

2026 Spring Semester

EAP

Master's Thesis

The effect of Japan's Ageing population
on Wage Trajectories of Young Workers

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September 2026

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Abstract

JEL classification: J11, J13, J21, J82

Keywords: ageing population, wage trajectories, cohort size, labor market, Japan

This research examines how Japan's ageing population affects wage trajectories within the regular employment tier, focusing on whether early-career workers experience a systematically different wage response to the growing presence of elderly workers in the regular employment tier. Using individual-level panel data from the Japanese Panel Study of Employment Dynamics (2016–2024) combined with education-group-specific demographic measures from the Labour Force Survey, I estimate mixed-effects models with a Mundlak correction. Results show that a one-standard-deviation increase in education-group-specific elderly regular employment share is associated with approximately a 1.5 percent wage reduction for the general regular workforce, consistent with systemic wage compression driven by institutional adjustment to population ageing. Workers aged 20–29, however, experience a net wage association of approximately 8.2 percent per standard deviation, consistent with a cohort-scarcity effect at labor market entry for scarce young entrants. These opposing associations reflect institutionally distinct channels, a structural wage compression across the seniority hierarchy and a youth-specific wage differential concentrated among early-career workers. The findings suggest that population ageing generates asymmetric wage consequences across the career, with implications for the design of wage-system and social security reform in Japan.

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1 Introduction

As of 2024, individuals aged 65 and over account for 29.3 percent of Japan’s total population, a share projected to reach roughly 38 percent by 2070, with one in four Japanese citizens expected to be aged 75 or older by that point (Cabinet Office, Government of Japan, 2025). Moriyama (2025) shows that the elderly population ratio rose from 9.1 percent in 1980 to 17.4 percent in 2000 and 28.7 percent in 2020, while the number of working-age persons aged 15–64 per older person fell from 7.4 to 2.1 over the same period. Since Japan operates a pay-as-you-go public pension system, this demographic shift has placed strong fiscal pressure on social security (Kondo & Shigeoka, 2017).

In response to labor shortages and pension pressure, the Japanese government amended the Act on Stabilization of Employment of Elderly Persons (EESL) in several stages (Government of Japan, 1971). The 2004 amendment, effective from 2006, made continued employment beyond mandatory retirement age a legal obligation, while the 2012 amendment extended this right to all workers wishing to remain employed (Moriyama, 2022). Over the same period, firms increasingly shifted from seniority-based wages toward performance- and job-based pay systems (Nishimura, 2016). Together, these institutional changes flattened the age-wage profile and compressed wages across the seniority hierarchy (Hamaaki et al., 2012).

Prior research has examined demographic wage effects mainly from two angles. One strand focuses on aggregate age-wage profile compression and cohort-size effects (Hamaaki et al., 2012), while another looks at the short-run employment effects of elderly employment policies (Kondo, 2016). No research, however, has directly examined whether early-career regular workers face a systematically different wage response to the long-run accumulation of elderly regular employment on the labor market.

This research addresses two questions: (1) whether growth in education-group-specific elderly regular employment share is associated with lower wages for the general regular workforce, and (2) whether workers aged 20–29 experience a differential wage response to the same demographic variation. To answer these questions, the analysis combines individual-level panel microdata from the Japanese Panel Study of Employment Dynamics (JPSED, 2016–2024) with education-group-specific demographic measures constructed from the Labour Force Survey Detailed Tabulation, and estimates mixed-effects models with a Mundlak correction.

The results show that a one-standard-deviation increase in education-group-specific elderly regular employment share is associated with roughly a 1.5 percent reduction in wages for the general regular workforce, consistent with system-wide wage compression. For workers aged 20–29, the same demographic variation is associated with roughly 8.2 percent higher wages, consistent with a cohort-scarcity effect at labor market entry for scarce young entrants. These findings suggest that population ageing generates opposing wage associations across the career hierarchy through institutionally distinct channels.

The remainder of this research is structured as follows: Section 2 reviews the literature on Japan’s employment system and demographic wage effects. Section 3 develops the theoretical framework and hypotheses. Section 4 describes the data and empirical strategy. Section 5 presents the estimation results and diagnostics. Section 6 discusses the findings, and Section 7 concludes.

2 Literature Review

2.1 Changes in Japan's Employment and Wage Systems Due to the Ageing Population

Lifetime employment (*shūshin koyō*), seniority wages (*nenkō chingin*), and internal labor markets are central to understanding the Japanese employment system. Koike (1988) describes how large Japanese corporations typically recruit graduates as generalists, offering seniority-based wage increases through their fifties and guaranteed employment until mandatory retirement at age 60. This lifetime employment arrangement is not merely a welfare provision but a productivity mechanism, firms build firm-specific skills through long-term investment in their workforce, while workers remain motivated by the prospect of rising wages tied to continued tenure. The micro-econometric underpinning of this pattern is examined by Fukao et al. (2006), who use matched employer-employee data constructed by linking the Census of Manufacturers and the Basic Survey of Wage Structure through the Establishment and Enterprise Census, covering Japanese manufacturing establishments over the period 1993–2003, estimates productivity-tenure and wage-tenure profiles simultaneously to distinguish between three competing theories of upward-sloping wage-tenure profiles which are firm-specific human capital accumulation, job matching quality, and deferred compensation under information asymmetry. Their results consistently show a steeper wage-tenure profile than productivity-tenure profile across manufacturing establishments. Workers receive approximately 10–15 percent less than their marginal product early in their careers, wages surpass productivity around 20–22 years of tenure, and by the time of mandatory retirement wages exceed productivity by roughly 20 percent. This pattern of backloaded wage payment is inconsistent with the human capital model, which predicts a steeper productivity profile than wage profile, and with the matching model, which predicts the two profiles to overlap, but is consistent with the deferred compensation contract proposed by Lazear (1979).

This wage structure operates within the broader framework of Japan's internal labor market. Date et al. (2024) confirms that the internal labor market within which this deferred-compensation structure operates remains empirically distinct from the external labor market in contemporary Japan. Using microdata from the Basic Survey on Wage Structure covering 1989–2021, estimating a finite mixture model that sorts full-time workers into two latent classes based on educational background, firm size, gender, and age, they identify an internal labor market class and an external labor market class with distinct wage-setting mechanisms. Workers in the internal class — disproportionately male, university-educated, and concentrated in large firms with 1,000 or more employees — receive wages that follow a seniority-based schedule insulated from short-term market conditions: in the 2013–2021 period, neither the output gap nor sectoral labor market tightness has any statistically significant effect on internal-class wage growth, while long-term growth expectations do. Workers in the external class — concentrated in small firms, with higher shares of female and lower-educated workers — face wages that respond directly to supply and demand, the output gap and employment conditions index both retain statistically significant positive effects on wage growth in the same period, consistent with market-driven wage setting. These differences were also visible during the COVID-19 shock, when both classes saw declining wage growth but the disruption was substantially larger in the external class. Kobayashi (1981) adds that within this internal labor market structure, external hiring is predominantly concentrated at the lowest-rung entry positions, with advancement occurring primarily through internal promotion.

As noted in the introduction, population ageing puts pressure on these institutions. To address labor shortages and pension pressure, the Japanese government amended the Act on Stabilization of Employment of Elderly Persons (EESL) in several stages (Government of Japan, 1971). Moriyama (2022) summarizes that the 1990 revision required employers to make efforts to ensure employment

until age 65, the 2004 amendment (enacted in 2006 (Kondo & Shigeoka (2017))), made continued employment a legal obligation, and the 2012 amendment extended this right to all workers wishing to continue rather than only a selected subset. Kondo & Shigeoka (2017) shows that the 2006 EESL revision raised the employment rate of men in their early 60s. Among men born in 1946 —the first cohort subject to the mandate —the share in salaried work rose by 2.4 percentage points at age 60 and 3.2 percentage points at age 61 relative to the unaffected 1945 cohort. These gains reflect the policy’s purpose: filling the gap between a mandatory retirement age of 60 and a gradually rising pension eligibility age.

Firms responded through several institutional measures. Ministry of Health, Labour and Welfare (2020) report that among firms with 31 or more employees, 2.7 percent had abolished mandatory retirement and 20.9 percent had raised the retirement age, while the majority (76.4 percent) relied on continued-employment programs. The most common arrangement was *sai koyō* (reemployment), under which the employment contract ends at age 60 and a new contract is offered, typically at reduced wages (Kondo, 2016). A related practice is *yakushoku teinen*, defined by Ohki (2018) as a system in which managers at each hierarchical level step down from line management positions and are reassigned to specialist or other equivalent roles. In large corporations, it is said that the system was largely introduced in two waves: first, from around the 1980s, in conjunction with the transition from a mandatory retirement age of 55 to 60, primarily with the aims of maintaining organizational renewal and vitality (for the purpose of developing the next generation) and suppressing rising labor costs; and second, from the 1990s onward, with the aim of resolving post shortages accompanying the ageing of the workforce composition. Drawing on Kondo and Shigeoka (2017), the 2004 amendment (enacted in 2006) to the EESL increased the number of employees continuing to work past the standard mandatory retirement age of 60 at large enterprises that had uniform retirement policies. Among small and medium-sized firms, however, take-up of the formal reemployment or employment-extension schemes mandated by the law was limited, largely because those firms were already employing older workers past age 60. Moriyama (2022) extends this picture by arguing that careers after mandatory retirement are stratified by enterprise size and pre-retirement employment status, with outcomes reflecting the cumulative accumulation of advantage and disadvantage over the working life. Notably, he finds that this stratification has a pronounced gender dimension, with men’s earnings exceeding women’s across all patterns of post-60 employment transition.

Nishimura (2016) traces a transition in Japanese wage-setting from seniority-based pay through competency-based pay (*shokkunō-kyū*) toward performance-based systems from the 1990s onward, driven by a shift in design philosophy from supply-side considerations —employee capability and motivation —to demand-side logic centred on corporate strategy and business performance, while retaining core features of the internal labor market. Evidence from the Japan Productivity Center survey, reported by The Japan Institute for Labour Policy and Training (2022), indicates that these changes became increasingly widespread during the 2000s and 2010s. Among management-level employees, the share of firms adopting job-based pay systems increased from 21.1 percent in 1999 to 78.5 percent in 2018. Among non-management employees, the prevalence of seniority-based wage systems declined steadily and fell below 50 percent after 2016. These trends suggest that the weight of age and years of service in determining wages is declining.

The most direct evidence on how population ageing has reshaped Japan’s wage structure comes from studies of the age-wage profile. Hamaaki et al. (2012) examine changes in the age–wage profile using annual microdata from the Basic Survey on Wage Structure for 1989–2008, focusing on male lifetime employees (those hired directly after graduation who remained with the same firm). They find that the seniority wage slope flattened gradually through the 1990s and developed a pronounced “kink” around age 40 by 2007–2008, especially for university graduates in the non-

manufacturing sector. To determine whether this flattening reflected a temporary cohort-specific phenomenon or a structural shift, they plot cohort-specific age–wage profiles and show that more recent cohorts also experience a flatter profile from around age 40, indicating that the pattern is an age effect rather than a cohort effect. They interpret this as long-run institutional adjustment to productivity slowdown and population ageing, not cyclical noise.

The institutional adjustments documented above establish that population ageing has fundamentally reshaped Japan’s internal wage-setting environment, yet whether the effect on the wages of younger workers entering and progressing through the same seniority hierarchy remains an unresolved question that the existing literature has not directly addressed.

2.2 Empirical Studies on the Effect of Population Ageing on Youth Wages

Okamura (2001) applies Welch (1979)’s Career-Phase Model to Japan using the Basic Survey on Wage Structure for 1988–1995, constructing industry-specific cohort-size measures for male standard workers in manufacturing, wholesale/retail, and finance/insurance. The model posits that workers enter as “learners” and transition to skilled “workers”; because the two phases are imperfect substitutes, a larger entering cohort depresses entry wages, an effect that fades as workers gain experience. Okamura’s findings confirm that these negative cohort-size effects are concentrated in the early-career learner phase, diminishing with experience and varying systematically across industries and education groups: significant wage-depressing effects emerge for both high school and university graduates in white-collar manufacturing, but for university graduates only in finance and insurance.

