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Abstract

The median wage in Japan has fallen nominally since 1999 due to a severe recession, while the statutory minimum wage has steadily increased over the same period. We used large micro-data sets from two government surveys to investigate how the minimum wage has affected wage distribution under the unusual circumstances of deflation. The compression of the lower tail of female wage distribution was largely explained by the increased real value of the minimum wage. The steady increases in the effective minimum wage reduced employment among low-skilled middle-aged female workers, but the mechanical effect associated with disemployment on wage compression was minimal. These results held even after controlling for composition effects. The minimum wage contributed to the reduction in the pay gap between full-time and part-time workers.

Key Words: Minimum Wage, Wage Distribution, Wage Inequality, Employment, Deflation

JEL Classification Code: J23 (Labor Demand), J31 (Wage Level and Structure; Wage Differentials), J38 (Wage Related Public Policy)

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1 Introduction

Wage distribution has evolved differently among advanced industrialized countries during the 1990s and early 2000s as these countries have shared the similar experience of rapid technological progress and increased exposure to international trade and outsourcing (Machin and Van Reenen, 1998; Koeniger, Leonardi, and Nunziata, 2007). The Japanese wage distribution has remained almost stable, with the exception of a compressed lower tail of the wage distribution among female workers (Genda, 1998; Shinozaki, 2002; Moriguchi and Saez, 2007; Kambayashi, Kawaguchi, and Yokoyama, 2008; Kawaguchi and Mori, 2008), in contrast to the wage dispersion observed in several Anglo-Saxon countries such as Canada, the United Kingdom, and the United States (Boudarbat, Limieux, and Riddell, 2006; Lemieux, 2006; Goos and Manning, 2007; Autor, Katz, Kearney, 2008).

Figure 1 displays the time series of the 10th, 50th, and 90th percentiles of the log nominal wage distribution in Japan and the United States. Panels A and B show Japan's unusual experience of nominal wage deflation, while panels C and D indicate the typical time series in Anglo-Saxon countries. The decline in the Japanese median wage after 1999 is in sharp contrast to the steady increase in the U.S. median wage for both male and female workers. During the period of deflation, the median wage was located at a higher position than the 10th and 90th percentile wages for Japanese male workers. The trends of 90/50 and 10/50 log wage differentials are at odds with a recent polarization in the U.S. labor market, in which employment in high-skilled and low-skilled jobs has expanded at the expense of medium-skilled jobs (Autor, Katz, Kearney, 2008). In contrast, the 10th and 90th percentiles of the wage distribution have diverged upwardly from the median for Japanese female workers. This trend implies dispersion at the right tail and compression at the left tail of the wage distribution. This paper focuses on the compression in the lower tail of the female wage distribution.[1]

This study assessed the importance of the minimum wage among labor market institutions as a determinant of the evolution of the wage distribution. We hypothesized that an increase in the real value of the minimum wage contributed to the compression of the wage distribution among low-skilled workers. Our hypothesis was largely motivated by previous studies in several countries, which documented the importance of the minimum wage as a determinant of the shape of the lower tail of

[1] The analysis of the upper tail of the wage dispersion will be left for a future study.

the wage distribution. DiNardo, Fortin, and Lemieux (1996) demonstrated that erosion of the real minimum-wage level during the 1980s contributed to the wage dispersion in the United States. Lee (1999) found that erosion of the real value of the minimum wage caused by general price inflation almost completely explained the wage dispersion over the corresponding period. Autor, Manning, and Smith (2008) confirmed that the minimum wage certainly plays a role in compressing the lower tail of the wage distribution after correcting for upward bias in Lee's (1999) results using an instrumental variable approach. Dustmann, Ludsteck, and Schönberg (2008) attributed the recent increase in the gap between the 15th and 50th percentile wages in Germany to a decline in the union coverage rate. In Germany, the contract between labor unions and firms that belong to an employer federation extends to nonunion workers for a specific group in a specific sector, while no statutory minimum wage exists. Some researchers have also provided relevant evidence in Japan. Abe and Tanaka (2007) pointed out that the prefectural minimum wage contributed to a reduction in the wage gap between full-time and part-time workers in rural areas. Abe and Tamada (2007) found that an increase in the minimum wage was associated with an increase in the wage level among part-time workers.[2] However, these studies examined only the effect on the level of the mean wage. Hori and Sakaguchi (2005) illustrated the wage distribution in 2003 by prefecture and industry separately for full-time and part-time workers but did not conduct a formal regression analysis for the relationship between minimum wage and wage distribution.[3]

In contrast, this study examined the evolution of Japanese wage distribution under conditions of wage deflation. Changes in the nominal minimum wage tend to lag behind general price inflation or deflation. Thus, the real value in minimum wage shifts toward the lower end of the wage distribution during a period of inflation, whereas the 'bite' of the minimum wage is greater during a period of deflation. In Japan, the nominal minimum wage has been revised and increased consistently almost every year. The rise in the minimum wage cumulated to about 20 percent between 1994 and 2003, despite the economic downturn. In fact, the wage distribution in rural areas with many low-wage

[2]In contrast to findings in Germany, Japan, and the United States, studies conducted in the United Kingdom have reported that the introduction of the British national minimum wage in 1999 did not contribute a great deal to wage compression because the minimum wage was low relative to the average wage and the fraction of workers affected by the minimum wage was very small (Dickens and Manning, 2004a, 2004b).

[3]A spike around the minimum-wage level is somewhat obscure in their illustration for several reasons. First, they did not take a logarithm of the hourly wage. Second, the sample was split into small subgroups. Finally, the bin width chosen was so narrow that the distribution was always uneven.

workers was vulnerable to deflation when combined with the increased minimum wage.

First, we used Lee's (1999) approach to quantify the contribution of the increased real minimum-wage level on wage compression among low-skilled workers between 1994 and 2003. Specifically, we examined how the minimum wage affected the shape of wage distribution by running a regression of the 10/50 log wage differential on the effective minimum wage. The hourly wage was calculated from the unusually precise data collected in the Basic Survey of Wage Structure (BSWS). Estimated regression coefficients were used to create the counterfactual wage distribution if the minimum wage stayed low in real terms during the 1990s and early 2000s. Then, we quantified the effect of the minimum wage on the lower tail wage inequality.

Our identification strategy was basically to exploit regional variation in the effective minimum wage over time, which is known as a difference-in-differences (or fixed-effects) approach. The 'effective' minimum wage can be measured using the distance between the log minimum wage and the log median wage. The minimum wage relative to the median wage varied significantly across prefectures because the nature of wage distribution differed by prefecture, as did the statutory minimum wage. Thus, the minimum wage effect on the wage distribution can be isolated from a common macroeconomic fluctuation. Our empirical model included prefecture-specific time trends as well as prefecture and time effects to allow for possible changes in the dispersion of the latent wage distribution. Instrumental variable methods were used to assess robustness against measurement errors and policy endogeneity. The kernel reweighting approach proposed by DiNardo, Fortin, and Lemieux (1996) was also employed to allow for changes in workforce composition.

A change in the minimum wage can affect the shape of the wage distribution via three channels: censoring, truncation, and spillover, as described by Lee (1999). Lee's approach is limited because it cannot differentiate among the three effects. Autor, Manning, and Smith (2008) decomposed the total effect into censoring and spillover effects under the assumptions of lognormal latent wage distribution and no disemployment effect. However, in light of evidence documented by Neumark and Wascher (2008), the assumption that the minimum wage will have no effect on employment may not be correct. Moreover, truncation could mechanically change the shape of the wage distribution. We developed an alternative method to evaluate the importance of the spillover effect on wage compression. This proposed method allows for a possible mechanical compression of the wage distribution associated with disemployment and does not require a distributional assumption about latent wage distribution. Although the spillover effect is not isolated from the censoring effect, it is evident from the comparison between actual and counterfactual wage changes in a range of percentiles during the sample period.

To develop our method, we analyzed how a binding minimum wage can affect employment among low-skilled workers. The effect of a minimum wage on employment is still vigorously debated, but both sides of the debate seem to agree that labor market friction determines whether a minimum wage has an adverse effect on employment among low-skilled workers (Card and Krueger, 1995; Neumark and Wascher, 2008). A few recent studies in Japan have investigated the disemployment effect (Kawaguchi and Yamada, 2006; Tachibanaki and Urakawa, 2007; Abe and Tamada, 2008) but no consensus has been reached, and the cross-sectional analyses conducted by Tachibanaki and Urakawa (2007) and Abe and Tamada (2008) did not control for either prefecture or year effects. We exploited the variations in effective minimum wage across prefectures over time to identify how the minimum wage affected employment. The employment rate was calculated using data from the Employment Status Survey (ESS).

In addition, we found that the minimum wage affected the number of new hires. Very few studies to date, with the exception of Portugal and Cardoso (2006), have explored this effect on worker flow, although many studies have investigated the effect on net employment.

Finally, we examined how the minimum wage affected the full-time/part-time wage differential. Some evidence suggests that an increase in the minimum wage may lower the pay gap between women working full-time and part-time. Manning and Petrongolo (2008) reported a faster wage growth at the bottom end of the hourly wage distribution for part-time workers compared to full-time workers after the introduction of a national minimum wage. Abe and Tanaka (2007) found that a minimum wage prevented wage erosion among part-time workers relative to full-time workers. We directly quantified how the minimum wage affected the pay gap between full-time and part-time workers using the counterfactual wage distribution without an increase in the effective minimum wage.

Our analysis revealed that an increase in the minimum wage relative to wage distribution significantly contributed to the compression of the wage distribution among low-skilled female workers. The 10/50 log wage differential was 0.35 in 1994 and narrowed to 0.32 in 2003 for male workers. If the distance between the minimum wage and the median of the wage distribution remained constant, the distance between the 10th and 50th percentiles of the wage distribution would have remained generally unchanged for the 10-year period. The 10/50 log wage differential stayed constant at around 0.51 between 1994 and 2003 for female workers but would have diverged by 0.05 without an increase in the real value of the minimum wage. These findings are consistent with the hypothesis that an increased real minimum wage contributes to wage compression among low-skilled workers. Moreover, the compression of the lower tail of the wage distribution was not attributable to the mechanical effect associated with disemployment, although a moderate adverse effect of minimum wage on employment was observed among middle-aged female workers. Furthermore, an increase in the real minimum-wage level contributed to a reduction in the pay gap between full-time and part-time workers by about 5 percentage points.

