluation Study (JAGES). In 2010, in JAGES we sent the questionnaires to 169,215 community-dwelling individuals over 65 years without long-term care insurance. From 31 municipalities in 12 out of 47 prefectures throughout Japan, participants were randomly selected from the public residence registries in 15 large municipalities; in the 16 smaller municipalities, all eligible residents got the mail-in survey. In total 112,123 subjects answered the questionnaire (response rate: 66.3%). After excluding the individuals with missing in demographic characteristics, 102,869 subjects were valid for analysis. Since JAGES exclude long-term care insurance takers from the study participants, we cannot compare the characteristics of the participants with the national census directly. However, the sex ratio of the total JAGES 2010 individuals was mostly comparable to that of the national census (national census: 42.6% men, 57.4% women; JAGES2010 data: 45.9% men, 54.1% women). JAGES female population was younger than that of the national census (national census: 32.2% age group ≥80; JAGES2010: 21.5% age group ≥80), whereas the age structure of the male population was mostly identical.[26]

In addition to these data, we obtained data of 9,893 JAGES participants with results of annual health checkups from five municipalities in Aichi Prefecture that participated in JAGES. After excluding participants with data missing for HbA1c, fasting blood glucose, casual blood glucose, or information of medication ($N = 306$) or SES variables (income, education and longest occupation) ($N = 2{,}774$), a total 6,813 (3,475 men and 3,338 women) participants were eligible for the analysis. We applied multiple imputation methods for the individuals having one or more missing data. Thus, the final study sample was 9,893 (4,471 men and 5,422 women). Approvals were received by the Ethics Committee in Research of Human Subjects at Nihon Fukushi University for the JAGES protocol (No. 10-05) and by the Ethics Committee of Chiba University, Faculty of Medicine for the use of the data (No. 1777).

### Measurement of diabetes and other metabolic risk factors

Annual health checkups are organized by local municipalities of Japan and performed at community centers or registered clinics or hospitals. Participant blood samples taken at annual checkups were analyzed following the standardized procedure of the Japan Society of Clinical Chemistry for HbA1c, fasting glucose, casual glucose, triglycerides (TG), and high-density lipoprotein (HDL) cholesterol. The HbA1c ratios were reported as the Japan Diabetes Society (JDS) values, then calculated for the values of the National Glycohemoglobin Standardization Program (NGSP) following a conversion formula.[27] Blood pressure was measured twice in the right upper arm with participants in the sitting position, and the mean of the two measurements was recorded.

### Definition of diabetes and other metabolic risk factors

Based on the report from the Committee of the JDS on the Diagnostic Criteria of Diabetes Mellitus,[28] we defined diabetes mellitus as having HbA1c of over 6.5% based on the NGSP and fasting blood sugar ≥126 mg/dL (≥7.0 mmol/L) and/or casual blood sugar ≥200 mg/dL (≥11.1 mmol/L). People regularly taking hypoglycemic agents or insulin were also considered to have diabetes. We used the following criteria of the Japanese Society of Internal Medicine[29] to define other metabolic risks: hypertriglyceridemia (TG ≥150 mg/dL), low HDL cholesterol (HDL <40 mg/dL), or taking appropriate medication and hypertension (systolic blood pressure ≥140 mm Hg and/or diastolic blood pressure ≥90 mm Hg or taking antihypertensive drugs).

### Socioeconomic status

Information of participants' annual household income, educational attainment, and longest occupation were collected using JAGES questionnaires. To adjust for differences in household size, we equivalized household income to per person in a household, dividing annual household income by the square root of the number of individuals per household. Because the income level of the participants was slightly higher than that of the entire JAGES population in 2010 (eTable 1), to apply the income level of the entire JAGES population, we categorized the study participants using quartiles of equivalized household income of all JAGES 2010 participants. We categorized the individuals into four groups: Q1 (low), JPY 1.25 million per year and below; Q2 (lower middle), JPY 1.251–1.944 million per year; Q3 (upper middle), JPY 1.945–3.061 million per year and Q4 (high), JPY 3.062 million per year and above. Educational status was categorized by the number of years of schooling (9 years or fewer, 10–12 years, and 13 years or more). Longest occupation was queried as follows: "What was the job that you did for most of your working life?" Responses included the following eight options: professional/technical, managerial, clerical, sales/service, skilled/manual, agriculture/forestry/fishery workers, other, and unemployed.[30]

## Covariates

We used categorical 5-year age groups, marital status, body mass index (BMI), other metabolic risks defined above, current smoking, current alcohol intake, physical activities, and dietary intake habits as covariates; these factors could be mediators of the association between SES and diabetes. Marital status was categorized as married, widowed, separated/unmarried, and other. BMI was classified into four groups: BMI <18.5, 18.5–24.9, 25.0–29.9, and ≥30.0. Other lifestyle factors included smoking status (nonsmoker or current smoker/ex-smoker), alcohol intake (nondrinker or drinker/ex-drinker), and walking time per day as physical activity (<30 min or ≥30 min). As for dietary intake habits, we included frequencies of the consumption of meat or fish (<1 servings/day or ≥1 servings/day) and fruit or vegetables (<1 servings/day or ≥1 servings/day).

## Statistical analysis

First, we calculated the prevalence of diabetes by the levels of socioeconomic indicators. Chi-squared test for sex was performed both in Table 1 and Table 2. Second, we performed multivariate Poisson regression analysis to calculate prevalence ratios (PRs) of diabetes and their 95% confidence intervals (CIs) across SES groups. To account for the potential biases due to the missing values, we used the multiple imputation techniques. All variables included in the analysis, such as the outcome variable, diabetes, explanatory variables, and covariates were imputed. Under a missing-at-random assumption, we created 10 imputed data using chained equation method, made analyses for each dataset, and combined the 10 results, using Rubin's combination method.[31,32] Under the chained equation method, we performed multinomial logistic regression for the categorical variables and ordinal logistic regression for the ordinal variables. We treated occupation and marital status as nominal variables and categorized diabetes, income, education, BMI, hypertension, triglyceridemia, smoking habit, alcohol intake, walking duration per day, and eating habit as ordinal variables, including dichotomous variables. Model 1 was adjusted for age and each SES indicators (income quartile, years of education, and longest occupation) separately. Model 2 was adjusted for age and all SES indicators. Model 3 was additionally adjusted for marital status, BMI, hypertension, low HDL, high TG, smoking status, alcohol intake, walking time per day, meat/fish intake, and fruit/vegetable intake.

Preliminary analysis showed that the interaction terms for sex and socioeconomic indicators were not statistically significant ($P$-value for the interaction term between income and sex = 0.18, $P$-value for the interaction term between education and sex = 0.20). However, as Hawkes et al mentioned,[33] irrespective of the statistical significance, the gender-stratified analysis is essential to address the determinants of ill health by gender. Accordingly, we decided to analyze the data stratified by sex. Also, to investigate the validity of our missing-at-random assumption for multiple imputations, we conducted a sensitivity analysis using the complete case dataset (eTable 2). We used Stata/SE version 13.1 (StataCorp LLC, College Station, TX, USA) for the analyses.

**Table 1.** Participant characteristics ($N = 6{,}813$)

| | | Men | Women | |
|---|---|---|---|---|
| | | ($N = 3{,}475$) | ($N = 3{,}338$) | $P$-value[a] |
| Diabetes | No | 2,948 (84.8%) | 2,996 (89.8%) | <0.001 |
| | Yes | 527 (15.2%) | 342 (10.2%) | |
| Age, years | 65–69 | 1,297 (37.3%) | 1,335 (40.0%) | 0.13 |
| | 70–74 | 1,213 (34.9%) | 1,093 (32.7%) | |
| | 75–79 | 606 (17.4%) | 572 (17.1%) | |
| | 80 and above | 359 (10.3%) | 338 (10.1%) | |
| Income quartile[b] | Q1 | 534 (15.4%) | 795 (23.8%) | <0.001 |
| | Q2 | 1,091 (31.4%) | 903 (27.1%) | |
| | Q3 | 1,054 (30.3%) | 897 (26.9%) | |
| | Q4 | 796 (22.9%) | 743 (22.3%) | |
| Years of education | 9 or less | 1,556 (44.8%) | 1,715 (51.4%) | <0.001 |
| | 10–12 | 1,284 (36.9%) | 1,259 (37.7%) | |
| | 13 and over | 635 (18.3%) | 364 (10.9%) | |
| Longest occupation | Professional/technical | 968 (27.9%) | 351 (10.5%) | <0.001 |
| | Managerial | 345 (9.9%) | 26 (0.8%) | |
| | Clerical | 371 (10.7%) | 812 (24.3%) | |
| | Sales/service | 309 (8.9%) | 646 (19.4%) | |
| | Skilled/manual | 970 (27.9%) | 413 (12.4%) | |
| | Agriculture/forestry/fishery worker | 202 (5.8%) | 217 (6.5%) | |
| | Other | 303 (8.7%) | 572 (17.1%) | |
| | Unemployed | 7 (0.2%) | 301 (9.0%) | |
| Marital status | Married | 3,159 (90.9%) | 2,329 (69.8%) | <0.001 |
| | Widowed | 227 (6.5%) | 861 (25.8%) | |
| | Separated/unmarried | 65 (1.9%) | 123 (3.7%) | |
| | Other/missing | 24 (0.7%) | 25 (0.7%) | |
| BMI, kg/m$^2$ | <18.5 | 77 (2.2%) | 160 (4.8%) | <0.001 |
| | 18.5–24.9 | 1,671 (48.1%) | 1,539 (46.1%) | |
| | 25.0–29.9 | 573 (16.5%) | 449 (13.5%) | |
| | ≥30.0 | 36 (1.0%) | 67 (2.0%) | |
| | Missing | 1,118 (32.2%) | 1,123 (33.6%) | |
| Hypertension | No | 1,500 (43.2%) | 1,406 (42.1%) | 0.040 |
| | Yes | 1,659 (47.7%) | 1,675 (50.2%) | |
| | Missing | 316 (9.1%) | 257 (7.7%) | |
| High TG | No | 2,208 (63.5%) | 2,145 (64.3%) | 0.54 |
| | Yes | 1,267 (36.5%) | 1,193 (35.7%) | |
| Low HDL | No | 2,682 (77.2%) | 2,673 (80.1%) | 0.004 |
| | Yes | 793 (22.8%) | 665 (19.9%) | |
| Smoking status | No | 824 (23.7%) | 2,851 (85.4%) | <0.001 |
| | Smoker/ex-smoker | 2,433 (70.0%) | 192 (5.8%) | |
| | Missing | 218 (6.3%) | 295 (8.8%) | |
| Alcohol intake | Drinker/ex-drinker | 2,248 (64.7%) | 631 (18.9%) | <0.001 |
| | None | 1,036 (29.8%) | 2,543 (76.2%) | |
| | Missing | 191 (5.5%) | 164 (4.9%) | |
| Walking time, min/day | <30 | 889 (25.6%) | 989 (29.6%) | <0.001 |
| | ≥30.0 | 2,447 (70.4%) | 2,183 (65.4%) | |
| | Missing | 139 (4.0%) | 166 (5.0%) | |
| Meat/fish intake, servings/day | ≥1 | 1,138 (32.7%) | 1,376 (41.2%) | <0.001 |
| | <1 | 2,133 (61.4%) | 1,796 (53.8%) | |
| | Missing | 204 (5.9%) | 166 (5.0%) | |
| Fruit/vegetable intake, servings/day | ≥1 | 2,478 (71.3%) | 2,773 (83.1%) | <0.001 |
| | <1 | 812 (23.4%) | 406 (12.2%) | |
| | Missing | 185 (5.3%) | 159 (4.8%) | |

BMI, body mass index; HDL, high-density lipoprotein; TG, triglyceride.
[a]Chi-squared test for sex.
[b]Income quartile calculated by all participants in JAGES2010 ('Low' −1.250, 'Middle-low' 1.251–1.944, 'Middle-high' 1.945–3.061, 'High' 3.062− million yen per year).

## RESULTS

Among our study participants, 15.2% of men and 10.2% of women had diabetes. Around 70% of both men and women were

**Table 2.** Prevalence of diabetes mellitus by socioeconomic status and sex ($N = 6{,}813$)

| | *N* | Men | *N* | Women |
|---|---|---|---|---|
| **Income quartile**[a] | | | | |
| Q1 | 534 | 94 (17.6%) | 795 | 93 (11.7%) |
| Q2 | 1,091 | 147 (13.5%) | 903 | 105 (11.6%) |
| Q3 | 1,054 | 166 (15.7%) | 897 | 86 (9.6%) |
| Q4 | 796 | 120 (15.1%) | 743 | 58 (7.8%) |
| *P*-value | | 0.16 | | 0.03 |
| **Years of formal education** | | | | |
| 9 or less | 1,556 | 237 (15.2%) | 1,715 | 187 (10.9%) |
| 10–12 | 1,284 | 196 (15.3%) | 1,259 | 113 (9.0%) |
| 13 and over | 635 | 94 (14.8%) | 364 | 42 (11.5%) |
| *P*-value | | 0.96 | | 0.16 |
| **Longest occupation** | | | | |
| Professional/technical | 968 | 145 (15.0%) | 351 | 43 (12.3%) |
| Managerial | 345 | 56 (16.2%) | 26 | 5 (19.2%) |
| Clerical | 371 | 59 (15.9%) | 812 | 71 (8.7%) |
| Sales/service | 309 | 51 (16.5%) | 646 | 55 (8.5%) |
| Skilled/manual | 970 | 132 (13.6%) | 413 | 47 (11.4%) |
| Agriculture/forestry/fishery workers | 202 | 28 (13.9%) | 217 | 21 (9.7%) |
| Other | 303 | 53 (17.5%) | 572 | 62 (10.8%) |
| Unemployed | 7 | 3 (42.9%) | 301 | 38 (12.6%) |

*P*-values were calculated using Chi-squared test.
[a]Income quartile calculated by all participants in JAGES2010 ('Low' –1.250, 'Middle-low' 1.251–1.944, 'Middle-high' 1.945–3.061, 'High' 3.062– million yen per year).

under 75 years old (Table 1). A total 15.4% of men and 23.8% of women were in the low-income quartile; these percentages were 22.9% and 22.3%, respectively, for the high-income quartile. With regard to years of formal education, 44.8% of men and 51.4% of women had nine years or fewer years of schooling, and 18.3% of men and 10.9% of women had 13 years or more. Distributions of longest occupation were entirely different between men and women. Compared with the entire JAGES 2010 population, our study participants were older; had slightly higher income and lower education levels; and there were more married and physically active participants, as well as more alcohol drinkers (eTable 1).

The prevalence of diabetes by income quartile among men was 17.6% in Q1 (lowest income), 13.5% in Q2, 15.7% in Q3, and 15.1% in Q4 (highest income) ($P = 0.16$); among women, prevalence values were 11.7% in Q1, 11.6% in Q2, 9.6% in Q3, and 7.8% in Q4 ($P = 0.03$) (Table 2). Education- and occupation-related gradients were not observed in the population.

The results of multivariate analysis showed that among women, an income-based gradient was observed in the prevalence of diabetes. Compared with Q4 (highest income category), PRs of diabetes for Q1, Q2, and Q3 were 1.43 (95% CI, 1.07–1.90), 1.33 (95% CI, 1.01–1.75), and 1.22 (95% CI, 0.91–1.64) (*P* for trend = 0.01; Table 3, women, model 1). After mutually adjusting for each SES factor, the PRs of Q1, Q2, and Q3 compared to Q4 were 1.42 (95% CI, 1.06–1.90), 1.33 (95% CI, 1.00–1.76), and 1.23 (95% CI, 0.91–1.65), respectively (*P* for trend = 0.016; Table 3, women, model 2). Even after adjustment for marital status, BMI, other metabolic risk factors, and lifestyle factors, the association was not attenuated (Q1: PR 1.43; 95% CI, 1.07–1.92, Q2: PR 1.32; 95% CI, 0.99–1.76; Q3: PR 1.22; 95% CI, 0.91–1.65; *P* for trend = 0.01; Table 3, women, model 3). No socioeconomic gradient was observed among men (Table 3).

The estimates based on our sensitivity analysis using complete case data were mostly identical to our original analysis with slightly smaller PRs and wider CIs (eTable 2).

## DISCUSSION

Using the large-scale data of Japanese older adults, we found two major findings on the social inequality in objectively measured

**Table 3.** Prevalence ratios and 95% confidence intervals for diabetes mellitus by sex with multiple imputation ($N = 9{,}893$)

| | Men ($N = 4{,}471$) | | | Women ($N = 5{,}422$) | | |
|---|---|---|---|---|---|---|
| | Model 1 | Model 2 | Model 3 | Model 1 | Model 2 | Model 3 |
| **Income quartile** | | | | | | |
| Q1 (lowest) | 1.16 (0.90–1.50) | 1.16 (0.88–1.52) | 1.18 (0.89–1.56) | 1.43 (1.07–1.90) | **1.42 (1.06–1.90)** | **1.43 (1.07–1.92)** |
| Q2 | 0.93 (0.74–1.17) | 0.94 (0.74–1.19) | 0.96 (0.76–1.22) | 1.33 (1.01–1.75) | **1.33 (1.00–1.76)** | **1.32 (0.99–1.76)** |
| Q3 | 1.02 (0.79–1.30) | 1.03 (0.80–1.32) | 1.02 (0.80–1.31) | 1.22 (0.91–1.64) | 1.23 (0.91–1.65) | 1.22 (0.91–1.65) |
| Q4 (Highest) | 1 (referent) | 1 (referent) | 1 (referent) | 1 (referent) | 1 (referent) | 1 (referent) |
| Trend *P* | 0.52 | 0.55 | 0.44 | **0.01** | **0.016** | **0.01** |
| **Years of formal education** | | | | | | |
| 9 years or less | 0.97 (0.78–1.20) | 0.99 (0.78–1.24) | 1.01 (0.80–1.27) | 1.03 (0.77–1.39) | 0.97 (0.72–1.33) | 0.94 (0.69–1.28) |
| 10–12 | 1.00 (0.81–1.25) | 1.02 (0.81–1.27) | 1.02 (0.81–1.27) | 0.96 (0.71–1.30) | 0.96 (0.71–1.31) | 0.95 (0.69–1.29) |
| 13+ | 1 (referent) | 1 (referent) | 1 (referent) | 1 (referent) | 1 (referent) | 1 (referent) |
| Trend *P* | 0.701 | 0.84 | 0.99 | 0.56 | 0.96 | 0.72 |
| **Longest occupation** | | | | | | |
| Professional/Technical | 1 (referent) | 1 (referent) | 1 (referent) | 1 (referent) | 1 (referent) | 1 (referent) |
| Managerial | 0.98 (0.73–1.32) | 0.98 (0.73–1.33) | 0.98 (0.73–1.33) | 1.20 (0.60–2.41) | 1.21 (0.60–2.41) | 1.20 (0.60–2.40) |
| Clerical | 1.03 (0.77–1.38) | 1.03 (0.77–1.38) | 1.05 (0.78–1.42) | 0.84 (0.59–1.20) | 0.86 (0.60–1.24) | 0.88 (0.62–1.26) |
| Sales/Service | 1.05 (0.79–1.41) | 1.04 (0.78–1.39) | 1.01 (0.75–1.35) | 0.79 (0.55–1.12) | 0.77 (0.53–1.10) | 0.75 (0.53–1.08) |
| Skilled/Manual | 0.86 (0.69–1.08) | 0.87 (0.69–1.09) | 0.86 (0.68–1.08) | 0.99 (0.69–1.43) | 0.97 (0.66–1.41) | 0.95 (0.65–1.39) |
| Agriculture/Forestry/Fishery workers | 0.91 (0.63–1.31) | 0.89 (0.62–1.30) | 0.88 (0.61–1.29) | 0.86 (0.55–1.33) | 0.83 (0.53–1.29) | 0.81 (0.52–1.27) |
| Others | 1.12 (0.84–1.48) | 1.10 (0.82–1.47) | 1.08 (0.81–1.45) | 1.03 (0.72–1.46) | 0.99 (0.69–1.42) | 0.98 (0.68–1.41) |
| Unemployed | 1.64 (0.68–3.94) | 1.60 (0.66–3.90) | 1.59 (0.66–3.86) | 1.03 (0.68–1.57) | 1.00 (0.65–1.54) | 0.98 (0.64–1.50) |

Model 1 was adjusted for adjusted for income quartile, years of formal education and longest occupation separately with age.
Model 2 was adjusted for income quartile, years of formal education, longest occupation, and age.
Model 3 was additionally adjusted for marital status, BMI, hypertension, low HDL, high TG, smoking status, alcohol drinking habit, walking time per day, and meat/fish intake and vegetable intake.

diabetes: 1) the clear income gradient in diabetes prevalence was only observed among women but not among men; and 2) among men and women, there was no clear gradient in diabetes prevalence by years of education and longest occupation.

The socioeconomic gradient was potentially more marked among women, which was consistent with recent studies in other countries.[12,21] Robbins et al have proposed, as potential reasons, that women culturally have difficulties in health care access than men, fewer opportunities for regular exercise, unhealthy lifestyle behaviors, disadvantaged nutritional factors, more psychological stress, more depression, and more negative pre- or peri-natal environmental factors.[18] Other scholars have suggested the different roles of obesity in the association between income and diabetes by sex. In a Swedish study, Agardh et al found that among the low-income group, BMI explained their excess risk for subjectively diagnosed type 2 diabetes by 21% among men and 35% among women.[3] Nonetheless, a study from Canada that investigated the association between self-reported diabetes and SES found that BMI did not explain the associations between income and diabetes both among men and women.[20] In the present study, further adjustment for covariates, including BMI, did not substantially alter the association between income and diabetes for both sexes. To clarify the reasons for the sex difference, further studies are needed.

We found a gradient in diabetes by income but not by education or occupation. In theory, income has both materialistic and psychosocial functions, and they may explain the income gradient in diabetes distinctively. First, low income means limited access to material goods and services to prevent diabetes, such as balanced diet and necessary preventive care.[34] Second, the access limitation also leads to the social isolation and exclusion because of the lack of opportunities for social interactions, leading to mental stresses. Stress science and endocrinological studies have suggested the direct effects of stress hormones on blood glucose levels and insulin intolerance, as well as health behaviors.[34] Potential gender differences in our result could be explained by the psychosocial functions of income, including health beliefs, attitudes, and lifestyles, which may differ between men and women even at the same income levels.[21] Specifically, as suggested by Saito et al, the loss of social interactions due to the lack of income might affect women more than men among Japanese older adults.[35] Lastly, although the detailed mechanisms are unknown, sex differences in the gene-related tolerance for diabetes may also explain the stronger association among women found in our study.[36,37]

Although we found a gradient in diabetes by income but not by education or occupation, these results were inconsistent with those among young or middle-aged adults[18] but consistent with results from older populations.[6] Socioeconomic status in older people should be interpreted differently from that at younger ages.[38] In many countries, older people are likely to have lower educational attainment. Among our study participants, the percentage of people with university or higher level educations was small in the age group investigated: 18.3% for men and 10.9% for women (Table 1). However, the university entrance rate in Japan was 56.6% among men and 57.1% among women in 2016.[39] Consequently, the number of older people with high educational attainment is small, resulting in less statistical power to capture the association between education level and diabetes. The null finding between longest occupation and diabetes among men and women may be explained by weak statistical power owing to small sample sizes of each occupational category. For example, among men, the PR of diabetes among unemployed compared with professional/technical workers was large (PR 1.64; 95% CI, 0.68–3.94), which is in line with known occupation-based health disparities around the world (Table 3).[40] Alternatively, the survivor effect could alter the association between education, previous occupation, and diabetes, given that those who are socioeconomically deprived are less likely to survive; this tendency could be stronger in Japan, where many people experienced the life-threatening post-war period.[41]

Apart from those discussed above, four additional limitations in our study should be mentioned. First and foremost, owing to the cross-sectional nature of the study, we cannot exclude the possibility of reverse causation (ie, diabetes causes reduced income). Second, generalizability is limited, as our data were obtained only from regions of central Japan and the study does not include older people with long-term care insurance. Third, selection bias should also be discussed. Our study participants were more health conscious than the general population, as participants were limited to those who underwent health checkups. In Japan, about 38% of the population received health checkups in 2010.[42] Underestimation of the magnitude of SES-related health associations found in this study may be owing to this bias. Nonetheless, our sensitivity analysis using complete cases only showed the same income-based gradient, suggesting the missing did not induce a critical bias to the levels of the income-based gradient in diabetes. Finally, we did not evaluate the health gradient stemming from other SES indicators, including previously suggested indicators associated with health: wealth,[5] relative deprivation,[43] and social exclusion.[35] Specifically, future studies should evaluate the wealth-based gradient given that older adults are more likely to rely on savings or other similar financial resources rather than regular income, which mostly consists of a government pension.

In conclusion, we found a clear income-based gradient in diabetes among Japanese older adults and the gradient was potentially more remarkable among women, but this was not the case for education and longest occupation. This was the first large-scale study clarifying the socioeconomic disparity in diabetes among Japanese older population. Given the findings of this study, monitoring income gradient in diabetes is important in public health actions, even in older populations. Future longitudinal and intervention studies should evaluate the causal link of income to diabetes onset, determine the mechanisms of the potential sex differences in the income/diabetes association, and identify ways to mitigate the income-based inequality.

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## APPENDIX A. SUPPLEMENTARY DATA

Supplementary data related to this article can be found at https://doi.org/10.2188/jea.JE20170206.

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# 地域在住高齢者が転出に至る要因の研究
## －望まない転出を予防するために－

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目的　本研究は，地域在住高齢者を対象として転出する前の状況から転出に至った経緯を縦断的に分析して，転出に至る要因を明らかにし，望まない転出を予防する要因の示唆を得ることを目的とした。
方法　調査は2010年と2013年に全国24市町村にて実施した。65歳以上の要介護認定を受けていない地域在住高齢者131,468人を対象に郵送調査を行い，86,005人（回収率65.4%）から回答が得られた。このうち，最長1,374日間（平均1,152日間）追跡ができた81,810人の中から，767人（男性0.84%，女性1.0%）が転出した。身体状況や機能状態，心理・社会的，社会経済的，社会参加，地域環境や外出頻度を自記式質問紙で調査した。Cox比例ハザードモデルを用いて，転出と上記調査項目との関連について，ハザード比と95%信頼区間（以下，95%CI）を算出した。
結果　分析の結果，人口密度が10分の1になるにつれて1.32（95%CI：1.19-1.47）倍の転出をしていた。環境要因では，１人暮らしだと2.22（95%CI：1.85-2.68）倍，等価所得が200万円未満（400万円以上と比較）だと1.35（95%CI：1.00-1.82）倍のリスクがあった。社会参加・活動の要因では，老人クラブに参加していないと2.27（95%CI：1.79-3.53）倍，スポーツの会に参加していないと1.53（95%CI：1.24-2.68）倍，趣味の会に参加していないと1.32（95%CI：1.11-2.87）倍，趣味がないと1.44（95%CI：1.23-1.89）倍のリスクだった。生活状況の要因では，野菜果物の摂取が週1回未満であると2.20（95%CI：1.14-4.24）倍，肉魚の摂取が週1回未満であると1.55（95%CI：1.03-2.34）倍のリスクだった。主観的な要因では，主観的健康感が悪いと1.40（95%CI：1.18-1.66）倍，地域への愛着がないと3.02（95%CI：2.59-3.53）倍のリスクだった。健康状態の要因では，半年以内に体重減少があると1.36（95%CI：1.12-1.65）倍のリスク，過去1年以内に転倒した経験があると1.45（95%CI：1.14-1.85）倍のリスクだった。し好品の要因では，タバコを吸っていると1.39（95%CI：1.05-1.85）倍のリスクだった。疾病状況では，がんの治療中だと1.43（95%CI：1.03-2.00）倍，心臓病治療中だと1.38（95%CI：1.10-1.72）倍，糖尿病治療中だと1.28（95%CI：1.03-1.60）倍のリスクだった。
結論　高齢者が住み慣れた地域で暮らし続けるためには，積極的な社会参加をすること，地域に愛着をもってもらうこと，健康状態が主観的にも客観的にも保たれていること，食事や所得など生活状態が安定していることが示唆された。
キーワード　地域在住高齢者，転出，社会参加，縦断分析，生活状態

厚生の指標　第65巻第4号　2018年4月

# 個人および地域レベルにおける要介護リスク指標とソーシャルキャピタル指標の関連の違い－JAGES2010横断研究－

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目的　地域づくりによる介護予防を推進する上で地域診断が重要とされ，ソーシャルキャピタル（Social Capital，以下，SC）が注目されている。地域診断指標の課題に生態学的錯誤（地域レベルの変数間の関連から個人レベルの関連を誤って推論），個人主義的錯誤（個人レベルの変数間の関連から地域レベルの関連を誤って推論）が挙げられる。地域診断に用いるSC指標にはこの両者がないことが望ましい。本研究の目的は，個人・地域の両レベルにおいて2つの錯誤がない要介護リスクと関連を示すSC指標を抽出することを目的とした。
方法　本研究は，日本老年学的評価研究（JAGES）2010に参加した25保険者31市町村の要介護認定を受けていない65歳以上の高齢者98,744名を分析対象とした。個人レベルのロジスティック回帰分析（有意水準5%）の目的変数には，基本チェックリストの要介護リスク指標である生活機能低下，フレイル，運動機能低下，低栄養，口腔機能低下，閉じこもり，認知機能低下，うつの8指標を使用した。説明変数は，SC指標（社会的サポート，社会参加，社会的ネットワーク，SaitoのSC指標）の頻度別35指標（280モデル）とした。調整変数は，年齢，性別，教育歴，等価所得，疾病の有無，主観的健康感，婚姻状態，家族構成とした。地域レベルの分析単位は校区とし，Spearmanの順位相関分析（有意水準5%）を実施した。変数は個人レベルと同様とし，年齢（前期・後期高齢者）による層別化を実施し，1校区あたり30名以上の前期349校区，後期287校区を分析対象とした。
結果　個人レベルではSC指標が高いほど要介護リスク全8指標が有意に低い保護的な関連が28／35指標（80.0%）でみられた。しかし，地域レベルではSC指標が高いほど要介護リスクが高い非保護的な関連が20／35指標（57.1%）で1つ以上みられた。生態学・個人主義的錯誤がなく，要介護リスクに保護的なSC指標は，社会的サポート，社会参加のうち，ボランティア（週1回，月1～2回），スポーツ・趣味（週1回，月1～2回，年数回），就労ありとSaitoのSC指標（社会参加，連帯感）の15／35指標（42.9%）に留まった。
結論　生態学・個人主義的錯誤は20／35指標（57.1%）でみられ，2つの錯誤がなく要介護リスク抑制を示唆する地域診断指標は，社会的サポートやボランティア・趣味・スポーツ・就労など一部（42.9%）のSC指標に留まった。
キーワード　ソーシャルキャピタル，介護予防，地域づくり，地域診断，生態学的錯誤，個人主義的錯誤

総合リハビリテーション 46 巻 12 号 （2018 年 12 月）

# 介護予防のための地域診断指標—文献レビューと 6 基準を用いた量的指標の評価

井手 一茂 ， 鄭 丞媛 ， 村山 洋史 ， 宮國 康弘 ， 中村 恒穂 ， 尾島 俊之 ， 近藤 克則

要旨 【背景】地域づくりによる介護予防に有用な地域診断の量的指標と今後の課題を明らかにすることを目的とした．【方法】医学中央雑誌 Web，PubMed で検索し入手した日本における 31 論文を対象に，① 研究デザイン，② 地域単位，③ 介護予防アウトカム指標，④関連指標を抽出した．2 つ以上の論文で指標間に有意な関連（再現性）があった指標について，5 人の評価者が相談せずに量的指標に必要な 6 基準を満たすか評価した．【結果】横断研究による市町村・校区レベルを地域単位とした研究が多く，アウトカム 28 指標，関連 69 指標が報告されていた．再現性があった 27 指標のうち 3 人以上が 6 基準を満たすと評価したのは 14 指標で，社会参加やサポートあり割合などが高い地域ほど，うつ，閉じこもり，転倒，残存歯数少ない，要支援・介護認定の割合が低かった．【結語】14 指標が地域診断に有用と思われた．今後は，低栄養，認知機能低下などにかかわる指標開発や縦断研究による予測妥当性の検証が望まれる．

医学のあゆみ 264 巻 4 号　2018 年 1 月 27 日　p.319-322

ライフコースの観点からみたコホート研究とその成果
JAGES（日本老年学的評価研究）

長嶺由衣子・近藤克則　千葉大学予防医学センター

サマリー　　JAGES（Japan Gerontological Evaluation Study ; 日本老年学的評価研究）プロジェクトは，1999 年にはじめられた高齢者を対象としたコホート研究である．プロジェクトの所期の目的は，世界でトップレベルの健康長寿を誇る日本の高齢者を身体・心理・社会的な側面から多面的に描きだすことにあり，現在も時代に応じて随時目的を検討・修正しながら調査・研究を続けている．JAGES は，市町村が 3 年ごとに策定する介護保険事業計画に向けた調査を共同で行う形で，毎回市町村を増やしながら調査を実施してきた．2010 年調査以降，子ども期の生活程度や逆境体験を調査項目に入れ，高齢期の健康指標との関連を検討してきた．今後，生活習慣病へのライフコースの影響を検証するため，すでに健診データと結合するなどの改善を行っている．さらに，質問紙調査と要介護認定，死亡データにとどまらない医療レセプトや介護レセプトなどのデータと結合することができれば，より詳細な研究が可能となるであろう．

キーワード　JAGES，高齢期，個人要因，地域要因，ライフコース

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Featured Article

# Social interaction and cognitive decline: Results of a 7-year community intervention

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**Abstract**

**Introduction:** There are few intervention studies that demonstrated linking social participation to lower risk of cognitive decline. We examined prospectively the protective effect of a community intervention program promoting social participation on the incidence of cognitive disability.

**Methods:** The baseline was established in a survey of community-dwelling older people aged 65 years old or more in July 2006 (2793 respondents, response rate 48.5%). The setting was Taketoyo town in Japan, where municipal authorities launched an intervention that was based on the establishment of community-based centers called "salons," where the town's senior residents could congregate and participate in social activities, ranging from arts and crafts, games, and interactive activities with preschool children. Three salons were established in May 2010, and a total of 10 salons were in operation by 2013. We recorded the frequency of salon participation among survey respondents till 2013 and conducted two follow-up surveys (in 2010 and 2013) to collect information about health status and behaviors. The onset of cognitive disability was followed from May 2007 to January 2014. We used the marginal structural models to evaluate the effect of program.

**Results:** The range of prevalence of cognitive disability was from 0.2% to 2.5% during the observation period. The proportion of respondents who participates to salons increased over time to about 11.7%. The frequency of salon participation was protectively associated with cognitive decline, even after adjusting for time-dependent covariates and attrition (odds ratio = 0.73, 95% confidence interval: 0.54–0.99).

**Discussion:** Our study suggests that operating community salons that encourage social interactions, light physical activity, and cognitive activities among older participants may be effective for preventing cognitive decline. In future studies, we need to understand what sorts of activities (e.g., those involving light physical activity vs. purely intellectual activities) are most effective in maintaining cognitive function.



*Keywords:* Prevention; Community intervention; Social participation; Japan; Marginal structural models

## 1. Introduction

Dementia is a major cause of disability and dependency among older people. Worldwide, an estimated 47.5 million people suffer from dementia while 7.7 million new cases are added each year [1]. The number of people living with dementia is projected to triple by 2050 [2].

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Japanese society is confronted with the fastest pace of population aging in the world, with a population prevalence of dementia that is nearly double the world average (Japan: 15.0% in 2012 [3] vs. world average: 5.2% in 2015 [4]). The population with dementia is forecast to reach 7 million by 2025 [5]. The prevention of dementia is therefore a top priority for Japanese public health policy [6]. One approach advocated by Japanese government to prevent cognitive decline in older adults has been to encourage more social participation [7]. Observational studies suggested that social participation is associated with lower risk of cognitive decline [8]. However, these studies are prone to confounding bias due to their observational nature; specifically, the selective participation of cognitively healthier individuals in community-based programs encouraging social participation.

Since 2007, the municipality of Taketoyo (population 41,000) in Aichi Prefecture, Japan, has been engaged in a community intervention program designed to facilitate social participation among older residents, with the aim of preventing cognitive and physical functional decline. The intervention program is based on opening 10 community-based centers (referred to as "salons"), where seniors can congregate to engage in a variety of social programs and activities [9,10].

In the present study, we report on the 7-year evaluation of this intervention program. Our design is quasi-experimental in the sense that the community salons were newly established by the Taketoyo municipality (where none previously existed), and we have information (from an ongoing cohort study) about the health status of individuals before and after the salons were opened. Because of the repeated assessments of salon participation and covariates over time, there is a possibility of time-varying confounding. For example, social participation could improve health status (covariates), which also may influence the probability of social participation in subsequent time periods. Therefore, we attempted to address time-varying confounding through marginal structural modeling with inverse probability weighting.

## 2. Methods

### 2.1. Study population

The study population consisted of participants in the Aichi Gerontological Evaluation Study (AGES), which was established in 1999 in Aichi prefecture. One of the field sties of the AGES cohort was in the town of Taketoyo (population 41,531 in 2006) [11]. We conducted a mail-in questionnaire survey of all community-dwelling older people who were physically and cognitively independent and aged 65 years or older ($n$ = 5759) in July 2006.

The questionnaire survey inquired about personal characteristics, health status, and health habits of the respondents. As shown in Fig. 1, the response rate to the baseline survey was 48.5% ($n$ = 2793). An additional 200 subjects were removed from the baseline of the present study because of missing/invalid information, relocation out of the area, or death/incident disability.

Of the 2593 eligible participants from the baseline survey, we lost 326 subjects due to death/functional decline and relocation during the 3-year and 3-month follow-up period. In the second survey in August 2010, we recontacted 1769 individuals (participation rate: 78.0%). During an additional 3-year and 2-month follow-up term, 268 respondents dropped out of our cohort. We obtained 1352 responses from 1501 eligible subjects in the third survey conducted in October 2013 (participation rate: 90.1%). The cumulative follow-up rate during the total period was therefore 73.9%. We also collected information on their frequency of participating in salons until the end of March 2014. In addition, the onset of functional and cognitive disability was followed from May 1, 2007 to January 7, 2014. (The observation period was 2443 days.) Our study protocol was approved by the Ethics Committee at Nihon Fukushi University and Seijoh University.

### 2.2. The intervention

Taketoyo town is located approximately 35 km south of Nagoya in Aichi Prefecture, Japan. The community salon project was launched in May 2007 when the municipal authorities began to open community-based centers where the town's senior residents could congregate and participate in social activities. Initially, three salons were established, and by 2013, a total of 10 salons were in operation, staffed by community volunteers (Fig. 1).

In September 2013, we conducted a survey of a sample of 152 volunteers in the Taketoyo program. Of the 91 volunteers who responded to the survey (26 male and 65 female), the average age was 71 years, and the average duration of volunteering was 4.3 years. Almost all of them were recruited to volunteer by their friends who were participating in the salon activities. They were all required to take a training course conducted by two occupational therapists.

