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Higher Minimum Wage, Stagnant Income? The case of women's work hours in Japan*

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Abstract

The margin of adjustment in work hours has received relatively less attention in the minimum wage literature, despite its potentially significant implications for distributional consequences. By applying the frequency distribution approach to a quasi-exogenous policy event in Japan, we find that a minimum wage increase reduced long working-hour jobs while increasing short working-hour jobs disproportionately among women. While the minimum wage had a positive compression effect on the wage distribution for women, its impact on their income inequality was much smaller. This reduced effect was driven by substantial reductions in work hours among women with annual incomes near institutional thresholds set by tax and social benefit provisions. The minimum wage, together with these income-based cutoffs, led women in Japan to work shorter hours.

Keywords: minimum wage, work hours, inequality, benefit cliffs, spousal tax deduction.

JEL classification: J20; J31; J38; K31.

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1 Introduction

Numerous studies have examined the consequences of raising the minimum wage. The literature generally agrees on two key points. First, an increase in the minimum wage compresses the hourly wage distribution at the lower end (DiNardo et al., 1996; Lee, 1999; Teulings et al., 2003; Autor et al., 2016; Fortin et al., 2021, etc.).[1] Second, despite some heterogeneity, recent studies report little to no overall disemployment effect (Dube et al., 2010; Allegretto et al., 2017; Harasztosi and Lindner, 2019; Cengiz et al., 2019, etc.).[2] Taken together, these findings suggest that the minimum wage is likely to increase the income of low-wage workers without causing substantial job losses. An important presumption underlying this distributional consequence is that low-wage workers remain working for the similar hours of work. However, this presumption may not hold if individuals adjust their working hours to keep their income below a specific cutoff, whether due to tax and social benefit institutions or other behavioral reasons.

This study draws on the case of women in Japan and provides causal evidence on the unintended effect of the minimum wage on work-hour reductions and its subsequent effect on the earnings distribution. In Japan, a significant proportion of women has been reported to maintain their annual income around ¥1 million (approximately $6,400 as of July 2024), at least since the late 1980s (Abe and Otake, 1995). Such bunching behavior was arguably driven by a set of institutional thresholds set by tax and social benefit provisions. The persistent behavioral response raises concerns that women may reduce their working hours in response to minimum wage increases, thereby limiting gains in their annual income.

To examine this possibility, we begin by estimating the extent to which the minimum wage increase reduced jobs with specific hours, following the frequency distribution approach proposed by Cengiz et al. (2019). Relying on granular variation across hourly wage bins enables us to detect marginal changes in employment headcounts near the minimum wage. The effect on each wage bin is identified using the quasi-exogenous 2007 policy event in Japan, which disproportionately

[1] In addition to the aforementioned studies in the U.S., some evidence of the wage compression effects of the minimum wage has been reported in countries such as Germany (Bossler and Schank, 2023), Canada (Fortin and Lemieux, 2015), the U.K. (Dickens and Manning, 2004), urban China (Howell, 2020), urban Mexico (Bosch and Manacorda, 2010), and Japan (Tachibanaki and Urakawa, 2006; Abe and Tanaka, 2007; Higuchi, 2013; Kambayashi et al., 2013; Yugami, 2016).

[2] Recent studies have reported either insignificant or only small negative employment effects of the minimum wage (Card and Krueger, 1994, 2000; Dube et al., 2010; Allegretto et al., 2017; Harasztosi and Lindner, 2019; Cengiz et al., 2019), while other studies report opposing evidence (Neumark and Wascher, 1992; Neumark et al., 2014; Neumark and Shirley, 2022). See Dube and Lindner (2024) and Clemens (2021) for comprehensive reviews in the literatures.

raised the minimum wage in affected regions to close the gap between welfare benefits and minimum wage earnings. We extend the stack-by-event approach in Cengiz et al. (2019) and compare employment headcounts in each wage bin relative to the new minimum wage across treated and control regions. To ensure comparability of wage bins across regions, we normalize wages by proxies for local wage premia.

Based on large-scale administrative payroll data from Japan, we find that the minimum wage increase led to a rise in short-hour jobs (25 hours or less per week), while reducing the headcount of long-hour jobs (more than 25 hours per week). The increase in short-hour jobs is driven by employment effects concentrated in wage bins up to ¥400 (approximately 2.6 USD as of July 2024) above the post-policy minimum wage, with no significant impact observed in higher wage bins, supporting the credibility of our identification strategy. We also examine the significance of pre-trends for both excess and missing jobs above and below the post-policy minimum wage and find that normalizing bins using proxies for wage premia is essential for constructing comparable wage bins across treated and control regions with respect to pre-policy employment trends. Importantly, the contrasting effects between short-hour and long-hour jobs are predominantly observed among women. These effects are even more pronounced among incumbent female workers than among those with less than one year of tenure, suggesting that work-hour adjustments occurred at the intensive rather than the extensive margin. Such intensive adjustments are stronger among women aged 35 to 54 than among those aged 25 to 34. Our analyses also reveal that the overall employment effect is significantly positive in the retail and restaurant sectors, while it is significantly negative in the manufacturing sector, consistent with previous evidence showing divergent effects across tradable and non-tradable sectors (Harasztosi and Lindner, 2019; Gopalan et al., 2021, etc.). The overall employment elasticity with respect to own wage is estimated to be 0.63 within our sample (s.e. = 0.31), a magnitude comparable to that reported for the U.S. (Cengiz et al., 2019).

Given a substantial rise in short-hour jobs, this study also examines the subsequent consequences on the distributions of hourly wages and annual earnings. In particular, we investigate whether the reduction in wage inequality after the minimum wage hike translates into a reduction in income inequality among women. Using Unconditional Quantile Partial Effects (UQPE) approach (Firpo et al., 2009; Dube, 2019) on the same administrative payroll data, we estimated the minimum wage elasticities for quantiles of hourly wage or annual income distributions. The UQPE approach has an advantage in evaluating overall distributional consequence of the minimum wage hike in the labor market without conditioning on specific groups, yet purging out the covariates from the minimum wage estimates (Firpo et al., 2009; Dube, 2019), under reasonable stability assumptions.[3] To retrieve the elasticity estimates on each quantiles, we need to identify the minimum wage impact on individual worker's probability to fall below threshold values corresponding to each quantile. To address the endogeneity concern, again, we rely on the identification variation arising from the policy change in 2007 to estimate the instrumental variable model proposed by Kawaguchi and Mori (2021). Our analyses reveal contrasting distributional effects of the minimum wage on hourly wages and annual income. UQPE estimates show that while the minimum wage increase raised hourly wages for women up to the 30th percentile of the wage distribution, the gains in annual income were much smaller and even insignificant at the 10th percentile or lower, consistent with the evidence of a shift from long-hour to short-hour jobs.

Work-hour reductions among women can be driven by several factors. The reduction in job size can occur, owing to the employer-side substitution from work hours per employee to employment headcount in the presence of quasi-fixed cost of labor. Workers facing a standard labor supply decision may also reduce work hours if the income effect outweighs the substitution effect. Moreover, institutional cutoffs for tax deductions and social contributions may create incentives to reduce work hours on both the demand and supply sides. To explore the underlying mechanisms, we exploited minimum wage elasticities estimated from UQPE analysis and the fact that individual female employees are classified into one of the 100×100 wage-income quantile cells in UQPE estimations. Specifically, we simulate a scenario in which the minimum wage is hypothetically increased from its 2006 level to the 2014 level, and compute the resulting changes in work hours for each individual observed in 2006. We find that work-hour adjustments are substantially larger among women with annual incomes near the institutional cutoff levels. The sharp pattern observed particularly around ¥1 million is unlikely to be driven by labor demand factors and is more consistent with worker-initiated adjustments in working hours. Taken together, our results

[3] One important presumption in this approach is the stability of distribution over the sample period. We obtained multiple evidence to support this presumption. Most importantly, previous studies have already shown that the minimum wage had no discernible impact on female labor participation by applying the same identification variations (Kawaguchi and Mori, 2021).

suggest that the minimum wage has limited effects on increasing women's earnings in Japan, most likely due to institutional thresholds or related behavioral responses.

The current study contributes to the literature in several ways. In a broader context, we contribute to the literature examining the margins of adjustment, in other words, mechanisms behind zero or small disemployment effects of the minimum wage.[4] The margin of adjustment is particularly important for understanding the distributional implications—specifically, who bears the cost of increased labor costs. Among various adjustment margins, changes in work hours have received comparatively less attention, especially in relation to their welfare consequences. Most recent studies in the U.S. report insignificant or at most modest reductions in work hours following minimum wage increases (Cengiz et al., 2019; Gopalan et al., 2021; Godoey and Reich, 2021; Cengiz et al., 2022; Jardim et al., 2022; Hampton and Totty, 2023; Godøy et al., 2025).[5] Our findings stand in contrast to the U.S. evidence, revealing a substantial decline in long-hour jobs and a concurrent rise in short-hour jobs. This pattern is revealed by examining employment counts across job types with different work hours—details that may be be less visible in aggregated measures such as average hours or full-time equivalent (FTE) employment. We further show that these changes have significant distributional implications, as demonstrated by our UQPE estimates.

Our study also contributes to the literature by providing implications for the growing interest in the interaction effect of minimum wage and other institutions. This interaction is crucial when considering the distributional effects of the minimum wage, as income gains from higher wages may be offset by reductions in means-tested tax credits and transfers. Recent studies have shown that such offsets can attenuate the impact of minimum wage increases on household income equality in the U.S. and U.K. (Dube, 2019; Giupponi et al., 2024).[6] However, direct empirical

[4]To account for the mixed evidence on employment effects, numerous studies have examined alternative adjustment channels on firm side, including increasing product prices (Leung, 2021; Harasztosi and Lindner, 2019; Aaronson, 2007), reduction in firm profit or firm value (Bell and Machin, 2018; Draca et al., 2011), decline in the provision of benefits (Clemens et al., 2018), labor-to-capital substitution (Aaronson and Phelan, 2019; Aaronson et al., 2018; Lordan and Neumark, 2018), labor-to-labor substitution (Clemens and Wither, 2019; Horton, 2017; Giuliano, 2013), productivity gains owing to intensified worker effort (Coviello et al., 2022; Ku, 2022; Hill and E., 2018), exit of low-quality services (Luca et al., 2019), and raising hiring standards (Butschek, 2022; Clemens et al., 2021). See Clemens (2021) and Dube and Lindner (2024) for recent reviews on this strand of the literature. Additionally, some studies identified the heterogeneous employment effects across labor markets with different extent of employer power (Okudaira et al., 2019; Azar et al., 2024).

[5]Outside of the U.S., the evidence is largely mixed. See, for example, Stewart and Swaffield (2008) and Connolly and Gregory (2002) in the U.K., McGuiness and Redmond (2018) in Ireland, Kim et al. (2023) in South Korea, and Kawaguchi and Mori (2021) in Japan.

[6]Related research has also examined effects on household consumption inequality. For example, Yamada (2016) finds that a minimum wage hike in Indonesia increased informal employment, which heightened income uncertainty and limited the reduction in consumption inequality.

evidence on behavioral responses remains relatively limited, with only a handful of studies exploring these mechanisms (Crépon and Kramarz, 2002; Neumark and Wascher, 2011; Kim et al., 2023). Among these, Kim et al. (2023) examined unintended consequences of the social security mandate in South Korea. They found that minimum wage hikes led to an increase in the proportion of workers whose hours fell just below the 60-hour-per-month threshold for mandatory social insurance contributions, suggesting that the minimum wage contributed to the rise in workers exempt from the mandate. More relevant to the present study is Godøy et al. (2025), who applied the frequency-distribution approach to the U.S. Current Population Survey and found that minimum wage increases raised the labor supply of single mothers with small children, in the presence of multiple public assistance programs including the ones with benefit cliffs. This finding contrasts with our results in Japan, where the reduction in work hours occurred primarily among incumbent female workers, with limited responses on the extensive margin of labor supply. One likely welfare consequence beyond income gains is increased time spent with children among parents who reduce their work hours in response to the minimum wage increase. However, the fact that we observe a stronger reduction in work hours among women aged 35 to 54 than among those aged 25 to 34 suggests that this shift contributed less to increased time spent on childcare among mothers with small children. Further research is needed to identify which institutional contexts—particularly within tax and social benefit systems—can enhance the potential of minimum wage policies to support job ladder progression for low-wage workers, especially women.

Finally, this study also adds another piece of causal evidence to the wage compression effect of the minimum wage. Previous studies show spillover effects as an important mechanism for reducing lower-tail wage inequality, though the magnitude of their impact on the wage distribution has been debated (Lee, 1999; Autor et al., 2016; Fortin et al., 2021). Accounting for job loss is also critical when analyzing wage inequality. Cengiz et al. (2019) addressed this issue by applying a frequency-distribution approach to estimate the minimum wage effects on job counts across wage bins. Their analysis detected moderate spillover effects above the minimum wage. In our study, we also detect spillover effects extending up to ¥400 above the minimum wage, which are comparable in size to those reported for the U.S. context (Cengiz et al., 2019; Gopalan et al., 2021). Previous studies in Japan have also reported wage compression effect of the minimum wage (Tachibanaki and Urakawa, 2006; Abe and Tanaka, 2007; Kambayashi et al., 2013; Higuchi, 2013; Yugami, 2016). This study builds on these earlier studies by using more recent data and incorporating a quasi-exogenous policy event for identification, alongside the frequency distribution approach.

The remainder of the paper is organized as follows: Section 2 summarizes the institutional background. Section 3 introduces the dataset utilized in this study and describes the estimating strategy. Section 4 presents the main results from frequency distribution approach. Section 5 presents additional complementary evidence on the distributional consequences of the minimum wage, along with analyses of the underlying mechanisms. Section 6 offers concluding remarks.

2 Institutional Background

2.1 Minimum Wage in Japan

The minimum wage in Japan is largely determined at the regional level. Japan consists of 47 prefectures, each of which sets its own regional minimum wage and considers revisions annually. Historically, minimum wages in Japan originated from inter-firm agreements. However, because enterprise-based negotiations dominated in Japan, some industries and regions were left uncovered, limiting the effectiveness of the minimum wage. To address this issue, the 1968 revision of the Minimum Wage Act introduced regionally applicable minimum wages determined through deliberations at local minimum wage committees, ensuring universal coverage. Consequently, the institution now provides almost complete coverage of workers, applying to both regular and non-regular employees, with only a limited number of exceptions.[7]

Regional minimum wages are deliberated by central and local minimum wage committees every year, which consider workers' living costs, prevailing wages, and the wage-paying capacity of typical firm owners in the region (Art.9, Part 2, Minimum Wage Act). This mechanism ties regional minimum wages to local economic conditions, which could introduce potential endogeneity concerns when analyzing the minimum wage effect. To address this issue, this study focuses on variations in regional minimum wages, leveraging the 2007 policy event to identify the causal impacts of these changes, as explained below.

[7]Exceptions were admitted for individuals with significantly reduced working capacity due to mental or physical disabilities, those undergoing a trial employment period, etc. In addition to regional minimum wages, local labor bureaus may permit small increments for specific industries (referred to as "Tokutei Saitei Chingin"), although these supplementary minimum wages are voluntary and carry no legal penalties.

