unmarried, and separated/divorced). Age groups were divided into "young" (aged 25–39), "middle" (40–59), and "old" (60–79).

We also collected the following four aspects of an individual's relationship with or assessment of social capital from the JGSS, reflecting the previous analysis of the association between social capital and self-rated health (Fujiwara & Kawachi, 2008; Kim & Kawachi, 2006; Subramanian, Kim, & Kawachi, 2002): (i) whether he/she is satisfied with his/her relationships with friends (yes = 1); (ii) whether he/she is satisfied with the place where he/she lives (yes = 1); (iii) whether he/she thinks that most people can be trusted (yes = 1); and (iv) whether he/she belongs to any hobby group or club (yes = 1). For the first two aspects, the JGSS asked respondents to choose from among 1 (= satisfied), 2, 3, 4, and 5 (= dissatisfied). We categorized 1 and 2 as "yes." To avoid multicolinearity from incorporating each of these similar variables into models separately, we tentatively used the mean of four binary answers as the composite index of social capital.

Finally, we considered the size of the area where a respondent lives. The JGSS asked each respondent to choose from among 1 (= small), 2 (= medium), and 3 (= large) in terms of the area, and we used these answers as three categorized variables.

With respect to the second group of variables that were used only to predict perceived happiness, occupational status is potentially most important. Many economic researches have observed that unemployment or an unstable occupational status reduces subjective well-being even after controlling for income (Clark & Oswald, 1994; Korpi, 1997; Winkelmann & Winkelmann, 1998; Di Tella, MacCulloch, & Oswald, 2001). The JGSS asked each respondent about his/her occupational status. We summarized the answers into eight categories: regular employee (including a management executive), non-regular employee, self-employment, family worker, unemployed, retired, homemaker, and other. Retired people and homemakers are not a part of the labor force and are not paid.

Finally, we considered the number of children, which has also been widely used as a predictor of perceived happiness. We also used its squared value as an explanatory variable, considering the possibility of its nonlinear associations with perceived happiness.

Prefecture-level predictors. The most important variable at a prefecture level is the Gini coefficient, which was calculated from the CSLCPHW. The Gini coefficient is one of the most widely used inequality measures, and Kawachi and Kennedy (1997) showed that the choice of inequality measures does not significantly affect the relationship between income inequality and health. We also controlled for (log-transformed) prefecture mean income, the proportion of people aged 60 and above (in both perceived happiness and self-rated health models), and per capita budget expenditure of the local government (in perceived happiness models only).
Additionally, we included indicator variables for 12 regional blocks, each of which comprised three to six prefectures (except Hokkaido) in order to control for the unspecified characteristics of a region wider than a prefecture and those for three years to control for year-specific factors.

Analytic strategy

We started with a one-way analysis of variance (ANOVA) for the means of perceived happiness and self-rated health using six key categories—gender, age, educational attainment, household income, occupational status, and political view—and conducted Holm multiple comparison tests for each category. For occupational status, we categorized it into "stable" (regular employee and self-employed), "unstable" (non-regular employee, family worker, unemployed, and other) and "not a part of labor force" (retired and homemaker). Since it is arguable whether the self-employed should be categorized as "stable" or "unstable," we considered two types of groupings: occupational status (A) in which the self-employed were categorized as "stable" and occupational status (B) in which they were categorized as "unstable." With respect to political views, the JGSS asked a respondent to choose from among five categories (1 = conservative to 5 = progressive) to the question, "Where would you place your political views on a 5-point scale?" We categorized the answers into "conservative" (= 1, 2), "neutral" (= 3), and "progressive" (= 4, 5).

Second, we employed regression analyses to assess the associations of perceived happiness and self-rated happiness with regional inequality and socioeconomic variables. We assumed that the subjective assessments of happiness and health are correlated and ran an ordered bivariate probit model of the form:

$$y_1^* = x_1'\beta_1 + e_1;\ y_1 = 1, \text{ if } y_1 < \mu_{11}, = 2, \text{ if } \mu_{11} < y_1 < \mu_{12}, \ldots, = 5, \text{ if } \mu_{14} < y_1,$$

$$y_2^* = x_2'\beta_2 + e_2;\ y_2 = 1, \text{ if } y_2 < \mu_{21}, = 2, \text{ if } \mu_{21} < y_2 < \mu_{22}, \ldots, = 5, \text{ if } \mu_{24} < y_2.$$

These two equations are correlated and were jointly estimated on the assumption that two disturbances, e_1 and e_2, have the binomial standard normal distribution:

$$\begin{pmatrix}\varepsilon_1\\ \varepsilon_2\end{pmatrix} \sim N\left(\begin{pmatrix}0\\ 0\end{pmatrix}, \begin{pmatrix}1 & \rho\\ \rho & 1\end{pmatrix}\right),$$

with ρ being the covariance of disturbances. Here, y_1 and y_2 are the outcomes for perceived happiness and self-rated health, respectively; y_1^* and y_2^* are their latent variables; x_1 and x_2 are the vectors of predictors; $\mu_{11},\ldots, \mu_{14}$ and $\mu_{21},\ldots, \mu_{24}$ are the threshold parameters; β_1 and β_2 are the vectors of coefficients; and e_1 and e_2 are the disturbances. For multilevel analysis, predictors x_1 and x_2 include individual- and prefecture-level factors as well as indicator variables for regional blocks and years.

We estimated the ordered bivariate probit model to simultaneously estimate β_1, β_2, $\mu_{11},\ldots,$ μ_{14}, $\mu_{21},\ldots, \mu_{24}$, and ρ using the maximum likelihood method. If the null hypothesis that ρ equals zero cannot be rejected, running two ordered probit models separately leads to biased estimation results. Furthermore, we estimated the two ordered probit models for perceived happiness and self-rated health separately (assuming that ρ equals zero) and compared the estimated coefficients with those obtained from the bivariate probit model. In all estimations, we used JGSS-provided sampling weights and computed robust standard errors to correct for potential heteroscedasticity.

