# Lingering Impact of Starting Working Life During a Recession: Health Outcomes of Survivors of the "Employment Ice Age" (1993–2004) in Japan

Takashi Oshio

Institute of Economic Research, Hitotsubashi University, Tokyo, Japan

## ABSTRACT

**Background:** A growing amount of evidence demonstrates the adverse impacts of economic downturns on population health. However, the extent to which the macroeconomic conditions at labor market entry affect health outcomes in later life remains relatively understudied. This study focused on the health outcomes of the cohort who entered the labor market during the "employment ice age" (EIA; 1993–2004) in Japan, when young people had difficulty finding jobs after graduating from college or high school.

**Methods:** We used repeated cross-sectional data ($N = 3,054,782$; 1,500,618 men and 1,554,164 women) obtained from an 11-wave population-based nationwide survey conducted every 3 years from 1986 through 2016. We considered three health outcomes: being in hospital, subjective symptoms, and self-rated health (SRH). We employed two types of statistical analyses: an age-period-cohort (APC) analysis, which controlled for age and period (wave) effects, and a difference-in-differences (DiD) analysis, in which the EIA experience was regarded as a treatment.

**Results:** The APC analysis confirmed the relative disadvantage of the EIA cohort for all three outcomes; for instance, the odds ratio of poor SRH for the EIA cohort was 1.29 (95% confidence interval [CI], 1.21–1.38) for men and 1.25 (95% CI, 1.17–1.34) for women. The DiD analysis confirmed the robustness of these results, especially for men.

**Conclusions:** The results underscored the lingering impact of the macroeconomic conditions at labor market entry on health outcomes in later life in Japan.

**Key words:** age-period-cohort analysis; difference-in-differences analysis; employment ice age; self-rated health; subjective symptoms

## INTRODUCTION

A growing amount of evidence has demonstrated the adverse impact of macroeconomic downturns on population health. Studies have shown that the Great Recession (2008–2009) associated with the global financial crisis had negative effects on mental health, mortality, self-rated health (SRH), and other health outcomes.[1–4] The possible key channels that link macroeconomic conditions and individual health include employment and financial strain; higher chances of unemployment and reduced income may raise health risks, especially among socioeconomically disadvantaged individuals.[5–8] Studies also observed the widening socioeconomic inequalities regarding health during the economic crisis, indicating that economic shocks have a stronger impact on socioeconomically disadvantaged individuals.[9–11]

Most preceding studies have compared population health during the pre- and post-crisis periods or focused on changes in population health during the crisis,[12,13] leaving the lingering impact of the economic crisis largely understudied. Difficulties in finding good jobs during the recession upon graduating from college and high school may have long-lasting adverse impacts on health in later life. The impact of an unsuccessful start of working life is likely serious and lasting in a society that provides limited chances to move to better jobs.[14–16] In such cases, comparing health outcomes of the cohort that started its working life during the economic crisis with those of other cohorts will provide new insights about the impact of macroeconomic conditions on population health.

The current study aims to examine the cohort-specific lingering impact of the recession in Japan, using repeated cross-sectional data. Japan experienced the "employment ice age" (EIA) during 1993–2004, when young people had difficulty in finding jobs after graduation and started their working lives under unfavorable wage conditions. The job openings-to-applicants ratio for those aged 20–24 years was less than one during this period,[17] and the growth in average starting salary for new college graduates dropped sharply to about 0.4% per year during 1993–2004 from 4.2% during the previous 12 years.[18] We focused on health outcomes of the cohort that entered labor force during the EIA.

Address for correspondence. Takashi Oshio, Institute of Economic Research, Hitotsubashi University, 2-1 Naka, Kunitachi, Tokyo 186-8603, Japan (e-mail: oshio@ier.hit-u.ac.jp).

This cohort, referred to as the "EIA cohort" hereafter, was born during 1970–1985 based on year of graduation from college or high school. Studies on labor economics provided evidence that this cohort had more limited chances of becoming full-time, regular workers and obtained lower wages than other cohorts.[19] Hence, it can be hypothesized that the EIA cohort faces poorer health outcomes in working life.

## MATERIALS AND METHODS

### Study sample
We used a large dataset obtained from the "Comprehensive Survey of Living Conditions" (CSLC), a nationwide population-based survey conducted by the Japanese government's Ministry of Health, Labour and Welfare (MHLW). The CSLC, conducted since 1986, comprises an annual household survey and a triennial health and income/savings survey. Samples of the CSLC are collected nationwide through a two-stage random sampling procedure. First, about 5,400 districts are selected randomly from about 940,000 national census districts. Second, about 290,000 households are selected randomly from each selected district, according to its population size.

We used the data collected from each of eleven waves of the CSLC conducted from 1986 through 2016. We restricted the study sample to individuals aged between 30 years (when most individuals have already started working after graduating from college or high school) and 59 years (1 year prior to the most common mandatory retirement age of 60 years) to assess the impact of the experience of EIA on later working life. After excluding respondents for whom essential information was missing, this study used data on 3,054,782 individuals (1,500,618 men and 1,554,164 women). Table 1 presents the numbers of households and their members (individuals) who responded to each survey, the response rates at the household level, and the number of individual participants in this study.

We obtained the CSLC data with permission from the MHLW. The CSLC was authorized by the Ministry of Internal Affairs and Communications, which is in charge of all government surveys in Japan, from the statistical, legal, ethical, and other viewpoints in accordance with the Statistics Law in Japan. Hence, ethics approval was not required for the current study.

### EIA and the EIA cohort
The EIA is defined as the period from 1993 through 2004 by the Cabinet Office of the Japanese Government.[20] The EIA cohort, which started its working life in this period, was born during 1970–1985 in most cases, depending on the year of graduation from college or high school. The structure of the study sample is illustrated in Figure 1. The shadowed area indicates the appearance of the EIA cohort in the study sample. This cohort did not appear during waves 1–5 (1986–1998), but entered the study sample during wave 6 (2001), when individuals born in 1970 and 1971 were 31 and 30 years old, respectively. The number of individuals in the EIA cohort has been gradually increasing since then. By wave 11 (2016), the entire EIA cohort had entered the study sample, with members' ages ranging from 31 to 46 years. In this wave, those aged 47 years or above (born before 1969) belonged to the pre-EIA cohort, and those aged 31 (born in 1986) were in the post-EIA cohort.