2.2 Policy Change in 2007

The 2007 amendment to the Minimum Wage Act provides a unique opportunity to identify the effects of minimum wage increases. The amendment addressed long-standing concerns about the "reversal phenomenon" of welfare benefits exceeding minimum wage earnings in some regions.[8] The discrepancies created situations where receiving welfare benefits provided greater financial incentives than working at minimum wage jobs, raising concerns about moral hazard. By the 2000s, the Labor Policy Council of the Ministry of Health, Labor, and Welfare (MHLW) began formal discussions on the need for legal reforms to align regional minimum wages with welfare benefit levels. Japan's minimum wage was also considered low compared to other developed countries, adding further pressure for change. This process resulted in the 2007 amendment, which mandated that regional minimum wages be consistent with welfare benefits (Art. 9, Part 3). This institutional reform led to significant increases in minimum wages, especially in the 12 regions where a positive gap existed prior to the amendment. Because prices remained stable during this period in Japan, these increases occurred in both nominal and real terms.

The continuous increases in the minimum wage following the 2007 amendment significantly raised the proportion of workers affected, particularly in regions with initially higher welfare benefits (i.e., affected regions). Figure 1 shows that affected regions experienced larger increases in the proportion of minimum wage workers compared to unaffected regions after the amendment. Importantly, the proportion of affected employees was nearly identical across the two groups before 2007, suggesting that the policy change introduced sizable and exogenous increases in minimum wage exposure. This variation serves as the basis for identifying the effects of the minimum wage in this study.

For identification to be valid, it is crucial that policy exposure status is not correlated with pre-existing regional economic trends. In affected regions, the relatively high welfare benefits were partly driven by higher housing allowances, which are typically granted in urban areas, or by additional winter heating allowances in colder regions. While these factors may raise concerns about correlation with regional trends, we argue that identification remains valid for two reasons. First, unaffected regions also include urban areas such as Aichi, home to Toyota's headquarters, demonstrating that urban characteristics are not unique to the affected regions. Second,

[8] The reversal phenomenon had been already pointed out by labor unions in the early 1990s, particularly in Toyama and Tokushima Prefectures.

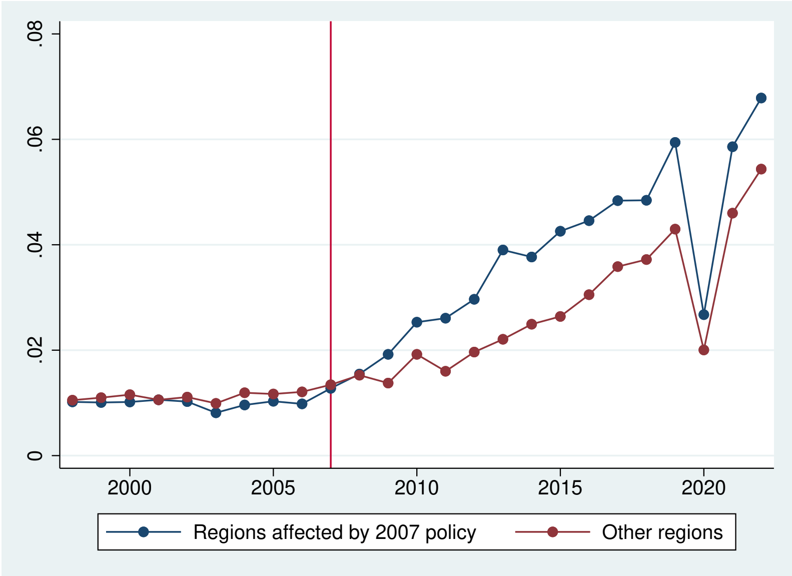

Figure 1: Proportion of Workers with Hourly Wage below New Minimum Wage

Note: The figure shows the proportions of workers with hourly wage rates below the new minimum wage, which takes effect in the following fall. "Regions affected by 2007 policy" are those affected by the 2007 revision of the Minimum Wage Act, specifically prefectures where welfare benefit levels exceeded minimum wage earnings. A red line indicates year of 2007. The data is drawn from Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). The figure adjusts for sampling weights.

as shown in subsequent analyses, there were no significant differences in pre-policy outcome trends between affected and unaffected regions after controlling for regional attributes. Similar identification variations have been used in international publications to examine the effects of minimum wage increases on training (Hara, 2017), housing rents (Yamagishi, 2021), manufacturing sector employment (Okudaira et al., 2019), labor market outcomes for less-educated workers (Kawaguchi and Mori, 2021), and hiring standards and recruiting channels (**?**). In addition to leveraging regional identification variation, this study incorporates granular variations across hourly wage bins, enabling us to account for whether the minimum wage is binding at the bin level.

2.3 Shifts in Wage and Income Distributions

The 2007 revision to the Minimum Wage Act had a significant impact on the distributions of hourly wage and annual earnings. Figure 2 illustrates this point by drawing the frequency distributions of hourly wages and annual income by gender over the period of substantial regional minimum wage increases. Panel A of Figure 2 suggests sizable wage compression effects, consistent with existing studies (DiNardo et al., 1996; Lee, 1999; Teulings et al., 2003; Kambayashi et al., 2013; Autor et al., 2016; Fortin et al., 2021, etc.). On the other hand, Panel B shows much lesser shifts in annual income among women even after the minimum wage hike. One striking pattern is that bunching around the annual income of approximately ¥1 million (6400 USD as of July 2024) became more pronounced after the 2007 policy change, predominantly among women. This pattern implies that changes in hourly wages did not necessarily translate into proportional increases in annual income for women, as many appear to have adjusted their work hours to keep their earnings around the ¥1 million.[9]

2.4 Institutional Cutoffs

The phenomenon of bunching around ¥1 million in annual earnings among women, as illustrated in Figure 2, has been persistently observed since at least the late 1980s in Japan across various datasets (Abe and Otake, 1995; Oishi, 2003; Akabayashi, 2006; Abe, 2009; Takahashi, 2010; Yokoyama and Kodama, 2016; Yokoyama, 2018; Kondo and Fukai, 2023; Hayashi, 2024, etc.). For instance, Abe and Otake (1995) demonstrated bunching just below ¥1 million among part-time workers in 1989, using administrative data, and found that this pattern was evident across all educational levels, including those with university degrees.[10] More recently, Kondo and Fukai (2023) used detailed tax records up to 2021 to show that the distribution of annual earnings continues to exhibit bunching particularly around ¥1.03 million among Japanese women. Importantly, their findings indicate that this bunching behavior becomes pronounced after marriage and persists even as children reach high school age or adulthood (Kondo and Fukai, 2023).

Such persistent bunching behavior among women is arguably driven by a set of tax and so-

[9] Kodama and Momota (2024, in Japanese) used microdata on part-time workers in Japan to examine whether minimum wage increases reduced work hours by estimating labor supply functions. To address the same question, we apply the frequency distribution and UQPE approaches to a large representative dataset, under a flexible set of assumptions.

[10] Akabayashi (2006) used the General Survey of Part-Time Workers in 1995 and reported the bunching below ¥1 million. Similarly, Abe (2009) analyzed microdata from the 2003 National Survey of Family Income and Expenditure, finding that this pattern was prevalent among women aged 30 to 50. Yokoyama (2018) also provided consistent evidence using the 2004 KHPS dataset. Additionally, several studies in Japanese publications observed the same phenomenon. For instance, Oishi (2003) used 1998 administrative micro data from the Comprehensive Survey of Living Conditions to show that married women's annual earnings were concentrated around ¥1.03 million. Takahashi (2010) uses panel data from the Japanese Panel Survey of Consumers (JPSC) covering the years 1994 to 2003 and observes a clustered observations of married women with annual earnings between ¥0.9 and 1 million. Yokoyama and Kodama (2016), using the same Basic Wage Structure Survey data as this study, documented an increasing prevalence of bunching between 1995 and 2013. Hayashi (2024) also analyzed KHPS/JHPS data in 2021 to confirm the bunching below ¥1 million.

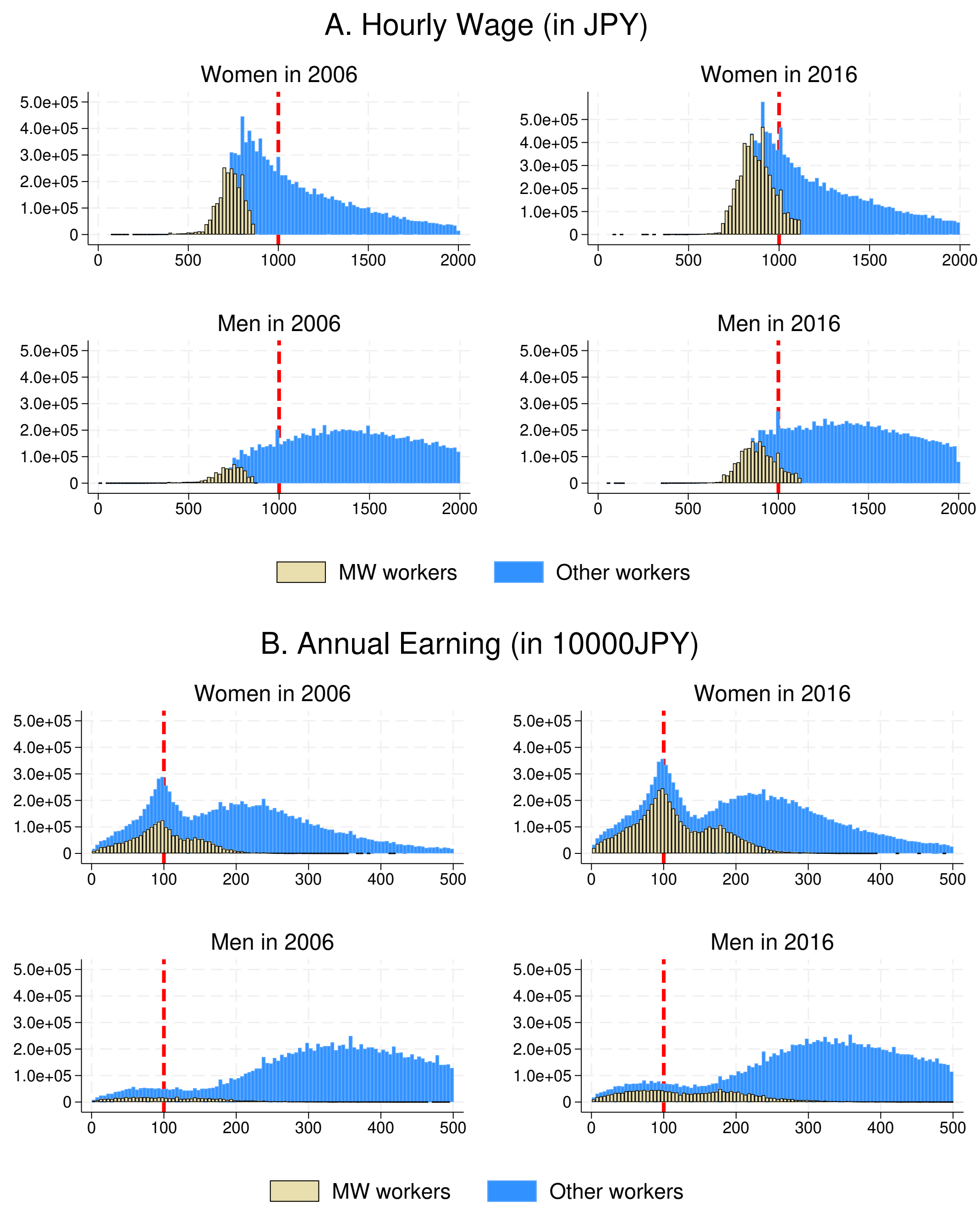

Figure 2: Frequency Distribution by Gender

Note: The figures display frequency distributions for the hourly wage rate (Panel A) and annual earnings (Panel B), based on employee wage records from the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). The hourly wage rate is calculated by dividing the monthly cash earnings in BSWS by actual monthly hours worked, while annual earnings are derived by multiplying monthly cash earnings by an annual conversion weight constructed from the Monthly Labour Survey (Japanese Ministry of Health, Labour and Welfare). Minimum wage (MW) workers are defined as employees whose hourly wage rate is equal to or less than 1.2 times the new minimum wage effective in the following fall. Minimum wages in Japan are set by region. Red dashed lines represent 1,000 JPY in Panel A and 1 million JPY in Panel B. Frequency counts are weighted using sampling weights.

cial benefit thresholds that have created incentives to keep annual income below specific levels. For example, Japan's income tax system is based on individual taxation, but it includes a spousal deduction that allows the partner to receive additional tax deductions as long as the spouse's income remains below a certain threshold. Similarly, under the social insurance system, a spouse of full-time workers are exempt from paying insurance contributions if the spouse's annual earnings fall below a specific threshold, thereby receiving pension and health insurance coverage at no additional cost. To mitigate institutional benefit cliffs faced by married women, the Japanese government has implemented several reforms. While these reforms reduced large cliffs, key programs such as the spousal deduction have remained in place. As a result, the system has become highly complex, introducing convexities into the budget constraints faced particularly by married women. During the period over which our estimating model is identified, the only threshold that generated a substantial earnings cliff for a majority of married women was the ¥1.3 million cutoff, which triggered mandatory contributions to social insurance.[11] Puzzlingly, however, we observe substantial bunching around ¥1 million—rather than ¥1.3 million—in Figure 2 and in other recent studies cited earlier. In particular, the ¥1.03 million threshold has been cited as a point below which married women adjust their income, even though this threshold does not create a cliff in household take-home pay.[12] It merely marks the basic tax exemption level and the starting point for the phase-out of the spousal deduction. This seemingly irrational bunching may be attributed to misperceptions about the complex tax and social insurance system. Kitao and Mikoshiba (2022) argues that policies such as the spousal deduction discourage women from making long-term investments in human capital, increasing the likelihood of entering non-regular employment. It is possible that such long-term decisions, combined with persistent misperceptions and institutional complexities, have contributed to the observed bunching around ¥1 million.

[11]One exception for the benefit cliff is the threshold for women aged 60 or older in the social insurance system. Note that the ¥1.03 million threshold for the special dependent deduction also creates a large earning cliff for those aged 19 to 22, usually unmarried college students. Other thresholds caused only minor discontinuities. For example, ¥1.03 million corresponded to the tax exemption floor, while ¥1.03 million was the point at which Spousal Deduction is replaced by Special Spousal Deduction. ¥1.05 million was also the point at which the Special Spousal Deduction began to phase out. There was also a small cliff at around ¥1 million, created by the threshold for residential tax floor depending on regional authorities.

[12]It is important to note that employers are also reported to offer spousal benefits tied to the ¥1.03 million threshold.

3 Data and Research Design

3.1 Main Data

Our main analysis is based on employee-level payroll data from the Basic Survey on Wage Structure (BSWS), conducted annually in June by the Ministry of Health, Labour and Welfare (MHLW). The BSWS is one of Japan's largest administrative datasets, covering private establishments with five or more employees.[13] Using employee questionnaires from the BSWS from 1998 to 2022, we created a repeated cross-section of payroll records, calculating hourly wage rates and annual earnings for each employee before tax and social insurance contribution. Our sample includes both full-time and part-time employees, but excludes those hired on very short-term contracts of less than one month. Hourly wages were calculated by dividing scheduled cash earnings by actual hours worked, while annual earnings were derived by multiplying scheduled cash earnings in June by conversion weights obtained from administrative monthly wage records.[14] Hourly wages and annual earnings were deflated to 2020 prices using the Consumer Price Index from the Ministry of Internal Affairs and Communications. As the BSWS is not a population survey and samples employees randomly within surveyed establishments, we adjusted the data with sampling weights.[15]

In the frequency-distribution approach below, we calculate employment headcounts per capita for each year-region-wage bin, which requires compiling survey years with comparable headcounts. For this reason, we restrict our sample period to those survey years with consistent and updated population registers, as will be explained later. This ensures that our headcount data remains consistent and accurately reflects the employee population in Japan. For the UQPE analysis in section 5, we use payroll records from 1998 to 2022 to estimate the effects on wage and income distributions.