Inoue (2022) uses the Statistical Survey of Actual Status for Salary in the Private Sector for wage data and the Labour Force Survey for labor inputs, covering the period from the 1990s through the 2010s with the core decomposition spanning 2000 to 2019. Initial OLS and IV estimates proved unstable across sample periods, likely due to omitted productivity change, so Inoue adopts the Feenstra (1994) identification strategy, which exploits the orthogonality of demand and supply shocks in a simultaneous-equation framework to recover a robust estimate. The estimated elasticity of substitution across age groups is approximately 4.2 for male workers (covering both regular and non-regular employees, assuming the elasticity of substitution is constant across groups of labor inputs), implying that a 1 percent increase in an age group’s relative labor supply reduces its relative wage by about 0.24 percent. In a robustness check restricting the sample to male regular workers employed in establishments with at least 10 employees, the estimated elasticity declines to 2.49, suggesting that age groups are somewhat less substitutable within the regular employment sector (Inoue, 2022). Using these structural parameters, Inoue decomposes the observed flattening where the wage ratio of men aged 35–39 to those aged 25–29 fell from roughly 1.4 in the 1990s–2000s to 1.27 in the 2010s, and the ratio of men aged 45–49 to those aged 25–29 declined from about 1.7 to 1.57. Crucially, the demographic contribution differs sharply by period: ageing accounts for only about one-third of the wage-profile change from 1980 to 2000, when cohort replacement (the wartime generation giving way to the postwar generation) increased the education and work experience of workers in the 45–59 age group and steepened the profile. After 2000, ageing explains more than 80 percent of the flattening, as the demographic channel became dominant while productivity-driven steepening subsided. The research’s central finding is that the flattening of the age–wage profile after 2000 is primarily a structural response to workforce ageing.

An important methodological implication of the cohort-size literature is that demographic wage effects are concentrated within education groups. Kawaguchi & Mori (2014) apply the double CES framework of Card & Lemieux (2001) to Japan using Labour Force Survey data covering 1986–2008. The production function contains two nested imperfect-substitution relationships. At the upper tier, college and high-school labor are combined imperfectly, with the elasticity of substitu-

tion σ_e measuring how easily firms can substitute one education group for the other. At the lower tier, workers of different ages within each education group are also imperfect substitutes, with the elasticity σ_a measuring how easily firms can substitute younger and older workers within the same educational category. Because substitution is imperfect at both levels, changes in birth-cohort size primarily affect wages within the corresponding education–age cell rather than uniformly across the labor market. Estimation for Japan yields an age-group elasticity of approximately $\sigma_a \approx 5.7$, which is close to the estimate of about 5 reported by Card & Lemieux (2001) for the United States. This similarity suggests that demographic supply shifts operate through comparable mechanisms in both countries and confirms that college and high-school graduates are imperfect substitutes in Japan (Kawaguchi & Mori, 2014).

Beyond the education-group framework, cohort-size effects also depend on institutional context. Brunello (2010) find that a 10 percent increase in cohort size reduces hourly earnings by 0.7 percent for high school graduates and 1.7 percent for college graduates, with effects significantly larger in Southern Europe. The mechanism is substitutability across age groups. When a large cohort enters, its wages fall — but how much depends on whether firms can hire a different age group instead. Strong employment protection blocks that substitution by making it costly to replace older insiders with younger outsiders. Lower substitutability means any given demographic supply shock produces a larger wage effect, because the excess supply cannot be absorbed by shifting demand across age groups. This is why the same demographic shift hits Southern European wages harder than Northern. Any institution that reduces age-group substitutability — such as deferred-compensation contracts and lifetime employment in Japan — similarly amplifies cohort-size wage effects.

While the studies above track demographic wage effects operating over long horizons, a separate line of evidence examines the shorter-run employment effects of specific policies that retain older workers in Japan. Kondo (2016) studies the 2006 EESL amendment, which mandated continued employment until the pension eligibility age, using a difference-in-differences design comparing establishments with higher versus lower shares of workers aged 55–59 in 2003, interacted with year fixed effects. The employment analysis uses an establishment panel from the Employment Trend Survey (2002–2011), while the earnings analysis pools worker-level microdata from the Basic Survey on Wage Structure (1998–2011). Across all specifications, Kondo (2016) finds no evidence of substitution between elderly and young full-time workers. In fact, establishments with a higher pre-reform share of workers aged 55–59 show a significantly positive effect on full-time employees under 25. This reflects a mechanical pattern: as incumbent workers reach mandatory retirement age, some exit even under the mandate, and firms hire young replacements. The earnings analysis reveals a separate adjustment channel, where the earnings drop at age 60 widened substantially for baby-boomer cohorts reaching 60 after 2006. Relative to the 1945 cohort, the decline was 4.5 percentage points larger for the 1948 cohort and 7.7 points for the 1949 cohort, concentrated in large firms (10.4 points for the 1949 cohort in firms with 500 or more employees). Critically, Kondo (2016) tests whether these wage cuts fed back into youth hiring and finds no such link, the two adjustments operated as independent responses to the mandate, not as a trade-off.

The studies reviewed above leave one question directly unaddressed. The literature on age-wage profile compression documents that population ageing reshapes wage structures across the entire seniority hierarchy, while the cohort-size literature establishes that demographic supply shifts affect wages most strongly at labor market entry. Yet no research has directly examined whether these two forces generate asymmetric wage responses across the career — specifically, whether early-career regular workers experience a systematically different wage trajectory from the broader regular workforce as elderly regular employment expands over the long run within Japan’s internal employment system.

3 Research Framework and Hypotheses

The studies reviewed above establish that population ageing has reshaped Japan’s employment and wage institutions, yet they leave a central question open. Kondo (2016) examines the short-run employment effects of the EESL amendments and finds no displacement of younger workers within five to six years of the reform, but this does not rule out a wage effect — short-run policy evaluations precede the cumulative demographic pressure that only becomes visible over longer horizons, as the gradual unfolding of age-wage profile flattening over two decades documented by Hamaaki et al. (2012) and the finding that ageing explains more than 80 percent of the flattening after 2000 (Inoue, 2022) suggest. The present research therefore examines the long-run relationship between the expansion of elderly regular employment and the wages of younger regular workers within Japan’s internal labor markets.

The analysis restricts attention to regular employees. The deferred-compensation framework is a feature exclusive to regular employment. Non-regular workers are generally excluded from seniority-based pay and face minimal wage growth over their careers, meaning there is no steep age-wage profile to compress (Fukao et al., 2006; Date et al., 2024). The entry-level wage-setting mechanism similarly belongs to the regular employment tier. As youth cohorts shrink, firms compete more intensely to recruit scarce entrants into career-track regular positions. Because these hires represent long-term deferred-compensation relationships, firms have incentives to attract and retain young workers, creating a bidding process that has no direct equivalent for fixed-term non-regular employment. The institutional pressure from population ageing is likewise concentrated in regular employment, since the EESL mandates continued employment specifically for workers already embedded in the regular employment system, and the resulting cost pressure falls on firms’ internal wage-setting structures governing regular employment (Moriyama, 2022; Kondo, 2016).

Within the regular employment sector, demographic wage effects are best measured at the level of education groups. Card & Lemieux (2001) show that demographic wage effects are concentrated within educational tracks because workers of the same education but different ages are imperfect substitutes, while workers across education levels are also imperfect substitutes. Kawaguchi & Mori (2014) confirm that this structure holds in Japan, finding an elasticity of substitution across age groups of approximately 5.7 and that college and high-school graduates are imperfect substitutes. A key implication is that a relative increase in older workers within a given education group compresses the wages of younger workers in the same group, because imperfect substitutability prevents the two age groups from being fully substituted for one another. Education therefore enters the analysis in two distinct roles: as an individual-level control capturing human-capital differences in wages, and as the dimension along which demographic variation is measured. The wage effects described below are average effects for the pooled regular workforce conditional on individual-level education controls — the fact that identification exploits cross-education variation does not imply that demographic effects are expected to differ systematically across education groups.

The first mechanism concerns the general effect on the regular workforce. A standard supply-side account would predict that an increase in elderly workers lowers only the wages of their closest substitutes. The mechanism operating here instead draws on a firm-level institutional logic rooted in deferred compensation, which predicts that an ageing workforce compresses wages across the entire seniority hierarchy rather than only for the elderly.

The proposed mechanism runs as follows. Deferred compensation creates wage-tenure profiles that rise faster than productivity: younger workers are paid below and older workers above their marginal product (Fukao et al., 2006). As the share of elderly regular workers grows, a larger proportion of the workforce enters the overpaid phase of this life-cycle contract, raising aggregate

labor costs at the firm level. The EESL mandates continued employment until age 65, but it does not specify the terms, and most firms comply through a reemployment system (*sai koyō*) rather than by extending the retirement age. Under this system, workers undergo mandatory retirement at 60 and are rehired on fixed-term contracts at substantially reduced wages (Kondo, 2016). Not all elderly workers transition to reemployment, however. A small share of elderly who remain on regular contracts (rather than transitioning to reemployment) continue to receive seniority-based wages (Moriyama, 2022). These retained elderly regular workers continue to receive seniority-based wages reflecting their position in the deferred-compensation schedule. Because cutting their nominal wages directly would breach the implicit contract and undermine motivation across the firm, firms instead reform the wage structure itself (Nishimura, 2016). The share of firms using age and tenure as key wage determinants has declined substantially, accompanied by a shift toward role-based pay systems (The Japan Institute for Labour Policy and Training, 2022). This restructuring compresses wages across the hierarchy simultaneously: the age-wage profile flattens and returns to tenure decline (Hamaaki et al., 2012). The mechanism therefore affects all workers in the internal hierarchy, not only the elderly cohort whose growing share triggered the reform. The effect accumulates gradually, becoming substantial only after decades of sustained demographic change (Inoue, 2022), which explains why it does not appear in research exploiting short-run policy variation (Hamaaki et al., 2012). Based on this mechanism, the following hypothesis is derived: **H1** (*General internal labor market effect*): Conditional on individual characteristics including education, an increase in elderly regular employment share within an education group is associated with lower wages for the general regular working-age population.

The second mechanism concerns early-career regular workers specifically. Kobayashi (1981) establishes that internal labor markets predominantly hire at the lowest job-ladder rung, making entry the primary point of contact between workers and the external labor market. Date et al. (2024) confirm empirically that at this contact point wages are determined by supply and demand rather than internal seniority schedules, while wages of workers already inside the internal market are insulated from external conditions. Two conditions jointly determine why this generates a wage differential specific to young regular workers and does not extend to mid-career or older workers.

First, cohort scarcity operates only at the point of hire. Starting wages for new graduates are determined through external market competition rather than internal seniority schedules. When youth cohorts shrink, firms compete more strongly for scarce young entrants, bidding up starting wages (Okamura, 2001). This mechanism is amplified by Japan's deferred-compensation institutions: Brunello (2010) shows that employment protection and deferred-compensation contracts reduce substitutability across age groups, intensifying cohort-size wage effects precisely in the institutional context that characterises Japan's internal labor market. This mechanism does not apply to mid-career workers because once a worker has been hired and wage progression becomes governed by the internal seniority schedule and firm-specific human capital accumulation, external labor supply no longer resets their wage (Date et al., 2024). The economic rationale aligns with this convention: using OECD panel data spanning 1960-2010, Kawaguchi & Murao (2014) demonstrate that entry-condition effects follow a systematic duration pattern—the scarring penalty on subsequent employment is strongest during the first decade after entry, then attenuates. Their framework further shows that this persistence is amplified in countries with stricter employment protection legislation, implying that Japan's internal labor markets, which insulate regular workers from external competition and slow job-to-job reallocation, are precisely the institutional environment in which initial conditions cast the longest shadow. A worker entering at age 22 therefore carries the initial wage stamp until approximately age 30-32, and the 20-29 window captures the years during which the entry-cohort effect has not yet been eroded by internal progression. Using age 29 as the upper bound is conservative and aligns with the definition of youth (OECD, 2017).

