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Research Article

Decomposing delayed first marriage and birth across cohorts: The role of increased employment instability among men in Japan

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Contents

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Decomposing delayed first marriage and birth across cohorts: The role of increased employment instability among men in Japan

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Abstract

BACKGROUND

Increased employment instability over recent decades has been argued to contribute to delayed family formation. Although previous research has shown negative associations between employment instability – such as unemployment, nonstandard employment, or unstable employment trajectories – and entry into marriage and childbirth, direct evidence on how increased employment instability to delayed marriage and childbirth across cohorts remains limited.

OBJECTIVE

This study examines the extent to which cross-cohort changes in employment status and work experience account for delays in first marriage and first birth among men in Japan born between 1945 and 1984.

METHODS

Discrete-time hazard models are applied to nationally representative retrospective survey data from Japan. Additionally, age-specific survival rates are calculated based on these estimates.

RESULTS

The hazard model results indicate that compositional changes in employment-related factors account for 26% of the decline in first marriage rates and 27% of the decline in first birth rates between the earliest and latest cohorts. Counterfactual simulations further reveal that these factors account for 31% of the increase in the proportion of individuals who never married and 45% of the increase in the proportion of childless at age 30.

CONCLUSIONS

Increased employment instability has contributed to the delay in first marriage and first birth, supporting theoretical claims that structural changes in the economy and labor market have driven shifts in family formation behavior.

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CONTRIBUTION
This study provides evidence of the direct link between cross-cohort delays in family formation and increased exposure to precarious employment, which corroborates the arguments regarding employment uncertainty.

1. Introduction

Increased employment instability has been identified as an important source of the postponement of family formation. The timing of union entry and childbearing has been delayed since the 1970s in Europe and the United States, and more recently in East Asia (Goldstein and Kenney 2001; Goldstein, Sobotka, and Jasilioniene 2009; Kohler, Billari, and Ortega 2002; Raymo et al. 2015). This delay is suggested to result from increased employment instability associated with economic downturns and labor market deregulation in these countries (Kohler, Billari, and Ortega 2002; Mills and Blossfeld 2013; Mills, Blossfeld, and Klijzing 2005; Oppenheimer 1988, 1994). Many studies have examined the relationship between individual exposure to employment instability – measured as spells of unemployment, nonstandard employment, or cumulative years of unemployment or nonstandard employment – and the timing of union formation, marriage, or childbirth (Alderotti et al. 2021; Mills, Blossfeld, and Klijzing 2005; Sassler and Lichter 2020).

In contrast, few studies have investigated the extent to which increased employment instability over time has contributed to cross-cohort delays in family formation. Consequently, the validity of theoretical claims that increased employment instability is responsible for these delays remains unclear (Kohler, Billari, and Ortega 2002; Mills and Blossfeld 2013; Mills, Blossfeld, and Klijzing 2005; Oppenheimer 1988, 1994). Given the negative associations observed at the individual level between employment instability and entry into union, marriage, or childbirth, it is reasonable to expect that increased employment instability contributes, at least some part, to the delay. However, if its contribution is small, we should be cautious about emphasizing its impact. Thus, understanding the precise macro-level contribution of individuals' increased employment instability across cohorts to delayed family formation is essential for evaluating these theoretical claims.

Some studies analyze the relationship between changing labor market conditions and the decline in marriage or fertility rates using period measures. Schneider, Harknett, and Stimpson (2018) demonstrate that men's reduced earnings and increased risk of incarceration account for the decrease in marriage rate over time in the United States. Raymo and Shibata (2017) investigate the extent to which changes in individual and aggregate employment conditions have contributed to trends in the total fertility rate in Japan. Unlike those studies, this study focuses on a cohort comparison, which allows us to examine the extent to which individuals have postponed the timing of marriage or childbirth and how increased employment instability has contributed to the delays.

I examine how increased employment instability across cohorts has contributed to delayed marriage and childbirth in Japan. Like many other high-income countries, Japan has experienced several decades of delayed marriage and childbirth, resulting in a persistently low total fertility rate of about 1.3 or less (Goldstein, Sobotka, and Jasilioniene 2009; Kohler, Billari, and Ortega 2002). In line with this trend, employment instability has increased significantly from the late 1990s to the 2000s, as reflected in the rise of nonstandard employment, unemployment, and inactivity among young people (Statistics Bureau of Japan 2023a). Meanwhile, cohabitation or nonmarital childbearing as alternative forms of family formation have yet to emerge (Mogi et al. 2023; Raymo, Iwasawa, and Bumpass 2009), nor have substantial shifts from traditional to individualistic values occurred across cohorts (Atoh 2001; Choe et al. 2014; Piotrowski et al. 2019) as expected in the second demographic transition (Lesthaeghe 1995; Raymo 2022; Van de Kaa 1987). In this context, increased employment instability is expected to account for part of the postponement of marriage and childbirth.

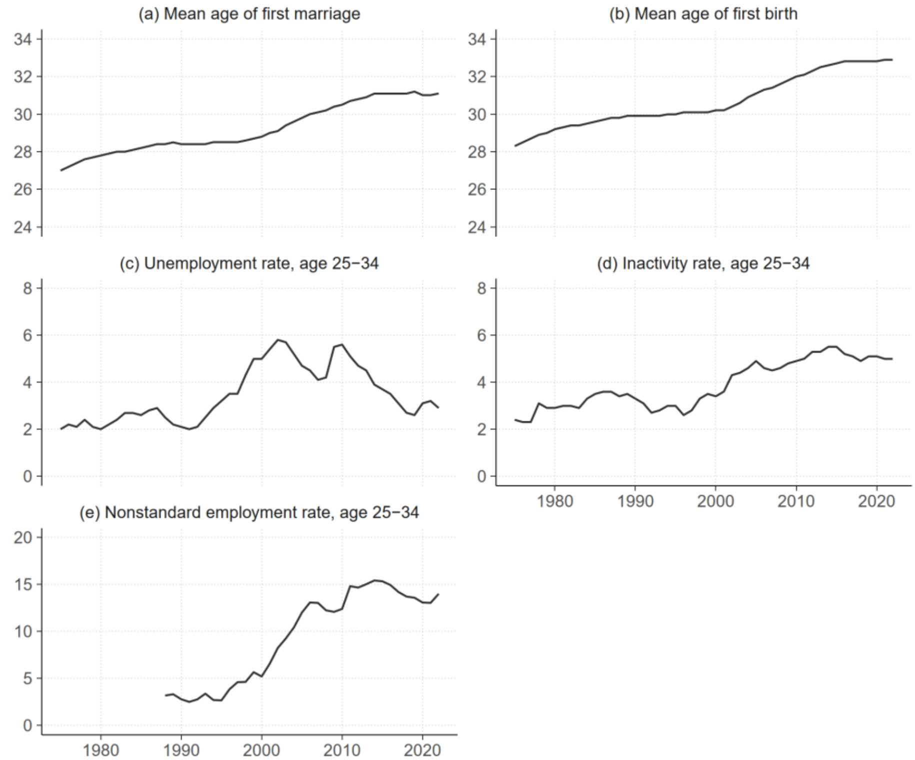
Specifically, this study investigates the extent to which the cross-cohort increase in men's exposure to employment instability – measured by nonstandard employment, nonemployment, and prior work experience – accounts for delayed first marriage and first birth among men born between 1945 and 1984, cohorts that simultaneously experienced increased employment instability alongside delayed marriage and childbirth. This study focuses on men since expectations for men's breadwinner roles remain strong, as reflected in the continued emphasis on their earning capacity (Brinton, Mun, and Hertog 2021; National Institute of Population and Social Security Research 2023), as well as the robust negative associations between nonstandard employment or nonemployment and marriage entry (Esteban-Pretel and Fujimoto 2022; Lim 2018; Matsuda and Sasaki 2020; Mizuocchi 2006; Mugiyama 2017; Piotrowski, Kalleberg, and Rindfuss 2015; Sakai and Higuchi 2005; Tsuya 2024). I focus on nonemployment, which includes both unemployment and inactivity, rather than unemployment alone because the rise in employment instability in Japan has been characterized by increases in both unemployment and inactive population, as discussed later. I apply discrete-time hazard models and counterfactual simulations to assess how cross-cohort changes in the composition of employment-related factors account for men's hazard of entering first marriage and first birth, as well as age-specific proportion of never having been married and childlessness. In doing so, this study provides evidence of the contribution of increased employment instability to cross-cohort delays in family formation.

Despite the importance of analyzing women, I restrict the target population to men for two reasons. First, the relationship between employment and the transition to first marriage or first birth among women has varied across cohorts (Fukuda 2013; Kimura 2022). Normative expectations for women have gradually shifted over decades from being primarily homemakers to balancing paid employment with housework and childcare (National Institute of Population and Social Security Research 2023), a shift described as "pro-work conservatism" (Brinton and Lee 2016). The changing relationship makes it difficult to assess the contribution of increased employment instability across cohorts. Second, the relationship is sensitive to the length of point at which employment is measured. In Japan, many women leave employment not only upon pregnancy and childbirth but also upon marriage, with the timing of employment exit shifting across cohorts (National Institute of Population and Social Security Research 2023). This complicates the relationship between employment and the transition to marriage and childbirth for women, as research has shown that the choice of lag years significantly affects the strength or direction of the association (Piotrowski, Kalleberg, and Rindfuss 2015). Given these complexities, I focus on the impact of increased employment instability among men on their delayed first marriage and first birth across cohorts.

2. The Japanese context

The timing of first marriage and first birth in Japan has been significantly delayed over the decades. Historically, the vast majority of people married at some point in their lifetime. In 1990, men typically married in their twenties (Brinton 1993), and only 5.6% never marry by age 50 (Raymo et al. 2015). Since then, both the average age at first marriage and the rate at which men never get married have increased. By 2015, the average age at first marriage for men had risen to 31.1 years (see Panel A in Figure 1) and the rate at which men never get married by age 50 had reached 24.2% (Statistics Bureau of Japan 2023b). Similarly, the age at first birth also increased (see Panel B), reflecting the trend of postponing childbearing (Mogi, Nisén, and Canudas-Romo 2021). Alternative partnership behaviors have not increased: Cohabitation remains uncommon and is typically brief before marriage (Mogi et al. 2023; Raymo, Iwasawa, and Bumpass 2009). Childbirth remains closely linked with marriage (Bumpass et al. 2009; Raymo et al. 2015), as evidenced by the fact that most births occur among married couples rather than unmarried ones (Raymo and Iwasawa 2008). As a result of this close linkage, the increase in the age at first birth has largely mirrored the rise in the age at first marriage (Iwasawa 2008; Jones 2007).

Figure 1: Trends in delayed first marriage and first birth and increasing unemployment, inactivity, and nonstandard employment among men in Japan, 1975–2022



Notes: Data in Panels A and B were retrieved from the Vital Statistics (Statistics Bureau of Japan 2023c). Data in Panels C, D, and E were from the Labor Force Survey (Statistics Bureau of Japan 2023a). The unemployment rate (%) and inactivity rate (%) are the annual average. Nonstandard employment rate (%) is calculated by the number of nonstandard employees divided by the total number of workers. February values are reported before 2001, and January–March averages are reported after 2002. Nonstandard employees include part-time, temporary, dispatched, contract, and entrusted employees as well as employees who are neither classified as regular employees nor as other employees. Note that the inclusion of the last category does not affect the trend (see Appendix Figure A-1).

