厚生労働科学研究費補助金（政策科学総合研究事業（統計情報総合研究））
総合研究報告書

縦断調査を用いた中高年者の生活実態の変化とその要因に関する研究

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研究要旨

本研究の目的は、「中高年者縦断調査」を用いて中高年者の就業、健康、社会活動等の実態を把握し、退職前後の行動変容等についてパネル・データの特性を生かした実証分析を行い、全世代型社会保障に向けた施策に資する基礎資料を得ると共に縦断調査の利活用を進めることである。得られた主要な結論は、次の４つである。第１に、家族介護の女性の労働供給に及ぼす影響は統計的に有意でマイナスだが、かなり限定的である。第２に、学歴による健康格差は加齢によって拡大し、社会参加や余暇時間での運動などいくつかの要因が無視できない程度でその影響を媒介している。第３に、引退は多くの健康変数に異なる形で影響を及ぼすことが明らかになった。例えば、男性の場合、引退が良好な形で即座に影響を及ぼすものとして、余暇時間での運動や主観的健康感、心理的ディストレスが挙げられる。また、引退後の変化のペースが望ましい方向に変化するものとしては、喫煙や余暇時間での運動、主観的健康感、有症が挙げられる。第４に、自分の親または配偶者の親が要介護状態になった場合、中高年女性が介護者になる確率は３０％前後だが、親との同居や労働供給への影響はいずれも軽微か、あるいは統計的に有意ではない。その結果、親が要介護状態になることによる中高年女性のメンタルヘルス悪化のうち、介護者になることで説明される部分は全体の４割前後とかなり高めとなる。

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A．研究目的

「ニッポン一億総活躍プラン」（平成 28 年６月）に示された高齢者の就労促進・社会参加が確保される社会、高齢者と現役世代共に安心して生活できる社会保障制度を構築することが課題となっている。高齢化の進行は、年金受給者の増加と労働力需給のみならず中高年者の世帯構造・介護状況、所得・資産、医療・介護サービス利用など、さまざまな影響を及ぼしている。したがって、全世代が安心して生活できる社会保障制度を構築するためには、中高年者の生活実態・健康状態を把握し、社会保障制度や社会経済的要因を考慮しながら退職前後の行動変容を分析

し、政策に資する基礎的資料を得ることが必要である。本研究の目的は、「中高年者縦断調査」を用いて中高年者の就業、健康、社会活動等の実態を把握し、退職前後の行動変容等についてパネル・データの特性を生かした実証分析を行い、全世代型社会保障に向けた施策に資する基礎資料を得ると共に縦断調査の利活用を進めることである。

B．研究方法

1 年目の平成 28 年度においては、縦断調査を用いた分析の論点整理のために先行研究の文献研究を行うとともに、既存パネル・データを用いた分析を幾つか行った。後者に関しては、「全国消費実態調査」都道府県別・要介護者の有無別・世帯 1 ヶ月当たり消費額・有業人員数のデータを用いた介護状況と就業・医療支出との関係に関する分析、中高年者を対象とする「くらしと仕事に関する中高年インターネット特別調査」を用いた公的年金の受給開始年齢の引き上げと在職老齢年金制度の就労に関する分析を行った。

2 年目の平成 29 年度においては、「中高年者縦断調査」のパネル・データが利用可能になったため、本格的なパネル分析を展開し、その代表的な成果を学術論文の形にまとめ、投稿した。

（倫理面への配慮）

政府の公的統計の二次利用に基づく分析であり、倫理面への追加的な配慮は不要。

C．研究結果

今回の研究成果のうち、「中高年者縦断調査」を用いて学術論文という形で結実した主要なものは次の４つである。

第１に、家族介護の女性の労働供給に及ぼす影響を固定効果分析で推計すると、統計的に有意なマイナスの影響が確認できるが、その影響はかなり小さめであり、労働供給を 3.2%減少させるにとどまっている。また、労働供給を続ける場合も、家族介護によって労働日数や労働時間はほとんど変化していない。一方、メンタルヘルスに及ぼす影響を見ると、家族介護はマイナス、雇用はプラスとなっているが、両者の交絡項の係数は有意ではないことが分かった。

第２に、学歴による健康格差が加齢によって拡大するという仮説「累積的不利仮説」(cumulative disadvantage hypothesis)を支持する結果が得られた。また、学歴による健康格差拡大においては、社会参加や余暇時間での運動などいくつかの要因が無視できない程度で媒介していることも明らかとなった。

第３に、引退が多くの健康変数に異なる形で影響を及ぼすことが明らかになった。男性の場合、引退が良好な形で即座に影響を及ぼすものとして、余暇時間での運動や主観的健康感、心理的ディストレスが挙げられる。また、引退後の変化のペースが望ましい方向に変化するものとしては、喫煙や余暇時間での運動、主観的健康感、有症が挙げられる。しかし、過度な飲酒は加齢や引退の影響を受けないことが分かった。

第４に、自分の親または配偶者の親が要介護状態になった場合、中高年女性が介護者になる確率はそれぞれ 30.9%、30.3%となるが、親との同居や労働供給への影響はいずれも軽微か、あるいは統計的に有意ではないことが分かる。その結果、親が要介護状態になることによる中高年女性のメンタルヘルス悪化のうち、介護者になることで説明される部分は全体の 4 割前後とかなり高めとなる。対照的に、同居や労働供給面での調整による媒介効果は限定的である。

D．考察

「中高年者縦断調査」は、豊富な調査項目

を含むこと、調査期間が長い（10 年以上）こと、サンプルの脱落率が比較的少なく（脱落率：毎回平均４%）パネル・データとしてのサンプル数が多いことなど多くのメリットがある（既存の個票レベルのパネル・データとして、JSTAR と JAGES があるが、前者は 10 地域に限られており、後者はやや高齢層に偏っておりまだ全国ではない）。

　したがって、クロス集計や年齢階層別のデータごとの回帰分析に加えて、変数間の相互影響・内生性を考慮した複数の推定方法による比較分析、生存時間分析、パネル・データ分析など新しい推定方法を応用して、中高年者の生活実態の把握と引退過程における政策と行動変容に関わる実証分析ができるというメリットがある。

　特に、健康関連変数と社会経済変数がともに豊富で学際的研究に有効である（例１：社会疫学・公衆衛生・老年学における学際的研究、例２：労働経済学・家計経済学・家族社会学での学際的研究）。ただし、居住地情報は第 1 回調査のみであるため、少子高齢化と都市部と地方との格差是正のために関心がもたれている人口移動に関する研究や最近発展している GIS を応用した研究については、同調査を用いた分析が難しいという課題が残されている。

E．結論

　本研究の結果は、「中高年者縦断調査」を駆使した実証分析によって、厚生労働行政の各分野の政策立案に資する基礎的資料のみならず、各分野の政策の連携に資する基礎的資料を提供することが可能になることを示唆している。

F．健康危険情報

G．研究発表

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H．知的財産権の出願・登録状況

１．特許取得

なし

２．実用新案登録

なし

３．その他

なし

（資料）

Takashi Oshio and Emiko Usui, "How does informal caregiving affect daughters' employment and mental health in Japan?" *Journal of the Japanese and International Economies*, 2018, in press.

Abstract
We examine the association of informal caregiving with daughters' employment and mental health in Japan, using the 2008–2014 waves of the Longitudinal Survey of Middle-Aged and Older Adults, a large and nationally representative panel survey of middle-aged Japanese individuals. We find that caregiving reduces the probability of employment by only 3.2 percent, after controlling for time-invariant individual heterogeneity, and is not associated with either the hours or days worked per week by working caregivers. We further observe that employment does not add to the psychological distress already being experienced by the caregivers as a result of their caregiving role.

1. Introduction

The use of female labor is currently a major policy challenge in Japan due to the declining prime working-age population and the rapidly increasing elderly population, due to reduced fertility and longevity of the elderly. Increasing the participation of women in the labor market is crucial for the growth of Japan's economy. However, Japan is a country in which approximately 70 percent of elderly care is provided at home, mainly by women (Cabinet Office, 2015). Therefore, it is essential to investigate whether and how informal caregiving by women might negatively affect their level of employment.

As discussed by Bauer and Sousa-Poza (2015) and Lilly et al. (2007), many previous studies conducted in other advanced countries— mainly in the United States and Europe—have shown that the effect of informal caregiving on employment is relatively limited, even though caregiving and low levels of employment are combinedly prevalent. However, the association between caregiving for elderly parents and the female labor supply in Japan has not yet been fully investigated.

We use the 2008–2014 waves of the Longitudinal Survey of Middle-Aged and Older Adults, a large and nationally representative panel survey of middle-aged Japanese individuals. From the cross-sectional dataset, we find a negative association between caregiving for elderly parents and women's labor supply at both the extensive margin (employment probability) and the intensive margin (hours worked conditional on employment). However, after controlling for time-invariant individual heterogeneity by fixed effects, we observe that informal caregiving reduces the probability of employment only modestly—by 3.2 percent. Furthermore, working women do not reduce their hours or days worked per week at the onset of caregiving for their elderly parents.

We further investigate how work affects the association between informal caregiving and caregivers' mental health. It is well known that informal caregiving has an adverse impact on caregivers' mental health (Coe and Van Houtven, 2009; Hiel et al., 2015; Oshio, 2014; Pinquart and Sörensen, 2003). However, whether employment exacerbates the adverse impact of caregiving has not been sufficiently studied either within or outside Japan. We find that work neither increases nor decreases the adverse impact of caregiving on the mental health of caregivers.

Overall, our results suggest that informal caregiving does not appear to be a significant deterrent to employment among middle-aged women in Japan. This may be because Japanese women tend to work shorter hours and have limited responsibility at work; in many cases, they can participate in informal caregiving without needing to significantly adjust their labor force participation. This situation is consistent with our observation that employment does not add to the caregivers' psychological distress.

The paper proceeds as follows. Section 2 reviews the literature on how providing informal caregiving affects caretakers' level of employment and their mental health. Section 3 provides details about the data and descriptive statistics of the sample. Section 4 presents the main estimation results, including the effect of informal caregiving on (1) employment; (2) hours of work conditional on working; and (3) caregivers' mental health. Section 5 concludes the paper.

2. Background

Many studies in the United States and Europe have examined the effect of informal caregiving on employment. These studies have raised the possibility that the observed large negative association between caregiving and employment may be biased for two reasons. The first reason is endogenous selection into caregiving, as women with a weaker attachment to the labor market are more likely to take on the caregiving role. To control for the potential endogeneity of caregiving, we applied the instrumental variable (IV) approach. Previous studies have used measures of parental health, such as health status and/or daily activities, as instruments for informal caregiving (Crespo and Mira, 2014; Meng, 2012; Nguyen and Connelly, 2014; Van Houtven et al., 2013), as well as the number of the woman's siblings (Coe and Van Houtven, 2009). Second, researchers have been concerned that time-invariant unobserved individual heterogeneity may be related to caregiving because caregivers may differ in human capital investment or experience. To control for individual heterogeneity, previous studies have used a fixed-effects (FE) approach (Leigh, 2010; Meng, 2012; Van Houtven et al., 2013).

Studies in the United States and European countries that have used these two approaches have found a limited association between caregiving and women's probability of working. These studies have also found that caregiving is associated with a relatively moderate reduction in work hours (Bolin et al., 2008; Lilly et al., 2010; Meng, 2012; Van Houtven et al., 2013). Therefore, studies from the United States and European countries imply that caregivers may be able to adjust their working hours and may not have to exit the labor force to care for elderly parents.

However, the link between informal caregiving and work has not been studied extensively in Japan. Using repeated cross-sectional data from the Comprehensive Survey of Living Conditions released by the Ministry of Health, Labour and Welfare, Sugawara and Nakamura (2014) show that the presence of coresiding elderly parents who require care reduces the probability of coresiding, middle-aged women continuing as regular workers. Using repeated cross-sectional data from the Labor Force Survey and the Employment Status Survey, Kondo (2016) finds that the availability of long-term care (LTC) facilities is not related to the labor force participation of middle-aged women. However, neither of these studies focuses directly on the way that caregivers' employment decisions are affected by caregiving activities because the data used in these two studies lack information on (i) whether all of the elderly parents (namely, father, mother, father-in-law, and mother-in-law) are alive, and (ii) whether middle-aged people who have living elderly parents actually provide them with care.

Two studies use panel data to control for individual heterogeneity in Japan. Shimizutani et al. (2008) observe that the introduction of a public long-term care insurance (LTCI) scheme in 2000 increased the probability of female caregivers being employed and increased the number of days per week and hours per day worked by female caregivers (Tamiya et al., 2011). In contrast, Fukahori et al. (2015) find that the LTCI system does not mitigate the adverse impact on the employment of middle-aged individuals who reside with an elderly person needing care. Because of these contrasting results regarding the impact of informal caregiving on caregivers' employment, this issue should be investigated using a large and nationally representative sample in Japan.

Meanwhile, as mentioned above, a growing number of studies have demonstrated that informal caregiving increases the psychological distress experienced by caretakers (Coe and Van Houtven, 2009; Hiel et al., 2015; Oshio, 2014; Pinquart and Sörensen, 2003; Sugihara et al., 2004). However, these studies have not examined the effects, if any, that working could have on caregivers' mental health. One might assume that caregivers would feel more stressed if they continue to work because of reduced leisure and personal time. However, the multiple roles performed by people may just as likely have positive mental health outcomes (Adelmann, 1994; Moen et al., 1992). Particularly, participating in the labor force has been shown to have a favorable impact on the mental health of middle-aged and elderly individuals (Hao, 2008), and retirement tends to have a negative effect on one's health (Kim and Moen, 2002). Hence, it is interesting to examine whether work adds to, or reduces, caregivers' psychological distress. Caregiving and continuing to work in the labor market may exacerbate psychological distress due to a decrease in leisure time; however, the performance of multiple fulfilling roles may also reduce psychological distress.

3. Data and descriptive statistics

3.1 Data

We use panel data from the Longitudinal Survey of Middle-Aged and Older Adults, conducted by the Japanese Ministry of Health, Labour and Welfare. The survey began in early November 2005 with a sample of 34,240 individuals aged 50 to 59 years, and these individuals have been surveyed every November in subsequent years. The initial response rate of the survey was 83.8 percent, with a subsequent attrition rate ranging from 1.2 percent to 9.8 percent. Because of the large sample size and low attrition rate, as well as the availability of information on (i) parent(s) or parent(s)-in-law who are still living; (ii) care needs of those alive; and (iii) which of those elderly parents are being cared for by the respondent, this survey is one of the most effective ways to study the association between informal caregiving and the employment and mental health of middle-aged women in Japan.

We focus on women, who are usually considered reliable resources for providing informal care for elderly parents, especially in Japan. Japanese women often face a situation of having to decide whether to (i) provide care for their elderly parents while continuing to participate in the labor market or (ii) stop doing one in order to focus on the other. We restrict our sample to female respondents between the ages of 50 and 59 who have at least one living parent or parent-in-law. We exclude women over age 60 from our sample, considering that their work decisions are likely to be affected by pension and retirement policies: workers in Japan can claim pensions starting at age 60, and the mandatory retirement age is often between the ages of 60 and 65.[1] We also limit our sample to the years 2008–2014 because the data from the earlier waves—between 2005 and 2007— do not include information on the family member(s) requiring care. We are left with a total of 21,788 observations for the 7,415 female respondents in the sample.

Regarding employment, the respondents are asked whether they have a paid job. The indicator variable for employment is defined as 1 if the respondent has a paid job and 0 otherwise. Those who have a paid job are then asked about (i) their average hours worked per week and (ii) their average days worked per week during October— the most recent month because the survey is conducted in early November—of the survey year. Regarding informal caregiving, the survey asks whether the respondents provide care to their immediate family (including father, mother, father-in-law, and mother-in-law), and if they do so, the family member(s) who receive care. We consider a respondent an informal caregiver if she cares for at least one of her parent(s) or parent(s)-in-law or both.

As instrumental variables for the caregiving decision, we use four indicator variables for the demand for care for the father, mother, father-in-law, and mother-in-law. "Care" in this survey means all activities, such as formal, informal, and at-home or institutionalized care, although these are not specified in detail in the questionnaire given to the respondents. The elderly parent's need for care is negatively related to how healthy that parent is, and this is likely to affect the respondent's involvement in informal caregiving in a largely exogenous way.