Second, elderly retained workers and new entrants occupy institutionally distinct positions within the firm hierarchy. Fukao et al. (2006) show that firms using life-cycle contracts can set lower entry wages because the prospect of future above-productivity pay deters shirking; since younger workers have longer expected tenure, deferred compensation provides stronger incentives for them, making young entrants preferred hiring targets. Research exploiting policy-induced variation in elderly employment finds no evidence of substitution between young and elderly workers (Kondo, 2016), confirming limited direct competition at the entry level. Workers already embedded within the seniority hierarchy instead face the full compression pressure operating across adjacent hierarchical levels. This institutional separation explains why the scarcity premium can materialise without being offset by the wage compression that operates across the broader seniority hierarchy. Because cohort scarcity operates only at the point of hire and institutional separation holds only between entrants and retained elderly workers — not between mid-career workers and elderly workers, who are closer substitutes within the internal hierarchy — neither condition extends to workers in their thirties or beyond. The differential wage response is therefore specific to early-career workers at labor market entry. Based on this mechanism, the following hypothesis is derived:

H2 (Youth-specific wage differential): Conditional on the general internal labor market effect, an increase in elderly regular employment share within an education group is associated with a positive wage differential for early-career regular workers aged 20–29 relative to the broader regular working-age population within the same education group.

4 Data and Model

4.1 Model Specification

The hypotheses derived in Section 3 are tested using Equation 1, which is constructed as follows.

$$\ln(\text{Income}_{iet}) = \beta_0 + \gamma \text{Young}_{it} + \beta_1 \text{cEES}_{e,t}^{\perp \text{edu}} + \delta \left(\text{Young}_{it} \times \text{cEES}_{e,t}^{\perp \text{edu}} \right) + \beta_2 t + \boldsymbol{\theta}' \bar{\mathbf{X}}_i + \boldsymbol{\lambda}' \mathbf{Z}_{it} + \alpha_i + \varepsilon_{it} \quad (1)$$

To test **Hypothesis 1**, the analysis requires a measure of elderly regular employment that varies across observable dimensions beyond pure time-series variation. A national aggregate measure varying only over time would be collinear with the secular time trend included in the specification. Therefore, the identifying variation is drawn from cross-education differences in elderly regular employment shares using the Labour Force Survey Detailed Tabulation.

Following Card & Lemieux (2001) and Kawaguchi & Mori (2014), this approach exploits the fact that demographic wage effects are concentrated within educational tracks because workers across education levels are imperfect substitutes, while workers of the same education but different ages are also imperfect substitutes. The education-group level therefore represents the appropriate dimension along which to measure demographic variation in Japan’s labor market.

The demographic variable is constructed in five steps. Throughout this section, e indexes education groups and t indexes survey years. The estimation sample is restricted to workers aged 20–64.

The elderly population is defined as workers aged 60–64. This age group is chosen because the Act on Stabilization of Employment of Elderly Persons sets the minimum mandatory retirement age at 60 while requiring employers to secure continued employment until age 65, making this five-year band the population directly affected by the law’s employment protection provisions (Government of Japan, 1971).

First, the number of elderly regular workers is defined as the population aged 60–64 within education group e and year t :

$$\text{Elderly}_{e,t} = N_{e,60-64,t} \quad (2)$$

Second, the working-age regular workforce is defined as the sum of all workers aged 20–59 within the same education group and year:

$$\text{WorkingAge}_{e,t} = \sum_{a \in \{20-24, 25-29, 30-34, 35-39, 40-44, 45-49, 50-54, 55-59\}} N_{e,a,t} \quad (3)$$

Third, the raw elderly employment share is computed as the ratio of elderly regular workers aged 60–64 to the working-age regular workforce aged 20–59:

$$\text{EES}_{e,t} = \frac{\text{Elderly}_{e,t}}{\text{WorkingAge}_{e,t}} \quad (4)$$

Fourth, the raw elderly employment share is orthogonalized with respect to education-group fixed effects. This is done by estimating the following auxiliary regression at the education-group-by-year level:

$$\text{EES}_{e,t} = \mu + \eta_e + u_{e,t} \quad (5)$$

where η_e denotes education-group fixed effects. These fixed effects absorb the average level of elderly employment share specific to each education group. The residual $u_{e,t}$ therefore captures the deviation of an education-group-year observation from its own education-group average. The education-adjusted elderly employment share is defined as this residual:

$$\text{EES}_{e,t}^{\perp \text{edu}} = \hat{u}_{e,t} = \text{EES}_{e,t} - \hat{\mu} - \hat{\eta}_e \quad (6)$$

The residual in Equation 6 removes all time-invariant differences in elderly employment shares across education groups while retaining within-education-group variation over time. Finally, the residualized measure is centered and standardized:

$$\text{cEES}_{e,t}^{\perp \text{edu}} = \frac{\text{EES}_{e,t}^{\perp \text{edu}} - \overline{\text{EES}^{\perp \text{edu}}}}{\hat{\sigma}_{\text{EES}^{\perp \text{edu}}}} \quad (7)$$

This combination of individual-level outcome data with a regressor varying at the education-group-by-year level follows the grouped-data framework of Angrist & Pischke (2010), which shows that this multi-level structure is a valid identification strategy provided that individual-level covariates purge compositional shifts and that inference accounts for within-group error correlation.

The identifying assumption underlying this strategy is that within-education-group annual deviations in elderly regular employment share are exogenous to individual wages conditional on the included controls, random intercepts, and Mundlak correction terms (Mundlak, 1978). This assumption would be violated if education-group-specific wage shocks simultaneously affected both wages and firms' decisions to retain elderly regular workers within the same year. Although this assumption cannot be tested directly, two considerations support its plausibility: First, the expansion of elderly regular employment is driven primarily by EESL compliance obligations determined by legislation and pension eligibility ages rather than by contemporaneous wage conditions. Second, the Mundlak correction limits the scope for individual-level confounding by controlling for correlation between unobserved heterogeneity and time-varying regressors. Nonetheless, be-

cause no valid external instrument is available, the estimates should be interpreted as conditional associations rather than causal effects.

Hypothesis 2 predicts that this general effect is accompanied by a positive wage differential for early-career workers. This is tested through the interaction term:

$$\delta \left(\text{Young}_{it} \times \text{cEES}_{e,t}^{\perp \text{edu}} \right), \quad (8)$$

whose coefficient captures whether workers aged 20–29 exhibit a systematically different wage response to within-education-group variation over time in elderly employment share relative to the broader regular workforce.

This specification is chosen for two reasons: first, the mixed-effects framework naturally accommodates the hierarchical structure of the data — individual-level outcomes combined with a group-level regressor — through individual-specific random intercepts α_i that absorb persistent unobserved differences in earning capacity across workers, second, the unbalanced panel structure of the JPSED, in which individuals enter and exit the sample across survey waves, can be handled without listwise deletion, avoiding disproportionate exclusion of the most demographically dynamic observations.

Individual-level covariates — gender, age, education, weekly working hours, firm size, industry, and occupational position — increase precision and absorb compositional differences across education-by-year cells, while standard errors are clustered at the education-by-year level because the effective sample size is determined by the number of such cells rather than the number of individual observations (Angrist & Pischke, 2010).

To address potential correlation between time-varying covariates and individual-specific random effects, the specification applies a selective Mundlak correction (Mundlak, 1978). Mundlak (1978) demonstrates that the within estimator remains valid because it exploits within-unit variation that is orthogonal to unit-level confounds by construction. The GLS estimator, when augmented with individual-level means of time-varying regressors, collapses to the within estimator and separates within-individual wage responses from persistent between-individual differences in career trajectories, firm sorting, and labor-supply behaviour.

4.2 Definition of Variables

The dependent variable, as defined in Equation 1, $\ln(\text{Income}_{iet})$, represents the natural logarithm of real labor income for individual i , belonging to education group e , observed at time t . The logarithmic transformation allows the estimated coefficients to be interpreted approximately as percentage changes in income while also reducing skewness and stabilizing the variance of the income distribution across individuals and over time. Income values are deflated using the Consumer Price Index (CPI) to account for changes in the general price level over the observation period (Statistics Bureau of Japan, Ministry of Internal Affairs and Communications, 2026). The constant term, β_0 , represents the intercept and captures the baseline level of log income for the reference individual, holding all other covariates at their reference categories or zero values. The term γ , Young_{it} is a binary indicator that equals 1 if individual i belongs to the young cohort (ages 20 to 29) at time t , and 0 otherwise. The coefficient γ therefore captures the baseline difference in log income between young workers and the reference age group. The term $\beta_1 \text{cEES}_{e,t}^{\perp \text{edu}}$ captures the association between the elderly regular employment share and wages for the general regular working-age population. The interaction term $\delta(\text{Young}_{it} \times \text{cEES}_{e,t}^{\perp \text{edu}})$ captures the differential association between the $\text{cEES}_{e,t}^{\perp \text{edu}}$ variable and log income for young workers relative to the reference group. This variable corresponds to Hypothesis 2. The term $\beta_2 t$ denotes a centered linear time trend, constructed as the deviation of each survey year from the sample mean year ($t = \text{Year}_t - \bar{\text{Year}}$). This transfor-

mation preserves the temporal ordering of the panel while improving interpretability and reducing collinearity between the intercept and time-related regressors. Consequently, the coefficient on t captures the average linear change in log income over time relative to the sample mean year rather than relative to an arbitrary base year.

The vector $\bar{\mathbf{X}}_i$ contains individual-specific means of time-varying regressors, included following the Mundlak (1978) approach to control for unobserved individual heterogeneity that may be correlated with time-varying covariates, allowing the model to separate within-individual wage variation from persistent between-individual differences in career trajectories, firm sorting, and labor supply behavior:

$$\bar{\mathbf{X}}_i = (\overline{\text{WorkingHours}}_i, \overline{\text{Position}}_i, \overline{\text{FirmSize}}_i)' \quad (9)$$

where $\overline{\text{WorkingHours}}_i$ captures persistent between-individual differences in labor supply intensity. Pistaferri (2003) uses Italian panel data on subjective wage expectations to decompose wage growth into anticipated, unanticipated, and risk components, and shows that each component affects individual labor supply decisions. This confirms that hours of work are not exogenously determined but instead reflect the same individual-specific preferences and expectations that influence wage outcomes—a joint determination that the Mundlak correlated random effects framework addresses by controlling for time-invariant individual heterogeneity. $\overline{\text{Position}}_i$ absorbs the persistent component of hierarchical rank that determines the pay scale applicable to each worker. In Japan’s internal labor markets, promotion reflects the assumption of greater responsibilities and higher skill demands rather than tenure accumulation alone (Kobayashi, 1981). Failing to control for position would therefore conflate two distinct wage-setting channels: the deferred-compensation schedule that governs within-rank wage progression over tenure, and the step change in pay scale that accompanies movement to a higher hierarchical rank. The Mundlak correction separates these by controlling for the individual’s average position over the panel, isolating within-rank seniority growth from career advancement effects. $\overline{\text{FirmSize}}_i$ captures persistent worker-firm sorting. In Japan’s internal labor market system, workers who enter large firms at graduation remain on a steeper seniority track, while the external market class is concentrated in smaller firms (Koike, 1988; Date et al., 2024). Industry is not included in the Mundlak correction because prior evidence shows that industry wage differentials in Japan are largely explained by the very variables already captured in $\bar{\mathbf{X}}_i$ —tenure, education, and firm size account for most apparent spurious variation across industries (Nishimura, 2016), and the declining role of firm-level seniority in wage determination documented by The Japan Institute for Labour Policy and Training (2022) further confirms that individual-level attributes dominate industry affiliation. Education is excluded from the Mundlak correction because it already enters the model as the dimension along which $\text{cEES}_{e,t}^{\perp \text{edu}}$ varies, meaning its individual mean would conflate the identifying cross-education variation with the between-individual heterogeneity correction. Individual-specific means are calculated as $\bar{X}_i = \frac{1}{T_i} \sum_{t=1}^{T_i} X_{it}$, where T_i denotes the total number of observations available for individual i . For categorical variables, the mean represents the proportion of waves in which individual i is observed in each category. Age is excluded from the correction because it increments deterministically by one unit per wave, making its individual mean collinear with entry timing rather than informative about stable unobserved heterogeneity.

$$\mathbf{Z}_{it} = \left(\begin{array}{c} \text{Gender}_i, \text{Age}_{it}, \text{Education}_{it}, \\ \text{WorkingHours}_{it}, \text{FirmSize}_{it}, \text{Industry}_{it}, \text{Position}_{it} \end{array} \right)' \quad (10)$$