Although the salon programs were not standardized across locations, popular activities included dance classes, chatting with other participants, arts and crafts (calligraphy, origami, and poetry recitation), singing, playing musical instruments, quizzes and games (e.g., bingo, cards, Japanese chess), as well as interactive activities with preschool children. In each salon, 90–120 minutes of programming were scheduled between 1 to 3 times per month.

Any resident aged 65 years or older was eligible to participate for a nominal fee of 100 Japanese Yen (roughly 1 US dollar) per visit.

### 2.3. Outcome variable

Our primary outcome is the onset of cognitive disability assessed by a standardized in-home assessment. The Japanese government established a national long-term care insurance scheme in 2000 [12]. Under this system, a certification

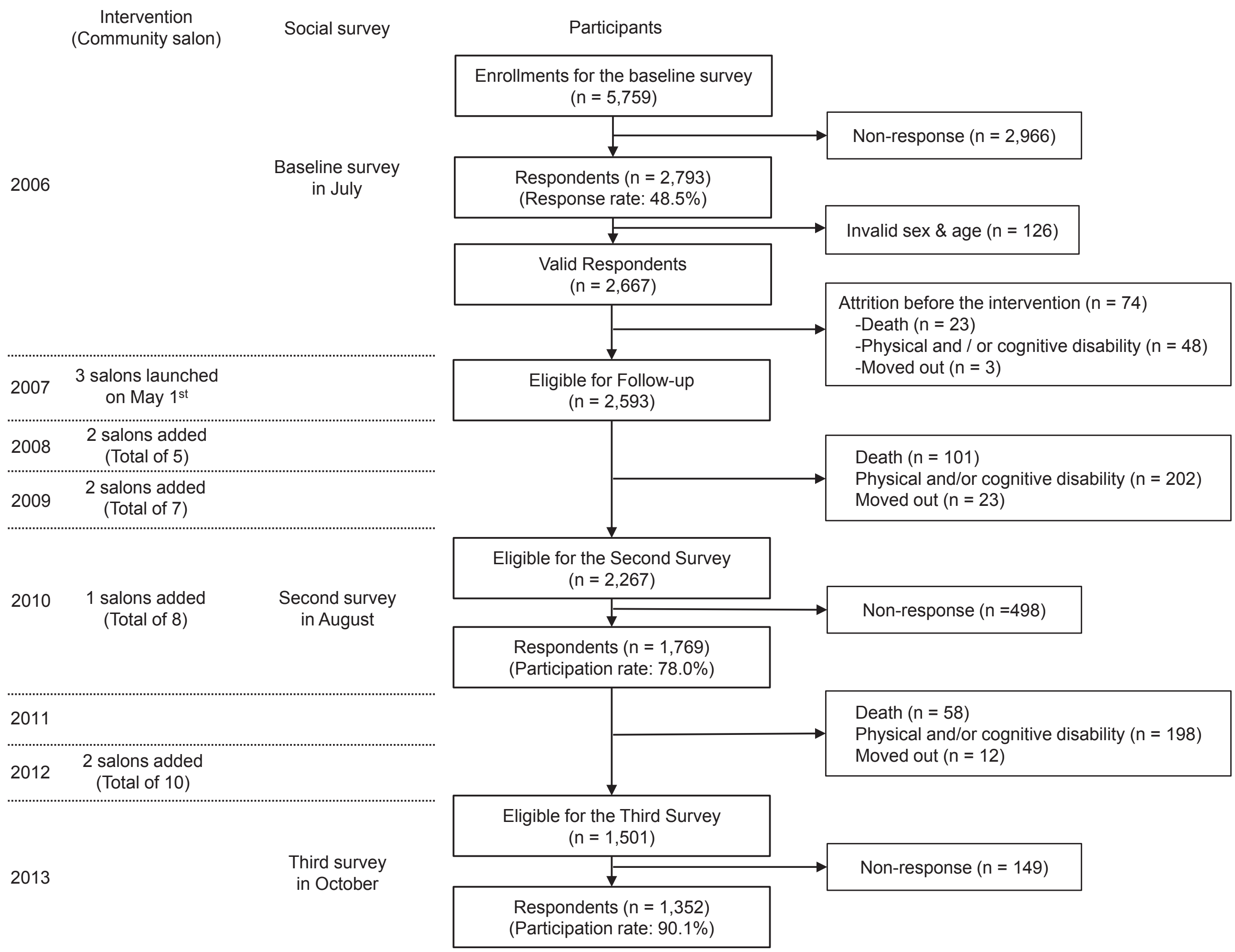


Fig. 1. Participants flow during the follow-up period.

committee in each municipality dispatches trained investigators to applicants' homes to evaluate their eligibility for nursing care (e.g., home helper).

During the home visit, each individual is assessed with regard to their activities of daily living and instrumental activities of daily living status, cognitive function (e.g., short-term memory, orientation, and communication), as well as mental and behavioral disorders (e.g., delusions of persecution, confabulation, and soliloquy) using a standardized protocol. Following the assessment, the applicants are classified into one of eight levels according to the severity of their cognitive disability (Supplementary Table 1).

The level of cognitive decline is strongly correlated with the Mini–Mental State Examination (Spearman rank correlation $\rho = -0.73, P < .001$), according to a previous validation study among 116 institutionalized older residents and 66 older individuals requiring outpatient day long-term care (The age range was between 60 to 101 years) [13]. Another previous study demonstrated that level I of the home assessment scale corresponds to a 0.5-point rating on the clinical dementia rating (suspected dementia) for 590 older individuals aged 75 years and more (specificity and sensitivity, both 0.88) [14]. On the basis of these validation studies, a cognitive decline rating of level I or higher was defined as incident cognitive disability in this study.

We matched data from the AGES participants to the long-term care insurance registry for the follow-up period from May 1, 2007 to January 7, 2014. Respondents who died, or who developed incident functional disability without cognitive decline, or moved out of the area, or did not respond to the survey during the follow-up period, were censored.

### 2.4. Explanatory variable

Our primary treatment variable was frequency of participation in community salons in each year. Taketoyo town keeps a participant register that recorded the participant's name and frequency of salon participation. Among participants during the 7-year follow-up, 10%–50% of them participated only once to 3 times in any given year. Our definition of "participant" was restricted to those who participated more than 3 times a year (numbers of the participants: 99, 159, 152, 162, 162, 175, and 158, for each year) (Supplementary Table 2) because we hypothesized that participation on fewer occasions could not be plausibly expected to prevent functional disability. Therefore, anyone who participated on fewer than four occasions was classified as nonparticipants. Because the distribution was right-skewed in each year, we log-transformed the values (arbitrarily inserting "1" instead of "0" for those who did not participate at all).

### 2.5. Covariates

We selected as potential confounding variables sex, age, educational attainment, and equalized household income [15], depressive symptoms measured by the Geriatric Depression Scale-15 (GDS-15) [16], low cognitive function [17], comorbidities including stroke, heart disease, and hypertension [18], drinking and smoking [19], walking time [20], and frequency of participating in sports clubs [21].

We did not perform a detailed assessment of mild cognitive impairment; instead we asked three items on the survey: "Have people around you noticed that you are forgetful, for example, by telling you that you often ask the same thing?" "Can you look up phone numbers and make phone calls by yourself" and "Do you sometimes forget what date it is today?" [17].

Age was grouped into the following categories: 65–74 years and 75 years or over. Educational attainment was categorized as 9 years or under and 10 years or over. Household income was equalized by the square root of the number of household members and categorized into "Under 2.0 million JPY" and "2.0 million JPY and over." Depressive symptoms were categorized into lower risk (4 points and under) versus higher risk (5 points and over) [22]. Low cognitive function was categorized as negative condition (1–3 points) versus neutral condition (0 point) [17]. As shown in Supplementary Fig. 1, our covariates were updated by the surveys conducted in 2010 and 2013.

### 2.6. Statistical analysis

Initially, we conducted a discrete time logistic regression analysis adjusting for all covariates to examine the association between frequency of salon participation and onset of cognitive disability.

In longitudinal data, exposures and covariates can have complex bidirectional associations. For example, salon participation (our treatment of interest) could affect the onset of diseases. Incident chronic conditions can therefore simultaneously act as a confounder of the relation between salon participation and cognitive function as well as a potential mediating variable between future salon participation and cognitive function. This scenario of time-varying confounding cannot be addressed by conventional methods such as restriction, stratification, or covariance adjustment [23]. In addition, our data had a high attrition rate in which 1241 respondents dropped out before the third wave survey, which may induce selection bias [24].

To address these potential biases, we used Marginal Structural Models (MSMs) that estimated the stabilized inverse probability of receiving the treatment (i.e., salon participation) $SW(t)$ and the probability of remaining uncensored up to time t $SW(c)$ to create a pseudo-population to balance the distribution of potential confounders across exposure levels and uncensored cases [25]. $SW(t)$ and $SW(c)$ are defined as

$$SW(t)=\prod_{k=0}^{t}\frac{f\{A(k)|\overline{A}(k-1),\ V,\ \overline{C}(k)=0\}}{f\{A(k)|\overline{A}(k-1),\overline{L}(k),\ \overline{C}(k)=0\ \}}$$

and

$$SW(c)=\prod_{k=0}^{t}\frac{\Pr\{C(k)=0|\overline{C}(k-1)=0,\ \overline{A}(k-1),\ V\}}{\Pr\{C(k)=0|\overline{C}(k-1)=0,\overline{A}(k-1),\ \overline{L}(k-1)\ \}}$$

where $A(k)$ denotes the exposure at year $k$, $\overline{A}(k-1)$ represents the exposure history prior year $k$, $V$ is baseline covariates, $\overline{L}(k-1)$ represents the covariates history including $V$, $C(k)$ is the incident censoring at year $k$, $\overline{C}(k-1)$ is uncensored history until year $k-1$.

Because our treatment—that is, frequency of salon participation—is a continuous measure, we estimated each person's probability density of receiving treatment in each year [26]. We used Poisson regression to create $SW(t)$ because the exposure is a zero-inflated variable [27]. The $SW(c)$ was estimated using logistic regression. And, we obtained an overall weight $SW=SW(t)\times SW(c)$ [28].

To estimate the odds ratios of cognitive decline as a result of salon participation, we used all baseline covariates $V$ as well as time-varying covariate history $\overline{L}$ in the weighted MSMs model to control for potential confounding in the pseudo-population [29].

In addition, the data obtained from the questionnaires included missing answers in each wave, which may cause loss of power and biased estimations [30]. To address potential bias due to missing data, we used multiple imputation by Markov Chain Monte Carlo method for covariates in each year [31]. We created 20 imputed data sets and combined each result of analysis using the Stata command "mi estimate."

All analyses were performed using STATA version 14.0 (STATA Corp LP., College Station, Texas, USA).

## 3. Results

Descriptive statistics are shown in Table 1. The incidence of cognitive disability tended to increase over time. The range of prevalence of cognitive disability was from 0.2% to 2.5% during the observation period. The proportion of respondents who participates to salons increased over time to about 11.7%. The incidence rate ratio for salon participation versus nonsalon participation was 0.61 (95% confidence interval [CI]: 0.30–1.10). The averaged frequency of

Table 1
Characteristics of analytic samples during follow-up period

| | 2007 | | 2008 | | 2009 | | 2010 | | 2011 | | 2012 | | 2013 | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | *n* | % | *n* | % | *n* | % | *n* | % | *n* | % | *n* | % | *n* | % |
| Cognitive disability | | | | | | | | | | | | | | |
| Nonincidence | 2556 | 98.6 | 2524 | 98.8 | 2494 | 98.8 | 1753 | 99.1 | 1715 | 97.8 | 1673 | 97.5 | 1349 | 99.8 |
| Incidence | 37 | 1.4 | 32 | 1.2 | 30 | 1.2 | 16 | 0.9 | 38 | 2.2 | 42 | 2.5 | 3 | 0.2 |
| Total | 2593 | 100 | 2556 | 100 | 2524 | 100 | 1769 | 100 | 1753 | 100 | 1715 | 100 | 1352 | 100 |
| Participation* | | | | | | | | | | | | | | |
| Nonparticipants | 2494 | 96.2 | 2397 | 93.8 | 2372 | 94.0 | 1607 | 90.8 | 1591 | 90.8 | 1540 | 89.8 | 1194 | 88.3 |
| Participants | 99 | 3.8 | 159 | 6.2 | 152 | 6.0 | 162 | 9.2 | 162 | 9.2 | 175 | 10.2 | 158 | 11.7 |
| Total | 2593 | 100 | 2556 | 100 | 2524 | 100 | 1769 | 100 | 1753 | 100 | 1715 | 100 | 1352 | 100 |
| Sex | | | | | | | | | | | | | | |
| Male | 1306 | 50.4 | - | - | - | - | 874 | 49.4 | - | - | - | - | 665 | 49.2 |
| Female | 1287 | 49.6 | - | - | - | - | 895 | 50.6 | - | - | - | - | 687 | 50.8 |
| Total | 2593 | 100 | - | - | - | - | 1769 | 100 | - | - | - | - | 1352 | 100 |
| Age | | | | | | | | | | | | | | |
| 65–74 years | 1770 | 68.3 | - | - | - | - | 910 | 51.4 | - | - | - | - | 387 | 28.6 |
| 75 years and over | 823 | 31.7 | - | - | - | - | 859 | 48.6 | - | - | - | - | 965 | 71.4 |
| Total | 2593 | 100 | - | - | - | - | 1769 | 100 | - | - | - | - | 1352 | 100 |
| Educational attainment | | | | | | | | | | | | | | |
| 9 years and under | 1405 | 54.8 | - | - | - | - | 925 | 52.7 | - | - | - | - | 701 | 52.2 |
| 10 years and over | 1158 | 45.2 | - | - | - | - | 831 | 47.3 | - | - | - | - | 643 | 47.8 |
| Total | 2563 | 100 | - | - | - | - | 1756 | 100 | - | - | - | - | 1344 | 100 |
| Equalized income | | | | | | | | | | | | | | |
| <2.0 million JPY | 921 | 47.1 | - | - | - | - | 682 | 48.3 | - | - | - | - | 574 | 53.5 |
| ≥2.0 million JPY | 1034 | 52.9 | - | - | - | - | 730 | 51.7 | - | - | - | - | 498 | 46.5 |
| Total | 1955 | 100 | - | - | - | - | 1412 | 100 | - | - | - | - | 1072 | 100 |
| Number of comorbidities | | | | | | | | | | | | | | |
| 0 | 756 | 41.6 | - | - | - | - | 530 | 36.8 | - | - | - | - | 368 | 32.5 |
| 1 | 899 | 49.4 | - | - | - | - | 790 | 54.9 | - | - | - | - | 647 | 57.2 |
| 2 | 150 | 8.3 | - | - | - | - | 117 | 8.1 | - | - | - | - | 109 | 9.6 |
| 3 | 14 | 0.7 | - | - | - | - | 2 | 0.2 | - | - | - | - | 7 | 0.7 |
| Total | 1819 | 100 | - | - | - | - | 1439 | 100 | - | - | - | - | 1131 | 100 |
| Depressive symptoms (GDS-15) | | | | | | | | | | | | | | |
| 4 points and under | 1617 | 72.3 | - | - | - | - | 1104 | 72.9 | - | - | - | - | 838 | 74.7 |
| 5 points and over | 620 | 27.7 | - | - | - | - | 411 | 27.1 | - | - | - | - | 284 | 25.3 |
| Total | 2237 | 100 | - | - | - | - | 1515 | 100 | - | - | - | - | 1122 | 100 |
| Low cognitive function | | | | | | | | | | | | | | |
| 0 point | 1486 | 59.7 | - | - | - | - | 1027 | 61.5 | - | - | - | - | 846 | 64.2 |
| 1 point and over | 1002 | 40.3 | - | - | - | - | 644 | 38.5 | - | - | - | - | 471 | 35.8 |
| Total | 2488 | 100 | - | - | - | - | 1671 | 100 | - | - | - | - | 1317 | 100 |
| Current drinking | | | | | | | | | | | | | | |
| No | 1535 | 60.3 | - | - | - | - | 1102 | 65.6 | - | - | - | - | 900 | 68.1 |
| Yes | 1010 | 39.7 | - | - | - | - | 579 | 34.4 | - | - | - | - | 422 | 31.9 |
| Total | 2545 | 100 | - | - | - | - | 1681 | 100 | - | - | - | - | 1322 | 100 |
| Current smoking | | | | | | | | | | | | | | |
| No | 1946 | 86.0 | - | - | - | - | 1459 | 91.0 | - | - | - | - | 1254 | 94.9 |
| Yes | 318 | 14.0 | - | - | - | - | 144 | 9.0 | - | - | - | - | 67 | 5.1 |
| Total | 2264 | 100 | - | - | - | - | 1603 | 100 | - | - | - | - | 1321 | 100 |
| Walking time (continuous) | | | | | | | | | | | | | | |
| Less than 30 minutes | 759 | 30.7 | - | - | - | - | 600 | 35.7 | - | - | - | - | 351 | 26.6 |
| 30–60 minutes | 909 | 36.7 | - | - | - | - | 571 | 34.0 | - | - | - | - | 518 | 39.2 |
| 60–90 minutes | 405 | 16.4 | - | - | - | - | 259 | 15.4 | - | - | - | - | 226 | 17.1 |
| 90 minutes and more | 402 | 16.2 | - | - | - | - | 250 | 14.9 | - | - | - | - | 226 | 17.1 |
| Total | 2475 | 100 | - | - | - | - | 1680 | 100 | - | - | - | - | 1321 | 100 |
| Frequency participating to sports club | | | | | | | | | | | | | | |
| Never | 1763 | 77.0 | - | - | - | - | 1034 | 71.2 | - | - | - | - | 756 | 67.6 |
| A few times a year | 59 | 2.6 | - | - | - | - | 57 | 3.9 | - | - | - | - | 39 | 3.5 |
| 1 to 3 times a month | 52 | 2.3 | - | - | - | - | 52 | 3.6 | - | - | - | - | 51 | 4.6 |
| Once a week | 179 | 7.8 | - | - | - | - | 109 | 7.5 | - | - | - | - | 89 | 8.0 |
| 2 or 3 times a week | 190 | 8.3 | - | - | - | - | 151 | 10.4 | - | - | - | - | 105 | 9.3 |
| Almost everyday | 47 | 2.0 | - | - | - | - | 49 | 3.4 | - | - | - | - | 79 | 7.0 |
| Total | 2290 | 100 | - | - | - | - | 1452 | 100 | - | - | - | - | 1119 | 100 |

Abbreviation: JPY, Japanese Yen.
*We defined more than 3-time visitors as "participants."

participation also increased from 9.02 to 18.67 times per year (Supplementary Table 2). Equalized income gradually decreased during the follow-up period. More than half of the sample reported household income under 2 million Japanese Yen (about 20,000 US dollars) by 2013.

Systematic differences were observed in who participated in the salons (Supplementary Table 3). Women were more likely to participate than men, as well as those who also participated in sports clubs. By contrast, individuals from higher SES background (higher educational attainment and equalized income), higher depression scores, sedentary individuals, as well as smokers and regular drinkers were less likely to participate. That is, healthier and more sociable persons at baseline were more likely to participate in salons.

As shown in Table 2, the results of the multivariate model showed that salon participation frequency (continuous variable) was not significantly associated with lower risk of cognitive disability (odds ratio [OR] = 0.96, 95% CI: 0.91–1.01). The log-transformed frequency of salon participation was also associated with lower risk of incident cognitive disability (OR = 0.72, 95% CI: 0.54–0.98). In addition, the MSMs results indicated that frequency of salon participation was protectively associated with cognitive disability, even after adjusting for time-dependent covariates and attrition (OR = 0.73, 95% CI: 0.54–0.99) (Table 3). This model also showed that older age (75 years or older) is an important risk factor for incident cognitive decline (OR = 3.25, 95% CI: 1.86–5.69), while increasing walking time was protective (OR = 0.69, 95% CI: 0.55–0.87).

## 4. Discussion

To our knowledge, this is the first study to demonstrate that a community-based intervention encouraging social participations is effective for the prevention of incident cognitive disability. The association remained after controlling for time-dependent confounding and attrition during the follow-up period, using MSMs. The association between frequency of salon participation and incident cognitive disability appears to be statistically and clinically important. For example, salon participation was associated with the prevention of cognitive disability (OR = 0.73, 95% CI: 0.54–0.99, in model 2 of MSMs) as well as the daily walking (OR = 0.69, 95% CI: 0.55–0.87).

As shown in Supplementary Table 3, people who participated in the salons were more likely to be female, from lower socioeconomic background (lower educational attainment and equalized income), less likely to report depressive symptoms, less physically active, and less likely to smoke or drink. Therefore, an important challenge in

Table 2
Odds ratio and 95% CI for the risk of cognitive decline in the results of multivariate model

| | Multivariate model | | | |
|---|---|---|---|---|
| | Model 1 | | Model 2 | |
| | Odds (95% CI) | *P* | Odds (95% CI) | *P* |
| Frequency of participation | | | | |
| ln (X + 1) | 0.96 (0.91–1.01) | .12 | 0.72 (0.54–0.98) | .04 |
| Sex (Ref.: male) | | | | |
| Female | 0.93 (0.67–1.29) | .68 | 0.95 (0.68–1.31) | .73 |
| Educational attainment (Ref.: 9 years) | | | | |
| 10 years and over | 0.81 (0.60–1.10) | .17 | 0.81 (0.59–1.09) | .16 |
| Age (Ref: 65–74 years) | | | | |
| 75 years and over | 6.21 (4.14–9.32) | <.01 | 6.23 (4.15–9.34) | <.01 |
| Equivalent income (Ref.: <2 million JPY) | | | | |
| 2 million JPY and more | 0.97 (0.68–1.40) | .88 | 0.97 (0.68–1.40) | .88 |
| Number of comorbidities | | | | |
| 0–3 | 0.93 (0.73–1.17) | .52 | 0.92 (0.73–1.17) | .51 |
| Depressive symptoms (Ref.: ≤4 points) | | | | |
| ≥5 points | 1.52 (1.09–2.13) | .01 | 1.51 (1.08–2.12) | .02 |
| Low cognition score (Ref.: 0 point) | | | | |
| ≥1 point | 2.36 (1.72–3.24) | <.01 | 2.36 (1.71–3.24) | <.01 |
| Current drinking | | | | |
| Yes (Ref.: no) | 0.75 (0.51–1.09) | .13 | 0.75 (0.51–1.09) | .13 |
| Current smoking | | | | |
| Yes (Ref.: no) | 1.52 (0.98–2.33) | .06 | 1.51 (0.98–2.33) | .06 |
| Walking time | | | | |
| 1: <30 minutes to 4: ≥90 minutes | 0.75 (0.63–0.89) | <.01 | 0.75 (0.63–0.89) | <.01 |
| Frequency participating in sports club | | | | |
| 1: a few times to 6: almost everyday | 0.98 (0.93–1.05) | .63 | 0.98 (0.93–1.05) | .63 |
| Cons. | 0.01 (0.01–0.02) | <.01 | 0.01 (0.01–0.02) | <.01 |

Abbreviations: CI, confidence interval; Odds, odds ratio; Ref., reference; JPY = Japanese Yen.

Table 3
Odds ratio and 95% CI for the risk of cognitive decline in the results of marginal structural model

| | Marginal structural models | | | |
|---|---|---|---|---|
| | Model 1 | | Model 2 | |
| | Odds (95% CI) | *P* | Odds (95% CI) | *P* |
| Frequency of participation | | | | |
| ln (X + 1) | 0.72 (0.53–0.98) | .04 | 0.73 (0.54–0.99) | .04 |
| Time-invariant covariates | | | | |
| Sex (Ref.: male) | | | | |
| Female | 1.08 (0.78–1.51) | .64 | 1.04 (0.74–1.46) | .82 |
| Educational attainment (Ref.: 9 years) | | | | |
| 10 years and over | 0.82 (0.59–1.13) | .22 | 0.86 (0.62–1.18) | .36 |
| Time-variant covariates at baseline | | | | |
| Age (Ref: 65–74 years) | | | | |
| 75 years and over | 4.89 (3.5–6.82) | <.01 | 2.30 (1.54–3.45) | <.01 |
| Equivalent income (Ref.: <2 million JPY) | | | | |
| 2 million JPY and more | 0.90 (0.63–1.3) | .58 | 0.92 (0.58–1.45) | .71 |
| Number of comorbidities | | | | |
| 0–3 | 1.01 (0.77–1.31) | .97 | 1.06 (0.74–1.52) | .74 |
| Depressive symptoms (Ref.: ≤4 points) | | | | |
| ≥5 points | 1.23 (0.83–1.80) | .30 | 0.84 (0.49–1.45) | .53 |
| Low cognition score (Ref.: 0 point) | | | | |
| ≥1 points | 1.67 (1.22–2.29) | <.01 | 1.04 (0.68–1.59) | .85 |
| Current drinking | | | | |
| Yes (Ref.: no) | 0.90 (0.63–1.30) | .58 | 1.42 (0.78–2.60) | .25 |
| Current smoking | | | | |
| Yes (Ref.: no) | 1.26 (0.82–1.93) | .29 | 0.86 (0.31–2.41) | .78 |
| Walking time (continuous) | | | | |
| 1: <30 minutes to 4: ≥90 minutes | 0.87 (0.73–1.02) | .09 | 1.16 (0.93–1.44) | .19 |
| Frequency participating in sports club | | | | |
| 1: a few times to 6: almost everyday | 1.09 (0.95–1.25) | .21 | 1.10 (0.96–1.26) | .19 |
| Time-variant covariates during follow-up term | | | | |
| Age (Ref: 65–74 years) | | | | |
| 75 years and over | | | 3.25 (1.86–5.69) | <.01 |
| Equivalent income (Ref.: <2 million JPY) | | | | |
| 2 million JPY and more | | | 1.01 (0.64–1.61) | .96 |
| Number of comorbidities | | | | |
| 0–3 | | | 0.91 (0.67–1.25) | .58 |
| Depressive symptoms (Ref.: ≤4 points) | | | | |
| ≥5 points | | | 1.68 (1.03–2.74) | .04 |
| Low cognition score (Ref.: 0 point) | | | | |
| ≥1 point | | | 2.26 (1.44–3.53) | <.01 |
| Current drinking | | | | |
| Yes (Ref.: no) | | | 0.61 (0.31–1.18) | .14 |
| Current smoking | | | | |
| Yes (Ref.: no) | | | 1.68 (0.59–4.81) | .33 |
| Walking time | | | | |
| 1: <30 minutes to 4: ≥90 minutes | | | 0.69 (0.55–0.87) | <.01 |
| Frequency participating in sports club | | | | |
| 1: a few times to 6: almost everyday | | | 0.94 (0.88–1.01) | .11 |
| Cons. | 0.01 (0.01–0.02) | <.01 | 0.01 (0.01–0.02) | <.01 |

Abbreviations: CI, confidence interval; Odds, odds ratio; Ref., reference; JPY, Japanese Yen.

scaling up the salon intervention to the general population is to understand how to design the programs in such a way as to appeal to broader segments of the population. One possibility is to adopt a social marketing approach consisting of outreach, education, and networking efforts within a targeted area [32].

There are plausible mechanisms by which salon participation might assist in the prevention of cognitive disability. In the interim appraisal that was conducted 2 years after the start of this intervention program, the salon participants reported that they had joined other local organizations such as sports or hobby clubs. They also reported increased social support (both receipt and provision) after joining the community salons [33]. That is, participants appear to have expanded their social network through salon participation, which in turn helped to preserve cognitive function [34]. In addition, some salon activities included light physical activity (e.g., stretching exercises) and other activities such as handicrafts that might have helped in the maintenance of cognitive functions [35,36]. Salon participation may also have helped to reduce the risk of depression (itself a risk factor for cognitive decline) by boosting people's sense of purpose in life [37].

In an interim appraisal of 100 salon participants conducted 3 years after the start of the salon project, the participants were more likely to perform well on cognitive tests such as the three-words delayed recall test and word fluency test [38]. However, this study did not compare the cognitive function of nonparticipants or control for confounding factors. In a previous analysis from the same intervention study—based on 5 years of follow-up—we reported that salon participation was associated with lower risk of functional disability [9]. However, we were able to control for baseline covariates only and did not take account of time-varying risk factors during follow-up.

In the United States, the Experience Corps was an intervention program that recruited retired seniors in Baltimore to serve as a volunteer teacher's aides in local schools. The program was designed to support the academic success of children and to promote the health of older volunteers by enhancing their physical, social, and cognitive functioning [39]. Their intervention found that brain cortical and hippocampal volumes were increased among male participants in the treatment group during 2-year follow-up [40]. In another intervention study conducted in Oregon, an online conversation program was provided to 83 seniors living in retirement communities or senior centers [41]. The intervention sought to increase social interaction among seniors through online conversations with trained interviewers for 6 weeks. The intervention group showed improved semantic and phonemic fluency compared with the control group. Multimodal interventions incorporating social engagement are therefore likely to be effective for the prevention of cognitive decline; however, evidences remain extremely scarce [42]. Our study suggests that operating community salons that encourage cognitive activities, social interactions, and light physical activity among older participants may be effective for preventing cognitive decline.

An important strength of this study was to reduce estimation bias caused by time-dependent variables. Although about 50% subjects dropped out from our cohort during 7-year observation terms, the proportion of salon participants increased year by year (Table 1). It causes a selection bias: Healthy subjects (time $t-1$) are more likely to participate in the community salons (time $t-1$ and t), thereby maintaining their health conditions as well as the lower risk of cognitive decline (time $t+1$). The MSMs created a pseudo-population in which the exposure is independent of measured confounders [29]. An additional strength is our objective measures of the frequency of salon participation as well as cognitive disability based on inhome assessment.

A limitation of our study is the response rate at baseline survey that was 48.5%. This may limit the generalizability of our findings, although our response rate is comparable to other surveys of community-based residents [43]. The demographic profile of our analytic sample is also similar to the independent census data for Taketoyo residents aged 65 years or older (Supplementary Table 4). In addition, the follow-up rates for the two follow-up surveys were quiet high (78.0% and 90.1%, for respectively). On the other hand, because we did not perform a detailed assessment of early presymptoms of cognitive decline, we cannot rule out the possibility of reverse causation, for example, that individuals with early symptoms of cognitive decline were less likely to participate in salon activities. In addition, there is a possibility that differential misclassification occurred in our assessment of outcomes because older persons who have mild cognitive impairment may tend to avoid medical diagnosis. We also lacked information about incident diabetes mellitus which may have affected both salon participation and cognitive decline. Our outcome variable (an 8-point scale of cognitive decline obtained from in-home assessment) does not amount to a clinically based diagnosis of dementia, although it has been demonstrated to be highly correlated with MMSE scores and CDR scores in previous validation studies [13,14].

In conclusion, opening and operating community-based salons may be an effective intervention for the prevention of cognitive decline in the older Japanese population. In future studies, the assessment of cognitive decline under the long-term care insurance scheme should be validated against *The Fourth Edition of Diagnostic and Statistical Manual of Mental Disorders* (*DSM-IV*) criteria. We need to also understand what sorts of activities and programs (e.g., those involving light physical activity vs. purely intellectual activities) are most effective in maintaining cognitive function.

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There are no conflicts of interest to declare.

This study has not been submitted elsewhere nor is it being considered elsewhere for publication.

## Supplementary data

Supplementary data related to this article can be found at http://dx.doi.org/10.1016/j.trci.2016.11.003.

### RESEARCH IN CONTEXT

1. Systematic review: The authors reviewed the published literature using PubMed. Although many observational studies have shown that social participation in older people is protectively associated with incident cognitive decline, no previous studies have demonstrated that social participation is effective for prevention of incident cognitive disability in the context of a community intervention study.
2. Interpretation: Our findings suggest that operating a "community salon" program can contribute to the prevention of the onset of cognitive decline through increasing the social participation of the participants.
3. Future directions: Further studies are needed to understand what sorts of activities and programs (e.g., those involving light physical activity vs. purely intellectual activities) are most effective in maintaining cognitive function.

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# Retirement and mental health: dose social participation mitigate the association? A fixed-effects longitudinal analysis

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## Abstract

**Background:** Empirical evidence investigating heterogeneous impact of retirement on mental health depending on social backgrounds is lacking, especially among older adults.

**Methods:** We examined the impact of changes in working status on changes in mental health using Japanese community-dwelling adults aged ≥65 years participating in the Japan Gerontological Evaluation Study between 2010 and 2013 ($N$ = 62,438). Between-waves changes in working status ("Kept working", "Retired", "Started work", or "Continuously retired") were used to predict changes in depressive symptoms measured by the Geriatric Depression Scale. First-difference regression models were stratified by gender, controlling for changes in time-varying confounding actors including equivalised household income, marital status, instrumental activities of daily living, incidence of serious illnesses and family caregiving. We then examined the interactions between changes in working status and occupational class, changes in marital status, and post-retirement social participation.

**Results:** Participants who transitioned to retirement reported significantly increased depressive symptoms ($\beta$ = 0.33, 95% CI: 0.21–0.45 for men, and $\beta$ = 0.29, 95% CI: 0.13–0.45 for women) compared to those who kept working. Men who were continuously retired reported increased depressive symptoms ($\beta$ = 0.13, 95% CI: 0.05–0.20), whereas males who started work reported decreased depressive symptoms ($\beta$ = −0.20, 95% CI: -0.38–-0.02). Men from lower occupational class (compared to men from higher class) reported more increase in depressive symptoms when continuously retired ($\beta$ = −0.16, 95% CI: -0.25–-0.08). Those reporting recreational social participation after retirement appeared to be less influenced by transition to retirement.

**Conclusions:** Retirement may increase depressive symptoms among Japanese older adults, particularly men from lower occupational class backgrounds. Encouraging recreational social participation may mitigate the adverse effects of retirement on mental health of Japanese older men.

**Keywords:** Retirement, Mental health, Social participation, Fixed-effects, Japan, Older adults

## Background

Population aging is occurring worldwide and an increasing number of older adults are engaging in work in late life. Japan is experiencing the fastest rate of population ageing in the world. In Japan, the labor force participation rate among people aged 65 years or older was 21.3% in 2014, which ranked 7th among all OECD countries [1], and the population of older workers increased from 4.8 million in 2004 to 6.8 million in 2014 [2]. To maintain both the quality of life of each older individual and the financial sustainability of healthcare and long-term care systems, preventing physical and mental impairments is the key goals of public health measures for the aged [3]. Specifically, the maintenance of mental health counts as it could be strongly affected by the stressful life events that are likely to happen in older ages. Among those life events, in this study, we focused on retirement, the major events that most people experience.

Transition from work to retirement is a major life event in most people's lives. However, there is controversy about

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the impact of retirement on mental health [4]. While many studies done in European countries suggest that retirement is beneficial to mental health [5–8] or has no health impact [9–11], results from some studies in Asian populations indicate the opposite, i.e. the transition to retirement is detrimental to mental health [12, 13]. These geographic differences in the association between retirement and mental health may be partly due to the differences in employment conditions and how retirements are culturally perceived by people in each region.

There are several potential mechanisms through which retirement may have adverse impact on mental health. Above and beyond the immediate reduction in income, social role theory posits that retirement results in the loss of many non-financial benefits of work, such as opportunities for health-promoting social contacts and access to social support that can buffer the adverse effects of stress [14–16]. Activity theory also posits that the transition to retirement results in reduced wellbeing for individuals for whom work provided meaning in life [17].

Moreover, the impact of retirement on mental health may not be universal and could vary between individuals depending on his/her occupational class or social situation (e.g. the presence of social support in the home environment from a marital partner) [4, 6, 18, 19]. In addition, social participation after retirement may compensate for any adverse effect of retirement stemming from loss of sense of meaning in life, which has been suggested to be linked to mental health [4, 12, 13, 20]. However, the potential impact of other types of social participation such as hobby clubs remains unknown.

In the present study, we sought to evaluate [1] the association between changes in working status and changes in depressive symptoms in a cohort of community-dwelling Japanese older adults, [2] whether the association can be mediated by changes in social contacts and social support, [3] the interaction between changes in working status and occupational class, changes in marital status, and social participation after retirement.

## Methods

### Data

We used data from the 2010 and 2013 waves of Japan Gerontological Evaluation Study (JAGES) project. Detailed descriptions of JAGES are available elsewhere [21, 22]. In brief, the JAGES conducted postal surveys targeting physically and cognitively independent community-dwelling older adults aged ≥65 years in Japan. The baseline survey was conducted in 2010 and we sampled 112,123 older adults residing in 31 municipalities in Japan (Response rate = 66.3%). The second survey was conducted in 2013 in which we sampled 137,736 older adults residing in 30 municipalities in Japan (Response rate = 71.1%). Twenty-four municipalities participated in both years. Sixty-two thousand four hundred thirty-eight participants responded to both surveys.

The 2013 wave of JAGES dataset consisted of five modules. The question inquiring about participants' sense of meaning in life was measured only in one of the five modules, which was only mailed to one-fifth of the total sample. We used respondents to this module ($N$ = 12,487) to test the interaction between retirement and two types of social participation, as well as whether a sense of meaning in life mediated this relationship. All variables used in this study were self-reported. Ethical approval of this study was obtained from the ethics review board of the University of Tokyo Medical School.

### Measurement

#### *Changes in depressive symptoms*

Depressive symptoms were measured at both waves using the validated Japanese short version of the Geriatric Depression Scale (GDS-15) [23] Total scores could range from 0 to15, where higher scores indicate more depressive symptoms. GDS-15 has been commonly used as a screening tool for depression among older adults with a cutoff of 5 or above to indicate clinical depression [23, 24]. We used changes in total GDS-15 scores from 2010 to 2013 as the continuous outcome variable in this study.

#### *Changes in working status*

Information on working status (either "currently working" or "retired and not currently working") was collected in both 2010 and 2013. Thus, no one among the subjects was considered to be partially retired. We categorized subjects as "Kept working (working/working)", "Retired (working → not working)", "Started work (not working → working)", or "Continuously retired (not working → not working)" based on patterns of changes in working status from 2010 to 2013.

#### *Changes in social contacts and social support*

We used two variables to tap changes in social contacts. First, we asked the number of friends and acquaintances the subjects met in the past month at both waves, and then, changes in the answers were modeled as a continuous variable. Second, we asked subjects "What is the relationship between you and the person you often meet?" At each wave, we focused on individuals who maintained social contacts with colleagues in the workplace. As for social support, we collected information at both waves on whether or not subjects received emotional ('Do you have anyone who will listen to you when you have worries and complains?') and instrumental

('Do you have anyone who can take care of you when you are sick in bed for a few days?') social support from their friends. Between-wave changes in perception of social supports from friends were made into a categorical variable.

#### *Occupational class*

The occupation that subjects had engaged in for the longest period of time was asked at both waves. Since it is unlikely that the longest job of older people aged 65 years or older change within 3 years (from 2010 to 2013), we treated this as a time-invariant variable. Professionals and managers were categorized as "higher occupational class". Clerical support workers, service and sales workers, craft and related trades workers, skilled agricultural, forestry, and fishery workers, and others were categorized as "lower occupational class".

#### *Social participation/a sense of meaning in life*

Participation in different clubs/groups was asked at both waves. We defined "social participation with roles" as participation in volunteering clubs or neighborhood councils. We defined "recreational social participation" as participation in sports organizations, hobby clubs, or older adults clubs. A sense of meaning in life was measured in 2013 by asking "Do you feel a sense of meaning in life" (yes/no).