The remainder of this paper is organized as follows. Section 2 introduces the minimum-wage system in Japan. Section 3 describes the data used in our analysis. Section 4 examines how the minimum wage affected the lower tail of the wage distribution separately for male and female workers. We quantified the effect by comparing the actual wage distribution to the counterfactual wage distribution without an increase in the effective minimum wage. Section 5 reexamines the relationship between minimum wage and wage compression using a counterfactual sample in the absence of disemployment. Section 6 re-analyzes wage compression and disemployment, with a constant workforce composition in terms of workers' observable attributes. Section 7 analyzes the effect of the minimum wage on the full-time/part-time wage differential. The last section presents our conclusions.

2 Statutory Minimum Wage in Japan

The minimum wage in Japan is based on the Minimum Wages Law, which was enacted in 1959 and substantially revised in 1967. The current law defines two types of minimum wages: a regional minimum wage based on collective agreement and a minimum wage based on the research and deliberations of minimum-wage councils. The first system assumes that the minimum wage agreed upon by craft-wide or industry-wide bargaining will be extended to nonunionized workers within the same sector. However, such bargaining does not really exist under the Japanese enterprise union system; in practice, all minimum wages in Japan are currently of the second type.[4]

Under the current system, the chief of the prefectural labor bureau determines the level of the prefectural minimum wage based on the regional minimum-wage councils' deliberations. Deliberations are largely influenced by decisions regarding the amounts of any minimum-wage increases, which are set annually by the central minimum-wage council. The central minimum-wage council consists of representatives of public interest (academics and a retired bureaucrat), employers, and employees. The central council classifies all Japanese prefectures into four ranks by actual wage levels and the standard cost of living. The central minimum-wage council then decides the minimum-wage increase (*meyasu*) for each rank.

The *meyasu* system was introduced in 1978 to moderate regional disparity in the minimum wage. Therefore, the current level of the prefectural minimum wage is the 1977 level of the prefectural minimum wage plus the accumulated total of *meyasu* since 1978. The current minimum wage level is thus affected by the prefectural minimum wage level several decades ago when the *meyasu* system had not yet been established; this has generated regional variation in the minimum wage, even within a rank. The council rarely alters their classification of the 47 prefectures into four ranks. The economic climate naturally differs across prefectures within a rank, which provides an additional source of independent variation in the effective minimum wage. However, the revised value of prefectural minimum wage is not exactly the same as that indicated by the central minimum-wage council. We can address concerns about policy endogeneity by using the 'indicated' (*meyasu*) minimum wage as an instrument.

The political process of the minimum-wage determination described above tends to be biased towards the status quo. This causes the real value of the minimum wage to creep up during a period of wage deflation, as shown in Figure 2: a time series of nominal and real minimum wages indicates a steady increase in the real value of the minimum wage. A related matter is an upward bias in the consumer price index (CPI). The *meyasu* minimum wage is supposed to be indexed to price inflation. However, if living costs are overestimated, the real minimum wage can remain stable at a high level.

[4]In Japan, the statutory minimum wage is not coordinated with any measures to promote small and medium enterprise.

The real minimum wage displayed here is calculated by dividing the nominal minimum wage by the Japanese CPI or the bias-corrected CPI.[5][6] The real increment is higher than the nominal increment.

The political climate also creates a bias towards the equalization of minimum-wage levels across prefectures. In 2003, the hourly minimum wage was 708 yen in Tokyo and 605 yen in Aomori. Tokyo was classified as Rank A (with the highest minimum wage), while Aomori was classified as Rank D (with the lowest minimum wage). Apparently, regional minimum wages are much less heterogeneous across prefectures than wage distributions. Thus, the regional minimum wage appears not to have changed in response to economic shocks to local labor markets. Thus, the degree to which the wage distribution is affected by the minimum wage differs significantly across prefectures. During a recession, the minimum-wage bite may be severe in rural areas.

Figure 3 plots the minimum wage denominated by the median wage in 1994 and 2003 by prefecture. Tokyo is located in the bottom-left corner because the real value of its minimum wage was low for both years. In contrast, Aomori, Akita, Miyazaki, and Okinawa are located in the top-right corner because they had relatively high minimum wages compared to the median wage for both years. All prefectures experienced an increase in levels of the real minimum-wage during this 10-year period, as evidenced by the fact that all the prefectures lie above the 45-degree line. The vertical distance from the 45-degree line indicates that increases in minimum wages differed across prefectures in real terms.

Legal enforcement of the minimum wage is weak in Japan. The prefectural labor bureau is in charge of enforcement. When an employer's noncompliance is detected, the labor bureau may institute a fine of up to 20,000 yen (about 200 U.S. dollars). Employers who violate the minimum-wage law must also compensate employees for the difference between the minimum wage and the actual wage. However, in practice, the minimum wage seems to be enforced mostly through public pressure on employers. In particular, the reputations of larger companies would be damaged if the public were aware that they paid workers less than the minimum wage.

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[5] Broda and Weinstein (2007) found that the average rate of deflation was 1.2 percent per year between 1998 and 2006 when the substitution bias and quality upgrading were taken into account in computing the Japanese CPI.

[6] The minimum wage is normalized by the median wage in the subsequent regression analysis. Our measure of the effective minimum wage is free of the CPI bias.

3 Data

This analysis used 1994–2003 micro data from the Basic Survey of Wage Structure (BSWS), which is compiled annually by the Japanese government. The survey covers private establishments with 5 or more regular employees and public establishments with 10 or more regular employees in almost all regions and industries in Japan, with the exception of agriculture. Approximately 1.5 million workers have been surveyed every year from 60,000–70,000 establishments. Establishments are randomly sampled in proportion to prefecture and industry size and the number of employees according to the Establishment and Enterprise Census, which lists all establishments in Japan. For the survey, randomly selected establishments are asked to extract employee information from payroll records,[7] and establishments and individual files are then merged using an establishment identification number.

The cross-sectional unit in the analysis is an individual worker whose relevant information is available from the establishment. Both full-time and part-time workers are included in the sample when they are directly hired by employers and accordingly appear on the establishment's payroll record. However, the BSWS does not cover workers who are employed by temporary agent firms and dispatched to establishments. The available information includes each worker's wages, age, sex, educational attainment only for full-time workers, full-time/part-time status, type of work or job, and working days/hours, as well as the firm's attributes, such as the number of regular workers (*joyo rodo sha*),[8] the number of new graduates hired, firm size, industry, and location. Data about wages include individuals' contracted hours of work and overtime hours between June 1 and June 30, contracted pay, overtime pay, and allowances (e.g., for family and transportation) over the corresponding period. Japanese minimum wage laws apply to the straight wage rate excluding allowances. We defined hourly wage as (wages for contracted hour – commutation allowance – perfect attendance allowance

[7] A person in charge of personnel at each establishment is asked to randomly choose a number of workers from the pool of employees using specific instructions for random sampling, including the sampling probability, which is dependent on the industry and establishment size. The sample does not include board members whose wage is set at a general meeting of shareholders.

[8] Workers who meet one of the following three criteria are classified as regular workers: 1. On contracts that do not clearly specify a contractual time period; 2. On contracts that last more than one month; or 3. On contracts that last less than one month, but on which the workers worked 18 or more days in the last two months. This classification includes part-time workers if one of the above criteria is satisfied.

– family allowance)/contracted hours of work, which is consistent with the minimum-wage law.[9][10]

Our analysis on how the minimum wage affects employment also included data from a household survey that covers non-employed as well as employed individuals. We used the Employment Status Survey (ESS) for the years 1997 and 2002. The ESS is distributed every 5 years to approximately 440,000 households in sampled units that cover the complete population.[11] The survey collects information about the number of household members and labor force status for household members aged 15 and older as of October 1 of each survey year. Our study drew on micro data about employment status, educational attainment, age, sex, and residential area. Overall, the sample included approximately 1 million individuals, with a half-million males and a half-million females for each year that the survey was conducted. The sample was restricted to data with valid age, educational background, and employment status.

4 The Role of the Minimum Wage

4.1 The Evolution of the Wage Distribution under Deflation

Increases in the minimum wage can affect the wage distribution through three channels. First, the wage distribution may be censored by the minimum wage. In this case, the wage distribution spikes around the minimum wage. Second, an increased minimum wage may result in the truncation of the wage distribution associated with disemployment. The disappearance of the bottom end of the wage distribution can mechanically change the distance between the 10th and 50th percentiles of the wage distribution. Finally, a rise in the minimum wage may exert a spillover effect on workers who earn more than the minimum wage. In a competitive labor market, substitution between workers with different skill levels can affect the wages paid to workers who are not directly affected by the minimum wage (Teulings, 2000, 2003). In a monopsonic labor market, spillovers can occur when the labor supply curve facing an employer is shifted by increases in the reservation wage for unemployed

[9]A change in the minimum wage conceivably may affect the level of allowances. However, the results obtained in our analysis are unchanged even when hourly wage is defined as wages (including allowances) for contracted hours divided by contracted hours of work.

[10]The custom of tipping is not practiced in Japan.

[11]The sample does not include foreign diplomats, foreign military personnel and their dependents, persons dwelling in Self Defense Force camps or ships, and persons serving sentences in correctional institutions.

workers (Manning, 2003).

Figures 4A and 4B illustrate the log wage distribution in low-wage and high-wage prefectures in 1994 and 2003 for male and female workers with hourly wages between 400 and 3,500 Japanese yen. The horizontal axis is the level of hourly wage. Aomori and Tokyo are examples of low-wage and high-wage prefectures, respectively. The wage distribution moved dramatically toward the lower end in Aomori from 1994 to 2003. A spike emerged around the minimum-wage level in the male wage distribution, while the female wage distribution was skewed and flattened at the minimum-wage level. The wage density is the highest at the minimum wage for female workers. In fact, more than 5 percent of female workers earned the minimum wage or less in 2003. As indicated by the density of hourly wages below 1,000 Japanese yen, the proportion of male low-wage workers also increased in Tokyo, although not to the same extent as in Aomori. However, the minimum wage did not seem to bind workers in Tokyo over the sample period.