The vector $\mathbf{Z}_{it}$ in Equation 10 contains individual-level control variables including gender, age, education, weekly working hours, firm size, industry and occupational position. These vari-

ables capture standard determinants of wage variation related to demographic characteristics, human capital, job characteristics, and firm attributes. Their inclusion ensures that the estimated coefficients on the macro-level variables reflect demographic labor market pressure rather than individual-level differences. Gender_i is included as a control for the gender wage gap, where the indicator equals 1 for female workers and 0 for male workers. Empirical evidence consistently shows that female workers earn lower wages on average than male workers in Japan, implying a negative association between the female indicator and income (Onozuka, 2016). Age_{it} is included as a proxy for labor market experience. Using Japanese microdata, Kimura et al. (2022) document a positive and concave relationship between job experience and wages: the tenure coefficient is positive and significant (0.034 for 2005–2008), while the squared term is negative, consistent with diminishing returns to additional years of experience. The research further shows that cumulative wage returns for 30 years of tenure were approximately 67.6 percent in 2005–2008, confirming the steep experience-wage profile traditionally observed in Japan. Education_{it} captures differences in human capital across individuals. The same specification in Kimura et al. (2022) includes a college dummy whose coefficient is positive and highly significant across all periods (0.155 in 2005–2008, 0.150 in 2009–2012, and 0.135 in 2013–2017), confirming that higher educational attainment is associated with a substantial wage premium in the Japanese labor market. Current students and individuals belonging to educational categories without a directly comparable classification in the Labour Force Survey Detailed Tabulation are excluded from the estimation sample to maintain consistency between the macro-level demographic indicators and the individual-level panel data. WorkingHours_{it} captures variation in weekly working hours. Increases in working hours are associated with higher total earnings and compensation growth. Empirical evidence from Japan suggests a positive relationship between working hours and income, implying a positive association between weekly working hours and income (Date et al., 2024). FirmSize_{it} captures differences in firm size across individuals. Using Japanese microdata from the Basic Survey on Wage Structure, Date et al. (2024) find that full-time workers at firms with 1,000 or more employees account for over 50 percent of the internal labor market class — characterized by seniority-based wages and long-term employment — compared with less than 11 percent in the external labor market class. Workers classified into the internal labor market class, which is dominated by large firms, earn substantially higher wages with steeper wage-tenure profiles than those in the external labor market class. Accordingly, a positive association between firm size and income is expected. Industry_{it} captures structural differences in wage determination across sectors within the Japanese labor market. Industry controls account for persistent wage variation between sectors that cannot be explained solely by individual worker characteristics. Industry categories were grouped according to the Japan Standard Industrial Classification (JSIC) defined by the Ministry of Internal Affairs and Communications of Japan (Ministry of Internal Affairs and Communications, Japan, 2026). Position_{it} captures hierarchical differences within firms, where higher-ranked positions are associated with higher wages through seniority-based promotion and internal labor market structures. In the Japanese employment system, promotion is closely linked to tenure accumulation and long-term employment relationships, meaning that workers occupying higher organizational positions generally receive substantially higher compensation (Tachibanaki & Ohta, 1994).

Furthermore Equation 1 also includes α_i which represents individual-specific random intercepts that account for persistent differences in earning capacity across individuals and ε_{it} which represents the idiosyncratic error term, capturing shocks to individual income that are not explained by the model.

4.3 Data source

The model specified in Section 4.1 requires two distinct types of data. The individual-level outcome equation requires panel microdata on wages, employment contract type, age, education, working hours, firm size, industry, and occupational position for regular workers across multiple years, so that the individual-level controls in $\mathbf{Z}_{it}$ can absorb compositional differences across education-by-year cells and the random intercepts α_i can capture persistent unobserved earning capacity. The demographic regressor $\text{cEES}_{e,t}^{\perp\text{edu}}$ additionally requires a macro-level tabulation of employed persons classified simultaneously by age group, education group, and employment type, so that the elderly regular employment share can be constructed separately for each educational track and survey year. These two requirements are met by separate data sources described below.

Individual-level data come from the Japanese Panel Study of Employment Dynamics (JPSED) microdata, publicly available through the Social Science Japan Data Archive, Center for Social Research and Data Archives, Institute of Social Science, The University of Tokyo (Recruit Works Institute, 2024). The JPSED covers 2016 through 2024 and provides over 30,000 observations per year after data cleaning. Its detailed individual-level information on wages, employment contract type, occupations, industries, age, education, working hours, and firm characteristics directly supplies all variables entering $\mathbf{Z}_{it}$ and the Mundlak correction vector $\bar{\mathbf{X}}_i$. The survey’s distinction between regular and non-regular employment contracts allows the analysis to be restricted to the regular employment tier where seniority-based wage progression operates, consistent with the model restriction described in Section 4.1. The longitudinal structure provides the within-person wage variation needed to separate demographic wage effects from persistent between-individual differences in earning capacity through the random intercepts and Mundlak correction. The panel also covers the post-EESL amendment period during which elderly regular employment expanded substantially, supplying the temporal dimension over which $\text{cEES}_{e,t}^{\perp\text{edu}}$ varies.

The following sample restrictions are applied to the JPSED microdata prior to estimation. The sample is restricted to regular workers, with current students and individuals belonging to educational categories without a directly comparable classification in the Labour Force Survey Detailed Tabulation excluded to maintain consistency between the macro-level demographic indicators and the individual-level panel data. All income variables are deflated using the Consumer Price Index (Statistics Bureau of Japan, Ministry of Internal Affairs and Communications, 2026). The income variable refers to annual earnings from the respondent’s main job and is restricted to individuals employed in the previous year. Prior to logarithmic transformation, the raw income distribution contains a small number of extremely low values that are difficult to reconcile with regular full-time employment in Japan. Such observations are likely to reflect reporting error, temporary labor-market exits, or employment situations that are not representative of stable regular employment. To reduce the influence of these extreme lower-tail observations, incomes below the first percentile of the real income distribution are excluded before further sample construction. This trimming procedure follows a common approach in wage regression analysis and is intended to remove implausibly low earnings values while preserving the overwhelming majority of observations. After trimming, logarithmic transformation is applied to the inflation-adjusted income measure. The resulting log-income distribution is substantially more symmetric and therefore better suited for linear panel-data estimation. Working hours per week are restricted to the range of 8 to 50 hours. The lower bound excludes observations inconsistent with regular employment intensity, while the upper bound follows from the Labour Standards Act of Japan (Government of Japan, 1947), which establishes a 40-hour standard workweek under Article 32(1) and permits up to 45 overtime hours per month under Article 36, implying a weekly ceiling of roughly 50 hours. Current students are excluded because individuals enrolled in education at the time of the survey may report incidental employment alongside their studies, which is institutionally distinct from the regular employment

relationships the model is designed to capture and would introduce wage observations not comparable to those of workers fully embedded in the regular employment tier. Position is measured using the respondent’s occupational rank within the firm’s hierarchy. The original survey distinguishes eight categories ranging from non-managerial employees to top executives. For ease of interpretation, these categories are recoded into simplified position levels. Employees without formal supervisory responsibilities are classified as *No Managerial Position*. Technical and professional positions receiving employment conditions equivalent to the *kakaricho* or *shunin* level are classified as *Unit Chief/Supervisor-equivalent professional*, while employees formally holding these positions are classified as *Unit Chief (Kakaricho)/Supervisor (Shunin)*. Technical and professional positions receiving employment conditions equivalent to the *kacho* level are classified as *Section Head-equivalent professional*, whereas employees formally holding the *kacho* position are classified as *Section Head (Kacho)*. Similarly, technical and professional positions receiving employment conditions equivalent to the *bucho* level are classified as *Division Head-equivalent professional*, while employees formally holding the *bucho* position are classified as *Division Head (Bucho)*. Finally, representative directors, executive officers, and senior advisors are grouped into the highest category, referred to as *Representative/Executive/Senior Advisor*.

Data for constructing $cEES_{e,t}^{\perp \text{edu}}$ were obtained from the Labour Force Survey Detailed Tabulation, Table 1-5-1, published by the Statistics Bureau of Japan via its e-Stat database (Statistics Bureau of Japan, 2025). This table has been available annually at the national level from 2002 onwards and provides a three-way cross-tabulation of employed persons by age group, educational attainment, and employment type, with regular employees (*seishain*) identified as a distinct category. Age groups are defined over five-year intervals from 15–19 up to and above 75, and the education variable covers graduated levels from junior high school through graduate degrees. To maintain consistency with the regular employment tier, the working-age denominator is restricted to individuals aged 20–59, capturing the core seniority-track workforce against which elderly employment pressure operates. The elderly numerator covers workers aged 60–64, the group directly targeted by the EESL continued-employment obligation. The estimation sample is restricted to workers aged 20–64 so that the sample corresponds to the age groups used to construct the elderly employment share. Workers aged 65 and above are excluded because post-retirement re-employment arrangements place them outside the core seniority hierarchy represented in both the numerator and denominator of the demographic measure (Moriyama, 2025; Kondo, 2016).

4.4 Descriptive Statistics

Table 1 presents summary statistics. The preferred demographic regressor, $cEES_{e,t}^{\perp \text{edu}}$, is standardized to mean zero and unit standard deviation by construction, so that estimated coefficients can be interpreted as the wage response to a one-standard-deviation shift in elderly regular employment share within an educational track. The range of $cEES_{e,t}^{\perp \text{edu}}$ from -1.44 to 2.48 standard deviations reflects genuine within-year cross-education variation in elderly regular employment share across the four education groups and nine survey years, confirming that the orthogonalization against education-group means successfully removes stable between-group level differences while preserving the time trend that constitutes the identifying variation for the hypothesis tests.

Table 1: Descriptive Statistics for the Estimation Sample

Variable	Mean	SD	Min	Max
Log income	6.1384	0.5011	3.4200	8.5221
Young cohort (20–29)	0.0450	0.2070	0	1
$cEES_{e,t}^{\perp edu}$	0.0000	1.0000	−1.4368	2.4810
Age	47.1	9.6	20	64
Working hours	41.5	6.2	8	50

Notes: Author’s own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

Table 2 describes the longitudinal structure of the total estimation sample of 33,713 observations and 5,653 individuals. Individuals are observed in 5.96 waves on average, providing a sufficiently long panel that allows disentangling within-person variation in wages from between-person heterogeneity using the random intercept approach and Mundlak correction.

Table 2: Panel Structure of the Estimation Sample

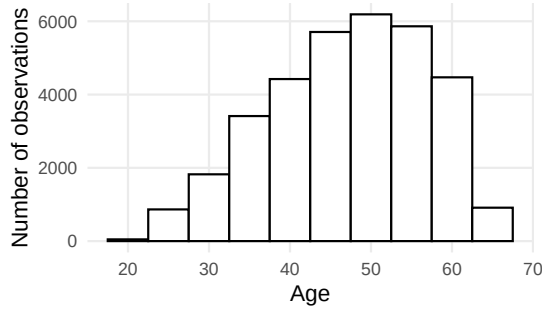
Statistic	Value
Total observations in estimation sample	33,713
Total individuals	5,653
Mean observations per individual	5.96
Individuals ever observed aged 20–29	436
Individuals observed aged 20–29 in 2+ waves	328
Mean within-person SD of log income	0.1595
Median within-person SD of log income	0.0929
Mean within-person SD of log income (Young)	0.2029
Median within-person SD of log income (Young)	0.1536

Notes: Within-person standard deviations are calculated using repeated observations within the panel. “Young” refers to individuals observed at least once in the 20–29 age cohort. Author’s own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

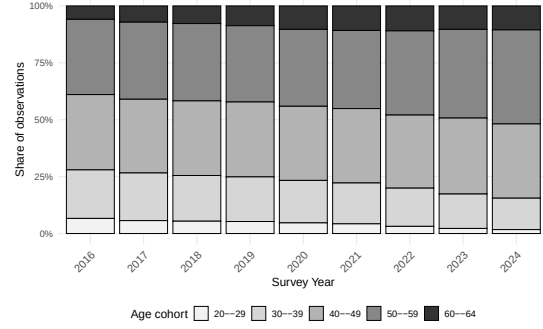
The age structure of the estimation sample aligns with the broader Japanese demographic trends described above. Based on Table 1 and Figure 1a, the mean age in the estimation sample is 47.1 years with the median equal to 48 years and standard deviation equal to 9.6 years. This suggests that the estimation sample is composed mainly of prime-age and elderly regular employees, as expected based on the discussion of ageing regular workforce in Section 2.1. The young age group of workers aged 20–29 represents 4.50 percent of the observations in the estimation sample. According to Figure 1b, the share of young workers declines substantially over time from 6.64 percent of annual observations in 2016 to 1.72 percent in 2024. However, the observed decline in the young worker share should be interpreted with caution because, as indicated in Table 2, it is at least partially a result of the longitudinal structure of the estimation sample: individuals who are observed as young workers at the earlier stages of the estimation sample age out of the age range of 20–29 and enter into older age categories in subsequent waves. Nevertheless, this does not preclude the possibility of identifying the effect of interest since the research problem is whether workers observed as young regular employees face different wages when elderly regular employment increases rather than how many individuals in the sample belong to the category of young workers at each wave of the estimation sample. The interaction effect between the dummy for the young worker and the measure of elderly regular employment is thus identified based on the difference in the wages associated with being observed as a young regular worker in a certain year.