#### *Covariates*

Changes in equivalised household income, marital status, instrumental activities of daily living, incidence of serious illnesses and family caregiving were treated as time-varying covariates. These variables are potential confounding factors, which could have influence on both retirement decision and mental health. Household income at each wave was equivalised in order to adjust for the number of members within households. Marital status of subjects at each wave was defined either "married" or "non-married (including being widowed, divorced, or single)". Subjects' instrumental activities of daily living (IADL) were measured using the Tokyo Metropolitan Institute of Gerontology (TMIG) Index of Competence as a broad indicator of physical health. TMIG Index of Competence was validated using the data of Japanese older community residents [25] and its total scores could range from 0 to13, where higher scores indicates more independency in daily livings. Incidence of serious illnesses and need for caregiving was assessed in 2013 by administering a checklist for events occurring in the past year.

### Statistical analysis

As for missing values in some variables used in the analysis, we executed multiple imputation by chained equation method to create 100 data sets without missing value. We used the FCS statement of the MI procedure in SAS version 9.3 to create imputed data sets. Then, we excluded those who indicated depression in 2010 (GDS score of 5 or more as described below), those who never worked, and those who answered inconsistently to the question about the occupational class at the two time points. We ran all models using each data set and, using MIANALYZE procedure, we obtained the final estimates combining 100 estimates derived from 100 analyses.

We used first-difference models stratified by gender to investigate the main effects of changes in working status on changes in depressive symptoms with full-sample data (**Model 1**). The first-difference models tool the changes in all variables in the models, which enable to can control for all observed and unobserved time-invariant confounders (e.g. age and education) by cancelling out the estimates of time-invariant confounders. In model 2, we added variables regarding changes in social contacts and social support to test whether these changes could explain the association between changes in working status and changes in depressive symptoms. In model 3, we added an interaction term between changes in working status and occupational class. In model 4, we added an interaction term between transition to retirement and changes in marital status.

In the analysis involving a sense of meaning in life, we analyzed only one of the five sub-samples of JAGES data because it is the only module that contains information on participants' sense of purpose in life. In model 5, we used the first-difference model again to test the main effect of changes in working status on depressive symptoms among this sub-population. In the model 6, we added the interaction term between transition to retirement and two types of social participation. Lastly, in model 7, we added sense of meaning in life as a covariate to check whether it mediated the interaction between retirement and social participation.

We stratified data by gender instead of modeling gender interaction effects to make it easy to interpret our results and prevent the problems of multicollinearity on interaction terms.

## Results

Descriptive analysis showed that subjects in the full-sample dataset and the sub-sample dataset were broadly similar in terms of their background characteristics (Table 1). Nearly a quarter of male and female subjects indicated depressive symptoms at baseline.

For men, we observed significant increases in GDS score among those who transitioned to retirement ($\beta$ = 0.33, 95% CI: 0.21–0.45) and those who were continuously retired ($\beta$ = 0.13, 95% CI: 0.05–0.20) compared to those who kept working (**Model 1** in Table 2). For

**Table 1** Demographic Characteristics of Subjects from Full-dataset and Sub-dataset by Gender

| | Full-sample | | | | | Sub-sample | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| | Men | | Women | | | Men | Women | | | |
| | *N* = 28,868 | | *N* = 33,569 | | | *N* = 5866 | *N* = 6621 | | | |
| | Mean | SD | Mean | SD | *p*-value | Mean | SD | Mean | SD | *p*-value |
| Age in 2010 | 72.9 | 5.6 | 73.1 | 5.6 | <.0001 | 72.9 | 5.6 | 73.1 | 5.6 | 0.02 |
| GDS-15 score change (2010 → 2013) | 0.1 | 2.4 | 0.03 | 2.4 | 0.0005 | 0.1 | 2.5 | −0.01 | 2.3 | 0.02 |
| | n | % | n | % | | n | % | n | % | |
| Working status in 2010 | | | | | | | | | | |
| Working status | 8095 | 30.3 | 5287 | 19.2 | <.0001 | 1687 | 31.1 | 1058 | 19.5 | <.0001 |
| Not working | 17,530 | 65.6 | 16,556 | 60 | | 3510 | 64.7 | 3231 | 59.6 | |
| Never worked | 1087 | 4.1 | 5730 | 20.8 | | 232 | 4.3 | 1133 | 20.9 | |
| Missing | 2156 | | 5996 | | | 437 | | 1199 | | |
| Depression in 2010 | | | | | | | | | | |
| Having depression in 2010 | 6353 | 25.5 | 6816 | 25.5 | 0.84 | 1340 | 26.4 | 1354 | 25.7 | 0.4 |
| Not having depression in 2010 | 18,602 | 74.5 | 19,878 | 74.5 | | 3730 | 73.6 | 3913 | 74.3 | |
| Missing | 3913 | | 6875 | | | 796 | | 1354 | | |
| Household income in 2010 (10,000 Japanese Yen) | | | | | | | | | | |
| < 300 | 12,129 | 45.6 | 15,051 | 54 | <.0001 | 2429 | 44.8 | 2921 | 53.2 | <.0001 |
| 300–600 | 9759 | 36.7 | 8427 | 30.2 | | 2027 | 37.4 | 1676 | 30.5 | |
| < 600 | 4711 | 17.7 | 4393 | 15.8 | | 969 | 17.9 | 897 | 16.3 | |
| Missing | 2269 | | 5698 | | | 441 | | 1127 | | |
| Marital status in 2010 | | | | | | | | | | |
| Married | 24,608 | 87.9 | 19,790 | 61.3 | <.0001 | 5007 | 88 | 3894 | 61.1 | <.0001 |
| Non-married | 3374 | 12.1 | 12,520 | 38.8 | | 685 | 12 | 2484 | 39 | |
| Missing | 886 | | 1259 | | | 174 | | 243 | | |
| IADL in 2010 | | | | | | | | | | |
| 13 | 9401 | 36.6 | 15,590 | 53.1 | <.0001 | 1882 | 36 | 3052 | 52.7 | <.0001 |
| 10, 12 | 12,959 | 50.4 | 11,548 | 39.3 | | 2642 | 50.5 | 2313 | 40 | |
| ≤ 9 | 3351 | 13 | 2250 | 7.7 | | 710 | 13.6 | 423 | 7.3 | |
| Missing | 3157 | | 4181 | | | 632 | | 833 | | |
| The number of friends subjects met in the past 1 month in 2010 | | | | | | | | | | |
| 0 | 2041 | 7.6 | 1065 | 3.5 | <.0001 | 419 | 7.7 | 198 | 3.3 | <.0001 |
| 1, 5 | 11,426 | 42.5 | 12,589 | 40.8 | | 2302 | 42.3 | 2495 | 41.1 | |
| 6- | 13,396 | 49.9 | 17,236 | 55.8 | | 2728 | 50.1 | 3385 | 55.7 | |
| Missing | 2005 | | 2679 | | | 417 | | 543 | | |
| Colleague as a friend who subjects met frequently in 2010 | | | | | | | | | | |
| Having colleague as a friend | 11,439 | 42.4 | 8027 | 25.4 | <.0001 | 2365 | 43.2 | 1522 | 24.4 | <.0001 |
| Not having colleague as a friend | 15,563 | 57.6 | 23,625 | 74.6 | | 3116 | 56.9 | 4713 | 75.6 | |
| Missing | 1866 | | 1917 | | | 385 | | 386 | | |
| Emotional social support from friends in 2010 | | | | | | | | | | |
| Having emotional social support | 7233 | 26.6 | 15,156 | 47.9 | <.0001 | 1425 | 25.7 | 2974 | 47.9 | <.0001 |
| Not having emotional social support | 19,950 | 73.4 | 16,482 | 52.1 | | 4112 | 74.3 | 3234 | 52.1 | |
| Missing | 1685 | | 1931 | | | 329 | | 413 | | |

**Table 1** Demographic Characteristics of Subjects from Full-dataset and Sub-dataset by Gender *(Continued)*

| | | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| Instrumental social support from friends in 2010 | | | | | | | | | | |
| Having instrumental social support | 632 | 2.3 | 2144 | 6.8 | <.0001 | 114 | 2.1 | 400 | 6.4 | <.0001 |
| Not having instrumental social support | 26,793 | 97.7 | 29,539 | 93.2 | | 5454 | 97.8 | 5844 | 93.6 | |
| Missing | 1444 | | 1886 | | | 298 | | 377 | | |
| Occupational class | | | | | | | | | | |
| Higher | 9265 | 36.2 | 3031 | 12.6 | <.0001 | 1879 | 36.5 | 605 | 12.9 | <.0001 |
| Lower | 16,325 | 63.8 | 20,940 | 87.4 | | 3268 | 63.5 | 4083 | 87.1 | |
| missing | 3153 | | 6394 | | | 698 | | 1326 | | |
| Social participation with roles in 2013 | | | | | | | | | | |
| Participated in social activities with roles | 6536 | 26.3 | 7361 | 27.8 | 0.001 | 1292 | 25.6 | 1461 | 28 | 0.005 |
| Not participated in social activities with roles | 18,346 | 73.7 | 19,132 | 72.2 | | 3764 | 74.5 | 3754 | 72 | |
| missing | 3986 | | 7076 | | | 810 | | 1406 | | |
| Recreational social participation in 2013 | | | | | | | | | | |
| Participated in recreational social activities | 11,714 | 46.5 | 16,067 | 58 | <.0001 | 2367 | 46.3 | 3167 | 57.9 | <.0001 |
| Not participated in recreational social activities | 13,458 | 53.5 | 11,657 | 42.1 | | 2751 | 53.8 | 2304 | 42.1 | |
| missing | 3696 | | 5845 | | | 748 | | 1150 | | |
| A sense of meaning in life in 2013 | | | | | | | | | | |
| Having a sense of meaning in life in 2013 | | | | | | 4666 | 86.2 | 5036 | 84.2 | 0.003 |
| Not having a sense of meaning in life in 2013 | | | | | | 748 | 13.8 | 945 | 15.8 | |
| missing | | | | | | 452 | | 640 | | |

*GDS-15* the short version of the Geriatric Depression Scale (ranging from 0 to 15, higher score indicates more depressive symptoms). Depression in 2010 was defined by GDS-15 score of 5 or above. Non-married includes being divorced, widowed, and single. *IADL* Instrumental activities of daily living (ranging from 0 to 13, higher score indicates more independency in daily livings). Higher occupational class = professionals and engineering, managers. Lower occupational class = clerical support workers, service and sales workers, craft and related trades workers, skilled agricultural, forestry, and fishery workers, others. Social participation with roles = volunteering clubs, neighborhood councils. Recreational social participation = sports organizations, hobby clubs, or older adults clubs. Full-sample: subjects from all of the five modules of JAGES. Sub-sample: subjects from only one of the five modules from JAGES that includes information on a sense of meaning in life. We used t-test for age and changes in GDS score and chi-square test for other variables to calculate *p*-values

women, we observed significant increases in GDS score among those who transitioned to retirement (β = 0.29, 95% CI: 0.13–0.45) compared to those who kept working. After including changes in social contacts and social support in **Model 2**, significant associations observed in **Model 1** were overall unchanged and remained statistically significant.

When including interaction terms (Additional file 1: Table S1), we did not find any statistically clear evidence on that the association between changes in working status and GDS scores were altered by occupational class (**Model 3**), though there were tendency that the relative increase in GDS score among men who were continuously retired were smaller among high occupational classes than lower occupational class (β = −0.08, 95% CI: -0.22–0.07) (Fig. 1). There was no significant interaction between changes in working status and occupational class among women. We observed no significant interactions between transition to retirement and changes in marital status among men and women (**Model 4**).

Analyses using the sub-dataset showed consistent results: significant increases in GDS score among men who transitioned to retirement (β = 0.40, 95% CI: 0.12–0.68) as well as men who were continuously retired (β = 0.29, 95% CI: 0.11–0.47) compared to men who kept working (**Model 5** in Additional file 2: Table S2). Women who transitioned to retirement reported marginally significant increases in GDS scores (β = 0.32, 95% CI: -0.04–0.68, $p$ = 0.09).

Although statistical evidence was weak but there was a trend among men and women that increase in depressive symptoms was less among those who participated in recreational social participation (β = −0.22, 95% CI: -0.72–0.29 for men, and β = −0.28, 95% CI: -0.92–0.37 for women) (**Model 6** in Additional file 3: Table S3, Fig. 2). After additional adjustment for whether or not subjects have a sense of meaning in life (**Model 7** in Additional file 3: Table S3), the magnitude of interaction effect between transition to retirement and recreational social participation among men considerably reduced (β = −0.05, 95% CI: -0.54–0.44), whereas it did not change among women (β = −0.27, 95% CI: -0.89–0.35). Notably, having a sense of meaning in life was strongly associated with increased depressive symptoms among men and women (β = −2.66, 95% CI: -3.00–-2.32 for men, and β = −1.84, 95% CI: -2.11-1.57 for women).

**Table 2** Multiple linear regression of Changes in GDS score on change in working status by gender (full-sample)

| Independent variables | Dependent Variable: Changes in GDS-15 score from 2010 to 2013 | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Model 1 | | | | Model 2 | | | |
| | Men | | Women | | Men | | Women | |
| | β coefficient (95%CI) | | | | | | | |
| Changes in working status (2010–2013) | | | | | | | | |
| Kept working | Ref. | Ref. | Ref. | Ref. | Ref. | Ref. | Ref. | Ref. |
| Retired | 0.33 | (0.21, 0.45) | 0.29 | (0.13,0.45) | 0.33 | (0.21,0.45) | 0.28 | (0.12,0.44) |
| Started work | −0.20 | (−0.38, −0.02) | 0.10 | (−0.13,0.33) | −0.21 | (−0.39,-0.03) | 0.10 | (−0.13,0.33) |
| Continuously retired | 0.13 | (0.05, 0.20) | 0.05 | (−0.04,0.14) | 0.10 | (0.02,0.18) | 0.05 | (−0.03,0.14) |
| Changes in Equivalized household income (10,000 JPY) | −0.0004 | (−0.0007, −0.0002) | −0.0005 | (−0.0007, -0.0002) | −0.0004 | (−0.0007, -0.0002) | −0.0004 | (−0.0007, -0.0002) |
| Changes in IADL | −0.24 | (−0.26, −0.21) | −0.28 | (−0.31,-0.25) | −0.23 | (−0.25,-0.2) | −0.26 | (−0.3,-0.23) |
| Incidence of stressful life events in the past 1 year | | | | | | | | |
| Serious illnesses | 0.59 | (0.47, 0.70) | 0.70 | (0.54,0.87) | 0.58 | (0.46,0.7) | 0.70 | (0.53,0.87) |
| Started family caregiving | 0.54 | (0.35, 0.73) | 0.55 | (0.38,0.72) | 0.54 | (0.34,0.73) | 0.55 | (0.38,0.72) |
| Changes in marital status | | | | | | | | |
| Married – Married | Ref. | Ref. | Ref. | Ref. | Ref. | Ref. | Ref. | Ref. |
| Married - Not married | 0.48 | (0.26, 0.70) | 0.15 | (−0.02,0.32) | 0.48 | (0.26,0.7) | 0.16 | (−0.01,0.33) |
| Not married – Married | −0.04 | (−0.45, 0.36) | 0.54 | (0.13,0.95) | −0.04 | (−0.44,0.37) | 0.51 | (0.11,0.92) |
| Not married - Not married | 0.20 | (0.08,0.31) | 0.11 | (0.04,0.18) | 0.21 | (0.09,0.33) | 0.11 | (0.04,0.18) |
| The number of friends subjects met in the past 1 month | | | | | −0.02 | (−0.03,-0.01) | −0.03 | (−0.04,-0.02) |
| Colleagues as a friend who subjects met frequently | | | | | | | | |
| Continuously having colleagues as a friend | | | | | Ref. | Ref. | Ref. | Ref. |
| Lost colleagues as a friend | | | | | 0.01 | (−0.09,0.11) | 0.06 | (−0.06,0.19) |
| Continuously not having colleagues as a friend | | | | | 0.08 | (0.004,0.16) | −0.003 | (−0.09,0.09) |
| Newly had colleagues as a friend | | | | | 0.04 | (−0.08,0.15) | −0.03 | (−0.17,0.1) |
| Emotional support from friends | | | | | | | | |
| Continuously having emotinal social support from friends | | | | | Ref. | Ref. | Ref. | Ref. |
| Lost emotional social support from friends | | | | | 0.16 | (0.04,0.28) | 0.16 | (0.04,0.27) |
| Continuously not having emotional social support from friends | | | | | 0.21 | (0.12,0.3) | 0.12 | (0.04,0.2) |
| Newly had emotional support from friends | | | | | 0.08 | (−0.03,0.19) | −0.01 | (−0.12,0.1) |
| Instrumental support from friends | | | | | | | | |
| Continuously having instrumental social support from friends | | | | | Ref. | Ref. | Ref. | Ref. |
| Lost instrumental social support from friends | | | | | 0.23 | (−0.2,0.66) | 0.19 | (−0.05,0.42) |
| Continuously not having instrumental social support from friends | | | | | 0.05 | (−0.32,0.42) | 0.06 | (−0.13,0.25) |
| Newly had instrumental support from friends | | | | | −0.01 | (−0.42,0.41) | −0.08 | (−0.33,0.16) |

*GDS-15* the short version of the Geriatric Depression Scale (ranging from 0 to 15, higher score indicates more depressive symptoms). Subjects are those who did not show depression at baseline (GDS score < 5). *IADL* Instrumental activities of daily living (ranging from 0 to 13, higher score indicates more independency in daily livings). Non-married includes being divorced, widowed, and single. Data of full-sample (all of the 5 sub-versions of JAGES datasets) was used for these analysis

## Discussion

We summarize our key findings. First, the transition to retirement in both men and women and sustained retirement in men were associated with increased depressive symptoms among Japanese older adults. Second, the associations between retirement and increased depressive symptoms remained statistically significant after additionally adjusting for changes in social contacts and social support, suggesting that they did not mediate the impact of retirement on mental health at least among Japanese older adults. Third, there was a statistically non-significant trend that men from higher occupational class backgrounds appeared to be less influenced by sustained retirement, whereas we did not find such a trend

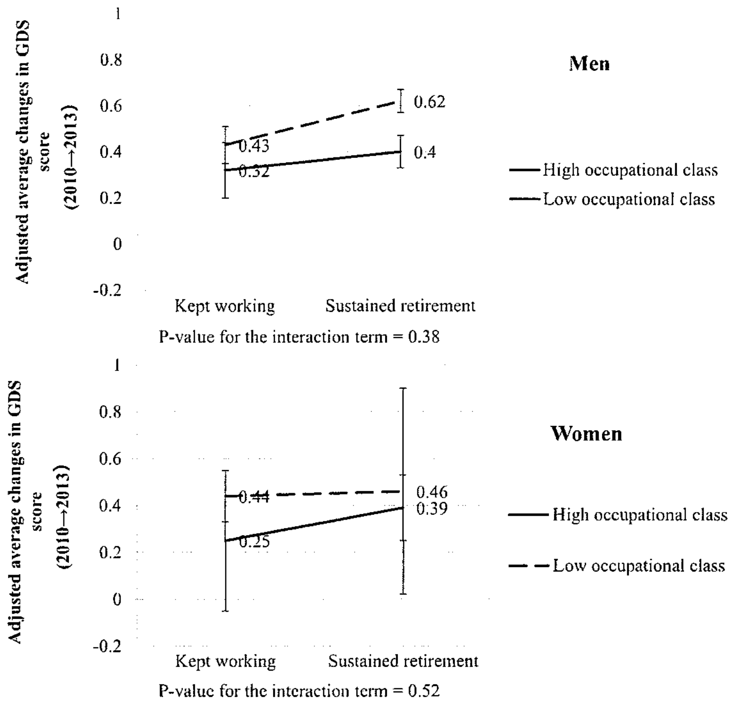


**Fig. 1** Adjusted average changes in GDS score by changes in working status and occupational class.
GDS: the short version of the geriatric depression scale (ranging from 0 to 15, higher score indicates more depressive symptoms). Adjusted for changes in time varying confounding factors including equivalised houshold income, IADL limitation, marital status, stressful life events, social relationships. Subjects are those who did not show depression at baseline (GDS score < 5). Occupational status: 1 = high (professionals and engineering, managers), 0 = low (clerical support workers, service and sales workers, craft and related trades workers, skilled agricultural, forestry, and fishery workers, others)

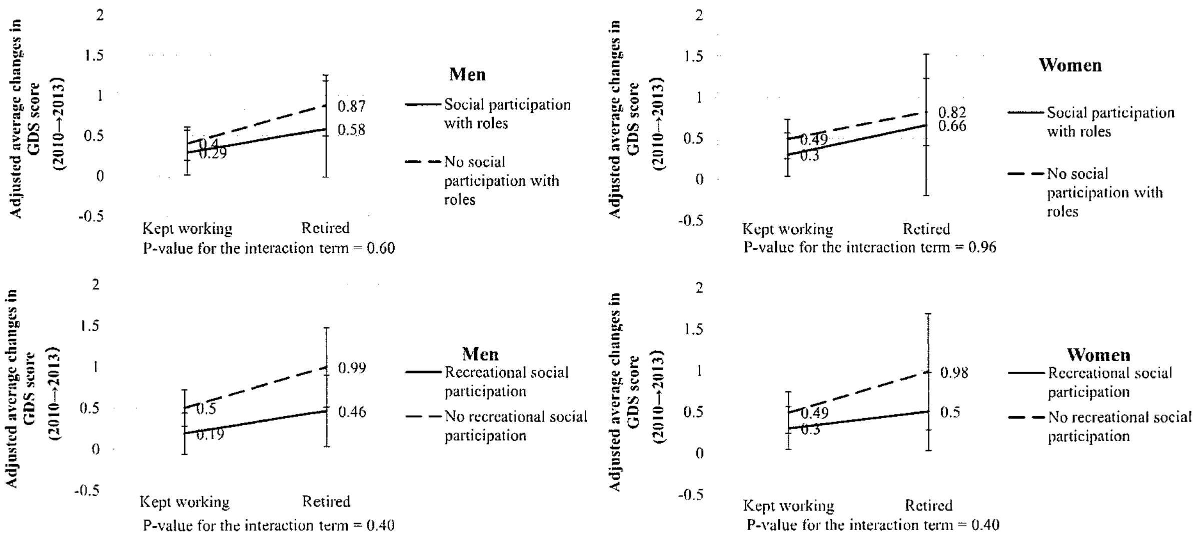


**Fig. 2** Adjusted average changes in GDS score by changes in working status and post-retirement social participation.
GDS: the short version of the geriatric depression scale (ranging from 0 to 15, higher score indicates more depressive symptoms). Adjusted for changes in time varying confounding factors including equivalised houshold income, IADL limitation, marital status, stressful life events, social relationships. Subjects are those who did not show depression at baseline (GDS score < 5). Social participation with roles includes volunteering clubs and neighborhood councils. Recreational social participation includes sports organizations, hobby clubs, and older adults clubs

among women. Fourth, we found no interaction between changes in working status and changes in marital status. Finally, there was a statistically non-significant trend that men and women who reported recreational social participation after retirement appeared to be less adversely impacted by retirement. Importantly, the magnitude of interaction effect between transition to retirement and post-retirement recreational social participation substantially reduced only among men after additionally adjusting for a sense of meaning in life. This, if causal, may suggest that having a sense of meaning in life through social participation may be crucial in mitigating the adverse impact of retirement on mental health – particularly for Japanese men.

In our data, nearly a quarter of male and female subjects indicated depressive symptoms at baseline defined by the GDS score of 5 or more. Another population-based study in Japan targeting people over the age of 64 reported that prevalence of those who scored more than 5 in GDS-15 was 23.8%, which is compatible to our data [26].

Our finding that retirement may increase depressive symptoms is inconsistent with many studies conducted in European countries [5–11], whereas a previous study in another Asian population supports our result [12, 13]. We speculate that, for Japanese workers (men especially), the loss of non-financial benefits of work outweighs the other benefits of retirement such as a relief from work-related stress overall. Interestingly, we found that changes in social contacts and social supports failed to explain the association between retirement and depressive symptoms. Although convincing evidence is lacking, the adverse effect of retirement may be attributable to a loss of other benefits of work including sense of meaning in life. Indeed, the literature suggests that many Japanese workers tend to find their purpose in life from work [27]. Retirement among Japanese workers may thus result in a loss of a sense of meaning in life.

Among Japanese workers who experienced the post-War period of rapid economic growth, employment was characterized by long hours of work as well as a shared sense of sacrifice with the goal of rebuilding the nation in the aftermath of the destruction wrought by the Second World War [28]. The participants in JAGES study – all of whom were aged over 65 years at baseline in 2010 – belong to this cohort. In return for their sacrifice of leisure and family life, Japanese workers of this generation were rewarded with a life-time employment guarantee. Commitment to the company thereby molded the identity of an entire generation of Japanese workers. It is in this context that we may interpret the findings of the present study. The transition to retirement for many Japanese workers entails the severing of links with work colleagues as well as loss of a sense of meaning in life. It is important to note at the same time, however, that the post-War lifetime employment guarantee has broken down in Japanese society in the aftermath of the 1991 economic bubble collapse. Our findings may therefore not apply to future generations of Japanese workers.

Our results also suggest that retirement may be more toxic among Japanese men from lower occupational class backgrounds compared to managerial/professional occupations. One explanation is that men from high SES backgrounds may have extra social roles besides work-related social roles, and therefore, may be less affected by a loss of work-related role due to retirement. Additionally, there may be some health benefits of retirement for the managerial/professional class. Indeed, mortality trends in Japan by occupational class indicate that suicide rates in these groups have risen sharply in the past two decades, possibly as a result of the pressures stemming from economic stagnation [29, 30]. It is possible that retirement can be somewhat beneficial for those workers as it relieves them of the work-related distress. The observed wide confidence intervals for the interaction terms between changes in working status and occupational class among women may be due to the small number of women who were from high-occupational background. Indeed, among female JAGES subjects in 2010, only 12.6% were categorized as high occupational class.

Our findings also indicate the trend that participation in recreational activity could mitigate the increase in depressive symptoms among Japanese retired men and women but perhaps through different mechanisms. Results from the model additionally adjusting for a sense of meaning in life suggests that recreational social participation may benefit retired Japanese men by compensating for a loss of a sense of meaning in life due to retirement. However, this was not the case for retired women. We speculate that recreational social participation for these retired women had differential meaning such as increased opportunities for informal socializing, rather than provision of a sense of meaning in life. Indeed, Takagi et al. also argued that there was a gender difference in mechanisms through which social participation enhances mental health [24]. We found no significant interaction between transition to retirement and social participation with roles including volunteerism, which was inconsistent with preceding findings [12, 13]. Participation in activities involving roles may not be completely voluntary and some participants might have been feeling a psychological distress due to "role strain", namely a pressure to accomplish their roles [20, 31]. This may explain the non-significant interaction effect of social participation with roles and its wide variance.

Ours is, to our knowledge, the first study to evaluate the association between changes in working status and changes in depressive symptoms using data of Japanese older people. Strengths of our study include the large sample size of older adults with repeated measures of both outcomes and exposures, and the use of first-difference models to account for all time-invariant confounding factors. Nonetheless, several limitations should be noted. First, there is a possibility of reverse causation, such that workers are more likely to seek retirement when they feel depressed [32]. Although we partly addressed this issue by excluding subjects who showed depression at baseline from the analyses, future studies should specify reasons for retirement to permit causal inference. Second, unmeasured confounding factors may also exist. Even though we controlled for all observed/ unobserved time-invariant confounding factors and several possible time-variant confounding factors, there may be other stressful life events or circumstances, which causes changes in working status and changes in depressive symptoms. Third, some of our measurements were admittedly crude. We used changes in received social supports from "friends" as a measurement of changes in work-related social relationships. However, it is obvious that social relationships people obtain from work are not necessarily reported as "friends". Fourth, we evaluated social participation at the same time with our final outcome. Since we hypothesized that post-retirement social participation could mitigate the possibly adverse impact of retirement on mental health, it is ideal to measure social participation before measuring post-retirement mental health. Fifth, generalizability of our findings is limited. As the effect of retirement on mental health is presumably culture dependent, our findings may not be generalizable to other countries or future generations. Lastly, our single item measurement of "a sense of meaning in life" is not validated, and therefore, it may not successfully capture a subject's sense of meaning in life.

## Conclusions

Retirement may increase depressive symptoms among Japanese older people, possibly men with lower occupational class backgrounds in particular. An important public health implication of our study is that there is a possibility that encouraging recreational social participation may compensate for a loss of a sense of meaning in life due to retirement and mitigate the potentially adverse effects of retirement on mental health of Japanese older men.

## Additional files

**Additional file 1: Table S1.** Regression Coefficients of Changes in Working Status and Interactions with occupational class (Model 3) and changes in marital status (Model 4) by gender (full-sample). (DOCX 17 kb)

**Additional file 2: Table S2.** Multiple linear regression of Changes in GDS score on changes in working status by gender (sub-sample). (DOCX 17 kb)

**Additional file 3: Table S3.** Regression Coefficients of Changes in Working Status and Interactions with social participation by gender (sub-sample). (DOCX 14 kb)

### Abbreviations

GDS-15: Short version of the Geriatric Depression Scale; IADL: Instrumental activities of daily living; JAGES: Japan gerontological evaluation study; TMIG: Tokyo Metropolitan Institute of Gerontology


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### Availability of data and materials

The dataset supporting the conclusions of this article is available in response to the request from the researchers admitted by the JAGES committee.

### Authors' contributions

SK: Conception and design, data analysis and interpretation, drafting the manuscript. KN: Conception and design, data interpretation, critical review of the manuscript. KK: data collection, critical review of the manuscript. KI: Conception and design, data interpretation, critical review of the manuscript.

### Competing interests

The authors declare that they have no competing interests.

### Consent for publication

Not applicable.

### Ethics approval and consent to participate

Participants were informed that their voluntary participation in the study and returning the self-administered questionnaire via mail indicated their consent to participate in the study. The study protocol and informed consent procedure were approved by the Ethics Committee in Research of Human Subjects at Nihon Fukushi University.

## Publisher's Note




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Original Article

# Development of an instrument for community-level health related social capital among Japanese older people: The JAGES Project

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ABSTRACT

*Background:* We developed and validated an instrument to measure community-level social capital based on data derived from older community dwellers in Japan.
*Methods:* We used cross-sectional data from the Japan Gerontological Evaluation Study, a nationwide survey involving 123,760 functionally independent older people nested within 702 communities (i.e., school districts). We conducted exploratory and confirmatory factor analyses on survey items to determine the items in a multi-dimensional scale to measure community social capital. Internal consistency was checked with Cronbach's alpha. Convergent construct validity was assessed via correlating the scale with health outcomes.
*Results:* From 53 candidate variables, 11 community-level variables were extracted: participation in volunteer groups, sports groups, hobby activities, study or cultural groups, and activities for teaching specific skills; trust, norms of reciprocity, and attachment to one's community; received emotional support; provided emotional support; and received instrumental support. Using factor analysis, these variables were determined to belong to three sub-scales: civic participation (eigenvalue = 3.317, $\alpha$ = 0.797), social cohesion (eigenvalue = 2.633, $\alpha$ = 0.853), and reciprocity (eigenvalue = 1.424, $\alpha$ = 0.732). Confirmatory factor analysis indicated the goodness of fit of this model. Multilevel Poisson regression analysis revealed that civic participation score was robustly associated with individual subjective health (Self-Rated Health: prevalence ratio [PR] 0.96; 95% confidence interval [CI], 0.94–0.98; Geriatric Depression Scale [GDS]: PR 0.95; 95% CI, 0.93–0.97). Reciprocity score was also associated with individual GDS (PR 0.98; 95% CI, 0.96–1.00). Social cohesion score was not consistently associated with individual health indicators.
*Conclusions:* Our scale for measuring social capital at the community level might be useful for future studies of older community dwellers.



## Introduction

A growing body of studies suggests that social capital has a significant influence on health and health behaviors.[1–8] The concept of social capital is used in two distinct approaches: the network-based approach and the social cohesion approach.[9] Most public health research adopts the latter social cohesion approach,

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which clarifies the contextual effect of community-level social capital as a group attribute or collective property. Community-level social capital is important for older adults to maintain health and well-being, as they are likely to spend many hours in the community. Japan is a global leader among developed countries that are experiencing rapid population aging. The proportion of older people is currently 26.0% and is predicted to reach 30.3% by 2025. To tackle issues associated with this situation, the Japanese government started a novel public health agenda for the health of older adults called Integrated Community Care for older adults.[10] This agenda aims to build social capital at the community level, improve local healthcare governance, and enrich local resources/environments supporting older residents. Therefore, interest in measuring and monitoring social capital at the community level has increased among central and local governments.

To date, several scales have been developed to measure social capital, including scales that can be used in the workplace,[11] in the school setting,[12,13] and for caregivers of children requiring special care needs,[14] as well as for trainees in clinical and translational science.[15] However, to our knowledge, no community social capital scale is available that is useful for studies of older people in industrialized countries like Japan. The generalizability of existing social capital scales might be limited, as most of them have been developed in only a few or single communities. Information on the validity and reliability of those scales is widely lacking.[2,16] Available scales also fail to capture multiple dimensions of community-level social capital, such as cognitive and structural social capital.[2,16]

In this paper, using large-scale data from a survey of community-dwelling older adults, we developed and validated an instrument to measure community social capital in older community-dwelling populations. Various definitions of community social capital have been offered.[16–21] Of these, influential definitions in the fields of epidemiology and public health include the definition by Coleman[22]: “a variety of different entities having two characteristics in common: they all consist of some aspect of social structure, and they facilitate certain actions of individuals who are within the structure”. Putnam's definition is also well known: “features of social organization such as networks, norms, and social trust that facilitate coordination and cooperation for mutual benefit”.[23] In social epidemiology, Kawachi and Berkman introduced a more straightfoward definition that is useful in public health settings: “resources that are accessed by individuals as a result of their membership of a network or a group”.[17]

Referring to these definitions, we have developed a health-related social capital scale at the community level because we assumed that our scale would be used to conduct community diagnosis (i.e., to evaluate the characteristics of the community and individual residents). Evaluating the contextual effects of community characteristics on individual health is a key interest of studies and activities of public health.

## Methods

### *Data*

We analyzed cross-sectional data derived from the year 2013 wave of the Japan Gerontological Evaluation Study (JAGES) Project. JAGES investigated people aged 65 years or older who did not have physical or cognitive disabilities, a state which was defined as not receiving public long-term care insurance benefits in 30 municipalities. The municipalities were not randomly selected but covered a wide range of characteristics in terms of regions and population sizes in Japan. In 13 relatively small municipalities, self-administered questionnaires were mailed to all functionally independent older adults, and in 17 municipalities, questionnaires were mailed to randomly selected older adults based on the official residential registers (response rate, 71.1%). The respondents were 129,739 residents nested in 832 communities that were primarily based on school districts, with some municipal exceptions. Communities containing <50 respondents were excluded to avoid non-precise values due to small samples. Ultimately, we derived data from 702 communities comprising 123,760 individuals. The mean number of observations per community was 176 (standard deviation [SD], 226).

We aggregated individual responses into small areas (i.e., school districts) to assess social capital at the community level. Although social capital could be evaluated at various levels of aggregations, such as municipality, prefecture, and country levels,[24] we selected the school district as the unit of community in this paper for the following reasons. First, in most regions, school district could represent a geographical scale in which older adults can travel easily by foot or bicycle, and many local activities by community organizations, such as senior citizens club and sports clubs, are performed within each school district. Second, school district is valuable unit for considering local public health activities. Using school districts as the sampling unit, we could evaluate regional variability in social capital within each municipality, which may help local public health practitioners in conducting their activities. Third, it is the smallest area size in which we could maintain sufficient precision of the aggregated information, in terms of the number of samples within each community.

### *Selection of candidate variables for social capital scale*

Referring to available concepts of social capital,[16–21] we selected 53 indicators that were potentially associated with social capital (See eTable 1 for full lists of variables). For example, selected variables included the proportion of residents in each community who reported participating in community-based activities (e.g., volunteer groups; sports groups or clubs; hobby activity groups; senior citizen clubs; community associations; study or cultural groups; nursing care prevention activities; activities to teach skills or pass on experiences to others; local events, including festivals and dances; activities to support older people requiring protection; activities to support older people requiring nursing care; activities to support parents raising children; and local living arrangement improvement or beautification activities). Items also assessed social integration (e.g., average number of friends and frequency of contact with them, the proportion of people who received or provided social support, and interactions with neighbors), trust, norms of reciprocity, and attachment.

### *Subjective health indicators*

The concurrent validity of our social capital scale was evaluated using the health indicators of self-rated health (SRH) and depressive symptoms. These are valid predictors of mortality regardless of other medical, behavioral, or psychosocial factors.[25,26] We measured SRH using the question “How do you feel about your current health status: excellent, good, fair, or poor?” Depressive symptoms were assessed using the 15-item Geriatric Depression Scale (GDS) that was developed for self-administration in the community using a simple binary (yes/no) format.[27,28] Scores $\geq$5 on the GDS indicate mild to severe depression.[28]

### *Statistical analysis*

#### *Selection of variables for social capital scale*

First, we aggregated each selected variable into the community (school district) level because a community social capital scale

**Table 1**
Correlation between social capital candidate indicators and subjective health at the community level (n = 702).

| | Partial correlation[a] | |
|---|---|---|
| | SRH (fair/poor) | GDS (≥5) |
| Volunteer group (≥once a month) | −0.093** | −0.193*** |
| Sports group (≥once a month) | −0.233*** | −0.355*** |
| Hobby activity (≥once a month) | −0.256*** | −0.332*** |
| Study or cultural group (≥once a month) | −0.144*** | −0.205*** |
| Skills teaching (≥once a month) | −0.106** | −0.200*** |
| Frequency of contact with friends (rarely) | 0.272*** | 0.372*** |
| Number of friends (≥10) | −0.140** | −0.255*** |
| Receive emotional support (any one or more) | −0.049 | −0.189*** |
| Provide emotional support (any one or more) | −0.154** | −0.265*** |
| Receive instrumental support (any one or more) | −0.186*** | −0.292*** |
| Community trust (strongly & moderately trusted) | −0.204*** | −0.373*** |
| Norms of reciprocity (agree strongly & agree) | −0.144*** | −0.331*** |
| Community attachment (strongly & moderately attached) | −0.136*** | −0.315*** |
| Facilities you feel free to drop in | −0.102** | −0.206** |

GDS, geriatric depression scale; SRH, self-rated health.
***$p < 0.001$, **$p < 0.01$, *$p < 0.05$.
[a] Population density and elderly proportion at municipality level were controlled.

should be created from the multiple indicators representing community-level characteristics. For example, when perceptions about trust are aggregated to the group level, it is no longer a measure of personal perceptions but a measure of the trustworthiness of people in the group.[9] Second, to extract the variables related to health outcomes, we calculated partial correlations between each candidate variable and the health indicators of SRH and GDS, after controlling for population density and the proportion of older individuals (ecological analysis). Candidate variables with moderate or strong correlations with either SRH or GDS were then extracted (r > 0.150). When several variables were conceptually similar, we adopted the variable with the closest relationship to be the health indicators. Third, we conducted exploratory factor analysis and eliminated low-communality variables so that the remaining variables maximized internal consistency, as evaluated based on Cronbach's alpha test. Fourth, we applied the maximum likelihood method with promax rotations for these factor analyses to account for the correlations among the factors identified. The utilization of multiple community indicators rather than a single indicator in creating a community social capital scale increases the reliability of the scale created. We then performed confirmatory factor analysis. We did not attempt to improve the fit index of our confirmatory factor analysis model via basing the analysis on residual covariance matrices.