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[1] The results remain largely unchanged even if we include women aged 60 years or above.

We also consider the respondents' mental health problems using the Kessler Screening Scale for Psychological Distress (K6). K6 score is a standardized and validated measure of nonspecific psychological distress (Kessler et al., 2002, 2010). The K6 contains six questions that ask whether the following feelings have been experienced in the past 30 days: (a) nervousness; (b) hopelessness; (c) restlessness or fidgeting; (d) depression; (e) feeling that everything was an effort; and (f) worthlessness. These items are rated on a five-point scale, ranging from 0 (none of the time) to 4 (all of the time). The items are summed to provide a score that ranges from 0 to 24. The reliability and validity of this tool have been demonstrated for a Japanese sample (Furukawa et al., 2008; Sakurai et al., 2011). Higher K6 scores indicate higher levels of psychological distress in the respondent.

3.2 Descriptive Statistics

Table 1 provides summary statistics of the key variables by caregiving status using the pooled sample of the 2008–2014 waves. Among women who have at least one living parent or parent-in-law or both, 18.0 percent (= 3,914/21,788) provide informal care to at least one parent or parent-in-law. When caregivers and non-caregivers are compared, caregivers tend to have somewhat poorer health and fewer children younger than 18 years old.

We then compare whether the employment and mental health variables differ by caregiving status in the upper panel of Table 2. The proportion of caregivers who have paid jobs is 68.8 percent, which is 6.8 percentage points lower than that among non-caregivers. Caregivers who have paid jobs work an average of 31.6 hours per week and 4.7 days per week—both values being somewhat less than those among non-caregivers (33.4 hours and 4.8 days). Meanwhile, the average K6 score was more among caregivers (4.74) than non-caregivers (3.53).

The lower panel of Table 2 shows the relationship between care demand and the prevalence of actual caregiving for each of the parents and parents-in-law. Having parent (s) or parent (s)-in-law who need care is positively related to the daughter or daughter-in-law becoming a caregiver. However, it should be noted that this relationship is not one-to-one; among non-caregivers, 4.5, 10.4, 3.1, and 10.6 percent have a father, mother, father-in-law, and mother-in-law, respectively, who requires care. This finding implies that caregiving is provided not only by women but also by other family members and institutions.

4. Estimation Results

4.1 Caregiving and work on the extensive margin: employment probability

We estimate a linear probability model in which the dependent variable is the indicator of having a paid job. The independent variables include an indicator of providing care to at least one parent or parent-in-law, in addition to a set of control variables. In line with the literature, the control variables consist of the woman's age and its square, self-assessed health, physical functional limitations, education, marital status, the number of children, whether the respondent is living with children younger than 18 years old, whether the household has a home mortgage, and survey years.

First, we estimate the model by ordinary least squares (OLS). Second, we estimate the instrumental variable (IV) model treating informal caregiving as endogenous. We use four indicator variables of each parent and parent-in-law's need for care as the instrument to explain caregiving. Third, we estimate the fixed-effects (FE) model to control for time-invariant individual heterogeneity. Finally, we estimate the fixed-effects models with instrumental variables (FE-IV) to control for both endogeneity of informal caregiving and time-invariant individual heterogeneity. Table 3 summarizes the main estimation results, and Table 4 provides the first-stage estimation results in IV and FE-IV models.

In these regression analyses, probit or logit models rather than linear probability ones could be an alternative approach because the dependent variable is binary. We choose linear probability models for two reasons in addition to its interpretability. First, many preceding studies (e.g., Crespo and Mira, 2014; Heitmueller, 2007; Leigh, 2010; Leigh, 2010) have employed linear probability models, facilitating the comparison of our estimation results with theirs.[1] Second, the sample size in the FE logit models would be substantially reduced because the respondents whose job statuses were unchanged throughout the sample periods were dropped from regressions. It should be also noted that unbiased FE probit models cannot be constructed. However, we should be cautious in interpreting the results of linear probability models because the estimated coefficients can imply probabilities outside the unit interval [0, 1].

As seen in Table 3, the OLS estimate of the coefficient on caregiving is –0.054, which is negative and significant, a result consistent with the finding that the proportion of workers among caregivers is 6.8 percentage points lower than that among non-caregivers (see Table 2). After controlling for the potential endogeneity of caregiving, the IV estimate of the coefficient on caregiving is –0.072, which is significant and slightly larger compared to the OLS estimate.

Nevertheless, we should be cautious in interpreting the validity of the IV model. The left columns of Table 4 present the first-stage regression results of the IV model. We found that the instruments used in the first-stage regression—that is, the four variables of the demand for care—are significantly and positively associated with caregiving. However, the *p*-value of the endogeneity test is 0.320, indicating that the null hypothesis that caregiving is exogenous cannot be rejected. Hence, we conclude that informal caregiving is largely exogenous in terms of the relationship with employment status among Japanese middle-aged women, indicating that potential endogeneity is not a serious concern.

Turning to the FE model, we find its estimate of the coefficient on caregiving (–0.032), despite being significant, is

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[1] We have also estimated probit and logit models and obtained similar marginal effects as compared to the linear probability model. Specifically, the marginal effects are –0.054 (0.012) for the probit model and –0.053 (0.012) for the logit model. These are very similar to estimates as compared to the estimates from the linear probability model in Table 3, –0.054 (0.012).

somewhat smaller in magnitude than the OLS and IV estimates. We also find that the *F*-test of the null hypothesis that all individual-level error terms in the FE model are equal to zero can be rejected (*p*-value <0.001; not reported in the table), confirming that the FE model is preferred to the OLS model. Another finding is that the FE-IV estimate is somewhat higher higher than the FE one, but the validity of the FE-IV model is questionable, because Table 4 shows that the exogeneity of informal caregiving cannot be rejected, as was the case with the FE model; the *p*-value of the endogeneity test is 0.223.

The key results obtained from Tables 3 and 4 can be summarized as follows. First, similar to the findings in many studies that estimate the IV models (e.g., Bolin et al., 2008; Crespo and Mira, 2014; Nguyen and Connelly, 2014; Van Houtven et al., 2013), the endogeneity of informal caregiving is less of a concern in our results in both the cross-sectional and FE models. Second, due to the association between the time-invariant individual heterogeneity and the regressors, the negative association between informal caregiving and employment is overestimated when not accounting for the individual heterogeneity.

In addition to these key results for caregiving, we obtain noteworthy findings about the associations between other variables and employment in Table 3. First, employment is negatively related to having two or more physical functional limitations and positively associated with having a home mortgage, even after controlling for time-invariant individual heterogeneity. Second, lower levels of self-assessed health are negatively associated with employment in the OLS and IV models, but their associations become insignificant in the FE and FE-IV models, suggesting that the cross-sectional correlation between employment and health is confounded by common time-invariant factors. Third, the confounding effects of time-invariant factors matter also for the associations of employment with divorced/widowed and never married status, both of which are positive in the OLS and IV models but negative in the FE and FE-IV models.

4.2 Caregiving and work on the intensive margin: hours and days worked conditional on employment

We further examine how informal caregiving is associated with the labor supply on the intensive margin. Specifically, for individuals who have paid jobs, we regress informal caregiving on hours worked per week and days worked per week separately, along with a set of covariates described in Section 4.1. Table 5 reports the results, focusing on the estimated coefficients on caregiving in each model. Caregiving reduces hours worked per week by 1.92 hours in the OLS model. This is largely consistent with the results from Table 2, which show that caregivers work 1.8 hours fewer per week than non-caregivers. The IV estimate provides a very close estimate—a reduction of 1.91 hours—although we confirm that the hypothesis that caregiving is exogenous cannot be rejected (not reported in the table), as in the case of employment models. By contrast, the impact of caregiving on hours worked per week is –0.31 and 0.13 in the FE and FE-IV models, respectively, which are small and insignificant.

We obtain similar results for the relationship between caregiving and days worked per week, as shown in the bottom panel of Table 5. Caregiving reduces 0.12 and 0.16 days worked per week in the OLS and IV models, respectively, both in line with the result in Table 2, which shows that the work week of caregivers is 0.14 days shorter than that of non-caregivers. The FE and FE-IV estimates are both 0.03, which reveal little association between caregiving and days worked per week.

Limited association between caregiving and working hours or days among working individuals—combined with a significantly negative, albeit small, association between caregiving and employments—suggests that caregivers may choose to remain in the labor force with the same working hours as before or leave the labor force completely without the opportunity of reducing working hours or days to adapt to caregiving. This may be due to the inflexibility of working hours or days in Japan, where workers are not allowed to adjust their working hours or days in response to family circumstances.[2] If the need of caregiving is too heavy to be met by reducing leisure time, women tend to stop working outside the home rather than reduce working hours. It is somewhat surprising, however, to see limited association between caregiving and working hours or days, considering that part-time workers constitute about 70 percent of middle-aged working women. The results suggest even part-time workers may have difficulty in adjusting working hours or days in accordance with the need of caregiving. More in-depth analysis is needed to explain why caregiving has limited association with work on both the extensive and intensive margins.

4.3 Impact of work on the association between caregiving and mental health

Last, we investigate how caregiving is associated with mental health and examine whether employment worsens the impact of caregiving on psychological distress. We regress psychological distress, measured by the K6 scores, on caregiving, employment, and the interaction between caregiving and employment, along with a set of control variables described in Section 4.1. We exclude self-assessed health, which is based on the respondent's subjective assessment and tends to overlap with psychological distress measured by K6 scores. Although many studies find a positive association between psychological distress and caregiving, few studies have examined how psychological distress is related to the situation in which employment and caregiving coexist. If the estimate of the coefficient on the interaction between employment and caregiving is positive, employment exacerbates caregivers' psychological distress; however, if it is negative, employment alleviates caregivers' psychological distress.

Table 6 presents the estimation results. The OLS model shows that psychological distress is associated positively with caregiving and negatively with employment; particularly, it indicates that psychological distress is not associated with the interaction between caregiving and employment. The IV model also gives similar results. In the FE and FE-IV models, psychological distress is positively associated with caregiving but not associated with employment or the interaction between

[2] Constructing the overemployment and underemployment indicators— as in Altonji and Paxson, 1988, 1992; Altonji and Usui, 2007; Usui, 2016; and Usui et al., 2016—shows that a significant proportion of Japanese workers are not satisfied with their working hours and that they are either overemployed or underemployed.

caregiving and employment.

Thus, regardless of the model specifications, our results confirm that work does not exacerbate the negative impact of caregiving on mental health. One plausible reason is that the positive mental health effect of performing multiple roles, which has been reported by Adelmann (1994), Hao (2008), and Moen et al. (1992), offsets the negative mental health effect of reduced leisure time and/or additional psychological pressures.

5. Conclusions

Based on the data from a large and nationally representative panel survey of middle-aged Japanese, we have obtained three noteworthy findings. First, the association between caregiving and employment is small in magnitude, albeit negative, after controlling for time-invariant individual heterogeneity. Second, caregiving is not related to either hours or days worked per week by the caregiver. Third, even though a negative association is found between caregiving and caregivers' mental health, employment does not increase the psychological distress experienced by the caregivers due to their caregiving role.

These results suggest that informal caregiving does not seriously harm employment for middle-aged women and that female caregivers can remain in the Japanese labor force without feeling additional psychological pressure from work. These findings may reflect the features of female employment in Japan. Women with paid jobs tend to work relatively short hours and tend to have jobs with limited responsibility, regardless of their caregiving status. In the sample of the current study, the average hours worked per week among working women is 31.6 hours for caregivers and 33.4 hours for non-caregivers (see Table 2). These hours are longer in the United States: 36.9 hours for those who have been caregivers at least once and 36.4 hours for those who have never been caregivers (Van Houtven et al., 2013). The hours are also longer in Europe: 36.5 hours for caregivers and 37.9 hours for non-caregivers (Sugano, 2015). Women in Japan also tend to be engaged in jobs with limited responsibility. Among the working women in our sample, only 2.9 percent hold managerial positions whereas 20.0 percent and 20.6 percent hold clerical and service positions, respectively. By comparison, the share of managerial, clerical, and service positions among working men in the sample is 18.4 percent, 8.5 percent, and 7.5 percent, respectively. Therefore, if middle-aged women were given the same roles and responsibilities at work as men, caregiving could have a larger impact on their employment.

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Table 1. Key features of respondents

		All	Caregivers	Non-caregivers
Age	*M* (*SD*)	56.9 (1.8)	57.0 (1.7)	56.8 (1.8)
Number of living children	*M* (*SD*)	2.1 (1.0)	2.0 (1.0)	2.1 (1.0)
Proportions (%)				
Marital status	Married	89.1	89.2	89.1
	Separated	2.7	2.3	2.8
	Divorced/widowed	5.8	5.2	6.0
	Never married	2.2	3.4	2.0
Educational attainment	Less than high school	9.5	7.1	10.0
	High school	51.1	47.3	51.9
	Some college	28.8	32.5	28.0
	University	10.1	12.8	9.5
	Other	0.5	0.3	0.5
Self-assessed health	Excellent	4.8	3.2	5.1
	Very good	31.8	26.5	33.0
	Good	47.0	49.1	46.5
	Fair	13.3	17.6	12.4
	Poor	2.5	2.8	2.5
	Very poor	0.6	0.8	0.5
Physical functional limitation	One	3.8	5.6	3.4
	Two or more	5.1	5.5	5.0
Having children younger than 18 years old		2.6	1.8	2.8
Home mortgage		27.2	25.1	27.6
N		21,788	3,914	17,874

Table 2. Labor supply variables and K6 scores by caregiving status

	Caregivers (A)		Non-caregivers (B)		Difference[a] (A–B)	
	M	(*SD*)	*M*	(*SD*)	*M*	(*SE*)
Employment	0.621	(0.485)	0.688	(0.463)	-0.068	(0.009)
Hours worked per week	31.59	(14.81)	33.39	(14.47)	-1.80	(0.33)
Days worked per week	4.69	(1.35)	4.83	(1.18)	-0.14	(0.03)
K6 score (range: 0–24)	4.74	(4.54)	3.53	(4.11)	1.21	(0.08)
Father needs care	0.179	(0.383)	0.045	(0.207)	0.134	(0.006)
Mother needs care	0.506	(0.500)	0.106	(0.307)	0.400	(0.008)
Father-in-law needs care	0.122	(0.327)	0.032	(0.175)	0.090	(0.005)
Mother-in-law needs care	0.391	(0.488)	0.107	(0.309)	0.284	(0.008)
N		3,914		17,874		

[a] All significant at the 0.1% significance level.

Table 3. The estimated association between informal caregiving and employment (N = 21,788)[a]

Dependent variable = Employment

	OLS		IV[b]		FE[c]		FE-IV[b]	
	Coef.	(SE)	Coef.	(SE)	Coef.	(SE)	Coef.	(SE)
Caregiving	-0.054***	(0.012)	-0.072**	(0.023)	-0.032***	(0.009)	-0.041*	(0.020)
Age	0.119	(0.102)	0.119	(0.102)	0.288***	(0.086)	0.286***	(0.085)
Age square	-0.012	(0.009)	-0.012	(0.009)	-0.026***	(0.007)	-0.026***	(0.007)
Marital status (ref. = married)								
Separated	-0.012	(0.031)	-0.012	(0.031)	-0.003	(0.019)	-0.003	(0.019)
Divorced/widowed	0.167***	(0.019)	0.166***	(0.019)	-0.109*	(0.048)	-0.108*	(0.048)
Never married	0.189***	(0.030)	0.190***	(0.030)	-0.067***	(0.008)	-0.068***	(0.009)
Educational attainment (ref. = high school)								
Less than high school	0.031	(0.018)	0.030	(0.018)				
Some college	0.001	(0.013)	0.002	(0.013)				
University	-0.017	(0.021)	-0.016	(0.021)				
Other	-0.051	(0.081)	-0.053	(0.081)				
Self-assessed health (ref. = fair)								
Excellent	0.015	(0.020)	0.014	(0.020)	-0.010	(0.013)	-0.010	(0.013)
Very good	0.011	(0.009)	0.010	(0.009)	-0.002	(0.005)	-0.002	(0.005)
Good	-0.086***	(0.014)	-0.086***	(0.014)	-0.014	(0.008)	-0.014	(0.008)
Poor	-0.157***	(0.026)	-0.156***	(0.026)	-0.024	(0.018)	-0.024	(0.018)
Very poor	-0.272***	(0.054)	-0.271***	(0.054)	-0.017	(0.042)	-0.016	(0.043)
Physical functional limitation								
One	-0.053*	(0.021)	-0.052*	(0.021)	-0.002	(0.013)	-0.002	(0.013)
Two or more	-0.120***	(0.022)	-0.120***	(0.022)	-0.037**	(0.013)	-0.037**	(0.013)
Number of living children	0.030***	(0.006)	0.030***	(0.006)	0.002	(0.005)	0.002	(0.005)
Children aged < 18 years	0.004	(0.030)	0.003	(0.030)	0.001	(0.019)	0.001	(0.019)
Home mortgage	0.067***	(0.011)	0.067***	(0.011)	0.026*	(0.011)	0.026*	(0.011)

[a] Adjusted for survey years. Robust standard errors clustered at the individual level are shown in parentheses.
[b] See Table 4 for the results of first-stage estimation.
*** $p < 0.001$, ** $p < 0.01$, * $p < 0.05$.