Figure 5 illustrates the relationship between minimum wage and wage compression. The figure plots the 10th percentile wage relative to the median wage along with the log of the minimum wage relative to the median wage by sex. The slope of the fitted line is positive both in panels A and B and greater in panel A for male workers than panel B for female workers. The plotted points are located at a higher position in panel B than in panel A because the distance between the 10th and 50th percentiles of the wage distribution is shorter for female workers than male workers. The plotted points are slightly mixed up in panel A, but are separated by year into the upper right and lower left in Panel B. The increase in the real value of the minimum wage appears to be an important cause of the compression of the female 10/50 log wage differential.

4.2 Effect on the Wage Distribution

We conducted a regression analysis to investigate the cause for this wage compression. Based on Lee's (1999) approach, we examined to what extent the minimum-wage bite can explain the compression between the 10th and 50th percentiles of the wage distribution separately for male and female workers.

Our estimation applied a model of the form

$$\ln\left(\frac{w_{it}^{p}}{w_{it}^{50}}\right) = \beta_p \ln\left(\frac{mw_{it}}{w_{it}^{50}}\right) + d_t\gamma + d_i\delta + u_{it}, \tag{1}$$

where i is the index for prefecture and t is the index for year.[12] The variable w^p is the pth percentile of the wage distribution, mw is the minimum wage, d_t is a vector of year dummies, and d_i is a vector of prefecture dummies. The minimum-wage bite is measured by the log of the minimum wage relative to the median wage. Parameter β_p represents the percentage change in the pth percentile wage relative to the median wage caused by a one percent increase in the effective minimum wage. Higher-order terms of the effective minimum wage might be included as additional regressors to capture the nonlinear relationship between the minimum wage and wage compression. However, neither square nor cubic terms were statistically significant in the model without prefecture effects, as can be expected from Figure 5. Year effects represent the evolution of the wage distribution over time, the real minimum-wage level being constant. Prefectural fixed effects (FEs) were added to allow for unobserved heterogeneity in the dispersion of the latent wage distribution across prefectures in some specifications.

Estimated results were produced using ordinary least-squares (OLS) and two-stage least squares (2SLS), and standard errors were clustered at the prefecture level. Tables 1A and 1B list the results for male and female workers, respectively. We began with an OLS regression of the 10/50 log wage differential on year dummies. The estimated year effects represent the unconditional evolution of the wage distribution over time. Column 1 in Table 1A shows that the 10/50 log wage differential was almost stable between 1994 and 2000 and started to erode slightly thereafter. Next, we added the effective minimum wage as an additional regressor. Column 2 shows that the coefficient for the effective minimum wage was positive and significant and that R^2 rose from 0.06 to 0.38. These results imply that the minimum-wage bite contributed to the reduction in the distance between the 10th and 50th percentiles of the wage distribution. The coefficients for year dummies shrank when the effective minimum wage was fixed. Thus, the 10/50 log wage differential would have diverged if the real minimum-wage level were unchanged.

[12] See Lee (1999) and Autor, Manning, and Smith (2008) for an explanation of its derivation and justification.

However, the OLS may suffer from an upward bias associated with sampling errors. Because the median wage appears in both sides as a denominator, the wage differential is automatically positively correlated with the effective minimum wage when the median wage is measured with errors. We extended an instrumental variable approach employed by Autor, Manning, and Smith (2008) to work around the potential bias. The instruments for the effective minimum wage were the minimum wage and the median of the log wage within a prefecture over the sample period. However, the minimum wage used here is not the actual but the 'indicated' (*meyasu*) minimum wage, to allow for policy endogeneity. As shown in column 3, the 2SLS estimates were almost identical to the OLS estimates. Given the large sample size of the BSWS and the Japanese minimum-wage setting, an identical result between the OLS and 2SLS seems quite natural. Hence, the potential bias due to sampling errors or policy endogeneity was negligible in our analysis.

Column 4 shows that the minimum wage had a greater effect on wage compression when we controlled for prefecture effects. Thus, the results obtained in columns 1–3 were not driven by the unobserved heterogeneity across prefectures. Column 5 shows that the results are robust to the inclusion of prefecture-specific time trends.

Columns 7–11 report the results of the 90/50 log wage differential, which declined slightly during the sample period. Contrary to our expectation, the effective minimum wage had a positive and significant effect on the 90/50 log wage differential. R^2 rose from 0.04 to 0.40. Again, the 2SLS estimates were identical to the OLS estimates. The minimum wage effect became smaller but still remained after controlling for prefecture effects and its interaction terms with time trends. However, that an increase in the minimum wage would push up the 90th percentile wage is not very realistic.

Our concern is a spurious correlation arising from an omitted variable. In the absence of a minimum wage, the 90/50 log wage differential should be larger in a prefecture with a higher level of wage dispersion. If the effective minimum wage is positively correlated with the dispersion of latent wage distribution, the effect of the minimum wage on the 90/50 log wage differential will be biased upward. In contrast, the effect on the 10/50 log wage differential will be biased downward because the sign of the correlation with the wage dispersion is reversed. We added the 75/50 log wage differential as an additional regressor to circumvent this potential bias.[13] As expected, column 6 reveals that the

[13] The results obtained from this analysis were identical after the 75/50 log wage differential was replaced with the

minimum wage had a greater effect on the 10/50 log wage differential, and column 12 shows that the minimum wage had a lesser effect on the 90/50 log wage differential. These results can be interpreted as correcting for the omitted-variable bias. However, the effect on the 90/50 log wage differential remains significant. Therefore, our estimate for the effect of the minimum wage on the lower tail wage inequality for male workers may be biased downward.

Columns 1–6 in Table 1B list the results of the 10/50 log wage differential for female workers. Column 1 shows the increasing trend of the 10th percentile wage relative to the median wage. Column 2 indicates that the coefficient for the effective minimum wage is positive and significant. R^2 rose from 0.06 to 0.56 after the effective minimum wage was added. Moreover, the estimated year effects were virtually zero or sometimes negative, conditional on the effective minimum wage. These results suggest that wage compression can be entirely explained by increases in the effective minimum wage over the sample period. The 2SLS estimates were identical to OLS estimates. The effective minimum wage had an even stronger effect after controlling for prefectural fixed effects, prefecture-specific trends, and wage dispersion. The minimum wage played a more significant role in pushing up the lower tail of the wage distribution for female workers in terms of estimated coefficients and R^2.

Columns 7–12 report the results of the 90/50 log wage differential. Column 7 shows that the 90/50 log wage differential was almost stable during the sample period. Column 8 reveals a positive and significant effect of minimum wage on the 90/50 log wage differential but R^2 increased only 6 percentage points. In this sense, the effective minimum wage does not significantly explain the 90/50 wage differential. The OLS estimates were identical to 2SLS estimates and similar to the fixed-effects estimates. However, the effect of the minimum wage plummeted after we controlled for prefecture trends or the 75/50 log wage differential. Overall, the minimum wage played a significant role in compressing the lower tail of the wage distribution but did not account for the change in the upper tail of the wage distribution for female workers.

70/60 log wage differential.

4.3 Counterfactual Wage Distribution without an Increase in the Effective Minimum Wage

Increases in the real value of the minimum wage contributed to the compression in the lower tail of the wage distribution from 1994 to 2003, especially among female workers, as described thus far. In other words, wage compression might not have occurred if the effective minimum wage had remained unchanged over the 10-year period. Following Lee's (1999) procedure, we constructed a counterfactual wage distribution without any increase in the effective minimum wage to quantify the relationship between the minimum wage and wage compression in more detail. The counterfactual wage in 2003 was calculated by subtracting the effect of the 10-year difference in the effective minimum wage from the actual wage in 2003. Specifically, for a worker k whose hourly wage ranks at pth percentile in prefecture i, the counterfactual wage in 2003 was simulated as follows:

$$\ln \widetilde{w_{k,i,2003}^{p}} = \ln w_{k,i,2003}^{p} - \widehat{\beta_p} \left\{ \ln\left(\frac{mw_{i,2003}}{w_{i,2003}^{50}}\right) - \ln\left(\frac{mw_{i,1994}}{w_{i,1994}^{50}}\right) \right\}, \tag{2}$$

where $\widehat{\beta_p}$ is the estimated coefficient obtained from the regression of the percentile wage differential on the effective minimum wage, year dummies, and prefecture dummies.[14] The pth percentile varies with prefecture, year, and sex.

Figure 6 displays the actual and counterfactual wage distributions in 1994 and 2003. The horizontal axis is the log hourly wage. Panels A and B show the wage distribution for male and female workers, respectively. The actual wage distribution in 1994 nearly overlaps with that in 2003 for male workers. However, the counterfactual wage distribution suggests that the lower tail of the wage distribution would have eroded if no increase in the minimum wage had occurred. The lower tail of the actual wage distribution in 2003 is compressed for female workers. The compression is displayed by the spike in the lower tail of the wage distribution. However, the lower tail of the counterfactual wage distribution in 2003 overlaps with that of the actual wage distribution in 1994. Thus, the compression of the lower tail of the wage distribution can be mostly attributed to the minimum wage increase.

Figure 7 displays the actual and counterfactual changes in the log hourly wage by percentile be-

[14] The minimum wage effect simulated here should be conservative relative to the one based on a fixed-effects model with prefecture-specific time trends.

tween 1994 and 2003. Panels A and B show that the actual 10th percentile wage remained unchanged for male workers and that lower percentile wages increased considerably for female workers between 1993 and 2004. However, if the effective minimum wage remained at the 1994 level, the 10th percentile wage would have fallen by 4.4 percentage points for male workers. Moreover, the actual rise in the lower percentiles of the female wage distribution can be largely attributed to the minimum-wage hike. Indeed, the simulated change in the log hourly wage was close to one percentage point from the 10th to 35th wage percentiles for female workers. The difference between the actual and counterfactual wage changes indicates a significant spillover effect on workers who earn more than the minimum wage.

5 Wage Compression or Employment Loss?

5.1 Effect on Employment

The minimum wage provided a "wage floor" during the period of deflation, as seen above. This brings up the question of how the wage floor affected employment during the corresponding period. To examine how the minimum wage affected employment, we conducted a standard pseudo-panel data analysis as set forth by Neumark and Wascher (1992) and Card and Krueger (1995) among others.