Figure 1c depicts the gender structure of the estimation sample. Men compose 72.5 percent

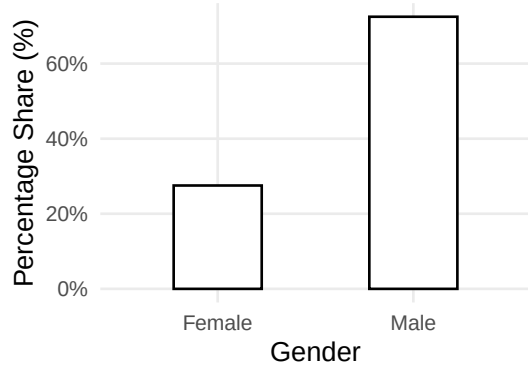
Figure 1: Descriptive Distributions of the Estimation Sample (Part I)



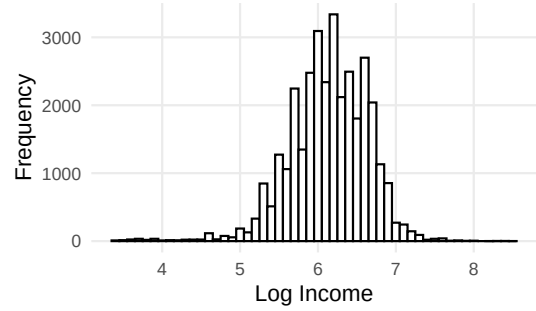
(a) Age Distribution



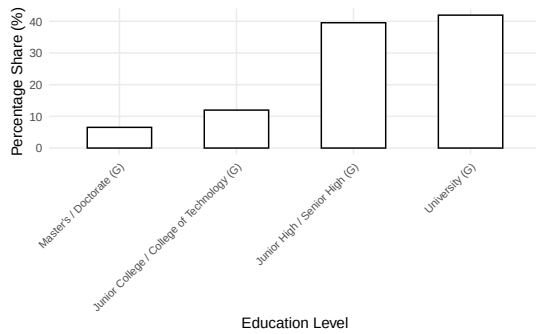
(b) Distribution of Young Cohort Observations Across Survey Years



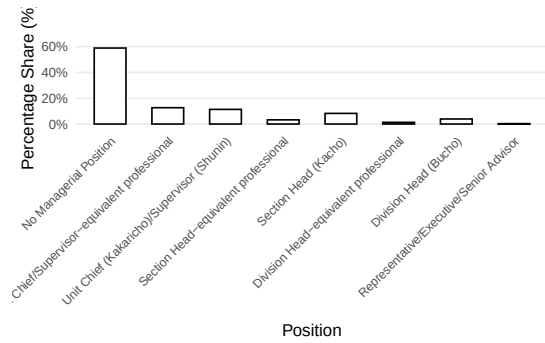
(c) Gender Distribution



(d) Log Income Distribution



(e) Education Distribution



(f) Position Distribution

Notes: Author's own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

(24,433 observations) of the estimation sample while women represent 27.5 percent (9,280 observations). This imbalance reflects the gender composition of Japan's regular employment sector, where women are disproportionately represented in non-regular employment arrangements (OECD, 2018). In all specifications analyzed in this research, control for the unequal gender composition of the regular employment tier is included through the gender dummy variable in Z_{it} .

The dependent variable $\ln(\text{Income}_{iet})$ has a mean value of 6.14 and standard deviation of 0.50 (see Table 1). Figure 1d shows that the distribution of log income is approximately normal, centered close to the sample mean, and does not exhibit extreme values that could substantially distort the estimation results. In addition to cross-sectional variation, the panel structure provides meaningful within-person variation in wages over time. As reported in Table 2, the mean within-person standard deviation of log income is 0.1595, indicating that individual earnings change sufficiently across survey waves to allow identification of the effect of interest. This longitudinal variation is essential for the mixed-effects framework because it ensures that the estimates are not driven solely by time-invariant differences between individuals but also reflect changes in wages observed within the same individual over time.

Figure 1e shows the educational structure of the estimation sample. University graduates constitute the largest educational group, accounting for 42.0 percent (14,141 observations) of the estimation sample, followed by individuals with junior and senior high school diplomas at 39.6 percent (13,333 observations), junior college and college of technology graduates at 12.0 percent (4,037 observations), and workers holding master's or doctorate degrees at 6.53 percent (2,202 observations). It is worth noting that this distribution is relevant to the empirical identification since, as noted in Section 2.1, there are significant differences in the extent of elderly regular employment growth across the educational categories considered in this research.

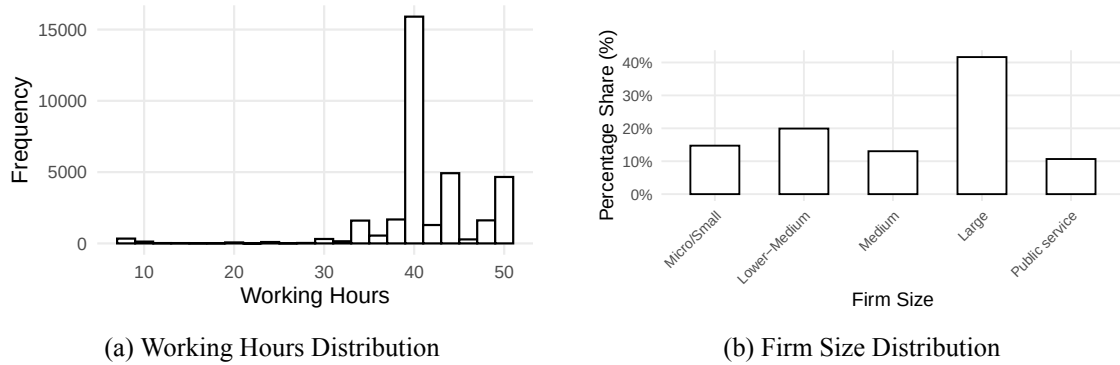
The occupational position distribution, shown in Figure 1f, reflects the seniority hierarchy represented in the estimation sample. Observations corresponding to workers without managerial responsibilities account for 59.0 percent of the sample (19,876 observations), representing entry-level and other regular employees without supervisory roles. The remaining 41.0 percent of observations are distributed across progressively more senior positions, beginning with unit chief/supervisor-equivalent professional positions and *kakaricho/shunin*-level supervisors, and extending through section heads (*kacho*), division heads (*bucho*), executives, and senior advisors. This distribution indicates that the estimation sample spans the full hierarchy of regular employment, from entry-level employees to senior management. Because different hierarchical positions are typically associated with different career stages, the sample contains workers from a broad range of age groups.

The average working hours per week are reported in Table 1 as 41.5 hours with the standard deviation of 6.2 hours, indicating the regular employment of the estimation sample. This can be seen from the distribution of working hours reported in Figure 2a. The primary peak of the distribution corresponds to the standard workweek, and the second corresponds to the standard workweek with a moderate amount of overtime. Working hours are incorporated in the specification as a time-varying variable Z_{it} as well as the individual mean of the same variable in the Mundlak correction vector $\bar{X}_i$ in order to separate the impact of changes in labor intensity on wages and the individual stable difference in labor supply decisions due to productivity variation.

The breakdown of the estimation sample by firm sizes is provided in Figure 2b. Observations from large firms constitute 41.6 percent of the estimation sample, with the second highest contribution being made by lower-medium firms with 19.9 percent share. Micro- and small firms account for 14.7 percent and medium firms for 13.0 percent of the estimation sample. Government public service accounts for another 10.7 percent of the sample. The preponderance of large firms in the estimation sample results from the nature of the Japanese regular employment institutions: large

firms and governmental employers develop seniority-based wage systems and continued employment more actively than other firms. It is critical to include controls for firm sizes, since wages and elderly workers' employment retention policies vary significantly across different types of firms. Hence, including firm-size fixed effects may be essential in investigating the relation between expansion in elderly workers' employment and wage changes.

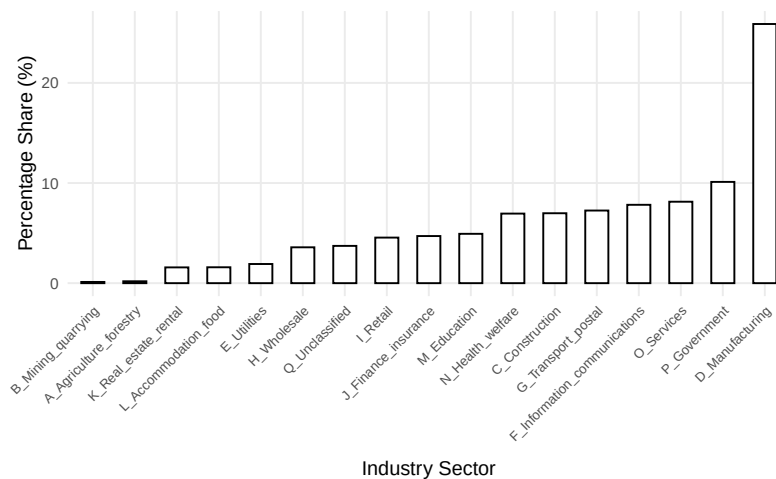
Figure 2: Descriptive Distributions of the Estimation Sample (Part II)



Notes: Author's own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

The predominance of manufacturing observations is significant. As shown in Figure 3, manufacturing accounts for 25.9 percent of all observations in the estimation sample (8,725 observations), making it the largest industry represented in the data. According to Japan's 2022 *Monozukuri Hakusho* (White Paper on Manufacturing), published jointly by the Ministry of Economy, Trade and Industry (METI), the Ministry of Health, Labour and Welfare (MHLW), and the Ministry of Education, Culture, Sports, Science and Technology (MEXT), the share of regular employees in manufacturing is 15.1 percentage points higher than the all-industry average, while the number of elderly workers (aged 65+) in the sector grew by approximately 330,000 over the 20 years to 2021 (Ministry of Economy, Trade and Industry et al., 2022). Manufacturing therefore combines a high concentration of regular employment with a pronounced long-term increase in its elderly workforce.

Figure 3: Descriptive Distributions of the Estimation Sample (Part III)



Notes: Author's own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

5 Results

5.1 Specification diagnostics

Table 3 reports generalized variance inflation factors (GVIFs) for the preferred specification using the adjusted measure $\text{GVIF}^{1/(2 \times \text{Df})}$, where values substantially below 5 indicate the absence of problematic multicollinearity. Most regressors show values close to 1, including the early-career worker dummy (1.154), gender (1.134), age (1.349), working hours (1.047), mean working hours (1.080), education (1.034), position controls (1.100), industry controls (1.034), and the $\text{Young} \times \text{cEES}_{e,t}^{\perp \text{edu}}$ interaction (1.123), indicating negligible collinearity. The Mundlak firm-size means are moderately higher, with a maximum adjusted GVIF of 1.980, reflecting the expected correlation between group means and time-varying firm-size controls under the decomposition proposed by Mundlak (1978) rather than a specification problem. The largest values appear for $\text{cEES}_{e,t}^{\perp \text{edu}}$ (2.181) and the linear time trend (2.302), consistent with their high partial correlation, which arises because both variables capture the long-run upward demographic trend in elderly employment over the panel period.