Table 4. The first stage estimation results in IV and FE-IV models (*N* = 21,788)[a]

Dependent variable = Caregiving

	IV			FE-IV		
	Coef.		(SE)	Coef.		(SE)
Father's need for care	0.237	***	(0.017)	0.185	***	(0.017)
Mother's need for care	0.368	***	(0.012)	0.254	***	(0.013)
Father-in-law's need for care	0.249	***	(0.022)	0.158	***	(0.021)
Mother-in-law's need for care	0.288	***	(0.012)	0.214	***	(0.014)
Age	-0.006		(0.080)	-0.139		(0.086)
Age square	0.001		(0.007)	0.01		(0.007)
Marital status (ref. = married)						
Separated	-0.031	*	(0.015)	-0.001		(0.018)
Divorced/widowed	0.002		(0.012)	0.102	*	(0.059)
Never married	0.095	***	(0.023)	-0.038		(0.090)
Educational attainment (ref. = high school)						
Less than high school	-0.029	**	(0.010)			
Some college	0.023	**	(0.008)			
University	0.028	*	(0.012)			
Other	-0.044		(0.038)			
Self-assessed health (ref. = fair)						
Excellent	-0.031	**	(0.012)	-0.014		(0.012)
Very good	-0.018	***	(0.006)	0.000		(0.006)
Good	0.029	***	(0.009)	0.019	*	(0.009)
Poor	-0.005		(0.018)	-0.002		(0.018)
Very poor	0.022		(0.037)	0.102	*	(0.043)
Physical functional limitation						
One	0.037	*	(0.015)	0.026		(0.014)
Two or more	-0.030	*	(0.013)	0.002		(0.013)
Number of living children	-0.001		(0.003)	-0.002		(0.007)
Children aged < 18 years	-0.028		(0.015)	-0.033		(0.022)
Home mortgage	-0.016	*	(0.007)	0.023	*	(0.012)
Endogeneity test[b]	0.320			0.223		
Joint significance of instruments[c]	< .001			< .001		
Overidentification test[d]	0.821			0.689		

[a] Adjusted for survey years. Robust standard errors clustered at the individual level are shown in parentheses.[b] *p*-value of the null hypothesis of exogeneity.[c] *p*-value of the null hypothesis of no joint significance. [d] *p*-value of the null hypothesis of valid exclusion restrictions.
*** $p < 0.001$, ** $p < 0.01$, * $p < 0.05$.

Table 5. The estimated associations of informal caregiving with hours worked per day and days worked per week[a] (N = 14,384)

	OLS			IV			FE		FE-IV	
	Coef.		(SE)	Coef.		(SE)	Coef.	(SE)	Coef.	(SE)
Dependent variable = hours worked per day										
Caregiving	-1.92	***	(0.46)	-1.91	*	(0.89)	-0.31	(0.32)	0.13	(0.82)
Dependent variable = days worked per week										
Caregiving	-0.12	**	(0.04)	-0.16	*	(0.08)	0.03	(0.03)	0.03	(0.07)

[a] Adjusted for age, age square, marital status, educational attainment, self-assessed health, physical functional limitation, number of living children, living with children younger than 18 years old, having a home mortgage, and survey years. Robust standard errors clustered at the individual level are shown in parentheses. The complete results (including the first-stage estimation results for IV and FE-IV models) are available upon request from the author.
*** $p < 0.001$, ** $p < 0.01$, * $p < 0.05$.

Table 6. The association across informal caregiving, employment, and psychological distress[a] (N = 20,959)

Dependent variable = K6 score (0–24)

	OLS			IV			FE			FE-IV		
	Coef.		(SE)	Coef.		(SE)	Coef.		(SE)	Coef.		(SE)
Caregiving	1.09	***	(0.17)	1.84	***	(0.31)	0.69	***	(0.15)	1.14	***	(0.34)
Employment	-0.36	***	(0.10)	-0.29	*	(0.12)	-0.10		(0.11)	-0.12		(0.14)
Caregiving × Employment	0.07		(0.21)	-0.17		(0.39)	-0.13		(0.17)	0.02		(0.38)

[a] Adjusted for age, age square, marital status, educational attainment, physical functional limitation, number of living children, living with children younger than 18 years old, having a home mortgage, and survey years. Robust standard errors clustered at the individual level are shown in parentheses. The complete results (including the first-stage estimation results for IV and FE-IV models) are available upon request from the author.
*** $p < 0.001$, ** $p < 0.01$, * $p < 0.05$.

Table 4. Estimated proportions of the associations between education and health mediated by each mediator[a] (N = 20,024)

	Men (n = 9,320)		Women (n = 10,704)	
	Proportion (%)	95% CI[b]	Proportion (%)	95% CI
Self-rated health				
Low household spending	1.0	(-1.6, 3.5)	-3.6	(-6.1, -1.2)
Low social participation	31.1 †[c]	(24.4, 37.7)	30.7 †	(24.3, 37.2)
Low physical activity	15.0 †	(6.8, 23.2)	9.3 †	(4.2, 14.4)
Smoking	3.1 †	(0.5, 5.7)	2.0	(-1.2, 5.1)
Problem drinking	1.2	(-0.3, 2.6)	-0.8	(-2.2, 0.6)
No regular health check-up	6.0 †	(2.2, 9.7)	2.0 †	(0.1, 3.9)
Total	57.3 †	(28.9, 85.6)	39.5 †	(12.3, 66.8)
Total for significant mediators[d]	55.2 †	(44.2, 66.1)	42.0 †	(33.8, 50.1)
Functional limitations				
Low household spending	0.3	(-3.2, 3.8)	-2.3	(-5.7, 1.2)
Low social participation	41.8 †	(32.3, 51.2)	27.4 †	(18.7, 36.2)
Low physical activity	12.1 †	(7.4, 16.8)	7.3 †	(3.3, 11.3)
Smoking	3.9 †	(0.3, 7.6)	9.8 †	(-4.5, 15.1)
Problem drinking	0.8	(-1.0, 2.6)	-0.9	(-2.6, 0.9)
No regular health check-up	6.5 †	(3.0, 10.0)	4.9 †	(2.0, 7.9)
Total	65.4 †	(21.3, 109.5)	46.4 †	(4.3, 88.4)
Total for significant mediators[d]	64.3 †	(53.6, 74.9)	49.5 †	(38.8, 60.3)
Psychological distress				
Low household spending	2.4	(-0.8, 5.5)	-1.6	(-6.3, 3.1)
Low social participation	32.2 †	(24.1, 40.3)	46.2 †	(33.2, 59.1)
Low physical activity	7.7 †	(1.9, 13.5)	6.5 †	(0.2, 12.8)
Smoking	0.6	(-2.4, 3.6)	4.7	(-1.9, 11.4)
Problem drinking	1.3	(-0.5, 3.1)	0.2	(-2.1, 2.5)
No regular health check-up	7.4 †	(2.8, 11.9)	6.1 †	(2.0, 10.3)
Total	51.6 †	(16.0, 87.3)	62.1 †	(7.9, 116.4)
Total for significant mediators[d]	47.3 †	(36.9, 57.6)	58.8 †	(44.9, 72.7)

[a] Controlled for baseline value of each health variable as well as ages at baseline.

[b] 95% confidence intervals obtained via bootstrap estimation with 2,000 replications.

[c] Indicates that bootstrap-estimated 95% CI was above zero.

[d] Mediators with †

（資料）

Takashi Oshio, “Widening disparities in health between educational levels and their determinants in later life: evidence from a nine-year cohort study,” ***BMC Public Health*****,** ***BMC Public Health*****, 2018, 18:278**

Abstract

Background: Education has attracted more attention as a key determinant of health in later life. In this study, the hypothesis that widened educational disparities in health can be observed in later life was investigated, and the factors that mediated the association between education and changes in health were also assessed.

Methods: Using the 10-wave longitudinal data of 20,024 individuals (9,320 men and 10,704 women) aged 50–59 years at baseline, collected from a nationwide population survey in Japan (2005–2014), the changes in self-rated health, functional limitations, and psychological distress between educational levels were compared. Mediation analysis was further conducted to assess the factors that mediated the association between education and changes in health, with reference to six types of potential mediators (household spending, social participation, leisure-time physical activity, smoking, problem drinking, and regular health check-ups). The analyses were conducted separately for men and women.

Results: All three health variables rapidly deteriorated among lower-educated men and women. For men, the six potential mediators mediated 55.2%, 64.3%, and 47.3% of the associations between educational levels and changes in self-rated health, functional limitations, and psychological distress, respectively. The proportions for women were 42.0%, 49.5%, and 58.8%, respectively. Social participation was the primary mediator, followed by physical activity, regular health check-ups, and smoking. In general, no substantial or consistent differences were observed between men and women.

Conclusions: The results suggested that policy measures that encourage social participation and promote healthy behaviors can improve educational disparities in health in later life.

Background

Education as a key determinant of health in later life has attracted more attention because it is one of the most stable indicators of one’s socioeconomic status after young adulthood [1, 2]. Education is also likely to affect other aspects of socioeconomic status that are associated with health [3, 4]. A well-established view is that health differences between educational levels increase with age. Health is predicted to deteriorate more rapidly with age for lower-educated individuals than for higher-educated individuals, which is known as the cumulative disadvantage hypothesis [5]. In line with this hypothesis, several studies have demonstrated that educational level is a key determinant of health disparities in later life among other several aspects of health, including mortality, disability, frailty, chronic diseases, mental health, self-rated health, or other health variables [5–13].

However, two key challenges must be addressed for the further understanding of the association between educational levels and health. First, more information is needed about the long-term changes in health at an individual level, particularly if the focus is on how health disparities will accumulate with age over time. Previous studies have often been based on cross-sectional or repeated cross-sectional data [8, 12, 13], and even if longitudinal data were used, analyses have often been limited to comparisons between a couple of survey waves with relatively short intervals [5–7, 10, 11, 14], with a recent exception that used longer longitudinal data [9]. Further evidence based on large-scale and extended longitudinal data must be obtained to examine the validity of the cumulative disadvantage hypothesis at an individual level [7].

Second and more importantly, the mechanism that explains the relationship between educational levels and widening health disparities with age has not been fully elucidated. Numerous studies have examined the possible explanations for the general relationship between education and health [3, 15]. In addition, results showed several potential mediators of this relationship. For example, a lower educational level is likely to cause material disadvantages, particularly in terms of income, which can reduce the access to healthy food and the chances of living in healthy conditions [16, 17]. Lower-educated individuals may also undertake an unhealthy lifestyle or behaviour, resulting in higher risks of worsening health [4, 18]. In this respect, how health behaviours and lifestyle habits, such as leisure-time physical activity, smoking, and problem drinking can link education to health should be assessed. In addition, social participation may be a potential mediator if it is positively associated with educational level [14], given that studies have demonstrated a positive association between social participation and health [19, 20]. However, existing observations about these mediating effects are conflicting, and as suggested by Chandola et al. [21], multiple pathways that link education and health must be considered, rather than focusing on a single potential mediator.

To address these challenges and further understand the association between educational levels and health, the validity of the cumulative disadvantage hypothesis was assessed at an individual level, and the factors that mediated the association between education and changes in health were investigated. For these purposes, the 10-wave (9-year) longitudinal data, obtained from a nationwide social survey, of 20,024 individuals (9,320 men and 10,704 women) aged 50–59 years at baseline were used. The changes in health and its evolution over the 9-year period were compared between the lower- and middle-/higher-educated individuals, with a focus on three health variables (self-rated health [SRH], functional limitations, and psychological distress).

Furthermore, the factors that mediated the association between educational levels and changes in health over the 9-year period were assessed. Six potential mediators were considered: household spending (as an alternate for income), social participation, leisure-time physical activity, smoking, problem drinking, and regular health check-ups. A mediation analysis was conducted to evaluate the mediating effect of each of these six variables on the association between education and health.

All of these analyses were conducted separately for men and women. It might be possible that educational differences in health may increase differently among men and women, and that the mediating mechanisms might operate in different ways by gender.

Methods

Study sample

Data from the Longitudinal Survey of Middle-Aged and Older Adults, a nationwide 9-year panel survey, that was conducted by the Japanese Ministry of Health, Labour and Welfare (MHLW) each year between 2005 and 2014, were obtained. Samples in the first wave were collected nationwide in November 2005 through a two-stage random sampling procedure. First, 2,515 districts were randomly selected from 5,280 districts used in the MHLW's nationwide, population-based "Comprehensive Survey of the Living Conditions of People on Health and Welfare," which was conducted in 2004. The 5,280 districts were, in turn, randomly selected from about 940,000 national census districts. Second, 40,877 residents aged 50–59 years as of October 30, 2005 were randomly selected from each selected district, according to its population size. A total of 34,240 individuals responded (response rate: 83.8%). The second to tenth waves of the survey were conducted in early November of each year from 2006 to 2014, and 22,748 individuals remained until the tenth wave (with an average attrition rate of 4.0% in each wave). No new respondents were added after the first wave.

Data of the 20,024 individuals (9,320 men and 10,704 women), who participated for 9 years were used, and all information required in the present study were provided. The respondents were divided into lower-educated individuals (whose educational attainment was below high school, that is, less than 12 years of schooling in total) and middle-/higher-educated ones (who had graduated from high school or above). Lower-educated individuals, including those who had not completed high school, comprised 15.4% of the entire sample. The study sample consisted of 58.4% of the individuals who participated in the first wave. The key attributes between this study sample and dropouts were compared to assess the potential bias in the estimation results.

Measurements

Health variables

Three health variables were considered: SRH, functional limitations, and psychological distress. SRH has often been used as a comprehensive alternative for general health conditions in social epidemiology because it has been repeatedly found to be a valid predictor of health outcomes, including mortality, physical and cognitive functioning, and morbidity [22–24]. In terms of SRH, the respondents were asked to choose 1 (*very good*), 2 (*good*), 3 (s*omewhat good*), 4 (*somewhat poor*), 5 (*poor*), or 6 (*very poor*) regarding their current health condition.

These categorical answers were used as a continuous variable with higher values that indicate poorer SRH. In terms of functional limitations, the respondents were asked whether they had any difficulty in each of the 10 activities of daily living (walking, getting out of bed, getting in/out of a chair, dressing, washing their face and hands, eating, toileting, bathing, ascending and descending stairs, and carrying purchased items). The degrees of functional limitations were also evaluated using the sum of items in which the respondents had difficulty performing.

Kessler 6 (K6) scores were established to measure psychological distress [25, 26]. From the survey, the respondents' assessments of psychological distress were first obtained using a 6-item psychological distress questionnaire—"During the past 30 days, about how often did you feel a) nervous, b) hopeless, c) restless or fidgety, d) so depressed that nothing could cheer you up, e) that everything was an effort, and f) worthless?" The questionnaire was rated on a 5-point scale (0 = *none of the time* to 4 = *all of the time*). The sum of the reported scores were then calculated (range: 0–24) and defined as the K6 score. Higher K6 scores reflected higher levels of psychological distress.

Potential mediators

Six types of potential mediators were considered for the association between education and health (household spending, social participation, leisure-time physical activity, smoking, alcohol drinking, and regular health check-ups). Each variable was completely evaluated throughout the 10 waves. Household spending was considered as a key factor that represents the material conditions rather than income because the number of respondents who did not report household spending was significantly lower compared to those who did not own a household or have their own income and because dependent wives who did not work outside their house did not have income. Reported household spending throughout the 10 waves were summarized, and a binary variable of low household spending was established by allocating one to the lowest tertile of the sum and zero if otherwise.