Table 1 reports summary statistics from the BSWS, disaggregated by job size. Job size is defined based on weekly working hours, specifically using a threshold of 25 hours per week. While

[13]BSWS employs a stratified two-stage sampling design, with establishments as the primary sampling units and employees as the secondary sampling units.

[14]The conversion weights were calculated from Monthly Labour Statistics (MHLW) to represent the ratio of total annual cash earnings (January to December) to monthly cash earnings in June, stratified by employment status (full-time and part-time) and establishment size. Bonuses were excluded as BSWS records bonuses from the previous year. Including them has minimal impact on the earnings distribution for lower-wage female workers.

[15]The employee sampling rate is determined by establishment size and industry classification, ranging from 1 (total enumeration) to 1/5 for establishments with 499 or fewer employees, and between 1/10 and 1/90 for establishments with 500 or more employees. Larger establishments have lower sampling rates.

this study adheres to this definition, we will later relax this assumption in the mechanism section. The summary statistics indicate that approximately 14.3% of workers worked 25 hours or fewer per week during the sample period. Short-hour jobs are predominantly held by women and are overrepresented in the retail and restaurant sectors. The proportion of workers earning below the new minimum wage, which takes effect in the following fall, increased substantially among short-hour jobs after the 2007 revision to the Minimum Wage Act. Before the policy change, approximately 2.5% of short-hour workers earned the minimum wage; this figure rose to 10.5% by 2015. [16]

3.2 Stacked Event Study Model

To evaluate the impact of minimum wages on work hours and employment, we exploit granular variations across wage bins by adopting the frequency-distribution approach proposed by Cengiz et al. (2019). This approach allows us to detect marginal changes in wage bins near the minimum wage. It also incorporates built-in placebo tests by estimating the impacts of the minimum wage on wage bins that are significantly higher than the minimum wage and therefore unlikely to be influenced. We extend the stack-by-event approach outlined in Cengiz et al. (2019, Appendix D) to identify the effects of minimum wages on the headcounts of workers with specific weekly work hours in each wage bin in relation to the new minimum wage.

Our analysis begins by defining each region affected by the 2007 policy change as a separate treatment event. For each event, we construct an event-specific panel dataset that includes employment frequency information across wage bins for both a treated region and its corresponding control regions. A treated region is defined as one that was exposed to the policy change in 2007, where a positive gap between welfare benefits and the minimum wage was initially observed. Among the 12 treated regions affected by the policy change, this study focuses on 10 policy events, excluding Aomori and Akita Prefectures due to insufficient increases in the minimum wage after 2007.[17] Control regions include all other regions that had no such gap and were

[16] Appendix Table A.1 presents the proportions of employees by hourly wage and annual income deciles in our BSWS sample. We observe important gender gap. First, in both wage and income deciles, women are distributed at disproportionately lower end of the deciles when compare against men. This aligns with the fact that Japan still lags behind to close the gender wage gap. Second, women are slightly more densely distributed at the lower end of income deciles than they are in the wage deciles. The proportions of women are higher in the second to fifth deciles of annual income distribution than they are in the hourly wage distribution.

[17] Appendix Figure A.3 presents the actual changes in nominal minimum wages, including Aomori and Akita.

Table 1: Summary Statistics by Job Size

	Short-hour job ($\leq$ 25h/week)	Long-hour job ($>$ 25h/week)
Female	0.719	0.366
Age	47.01	45.79
New entrant (tenure $<$ 1 year)	0.236	0.0857
No bonus in the previous year	0.726	0.205
Industry		
Retail	0.274	0.0853
Restaurant	0.154	0.0185
Accommodation/hotel	0.0180	0.0114
Manufacturing	0.0799	0.265
Mean		
Hourly wage (JPY)	1475.1	1827.7
Annual earning (10000JPY)	96.99	383.6
Annual earning with bonus (10000JPY)	102.5	468.3
Median		
Hourly wage (JPY)	994.0	1567.2
Annual earning (10000JPY)	86.03	344.3
Annual earning with bonus (10000JPY)	87.59	408.5
Proportion of workers below new minimum wage		
in 2006	0.0249	0.00862
in 2008	0.0367	0.0113
in 2014	0.0921	0.0183
in 2015	0.105	0.0193
N of observations	4,377,163	26,269,890

Note: The dataset comprises employee payroll data from 1998 to 2022, drawn from the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). Sample statistics are calculated by using sampling weights. Hourly wages and annual earnings are deflated using the Consumer Price Index (CPI) with 2020 as the base year. "Annual earning with bonus" indicates a sum of annual earnings and a bonus from the previous year. "Proportion of workers below new minimum wage" indicates the proportions of workers with hourly wage rates below the new minimum wage, which takes effect in the following fall.

therefore unaffected by the policy change. Ideally, control regions should be otherwise similar and have identical wage distributions, with each wage bin containing jobs comparable to those in the corresponding wage bin in the treated region. However, ensuring this assumption is not straightforward, as wage distributions may vary across regions: for example, ¥1000-jobs in urban regions may differ from ¥1000-jobs in less urbanized regions in terms of job characteristics and skill levels.

To address this concern, we normalize hourly wages in each region by the wage premia in that region to ensure that the normalized wage bins are comparable across treated and control regions. In our baseline specification, we define this wage premia as the pre-policy minimum wage in each region, assuming that minimum wage jobs were comparable across treated and control regions before the policy change. This is a reasonable assumption, as the minimum wage was previously determined using similar statistics across regions, such as regional living costs. In fact, the proportion of minimum wage workers in treated and control regions was nearly identical before the 2007 revision of the Minimum Wage Act, as shown in Figure 1. More specifically, we subtract the 2006 minimum wage in region p (w_p^*) from an hourly wage of an individual i living in region p in year t:

$$w_{ipt} = w'_{ipt} - w_p^* \tag{1}$$

where w' is the raw hourly wage before normalization, and w is normalized hourly wage. In the following estimation model, we compare employment headcounts in a treated region across wage bins defined over these normalized hourly wages with those wage bins in control regions within each event-specific dataset. As we will show later, our proxy for the wage premia sufficiently accounts for potential pre-existing bin-specific employment trends. To test the sensitivity of our results, we also modify the wage premia by incorporating estimated region-specific effects, following an approach similar in spirit to Giupponi et al. (2024). The results remain robust to this specification change.[18] Unless otherwise stated, our discussion below will always refer to wages, wage bins, or minimum wages after normalization by wage premia.

For each event dataset h, we calculate per-capita employment counts in each region-year-¥25 wage bins relative to the new minimum wage.[19] The new minimum wage is defined as the effective minimum wage in the treated region as of 2014, which approximately corresponds to the original minimum wage target set by the government specifically for that treated region to close the gap between welfare benefit and the minimum wage-level earning. Thus, wage bins for both treated and control regions are defined by the distance from the new minimum wage in the treated region. This approach compares per-capita employment counts in the same wage bins, defined by their distance from the new minimum wage in the treated region, across treated and

[18] We chose not to normalize by the estimated local wage premia, as we found slightly pronounced pre-trends in excess and missing jobs, as shown in Appendix Figure A.6.

[19] ¥25 is approximately equal to 0.17 USD as of November 2024. Per-capita employment counts are calculated by dividing the employment counts in the BSWS by the regional population, based on the Population Estimates (the Ministry of Internal Affairs and Communications). When collapsing the original employee data into a bin-region-year panel, we adjusted the data using sampling weights.

control regions. Because wages and wage bins in each region were normalized by subtracting the pre-policy minimum wage, as discussed above, the types of workers in these wage bins are again assumed to be otherwise similar between the two groups.

We then stacked all the 10 event-specific datasets to estimate the following event study model on changes in per-capita employment count since 2006 for region p in year t in event h dataset:

$$\Delta E_{pjth} = \sum_{k \in K} \alpha_{tk} I^k_{pjth} + \delta_t + t_h + t_j + \Omega_{pjth} + \epsilon_{pjth}. \tag{2}$$

where j indicates the jth ¥25 bin relative to the new minimum wage. ΔE_{pjth} represents the change in per-capita numbers of jobs since 2006 (i.e., $\Delta E_{pjth} = E_{pjth} - E^{2006}_{pjh}$). In the main analysis below, we estimate this model, separately for short-hour (≤ 25 hours per week) and long-hour jobs (> 25 hours per week). I^k_{pjth} is a treatment indicator that equals 1 if the jth bin in the treated region falls between k and $k + 100$ yens relative to the new minimum wage (MW). This treatment variable allows us to capture various minimum wage effects across wage bins, particularly those above or below the minimum wage target set for the treated region. For instance, $k = 0$ indicates four ¥25 bins at or above the new minimum wage in the treated region, specifically the bins within the interval $[\text{MW}, \text{MW} + ¥100)$. Similarly, $k = -100$ indicates four bins below the new minimum wage in the treated region, specifically bins within the interval $[\text{MW} - ¥100, \text{MW})$. In this study, we set $k \in K = \{-\infty, -200, -100, 0, 100, \ldots, 800, 900, \infty\}$, with $k = -\infty$ or $k = \infty$ indicating the wage bins below (MW $-$ 200) yen or the wage bins at or above (MW + 1000) yen, respectively, in the treated region. By stacking all events, we estimate the average treatment effects of the event-specific impacts of minimum wage increases resulting from the 2007 policy change. The term δ_t absorbs common year effects on overall wage distributions. The terms t_h and t_j controls for linear trends specific to each event and each wage bins.[20] The error term ϵ_{pjth} is clustered at the event-region level to account for potential within-group correlations. The estimation is weighted by regional population.

One identification concern in the context of Japan is the potential confounding effect of small minimum wage increases in control regions. Since minimum wages increased moderately across all regions, it is essential that the impacts of the 2007 policy-induced minimum wage increase are identified separately after accounting for small minimum wage increases in the control regions.

[20] Note that we take a change in outcome variable since 2006, rather than a first-difference in the outcome variable.

We achieve this by relying on granular bin-level information in both treated and control regions. To account for small increases in the minimum wage within the control group, we create a dummy variable for each wage bin indicating a small minimum wage increase and interact it with a post-policy period dummy. In the above equation, we control for Ω_{pjth}, a set of dummies indicating bins in the control regions affected by the new minimum wage increases in the control regions within each event dataset h, interacted with the post-policy period dummies.[21] By including these dummies, we absorb common impacts on a subset of wage bins affected by small minimum wage increases in control regions. This approach follows Cengiz et al. (2019), who addressed similar small increases in state or federal minimum wages in the U.S.

In our initial specification, we allow the estimated treatment effects (α_{tk}) to vary across years t. Thus, the estimate for α_{tk} represents the treatment effect on the cumulative change in employment counts from 2006 to year t. By differencing the outcome variable, we can control for unobserved time-invariant effects specific to each region, each wage bin, each event, and any combinations of these dimensions. For the above stacked regression to identify the minimum wage effect, we assume that both treated and control regions would have followed similar employment trends in the absence of the policy change. To assess the validity of this assumption, we examine the significance of the pre-existing trends by checking the estimates for α_{tk} with $t < 2006$.

3.3 Sample Construction and Empirical Implementation

To construct the event-specific datasets, we use the individual payroll records from the BSWS. Because the frequency distribution approach relies on employee headcounts, it is crucial to ensure that a consistent sampling approach was adopted across years. This study focuses on selected survey years from the BSWS to align with its population register. Specifically, we use payroll records from years in which the establishment population register was constructed within two years prior to the survey (t = 1998, 2006, 2008, 2014 and 2015). Focusing on these survey years allow us to eliminate some survey years in which population register was not updated and thus employees from relatively new establishments were omitted from the sample. In short, the final

[21] Specifically, separately for each event, we constructed six variables to indicate bins that fall above or below the minimum wage in the control regions by interacting $\{POST\} \times \{(mw_c - 200, mw_c - 100], (mw_c - 100, mw_c], (mw_c, mw_c + 100], (mw_c + 100, mw_c + 200], (mw_c + 200, mw_c + 300], (mw_c + 300, mw_c + 400]\}$, where $(mw_c - 200, mw_c - 100]$ is a dummy that takes the value of 1 for bins in the control region that fall between -200 and -100 from the new minimum wage in the control region (mw_c), for instance, and $POST$ is a dummy that takes the value of 1 after 2008. The new minimum wage here refers to the minimum wage in 2014 in control regions.

sample in the frequency approach captures employment headcounts from a representative sample of establishments in Japan, except for the new establishments opened less than a year. Appendix Table A.2 compares summary statistics for these selected survey years (Panel A) with those for the entire sample period (Panel B), which will be used in the inequality analysis later in this study. Limiting the survey years does not alter the summary statistics at all, despite discrepancies from the population register.

In the following analysis, we examine the treatment estimates separately by survey years (t = 1998, 2008, 2014, and 2015; 2006 as a base year) to investigate the dynamic evolution of the effects, as well as potential pre-policy trends. We also present the aggregated estimated treatment effects by calculating the average treatment effects for 2014 and 2015, since the positive gap between welfare benefits and minimum wage earnings was fully closed by 2015. We divide this average treatment effect between 2014 and 2015 by the sample average of pre-policy employment-to-population ratio in treated regions ($\overline{EPOP_{-1}}$) to obtain the main treatment estimates, following the approach of Cengiz et al. (2019).[22] Thus, for each wage bin relative to the new minimum wage, we define the main treatment effect as $\alpha_k = (\alpha_{k2014} + \alpha_{k2015}) \times 0.5/\overline{EPOP_{-1}}$. This procedure normalizes the average treatment effects by the pre-policy employment level rather than the population. Consequently, our estimates for excess jobs ($\Delta a = \sum_{k=0}^{300} \alpha_k$) and missing jobs ($\Delta b = \sum_{k=-\infty}^{-100} \alpha_k$) are also normalized to reflect the differences between actual and counterfactual employment counts relative to the pre-treatment total employment. The sum of Δa and Δb ($\Delta e = \Delta a + \Delta b$) represents the estimated percentage change in total employment driven by the minimum wage increase. We also estimate employment elasticities with respect to the minimum wage and own-wage. The latter is calculated by dividing the estimated percentage change in affected employment by the percentage change in affected wages.[23]

[22]We obtain the pre-policy employment-to-population ratio by dividing the number of workers in 2005 by the population counts in 2005. This information was derived from the Population Census and the Population Estimates provided by the Ministry of Internal Affairs and Communications.

[23]We followed the procedure outlined in Cengiz et al. (2019) to construct these estimates.