However, this does not invalidate inference. Wooldridge (2019) emphasizes that imperfect multicollinearity does not bias estimates and primarily inflates standard errors, implying that statistically significant coefficients surviving collinearity satisfy a stricter inferential standard. Moreover, even VIF values above 10 do not automatically invalidate inference when sample size and regressor variation are sufficiently large. In the present case, the estimation sample contains 33,713 observations. Following Mundlak (1978, eqs. 2.3–2.7, 3.4–3.7), the observed collinearity operates mainly in the between-education-group dimension, because the common time trend is absorbed by the between-projection operator K , whereas $\text{cEES}_{e,t}^{\perp \text{edu}}$ retains within-group identifying variation through $M = I - K$. As shown by Mundlak (1978), the within-estimator variance $V_w = \sigma^2(X'MX)^{-1}$ depends only on within-group variation, while between-dimension collinearity inflates only $V_b = (\sigma^2 + T\omega_1^2)(X'KX)^{-1}$. Similarly, Kropko & Kubinec (2020, pp. 7–8, eqs. 6 and 8) show that time fixed-effects estimators exploit cross-sectional variation within periods rather than common aggregate trends.

Empirical diagnostics further confirm that collinearity does not compromise estimation. The likelihood-ratio test reported in Table 4 shows that excluding the interaction term substantially worsens model fit ($\chi^2 = 14.51$, $p < 0.001$). The confidence intervals reported in Table 5 remain narrow and exclude zero for both $\text{cEES}_{e,t}^{\perp \text{edu}}$ ($[-0.0207, -0.0084]$) and the interaction term ($[0.0782, 0.1141]$). Furthermore, the year jackknife estimates reported in Table 6 remain stable across all survey waves, ranging from -0.0080 to -0.0164 for $\text{cEES}_{e,t}^{\perp \text{edu}}$, while the interaction coefficient remains positive in every specification, ranging from 0.0896 to 0.1025 . Taken together, these results indicate that the observed collinearity modestly inflates standard errors but does not threaten the reliability of the main estimates, particularly for the youth-specific interaction central to Hypothesis 2.

Table 3: Generalized Variance Inflation Factors (GVIF) for the Preferred Specification

Variable	GVIF	Df	GVIF ^{1/(2×Df)}
Young worker (20–29)	1.3308	1	1.1536
cEES _{e,t} ^{⊥edu}	4.7586	1	2.1814
Year trend (<i>t</i>)	5.2992	1	2.3020
Mean working hours (Mundlak)	1.1673	1	1.0804
Education	1.2211	3	1.0339
Gender	1.2860	1	1.1340
Age	1.8189	1	1.3487
Working hours	1.0952	1	1.0465
Firm size	9.5896	4	1.3266
Industry	2.9509	16	1.0344
Position	3.8183	7	1.1004
<i>Mundlak terms</i>			
Mean size: Lower-medium	2.5652	1	1.6016
Mean size: Medium	2.2038	1	1.4845
Mean size: Large	3.9192	1	1.9797
Mean size: Public service	3.5213	1	1.8765
Mean position: Unit Chief/Supervisor-equivalent professional	1.2287	1	1.1085
Mean position: Unit Chief (Kakaricho)/Supervisor (Shunin)	1.2422	1	1.1146
Mean position: Section Head-equivalent professional	1.1627	1	1.0783
Mean position: Section Head (Kacho)	1.3803	1	1.1749
Mean position: Division Head-equivalent professional	1.1624	1	1.0782
Mean position: Division Head (Bucho)	1.5080	1	1.2280
Mean position: Representative Executive/Senior Advisor	1.3496	1	1.1617
Young × cEES _{e,t} ^{⊥edu}	1.2607	1	1.1228

Notes: Since raw GVIF values are not directly comparable across variables with different degrees of freedom, the adjusted measure GVIF^{1/(2×Df)} is reported. Values substantially below the conventional threshold of 5 indicate that multicollinearity is not a concern. The highest adjusted GVIF is 2.3020 for the linear year trend, well below commonly used cutoffs. Author’s own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

Table 4: Likelihood Ratio Tests for cEES_{e,t}^{⊥edu} Specifications

Model comparison	χ^2	df	<i>p</i> -value
Without cEES _{e,t} ^{⊥edu} vs. baseline model	124.1900	2	< 0.001
Without interaction vs. baseline model	14.5100	1	< 0.001
Without cEES _{e,t} ^{⊥edu} vs. without interaction	109.6800	1	< 0.001

Notes: All models estimated by maximum likelihood. Author’s own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

Table 5: Confidence Intervals for Key Coefficients

Variable	Estimate	2.5%	97.5%
Young worker (20–29)	−0.0810	−0.1057	−0.0566
$\text{cEES}_{e,t}^{\perp\text{edu}}$	−0.0150	−0.0207	−0.0084
Young $\times$ $\text{cEES}_{e,t}^{\perp\text{edu}}$	+0.0960	+0.0782	+0.1141

Notes: Wald confidence intervals. Author’s own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

Table 6: Jackknife Stability by Survey Year

Dropped year	$\text{cEES}_{e,t}^{\perp\text{edu}}$			Young $\times$ $\text{cEES}_{e,t}^{\perp\text{edu}}$		
	Estimate	SE	<i>p</i> -value	Estimate	SE	<i>p</i> -value
2016	−0.0131	0.0036	< 0.001	0.0921	0.0097	< 0.001
2017	−0.0151	0.0031	< 0.001	0.0900	0.0097	< 0.001
2018	−0.0164	0.0034	< 0.001	0.0982	0.0098	< 0.001
2019	−0.0158	0.0035	< 0.001	0.0962	0.0097	< 0.001
2020	−0.0158	0.0032	< 0.001	0.0896	0.0098	< 0.001
2021	−0.0148	0.0035	< 0.001	0.0997	0.0101	< 0.001
2022	−0.0157	0.0034	< 0.001	0.1025	0.0101	< 0.001
2023	−0.0134	0.0033	< 0.001	0.1011	0.0097	< 0.001
2024	−0.0080	0.0034	0.019	0.0896	0.0097	< 0.001

Notes: Author’s own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

5.2 Hypothesis tests

Results from the main theoretical findings are presented in Table 7.

The coefficient on the young-worker indicator is negative and statistically significant ($\hat{\gamma} = -0.0812$, $t = -6.485$, $p < 0.001$). When $\text{cEES}_{e,t}^{\perp\text{edu}} = 0$, workers aged 20–29 are associated with approximately 8.1 percent lower wages than otherwise comparable workers in the reference age group. This baseline wage deficit is consistent with the position of early-career workers on the deferred-compensation schedule of Japanese internal labor markets.

Hypothesis 1 predicts that a higher elderly regular employment share is associated with lower wages for the general regular workforce. The estimated coefficient on $\text{cEES}_{e,t}^{\perp\text{edu}}$ is negative and statistically significant ($\hat{\beta}_1 = -0.0146$, $t = -4.638$, $p < 0.001$). A one-standard-deviation increase in the education-group-specific elderly regular employment share is associated with approximately 1.5 percent lower wages among regular workers in general. Hypothesis 1 is supported by the evidence.

The modest absolute magnitude of $\hat{\beta}_1$ is consistent with the aggregation level at which $\text{cEES}_{e,t}^{\perp\text{edu}}$ is constructed. Because the compression mechanism operates partly through firm-internal wage-setting structures, assigning a uniform education-group-level elderly employment share to all individuals within a group introduces measurement error that attenuates the estimate toward zero. The coefficient should therefore be interpreted as a lower bound on the underlying compression association. The coefficient remains negative and statistically significant in all nine year-jackknife specifications, ranging from -0.0080 to -0.0164 (Table 6), indicating that the estimated wage compression effect is not driven by any single survey wave.

Hypothesis 2 predicts that the wage association with elderly regular employment differs for early-career workers. The interaction between the young-worker dummy and $\text{cEES}_{e,t}^{\perp\text{edu}}$ is positive and statistically significant ($\hat{\delta} = 0.0962$, $t = 10.484$, $p < 0.001$). The wage association linked to

Table 7: Comprehensive Mixed-Effects Estimation Results

Variable	Estimate	Std. Error	t-value	p-value	Sign.
<i>Main theoretical variables</i>					
Young worker (20–29)	−0.0812	0.0125	−6.4850	< 0.001	***
$cEES_{e,t}^{\perp edu}$	−0.0146	0.0031	−4.6380	< 0.001	***
Young $\times$ $cEES_{e,t}^{\perp edu}$	0.0962	0.0092	10.4840	< 0.001	***
<i>Continuous demographic and labor variables</i>					
Time trend	0.0025	0.0013	1.8920	0.0585	*
Age	0.0033	0.0005	6.6710	< 0.001	***
Female	−0.2058	0.0113	−18.1880	< 0.001	***
Working hours	0.0028	0.0003	9.6080	< 0.001	***
<i>Education controls</i>					
Junior/Senior High	−0.0429	0.0131	−3.2660	0.0011	***
University	0.0378	0.0135	2.7900	0.0053	***
Master's/Doctorate	0.1410	0.0205	6.8680	< 0.001	***
<i>Firm size controls</i>					
Lower-Medium	0.0279	0.0104	2.6920	0.0071	***
Medium	0.0629	0.0124	5.0850	< 0.001	***
Large	0.0907	0.0122	7.4590	< 0.001	***
Public service	0.1327	0.0206	6.4430	< 0.001	***
<i>Industry controls</i>					
Mining/Quarrying	0.0208	0.0963	0.2160	0.8292	
Construction	0.1158	0.0577	2.0050	0.0450	**
Manufacturing	0.1470	0.0569	2.5840	0.0098	***
Utilities	0.1672	0.0612	2.7340	0.0063	***
Information/Communications	0.1369	0.0577	2.3730	0.0177	**
Transport/Postal	0.0971	0.0574	1.6930	0.0904	*
Wholesale	0.1080	0.0577	1.8720	0.0613	*
Retail	0.0908	0.0573	1.5840	0.1131	
Finance/Insurance	0.1024	0.0574	1.7840	0.0745	*
Real Estate/Rental	0.1169	0.0591	1.9790	0.0478	**
Accommodation/Food	0.0884	0.0586	1.5090	0.1312	
Education	0.1578	0.0575	2.7440	0.0061	***
Health/Welfare	0.0962	0.0574	1.6750	0.0939	*
Services	0.0815	0.0569	1.4320	0.1521	
Government	0.0974	0.0587	1.6580	0.0974	*
Unclassified	0.0937	0.0572	1.6400	0.1011	
<i>Position controls</i>					
Unit Chief/Supervisor-equivalent professional	0.0833	0.0068	12.2140	< 0.001	***
Unit Chief (Kakaricho)/Supervisor (Shunin)	0.1038	0.0077	13.4380	< 0.001	***
Section Head-equivalent professional	0.1614	0.0120	13.4980	< 0.001	***
Section Head (Kacho)	0.2138	0.0103	20.7790	< 0.001	***
Division Head-equivalent professional	0.2519	0.0182	13.8370	< 0.001	***
Division Head (Bucho)	0.2859	0.0161	17.7780	< 0.001	***
Representative/Executive/Senior Advisor	0.2409	0.0396	6.0840	< 0.001	***

Notes: t-statistics are calculated using Satterthwaite approximated degrees of freedom. Omitted reference categories are: male, micro/small firms, agriculture/forestry/fishery industry, no managerial position, and junior college/college of technology education group. Notations: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. Author's own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

elderly regular employment is approximately 9.6 percentage points more positive for workers aged 20–29 than for the general regular workforce. The total net association for workers aged 20–29 is $\hat{\beta}_1 + \hat{\delta} = -0.0146 + 0.0962 = +0.0816$: a one-standard-deviation increase in elderly regular employment share is associated with approximately 8.2 percent higher wages among workers aged 20–29, whereas the corresponding association for the general workforce is negative. Hypothesis 2 is supported by the evidence.

The two associations operate through institutionally distinct channels. The general compression association reflects wage-structure adjustment distributed across the full seniority hierarchy, as the growing share of workers in the overpaid phase of the deferred-compensation schedule raises aggregate labor costs that firms absorb through structural wage-system reform rather than through direct nominal wage cuts. The youth association operates through the external recruitment margin: as younger cohorts shrink, firms compete over a smaller pool of potential regular entrants, and the relative value of young workers in external recruitment rises even while compression pressures bear down on the broader regular workforce.