The employment rate for demographic group j in prefecture i in year t can be specified as

$$\frac{emp_{jit}}{pop_{jit}} = \beta_{0j} + \beta_{1j} \ln\left(\frac{mw_{it}}{w_{it}^{50}}\right) + d_t\gamma_j + d_i\delta_j + u_{jit}, \tag{3}$$

where emp is the number of employed individuals, and pop is population size. Again, the effective minimum wage is measured by the log of the minimum wage relative to the median wage.[15] A common macroeconomic shock is flexibly captured by year dummies in this specification. The results obtained in this study changed only marginally, even after the employment rate for male college graduates aged between 31 and 59 was included as an additional regressor to enable further control of aggregate fluctuations in employment. In light of the criticism by Card and Krueger (1995), we did

[15] The time trend in our measure of the effective minimum wage is similar to the Kaitz index.

not include the college enrollment rate as a regressor. We included prefecture dummies to allow for an unobserved prefecture effect in some specifications. Thus, the regressors include only exogenous variables in our preferred reduced-form specification of labor demand.

Our analysis focused on low-skilled workers who had completed high school or less. This low-skilled group tends to be most affected by increases in the minimum wage. Typical low-wage workers are young or middle-aged women with part-time jobs. If parameter β_1 is negative, an increase in the minimum wage reduces the employment rate. The elasticity of the employment rate can be calculated via the estimated parameter β_1 divided by a national average of employment rate for group j. The employment rate was calculated from ESS data; these surveys were conducted only in the years 1997 and 2002, during the sample period of 1994–2003. The effective minimum wage was calculated from the BSWS, as in the previous analysis.

The model was estimated using weighted least squares (WLS). We used the square root of the sample variance in the employment rate as the weight. This approach is also known as a minimum χ^2 method for the analysis of grouped data. Table 2A reports the results of the disemployment effect by age group among male workers. Columns 1–4 report the cross-sectional estimates. The estimated year effect was negative for all age groups, indicating a decline in the labor force attachment. The disemployment effect was small but significant for males aged 31–59 years. However, it became non-significant after we controlled for the prefecture effect. Indeed, we found no statistically significant effects in any age group in fixed-effects specifications.

Table 2B reports how the minimum wage affected female employment by age group. The minimum wage effect was nonsignificant except for females aged 31–59 years. However, column 7 reveals a moderate and significant disemployment effect for females aged 31–59 years. The implied elasticity is –0.320. This result seems plausible, given the high proportion of part-time workers and the fact that the minimum-wage bite is considerable for this demographic group.[16]

[16] When the number of hours of work was used as a dependent variable instead of the employment rate, the estimated minimum wage effects were generally statistically non-significant for all demographic groups after controlling for prefectural fixed effects. These results are not surprising because a decreased number of employees can reduce firms' labor costs more effectively than the number of hours of work when fixed costs arise in employment. Most previous studies have focused on the effect on employment.

5.2 Effect on New Hires

The costs of employment adjustment are asymmetric between hiring and firing. Hiring is less costly than firing because employment regulations levy high firing costs on firms.[17] Estimated employment elasticity was imprecise but greatest for those aged 22 or younger, as shown in column 5 of Tables 2A and 2B. Given the costs incurred by firing, including legal costs and sunk costs for training, the disemployment effect is presumably pronounced at the margin of new hires.

We examined how the minimum-wage hike affected the number of new hires conditional on the number of regular workers, similar to the pseudo-panel data analysis of net employment. The number of new graduates hired is available from the BSWS for every year between 1994 and 2003, whereas the employment rate is available from the ESS every five years. The results for female new graduates are outlined below.

$$\overline{\ln\left(NewHire_{it}\right)} = \underset{(0.554)}{-1.470}\ln\left(\frac{mw_{it}}{w_{it}^{50}}\right) + \underset{(0.238)}{0.677}\ln\left(Employee_{it}\right) + d_t\widehat{\gamma} + d_i\widehat{\delta} + \left(d_i \cdot t\right)\widehat{\zeta}, \quad (4)$$
$$R^2 = 0.97$$

where $NewHire$ is the number of new graduates hired, $Employee$ is the number of regular workers, and the last three terms are prefecture dummies, year dummies, and prefecture-specific time trends, respectively. The bar represents the predicted value, and the hat represents the OLS estimator. A total of 470 observations are included. Standard errors in parenthesis are clustered at the prefecture level.

As shown above, a one-percent increase in the effective minimum wage leads to a 1.47-percent decrease in female new graduates hired. For male new graduates, however, the effect on new hires was statistically non-significant and smaller than that for female new graduates. The estimated coefficient on the effective minimum wage was –0.293 with a standard error of 0.822, and R^2 was 0.94. If the interpretation can be extended to other demographic groups, the minimum-wage hike is also considered to have reduced new hires among middle-aged female (part-time) workers.

[17] In most cases, Japanese employment regulations are not put into statutory form but are established by court precedents (Sugeno 2002).

5.3 Counterfactual Wage Distribution in the Absence of Disemployment

Up to this point, the results have confirmed that an increase in the effective minimum wage compresses the lower tail of the wage distribution but reduces employment for low-skilled middle-aged female workers. Our concern is that the lower tail of the wage distribution may be mechanically compressed by the truncation of the bottom end of the wage distribution associated with disemployment. The distance between the 10th and 50th percentiles of the wage distribution should mechanically shrink after the wage distribution is truncated at the minimum wage.[18]

We constructed a counterfactual wage distribution to quantify the mechanical effect. The estimated disemployment effect was used to recover the counterfactual wage distribution if employment loss did not occur. The change in the employment rate caused by changes in the minimum wage between years $t-1$ and t can be expressed as $\Delta\left(\frac{emp_{jit}}{pop_{jit}}\right) = \widehat{\beta_{1j}}\Delta\log\left(\frac{mw_{it}}{w_{it}^{50}}\right)$, where $\widehat{\beta_{1j}}$ is the estimated coefficient for the effective minimum wage in the fixed-effect estimates of the employment equation for group j. The change in the number of employed can be expressed as

$$\Delta emp_{jit} = \widehat{\beta_{1j}}\Delta\ln\left(\frac{mw_{it}}{w_{it}^{50}}\right)\cdot pop_{jit} + \frac{\Delta pop_{jit}}{pop_{ji,t-1}}\cdot emp_{t-1}, \tag{5}$$

where Δpop_{jit} is assumed to be constant over time. The first term represents the reduction in the number of employed caused by the minimum-wage hike, and the second term represents the change in the number of employed due to the population change. Using this formula, the counterfactual wage distribution for group j in prefecture i in year t can be constructed using the following steps.

1. Substituting the actual change in the effective minimum wage yields the number of workers who lost their job by group and prefecture between years $t-1$ and t. Then, calculate the total number of unemployed workers between 1995 and 2003, $N_{it}^{add} = \sum_j(-\Delta emp_{jit})$.

2. Adding N_{it}^{add} workers into the lowest end of the wage distribution yields the counterfactual wage distribution in the absence of disemployment. The counterfactual wage distribution is

[18] An arbitrary continuous distribution requires that $\int_{\omega^{10}}^{\omega^{50}} f(\omega)\,d\omega = 0.4$, where ω is the log hourly wage, ω^{10} and ω^{50} are the 10th and 50th percentiles of the log wage distribution and $f(\cdot)$ is the probability density function. The distribution truncated at mw also requires that $\int_{\omega_*^{10}}^{\omega_*^{50}} f_*(\omega)\,d\omega = 0.4$, where ω_*^{10} and ω_*^{50} are the 10th and 50th percentiles of the truncated distribution Then, $f_*(\omega) = f(\omega\,|\omega \geq mw) = \frac{f(\omega)}{1-\Pr(\omega\geq mw)} \geq f(\omega)$. Thus, $\omega^{50}-\omega^{10} \geq \omega_*^{50}-\omega_*^{10}$.

produced by the wage data on $N_{it} + N_{it}^{add}$ workers, where zero log wage is assigned to N_{it}^{add} unemployed workers.

We constructed the counterfactual wage distribution only for female workers because we found no significant disemployment effect for male workers. Among low-skilled workers, the minimum-wage hike reduced employment for females aged 31–59 years. Thus, we recovered the wage distribution in the absence of disemployment for this demographic group. In the process of creating the counterfactual sample, we lost the observations from the first year of the sample period. The results of the 10/50 log wage differential in Table 1B were reproduced for the counterfactual sample in Table 3. Estimation results differed only marginally. Column 2 shows the estimated coefficient for the effective minimum wage, which is almost identical to that in Table 1B. Hence, the mechanical effect associated with disemployment is negligible. After we added the effective minimum wage, R^2 rose from 0.06 to 0.59, and the estimated year effects became virtually zero or sometimes negative. Columns 3–6 confirm the robustness of the results. The reduction in the distance between the 10th and 50th percentiles of the wage distribution can still be generally explained by the increase in the effective minimum wage. Thus, the compression of the wage distribution cannot be attributed to the truncation of the wage distribution, but to the censoring and spillover arising from the minimum-wage hike.

6 Minimum Wage Effects or Composition Effects?

In Japan, the labor force has been aging, and job tenure has increased for female workers in the period between 1994 and 2003. These shifts in workforce composition may have mechanically raised or lowered wage inequality. To isolate the minimum wage effect from the composition effect, we employed the kernel reweighting approach proposed by DiNardo, Fortin, and Lemieux (1996, hereafter DFL). The DFL approach enables us to estimate counterfactual 10/50 and 90/50 log wage differentials and the counterfactual effective minimum wage, the observed attributes being fixed at 1994 levels. Educational information is not available for part-time workers in BSWS, while job tenure is available for all workers. In fact, job tenure is the key determinant of wages in the Japanese labor market. Thus, a full set of dummy variables for age and job tenure are used as the attributes to calculate the reweighting function. Next, we reexamined Lee's (1999) model of wage compression using the counterfactual wage data without changes in workforce composition.