Increasing employment instability has been identified as one of the primary sources of delayed marriage and childbirth (Raymo et al. 2015). Since the 1990s, the employment situation of young and middle-aged workers has worsened (Houseman and Osawa 2003; Kalleberg, Hewison, and Shin 2021). The unemployment rate for men aged 25 to 34 rose from about 2% in the 1980s to 6% in the 2000s (see Panel C in Figure 1). Furthermore, the population of inactive youths also increased (see Panel D). In the 2000s, the increase in young people not in employment, education, or training (NEETs) attracted public attention as a pressing youths employment issue. Unlike the original definition by the International Labour Organization, the Japanese version of NEET was defined as inactive individuals aged 15 to 34 who are neither primarily enrolled in school nor engaged in housework (Genda and Maganuma 2004; Kosugi 2004). Such inactive youths share similar socioeconomic backgrounds with the unemployed, and most of them had been employed in the past year (Genda 2007). Young people who have experienced at least one month of nonemployment frequently move between employment and nonemployment states over the course of their early careers, while permanent nonemployment is rare and more common among women who have had a child (Kagawa et al. 2022). These studies suggest that, in the Japanese context, nonemployment – including both unemployment and inactivity – is considered a marker of employment instability.

Furthermore, the rise of nonstandard employment is recognized as an increase in employment instability in Japan. As shown in Panel E in Figure 1, the share of men in this age group with nonstandard employment rose from 3% in 1990 to 15% in 2015. The Japanese labor market is characterized by an insider-outsider segmentation between nonstandard employees and their regular counterparts (Keizer 2008). Regular employment offers indefinite full-time contracts, extensive on-the-job training, and opportunities for career advancement, representing the ‘ideal type’ of standard employment relationship (Kalleberg 2018; Schoppa 2006). Opportunities to enter regular employment are primarily concentrated at the time of graduation: Most students begin applying for jobs while still enrolled in school, receive regular employment job offers by the time they graduate, and are expected to remain with the same firm for long periods of time (Brinton and Kariya 1998; Kariya 1998; Ryan 2001). However, in the 1990s and 2000s, many graduates were unable to secure regular employment and were hired into nonstandard employment jobs. Nonstandard employment, which includes various types of employment such as part-time, temporary, dispatched, contract, and entrusted work, lacks the employment benefits typically associated with regular employment. Regardless of contract type, these nonstandard employment jobs offer lower wages, greater employment insecurity, and limited training opportunities (Houseman and Osawa 2003; Kambayashi and Kato 2016). Moreover, nonstandard employment experiences have long-lasting negative effects on career prospects and earnings. Individuals who begin their careers in nonstandard employment are less likely to obtain regular employment (Hamaaki et al. 2013; Kondo 2007). Additionally, years of experience in nonstandard employment are negatively associated with future earnings when compared to experience in regular employment (Yu 2012).

These employment instabilities would have a detrimental effect on future certainty, especially for men. The role expectations for married couples in Japan are highly differentiated by gender: Men are expected to be the primary breadwinners of their households, with minimal contributions to domestic work (Fuwa 2004; Kan et al. 2022), whereas women are expected to devote their efforts to housework, childcare, and parenting (Brinton 2001; Koyama 2012; Tsuya and Bumpass 2004). Because of these gendered role expectations, men's employment and earnings are far more important criteria than women's when choosing a marital partner (Brinton, Mun, and Hertog 2021; Yu and Hertog 2018). More than 90% of women who have never married report that 'income and earning capacity' is important or at least a consideration when selecting a marital partner, and this proportion has changed little since the 1990s (National Institute of Population and Social Security Research 2023). For men who have never married, this proportion has increased over time but remained below 50% in the most recent period (National Institute of Population and Social Security Research 2023). The strong and persistent gendered expectation regarding men's breadwinning role reinforces the delaying effect of men's employment instability on marriage and childbearing.

3. Literature review

3.1 Theoretical background

Employment instability has risen in many countries since the 1980s, alongside globalization and labor market deregulation (Blossfeld et al. 2005; Mills and Blossfeld 2013). Employment instability is often measured by an individual's unemployment, fixed-term employment contract, or other nonstandard employment contract (Alderotti et al. 2021; Barbieri et al. 2015; Kalleberg 2018; Scherer and Brini 2023). Unemployed individuals face losses of earned income, challenges in daily living, and uncertain future career prospects. Fixed-term workers hold jobs but lack long-term job stability due to the limited duration of their contracts. In labor markets characterized by a strong insider-outsider divide, fixed-term workers find it particularly difficult to move into more stable employment jobs (Barbieri and Cutuli 2016; Barbieri and Scherer 2009). In addition to fixed-term employment, various nonstandard employment arrangements, such as temporary, contract, or part-time employment, are considered unstable forms of employment because they lack the job security associated with standard employment, including indefinite contracts, job protection, and clear future career prospects (Kalleberg 2018; Standing 2011).

Employment instability is expected to be associated with delayed entry into marriage and childbearing, especially for men. Men's higher income derived from employment promotes marriage by increasing household utility through specialization, based on the premise that men have a comparative advantage over women in wage labor (Becker

1981). Higher income is also linked to a greater likelihood of childbearing by ensuring the capacity to meet the financial costs of having children (Becker 1981). The argument that there is an economic bar to marriage also claims that individuals with insecure employment or unemployment refrain from marriage or are viewed as unsuitable marital partners because they cannot meet the expected standard of living after marriage (Gibson-Davis, Edin, and McLanahan 2005; Gibson-Davis, Gassman-Pines, and Lehrman 2018; Ishizuka 2018). Similarly, childbirth, which involves additional financial expenses, is also expected to be postponed when economic resources are insufficient (Happel, Hill, and Low 1984).

Beyond the current level of economic resources, employment instability also undermines the expectation of a future breadwinner role (Oppenheimer 1988, 1994). The future uncertainty triggered by employment instability discourages individuals from marrying and having children, as marriage and childrearing require long-term binding commitments (Mills and Blossfeld 2005, 2013; Oppenheimer 1988, 1994). This theory suggests that future uncertainty itself influences the decision to marry and have children, even when current levels of economic resources are similar. In addition, the connection between employment instability and entry into marriage and childbirth is more pronounced for men in the context of gendered breadwinner role expectations (Esping-Andersen and Billari 2015; Goldscheider, Bernhardt, and Lappegård 2015; Oppenheimer 1988, 1994). While the relationship between stable employment or higher earning capacity and the transition to marriage has become more positive for women, reflecting shifting gender norms (Dominguez-Folgueras and Castro-Martín 2008; Fukuda 2013; Kim 2017; Kimura 2022; Sweeney 2002; Sweeney and Cancian 2004), the relationship has remained stable for men across cohorts (Kim 2017).

Past employment trajectories are also important in determining marriage and childbirth. The life course perspective posits that people make decisions based not only on their current situation but also on their past events, experiences, or trajectories (Huinink and Kohli 2014; Mayer 2009). A lack of stable work experience may signal career immaturity and, thus, an inability to take on the breadwinner role in the future, leading to delays in family formation (Oppenheimer 2003; Oppenheimer, Kalmijn, and Lim 1997). Moreover, experiences of unemployment or fixed-term employment may hinder the accumulation of financial resources (Kravdal 2002) and human capital (Adserà 2004) or lead to negative psychological outcomes (Knabe and Rätzel 2011). This effect of past employment experiences can also be understood as a ‘scar effect’: Unemployment deteriorates subsequent earnings, job quality, subjective well-being, health, pension benefits, or wealth (Brand 2015), all of which contribute to the postponement of marriage and childbirth (Barclay and Kolk 2020; Fu and Goldman 1996; Schneider 2011). These arguments suggest that past employment trajectories are associated with individuals’ decisions to marry or have children, even when current employment conditions are similar.

3.2 Previous empirical findings

Studies have shown that men's unemployment is negatively associated with the entry into first marriage and first birth. Unemployment is used as an indicator of individuals' employment instability and is expected to be associated with lower economic resources and increased future uncertainty. Consistent with this expectation, men's unemployment has been shown to be associated with delayed transitions to first marriage (Ahn and Mira 2001; Alderotti et al. 2021; Kalmijn 2011; Kalmijn and Luijkx 2005; Kim 2017; Oppenheimer 2003; Oppenheimer, Kalmijn, and Lim 1997; Piotrowski, Kalleberg, and Rindfuss 2015; de la Rica and Iza 2005; Yoon, Lim, and Kim 2022). Moreover, men's unemployment is also negatively associated with the timing of first birth (Alderotti et al. 2021; Barbieri et al. 2015; Dupray and Pailhé 2018; Kravdal 2002; Lundström and Andersson 2012; Miettinen and Jalovaara 2020; Pailhé and Solaz 2012; Scherer and Brini 2023; van Wijk, de Valk, and Liefbroer 2022).

The relationship between men's nonstandard employment and the transition to marriage or childbirth has also been studied. In Europe and the United States, fixed-term or part-time employment is commonly used as a measure of nonstandard employment. Compared to permanent or full-time employment, fixed-term or part-time employment is negatively associated with entry into first marriage (Ahn and Mira 2001; Kalmijn 2011; Oppenheimer 2003; Oppenheimer, Kalmijn, and Lim 1997) and first birth (Barbieri et al. 2015; Dupray and Pailhé 2018; Lundström and Andersson 2012; Scherer and Brini 2023; Schmitt 2021; Sutela 2012; Vignoli, Tocchioni, and Mattei 2020), while other studies have found no significant association (de Lange et al. 2014; van Wijk, de Valk, and Liefbroer 2021). Studies suggest that either greater segmentation of the labor market between fixed-term and permanent employment or weak welfare support systems reinforce the negative association between fixed-term employment and entry into marriage and childbirth for men (Alderotti et al. 2021; Barbieri et al. 2015). In East Asia, nonstandard employment is considered a nominal status that includes various forms of precarious employment, as opposed to 'standard' employment contracts (Kalleberg 2018). For men in particular, nonstandard employment has been shown to be negatively associated with the transition to marriage (Esteban-Pretel and Fujimoto 2022; Kim 2017; Matsuda and Sasaki 2020; Mugiyama 2017; Piotrowski, Kalleberg, and Rindfuss 2015; Yoon, Lim, and Kim 2022).

Work experience has also been used as a measure of individuals' employment instability. Previous unstable employment trajectories, such as long periods of unemployment or nonstandard employment, are associated with delayed transitions to first marriage and first birth and weaker fertility intentions for men (Bukodi 2012; Busetta, Mendola, and Vignoli 2019; Ciganda 2015; Dupray and Pailhé 2018; Kalmijn and Luijkx 2005; Mugiyama 2017; van Wijk, de Valk, and Liefbroer 2022). Nonstandard employment at the point of labor market entry is also negatively associated with the transition to first marriage (Mizuocchi 2006; Sakai and Higuchi 2005; Tsuya 2024; Wolbers 2007) and first birth (Sakai and Higuchi 2005). In addition, past experiences of unemployment or nonstandard employment have a negative impact even after controlling for current employment status or economic conditions (Clarkberg 1999; Jalovaara 2012; Kalmijn 2011; van Wijk, de Valk, and Liefbroer 2022). These studies suggest that both past employment experience and current employment status are important in predicting delayed entry into first marriage and first birth.

To sum up, theories and evidence demonstrate the individual-level associations between men's employment instability, such as unemployment, nonstandard employment, or previous employment experience, and later entry into first marriage and first birth. Given these associations, we can expect that increased employment instability across cohorts has contributed to cross-cohort delays in first marriage and first birth at the macro level. This paper examines the extent to which compositional changes in employment status and work experience across cohorts account for the delays in entry into first marriage and first birth for men in Japan.

4. Methods

4.1 Data and sample

I used data from the Social Stratification and Mobility Survey of Japan (SSM), a nationally representative survey of men and women aged 20 to 79 years conducted in 2015. The survey collects retrospective, year-by-year information on respondents' job histories, from their first to present job, as well as details about their family histories, including the timing of first marriage and first birth. The response rate was 50.1%, but the characteristics of the SSM sample closely resemble those of national statistics (Shirahase and Miwa 2021).[2] Although the sample size is modest, the 2015 SSM

[2] Shirahase and Miwa (2021) report that the SSM sample underrepresents individuals in their 20s (born in 1985–1994) who are employed, compared to the 2015 Population Census and the Labor Force Survey. In addition, Yoda (2018) notes that respondents who were in their 70s (born between 1935–1944) at the time of the survey should be interpreted with caution when considering cohort representativeness, as more than 20% of them were already deceased. These groups are not included in this study. To further confirm the representativeness of the sample, I compared the distribution of employment status by age at the time of the survey with the Labor Force

represents the most comprehensive dataset available in Japan, covering a wide range of cohorts and providing detailed longitudinal job and family histories.