As seen in Figure 1, the EIA cohort was relatively young in the study sample and had been increasing its share in more recent waves. In the entire study sample, the EIA cohort consisted of 417,988 individuals, accounting for 13.7% of the total. In wave 11, the EIA cohort consisted of 109,467 individuals, accounting for 52.5% of the sample in this wave.

### Measures
#### Health outcomes
We considered three health outcomes—being in hospital, subjective symptoms, and SRH—all of which were obtained from responses in the CSLC. The CSLC asked the respondents whether they were in hospital during the survey period. We constructed a binary variable for being in hospital, allocating 1 to the respondents who answered yes and 0 otherwise. Those who were not hospitalized were questioned about the other two outcomes; in other words, those in hospital were excluded from the analysis of these outcomes. As for subjective symptoms, the survey asked, "Have you been feeling ill due to sickness or injury [subjective symptoms] for the past few days?" We constructed a binary variable for having any subjective symptoms, allocating 1 to those who answered yes to this question and 0 otherwise. Regarding SRH, the respondents were asked, "What is your current health status? Is it excellent, very good, good, fair, or poor?" We constructed two types of binary variables for SRH: (i) allocating 1 to *fair* or *poor* and 0 otherwise or (ii) allocating 1 to *poor* and 0 otherwise.

**Table 1.** Numbers of respondents in the Comprehensive Survey of Living Conditions (CSLC) and of individual participants in this study

| Wave | Survey year | Number of CSLC respondents (households) | Response rate for households (%) | Number of CSLC respondents (individuals) | Individuals used in this study | | |
|---|---|---|---|---|---|---|---|
| | | | | | Total | Men | Women |
| 1 | 1986 | 240,283 | 95.7 | 803,807 | 344,025 | 169,048 | 174,977 |
| 2 | 1989 | 250,609 | 93.1 | 803,228 | 339,061 | 167,073 | 171,988 |
| 3 | 1992 | 253,653 | 92.9 | 783,095 | 325,783 | 160,047 | 165,736 |
| 4 | 1995 | 247,229 | 91.0 | 746,592 | 305,708 | 150,258 | 155,450 |
| 5 | 1998 | 247,882 | 89.7 | 721,478 | 292,052 | 143,859 | 148,193 |
| 6 | 2001 | 247,278 | 87.4 | 703,399 | 283,842 | 139,734 | 144,108 |
| 7 | 2004 | 220,948 | 79.9 | 619,573 | 246,983 | 120,978 | 126,005 |
| 8 | 2007 | 230,596 | 80.1 | 624,171 | 249,800 | 122,409 | 127,391 |
| 9 | 2010 | 229,785 | 79.4 | 609,018 | 232,498 | 113,957 | 118,541 |
| 10 | 2013 | 235,012 | 79.6 | 603,211 | 226,596 | 110,800 | 115,796 |
| 11 | 2016 | 224,641 | 77.6 | 568,425 | 208,434 | 102,455 | 105,979 |
| Total | | 2,627,916 | — | 7,585,997 | 3,054,782 | 1,500,618 | 1,554,164 |

This study was restricted to individuals aged 30–59 years for whom essential information was available.

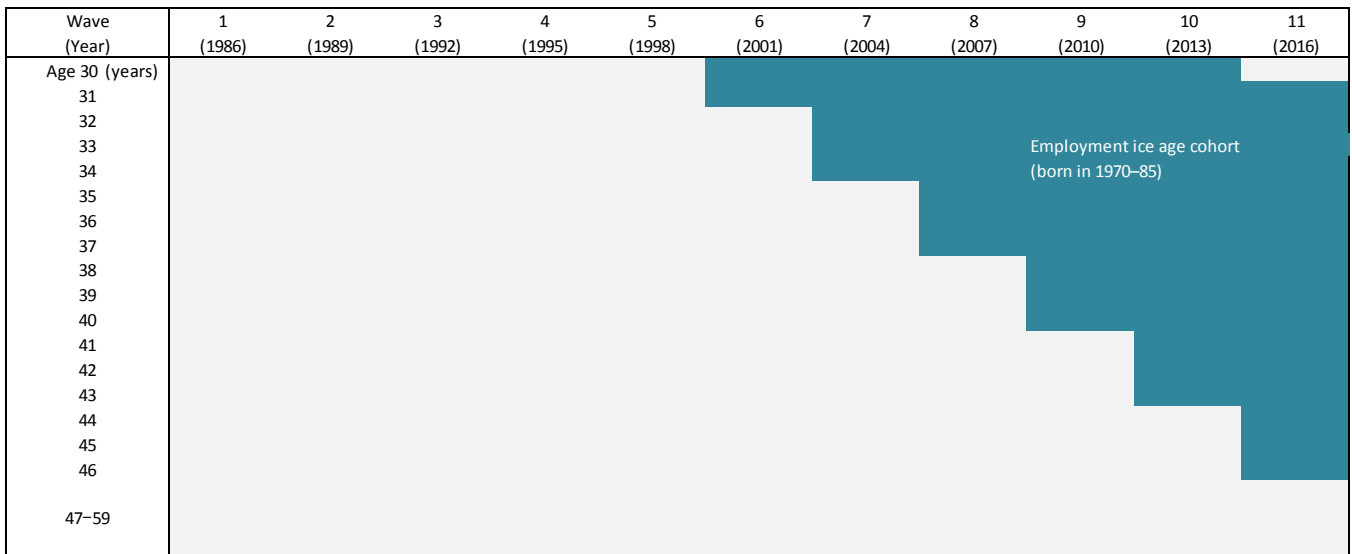| Wave | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 | 9 | 10 | 11 |
|---|---|---|---|---|---|---|---|---|---|---|---|
| (Year) | (1986) | (1989) | (1992) | (1995) | (1998) | (2001) | (2004) | (2007) | (2010) | (2013) | (2016) |



**Figure 1.** Appearance of the employment-ice-age cohort in the study sample

*Covariates*

We considered four covariates: having a spouse, having a paid job, household spending, and regional areas. For household spending, which was used as a proxy for household income, we first adjusted its reported value by the consumer price index (with the base year of 2015).[21] Then, we adjusted it for household size by dividing it by the square root of the number of household members[22] and categorized it into quartiles and binary variables of each quartile. We also constructed a binary variable of those who did not answer the question regarding household spending. For regional areas, we constructed binary variables for each of the eight areas (Hokkaido, Tohoku, Kanto, Chubu, Kinki, Chugoku, Shikoku, and Kyushu).