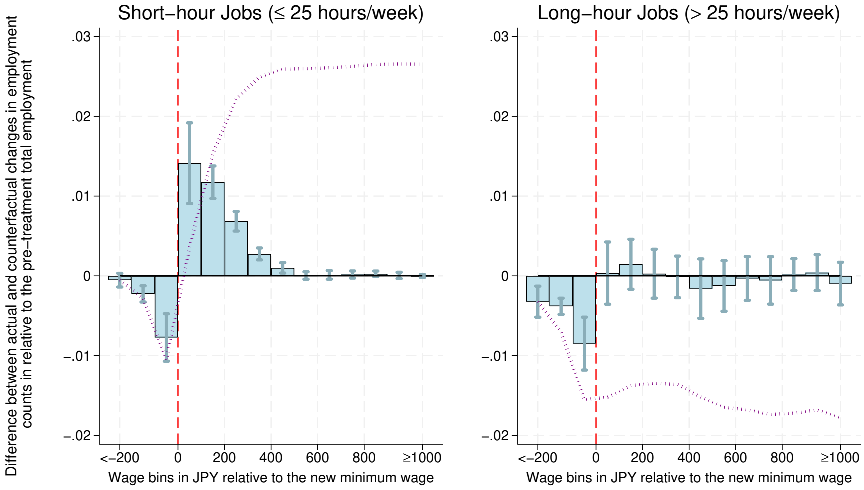

Figure 3: Impact of Minimum Wage on the Wage Distribution by Size of the Job

Note: The figures plot the estimated minimum wage effects on employment counts from an event-study analysis. We exploit the 10 events of the minimum wage increases after the 2007 revision to the Minimum Wage Act in Japan. We stack all event-specific region-bin level datasets to estimate an average effect across 10 events using a single set of treatment effects. The dataset is drawn from employee observations in 1998, 2006, 2008, 2014, and 2015 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). The ploted estimates are the averages of the treatment effects in 2014 and 2015. All the estimations control for year dummies, linear trends specific to events and wage bins, and dummies to indicate small minimum wage increases in the control regions after 2007. All the estimations are weighted by the size of regional population. Standard errors are cluster at region and event level. Purple dashed line indicates a running sum of the estimated impacts at each bin.

4 Main Results

4.1 Frequency Distribution Estimates by Job Size

The frequency distribution approach has advantages over other methods, particularly in identifying the minimum wage impacts at the wage-bin level before aggregating them into overall employment effects. We begin by presenting the estimated impact at each bin relative to the new minimum wage, focusing on the average treatment effects in 2014 and 2015 ($\hat{\alpha_k}$). Figure 3 shows the bin-wise treatment effects separately for short-hour and long-hour jobs, including both men and women. The treatment effects on the vertical axis represent the difference between actual and counterfactual employment counts relative to pre-treatment total employment. The purple dashed line shows the cumulative sum of the estimated impacts across bins, essentially reflecting the total employment effect within each category.

The two graphs in Figure 3 suggest contrasting effects of the minimum wage by work-hour category. Consistent with previous studies using the frequency approach (Cengiz et al., 2019;

Gopalan et al., 2021; Azar et al., 2024; Godøy et al., 2025), we observe significant reductions in jobs below the new minimum wage for both short-hour and long-hour jobs. However, while short-hour jobs show large and positive increases above the new minimum wage, no significant increase is observed for long-hour jobs. We also observe spillover effects among short-hour jobs, extending up to ¥400 (approximately 2.6 USD as of July 2024), which is consistent with previous findings in the U.S. (Cengiz et al., 2019; Gopalan et al., 2021). Importantly, we find no significant treatment effects on wage bins that are unlikely to be affected by the minimum wage increase, such as those more than ¥1000 above the minimum wage, supporting the credibility of our identification strategy.

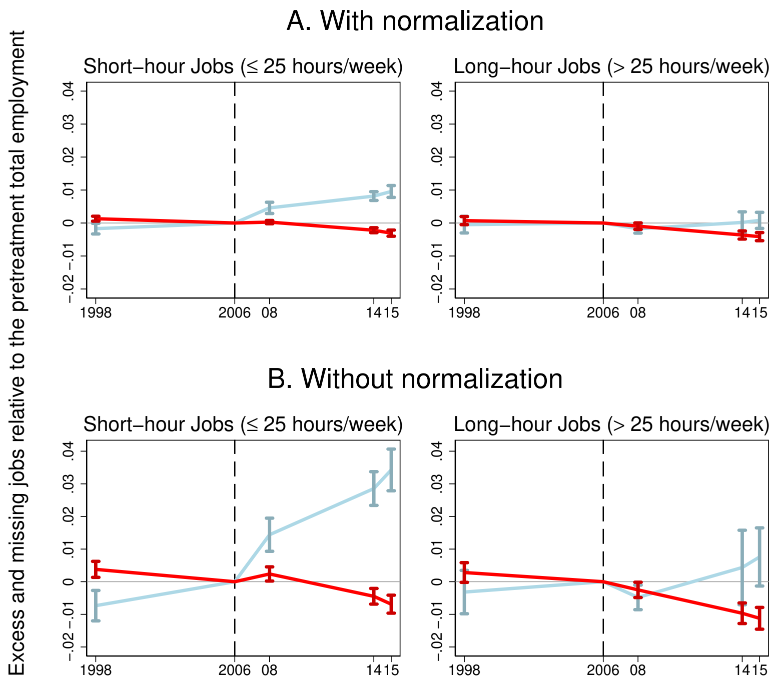

Figure 4: Impact of Minimum Wage on Missing (Red) and Excess (Blue) Jobs Over Time

Note: The figures plot the estimated minimum wage effects on employment counts from an event-study analysis, separately for missing and excess jobs. Panel A shows the estimates with wage bins normalized by the minimum wage effective in 2006. Panel B shows the estimates with wage bins without normalization. In both estimations, we exploit the 10 events of the minimum wage increases after the 2007 revision to the Minimum Wage Act in Japan. We stack all event-specific region-bin level datasets to estimate an average effect across 10 events using a single set of treatment effects. The dataset is drawn from employee observations in 1998, 2006, 2008, 2014, and 2015 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). All the estimations control for year dummies, linear trends specific to events and wage bins, and dummies to indicate small minimum wage increases in the control regions after 2007. All the estimations are weighted by the size of regional population. Standard errors are cluster at region and event level.

The estimated treatment effects on each bin are then aggregated into the treatment effects on missing and excess jobs. Figure 4 shows, Δa and Δb, separately by survey years (base year = 2006). Panel A presents the estimates from our baseline specification in equation (2), which includes controls for bin-specific linear trends, after normalizing the wage bins by proxies for the wage premium in equation (1). Panel B shows the estimates without normalization but with a same set of control variables. A comparison between the two panels indicates that normalization of wage bins eliminates differences in employment trends prior to the 2007 policy change, allowing for the construction of valid counterfactual bins across treated and control regions. In the main analysis here, we used the pre-policy minimum wage as a proxy for wage premia. To test the sensitivity of our results, we also replace the proxy by the estimated region-specific wage effects, following the approach of Giupponi et al. (2024). Specifically, we normalize each wage bin using region fixed effects obtained from a Mincer-type wage regression that controls for individual traits (**check**) in addition to region dummies. Appendix Figure A.5 presents the estimation results using this alternative specification. The results remain robust to this change. However, we chose not to adopt this normalization in the main analysis, as we found slightly pronounced pre-trends in excess and missing jobs under this specification, as shown in Appendix Figure A.6. Appendix Figure A.4 also examines the sensitivity against excluding bin-specific linear trends from Panel A of Figure 4 and confirms that controlling for bin-specific linear trends play some roles in eliminating the pre-trends. Given the 2007 policy change and the normalization of wage bins, we consider that the common trend assumption is likely to hold.

Histograms for annual earnings shown in Figure 2 implied gender differences in work-hour adjustments. To examine this further, Figure 5 presents bin-wise treatment effects, averaged over 2014 and 2015, by gender and job category. The results indicate that the contrasting pattern observed in Figure 3 is primarily driven by women. Table 2 evaluates the total employment effect by calculating employment elasticities with respect to minimum wage, in addition to the relevant statistics. Comparing the total employment effect with the separate effects on short-hour and long-hour jobs suggests that focusing only on the total effect masks important heterogeneity in work-hour distributions. The employment elasticity for short-hour jobs is 0.14 for women (affected share = 0.25), while it is 0.06 for men (affected share = 0.19). On the other hand, the employment elasticity for long-hour jobs is -0.13 for women (affected share = 0.09), while it is small

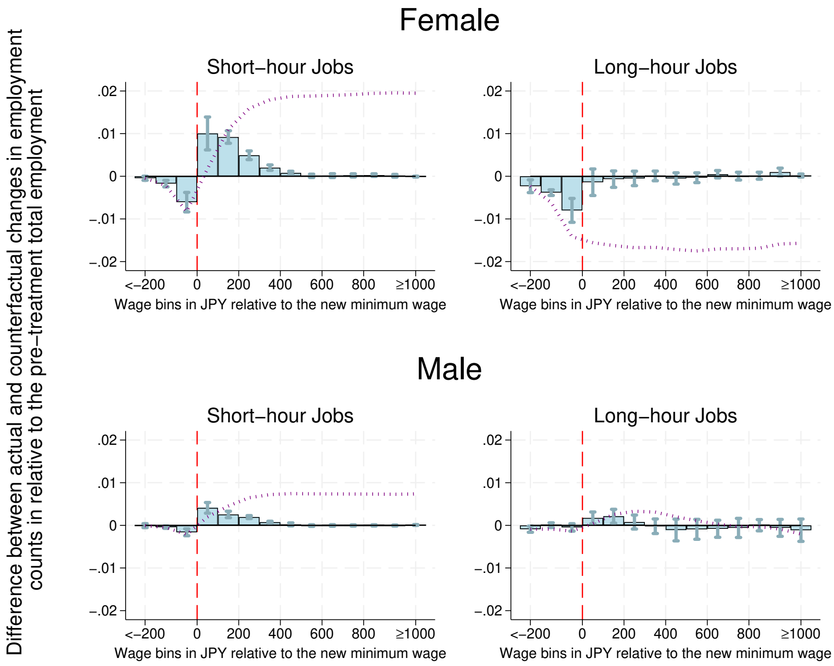

Figure 5: Impact of Minimum Wage on the Wage Distribution by Job Size and Gender

Note: The figures plot the estimated minimum wage effects on employment counts from an event-study analysis. We exploit the 10 events of the minimum wage increases after the 2007 revision to the Minimum Wage Act in Japan. We stack all event-specific region-bin level datasets to estimate an average effect across 10 events using a single set of treatment effects. The dataset is drawn from employee observations in 1998, 2006, 2008, 2014, and 2015 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). The ploted estimates are the averages of the treatment effects in 2014 and 2015. All the estimations control for year dummies, linear trends specific to events and wage bins, and dummies to indicate small minimum wage increases in the control regions after 2007. All the estimations are weighted by the size of regional population. Standard errors are cluster at region and event level. Purple dashed line indicates a running sum of the estimated impacts at each bin.

and statistically insignificant for men.

The overall employment elasticity with respect to own wage (OWE) is estimated to be 0.63 in our sample (s.e. = 0.31), a magnitude comparable to that reported for the U.S. Using the frequency distribution approach, Cengiz et al. (2019) estimated the overall OWE in the U.S. dataset to be 0.41 (s.e. = 0.43). Our estimate is higher than the median of OWE estimates reported in the repository proposed in Dube and Zipperer (2024). A likely explanation is that our payroll data exclude establishments with fewer than five employees and those within their first year of operation, understating the extent of labor-to-capital substitution through firm entry and exit dynamics (Aaronson et al., 2018).

Table 2: Impact of Minimum Wage on Employment and Wages by Gender and Job Size

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				Short-hour jobs		Long-hour jobs	
	All	Female	Male	Female	Male	Female	Male
Missing jobs below new MW (Δ b)	-0.026***	-0.023***	-0.004***	-0.008***	-0.002***	-0.014***	-0.001
	(0.003)	(0.002)	(0.001)	(0.001)	(0.000)	(0.002)	(0.001)
Excess jobs above new MW (Δ a)	0.037***	0.024***	0.014***	0.026***	0.009***	-0.002	0.004
	(0.007)	(0.004)	(0.003)	(0.002)	(0.001)	(0.003)	(0.003)
%Δ affected employment	0.149	0.008	0.318**	0.072***	0.038***	-0.182***	0.166
	(0.098)	(0.029)	(0.118)	(0.008)	(0.005)	(0.027)	(0.167)
%Δ affected wage	0.235***	0.083***	0.196***	-	-	-	-
	(0.042)	(0.018)	(0.043)				
Employment elasticity w.r.t. MW	0.086	0.008	0.079**	0.139***	0.056***	-0.129***	0.024
	(0.057)	(0.031)	(0.029)	(0.016)	(0.008)	(0.019)	(0.024)
Employment elasticity w.r.t. affected wage	0.633*	0.091	1.627***	-	-	-	-
	(0.313)	(0.329)	(0.271)				
Proportion affected in 2006	.075	.137	.032	.248	.191	.091	.019
N of observations	165600	165600	165600	165600	165600	165600	165600

Note: The table reports the estimated impacts of the Minimum Wage increases from event-study estimates based on the 2007 revision to the Minimum Wage Act in Japan. We stack all event-specific region-bin level datasets to estimate an average effect across 10 events using a single set of treatment effects. The dataset is drawn from employee observations in 1998, 2006, 2008, 2014, and 2015 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). "Short-hour jobs" are those employment with less than or equal to 25 hours of work per week."Long-hour jobs" are those employment with more than 25 hours of work per week. All the estimations control for year dummies, linear trends specific to events and wage bins, and dummies to indicate small minimum wage increases in the control regionals after 2007. All the estimations are weighted by the size of regional population. Standard errors are clustered at region and event level.

OWE is a useful metric for interpreting employment responses to minimum wage changes, taking into account the degree to which workers are bound by the minimum wage. Following Cengiz et al. (2019), we compute OWE in Table 2 by aggregating bin-level employment changes and the resulting changes in the hourly wage bill, relative to the pre-policy total wage bill among workers initially earning below the post-policy minimum wage. However, because our analysis is stratified by weekly hours, this approach is not suitable for calculating OWE separately for short- and long-hour jobs. We therefore do not report OWE in the following analyses. Since the policy interpretation of employment elasticities depends on the share of workers affected, we report both the estimated elasticity and the pre-policy proportion of workers below the new minimum wage in the tables that follow.

4.2 Heterogeneity and Sensitivity

This subsection presents additional analyses to examine the heterogeneity and robustness of our main results. In summary, we find substantial heterogeneity by industry, worker status (new entrants vs. incumbents), and age among female workers. We also conduct some sensitivity tests to assess the robustness of our findings. The results also remain robust across several sensitivity checks.

Columns 1 to 4 of Table 3 explore heterogeneity in the overall employment effects by industry. These estimates include employment headcounts across both genders and both work-hour categories. Our analyses reveal that the overall employment effect is significantly positive in retail and restaurant sectors, while it is significantly negative in manufacturing sector, consistent with previous evidence showing the divergent effects across tradable and non-tradable sectors (Harasztosi and Lindner, 2019; Gopalan et al., 2021, etc.). A significant and negative employment effect is consistent with a previous finding in Okudaira et al. (2019).

Identifying whether the shift from long-hour to short-hour jobs occurs along the extensive or intensive margin is crucial for understanding labor supply responses to minimum wage increases. Godøy et al. (2025) applied the frequency-distribution approach to the U.S. Current Population Survey and found that minimum wage hikes increased the labor supply of single mothers with small children. Columns 5 to 8 of Table 3 examine this issue in the context of Japan. The results show that the contrasting effects between short-hour and long-hour jobs are more pronounced

Table 3: Heterogeneity of Employment Effects

	(1) Retail	(2) Food	(3) Hotel	(4) Manuf.
Employment elasticity w.r.t. MW	0.061***	0.055***	-0.000	-0.057***
	(0.018)	(0.010)	(0.003)	(0.013)
Proportion affected in 2006	.184	.223	.079	.062
	(5)	(6)	(7)	(8)
	New entrants (female)		Incumbents (female)	
	Short-hour	Long-hour	Short-hour	Long-hour
Employment elasticity w.r.t. MW	0.028***	-0.027***	0.114***	-0.098***
	(0.005)	(0.007)	(0.012)	(0.014)
Proportion affected in 2006	.35	.141	.214	.084

Note: The table reports the estimated impacts of the Minimum Wage increases from event-study estimates based on the 2007 revision to the Minimum Wage Act in Japan. We stack all event-specific region-bin level datasets to estimate an average effect across 10 events using a single set of treatment effects. The dataset is drawn from employee observations in 1998, 2006, 2008, 2014, and 2015 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). "Short-hour jobs" are those employment with less than or equal to 25 hours of work per week."Long-hour jobs" are those employment with more than 25 hours of work per week.All the estimations control for year dummies, linear trends specific to events and wage bins, and dummies to indicate small minimum wage increases in the control prefectures after 2007. N = 165600. All the estimations are weighted by the size of prefecture population. Standard errors are clustered at region and event level.

among incumbent female workers than among those with less than one year of tenure, suggesting that work-hour adjustments occurred primarily along the intensive margin. This contrasts with the findings of Godøy et al. (2025), but aligns with Kawaguchi and Mori (2021), who use Japanese Labor Force Survey data and find a positive but statistically insignificant effect of the minimum wage on labor force participation among less-educated women.