*Evaluating concurrent validity*

The concurrent validity of our scale was determined using multilevel Poisson regression predicting individual SRH and GDS. To model contextual effects of community social capital, we used multilevel analysis to account for the variability in health outcomes due to individual compositions (i.e., individual's sociodemographic backgrounds and the responses to the questions used for making our community social capital scale).[9] To model potentially different associations between community social capital and individual health across individual characteristics, we also applied a cross-level interaction term. We used Stata 12.1 (StataCorp, College Station, TX, USA) and MLwiN 2.32 (Centre for Multilevel Modelling, Bristol University, Bristol, UK) for statistical analysis.

*Ethical considerations*

JAGES participants were informed that participation in the present study was voluntary and that completing and returning the self-administered questionnaire via mail indicated their consent to participate in the study. Ethics approval was obtained from the Ethics Committee at Nihon Fukushi University (13-14).

**Table 2**
Extraction of social capital candidate indicators based on reliability.

| | Communalities (Factor analysis) | |
|---|---|---|
| Number of items | 14 | 11 |
| Volunteer group | 0.325 | 0.315 |
| Sports group | 0.638 | 0.640 |
| Hobby activity | 0.743 | 0.752 |
| Study or cultural group | 0.500 | 0.495 |
| Skills teaching | 0.295 | 0.288 |
| Less frequency of contact with friends | 0.273 | |
| Number of friends | 0.229 | |
| Receive emotional support | 0.679 | 0.687 |
| Provide emotional support | 0.538 | 0.533 |
| Receive instrumental support | 0.392 | 0.394 |
| Community trust | 0.795 | 0.883 |
| Norms of reciprocity | 0.720 | 0.650 |
| Community attachment | 0.534 | 0.529 |
| Facilities you feel free to drop in | 0.273 | |
| Cronbach's alphas | 0.728 | 0.752 |

## Results

Based on the results of the correlation analysis, we selected 14 of the 53 candidate variables that were strongly or moderately associated with health indicators (Table 1). We excluded three variables to improve communalities and ultimately adopted 11 for inclusion in our health-related community social capital scale based on internal consistency ($\alpha = 0.752$) (Table 2).

Exploratory factor analysis (Table 3) suggested that three factors (eigenvalues: 3.317, 2.633, and 1.424) composed of the 11 variables, with cumulative contribution of 67.0%. The first factor was mainly associated with the participation in volunteer groups, sports groups, hobby activities, study or cultural groups, and activities for teaching skills ($\alpha = 0.797$). We collectively named this factor “civic participation”. The second factor that was strongly associated with trust, community trust and attachment ($\alpha = 0.853$), was named “social cohesion”. The third factor that was strongly associated with receiving and providing emotional support and receiving instrumental support ($\alpha = 0.732$) was named “reciprocity”. Social cohesion score significantly correlated with reciprocity score (r = 0.436, $p < 0.001$). Confirmatory factor analysis showed that the root mean square error of approximation was 0.089, the comparative fit index

**Table 3**
Factor loadings of community-level social capital scale.

| | Exploratory factor analysis[a] | | | Confirmatory factor analysis[b] | | |
|---|---|---|---|---|---|---|
| | Civic participation (F1) | Social cohesion (F2) | Reciprocity (F3) | Civic participation (F1) | Social cohesion (F2) | Reciprocity (F3) |
| Volunteer group | 0.536 | 0.119 | −0.029 | 0.557 | – | – |
| Sports group | 0.791 | −0.015 | 0.100 | 0.796 | – | – |
| Hobby activity | 0.868 | −0.020 | 0.021 | 0.867 | – | – |
| Study or cultural group | 0.706 | −0.023 | −0.051 | 0.693 | – | – |
| Skills teaching | 0.536 | 0.003 | −0.060 | 0.532 | – | – |
| Community trust | 0.055 | 0.934 | −0.009 | – | 0.947 | – |
| Norms of reciprocity | −0.058 | 0.817 | −0.015 | – | 0.790 | – |
| Community attachment | 0.055 | 0.716 | 0.007 | – | 0.727 | – |
| Received emotional support | −0.092 | −0.005 | 0.831 | – | – | 0.828 |
| Provided emotional support | 0.104 | −0.097 | 0.750 | – | – | 0.682 |
| Received instrumental support | −0.061 | 0.257 | 0.486 | – | – | 0.603 |
| Correlation coefficient | | | | | | |
| F1 & F2 | 0.154 ($p = 0.000$) | | | 0.178 ($p = 0.000$) | | |
| F1 & F3 | 0.065 ($p = 0.087$) | | | 0.031 ($p = 0.495$) | | |
| F2 & F3 | 0.436 ($p = 0.000$) | | | 0.392 ($p = 0.000$) | | |

[a] Exploratory factor analysis was applied promax rotation and maximum likelihood method.
[b] Model fit indicators of confirmatory factor analysis were as follows: Chi-square (df) = 271.2(41), $p < 0.001$, RMSEA = 0.089, CFI = 0.925, TLI = 0.899, SRMR = 0.058.

was 0.925, the Tucker–Lewis index was 0.899, and the standardized root mean square residual was 0.058, which were almost comparable to the criterion of the fit index.

Table 4 shows descriptive statistics for the variables in the multilevel Poisson regression model. Individual-level civic participation was calculated via summing the number of civic groups (up to five) in which respondents participated once or more per month. Individual-level social cohesion and reciprocity were dichotomized into those who responded “strongly/moderately agree” and “any one or more” compared with all other responses. Even after controlling for individual socio-demographic status (i.e., age, gender, marital status, education, and annual household income), all community-level social capital scores were significantly associated with depressive symptoms (Table 5). The prevalence ratio (PR) was 0.94 (95% confidence interval [CI], 0.92–0.95) per 1 SD increase in the score for civic participation. The PRs for social cohesion and reciprocity were 0.97 (95% CI, 0.95–0.99) and 0.96 (95% CI, 0.95–0.98), respectively. When the outcome was changed to self-rated health, the PRs of social capital scores were similar to the associations with GDS scores, although statistical significance was marginal for social cohesion and reciprocity.

Additional adjustments for individual-level responses to the questions used to form our scale for evaluating social capital did not affect the PRs for civic participation (PR for poor/fair SRH, 0.96; 95% CI, 0.94–0.98; PR for GDS, 0.95; 95% CI, 0.93–0.97; Model 2). Reciprocity was also associated with individual GDS (PR 0.98; 95% CI, 0.96–1.00). On the other hand, the same adjustment attenuated the association between community-level social cohesion and individual health indicators. Most cross-level (individual and community) interactions were not statistically significant (Model 3). Applying an indicator method using a dummy variable as categorical data, which was done to account for non-normal distribution and missing data, did not alter the major results and trends (eTable 2). In addition, correlation coefficients between each community-level social capital score and individual responses were not high (0.09–0.17).

**Table 4**
Descriptive statistics of final sample.

| | |
|---|---|
| Outcome | |
| Fair/poor health | No (n = 97,324, 78.6%),[a] |
| | Yes (n = 22,134, 17.9%) |
| | Unknown (n = 4,302, 3.5%)[b] |
| Depressive symptoms | No (n = 74,648, 60.3%),[a] |
| | Yes (n = 26,700, 21.6%) |
| | Unknown (n = 22,414, 18.1%)[b] |
| Level 1 (individuals, n = 123,760) | |
| Age, years | Mean = 74.0, Range = 65–106 |
| Gender | Male (46.3%)[a] |
| | Female (53.7%) |
| Marital status | Married (70.0%)[a] |
| | Divorced (20.9%) |
| | Separated (3.3%) |
| | Never married (2.2%) |
| | Unknown (3.6%) |
| Education | ≥10 years (n = 56.0%)[a] |
| | <10 years (41.7%) |
| | Unknown (2.3%) |
| Annual household income (Equivalent income) | ≥\4,000,000 (8.4%)[a] |
| | \2,000,000–\3999.999 (30.0%) |
| | <\2,000,000 (41.4%) |
| | Unknown (20.2%) |
| Civic participation (number of groups which participated once or more per month in five indicators) | None (45.1%) |
| | One (15.5%) |
| | Two (10.3%) |
| | Over three (6.9%) |
| | Unknown (22.1%)[b] |
| Social cohesion (number of “strongly/moderately agree” in three indicators) | None (12.9%) |
| | One (17.1%) |
| | Two (20.6%) |
| | Three (44.6%) |
| | Unknown (4.7%)[b] |
| Reciprocity (number of “any one or more” in three indicators) | None (1.3%) |
| | One (3.0%) |
| | Two (6.7%) |
| | Three (82.8%) |
| | Unknown (6.2%)[b] |
| Level 2 (communities, n = 702) | |
| Civic participation (factor score) | Mean = 0, Range = −2.79 to 3.66 |
| Social cohesion (factor score) | Mean = 0, Range = −3.82 to 2.86 |
| Reciprocity (factor score) | Mean = 0, Range = −4.55 to 2.26 |

[a] Reference categories used for subsequent regression analyses.
[b] Unknown cases in these major variables were eliminated in subsequent regression analyses.

## Discussion

We developed and validated an 11-item scale, which was comprised of sub-dimensions of civic participation, social cohesion,

**Table 5**
Estimated prevalence ratios from multilevel Poisson regression analysis.

| | Self-rated health (fair/poor) | | | Depressive symptoms (GDS ≥5) | | |
|---|---|---|---|---|---|---|
| | Model 1<br>*PR* (95% CI) | Model 2<br>*PR* (95% CI) | Model 3<br>*PR* (95% CI) | Model 1<br>*PR* (95% CI) | Model 2<br>*PR* (95% CI) | Model 3<br>*PR* (95% CI) |
| *Fixed parameters* | | | | | | |
| **Community level variables** | | | | | | |
| Civic participation (factor score) | 0.94*** (0.92–0.96) | 0.96*** (0.94–0.98) | 0.97** (0.95–0.99) | 0.94*** (0.92–0.95) | 0.95*** (0.93–0.97) | 0.95*** (0.93–0.97) |
| Social cohesion (factor score) | 0.99 (0.97–1.01) | 1.03* (1.01–1.05) | 1.03 (0.99–1.07) | 0.97** (0.95–0.99) | 1.02* (1.00–1.04) | 1.03 (1.00–1.06) |
| Reciprocity (factor score) | 0.98 (0.97–1.01) | 1.00 (0.98–1.03) | 1.02 (0.95–1.09) | 0.96*** (0.95–0.98) | 0.98* (0.96–1.00) | 1.03 (0.98–1.09) |
| **Individual-level variables** | | | | | | |
| Civic participation (0–3) | | 0.74*** (0.72–0.76) | 0.74*** (0.72–0.76) | | 0.76*** (0.75–0.78) | 0.76*** (0.75–0.78) |
| Social cohesion (0–3) | | 0.84*** (0.83–0.85) | 0.84*** (0.83–0.85) | | 0.77*** (0.76–0.77) | 0.77*** (0.76–0.77) |
| Reciprocity (0–3) | | 0.86*** (0.84–0.88) | 0.86*** (0.84–0.88) | | 0.82*** (0.80–0.83) | 0.82*** (0.80–0.83) |
| **Cross-level interactions** | | | | | | |
| Community level civic participation × Individual level civic participation | | | 0.98 (0.95–1.00) | | | 1.01 (0.98–1.03) |
| Community level social cohesion × Individual level social cohesion | | | 1.00 (0.98–1.02) | | | 0.99 (0.98–1.01) |
| Community level reciprocity × Individual level reciprocity | | | 0.99 (0.97–1.02) | | | 0.98* (0.96–1.00) |
| *Random parameters* | | | | | | |
| Community level intercept variance (standard error) | 0.004 (0.002) | 0.004 (0.002) | 0.003 (0.002) | 0.001 (0.001) | <0.001 | <0.001 |
| Community level civic participation slope variance (standard error) | <0.001 | <0.001 | <0.001 | <0.001 | <0.001 | <0.001 |
| Community level social cohesion slope variance (standard error) | <0.001 | <0.001 | <0.001 | <0.001 | <0.001 | <0.001 |
| Community level reciprocity slope variance (standard error) | <0.001 | <0.001 | <0.001 | <0.001 | <0.001 | <0.001 |

CI, confidence interval; GDS, geriatric depression scale; PR, prevalence ratio.
***$p < 0.001$, **$p < 0.01$, *$p < 0.05$.
All models are also adjusted for individual-level age, gender, marital status, education, and annual house income. Unknown cases in social capital variables and dependent variables were eliminated in this analysis. Sample size of model 1 was as follows: SRH; individual = 119,458, community = 702, GDS; individual = 101,348, community = 702. Sample size of model 2 and model 3 were as follows: SRH; individual = 88,436, community = 702, GDS; individual = 79,081, community = 702.

and reciprocity, to measure social capital at the community level. Evaluations of communalities of factors and internal consistency, as well as confirmatory factor analysis, demonstrated that these 11 indicators formed a reliable scale. We also found evidence to support convergent validity: the indicators were correlated with health outcomes in expected directions. Our instrument might be useful for gerontological studies and activities in Japan, and, although further studies are needed, the scale may be useful in other countries with a similar context to that of Japan.

The distinction among civic participation, social cohesion, and reciprocity dimensions are fundamental in social capital theory.[3,16,17,29] Our factor analysis statistically identified these three components. According to Islam et al, the structural dimension of social capital includes externally observable aspects of social organization and is characterized by behavioral manifestations of network connections or civic engagement.[29] This concept was reflected in the variables included in our civic participation and reciprocity variables. That the factors reflecting subjective attitudes, such as trust, norms of reciprocity, and attachment within the community, were collectively named "social cohesion" was also theoretically reasonable. The "reciprocity" component of our social capital scale may capture the dimension of community social capital that promotes the exchange of individual social supports within the community. Our results, which showed strong correlation between "reciprocity" and "social cohesion", are also consistent with the theoretical framework. We did not find remarkable inflations of confidence intervals or strong correlation between community-level social capital scores and individual responses after adjusting for individual measures. These findings suggest that the potential for multicollinearity was not large.

Social capital studies in public health have so far yielded mixed findings, potentially as a result of inconsistencies in the manner in which investigators have operationalized and measured the concept of "social capital", often by resorting to proxy variables available through secondary data sets. Previous studies have used indicators, such as community voting rates in elections or the local crime rate, as proxies for social capital. Although these variables could be viewed as either antecedents or consequences of social capital, they do not directly capture its core concepts. In addition, differences between study findings might relate to variations in culture, region, units of analysis, and age cohorts.

Although communities in Japan might have closer social ties than those in the United States, generalized trust, which is the central concept of social cohesion or cognitive social capital, is low in cohesive societies because human relations in such a society are based on mutual "assurance" rather than "trust".[30] This has resulted in an unfavorable effect of trust on health in some Japanese studies.[31,32] Nonetheless, studies to date are difficult to compare due to their use of conventionally created alternative scales based on an assortment of concepts. Hence, we believe that our novel scale for measuring community social capital could provide a useful option for such studies. Using a common scale would contribute to the discussion about between-study differences of the "true" interests of investigators.

Our findings suggest that community-level civic participation is more closely correlated with health outcomes than social cohesion. This might be because the indicators of civic participation applied in our study were more objective than those used to determine social cohesion. Our analysis also showed a reverse predictive value of community-level social cohesion for health indicators. Further studies should be conducted to identify the possible reasons for those results. Portes's concepts of the dark side of social capital may help understanding the reasons.[33] Portes has pointed out four potentially harmful characteristics of group-level social capital: the exclusion of outsiders, excess claims on group members, restrictions on individual freedoms, and downward-leveling norms. A Japanese empirical study demonstrated that stronger social cohesion was associated with depressive symptoms in residents whose hometown of origin differed from the communities where they currently resided.[34] Alternatively, the weaker predictive value of community-level social cohesion might be explained via measurement bias due to the potential non-participation of those who trust others less. Information bias when asking about personal perceptions might be another explanation. Compared with information about individual memberships in community groups, individual perceptions regarding trust or reciprocity could be more easily influenced by various temporary conditions. This might also reflect some degree of reverse causation, in which healthier people are more likely to participate. Although individual-level perceptions of social capital were more strongly associated with individual health compared with community-level social capital,[35] this is not unexpected, since more proximal exposure to the individual is likely to correlate with individual-level health outcomes.

### *Strengths and limitations*

This study has several important and inherent strengths. The survey was originally designed to measure social capital, which allowed the use of various conceptually appropriate candidate variables. The large sample size in terms of numbers of individual participants and communities in particular is an important strength, as our analysis has sufficient power to create a community-level scale. We created a social capital scale specifically for the elderly population, which is the first of its kind. Nevertheless, the study also had some key limitations. The cross-sectional design might have included reverse causality that potentially biased the results of our evaluation of concurrent validity. Moreover, it would be better to conduct further validations. For example, criterion validity could be evaluated using more objectively measured community-level variables, such as the proportion of participation in each organization and the voting rate. Validation studies using hard outcomes, such as mortality, are also required. Although the response rate to our survey was relatively high (71.1%), selection bias cannot be fully excluded. Generalizability might be limited because our dataset was not a nationally representative sample and was created for older adults. Caution is needed when applying our community social capital scale to data obtained from alternative contexts, such as younger populations or using data created via alternative survey methods. However, the geographic and cultural variations of the municipalities included in our sample were high, and the municipalities included metropolitan and rural areas. Moreover, we used school districts as community units, but we do not know whether our scale would be similarly valid when evaluating a community defined using an alternative area unit. Ideally, a social capital index that can be applicable at any level of aggregation should be created. However, we found that the three factors based on our alternative factor analyses using individual-level data also showed the same items in each factor, suggesting the potential generalizability of our social capital scores to alternative units of aggregation.

## Conclusion

We developed a health-related community social capital scale, which was composed of 11 items assessing civic participation, social cohesion, and reciprocity. The scale was designed to allow calculations based on data derived from older populations. This new standard social capital measure could shed light on public health and gerontological issues, as well as other matters associated with community social capital.

## Author's contributions

Conceived and designed the analysis: MS, NK, JA. Performed the survey: KK, the JAGES Group. Analyzed the data: MS, SK, JA. Contributed to the writing of the manuscript: MS, NK, JA, IK, KK, TO. All authors read and approved the final version of the manuscript.

## Conflicts of interest

None declared.

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## Appendix A. Supplementary data

Supplementary data related to this article can be found at http://dx.doi.org/10.1016/j.je.2016.06.005.

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# Is a hilly neighborhood environment associated with diabetes mellitus among older people? Results from the JAGES 2010 study

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ABSTRACT

*Background:* Although living in a hilly environment may promote muscular activity in the daily lives of residents, and such activity may prevent diabetes mellitus, few studies have focused on the impact of living in a hilly environment on diabetes mellitus. The purpose of this study was to investigate the impact of a hilly neighborhood environment on DM in older people.

*Methods:* We used data from the Japan Gerontological Evaluation Study, a population-based, cross-sectional study of individuals aged 65 or older without long-term care needs in Japan, which was conducted in 2010. A total of 8904 participants in 46 neighborhoods had responded to the questionnaire and undergone a health check. Diabetes mellitus was diagnosed as $HbA_{1c} \geq 6.5\%$ and those undergoing treatment for diabetes mellitus. Poorly controlled diabetes mellitus was diagnosed in those without other chronic diseases who had an $HbA_{1c} > 7.5\%$, and in those with other chronic diseases if their $HbA_{1c}$ was >8.0%. Neighborhood environment was evaluated based on the percentage of positive responses in the questionnaire and geographical information system data. A multilevel analysis was performed, adjusted for individual-level risk factors. Furthermore, sensitivity analysis was conducted for those who were undergoing treatment for diabetes mellitus ($n = 1007$).

*Results:* After adjustment for other physical environmental and individual covariates, a 1 interquartile range increase (1.48°) in slope in the neighborhood decreased the risk of poorly controlled diabetes mellitus by 18% (odds ratio [OR]: 0.82, 95% confidence interval [CI]: 0.70–0.97). Sensitivity analysis confirmed that larger slopes in the neighborhood showed a significant protective effect against diabetes mellitus among those who were undergoing treatment for diabetes mellitus (OR: 0.73, 95% CI: 0.59–0.90).

*Conclusion:* A hilly neighborhood environment was not associated with diabetes mellitus, but was protective against poorly controlled diabetes mellitus.



## 1. Introduction

The control of type 2 diabetes mellitus (DM) is one of the most important public health issues in our aging society (Zimmet et al., 2001). The number of patients with type 2 DM worldwide is

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estimated to reach 266 million in 2030 (Qi et al., 2008), a trend that is partially attributed to the increasing older population (Charvat et al., 2015). Although several individual-level lifestyle risk factors (such as healthy diet, physical activity, and obesity (Zimmet et al., 2001; Zimmet, 1992)) have been identified, and individual-level interventions are reported to be effective (Colagiuri et al., 2010; Aguiar et al., 2016), changes to the neighborhood environments that influence lifestyle behaviors (Feng and Astell-Burt, 2013) and other such interventions might also be effective as a preventative public health strategy (Rose, 1992).

Previous studies have reported the preventive effect of changes to neighborhood environments on the development of type 2 DM, including changes to deprived neighborhoods (Muller et al., 2013; Williams et al., 2012), as well as access to exercise facilities, grocery stores (Christine et al., 2015; Auchincloss et al., 2008), and green spaces (Astell-Burt et al., 2014). It is known that the neighborhood environment related to walkability, such as the number of parks, intersection density, public transportation density, and net residential density, is associated with physical activity (Sallis et al., 2016). Moreover, perceived neighborhood walkability has shown an inverse association with sedentary behavior (Van Dyck et al., 2012). However, few studies have focused on the impact of a hilly environment on DM. A hilly environment may promote muscular activity in the daily lives of residents, and such activity may prevent DM (Minetti et al., 2002). Interestingly, Villanueva et al. reported that the rate of self-reported DM in adults living in a neighborhood with steeper slopes was lower than that of adults living in flatter areas (Villanueva et al., 2013). Studies using a more objective assessment of DM status, such as $HbA_{1c}$ levels, rather than self-reported assessment studies, are needed to investigate the link between a hilly environment and DM.

In older people, the control of DM is important to the prevention of complications (The effect of intensive t, 1993; Intensive blood-glucose c, 1998; Ohkubo et al., 1995) such as diabetic nephropathy, retinopathy, neuropathy, and microvascular problems (Zimmet, 1999). Management of these complications involves significant medical costs (Liebl et al., 2015). Furthermore, intensive glycemic control to prevent the onset or deterioration of DM may be harmful in frail older people who are more likely to have other diseases (Lipska et al., 2015; Vijan et al., 2014). Therefore, close attention should be paid to those with poorly controlled DM, which is defined as $HbA_{1c}$ levels of 7.5% or greater in healthy individuals with no comorbidities other than DM, or 8.0% or greater in elderly people with chronic disease, mild cognitive impairment, or an inability to perform two or more daily tasks (Kirkman et al., 2012). To our knowledge, however, no study has focused on the association between the neighborhood environment and poorly controlled DM in older people.

In this study, we used large population-based data on older people (i.e., 65 years or over) without long-term care needs to investigate the link between the neighborhood physical environment—assessed by both survey-based and geographic information system (GIS)-based data—and DM, which was assessed by $HbA_{1c}$ levels obtained from community-based health checks. We hypothesized that older people living in a hilly environment, that is, living in an area with steep slopes, may be less likely to have DM, especially poorly controlled DM, after accounting for potential lifestyle and other such factors in the neighborhood physical environment.

## 2. Methods

### 2.1. Sample

We used part of the data from the Japan Gerontological Evaluation Study (JAGES) project, a large, ongoing, prospective epidemiological study of the Japanese population aged 65 years or older from 31 municipalities across all areas of Japan without long-term care needs ($n = 112{,}123$; response rate, 66.3%). For the current study, we used data from individuals in six municipalities in the Chita peninsula, Aichi, for whom we had health check data ($n = 9893$). The data were collected in August 2010 to January 2012. We excluded individuals for whom information on HbA1c levels ($n = 393$) and neighborhood code ($n = 7$) were missing. After further limiting the sample to those who had lived in the same city for 10 years or more, a total of 8904 participants were included in analyses to determine the impact of living in a hilly environment on DM. The JAGES protocol was approved by the Ethics Committee for Research of Human Subjects at Nihon Fukushi University (No. 10-05).

### 2.2. $HbA_{1c}$ measurement and blood glucose level

$HbA_{1c}$ was measured at community health centers or registered hospitals in health checks organized by the municipality and analyzed by local hospital laboratories using the latex agglutination method, and reported by the Japan Diabetes Society (JDS). The measured ratio was converted to a National Glycohemoglobin Standardization Program (NGSP) value based on the transform equation (Geistanger et al., 2008). Fasting blood glucose level was also measured in around half of the participants ($N = 4623$, 51.9%), and casual blood glucose level was measured in another one-fifth of participants ($N = 1960$, 20.2%).

DM was diagnosed if $HbA_{1c}$ was $\geq$6.5%, fasting glucose level was $\geq$126 mg/dl, or casual glucose level was $\geq$200 mg/dl, or if the patient was undergoing treatment for DM regardless of the level of HbA1c. Poorly controlled DM was defined as $HbA_{1c} > 7.5\%$ for those without other diseases or >8.0% for those with other chronic diseases, including cancer, heart disease, stroke, hypertension, or dyslipidemia, based on the latest consensus report on DM among older people from American Diabetes Association (Kirkman et al., 2012).

### 2.3. Objective neighborhood environment (slope, population density, and land value)

We used the ArcGIS 10.1 software to assess the neighborhood physical environment as a special calculation, and included the presence of slopes, grocery stores, and parks. In accordance to a previous study using JAGES data (Takagi et al., 2013a), we used the elementary school district as the neighborhood unit, which is a proxy for a geographical area that is easy for elderly people to navigate (Hanibuchi et al., 2008). The average area of an elementary school district was 6.34 $km^2$ ($SD = 3.86$).

The average land slope of school districts was calculated using the Elevation, Degree of Slope 5th Mesh Data (as of 2011) of the National Land Numerical Information from the Ministry of Land, Infrastructure, Transport and Tourism in Japan, based on the Digital Map 50 m Grid (Elevation) from the Geospatial Information Authority of Japan (GSI) (Ministry of Land I and Transport and Tourism, 2016a). The range of slopes was from 0.95 to 9.79°, with a mean of 3.03 ($SD = 1.82$).

The number of grocery stores in each school district was calculated using the 500-m mesh data from the Census of Commerce of 2007 conducted by the Ministry of Economy, Trade and Industry. In this study, a grocery store was defined as a department store, general merchandise store, specialized supermarket, daily commodities store, or convenience store. The average count of parks was calculated using the City Park Data (as of 2011) of National Land Numerical Information from the Ministry of Land, Infrastructure, Transport and Tourism in Japan, from GSI. The city park parameter included open spaces, athletic grounds, and ball parks that were built based on the Urban Park Act across the country (Ministry of Land I and Transport and Tourism, 2016b). Similarly, the number of hospitals in each school district was calculated using Medical Institution Data (as of 2010) of the National Land Numerical Information from the Ministry of Land, Infrastructure, Transport and Tourism in Japan, from GSI.

The population density of school districts was calculated using the 2010 census and Land Utilization Tertiary Mesh Data (as of 2010) of the National Land Numerical Information from the Ministry of Land, Infrastructure, Transport and Tourism in Japan based on the 1:25,000 Topographic Map of Japan. These calculations excluded non-developed areas (e.g., rivers, lakes, forest, and wasteland) (Ministry of Land I and Transport and Tourism, 2016c).

Further, to assess neighborhood socioeconomic status, land value was obtained from the GSI. The mean land value reported within each school district was used as a proxy for neighborhood socioeconomic status, according to previous studies in Japan (Takagi et al., 2013a, 2013b; Aida et al., 2013).

### 2.4. Perception of physical environments

To assess the psychological barriers to walking around the neighborhood due to slopes or stairs, we used the following question: “How do you feel about the difficulty in walking within 1 km around your home due to slopes or steps?” A Likert scale with four possible responses was used (1: a lot; 2: to some extent; 3: not so much; 4: not at all), along with “unknown” as a response item. The response was recorded as 1 for a positive response and 0 for the others. Similarly, the perception of ease of access to grocery stores to obtain fresh vegetables and fruits, and to parks or sidewalks, was also determined in the questionnaire by asking “Do you feel that it is easy to access grocery stores within 1 km around your home to obtain fresh vegetables or fruits?” and “Do you feel that the park or sidewalks are appropriate for physical activity or walking?” which were answered with the same response items as above. The proportion of positive responses for each school district was calculated.

### 2.5. Covariates

For individual covariates, the questionnaire inquired about the duration of residence in the city, marital status (married, widowed, divorced, never married, other), number of household members, annual household income, education, working status (currently working, retired, never worked), drinking (current drinker, ex-drinker, never drinker), smoking (current smoker, ex-smoker, never smoker), vegetable consumption (2 times/day, once a day, 4–6 times/week, $\leq$ 3 times/week), walking habits (less than 60 min/day or $\geq$ 60 min/day), frequency of meeting with friends (less than once per month or 1 or more times per month), going-out behavior (less than once per month or 1 or more times per month), height, weight, and depression status using the Geriatric Depression Scale (GDS) (Wada et al., 2003) with a validated cut-off (5 or more) (Schreiner et al., 2003; Tani et al., 2015; Murata et al., 2008). Age, sex, height and weight were retrieved from the health check records, and body mass index (BMI) was calculated.

### 2.6. Analysis

For neighborhood assessments that were obtained from objective and questionnaire sources, except for population density, the interquartile range (IQR) was calculated to facilitate comparison of the strength of the association with DM. Before the main analysis, the association between each neighborhood assessment was calculated using Spearman's correlation coefficient.

The multilevel logistic regression with a random intercept model was used to investigate the association between neighborhood slope and DM or poorly controlled DM. Model 1 was adjusted for other neighborhood measurements that were associated with slope: perception of access to grocery stores (unit: IQR), number of parks (unit: IQR), number of hospitals (unit: IQR), population density (tertile), and land value (unit: IQR). Model 2 was adjusted for individual covariates that may be associated with DM: age; sex; marital status; number of household members; income; working, drinking, and smoking status; vegetable consumption; walking; going-out behavior; frequency of meeting with friends; categorized BMI (<18.5, 18.5–25, 25+); and depression. Model 3 was adjusted for the covariates in Models 1 and 2 simultaneously. For sensitivity analysis, we performed the same analyses on patients who were receiving DM treatment to avoid possible confounding biases in the association between a hilly environment and poorly controlled DM.

For the sensitivity analysis, we performed the same analysis but excluded participants who left the house less than once a month and who had a BMI of 30 or more (N = 526), as it can be considered that they are not exposed to a hilly environment.

## 3. Results

Table 1 shows the characteristics of participants with DM and poorly controlled DM. The prevalence of DM and poorly controlled DM was 17.7% (n = 1573) and 2.5% (n = 223), respectively. The average age was similar between participants with DM (72.5 years, standard deviation [SD] = 5.2), poorly controlled DM (72.3 years, SD = 5.0), and all participants (72.6 years, SD = 5.5) (72.3 years, SD = 5.0). More than 96% of all participants had lived in the current city for 20 years or more. Around one-third of participants were current drinkers, around 10% were current smokers, and 25% ate vegetables less than once a day. Two-thirds of participants walked less than 60 min per day, 20% met friends less than once per month, and 5% went out less than once per month. These unhealthy behaviors were more likely to be found among participants with DM or poorly controlled DM. Furthermore, DM or poorly controlled DM participants were more likely to be depressed and overweight (i.e., BMI $\geq$ 25.0). For DM and poorly controlled DM participants, the mean $HbA_{1c}$ level was 6.5% and 8.2%, and the proportion of those undergoing treatment for DM was 64.0% and 77.6%, respectively.

Table 2 shows the Spearman's correlation coefficients for community-level measures. GIS-slope was inversely correlated with a low perception of access to grocery stores to obtain fresh vegetables and fruits in the community, GIS-number of parks, and land value (r = −0.49, −0.30, and −0.48, respectively). No statistically significant correlation was observed between GIS-slope and other community-level measures.

Results of the multilevel logistic regression analysis for DM are shown in Table 3. The crude model revealed a protective association, although the significance was marginal (p = 0.085), between GIS-slopes and DM (odds ratio [OR] = 0.93, 95% confidence interval [CI] = 0.89–1.01). After adjustment for neighborhood environment covariates (Model 1), individual covariates (Model 2), and neighborhood and individual covariates (Model 3), the association became null, but the point estimate was similar (OR = 0.98, 95% CI = 0.88–1.09; OR = 0.93, 95% CI = 0.85–1.01; OR = 0.98, 95% CI = 0.88–1.10, respectively). No statistically significant association was observed between other community-level measures and DM in the fully adjusted model.

Table 4 shows the results of the multilevel logistic regression analysis for poorly controlled DM. The crude model showed a significant association between GIS-slope, and poorly controlled DM (OR = 0.86, 95% CI = 0.75–0.98). After adjustment for neighborhood environment covariates (Model 1), individual covariates (Model 2), and neighborhood and individual covariates (Model 3), the same trend remained (OR = 0.83, 95% CI = 0.70–0.99; OR = 0.85, 95% CI = 0.74–0.97; OR = 0.82, 95% CI = 0.70–0.97, respectively). Similarly to DM, no statistically significant

**Table 1**
Characteristics of total participants, and participants with diabetes mellitus and poorly controlled diabetes mellitus (N = 8904).

| | | Total | | DM (n = 1,573, 17.7%) | | Poorly controlled DM (n = 223. 2.5%) | |
|---|---|---|---|---|---|---|---|
| | | N or Mean | % or SD | N or Mean | % or SD | N or Mean | % or SD |
| Sociodemographics | | | | | | | |
| Age | mean (y.o.) | 72.6 | 5.5 | 72.5 | 5.2 | 72.3 | 5.0 |
| Sex | Men | 4072 | 45.7 | 888 | 56.5 | 125 | 56.1 |
| | Women | 4832 | 54.3 | 685 | 43.6 | 98 | 44.0 |
| Duration of living in the same city | 10-19 years | 311 | 3.5 | 58 | 3.7 | 8 | 3.6 |
| | 20-29 years | 523 | 5.9 | 113 | 7.2 | 12 | 5.4 |
| | 30-39 years | 1522 | 17.1 | 312 | 19.8 | 45 | 20.2 |
| | 40-49 years | 2177 | 24.5 | 378 | 24.0 | 56 | 25.1 |
| | 50+ years | 4371 | 49.1 | 712 | 45.3 | 102 | 45.7 |
| Marital status | Married | 6858 | 77.0 | 1253 | 79.7 | 178 | 79.8 |
| | Widow | 1683 | 18.9 | 259 | 16.5 | 41 | 18.4 |
| | Divorce/Never married/Other | 257 | 2.9 | 45 | 2.9 | 2 | 0.9 |
| | Missing | 106 | 1.2 | 16 | 1.0 | 2 | 0.9 |
| Number of household members | 1 | 736 | 8.3 | 112 | 7.1 | 13 | 5.8 |
| | 2 | 4199 | 47.2 | 740 | 47.0 | 106 | 47.5 |
| | 3+ | 3388 | 38.1 | 614 | 39.0 | 82 | 36.8 |
| | Missing | 581 | 6.5 | 107 | 6.8 | 22 | 9.9 |
| Education | <10 years | 4522 | 50.8 | 787 | 50.0 | 116 | 52.0 |
| | 10-12 years | 3062 | 34.4 | 552 | 35.1 | 71 | 31.8 |
| | 13+ years | 1217 | 13.7 | 218 | 13.9 | 35 | 15.7 |
| | Missing | 103 | 1.2 | 16 | 1.0 | 1 | 0.5 |
| Annual income | <1.5 million | 1040 | 11.7 | 193 | 12.3 | 26 | 11.7 |
| | 1.5–2.9 million | 3027 | 34.0 | 549 | 34.9 | 85 | 38.1 |
| | 3.0–5.9 million | 2835 | 31.8 | 513 | 32.6 | 74 | 33.2 |
| | 5.0 + million | 1027 | 11.5 | 164 | 10.4 | 19 | 8.5 |
| | Missing | 975 | 11.0 | 154 | 9.8 | 19 | 8.5 |
| Working status | Currently working | 1631 | 18.3 | 280 | 17.8 | 43 | 19.3 |
| | Retired | 5268 | 59.2 | 928 | 59.0 | 125 | 56.1 |
| | Never worked | 895 | 10.1 | 167 | 10.6 | 27 | 12.1 |
| | Missing | 1110 | 12.5 | 198 | 12.6 | 28 | 12.6 |
| Behaviors | | | | | | | |
| Drinking | Current | 3147 | 35.3 | 582 | 37.0 | 72 | 32.3 |
| | Quit | 251 | 2.8 | 63 | 4.0 | 15 | 6.7 |
| | Never | 4952 | 55.6 | 815 | 51.8 | 119 | 53.4 |
| | Missing | 554 | 6.2 | 113 | 7.2 | 17 | 7.6 |
| Smoking | Never | 4984 | 56.0 | 744 | 47.3 | 101 | 45.3 |
| | Quit | 2319 | 26.0 | 488 | 31.0 | 68 | 30.5 |
| | Current | 748 | 8.4 | 166 | 10.6 | 25 | 11.2 |
| | Missing | 853 | 9.6 | 175 | 11.1 | 29 | 13.0 |
| Vegetable eating | 2 times/d | 4009 | 45.0 | 633 | 40.2 | 84 | 37.7 |
| | 1 time/d | 2793 | 31.4 | 527 | 33.5 | 71 | 31.8 |
| | 4-6 times/w | 941 | 10.6 | 161 | 10.2 | 25 | 11.2 |
| | 3 or less times/w | 648 | 7.3 | 132 | 8.4 | 24 | 10.8 |
| | Missing | 513 | 5.8 | 120 | 7.6 | 19 | 8.5 |
| Walking time per day | <60 min/d | 5591 | 62.8 | 993 | 63.1 | 147 | 65.9 |
| | 60+ min/d | 2812 | 31.6 | 488 | 31.0 | 61 | 27.4 |
| | Missing | 501 | 5.6 | 92 | 5.9 | 15 | 6.7 |
| Frequency of meeting with friends | <1 time per month | 1720 | 19.3 | 335 | 21.3 | 49 | 22.0 |
| | 1+ time per month | 6655 | 74.7 | 1136 | 72.2 | 154 | 69.1 |
| | Missing | 529 | 5.9 | 102 | 6.5 | 20 | 9.0 |
| Frequency of going out | <1 time/m | 375 | 4.2 | 65 | 4.1 | 16 | 7.2 |
| | 1+ times/m | 8038 | 90.3 | 1406 | 89.4 | 189 | 84.8 |
| | Missing | 491 | 5.5 | 102 | 6.5 | 18 | 8.1 |
| Health status | | | | | | | |
| GDS | <5 | 5236 | 58.8 | 863 | 54.9 | 120 | 53.8 |
| | 5+ | 2263 | 25.4 | 456 | 29.0 | 67 | 30.0 |
| | Missing | 1405 | 15.8 | 254 | 16.2 | 36 | 16.1 |
| BMI | mean | 23.0 | 3.1 | 23.8 | 3.4 | 23.8 | 3.3 |
| | <18.5 | 575 | 6.5 | 74 | 4.7 | 14 | 6.3 |
| | 18.5–24.9 | 6229 | 70.0 | 994 | 63.2 | 138 | 61.9 |
| | 25.0+ | 2100 | 23.6 | 505 | 32.1 | 71 | 31.8 |
| HbA1c | mean (%) | 5.54 | 0.67 | 6.49 | 1.00 | 8.20 | 1.10 |
| DM treatment | not treated | 5766 | 64.8 | 417 | 26.5 | 23 | 10.3 |
| | treated | 1007 | 11.3 | 1007 | 64.0 | 173 | 77.6 |
| | Missing | 2131 | 23.9 | 149 | 9.5 | 27 | 12.1 |

Abbreviations: BMI, body mass index (calculated as weight in kilograms divided by height in square meters); GDS, Geriatric Depression Scale; DM, diabetes mellitus.

association was observed between other community-level measures and poorly controlled DM.