Finally, we evaluated effect modification to the sensitivities to regional inequality of perceived happiness and self-rated health using the categories of key individual attributes defined above. For each category group, we estimated the bivariate ordered probit model (Model 1) separately by each category and compared how the sensitivities to regional inequality differed across different categories. For example, we estimated the coefficients on the Gini coefficient separately for males and females to evaluate effect modification by gender. We also examined the statistical significance of the sensitivity to inequality by category. For all of these empirical analyses, we basically used the five percent level as a criterion to assess statistical significance.

3. Results

Overview of descriptive statistics

Before employing the one-way ANOVA and regression analyses, we present an overview of perceived happiness and health on an aggregated basis. Table 2 shows the joint frequency distribution of perceived happiness and self-rated health. As seen in the rightmost column and the bottom row of this table, the proportion of the top two categories is 48 percent for perceived happiness and 63 percent for self-rated health, while the proportion of the bottom two categories is limited to 20 percent and 6 percent, respectively. We also note that the cells at the diagonal have higher frequencies than others, indicating that happier individuals tend to feel healthier and vice versa. Indeed, Spearman's ρ between the two outcomes is calculated as 0.356, which is significantly positive. However, two points should be noted. First, this positive correlation could be accounted for by their associations with the common socioeconomic variables, and not by any causality. Second, a strong correlation does not imply a tight correspondence; healthy people are not necessarily happy and vice versa. Indeed, only 38 percent of the respondents lie at the diagonal of this matrix.

Table 3 compares the means of reported perceived happiness and self-rated health by category and summarizes the results of a multiple-comparison test with Holm-adjusted p-values. Using the five percent significance level, we obtain six findings: (i) females feel healthier than males; (ii) the young feel healthier and somewhat happier than the middle-aged and old; (iii) higher educational attainment makes people feel happier and healthier; (iv) money makes people happier, but its impact on self-rated health diminishes as it increases; (v) an unstable occupational status makes people unhappier; and (vi) the conservative feel happier. However, it is difficult to interpret the results, in particular (v) and (vi), because we cannot rule out reverse causation for each.

Results of ordered probit models

Following the results of the descriptive analysis reported in Tables 2 and 3, Table 4 presents the estimation results of the bivariate ordered probit model for perceived happiness and self-rated health (Model 1) and two ordered probit models that are separately estimated for each outcome (Models 2 and 3). The table summarizes the estimated coefficients on each predictor and their robust standard errors. The results of the bivariate ordered probit model

(Model 1) are divided into the top part (perceived happiness) and the bottom part (self-rated health).

We first note that the coefficient on the Gini coefficient is significantly negative for both perceived happiness and self-rated health (–2.00 and –1.77, respectively), indicating that regional inequality is negatively associated with both outcomes. The estimate of ρ is 0.39, with a standard error of 0.02. The Wald statistic for the test of the null hypothesis that ρ equals zero is 391.44, which is well above 6.63—the critical value of the chi-squared with a single restriction at the one percent level. Hence, we can reject this hypothesis and conclude that a correlation between omitted variables after the influences of key factors in the two equations is significantly positive.

The table also shows the results of the ordered probit models, which are estimated separately for perceived happiness and self-rated health (Models 2 and 3). The pattern of significance of each variable is mostly unchanged from the bivariate probit model (Model 1). However, the absolute value of the coefficient on the Gini coefficient declines modestly for both the outcomes (from 2.00 to 1.60 for perceived happiness and from 1.77 to 1.64 for self-rated health), and their statistical significance decreases (with p-values increasing to above five percent from below it for both outcomes).

Effect modification using key individual attributes

Next, we compared the estimation results across key groups of individuals. Table 5 compares the estimated coefficients on the Gini coefficient in the ordered bivariate models for perceived happiness and self-rated health (Models 1), as well as their robust standard errors and p-values. The first row presents the results obtained from the entire sample (reported in Table 4) as a benchmark for comparisons.

We found substantial differences of effect sizes within all category groups, clearly indicating effect modification by each category. Females are more sensitive to regional inequality than males for perceived happiness. The young are most sensitive to inequality when assessing happiness, while in terms of self-rated health, the middle-aged are more sensitive to inequality than other age groups. Individuals who graduated from college or higher are more sensitive to inequality for happiness than others. Individuals in the highest income class are more sensitive to inequality for perceived happiness than others, while low-income individuals are most sensitive to inequality for self-rated health. Individuals with an unstable occupational status are much more affected by inequality, when assessing both happiness and health, than those in a stable occupational status and those out of the labor force. This result holds regardless of categorizing the self-employed as stable or unstable. Finally, those who are politically neutral are most sensitive to inequality for perceived happiness, while progressive individuals are most sensitive to self-rated health.

However, we should be cautious in interpreting these results because comparisons of estimated coefficients on the Gini coefficient do not make sense if the Gini coefficient is distributed differently between categories. To check this, we applied the Kolmogorov-Smirnov tests between each category and the remaining one or two categories in each category group. We found that the null hypothesis that the Gini coefficient is distributed differently between categories cannot be rejected at the five percent significance level for two cases: between individuals who graduated from colleges or higher and others and between low income-individuals and others.

4. Discussion

Our estimation results confirmed the negative impact of regional inequality on both perceived happiness and self-rated health, a result consistent with those of the previous studies that analyzed the two outcomes separately. The novelty of our analysis is that it obtained this result even after controlling for various individual and regional characteristics and taking into account the correlation between the two subjective outcomes.

We also found that key socioeconomic factors tend to affect perceived happiness and self-rated health in the same directions with statistical significance. More specifically, both our ANOVA and regression analyses confirm that higher-income, younger age, and a higher level of social capital make people feel both happier and healthier. It is likely that these relations account for the observed positive correlation between the two subjective outcomes. This result also suggests that it is incorrect to view the relation between the two subjective outcomes in a unidirectional manner. Furthermore, the comparisons between the bivariate ordered probit model (which estimated perceived happiness and self-rated health jointly) and two ordered probit models (which estimated the two outcomes separately) revealed that separate estimations tend to underestimate the magnitude of the association of regional inequality.