**Statistical analysis**

We utilized two approaches to examine how health outcomes differed between the EIA cohort and other cohorts. First, we conducted the age-period-cohort (APC) analysis to capture the cohort effect.[23] Specifically, we estimated logistic regression models to explain the binary variable of each health outcome, controlling for binary variables for each age and wave. The so-called identification problem, which is caused by the linear identity (age = period [wave] − cohort) and has been a key issue in the APC analysis,[24,25] could be largely avoided in this study. This is because using repeated cross-sectional data, we considered binary variables, rather than continuous ones, for ages, waves, and cohorts with different year intervals—1 year for ages, 3 years for waves, and only two cohorts (EIA cohort vs other cohorts).[23] We estimated two types of logistic models—with and without covariates (marital and work statuses, household spending, and regional areas)—for each health outcome, separately for male and female participants. Specifically, denoting the probability of each health outcome by $p$, we estimated the logistic regression model for individual $i$:

$$\log(p_i/(1-p_i)) = constant + \beta EIA\ cohort_i$$
$$+ \sum_{a=30}^{59} \gamma_a age_{ai} + \sum_{w=1}^{11} \delta_w wave_{wi} + (covariates_i) + \varepsilon_i$$

where *EIA cohort*, *age*, and *wave* indicate the binary variables of the EIA cohort, each age, and each wave, respectively, and $\varepsilon$ is an

error term. Under the framework of mediation analysis,[26,27] the difference in the estimated coefficient of the EIA cohort between the models with and without covariates is expected to reflect the EIA's impact mediated by sociodemographic and socioeconomic factors.

Second, we conducted a difference-in-differences (DiD) analysis.[28,29] We compared changes in health outcomes during the pre- and post-EIA periods between the two age groups. Based on Figure 2, we chose waves 1–5 (ie, 1986–1998) and wave 11 (2016) as the pre-EIA and post-EIA periods, respectively. The EIA cohort did not appear in waves 1–5 at all, whereas it fully appeared in wave 11. As the two age groups, we chose individuals aged 31–46 years and others (ie, aged 30 or 47–59 years). Individuals aged 31–46 years supposedly faced the EIA if they appeared in wave 11 but did not if they appeared in waves 1–5. In contrast, individuals aged 30 or 47–59 years, regardless of whether or not they appeared in waves 1–5 or wave 11, are considered to have never faced such an experience. We thus treated those aged 31–46 years as the treatment group and others as the control group. We did not use the data of waves 6–10, in which the age group of 31–46 years included both the EIA and other cohorts, as seen in Figure 2.

To compare changes in health outcomes during the pre- to the post-EIA periods between two age groups, we estimated the logistic regression model to explain health outcomes by using (i) a binary variable of those aged 31–46 years, (ii) a binary variable of wave 11, and (iii) their interaction, with/without covariates separately for men and women. We also controlled for each age- and wave-specific factor by including binary variables for each age and wave in the regression models. Specifically, we estimated the following logistic regression model:

$$\log(p_i/(1-p_i)) = constant + \beta_1 Aged\ 31\text{–}46\ years_i$$
$$+ \beta_2 Wave\ 11 + \beta_3 Aged\ 31\text{–}46\ years_i \times Wave\ 11$$
$$+ \sum_{a=30}^{59} \gamma_a age_{ai} + \sum_{w=1}^{4} \delta_w wave_{wi} + (covariates_i) + \varepsilon_i$$

where the second to forth terms on the right-hand side correspond to (i), (ii), and (iii), respectively, and $\varepsilon$ is an error term. The estimated coefficient of the interaction term ($\beta_3$) is expected to indicate the EIA's impact. This type of regression model
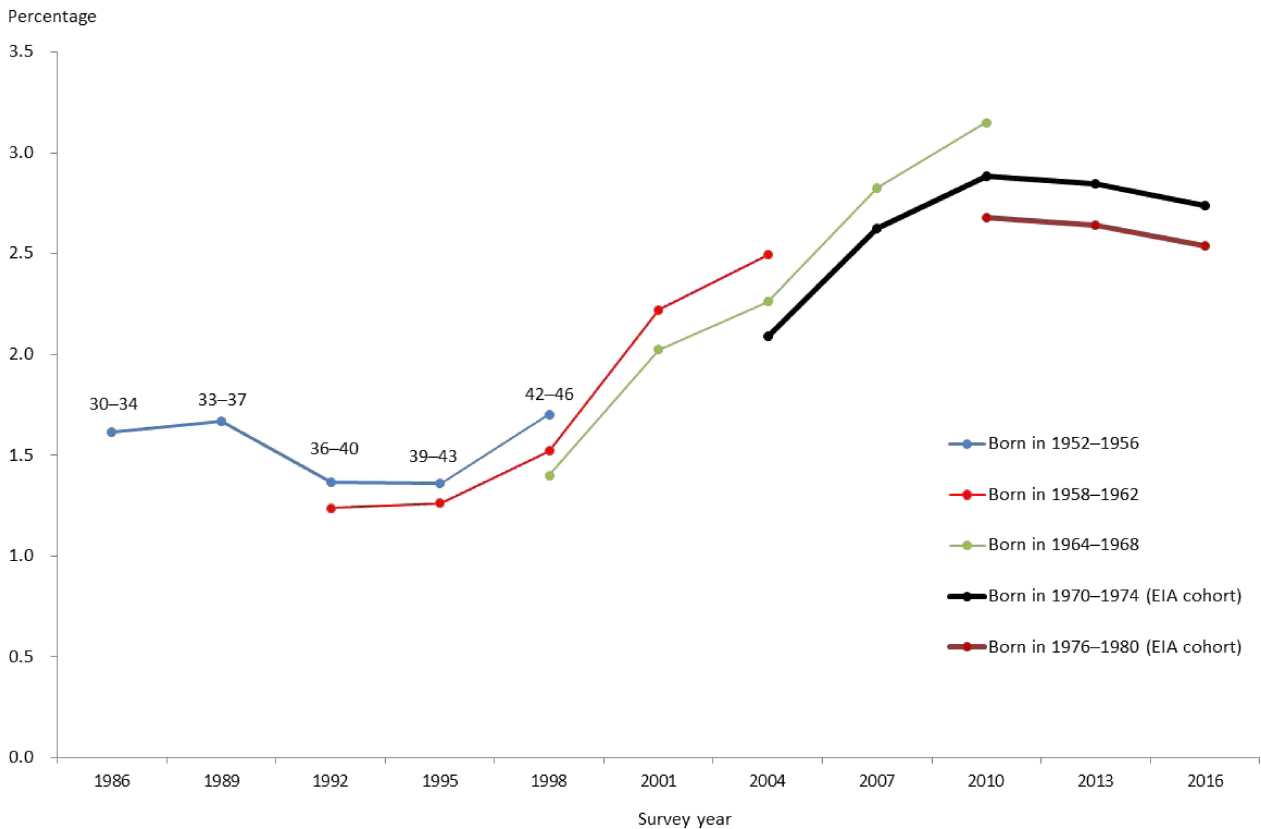
**Figure 2.** Age evolution of the prevalence of poor self-rated health by cohort: men. The dots in each curve indicate age (year) bands of 30–34, 33–37, 36–40, 39–43, and 42–46 years, respectively. Only the first three age bands for the cohort born in 1976–1980. EIA, employment ice age.