Similar results were obtained for male workers (see Tables 1A and 4A), while results for female workers differed slightly (see Tables 1B and 4B). The minimum wage effect became greater for the female lower-tail (10/50) inequality and lesser for the female upper-tail (90/50) inequality. The former suggests that wage compression caused by the minimum-wage hike would be more pronounced if neither aging nor increase in job tenure occurred. The latter implies that the expansion of female upper tail inequality can be attributed to aging and lengthening job tenure in the labor force. These results reinforce our findings in Section 4.

We used a similar approach to reexamine the disemployment effect. When the counterfactual employment rate was estimated using the DFL approach, job tenure was replaced with years of potential work experience because job tenure information was not available for non-employed individuals in ESS, while educational information is available for both employed and non-employed individuals. Therefore, it is not possible to conduct a simple comparison of the results. However, Table 5 shows that our findings in Section 5 are robust to the change in workforce composition. A moderate disemployment effect was observed for females aged 31–59.

7 The Part-time Pay Penalty

The effect of the minimum wage on wage compression has an implication for the part-time penalty, i.e., the pay gap between full-time and part-time workers. Employees who are paid the minimum wage are typically part-time workers. A reduction in dispersion in the bottom end of the wage distribution may cause a reduction in the full-time/part-time wage differential. Again, the counterfactual wage distribution without an increase in the effective minimum wage was used to quantify the effect of the minimum wage on the pay gap between full-time and part-time workers. Our analysis focused on female workers because the proportion of male part-time workers was very small.[19]

Table 6 reports the actual and counterfactual pay gaps between full-time and part-time workers. The fraction of part-time workers in the workforce increased from 21.5 to 32.0 percent between 1994

[19] The fraction of part-time male workers was 1.8 percent in 1994 and 4.0 percent in 2003.

and 2003. The actual pay gap was 36 percent in 1994 and increased to 38 percent in 2003. However, the pay gap would have been 40 percent in 2003 if the minimum wage had remained at the 1994 level. These results imply that the minimum wage contributed to the reduction in the full-time/part-time wage differential by 2 percentage points at the mean.

Figure 8 illustrates the full-time/part-time log wage differential by wage percentile. The pay gap between full-time and part-time workers increases from the lower to the upper tail of the wage distribution. The actual pay gap did not change below the 30th percentile between 1994 and 2003. However, the pay gap would have expanded if no increase in the minimum wage had occurred. The minimum wage had a greater effect in the lower tail of the wage distribution. For example, the simulated pay gap without the minimum wage increase is about 5 percentage points at the 25th percentile.

8 Conclusions

This study empirically examined how the minimum wage affected the wage distribution between 1994 and 2003 in Japan, the world's second largest economy. Japan's experience after the late 1990s differed from that of the United States in the 1980s and 1990s. The median wage fell in a deflationary economy, and the statutory minimum wage steadily increased despite the recession. The combination of the declines in the median wage and increases in the minimum wage substantially raised the minimum wage relative to median wage between 1994 and 2003. Indeed, the minimum-wage hike compressed the lower tail of the wage distribution in Japan, whereas a fall in the effective minimum wage resulted in an increased wage inequality in the United States.

Our analysis revealed that the minimum wage had a significant effect on wage compression for female workers. The decline in the 10/50 wage differential among female workers between 1994 and 2003 was largely explained by the increase in the minimum wage relative to the median wage. Without this increase in the effective minimum wage, only small increases in hourly wages in the lower half of the distribution would have occurred for female workers. The minimum-wage hike reduced employment for low-skilled middle-aged female workers. The disemployment effect was –0.32 in elasticity terms. However, we obtained similar results for wage compression after recovering the wage distribution in the absence of disemployment. We also found that the increase in the effective minimum wage decreased the full-time/part-time wage differential by 5 percentage points in the lower tail of the wage distribution among female workers. These results held even after controlling for composition effects.

To summarize, the minimum wage provided a wage floor for female workers in Japan's deflationary economy. However, this benefit of the minimum-wage system came at the cost of moderate employment loss among low-skilled middle-aged female workers. The findings imply a policy trade-off between the reduction in wage inequality and disemployment of workers who are weakly attached to the labor market.

Some issues remain for future research. First, it would be helpful to address the issue of employment in more detail by using unique data about job flow at the establishment level. Data from the Survey on Employment Trends (*Koyou Doukou Chousa*) could be used to analyze how the minimum wage affects job flow for various demographic groups. Second, the minimum-wage hike may affect college enrollment and occupational choices. Moreover, constructing a model of educational and occupational choices would be helpful to examine how the minimum wage affects complex individual choices.

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Table 1A: How the real value of the minimum wage affected the wage distribution.
Sample: Males, 1994–2003

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Estimation Methods	OLS	OLS	2SLS	FE	FE	OLS	OLS	OLS	2SLS	FE	FE	OLS
Dependent Variables	10/50 log Wage Differential						90/50 log Wage Differential					
ln(MW/W50)	–	0.29	0.28	0.49	0.50	0.41	–	0.42	0.42	0.32	0.27	0.21
		(0.07)	(0.07)	(0.06)	(0.06)	(0.05)		(0.12)	(0.13)	(0.06)	(0.06)	(0.04)
ln(W75/W50)	–	–	–	–	–	-0.80	–	–	–	–	–	1.46
						(0.13)						(0.09)
Year 1995	-0.00	-0.00	-0.00	-0.01	-0.01	-0.01	-0.01	-0.01	-0.01	-0.01	-0.01	-0.01
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Year 1996	0.00	-0.01	-0.01	-0.01	-0.01	-0.01	-0.00	-0.01	-0.01	-0.01	-0.00	-0.01
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)
Year 1997	0.01	-0.00	-0.00	-0.01	-0.01	-0.01	-0.02	-0.03	-0.03	-0.03	-0.02	-0.01
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)
Year 1998	0.00	-0.01	-0.01	-0.02	-0.02	-0.02	-0.02	-0.04	-0.04	-0.04	-0.02	-0.02
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.00)	(0.00)	(0.00)
Year 1999	0.00	-0.01	-0.01	-0.02	-0.03	-0.03	-0.02	-0.04	-0.04	-0.04	-0.02	-0.02
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)
Year 2000	-0.00	-0.02	-0.02	-0.03	-0.04	-0.04	-0.02	-0.05	-0.05	-0.05	-0.02	-0.02
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)
Year 2001	-0.01	-0.03	-0.03	-0.04	-0.05	-0.04	-0.01	-0.04	-0.05	-0.04	-0.01	-0.02
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)
Year 2002	-0.02	-0.04	-0.04	-0.06	-0.07	-0.05	-0.01	-0.04	-0.04	-0.03	-0.00	-0.03
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)
Year 2003	-0.02	-0.05	-0.05	-0.07	-0.07	-0.06	-0.02	-0.06	-0.06	-0.05	-0.02	-0.03
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)
Constant	-0.51	-0.20	-0.21	-0.04	-0.01	0.19	0.62	1.06	1.07	0.98	-0.01	0.36
	(0.00)	(0.08)	(0.08)	(0.06)	(0.00)	(0.08)	(0.01)	(0.13)	(0.13)	(0.06)	(0.00)	(0.06)
Prefecture trends	No	No	No	No	Yes	No	No	No	No	No	Yes	No
R^2	0.06	0.38	–	0.91	0.94	0.65	0.04	0.40	–	0.93	0.96	0.88

Notes: A total of 470 observations are included. Standard errors in parentheses are clustered at the prefecture level. MW, W50, and W75 represent minimum wage, median wage, and 75th percentile wage, respectively. The base year is 1994. Instrumental variables are the *meyasu* minimum wage and the median of the log wage within a prefecture between 1994 and 2003. The first-stage *F*-statistic is 170,000.

Table 1B: How the real value of the minimum wage affected the wage distribution.
Sample: Females, 1994–2003

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Estimation Methods	OLS	OLS	2SLS	FE	FE	OLS	OLS	OLS	2SLS	FE	FE	OLS
Dependent Variables	10/50 log Wage Differential						90/50 log Wage Differential					
ln(MW/W50)	–	0.39	0.38	0.54	0.61	0.42	–	0.21	0.21	0.27	0.11	0.09
		(0.04)	(0.04)	(0.10)	(0.06)	(0.03)		(0.11)	(0.12)	(0.13)	(0.10)	(0.04)
ln(W75/W50)	–	–	–	–	–	-0.49	–	–	–	–	–	1.76
						(0.07)						(0.06)
Year 1995	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00	-0.01	-0.01	-0.01	-0.01	-0.00	-0.01
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Year 1996	0.00	-0.01	-0.01	-0.01	-0.01	-0.01	-0.02	-0.02	-0.02	-0.02	-0.01	-0.02
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)
Year 1997	0.00	-0.01	-0.01	-0.01	-0.02	-0.01	-0.02	-0.03	-0.03	-0.03	-0.01	-0.02
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)
Year 1998	0.01	-0.01	-0.01	-0.02	-0.02	-0.02	-0.03	-0.04	-0.04	-0.04	-0.01	-0.03
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)
Year 1999	0.00	-0.02	-0.02	-0.03	-0.03	-0.02	-0.02	-0.03	-0.03	-0.04	0.01	-0.03
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)
Year 2000	0.01	-0.01	-0.01	-0.02	-0.03	-0.02	-0.02	-0.04	-0.04	-0.04	0.01	-0.03
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Year 2001	0.02	-0.01	-0.01	-0.02	-0.03	-0.01	-0.01	-0.03	-0.03	-0.03	0.03	-0.03
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Year 2002	0.02	-0.01	-0.01	-0.03	-0.03	-0.01	-0.00	-0.02	-0.02	-0.03	0.04	-0.04
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
Year 2003	0.03	-0.01	-0.01	-0.02	-0.03	-0.01	-0.01	-0.03	-0.03	-0.03	0.04	-0.04
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
Constant	-0.35	-0.14	-0.15	-0.10	-0.00	0.01	0.60	0.71	0.71	0.77	-0.00	0.17
	(0.00)	(0.02)	(0.02)	(0.05)	(0.00)	(0.03)	(0.01)	(0.06)	(0.07)	(0.07)	(0.00)	(0.03)
Prefecture trends	No	No	No	No	Yes	No	No	No	No	No	Yes	No
R^2	0.07	0.56	–	0.91	0.96	0.71	0.03	0.09	–	0.83	0.91	0.84

Notes: A total of 470 observations are included. Standard errors in parentheses are clustered at the prefecture level. MW, W50, and W75 represent minimum wage, median wage, and 75th percentile wage, respectively. The base year is 1994. Instrumental variables are the *meyasu* minimum wage and the median of the log wage within a prefecture between 1994 and 2003. The first-stage *F*-statistic is 26,143.