Another notable finding is that key individual attributes modify the association of regional inequality with subjective assessments of happiness and health. Notably, the widening inequality most directly reduces the well-being of those in an unstable status and who face the most serious uncertainty about future employment and income. Alesina et al. (2004) found that the rich and the right-wingers are largely unaffected by inequality, while inequality has strong negative effects on the perceived happiness of the poor and left-wingers in Europe. They also observed that the poor and the left-wingers are not affected by inequality, while the effect on the rich is negative and well-defined in the United States. The case of Japan differs from that of both Europe and the United States; while the rich Japanese are affected by inequality, and the politically neutral rather than progressive or conservative ones are most sensitive to it. In addition, the association of inequality tends to show different patterns across individual features between perceived happiness and self-rated health.

It should be noted, however, that those with an unstable occupational status are strongly affected by inequality in terms of both perceived happiness and health. Along with the fact that they are unhappier than others, as seen from Table 3, it suggests that occupational status is one of the key determinants of individual well-being. These facts should be taken seriously, given the steadily declining proportion of regular employees in the labor market in Japan.

We recognize that this analysis has various limitations. First, we dealt with happiness as only a single item based on the survey results of its subjective assessment. This approach followed many empirical studies of happiness and was also parallel to the analysis of self-rated health. However, it cannot be free from criticism that happiness should be multi-dimensional and that it cannot be summarized as a single item. Given this feature of happiness, the validity of perceived happiness observed from surveys should be addressed further.

Second, while we took into account the correlation between perceived happiness and health when estimating regression models, the manner in which these two aspects of individual welfare interact with each other remains to be addressed. Third, as is often the case with a multilevel analysis of this type, pathways or mediation process from income inequality in society with respect to subjective outcomes at an individual level should be further investigated. Fourth, we disregarded the possibility that subjective outcomes change individual characteristics, which we assumed to be exogenous. These issues should be explicitly addressed in future research.

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Table 1. Selected descriptive statistics (pooled data for 2000, 03, 06)

		Mean	S.D.	Min	Max
(1) Prefecture-level variables: N = 141 (47 prefectures * 3 years, not weighted)					
Gini coefficient		0.370	0.027	0.308	0.436
Mean household income (million yen)[a]		3.104	0.496	1.677	4.437
Per capita budget expenditure (million yen)[b]		0.451	0.128	0.195	0.873
Proportion of people aged 65 and above (%)		20.8	3.1	12.8	27.6
(2) Individual-level variables: N = 4,466 (1,872 in 2000; 1,237 in 2003; 1,357 in 2006)					
Household income (thousand yen)[a]		3,683	2,543	0	32,200
Age		52.7	14.4	25	80
Number of children		1.83	1.08	0	10
Social capital		0.56	0.21	0	1
Categorical variables		Percentage			
Gender	Males	49.4	(reference)[d]		
	Females	50.6			
Age group	Young (aged 25-39)	22.7	(reference)		
	Middle (aged 40-59)	41.2			
	Old (aged 60-79)	36.1			
Marital status	Married	81.1	(reference)		
	Unmarried	8.7			
	Divorced/widowed	10.2			
Educational attainment	Junior high school or lower	21.2	(reference)		
	High school or lower	48.5			
	College or higher	30.3			
Income class	Low	33.3	(reference)		
	Middle	32.9			
	High	33.7			
Occupational status	Regular employee[c]	37.2	(reference)		
	Non-regular employee	14.2			
	Self-employed	9.4			
	Family worker	2.8			
	Unemployed	1.3			
	Retired	9.7			
	Homemaker	19.9			
	Other	4.5			
Size of residential area	Small	31.8	(reference)		
	Medium	40.7			
	Large	27.5			
(3) Regional blocks	Hokkaido, Tohoku, Kanto 1&2, Hokuriku, Tokai, Kinki 1&2, Chugoku, Shikoku, Kyushu 1&2				

Note: a. household size adjusted, pre-tax, and evaluated at 2005 prices.
b. evaluated at 2005 prices.
c. includes a management executive.
d. indicates the reference group for each category group in regression models.

Table 2. Joint frequency distribution of perceived hapiness and self-rated health

(percent)

Happiness	Self-rated health					
	1 (= poor)	2	3	4	5 (= excellent)	Total
1 (= unhappy)	0.3	0.8	1.4	0.6	0.7	3.8
2	0.3	2.0	6.7	4.6	2.2	15.9
3	0.3	1.3	12.6	10.6	7.0	31.9
4	0.1	0.7	5.9	10.3	7.4	24.4
5 (= happy)	0.2	0.3	3.7	7.0	12.8	24.0
Total	1.3	5.1	30.3	33.2	30.1	100.0

Note: Spearman's ρ = 0.356 (p <0.001).