The asymmetry in magnitudes — $|\hat{\delta}|$ is roughly six times larger than $|\hat{\beta}_1|$ — reflects the different margins on which the two effects operate rather than any inconsistency in the underlying mechanisms. The wage-compression mechanism underlying H1 operates primarily within firms because the cost pressure generated by elderly regular employment depends on each firm’s internal age structure and wage system. By contrast, the empirical measure varies only at the education-group-by-year level and assigns the same elderly employment share to all workers within the same educational category. This aggregation introduces measurement error and likely attenuates the estimated compression effect. The scarcity mechanism underlying H2 differs in this respect. Competition for young entrants occurs across the external labor market within each education group rather than within individual firms. Because cohort scarcity itself is determined at the education-group level, the aggregate measure corresponds more closely to the level at which the mechanism operates and is therefore less affected by this source of attenuation.

Taken together, the three main coefficients describe a labor market in which a larger elderly regular-employment share is associated with wage compression across the established regular workforce and a more favourable wage differential for young workers entering that same system. The negative $\hat{\beta}_1$ and the positive $\hat{\delta}$ operate in opposite directions on the same demographic variation, consistent with population ageing generating asymmetric wage consequences across the career hierarchy through institutionally distinct channels.

Table 7 also reports estimates for the control variables. The coefficient on age is positive (0.0033, $p < 0.001$), reflecting gradual human capital accumulation over the career. Female workers are associated with substantially lower wages than male workers (-0.2058 , $p < 0.001$), consistent with evidence of a persistent gender wage gap among Japanese regular employees (Onozuka, 2016). Working hours are positively associated with wages (0.0028, $p < 0.001$). The education coefficients follow the expected human-capital ordering. Relative to workers with junior college or college of technology education, workers with junior or senior high school education are associated with lower wages (-0.0429 , $p < 0.01$), university graduates with higher wages (0.0378, $p < 0.01$), and workers holding master’s or doctorate degrees with the largest educational premium (0.1410, $p < 0.001$), consistent with the continued importance of educational attainment in determining wages in Japan (Araki, 2025). Firm-size coefficients are positive and increase monotonically with employer size. Relative to workers in micro and small firms, workers in lower-medium firms (0.0279, $p < 0.01$), medium firms (0.0629, $p < 0.001$), large firms (0.0907, $p < 0.001$), and public-service organisations (0.1327, $p < 0.001$) are associated with progressively higher wages, consistent with the well-established firm-size wage premium in Japan (Tachibanaki & Ohta, 1994). Position controls display substantial wage differentials across the

seniority hierarchy relative to workers with no managerial position. Estimated premiums increase with hierarchical rank, ranging from approximately 8.3 percent for unit chief/supervisor-equivalent professional positions (0.0833, $p < 0.001$) and 10.4 percent for formal *kakaricho/shunin*-level supervisors (0.1038, $p < 0.001$), to 21.4 percent for section heads (*kacho*) (0.2138, $p < 0.001$) and 28.6 percent for division heads (*bucho*) (0.2859, $p < 0.001$). Equivalent professional positions at the section-head and division-head levels also exhibit sizeable positive wage differentials. Industry coefficients are generally positive relative to the omitted agriculture, forestry, and fisheries category, with the largest premiums observed in utilities (0.1672, $p < 0.01$), education (0.1578, $p < 0.01$), and manufacturing (0.1470, $p < 0.01$).

Table 8 reports the individual-specific means of time-varying firm-size and position indicators included following Mundlak (1978). The Mundlak means are uniformly larger than their corresponding time-varying coefficients, with firm-size means ranging from 0.0900 to 0.2955 (versus 0.0279 to 0.1327) and position means ranging from 0.1729 to 0.4818 (versus 0.0833 to 0.2859). The individual-specific mean of working hours equals 0.0354 compared to the contemporaneous coefficient of 0.0028. This pattern confirms that the Mundlak correction successfully separates persistent between-individual heterogeneity in firm sorting, career trajectories, and labor-supply patterns from the within-individual wage variation that identifies $\hat{\beta}_1$ and $\hat{\delta}$.

Table 8: Mundlak Correction Terms

Variable	Estimate	Std.Error	t-value	p-value	Sign.
<i>Working hours Mundlak mean</i>					
Mean working hours	0.0354	0.0043	8.3080	< 0.001	***
<i>Firm size Mundlak means</i>					
Mean: Lower-Medium	0.0900	0.0199	4.5180	< 0.001	***
Mean: Medium	0.1089	0.0226	4.8130	< 0.001	***
Mean: Large	0.2869	0.0193	14.8860	< 0.001	***
Mean: Public service	0.2955	0.0273	10.8180	< 0.001	***
<i>Position Mundlak means</i>					
Mean: Unit Chief/Supervisor-equivalent professional	0.2029	0.0213	9.5390	< 0.001	***
Mean: Unit Chief (Kakaricho)/Supervisor (Shunin)	0.1729	0.0215	8.0520	< 0.001	***
Mean: Section Head-equivalent professional	0.2668	0.0393	6.7860	< 0.001	***
Mean: Section Head (Kacho)	0.2112	0.0236	8.9410	< 0.001	***
Mean: Division Head-equivalent professional	0.3185	0.0570	5.5910	< 0.001	***
Mean: Division Head (Bucho)	0.2948	0.0308	9.5830	< 0.001	***
Mean: Representative/Executive/Senior Advisor	0.4818	0.0804	5.9890	< 0.001	***

Notes: Omitted reference categories are: micro/small firms and no managerial position. Author's own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

5.3 Robustness Check

The robustness checks address two distinct concerns about the main results. First, whether sparse education-by-year cells containing young workers generate artificial significance, and whether the youth-specific interaction estimate is driven by the concentration of young workers in earlier survey waves due to ageing-out.

Table 9 reports cell-size robustness checks for both key coefficients. The preferred specification is re-estimated after dropping education-by-year cells with fewer than 10 and 20 young-worker observations respectively, while retaining all non-young observations. The interaction coefficient

$\hat{\delta}$ remains stable across all three specifications, ranging from 0.096 to 0.111 and retaining statistical significance in every case, confirming that the youth-specific wage differential is not an artefact of small-cell noise in later waves.

The general compression coefficient $\hat{\beta}_1$ is more sensitive to the restriction, declining in magnitude from -0.0146 in the full sample to -0.0068 and -0.0076 in the restricted samples. This attenuation is consistent with a reduction in identifying variation rather than evidence against the wage-compression mechanism. The identifying variation for $\hat{\beta}_1$ comes from the divergence in elderly employment growth across education groups, which accumulated progressively over the panel period and is therefore most pronounced in later waves. Later waves also contain the thinnest Young cells due to ageing-out. Dropping thin cells therefore removes a disproportionate share of the cross-education variation that identifies $\hat{\beta}_1$, mechanically attenuating the estimate.

Table 9: Cell-Size Robustness Check

Specification	$\hat{\beta}_1$ (cEES)	$\hat{\delta}$ (Young $\times$ cEES)
Full sample	-0.0146^{***}	0.0962^{***}
Cells $n \geq 10$	-0.0068^*	0.0955^{***}
Cells $n \geq 20$	-0.0076^{**}	0.1105^{***}

Notes: Preferred specification re-estimated dropping education-by-year cells with fewer than 10 and 20 young-worker observations respectively. Non-young observations are retained in all samples. $^*p < 0.10$, $^{**}p < 0.05$, $^{***}p < 0.01$. Author's own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

The second concern is whether the interaction estimate is disproportionately influenced by the earlier survey waves, when the number of observations classified as young workers is larger, rather than reflecting a stable relationship across the full panel period. As shown in the descriptive statistics, the share of observations belonging to the 20–29 age group declines over time, implying that young-worker observations are more concentrated in the earlier survey waves. This raises the possibility that the estimated interaction effect is driven primarily by the early part of the sample.

To examine this possibility, the model is re-estimated separately for the early (2016–2019) and late (2020–2024) sub-periods, with results reported in Table 10. The interaction coefficient $\hat{\delta}$ is highly stable across the two sub-periods, taking values of 0.072 and 0.073 respectively, and remains statistically significant in both cases. This indicates that the youth-specific wage differential is not driven by the concentration of young-worker observations in the earlier survey waves.

Table 10: Interaction Effect Across Early and Late Survey Waves

Period	Estimate	Std. Error	t -value	p -value
2016–2019	0.0717	0.0232	3.092	0.002
2020–2024	0.0734	0.0194	3.788	< 0.001

Notes: Estimates correspond to the interaction coefficient $\text{Young}_{it} \times cEES_{e,t}^{\perp edu}$ from the preferred specification re-estimated separately for early and late survey waves. Standard errors clustered at the education-by-year level. Author's own compilation from Recruit Works Institute (2024) and Statistics Bureau of Japan (2025).

Taken together, the cell-size restrictions confirm that sparse cells do not generate the significance of $\hat{\delta}$, while the sub-period split confirms that the interaction is identified consistently across the full panel period and is not driven by the data-rich early waves. The theoretical and empirical

evidence jointly support the validity of the unweighted estimator under the ageing-out structure of the panel.

6 Discussion

Both hypotheses are supported by the evidence. Hypothesis 1, predicting that an increase in education-group-specific elderly regular employment share is associated with lower wages for the general regular workforce, is confirmed by the negative and statistically significant $\hat{\beta}_1 = -0.0146$ ($p < 0.001$). Hypothesis 2, predicting a positive youth-specific wage differential, is confirmed by the positive and highly significant interaction term ($\hat{\delta} = +0.0962$, $p < 0.001$), yielding a net wage association of approximately +8.2 percent per standard deviation for workers aged 20–29. The estimates reflect conditional associations rather than causal effects, since no external instrument is available to rule out education-group-level time-varying confounders.

The general compression finding is consistent with the theoretical expectation developed in Section 3. As the share of regular workers in the overpaid phase of the deferred-compensation schedule grows, firms face rising aggregate labor costs that cannot be resolved through direct nominal wage cuts to incumbent workers without undermining the implicit contract, and instead adjust through structural wage-system reform that compresses wages across the full seniority hierarchy. This direction aligns with Hamaaki et al. (2012), who document a pronounced and progressive flattening of the age–wage profile for male lifetime employees between 1989 and 2008, interpreting it as long-run institutional adjustment to population ageing rather than cyclical noise. Inoue (2022) strengthens this interpretation by showing that after 2000 workforce ageing explains more than 80 percent of the observed flattening of the age–wage profile. The institutional channel through which this compression operates is further documented by Nishimura (2016) and The Japan Institute for Labour Policy and Training (2022): the share of firms adopting job-based pay for management-level employees rose from 21.1 percent in 1999 to 78.5 percent in 2018, while the prevalence of seniority-based wage systems among non-management employees fell below 50 percent after 2016. The present finding extends this aggregate and firm-level evidence by showing that the compression association is detectable at the level of individual wages within education groups, conditional on individual characteristics and persistent unobserved heterogeneity.

The absence of a compression effect in Kondo (2016) does not contradict this finding. Kondo’s design exploits short-run variation around the 2006 EESL amendment and identifies effects within five to six years of the reform. The compression mechanism documented by Hamaaki et al. (2012) and Inoue (2022) accumulates gradually over decades, becoming statistically detectable only over longer horizons than discrete policy evaluations can capture. The present panel covers 2016–2024, allowing more time for the long-run demographic pressures identified by Inoue (2022) to affect individual wages than in the shorter post-2006 period examined by Kondo (2016).

The youth-specific differential is also consistent with the theoretical expectation and with prior evidence on cohort-size effects in Japan. The positive $\hat{\delta}$ reflects the external recruitment margin identified in Section 3: Kobayashi (1981) establishes that internal labor markets in Japan hire almost exclusively at the lowest job-ladder rung, making entry the primary contact point between workers and the external labor market, and Date et al. (2024) confirm empirically that wages at this margin respond to external supply and demand while wages of workers already embedded in the internal hierarchy are insulated. As youth cohorts contract, firms therefore compete more intensively over a smaller pool of potential regular entrants, bidding up starting wages. Okamura (2001) documents this cohort-size mechanism directly for Japan, finding that larger entering cohorts depress entry wages with effects concentrated in the early-career phase and attenuating with experience. Brunello (2010) provides the cross-national institutional context: cohort-size effects

on entry wages are significantly larger under stricter employment protection, because protection reduces substitutability across age groups and concentrates supply-shift adjustment in wages. Japan’s internal labor market institutions, which produce strong employment protection norms and slow job-to-job reallocation for regular workers, are precisely the environment in which this amplification operates. The duration of the effect is consistent with Kawaguchi & Murao (2014), who show that entry-condition effects are strongest during the first decade after labor market entry and are amplified in countries with stricter employment protection legislation. Oreopoulos et al. (2012) similarly show that entry-condition effects on wages persist for approximately ten years and are driven by conditions at the very first year of entry.