In terms of social participation, respondents were asked whether they participated in 6 types of social participation (hobbies or cultural activities, exercise or sports, community events, support for children, support for the elderly, and other activities) within the past year from the date of the survey. The answers regarding social participation were summarized, showing that the respondents were engaged in each wave throughout the 10 waves (range: 0–60), and a binary variable of low social participation was established by allocating one to the lowest tertile of the sum and zero if otherwise.

Physical activity, smoking, alcohol drinking, and regular health check-ups were considered as key behaviours that are associated with health. Respondents were asked how they were engaged in leisure-time physical activity. A binary variable of low physical activity was then established by allocating one to those who did not engage in moderate (without breathlessness or heart palpitations) or more intense exercise at least few days per week throughout the 10 waves. The respondent was considered as a smoker if he/she answered that he/she was currently smoking all throughout the waves. Problem drinking was defined as an intake of more than two *go* (360 ml) per day of Japanese *sake* or an equivalent amount of alcohol, which corresponds to about 40 g of pure alcohol. This threshold was based on a study showing that maintaining alcohol consumption below 46 g/day minimized the risks of mortality in a Japanese population [27]. Those who drank above this threshold in at least one wave were considered as problem drinkers. Lastly, a binary variable for those with no regular health check-ups was established by allocating one to those who reported that they did not have a health check-up in at least one wave. In addition to these variables, binary variables of sex and each age (50–59 years old) at baseline and the baseline values of each health variable as covariates were used.

Statistical analyses

For the descriptive analysis, the baseline values and changes in the three health variables over the 9-year period between the lower- and middle-/higher-educated individuals were compared for both men and women. Then, two types of linear regression models (Models A and B) were estimated separately for men and women to explain the change in each health variable between baseline and each wave, allowing random effects to consider error terms at an individual level. In Model A, the wave was used as a continuous variable and the binary variable of low educational level as key explanatory variables, along with the covariates. The coefficients of the wave and low educational level were both expected to be positive. In Model B, the interaction term of the wave and low educational level were added to Model A. The coefficient of this interaction term was expected to be positive if low educational level adds to the pace of deterioration in health with increasing age.

In these regression models, each health variable was normalised by its mean and standard deviation to help assess and compare the substantive degrees of association between health and other variables. In addition, inverse probability weighting was used to mitigate potential sources of attrition bias [28, 29]. Specifically, the probit model was first estimated to predict observation presence through wave 10, using the baseline values of each health variable and the binary variables of lower education and each age at baseline. Then, the inverse of the predicted probability of presence was used as the weight when estimating the regression models.

Subsequently, a mediation analysis was performed separately for men and women, with the conventional three-step estimation procedure along with bootstrapping to assess the significance of the mediating effects [30, 31]. Changes in the three health variables over the 9-year period were the focus. In the first step, Model 1 was used to explain the change in each health variable between baseline and the tenth wave by the binary variable of low educational level. In the second step, Model 2 was used to explain each potential mediator by the binary variable of low educational level. In the third step, Model 3 was utilized to explain the change in each health variable by low educational level. In each of Models 1, 2, and 3, health variables at baseline as well as other covariates were controlled for.

For each potential mediator an actual mediator was suspected if the estimated coefficients of low educational level in Models 1 and 2 and the estimated coefficients for the potential mediators were all statistically significant. To examine the statistical significance of the mediating effect, the 95% confidence interval (CI) of the proportion of the association between education and the change in each health variable were subsequently estimated via bootstrap estimation with 2,000 replications.

Results

Widening disparities in health

The first half of Table 1 shows the comparison (1) of the values at baseline and (2) the changes over the 9-year period for the lower- and middle-/higher-educated individuals in terms of SRH, functional limitations, and psychological distress, between lower- and middle-/higher-educated individuals. For both men and women, SRH and functional limitations at baseline were worse among lower-educated individuals than middle-/higher-educated ones, whereas no difference was observed in terms of psychological distress. Over the 9-year period, self-rated health and psychological distress deteriorated among lower-educated men, while functional limitations and psychological distress deteriorated among lower-educated women. Deterioration in functional limitations or self-rated health showed no difference among men and women with varying educational backgrounds. However, it should be noted that baseline values of health variables or other covariates were not controlled for in Table 1.

The second half of Table 1 shows the comparison of the six potential mediators in terms of educational level over the 9-year period. Lower-educated individuals were at significantly higher risks of low household spending, low social participation, low physical activity, smoking, problem drinking, and no regular check-ups compared to middle-/higher-educated individuals, while the difference in the proportion of problem drinking was small for both men and women and significant only at the 10% level for women.

To confirm the widening educational disparities in health with age, Table 2 presents the estimation results of the regression models to explain the change in each health variable between baseline and each wave after controlling for sex and age at baseline. In Model A, results showed that, for both men and women, low educational level accelerated deterioration in health. This result was obtained even after controlling for (i) the adverse effect of aging on health (which is captured by a positive coefficient of a continuous variable of the wave), and (ii) the initial level effect (which means that higher initial levels reduced additional increases in subsequent waves and is indicated by a negative coefficient of the health variable at baseline).

By adding the interaction term between low educational level and the wave in Model B, the coefficients of the interaction terms were positive and significant in all models, except for self-rated health for women. This observation indicates that low educational level generally accelerated the deterioration in health with age for both men and men.

Mediation analysis

The estimation results of Models 1 and 3 based on the mediation analysis are presented in Table 3, which focuses on the change in health variables between baseline and the tenth wave. The results of Model 1 confirmed the adverse effect of low educational level on the changes in all the three types of health variables. The results of Model 2 are not presented to conserve space (available upon request), but it was confirmed that all potential mediators were significantly associated with low educational level ($p < 0.001$).

The results of Models 3 help understand the mediating mechanism. For example, in the case of SRH for men, the estimated coefficient of low educational level was substantially attenuated to 0.08 from 0.16 in Model 1, after controlling for the six potential mediators, suggesting that a substantial portion of the association between education and the change in SRH was influenced by those mediators. Among the six variables, low social participation, smoking, problem drinking, and no regular health check-ups were positively associated with deteriorated SRH. Household spending was not related to SRH. A reduction in the estimated coefficient of low educational level from Model 1 to 3 was commonly observed in all models, while the levels of the coefficient were somewhat different between men and women. Another finding was that estimation results of the six mediators were not much different between men and women in terms of the magnitudes and statistical significance of their estimated coefficients; notably, low social participation and physical activity were most closely associated with the changes in health variables in all models.

Table 4 shows the comparison of the magnitude of each variable's mediating effects as well as statistical significance. In the case of men's self-rated health, social participation had the largest mediating effect, which accounted for 31.1% of the association between educational levels and SRH. The magnitude of the mediating effect of social participation was remarkably higher than that of physical activity (15.0%), regular health check-ups (6.0%), and smoking (3.1%). The mediating effects of these four variables were all significant, given that the bootstrap-estimated 95% CI did not include zero. By contrast, the mediating effect of household spending or problem drinking was not significant. The mediating effect of these six potential mediators accounted for 57.3% (95% CI: 28.9—58.6%) of the association between low educational level and SRH. If limited to four significant mediators, the mediating effect was 55.2% (95% CI: 44.2—66.1%) in total.

Largely similar results were obtained for other combinations of the health variable and gender. For men, the six potential mediators accounted for 64.3% and 47.3% of functional limitations and psychological distress, respectively. The proportions for women were 42.0%, 49.5%, and 58.8% for the three health variables, respectively, not much different from those for men. For both men and women, social participation was the primary mediator for all health variables. Albeit to a lesser extent, leisure-time physical activity and regular health check-up, and smoking in some cases, were found to be important mediators for all health variables for both genders.

Discussion

In the present study, the association between the changes in health and educational levels were investigated using the 10-wave longitudinal data of the individuals aged 50–59 years old at baseline. The estimation results clearly support the hypothesis that educational disparities in health would accumulate with age in terms of SRH, functional limitations, and psychological distress. These results were generally in accordance with those in previous studies that demonstrated educational disparities in health [5–13], although the present study additionally revealed the changes in disparities over the 9-year period.

The results of the mediating analysis highlighted the importance of the pathways that link education to health in later life. The proportions of the association between educational levels and the change in heath mediated by a set of six factors (household spending, social participation, leisure-time physical activity, smoking, problem drinking, and regular check-ups) were in the range from 47.3% to 64.3% and from 42.0% to 58.8% for men and women, respectively, depending on health variables (self-rated health, functional limitations, and psychological distress). These results suggested that we can construct policy measures to alleviate the accumulation of educational disparities in health by blocking the pathways that link low educational level to health.

In this respect, the key mediators for the association between education and health must be identified. Moreover, the prediction of health behaviors as key mediators is also important, as already suggested by previous studies [4, 18]. Indeed, estimation results confirmed that leisure-time physical activity and, to a lesser extent, smoking mediated the effect of education on health, whereas problem drinking did not. In addition, regular health check-up, which is not a narrowly defined heath behaviour, was also an important mediator. This result is also consistent with the assumption that health literacy mediates the effect of education on health [32] because it is reasonable to argue that individuals with a higher level of health literacy are more inclined to have regular heath check-ups.

Another significant finding is that social participation was the primary mediator of the association between education and health because of the magnitude of its mediating effect was well above those of other factors for both men and women. Numerous studies have demonstrated that social participation has a favourable effect on health [19, 20]. The results of the present study suggested that lower-educated individuals are at high risk in failing (or be reluctant) to engage in social participation, which in turn affects the health of lower-educated individuals.

However, a one-way causation from social participation (as well as other mediators) may not affect health. Rather, a two-way causation between the two variables may be assumed, considering that healthier individuals are more likely to engage intensively in social participation, which in turn further enhances their health. This two-way causation between social participation and health may result in the accumulation of the mediating effect of social participation over time. Compared to the present study, Ettman et al. [14] indicated a more limited mediating effect of social participation between educational levels on frailty. The difference was probably attributed to the difference in the time intervals in observing the change in health: 2 years in the study by Ettman et al. versus 9 years in the present study.

In contrast to social participation and health behaviours, a somewhat surprising result was that household spending, which was used as an alternative for income, did not have any mediating effect on the association between educational levels and health. Two remarks should be made on this result. First, it may probably be wrong to argue for a limited mediating effect of income, because income may likely provide material sources to health-promoting behaviours, access to health service, and healthy lifestyle. In this sense, income may possibly arbitrate the mediating effects of other factors. Second, the present study, which focused on how income (along with other factors) mediated the effects of education on health, did not address the differential effects of education versus income on health, which should be addressed in another analytic framework [1, 33].

Finally, the results did not show any substantial differences between men and women, and gender differences depended on the types of health variable. For both men and women, educational disparities in health widened at a largely similar pace, albeit somewhat differently across health variables. In addition, the proportion of the association between education and health mediated by six potential mediators was in the range from 47.3% to 64.3% and 42.0% to 58.8% for men and women, respectively, which were largely overlapped. Moreover, the key mediator was social participation for both men and women, and physical activity, regular health check-up, and smoking worked as important mediators commonly for both genders. However, we should be cautious in any generalization, because the results may depend on socio-institutional backgrounds.

The present study has several limitations, in addition to the limited coverage of health variables: for instance, it did not analyse the educational difference in mortality due to lack of data availability from the current dataset. First, attrition biases were not fully controlled, although the study sample and dropouts were compared. Hence, as mentioned above, the association between educational levels and changes in health observed in the present study might have been underestimated. Second, potential mediators for the association between education and health were not comprehensively explored, although their association was significantly mediated by the six factors that were considered in the present study. Hence, one should be cautious in interpreting the proportion of the mediated association in Table 4. The remaining proportion did not indicate the magnitude of the direct unmediated effect of education on health. Third and most importantly, the possibility that a third unobserved factor that affects both education and the mediators exist was not ruled out. For instance, some genetic characteristics or personality trait can make an individual more inclined to both continue his/her education and participate in social activities. If that is the case, caution should be undertaken in interpreting the observed association between education and the mediators as well as its effects on the association between education and health.

Conclusions

Based on the statistical analyses using the 10-wave cohort data of the nationwide survey in Japan, educational disparities tended to widen with age in later life. In addition, a substantial portion of the associations between educational levels and changes in health was mediated by social participation and health-related activities, which contributed to a cumulative disadvantage of low educational level. These results suggested that policy measures that encourage social participation and promote healthy behaviours can improve educational disparities in health in later life.

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Table 1. Comparing health and other variables by educational level[a] (N = 20,024)

Educational level	Men (n = 9,320)				Women (n = 10,704)			
	Lower	Middle /higher	Difference		Lower	Middle /higher	Difference	
	(A)	(B)	A − B	p-value	(A)	(B)	A − B	p-value
Health variables								
(1) Values at baseline								
Self-rated health	2.84	2.68	0.16	< 0.001	2.91	2.68	0.23	< 0.001
(range: 1-6)	(0.97)	(0.94)			(0.96)	(0.90)		
Functional limitations	0.22	0.14	0.08	0.007	0.32	0.21	0.11	< 0.001
(range: 0-10)	(1.27)	(0.99)			(1.44)	(1.07)		
Psychological distress	2.66	2.72	-0.06	0.584	3.16	3.05	0.11	0.301
(range: 0-24)	(3.71)	(3.77)			(4.04)	(3.82)		
(2) Changes over ten waves								
Self-rated health	0.21	0.15	0.06	0.042	0.16	0.14	0.02	0.404
	(1.05)	(1.00)			(1.00)	(0.94)		
Functional limitations	0.27	0.24	0.03	0.505	0.47	0.26	0.21	< 0.001
	(1.94)	(1.66)			(2.19)	(1.72)		
Psychological distress	0.58	0.05	0.53	< 0.001	0.60	0.37	0.23	0.029
	(4.22)	(3.85)			(4.21)	(3.95)		
Potential mediators over ten waves								
Low household spending[b]	0.45	0.34	0.11	< 0.001	0.49	0.37	0.12	< 0.001
	(0.50)	(0.47)			(0.50)	(0.48)		
Low social activity[b]	0.53	0.30	0.24	< 0.001	0.53	0.28	0.24	< 0.001
	(0.50)	(0.46)			(0.50)	(0.45)		
Low physical activity	0.98	0.95	0.04	< 0.001	0.97	0.94	0.03	< 0.001
	(0.14)	(0.23)			(0.18)	(0.24)		
Smoking	0.61	0.51	0.10	< 0.001	0.29	0.19	0.10	< 0.001
	(0.49)	(0.50)			(0.46)	(0.40)		
Problem drinking	0.19	0.17	0.02	0.022	0.12	0.10	0.02	0.063
	(0.39)	(0.37)			(0.32)	(0.30)		
No regular health check-up	0.80	0.66	0.14	< 0.001	0.82	0.75	0.06	< 0.001
	(0.40)	(0.48)			(0.39)	(0.43)		
Number of individuals	1,452	7,868			1,631	9,073		

[a] Figures in the parentheses are standard deviations. [b] Lowest tertile.

*** $p < 0.001$, ** $p < 0.01$

Table 2. Estimated associations with changes in normalised health variables between baseline and each wave[a]

	Men (n = 9,320 × 10 waves)						Women (n = 10,704 × 10 waves)					
	Model A			Model B			Model A			Model B		
	Coef.[b]		SE[c]	Coef.		SE	Coef.		SE	Coef.		SE
(1) Self-rated health[d]												
Low-educated	0.116	***	(0.012)	0.081	***	(0.015)	0.105	***	(0.011)	0.092	***	(0.014)
Low-educated × Wave				0.006	***	(0.002)				0.002		(0.002)
Wave	0.017	***	(0.001)	0.016	***	(0.001)	0.015	***	(0.001)	0.014	***	(0.001)
Self-rated health at baseline	-0.467	***	(0.004)	-0.467	***	(0.004)	-0.474	***	(0.004)	-0.474	***	(0.004)
(2) Functional limitations[d]												
Low-educated	0.039	***	(0.011)	0.020		(0.015)	0.089	***	(0.012)	0.029		(0.016)
Low-educated × Wave				0.004	*	(0.002)				0.011	***	(0.002)
Wave	0.015	***	(0.001)	0.015	***	(0.001)	0.019	***	(0.001)	0.017	***	(0.001)
Functional limitations at baseline	-0.538	***	(0.005)	-0.538	***	(0.005)	-0.440	***	(0.005)	-0.440	***	(0.005)
(3) Psychological distress[d]												
Low-educated	0.092	***	(0.012)	0.018		(0.014)	0.052	***	(0.012)	0.021		(0.015)
Low-educated × Wave				0.014	***	(0.002)				0.006	***	(0.002)
Wave	-0.001		(0.001)	-0.003	***	(0.001)	0.006	***	(0.001)	0.005	***	(0.001)
Psychological distress at baseline	-0.431	***	(0.005)	-0.431	***	(0.005)	-0.380	***	(0.005)	-0.380	***	(0.005)

[a] Controlled for ages at baseline. [b] Coefficient. [c] Standard error. [d] Normalised by mean and standard deviation.