The analytical sample consists of male respondents born between 1945 and 1984. These cohorts cover the period during which delays in first marriage, and likely first birth, progressed, as indicated in Appendix Figure A-2. The timing of first marriage was delayed from the earliest cohort to the 1955–1964 cohort, further delayed in the 1965–1974 cohort, and then stabilized at a similar level in the 1975–1984 cohort. After excluding respondents with missing data (92 respondents) and those who were married or had a first child before age 18 (2 respondents), 2,507 respondents remained for the analysis.

I transformed the data into person-year files as the SSM provides annual information on work and family histories. Respondents were included in the risk starting at age 18, which is the legal age of marriage for men in Japan, and were treated as censored cases at age 49, or at the time of the survey if they were younger than 49.[3] The first analytical sample consists of person-years for individuals who have never married, comprising 33,824 person-year observations. The second sample consists of 40,779 person-year observations for individuals who had not yet had their first child, regardless of their marital status.[4]

4.2 Variables

The dependent variables are the transitions to first marriage and first birth. The transition to first marriage was constructed using the reported age at first marriage, measured as entering marriage at age t, given that the respondent was never married at age $t – 1$. The transition to first birth was measured as the respondent having a biological child at age t if they did not have one at age $t – 1$, based on the birth year of their first child.

The focal independent variable is the respondents' birth cohort, separated into 10-year intervals: 1945–1954, 1955–1964, 1965–1974, and 1975–1984.

Survey, as shown in Appendix Table A-1. The distributions are generally similar between the two surveys, except for the shares of self-employment and nonemployment. The share of nonemployment is lower, and the share of self-employment is higher in the SSM than in the Labor Force Survey. This discrepancy is partly due to differences in how employment is measured: The SSM asks about the respondent's usual job, while the Labor Force Survey measures whether the respondent worked during the last week of the month, which may underestimate casual work. Therefore, while the SSM survey may underestimate the nonemployment, at least at the time of the survey, the sample is well representative of the population.

[3] Some respondents may get married or have a child after age 50. Although I also conducted the analyses including observations from age 50 to 59, the results were not much different (available upon request).

[4] The SSM survey asked individuals who had children if those children were their biological children or stepchildren. I considered the oldest biological child to be the respondent's first child and coded the year of birth as the year of transition to first birth.

Another important independent variable, employment instability, is measured by using two variables: employment status and cumulative years spent in each employment status. First, employment status is categorized into regular employment, nonstandard employment, self-employment, nonemployment, and enrollment in school. Nonstandard employment includes part-time work, temporary work, dispatched work, contract work, and entrusted work, following previous studies (Houseman and Osawa 2003; Kalleberg, Hewison, and Shin 2021). Nonemployment encompasses both unemployment and inactivity, as these categories are not distinguished in the retrospective job history data. However, given that inactive youths face similar unstable employment situations as the unemployed in the Japanese context, as discussed earlier, the nonemployment category still provides meaningful insight into the impact of employment instability. Respondents enrolled in school are categorized as "enrollment in school" regardless of their employment status, as student employment has different implications than nonstudent employment and is generally not considered unstable. In the multivariate analyses, this category is not reported due to complete multicollinearity with educational attainment variable, which is explained later. Second, cumulative years in each employment status are defined as the number of years worked in each type of employment – that is, cumulative years in regular employment, nonstandard employment, and self-employment – following previous studies on the scar effect of nonstandard employment in Japan (Yu 2012).[5]

The control variables include age and educational attainment. Age is measured as a time-varying continuous variable. To account flexibly for changes in hazards over the life course, linear, quadratic, and cubic terms for age are included. I chose a polynomial with a third-order term based on both theoretical and empirical considerations. Hazards for both first marriage and first birth have been shown to rise sharply and then decline,

[5] A possible concern is that cumulative years, by definition, increase with age, which may lead to multicollinearity with age. Appendix Table A-5 shows that the correlation coefficient between cumulative years in regular employment and age is just under 0.8. An alternative way to measure work experience in each employment status is to use the share of years. For example, the share of years in regular employment is calculated by dividing the cumulative number of years in regular employment by the total number of years worked and not employed, resulting in a value ranging from 0 to 1. The value is set to zero if the individual has never been employed and has not left an educational institution. While these variables are only weakly correlated with age, they are highly correlated with the employment status variable. Specifically, the correlation coefficients between the share of years in regular employment and the regular employment dummy are around 0.87; similar correlations are observed for self-employment and nonstandard employment. The correlation coefficients between cumulative years and these dummy variables are lower than those for the share of years. Although neither measure is perfect, multicollinearity between employment status and work experience is a more serious issue, given the interest to accurately estimate the coefficients of both variables. Consequently, I decided to use the cumulative years variable rather than the share variable. It should be noted that the main conclusions of this paper are not significantly affected when using the share of years spent in each employment status (see Appendix Table A-6).

patterns often approximated by quadratic polynomials (Barbieri et al. 2015; Kalmijn and Luijkx 2005; Kim 2017; Mugiyama 2017; Piotrowski, Kalleberg, and Rindfuss 2015; Tsuya 2024; Yoon, Lim, and Kim 2022). In the current sample, however, the age-specific hazard rises sharply until the late 20s and then declines slowly, resulting in right-skewed distributions (shown in Figure A-3). To account for these age patterns, a cubic polynomial is more appropriate. Some studies show that the cubic polynomials better fit age-specific fertility patterns (Raftery et al. 1996). Singer and Willett (2003) also suggest comparing models with multiple quadratic functions of time when using discrete-time hazard models. The results indicate that a cubic polynomial function substantially improves the model fit (see Appendix Table A-2). Based on these considerations, the cubic term for age was introduced.[6]

Educational attainment is defined as the highest level of education completed and is measured as a time-varying variable. I also added a category indicating enrollment in school for respondents who were still in school. The likelihood of entry into marriage (and likely childbirth) is quite low during this period (Raymo 2003). Because more individuals remain in school longer due to educational expansion, the impact of increased nonemployment due to labor market changes would be overestimated if periods of school enrollment were not distinguished. The resulting categories of educational attainment were junior high school, high school, vocational school, junior college, university (or higher), and enrollment in school.

Descriptive statistics of the variables used are presented in Table 1.

[6] Higher order polynomials may provide better model fits. Appendix Table A-2 presents the fit statistics for both quadratic and quintic polynomials of age (see 1c and 1d). For the transition to first marriage, the quadratic polynomial shows a better fit than the cubic; however, the improvement in goodness of fit is not as substantial as the improvement from quadratic to cubic. Thus, I chose the cubic polynomial based on substantive considerations. Furthermore, the shape of the age-specific hazard may differ across cohorts (see Figure A-3). To account for this variation, I also estimated models that include interactions between cohort and the linear, quadratic, and cubic terms of age (see 2a–2c). These models slightly improved the fit, as confirmed by log-likelihood tests, although the BIC worsened. For the sake of simplicity, I chose the cubic polynomial of age without interaction with cohort for both the transition to first marriage and first birth. The results did not differ substantially when using the more flexible models (2c), as shown in Appendix Table A-4.

Table 1: Descriptive statistics

	Transition to first marriage		Transition to first birth	
	Person-year	At last observation	Person-year	At last observation
Transition to first marriage/first birth	0.060	0.813	0.044	0.712
Cohort				
1945–1954	0.285	0.323	0.298	0.323
1955–1964	0.251	0.235	0.250	0.235
1965–1974	0.274	0.236	0.272	0.236
1975–1984	0.191	0.205	0.180	0.205
Employment status				
Regular employment	0.674	0.817	0.700	0.809
Self-employment	0.065	0.093	0.075	0.103
Nonstandard employment	0.071	0.054	0.064	0.055
Nonemployment	0.052	0.024	0.045	0.024
Enrollment in school	0.138	0.012	0.116	0.009
Cumulative years in regular employment	6.010 (6.534)	9.396 (7.133)	7.110 (7.116)	11.694 (7.956)
Cumulative years in self-employment	0.563 (2.520)	0.903 (3.357)	0.739 (2.984)	1.247 (4.170)
Cumulative years in nonstandard employment	0.791 (2.643)	1.010 (3.308)	0.774 (2.604)	1.101 (3.477)
Age	26.260 (7.039)	30.492 (7.373)	27.539 (7.529)	33.266 (7.877)
Educational attainment				
Junior high	0.110	0.109	0.112	0.109
High school	0.437	0.437	0.437	0.437
Vocational school	0.087	0.101	0.088	0.100
Junior college	0.016	0.022	0.015	0.022
University	0.212	0.320	0.232	0.323
Enrollment in school	0.138	0.012	0.116	0.009
N	33,824	2,507	40,779	2,507

Notes: Means and proportions are reported. Standard deviations are shown in parentheses. The "person-year" columns show summary statistics in person-year observations. The "at last observation" columns show summary statistics at the time of entry into first marriage or first birth for those who have ever experienced these transitions and at the last observation period for those who have never experienced these transitions.

4.3 Statistical methods

I used discrete-time event history analysis with logit models to examine the relationship between transitions and the independent variables (Allison 2014; Singer and Willett 2003). The discrete-time hazard rate was defined as the conditional probability that a transition occurred for individual i at time t, given that it had not occurred at time $t - 1$, represented as $h_{it} = \Pr(T = t \mid T > t - 1)$.[7] In Model 1, I estimated the following model for predicting the hazard rate h_{it}:

[7] This implies that the transition is predicted by the state of the variables in the same year. Piotrowski, Kalleberg, and Rindfuss (2015) show that lagging employment status does not significantly change its effect on men's transition to first marriage. I also confirmed that lagging the independent variables by one or two years does not significantly alter the main results (see Appendix Table A-7).

$$\log\frac{h_{it}}{1-h_{it}} = \boldsymbol{X}_{it}\boldsymbol{\beta}, \tag{1}$$

where $\boldsymbol{X}_{it}$ refers to the variables of age, cohort, and educational attainment.

In Model 2, I introduced variables for employment status and cumulative years in each employment status, denoted as $\boldsymbol{W}_{it}$, as follows:

$$\log\frac{h_{it}}{1-h_{it}} = \boldsymbol{X}_{it}\boldsymbol{\beta}' + \boldsymbol{W}_{it}\boldsymbol{\delta}. \tag{2}$$

I present average marginal effects (AMEs) of each independent variable instead of logit coefficients. Logit coefficients cannot be compared across models when the models include different sets of independent variables, whereas AMEs, which measure the magnitude by which a one-unit change in an independent variable affects the probability of an outcome (e.g., the annual hazard rate in discrete-time event history analysis), can be directly compared across models (Long and Freese 2014; Mood 2010).[8] I also applied a seemingly unrelated estimation technique to determine whether difference in AMEs varied across models (Mize, Doan, and Long 2019).[9] This technique uses the z-statistic of the difference in AMEs by estimating cross-model covariances to formally test differences across models.