Table 3. Comparing perceived happiness and self-rated health by key individual attributes

	Mean	S.E.	Difference in means		Holm-adjusted p-value	Number of observations
Perceived happiness (1 = unhappy, 2, 3, 4, 5 = happy)						
Gender						
(1) Male	3.83	(0.94)				2,196
(2) Female	3.88	(0.96)	(2) - (1) =	0.04	0.122	2,247
Age						
(1) Middle	3.82	(0.94)				1,008
(2) Old	3.85	(0.96)	(2) - (1) =	0.04	0.792	1,831
(3) Young	3.94	(0.94)	(3) - (2) =	0.09	0.075	1,604
Educational attainment						
(1) Junior high school or lower	3.76	(1.02)				940
(2) High school	3.82	(0.98)	(2) - (1) =	0.08	0.165	2,158
(3) College or higher	3.95	(0.87)	(3) - (2) =	0.12	<0.001	1,354
Income						
(1) Low	3.72	(1.03)				1,477
(2) Middle	3.83	(0.91)	(2) - (1) =	0.11	0.006	1,466
(3) High	4.03	(0.88)	(3) - (2) =	0.20	<0.001	1,500
Occupational status (A)						
(1) Unstable	3.74	(1.03)				2,073
(2) Stable	3.87	(0.91)	(2) - (1) =	0.14	<0.001	1,052
(3) Out of labor force	3.93	(0.94)	(3) - (2) =	0.05	0.344	1,318
Occupational status (B)						
(1) Unstable	3.78	(1.01)				1,655
(2) Stable	3.87	(0.90)	(2) - (1) =	0.10	0.015	1,470
(3) Out of labor force	3.93	(0.94)	(3) - (2) =	0.05	0.368	1,318
Political view						
(1) Progressive	3.81	(0.95)				2,169
(2) Neutral	3.83	(0.93)	(2) - (1) =	0.02	1.000	991
(3) Conservative	3.97	(0.94)	(3) - (2) =	0.14	0.002	1,127
Self-rated health (1 = poor, 2, 3, 4, 5 = excellent)						
Gender						
(1) Male	3.44	(1.12)				2,205
(2) Female	3.53	(1.14)	(2) - (1) =	0.09	0.010	2,262
Age						
(1) Old	3.33	(1.20)				1,614
(2) Middle	3.53	(1.08)	(2) - (1) =	0.20	<0.001	1,841
(3) Young	3.67	(1.08)	(3) - (2) =	0.15	0.002	1,012
Educational attainment						
(1) Junior high school or lower	3.23	(1.22)				947
(2) High school	3.52	(1.12)	(2) - (1) =	0.29	<0.001	2,164
(3) College or higher	3.61	(1.06)	(3) - (2) =	0.09	0.044	1,355
Income						
(1) Low	3.31	(1.19)				1,489
(2) Middle	3.55	(1.11)	(2) - (1) =	0.25	<0.001	1,471
(3) High	3.60	(1.07)	(3) - (2) =	0.05	0.678	1,507
Occupational status (A)						
(1) Unstable	3.35	(1.20)				1,323
(2) Stable	3.51	(1.16)	(2) - (1) =	0.16	0.004	1,061
(3) Out of labor force	3.57	(1.06)	(3) - (2) =	0.06	0.430	2,083
Occupational status (B)						
(1) Unstable	3.35	(1.20)				1,323
(2) Stable	3.52	(1.16)	(2) - (1) =	0.17	<0.001	1,481
(3) Out of labor force	3.57	(1.04)	(3) - (2) =	0.05	0.513	1,663
Political view						
(1) Progressive	3.44	(1.14)				993
(2) Neutral	3.48	(1.10)	(2) - (1) =	0.05	0.811	2,173
(3) Conservative	3.55	(1.17)	(3) - (2) =	0.07	0.235	1,228

Note: The self-employed are categorized as "stable" and "unstable" in Occupational status (A) and (B), respectively.

Table 4. Estimated associations of independent variables with perceived happiness and self-rated health

		Bivariate ordered probit model			Ordered probit models		
		Coef.	Robust S.E.	p-value	Coef.	Robust S.E.	p-value
Perceived happiness (1 = unhappy, 2, 3, 4, 5 = happy)		*Model 1*			*Model 2*		
Gini		-2.00	(0.94)	0.033	-1.60	(0.94)	0.089
Income class:	Middle	0.01	(0.05)	0.883	0.00	(0.05)	0.995
	High	0.26	(0.05)	<0.001	0.25	(0.05)	<0.001
Gender:	Female	0.09	(0.05)	0.064	0.09	(0.05)	0.056
Age group:	Middle	-0.32	(0.05)	<0.001	-0.33	(0.05)	<0.001
	Old	-0.24	(0.06)	<0.001	-0.23	(0.06)	<0.001
Marital status:	Unmarried	-0.67	(0.08)	<0.001	-0.64	(0.08)	<0.001
	Divorced/widowed	-0.39	(0.06)	<0.001	-0.38	(0.06)	<0.001
Educational attainment:	High school	-0.03	(0.05)	0.548	-0.03	(0.05)	0.499
	College or higher	0.01	(0.06)	0.821	0.02	(0.06)	0.773
Occupational status:	Non-regular employee	-0.08	(0.06)	0.161	-0.08	(0.06)	0.187
	Self-employed	0.00	(0.06)	0.992	0.02	(0.07)	0.788
	Family worker	-0.12	(0.10)	0.259	-0.10	(0.10)	0.325
	Unemployed	-0.67	(0.18)	<0.001	-0.64	(0.19)	0.001
	Retired	0.10	(0.07)	0.171	0.03	(0.07)	0.653
	Homemaker	0.09	(0.06)	0.151	0.05	(0.06)	0.394
	Other	0.07	(0.09)	0.470	-0.01	(0.10)	0.946
Size of residential area:	Medium	0.10	(0.05)	0.026	0.10	(0.05)	0.026
	Large	0.03	(0.05)	0.474	0.04	(0.05)	0.388
Number of children		0.03	(0.04)	0.456	0.07	(0.05)	0.148
Number of children squared		-0.01	(0.01)	0.518	-0.01	(0.01)	0.219
Social capital		1.15	(0.09)	<0.001	1.15	(0.09)	<0.001
Log of mean household income		0.10	(0.26)	0.706	0.03	(0.26)	0.908
Proportion of people aged 65 and above		-0.02	(0.01)	0.120	-0.01	(0.01)	0.401
Per capita budget expenditure		0.78	(0.24)	0.001	0.51	(0.25)	0.046
Self-rated health (1 = poor, 2, 3, 4, 5 = excellent)					*Model 3*		
Gini		-1.77	(0.85)	0.037	-1.64	(0.85)	0.053
Income class:	Middle	0.10	(0.04)	0.023	0.11	(0.04)	0.015
	High	0.11	(0.05)	0.021	0.12	(0.05)	0.011
Gender:	Female	0.12	(0.03)	<0.001	0.12	(0.03)	<0.001
Age group:	Middle	-0.15	(0.05)	0.001	-0.15	(0.05)	0.001
	Old	-0.32	(0.05)	<0.001	-0.32	(0.05)	<0.001
Marital status:	Never married	-0.05	(0.06)	0.385	-0.06	(0.06)	0.365
	Divorced/widowed	0.00	(0.06)	0.984	-0.01	(0.06)	0.936
Educational attainment:	High school	0.11	(0.05)	0.027	0.11	(0.05)	0.024
	College or higher	0.11	(0.06)	0.048	0.11	(0.06)	0.049
Size of residential area:	Medium	0.02	(0.05)	0.639	0.02	(0.05)	0.657
	Large	0.02	(0.05)	0.648	0.02	(0.05)	0.643
Social capital		1.06	(0.09)	<0.001	1.05	(0.08)	<0.001
Log of mean household income		0.21	(0.25)	0.402	0.21	(0.25)	0.407
Proportion of people aged 65 and above		0.02	(0.01)	0.079	0.02	(0.01)	0.107
atnhρ		0.39	(0.02)	<0.001			
ρ		0.37	(0.02)				