*** $p < 0.001$.

Table 3. Estimated associations with changes in normalised health variables over the 9-year period[a]

	Men (n = 9,320)						Women (n = 10,704)					
	Model 1			Model 3			Model 1			Model 3		
	Coef.[b]		SE[c]	Coef.		SE	Coef.		SE	Coef.		SE
(1) Self-rated health[d]												
Low-educated	0.16	***	(0.02)	0.08	***	(0.02)	0.17	***	(0.02)	0.11	***	(0.02)
Low household spending				0.02		(0.02)				-0.05	***	(0.01)
Low social participation				0.23	***	(0.02)				0.21	***	(0.02)
Low physical activity				0.23	***	(0.04)				0.21	***	(0.03)
Smoking				0.05	**	(0.02)				0.03		(0.02)
Problem drinking				0.05	**	(0.02)				-0.04		(0.02)
No regular health check-up				0.07	***	(0.02)				0.04	**	(0.02)
Self-rated health at baseline	-0.55	***	(0.01)	-0.57	***	(0.01)	-0.58	***	(0.01)	-0.59	***	(0.01)
(2) Functional limitations[d]												
Low-educated	0.05	*	(0.02)	-0.01		(0.03)	0.16	***	(0.03)	0.10	***	(0.03)
Low household spending				0.05	*	(0.02)				-0.03		(0.02)
Low social participation				0.20	***	(0.02)				0.19	***	(0.02)
Low physical activity				0.13	**	(0.05)				0.14	***	(0.04)
Smoking				0.01		(0.02)				0.13	***	(0.02)
Problem drinking				0.08	**	(0.03)				-0.05		(0.03)
No regular health check-up				0.02		(0.02)				0.11	***	(0.02)
Functional limitations at baseline	-0.58	***	(0.01)	-0.58	***	(0.01)	-0.50	***	(0.01)	-0.51	***	(0.01)
(3) Psychological distress[d]												
Low-educated	0.14	***	(0.02)	0.07	***	(0.02)	0.09	***	(0.02)	0.04		(0.02)
Low household spending				0.03	*	(0.02)				-0.01		(0.01)
Low social participation				0.19	***	(0.02)				0.15	***	(0.02)
Low physical activity				0.08	*	(0.03)				0.06	*	(0.03)
Smoking				0.00		(0.01)				0.04	*	(0.02)
Problem drinking				0.05	*	(0.02)				0.00		(0.02)
No regular health check-up				0.06	***	(0.02)				0.07	***	(0.02)
Psychological distress at baseline	-0.52	***	(0.01)	-0.54	***	(0.01)	-0.47	***	(0.01)	-0.48	***	(0.01)

[a] Controlled for ages at baseline. [b] Coefficient. [c] Standard error. [d] Normalised by mean and standard deviation

*** $p < 0.001$, ** $p < 0.01$.

Table 4. Estimated proportions of the associations between education and health mediated by each mediator[a] (N = 20,024)

	Men (n = 9,320)		Women (n = 10,704)	
	Proportion (%)	95% CI[b]	Proportion (%)	95% CI
Self-rated health				
Low household spending	1.0	(-1.6, 3.5)	-3.6	(-6.1, -1.2)
Low social participation	31.1 †[c]	(24.4, 37.7)	30.7 †	(24.3, 37.2)
Low physical activity	15.0 †	(6.8, 23.2)	9.3 †	(4.2, 14.4)
Smoking	3.1 †	(0.5, 5.7)	2.0	(-1.2, 5.1)
Problem drinking	1.2	(-0.3, 2.6)	-0.8	(-2.2, 0.6)
No regular health check-up	6.0 †	(2.2, 9.7)	2.0 †	(0.1, 3.9)
Total	57.3 †	(28.9, 85.6)	39.5 †	(12.3, 66.8)
Total for significant mediators[d]	55.2 †	(44.2, 66.1)	42.0 †	(33.8, 50.1)
Functional limitations				
Low household spending	0.3	(-3.2, 3.8)	-2.3	(-5.7, 1.2)
Low social participation	41.8 †	(32.3, 51.2)	27.4 †	(18.7, 36.2)
Low physical activity	12.1 †	(7.4, 16.8)	7.3 †	(3.3, 11.3)
Smoking	3.9 †	(0.3, 7.6)	9.8 †	(-4.5, 15.1)
Problem drinking	0.8	(-1.0, 2.6)	-0.9	(-2.6, 0.9)
No regular health check-up	6.5 †	(3.0, 10.0)	4.9 †	(2.0, 7.9)
Total	65.4 †	(21.3, 109.5)	46.4 †	(4.3, 88.4)
Total for significant mediators[d]	64.3 †	(53.6, 74.9)	49.5 †	(38.8, 60.3)
Psychological distress				
Low household spending	2.4	(-0.8, 5.5)	-1.6	(-6.3, 3.1)
Low social participation	32.2 †	(24.1, 40.3)	46.2 †	(33.2, 59.1)
Low physical activity	7.7 †	(1.9, 13.5)	6.5 †	(0.2, 12.8)
Smoking	0.6	(-2.4, 3.6)	4.7	(-1.9, 11.4)
Problem drinking	1.3	(-0.5, 3.1)	0.2	(-2.1, 2.5)
No regular health check-up	7.4 †	(2.8, 11.9)	6.1 †	(2.0, 10.3)
Total	51.6 †	(16.0, 87.3)	62.1 †	(7.9, 116.4)
Total for significant mediators[d]	47.3 †	(36.9, 57.6)	58.8 †	(44.9, 72.7)

[a] Controlled for baseline value of each health variable as well as ages at baseline.

[b] 95% confidence intervals obtained via bootstrap estimation with 2,000 replications.

[c] Indicates that bootstrap-estimated 95% CI was above zero.

[d] Mediators with †

（資料）

Takashi Oshio and Mari Kan, "The dynamic impact of retirement on health: evidence from a nationwide ten-year panel survey in Japan," ***Preventive Medicine*****, 2017,100, 287-293.**

ABSTRACT

Retirement is a major life-course transition that is closely related to changes in health. This study examined the dynamic impact of retirement on health and health behaviors, distinguishing an immediate change in the level of health at retirement and a change in the rate of change after retirement. We used panel data from 9,283 individuals (4,441 men and 4,842 women) who had retired during a nationwide ten-year panel survey in Japan conducted in 2005–2014. We focused on three health behaviors (current smoking, heavy alcohol drinking, and leisure-time physical activity) and two health indicators (self-rated health and psychological distress). We estimated regression models that controlled for both time-invariant individual attributes and the endogeneity of retirement, using panel data collected during the five years before and after retirement. Results generally confirmed that the transition was accompanied by favorable changes in health and health behaviors with some gender differences. Among men, retirement immediately promoted leisure-time physical activity and reduced poor self-rated health and psychological distress. Retirement also accelerated smoking cessation and leisure-time physical activity and decelerated reporting poor health. Among women, retirement immediately promoted leisure-time physical activity and reduced psychological distress, while it did not affect the rate of change in any health variable after retirement. The current study underscores the need for more in-depth knowledge of the dynamic impact of retirement on health. This will assist in developing policy measures to help the middle-aged population make healthy transitions from work to retirement.

INTRODUCTION

Retirement is a major transition in later life that is closely related to changes in health. The impact of retirement on health is potentially a key determinant of quality of life among middle-aged and elderly individuals (van der Heide et al., 2013; Zantinge et al., 2014). Additionally, the association between retirement and health is a central issue for public policy in developed countries, because retirement is closely related to public pension schemes (Gruber and Wise, 1999) and health and long-term care for the elderly are expected to continue to increase public spending (de la Maisonneuve and Oliveira Martins, 2013).

It is reasonable to predict that retirement would have a favorable impact on health, considering the stressful influence of work. Indeed, many studies have attempted to confirm this, focusing on various types of health behaviors such as smoking (Celidoni and Rebba, 2016; Ding et al., 2016; Lang et al., 2007), alcohol consumption (Brennan et al., 2010; Celidoni and Rebba, 2016; Ding et al., 2016; Zins et al., 2011), and physical activity (Chung et al., 2009; Ding et al., 2016; Feng et al., 2016; Slingerland et al., 2007; Stenholm et al., 2016). Studies have also considered overall health variables measured by self-rated health and mental health indicators (Behncke, 2012; Coe and Zamarro, 2011; Hesel, 2016; Neuman, 2008; Westerlund et al., 2009; Westerlund et al., 2010; Zhu, 2016). As surveyed by van der Heide et al. (2013) and Zantinge et al. (2014), many studies have confirmed that retirement has a beneficial effect on health, while several other studies have obtained opposing or inconsistent results. Indeed, there are many reasons to assume the negative effects of retirement on health, through life-course disruptions, loss of key social role, income loss, and others.

There are at least three factors that may result in mixed and inconsistent observations about the positive effects of retirement, besides differences inherent to datasets collected from different countries and study groups. First, results may be biased as studies have not fully considered individual differences such as personality traits and inherent characteristics. Prospective cohort studies have usually compared health variables between participants who had retired during baseline and follow-up and those who continued to work throughout the study (e.g., Feng et al., 2016; Lang et al., 2007; Slingerland et al., 2007). These studies did control for sociodemographic and socioeconomic attributes observed through surveys, but they could not control for unobserved individual attributes, making it difficult to identify the causal effect of retirement on health. Fixed-effects (FE) regression models have often been used to control for time-invariant individual attributes, both observed and unobserved (Celidoni and Rebba, 2016; Chung et al. 2009; Zhu, 2016).

Second, retirement must be endogenous in general; it may be a choice made by an individual, at least to some extent. To alleviate the endogeneity biases, an increasing number of studies have been utilizing the instrumental variable (IV) method (Behncke, 2012; Coe and Zamarro, 2011; Hessel, 2016; Zhu, 2016). In the first stage, this method estimates retirement through an IV expected to affect retirement but not health directly. In the second stage, the model explains health by the retirement predicted in the first stage. Many studies have used eligibility for public pension benefits as an IV (Coe and Zamarro, 2011; Hessel, 2016; Neuman, 2008; Zhu, 2016), because it is institutionally fixed and expected to affect an individual's decision to retire but not his/her health directly. In recent years, FE-IV models, which are a combination of an FE model and an IV method, have often been used to address biases due to both individual time-invariant attributes and the endogeneity of retirement (Bonsang et al., 2012; Godard, 2016; Zhu, 2016).

Third, retirement is likely to affect health in two different ways: (i) an immediate change in the level at retirement and (ii) a change in the rate of change after retirement. For example, it might be that even if health keeps deteriorating after retirement, retirement reduces its rate of deterioration. A simple comparison between pre- and post-retirement levels of the health outcome may fail to capture this type of beneficial impact of retirement on health, even if the endogeneity of retirement is successfully controlled for. Indeed, studies have found that the health effect of retirement tends to change over time (Stenholm et al., 2016; Zhu, 2016), suggesting the need for examining the dynamic effect of retirement on health.

In the current study, we examined how retirement affects the dynamics of health and health behaviors, explicitly considering the above-mentioned issues—that is, (i) controlling for individual heterogeneity, (ii) alleviating endogeneity biases of retirement, and (iii) distinguishing two types of health effects of retirement. We estimated FE-IV models to examine both types of health effects of retirement separately for three health behaviors (current smoking, heavy alcohol drinking, and leisure-time physical activity) and two health indicators (self-rated health and psychological distress). We also considered gender differences in health effects of retirement, assuming that socio-institutional backgrounds of retirement and their implications for health may differ between men and women.

The present study is also expected to shed new light on the understanding of the impact of retirement on health; it used a nationwide dataset in Japan, contrary to previous studies, most of which have used data from Europe, the U.S., and other Western countries. Japan is characterized not only by a high level of labor force participation and long life-expectancy among the elderly but also by a gradual and less straightforward transition from work to retirement (Shimizutani and Oshio, 2010). In addition, a lower share of full-time employees among middle-aged women is expected to lead to more limited impact of retirement on women's health in Japan.

METHODS

Study sample

We used data obtained from a nationwide, ten-wave panel survey, "The Longitudinal Survey of Middle-Aged and Older Adults," which was conducted by the Japanese Ministry of Health, Labour and Welfare (MHLW) each year between 2005 and 2014. Japan's Statistics Law required the survey to be reviewed from statistical, legal, ethical, and other viewpoints. We obtained the survey data from the MHLW with its official permission, so the current study did not require ethical approval.

Samples in the first wave were limited to those aged 50–59 years and were collected nationwide in November of 2005 through a two-stage random sampling procedure. A total of 34,240 individuals responded (response rate: 83.8%). The second to tenth waves of the survey were conducted in early November of each year from 2006 to 2014, and 22,748 individuals remained in the tenth wave (average attrition rate of 4.0% in each wave). No new respondents were added after the first wave.

To capture the impact of retirement as precisely as possible, we focused exclusively on the observations of the respondents who had been working continuously since the first wave and retired during the second and tenth waves (assuming that they had been working until the first wave). We excluded the data of participants when and after they resumed working after the first retirement. We also considered the observations at most five years before and after retirement; for example, we concentrated on the observations between waves 1 and 9 for the respondents who retired in wave 4 and on the observations between waves 3 and 10 for the respondents who retired in wave 8. This is because too long a period from retirement may make it difficult to distinguish the effects of retirement from other factors. Excluding further respondents who were missing key variables, we used the data of 9,283 individuals (4,441 men and 4,842 women). The total number of observations was 54,113 (25,833 for men and 28,280 for women).

Measures

Health behaviors

We considered three health behaviors: current smoking, heavy alcohol drinking, and leisure-time physical activity, each of which was expressed as a binary variable. We considered a participant who answered "yes" to the question "do you smoke currently?" to be a current smoker. We defined heavy problem drinking as an intake of more than three *go* (540 ml) of Japanese sake or an equivalent amount of alcohol every day, which corresponds to about 60 g of pure alcohol. This threshold was based on a study that showed that maintaining alcohol consumption below 46 g/day appeared to minimize the risks of mortality in a Japanese population (Inoue et al., 2012). We considered respondents to have engaged in leisure-time physical activity if they reported that they were doing moderate-intensity or vigorous aerobic activity at least two days per week. This threshold was roughly consistent with the guideline proposed by the MHLW (2013).

Health

We considered two health indicators—poor self-rated health and psychological distress, each of which was expressed as a binary variable. Regarding self-rated health, the respondents were asked to indicate their current health condition on a 6-point scale: 1 (*very good*), 2 (*good*), 3 (s*omewhat good*), 4 (*somewhat poor*), 5 (*poor*), and 6 (*very poor*). A binary variable for poor self-rated health was constructed by assigning the value 1 to those who indicated 4, 5, or 6 on the scale, and zero to those who indicated 1, 2, or 3 on the scale.

We measured psychological distress using the Kessler Psychological Distress Scale (K6; Kessler et al., 2002; Kessler et al., 2010). The respondents were asked to answer a six-item questionnaire that included items such as, "During the past 30 days, about how often did you feel a) nervous, b) hopeless, c) restless or fidgety, d) so depressed that nothing could cheer you up, e) that everything was an effort, and f) worthless?" The questions were rated on a 5-point scale (0 = *none of the time* to 4 = *all of the time*). Then, the sum of the reported scores (range: 0–24) was calculated and defined as the K6 score. Higher K6 scores reflect higher levels of psychological distress. K6 scores ≥ 5 indicate mood/anxiety disorder in a Japanese sample, as validated by preceding studies (Furukawa et al., 2008; Sakurai et al., 2011). A binary variable for psychological distress was constructed by assigning the value 1 to those with K6 scores ≥ 5 and the value zero to those with K6 scores below 5.