Based on these estimates, I calculated survival rates – that is, the age-specific proportions of individuals who never married or remained childless, which is a more consistent measure of the phenomena of delayed marriage and delayed birth (Bloome and Ang 2020). To obtain the survival rates, I first calculated the predicted age-specific hazard rate of cohort c using the results of the discrete-time logit models in Model 1 (Equation 1):

$$\hat{h}_t(C = c, \boldsymbol{X}_t) = \frac{1}{N}\sum_{i=1}^{N}\hat{h}_{it}(\boldsymbol{C}_i = \boldsymbol{c}, \boldsymbol{X}_{it} = \boldsymbol{x}_{it}), \text{ and} \tag{3}$$

The baseline survival rates were then calculated as

$$\hat{S}(T = t \mid C = c) = \prod_{k=18}^{t}[1 - h_k(C = c, \boldsymbol{X}_k)]. \tag{4}$$

[8] Notably, the baseline estimates (Model 1) are derived from the model controlling for educational attainment. This indicates that the 'reduced' portion of the associations between cohorts and transitions should be interpreted as the one conditioned on the effects of compositional changes in educational attainment across cohorts.

[9] Several studies also apply this technique to discrete-time event history analysis (Pessin, Rutigliano, and Potter 2022; Rackin and Gibson-Davis 2022).

Next, I obtained the predicted age-specific hazard rate of cohort c, net of compositional changes in employment status and work experience, using the results from Model 2 (Equation 2):

$$\hat{h}_t(C = c, \boldsymbol{X}_t, \boldsymbol{W}_t) = \frac{1}{N}\sum_{i=1}^{N} \hat{h}_{it}(\boldsymbol{C}_i = \boldsymbol{c}, \boldsymbol{X}_{it} = \boldsymbol{x}_{it}, \boldsymbol{W}_{it} = \boldsymbol{w}_{it}). \tag{5}$$

The difference in hazard rates between cohort c and the earliest cohort (1945–1954) was calculated as $\hat{h}_t(C = c, \boldsymbol{X}_t, \boldsymbol{W}_t) - \hat{h}_t$ (C = 1945–1954, $\boldsymbol{X}_t$, $\boldsymbol{W}_t$). This difference gauges the increased hazard if employment status and work experience are held constant. By adding this increased hazard to the baseline hazard rates for each cohort, I obtained the standardized survival rate, which indicates what would have been observed if there had been no compositional changes in employment-related variables across cohorts:

$$\begin{gathered}\hat{S}_{\text{standardized}}(T = t \mid C = c) \\ = \prod_{k=18}^{t}\left[1 - \begin{bmatrix}\hat{h}_k(C = 1945\text{–}1954, \boldsymbol{X}_k) + \hat{h}_k(C = c, \boldsymbol{X}_k, \boldsymbol{W}_k) \\ -\hat{h}_k(C = 1945\text{–}1954, \boldsymbol{X}_k, \boldsymbol{W}_k)\end{bmatrix}\right].\end{gathered} \tag{6}$$

The comparison between Equations (4) and (6) allows us to assess the extent to which compositional changes in employment-related variables contributed to cross-cohort differences in survival rates. The observed difference between the 1945–1954 cohort and cohort c can be decomposed as

$$\begin{aligned}&\hat{S}(T = t \mid C = c) - \hat{S}(T = t \mid C = 1945\text{–}1954) = \\ &\left[\hat{S}(T = t \mid C = c) - \hat{S}_{\text{standardized}}(T = t \mid C = c)\right] + \\ &\qquad +\left[\hat{S}_{\text{standardized}}(T = t \mid C = c) - \hat{S}(T = t \mid C = 1945\text{–}1954)\right].\end{aligned} \tag{7}$$

Here, the first term captures the contribution of changing compositions of employment status and work experience, while the second term represents other factors unrelated to employment compositions that account for cross-cohort differences.

This method extends related approaches used in previous studies. The methods are developed to calculate survival estimates and summary fertility measures, such as cohort or period total fertility rate or parity progression ratios, using individual-level event history analysis estimates (Retherford et al. 2010; Van Hook and Altman 2013; Zang 2019). Unlike these studies, I focus on survival rates – that is, age-specific proportions of individuals who never married or remained childless – to assess delays in first marriage and first birth. Furthermore, I decompose the extent to which employment-related variables account for cross-cohort changes (Equation 7) by comparing unstandardized

(Equation 4) and standardized (Equation 6) survival rates, an approach not explicitly conducted in previous studies.

5. Results

5.1 Descriptive results

Figure 2 shows the Kaplan–Meier survival plots for the timing of first marriage and first birth across cohorts. In the 1945–1954 cohort, most men married in their 20s. The proportion of individuals who never married declined sharply, with about 80% of men married by age 30. The pace of marriage slowed thereafter, but over 90% of men were eventually married by age 49. Similarly, about 60% of men had their first child by age 30, and approximately 85% had their first child by age 40. In subsequent cohorts, the timing of first marriage and first birth was consistently delayed. About 35% of men in the 1955–1964 cohort remained unmarried at age 30, and this proportion increased to approximately 50% in later cohorts.[10]

The timing of first birth has also been delayed. In the earliest cohort, over 60% of men at age 30, and around 80% of men had a child by age 35. In contrast, in the latest cohort, 65% of men had not had a child by age 30, and this proportion remained higher at later ages. Compared with transition to first marriage, the proportions of individuals who had not experienced first birth are higher. This indicates that not all men have a child after entering their first marriage.

[10] Although the timing of first marriage and first birth appeared to be earlier in the 1975–1984 cohort than in the previous cohort, the differences were not statistically significant according to Wilcoxon tests ($p = 0.264$ and $p = 0.074$). As shown in Appendix Figure A-2, the age-specific proportion of individuals who never married in the 1965–1974 and 1975–1984 cohorts almost overlap, so these results are consistent with the Population Statistics.

Figure 2: Kaplan–Meier survival estimates for the transition to first marriage and first birth by cohort

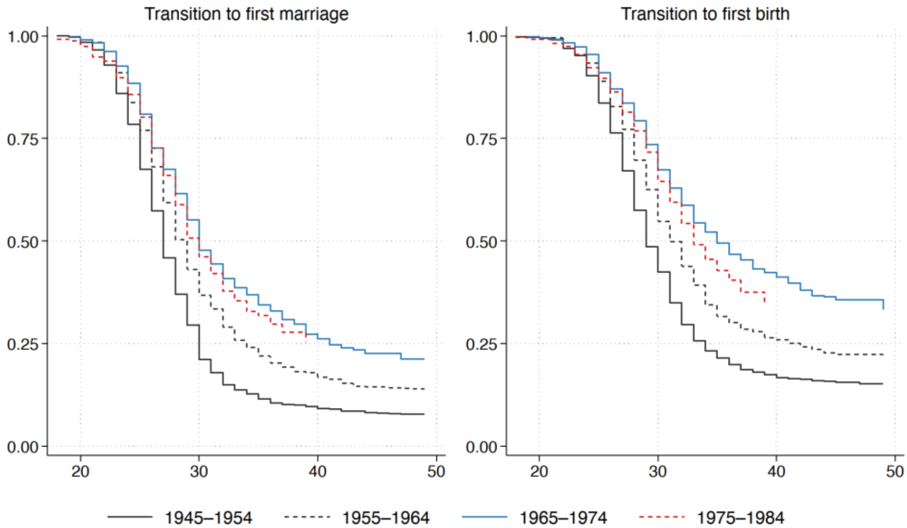
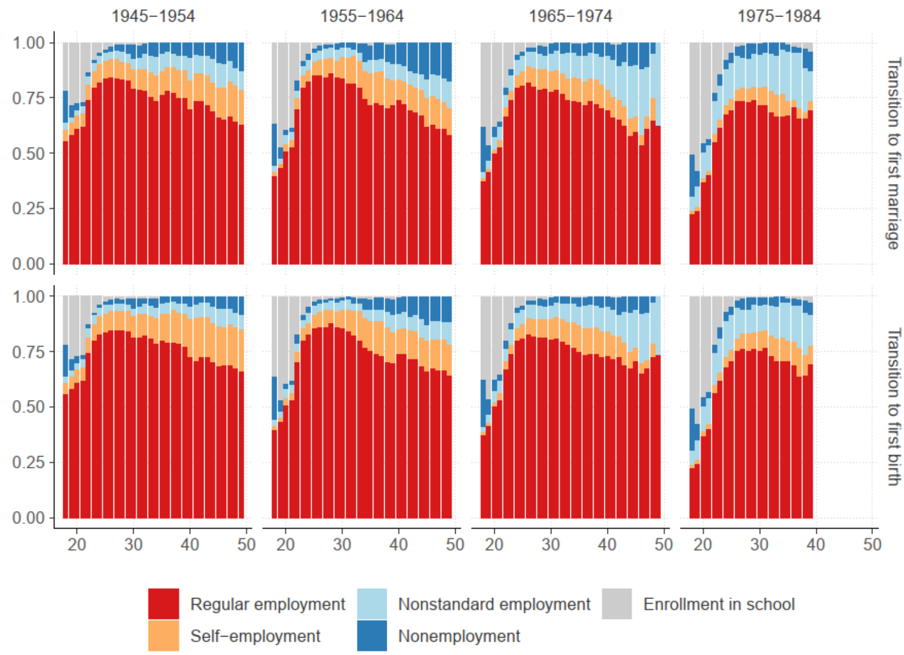


Figure 3 shows the distribution of employment status among those who never married (upper panels) or remained childless (lower panels) at each age. For both samples, in the 1945–1954 cohort, most men spent their post-school careers in either regular employment or self-employment, with few experiencing nonstandard employment or nonemployment. In later cohorts, a larger proportion of men were exposed to nonstandard employment or nonemployment. In the 1965–1974 cohort, about 13% of men who never married by age 30 were in nonstandard employment or nonemployment. In the 1975–1984 cohort, this figure rose to 20%. They indicate that more individuals spend a greater number of years in nonstandard employment or nonemployment in later cohorts. Moreover, the share of those in nonstandard employment or nonemployment increased with age, suggesting that individuals in nonstandard employment or nonemployment are more likely to never married and remain childless.

Figure 3: Distribution of employment status of those who never married or remained childless by age and cohort



Notes: The upper panels show the share of each employment status among those not yet married at each age, while the lower panels show the share among those without children at each age.

5.2 Estimated results for discrete-time hazard models

Table 2 shows the AMEs estimated from the discrete-time logit models (Equations 1 and 2). The coefficients indicate the difference in percentage points (pp) of the annual hazard rate compared to the reference category. Regarding the transition to first marriage, Model 1 reveals that men in later cohorts were less likely to enter marriage. Compared to the 1945–1954 cohort, men in the 1955–1964 cohort were 2.5 pp less likely to marry per year, while this value was 3.9 pp for the 1965–1974 cohort and 3.7 pp for the 1975–1984 cohort.

The transition to first birth also became less likely in the later cohorts. Compared with the 1945–1954 cohort, the annual hazard rate of transition to first birth decreased by 1.7 pp for the 1955–1964 cohort, 3.1 pp for the 1965–1974 cohort, and 2.7 pp for the

1975–1984 cohort (see Model 1). The trends are similar to those for the transition to first marriage, but the change in the magnitude of the hazard rates is relatively smaller, reflecting the lower baseline hazard rates for the transition to first birth (see Figure 2 and Figure A-3).

Table 2: Average marginal effects of discrete-time logit models predicting transitions to first marriage and first birth

	Transition to first marriage			Transition to first birth		
	Model 1	Model 2	Diff.	Model 1	Model 2	Diff.
Cohort (ref: 1945–1954)						
1955–1964	–.0248	–.0223	–.0025	–.0168	–.0146	–.0022
	(.0039)	(.0038)	(.0004)	(.0031)	(.0030)	(.0003)
1965–1974	–.0387	–.0339	–.0048	–.0311	–.0272	–.0039
	(.0037)	(.0036)	(.0005)	(.0029)	(.0028)	(.0004)
1975–1984	–.0371	–.0273	–.0098	–.0273	–.0198	–.0075
	(.0040)	(.0041)	(.0010)	(.0032)	(.0033)	(.0008)
Employment status (ref: regular employment)						
Self-employment		.0179			.0091	
		(.0092)			(.0065)	
Nonstandard employment		–.0303			–.0150	
		(.0056)			(.0055)	
Nonemployment		–.0545			–.0351	
		(.0037)			(.0038)	
Cumulative years in regular employment		.0023			.0032	
		(.0009)			(.0007)	
Cumulative years in self-employment		.0006			.0019	
		(.0012)			(.0009)	
Cumulative years in nonstandard employment		–.0007			–.0001	
		(.0013)			(.0010)	
N	33,824	33,824		40,779	40,779	

Notes: Standard errors in parentheses. Controls for educational attainment and cubic polynomials of age. The log-odds coefficients of all independent variables are reported in Appendix Table A-3. The “Diff.” columns show the differences in AMEs between Models 1 and 2, following Mize, Doan, and Long (2019).