Note. 1. Italics denotes category. See Table 1 for the reference group for each category.

2. All models include indicator variables for regional blocks and survey years.

3. For the bivariate ordered probit model, Wald test of $\rho = 0$: $\chi^2(1) = 391.44$.

4. The numbers of observations are 4,442, 4,442, and 4,466 for Models 1-3, respectively.

Table 5. Comparing estimated sensitivities to inequality of perceived happiness and self-rated health by key individual attributes

	Perceived happiness			Self-rated health		
	Coef.	Robust S.E.	*p* -value	Coef.	Robust S.E.	*p* -value
Total	-2.00	(0.94)	0.033	-1.77	(0.85)	0.037
Gender						
Male	-0.71	(1.36)	0.604	-1.77	(1.21)	0.142
Female	-3.49	(1.32)	0.008	-1.96	(1.22)	0.107
Age						
Young	-3.66	(2.02)	0.070	-1.20	(1.90)	0.528
Middle	-1.39	(1.47)	0.344	-2.95	(1.29)	0.022
Old	-2.02	(1.50)	0.178	-1.09	(1.41)	0.438
Educational attainment						
Junior high school or lower	-1.57	(1.07)	0.143	-1.76	(0.96)	0.066
High school	-0.73	(1.34)	0.586	-1.89	(1.21)	0.118
College or higher*	-3.25	(1.73)	0.061	-1.88	(1.58)	0.234
Household income						
Low*	-1.47	(1.49)	0.326	-2.70	(1.45)	0.062
Middle	-0.78	(1.62)	0.631	-0.05	(1.58)	0.977
High	-3.20	(1.77)	0.071	-1.85	(1.45)	0.202
Occupational status (A)						
Stable	-0.60	(1.40)	0.671	-1.36	(1.26)	0.280
Unstable	-5.70	(1.99)	0.004	-4.00	(1.85)	0.031
Out of labor force	-1.42	(1.67)	0.393	-0.30	(1.55)	0.845
Occupational status (B)						
Stable	0.03	(1.58)	0.986	-0.50	(1.39)	0.722
Unstable	-5.09	(1.68)	0.002	-4.22	(1.57)	0.007
Out of labor force	-1.42	(1.67)	0.393	-0.30	(1.55)	0.845
Political view						
Progressive	0.40	(2.08)	0.847	-3.31	(1.85)	0.074
Neutral	-3.37	(1.38)	0.015	-0.93	(1.21)	0.444
Conservative	-2.19	(1.72)	0.204	-1.95	(1.62)	0.228

Note. 1. This table compares the estimated coefficients on the Gini coefficient in the bivariate ordered probit models for each category.

2. The self-employed are categorized as "Stable" and "Unstable" in Work status (A) and (B), respectively.

3. The null hypothesis that the distribution of the Gini coefficient differs between the category with * and the other two categories in the same category group cannot be rejected at the five percent significance level.

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分担研究報告書

Does Health Status Matter to People's Retirement Decision in Japan?: An Evaluation of "Justification Hypothesis" and Measurement Errors in Subjective Health Evidence from the Fiscal 2008 Survey on Health and Retirement

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研究要旨

この論文では，2008 年度の「健康と引退に関する調査」の個票データを用いて，主観的健康指標に含まれうる内生バイアスの大きさの評価を行う。内生バイアスの原因として，(1) 回答者が，自分が働いていないことを正当化するために，健康状態を低めに申告する行動(正当化仮説)と，(2) 健康指標に含まれる測定誤差，の二つに注目する。

我々は，主に以下の三つの方法でバイアスの大きさとその有意性を評価した。一つ目は，主観的及び客観的な健康指標のカーネル分布の特性や，就業状態へ与える効果の比較である。二つ目は，操作変数を用いた場合の健康が就業状態に与える効果と，それを用いなかった場合の効果の比較である。三つ目は，正当化仮説を検証するために，本来は相関を持たないと考えられる，不健康を理由とする引退の確率と有効求人倍率の時系列的な相関の有無を分析することである。この論文では，主観的健康指標として，回答者が自分の健康状態を 5 段階評価した指標と，過去一年間における，体調あるいは気分がすぐれないことによる日常生活や仕事への支障の有無を 4 段階評価した指標の二つを用い(両者とも二値変数に変換して利用)，客観的健康指標としては，各回答者の慢性疾患の数と，主成分分析に基づいて指標化された健康状態の二つを用いた。上記の二番目の分析方法では，操作変数として，回答者の居住地(同一の郵便番号の地域の重心となる地点)から最寄りの小規模病院(病床数 20 以上 100 以下)までの最短距離，回答者が属している二次医療圏の診療所密度，30 歳時点の BMI (Body Mass Index) の値の三つを用いた。