Table 4 examines which age groups of women experienced the strongest shift toward short-hour jobs. The results indicate a clear substitution from long-hour to short-hour jobs among mid-aged women, suggesting that the patterns observed among incumbent workers in the previous table are likely concentrated in this age group. In contrast, the effect is less pronounced among women aged 25 to 34, who are more likely to be of childbearing age or have small children. Among older women, the increase in short-hour jobs appears to reflect new entries into part-time employment. For women under 25, the strong substitution away from long-hour jobs may be partly driven by income thresholds associated with the dependent tax deduction, which allows parents to claim a sizable deduction from their taxable income if the dependent, who is under the

Table 4: Heterogeneity of Employment Effects by Age Group (Women)

	(1) under 25	(2) 25 to 34	(3) 35 to 44	(4) 45 to 54	(5) 55 to 64
A. Short-hour jobs (≤ 25 h/week)					
Employment elasticity w.r.t. MW	0.033***	0.008**	0.032***	0.028***	0.033***
	(0.006)	(0.003)	(0.004)	(0.004)	(0.008)
Proportion affected in 2006	.404	.215	.21	.206	.235
B. Long-hour jobs (> 25 h/week)					
Employment elasticity w.r.t. MW	-0.059***	-0.013**	-0.022***	-0.027***	-0.009
	(0.008)	(0.005)	(0.004)	(0.005)	(0.005)
Proportion affected in 2006	.068	.052	.087	.113	.163

Note: The table reports the estimated impacts of the Minimum Wage increases from event-study estimates based on the 2007 revision to the Minimum Wage Act in Japan. We stack all event-specific region-bin level datasets to estimate an average effect across 10 events using a single set of treatment effects. The dataset is drawn from employee observations in 1998, 2006, 2008, 2014, and 2015 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). "Short-hour jobs" are those employment with less than or equal to 25 hours of work per week."Long-hour jobs" are those employment with more than 25 hours of work per week.All the estimations control for year dummies, linear trends specific to events and wage bins, and dummies to indicate small minimum wage increases in the control regions after 2007. N = 165600. All the estimations are weighted by the size of regional population. Standard errors are clustered at region and event level.

age of 23, has annual earnings below ¥1.03 million. Appendix Table A.5 shows that restricting the sample to women aged 23 to 59 reduces the estimated effect observed in Table 2 by about half, suggesting that the dependent tax deduction may play a role in the observed patterns.

One likely welfare consequence beyond income gains is increased time spent with children among parents who reduce their work hours in response to minimum wage increases. However, the fact that work-hour reductions are stronger among women aged 35 to 54 than among those aged 25 to 34 suggests that the shift contributed less to increased time allocation to childcare among mothers with small children.

Finally, we conducted several sensitivity tests on the main results reported in Table 2. First, we added additional regional control variables to the baseline specification. Specifically, we controlled for the population shares of six demographic groups—defined by gender and three age categories—within each prefecture. The results remain largely unchanged, as reported in Appendix Table A.3. Second, we tested the robustness of our results to alternative thresholds for defining short-hour and long-hour jobs. While the baseline analysis uses a cutoff of 25 hours per week, Appendix Table A.4 presents results using thresholds of 20 and 30 hours. As expected, the estimated employment elasticity for short-hour jobs becomes more positive as the threshold increases, although the pre-policy proportion affected remains largely stable. In contrast, for long-hour jobs, the elasticity becomes more negative as the threshold increases, while the pre-policy proportion affected declines.

5 Distributional Consequence and Mechanism

5.1 Unconditional Quantile Partial Effect Estimation

Previous research has shown that increases in the minimum wage often lead to a compression of the wage distribution (DiNardo et al., 1996; Lee, 1999; Teulings et al., 2003; Kambayashi et al., 2013; Autor et al., 2016; Fortin et al., 2021, etc.). However, since the minimum wage increase disproportionately led to a rise in short-hour jobs among women as shown in the previous section, the overall annual earnings of women may not have been increased substantially despite a compression effect on hourly wage distribution. Consequently, the impact of the minimum wage on annual income inequality may be limited. This section examines the consequence of work-hour adjustments by comparing the effects of the minimum wage on income distribution with those on wage distribution using BSWS payroll records. Although the BSWS is the largest administrative payroll dataset with broad coverage and establishment size is adjusted using sampling weights, its employee payroll records primarily sample workers in establishments with five or more employees and do not include individuals out of employment. Thus, the results in this section should be interpreted as reflecting trends in distributional consequences within the scope of our sample. For a more comprehensive analyses on inequality, our companion paper, Mori and Okudaira (2025), examines the impact of the minimum wage using detailed administrative household-level data.

We apply the unconditional quantile partial effect (UQPE) approach proposed by Firpo et al. (2009) to examine the distributional consequences of the minimum wage. Specifically, we follow a version of the approach taken by Dube (2019) in his study on the minimum wage and household income inequality in the United States. Intuitively, this approach first estimates the extent to which the minimum wage increase shifted the cumulative density function (CDF) of an outcome variable (i.e., hourly wage rate or annual income in this study). By inverting the estimated impact on the CDF, it then identifies its impact on specific quantiles of the unconditional distribution, such as changes in the median wage or income before and after the 2007 policy change. The approach is built on the recentered influence function (RIF), which quantifies the incremental effect of adding an individual observation to the original distributional statistic (Firpo et al., 2009).[24] It is relatively computationally simple, as it primarily involves running a series of regressions on a dummy variable indicating whether the outcome variable falls below a specific cutoff value and using density estimates to derive the impact on wage or income quantiles. The UQPE estimation also has advantages in examining the distributional consequence of the policy, because it estimates the impact on the entire distribution without conditioning on specific attributes, such as family background or educational level. Moreover, it still allows for the estimation of partial effects, as confounding factors are accounted for when running the regressions on the CDF. To accurately estimate the impact of the minimum wage on the CDF using the UQPE approach, we must ensure the validity of several underlying assumptions. We consider these conditions to be met in the context of this study.[25]

We implement our procedure following Firpo et al. (2009) and Dube (2019). We first construct a dummy variable that takes the value of 1 if wage or income falls below a specific threshold value, c. We then specify a model where this dummy variable is regressed on the minimum wage:

$$I(y_{it} < c) = \alpha_c \ln(MW_{pt-1}) + X_{it}\gamma_c + t \cdot \mu_{cp} + \tau_{cp} + \theta_{ct} + \epsilon_{it} \tag{3}$$

where the model controls for year effects, region-specific trends, in addition to polynomials of individual age (X_{it}). The regression is weighted by sampling rates. Standard errors are clustered at the region level. The estimates for α_c are used to obtain the UQPE by inverting the cumulative

[24]More detailed formalization is provided in Appendix A.

[25]There are three main assumptions that support the validity of our approach. First, the minimum wage should change continuously rather than discontinuously. Second, the density function should be approximately linear in the lower half of the distribution. Finally, the wage or income quantiles in the lower part of the distribution should remain stable over time. The first condition is satisfied since the Japanese minimum wage is revised continuously across regions rather than experiencing abrupt changes (see Section 2). As discussed in the previous section, we exploit the exogenous variation in minimum wages. However, the identified first-stage estimates still suggest continuous rather than discrete variation, as presented in Appendix Table A.6. We also demonstrate in Appendix A that the second assumption is likely to hold in this study. Additionally, it is important to note that previous research applying the same identification strategy to individual-level microdata from the Labour Force Survey (Ministry of Internal Affairs and Communications) has shown that the minimum wage did not affect female labor force participation (Kawaguchi and Mori, 2021) . This suggests that the unconditional wage and income distributions remain relatively stable over the sample period.

distribution function (CDF) of the outcome variable, allowing us to derive the marginal effects of minimum wage changes at different quantiles in wage or income distributions. In this inversion, we use a kernel density estimate at each quantile, applying a local linear approximation to the counterfactual CDF. By performing this process for various threshold value c, we can map out the impacts across different quantiles in the wage or income distribution.

To identify equation (3), we need to ensure that the minimum wage is determined exogenously, conditional on covariates. We rely on the same exogenous increases in the minimum wage after the policy change in 2007 but apply the instrumental variable approach proposed by Kawaguchi and Mori (2021) this time. Specifically, we estimate the following first-stage model, in which the regional minimum wage is instrumented using the policy change:

$$\ln\left(MW_{pt-1}\right) = \sum_{\tau=2001}^{2022} \beta_\tau \cdot \min\left(\ln\left(\frac{WB_{p2006}}{MW_{p2006}}\right),\ 0\right) + X_{it}\Gamma + t \cdot \omega_p + \delta_t + \eta_t + \nu_{it} \tag{4}$$

where $\ln\left(\frac{WB_{p2006}}{MW_{p2006}}\right)$ is the log point difference between the welfare benefit and minimum wage earning in region p in 2006. In so doing, we aim to isolate exogenous variations in minimum wage changes following the 2007 amendment to the Minimum Wage Act.

Appendix Table A.6 reports the estimated coefficients for β_τ in equation (4) from the first-stage of model of UQPE. Estimation results are consistent with the previous work (Kawaguchi and Mori, 2021, Figure 7), in that the gap between the welfare benefit and the 2006 minimum wage had substantial power to predict the minimum wage increase after 2007.[26] First-stage F statistic is sufficiently high, supporting the relevance of the instruments. Importantly, the coefficient estimates for β_τ prior to the policy change are statistically significant but economically insignificant, supporting the validity of the exclusion restriction.

5.2 Effects on Wage and Income Inequalities

Figure 6 shows the results for the effects of the minimum wage on women's wage distribution by each percentile based on the UQPE approach. The estimated elasticities are positive and statistically significant up to the 30th percentile, with a peak around the 19th percentile, corresponding to approximately 880 yen. This implies that a minimum wage hike increases the lower tail of

[26] Particularly, Kawaguchi and Mori (2021) adopted the same identification strategy in this study and found no economically significant pre-trends in the employment rate, labor force participation rate, and hours of work of the less educated.

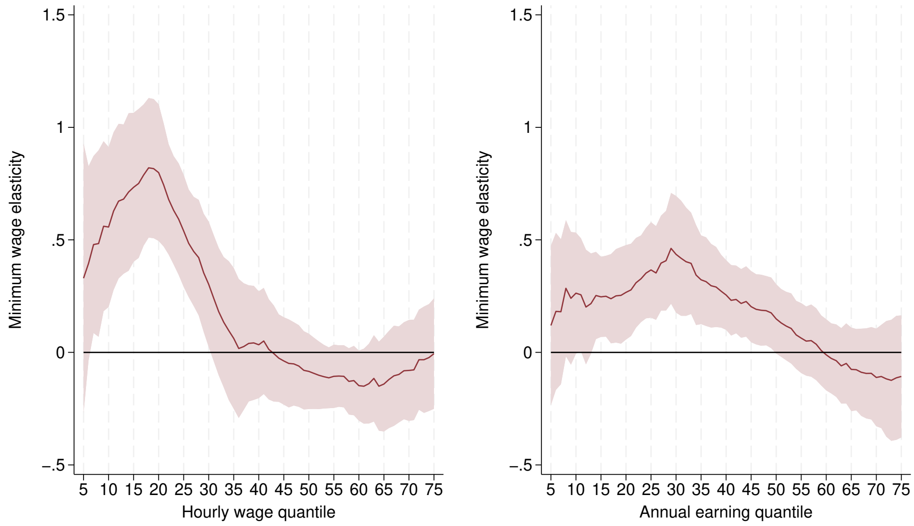

Figure 6: Minimum Wage Elasticities for Unconditional Wage and Earning Quantiles (Women)

Note: The reported estimates represent minimum wage elasticities by unconditional quantiles of hourly wage rates (left) and annual earnings (right) for female workers. Both hourly wages and earnings are adjusted to 2020 prices. The estimates are derived from a series of IV estimation models, regressing an indicator for having an hourly wage (or annual earning) below quantile-based cutoffs (ranging from the 5th to 50th percentiles) on the log minimum wage and covariates. The sample size is 12,273,105. Unconditional quantile partial effects (UQPE) are calculated by dividing the coefficient on the log minimum wage by the negative of the hourly wage (or annual earning) density at each quantile. These UQPE estimates are then normalized by dividing them by the quantile-specific hourly wage (or annual earning) cutoff to obtain elasticities. All estimations include controls for individual age (and its polynomials), year dummies, and region-specific linear time trends. The shaded areas indicate 95 percent confidence intervals. Standard errors are clustered at the region level, and the estimations are weighted by sampling weights.

wages, as shown in previous studies. The results are particularly consistent with evidence from Japan (Kambayashi et al., 2013), which confirmed a significant wage compression effect among women using BSWS data from 1994 to 2003. Figure 6 also presents the results for women's income distribution. The estimated elasticities are positive but not statistically significant below the 12th percentile. They become statistically significant from the 12th to the 52nd percentile, with the slope of increasing effects becoming steeper around the 29th percentile.

It is important to emphasize that individuals in the same quantiles in the wage and income distributions do not necessarily correspond to the same group of workers. Appendix Figure A.9 presents a heatmap illustrating this point by comparing the proportions of observations in specific wage and income percentiles. The heatmap indicates that individuals with the same hourly wage can have different income levels, particularly in the lower tail of the wage quantiles. This suggests substantial variability in working hours among low-wage workers. Some female workers with low hourly wages likely reduced their working hours in response to the minimum wage hike. In the next subsection, we exploit this pattern to simulate work-hour adjustments across different annual earning quantiles.

Appendix Figure A.7 presents the corresponding estimates from a robustness test, in which we include additional regional control variables in equation (3)—specifically, six variables measuring the population share by gender and three age groups. Although the impact becomes somewhat larger at the lower tail of both hourly wages and annual earnings, the main implication remains: the gains in annual income are much smaller—and even insignificant—at the 10th percentile or lower, despite the observed wage compression. Finally, Appendix Figure A.8 presents the UQPE results for men. Unlike women, men exhibit a much weaker wage compression effect, with the quantile elasticity estimated at approximately 0.75 at the 5th percentile but declining sharply by the 10th percentile. Similar to the case of women, the increase in hourly wages does not clearly translate into higher annual income.

5.3 Mechanism

Work-hour reductions among women can be driven by several factors. In general, an increase in wages can reduce working hours through both labor supply and labor demand channels. On the demand side, one potential mechanism is employer-side substitution from work hours per employee to employment headcount in the presence of quasi-fixed labor costs. On the supply side, workers may reduce their hours if the income effect outweighs the substitution effect. Institutional cutoffs for tax deductions and social contributions can also create incentives to reduce work hours on both the demand and supply sides. As discussed in Section 2.4, Japan's tax and social security systems have featured a complex set of benefit cliffs and convexities in effective marginal tax rate. It is possible that women's long-term decisions to invest in human capital, combined with persistent misperceptions, have contributed to the observed bunching around ¥1 million. Complexity of tax and social benefit system may also raise compliance costs, creating a wedge between optimal and actual labor supply decisions.