Table 2A: How the minimum wage affected the employment rate.
Dependent variable: Employment rate
Sample: Males, High School Education or Less, 1997 and 2002

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Estimation Methods	WLS				FE			
Age Groups	≤22	23–30	31–59	≥60	≤22	23–30	31–59	≥60
ln(MW/W50)	0.226	-0.008	-0.047	-0.040	-0.400	0.044	-0.052	-0.095
	(0.086)	(0.034)	(0.024)	(0.062)	(0.334)	(0.142)	(0.077)	(0.173)
Year 2002	-0.081	-0.047	-0.034	-0.056	-0.042	-0.050	-0.034	-0.053
	(0.013)	(0.005)	(0.004)	(0.009)	(0.021)	(0.009)	(0.005)	(0.011)
Constant	1.063	0.925	0.901	0.447	0.413	0.979	0.896	0.389
	(0.089)	(0.036)	(0.025)	(0.064)	(0.346)	(0.148)	(0.080)	(0.179)
Average Employment Rate	0.800	0.914	0.935	0.458	0.800	0.914	0.935	0.458
Elasticity	0.283	-0.009	-0.050	-0.087	-0.500	0.048	-0.056	-0.208
R^2	0.305	0.532	0.600	0.354	0.688	0.865	0.928	0.862

Notes: A total of 94 observations are included. Standard errors are in parentheses. The square root of the sample variance in the employment is used as the weight. MW and W50 represent minimum wage and median wage, respectively.

Table 2B: How the minimum wage affected the employment rate.
Dependent variable: Employment rate
Sample: Females, High School Education or Less, 1997 and 2002

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Estimation Methods	WLS				FE			
Age Groups	≤22	23–30	31–59	≥60	≤22	23–30	31–59	≥60
ln(MW/W50)	0.183	0.128	0.225	-0.021	-0.543	-0.213	-0.220	-0.126
	(0.096)	(0.091)	(0.100)	(0.054)	(0.387)	(0.195)	(0.118)	(0.113)
Year 2002	-0.072	0.002	-0.036	-0.025	-0.032	0.019	-0.012	-0.020
	(0.013)	(0.013)	(0.013)	(0.007)	(0.022)	(0.011)	(0.007)	(0.006)
Constant	0.850	0.693	0.811	0.227	0.485	0.523	0.588	0.175
	(0.049)	(0.047)	(0.051)	(0.028)	(0.194)	(0.098)	(0.059)	(0.057)
Average Employment Rate	0.731	0.633	0.687	0.224	0.731	0.633	0.687	0.224
Elasticity	-0.250	-0.202	0.327	-0.092	-0.743	-0.336	-0.320	-0.561
R^2	0.249	0.028	0.087	0.145	0.611	0.117	0.717	0.777

Notes: A total of 94 observations are included. Standard errors are in parentheses. The square root of the sample variance in the employment is used as the weight. MW and W50 represent minimum wage and median wage, respectively.

Table 3: How the real value of the minimum wage would affect wage distribution, in the absence of disemployment.

Sample: Females, Counterfactual Constructed Sample without Employment Loss, 1994–2003

	(1)	(2)	(3)	(4)	(5)	(6)
Estimation Methods	OLS	OLS	2SLS	FE	FE	OLS
Dependent Variables	10/50 log Wage Differential					
ln(MW/W50)	–	0.40	0.39	0.56	0.59	0.42
		(0.04)	(0.04)	(0.09)	(0.06)	(0.03)
ln(W75/W50)	–	–	–	–	–	-0.48
						(0.08)
Year 1996	0.00	-0.00	-0.00	-0.01	-0.01	-0.00
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Year 1997	0.00	-0.01	-0.01	-0.01	-0.01	-0.01
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Year 1998	0.01	-0.01	-0.01	-0.01	-0.02	-0.01
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)
Year 1999	0.00	-0.02	-0.02	-0.02	-0.03	-0.02
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)
Year 2000	0.01	-0.01	-0.01	-0.02	-0.02	-0.01
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.00)
Year 2001	0.02	-0.01	-0.01	-0.02	-0.02	-0.01
	(0.00)	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)
Year 2002	0.02	-0.01	-0.01	-0.02	-0.03	-0.00
	(0.00)	(0.01)	(0.01)	(0.01)	(0.00)	(0.00)
Year 2003	0.03	-0.01	-0.01	-0.02	-0.02	-0.00
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)
Constant	-0.35	-0.14	-0.14	-0.10	-0.01	0.00
	(0.00)	(0.02)	(0.02)	(0.04)	(0.00)	(0.03)
Prefecture trends	No	No	No	No	Yes	No
R^2	0.06	0.59	0.59	0.92	0.96	0.72

Notes: A total of 423 observations are included. Standard errors in parentheses are clustered at the prefecture level. MW, W50, and W75 represent minimum wage, median wage and 75th percentile wage, respectively. The base year is 1994. Instrumental variables are the *meyasu* minimum wage and the median of the log wage within a prefecture between 1994 and 2003. The first-stage *F*-statistic is 98,275.

Table 4: How the real value of the minimum wage would affect wage distribution, without changes in workforce composition.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Estimation Methods	OLS	2SLS	FE	FE	OLS	OLS	2SLS	FE	FE	OLS
Dependent Variables	10/50 log Wage Differential					90/50 log Wage Differential				
Sample	A. Males									
ln(MW/W50)	0.28	0.27	0.46	0.42	0.45	0.44	0.45	0.30	0.20	0.26
	(0.05)	(0.06)	(0.05)	(0.04)	(0.06)	(0.10)	(0.11)	(0.05)	(0.04)	(0.06)
R^2	0.36	–	0.88	0.67	0.92	0.41	–	0.93	0.88	0.95
Sample	B. Females									
ln(MW/W50)	0.46	0.45	0.63	0.49	0.65	0.20	0.22	0.03	0.11	-0.12
	(0.04)	(0.04)	(0.07)	(0.03)	(0.04)	(0.11)	(0.11)	(0.09)	(0.05)	(0.14)
R^2	0.64	–	0.92	0.76	0.96	0.07	–	0.81	0.82	0.87

Notes: A total of 470 observations are included. Standard errors in parentheses are clustered at the prefecture level. MW, W50, and W75 represent minimum wage, median wage, and 75th percentile wage, respectively. Other covariates include year dummies in all columns, prefecture-specific trends in columns 4 and 9, and ln(W75/W50) in columns 5 and 10. Instrumental variables are the *meyasu* minimum wage and the median of the log wage within a prefecture between 1995 and 2003. First-stage *F*-statistics are 65,548 and 68,858 for males and females, respectively.

Table 5: How the minimum wage would affect the employment rate, without changes in workforce composition.
Dependent variable: Employment rate
Sample: High school graduates or less, 1997 and 2002

	(1)	(2)	(3)	(4)
Estimation Methods	FE			
Age Groups	≤22	23–30	31–59	≥60
Sample	A. Males			
ln(MW/W50)	-0.248	-0.021	-0.088	-0.074
	(0.323)	(0.113)	(0.060)	(0.153)
Average Employment Rate	0.798	0.915	0.935	0.443
Elasticity	-0.310	-0.023	-0.095	-0.167
R^2	0.637	0.858	0.926	0.915
Sample	B. Females			
ln(MW/W50)	-0.271	-0.161	-0.164	-0.267
	(0.296)	(0.144)	(0.080)	(0.611)
Average Employment Rate	0.727	0.632	0.686	0.133
Elasticity	-0.374	-0.255	-0.240	-2.002
R^2	0.627	0.064	0.765	0.622

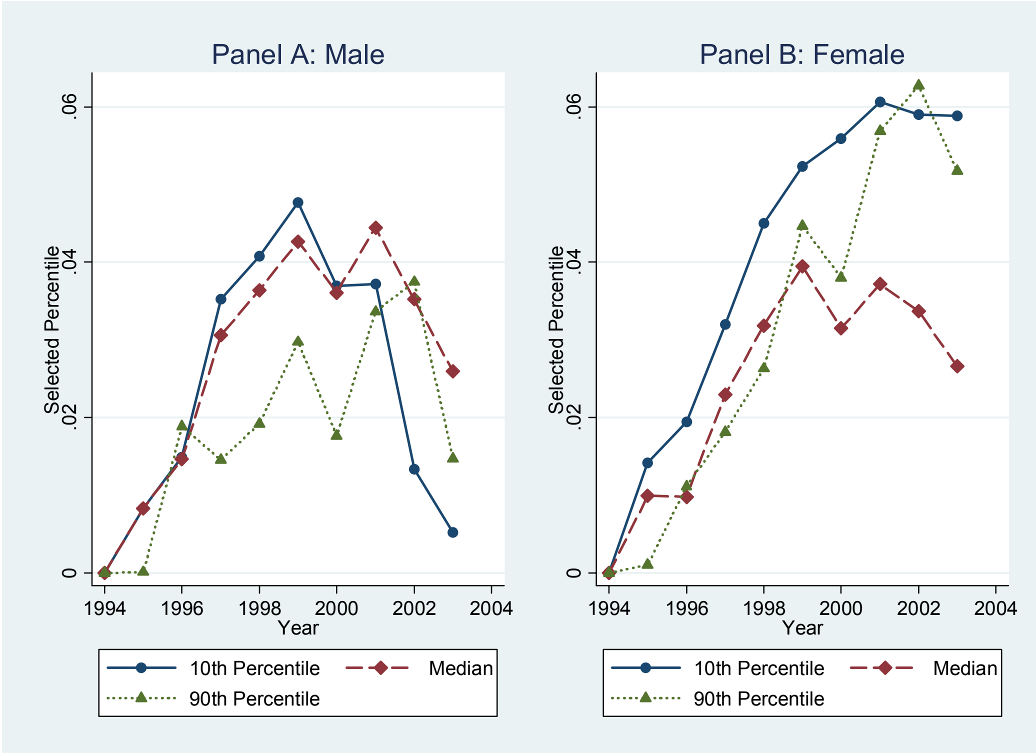
Notes: A total of 94 observations are included. Standard errors are in parentheses. The square root of the sample variance in the employment is used as the weight. MW and W50 represent minimum wage and median wage, respectively. Other covariates include year dummies in all columns.