これらの分析の結果，健康指標の内生バイアスの存在を示唆するいくつかのエビデンスが得られた。まず，一つ目の分析では，主観的な健康状態が良い回答者の中にも客観的な健康状態の悪い人がいることが分かった。このことは，自分の健康状態を正確に評価することが難しいことを意味しているのかもしれない。また，逆に，主観的な健康状態が悪い回答者の中に客観的な健康状態の良い人が相当数いる結果も得られたが，このことは我が国における正当化仮説の妥当性を示唆している可能性がある。二つ目の分析では，日常生活や仕事への支障の有無は，健康状態に影響を与えると考えられる外生変数(操作変数)と弱い相関しか持たないにもかかわらず，回答者の就業状態には強い影響を与えるという結果が得られた。このことは，既に退職した回答者がこのことを健康の悪化で正当化するために自分の健康状態を低めに申告しているか，あるいは，自分の健康状態を正確に評価することが難しいことを示唆しているのかもしれない。さらに，操作変数法で推定した健康の効果は OLS 推定の結果よりも大きいことから，健康指標の測定誤差による 0 方向のバイアスが緩和されたと考えることもできるが，操作変数の妥当性がそれほど高くないと判断されるため，結果の解釈には注意が必要である。最後に，三つ目の分析では，不健康を

理由とする引退と有効求人倍率の間に正の相関が見られた。このことは，好況期の引退を不健康によって正当化していると考えることができる。

A. 研究目的

本論文の目的は，主観的健康指標に含まれうる内生バイアスの大きさの評価を行うことである。内生バイアスの原因として，(1) 回答者が，自分が働いていないことを正当化するために，健康状態を低めに申告する行動 (正当化仮説)，(2) 健康指標に含まれる測定誤差，の二つに注目する。

B. 研究方法

2008 年度「健康と引退に関する調査」の個票データを用いて，(1) 主観的健康指標と客観的健康指標のカーネル分布の特性や，それらが就業状態へ与える効果の比較，(2) 操作変数を用いた場合の健康が就業状態に与える効果と，それを用いなかった場合の効果の比較，(3) 正当化仮説を検証するために，本来は相関を持たないと考えられる，不健康を理由とする引退の確率と有効求人倍率の時系列的な相関の有無の分析，の三つの方法で研究を行った。

C. 研究結果

一つ目の分析では，主観的な健康状態が良い回答者の中にも客観的な健康状態の悪い人がいることが分かった。また，逆に，主観的な健康状態が悪い回答者の中に客観的な健康状態の良い人が相当数存在するという結果も得られた。二つ目の分析では，日常生活や仕事への支障の有無は，健康状態に影響を与えると考えられる外生変数 (操作変数) と弱い相関しか持たないにもかかわらず，回答者の就業状態には強い影響を与えるという結果が得られた。三つ目の分析では，不健康を理由とする引退と有効求人倍率の間に正の相関が見られた。

D. 考察

一つ目の分析結果から，自分の健康状態を正確に評価することが難しいことと，正当化仮説の妥当性が示唆される。二つ目の分析結果は，既に退職した回答者がこのことを正当化するために自分の健康状態を低めに申告しているか，自分の健康状態を正確に評価することが難しいことを意味している可能性がある。三つ目の分析結果は，再就職先を見つけるのが相対的に容易と考えられる好況期における引退を，自らの健康状態の悪化によって正当化していると解釈できる。

E. 結論

上記の三つの分析の結果から，我が国においては，主観的健康指標に内生バイアスが生じていることが疑われる。

F. 健康危険情報

なし

G. 研究発表

1.論文発表

なし

2.学会発表

・Does Health Status Matter to People's Retirement Decision in Japan?: An Evaluation of "Justification Hypothesis" and Measurement Errors in Subjective Health, International Health Economics Association, 7th World Congress on Health Economics（於北京，2009 年 7 月 12-15 日)。

・Does Health Status Matter to People's Retirement Decision in Japan?: An Evaluation of "Justification Hypothesis" and Measurement Errors in Subjective Health, Econometric Society, 2009 Far East and South Asia Meeting of the Econometric Society (於東京大学, 2009 年 8 月 3-5 日)。

・Does Health Status Matter to People's Retirement Decision in Japan?: An Evaluation of "Justification Hypothesis" and Measurement Errors in Subjective Health, 日本経済学会, 2009 年度日本経済学会秋季大会(於専修大学, 2009 年 10 月 10-11 日)。

・Does Health Status Matter to People's Retirement Decision in Japan?: An Evaluation of "Justification Hypothesis" and Measurement Errors in Subjective Health, 統計研究会, 労働市場研究委員会, (於構造計画研究所新橋イノベーションサイト, 2009 年 10 月 30 日)

H. 知的所有権の取得状況の出願・登録状況

1.特許取得

なし

2.実用新案登録

なし

3.その他

The studies on the design of social security benefit and contribution schemes with attention to the relations between income, assets, consumption and the burdens of social security premium and tax:
Report for Fiscal 2009 (Study Supported by the Health Science Research Grants from the Ministry of Health, Labour and Welfare (Study Project for Promotion of Policy Sciences))

Does Health Status Matter to People's Retirement Decision in Japan?: An Evaluation of "Justification Hypothesis" and Measurement Errors in Subjective Health Evidence from the Fiscal 2008 Survey on Health and Retirement[1]

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1 Introduction

The Japanese society has been facing a rapid aging and a decrease in a birthrate for the last couple of decades. The large shortage of workforce will be one of the most critical socio-economic issues. Lately, numbers of health economists focus on the relation between health status and job continuation around retirement age, since healthy elderly persons are expected to offset the lack of labor force. Health status is often one of major reasons for the retirement of workers in Japan. For example, the basic statistics based on the data using in this study shows that bad health and/or deteriorating health is the second significant reasons for males and females to leave the labor market which follows the mandatory retirement by the employers and the retirement after marriage, respectively (Table 1). However, since self-reported health status which is sometimes unreliable would cause statistical bias, we have to use this variable very carefully in econometric analysis.