The results of analysis for those with poorly controlled DM undergoing treatment are shown in Table 5. Overall, a high GIS-slope was inversely associated with poorly controlled DM participants undergoing treatment, with the point estimate of the GIS-

**Table 2**
Summary of neighborhood measures and their Spearman correlations (n = 46).

| | Mean | SD | Median | Min | Max | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| 1. GIS-Slopes (degree) | 3.03 | 1.82 | 2.63 | 0.95 | 9.79 | 1.00 | | | | | | | |
| 2. Perception of difficulty to walk due to slopes and stairs in the community (%) | 49.2 | 12.2 | 48.1 | 25.4 | 80.3 | 0.21 | 1.00 | | | | | | |
| 3. GIS-Number of grocery stores (n) | 23.5 | 16.5 | 18.5 | 6 | 71 | −0.33 | −0.18 | 1.00 | | | | | |
| 4. Perception of access to grocery stores for fresh vegetables and fruits in the community (%) | 73.7 | 12.9 | 75.4 | 42.2 | 95.1 | **−0.49** | 0.19 | 0.25 | 1.00 | | | | |
| 5. GIS -Number of parks (n) | 4.1 | 3.8 | 3 | 0 | 14 | **−0.30** | −0.17 | **0.42** | **0.42** | 1.00 | | | |
| 6. Perception of access to park and sidewalk (%) | 73.8 | 11.7 | 75.7 | 52.6 | 94.6 | −0.21 | 0.29 | 0.15 | **0.51** | **0.57** | 1.00 | | |
| 7. GIS-number of hospital | 18.4 | 13.2 | 16 | 3 | 67 | 0.17 | −0.19 | 0.22 | −0.13 | 0.10 | −0.12 | 1.00 | |
| 8. Population density (population/km$^2$) | 2577.3 | 1768.9 | 2211.0 | 405.8 | 10,227.9 | −0.29 | 0.10 | 0.14 | **0.38** | 0.29 | **0.41** | −0.28 | 1.00 |
| 9. Land value (yen/m$^2$) | 67,958 | 21,644 | 66,356 | 29,600 | 109,800 | **−0.48** | 0.07 | 0.20 | **0.43** | **0.41** | **0.59** | **−0.32** | **0.57** |

*Bold shows significance $p < 0.05$.

**Table 3**
Association between neighborhood environment and diabetes mellitus (N = 8904).

| | | Crude | | Model 1 (adjusted for neighborhood environment covariates) | | Model 2 (adjusted for individual covariates) | | Model 3 (adjusted for neighborhood and individual covariates) | |
|---|---|---|---|---|---|---|---|---|---|
| | | OR | 95% CI | OR | 95% CI | OR | 95% CI | OR | 95% CI |
| GIS-Slopes (IQR: 1.48°) | | 0.93 | 0.86–1.01 | 0.98 | 0.88–1.09 | 0.93 | 0.85–1.01 | 0.98 | 0.88–1.10 |
| Perception of difficulty to walk due to slopes and stairs in the community (IQR: 14.5%) | | 1.07 | 0.94–1.22 | 1.01 | 0.88–1.17 | 1.07 | 0.94–1.22 | 1.05 | 0.92–1.20 |
| GIS-Number of grocery stores (IQR:16) | | 0.99 | 0.89–1.09 | | | 0.97 | 0.88–1.08 | | |
| Perception of access to grocery stores for fresh vegetables and fruits in the community (IQR:16.3%) | | 1.11 | 0.98–1.27 | 0.99 | 0.85–1.16 | 1.10 | 0.96–1.26 | 0.99 | 0.84–1.16 |
| GIS-Number of parks (IQR:5) | | 1.16 | 1.01–1.33 | 1.07 | 0.89–1.57 | 1.16 | 1.002–1.34 | 1.15 | 0.98–1.35 |
| Perception of access to park and sidewalk (IQR: 19.4%) | | 1.26 | 1.06–1.50 | | | 1.25 | 1.04–1.49 | | |
| GIS-Number of hospital (IQR:16) | | 0.89 | 0.78–1.00 | 0.92 | 0.76–1.11 | 0.88 | 0.77–0.999 | 0.91 | 0.75–1.11 |
| Population density (ref: low) | Middle | 1.22 | 0.94–1.57 | 1.09 | 0.81–1.48 | 1.19 | 0.92–1.56 | 1.02 | 0.75–1.38 |
| | High | 1.23 | 0.96–1.59 | 0.98 | 0.69–1.39 | 1.23 | 0.95–1.60 | 0.99 | 0.68–1.42 |
| Land value (IQR:35,017 yen/m$^2$) | | 1.21 | 1.02–1.44 | 0.98 | 0.76–1.27 | 1.20 | 1.01–1.44 | 1.06 | 0.84–1.33 |

Unit: interquartile range (IQR).
Model 1 adjusted for each hilly assessment, perception of access to grocery stores (unit: IQR), number of parks (unit: IQR), number of hospitals (unit: IQR), population density (tertile), and land value (unit: IQR).
Model 2 adjusted for age, sex, marital status, household number, income, working status, drinking, smoking, vegetable consumption, walking, going-out behavior, frequency of meeting, BMI (category), and depression.
Model 3 adjusted for covariates in Model 1 and 2.

**Table 4**
Association between neighborhood environment and poorly controlled diabetes mellitus (N = 8904).

| | | Crude | | Model 1 (adjusted for neighborhood environment covariates) | | Model 2 (adjusted for individual covariates) | | Model 3 (adjusted for neighborhood and individual covariates) | |
|---|---|---|---|---|---|---|---|---|---|
| | | OR | 95% CI | OR | 95% CI | OR | 95% CI | OR | 95% CI |
| GIS-Slopes (IQR: 1.48°) | | 0.86 | 0.75–0.98 | 0.83 | 0.70–0.99 | 0.85 | 0.74–0.97 | 0.82 | 0.70–0.97 |
| Perception of difficulty to walk due to slopes and stairs in the community (IQR: 14.5%) | | 1.11 | 0.90–1.36 | 1.11 | 0.89–1.37 | 1.08 | 0.88–1.33 | 1.08 | 0.88–1.33 |
| GIS-Number of grocery stores (IQR:16) | | 0.98 | 0.84–1.15 | | | 0.98 | 0.84–1.14 | | |
| Perception of access to grocery stores for fresh vegetables and fruits in the community (IQR:16.3%) | | 1.13 | 0.93–1.38 | 1.05 | 0.84–1.32 | 1.11 | 0.91–1.35 | 1.03 | 0.83–1.28 |
| GIS -Number of parks (IQR:5) | | 0.99 | 0.77–1.26 | 0.91 | 0.72–1.16 | 0.99 | 0.78–1.27 | 0.93 | 0.73–1.17 |
| Perception of access to park and sidewalk (IQR: 19.4%) | | 1.17 | 0.88–1.56 | | | 1.14 | 0.85–1.52 | | |
| GIS-Number of hospital (IQR:16) | | 0.88 | 0.72–1.08 | 1.05 | 0.78–1.40 | 0.88 | 0.73–1.07 | 1.03 | 0.77–1.36 |
| Population density (ref: low) | Middle | 1.36 | 0.91–2.02 | 1.20 | 0.76–1.90 | 1.34 | 0.90–1.99 | 1.18 | 0.75–1.84 |
| | High | 1.23 | 0.83–1.83 | 1.07 | 0.61–1.90 | 1.20 | 0.80–1.78 | 1.00 | 0.57–1.76 |
| Land value (IQR:35,017 yen/m$^2$) | | 1.16 | 0.87–1.55 | 0.90 | 0.64–1.28 | 1.17 | 0.88–1.57 | 0.92 | 0.66–1.28 |

Unit: interquartile range (IQR).
Model 1 adjusted for each hilly assessment, perception of access to grocery stores (unit: IQR), number of parks (unit: IQR), perception of access to home doctor (unit: IQR), and population density (tertile).
Model 2 adjusted for age, sex, marital status, household number, income, working status, drinking, smoking, vegetable consumption, walking, going-out behavior, frequency of meeting, BMI (category), and depression.
Model 3 adjusted for covariates in Model 1 and 2.

**Table 5**
Association between neighborhood environment and poorly uncontrolled diabetes mellitus among participants of under-treatment of DM (N = 1007).

| | | Crude | | Model 1 (adjusted for neighborhood environment covariates) | | Model 2 (adjusted for individual covariates) | | Model 3 (adjusted for neighborhood and individual covariates) | |
|---|---|---|---|---|---|---|---|---|---|
| | | OR | 95% CI | OR | 95% CI | OR | 95% CI | OR | 95% CI |
| GIS-Slopes (IQR: 1.48°) | | 0.85 | 0.73–0.99 | 0.76 | 0.62–0.94 | 0.84 | 0.72–0.99 | 0.73 | 0.59–0.90 |
| Perception of difficulty to walk due to slopes and stairs in the community (IQR: 14.5%) | | 1.03 | 0.83–1.29 | 1.10 | 0.84–1.44 | 1.04 | 0.83–1.31 | 1.14 | 0.86–1.51 |
| GIS-Number of grocery stores (IQR:16) | | 1.05 | 0.90–1.23 | | | 1.05 | 0.90–1.22 | | |
| Perception of access to grocery stores for fresh vegetables and fruits in the community (IQR:16.3%) | | 1.14 | 0.93–1.38 | 1.23 | 0.95–1.59 | 1.12 | 0.92–1.38 | 1.24 | 0.94–1.62 |
| GIS -Number of parks (IQR:5) | | 0.90 | 0.69–1.17 | 0.80 | 0.60–1.07 | 0.88 | 0.67–1.14 | 0.78 | 0.58–1.05 |
| Perception of access to park and sidewalk (IQR: 19.4%) | | 0.97 | 0.70–1.34 | | | 0.92 | 0.66–1.28 | | |
| GIS-Number of hospital (IQR:16) | | 0.96 | 0.77–1.21 | 1.12 | 0.78–1.62 | 1.00 | 0.79–1.26 | 1.20 | 0.82–1.75 |
| Population density (ref: low) | Middle | 1.05 | 0.65–1.71 | 0.86 | 0.49–1.51 | 1.04 | 0.63–1.71 | 0.83 | 0.47–1.49 |
| | High | 0.98 | 0.62–1.55 | 0.80 | 0.39–1.67 | 0.90 | 0.56–1.44 | 0.74 | 0.35–1.58 |
| Land value (IQR:35,017 yen/m$^2$) | | 1.00 | 0.72–1.39 | 0.77 | 0.51–1.18 | 1.00 | 0.71–1.40 | 0.77 | 0.49–1.20 |

Unit: interquartile range (IQR).
Model 1 adjusted for each hilly assessment, perception of access to grocery stores (unit: IQR), number of parks (unit: IQR), perception of access to home doctor (unit: IQR), and population density (tertile).
Model 2 adjusted for age, sex, marital status, household number, income, education, working status, drinking, smoking, vegetable consumption, walking, going-out behavior, frequency of meeting, BMI (category), and depression.
Model 3 adjusted for covariates in Model 1 and 2.

slope impact being similar to that for those who were not undergoing treatment for DM.

As for the sensitivity analysis, excluding those who did not go out and were obese, we confirmed similar results. A high GIS-slope was not associated with DM (OR: 0.97, 95% CI: 0.89–1.05), but showed a protective effect for poorly controlled DM (OR: 0.83, 95% CI: 0.70–0.99), which was also confirmed among those with treated DM (OR: 0.70, 95% CI: 0.56–0.89) in the fully adjusted model.

## 4. Discussion

In this study, we found that a hilly neighborhood environment was not significantly associated with DM, but did show a protective effect, with a 1 interquartile range (IQR) increase (1.48°) in slope in the neighborhood associated with an 18% decrease in the risk of poorly controlled DM. This protective effect against poorly controlled DM was confirmed when we limited the analysis to those who were undergoing treatment for DM, suggesting that DM treatment status was not a confounder of the association.

The result of this study is consistent with that of a previous report showing lower rates of self-reported DM among adults living in a neighborhood with steeper slopes than those living in a flat area (Villanueva et al., 2013). We have added new evidence that living in a hilly environment has a protective effect on poorly controlled DM, but not DM per se, among older people whose DM status was determined by objective measurement (i.e., $HbA_{1c}$). Furthermore, reverse causation is unlikely, as analysis was restricted to those living in the same city for 10 years or more.

There are three possible preventive mechanisms of living in a hilly neighborhood environment against poorly controlled DM. First, daily walks around the neighborhood may enable “unconscious” physical training that protects against poorly controlled DM. It is known that physical training increases glucose uptake due to the redistribution of GLUT4 glucose transporters (Cartee, 2015; Richter and Hargreaves, 2013), which may function to prevent poorly controlled DM. Previous studies reported that a certain amount of motivation is needed to undertake walking to prevent DM (Kabeya et al., 2016; Honda et al., 2015). In our study, walking duration was not associated with poorly controlled DM, but a hilly environment showed a significant preventive effect, suggesting that the hilliness of a walking path, and not walking duration per se, is important in the prevention of poorly controlled DM. Second, older people living in a hilly environment may be reluctant to use a bicycle, and so might walk than with those living in a flatter area. Third, a hilly environment tends to have beautiful scenery that may encourage older people to walk. We did not find an association between a hilly environment and walking duration or going-out behavior, possibly because walking behaviors were self-reported. A further study is needed to elucidate how a hilly environment is protective against poorly controlled DM using an objective measurement of physical activity, such as an actigraph (Lee et al., 2015).

Several limitations of this study need to be considered. First, the sample may not be representative of all populations as we limited the analysis to those who received a health check organized by the municipality under Japan's National Health Insurance system. Accordingly, the study findings may not be generalizable. Second, the school district examined may be too large for older people; that is, the area of daily activity may be smaller and may not cover the hilly environment within the school district. To account for this limitation, a further study using GPS tracking is needed. Third, $HbA_{1c}$ was assessed only once. However, examination of the trajectory of HbA1c could provide information on the potential time-dependent effect of a hilly environment on $HbA_{1c}$. Fourth, we could not distinguish between type 1 and type 2 DM, and it is therefore unclear whether the protective effect on poorly controlled DM is applicable to both types of DM. However, we considered that most of the cases in this study were type 2 DM as prevalence of type 1 DM was low (Maahs et al., 2010). Because type 1 DM involves the autoimmune destruction of pancreatic β-cells leading to insulin deficiency, which is treated with insulin injections, living in a hilly environment would not be associated with the prevention of poorly controlled DM.

Nevertheless, the current finding has possible health policy implications in the prevention of poorly controlled DM among older people. City planners should be advised to build slopes or steps within the community to prevent poorly controlled DM since it is not feasible to create a hilly landscape, and moving people to a hilly environment may have adverse effects on social networks.

In conclusion, a hilly neighborhood environment showed a protective effect on poorly controlled DM, but not DM per se, among older people in Japan. These findings can be used in residential planning policies with the goal of protecting against a rise in poorly controlled DM among older people.

## Competing interests

Nothing to declare.

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# Dental Status and Compression of Life Expectancy with Disability.

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## Abstract

This study examined whether the number of teeth contributes to the compression of morbidity, measured as a shortening of life expectancy with disability, an extension of healthy life expectancy, and overall life expectancy. A prospective cohort study was conducted. A self-reported baseline survey was given to 126,438 community-dwelling older people aged ≥65 y in Japan in 2010, and 85,161 (67.4%) responded. The onset of functional disability and all-cause mortality were followed up for 1,374 d (follow-up rate = 96.1%). A sex-stratified illness-death model was applied to estimate the adjusted hazard ratios (HRs) for 3 health transitions (healthy to dead, healthy to disabled, and disabled to dead). Absolute differences in life expectancy, healthy life expectancy, and life expectancy with disability according to the number of teeth were also estimated. Age, denture use, socioeconomic status, health status, and health behavior were adjusted. Compared with the edentulous participants, participants with ≥20 teeth had lower risks of transitioning from healthy to dead (adjusted HR, 0.58 [95% confidence interval (CI), 0.50-0.68] for men and 0.70 [95% CI, 0.57-0.85] for women) and from healthy to disabled (adjusted HR, 0.52 [95% CI, 0.44-0.61] for men and 0.58 [95% CI, 0.49-0.68] for women). They also transitioned from disabled to dead earlier (adjusted HR, 1.26 [95% CI, 0.99-1.60] for men and 2.42 [95% CI, 1.72-3.38] for women). Among the participants aged ≥85 y, those with ≥20 teeth had a longer life expectancy (men: +57 d; women: +15 d) and healthy life expectancy (men: +92 d; women: +70 d) and a shorter life expectancy with disability (men: -35 d; women: -55 d) compared with the edentulous participants. Similar associations were observed among the younger participants and those with 1 to 9 or 10 to 19 teeth. The presence of remaining teeth was associated with a significant compression of morbidity: older Japanese adults' life expectancy with disability was compressed by 35 to 55 d within the follow-up of 1,374 d.

**KEYWORDS:**

aging; dentition; epidemiology; longevity; oral health; survival analysis

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BMC Geriatrics

# An additive effect of leading role in the organization between social participation and dementia onset among Japanese older adults: the AGES cohort study

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**Abstract**

**Background:** Several previous studies reported social participation may reduce the incident of dementia; therefore, the type of positions held in the organization may relate to dementia onset. However, this hypothesis remains largely unknown. The purpose of the present study was to examine the additive effect of a leadership position in the organization on dementia onset and social participation among elderly people in a local community, according to data from a Japanese older adults cohort study.

**Methods:** Of 29,374 community-dwelling elderly, a total of 15,313 subjects responded to the baseline survey and were followed-up from November 2003 to March 2013. To evaluate the association between dementia onset and social participation as well as the role in the organization, we conducted Cox proportional hazard regression analysis with multiple imputation by age group (aged 75 years older or younger). The dependent variable was dementia onset, which was obtained from long-term care insurance data in Japan; independent variables were social participation and the role in the organization to which they belonged (head, manager, or treasurer). Covariates were sex, age, educational level, marriage status, job status, residence status, alcohol consumption, smoking status, and walking time, instrumental activities of daily living, depression, and medical history.

**Results:** During the follow-up period, 708 young-old elderly people (7.7%) and 1289 old-old elderly people (27.9%) developed dementia. In young-old elderly, relative to social non-participants, adjusted Hazard Ratio (HR) for dementia onset for participants (regular members + leadership positions) was 0.75 (95% confidence interval (CI), 0.64–0.88). Relative to regular members, adjusted HR for dementia onset for non-participants was 1.22 (95% CI, 1. 02–1.46), for leadership positions 0.81 (95% CI, 0.65–0.99). The results for old-old elderly participants did not show that any significantly adjusted HR between dementia onset and social participation, the role in the organization.

**Conclusions:** In young-old elderly people, social participation might have a positive effect on dementia onset, and holding leadership positions in organization could lead to a decrease in risk of dementia onset by almost 20% than regular members.

**Keywords:** Japan, Social participation, Leadership role, Dementia onset, Cohort study

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## Background

The number of dementia patients has increased dramatically because of the aging population worldwide. In 2010, more than 35 million people developed dementia and it is estimated that increase to 115 million people in 2050 [1]. The population aging rate of Japan is 26.7% in 2015, and prevalence of dementia will increase from 2.8 million (9.5%) in 2010 to 4.7 million (12.8%) in 2025 [2].

Identifying factors related to dementia onset is fundamental for improving preventive strategies; several systematic reviews and meta-analyses have identified some modifiable factors related to cognitive function or dementia onset [3–5], and social participation is one of the factors related to dementia onset [5]. The following, which are promoted by social participation, decrease risk of dementia: increasing physical activity (leaving one's home), accessing emotional support by expanding social networks, and increasing frequency of cognitive activity by obtaining a social role [6]; however, most of them only focused on absence of social participation and dementia onset or cognitive function, and the additive effect of leadership positions remains largely unknown.

Some observational studies investigated the relationship between leadership positions and health status. According to Ishikawa et al. [7], holding leadership positions on the association was related to a 12% risk reduction of mortality. Takagi et al. [8] suggested that performing leadership positions was significantly related to low odds ratio (OR) for depression for women (OR, 0.57; 95% CI, 0.37–0.88). Having leadership positions within civic groups may decrease the risk of dementia considerably; elderly people who manage the organization to which they belong perform various tasks or acquire roles that stimulate brain function or are beneficial to their health more so than compared with regular members, and this positively affects cognitive function.

The degrees of relationship between social participation and dementia onset may be different according to age group; in old-old (aged 75 or over), the age-related change has a greater effect on physical or mental health than in the young-old (aged 65–74) [9], and social participation can be a burden to the old-old. Therefore, to examine the relationship between social participation and dementia onset by age group is needed.

The purpose of the present study was to assess the additive effect of leadership positions in civic groups on the association between dementia onset and social participation among older adults in a local community, using data from a large cohort study (the Aichi Gerontological Evaluation Study: AGES).

## Methods

### Data

This study was based on data from the Aichi Gerontological Evaluation Study (AGES) project as a part of the Japan Gerontological Evaluation Study (JAGES). JAGES is a largely Japanese prospective cohort study aiming to find out the details of related factors for major health problems among the older adults, such as depression, dementia, or functional deterioration [10, 11].

### Participants

Participants were chosen from within six municipalities in Aichi prefecture, consisting of urban, semi-urban, and rural settings. Out of 49,707 older adults aged 65 years and older in a local community who did not receive public long-term care insurance benefits, 29,374 individuals were selected using two methods: random sampling in two larger municipalities, and a complete survey in four semi-urban or rural municipalities. In October 2003, we conducted the baseline mail survey. A total of 15,313 individuals completed the baseline self-administrated questionnaire and followed up from November 2003 to March 2013. To identify the predictive factors for dementia onset, we involved relatively healthy older adults, and excluded individuals with any premonitory symptoms of dementia, such as being unable to walk, take a bath or use a toilet independently. Individuals who developed dementia within two years of the baseline were also excluded to clarify the relationship between dementia onset and initial conditions (Fig. 1).

The Ethics Committee on Research of Human Subjects at Nihon Fukushi University approved this study protocol.

### Measurements

#### *Incident dementia*

Dementia onset was determined using disabling dementia, which is defined as incident functional disability with dementia. This was obtained from long-term care insurance data managed by local municipalities, as described previously [12]. Briefly, the degree of functional disability was evaluated according to a two-step procedure: on-site assessment of physical and mental condition by an agent from the home care provider, and further assessment by the Long-term Care Approval Board, consisting of health care professionals (doctors, nurses, caseworkers, or others) that referenced the results of on-site assessment and the primary physician's report, which is a standard form for assessing medical conditions and physical functions by a home physician [2]. Dementia was determined according to the Degree of Independence in Daily Living for Elderly with Dementia (Dementia Scale) [13, 14]. This scale was developed by the Japanese Ministry of Health, Labour

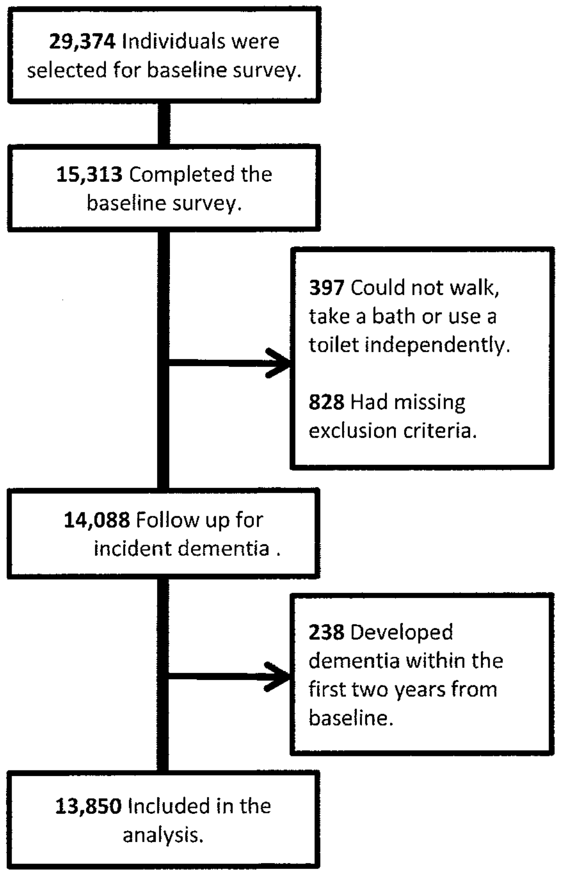


**Fig. 1** Flow of participants thorough the study

and Welfare, and health professionals in Japan use it to assess physical and cognitive function and classify individuals into levels I–IV and M. Level I means that the individuals have symptoms of dementia, but will be able to maintain an independent daily life. Level II indicates that the individuals show some symptoms and behaviors causing trouble in their daily life or some difficulties with communication, but could continue to live independently if monitored. Level III indicates that the individuals have the same symptoms as Level II patients, but more frequently, and sometimes require care to support their daily lives. Level IV indicates that the individuals have the same symptoms as in Level III, but more frequently, and always need care in their daily lives. Level M indicates individuals with severe mental or physical diseases and behavioral disorders, who require specialized medical care. We defined individuals scoring levels II to IV or M as having dementia. A previous study has shown that the Dementia Scale is well correlated with the Mini Mental State Exam score [15].

### *Social participation and leadership positions in an organization*

The scale of social participation was taken from the Japanese General Social Survey [16], and categorized organizations into following eight types: neighborhood association, senior citizen club/fire-fighting team, religious group, political organization or group, industrial or trade association, volunteer group, citizen or consumer group, hobby group, and sports group or club. Participants were asked whether they were members of each association and their frequency of participation; those who answered "I do not participate in any organization" and "participate in the organization" but "very little" for frequency of participation were classified as "non-members". Therefore, the individuals who belonged to one or more associations were asked their position in the organization; those who serve as head, manager or treasurer were categorized as having "leadership positions", while others were classified as "regular members".

### *Covariates*

In this study, demographic variables, health behavior, and health status were included as covariates.

Demographic variables consisted of sex, age (65–69, 70–74, 75–79, 80–84, 85 years and over), educational attainment (less than 10 years, 10 or more years), marital status (married, other), residential status (solitary, other), employment (worker, non-worker), health behavior including alcohol status (drinker, non-drinker), smoking status (smoker/ former smoker, never-smoker), walking time (less than 30 min/day, 30 min/day and longer), health status included instrumental ADLs (IADLs) (the subscale of Tokyo Metropolitan Institute of Gerontology Index of Competence: TMIG-IC [17]), medical history (heart disease, stroke, hypertension, diabetes), depression (Geriatric Depression Scale - Short Version, GDS-SV [18]). Those who earned full score for TMIG-IC were categorized as "high", the GDS-SV cut-off was 5, as in a previous study [19], and subjects who scored above the cut-off were categorized as "depressed".

### Statistical analysis

To handle missing data, we carried out multiple imputation with full conditional specification, and created 50 multiply imputed datasets [20]. Imputed model included incident of dementia, social participation and leading positions, demographic variables, health behavior, and health status. Therefore, Cox proportional hazard models were used on these datasets. These estimates and their standard errors were combined using Rubin's rules [21], and Hazard Ratio (HR) or confidence interval (CI) was calculated. For comparison, Cox proportional hazards model was used on the subset of complete case data.

We calculated HRs for incident of dementia according to social participation and by age group (young-old, old-old) using the Cox proportional hazards model to examine the relationship between these factors, and carried out a similar analysis model that exchanged social participation and leading role variables to assess the additive effects of leading positions. We used a level of significance of less than 5% in all analyses. SAS 9.4 (SAS Institute, Cary, NC) was used for all calculations.

## Results

Of 29,374 individuals, 15,313 completed the baseline survey (response rate, 52.1%). Non-responders were younger, and there was no difference between sexes. Of the 15,313 subjects, 13,850 were included in the analysis. A total of 1463 individuals were excluded from analysis; 397 could not walk, take a bath or use a toilet independently, 828 had missing exclusion criteria, and 238 developed dementia within the first two years from baseline (Fig. 1). The mean follow-up period was 7.9 years (standard deviation, 2.4 years), and the number of missing values across each variable varied between 0 (0%) and 933 (10.1%) in young-old, 0 (0%) and 721 (15.6%) in old-old; the total number of individuals who had incomplete data among the all variables was 2629 (28.5%) in young-old and 1663 (36.0%) in old-old. The number of individuals who died during follow-up was 1611 (17.5%) in young-old and 1363 (29.5%) in old-old.

Of the 13,850 subjects of the analysis, 9234 (66.7%) were young-old and 4616 (33.3%) were old-old. Of these young-old, 708 (7.7%) developed dementia, 3003 (32.5%) were non-members, 2514 (27.2%) were regular members, 2784 (30.1%) were in leadership positions, whereas in old-old, 1289 (27.9%) developed dementia, 1774 (38.4%) were non-members, 1289 (27.9%) were regular members, and 832 (18.0%) were in leadership positions (Table 1). Table 2 shows that the incidence of dementia onset increased with age. The incidence in each category of old-old individuals was much higher than in young-old participants.

The results of Cox proportional hazards model on the imputed data indicated that the crude HR for dementia onset for regular members or those holding leadership positions, compared with non-members, was 0.65 (95% CI, 0.55–0.75), and adjusted HR was 0.75 (95%CI, 0.64–0.88) in young-old, whereas crude HR was 0.73 (95% CI, 0.64–0.82), but adjusted HR was non-significant in old-old (Table 3).

Table 4 shows the relationship between having a leading role and dementia onset. In young-old, both crude HR and adjusted HR for dementia onset for non-members, relative to regular members, were significant (crude HR, 1.38; 95%CI, 1.15–1.65, adjusted HR, 1.22; 95% CI, 1.02–1.46), and crude HR or adjusted HR for leadership positions were also significant (crude HR, 0.76; 95% CI, 0.61–0.94, adjusted HR, 0.76; 95% CI, 0.65–0.99); however, in the old-old group, there was not significant adjusted HR.

## Discussion

The present study showed that social activity non-members have a greater risk of incident dementia than social activity members, and members in leadership positions have a significantly lower risk compared with the non-leading members in the young-old group. However, in the old-old group, non-significant differences in dementia risk were observed. These findings seem to suggest that social participation might be effective for prevention of dementia, and this preventive effect could become stronger in the young-old group if leadership positions are taken.

Our findings are broadly consistent with those of previous studies. Kuiper et al. [4] assessed the relationship between social participation and incidence of dementia through meta-analysis. The results of this analysis revealed that individuals with less social participation had a higher risk of dementia onset relative to subjects with higher levels of social participation (RR, 1.41; 95% CI, 1.13–1.75). Although the mechanism underlying the association between social participation and incidence of dementia was not identified, the following pathways were possible: 1) higher level of physical activity due to leaving the home may promote cognitive reserve [6], 2) frequent contact with others may cause positive emotional states such as increased self-esteem, social competence, and adequate mood, which lead to lower stress levels [22], 3) performing various activities (e.g., engaging in a hobby, calculating the scores of games) that stimulate cognitive function serves to prevent a cognitive decline ("use it or lose it" theory) [23]. The present study implies that social participation might have a suppressive effect on the incidence of dementia, but the effect may be different based on participation in social activities. Although the reasons for the additive effect of a leadership role on incidence of dementia are not fully understood, one reason might be the difference in the frequency of social participation. Compared with regular members, individuals who take on leadership roles such as president, facilitator or treasurer have more frequent opportunities for social participation, and also take responsibility for actions to manage group activities (e.g., holding meetings, planning activities, and communicating with regular members). In this study, the proportion of individuals engaging in group activities more than once a month was higher among those in leadership positions than regular members (81.7% vs 64.8%, data not shown). Higher frequency of

**Table 1** Initial characteristics of the participants

| | | Young-old (n = 9234) | | Old-old (n = 4616) | |
|---|---|---|---|---|---|
| | | n | % | n | % |
| Dementia onset | No-dementia | 8526 | 92.3 | 3327 | 72.1 |
| | Dementia | 708 | 7.7 | 1289 | 27.9 |
| Sex | Male | 4714 | 51.1 | 2080 | 45.1 |
| | Female | 4520 | 48.9 | 2536 | 54.9 |
| Social participation | Non-participants | 3003 | 32.5 | 1774 | 38.4 |
| | Regular-members | 2514 | 27.2 | 1289 | 27.9 |
| | Leadership positions | 2784 | 30.1 | 832 | 18.0 |
| | Missing | 933 | 10.1 | 721 | 15.6 |
| Age | 65–69 | 5082 | 55.0 | – | – |
| | 70–74 | 4152 | 45.0 | – | – |
| | 75–79 | – | – | 2827 | 61.2 |
| | 80–84 | – | – | 1269 | 27.5 |
| | ≥ 85 | – | – | 520 | 11.3 |
| Educational attainment | < 10 yrs | 5286 | 57.2 | 2849 | 61.7 |
| | ≥ 10 yrs | 3896 | 42.2 | 1712 | 37.1 |
| | Missing | 52 | 0.6 | 55 | 1.2 |
| Marital status | Married | 7343 | 79.5 | 2647 | 57.3 |
| | Single | 1766 | 19.1 | 1891 | 41.0 |
| | Missing | 125 | 1.4 | 78 | 1.7 |
| Living arrangement | Living with others | 8294 | 89.8 | 3900 | 84.5 |
| | Living alone | 779 | 8.4 | 569 | 12.3 |
| | Missing | 161 | 1.7 | 147 | 3.2 |
| Occupational status | Employed | 2806 | 30.4 | 608 | 13.2 |
| | Not employed | 6296 | 68.2 | 3916 | 84.8 |
| | Missing | 132 | 1.4 | 92 | 2.0 |
| Walking time (per day) | < 30 min | 2794 | 30.3 | 1586 | 34.4 |
| | ≥ 30 min | 5523 | 59.8 | 2548 | 55.2 |
| | Missing | 917 | 9.9 | 482 | 10.4 |
| Medical history | | | | | |
| Heart disease | No | 8164 | 88.4 | 3809 | 82.5 |
| | Yes | 1070 | 11.6 | 807 | 17.5 |
| Stroke | No | 9119 | 98.8 | 4520 | 97.9 |
| | Yes | 115 | 1.2 | 96 | 2.1 |
| Hypertension | No | 6266 | 67.9 | 2905 | 62.9 |
| | Yes | 2968 | 32.1 | 1711 | 37.1 |
| Diabetes | No | 8168 | 88.5 | 4157 | 90.1 |
| | Yes | 1066 | 11.5 | 459 | 9.9 |
| Alcohol consumption | Non-drinker | 5535 | 59.9 | 3317 | 71.9 |
| | Drinker | 3582 | 38.8 | 1181 | 25.6 |
| | Missing | 117 | 1.3 | 118 | 2.6 |
| Smoking | Never smoked | 5312 | 57.5 | 2800 | 60.7 |
| | Past smoker/smoker | 3615 | 39.1 | 1601 | 34.7 |
| | Missing | 307 | 3.3 | 215 | 4.7 |

**Table 1** Initial characteristics of the participants *(Continued)*

| | | Young-old (n = 9234) | | Old-old (n = 4616) | |
|---|---|---|---|---|---|
| | | n | % | n | % |
| Depression | normal | 6004 | 65.0 | 2591 | 56.1 |
| | depressed | 2304 | 25.0 | 1316 | 28.5 |
| | Missing | 926 | 10.0 | 709 | 15.4 |
| IADL | High | 7649 | 82.8 | 3196 | 69.2 |
| | Low | 1335 | 14.5 | 1182 | 25.6 |
| | Missing | 250 | 2.7 | 238 | 5.2 |

social participation may help to strengthen the health benefits of social participation [24], or enable individuals to obtain information that supports a healthy lifestyle [25]. Socially-responsible activities may improve the quantity or quality of stimulation of the brain's cognitive function, or maintain better mental health [8]. However, we did not investigate the type of activity, or use laboratory data, so this is only speculation. As little is known about the mechanism behind the increased positive effect of leadership on cognitive function, further investigation is needed.

In contrast to the young-old group, there were no significant relationships between social participation or leadership and dementia onset in the old-old group. These results support the findings of previous studies [11, 26, 27]. Iwasa et al. [26] suggested that social participation was not attributed to the prevention of cognitive decline among Japanese community-dwelling elderly aged 70 years and over, based on the data from a five-year prospective cohort study. One possible explanation is that as the prevalence of individuals with health problems is much higher in this group than in the young-old

**Table 2** Incidence rates (1000 person-years) of dementia onset by sex, age, and educational attainment

| | Young-old | | Old-old | |
|---|---|---|---|---|
| | Incidence rate | 95% CI | Incidence rate | 95% CI |
| Sex | | | | |
| Male | 9.3 | 8.3–10.4 | 36.2 | 32.8–39.9 |
| Female | 9.3 | 8.4–10.5 | 43.9 | 40.7–47.4 |
| Age | | | | |
| 65–69 | 5.5 | 4.8–6.4 | – | – |
| 70–74 | 14.3 | 13.0–15.8 | – | – |
| 75–79 | – | – | 31.3 | 28.8–34.1 |
| 80–84 | – | – | 51.5 | 46.3–57.3 |
| ≥ 85 | – | – | 83.3 | 71.8–96.8 |
| Educational attainment | | | | |
| < 10 yrs | 10.2 | 9.2–11.3 | 42.6 | 39.4–45.9 |
| ≥ 10 yrs | 8.9 | 7.8–10.1 | 36.2 | 32.6–40.1 |

**Table 3** Relationship between social participation and dementia onset

| | Crude | | Adjusted | |
|---|---|---|---|---|
| | HR | 95% CI | HR | 95% CI |
| Young-old(*n* = 9234) | | | | |
| Social participation | | | | |
| Non-participants | reference | | reference | |
| Participants | 0.64 | 0.55–0.75 | 0.75 | 0.64–0.88 |
| Old-old(*n* = 4616) | | | | |
| Social participation | | | | |
| Non-participants | reference | | reference | |
| Participants | 0.73 | 0.65–0.82 | 0.91 | 0.81–1.03 |

Adjusted for sex, age, educational attainment, marital status, living arrangement, occupational status, walking time, medical history, alcohol consumption, smoking, depression, and IADL

group, the relationship between social participation and dementia onset in old-old elderly may be relatively weaker than that in the young-old group. In the present study, health status such as diabetes, depression, and Independent Activity of Daily Living were strongly related to dementia onset (Additional file 1); therefore, these health problems may be the major correlated factors of incidence of dementia in old-old elderly. However, we may have underestimated this relationship in the old-old group for several reasons. First, the percentage of individuals who had died or moved out during follow-up period was much higher (29.5%) among the old-old than young-old (17.5%), which means that about 30% of the old-old participants had died or moved out before developing dementia. Secondly, the presence or absence of social participation and leadership were assessed at baseline, but prior experience was not assessed; therefore, old-old participants who had experienced social participation or leadership before the assessment but had already retired from these activities at the time of baseline assessment were categorized as non-members. Thus, in this study, the category of non-member in the old-old group contained those who were non-members later in life and those who were members before the study period. These reasons can be attributed to underestimation of the association between social participation and incidence of dementia. Further studies of the association of social participation with dementia onset in old-old elderly people are needed.