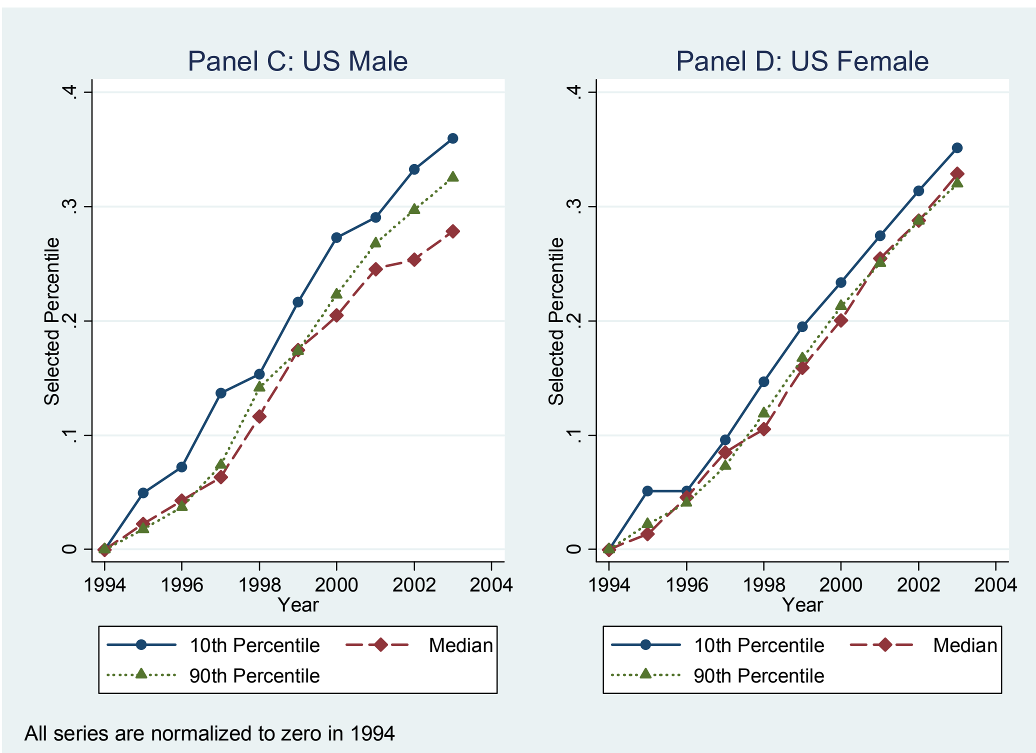
Table 6: Actual and counterfactual pay gaps between full-time and part-time female workers.

	(1)	(2)	(3)	(4)	(5)	(6)
Sample	1994 Actual		2003 Actual		2003 Counterfactual	
	Full-time	Part-time	Full-time	Part-time	Full-time	Part-time
log Wage	7.05	6.69	7.14	6.75	7.13	6.73
	(0.0007)	(0.0008)	(0.0008)	(0.0007)	(0.0008)	(0.0008)
log Wage Differentials	0.36		0.38		0.40	
	(0.001)		(0.001)		(0.001)	
Observations	384801	105210	283943	133475	283943	133475
		[21.5%]		[32.0%]		[32.0%]

Notes: Standard errors are in parentheses. The proportion of part-time workers is in square brackets.

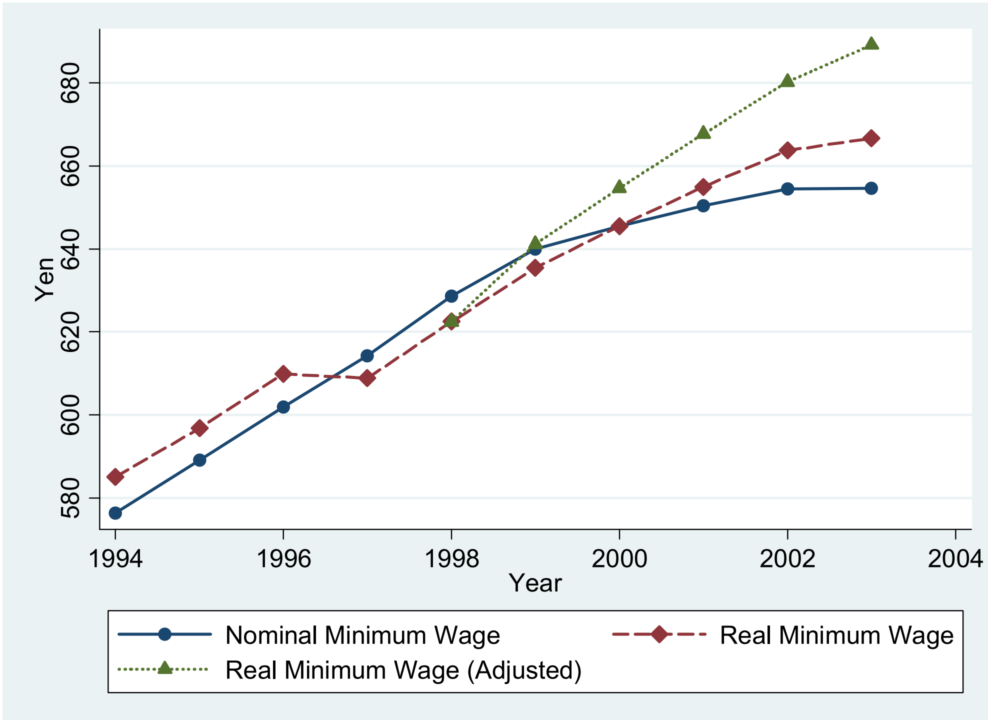
Figure 1: Trends in selected percentiles of the log hourly wage.





Notes: Data about American workers are taken from the Merged Outgoing Rotation Groups 1994–2003. Following Feenberg and Roth's (2007) recommendation, the hourly wage is calculated by "earnwke" divided by "uhourse".

Figure 2: Nominal and real minimum wages.



Notes: Minimum wages are weighted averages of regional minimum wages. The weight is the number of workers in the BSWS. The real minimum wage is calculated by dividing the nominal minimum wage by the CPI. The adjusted real minimum wage between 1999 and 2004 is calculated using the CPI corrected according to the results of Broda and Weinstein (2007).

Figure 3: The ratio of the minimum wage to the median wage by prefecture in 1994 and 2003.

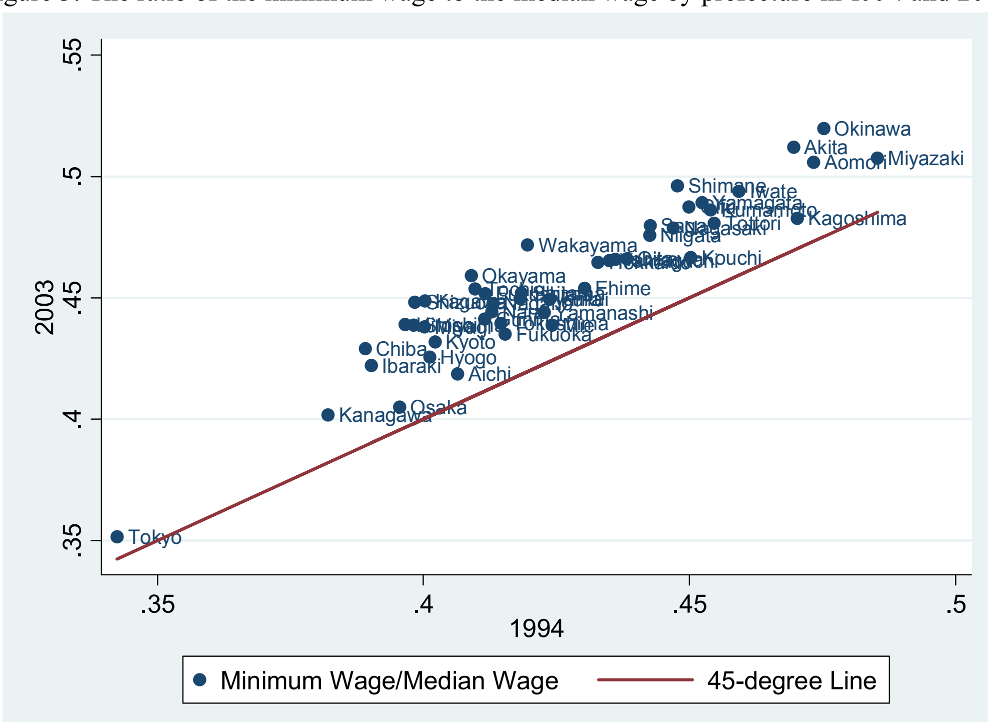


Figure 4A: Male log wage distribution by selected prefecture and year.

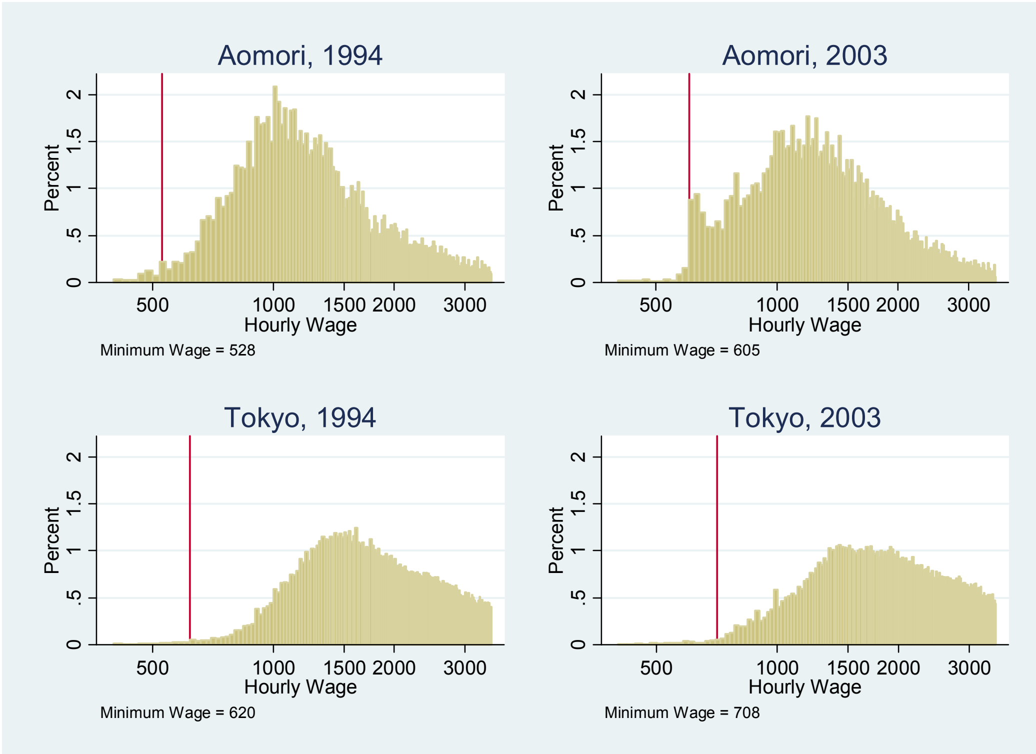


Figure 4B: Female log wage distribution by selected prefecture and year.