Model 2 adds employment status and cumulative years of experience in each employment status to Model 1. Nonstandard employment and nonemployment are negatively associated with both transitions. Nonstandard employment is associated with a 3.0 pp lower annual hazard rate of first marriage, while nonemployment with a 5.5 pp lower rate compared to regular employment. In addition, cumulative years in regular employment are positively associated with both transitions. Each additional year of experience in regular employment is associated with a 0.23 pp higher transition rate to first marriage. These results suggest that employment, but not nonstandard employment, and a stable career trajectory in regular employment are associated with higher entry rates into first marriage and first birth.

In addition, employment status and cumulative years of experience partially account for the cross-cohort decline in hazard rates. The third column shows the changes in the

AMEs of the cohorts between Models 1 and 2, along with the standard errors. For example, for the transition to first marriage, the AMEs for the 1955–1964 cohort decreased by 0.25 pp after controlling for employment-related variables. Calculating the proportion of the reduction in AMEs shows that the magnitude of the reduction is greater in the later cohorts: the reduction is 10% (= 0.0025 / 0.0248) for the 1955–1964 cohort, 12% (= 0.0048 / 0.0387) for the 1965–1974 cohort, and 26% (= 0.0098 / 0.0371) for the 1975–1984 cohort.

For the transition to first birth, the rates are 1.5 pp lower for nonstandard employment and 3.5 pp lower for nonemployment than for regular employment. Longer experience in regular employment is also associated with a higher transition rate to first birth, with a 0.32 pp higher rate per additional year. In contrast, experience in nonstandard employment is not positively associated with both transitions.

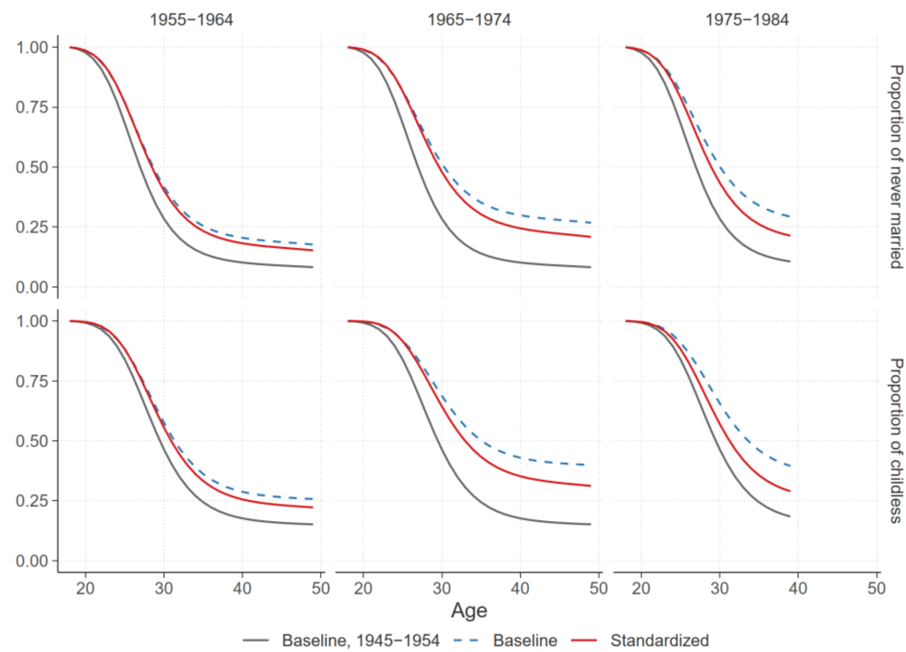
The cohort differences in the hazard rate of first birth also decline after introducing employment-related variables. The reduction was 13% (= 0.0022 / 0.0168) for the 1955–1964 cohort and 13% (= 0.0039 / 0.0311) for the 1965–1974 cohort, whereas it was 27% (= 0.0075 / 0.0273) for the 1975–1984 cohort. For both transitions, the reduction is greatest in the most recent cohort. This reflects greater exposure to nonstandard employment or nonemployment and longer duration spent outside of regular employment for the most recent cohort, as illustrated in Figure 3.

5.3 Contribution of increased employment instability to age-specific proportions of individuals who never married and remained childless

Figure 4 shows the estimated age-specific proportions of individuals who never married or remained childless by cohort. The baseline age-specific proportions of individuals who never married (blue dashed lines) in each cohort are generally higher than those in the 1945–1954 cohort (gray lines), with the gap being larger for the 1965–1974 and 1975–1984 cohorts, consistent with Figure 2. The standardized proportions (red lines), which represent the counterfactual proportions of individuals who never married without the compositional changes in employment-related factors, are lower than the baseline proportions across all ages. This indicates that the proportion of individuals who never married would have been lower if exposure to employment instability had not increased between cohorts.

Similar patterns are observed for the proportion of individuals who remained childless. When compositional changes in the employment-related variables are taken into account, the proportions of childless are lower than the baseline proportions and come closer to those of the earliest cohort.

Figure 4: Estimated baseline and standardized age-specific proportion of individuals who never married or remained childless by cohort

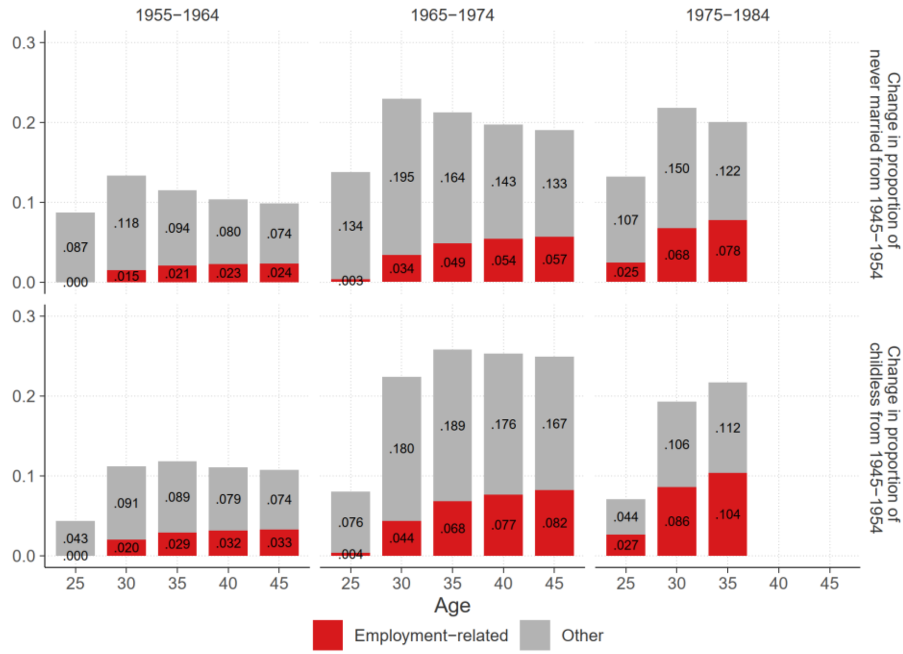


Notes: Baseline proportions were estimated using Equation (4) and indicate the predicted age-specific proportions of individuals who never married or remained childless (i.e., survival rates) of each cohort net of educational attainment. Standardized proportions were estimated using Equation (6) and indicate the proportions that would have been observed in the absence of compositional changes in employment status and work experience across cohorts.

Figure 5 shows the decomposition results of the extent to which changes in the age-specific proportions of individuals who never married or remained childless between the earliest and later cohorts are accounted for by changes in employment-related factors. The height of the bars indicates the cross-cohort difference in the proportions by age. From the 1945–1954 cohort to the 1955–1964 cohort, the proportion of men never married increased by 13.3 pp. Of this increase, 1.5 pp, or 11% (= 0.015 / (0.015 + 0.118)), is accounted for by employment-related changes. This contribution is higher at later ages. At age 45, 24% (= 0.024 / (0.024 + 0.074)) of the total increase in the proportion is accounted for by employment-related changes. In the later cohorts, employment-related factors play a larger role in the increase. For the 1965–1974 cohort, the proportion of individuals who had never married by age 30 increased by 22.9 pp compared with the 1945–1954 cohort, with employment-related changes contributing 3.4 pp, or 17%

(= 0.034 / (0.034 + 0.195)). For the 1975–1984 cohort, 6.8 pp of the increase at age 30 is accounted for by employment-related changes, representing 31% (= 0.068 / (0.150 + 0.068)) of the total increase of 21.8 pp.

Figure 5: Contribution of employment-related factors to the increasing age-specific proportion of individuals who never married or remained childless between the 1945–1954 cohort and the later cohort



Notes: The decomposition follows Equation (7). The employment-related component indicates the extent to which cross-cohort changes in proportion of individuals who never married or remained childless from the 1945–1954 cohort are accounted for by compositional shifts in employment status and cumulative years in each employment status. The other component reflects changes that are not accounted for by the changes in these employment-related characteristics.

The increase in the proportion of individuals who remained childless was also partly accounted for by changing employment-related factors, with a greater contribution from these factors than in the increase in proportion of those who never married. From the 1945–1954 cohort to the 1955–1964 cohort, the proportion increased by about 11 pp between ages 30 and 45. The contribution of employment-related factors increases with age: At age 45 they contributed 3.3 pp to the total increase. The contribution of employment-related factors is larger for the subsequent cohorts. For the 1965–1974 cohort, the proportion at age 30 was 23.9 pp higher than in the earliest cohort, and 20% (= 0.044 / (0.044 + 0.180)) of this difference is accounted for by employment-related factors. This proportion increases with age, reaching 7.7 pp at age 40, or 30% (= 0.077 / (0.077 + 0.176)) of the total increase. The contribution of employment-related factors is greatest for the 1975–1984 cohort. At age 30, these factors account for 45% (= 0.086 / (0.086 + 0.106)) of the total increase since the earliest cohort. Notably, this figure is greater than the corresponding contribution to the proportion of those who never married, which is 31%. This suggests that the compositional change in employment-related factors accounts for a larger share of the proportion of those who remained childless than for the increase in the proportion of those who never married.

In sum, these results suggest that changing exposure to employment-related factors, including increased nonstandard employment or nonemployment and decreased experience in regular employment, accounts for the increase in age-specific proportion of individuals who never married and remained childless across cohorts, with the contribution increasing with age. In addition, when comparing the proportions of individuals who never married and those who remained childless, the contribution of employment-related factors to the childless proportion is greater, particularly for the most recent cohort. At age 30, the compositional change in employment-related factors between the earliest and latest cohorts accounts for 31% of the increase in the proportion of individuals who never married, whereas this figure is 45% for the increase in the proportion of individuals who remained childless.

5.4 Supplementary analyses: Transition to first birth among married men

We have seen that increased employment instability has contributed to delayed entry into both first marriage and first birth across cohorts. Previous studies emphasize that most of the fertility decline in Japan is driven by the decline in marriage prevalence rather than a decline in marital fertility (Atoh, Kandiah, and Ivanov 2004; Iwasawa 2008; Jones 2007). This suggests that employment-related factors contribute little to delayed first birth among married individuals. However, Figure 5 shows that the contribution of changing employment-related factors accounts for a larger share of the increased proportion of childless than of never married. This difference suggests that increased employment instability contributes not only to delayed marriage but also to delayed marital birth.