This study identifies the significance of the endogeneity biases in the estimated health effects. We address the biases arose from the following two sources: (1) "justification hypothesis," wherein retired respondents are assumed to justify their leaving labor force (e.g. early retirement) by false poor health (Chirikos and Nestel, 1984; Anderson and Burkhauser, 1985; Bazzoli, 1985; Bound, 1991; Waidmann et al., 1995; Dwyer and

[1]We would like to thank Akiko Oishi, Charles Yuji Horioka, Yasushi Iwamoto, Hiroshi Ogawa, Ryo Kambayashi, and Yoshiaki Omori for their helpful comments, given at the 2009 Far East and South Asia Meeting of the Econometric Society, the 2009 Autumn Meeting of the Japanese Economic Association, and the meeting of the *Rodo Shijo Kenkyu Iinkai* held at October 30, 2009.

Mitchell, 1999), and (2) classical measurement errors in the health variables. If the hypothesis holds true, poorer health is more frequently reported by the retired respondents, all other things being equal. Hence, health effects on labor market participation can be overestimated, as opposed to the attenuation bias of the measurement error.

Since the retirement age of Japan is much higher than other countries, "justification hypothesis" may be more applicable to this nation. However, only a few works (Iwamoto, 2000; Oishi, 2000) have tackled this problem previously. Iwamoto (2000) and Oishi (2000) compare the effects of several health indicators on wage income, labor market participation, and retirement behavior, controlling for the endogeneity bias. Both studies find the different health effects between those measurements. In particular, Iwamoto (2000) points out that subjective health indicators (self-rated health and presence of work limitations) have more obvious effects on income and employment status than objective ones, suggesting that the measurement error in subjective health is not so severe that researchers can use it in empirical analysis. However, this clearer effect of subjective health itself may be an evidence of the severity of the justification bias.

This study therefore evaluates the endogeneity biases by the following three strategies: (1) comparing the properties and effects on employment status of various health measurements, (2) using three instrumental variables (IVs) for health status that have never been used in previous studies, and (3) analyzing a relation of seemingly unrelated variables to verify "justification hypothesis." We use several kinds of health measurements such as binary subjective health (self-reported poor health and limitations of daily activities at home and/or on the job), number of chronic diseases which have not been completely recovered by the latest timing of the survey, and our original health status scoring based on principle component analysis. Our IVs are distance in a straight line from respondent's home to the nearest low-volume hospital, variations in the number of clinics among different medical spheres, and a body mass index (BMI) in respondent's 30 years of age. First, we mainly evaluate how the results differ between objective and subjective health measures. Since the objective health is generally less affected by the justification behavior and measurement errors, its bias should be smaller than subjective one. Second, compared to being not-instrumented, instrumented health effects will decrease if people actually justify their unemployed status by poor health and will expand if the measurement error is a serious problem in a model. Finally, we examine the relation between the ratio of the retirement due to bad health status and a job openings-to-applications ratio. If "justification hypothesis" makes sense, those variables can be correlated because the respondents who retired in the period of the high job openings ratio may be more likely to justify their retirement by false poor health.

In this study, we apply a Japanese version of Health and Retirement Survey conducted by the National Institute of Population and Social Security Research in 2008 and 2009, which was funded by a research grant from the Ministry of Health, Labor and Welfare. The survey focuses on those who are around retirement age and includes detailed information on various objective and subjective health conditions, retirement behavior, job status, working hours, and financial status.

As a result, we obtain several evidences for "justification hypothesis" and the measurement errors in the health variables. First, our objective health measures widely distribute in the poorer subjective health. This indicates a possibility of the endogeneity problems in subjective health. Second, the limitation of daily activities is strongly correlated with one's employment status in spite of its weak correlation with exogenous factors determining health. A likely explanation of this result is that respondents report the poorer health status than true one to justify their employment status or they may not be able to assess their own health status accurately oneself. Third, instrumented health effects are larger than not-instrumented ones, as would be consistent with an alleviation of attenuation bias. Finally, the intrinsically unrelated variables are positively correlated with each other both in time series and cross section. This demonstrates a noticeable tendency of justifying the retirement in good economic condition.

This paper is organized as follows. Section 2 formulates our empirical specification. Section 3 describes the data source and our variables, including employment status variables and health measures. Section 4 looks for the exogenous determinants of health status. Section 5 shows the estimation results. Section 6 adduces an evidence supporting "justification hypothesis." Section 7 presents the conclusions of this paper.

2 Empirical specifications

This paper employs various empirical specifications, including probit, linear probability model (LPM), two-stage least squares (2SLS), and Tobit models, in order to accomplish the evaluation of the endogeneity problems. We can expect that attenuation bias will occur in LPM if measurement errors exist in the subjective health indicator. Meanwhile, the direction of the bias is theoretically ambiguous in the maximum likelihood estimate (MLE).[2] Hence, we compare the outcomes of LPM and 2SLS regressions to check the seriousness of attenuation bias. However, since LPM has some deficiencies (e.g. some of the LPM fitted value may be outside the unit interval), we also use the probit model in appraising "justification hypothesis."