specification has been often used in DiD analysis, including in recent studies in Japan.[30,31] To adjust for the potential bias due to the difference in age contributions between the pre-EIA (wave 1–5) and post-EIA (wave 11) periods, we weighted the data of each age in wave 11 with the ratio of the share of respondents of that age in wave 1–5 to the share of respondents of the same age in wave 11 in the DiD regression models.

This DiD approach assumes that if there is no EIA, health outcomes would change from the pre-EIA to the post-EIA period in parallel ways, resulting in a non-significant coefficient of the interaction term.[28,29] To evaluate the validity of this equal trends assumption, we compared the prevalence trends of each health outcome between the control and treatment groups before the EIA. The DiD approach with cross-sectional data additionally assumes that whether an respondent is observed before or after the treatment is independent of his/her observed outcome, given that he/she belongs to the treatment or control group.[32] This additional assumption clearly held in this study, because the treatment and control groups were distinguished solely by age in each wave.

## RESULTS

### Descriptive analysis

Table 2 summarizes key features of all the respondents in this study. Note that the figures in the table are sample means not adjusted for ages. The propositions of those having a spouse and a paid job and belonging to higher quartiles of household spending are somewhat lower for the EIA cohort for both genders, probably reflecting the younger ages of its members. The

**Table 2.** Key features of the study sample

| | | Men | | | Women | | |
|---|---|---|---|---|---|---|---|
| | | EIA cohort | Other cohorts | Total | EIA cohort | Other cohorts | Total |
| Age, years | M | 36.0 | 46.2 | 44.8 | 36.0 | 46.2 | 44.8 |
| | SD | (4.3) | (8.1) | (8.5) | (4.3) | (8.2) | (8.5) |
| Proposition, % | | | | | | | |
| Having a spouse | | 60.6 | 82.5 | 79.5 | 68.1 | 83.8 | 81.7 |
| Having a paid job | | 91.5 | 94.6 | 94.1 | 52.3 | 55.9 | 55.4 |
| Household spending | | | | | | | |
|   1st quartile (lowest) | | 25.5 | 20.0 | 20.8 | 26.6 | 19.7 | 20.6 |
|   2nd quartile | | 23.3 | 22.8 | 22.9 | 23.5 | 22.6 | 22.8 |
|   3rd quartile | | 23.8 | 24.2 | 24.2 | 24.0 | 24.4 | 24.3 |
|   4th quartile | | 20.7 | 25.9 | 25.2 | 19.2 | 26.4 | 25.4 |
|   Not answered | | 6.8 | 7.0 | 7.0 | 6.8 | 6.9 | 6.9 |
| Health outcome | | | | | | | |
|   Being in hospital | | 1.0 | 1.1 | 1.1 | 0.9 | 0.9 | 0.9 |
|   Any subjective symptom | | 26.4 | 26.9 | 26.8 | 28.8 | 32.2 | 31.8 |
|   Self-rated health (poor or fair) | | 10.0 | 10.5 | 10.4 | 9.3 | 11.7 | 11.4 |
|   Self-rated health (poor) | | 1.2 | 1.1 | 1.1 | 0.9 | 1.0 | 1.0 |
| N | | 205,374 | 1,295,244 | 1,500,618 | 212,614 | 1,341,550 | 1,554,164 |

EIA, employment ice age.
Figures other than age were not adjusted for age.

prevalence of each poorer health outcome was somewhat lower for the EIA cohort than for other cohorts, again possibly reflecting the younger ages of its members. The prevalence of subjective symptom was somewhat higher among women, but other health outcomes did not show a substantial gender difference.