To complement these results, I also analyzed whether the transition to first birth among married men changed across cohorts and how employment-related factors contributed to this trend. The analytical sample includes person-year observations of men who were married (including first marriages and remarriages) but had not yet had their first child, comprising 8,101 person-year observations from 1,763 respondents.

Respondents who reported the same age at first marriage and age at first birth were excluded from the analysis.[11]

Table 3 shows the results of discrete-time logit models. Married men in the 1965–1974 cohort were 3.9 pp less likely to enter into first birth than those in the 1945–1954 cohort and were 2.7 pp less likely in the 1975–1984 cohort, indicating that the interval between marriage and childbearing has become longer in more recent cohorts (see Model 1). Nonemployment and shorter duration of regular employment were also associated with a delayed transition to marital birth (see Model 2). Importantly, controlling for employment status and work experience reduced the difference in annual hazard rate, particularly between the earliest and latest cohorts; 0.7 pp of the total change in the AMEs, or 27% (= 0.0072 / 0.0268), is accounted for by employment-related factors.

The result suggests that in the most recent cohort, increased employment instability contributes, in part, to delayed first births within marriage. I discuss this implication in more detail in the concluding section.

[11] Among respondents who had a first child in the analytical sample, 2.7% reported that the age of first birth was earlier than the age of first marriage, and 12.5% reported the same age of first birth and first marriage. Since respondents who reported the same age of first birth and first marriage are likely to be premarital pregnancies, I decided to remove these cases to see if employment instability delays the decision of married individuals to have a child. However, it is still possible that our sample includes some married individuals whose partner was pregnant before marriage and excludes some married individuals whose partner was not pregnant before marriage due to the way the timing of marriage and childbirth is collected in the SSM. The SSM asks for the year of birth of the first through fourth children and the age of the first and last marriage, not the month or date. For example, a respondent born on September 1, 1970, reported age at first marriage as 29 and year of birth of first child as 2000. In this case, I coded him as having married at age 29 and had his first child at age 30, but he could have married in April 2000 and had a first child in July 2000. We should consider this case as a premarital pregnancy. We can also guess that some of the removed respondents (i.e., those whose age at first marriage and age at first birth are the same) may have actually married before getting pregnant. Because of this limitation, this analysis should be interpreted carefully. Further studies are needed to determine the exact relationship between employment instability and delayed marital birth.

Table 3: Average marginal effects of discrete-time logit models predicting the transition to first birth among married men

	Model 1	Model 2	Diff.
Cohort (ref: 1945–1954)			
1955–1964	–.0040	–.0033	–.0007
	(.0118)	(.0118)	(.0011)
1965–1974	–.0389	–.0373	–.0016
	(.0115)	(.0115)	(.0011)
1975–1984	–.0268	–.0196	–.0072
	(.0132)	(.0135)	(.0025)
Employment status (ref: regular employment)			
Nonstandard employment		.0077	
		(.0223)	
Self-employment		.0328	
		(.0378)	
Nonemployment		–.0468	
		(.0561)	
Cumulative years in regular employment		.0093	
		(.0033)	
Cumulative years in self-employment		.0066	
		(.0039)	
Cumulative years in nonstandard employment		.0028	
		(.0044)	
N	8,101	8,101	

Notes: Standard errors in parentheses. Controls for educational attainment and cubic polynomials of age. The full log-odds coefficients and descriptive statistics are shown in Tables A-8 and A-9 in the Appendix. The “Diff.” columns show the differences in AMEs between Models 1 and 2, following Mize, Doan, and Long (2019).

6. Discussion and conclusion

Increased employment instability has been argued to contribute to the postponement of family formation in recent decades. Studies have shown that individuals’ exposure to employment instability is negatively associated with entry into unions, marriage, and childbirth. However, evidence directly demonstrating the macro-level contribution of increased individual’s employment instability has been limited. This study examined the extent to which increased individual exposure to nonstandard employment, nonemployment, and greater work experience outside regular employment accounts for cross-cohort delays in first marriage and first birth among men in Japan born between 1945 and 1984.

Results from discrete-time hazard models showed that increased employment instability contributed substantially to the decline in the rate of entry into first marriage and first birth among men across cohorts. Men in more recent cohorts delayed the timing of marriage and birth and were more likely to be exposed to nonstandard employment or nonemployment rather than regular employment or self-employment, leading to longer durations of unstable employment. The cross-cohort compositional changes in employment-related factors accounted for about 10% of the decline in the annual first marriage rate from the earliest cohort (1945–1954) to the middle cohorts (1955–1974), and 26% from the earliest cohort to the latest cohort (1975–1984). Employment-related changes also contributed to the decline in the annual first birth rate, accounting for 13% of the decline from the earliest to the middle cohorts and of 27% from the earliest to the latest cohort.

Counterfactual simulations of survival estimates revealed a more detailed picture. The gap in proportion of individuals who never married between the earliest and later cohorts increases around age 30 and then decreases slightly, but the increased gap persists thereafter, with the increased contribution of employment-related factors with age. For the latest cohort, 6.8 percentage points of the total increase, or 31%, in the proportion of individuals who never married at age 30 were accounted for by cross-cohort changes in employment status and work experience. A similar pattern was observed for the proportion of individuals who remained childless, where employment-related factors accounted for a larger share of the total increase: Changing employment-related factors contributed 8.6 percentage points, or 45%, of the increase in the proportion at age 30 from the earliest to the latest cohort.

Overall, the results suggest that increased employment instability has contributed significantly to the cross-cohort trend of delayed first marriage and first birth among men in Japan. The contributions are not negligible: Although depending on the measure, approximately 26% of the decline in the hazard rates of first marriage and first birth, 31% of the increase in the proportion of individuals who never married by age 30, and 45% of the increase in the proportion of individuals who remained childless can be accounted for by increased individual exposure to employment instability across cohorts. These findings support claims that increased employment instability, associated with economic downturns and labor market transformation, has led to the postponement of family formation (Blossfeld et al. 2005; Kohler, Billari, and Ortega 2002; Mills and Blossfeld 2013; Oppenheimer 1988, 1994). As suggested by Mills and Blossfeld (2013), structural changes in the economy and labor market can complement the second demographic transition argument, which emphasizes the rise of individualistic values driving changes in union formation and childbearing (Lesthaeghe 1995; Van de Kaa 1987), by shedding light on the role of structural factors.

At the same time, much of the cross-cohort delay in the timing of first marriage and first birth remains unaccounted. In addition to ideational transitions, several explanations have been proposed. In Japan, higher opportunity costs of marriage, marriage market mismatches, and extended parental coresidence have also been suggested as the potential sources contributing to delayed marriage and childbirth (Raymo et al. 2015). Moreover, the discrepancy between changes in individual values and the persistent gender essentialism may further suppress entry into marriage or childbirth (Brinton and Lee 2016). The analytical approach used in this paper can be applied to disentangle the remaining sources of the postponement of family formation.

Notably, the supplementary analyses showed that increased employment instability also contributed to delayed entry into first birth among married men in addition to delayed entry into first marriage. This finding challenges the prevailing assumption about the process of low fertility in Japan and potentially in other East Asian contexts. Although previous studies have attributed the primary driver of low fertility in Japan to the decline in marriage prevalence (Atoh, Kandiah, and Ivanov 2004; Iwasawa 2008; Jones 2007), the current results highlight that increased employment instability is also associated with delayed first births among married men. Moreover, the magnitude of the hazard rates also differs between the transition to first marriage and first birth (see Figure 2 and Figure A-3), suggesting that changes in the timing of first marriage cannot be directly equated with changes in the timing of first birth. Thus, focusing solely on delayed marriage may underestimate the overall contribution of employment instability to fertility, even within the context of the strong linkage between marriage and fertility in Japan (Bumpass et al. 2009; Raymo et al. 2015). Furthermore, the presence of the nonnegligible occurrence of premarital pregnancies (Mogi et al. 2023; Raymo and Iwasawa 2008) suggests that childbirth decisions may not always be made prior to marriage. Future studies should expand their focus beyond marriage entry alone to examine first birth entry without conditioning on marriage, marital birth, or premarital pregnancy in order to investigate the impact of employment instability or other factors on family formation.

Despite the sole focus on men in this paper, it is important to emphasize the gender differences in the role of increased employment instability in delayed family formation. Employment instability has been suggested to have only a weak delaying effect on women’s entry into marriage or childbirth, or even a positive effect (Friedman, Hecher, and Kanazawa 1994). This is particularly true in societies where stronger gender norms surrounding marriage and childrearing prevail (Oppenheimer 1988, 1994). Studies in Japan suggest that the association between employment instability and marriage entry for women is not strongly negative or even positive (Esteban-Pretel and Fujimoto 2022; Lim 2018; Matsuda and Sasaki 2020; Mugiyama 2017; Piotrowski, Kalleberg, and Rindfuss 2015). Additionally, increased employment instability has not been shown to lead to lower fertility among women (Hashimoto and Kondo 2012; Raymo and Shibata 2017). The current results, alongside previous studies, suggest that delays in marriage and childbirth have progressed more rapidly for men from lower socioeconomic backgrounds, who face greater exposure to employment instability, whereas the delay has progressed more evenly across socioeconomic backgrounds for women. Thus, the increase in employment instability has likely contributed primarily to delayed marriage or childbirth on the male side, consistent with the claim that increased men’s employment instability leads to delayed marriage and childbirth (Oppenheimer 1988, 1994).

Recently, even in Japan, which has a strong gendered division of labor and significant gender inequality in the labor market, normative expectations for marriage and childrearing have become more gender egalitarian. These shifting norms increasingly emphasize the role of women's earning potential (National Institute of Population and Social Security Research 2023) and are supported by the expansion of family-supportive policies (Nagase 2018; Yamaguchi 2017), which may lead to changes in the socioeconomic gradient of women's entry into marriage (Fukuda, Raymo, and Yoda 2020; Kimura 2022). The changing norms, alongside decreasing gender inequality in the labor market, are expected to shift the relationship between women's employment and marriage or childbirth entry in a more positive direction (Dominguez-Folgueras and Castro-Martín 2008; Esping-Andersen and Billari 2015; Goldscheider, Bernhardt, and Lappegård 2015; Sweeney 2002). Although more careful and detailed analyses are needed to assess the contribution of increased employment instability for women, as discussed in Section 1, a two-sided approach that considers both women and men will provide a more complete picture of the role of employment instability in the postponement of family formation among more recent cohorts.

In addition to the inclusion of women, there are several limitations and directions for future research. First, the sample size is relatively small, leading to some uncertainty in the estimates. Using additional data with large sample sizes would allow for a more accurate assessment. Second, the use of retrospective survey day may introduce recall bias, as retrospective data collection is more likely to be oversimplified compared to prospective methods (Manzoni et al. 2010). This could lead to an underestimation of employment instability. Third, while the counterfactual simulation assumes that increasing employment instability across cohorts is exogenous, some unobserved factors may induce confounding. For example, if individuals in recent cohorts are more likely to have attitudes avoiding marriage and childbearing and are also more likely to choose nonstandard employment or nonemployment, the contribution of increased employment instability would be overestimated. Fourth, economic resources, such as income or earnings, were not considered in this analysis. Including these variables would allow for disentangling the effects of employment into economic or uncertainty-related factors (Kalmijn 2011; van Wijk, de Valk, and Liefbroer 2021). Fifth, it was not possible to distinguish between unemployment and inactivity. While it is expected that inactive youths in Japan face employment challenges similar to those of the unemployed, there is currently insufficient evidence on how unemployment and inactivity differ in their relationship with family formation behavior.[12] In summary, future studies could extend the findings of this paper by utilizing prospective panel surveys or administrative data that include detailed information on individual attitudes, economic resources, and specific employment characteristics.