**Table 4** Relationship between having a leadership positions and dementia onset

| | Crude | | Adjusted | |
|---|---|---|---|---|
| | HR | 95% CI | HR | 95% CI |
| Young-old(*n* = 9234) | | | | |
| Regular-members | reference | | reference | |
| Non-participants | 1.38 | 1.15–1.64 | 1.22 | 1.02–1.46 |
| Leadership positions | 0.76 | 0.61–0.94 | 0.81 | 0.65–0.999 |
| Old-old(*n* = 4616) | | | | |
| Regular-members | reference | | reference | |
| Non-participants | 1.30 | 1.15–1.48 | 0.99 | 0.86–1.13 |
| Leadership positions | 0.86 | 0.72–1.02 | 0.98 | 0.83–1.14 |

Adjusted for sex, age, educational attainment, marital status, living arrangement, occupational status, walking time, medical history, alcohol consumption, smoking, depression, and IADL

This study has several limitations. First, the incidence of dementia in this study was obtained from the results of an examination and judgment by the Certification Committee of Needed Long-Term Care in the participant's municipality. Therefore, underestimation of dementia incidence might have occurred, because every dementia patient does not necessarily submit an application to the Certification Committee. Second, as the type of dementia was not assessed, such as Alzheimer disease, vascular dementia, or Lewy body dementia, the effect of social participation or leadership on each type of dementia remained unclear. Third, the response rate of the baseline survey was 52.1%, meaning that non-responders may have induced selection bias. In this study, the characteristics of non-responders were unknown, but we think it is possible that old-old people or those with lower health status may have been less likely to respond to the survey. There may therefore have been differences in baseline characteristics between study participants and non-participants. Fourth, as the experience of social participation or leadership before the baseline survey was not assessed, the relationship between social participation and dementia onset may be affected by the results of these factors, especially among old-old participants. Future studies evaluating this association should take into account the subject's experience of social participation and leadership before the baseline survey. Fifth, this study could not identify which types of social activity or leadership were related to the incidence of dementia. Further studies are needed to examine this issue, especially qualitative studies that assess the influence of social participation or a leadership role on older adults' daily lives. Finally, as the assessment of social participation and leadership were carried out only at the baseline survey, the influence of change in status of participation during the follow-up period on the relationship was not clear.

In summary, despite the above-mentioned limitations, this study revealed that social participation might have a repressive effect on the incidence of dementia and also leadership within the activity group might have stronger positive effect on dementia incidence among young-old adults. These finding should be used to encourage young-old adults to participate in and take leadership positions in social activity organizations.

## Conclusions

In young-old elderly people, social participation might have a positive effect on the prevention of dementia onset, and leadership within a group may lead to a reduction of risk of dementia onset of almost 20%, compared with regular members.

## Additional file

**Additional file 1:** Relationship between having a leadership positions and dementia onset in old-old. (XLSX 13 kb)

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### Availability of data and materials

The dataset is available to the researcher who submit research proposal to the JAGES committee (https://www.jages.net/) and permit to use this dataset.

### Authors' contributions

Study conception and design: YN TS SK TT KS HK KM TA KK. Acquisition of data: KK The JAGES Group. Analysis and interpretation of the data: YN KM TA. Drafting the paper: YN TA. Critical revision: TS SK TT KS HK MK TA KK. All authors read the manuscript and approved the submission.

### Ethics approval and consent to participate

We obtained written consent from the participants. The protocol of AGES project was approved by the Ethics Committee on Research of Human Subjects at Nihon Fukushi University.

### Consent for publication

Not Applicable.

### Competing interests

The authors declare that they have no competing interests.

## Publisher's Note



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原　著

# 「健康交流の家」開設による健康増進効果の検証

# Effectiveness of “Kenko-koryu-no-ie” in improving health

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## 抄 録

【目的】

まちづくりによる地域住民の健康増進を図る為 , 愛知県東海市では , 地域住民の交流と健康づくりの機能を併せ持つ「健康交流の家」の開設が進められている。本研究の目的は ,「健康交流の家」の開設に伴う , 地域住民の健康行動及び主観的健康感の変化を検証することである。

【方法】

2014 年 2 月 ,「健康交流の家」を利用している団体責任者 17 名および施設利用者 326 名の計 333 名に対し , 開設前後における施設利用状況 , 健康行動および主観的健康感の変化に関する自記式質問紙調査を実施した。

【結果】

施設利用者 280 名より回答を得 , 無効回答を除いた 221 名（有効回答率 66.4%）を分析対象とした。開設後 , 集団では 7 団体（41.2%）の施設の利用頻度が増加し , また , 個人の行動変化においても , 97 名（51.3%）の利用頻度の増加がみられた。健康行動の変化は , 歩行機会:84 名 (38.0%) , 外出の機会:88 名 (39.8%) , 会話の機会:117 名 (52.9%) , 趣味の会への参加機会：56 名（25.3%）, スポーツの会への参加機会：41 名（18.6%）で増加がみられた。また主観的健康感の変化は , 82 名（37.1%）が開設前後で , 良い方向に変化したと回答した。健康行動と主観的健康感の変化との関連を分析したところ , 健康行動が増加した者は , 主観的健康感も良い方向に変化した割合が有意に高かった。

【考察】

「健康交流の家」の開設は , 地域住民の身体活動や社会活動といった健康行動を促進し , 主観的健康感を改善させたと考えられた。本結果より,「健康交流の家」は , まちづくりによる一次予防に寄与できる可能性が示唆された。

## Abstract

Background: A “Kenko-koryu-no-ie” approach was started in Tokai city, Aichi Prefecture, to improve the health of the local residents through community development. This study examines the changes in health behavior of local residents and their subjective health, as brought about by the facility.

Methods: Participants included 17 organization representatives and 326 elderly people who used the facility. The participants responded to a self-reported questionnaire about facility usage conditions, health behavior and subjective health before and after the commencement of the facility, which was started one year ago.

Results: Answers were received from 280 facility users; however, due to invalid answers, only 221 respondents could be included in the study (effective response rate 66.4%). Among the respondents, 7 organizations

(41.2%) and 97 participants (51.3%) reported that the frequency of usage of the facility had increased. Changes in the participants' health behavior were observed with increased opportunities to walk (38.0%), opportunities to go out (39.8%), opportunities for conversation (52.9%), opportunities to participate in hobby meetings (25.3%), opportunities to participate in sports meetings (18.6%). Furthermore, 82 people (37.1%) reported positive changes in subjective health since the commencement of the facility. Analysis of the relevance of the changes between the health behavior and subjective health indicated that those who exhibited improvement in health behavior also reported an improvement in their subjective health.

Conclusion: The establishment of the "Kenko-koryu-no-ie" helped in the promotion of health behavior in local residents and improved their subjective health. Through community development, the concept of "Kenko-koryu-no-ie" could be used as a primary prevention program for the elderly.

キーワード：まちづくり , 高齢者 , 健康行動 , 社会参加 , 主観的健康感

Key words：community development, elderly people, health behavior, social participation, subjective health

## Ⅰ 緒言

介護予防において , 高齢者個人へのアプローチだけでなく , 生活環境の調整や , 生きがい・役割を持てる環境づくりなど , 地域全体に働きかけるポピュレーション戦略が重視されている[1) 2)]。健康日本 21（第二次）でも , 社会参加の機会の増加など地域におけるソーシャル・キャピタルに着目した「社会的環境の質向上」による健康格差の縮小が基本的な方向として掲げられ[3)], また , 介護予防マニュアル（改訂版）においても ,「まちづくり」による一次予防への方向転換が謳われている[4)]。しかし , その具体的な方法については未だ模索段階にあり , 多くの市町村でまちづくりによる健康づくりや介護予防のモデル事業が行われているが , それらの効果を検証した報告は少ない。

本研究の対象である愛知県東海市では , こうしたまちづくりによる住民の健康増進を図るため , 地域住民の交流の促進と健康行動の増進を目的とした「健康交流の家」の開設が進められている。東海市 O 地区における「健康交流の家」は , 2013 年 4 月に開設された。開設後 , 施設の利用者数が増加し , 地域住民への健康効果が期待されている。地域住民の交流活性とそれに伴う健康行動の変化によって健康増進効果が得られれば ,「健康交流の家」は , まちづくりによる一次予防の先駆的な事例と考えることができる。

そこで , 本研究は ,「健康交流の家」の開設による , 開設前（2012 年 12 月）と開設 1 年後（2013 年 12 月）の 1 年間における , 地域住民の「健康交流の家」の利用状況と , 健康行動および主観的健康感の変化を検証することを目的とした。

## Ⅱ 調査方法

### 1. 調査対象

愛知県東海市の「健康交流の家」を利用している 17 団体の施設利用者（団体責任者 17 名を含む）272 名と , 交流（サロン）スペースを主に利用する , 団体に所属しない施設利用者 61 名の計 333 名を対象とした。

「健康交流の家」の概要

東海市は , 人口約 11 万人 , 高齢化率 19.5%（2013 年 4 月）, 要介護認定率 14.6%（2012 年 3 月）の自治体である。「健康交流の家」は，「東海市立敬老の家の設置及び管理に関する条例」をもとに，敬老の家に異なる機能を持たせた施設として，従来の和室を基本とした施設内容から，運動ができる機能とテーブル・椅子での交流を可能とする施設整備が進められており，O 地区ではこの「健康交流の家」に自治会集会所を合築している。敬老の家（「健康交流の家」）は , 高齢者に対し , 教養の向上 , レクリエーション等のための場を提供し , 高齢者の心身の健康増進を図ることを目的として , 市町村が設置する施設である。一方 , 自治会集会所は , 自治会の活動の場を確保することを目的として , 自治会が主体となり , 市町村の一部費用補助を受けて設置する施設である。O 地区ではこれら両施設を合築することにより，地域住民相互の親睦や住みよい生活環境の維持向上を図り，高齢者をはじめとした利用者が集いやすい施設としている。このように敬老の家（「健康交流の家」）と異なる施設を合築することで , 利用対象や活動内容が多様化し , また , 経済的にも一施設に予算を集中投下でき , 利便性や快適性の面で , 住民の意向をより反映した施設整備をすることができる。合築する施設は自治会集会所だけでなく，他地域の「健

康交流の家」では津波避難施設を併設することで，地域防災の拠点としての機能も併せ持たせている。また，施設の運営は，地域住民による主体的な運営がされている。O 地区の「健康交流の家」では，自治会長が指定管理者となり，地域のボランティアが主体的に運営に携わっている。このボランティア団体は，もともと地域で地元の助け合いの活動していた住民であり，平成 24 年度より実施されている「地域支えあい体制づくり事業」に則って登録団体となった団体である。

O 地区の「健康交流の家」の主な特徴としては，建築総面積 275.4m$^2$ の中に広い活動スペースの確保，利用者の憩いの場となる交流（サロン）スペースの設置，そこで応対するボランティアスタッフの存在などが挙げられる。広い活動スペースの確保は，両施設を合築し，間仕切りを解放することにより実現し，それによって，スポーツ活動や趣味活動などの活動内容の幅が広がり，収容人数を増やすことができている。開設前（2012 年 12 月）は，施設の利用団体およびその人数は 13 団体 168 名であったが，開設 1 年後（2013 年 12 月）には，17 団体 272 名に増加した。また，交流（サロン）スペースの設置と，そこで活動するボランティアスタッフの配置により，利用者にとって快適な憩いの場が提供されている。交流（サロン）スペースは，地域住民が自由に利用でき，そこではボランティアスタッフが，コーヒーや紅茶，お菓子（福祉作業所からの委託販売品）などを提供してくれる。こうした交流（サロン）スペースの設置により，スポーツや趣味の会などの活動前後にも利用者同士が交流を楽しむ機会が増えたり，活動に参加しない地域住民も，散歩途中などに気軽に立ち寄り，住民同士の交流する機会が増えている。さらに，ボランティアスタッフ自身も地域の高齢者が中心となって活動しているため，スタッフ自身の地域での役割づくりにも寄与している。また，その他の特徴として，テラス席の設置や明るい室内の雰囲気を作り出す採光方法など，住民と自治体との話し合いがもたれ，設立指針に住民の意向が直接反映されている。また，費用については，自治会集会所部分の一部は，住民自身が負担をしている。こうした設立段階からの運営も含めた住民参加と費用の自己負担などにより，O 地区の「健康交流の家」は，住民自らが作り上げた施設となっている。

### 2. 調査項目

主な調査内容は，利用者特性，施設の利用状況，健康行動，主観的健康感である。利用者特性および施設の利用状況については，団体利用における集団の活動変化についても把握するため，施設利用者を対象に個人の行動変化を調査する [ 調査票 2] に加え，団体責任者を対象とする [ 調査票 1] を作成した。

#### [調査票 1]：集団の変化

1）利用者特性

利用者特性の変化を，開設前（2012 年 12 月）と開設 1 年後（2013 年 12 月）について，年齢，性別，自宅から施設までの所要時間［1. 徒歩 10 分未満，2. 徒歩 10 ～ 30 分，3. 徒歩 30 分～ 1 時間，4. 徒歩 1 時間以上］を尋ねた。

2）団体の施設利用状況

施設の利用頻度は，開設前（2012 年 12 月）と開設 1 年後（2013 年 12 月）について，［1. 2 ～ 3 ヵ月に 1 回，2. 月 1 ～ 3 回，3. 週 1 回，4. 週 2 回，5. 週 3 回，6. 週 4 回，7. 週 5 回］の 7 段階で尋ね変化を分析した。

施設の利用内容は，開設前（2012 年 12 月）と開設 1 年後（2013 年 12 月）の変化を総括的に，［1. 良い方向に変化した，2. 変わらない，3. 悪い方向に変化した］の 3 段階で尋ねた。また，その活動内容を，［1. 体操・太極拳，2. 舞踏・ダンス，3. 囲碁・将棋・麻雀，4. 楽器演奏，5. コーラス・民謡，6. カラオケ，7. 俳句，8. 書道，9. 茶道，10. 手工芸，11. 絵画・絵手紙，12. 会話，13. 会議］について尋ねた。

#### [調査票 2]：個人の変化

1）個人の施設利用状況

施設の利用頻度は，開設前（2012 年 12 月）と開設 1 年後（2013 年 12 月）について，［1. 利用していない，2. 2 ～ 3 ヵ月に 1 回，3. 月 1 ～ 3 回，4. 週 1 回，5. 週 2 回，6. 週 3 回，7. 週 4 回，8. 週 5 回］の 8 段階で尋ねた。

施設の利用内容は，開設前（2012 年 12 月）と開設 1 年後（2013 年 12 月）の変化を総括的に，［1. 良い方向に変化した，2. 変わらない，3. 悪い方向に変化した］の 3 段階で尋ねた。また，その活動内容を，［1. 体操・太極拳，2. 舞踏・ダンス，3. 囲碁・将棋・麻雀，4. 楽器演奏，5. コーラス・民謡，6. カラオケ，7. 俳句，8. 書道，9. 茶道，10. 手工芸，11. 絵画・絵手紙，12. 会話，13. 会議］について尋ねた。

3）健康行動

健康行動は，開設前（2012 年 12 月）と開設 1 年後（2013 年 12 月）における，歩行する機会，外出する機会，会話の機会，趣味の会の参加機会，スポーツの会の参加機会について，その機会の増減を，

［1. 増加した , 2. 変わらない , 3. 減少した］の 3 段階で尋ねた。

4）主観的健康感

主観的健康感は , 現時点の健康状態を ,［1. とてもよい , 2. まあよい , 3. あまりよくない , 4. よくない］の 4 段階で尋ね , また , 開設前（2012 年 12 月）と開設 1 年後（2013 年 12 月）の変化を ,［1. 良い方向に変化した , 2. 変わらない , 3. 悪い方向に変化した］の 3 段階で尋ねた。

### 3. 調査方法および調査時期

本調査は , 自記式質問紙調査である。2014 年 2 月に ,「健康交流の家」を利用する団体責任者に対し ,［調査票 1］および［調査票 2］を直接配布し , その団体に所属する施設利用者に対しては , 団体責任者による［調査票 2］の配付・回収を依頼した。また , 団体に所属しない施設利用者に対しては , 交流（サロン）スペースに［調査票 2］を設置し , ボランティアから協力を呼びかけてもらった。

### 4. 分析方法および分析対象

利用者特性について , 記述統計を算出した。次に ,「健康交流の家」の利用状況と健康行動の変化との関連 , 健康行動と主観的健康感の変化との関連を検証するため , $\chi^2$ 検定を実施した。有意水準は 5%とし , 統計解析には , SPSSver.22.0 を用いた。

### ［調査票 1］：集団の変化

17 団体の団体責任者 17 名中 17 名より有効回答を得られたため , 17 名（有効回答率 100.0%）を分析対象とした。

### ［調査票 2］：個人の変化

団体に所属する施設利用者 272 名中 239 名（回収率 87.9%）, および団体に所属しない施設利用者 61 名中 41 名（回収率 67.2%）の計 280 名（回収率 84.1%）より回答を得た。そのうち , 無効回答の 59 名を除く , 221 名（有効回答率 66.4%）を分析対象とした。

### 5. 倫理的配慮

本研究は , 愛知県東海市の協力を得て実施した。対象者には , 研究目的と意義 , 無記名・プライバシーの配慮 , 拒否権などについて説明会ならびに文書を用いて説明し , 質問紙の返答をもって研究参加への同意を得たものと見なした。

## Ⅲ 結果

### 1. 利用者の変化

団体責任者向けの [ 調査票 1] により評価した利用者の変化は , 以下の通りであった。

1）年齢（表 1）

全年代において利用者は増加しており ,「65 歳未満」では , 24 名から 36 名（増加率；50.0%）,「65 ～ 75 歳未満」では , 88 名から 123 名（増加率；39.8%）,「75 歳以上」では , 51 名から 66 名（増加率；29.4%）の増加がみられた。

2）性別（表 1）

利用者は , 男女ともに増加しており ,「男性」では , 57 名から 83 名（増加率；45.6%）,「女性」では , 106 名から 142 名（増加率；34.0%）の増加がみられた。

3）自宅から施設までの所要時間（表 1）

徒歩所要時間が短いほど , 利用者数の増加率が高い傾向がみられた。「徒歩 10 分未満」では , 97 名から 157 名（増加率；61.9%）,「徒歩 10 ～ 30 分」では , 42 名から 62 名（増加率；47.6%）,「徒歩 30 分～ 1 時間」では , 17 名から 20 名（増加率；17.6%）の増加がみられたが ,「徒歩 1 時間以上」では , 増減はみられなかった。

表 1.　利用者の変化

| | 開設前 n=168 | | 開設後 n=272 | | 増加率 |
|---|---|---|---|---|---|
| | 人数 | (%) | 人数 | (%) | % |
| [年齢] | | | | | |
| 65 歳未満 | 24 | (14.7) | 36 | (16.0) | 50.0 |
| 65ー75 歳未満 | 88 | (54.0) | 123 | (54.7) | 39.8 |
| 75 歳以上 | 51 | (31.3) | 66 | (29.3) | 29.4 |
| [性別] | | | | | |
| 男性 | 57 | (35.0) | 83 | (36.9) | 45.6 |
| 女性 | 106 | (65.0) | 142 | (63.1) | 34.0 |
| [施設から自宅までの所要時間] | | | | | |
| 徒歩 10 分未満 | 97 | (57.7) | 157 | (62.5) | 61.9 |
| 徒歩 10ー30 分未満 | 42 | (25.0) | 62 | (24.7) | 47.6 |
| 徒歩 30 分ー1 時間未満 | 17 | (10.1) | 20 | ( 8.0) | 17.6 |
| 徒歩 1 時間以上 | 12 | ( 7.1) | 12 | ( 4.8) | 0.0 |
| 合計 | 168 | (100.0) | 272 | (100.0) | 61.9 |

無回答を除く

### 2. 施設の利用頻度の変化

1）団体の利用頻度の変化（表 2, 図 1-1）

団体責任者を対象とした [ 調査票 1] によれば , 開設前後で利用頻度が増加した団体は 7 団体（41.2%）であり , そのうち開設後に新たに利用をはじめた団体は 4 団体（23.5%）であった。減少した団体はなかった。

2）個人の利用頻度の変化（表 2, 図 1-2）

利用者個人を対象とした [ 調査票 2] によれば , 開設前後で利用頻度が増加した者は 97 名（51.3%）であった。全利用者の利用頻度の変化は ,「2 ～ 3 ヵ月に 1 回」を週 0.1 回 ,「月 1 ～ 3 回」を週 0.5 回 ,「週

表 2.　施設の利用頻度の変化

| | 団体数 n=17 | (%) | 人数 n=221 | (%) |
|---|---|---|---|---|
| 増加 | 7 | (41.2) | 97 | (51.3) |
| 減少 | 0 | ( 0.0) | 3 | ( 1.6) |
| 変化なし | 10 | (58.8) | 89 | (47.1) |

無回答を除く

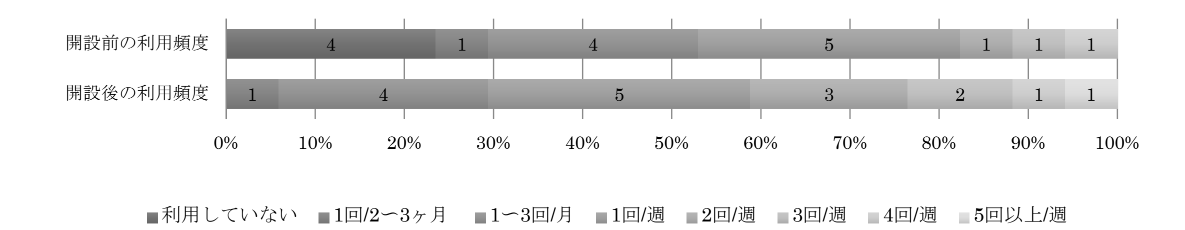


図1-1.　施設の利用頻度［団体の変化］　n=17（団体）

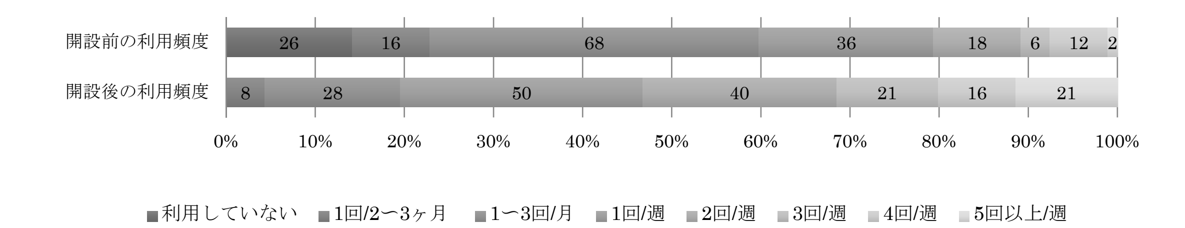


図1-2.　施設の利用頻度［個人の変化］　n=221（人）

5 回以上」を週 5 回として換算すると，開設前は週平均 0.9 ± 1.1 回であったのに対し，開設後は週平均 1.8 ± 2.1 回に増加しており，週平均で 1.0 ± 1.9 回の増加がみられた。また，自宅から施設までの徒歩所要時間と利用頻度の変化との関連を分析したところ，所要時間が短い利用者ほど，利用頻度の増加した割合が高かった。

3) 団体の利用頻度と個人の利用頻度との関連（図 1-3）

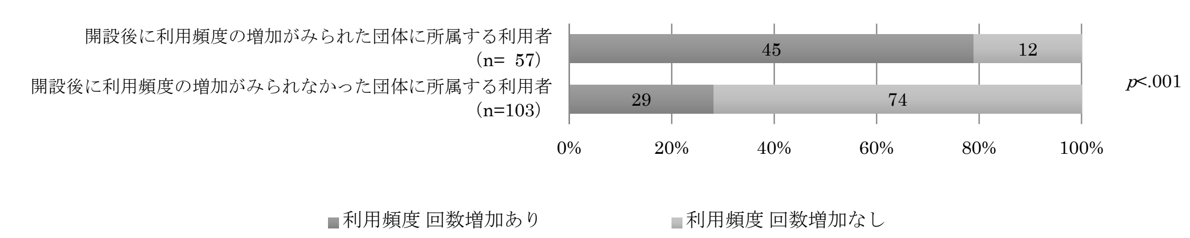


図1-3.　施設の利用頻度［団体の変化と個人の変化との関連］

施設の利用頻度における , 団体の変化と個人の変化との関連を分析したところ , 利用頻度の増加した団体に所属している者は , そうでない団体に所属している者に比べ , 利用頻度の増加した割合が有意に高かった。

4) 自宅から施設までの所要時間と利用頻度との関連（図 1-4）

自宅から施設までの所要時間と利用頻度との関連を分析したところ , 施設の近郊に居住している者ほど , 利用頻度の増加した割合が有意に高かった。

## 3. 施設利用内容の変化

1) 団体の利用内容の変化（表 3-1, 表 3-2）

開設前後で利用内容が総括的に良い方向に変化したと回答した団体（責任者）は , 12 団体（70.6%）であった。また , その活動内容は ,「体操・太極拳」で 4 団体から 5 団体 ,「囲碁・将棋・麻雀」で 1 団体から 2 団体 ,「会話」で 0 団体から 2 団体 ,「会議」で 2 団体から 4 団体の増加がみられた。

2) 個人の利用内容の変化（表 3-1, 表 3-2）

開設前後で利用内容が総括的に良い方向に変化

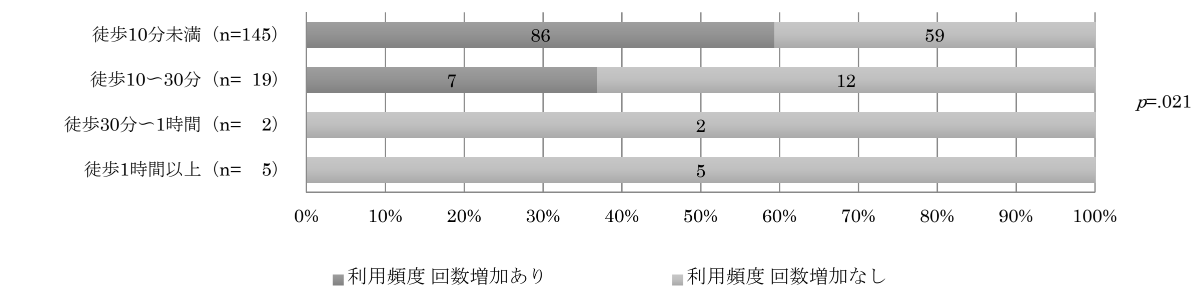


図1-4. 施設の利用頻度［自宅から施設までの所要時間と個人の変化との関連］

表 3-1. 施設の利用内容の変化

| | 団体数 n=17 | (%) | 人数 n=221 | (%) |
|---|---|---|---|---|
| 良い方向に変化（改善） | 12 | (70.6) | 115 | (52.0) |
| 悪い方向に変化（悪化） | 0 | ( 0.0) | 1 | ( 0.5) |
| 変化なし | 5 | (29.4) | 59 | (26.7) |

無回答を除く

表 3-2. 開設前後における施設の利用内容

| | 開設前 n=13 | | 開設後 n=17 | | 増加率 | 開設前 n=158 | | 開設後 n=221 | | 増加率 |
|---|---|---|---|---|---|---|---|---|---|---|
| | 団体数 | (%) | 団体数 | (%) | % | 人数 | (%) | 人数 | (%) | % |
| ［運動系］ | | | | | | | | | | |
| 体操・太極拳 | 4 | (30.8) | 5 | (29.4) | 25.0 | 31 | (19.6) | 56 | (25.3) | 80.6 |
| 舞踏・ダンス | 5 | (38.5) | 5 | (29.4) | 0.0 | 33 | (20.9) | 38 | (17.2) | 15.2 |
| ［芸術系］ | | | | | | | | | | |
| 囲碁・将棋・麻雀 | 1 | ( 7.7) | 2 | (11.8) | 100.0 | 21 | (13.3) | 26 | (11.8) | 23.8 |
| 楽器演奏 | 1 | ( 7.7) | 1 | ( 5.9) | 0.0 | 3 | ( 1.9) | 9 | ( 4.1) | 200.0 |
| コーラス・民謡 | 1 | ( 7.7) | 1 | ( 5.9) | 0.0 | 16 | (10.1) | 11 | ( 5.0) | -31.3 |
| カラオケ | 2 | (15.4) | 2 | (11.8) | 0.0 | 17 | (10.8) | 16 | ( 7.2) | -5.9 |
| 俳句 | 0 | ( 0.0) | 0 | ( 0.0) | 0.0 | 1 | ( 0.6) | 1 | ( 0.5) | 0.0 |
| 書道 | 0 | ( 0.0) | 0 | ( 0.0) | 0.0 | 2 | ( 1.3) | 5 | ( 2.3) | 150.0 |
| 茶道 | 0 | ( 0.0) | 0 | ( 0.0) | 0.0 | 8 | ( 5.1) | 8 | ( 3.6) | 0.0 |
| 手工芸 | 1 | ( 7.7) | 1 | ( 5.9) | 0.0 | 11 | ( 7.0) | 16 | ( 7.2) | 45.5 |
| 絵画・絵手紙 | 0 | ( 0.0) | 0 | ( 0.0) | 0.0 | 2 | ( 1.3) | 14 | ( 6.3) | 600.0 |
| ［その他］ | | | | | | | | | | |
| 会話 | 0 | ( 0.0) | 2 | (11.8) | − | 17 | (10.8) | 54 | (24.4) | 217.6 |
| 会議 | 2 | (15.4) | 4 | (23.5) | 100.0 | 44 | (27.8) | 41 | (18.6) | -6.8 |

無回答を除く

したと回答した利用者は，115 名（65.7%）であった。また，その活動内容は，「体操・太極拳」で 31 名から 56 名（増加率；80.6%），「舞踏・ダンス」で 33 名から 38 名（増加率；15.2%）「囲碁・将棋・麻雀」で 21 名から 26 名（増加率;23.8%），「楽器演奏」で 3 名から 9 名（増加率；200.0%），「書道」で 2 名から 5 名（増加率；150.0%），「手工芸」で 11 名から 16 名（増加率；45.5%），「絵画・絵手紙」で 2 名から 14 名（増加率；600.0%），「会話」で 17 名から 54 名（増加率；217.6%）の増加がみられた。

3）団体の利用内容と個人の利用内容との関連（図 2）

施設の利用内容における，団体の変化と個人の変化との関連を分析したところ，有意差は示されなかったものの，利用内容が改善した団体に所属している者は，そうでない団体に所属している者に比べ，利用内容の改善した割合は高い傾向であった。

## 4. 施設利用と健康行動との関連

1）歩行する機会（図 3, 図 4-1）

歩行する機会が増加したと回答した利用者は 84 名（44.4%）であった。また施設の利用変化と歩行する機会の変化との関連を分析したところ，施設の利用頻度が増加した者，および利用内容が良い方向に変化した者は，歩行する機会が有意に増加していた。

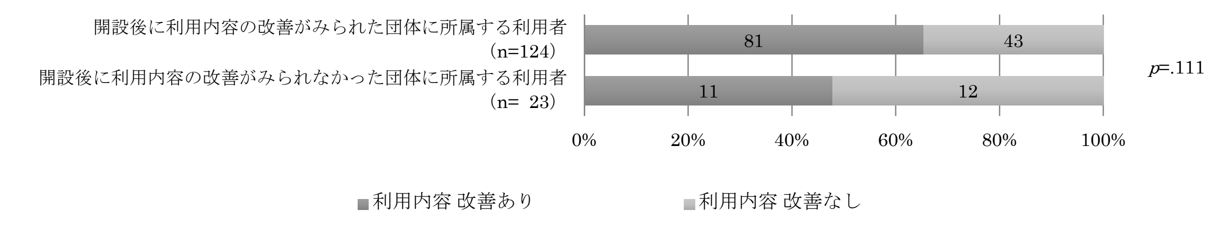


図2. 施設の利用内容［団体の変化と個人の変化との関連］

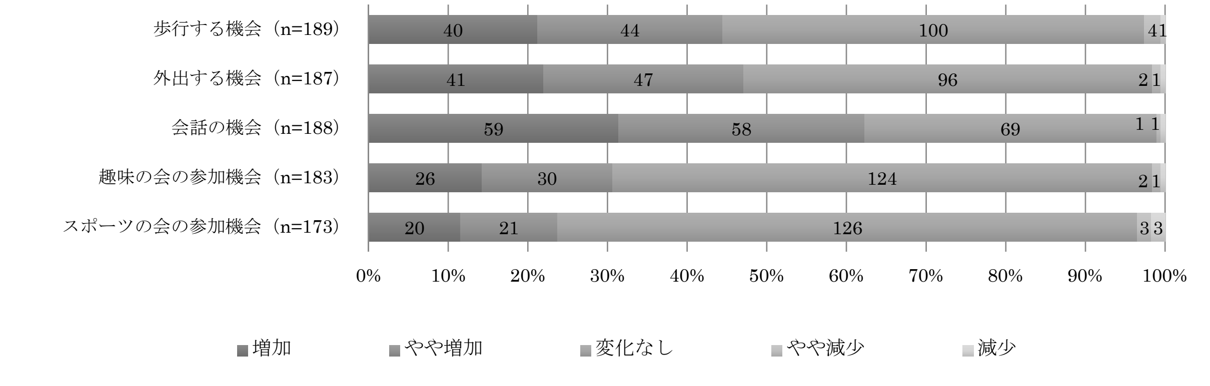


図3. 健康行動の変化

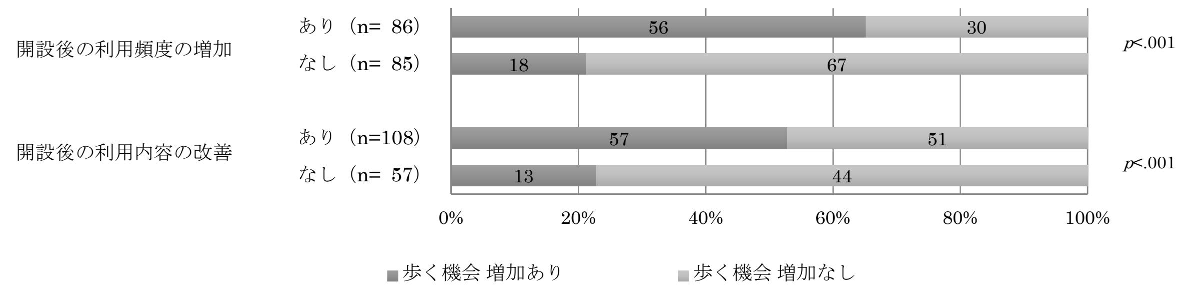


図4-1. 施設利用状況と健康行動の変化との関連［歩行の機会］

2) 外出する機会（図 3, 図 4-2）

外出する機会が増加したと回答した者は 88 名（47.1%）であった。また施設の利用変化と外出する機会の変化との関連を分析したところ，施設の利用頻度が増加した者，および利用内容が良い方向に変化した者は，外出する機会が有意に増加していた。

3) 会話の機会（図 3, 図 4-3）

会話の機会が増加したと回答した者は 117 名（62.2%）であった。また施設の利用変化と会話の機会の変化との関連を分析したところ，施設の利用頻度が増加した者，および利用内容が良い方向に変化した者は，会話の機会が有意に増加していた。

4) 趣味の会の参加機会（図 3, 図 4-4）

趣味の会の参加機会が増加したと回答した者は 56 名（30.6%）であった。また施設の利用変化と趣味の会の参加機会の変化との関連を分析したところ，施設の利用頻度が増加した者，および利用内容が良い方向に変化した者は，趣味の会の参加機会が有意に増加していた。

5) スポーツの会の参加機会（図 3, 図 4-5）

スポーツの会の参加機会が増加したと回答した者は 41 名（23.7%）であった。また施設の利用変化とスポーツの会の参加機会の変化との関連を分析したところ，施設の利用頻度が増加した者，および利用内容が良い方向に変化した者では，スポーツの会の参加機会が有意に増加していた。