To explore the underlying mechanisms—particularly to assess whether work-hour reductions are driven by supply-side or demand-side factors—we examine the impact of the minimum wage across annual income levels, where institutional thresholds are defined. Our analysis builds on

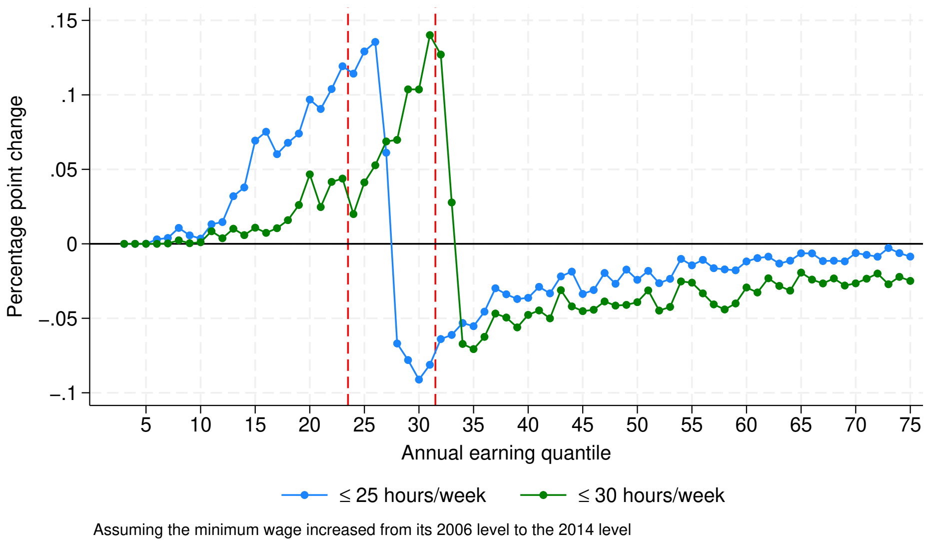

Figure 7: Predicted Change in Proportions of Short-hour Jobs by Income Quantiles among Women

Note: The figure reports the predicted proportions of female employees (aged 23 to 59) reducing their work hours following the minimum wage hike, across income quantiles. The calculation is based on the elasticity estimates from the Unconditional Quantile Partial Effects (UQPE) presented in Figure 6. For each of the 100×100 cells defined by wage and income quantiles, we assign the corresponding elasticity estimates from the UQPE and predict the impact of the minimum wage increase on each individual's hourly wage rate and annual income. We assume that the minimum wage increases to the levels observed between 2006 and 2014. For each individual within the cells, we then reverse-calculate the implied change in working hours. A dummy variable is created to indicate individuals predicted to work less than or equal to 25 or 30 hours per week. We then compute the predicted change in the proportion of short-hour jobs by using sampling weights. The red dashed lines correspond to annual income quantiles at 1.03 and 1.3 million JPY, reflecting thresholds for tax deductions and mandatory enrollment in additional social security programs.

the minimum wage elasticities estimated from the UQPE approach and the fact that individual female employees are classified into one of the 100×100 wage–income quantile cells in the UQPE estimations. Using this framework, we simulate a counterfactual scenario in which the minimum wage increases to the levels observed between 2006 and 2014. Specifically, for each female worker observed in 2006, we assign the corresponding elasticity estimates for hourly wage and income quantile values (i.e., estimates in Figure 6), and predict changes in hourly wage and annual income after the hypothetical change in the minimum wage.[27] Based on these predicted values, we then reverse-calculate the change in work hours for each individual. To shed light on behavioral responses, we create a dummy variable indicating whether an individual's weekly hours are predicted to fall below a specific threshold, such as 25 hours per week. This allows us to assess whether the observed adjustments are concentrated near institutional income thresholds.

[27]In assigning individuals to the 100×100 wage–income quantile cells, cases falling below the 5th percentile or above the 95th percentile in either wage or income were treated as missing.

Figure 7 presents the predicted proportion of female employees aged 23 to 59 reducing their work hours following the minimum wage hike, across income quantiles. The red dashed lines correspond to annual income quantiles at ¥1.03 and 1.3 million, reflecting thresholds for tax deductions and mandatory enrollment in additional social security programs. The figure shows that transitions into jobs with fewer than 25 weekly hours are concentrated near the lower threshold, while transitions into jobs with fewer than 30 hours per week tend to occur below the higher threshold. The sharp pattern around ￥1 million is unlikely to be explained by labor demand factors, as no additional employer-side social insurance contribution required at that income level during the time of the analyses. A simple labor demand model with quasi-fixed costs offers no clear rationale for this change in hours. These findings are therefore more consistent with worker-initiated adjustments, pointing to a supply-side response.

Finally, we also simulated potential earnings gains under the assumption that women did not reduce their work hours but continued working the same number of hours as observed in 2006. Women at income quantiles corresponding to ¥1.03 million are predicted to earn 3.7% more in annual income. Taken together, our results suggest that the minimum wage, combined with institutional cutoffs, has had limited effects on increasing women's earnings in Japan, most likely due to their behavioral responses to those thresholds.

6 Conclusion

This study examined the unintended effects of minimum wage increases on women's working hours and earnings in Japan, using large-scale administrative payroll data. Relying on the quasi-experimental variation generated by the 2007 policy reform, we applied a frequency-distribution approach to estimate employment changes across wage bins. The analysis revealed a clear shift from long-hour to short-hour jobs—particularly among incumbent female workers—which was concentrated in wage bins near the new minimum wage. We further estimated Unconditional Quantile Partial Effects (UQPE) to evaluate distributional impacts, finding that while hourly wages increased up to the 30th percentile, gains in annual income were modest or insignificant at the bottom of the income distribution. Simulations based on the UQPE estimates suggest that these limited income gains are driven in part by reductions in work hours among women with earnings near institutional income cutoffs.

The findings highlight the importance of recognizing working hours as a key margin of adjustment in minimum wage evaluations, particularly when institutional thresholds are present. Our results contrast with existing evidence from the U.S., where Godøy et al. (2025) found that minimum wage increases raised employment rate among single mothers with small children who are subject to means-tested public assistance programs. In contrast, our findings suggest that in institutional contexts like Japan, where complex tax and social benefit thresholds persist, minimum wage hikes may lead to reductions in work hours among affected workers. This points to the importance of considering institutional interactions when assessing the broader effects of minimum wage increases.

Finally, studying the effect of minimum wage increases on women's income or wage inequality alone is not sufficient to understand the full picture of their welfare consequences. First, as shown in Kawaguchi and Mori (2009), low-wage women are not necessarily from low-income households in Japan. Our companion paper Mori and Okudaira (2025) investigates this point using household-level microdata. Second, some women may be better off working fewer hours if doing so allows them to spend more time with their children. However, our results show that work-hour reductions are more pronounced among women aged 35 to 54 than among those aged 25 to 34, suggesting that the shift contributed less to increased time spent on childcare among mothers with small children. While shorter working hours may still improve well-being for mid-aged women whose children have grown up, evaluating such welfare gains is not straightforward. Further research is needed to assess the broader welfare effects beyond income, including how tax and social security systems have shaped women's long-term life course decisions.

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Appendix A. Unconditional Quantile Partial Effect

To estimate the impact of minimum wage policies on different quantiles of the income distribution, we use the Recentered Influence Function (RIF) regression as developed by Firpo et al. (2009).

$$RIF\left(y_{it},\ Q_t\right) = \left[Q_\tau + \ \frac{\tau}{f\left(Q_\tau\right)}\right] - \frac{I\left(y_{it} < Q_\tau\right)}{f\left(Q_\tau\right)} = k_\tau - \frac{I\left(y_{it} < Q_\tau\right)}{f\left(Q_\tau\right)} \tag{5}$$

where Q_τ is the annual income (hourly wage rate) of τth percentile, $I\left(y_{it} < Q_\tau\right)$ is an indicator function that equals 1 if the annual income (hourly wage rate) y_{it} of individual i at time t is below Q_τ, and 0 otherwise, and $f\left(Q_\tau\right)$ is the probability density function at Q_τ.

To determine the effect of the minimum wage on the RIF, we consider the following estimation equation:

$$RIF\left(y_{it},\ Q_t\right) = \gamma_\tau \ln\left(MW_{pt}\right) + X_{it}\Lambda_\tau + t \cdot \sigma_{\tau p} + \delta_{\tau t} + \nu_{\tau it} \tag{6}$$

The coefficient γ_τ obtained from this regression indicates the effect of the minimum wage on the specific quantile Q of the annual income (hourly wage rate) distribution.

Since the first term of $RIF\left(y_{it},\ Q_t\right), k_\tau$, is a constant and $I\left(y_{it} < Q_\tau\right)$ represents the CDF for any given threshold quantiles, this specification is directly related to the specifications shown in the main text as follows:

$$I\left(y_{it} < c\right) = \alpha_c \ln\left(MW_{pt}\right) + X_{it}\Gamma_c + t\ \cdot \mu_{cp} + \theta_{ct} + \epsilon_{cit} \tag{7}$$

The relationship between γ_τ and α_c is given by: $\gamma_\tau = -\frac{\alpha_{c(\tau)}}{f(c(\tau))}$.

Appendix B. Complementary Figures and Tables

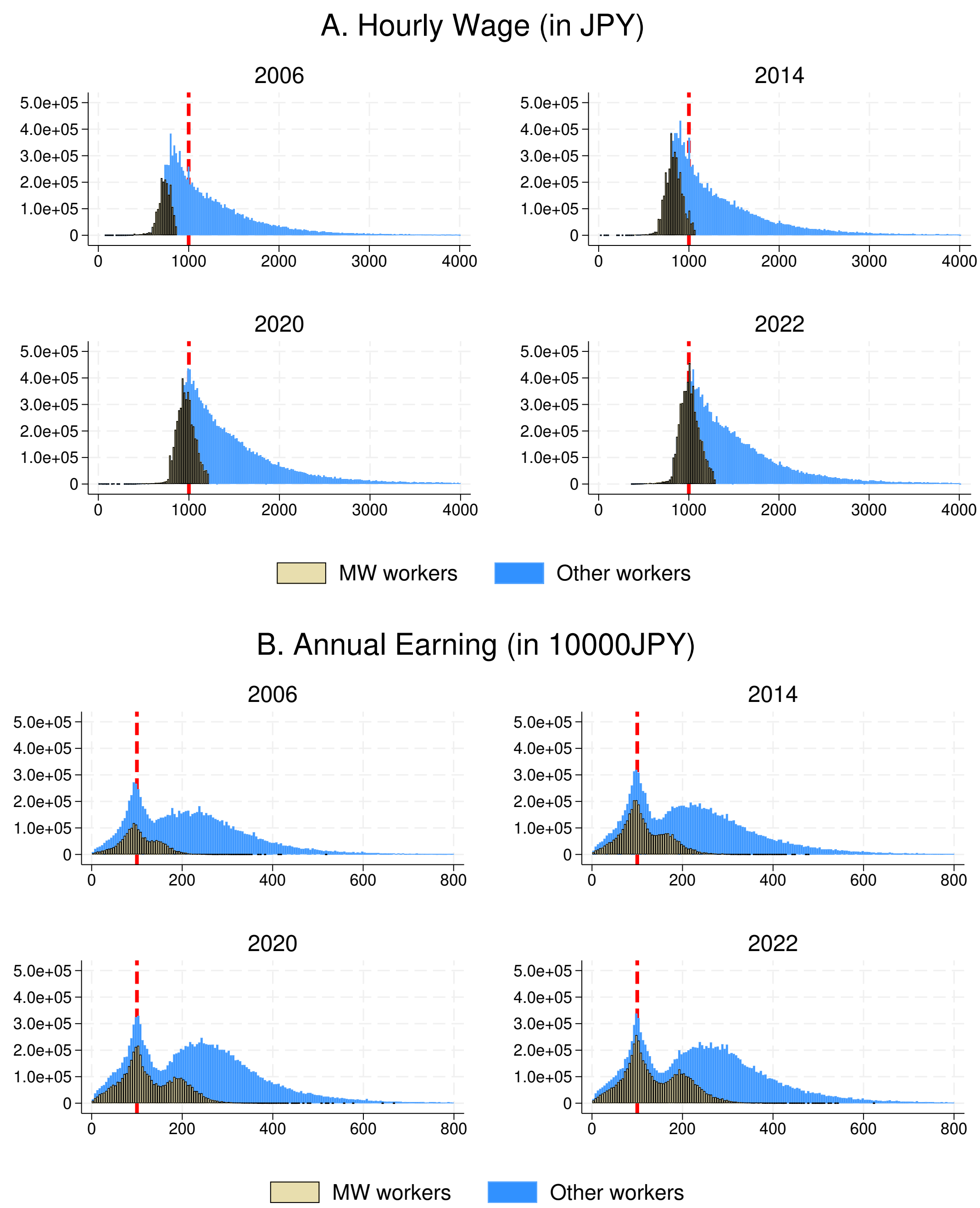

Figure A.1: Frequency Distribution (Women Aged 23 or Older)

Note: The figures display frequency distributions for the hourly wage rate (Panel A) and annual earnings (Panel B), based on employee wage records from the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). The hourly wage rate is calculated by dividing the monthly cash earnings in BSWS by actual monthly hours worked, while annual earnings are derived by multiplying monthly cash earnings by an annual conversion weight constructed from the Monthly Labour Survey (Japanese Ministry of Health, Labour and Welfare). Minimum wage (MW) workers are defined as employees whose hourly wage rate is equal to or less than 1.2 times the new minimum wage effective in the following fall. Minimum wages in Japan are set by region. Red dashed lines represent 1,000 JPY in Panel A and 1 million JPY in Panel B. Frequency counts are weighted using sampling weights.