Covariates

As covariates, we constructed three binary variables to indicate whether the respondent was living alone, had a spouse, and was providing informal care to any family member. It should be noted that these covariates are potentially endogenous and affected by both retirement and health; however, we confirmed that estimation results remained virtually intact even if omitting them in regressions. In addition, we used the indicator variables for each wave to control for wave-specific factors.

Analytic strategy

Following some descriptive analyses, we estimated regression models to explain each health variable separately. The benchmark model is given by the following:

$$Health_{it} = \alpha Retired_{it} + \theta(Age_{it} - Retirement\ age_i) + \theta\beta Retired_{it} \times (Age_{it} - Retirement\ age_i) + \gamma \boldsymbol{X}_{it} + \varepsilon_i + \xi_{it}, \qquad (1)$$

where *Health* indicates a binary variable of health, and *Age* and *Retirement age* indicate current age and retirement age, respectively. The subscripts *i* and *t* correspond to individual and wave, respectively. *Retired* is a binary variable, which is equal to one if age is equal to or higher than retirement age and zero otherwise. The value of ($Age_{it} - Retirement\ age_i$) is in the range between –5 and 5 and is negative before retirement, equal to zero at retirement, and positive after retirement. $\boldsymbol{X}$ is a set of time-variant covariates, ε_i is a time-invariant individual factor, and ξ_{it} is an error.

As illustrated in Figure 1, an immediate change in the level of health at retirement is indicated by α. The rate of change in health changes from θ before retirement to $(1+\beta)\theta$ after retirement. β indicates the proportion of a change in the rate of change in health after retirement with its positive and negative values corresponding to acceleration and deceleration, respectively. The value of β is implicitly computed by dividing the estimated value of $\beta\theta$ by that of θ.

In the actual regression analyses, we estimated

$$Health_{it} = (\alpha_1 + \alpha_2 Female_i) \times Retired_{it} + (\theta_1 + \theta_2 Female_i) \times (Age_{it} - Retirement\ age_i) + (\theta_1\beta_1 + \theta_2\beta_2 Female_i) \times Retired_{it} \times (Age_{it} - Retirement\ age_i) + \gamma \boldsymbol{X}_{it} + \varepsilon_i + \xi_{it}, \qquad (2)$$

for the entire sample to incorporate potential gender differences, instead of estimating eq. (1) separately for men and women. Eq. (2) includes three interaction terms with a binary variable, $Female_i$, which indicate female participants. An immediate impact on the level of health at retirement (denoted by α in eq. (1)) is given by α_1 for men and $\alpha_1 + \alpha_2$ for women, with the gender difference to be tested by the significance of estimated value of α_2. The proportion of change in the rate of change in health after retirement (denoted by β in eq. (1)) is calculated by dividing the estimated value of $\theta_1\beta_1$ by that of θ_1 for men and by dividing the estimated value of $(\theta_1\beta_1 + \theta_2\beta_2)$ by that of $(\theta_1 + \theta_2)$ for women. The gender difference can be tested by the significance of the difference between these two estimated proportions.

We first estimated eq. (2) as an FE model, in which all variables are mean-centered and, hence, a time-invariant individual factor (ε) is automatically removed from regression. To make the estimation results easily understood, we treated the regression model as a linear probability model (*Wooldridge, 2013*) rather than a logistic/probit model. Further, considering the potential endogeneity of retirement, we estimated two additional first-stage, linear FE models: (i) to explain *Retired* by *Eligible*, that is, a binary variable allocated as 1 if age is equal to or higher than the eligibility age for public pension benefits, and (ii) to explain (*Age − Retirement age*) by (*Age − Eligibility age*), which is the difference between the current age and the eligibility age for public pension benefits, along with the same covariates used in eq. (2). In the second stage, we estimated the FE model (2) by replacing *Retired* and (*Age – Retirement age*) with their predicted values obtained from the first-stage estimations.

For the eligibility ages of public pension benefits, we used those for the wage-proportional benefits of the Employees' Pension Insurance (EPI) program, which covers private-sector employees. This was relevant for public-sector employees as well, because they have a similar pension program to the EPI. EPI benefits consist of flat-rate and wage-proportional components. The eligibility age for the flat-rate benefit was raised gradually from age 60 in 2001 for men and 2006 for women. The eligibility age for the wage-proportional benefit was raised gradually from age 60 in 2013 for men but remained fixed at 60 until 2018 for women. We focused on the eligibility age for the wage-proportional benefit as the EPI insured participants were generally not eligible for any benefit before that age. It should be noted that the variation of the eligibility age was limited; the proportions of eligibility age 60 (for those born before April 2, 1953), 61 (for those born between April 2, 1953 and April 1, 1955), and 62 (between April 2, 1955 and April 1, 1957) were 88.0%, 10.1%, and 1.9%, respectively, among male participants, and the eligibility age was 60 for all female participants. However, both IVs (*Eligible* and *Age − Eligibility age*), had sufficiently large variation in the observations to make the first-stage estimations effective.

RESULTS

Figure 2 depicts the observed distribution of retirement age for men and women, confirming the spikes of retirement age at 60 for both genders; 21.5% and 16.6% of men and women, respectively, retired at age 60. This result is in line with the fact that most participants in this survey became eligible for public pension benefits at age 60.

Table 1 compares occupational status and hours worked per week between men and women one year before retirement, along with educational attainment. Compared to women, a larger proportion of men had been regular employees and executives and had been working for a longer time. Nearly half of female participants had been working as part-time or temporary workers. Table 2 shows how the level of each health variable changes from two years before to two years after retirement. Among both men and women, the prevalence of current smoking and heavy drinking decreases after retirement while that of leisure time activity increases. Self-rated health worsens after retirement while there is no significant change in psychological distress.

However, comparisons between only two time points cannot grasp the dynamics of health around retirement. Figures 3 and 4 compare evolutions of health and health behaviors around retirement among men and women, respectively. Remarkable jumps at retirement are observed for leisure-time physical activity among both men and women. By contrast, smoking secession accelerates after retirement especially among men. A trend in psychological distress turns from upward to downward at retirement, albeit not substantially, among both men and women.

Estimation results of FE models are summarized in Table 3. The key focuses are on (i) the estimated coefficient on *Retired* (α), i.e., the immediate impact of retirement, and (ii) the estimated proportion of the impact on the rate of change after retirement (β). The estimated values of α suggest that retirement immediately discouraged both men and women from smoking and prompted them to engage in leisure-time physical activity. Meanwhile, the estimated values of β suggest that retirement reduced a rising pace of reporting poor self-rated health and psychological distress among both men and women while it accelerated smoking cessation only among men. The gender difference was not significant in α or β for any health variable.

To examine how these estimation results are affected by controlling for the endogeneity of retirement, Table 4 summarizes the FE-IV results (with first-stage regression results available upon request). Retirement immediately encouraged both men and women to engage in leisure-time physical activity and reduce their probability of psychological distress. Meanwhile, retirement immediately reduced the probability of poor self-rated health only among men. Significant changes in the rate of change in health variables after retirement were observed only among men; retirement accelerated smoking secession and leisure-time physical activity and decelerated self-reporting poor health. A significant gender difference was observed in two cases; the immediate impact on leisure-time physical activity was higher and the post-retirement rate of reporting poor health declined more remarkably among men.

DISCUSSION

We investigated the dynamics of health around retirement and generally confirmed that the transition is accompanied by favorable changes in health and health behaviors. However, results were not fully consistent across health variables. The most remarkable and consistent impact was observed on leisure-time physical activity, in line with several preceding studies. Current smoking was another health behavior affected by retirement especially among men. By contrast, alcohol consumption was not related to retirement, adding to generally mixed results in preceding studies. Retirement had a generally positive impact on self-rated health and psychological distress, confirming general results in preceding studies.

Results also uncovered gender differences in the health effect of retirement. The effect of retirement on health was more limited for women than for men, although the differences were not statistically significant in most cases. We can speculate that our findings were related to the gender differences in occupational status before retirement. As shown in Table 1, female participants worked less than male participants before retirement, with a higher proportion of part-time and temporary workers and shorter hours worked, which may have resulted in a more limited impact of retirement on health for women.

Finally, our findings highlighted the importance of two methodological issues. First, controlling for endogeneity of retirement tended to affect substantially the estimation results, as already suggested by previous studies which utilized FE-IV methods. Second, an immediate change in the level of health at retirement and a change in its rate of change after retirement should be distinguished. These two types of impact differed across health variables as well as between genders, making simple comparisons between before and after retirement sometimes misleading.

Study limitations and strength

We recognize that the current study has several limitations. As suggested by Chung et al. (2009), job status before retirement is expected to confound the effect of retirement on health even among those of the same gender, an issue disregarded in the present study. More broadly, the relevance of retirement for health is likely affected by socio-institutional background. Notably, a gradual transition to retirement and a limited proportion of full-time employees among middle-aged women require us to be cautious in generalizing the results in this study to other countries.

Meanwhile, our analysis had two important features. First, it controlled for the endogeneity of retirement as well as time-invariant individual attributes. Second, it distinguished an immediate change in the level of health at retirement and a change in its rate of change after retirement. These two methodologies allowed us to provide new insights into the understanding of the dynamics of health around retirement.

Conclusions

The current study underscores the need for more in-depth knowledge of the dynamic impact of retirement on health. This will assist in developing policy measures to help the middle-aged population make healthy transitions from work to retirement.

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Figure 1. Dynamics of health around retirement: an illustrative example

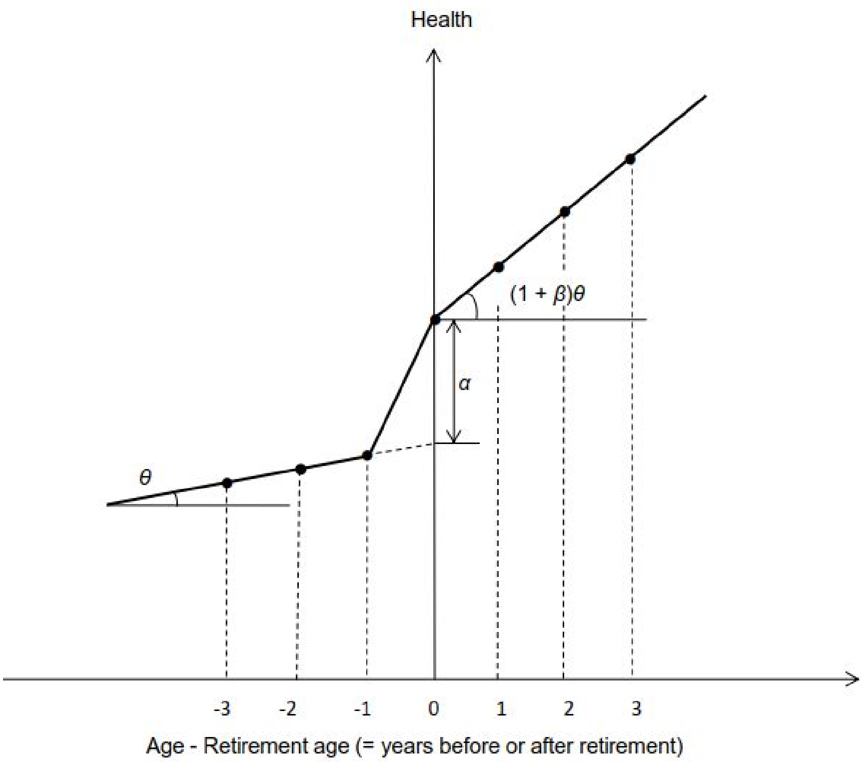

Figure 2. Distribution of retirement age

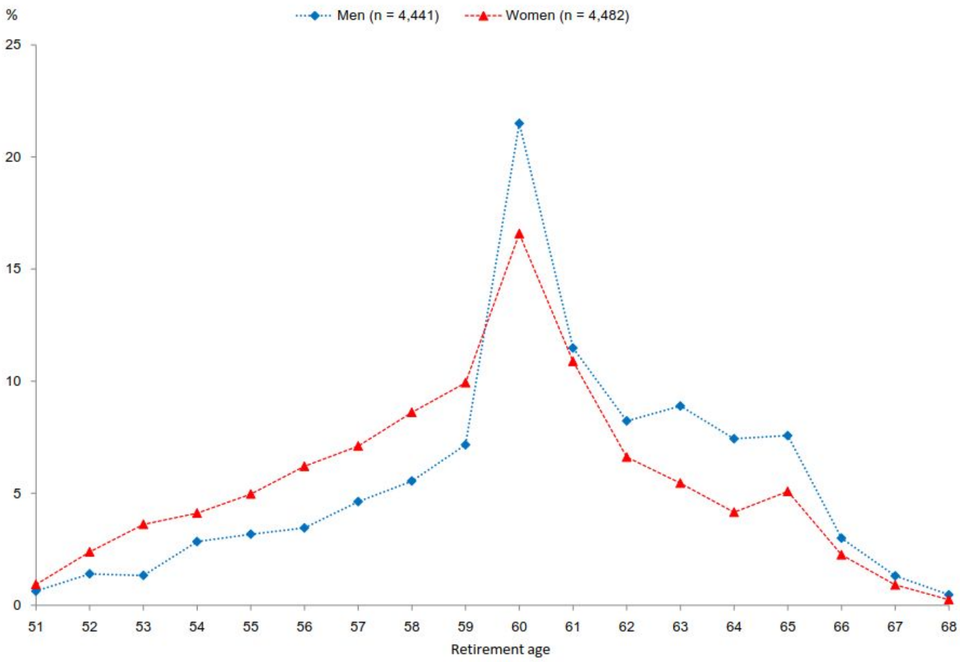

Figure 3. Evolution of health and health behavior among men around retirement (n = 25,833 of 4,441 individuals)

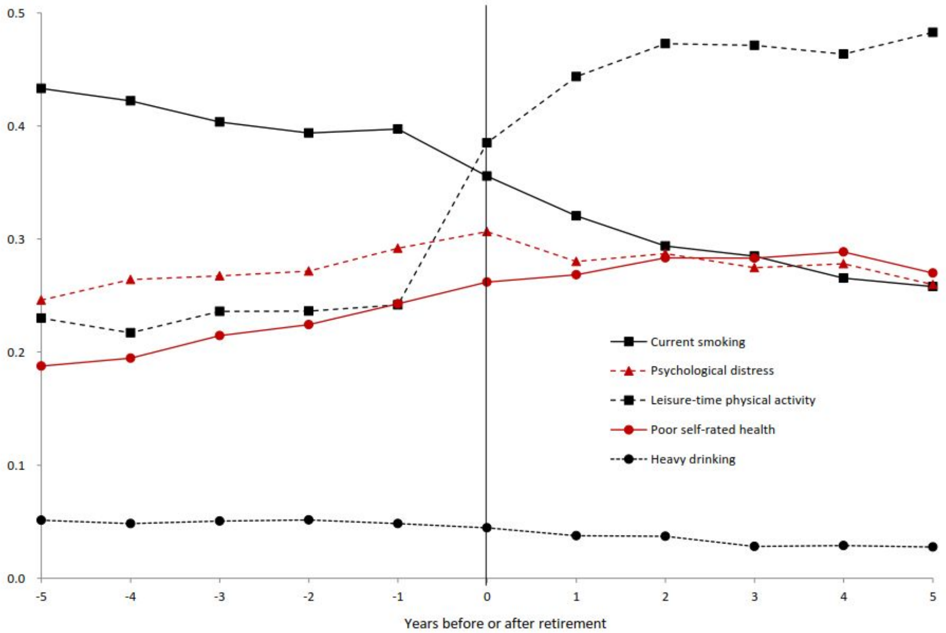

Figure 4. Evolution of health and health behavior among women around retirement
(n = 28, 280 of 4,842 individuals)

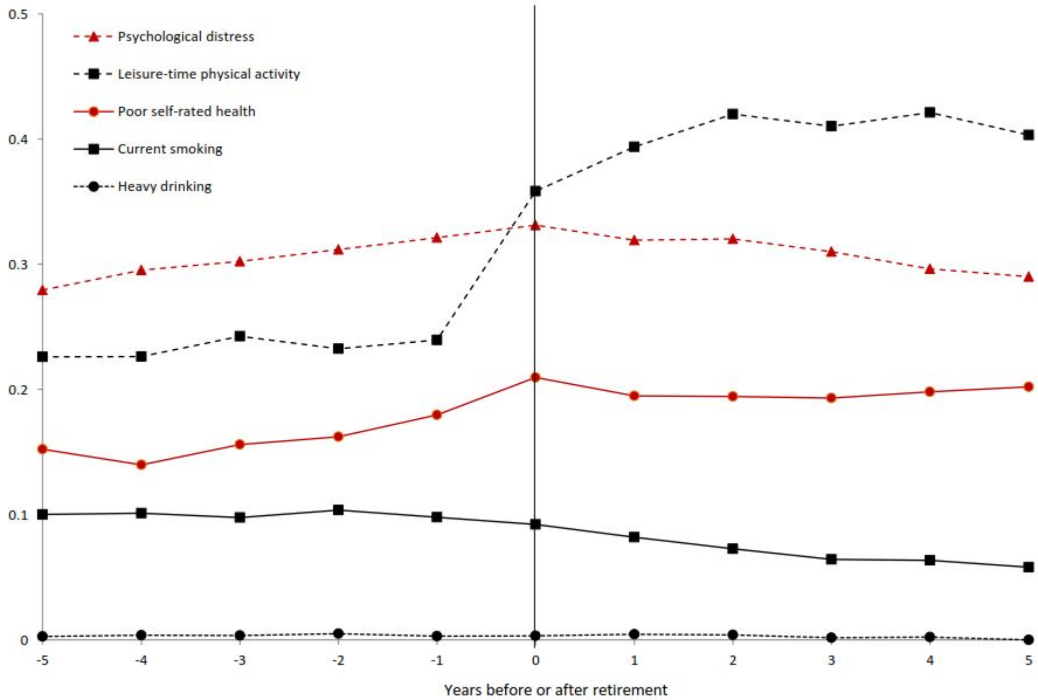

Table 1
Job status one year before retirement.