[12] As suggestive evidence, Mugiyama (2024) shows that the likelihood of a non-cohabiting partnership entry, which is the transitory phase to marriage entry, is almost as low for the unemployed and the inactive in Japan.

In conclusion, this study provides evidence that men's increased employment instability has contributed substantially to the delay in family formation across cohorts. Understanding the sources of changes in the timing of union formation, marriage, or childbearing is an important research agenda for demographic and broader social science research. The framework presented in this paper could be applied to different contexts beyond Japan, as well as to other causes and consequences beyond the relationship between increased employment instability and entry into first marriage and first childbirth. Further studies are needed to advance our understanding of the postponement of family formation and to explore the broader implications of shifting economic and labor market conditions.

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Appendix

Figure A-1: Trends in nonstandard employment rate (%) among men in Japan, 1988–2022

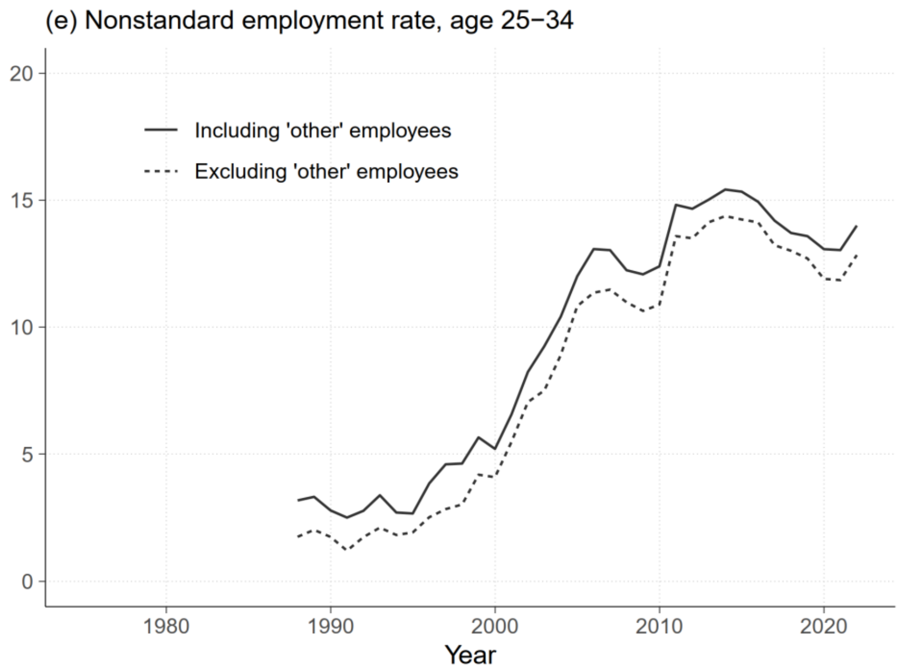
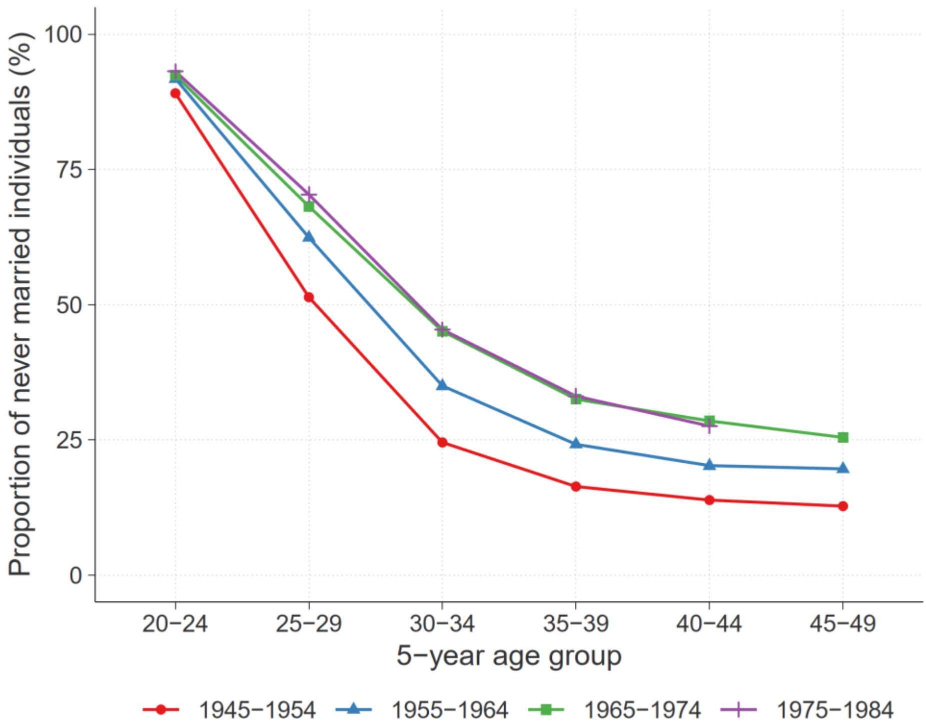


Table A-1: Distribution of employment status at the time of survey in 2015 Labor Force Survey and 2015 SSM

	10-year age group			
	30–39	40–49	50–59	60–69
2015 Labor Force Survey				
Executive	3.1	6.2	8.3	8.4
Regular employee	72.8	71.7	64.8	16.2
Part-time or temporary employee	4.3	2.6	3.1	9.6
Dispatched employee	1.6	1.2	0.8	1.0
Contract employee	3.4	2.3	2.9	6.0
Entrusted employee	0.3	0.2	0.7	6.5
Other employee	0.9	0.6	0.7	1.5
Self-employment	5.0	7.9	9.6	13.7
Family worker	1.0	0.6	0.4	0.2
Nonemployment	7.7	6.7	8.7	36.8
2015 SSM				
Executive	2.1	5.6	8.4	7.6
Regular employee	73.8	74.3	65.6	12.1
Part-time or temporary employee	3.3	3.4	2.7	11.9
Dispatched employee	2.1	0.7	1.4	1.3
Contract or entrusted employee	4.0	2.0	2.2	13.3
Self-employment	7.3	9.7	13.7	19.1
Family worker	4.0	1.6	0.2	0.0
Nonemployment	3.4	2.8	5.9	34.7
N in SSM	523	611	593	833

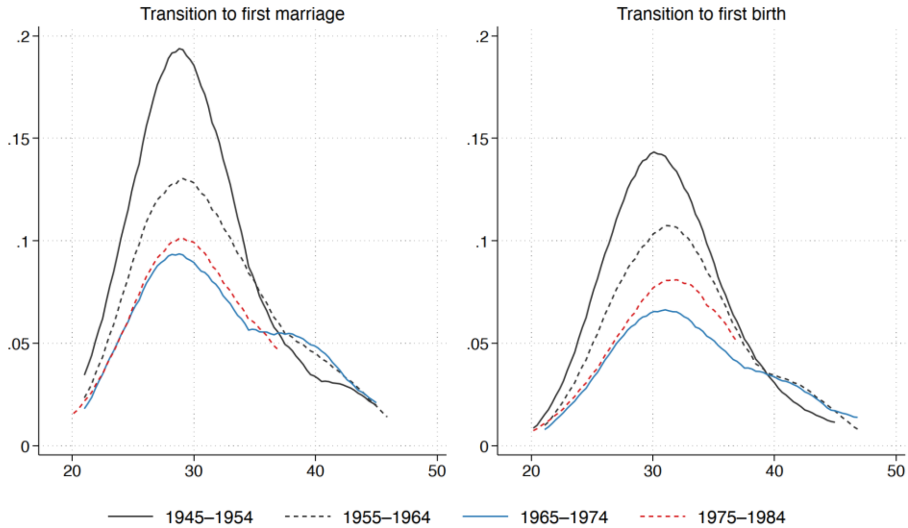
Notes: The data were retrieved from Labor Force Survey, 2015 (Statistics Bureau of Japan 2023) and 2015 SSM survey (author calculation).

Figure A-2: Proportion of individuals who never married by age and cohort in the Population Census



Notes: The data were retrieved from Population Census (Statistics Bureau of Japan 2023b). The proportion of individuals who never married is calculated from the corresponding year of the Population Census.

Figure A-3: Kernel density plots of hazard rate of transition to first marriage and first birth by cohort



Note: The bandwidth is set to two years.

Table A-2: Model comparisons in terms of age polynomials and interactions between cohort and age

	Log likelihood	*df*	AIC	BIC
Transition to first marriage (*N* = 33,824)				
(1a) C+A+A^2	–6932.78	6	13877.56	13928.13
(1b) C+A+A^2+A^3	–6857.14	7	13728.27	13787.28
(1c) C+A+A^2+A^3+A^4	–6848.85	8	13713.69	13781.13
(1d) C+A+A^2+A^3+A^4+A^5	–6842.72	9	13703.44	13779.30
(2a) C+A+A^2+A^3+CA	–6851.45	10	13722.90	13807.18
(2b) C+A+A^2+A^3+CA+CA^2	–6841.55	13	13709.09	13818.67
(2c) C+A+A^2+A^3+CA+CA^2+CA^3	–6839.62	16	13711.23	13846.09
Transition to first birth (*N* = 40,779)				
(1a) C+A+A^2	–6548.13	6	13108.26	13159.95
(1b) C+A+A^2+A^3	–6510.10	7	13034.20	13094.51
(1c) C+A+A^2+A^3+A^4	–6509.89	8	13035.77	13104.70
(1d) C+A+A^2+A^3+A^4+A^5	–6506.09	9	13030.17	13107.71
(2a) C+A+A^2+A^3+CA	–6504.33	10	13028.65	13114.81
(2b) C+A+A^2+A^3+CA+CA^2	–6494.60	13	13015.20	13127.21
(2c) C+A+A^2+A^3+CA+CA^2+CA^3	–6492.30	16	13016.59	13154.44

Notes: C represents cohort dummy variables, A represents age as a continuous variable, and CA represents the interaction term between cohort and age. AIC refers to Akaike information criterion and BIC refers to Bayesian information criterion.

Table A-3: Log-odds coefficients of discrete-time logit models predicting the transitions to first marriage and first birth (corresponding to Table 2)

	Transition to first marriage		Transition to first birth	
	Model 1	Model 2	Model 1	Model 2
Cohort (ref: 1945–1954)				
1955–1964	–.3922	–.3687	–.3468	–.3165
	(.0626)	(.0630)	(.0647)	(.0651)
1965–1974	–.6785	–.6126	–.7501	–.6757
	(.0652)	(.0659)	(.0693)	(.0700)
1975–1984	–.6431	–.4688	–.6293	–.4510
	(.0716)	(.0737)	(.0765)	(.0785)
Employment status (ref: regular employment)				
Self-employment		.2778		.2024
		(.1318)		(.1354)
Nonstandard employment		–.6890		–.4331
		(.1658)		(.1895)
Nonemployment		–1.9513		–1.5735
		(.3303)		(.3670)
Cumulative years in regular employment		.0426		.0795
		(.0168)		(.0184)
Cumulative years in self-employment		.0104		.0475
		(.0221)		(.0224)
Cumulative years in nonstandard employment		–.0125		–.0026
		(.0237)		(.0260)
Educational attainment (ref: university)				
Junior high	–.2156	–.3226	–.1982	–.5543
	(.0844)	(.1398)	(.0894)	(.1531)
High school	–.0804	–.2138	–.0249	–.3222
	(.0551)	(.0934)	(.0575)	(.1006)
Vocational school	–.0942	–.2073	–.0611	–.2646
	(.0871)	(.0981)	(.0936)	(.1062)
Junior college	.0615	–.0256	.3646	.1891
	(.1658)	(.1722)	(.1657)	(.1734)
Enrollment in school	–1.6736	–1.7001	–1.3250	–1.2848
	(.2395)	(.2403)	(.2991)	(.3008)
Age – 18	.8548	.8090	.9154	.8385
	(.0422)	(.0451)	(.0547)	(.0575)
$(\text{Age} - 18)^2$	–.0533	–.0524	–.0498	–.0492
	(.0033)	(.0034)	(.0040)	(.0041)
$(\text{Age} - 18)^3$	.0009	.0009	.0007	.0007
	(.0001)	(.0001)	(.0001)	(.0001)
Intercept	–5.5041	–5.3808	–6.8011	–6.5802
	(.1714)	(.1808)	(.2363)	(.2450)
N	33,824	33,824	40,779	40,779

Note: Standard errors in parentheses.