Figure 3 and Figure 4 illustratively compare the age evolution of the prevalence of poor SRH across five cohorts for men and
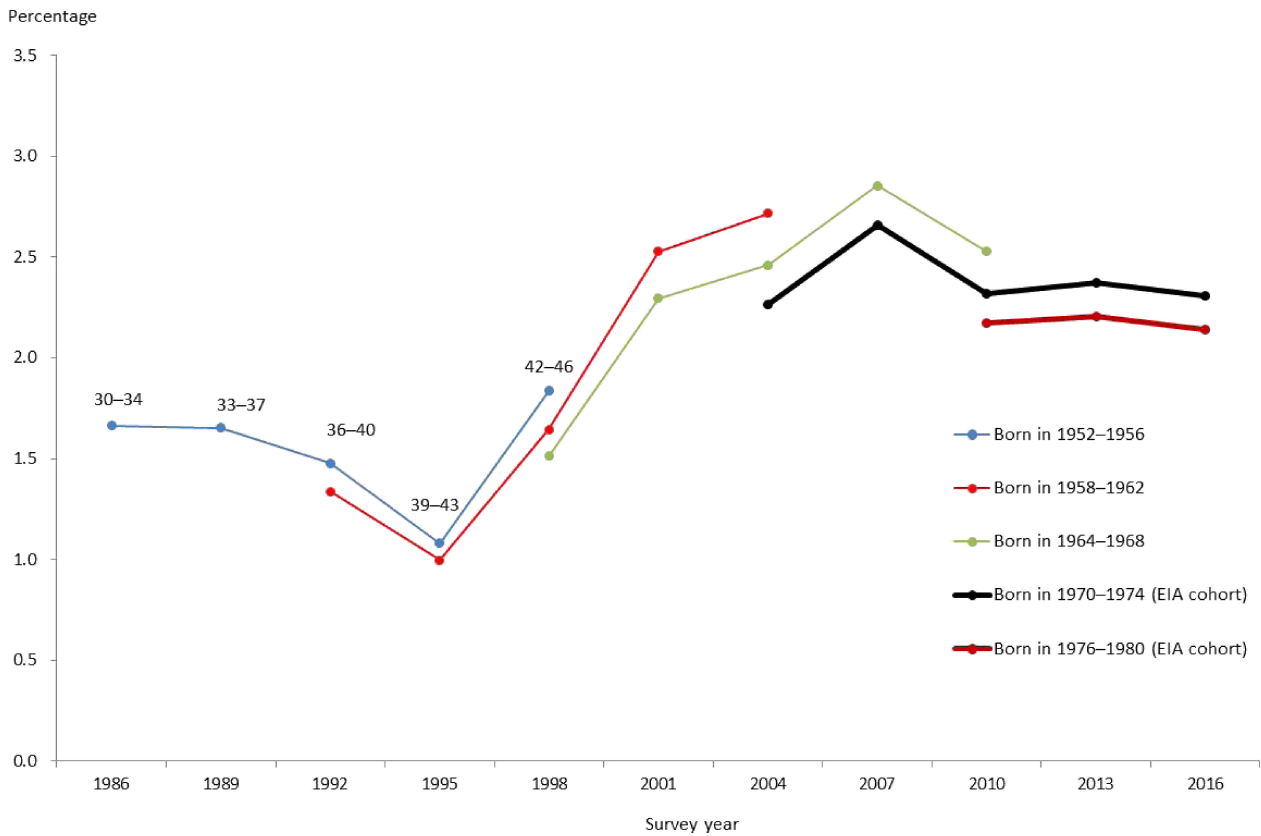
**Figure 3.** Age evolution of the prevalence of poor self-rated health by cohort: women. The dots in each curve indicate age (year) bands of 30–34, 33–37, 36–40, 39–43, and 42–46 years, respectively. Only the first three age bands for the cohort born in 1976–1980. EIA, employment ice age.
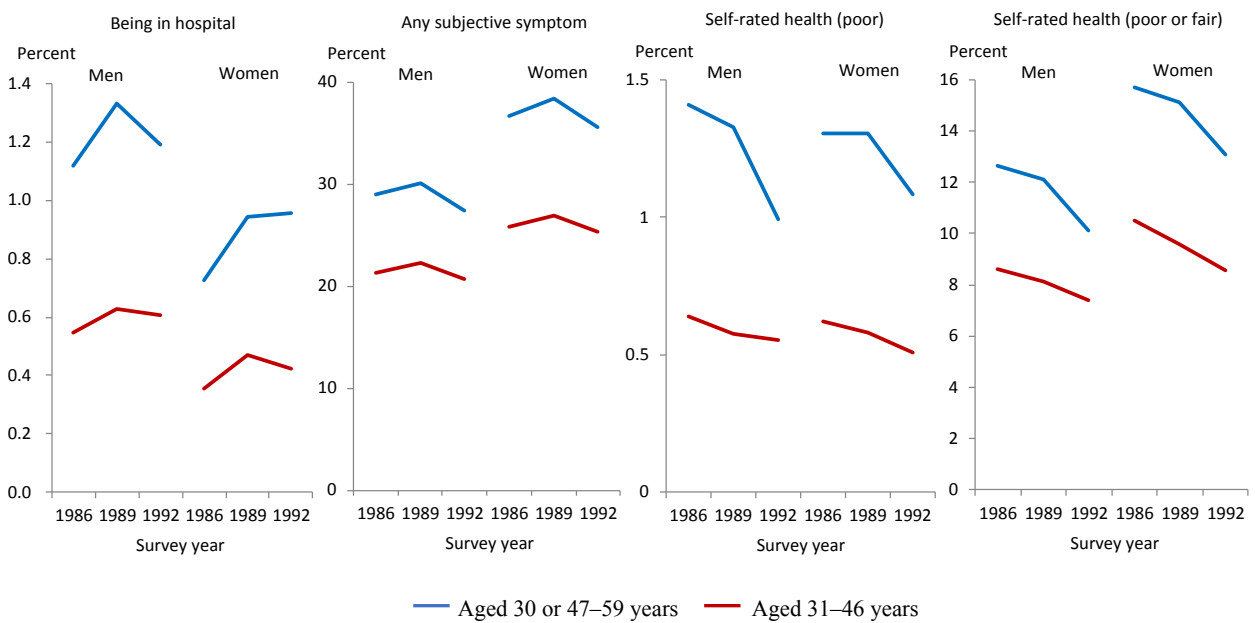


**Figure 4.** Prevalence trends of "any subjective symptom" and "poor self-rated health" before the employment ice age

women, respectively, focusing on the age range of 30–46 years (30–40 years for the cohort born in 1976–80). The two youngest cohorts were born during 1970–74 and 1976–80, respectively, and started their working life during the EIA. The three older cohorts, who were born during 1952–1956, 1958–1962, and 1964–1968, started their working life before the EIA.