We specify the following three econometric models: (1) univariate probit model, (2) IV probit model, and (3) Tobit model. Dichotomous employment status indicators and censored working hours are dependent variables in (1)-(2) and (3), respectively. Here, we omit a specification of LPM because it is a simple OLS which has a binary variable in the left-hand side.[3]

[2]Levine (1985) considers the measurement error bias in MLE, including probit and censorship type model estimates. He suggests that MLE is affected not only by the classical attenuation bias but also by the additional effects which determine the direction of the bias due to measurement error, differently from normal linear model. Hsiao (1991) and Wang (1998) explore identification conditions for binary choice and censored models, respectively. Two- or three-step procedures for estimating a consistent estimate and the corresponding asymptotic covariance matrices are proposed in their papers. Recently, Edgerton and Jochumzen (2003) reveal by the Monte Carlo and empirical studies that attenuation occurs in the coefficient of independent variable(s) of probit model that is measured with error. They also derive multi-step LIML estimator and find its consistency and good small-sample property under some assumptions.

[3]Since a binary response in the left-hand side is a Bernoulli random variable in LPM, its conditional variance is expressed as $X\beta(1 - X\beta)$, where X and β are a vector of covariates and its coefficients,

First, we explain the probit models. Let y_i be a binary employment status variable; for example, it takes unity if an individual has retired or has no jobs and zero otherwise. Consider the following binary choice model:

$$y_i = 1(y_i^* = \alpha h_i + \beta X_{1i} + \epsilon_i > 0), \tag{1}$$

where y_i^* denotes an unobserved latent variable; h_i, an observed health measure; X_{1i}, a vector of other household characteristics; and ϵ, a stochastic error term which has a standard normal distribution. If h_i is an exogenous health variable, α will be estimated to be consistent. However, if h_i is measured with error, the attenuation bias will occur in the estimate of α. Moreover, under the "justification hypothesis," the health effect on retirement can be overestimated because people try to justify their early retirement by false poor health.

In order to address those endogeneity biases, we employ IV probit model. This model is formulated as follows:

$$\begin{aligned} y_i &= 1(y_i^* = h_i\alpha + X_{1i}\beta + \epsilon_i > 0), & (2)\\ h_i &= X_{1i}\gamma + X_{2i}\delta + v_i, & (3) \end{aligned}$$

where X_{2i} is a vector of additional instruments and (ϵ_i, v_i) has a zero-mean and bivariate normal distribution. The error terms are permitted to be correlated one another, $Cov(\epsilon_i, v_i) = \rho$. On the other hand, this simultaneous model breaks into two parts for y_i and h_i when $\rho = 0$, implying that it is appropriate to use the univariate probit model, eq. (1). Even if h_i is a binary endogenous variable, the above simultaneous model will still generate a consistent estimate, but the estimate may not be efficient. In this case we have to use the recursive bivariate probit model, wherein the first-stage equation eq. (2) is a reduced form probit model for binary health indicator, in order to obtain an efficient estimate.

Next, we show the standard censored Tobit model that is adopted to estimate the health effect on hours worked. Let y_i denote working hours, and then we formulate it as follows:

$$\begin{aligned} y_i^* &= \alpha h_i + \beta X_{1i} + \epsilon_i, & (4)\\ y_i &= \max(0, y_i^*), & (5) \end{aligned}$$

where y_i^* is a latent variable which is observed for values greater than 0 and censored otherwise; and $\epsilon_i \sim N(0, \sigma^2)$. IV Tobit model allows h_i be endogenous through a correlation of the error terms in the health and working hours equations. However, we do not use it due to a severe weak identification problem in this model.

respectively. Apparently, heteroskedasticity is the case we have to consider in this variance unless all coefficients are zero; therefore, we use heteroskedasticity-robust standard errors in LPM to deal with this issue.

3 Data and variables

3.1 Data source and sample selection

The data in this paper is the *Survey on Health and Retirement*, conducted by the National Institute of Population and Social Security Research in March of 2008 and 2009. In order to examine various effects of people's health status on retirement behavior, the survey focuses on males and females who are 45 and older and younger than 80 years old. For the first wave of the survey in the year of 2008, 2,747 people are randomly extracted out of 39,311 monitoring samples owned by the *Central Research Services, Inc* (CRS). The monitoring samples are collected by the monthly omnibus survey conducted by CRS. The CRS extracts samples randomly from the residents' administrative registration records every month and creates the master sample including those who agree to be monitored for all kinds of surveys. For adjusting the distributions of respondents' sex and age to the National Census, the CRS carefully extracts the samples in a way that the number of respondents becomes proportional to the number of population in each sex and 5-year age group based on the residents' administrative registration records in each municipal city. The remuneration paid for respondents is a 500 yen coupon ticket for purchasing books. Out of those, 1,074 people responded the survey (valid response rates: 39%) in the first wave. Then, the second wave is a follow-up survey on these 1,074 respondents. Out of 1,074, 862 respondents (response rate: 80%) answered the survey and so 212 (approximately 20%) dropped out from the sample. Further, in the second wave, 578 people are newly chosen at random from CRS monitoring samples. Out of 578, 257 people (response rate: 44%) responded the survey.

This survey has a couple of unique characteristics which is different from the data used in previous studies such as Iwamoto (2000) and Oishi (2000). First, the survey asked respondents diseases in detail which were diagnosed by physicians. Hence, we can control for respondents' both subjective and relatively objective chronic health status more accurately than previous works. Second, the survey includes the data on a respondent's retirement and re-employment history in the past. Therefore, this study would distinguish respondents who have not been retired yet from those who have been re-employed either on full-time or part-time basis since the first compulsory retirement.

Among whole sample, we use only male respondents in our econometric analysis because of some complications of the female retirement behavior. Since a number of life-cycle related factors (e.g. getting married, baring children, and providing long-term care to family members) make female workers leave labor market more often compared to male workers, a simple analytic framework probably cannot describe the mechanism of female retirement behavior.[4] Moreover, compared to males, females are less likely to feel embarrassed by leaving labor market in young and having no job, which is not uncommon for women in Japan. Therefore, "justification hypothesis" probably does not matter to female workers. For the same reason, Iwamoto (2000) also does not include female work-

[4] 11.5% of female respondents choose "Other reasons" as a reason for retirement while only 5.3% male respondents choose it. This is an evidence of the variety in the causes of the female retirement.