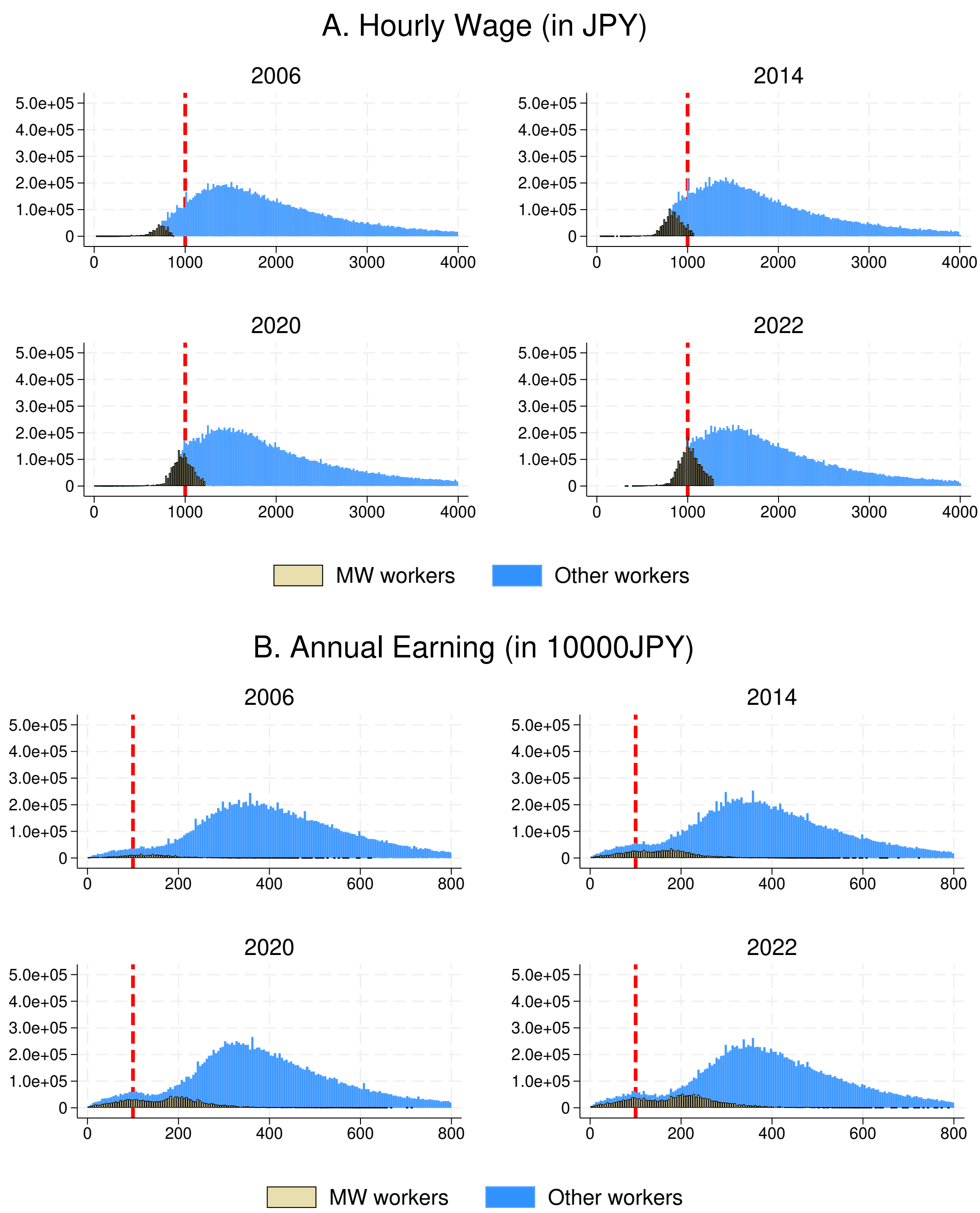

Figure A.2: Frequency Distribution (Men Aged 23 or Older)

Note: The figures display frequency distributions for the hourly wage rate (Panel A) and annual earnings (Panel B), based on employee wage records from the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). The hourly wage rate is calculated by dividing the monthly cash earnings in BSWS by actual monthly hours worked, while annual earnings are derived by multiplying monthly cash earnings by an annual conversion weight constructed from the Monthly Labour Survey (Japanese Ministry of Health, Labour and Welfare). Minimum wage (MW) workers are defined as employees whose hourly wage rate is equal to or less than 1.2 times the new minimum wage effective in the following fall. Minimum wages in Japan are set by region. Red dashed lines represent 1,000 JPY in Panel A and 1 million JPY in Panel B. Frequency counts are weighted using sampling weights.

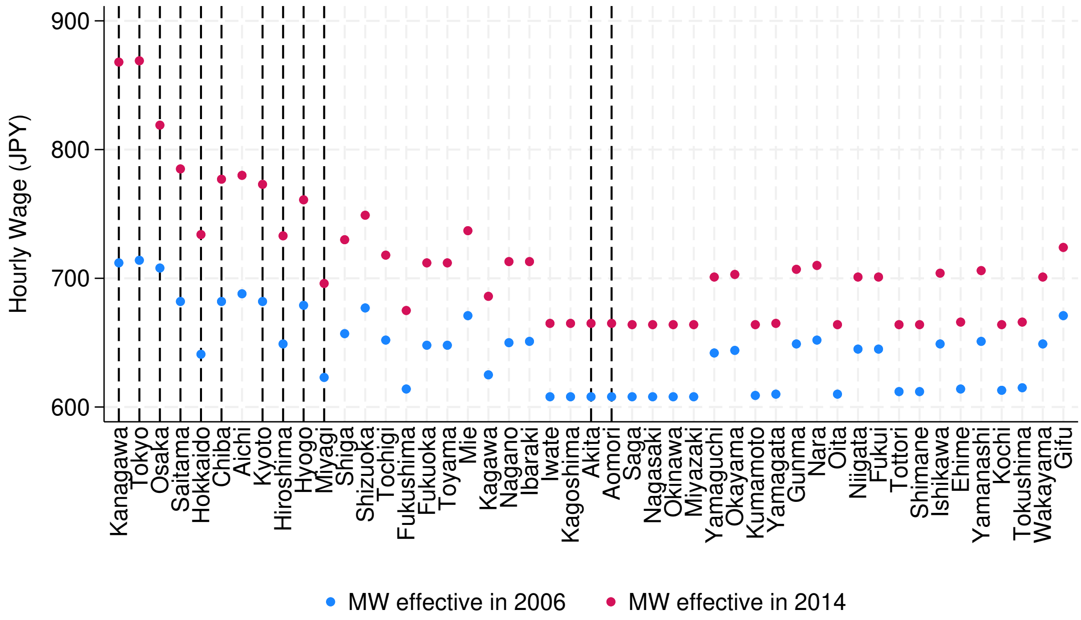

Figure A.3: Changes in Minimum Wage

Note: The plots are sorted by the rate of effective minimum wage increases between 2006 and 2014. Dashed lines represent prefectures affected by the policy in 2007 due to positive gaps between welfare benefits and minimum-wage earnings. "MW effective in 2006 (2014)" refers to the minimum wage revised in the fall of 2005 (2013). Note that the individual payroll data used in this study reflects information as of June each year.

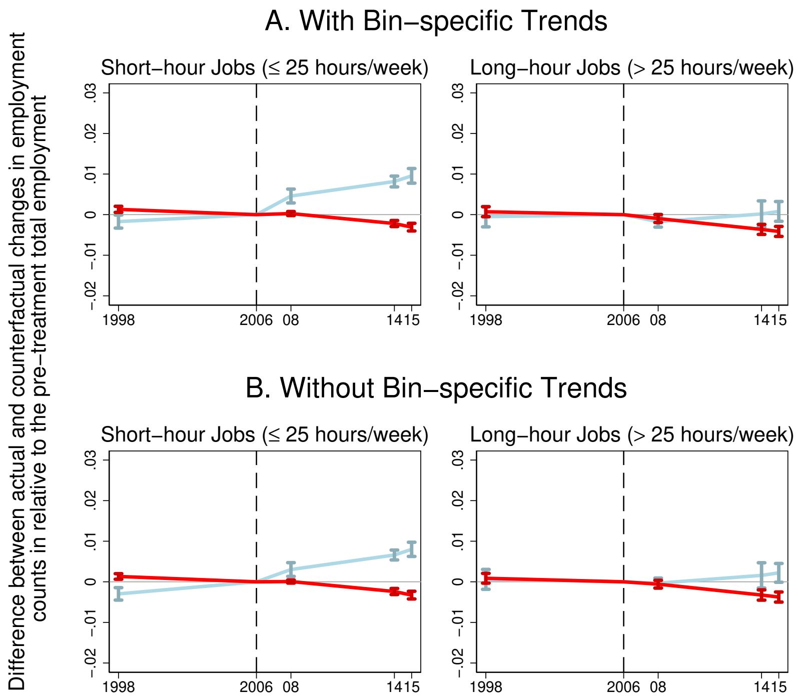

Figure A.4: Impact of Minimum Wage on the Missing (Red) and Excess (Blue) Jobs Over Time: Sensitivity against Excluding Bin-specific Trends

Note: The figures plot the estimated minimum wage effects on employment counts from an event-study analysis. We exploit the 10 events of the minimum wage increases after the 2007 revision to the Minimum Wage Act in Japan. We stack all event-specific region-bin level datasets to estimate an average effect across 10 events using a single set of treatment effects. The dataset is drawn from employee observations in 1998, 2006, 2008, 2014, and 2015 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). All the estimations control for event-specific year effects, and dummies to indicate small minimum wage increases after 2007. All the estimations are weighted by the size of regional population. Standard errors are cluster at region and event level.

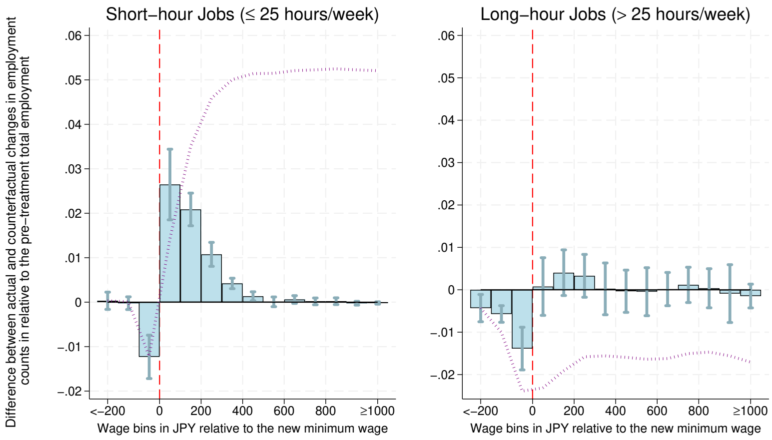

Figure A.5: Impact of Minimum Wage on the Wage Distribution by Size of the Job - Sensitivity Against Using Alternative Wage Premia

Note: The figures plot the estimated minimum wage effects on employment counts from an event-study analysis. Wage bins are adjusted using estimated local wage premia obtained from a Mincer-type wage regression. We exploit the 10 events of the minimum wage increases after the 2007 revision to the Minimum Wage Act in Japan. We stack all event-specific region-bin level datasets to estimate an average effect across 10 events using a single set of treatment effects. The dataset is drawn from employee observations in 1998, 2006, 2008, 2014, and 2015 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). The ploted estimates are the averages of the treatment effects in 2014 and 2015. All the estimations control for year dummies, linear trends specific to events and wage bins, and dummies to indicate small minimum wage increases in the control regions after 2007. All the estimations are weighted by the size of regional population. Standard errors are cluster at region and event level. Purple dashed line indicates a running sum of the estimated impacts at each bin.

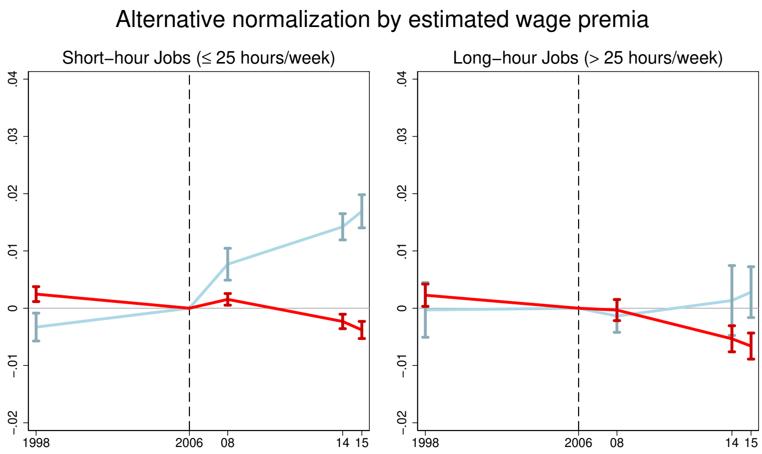

Figure A.6: Impact of Minimum Wage on the Missing (Red) and Excess (Blue) Jobs Over Time: Wage Bins Adjusted by Alternative Wage Premia

Note: The figures plot the estimated minimum wage effects on employment counts from an event-study analysis. Wage bins are adjusted using estimated local wage premia obtained from a Mincer-type wage regression. We exploit the 10 events of the minimum wage increases after the 2007 revision to the Minimum Wage Act in Japan. We stack all event-specific region-bin level datasets to estimate an average effect across 10 events using a single set of treatment effects. The dataset is drawn from employee observations in 1998, 2006, 2008, 2014, and 2015 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). All the estimations control for event-specific year effects, dummies to indicate small minimum wage increases after 2007. All the estimations are weighted by the size of regional population. Standard errors are cluster at region and event level.

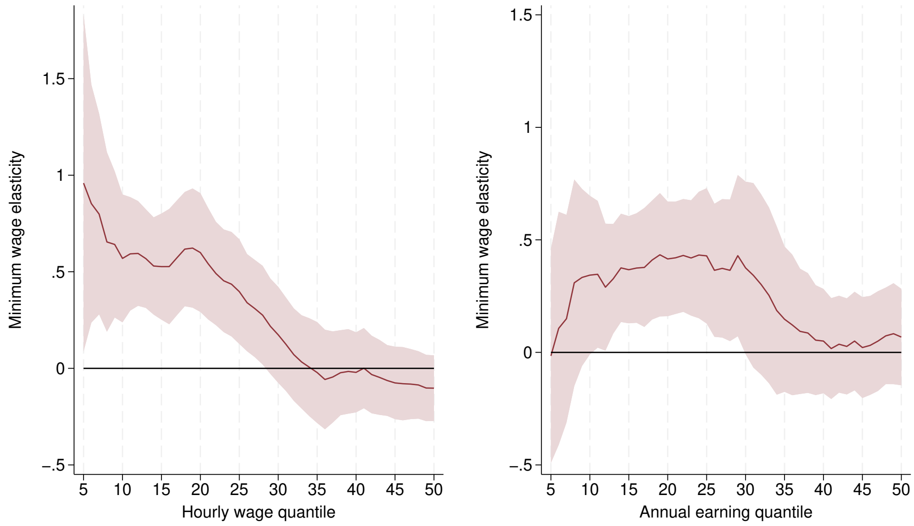

Figure A.7: Minimum Wage Elasticities for Unconditional Wage and Earning Quantiles (Women) with Additional Regional Control Variables

Note: The reported estimates represent minimum wage elasticities by unconditional quantiles of hourly wage rates (left) and annual earnings (right) for female workers. These estimations include additional regional controls, specifically the population shares by gender and three age categories (24 or younger, 25 to 64, 65 or older) within each prefecture. Both hourly wages and earnings are adjusted to 2020 prices. The estimates are derived from a series of IV estimation models, regressing an indicator for having an hourly wage (or annual earning) below quantile-based cutoffs (ranging from the 5th to 50th percentiles) on the log minimum wage and covariates. The sample size is 12,273,105. Unconditional quantile partial effects (UQPE) are calculated by dividing the coefficient on the log minimum wage by the negative of the hourly wage (or annual earning) density at each quantile. These UQPE estimates are then normalized by dividing them by the quantile-specific hourly wage (or annual earning) cutoff to obtain elasticities. All estimations include controls for individual age (and its polynomials), year dummies, and prefecture-specific linear time trends. The shaded areas indicate 95 percent confidence intervals. Standard errors are clustered at the prefecture level, and the estimations are weighted by sampling weights.