Table A-4: Average marginal effects of discrete-time logit models predicting transition to first marriage and first birth, using models taking interactions between cohort and age

	Transition to first marriage			Transition to first birth		
	Model 1	Model 2	Diff.	Model 1	Model 2	Diff.
Cohort (ref: 1945–1954)						
1955–1964	–.0246	–.0223	–.0024	–.0166	–.0144	–.0022
	(.0039)	(.0038)	(.0005)	(.0031)	(.0030)	(.0003)
1965–1974	–.0387	–.0340	–.0047	–.0309	–.0269	–.0040
	(.0037)	(.0037)	(.0006)	(.0029)	(.0028)	(.0004)
1975–1984	–.0373	–.0279	–.0095	–.0271	–.0195	–.0076
	(.0041)	(.0042)	(.0010)	(.0033)	(.0035)	(.0008)
Employment status (ref: regular employment)						
Self-employment		.0187			.0088	
		(.0094)			(.0066)	
Nonstandard employment		–.0314			–.0153	
		(.0055)			(.0055)	
Nonemployment		–.0551			–.0354	
		(.0036)			(.0038)	
Cumulative years in regular employment		.0021			.0031	
		(.0009)			(.0007)	
Cumulative years in self-employment		.0003			.0019	
		(.0012)			(.0009)	
Cumulative years in nonstandard employment		–.0007			–.0002	
		(.0013)			(.0010)	
N	33,824	33,824		40,779	40,779	

Notes: Standard errors in parentheses. Controls for educational attainment, cubic polynomials of age, and the interactions between polynomials of age and cohort. The "Diff." columns show the differences in AMEs between Models 1 and 2, following Mize, Doan, and Long (2019).

Table A-5: Correlation matrix between cumulative years in each employment status, share of years spent in each employment status, age and employment status dummies

	Cumulative years in regular employment	Cumulative years in self-employment	Cumulative years in nonstandard employment	Share of years in regular employment	Share of years in self-employment	Share of years in nonstandard employment	Age	Regular employment dummy (1/0)	Nonstandard employment dummy (1/0)	Self-employment dummy (1/0)
Sample of transition to first marriage										
Cumulative years in regular employment	1.000									
Cumulative years in self-employment	–.114	1.000								
Cumulative years in nonstandard employment	–.093	–.003	1.000							
Share of years in regular employment	.556	–.270	–.299	1.000						
Share of years in self-employment	–.174	.803	–.035	–.335	1.000					
Share of years in nonstandard employment	–.190	–.037	.769	–.384	–.056	1.000				
Age	.767	.241	.283	.191	.095	.087	1.000			
Regular employment dummy (1/0)	.458	–.276	–.260	.877	–.335	–.318	.141	1.000		
Self-employment dummy (1/0)	–.140	.708	–.010	–.291	.866	–.034	.110	–.380	1.000	
Nonstandard employment dummy (1/0)	–.137	–.035	.614	–.306	–.051	.774	.083	–.397	–.073	1.000
Sample of transition to first birth										
Cumulative years in regular employment	1.000									
Cumulative years in self-employment	–.121	1.000								
Cumulative years in nonstandard employment	–.111	–.002	1.000							
Share of years in regular employment	.561	–.308	–.314	1.000						
Share of years in self-employment	–.189	.824	–.030	–.377	1.000					
Share of years in nonstandard employment	–.207	–.039	.777	–.399	–.053	1.000				
Age	.781	.261	.238	.202	.115	.053	1.000			
Regular employment dummy (1/0)	.456	–.314	–.270	.867	–.376	–.324	.143	1.000		
Self-employment dummy (1/0)	–.146	.713	.001	–.321	.852	–.028	.130	–.436	1.000	
Nonstandard employment dummy (1/0)	–.147	–.037	.604	–.310	–.049	.756	.051	–.401	–.075	1.000

Note: Person-year data are used.

Table A-6: Average marginal effects of discrete-time logit models predicting transition to first marriage and first birth, using share of years spent in each employment status

	Transition to first marriage			Transition to first birth		
	Model 1	Model 2	Diff.	Model 1	Model 2	Diff.
Cohort (ref: 1945–1954)						
1955–1964	–.0248	–.0225	–.0024	–.0168	–.0150	–.0018
	(.0039)	(.0038)	(.0005)	(.0031)	(.0030)	(.0003)
1965–1974	–.0387	–.0343	–.0044	–.0311	–.0278	–.0033
	(.0037)	(.0036)	(.0005)	(.0029)	(.0028)	(.0004)
1975–1984	–.0371	–.0278	–.0093	–.0273	–.0204	–.0069
	(.0040)	(.0041)	(.0010)	(.0032)	(.0033)	(.0008)
Employment status (ref: regular employment)						
Self-employment		.0346			.0134	
		(.0126)			(.0079)	
Nonstandard employment		–.0256			–.0115	
		(.0063)			(.0062)	
Nonemployment		–.0484			–.0318	
		(.0046)			(.0046)	
Share of years spent in regular employment		.0586			.0599	
		(.0111)			(.0103)	
Share of years spent in self-employment		.0217			.0388	
		(.0155)			(.0131)	
Share of years spent in nonstandard employment		.0239			.0180	
		(.0149)			(.0135)	
N	33,824	33,824		40,779	40,779	

Notes: Standard errors in parentheses. Controls for educational attainment and cubic polynomials of age. The "Diff." columns show the differences in AMEs between Models 1 and 2, following Mize, Doan, and Long (2019).

Table A-7: Average marginal effects of discrete-time logit models predicting transition to first marriage and first birth, using lagged independent variables

	Transition to first marriage			Transition to first birth		
	Model 1	Model 2	Diff.	Model 1	Model 2	Diff.
One year lagged						
Cohort (ref: 1945–1954)						
1955–1964	–.0269	–.0243	–.0026	–.0178	–.0156	–.0021
	(.0042)	(.0041)	(.0004)	(.0033)	(.0032)	(.0003)
1965–1974	–.0420	–.0371	–.0049	–.0328	–.0290	–.0038
	(.0040)	(.0039)	(.0006)	(.0030)	(.0030)	(.0004)
1975–1984	–.0408	–.0309	–.0099	–.0291	–.0216	–.0076
	(.0043)	(.0044)	(.0011)	(.0033)	(.0035)	(.0008)
Employment status (ref: regular employment)						
Self-employment		.0037			.0113	
		(.0091)			(.0072)	
Nonstandard employment		–.0294			–.0131	
		(.0064)			(.0060)	
Nonemployment		–.0446			–.0361	
		(.0058)			(.0041)	
Cumulative years in regular employment		.0025			.0030	
		(.0010)			(.0008)	
Cumulative years in self-employment		.0020			.0014	
		(.0013)			(.0010)	
Cumulative years in nonstandard employment		–.0009			–.0008	
		(.0014)			(.0011)	
N	31,488	31,488		38,553	38,553	

	Transition to first marriage			Transition to first birth		
	Model 1	Model 2	Diff.	Model 1	Model 2	Diff.
Two years lagged						
Cohort (ref: 1945–1954)						
1955–1964	–.0296	–.0269	–.0027	–.0193	–.0171	–.0022
	(.0046)	(.0044)	(.0005)	(.0035)	(.0034)	(.0003)
1965–1974	–.0459	–.0410	–.0049	–.0354	–.0315	–.0039
	(.0043)	(.0042)	(.0006)	(.0032)	(.0031)	(.0004)
1975–1984	–.0447	–.0345	–.0102	–.0316	–.0238	–.0077
	(.0046)	(.0048)	(.0011)	(.0035)	(.0037)	(.0008)
Employment status (ref: regular employment)						
Self-employment		–.0051			.0048	
		(.0095)			(.0074)	
Nonstandard employment		–.0384			–.0198	
		(.0065)			(.0058)	
Nonemployment		–.0535			–.0355	
		(.0057)			(.0048)	
Cumulative years in regular employment		.0017			.0030	
		(.0011)			(.0009)	
Cumulative years in self-employment		.0018			.0017	
		(.0014)			(.0011)	
Cumulative years in nonstandard employment		–.0013			–.0006	
		(.0016)			(.0012)	
N	29,146	29,146		36,321	36,321	

Notes: Standard errors in parentheses. Controls for educational attainment and cubic polynomials of age. The “Diff.” columns show the differences in AMEs between Models 1 and 2, following Mize, Doan, and Long (2019).

Table A-8: Descriptive statistics for the sample of married men

	Person-year	At last observation
Transition to first birth	0.186	0.853
Cohort		
1945–1954	0.353	0.361
1955–1964	0.252	0.245
1965–1974	0.254	0.227
1975–1984	0.141	0.167
Employment status		
Regular employment	0.852	0.860
Nonstandard employment	0.105	0.102
Self-employment	0.030	0.028
Nonemployment	0.007	0.005
Enrollment in school	0.006	0.005
Cumulative years in regular employment	11.290 (6.965)	11.189 (6.679)
Cumulative years in nonstandard employment	1.165 (3.662)	1.071 (3.560)
Cumulative years in self-employment	0.628 (2.274)	0.589 (2.235)
Age	32.161 (6.515)	31.805 (6.435)
Educational attainment		
Junior high	0.102	0.102
High school	0.434	0.433
Vocational school	0.096	0.096
Junior college	0.013	0.022
University	0.349	0.342
Enrollment in school	0.006	0.005
N	8,101	1,763

Notes: Means and proportions are reported. Standard deviations are shown in parentheses. The “Person-year” column shows summary statistics in person-year observations. The “at last observation” column shows summary statistics at the time of entry into first birth for those who have ever experienced these transitions and at the last observation period for those who have never experienced these transitions.

Table A-9: Log-odds coefficients of discrete-time logit models predicting the transition to first birth among married men (corresponding to Table 3)

	1	Model 2
Cohort (ref: 1945–1954)		
1955–1964	–.0262	–.0217
	(.0775)	(.0778)
1965–1974	–.2731	–.2632
	(.0816)	(.0820)
1975–1984	–.1831	–.1328
	(.0917)	(.0933)
Employment status (ref: regular employment)		
Self-employment		.0522
		(.1504)
Nonstandard employment		.2141
		(.2339)
Nonemployment		–.3605
		(.4864)
Cumulative years in regular employment		.0642
		(.0230)
Cumulative years in self-employment		.0453
		(.0267)
Cumulative years in nonstandard employment		.0195
		(.0305)
Educational attainment (ref: university)		
Junior high	.0570	–.3224
	(.1099)	(.1904)
High school	.0612	–.2044
	(.0680)	(.1217)
Vocational school	.0193	–.1456
	(.1088)	(.1263)
Junior college	.7703	.6022
	(.2108)	(.2194)
Enrollment in school	–.4649	–.2186
	(.4131)	(.4204)
Age – 18	.3470	.2872
	(.0788)	(.0818)
$(\text{Age} – 18)^2$	–.0187	–.0184
	(.0056)	(.0056)
$(\text{Age} – 18)^3$	.0002	.0002
	(.0001)	(.0001)
Constant	–2.8649	–2.6968
	(.3555)	(.3626)
N	8,101	8,101

Note: Standard errors are in parentheses.