Figure 3 for men shows that the prevalence of poor SRH has been rising after bottoming out in the early 1990s through variations across cohorts, probably reflecting the changes in macroeconomic conditions; the trend of the unemployment rate turned from downward to upward, and the wage growth rate declined sharply in the early 1990s. By cohort, the prevalence of

**Table 3.** Estimated odds ratios of each health outcome for the employment-ice-age (EIA) cohort, after controlling for age and wave effects

| | Not controlling for covariates[a] | | Controlling for covariates | |
|---|---|---|---|---|
| | OR | 95% CI | OR | 95% CI |
| Men (N = 1,500,618) | | | | |
| Being in hospital | 1.29 | (1.21, 1.38) | 1.29 | (1.20, 1.37) |
| Any subjective symptom | 1.20 | (1.18, 1.21) | 1.18 | (1.17, 1.20) |
| Self-rated health (poor or fair) | 1.06 | (1.04, 1.09) | 1.04 | (1.02, 1.07) |
| Self-rated health (poor) | 1.29 | (1.21, 1.38) | 1.25 | (1.17, 1.33) |
| Women (N = 1,554,164) | | | | |
| Being in hospital | 1.19 | (1.12, 1.28) | 1.15 | (1.08, 1.23) |
| Any subjective symptom | 1.22 | (1.21, 1.24) | 1.21 | (1.19, 1.22) |
| Self-rated health (poor or fair) | 1.07 | (1.04, 1.09) | 1.02 | (1.00, 1.05) |
| Self-rated health (poor) | 1.25 | (1.17, 1.34) | 1.16 | (1.08, 1.24) |

CI, confidence interval; OR, odds ratio.
[a]The covariates include having a spouse and paid job, household spending, and regional areas.

**Table 4.** Comparison of school-age (16–22 years) health outcomes between the employment-ice-age (EIA) and other cohorts[a]

| Health outcome prevalence (%) | EIA cohort (A) | Other cohort (B) | Difference (A − B) | P-value |
|---|---|---|---|---|
| Men (N = 239,519) | | | | |
| Being in hospital | 0.52 | 0.40 | 0.12 | <0.001 |
| Any subjective symptom | 15.04 | 14.23 | 0.80 | <0.001 |
| Self-rated health (poor or fair) | 4.21 | 5.37 | −1.16 | <0.001 |
| Self-rated health (poor) | 0.40 | 0.52 | −0.12 | 0.003 |
| Women (N = 234,816) | | | | |
| Being in hospital | 0.41 | 0.36 | 0.05 | 0.114 |
| Any subjective symptom | 18.99 | 18.39 | 0.60 | 0.002 |
| Self-rated health (poor or fair) | 5.27 | 6.88 | −1.61 | <0.001 |
| Self-rated health (poor) | 0.39 | 0.42 | −0.03 | 0.348 |

[a]The data relate to respondents born during 1926–1986.

poor SRH for the two EIA cohorts started at high levels and, since then, stayed relatively high compared with the previous three cohorts. Figure 4 demonstrates similar results for women, showing higher prevalence of poor SRH for the two EIA cohorts, and this prevalence showed a clearer upturn for women than for men in 1995.

## Regression analysis

Table 3 summarizes the results of the regression models in the APC analysis, focusing on the estimated odds ratios (ORs) of each health outcome for the EIA cohort after controlling for ages and waves. The table compares the results with and without controlling for covariates. Based on the ORs, we found that the EIA cohort faced poorer health outcomes compared with other cohorts, for both men and women, regardless of whether or not covariates were controlled for. For instance, the OR of poor SRH for the male EIA cohort was 1.29 (95% confidence interval [CI], 1.21–1.38) without control for covariates and 1.25 (95% CI, 1.17–1.33) with control for covariates, while it was 1.25 (95% CI, 1.17–1.34) and 1.16 (95% CI, 1.08–1.24), respectively, for women. The ORs of poor/fair SRH were lower than those of other outcomes, but were significantly higher than one.

Another noticeable finding here is that the ORs were not much attenuated after controlling for covariates. Although not reported to conserve space, the regression analysis found that the EIA cohort had lower chances of having a spouse and paid job and higher household spending, all of which were negatively associated with poorer health outcomes. These results point to the possibility that sociodemographic and socioeconomic factors mediated the impact of EIA experience on health in later life. However, the results in Table 3 suggest that such mediating effects, if any, were relatively limited, especially among men.

We should be cautious in interpreting the results in Table 3 because the EIA cohort may have worse health outcomes due to some events that affected the individuals before they started their working lives. To assess this potential bias, Table 4 compares each health outcome during school age (16–22 years) between the EIA and other cohorts for both men and women. Out of eight gender and health outcome combinations, three yielded worse outcomes and three better for the EIA cohort, while the outcomes of two combinations were not significantly different between the

EIA and other cohorts. Table 3 presents the estimated EIA effects on being in hospital for men and on any subjective symptom for both genders; these effects are relatively substantial, but are probably somewhat overestimated.

Table 5 shows key results of the DiD analysis, presenting the estimated ORs of binary variables of individuals aged 31–46 years, wave 11, and their interaction, with and without controlling for covariates. The ORs of individuals aged 31–46 years were all well below one in all models, a reasonable result considering that those individuals were relatively young. The ORs of wave 11 were mixed, suggesting that the direction of wave effects depended on the type of health outcomes. More importantly, ORs of the interaction between individuals aged 31–46 years and wave 11 were all above one for all health outcomes for both men and women, before controlling for covariates. This result indicated that the EIA cohort had poorer health outcomes compared with the other cohorts. ORs of the interaction declined after controlling for covariates but remained above one for all health outcomes among men, while they declined to below one for being in hospital and poor/fair SRH among women.

Figure 4 compares the prevalence trends of each health outcome between the control and treatment groups before the EIA. The prevalence of each health outcome moved largely in tandem between the control and treatment groups in the pre-EIA waves, though to a greater or lesser degree depending on the gender and heath outcome combination, suggesting that the equal trends assumption largely held in this study sample.

## DISCUSSION

We examined whether the health outcomes differed between the cohorts who started working life during the recession and the other cohorts. Specifically, we focused on the cohort that experienced adverse macroeconomic conditions during 1994–2003 (the EIA) when they started their working life in Japan. The results in this study showed that starting working life during the recession had a long-lasting impact on health outcomes for both men and women. The APC analysis showed that the EIA cohort had poorer health outcomes, even after controlling for age and cohort effects. This observation was underscored by the DiD analysis, which treated the experience of EIA as a treatment implemented on those aged 31–46 years.