	Men	Women
Occupational status		(%)
Regular employee	50.2	24.3
Executive	5.8	1.7
Part-time or temporary worker	9.0	46.5
Dispatched employee	0.9	0.8
Contract worker	20.9	7.2
Self-employed	9.1	4.1
Family worker	0.6	8.4
Side job at home	0.1	3.5
Other	3.4	3.4
Total	100.0	100.0
Hours worked per week		
M	39.4	29.6
SD	(13.2)	(14.3)
Cf. Education level		(%)
Junior high school	17.7	17.2
High school	48.8	53.7
Junior college	6.8	22.1
College or above	25.9	6.6
Other	0.8	0.4
Total	100.0	100.0
N	4441	4842

Table 2
Changes in health and health behavior (prevalence) from two years before retirement to two years after retirement.

	Men ($n = 5094$)				Women ($n = 5599$)			
	Before	After	Change		Before	After	Change	
Current smoking	0.394	0.294	−0.100	***	0.104	0.073	−0.031	***
Heavy drinking	0.052	0.037	−0.015	*	0.005	0.004	−0.001	
Leisure-time physical activity	0.236	0.473	0.236	***	0.232	0.420	0.187	***
Poor self-rated health	0.224	0.283	0.059	***	0.162	0.194	0.032	**
Psychological distress	0.272	0.287	0.015		0.312	0.320	0.009	

* $p < 0.05$.
** $p < 0.01$.
*** $p < 0.001$.

Table 3
Estimated associations between retirement and health obtained from FE models (54,113 observations of 9283 individuals)[a].

		Retired[b]			(Age − Retirement age)[c]			Retired × (Age − Retirement age)			β		
		Coef. (α)		(SE)	Coef. (θ)		(SE)	Coef. ($\beta\theta$)		(SE)			(SE)
Current smoking	Men	−0.014	***	(0.003)	−0.021	***	(0.001)	−0.004	**	(0.002)	0.200	*	(0.081)
	Women	−0.013	***	(0.003)	−0.002		(0.001)	−0.002		(0.001)	0.983		(1.079)
	Difference	−0.001		(0.001)	−0.019	***	(0.001)	−0.002		(0.002)	−0.784		(1.075)
Heavy alcohol drinking	Men	−0.001		(0.002)	−0.003	***	(0.001)	0.000		(0.001)	0.161		(0.414)
	Women	−0.001		(0.002)	0.000		(0.001)	−0.001		(0.001)	4.681		(29.80)
	Difference	0.000		(0.001)	−0.003	**	(0.001)	0.000		(0.001)	−4.520		(29.76)
Leisure-time physical activity	Men	0.138	***	(0.007)	0.014	***	(0.002)	0.005		(0.003)	0.365		(0.278)
	Women	0.136	***	(0.006)	0.003		(0.003)	0.004		(0.003)	1.479		(2.167)
	Difference	0.002		(0.002)	0.011	***	(0.003)	0.001		(0.004)	−1.114		(2.155)
Poor self-rated health	Men	0.006		(0.006)	0.012	***	(0.002)	−0.012	***	(0.003)	−0.990	***	(0.187)
	Women	0.008		(0.005)	0.008	***	(0.002)	−0.009	***	(0.003)	−1.122	***	(0.304)
	Difference	−0.002		(0.002)	0.004		(0.003)	−0.003		(0.004)	0.132		(0.354)
Psychological distress	Men	0.003		(0.007)	0.008	***	(0.002)	−0.009	**	(0.003)	−1.219	***	(0.338)
	Women	0.002		(0.006)	0.012	***	(0.002)	−0.010	***	(0.003)	−0.837	***	(0.217)
	Difference	0.001		(0.002)	−0.004		(0.003)	0.001		(0.004)	−0.382		(0.403)

[a] Controlled for living alone, having a spouse, providing informal care to any family member, and waves.
[b] =1 if age ≥ retirement age; =0 otherwise.
[c] Range: from −5 to 5.
* $p < 0.05$.
** $p < 0.01$.

Table 4
Estimated associations between retirement and health obtained from FE-IV models (54,113 observations of 9283 individuals)[a].

		Retired[b]			(Age − Retirement age)[c]			Retired × (Age − Retirement age)			β		
		Coef. (α)		(SE)	Coef. (θ)		(SE)	Coef. ($\beta\theta$)		(SE)			(SE)
Current smoking	Men	−0.009		(0.030)	−0.023	***	(0.005)	−0.003	***	(0.001)	0.132	*	(0.055)
	Women	−0.047		(0.028)	0.005		(0.005)	−0.001		(0.001)	−0.257		(0.299)
	Difference	0.039		(0.033)	−0.027	***	(0.005)	−0.002		(0.001)	0.389		(0.298)
Heavy alcohol drinking	Men	−0.026		(0.019)	0.002		(0.003)	−0.001		(0.001)	−0.578		(1.032)
	Women	−0.019		(0.018)	0.003		(0.003)	−0.001		(0.000)	−0.231		(0.27)
	Difference	−0.007		(0.021)	−0.001		(0.004)	0.000		(0.001)	−0.347		(0.95)
Leisure-time physical activity	Men	0.475	***	(0.061)	−0.044	***	(0.011)	0.006	***	(0.002)	−0.138	**	(0.047)
	Women	0.283	***	(0.058)	−0.025	*	(0.010)	0.002		(0.001)	−0.070		(0.065)
	Difference	0.192	**	(0.068)	−0.019		(0.011)	0.004	*	(0.002)	−0.068		(0.065)
Poor self-rated health	Men	−0.136	**	(0.051)	0.034	***	(0.009)	−0.005	***	(0.002)	−0.153	**	(0.053)
	Women	−0.054		(0.049)	0.017	*	(0.008)	0.000		(0.001)	0.022		(0.071)
	Difference	−0.082		(0.057)	0.016		(0.009)	−0.006	***	(0.002)	−0.174	*	(0.079)
Psychological distress	Men	−0.204	***	(0.057)	0.039	***	(0.010)	0.002		(0.002)	0.040		(0.047)
	Women	−0.124	*	(0.055)	0.028	**	(0.009)	0.002		(0.001)	0.077		(0.056)
	Difference	−0.080		(0.063)	0.012		(0.011)	−0.001		(0.002)	−0.038		(0.064)

[a] Controlled for living alone, having a spouse, providing informal care to any family member, and waves.
[b] =1 if age ≥ retirement age; =0 otherwise.
[c] Range: from −5 to 5.
* $p < 0.05$.
** $p < 0.01$.
*** $p < 0.001$.

（資料）

Takashi Oshio and Mari Kan, "Impact of parents' need for care on middle-aged women's lifestyle and psychological distress: evidence from a nationwide longitudinal survey in Japan," ***Health and Quality of Life Outcomes*****, 2018, 16:63.**

Abstract

Background: Many studies have separately addressed the associations of informal caregiving with coresidence, a caregiver's work status, and health conditions, but not jointly. We examined how their parents' need for care affects middle-aged women's lifestyle and psychological distress, considering the potential simultaneity of decisions on caregiving and living adjustments.

Methods: We used 22,305 observations of 7,037 female participants (aged 54–67 years) from a nationwide longitudinal survey in Japan conducted during 2009 and 2013. We considered the occurrence of parents' need for care (OPNC) as an external event and estimated regression models to explain how it affected the probabilities of the participants becoming caregivers, coresiding with parents, and working outside the home. We further conducted the mediation analysis to examine how the impact of OPNC on participants' psychological distress measured by Kessler 6 (K6) scores was mediated by caregiving and living adjustments.

Results: OPNC made 30.9% and 30.3% of middle-aged women begin informal caregiving for parents and parents-in-law, respectively, whereas the impact on residential arrangement with parents or work status was non-significant or rather limited. OPNC raised middle-aged women' K6 scores (range: 0–24) by 0.368 (SE: 0.061) and 0.465 (SE: 0.073) for parents and parents-in-law, respectively, and informal caregiving mediated those impacts by 37.7% (95% CI: 15.6-68.2%) and 44.0% (95% CI: 22.2-75.4%), respectively. By contrast, the mediating effect of residential arrangement with parents or work status was non-significant.

Conclusions: Results underscore the fact that OPNC tends to promote middle-aged women to begin informal caregiving and worsen their psychological distress.

Background

Informal caregivers provide majority of the long-term care in many countries. Owing to longer life expectancy and a smaller number of siblings, we now face a higher probability of individuals having to provide informal care to old parents [1]. Hence, the occurrence of parents' need for care (OPNC) is a key driver of the change in the lifestyle of middle-aged individuals, especially women, who still tend to play a dominant role in informal care. If their parents happen to need care, adult children are probably forced to consider who will provide care to them, whether they will start coresiding with parents [2–5], whether a caregiver will stop work outside the home [6, 7], and so on.

Actually, many studies have already addressed the associations of informal caregiving with coresidence, a caregiver's work status, and health conditions, albeit not jointly. Poor health of parents tends to raise the probability of their coresidence with their adult children [2–5]. In comparison, mixed findings have been reported on the association between informal caregiving and a caregiver's work status. However, many studies have shown that the effect of informal caregiving on employment is relatively limited [6, 7]. One possible reason is the endogenous selection for assuming a caregiving role. Specifically, women, who tend to have a weaker attachment to the labor market, are more likely to take on the caregiving role [8].

A key limitation of previous studies is that they have often considered informal caregiving as an exogenous variable, thereby ignoring possible simultaneity biases. Further, most of these studies did not consider the simultaneity of decisions on informal caregiving and other behaviors, such as coresidence with parents and work outside the home, which are likely to interact with each other.

In the present study, we attempted to control for potential biases owing to endogeneity of informal caregiving and simultaneity of decisions on informal caregiving and other behaviors, in order to examine the relevance of informal caregiving to the life arrangements and well-being of middle-aged women more precisely. Therefore, we focused on OPNC, which was considered largely exogenous, and examined how the middle-aged women responded to it in terms of caregiving, residential relationship with parents, and work status, taking into account the impact of their pre-OPNC statuses as well as their interactions under the framework of a simultaneous regression model.

We further examined how the onset of caregiving and living adjustments mediated the impact of OPNC on the middle-aged women's psychological distress, based on the theoretical framework of the mediation analysis [9, 10]. It is reasonable to predict that these living adjustments, which are likely correlated with each other, will affect middle-aged women's psychological distress, especially if they become caregivers [11, 12]. Indeed, studies have evidenced that informal caregiving has a negative association with a caregiver's health and quality of life [13–14]. However, some studies suggest that conditions surrounding caregiving—such as coresidence with parents and employment status—tend to mediate the impact of informal caregiving on a caregiver's psychological distress [16–20]. In this study, we computed the mediating effects along with their statistical significance of caregiving and living adjustments.

Methods

Study sample

We used longitudinal data obtained from a nationwide, population-based longitudinal survey titled, "The Longitudinal Survey of Middle-Aged and Older Adults," conducted by the Japanese Ministry of Health, Labour and Welfare (MHLW). Samples in the first wave were collected nationwide from individuals between the ages of 50-59 years in November 2005 through a two-stage random-sampling procedure. First, 2,515 districts were randomly selected from 5,280 districts used in the MHLW's nationwide, population-based 'Comprehensive Survey of the Living Conditions of People on Health and Welfare', which was conducted in 2004. The 5,280 districts, in turn, had been randomly selected from about 940,000 national census districts. Second, depending on the population size of each district, 40,877 residents aged 50–59 years as of October 30, 2005 were randomly selected.A total of 34,240 individuals responded (response rate: 83.8%). The second to ninth waves were conducted in early November of each year from 2006 to 2013, with no additional sampling (average attrition rate in each wave: 4.3%).

We took full advantage of the longitudinal structure of the dataset to capture the timing of OPNC and how the middle-aged women's responses to it in terms of caregiving, residential relationship with parents, and work status, taking into account the impact of their pre-OPNC statuses. Specifically, we first compiled the data on female participants from the fourth to ninth waves (2008–13), because information on each individual's parents' need for care was collected only from the fourth wave. We then limited our analysis to the data of female participants who had not faced parents' need for care—and thus had not provided caregiving to parents—in the year prior to the survey year. It means that we excluded the data of participants who had already faced parents' need for care in the fourth wave (2008) and focused on the participants' data from the fifth wave (2009) onwards. This allowed us to capture the exogenous impact of OPNC in the survey year. We also excluded the data of participants whose parents died in the survey year. After further excluding participants with missing data, we used 22,305 observations of 7,303 women for the statistical analysis, in which we focused on their responses to OPNC at the survey year.

Measures

The survey asked respondents whether care was needed for each family member. We collected the data of parents' need for care and constructed a binary variable in which the emergence of parents' need for care in the survey year (after no need was reported in the previous year) was scored as "1" and other conditions were scored as "0," for the participants' parents (father and/or mother) and parents-in-law (father-in-law and/or mother-in-law). Similarly, we constructed binary variables for participants' provision of informal care to and coresidence with parents and parents-in-law, by allocating "1" if the participant was providing informal care and was residing with parents or parents-in-law, and "0" otherwise. We also constructed a binary variable for work outside the home by allocating "1" if the participant answered that she was engaged in any paid job and "0" otherwise.

Additionally, we focused on the impact of OPNC on women's psychological distress measured by K6 scores [21, 22]. The reliability and validity of K6 scores have been demonstrated in a Japanese population [22, 23]. Participants were asked to complete a six-item psychological distress questionnaire: "During the past 30 days, about how often did you feel a) nervous, b) hopeless, c) restless or fidgety, d) so depressed that nothing could cheer you up, e) that everything was an effort, or f) worthless?" Responses were rated on a 5-point scale (1 = none of the time to 5 = all of the time). K6 score (range: 0–24) was constructed by subtracting six from the sum of the responses. Higher K6 scores reflect higher levels of psychological distress. We additionally focused on the proportion of respondents with K6 score ≥ 5, which has been found to indicate mood or anxiety disorders in a Japanese population [23]. For the entire respondents in this study sample (n = 22, 307), Cronbach alpha coefficient for K6 scores was 0.90, K6 scores' mean and standard deviation were 3.4 and 4.0, respectively, and the proportion of those with K6 score ≥ 5 was 30.0%.

As for control variables, we used the respondent's age, educational attainment (junior high school, high school, college or above, other), having a spouse, and household expenditure as a proxy of household income. These factors were taken into account, because they were expected to affect the costs—both pecuniary and psychological—of informal caregiving and living arrangements, and correspondingly, their impact on psychological distress. Household expenditure was adjusted for household size by dividing the reported value of household expenditure by the square root of the number of members in the household, as was done in recent publications of the Organisation for Economic Co-operation and Development [24].