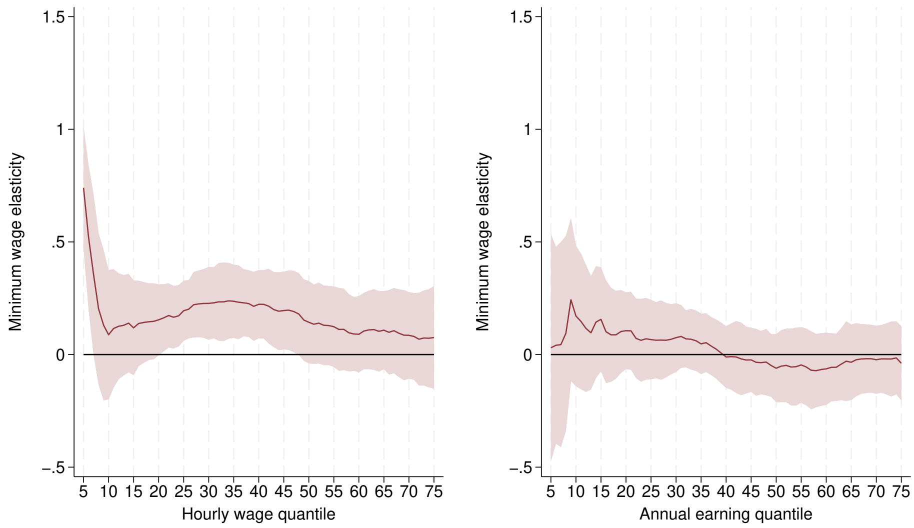

Figure A.8: Minimum Wage Elasticities for Unconditional Wage and Earning Quantiles (Men)

Note: The reported estimates represent minimum wage elasticities by unconditional quantiles of hourly wage rates (left) and annual earnings (right) for female workers. Both hourly wages and earnings are adjusted to 2020 prices. The estimates are derived from a series of IV estimation models, regressing an indicator for having an hourly wage (or annual earning) below quantile-based cutoffs (ranging from the 5th to 50th percentiles) on the log minimum wage and covariates. The sample size is 12,273,105. Unconditional quantile partial effects (UQPE) are calculated by dividing the coefficient on the log minimum wage by the negative of the hourly wage (or annual earning) density at each quantile. These UQPE estimates are then normalized by dividing them by the quantile-specific hourly wage (or annual earning) cutoff to obtain elasticities. All estimations include controls for individual age (and its polynomials), year dummies, and region-specific linear time trends. The shaded areas indicate 95 percent confidence intervals. Standard errors are clustered at the region level, and the estimations are weighted by sampling weights.

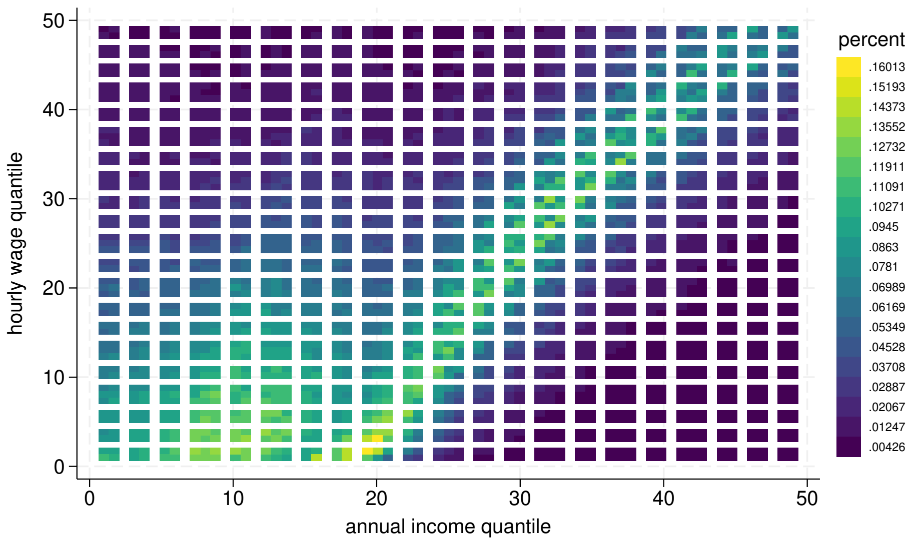

Figure A.9: Heatmap by Wage and Income Quantiles (Women)

Note: The figure reports proportions of observations in each cell defined by hourly wage and annual income quantiles among female employees. The dataset is drawn from employee observations from 2000 to 2022 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare).

Table A.1: Proprotion of Employees by Deciles of Wage and Income

	Deciles of real hourly wage										
	1	2	3	4	5	6	7	8	9	10	Total
Male	2.35	2.94	3.88	4.94	5.79	6.39	6.99	7.61	8.26	8.76	57.91
Female	7.65	7.06	6.12	5.06	4.21	3.61	3.01	2.39	1.74	1.24	42.09
Total	10.00	10.00	10.00	10.00	10.00	10.00	10.00	10.00	10.00	10.00	100.00

	Deciles of real annual income										
	1	2	3	4	5	6	7	8	9	10	Total
Male	2.69	2.13	3.07	4.27	5.54	6.64	7.49	8.15	8.74	9.20	57.91
Female	7.31	7.87	6.93	5.73	4.46	3.36	2.51	1.85	1.26	0.80	42.09
Total	10.00	10.00	10.00	10.00	10.00	10.00	10.00	10.00	10.00	10.00	100.00

Note: The table reports proportions of observations by gender in each group defined by deciles of hourly wage rate or annual income. The dataset is drawn from employee observations from 1998 to 2022 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). Table adjusts for sampling weights. Sample size is 12,273,105 for women, and 18,373,948 for men.

Table A.2: Summary Statistics by Methods

	mean	sd	p25	p50	p75
A. Frequency-distribution (N = 6,184,631)					
Short-hour jobs ($\leq$ 25h/week)	0.15	0.36	0.00	0.00	0.00
Female	0.42	0.49	0.00	0.00	1.00
Age	45.60	13.07	35.00	45.00	55.00
New entrant (tenure $<$ 1 year)	0.12	0.32	0.00	0.00	0.00
No bonus in the previous year	0.28	0.45	0.00	0.00	1.00
Hourly wage (JPY)	1749.38	1235.60	1035.46	1442.77	2077.41
Annual earning (10000JPY)	337.93	211.93	193.33	308.02	445.84
Annual earning with bonus (10000JPY)	412.56	290.79	205.21	361.15	550.38
Industry					
Retail	0.12	0.32	0.00	0.00	0.00
Restaurant	0.04	0.20	0.00	0.00	0.00
Accommodation/hotel	0.01	0.11	0.00	0.00	0.00
Manufacturing	0.24	0.43	0.00	0.00	0.00
B. UQPE (N = 30,647,053)					
Short-hour jobs ($\leq$ 25h/week)	0.16	0.36	0.00	0.00	0.00
Female	0.42	0.49	0.00	0.00	1.00
Age	45.98	13.07	35.00	45.00	55.00
New entrant (tenure $<$ 1 year)	0.11	0.31	0.00	0.00	0.00
No bonus in the previous year	0.29	0.45	0.00	0.00	1.00
Hourly wage (JPY)	1772.55	1372.59	1063.51	1462.81	2086.64
Annual earning (10000JPY)	338.80	212.06	196.42	310.01	444.57
Annual earning with bonus (10000JPY)	411.07	287.91	208.14	361.91	545.83
Industry					
Retail	0.11	0.32	0.00	0.00	0.00
Restaurant	0.04	0.20	0.00	0.00	0.00
Accommodation/hotel	0.01	0.11	0.00	0.00	0.00
Manufacturing	0.24	0.42	0.00	0.00	0.00

Note: The dataset comprises employee payroll data from 1998, 2006, 2008, 2014, and 2015 in Panel A, and from 1998 to 2022 in Panel B, drawn from the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). In frequency-distribution approach in the main text, we collapsed this dataset by event-year-prefecture-wage bin. Sample statistics are calculated by using sampling weights. Hourly wages and annual earnings are deflated using the Consumer Price Index (CPI) with 2020 as the base year. "Annual earning with bonus" indicates a sum of annual earnings and a bonus from the previous year.

Table A.3: Sensitivity Against Controlling for Additional Regional Attributes

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
				Short-hour jobs		Long-hour jobs	
	All	Female	Male	Female	Male	Female	Male
Missing jobs below new MW (Δ b)	-0.024***	-0.023***	-0.001	-0.009***	-0.002***	-0.014***	0.001
	(0.003)	(0.002)	(0.001)	(0.001)	(0.000)	(0.002)	(0.001)
Excess jobs above new MW (Δ a)	0.040***	0.023***	0.017***	0.026***	0.009***	-0.002	0.007***
	(0.006)	(0.004)	(0.002)	(0.002)	(0.001)	(0.003)	(0.002)
%Δ affected employment	0.210**	0.002	0.488***	0.069***	0.038***	-0.182***	0.453***
	(0.084)	(0.029)	(0.098)	(0.009)	(0.005)	(0.027)	(0.132)
%Δ affected wage	0.251***	0.080***	0.230***	-	-	-	-
	(0.033)	(0.019)	(0.027)	-	-	-	-
Employment elasticity w.r.t. MW	0.122**	0.002	0.121***	0.132***	0.056***	-0.129***	0.065***
	(0.049)	(0.031)	(0.024)	(0.017)	(0.008)	(0.019)	(0.019)
Employment elasticity w.r.t. affected wage	0.837***	0.021	2.122***	-	-	-	-
	(0.240)	(0.364)	(0.258)	-	-	-	-
Proportion affected in 2006	.075	.137	.032	.248	.191	.091	.019
N of observations	165600	165600	165600	165600	165600	165600	165600

Note: The table reports the estimated impacts of the Minimum Wage increases from event-study estimates based on the 2007 revision to the Minimum Wage Act in Japan. We stack all event-specific region-bin level datasets to estimate an average effect across 10 events using a single set of treatment effects. The dataset is drawn from employee observations in 1998, 2006, 2008, 2014, and 2015 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). "Short-hour jobs" are those employment with less than or equal to 25 hours of work per week."Long-hour jobs" are those employment with more than 25 hours of work per week. All the estimations control for year dummies, linear trends specific to events and wage bins, and dummies to indicate small minimum wage increases in the control regionals after 2007, in addition to six variables representing the proportions of the population by gender and three age groups (20-24, 25-59, 60 or older). All the estimations are weighted by the size of regional population. Standard errors are clustered at region and event level.

Table A.4: Defining Short/Long-Hour Jobs by Alternative Threshold Values

	(1)	(2)	(3)	(4)
	Short-hour jobs		Long-hour jobs	
	$\leq \theta$ hours/ week		$> \theta$ hours/ week	
	Female	Male	Female	Male
A. Threshold value (θ) = 20				
Missing jobs below new MW (Δ b)	-0.005***	-0.002***	-0.018***	-0.002*
	(0.001)	(0.000)	(0.002)	(0.001)
Excess jobs above new MW (Δ a)	0.018***	0.007***	0.006	0.007*
	(0.001)	(0.001)	(0.003)	(0.003)
%Δ affected employment	0.054***	0.027***	-0.108***	0.228
	(0.006)	(0.003)	(0.026)	(0.159)
Employment elasticity w.r.t. MW	0.110***	0.045***	-0.100***	0.040
	(0.012)	(0.006)	(0.024)	(0.028)
Proportion affected in 2006	.248	.201	.112	.021
B. Threshold value (θ) = 30				
Missing jobs below new MW (Δ b)	-0.011***	-0.002***	-0.011***	-0.001
	(0.001)	(0.001)	(0.001)	(0.001)
Excess jobs above new MW (Δ a)	0.030***	0.011***	-0.006**	0.003
	(0.002)	(0.001)	(0.002)	(0.003)
%Δ affected employment	0.074***	0.048***	-0.246***	0.123
	(0.009)	(0.006)	(0.030)	(0.188)
Employment elasticity w.r.t. MW	0.153***	0.069***	-0.143***	0.016
	(0.019)	(0.009)	(0.017)	(0.025)
Proportion affected in 2006	.249	.174	.07	.016

Note: The table reports the estimated impacts of the Minimum Wage increases from event-study estimates based on the 2007 revision to the Minimum Wage Act in Japan. We stack all event-specific region-bin level datasets to estimate an average effect across 10 events using a single set of treatment effects. The dataset is drawn from employee observations in 1998, 2006, 2008, 2014, and 2015 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). All the estimations control for year dummies, linear trends specific to events and wage bins, and dummies to indicate small minimum wage increases in the control regions after 2007. All the estimations are weighted by the size of regional population. Standard errors are clustered at region and event level.

Table A.5: Restricting Observations to Individuals Aged Between 23 and 59

	(1)	(2)	(3)	(4)
	Short-hour jobs		Long-hour jobs	
	Female	Male	Female	Male
Missing jobs below new MW (Δ b)	-0.005***	-0.000	-0.011***	0.001
	(0.001)	(0.000)	(0.001)	(0.001)
Excess jobs above new MW (Δ a)	0.014***	0.003***	-0.002	0.005**
	(0.001)	(0.000)	(0.002)	(0.002)
%Δ affected employment	0.042***	0.028***	-0.160***	0.563**
	(0.007)	(0.003)	(0.020)	(0.196)
Employment elasticity w.r.t. MW	0.063***	0.019***	-0.097***	0.047**
	(0.011)	(0.002)	(0.012)	(0.016)
Proportion affected in 2006	.202	.092	.081	.011
N of observations	165600	165600	165600	165600

Note: The table reports the estimated impacts of the Minimum Wage increases from event-study estimates based on the 2007 revision to the Minimum Wage Act in Japan. We stack all event-specific region-bin level datasets to estimate an average effect across 10 events using a single set of treatment effects. The dataset is drawn from employee observations in 1998, 2006, 2008, 2014, and 2015 in the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). "Short-hour jobs" are those employment with less than or equal to 25 hours of work per week."Long-hour jobs" are those employment with more than 25 hours of work per week.All the estimations control for year dummies, linear trends specific to events and wage bins, and dummies to indicate small minimum wage increases in the control regions after 2007. All the estimations are weighted by the size of regional population. Standard errors are clustered at region and event level.

Table A.6: First Stage Estimation Results (UQPE, Female Sample)

$\beta_{\tau=1999}$	0.009	$\beta_{\tau=2012}$	0.650
	(0.003)		(0.052)
$\beta_{\tau=2000}$	0.006	$\beta_{\tau=2013}$	0.698
	(0.003)		(0.045)
$\beta_{\tau=2001}$	-0.005	$\beta_{\tau=2014}$	0.723
	(0.003)		(0.051)
$\beta_{\tau=2002}$	-0.004	$\beta_{\tau=2015}$	0.726
	(0.004)		(0.053)
$\beta_{\tau=2003}$	-0.008	$\beta_{\tau=2016}$	0.707
	(0.005)		(0.053)
$\beta_{\tau=2004}$	-0.008	$\beta_{\tau=2017}$	0.685
	(0.007)		(0.051)
$\beta_{\tau=2005}$	-0.007	$\beta_{\tau=2018}$	0.662
	(0.005)		(0.047)
$\beta_{\tau=2006}$	-0.003	$\beta_{\tau=2019}$	0.631
	(0.003)		(0.044)
$\beta_{\tau=2008}$	0.065	$\beta_{\tau=2020}$	0.596
	(0.015)		(0.04)
$\beta_{\tau=2009}$	0.201	$\beta_{\tau=2021}$	0.581
	(0.027)		(0.038)
$\beta_{\tau=2010}$	0.409	$\beta_{\tau=2022}$	0.540
	(0.033)		(0.034)
$\beta_{\tau=2011}$	0.529		
	(0.051)		
F statistic		799.19	
N		12,273,105	

Note: The table reports the estimated coefficients for β_τ in equation (1) from the first stage of a single IV regression model of UQPE for women. The baseline year is 2007. The "F statistic" refers to the Sanderson-Windmeijer multivariate F-test for excluded instruments. All estimations include controls for individual age (and its polynomials), year dummies, and region-specific linear time trends. Standard errors are shown in parentheses and are clustered at the region level. The dataset comprises female employee observations from 1998 to 2022, drawn from the Basic Survey on Wage Structure (Japanese Ministry of Health, Labour and Welfare). Estimations are weighted using sampling weights.