The results of the current study are in line with previous observations about the impact of economic recessions on

**Table 5.** Relationship of health outcomes with "aged 31–46 years," "wave 11," and their interaction: a difference-in-differences analysis[a]

| | Aged 31–46 years[b] | | Wave 11 | | Aged 31–46 years × Wave 11 | |
|---|---|---|---|---|---|---|
| | OR | 95% CI | OR | 95% CI | OR | 95% CI |
| Not controlling for covariates[c] | | | | | | |
| Men (N = 892,740) | | | | | | |
| Being in hospital | 0.46 | (0.40, 0.53) | 1.15 | (1.04, 1.27) | 1.43 | (1.26, 1.62) |
| Any subjective symptom | 0.63 | (0.61, 0.65) | 1.02 | (0.99, 1.04) | 1.23 | (1.19, 1.27) |
| Self-rated health (poor or fair) | 0.63 | (0.59, 0.66) | 1.26 | (1.22, 1.30) | 1.07 | (1.02, 1.11) |
| Self-rated health (poor) | 0.49 | (0.42, 0.58) | 1.62 | (1.48, 1.78) | 1.39 | (1.24, 1.56) |
| Women (N = 922,323) | | | | | | |
| Being in hospital | 0.53 | (0.45, 0.62) | 1.52 | (1.37, 1.68) | 1.16 | (1.02, 1.32) |
| Any subjective symptom | 0.62 | (0.60, 0.64) | 0.74 | (0.72, 0.75) | 1.28 | (1.24, 1.31) |
| Self-rated health (poor or fair) | 0.69 | (0.66, 0.72) | 0.90 | (0.87, 0.93) | 1.06 | (1.01, 1.11) |
| Self-rated health (poor) | 0.46 | (0.39, 0.55) | 1.27 | (1.15, 1.41) | 1.42 | (1.24, 1.62) |
| Controlling for covariates | | | | | | |
| Men (N = 892,740) | | | | | | |
| Being in hospital | 0.56 | (0.48, 0.65) | 0.94 | (0.85, 1.04) | 1.40 | (1.24, 1.59) |
| Any subjective symptom | 0.67 | (0.64, 0.69) | 1.01 | (0.99, 1.03) | 1.22 | (1.18, 1.26) |
| Self-rated health (poor or fair) | 0.70 | (0.66, 0.73) | 0.96 | (0.92, 0.99) | 1.05 | (1.00, 1.09) |
| Self-rated health (poor) | 0.66 | (0.56, 0.77) | 0.96 | (0.87, 1.05) | 1.33 | (1.18, 1.49) |
| Women (N = 922,323) | | | | | | |
| Being in hospital | 0.65 | (0.55, 0.77) | 1.88 | (1.69, 2.08) | 1.06 | (0.93, 1.21) |
| Any subjective symptom | 0.64 | (0.61, 0.66) | 0.78 | (0.76, 0.80) | 1.26 | (1.23, 1.30) |
| Self-rated health (poor or fair) | 0.75 | (0.71, 0.79) | 0.72 | (0.70, 0.74) | 1.01 | (0.97, 1.05) |
| Self-rated health (poor) | 0.60 | (0.51, 0.70) | 0.95 | (0.86, 1.04) | 1.25 | (1.09, 1.42) |

CI, confidence interval; OR, odds ratio.

[a]The estimates are adjusted for differences in age distributions between those aged 31–46 years (treatment group) and others (control group).

[b]This age group supposedly faced the employment ice age at labor market entry only if they appeared in wave 11 (in 2016).

[c]The covariates include having a spouse and paid job, household spending, and regional areas.

population health,[5–8] as well as the disadvantage of the EIA cohort in terms of job status and wage earnings in Japan.[20] The results are also consistent with previous observations in Japan,[33] which confirmed that starting working life with a precarious job status tends to result in lower SES and poorer health outcomes. This implies that these individuals have socio-institutional backgrounds that generally limit chances to move to better jobs after entering the labor market as precarious employees.[34] The limited mediating effects of sociodemographic and socio-economic factors observed in this study, especially among men, may reflect the general perception that failure at the start of one's working life is difficult to overcome and presumably has a traumatic and lingering impact on health. The observations of the current study did not show any significant gender differences before controlling for covariates, but results of the DiD analysis after controlling for them suggested that the traumatic impact of adverse economic conditions at labor market entry is more relevant for men.

We should be cautious in making any generalizations based on the observations in this study. The impact of macroeconomic conditions at the start of one's working life on health outcomes in later life must depend on socio-institutional backgrounds. The impact is likely to be more mixed in societies with more flexible job markets, wherein there are sufficient chances to find better jobs, even if the economic downturn generally increases financial strain and health risks, as suggested by preceding studies in European countries.[14–16]

We acknowledge that this study has several limitations, in addition to concerns about the reliability of self-reported health outcomes and ignorance of intra-cohort variations. Our statistical analysis relied on repeated cross-sectional data, which made it impossible to track the impact of the recession experience on health outcomes in later life at an individual level. We controlled for age and wave cohorts by including binary variables for each of them in the APC analysis but ignored the interaction effects among ages, waves, and cohorts. We argued for limited mediating effects of sociodemographic and socioeconomic factors, but experiencing an economic downturn may have other mediators for its impact on health outcomes.

Despite these limitations, this study highlighted the lingering impact of starting working life during a recession on health outcomes in Japan. This finding, though dependent on socio-institutional backgrounds, provides new insights about the association between macroeconomic conditions and population health. It also implies the need for structural reforms in the labor market from the viewpoint of public health in Japan.

## ACKNOWLEDGMENTS

## REFERENCES

1. Frasquilho D, Matos MG, Salonna F, et al. Mental health outcomes in times of economic recession: a systematic literature review. *BMC Public Health*. 2016;16:115.

2. Karanikolos M, Heino P, McKee M, Stuckler D, Legido-Quigley H. Effects of the Global Financial Crisis on health in high-income OECD countries: a narrative review. *Int J Health Serv*. 2016;46:208–240.