The asymmetry in magnitudes — $|\hat{\delta}|$ is roughly six times larger than $|\hat{\beta}_1|$ — reflects the different aggregation properties of the two effects rather than any inconsistency in the underlying mechanisms. The wage-compression mechanism operates partly within firms because the cost pressure generated by elderly regular employment depends on each firm’s internal age composition and seniority structure. However, the empirical measure of elderly employment share is constructed at the education-group-by-year level and assigns the same value to all workers within a given education group. As a result, the measure cannot capture differences in elderly employment burdens across firms. This aggregation introduces aggregation error and likely attenuates the estimated compression effect, biasing $\hat{\beta}_1$ toward zero.

The scarcity mechanism differs in this respect. Competition for young regular entrants takes place across the external labor market rather than within individual firms. Firms compete for the limited supply of young workers available within each education group, so cohort scarcity itself is determined at the education-group level. Because the level of measurement corresponds more closely to the level at which the mechanism operates, the interaction effect is subject to less attenuation from aggregation. The asymmetry in coefficient magnitudes is therefore consistent with the different levels at which the two mechanisms operate.

7 Conclusion

7.1 Research Summary

This research investigated how Japan’s ageing population, operating through the education-group-specific accumulation of elderly regular workers aged 60–64, affects wage trajectories within the regular employment tier. Using JPSED panel data (2016–2024) combined with education-group-level demographic measures from the Labour Force Survey, and estimating mixed-effects models with a Mundlak correction following Mundlak (1978), the analysis addressed two research questions.

The general internal labor market hypothesis (H1) is supported by the evidence: a one-standard-deviation increase in education-group-specific elderly regular employment share is associated with approximately a 1.5 percent lower wage among the general regular workforce ($\hat{\beta}_1 = -0.0146$, $p < 0.001$). This finding is consistent with the systemic wage-compression mechanism predicted by the deferred-compensation framework and with the broader institutional adjustments associated with successive EESL amendments.

The youth-specific differential hypothesis (H2) is also supported by the evidence. Workers aged 20–29 exhibit a positive wage differential of +9.62 percentage points per standard deviation relative to the general workforce ($\hat{\delta} = +0.0962$, $p < 0.001$), yielding a net coefficient of +0.082 ($\hat{\beta}_1 + \hat{\delta} = -0.0146 + 0.0962 = +0.0816$). Thus, a one-standard-deviation increase in education-group-specific elderly regular employment share is associated with approximately 8.2 percent higher wages among workers aged 20–29. This result remains robust across alternative specifications, with the interaction coefficient ranging from 0.0896 to 0.1025 in the year jackknife

analysis and 0.0955 to 0.1105 in the cell-size robustness checks, retaining statistical significance throughout.

The two forces operate through institutionally distinct channels and on different time horizons. The general compression is a continuous, cumulative structural trend distributed across the full seniority hierarchy and operating over the full working life (Hamaaki et al., 2012). The youth premium is a discrete entry-level effect tied to labor market entry, concentrated at a single recruitment moment for a shrinking cohort, which erodes as workers accumulate tenure (Date et al., 2024; Kobayashi, 1981). Workers entering during high- $cEES_{e,t}^{\perp \text{edu}}$ years receive a starting-wage advantage, but the wage structure they face at peak career is simultaneously being flattened.

This research contributes to the literature in three respects. First, it provides direct evidence that early-career regular workers experience a wage association with education-group-specific elderly employment expansion that differs systematically from that of the general workforce. Prior research has examined either aggregate age-wage profile compression or short-run employment displacement effects, but the joint estimation of a general compression association and a youth-specific differential within the same empirical framework has been lacking. Second, the empirical framework itself constitutes a methodological contribution. The strategy of combining individual-level panel microdata with education-group-level demographic measures, estimated through mixed-effects models with a Mundlak correction, provides a replicable template for studying demographic wage effects in internal labor markets. This framework is not specific to Japan or to the 2016–2024 period: it can be applied directly to other national contexts where administrative or survey panel data can be matched to group-level demographic tabulations, making it straightforwardly transferable to comparative ageing research across OECD economies facing similar demographic pressures. Third, the finding that the entry premium is a time-limited effect while compression is a structural trend affecting every year of work after entry contributes to understanding how population ageing reshapes wage trajectories asymmetrically across the career, with direct implications for the design of wage-system and social security reform beyond Japan.

7.2 Policy Implications

The empirical results suggest that the costs of population ageing are distributed unevenly across the regular workforce. A one-standard-deviation increase in the education-group-specific elderly regular employment share is associated with approximately 1.5 percent lower wages for the general regular workforce ($\hat{\beta}_1 = -0.0146$, $p < 0.001$), while workers aged 20–29 exhibit a positive net wage association of approximately 8.2 percent, consistent with cohort scarcity operating at the point of labor market entry. Because this youth differential is the only offsetting association identified in the model, and because workers aged 30 and above account for over 95 percent of the estimation sample, the negative general coefficient is identified almost entirely from mid-career regular employees. The adjustment burden of population ageing therefore falls on workers who have aged past the entry margin where scarcity operates, yet remain decades away from the settlement phase of the deferred-compensation contract — fully exposed to the wage-compression process operating across the seniority hierarchy with no countervailing premium. This inference is independently corroborated by Hamaaki et al. (2012), who document that the post-2000 flattening of the age–wage profile produced a pronounced kink around age 40, precisely the career stage at which the present results locate the compression burden.

These distributional findings carry important policy implications. Workers in the middle stages of their careers often combine labor-market responsibilities with household formation, child-rearing, and investments in their children’s education. Gordon (2018) argues that wage stagnation and deteriorating employment prospects can delay family formation and child-rearing decisions, with consequences extending beyond individual welfare. A persistent reduction in wage growth for

mid-career workers may therefore affect household consumption, fertility decisions, and intergenerational living standards.

One possible response is to strengthen support for households in the child-rearing stage through measures such as childcare provision, family-related benefits, and policies that facilitate the reconciliation of work and family responsibilities. The OECD (2018) emphasizes the importance of family-friendly policies and improved work–life balance in promoting labor-force participation and supporting longer working lives. Policies that reduce the burden faced by working families may therefore help mitigate some of the wider consequences of wage compression among mid-career households.

A second response is to promote wage systems that rely less heavily on uniform compression across the seniority hierarchy and instead link compensation more closely to productivity and responsibilities. The OECD (2018) argues that steep seniority-based wage systems are increasingly difficult to reconcile with extended working lives and recommends greater flexibility in wage determination. Hwang (2013), examining Korea’s retirement-age extension, similarly argues that extending employment without corresponding wage-system reform risks increasing labor costs and shifting adjustment pressures elsewhere within firms. Because both Korea and Japan rely heavily on deferred-compensation arrangements, the Korean experience suggests that retirement-age extension policies are more sustainable when accompanied by reforms to wage-setting institutions. Such measures may help ensure that the costs of demographic adjustment are shared more evenly across generations rather than falling disproportionately on workers in the middle of the career ladder.

The results do not imply that continued employment for older workers should be reversed. Elderly employment remains an important response to labor shortages and population ageing, and the OECD (2018) argues that extending working lives is essential for sustaining economic growth and public finances in ageing societies. Rather, the findings suggest that the benefits of longer working lives should be accompanied by institutional arrangements that prevent the adjustment costs from becoming excessively concentrated on a single segment of the workforce.

More broadly, the results suggest that evaluations of ageing policies should consider not only the employment outcomes of older workers and the entry opportunities of younger workers, but also the distributional consequences for the intermediate generations that support both groups. While previous research has largely focused on the aggregate consequences of demographic change, the present findings suggest that ageing policies should also be assessed in terms of how adjustment costs are distributed across different stages of the career ladder.

7.3 Limitations

Three limitations of this research should be acknowledged, though none fundamentally undermines the core findings.

The most important limitation concerns the level at which the demographic regressor is constructed. Because education-group-level tabulations from the Labour Force Survey are used rather than firm-level employment records, the elderly employment share is assigned uniformly to all individuals within an education group and year. Since the wage compression mechanism operates partly through firm-internal wage-setting structures, this introduces aggregation error that likely attenuates the estimated coefficients toward zero. Crucially, this means the results reported here should be interpreted as a lower bound: the true compression and scarcity effects operating within firms are plausibly larger than what the education-group-level measure can detect. The fact that both hypotheses are nonetheless supported with high statistical significance, and that the interaction coefficient remains stable across the year jackknife and cell-size robustness checks, suggests that the effects are sufficiently strong to survive this measurement noise. Future research using

matched employer–employee data with firm-level elderly employment records would allow the mechanisms identified here to be tested at the appropriate level of aggregation and would permit separation of within-firm from between-firm demographic wage effects. Such data would also enable investigation of whether the compression and scarcity effects vary systematically by firm size, given the descriptive evidence in Section 4.4 on differential elderly employment retention across large and small firms.

Second, the panel covers nine survey waves spanning 2016–2024, which, while sufficient to identify the demographic associations that are the focus of this research, limits the ability to track workers across their full career and directly observe whether the entry premium erodes over time. The theoretical framework in Section 3 predicts that the cohort-scarcity premium is a time-limited entry effect concentrated in the first decade of working life, but the present data can only document the premium at the moment of observation rather than trace its full trajectory. Longer panels following young cohorts from entry through prime age would allow a direct test of this prediction.

Third, while the Mundlak correction addresses individual-level confounding by controlling for correlation between time-varying regressors and unobserved individual heterogeneity, the identification strategy cannot fully rule out bias from time-varying unobservables at the education-group-by-year level without an external instrument. The estimates should therefore be interpreted as conditional associations rather than causal effects. Nevertheless, these limitations do not alter the central contribution of the research. Despite relying on an aggregate demographic measure and a relatively short panel, the analysis consistently identifies the same qualitative patterns across multiple robustness checks.

Regardless of these limitations, the analysis consistently identifies the same qualitative patterns across multiple robustness checks and provides new evidence on how population ageing reshapes wage trajectories across different stages of the career within Japan’s regular employment system. The findings thus establish a foundation for future research using matched employer–employee data to test the mechanisms identified here at the firm level.

Acknowledgments

I would like to express my deepest and most sincere gratitude to Professor Murakami Yukiko for her invaluable guidance, encouragement, patience, and unwavering support throughout the entire process of preparing this thesis. From the earliest stages of developing the research question to the final revisions, Professor Murakami continuously provided detailed feedback, thoughtful advice, and insightful comments that greatly strengthened both the analysis and the overall presentation of this research. Her dedication to academic excellence and her willingness to devote considerable time and effort to supervising this research have been truly inspiring. I am profoundly grateful for the opportunity to learn under her guidance, and this thesis would not have been possible without her constant support.

I would like to express my sincere gratitude to Waseda University and the International Office, International Affairs Division, for their generous support through the Overseas Research Travel Grant Program for Upper-level Undergraduate and Graduate Students. This grant enabled me to present a draft of my master's thesis online at the 55th EBES (Eurasia Business and Economics Society) Conference, held in Budapest in April 2026, by covering the conference registration fee, and I am deeply thankful for the opportunity it provided to receive valuable international feedback on my research at this stage of my academic development.

Thank you to the Social Science Japan Data Archive, Center for Social Research and Data Archives, Institute of Social Science, The University of Tokyo, for allowing me to access the Japanese Panel Study of Employment Dynamics, which helped me carry out this research. As a master's student, opportunities to work with large-scale, high-quality panel data of this kind are rare, and I hope this study makes a meaningful contribution that reflects the trust extended through the provision of these data.

I would also like to extend my sincere appreciation to the members of my seminar, whose comments and feedback greatly contributed to the development of this thesis. I am especially grateful to my senior colleague, Wang Yijuan, for her insightful advice and guidance on my research, and to my classmates, Ouyang Yuni and An Jiaju, whose thoughtful comments during seminar discussions helped me refine my arguments and analysis. I am equally thankful to my junior colleague, Izumi Saito, whose questions and suggestions offered valuable perspectives on my work. Their constructive feedback throughout my graduate studies significantly improved the quality of this thesis.

Any remaining errors or omissions are solely my own.

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