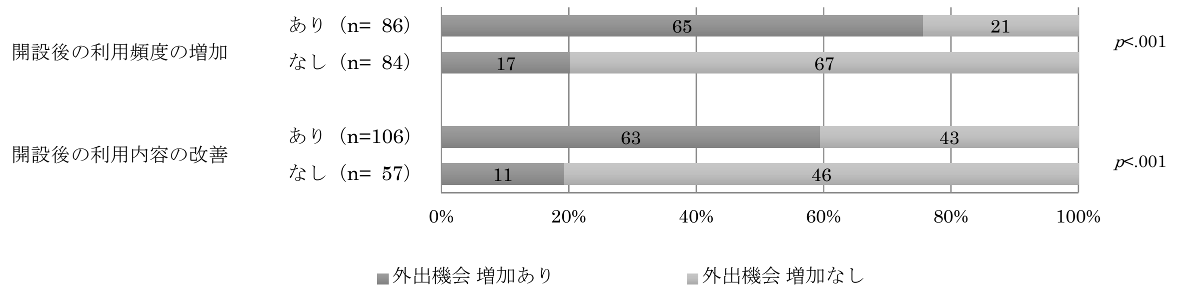


**図4-2．施設利用状況と健康行動の変化との関連［外出の機会］**

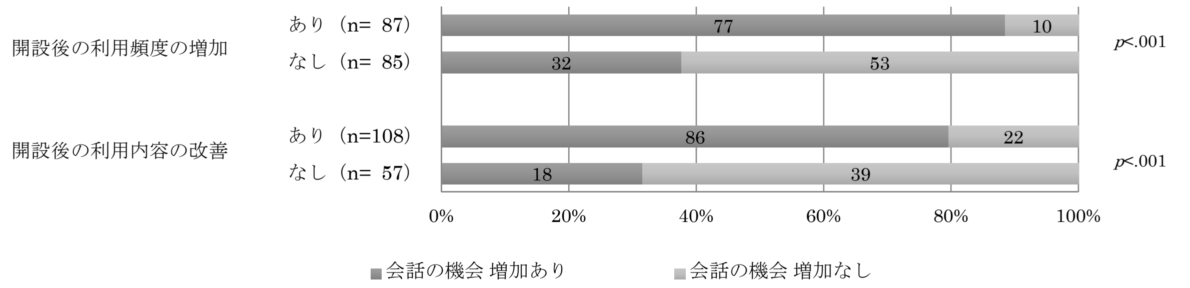


**図4-3．施設利用状況と健康行動の変化との関連［会話の機会］**

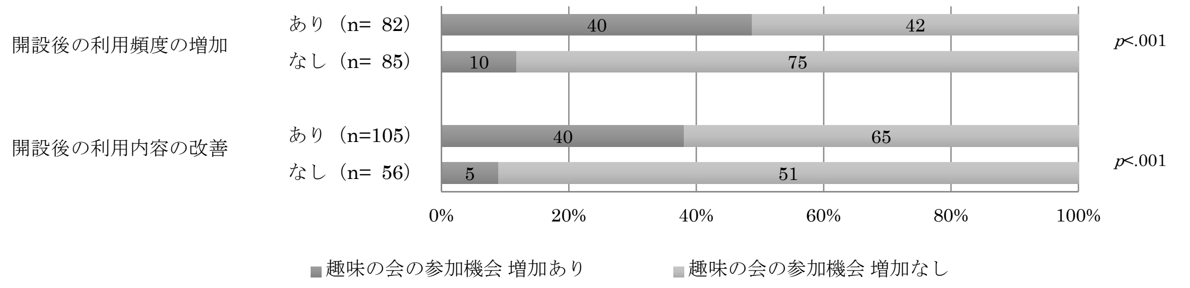


**図4-4．施設利用状況と健康行動の変化との関連［趣味の会の参加機会］**

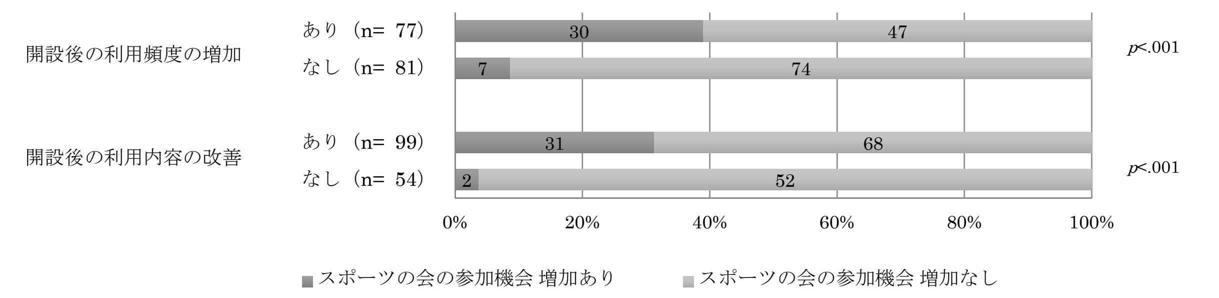


図4-5.　施設利用状況と健康行動の変化との関連［スポーツの会の参加機会］

表 4. 現在の主観的健康感と開設前後における変化

n=221

| | 人数 | (%) |
|---|---|---|
| ［主観的健康感］ | | |
| とてもよい | 32 | (17.7) |
| まあよい | 126 | (69.6) |
| あまりよくない | 18 | ( 9.9) |
| よくない | 5 | ( 2.8) |
| ［主観的健康感の変化］ | | |
| 良い方向に変化（改善） | 82 | (47.7) |
| 悪い方向に変化（悪化） | 0 | ( 0.0) |
| 変化なし | 90 | (52.3) |

無回答を除く

## 5. 健康行動と主観的健康感との関連

1）現在の主観的健康感と開設前後における変化（表 4）

現在の主観的健康感を尋ねたところ，「とてもよい」と回答した者は 32 名（17.7%），「まあよい」と回答した者は 126 名（69.6%）であり，施設の開設前後 1 年間における主観的健康感が良い方向に変化したと回答した者は 82 名（47.7%）であった。

2）健康行動と主観的健康感の変化との関連（図 5）

健康行動の変化と主観的健康感の変化との関連を分析したところ，歩行の機会，外出の機会，会話の機会，趣味の会への参加機会，スポーツの会への参加機会の健康行動が増加した者は，主観的健康感が良い方向に変化していた。

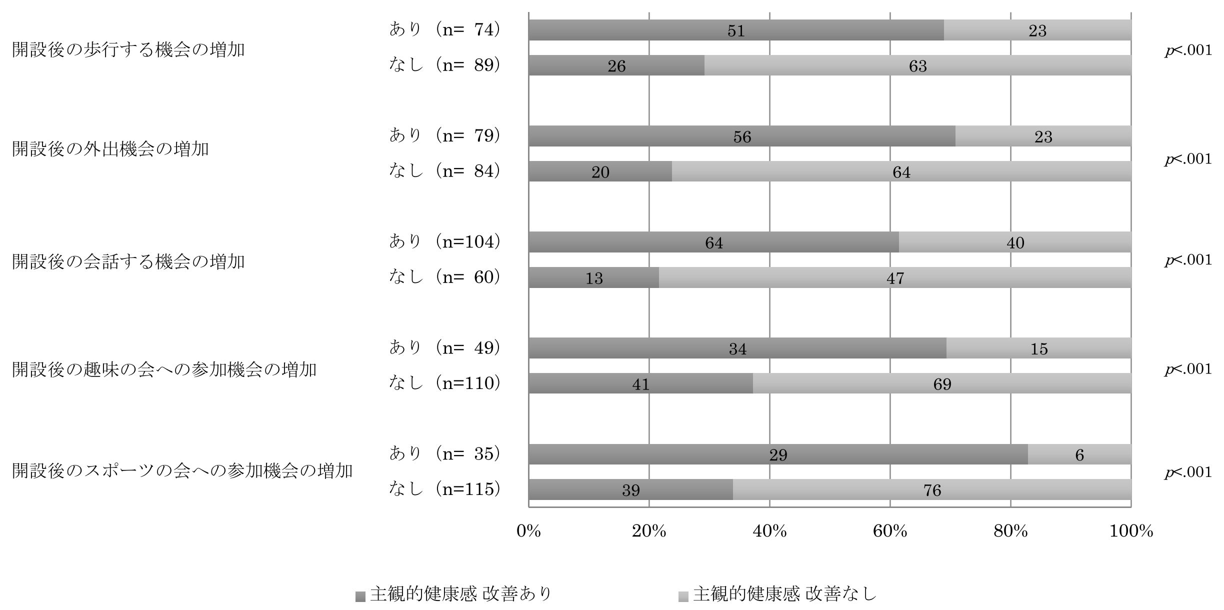


図5.　健康行動と主観的健康感の変化との関連

## Ⅳ 考察

### 1. 施設利用と健康行動との関連

利用者の変化では，すべての年代，性別において，利用者数の増加がみられた。利用者数の変化を居住地からの所要時間別にみると，自宅から施設までの徒歩所要時間が短いほど，利用者数および利用頻度が高まる傾向を示した。また，利用者の所属する集団の活動変化をみると，約 4 割の団体の活動頻度が増加しており，そのような団体に所属する利用者は，個人の行動レベルにおいても，施設の利用頻度が高くなる傾向を示した。

健康行動の変化は，歩行する機会，外出の機会，会話の機会，趣味の会への参加機会，スポーツの会への参加機会が増加している傾向が示された。また，施設の利用頻度と健康行動との関連を分析したところ，利用頻度が増加した者は，歩行，外出，会話，趣味の会への参加，スポーツの会への参加のすべての健康行動の機会において，増加割合が高かった。

近年，生活環境が健康へ与える影響に関心が高まり，国内外の研究において，居住地域の公園や緑地，公共施設などの物的環境が，住民の身体活動や社会活動に関連することが報告されている[5)6)]。本結果においても，「健康交流の家」の開設が，近隣住民の集団および個人の身体活動や社会活動の機会を増やしたと考えられた。

### 2. 健康行動と主観的健康感との関連

主観的健康感の変化では，約半数の利用者が，健康状態が良い方向に変化していた。また，健康行動と主観的健康感の変化との関連を分析したところ，歩行する機会，外出の機会，会話の機会，趣味の会への参加機会，スポーツの会への参加機会の健康行動が増加した者は，主観的健康感が改善している割合が高かった。

主観的健康感は，単なる主観に留まらず，死亡[7) 8)]や要介護認定[9)]の予測力があることが知られている。日常生活での歩行やスポーツなどの身体活動は，慢性疾患を予防し[10)]，要介護リスクや死亡リスクを低下させるなど[11) -13)]，多くの健康指標との関連が報告されている。また，身体活動は一人で実施するより，組織に参加して実施する方が，要介護認定を受けにくいことも指摘されている[14)]。スポーツ活動だけでなく，趣味の会などへの社会活動の参加は，社会的な交流や支え合いの機会が増えることで，要介護状態の予防や，認知症の発症を抑制することが報告さている[15) -18)]。本研究においても，利用者の身体活動および社会活動の増加が，主観的健康感に良い影響を与えた可能性が高いと考えられた。

## Ⅴ．本研究の限界

本研究の限界は，以下の四点である。

一点目は，研究デザインが横断研究であり，一年前の状態を思い出して，それと比較して現状を回答してもらった点である。このような調査法による回答には想起バイアスが起こりうることが知られており，今回の結果においてもバイアスが生じている可能性を排除できない。しかし，今回，参加頻度が増えた者ほど，健康行動が増えていたこと，そのような者で主観的健康感が改善していた者が多かったことから，全てが想起バイアスとは考えにくい。今後，この問題を避けるためには，開設前に調査を行い，その時点での状況を回答してもらうことが望ましい。

二点目は，研究対象者が「健康交流の家」を利用している住民に限られており，「健康交流の家」を利用していない住民との比較をしていない点である。「健康交流の家」の利用が，真に健康増進へ影響しているのかを明らかにするためには，「健康交流の家」の利用のない対照群との健康行動や健康状況を比較し，その効果を検証する必要がある。

三点目は，健康状態を評価する調査項目が，主観的健康感に限定されている点である。主観的健康感も将来の要介護リスクを予測する重要な指標[8) 9)]であるが，一次予防の効果をさらに検証していくためには，今後の要介護認定発生や心理的指標など，介護予防に重要な指標を多角的に評価する必要がある。

四点目は，研究対象地域が一地域である点である。本研究の対象である愛知県東海市の O 地区は，筆者らの過去の分析の結果から，元々，ソーシャル・キャピタルの高い地域であったために，健康増進効果がみられた可能性もある。今後，他地域での調査で再現性の検証や地域間比較などを実施する必要がある。

## Ⅵ．まとめ

本研究は，「健康交流の家」の開設に伴う，地域住民の健康行動および主観的健康感の変化を検証した。「健康交流の家」の開設 1 年後，地域住民の施設の利用頻度は増加し，身体活動および社会活動などの増加がみられた。さらに，これらの健康に望ましい行動が増加した者は，主観的健康感が良い方向に変化する傾向が示された。

したがって，従来の敬老の家から他の機能を併せ持たせた「健康交流の家」の開設は，地域住民の健康行動を促進し，主観的健康感を改善させたと考えられた。

本結果より，「健康交流の家」は，まちづくりによる一次予防に寄与できる可能性が示唆された。

## 謝辞

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# Exercising alone versus with others and associations with subjective health status in older Japanese: The JAGES Cohort Study

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**Although exercising with others may have extra health benefits compared to exercising alone, few studies have examined the differences. We sought to examine whether the association of regular exercise to subjective health status differs according to whether people exercise alone and/or with others, adjusting for frequency of exercise. The study was based on the Japan Gerontological Evaluation Study (JAGES) Cohort Study data. Participants were 21,684 subjects aged 65 or older. Multivariable logistic regression models were used to examine the association. The adjusted odds ratios (ORs) for poor self-rated health were significantly lower for people who exercised compared to non-exercisers. In analyses restricted to regular exercisers the ORs for poor health were 0.69 (95% confidence intervals: 0.60–0.79) for individuals exercising alone more often than with others, 0.74 (0.64–0.84) for people who were equally likely to exercise alone as with others, 0.57 (0.43–0.75) for individuals exercising with others more frequently than alone, and 0.79 (0.64–0.97) for individuals only exercising with others compared to individuals only exercising alone. Although exercising alone and exercising with others both seem to have health benefits, increased frequency of exercise with others has important health benefits regardless of the total frequency of exercise.**

Physical activity has been demonstrated to have various health benefits[1,2]. The benefits of physical activity apply regardless of the context, i.e. whether it occurs as part of work, leisure, transport, or housework[3]. However, it remains unclear whether exercise is more beneficial for those exercising with others, compared to exercising alone (e.g. on the basement treadmill).

This question has been previously discussed by distinguishing physical activity into exercising alone versus with others[4]. The mechanisms for health benefits from exercising with others may include not only physiological effects through physical activity, but also psychological and social factors. A systematic review focusing on the psychosocial benefits of exercising with others revealed that working out with others may enhance social connectedness, social support, and peer bonding[5]. These social relationships have been shown in turn to have potential health benefits[6,7], and exercising with others may therefore have extra health benefits compared to exercising alone.

However, few studies have examined the differences in health associations between exercising alone and exercising with others. One study conducted on middle-aged Japanese adults showed that there was statistically no difference in the incidence of poor mental health five years later between non-exercisers and those exercising mostly alone, while the incidence was lower among those exercising mostly with exercising others, compared to non-exercisers[8]. However, the study did not directly compare exercising alone and with others, and the analyses

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| | | Exercising with others (Ewo) | | | | | |
|---|---|---|---|---|---|---|---|
| | | None | A few times/yr | 1-3/mo | 1/wk | 2-3/wk | ≥4/wk |
| | None | Non-exercisers | | | Ewo-only | | |
| | A few times/yr | | | | | | |
| Exercising alone (Ea) | 1-3/mo | | | | | Ea<Ewo | |
| | 1/wk | Ea-only | | | Ea=Ewo | | |
| | 2-3/wk | | | Ea>Ewo | | | |
| | ≥4/wk | | | | | | |

**Figure 1. Patterns of exercise.** Ea-only: people who only exercised alone. Ea > Ewo: people who reported exercising more frequently alone than with others. Ea = Ewo: people who reported exercising alone or with others with equal frequency. Ea < Ewo: people who exercised with others more frequently than exercising alone. Ewo-only: people who only exercised with others.

did not adjust for differences in the frequency of exercise. One cohort study examining older Japanese adults showed a higher risk of incident functional disability (hazard ratio was 1.29 (95% confidence intervals: 1.02–1.64)) among those who did not participate in a sports organization compared to those who did, even though both groups reported regular exercise[9]. A cross-sectional study in Australian adults showed that sports club participants resulted in more positive benefits for various aspects of quality of life than gymnasium participants or walking participants[10]. These studies suggest the possibility that exercising with others has additional health effects over and above exercising alone. However, exercising alone and exercising with others were not directly compared. We therefore sought to address this gap using cross-sectional data from a cohort of older Japanese adults.

Self-rated health is one subjective indicator that reflects overall health status. Self-rated health is commonly used as a health outcome because of its established validity as a predictor of mortality, regardless of other medical, behavioral, or psychosocial factors[11]. Therefore, the aim of this study was to examine whether the association of subjective health status to exercise differs according to exercising alone and/or with others, adjusting for frequency of exercise. We hypothesized that there would be a lower prevalence of poor self-rated health among those performing exercising with others compared to those only exercising alone, even after adjusting for frequency of exercise. Although the existing guideline on physical activity mentions intensity and duration[3], it does not mention whether exercise should be performed alone or with others. If exercising with others is shown to have greater health benefits than exercising alone, this would suggest the importance of including a social interaction perspective in health promotion using physical activity.

## Methods

**Study sample.** We used cross-sectional data from the baseline wave of the Japan Gerontological Evaluation Study (JAGES), which is a population-based survey of community-dwelling seniors[12]. The JAGES sample includes only those who did not already have functional disabilities at the baseline survey. Those without functional disabilities were defined as those without eligibility for receiving long-term public care insurance benefits. The cohort was established in 2010 to examine prospectively the determinants of healthy aging in a sample of individuals aged 65 years and older. Subjects were selected by random sampling in each municipality, using the residential registry in each locality as the sampling frame. The present analysis was based on a sub-sample of the JAGES cohort study as a national sample of 137,736 people in 30 municipalities across Japan (response rate: 71.1%). Questionnaires were sent to 38,724 people and responses were received from 27,684 (response rate: 71.5%). We excluded 6,000 respondents who did not respond to the questions on age, sex, self-rated health, frequency of exercising alone and with others, or need of assistance in activities of daily living (ADL). The final study population consisted of 21,684 subjects. Subjects comprised 10,390 men (47.9%) and 11,294 women (52.1%), with a mean age of 73.5 ± 6.0 years.

**Measures.** *Subjective health status.* Subjects were asked, "How is your current health status?" with possible responses: excellent, good, fair, and poor. Dichotomisation of multinominal self-rated health is frequently used in studies and has been validated[13]. Based on the previous study, subjects who responded with "fair," or "poor," were combined to form our outcome variable. The test-retest reliability of self-rated health was shown to be good in a variety of subgroups by age and sex[14]. In addition, the criterion-related validity of self-rated health was shown to predict mortality in a review[11], and similar results were also observed in older Japanese adults, regardless of age, marital status, health behaviors, symptoms of depression, and chronic co-morbid conditions[15].

*Exercising alone and exercising with others.* To define exercising alone, respondents were asked, "How often do you exercise alone?" To define exercising with others, respondents were asked, "How often do you exercise with a relative, friend, or acquaintance?" For each question, possible responses were: four or more times a week, two or three times a week, once a week, one to three times a month, a few times a year, and none. Based on a previous study that examined the relationship between mortality and physical activity[16], the frequency of exercising alone and exercising with others was divided into six mutually exclusive categories: (1) non-exercisers, (2) people who only exercised alone (Ea-only), (3) people who reported exercising more frequently alone than with others (Ea > Ewo), (4) people who reported exercising alone or with others with equal frequency (Ea = Ewo), (5) people who exercised with others more frequently than exercising alone (Ea < Ewo); and (6) people who only exercised with others (Ewo-only) (Fig. 1). Next, the total frequency of exercise (combinations of two variable categories) was calculated and divided into six categories (see Supplementary Fig. S1). The higher the category, the greater the frequency of exercise. The categories were dichotomized into two groups: categories 1 to 3 reflected individuals who exercised less than twice a week, categories 4 and 5 exercised more than twice a week.

**Covariates.** Based on previous studies[9,17], age, sex, annual equivalized income (less than 2 million yen per year = "low", 2–3.99 million yen per year = "middle", 4 million yen or more per year = "high"), educational attainment (less than 10 years, more than 10 years), household composition (living alone, with others), occupational status (employed, not employed), self-reported medical conditions (no illness or disability, illness or disability), instrumental activities of daily living (IADL) (instrumental self-maintenance[18]; 5 points = "high", 0–4 points = "low"), depressive symptoms (Geriatric Depression Scale[19]; 0–4 points = "no depression", 5–9 points = "depressive tendency", 10 points or more = "depression"), and total frequency of exercise were included as covariates in our regression models. Furthermore, as exercising with others may reflect sociability; frequency of meeting friends (two or more times a week, once a month to once a week, less than once a month), receiving instrumental support, providing instrumental support, receiving emotional support, and providing emotional support (yes, no) were also included as covariates.

**Statistical analysis.** To examine whether the association of subjective health status to exercise differs according to exercising alone and/or with others, we performed multivariable logistic regression to calculate the odds ratios (ORs) for poor self-rated health. All variables were set as dummy variables. A "missing" category was used in analysis to account for missing values in response to questions.

The dependent variable was self-rated health and independent variables were the six groups characterized by frequency of exercising alone and exercising with others. In Model 1, age, sex, annual equivalized income, educational attainment, household composition, occupational status, self-reported medical conditions, IADL, depression, frequency of meeting friends, receiving instrumental support, providing instrumental support, receiving emotional support, and providing emotional support were added as covariates to the univariate model. In Model 2, total frequency of exercise was added to Model 1. In addition, to perform sensitivity analysis for examining whether the associations differ by total frequency of exercise, we conducted further analysis by stratifying the analyses into categories 4 and 5 (those who exercise at least twice a week) versus categories 1 to 3 (those who exercised less than twice a week).

SPSS 21.0 J was used for statistical analysis with a 2-tailed significance level set at 5%.

**Ethics statement.** Ethical approval for the study was obtained from the Nihon Fukushi University Ethics Committee (application number: 10–04) and Chiba University Ethics Committee (application number: 1777). This study was performed in accordance with the principles of the Declaration of Helsinki. Informed consent was obtained from all participants.

## Results

Table 1 shows characteristics of individuals according to their patterns of exercise. Those who exercised with others (Ea > Ewo, Ea = Ewo, Ea < Ewo and Ewo-only) tended to be younger, and this group had a higher proportion of people with a high equivalized income, high educational attainment, living with others, high IADL score, no depression, rich social relationships, and good self-rated health. Among exercisers (Ea-only, Ea > Ewo, Ea = Ewo, Ea < Ewo and Ewo-only), there was a higher proportion of people who exercised less than twice a week among individuals who only exercised with others (Ewo-only).

Table 2 shows the adjusted ORs for poor self-rated health according to patterns of exercise. In Model 1 for all participants, the ORs for poor health were significantly lower for individuals who exercised (regardless of whether alone or with others; (Ea-only, Ea > Ewo, Ea = Ewo, Ea < Ewo and Ewo-only)). In the next set of models, we excluded non-exercisers in order to draw comparisons just among the different types of people who performed regular exercise. In these analyses, individuals who exercised alone (Ea-only) became the reference group for all comparisons. In Model 1, the ORs were 0.67 (95% confidence intervals: 0.58–0.77) for people who exercised alone more often than with others (Ea > Ewo), 0.72 (0.63–0.82) among people who exercised equally frequently alone or with others (Ea = Ewo), 0.58 (0.44–0.76) for individuals who exercised more often with others compared to alone (Ea < Ewo), and 0.86 (0.70–1.05) for individuals who only exercised with others (Ewo-only). The category of individuals who exclusively exercised with others (Ea > Ewo, Ea = Ewo, Ea < Ewo and Ewo-only) was statistically indistinguishable from people who exercised alone (Ea-only). The covariates in Model 1 plus total frequency of exercise were included in Model 2; the corresponding ORs were 0.69 (0.60–0.79), 0.74 (0.64–0.84), 0.57 (0.43–0.75), 0.79 (0.64–0.97).

Stratified analysis was then performed by dichotomizing the sample according to frequency of exercise. In Model 2, the ORs for individuals only exercising with others (Ewo-only) were similar results of the analysis performed on all exercisers, even though these were not statistically significant in either stratum.

## Discussion

This study was the first to examine whether the association of subjective health status to exercise differs according to exercising alone and/or with others, adjusting for frequency of exercise. As expected, in the analysis of all participants, the ORs for poor self-rated health were significantly lower for all exercise groups (Ea-only, Ea > Ewo, Ea = Ewo, Ea < Ewo and Ewo-only) compared to non-exercisers. In the analysis excluding non-exercisers, the ORs for poor self-rated health were significantly lower for people who exercised both alone and with others (Ea > Ewo, Ea = Ewo and Ea < Ewo) and people who only exercised with others (Ewo-only) compared to people who only exercised alone (Ea-only), after adjusting for total frequency of exercise. Moreover, although the ORs were not significantly lower for people who only exercised with others (Ewo-only), similar results were found

| | | Ea-only | Ea > Ewo | Ea = Ewo | Ea < Ewo | Ewo-only | Non-exercisers |
|---|---|---|---|---|---|---|---|
| N | Mean ± SD | 6,018 | 3,685 | 3,895 | 760 | 1,131 | 6,195 |
| Age (years) | | 73.8 ± 6.1 | 72.6 ± 5.4 | 72.9 ± 5.4 | 72.3 ± 5.1 | 72.4 ± 5.5 | 74.5 ± 6.7 |
| Sex (%) | Males | 48.8 | 53.4 | 46.9 | 44.1 | 36.5 | 47.1 |
| Equivalized income (%) | Low | 44.3 | 38.2 | 41.7 | 41.1 | 35.2 | 44.1 |
| | Middle | 30.7 | 36.9 | 32.2 | 36.6 | 38.2 | 27.3 |
| | High | 7.9 | 10.6 | 9.2 | 10.3 | 10.8 | 8.5 |
| | Missing | 17.2 | 14.3 | 16.9 | 12.1 | 15.8 | 20.1 |
| Educational attainment (%) | ≤9 | 40.1 | 30.8 | 37.6 | 29.3 | 29.9 | 47.9 |
| | ≥10 | 58.5 | 68.1 | 61.2 | 69.5 | 68.3 | 50.1 |
| | Missing | 1.4 | 1.1 | 1.2 | 1.2 | 1.8 | 2.1 |
| Household composition (%) | Living alone | 16.8 | 12.0 | 11.1 | 11.2 | 11.4 | 13.6 |
| | With others | 79.1 | 84.8 | 85.0 | 85.7 | 85.5 | 81.3 |
| | Missing | 4.1 | 3.2 | 3.9 | 3.2 | 3.1 | 5.1 |
| Occupational status (%) | Employed | 22.2 | 20.8 | 22.2 | 17.4 | 22.9 | 26.1 |
| | Not employed | 70.1 | 72.5 | 70.1 | 76.1 | 70.5 | 63.6 |
| | Missing | 7.7 | 6.7 | 7.7 | 6.6 | 6.6 | 10.3 |
| Self-reported medical condition (%) | No illness or disability | 14.4 | 16.9 | 16.4 | 12.8 | 18.2 | 14.2 |
| | Illness or disability | 81.3 | 78.3 | 77.8 | 80.5 | 75.2 | 80.4 |
| | Missing | 4.3 | 4.8 | 5.8 | 6.7 | 6.6 | 5.4 |
| IADL (%) | High | 81.2 | 86.8 | 84.9 | 90.4 | 86.4 | 70.4 |
| | Low | 16.6 | 11.8 | 13.1 | 7.6 | 11.9 | 26.4 |
| | Missing | 2.2 | 1.4 | 2.1 | 2.0 | 1.7 | 3.2 |
| Depression (%) | No depression | 60.4 | 72.2 | 72.1 | 72.5 | 68.8 | 53.5 |
| | Depressive tendency | 18.5 | 11.3 | 12.2 | 12.1 | 13.4 | 20.9 |
| | Depression | 5.6 | 2.0 | 2.7 | 2.8 | 4.1 | 8.7 |
| | Missing | 15.4 | 14.4 | 13.0 | 12.6 | 13.8 | 16.9 |
| Frequency of meeting friends (%) | <1/mo | 31.9 | 16.1 | 16.3 | 13.8 | 16.9 | 35.6 |
| | 1/mo-1/wk | 35.4 | 38.7 | 29.4 | 29.7 | 33.6 | 31.6 |
| | ≥2/wk | 28.5 | 42.9 | 51.2 | 54.5 | 47.3 | 26.7 |
| | Missing | 4.2 | 2.3 | 3.1 | 2.0 | 2.2 | 6.2 |
| Receiving emotional support (%) | Yes | 91.3 | 95.7 | 95.7 | 97.1 | 95.8 | 89.3 |
| | No | 6.8 | 2.7 | 2.9 | 1.8 | 2.9 | 7.9 |
| | Missing | 1.9 | 1.6 | 1.4 | 1.1 | 1.3 | 2.8 |
| Providing emotional support (%) | Yes | 89.3 | 94.7 | 94.2 | 95.9 | 94.3 | 85.1 |
| | No | 7.6 | 3.0 | 3.5 | 2.1 | 3.7 | 10.6 |
| | Missing | 3.0 | 2.4 | 2.2 | 2.0 | 1.9 | 4.2 |
| Receiving instrumental support (%) | Yes | 91.7 | 95.6 | 96.1 | 97.0 | 95.5 | 91.8 |
| | No | 6.3 | 2.6 | 2.5 | 1.8 | 3.3 | 5.7 |
| | Missing | 2.0 | 1.8 | 1.4 | 1.2 | 1.2 | 2.4 |
| Providing instrumental support (%) | Yes | 75.3 | 82.5 | 81.5 | 82.2 | 83.3 | 72.1 |
| | No | 19.8 | 13.6 | 14.0 | 14.2 | 13.9 | 21.9 |
| | Missing | 4.9 | 3.9 | 4.5 | 3.6 | 2.8 | 6.0 |
| Frequency of exercising alone (%) | None | 0.0 | 0.0 | 0.0 | 0.0 | 100.0 | 100.0 |
| | A few times/yr | 7.8 | 0.0 | 13.4 | 24.3 | 0.0 | 0.0 |
| | 1–3/mo | 8.1 | 4.9 | 9.7 | 26.7 | 0.0 | 0.0 |
| | 1/wk | 11.4 | 9.4 | 12.3 | 29.6 | 0.0 | 0.0 |
| | 2–3/wk | 27.3 | 31.4 | 29.8 | 19.3 | 0.0 | 0.0 |
| | ≥4wk | 45.5 | 54.3 | 34.7 | 0.0 | 0.0 | 0.0 |
| Frequency of exercising with others (%) | None | 100.0 | 0.0 | 0.0 | 0.0 | 0.0 | 100.0 |
| | A few times/yr | 0.0 | 30.9 | 13.4 | 0.0 | 20.4 | 0.0 |
| | 1–3/mo | 0.0 | 28.0 | 9.7 | 7.0 | 16.0 | 0.0 |
| | 1/wk | 0.0 | 26.2 | 12.3 | 14.5 | 18.5 | 0.0 |
| | 2–3/wk | 0.0 | 14.8 | 29.8 | 35.9 | 22.9 | 0.0 |
| | ≥4wk | 0.0 | 0.0 | 34.7 | 42.6 | 22.2 | 0.0 |

Continued

| | | Ea-only | Ea > Ewo | Ea = Ewo | Ea < Ewo | Ewo-only | Non-exercisers |
|---|---|---|---|---|---|---|---|
| Total frequency of exercise (%) | Non-exercisers | 0.0 | 0.0 | 0.0 | 0.0 | 0.0 | 100.0 |
| | Category 1 | 7.8 | 0.0 | 0.0 | 0.0 | 20.4 | 0.0 |
| | Category 2 | 8.1 | 4.9 | 13.4 | 7.0 | 16.0 | 0.0 |
| | Category 3 | 11.4 | 9.4 | 9.7 | 14.5 | 18.5 | 0.0 |
| | Category 4 | 27.3 | 31.4 | 12.3 | 35.9 | 22.9 | 0.0 |
| | Category 5 | 45.5 | 54.3 | 64.5 | 42.6 | 22.2 | 0.0 |
| Self-rated health (%) | Poor | 18.1 | 10.0 | 11.0 | 8.9 | 12.6 | 24.7 |

**Table 1. Characteristics of individuals according to patterns of exercise.** Ea-only: people who only exercised alone. Ea > Ewo: people who reported exercising more frequently alone than with others. Ea = Ewo: people who reported exercising alone or with others with equal frequency. Ea < Ewo: people who exercised with others more frequently than exercising alone. Ewo-only: people who only exercised with others. Total frequency of exercise (categories 1 to 3): people who exercised less than twice a week. Total frequency of exercise (categories 4 and 5): people who exercised at least twice a week. Results are presented as mean ± SD for continuous variables and percentage (%) for categorical variables.

when stratified analysis was performed using the collapsed groups reflecting frequency of exercise per week. These results imply that increased frequency of exercise with others has important health benefits regardless of the total frequency of exercise, although exercising alone and exercising with others both seem to have health benefits.

In a previous study on middle-aged adults, there was no difference between those who did not perform exercise or play sports (the reference category) and those who exercised mostly alone, whereas there was a significantly lower OR of poor mental health later on among those who exercised mostly with others[8]. Similarly, in a study on older adults, even for those exercising once a week or more, the risk of incident functional disability was significantly lower among those who participated in a sports organization compared to those who did not[9]. The results of these previous studies are consistent with the finding in the present study that the OR of poor self-rated health was significantly lower among those exercising with others than those only exercising alone.

In those who exercised with others, the ORs for poor self-rated health seem to be smaller for those exercising both alone and with others (Ea > Ewo, Ea = Ewo and Ea < Ewo) than those who only exercised with others (Ewo-only). This could still be residual confounding by total MET-hours, even though we were only crudely able to adjust frequency of exercise (i.e. those performing both may be likely to be spending more total time exercising compared to those only exercising alone). In contrast, the above-mentioned study on the association with mental health did not find any differences in the risk of poor mental health between those exercising both alone and with others and those who did not exercise *et al.*[8]. Although it is true that the reference for comparison was not the same, the trend observed was different from that of the present study. One reason for this difference may be that those exercising both alone and with others accounted for over half of those who exercised in the present study, which includes representative samples, but accounted for only 3% in the previous study.

Social relationships may be one mechanism underlying the health benefits of exercising with others[4,20]. Reviews have indicated that poor social relationships can increase mortality risk[6,7], and similar results were also observed in older Japanese adults[21]. In addition, social connectedness while exercising contributes to exercise adherence[22]. Previous studies examining the mechanism underlying the relationship between exercising with others and health revealed the possibility that social relationships may contribute to the association between participation in a sports organization and incidence of functional disability[9,17]. In this research, we used a part of general social relationships (frequency of meeting friends, receiving instrumental support, providing instrumental support, receiving emotional support, and providing emotional support) as covariates which could serve as measures of sociability. As we could not use specific social relationships in exercising with others, future studies are needed to use specific social relationships in exercising with others to examine whether these social relationships mediate the association between exercising with others and health. Other possible mechanisms that may have a positive association with exercising with others are: adherence to exercise routines[23–25], self-esteem and other psychological factors[5], social capital[26] and other social factors[4]. For example, those who exercise with others may have continued to exercise for more years at the time of the survey than those who exercise alone. As we could not determine the roles of those factors in the present study, further studies are needed.

The present study had some limitations. Firstly, while we considered the frequency of exercise, which is an important point when investigating the association between exercise and health, we did not consider intensity or duration[3], or type of exercise[27]. The differences between exercising alone and exercising with others may be residually confounded by differences in these factors. The second limitation is that the phrase "exercise with others" did not differentiate between exercise with only one other person and exercise with two or more other people or in a group or organization. Associations with health may differ between the different forms of exercise with others. The third is that we used combinations of two variable categories for "total frequency of exercise", which may have resulted in a slight lack of accuracy. The fourth is that there may be a confounding effect from demographic and psychosocial factors related to exercising with others[28], which we did not examine. The fifth is that the study was cross-sectional, and therefore cannot determine causal relationships. Further studies are therefore also needed to consider these points.

| | | Crude | | Model 1 | | Model 2 | |
|---|---|---|---|---|---|---|---|
| | N | OR | 95%CI | OR | 95%CI | OR | 95%CI |
| **All participants** | | | | | | | |
| Non-exercisers | 6,195 | ref | — | ref | — | | |
| Ea-only | 6,018 | 0.68 | 0.62–0.74 | 0.75 | 0.69–0.83 | | |
| Ea > Ewo | 3,685 | 0.34 | 0.30–0.38 | 0.50 | 0.43–0.57 | | |
| Ea = Ewo | 3,895 | 0.38 | 0.34–0.42 | 0.54 | 0.48–0.61 | | |
| Ea < Ewo | 760 | 0.30 | 0.23–0.39 | 0.43 | 0.33–0.56 | | |
| Ewo-only | 1,131 | 0.44 | 0.37–0.53 | 0.64 | 0.52–0.78 | | |
| **Exercisers-only: all participants excluding non-exercisers** | | | | | | | |
| **Total frequency of exercise: Category 1–5 (all exercisers)** | | | | | | | |
| Ea-only | 6,018 | ref | — | Ref | — | ref | — |
| Ea > Ewo | 3,685 | 0.50 | 0.44–0.57 | 0.67 | 0.58–0.77 | 0.69 | 0.60–0.79 |
| Ea = Ewo | 3,895 | 0.56 | 0.50–0.63 | 0.72 | 0.63–0.82 | 0.74 | 0.64–0.84 |
| Ea < Ewo | 760 | 0.45 | 0.34–0.58 | 0.58 | 0.44–0.76 | 0.57 | 0.43–0.75 |
| Ewo-only | 1,131 | 0.66 | 0.54–0.79 | 0.86 | 0.70–1.05 | 0.79 | 0.64–0.97 |
| **Total frequency of exercise: Category 4–5 (exercisers ≥2/wk)** | | | | | | | |
| Ea-only | 4,381 | ref | — | Ref | — | ref | — |
| Ea > Ewo | 3,159 | 0.53 | 0.46–0.61 | 0.69 | 0.59–0.80 | 0.69 | 0.59–0.80 |
| Ea = Ewo | 2,995 | 0.57 | 0.50–0.65 | 0.73 | 0.62–0.85 | 0.76 | 0.65–0.90 |
| Ea < Ewo | 597 | 0.45 | 0.33–0.61 | 0.56 | 0.40–0.77 | 0.55 | 0.40–0.75 |
| Ewo-only | 510 | 0.60 | 0.45–0.80 | 0.80 | 0.58–1.09 | 0.78 | 0.57–1.06 |
| **Total frequency of exercise: Category 1–3 (exercisers <2/wk)** | | | | | | | |
| Ea-only | 1,637 | ref | — | ref | — | ref | — |
| Ea > Ewo | 526 | 0.54 | 0.41–0.71 | 0.72 | 0.54–0.97 | 0.69 | 0.51–0.93 |
| Ea = Ewo | 900 | 0.55 | 0.44–0.69 | 0.69 | 0.54–0.88 | 0.65 | 0.51–0.84 |
| Ea < Ewo | 163 | 0.46 | 0.28–0.76 | 0.65 | 0.39–1.10 | 0.62 | 0.37–1.05 |
| Ewo-only | 621 | 0.59 | 0.46–0.76 | 0.77 | 0.59–1.02 | 0.79 | 0.60–1.04 |

**Table 2. Odds ratios of poor self-rated health according to patterns of exercise.** Ea-only: people who only exercised alone. Ea > Ewo: people who reported exercising more frequently alone than with others. Ea = Ewo: people who reported exercising alone or with others with equal frequency. Ea < Ewo: people who exercised with others more frequently than exercising alone. Ewo-only: people who only exercised with others. Model 1 was adjusted for sex, age, equivalized income, educational attainment, household composition, occupational status, self-reported medical conditions, IADL, depression, and sociability. Model 2 was adjusted for the covariates in Model 1 plus total frequency of exercise.

## Conclusion

Among older Japanese adults, although exercising alone and exercising with others both seem to have health benefits, increased frequency of exercise with others has important health benefits regardless of the total frequency of exercise. A social interaction perspective may be useful to assist with promoting exercise benefits for older adults.

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## Author Contributions

Conceived and designed the experiments: S.K., T.T., S.I., Y.K. and K.K. Analyzed the data: S.K., T.T., S.I., Y.K., I.K. and K.K. Wrote the paper: S.K., T.T., S.I., Y.K., I.K. and K.K. Acquisition of data: S.K., Y.K. and K.K.