3. Karanikolos M, Mladovsky P, Cylus J, et al. Financial crisis,

austerity, and health in Europe. *Lancet*. 2013;381:1323–1331.

4. Parmar D, Stavropoulou C, Ioannidis JP. Health outcomes during the 2008 financial crisis in Europe: systematic literature review. *BMJ*. 2016;354:i4588.

5. Córdoba-Doña JA, Escolar-Pujolar A, San Sebastián M, Gustafsson PE. How are the employed and unemployed affected by the economic crisis in Spain? Educational inequalities, life conditions and mental health in a context of high unemployment. *BMC Public Health*. 2016;16:267.

6. Córdoba-Doña JA, San Sebastián M, Escolar-Pujolar A, Martínez-Faure JE, Gustafsson PE. Economic crisis and suicidal behaviour: the role of unemployment, sex and age in Andalusia, southern Spain. *Int J Equity Health*. 2014;13:55.

7. Riumallo-Herl C, Basu S, Stuckler D, Courtin E, Avendano M. Job loss, wealth and depression during the Great Recession in the USA and Europe. *Int J Epidemiol*. 2014;43:1508–1517.

8. Ruiz-Pérez I, Bermúdez-Tamayo C, Rodríguez-Barranco M. Socio-economic factors linked with mental health during the recession: a multilevel analysis. *Int J Equity Health*. 2017;16:45.

9. Bacigalupe A, Escolar-Pujolar A. The impact of economic crises on social inequalities in health: what do we know so far? *Int J Equity Health*. 2014;13:52.

10. Bacigalupe A, Esnaola S, Martín U. The impact of the Great Recession on mental health and its inequalities: the case of a Southern European region, 1997–2013. *Int J Equity Health*. 2016; 15:17.

11. Heggebø K, Tøge AG, Dahl E, Berg JE. Socioeconomic inequalities in health during the Great Recession: a scoping review of the research literature. *Scand J Public Health*. 2019;47:635–654.

12. Katikireddi SV, Niedzwiedz CL, Popham F. Trends in population mental health before and after the 2008 recession: a repeat cross-sectional analysis of the 1991–2010 Health Surveys of England. *BMJ Open*. 2012;2:e001790.

13. Thomson RM, Niedzwiedz CL, Katikireddi SV. Trends in gender and economic inequalities in mental health following the Great Recession and subsequent austerity policies: a repeat cross-sectional analysis of the Health Surveys for England. *BMJ Open*. 2018;8: e022924.

14. Gebel M. Early career consequences of temporary employment in Germany and the UK. *Work Employ Soc*. 2010;24:641–660.

15. McGinnity F, Mertens A, Gundert S. A bad start? Fixed-term contracts and the transition from education to work in West Germany. *Eur Sociol Rev*. 2005;21:359–374.

16. Steijn B, Need A, Gesthuizen M. Well begun, half done? Long-term effects of labour market entry in the Netherlands, 1950–2000. *Work Employ Soc*. 2006;20:453–472.

17. Ministry of Health, Labour and Welfare (MHLW). Employment referrals for general workers. https://www.mhlw.go.jp/english/

database/db-l/general_workers.html. 2019. Accessed 19.05.18.

18. MHLW. Starting salary for new graduates. https://www.mhlw.go.jp/english/database/db-l/starting-salary.html; 2019. Accessed 19.05.18.

19. Genda Y, Kondo A, Ohta S. Long-term effects of a recession at labor market entry in Japan and the United States. *J Hum Resour*. 2010; 45:157–196.

20. Cabinet Office. The 5th meeting of the Council on Economic and Fiscal Policy. https://www5.cao.go.jp/keizai-shimon/kaigi/minutes/2019/0410/agenda.html; 2019 Accessed 19.05.18.

21. Ministry of Internal Affairs and Communications. Consumer price index. https://www.stat.go.jp/english/data/cpi/index.html; 2019. Accessed 19.05.18.

22. Organisation for Economic Co-operation and Development, 2011. *Divided we stand: why inequality keeps rising*. Paris: OECD; 2011.

23. Yang Y, Land KC. *Age-Period-Cohort Analysis: New Models, Methods, and Empirical applications*. Boca Raton, FL: CRC Press; 2013.

24. Bell A, Jones K. The impossibility of separating age, period and cohort effects. *Soc Sci Med*. 2013;93:163–165.

25. Reither EN, Masters RK, Yang YC, Powers DA, Zheng H, Land KC. Should age-period-cohort studies return to the methodologies of the 1970s? *Soc Sci Med*. 2015;128:356–365.

26. MacKinnon DP. *Introduction to Statistical Mediation Analysis*. New York, NY: Erlbaum; 2008.

27. Mackinnon DP, Dwyer JH. Estimating mediated effects in prevention studies. *Eval Rev*. 1993;17:144–158.

28. Dimick JB, Ryan AM. Methods for evaluating changes in health care policy: the difference-in-differences approach. *JAMA*. 2014;312:2401–2402.

29. Gertler PJ, Martinez S, Premand P, Rawlings LB, Vermeersch CMJ, editors. *Impact evaluation in practice*. second edition (Ch. 7). Washington, DC: World Bank Group; 2016:129–142.

30. Miyawaki A, Noguchi H, Kobayashi Y. Impact of medical subsidy disqualification on children's healthcare utilization: A difference-in-differences analysis from Japan. *Soc Sci Med*. 2017;191:89–98.

31. Sasabuchi Y, Matsui H, Yasunaga H, Fushimi K. Increase in avoidable hospital admissions after the Great East Japan Earthquake. *J Epidemiol Community Health*. 2017;71:248–252.

32. Lee M, Kang C. Identification for difference in differences with cross-section and panel data. *Econ Lett*. 2006;92:270–276.

33. Oshio T, Inagaki S. The direct and indirect effects of initial job status on midlife psychological distress in Japan: evidence from a mediation analysis. *Ind Health*. 2015;53:311–321.

34. Hamaaki J, Hori M, Maeda S, Murata K. How does the first job matter for an individual's career life in Japan? *J Jpn Int Econ*. 2013;29:154–169.