Analytic Strategy

We compared the probabilities of three variables—becoming a caregiver, coresiding, and working outside the home—between those who faced parents' need for care and those who did not, without controlling for other variables. However, as explained in the Introduction, we had to control for potential biases owing to the endogeneity of informal caregiving and simultaneity of decisions on informal caregiving, coresidence, and working outside the home as well as their statuses prior to the survey year. Therefore, we jointly estimated a set of linear regression models within the framework of the seemingly unrelated regression (SUR) model [26]:

$$\text{Caregiving}_t = \alpha_1\text{OPNC}_t + \beta_1\text{Coresidence}_{t\text{-}1} + \gamma_1\text{Work}_{t\text{-}1} + \boldsymbol{Z}_t\boldsymbol{\delta}_1 + \varepsilon_{1t}$$
$$\text{Coresidence}_t = \alpha_2\text{OPNC}_t + \beta_2\text{Coresidence}_{t\text{-}1} + \gamma_2\text{Work}_{t\text{-}1} + \boldsymbol{Z}_t\boldsymbol{\delta}_2 + \varepsilon_{2t}$$
$$\text{Work}_t = \alpha_3\text{OPNC}_t + \beta_3\text{Coresidence}_{t\text{-}1} + \gamma_3\text{Work}_{t\text{-}1} + \boldsymbol{Z}_t\boldsymbol{\delta}_3 + \varepsilon_{3t}.$$

Here, the subscript t indicates year t (t =2009, 10, 11, 12, and 13), and $\boldsymbol{Z}$ and ε_i (i = 1, 2, and 3) indicate a set of control variables and an error term. We estimated α_i and β_i, which are coefficients of OPNC and Coresidence, respectively, as well as $\boldsymbol{\delta}_i$, which is a set of coefficients of each control variable included in $\boldsymbol{Z}$. This set of regression models attempted to capture the impacts of OPNC on caregiving, coresidence, and work, assuming that these three variables were affected by coresidence and work in the previous year, and that the error terms were correlated with each other.

The focus was on the estimated value of α_i, which indicates the impact of OPNC on caregiving, coresidence, and work. Because we limited the analysis to the respondents who did not face OPNC (and thus did not engage in caregiving) in the previous year, the estimated value of α_1 indicates the probability of newly becoming a caregiver in response to OPNC. $\boldsymbol{Z}$ included age, educational attainment, having a spouse, and household expenditure, as mentioned earlier.

One may be tempted to estimate a multivariate probit model rather than a set of linear regression models within the framework of the SUR model, considering that three dependent variables are all binary ones. However, we did not use a multivariate probit model because “no OPNC” (OPNC = 0) perfectly predicted “no caregiving” (caregiving = 0) in the first caregiving model, thus omitting OPNC from regression. It has been also known that linear probability models obtain results generally similar to those of probit or logistic models and that their theoretical flaws can be disregarded in most cases [26].

We further estimated regression models to explain the extent to which OPNC affected K6 scores and how its impact was confounded by caregiving, coresidence, and work. Specifically, we first estimated the benchmark model (Model 1), which explained K6 scores by OPNC. Next, we estimated three models (Models 2–4), each of which included caregiving, coresidence, and work as an additional predictor. Then, we examined how the results were affected by adding all of these variables in Model 5. In all these models, we included a set of control variables ($\boldsymbol{Z}$) as well as K6 scores, coresidence, and work status in the previous year.

Finally, we conducted the mediation analysis [9, 10] to examine how the impact of OPNC was mediated by three potential mediators: caregiving, coresidence, and work. Based on the results of (i) the SUR model (which examined the impacts of OPNC on each of three mediators) and (ii) Model 1 (which explained K6 scores by OPNC), and (iii) Model 5 (which explained K6 scores by OPNC and three mediators), we computed the mediating effects of each of three mediators. We examined their statistical significance by bootstrap estimating their 95% confidence intervals with 3,000 replications.

Results

Descriptive analyses

Table 1 summarizes the key characteristics of 7,037 participants at baseline (in 2009). Among the participants, 11.3%, and 22.0% were residing with parents and parents-in-law, respectively. We also observed that 63.5% of the participants were working outside the home.

Table 2 compares the probabilities of caregiving, coresidence, and work between women who faced OPNC and those who did not. It was found that 30.7% and 29.7% of the participants started caregiving in response to the OPNC of parents and parents-in-laws, respectively. The difference in the probabilities of caregiving in the right column of Table 3 indicates the probability of newly becoming a caregiver in response to OPNC, because the probability of caregiving was equal to zero among those who did not face OPNC. The probabilities of coresidence and work were lower among women who faced OPNC than those who did not, but their differences (ranging between 1.8–7.2%) were much more limited as compared to those with the probabilities of caregiving. Table 2 also shows that the mean K6 score and the proportion of those with a K6 score ≥ 5 was much higher among women who faced OPNC than among those who did not, for both parents and parents-in-law.

Regression analyses

Table 3 summarizes the estimated impact of OPNC on women’s behavior. We observed that 30.9% and 30.3% of women started caregiving in response to the OPNC of parents and parents-in-law, respectively. The magnitude of this impact was almost the same as those observed in Table 2 (30.7% and 29.7%). Coresidence in the previous year raised the probability of caregiving for both parents (6.2%) and parents-in-law (8.7%), whereas work in the previous year slightly reduced it for parents. As for coresidence, OPNC slightly raised the probability of coresidence for parents (1.3%) and it had no significant impact (0.6%) for parents-in-law. Instead, residential status in the previous year was a key determinant of the current residential status. A negative impact of OPNC on work (1.1% for parents and 2.4% for parents-in-law) was rather limited and smaller than that suggested by the descriptive comparisons in Table 2. We further observed that previous work status strongly determined the current one.

Table 4 presents the estimation results of Models 1–5, which explain how OPNC affected women’s K6 score. As the benchmark model, Model 1 showed that OPNC raised women’s K6 scores by 0.368 and 0.465 for parents and parents-in-law, respectively. These impacts were equivalent to 0.09 and 0.13 standard deviation of K6 scores. Model 2 showed that the impact of OPNC was substantially mediated by becoming a caregiver for both types of parents. The inclusion of caregiving substantially attenuated the association between OPNC and K6 scores—the coefficient declined by 38.1% to 0.227 for parents and by 43.6% to 0.263 for parents-in-law (0.220)—while caregiving had a significant, positive correlation with K6 scores for both parents (0.454) and parents-in-law (0.668).

Model 3 showed that coresidence or work did not have any positive association with K6 scores, leaving the impact of OPNC virtually intact, for both parents and parents-in-law, while work reduced K6 scores in the case of caring for parents-in-law. Finally, Model 5, which included all related variables, largely mirrored the results in Models 2–4; the coefficients for OPNC and caregiving remained close to those in Model 2, while the coefficients for coresidence and work remained almost intact from Models 3 and 4, respectively.

Lastly, Table 5 presents the results of the mediation analysis, based on the results of the SUR models presented in Table 3 and those of Models 1 and 5 presented in Table 4. For caregiving to parents, OPNC raised K6 score by 0.368, and 37.7% of this impact (i.e., 0.139) was mediated by caregiving. In contrast, coresidence or work did not significantly mediate the impact of OPNC on K6 scores. We found similar results for parents-in-low; caregiving mediated 44.0% of the impact of OPNC K6 scores, while coresidence or work did not work as a mediator.

Discussion

We examined how OPNC affects the lifestyle and psychological distress of middle-aged women, using the data obtained from a nationwide longitudinal survey in Japan. Unlike most previous studies, we examined the impact of OPNC on caregiving, coresidence, and work, adjusted for their potential interactions and the effects from their previous statuses.

Results confirmed that about 30% of women began caregiving for their parents or parents-in-law in response to their need for care during the survey period (2019–2013). We also observed that the probability of becoming a caregiver was positively associated with previous coresidence with parents, a finding which was consistent with the result of a previous study conducted outside Japan [27]. Compared to the impact on the probability of becoming a caregiver, the probability of coresidence with parents was less sensitive to OPNC. In line with the results of previous studies [2, 4, 5] we obtained some evidence that OPNC prompted individuals to coreside with their parents, but the impact was rather small. Women who have been residing separately from parents seem to prefer going to their parents' house to take care of them at least at the onset of the need for caregiving. The impact on work status was also limited, which was generally in line with the results of previous studies [6, 7].

Hence, we can argue that middle-aged women tend to respond to OPNC mainly by becoming a caregiver, at least initially, without substantial adjustments to coresidence with parents and work status. One possible explanation, which seems to be relevant in Japan, where intergenerational family setting is common, is that the parent-child coresidence, along with the wife's labor force participation, may reflect the implicit contract regarding informal care and other life arrangements, which is traditionally made between adult children and their parents before OPNC [28, 29].

At the same time, results underscore the fact that OPNC is a stressful event for middle-aged women. OPNC raised psychological distress and its adverse impact was substantially mediated by becoming a caregiver. Coresiding with parents and work did not explain the variations in women's psychological distress after including OPNC as an explanatory variable. This observation was consistent with the finding that women tended to become caregivers with limited adjustments to coresidence and work.

Additionally, the present study highlights that the kin relationship tends to confound the impact of caregiving on psychological distress. Compared to parents, the adverse impacts of both OPNC and caregiving on psychological distress were higher for parents-in-law. This observation confirmed the importance of kin relationship between caregivers and care recipients for a caregiver's psychological distress, as already evidenced by previous studies [28–30].

We recognize that the present study has several limitations. First, we did not assess caregiving burden in terms of time spent on caregiving or the level of care required in the statistical analysis. This requires us to be cautious in any generalization of the obtained results. Second, we ignored the impact of prolonged caregiving on women's lifestyle and psychological distress. As caregiving continues and the nursing care levels increase, women are more likely to adjust their lifestyle and feel more distressed, especially if the conflict between informal care and other roles becomes incompatible [20]. In this sense, it is likely that the present study may underestimate the impact of OPNC on women's lifestyle and psychological distress. Following previous longitudinal studies (e.g., [31–33]), the dynamics of caregiving and its associations with lifestyle and mental health of caregivers, care recipients, and their family members must be addressed using more detailed longitudinal data. Third, we must expand the analysis to address how wider aspects of women's multiple roles including interpersonal relations with others and other social ties are affected by OPNC [16, 19].

Conclusions

Overall, the results highlighted that the onset of caregiving tends to be a serious external event that affects middle-aged women's psychological distress, even if its impact on their lifestyle is relatively limited. If long-term care for the elderly keeps relying heavily on informal caregiving at home, policy measures to support informal caregivers are required. Providing a wider range of home-visit nursing care services to in-house care recipients and expanding institutional care services could be helpful in mitigating any psychological pressure and stress caused by informal caregiving at home.

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Table 1 Key characteristics of 7037 participants at baseline (in 2009)

Educational attainment		Proportion (%)
Junior high school		12.7
High school		64.1
Junior college		14.6
College or above		8.3
Having a spouse		88.0
Residing with parents		11.3
Residing with parents-in-law		22.0
Working outside home		63.5
Age (years)	*M*	58.0
	SD	(2.7)
Monthly household expenditure (thousand yen)[a]	*M*	230.0
	SD	(283.0)

[a]Adjusted for household size

Table 2 Behavioral probabilities and K6 scores of women with and without parental need for care[a]

	The need of care		Difference (A-B)	
	Occurred (A)	Not occurred (B)		
Parents ($n = 15{,}972$)				
Caregiving	30.7%	0	30.7%	***
			SD (0.4%)	
Coresidence	13.7%	15.5%	−1.8%	*
			SD (0.8%)	
Work	55.1%	60.5%	−5.4%	***
			SD (1.0%)	
K6 score (range: 0–24)	*M* 3.79	3.20	0.59	***
	SD (0.08)	(0.03)	(0.00)	
K6 score ≥ 5	33.3%	28.4%	4.9%	***
Number of observations	2648	13,324		
Parents-in-law ($n = 10{,}887$)				
Caregiving	29.7%	0	29.7%	***
			SD (0.5%)	
Coresidence	33.1%	40.3%	−7.2%	***
			SD (1.2%)	
Work	55.7%	62.0%	−6.3%	***
			SD (1.2%)	
K6 score (range: 0–24)	*M* 3.90	3.20	0.70	***
	SD (0.10)	(0.04)	(0.10)	
K6 score ≥ 5	34.4%	29.3%	5.1%	***
Number of observations	2086	8801		

[a]All variables were evaluated in each wave when parents' need for care emerged
***$p < 0.001$, *$p < 0.05$

Table 3 Estimated impact of the occurrence of parents' need for care on women's behavior[a]

Explanatory variables	Women's behavior								
	Caregiving			Coresidence			Work		
	Coef.		SE	Coef.		SE	Coef.		SE
Parents ($n = 15{,}972$)									
Occurrence of parents' need for care	0.309	***	(0.004)	0.013	***	(0.003)	−0.011		(0.006)
Coresidence in the previous year	0.062	***	(0.004)	0.934	***	(0.003)	−0.001		(0.006)
Work in the previous year	−0.012	***	(0.003)	0.002		(0.002)	0.819	***	(0.005)
Parents-in-law ($n = 10{,}887$)									
Occurrence of the need for care	0.303	***	(0.005)	−0.006		(0.004)	−0.024	***	(0.007)
Coresidence in the previous year	0.087	***	(0.004)	0.928	***	(0.004)	0.011	*	(0.006)
Work in the previous year	0.003		(0.004)	0.008	*	(0.004)	0.813	***	(0.006)

[a]Adjusted for age, educational attainment, having a spouse, household expenditure in all models
***$p < 0.001$, *$p < 0.05$

Table 4 Estimated impact of the occurrence of parents' need for care on women's K6 scores[a]

	Model 1			Model 2			Model 3			Model 4			Model 5		
	Coef.		SE	Coef.		SE	Coef.		SE	Coef.		SE	Coef.		SE
Parents ($n = 15{,}972$)															
Occurrence of parents' need for care	0.368	***	(0.061)	0.227	***	(0.071)	0.366	***	(0.061)	0.367	***	(0.061)	0.228	***	(0.071)
Caregiving				0.454	***	(0.121)							0.449	***	(0.121)
Coresidence							0.119		(0.164)			(0.080)	0.072		(0.165)
Work										−0.062			−0.057		(0.080)
Parents-in-law ($n = 10{,}887$)															
Occurrence of parents' need for care	0.465	***	(0.073)	0.263	**	(0.085)	0.465	***	(0.073)	0.460	***	(0.073)	0.255	**	(0.085)
Caregiving				0.668	***	(0.145)							0.674	***	(0.146)
Coresidence							−0.028		(0.158)				−0.109		(0.159)
Work										−0.211	*	(0.100)	−0.202	*	(0.100)

[a] Adjusted for age, educational attainment, having a spouse, household expenditure, coresidence, and work in the previous year in all models
***$p < 0.001$, **$p < 0.01$, *$p < 0.05$

Table 5 Estimated impact of the occurrence of parents' need for care on women's K6 scores mediated by their behavior[a]

	Coefficient	95% CI[b]	Proportion (%)	95% CI[b]
Parents ($n = 15{,}972$)				
Mediated by:				
Caregiving	0.139	(0.058, 0.217)	37.7	(15.6, 68.2)
Coresidence	0.001	(−0.004, 0.006)	0.3	(−1.0, 1.8)
Work	0.001	(−0.001, 0.003)	0.2	(−0.4, 0.9)
Total	0.140	(0.061, 0.219)	38.1	(16.1, 68.5)
Unmediated	0.228	(0.088, 0.368)	61.9	(31.5, 83.9)
Total	0.368	(0.249, 0.487)	100.0	
Parents-in-law ($n = 10{,}887$)				
Mediated by:				
Caregiving	0.162	(0.085, 0.239)	44.0	(22.2, 75.4)
Coresidence	0.001	(−0.002, 0.004)	0.1	(−0.4, 1.0)
Work	0.005	(−0.000, 0.011)	1.0	(−0.1, 2.5)
Total	0.210	(0.113, 0.309)	45.1	(23.1, 76.7)
Unmediated	0.255	(0.085, 0.428)	54.9	(23.3, 76.9)
Total	0.465	(0.323, 0.608)	100.0	

[a]Adjusted for age, educational attainment, having a spouse, household expenditure, coresidence, and work in the previous year in all models
[b]Bootstrap-estimated with 3000 replications