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RESEARCH ARTICLE

# Neighborhood Characteristics and Cardiovascular Risk among Older People in Japan: Findings from the JAGES Project

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## Abstract

Previous studies have found an association between neighborhood characteristics (i.e., aspects of the physical and social environment) and the incidence of cardiovascular disease (CVD) and elevated CVD risk. This study investigated the relationship between neighborhood characteristics and CVD risk among older people in Japan where research on this association is scarce. Data came from the Japan Gerontological Evaluation Study project; questionnaire data collected from 3,810 people aged 65 years or older living in 20 primary school districts in Aichi prefecture, Japan, was linked to a computed composite CVD risk score based on biomarker data (i.e., hemoglobin A1c, systolic blood pressure, diastolic blood pressure, high-density lipoprotein cholesterol, low-density lipoprotein cholesterol, and estimated glomerular filtration rate). A sex-stratified multilevel linear regression analysis revealed that for male participants, living in neighborhoods with a higher perceived occurrence of traffic accidents and reduced personal safety was associated with an elevated CVD risk (coefficient = 1.08 per interquartile range increase, 95% confidence interval [CI] = 0.30 to 1.86) whereas males living in neighborhoods with a higher perceived proximity of exercise facilities had a lower risk (coefficient = −1.00, 95% CI = −1.78 to −0.21). For females, there was no statistically significant association between neighborhood characteristics and CVD risk. This study suggests that aspects of the neighborhood environment might be important for CVD morbidity and mortality in Japan, particularly among men.

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## Introduction

In recent years there has been increasing interest in the role of neighborhood characteristics (i.e., aspects of the physical and social environment) in the emergence of cardiovascular disease (CVD)/CVD risk factors [1, 2] given their potential to influence individual behavioral factors that are associated with CVDs such as dietary habits [3–5], physical activity levels [6] and smoking [7, 8]. Previous studies have focused on a wide range of factors linked to the neighborhood physical environment including access to shops selling fresh fruits and vegetables [1, 2, 9], hilly residential areas [10], physical activity resources [1, 2, 11–13], safety from traffic [11, 12], street connectivity [11, 12, 14] and the degree of greenness [15–18]. Similarly, in terms of the social environment, several factors have been examined such as social cohesion [1, 2, 11], structural social capital [19], social support for pro-physical activity [11] and the perceived risk of crime [14, 20, 21]. In the main, these studies have shown that a favorable neighborhood environment is inversely associated with CVD/CVD risk factors, while detrimental neighborhood features have been linked to a greater CVD risk [16, 17].

Neighborhood environments might affect the cardio-metabolic risk of their residents in two main ways: (a) through the impact they have on the energy balance and (b) as a result of psychological stress [2]. For example, energy consumption or dietary composition is influenced by the availability of different food items in the immediate neighborhood environment. Moreover, residing in communities which have exercise and leisure facilities may be associated with a higher level of physical activity [11–13, 15–18], whereas a high crime rate may result in more time being spent indoors, and reduced physical activity. In addition, the detrimental features of communities (e.g. crime, traffic accidents and high unemployment) can also act as chronic psychological stressors, which can damage health both directly [14, 20, 21] and also indirectly through unhealthy behaviors induced by psychological stress [22–24]. In contrast, community-level social capital (i.e., resources accessed by individuals and groups within a social structure that facilitate cooperation, collective action, and the maintenance of norms [25]), might protect people from the effects of psychological stress as a result of greater social support [26, 27] or by underpinning the emergence of a health-promoting environment (e.g., from increased physical activity associated with participating in community organizations) [28].

Despite an increasing focus on the role of neighborhood characteristics in the emergence of CVD/CVD risk, there are still important research gaps. For example, although most studies have used various predictor variables (e.g., smoking, physical activity, body mass index (BMI), hypertension, fruit and vegetable consumption) as CVD risk factors, few have simultaneously assessed multiple biomarkers which would more accurately represent the overall risk for a future CVD event [1, 11, 20, 29–31]. Further, these associations have still not yet been studied in many parts of the world. For instance, to the best of our knowledge, there is no research on the association between aspects of the physical and social neighborhood environment and CVD risk in Japan, even though CVD is one of the leading causes of death in the country [32].

To address this issue the current study used data from the Japan Gerontological Evaluation Study (JAGES) project, an ongoing epidemiological study that focuses on identifying social determinants of health among Japanese people aged 65 years and older [33], to investigate the association between neighborhood characteristics and CVD risk as evaluated by multiple biomarkers.

## Methods

### Data

The study data were collected through a postal survey undertaken in August 2010 to January 2012 in six municipalities in Chita peninsula, Aichi prefecture, Japan. This was linked to data obtained in the municipality-organized voluntary health check-up (the Tokutei kenkou shinsa: Kenshin) in four of the six municipalities. Of the 20,432 people who participated in the JAGES project in Aichi prefecture, health check-up information was linked for 9,893 (48.4%). Study exclusion criteria included missing information for either: location or residence (at a primary school zone level) (n = 9), weight and height (n = 41), biomarkers (n = 5,530); current smoking status (n = 409). Individuals who had self-reported difficulty in daily living activities were also excluded (n = 293). In addition, one municipality did not provide information on blood pressure. A total of 3,810 participants residing in 20 primary school districts in 3 municipalities were thus included in the subsequent analyses.

The study protocol and questionnaire procedures were approved by the Ethics Committee for Research on Human Subjects at Nihon Fukushi University, Japan (No. 10–05) and the Ethics Committee for Medical Research at the University of Tokyo (No. 10555). Informed consent was assumed with the voluntary return of the questionnaire.

## Dependent variable

CVD risk was evaluated with the Suita Score [34] using information from six biomarkers measured in the health check-up (i.e., hemoglobin A1c (HbA1c), systolic and diastolic blood pressure, low-density (LDL) and high-density (HDL) lipoprotein cholesterol and the estimated glomerular filtration rate (eGFR)) together with age, sex and current smoking status. The Suita Score was developed by the National Cerebral and Cardiovascular Center, Japan, to predict possible future CVD events; it was developed based on the Framingham Risk Score [35], which is known to overestimate the risk of coronary heart disease among the Japanese population [34]. The score can range from 10 to 95 with higher values indicating a higher risk of a future cardiovascular event.

## Independent variables

The questionnaire obtained information on different aspects of the neighborhood environment. Following the lead of previous studies [1, 2], we examined: (1) the proximity of shops selling fresh fruits and vegetables; (2) the proximity of exercise facilities (i.e., places that are suitable for exercise or walking); (3) the presence of slopes or stairs in the neighborhood; (4) social capital; and (5) personal safety and the risk of traffic accidents.

## Physical environment

For the first three above-mentioned variables (1–3), participants were asked about the environment within 1 km of their residential location (i.e., about shops or facilities that sell fresh fruits and vegetables, parks or streets which are suitable for sports activities or walking, and places where they find it difficult to walk because of environmental obstacles such as slopes or stairs). They used one of five answer options to describe the extent to which these factors were present in their immediate environment (i.e., greatly, to some extent, a few, none and don't know). For the current study we calculated the proportion of those who answered either "greatly" or "to some extent" by primary school district.

These three neighborhood environmental features were also evaluated by geographic information system (GIS) analysis; the number of shops which purportedly sold fresh food (such as fruits, vegetables, meat and fish), the number of parks (i.e., exercise facilities), and the hilliness of the neighborhood were calculated for each primary school district. Information on the number of grocery stores at a 500-meter resolution was available from the 2007 Commerce Census conducted by the Ministry of Economy, Trade and Industry; in this study, the number of grocery stores was calculated for each school district, assuming that all of the stores were located in the centroids of each geographical unit (i.e., 500m×500m). A grocery store was defined as either a department store, general merchandise store, specialized supermarket, or daily commodities store. The average park count was calculated using National Land Numerical Information City Park Data (as of 2011) from the Ministry of Land, Infrastructure, Transport and Tourism (MLIT), Japan. The average hilliness of each neighborhood i.e. the land slope was calculated based on the Elevation, Degree of Slope 5th Mesh Data (as of 2011) which was also obtained from MLIT and was based on the Digital Map 50 m Grid (Elevation). ArcGIS 10.1 software was used for all spatial calculations.

## Social environment

Community-level social capital was operationalized in two distinct forms. Cognitive social capital, which includes norms, values, attitudes and beliefs [36], was measured with the following three questions, "generally speaking, do you trust people in your community?", "do people in your community try to be helpful to others?" and "how attached do you feel to your community?" Each question had five answer options (i.e., greatly, to some extent, cannot say, not very much, and not at all). The proportion of those who answered greatly or to some extent for each question was calculated, standardized to a mean of 0 and standard deviation of 1 and then averaged to create a composite score for each primary school district. Structural social capital, which is usually operationalized as the extent and intensity of associational links and activity in society [36], was measured with three questions on social participation (i.e., participating in volunteer groups, sports organizations and hobby groups at least once a month). The proportion of those who participated in each activity was calculated, standardized and then averaged to create a composite score. Perceived personal safety and the risk of neighborhood traffic accidents was evaluated with three questions; "Is there any road or intersection which is a high risk for traffic accidents?", "Is there any place where you think it is dangerous to walk at night?", and "how unsafe do you feel in your community?" The proportion of those who answered "greatly" or "to some extent" for each question was calculated and standardized to compute a composite score.

## Covariates

Information on participants' demographic and socioeconomic characteristics was also collected (i.e. age, sex, marital status, educational attainment, employment and household equivalent income). BMI was calculated based on self-reported height and weight. After excluding those whose height or weight was above or below 4 SD from the mean, BMI was divided into three categories: $< 18.5$; 18.5 to 24.9; $\geq 25.0$. Marital status referred to being either married or not married, or was unknown. There were three educational attainment levels based on the number of years spent in school, $\leq 9$ years; 10–12 years; and $\geq 13$ years. For employment status respondents were categorized as being either employed; retired and not employed; had never worked; or had missing data. Annual household equivalent income was calculated by dividing the total household income by the square root of the number of household members [37] and was divided into three categories: low (less than 2 million yen); middle (2 to 4 million yen); and high (more than 4 million yen). Mental health i.e., the presence of depression, was assessed using the 15-item Geriatric Depression Scale (GDS), a self-administered questionnaire. Participants were categorized into three groups i.e., non-depressed (a score of 0–4); mild depression (5–9); and severe depression (10–15) [38].

Lifestyle information was collected on (1) the frequency of consuming fruit and vegetables, (2) alcohol consumption, (3) current cigarette smoking, (4) time spent in daily walking and (5) the frequency of meeting friends. The frequency of fruit and vegetable consumption was assessed with the question, "How often did you eat fruit and vegetables in the previous month?" Responses were grouped into four categories, $\geq$ twice a day; once a day; four to six times a week; $\leq$ two to three times a week. Alcohol consumption was categorized with the answer options yes, no, and used to drink. Smoking status was assessed by the question, "Do you smoke?" with four possible answers, never; quit smoking more than 5 years ago; quit smoking 5 or less years ago; and currently smoke. To assess time spent in daily walking respondents were asked, "How long do you walk for each day on average?" with four response options, $< 30$ minutes; 30–59 minutes; 60–89 minutes; and $\geq 90$ minutes. Information on the frequency of meeting friends was obtained with the question "How often do you meet friends or acquaintances?" with answers divided into two categories, $\geq$ once a week; and $<$ once a week. Neighborhood socioeconomic status (SES) was evaluated by calculating the land price of each primary school zone for the year 2010, with information which was published by MLIT [39].

## Statistical analysis

Following the lead of a previous study which investigated the association between neighborhood SES and the Framingham Score [40], a multilevel linear regression analysis was conducted to investigate the association between neighborhood characteristics and the Suita Score as a continuous variable. Sex-specific analyses were undertaken as previous research has indicated that not only do male and female lifestyles differ in Japan [32, 41] but that the way the sexes interact with their surrounding neighborhood may differ [42]. For the multilevel analysis we defined individual participants and primary school districts as Level 1 and Level 2, respectively. To ensure comparability the nine neighborhood characteristics that were the focus of this study were standardized (i.e. divided by their inter-quartile range [IQR] values) and examined separately, with models being adjusted sequentially [2]. Model 1 adjusted for individual socio-demographic variables (i.e., age in years, age-squared, sex, BMI, marital status, educational status, employment status, equivalent household income), depression and neighborhood SES; Model 2 further adjusted for individual behavioral variables (i.e., time spent walking, fruit and vegetable consumption, alcohol consumption, smoking and frequency of meeting friends). To maximize statistical power, if information was missing for any of the covariates a "missing" category was created and included in the analysis. Additional analyses showed that the overall results did not change when the participants with missing information were excluded. When investigating the effect of neighborhood cognitive and structural social capital, we also included individual-level cognitive (i.e., trust, attachment and reciprocity) and structural (i.e., participation in sports organizations, volunteer groups and hobby clubs once a month or more frequently) social capital measures as covariates, respectively.

To examine the robustness of the results, we also performed a sensitivity analysis, where the analysis was restricted to those who had resided in the primary school districts for 10 years or longer while adjusting for the same covariates included in Model 2, as it is possible that the effects of the neighborhood environment may be cumulative and operate over a longer period of time. We chose 10 years as a cut-off principally because the Suita Score and several similar scores usually calculate the 10-year risk of CVD, which means 10 years is thought of as being a long enough period of time to differentiate those who are at risk and those who are not in terms of developing CVD.

Stata 13.0 (Stata Corp, College Station, TX) was used to conduct the statistical analyses. Results are shown in the form of coefficients with 95% confidence intervals (CI). The level of statistical significance was $p < 0.05$.

## Results

Descriptive statistics of the study participants are presented in Table 1. Participants' average age was 71.3 years and 52.6% were female. More than three quarters of the participants were married, while just under half of them (47.8%) had $\leq$ 9 years education, 36.1% had 10–12 years of education while 12.4% had obtained a further education ($\geq$ 13 years). Most households (42.2%) were in the lowest income category of less than 2 million yen. Just under half of the participants (46.4%) consumed vegetables twice a day or more often with 40% of them reporting that they drink alcohol although this proportion was more than three times higher among male participants (63.1% vs. 19.1%). A sex-based imbalance was also observed for current smoking with 19.0% of the male participants smoking compared to only 2.4% of females. Nearly one-third (32.7%) of the participants walked for at least one hour every day, while 75.3% of them met with friends one or more times a week. There was no statistically significant difference by sex in HbA1c, systolic blood pressure or eGFR, while diastolic blood pressure was higher among males and LDL and HDL cholesterols were statistically higher among females. The total Suita Score was significantly higher among male participants.

Details of the neighborhood characteristics evaluated at a primary school district level are presented in Table 2. Seventy-five percent of the participants in each school district answered that there were shops selling fresh fruit and vegetables in the vicinity (76.4, IQR = 69.3 to 79.9), while the proportion was almost identical for the perceived proximity of exercise facilities (76.1, IQR = 68.7 to 83.8). A much lower proportion of respondents perceived barriers to walking in their immediate environment (43.0, IQR = 34.2 to 53.9). A GIS-based analysis revealed there were 5.4 shops (median, IQR = 3.0 to 12.6) per square kilometer ($km^2$) that supposedly sold fresh fruit and vegetables, while there were 0.88 (median, IQR = 0.19 to 2.02) parks per $km^2$. The median value of the neighborhood slopes (a measure of hilliness) was 3.17 degrees (IQR = 2.30 to 4.44). The school district-level indices of cognitive and structural social capital were 0.03 (IQR = −0.38 to 0.23) and 0.00 (IQR = −0.25 to 0.61). The median value for perceived personal safety and risk of traffic accidents was 0.25 (IQR = −0.57 to 0.56). The median land price was 64,362.5 yen (IQR = 42,550 to 82,375) per square meter.

In the multilevel linear regression analysis male participants who lived in neighborhoods with a higher perceived proximity of exercise facilities had a lower CVD risk (Table 3). Each IQR increase in the index of the perceived proximity of exercise facilities resulted in a 0.77 (95% confidence interval [CI] = −1.48 to −0.07) reduction in the Suita Score in Model 1. Specifically, those who resided in communities with a greater number of perceived exercise facilities (i.e., the top 25%) had a 0.77 lower score compared with those in communities with fewer perceived exercise facilities (i.e., the bottom 25%). This remained statistically significant even after adjusting for individual behavioral factors (coefficient = −1.00, 95% CI = −1.78 to −0.21). In contrast, living in a neighborhood where the perceived risk of traffic accidents and reduced personal safety was greater was associated with a higher CVD risk (i.e., coefficient = 0.81, 95% CI = 0.11 to 1.52) among men, which specifically meant that those in communities with a higher perceived risk of traffic accidents and reduced personal safety (i.e., the top 25%) had 0.81 points higher score compared with those residing in communities with a lower perception; this also remained statistically significant after adjusting for individual behavioral factors in Model 2 (coefficient = 1.08, 95% CI = 0.30 to 1.86). Among the female participants, there was no statistically significant association between any of the neighborhood characteristics and CVD risk.

When the analysis was restricted to those participants who had resided in their current residential location for 10 years or longer, the associations observed among the male participants

**Table 1. Characteristics of the study participants in Chita peninsular, Aichi Prefecture, Japan in 2010 (n = 3,810).**

| | Total | Men | Women | p-value |
|---|---|---|---|---|
| | (n = 3,810) | (n = 1,805) | (n = 2,005) | |
| | *Mean [SD] / n (%)* | | | |
| Age (in years) | 71.3 [5.0] | 71.2 [4.8] | 71.4 [5.2] | 0.28 |
| Marital status | | | | |
| Married | 2,935 (77.0) | 1,586 (87.9) | 1,349 (67.3) | < 0.001 |
| Not married | 748 (19.6) | 164 (9.0) | 586 (29.2) | |
| Other | 127 (3.3) | 57 (3.2) | 70 (3.5) | |
| Educational attainment | | | | |
| ≤ 9 years | 1,822 (47.8) | 782 (43.3) | 1,040 (51.9) | < 0.001 |
| 10–12 years | 1,376 (36.1) | 678 (37.6) | 698 (34.8) | |
| ≥ 13 years | 474 (12.4) | 282 (15.6) | 192 (9.6) | |
| Missing | 138 (3.6) | 63 (3.5) | 75 (3.7) | |
| Annual equivalent household income (yen) | | | | |
| < 2 million | 1,607 (42.2) | 765 (42.4) | 842 (42.0) | < 0.001 |
| 2–4 million | 1,380 (36.2) | 732 (40.6) | 648 (32.3) | |
| > 4 million | 234 (6.1) | 112 (6.2) | 122 (6.1) | |
| Missing | 589 (15.5) | 196 (10.9) | 393 (19.6) | |
| Occupational status | | | | |
| Currently employed | 701 (18.4) | 411 (22.8) | 290 (14.3) | < 0.001 |
| Retired | 2,314 (60.7) | 1,228 (68.0) | 1,086 (54.2) | |
| Unemployed | 374 (9.8) | 64 (3.6) | 310 (15.5) | |
| Missing | 421 (11.1) | 102 (5.7) | 319 (15.9) | |
| BMI (kg/m$^2$) | | | | |
| Lean (< 18.5) | 195 (5.1) | 52 (2.9) | 143 (7.1) | 0.001 |
| Normal (18.5–24.9) | 2,704 (71.0) | 1,300 (72.0) | 1,404 (70.0) | |
| Overweight (25–29.9) | 827 (21.7) | 426 (23.6) | 401 (20.0) | |
| Obese (> 30) | 84 (2.2) | 27 (1.5) | 57 (2.8) | |
| GDS | | | | |
| Not depressed (0–4) | 2,260 (59.3) | 1,030 (57.1) | 1,230 (61.4) | 0.01 |
| Mild depression (5–9) | 720 (18.9) | 346 (19.2) | 374 (18.7) | |
| Depression (10–15) | 196 (5.1) | 93 (5.2) | 103 (5.1) | |
| Missing | 634 (16.6) | 336 (18.6) | 298 (14.9) | |
| Vegetable intake | | | | |
| > twice/day | 1,767 (46.4) | 685 (38.0) | 1,082 (54.0) | < 0.001 |
| Once/day | 1,252 (32.9) | 638 (35.4) | 614 (30.6) | |
| 4–6 times/week | 444 (11.7) | 275 (15.2) | 169 (8.4) | |
| 2–3 times/week or less | 304 (8.0) | 195 (10.8) | 109 (5.4) | |
| Missing | 43 (1.1) | 12 (0.7) | 31 (1.6) | |
| Alcohol consumption | | | | |
| Drink | 1,521 (39.9) | 1,138 (63.1) | 383 (19.1) | < 0.001 |
| Don't drink | 2,249 (59.0) | 656 (36.3) | 1,593 (79.5) | |
| Missing | 40 (1.1) | 11 (0.6) | 29 (1.5) | |
| Cigarette smoking | | | | |
| Never | 2,310 (60.6) | 448 (24.8) | 1,862 (92.9) | < 0.001 |
| Quit | 1,109 (29.1) | 1,014 (56.2) | 95 (4.7) | |
| Smoke | 391 (10.3) | 343 (19.0) | 48 (2.4) | |
| Walking hours | | | | |

(*Continued*)

Table 1. (*Continued*)

| | Total (n = 3,810) | Men (n = 1,805) | Women (n = 2,005) | p-value |
|---|---|---|---|---|
| < 30 min | 1,070 (28.1) | 458 (25.4) | 612 (30.5) | < 0.001 |
| 30–59 min | 1,297 (34.0) | 614 (34.0) | 683 (34.1) | |
| 60–89 min | 634 (16.6) | 329 (18.2) | 305 (15.2) | |
| ≥ 90 min | 612 (16.1) | 335 (18.6) | 277 (13.8) | |
| Missing | 197 (5.2) | 69 (3.8) | 128 (6.4) | |
| Meeting friends | | | | |
| < once/week | 745 (19.6) | 496 (27.5) | 249 (12.4) | < 0.001 |
| ≥ once/week | 2,869 (75.3) | 1,233 (68.3) | 1,636 (81.6) | |
| Missing | 196 (5.1) | 76 (4.2) | 120 (6.0) | |
| $HbA_{1c}$ (%) | 5.6 [0.7] | 5.6 [0.8] | 5.5 [0.6] | 0.29 |
| Systolic blood pressure (mmHg) | 131.7 [16.5] | 131.9 [16.4] | 131.5 [16.5] | 0.56 |
| Diastolic blood pressure (mmHg) | 73.7 [10.2] | 74.8 [10.5] | 72.8 [9.9] | < 0.001 |
| LDL cholesterol (mg/dL) | 119.7 [28.9] | 114.6 [29.0] | 124.2 [28.0] | < 0.001 |
| HDL cholesterol (mg/dL) | 58.9 [15.6] | 54.7 [15.0] | 62.6 [15.2] | < 0.001 |
| eGFR(mL/min/1.73 $m^2$) | 68.9 [13.9] | 68.3 [14.1] | 69.4 [13.6] | 0.17 |
| Suita Score | 49.9 [7.5] | 54.2 [6.7] | 46.1 [5.8] | < 0.001 |

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were generally more pronounced (Table 4), while there was no significant association between any of the neighborhood characteristics and CVD risk among the female participants.

## Discussion

### Summary of the study's main findings

The present study revealed that for males aged 65 years or older, living in a neighborhood with a higher perceived proximity to exercise facilities was associated with a reduced CVD risk,

**Table 2. Basic attributes of the neighborhood environment in the study locations in Chita peninsular, Aichi Prefecture, Japan (n = 20).**

| Attributes | Median | IQR |
|---|---|---|
| 1. Proximity of shops selling fresh vegetables and fruits | | |
| 1a. Perceived (%) | 76.4 | 69.3, 79.9 |
| 1b. GIS-based (per $km^2$) | 5.4 | 3.0, 12.6 |
| 2. Proximity of exercise facilities | | |
| 2a. Perceived (%) | 76.1 | 68.7, 83.8 |
| 2b. GIS-based (per $km^2$) | 0.88 | 0.19, 2.02 |
| 3. Walking environment in the neighborhood | | |
| 3a. Perceived (%) | 43.0 | 34.2, 53.9 |
| 3b. GIS-based hilliness (degree) | 3.17 | 2.30, 4.44 |
| 4. Social capital | | |
| 4a. Cognitive | 0.03 | −0.38, 0.23 |
| 4b. Structural | 0.00 | −0.25, 0.61 |
| 5. Personal safety and risk of traffic accidents | | |
| 5a. Perceived | 0.25 | −0.57, 0.56 |
| Neighborhood SES | | |
| Land price (yen per $m^2$) | 64,362.5 | 42,550, 82,375 |

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**Table 3. Multilevel linear regression analysis examining the association between neighborhood characteristics and cardiovascular risk among older people living in Chita peninsula, Aichi Prefecture, Japan.**

| | Male participants (n = 1,805) | | | | Female participants (n = 2,005) | | | |
|---|---|---|---|---|---|---|---|---|
| | Model 1 | | Model 2 | | Model 1 | | Model 2 | |
| | coef. | 95% CI | coef. | 95% CI | coef. | 95% CI | coef. | 95% CI |
| 1. Proximity of shops selling fresh fruits and vegetables | | | | | | | | |
| 1a. Perceived | −0.02 | −0.48, 0.45 | −0.06 | −0.63, 0.50 | −0.15 | −0.67, 0.37 | −0.15 | −0.66, 0.36 |
| 1b. GIS-based | 0.12 | −0.69, 0.92 | −0.04 | −0.95, 0.87 | 0.14 | −0.63, 0.92 | 0.07 | −0.70, 0.83 |
| 2. Proximity of exercise facilities | | | | | | | | |
| 2a. Perceived | −0.77 | −1.48, −0.07* | −1.00 | −1.78, −0.21* | −0.16 | −0.96, 0.64 | −0.22 | −1.01, 0.56 |
| 2b. GIS-based | −0.14 | −0.59, 0.30 | −0.09 | −0.61, 0.43 | 0.15 | −0.34, 0.64 | 0.10 | −0.38, 0.59 |
| 3. Walking environment in the neighborhood | | | | | | | | |
| 3a. Perceived | −0.58 | −1.25, 0.08† | −0.62 | −1.39, 0.14 | −0.30 | −0.98, 0.38 | −0.31 | −0.98, 0.35 |
| 3b. GIS-based | −0.21 | −0.76, 0.34 | −0.16 | −0.83, 0.51 | 0.08 | −0.55, 0.71 | 0.09 | −0.54, 0.71 |
| 4. Social Capital | | | | | | | | |
| 4a. Cognitive | −0.16 | −0.57, 0.25 | −0.26 | −0.71, 0.19 | −0.14 | −0.55, 0.27 | −0.17 | −0.57, 0.23 |
| 4b. Structural | −0.27 | −0.94, 0.41 | −0.02 | −0.77, 0.74 | 0.07 | −0.61, 0.74 | 0.08 | −0.59, 0.75 |
| 5. Personal safety and risk of traffic accidents | | | | | | | | |
| 5a. Perceived | 0.81 | 0.11, 1.52* | 1.08 | 0.30, 1.86** | 0.40 | −0.37, 1.17 | 0.44 | −0.31, 1.19 |

**: p < 0.01

*: p < 0.05

†: p < 0.10

Model 1 adjusted for basic individual socio-demographic variables (i.e., age in years, age-squared, sex, body mass index, marital status, educational status, employment status, equivalent household income and depression) and neighborhood SES. When investigating the effect of neighborhood cognitive and structural social capital (4a and 4b), we also included individual-level cognitive and structural social capital measures as covariates, respectively.

Model 2 further adjusted for individual behavioral variables (i.e., time spent walking, fruit and vegetable consumption, alcohol consumption, smoking and frequency of meeting friends).

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while those who lived in neighborhoods where traffic accidents and reduced safety were perceived as being more common had a higher CVD risk. Among female participants, no associations were observed between neighborhood characteristics and CVD risk. An additional analysis that focused only on those participants who had resided in their current residential location for at least 10 years produced the same results although the associations seen among men were more pronounced.

## Mechanisms linking neighborhood characteristics and cardiovascular risk

To the best of our knowledge, no previous study has examined the association between the perceived proximity of exercise facilities and CVD risk, although research on other health outcomes has produced similar findings to the associations seen among men in this study. For example, Auchincloss et al. [43] found that neighborhood physical activity resources were associated with a lower hazard ratio for diabetes (0.65) in the U.S. Mathis et al. [44] investigated the association between BMI and local recreational facilities and showed that obese older adults had a significantly lower probability of having a park in their neighborhood. Previous research has also highlighted the ways in which such an environment might be beneficial for health. For instance, an earlier study showed that people who live in areas where there are more parks are more likely to walk for at least 60 minutes per week, compared to those who live in areas with fewer parks [12]. In addition, in Canada, park features (the number of facilities in a park such

**Table 4. Sensitivity analysis restricted to those participants who had resided in their current location for ≥ 10 years in Chita peninsula, Aichi Prefecture, Japan.**

| | Male participants (n = 1,692) | | | | Female participants (n = 1,899) | | | |
|---|---|---|---|---|---|---|---|---|
| | Model 1 | | Model 2 | | Model 1 | | Model 2 | |
| | coef. | 95% CI | coef. | 95% CI | coef. | 95% CI | coef. | 95% CI |
| 1. Proximity of shops selling fresh fruits and vegetables | | | | | | | | |
| 1a. Perceived | −0.11 | −0.58, 0.37 | −0.16 | −0.72, 0.41 | −0.14 | −0.64, 0.37 | −0.12 | −0.63, 0.38 |
| 1b. GIS-based | 0.07 | −0.74, 0.89 | −0.07 | −0.99, 0.85 | 0.13 | −0.64, 0.90 | 0.05 | −0.70, 0.81 |
| 2. Proximity of exercise facilities | | | | | | | | |
| 2a. Perceived | −0.84 | −1.56, −0.13* | −1.07 | −1.86, −0.28** | −0.18 | −0.96, 0.61 | −0.22 | −0.99, 0.55 |
| 2b. GIS-based | −0.14 | −0.59, 0.31 | −0.11 | −0.62, 0.41 | 0.13 | −0.35, 0.61 | 0.09 | −0.39, 0.56 |
| 3. Walking environment in the neighborhood | | | | | | | | |
| 3a. Perceived | −0.53 | −1.22, 0.16 | −0.57 | −1.35, 0.21 | −0.27 | −0.94, 0.41 | −0.28 | −0.94, 0.39 |
| 3b. GIS-based | −0.29 | −0.83, 0.24 | −0.27 | −0.92, 0.38 | 0.04 | −0.58, 0.66 | 0.05 | −0.56, 0.66 |
| 4. Social Capital | | | | | | | | |
| 4a. Cognitive | −0.20 | −0.60, 0.21 | −0.32 | −0.77, 0.12 | −0.14 | −0.54, 0.26 | −0.15 | −0.54, 0.25 |
| 4b. Structural | −0.25 | −0.94, 0.43 | −0.01 | −0.77, 0.75 | 0.07 | −0.60, 0.75 | 0.10 | −0.56, 0.76 |
| 5. Personal safety and risk of traffic accidents | | | | | | | | |
| 5a. Perceived | 0.91 | 0.23, 1.60** | 1.18 | 0.43, 1.94** | 0.34 | −0.42, 1.11 | 0.39 | −0.36, 1.14 |

**: $p < 0.01$

*: $p < 0.05$

Models were adjusted for basic individual socio-demographic variables (i.e., age in years, age-squared, sex, body mass index, marital status, educational status, employment status, equivalent household income and depression) and neighborhood SES and individual behavioral variables (i.e., time spent walking, fruit and vegetable consumption, alcohol consumption, smoking and frequency of meeting friends). When investigating the effect of neighborhood cognitive and structural social capital (4a and 4b), we also included individual-level cognitive and structural social capital measures as covariates, respectively.

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as open spaces, paths, playgrounds, wooded areas, and unpaved trails) were a significant predictor of physical activity [13]. In England, the reported frequency of green space use among people declined with increasing distance whereas those individuals who lived closest to different types of green space tended to achieve the recommended level of physical activity [15]. The results of the current study suggest that such health-promoting facilities may also have the potential to reduce CVD risk in Japan. Future studies should obtain more detailed objective information about how respondents interact with the specific features (e.g. parks) of the surrounding environment to elucidate more precisely how these neighborhood characteristics might influence CVD risk.

The association between perceived neighborhood safety, risk of traffic accidents and CVD risk observed among men accords with the findings from earlier studies that have used the same/similar neighborhood predictor variables. For example, in Sweden high levels of neighborhood violent crime have been associated with coronary heart disease (CHD) (i.e., a first hospitalization for a fatal or non-fatal CHD event) [20]. Similarly, neighborhood safety (defined as being able to walk safely during the day and at night, crime and violence) has also been linked to the cardiovascular health (CVH) score, a concept recently introduced by the American Heart Association [1]. The perceived risk of traffic accidents and reduced personal safety have also been associated with other health indicators such as obesity [21, 44], mental well-being [45–47] and physical functioning [47]. Living in fear might result in more time being spent indoors, lower levels of physical activity, and higher levels of psychological stress, all of which have the possibility to impact on CVD risk. It should also be emphasized that we observed this association with perceived safety in a setting (Japan) where crime is not as common as in other countries [48]. This highlights the importance of perceived safety on CVD risk and the need for future research to determine the factors underlying this association more precisely (e.g. what level of crime has an effect, does personal victimization explain this association etc.).

Although previous studies showed that living in a favorable food environment was associated with a lower CVD risk, in the current study we did not observe this association. For example, Unger et al. [1] found that the presence of more healthily provisioned neighborhood food stores was positively associated with CVH scores in the U.S. One possible reason for the null finding in the current study is that the variation in the food environment was not large enough to be mirrored in the CVD risk. Another possibility comes from the fact that we did not consider the availability of fruit and vegetables from other sources such as mobile vendors or catering services or whether participants grew their own produce.

The association between the neighborhood walking environment (e.g., the perceived presence of slopes and stairs and a GIS measure of neighborhood 'hilliness') and CVD risk or health in general has not been extensively studied. An earlier study undertaken in Australia nonetheless showed that the odds of self-reported diabetes were lower among those who lived in neighborhoods with steeper slope levels [10]. As neighborhoods were hillier on average in our study (3.2 degrees) than the Australian study (3.3%, equivalent to 1.9 degrees), it would have been reasonable to expect that there would have also been an association with CVD risk in this study on the basis that hilliness is likely to be associated with an increase in energy expenditure; however, hilly environments might also prevent people from undertaking physical activity (especially those who are older).

To date, research on the association between neighborhood social capital and CVD risk has produced conflicting results. Unger et al. [1] showed for example, that social capital (termed as social cohesion in their study) was not associated with the CVH score. In contrast, Scheffler et al. [19] found that a one-standard deviation increase in a social capital index (i.e., the Petris Social Capital Index) was significantly associated with a decreased recurrence of acute coronary syndrome, and that this was observed among those living in areas with lower SES. In addition, one component of neighborhood cognitive social capital i.e., a lack of fairness, has also been linked with decreased systolic blood pressure [49]. In the present study we did not find an association when using the Suita Score. It is uncertain what underlies this difference, although as Unger et al. [1] pointed out, it is possible that the mechanisms linking social capital and CVD risk might operate more at the individual level. Indeed, in this study, those men who participated in volunteer groups had a lower CVD risk (coefficient = −1.35, p = 0.003; data not shown in table), while Yazawa et al. [50] who also used JAGES data found that individual-level social participation was inversely associated with hypertension.

## The sex difference in the association between neighborhood characteristics and CVD risk

Earlier research showed that there was a sex difference in the effects of the neighborhood environment on cardiovascular risk. However, it also highlighted that there was a more pronounced association among female participants, which is opposite to the finding of the current study. For example, the effect of inequity on CHD incidence and case fatality as measured by neighborhood deprivation has been found to be slightly stronger for women than men [51]. A sex difference has also been seen for other health outcomes. In a study undertaken in England [47], multiple aspects of the residential environment (e.g., trust, integration into the wider society, a left-wing political climate, the physical quality of the residential environment, and the unemployment rate) had a larger effect for females than males in terms of self-rated health.

Inoue et al. [52] also found that females were more affected by neighborhood socioeconomic characteristics (i.e., differences in the Gini coefficient) than men in terms of C-reactive protein concentration in rural China. Furthermore, while studying the French population, Vallée et al. [53] suggested that females might be more influenced by the place where they live since they tend to spend more time in their community surroundings.

One possible reason for these conflicting findings might relate to differences in the age range of the study participants. Our participants were aged 65 years or older. Yasunaga et al. [54] showed that among old people living in another location in Japan, the age-related reduction in habitual physical activity was reflected in a reduction in activity $\geq 3$ metabolic equivalents (METs; multiples of resting metabolic rate) among men while for women it was $< 3$ METs, where 3 METs is the moderate physical activity intensity level recommended by the World Health Organization to reduce non-communicable diseases [55]. Given this, the sex difference observed in the present study might have resulted from the different degrees to which sex moderated the age-related reduction in environmental physical activity. Another possibility comes from the fact that the mean CVD risk score was lower among females (i.e., females were generally healthier than males); it is possible that it might have been more difficult to detect a neighborhood effect among females, as they were generally healthier irrespective of their surrounding environment.

## Study limitations and future research

This study has several limitations. First, our study population might not have been fully representative of all older persons living in Japan in several respects. For example, as our survey was a postal survey which required a physical response to return the completed questionnaire it is possible that those who had mobility or other serious health problems may have been automatically excluded. Furthermore, as we used data from a non-compulsory health check-up it is possible that our participants might have differed systematically from those who did not undergo the medical examination in terms of such things as their socio-economic or health status or how much they interact with the neighborhood environment–potentially biasing our results. In addition, this study only focused on those who lived in Chita peninsular which limits the generalizability of our findings. Second, the cross-sectional nature of the study precluded the possibility of establishing causal relations. To further elucidate the impact of the neighborhood environment on CVD risk in this setting, future studies should investigate these associations using longitudinal data. Third, we lacked information on potentially important explanatory variables which might have affected both health and participants' perceptions of the neighborhood environment e.g. information on the purchasing of fresh fruits and vegetables (such as what was purchased and where it was purchased from), consumption of home-grown vegetables, availability and utilization of exercise facilities (not only parks but also other facilities such as recreational centers), incidence of crime or personal experience of victimization. In addition, as information on the number of grocery shops was obtained from the 2007 Commerce Census, it is also possible that the number of shops had changed by the time the study was undertaken (i.e., 2010). Fourth, as the Suita Score has been validated only among people residing in Suita City, Osaka, CVD risk estimation might not have been accurate in this study.

Future research should focus on how to combine different methods of assessing the neighborhood environment (e.g. GIS-based methods and participants' perceptions) and to determine what underlies any differences in outcomes such as the difference that was observed in the current study, where statistically significant associations were only observed for participant perceptions of the neighborhood environment. In addition, as there has been comparatively little research about the effects of the neighborhood environment on health and well-being among older age groups, this should also be a priority for future research, especially given the ongoing process of population aging in Japan and other countries.

## Conclusion

This study revealed that for older men living in a neighborhood with a higher perceived proximity to exercise facilities was inversely associated with CVD risk, while living in neighborhoods where traffic accidents and reduced safety were perceived as being more common was positively associated with CVD risk. This study suggests that efforts by the Japanese government to create a healthy living environment might be important in combating CVD morbidity and mortality in Japan, particularly among men.

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