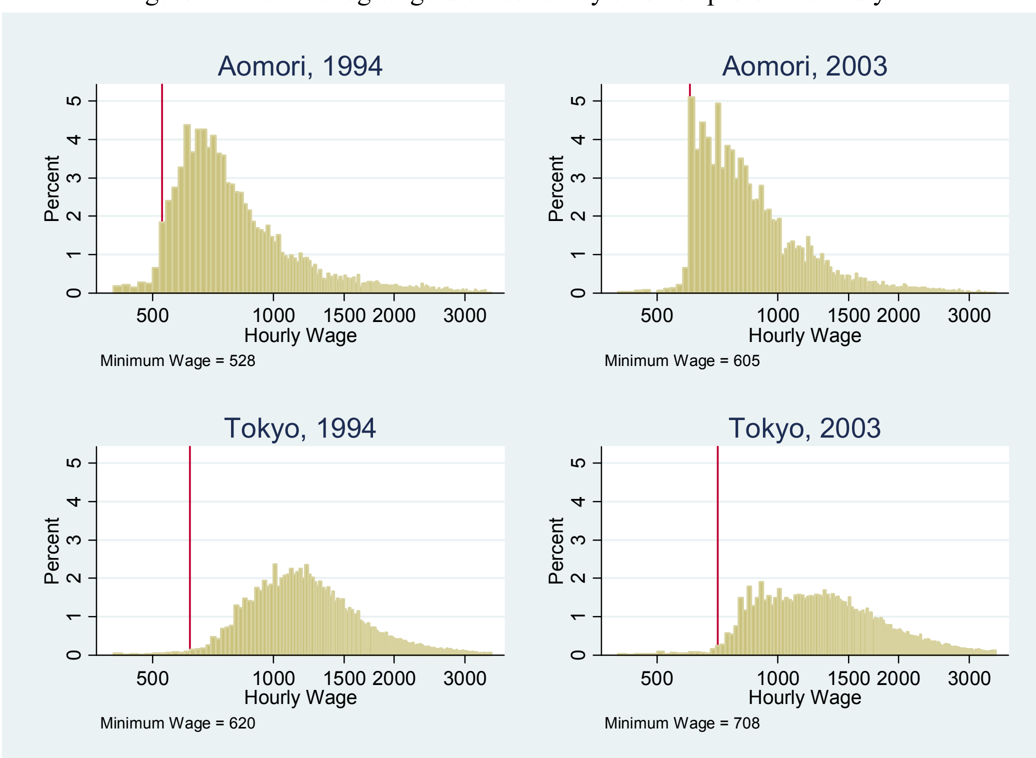


Figure 5: Wage compression and minimum wage.

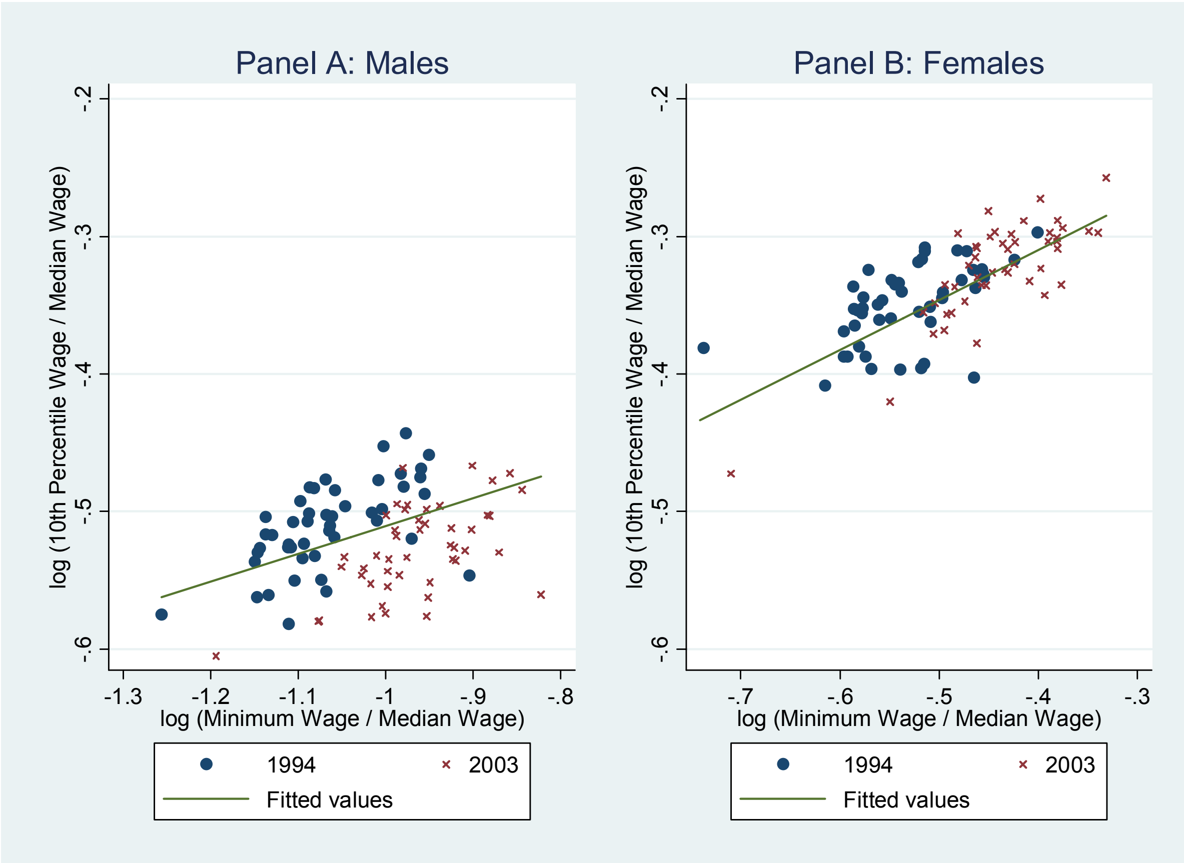


Figure 6: Actual and counterfactual log wage distributions.

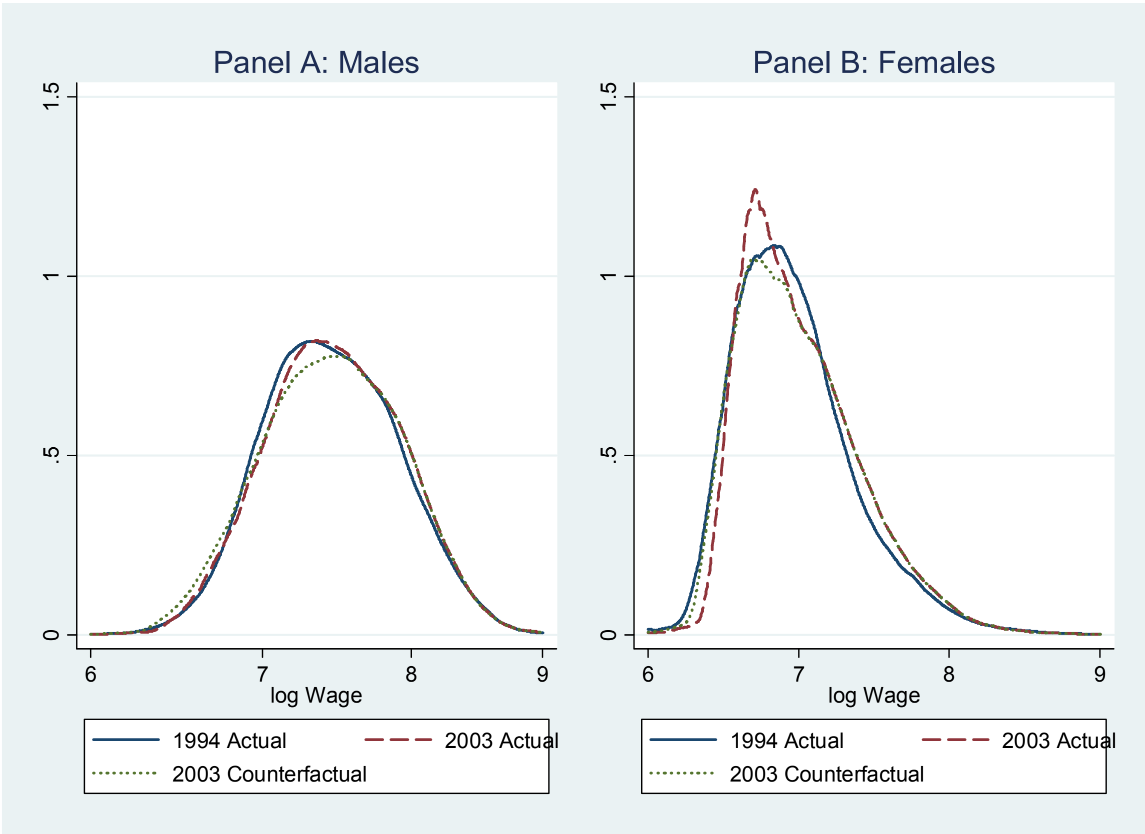


Figure 7: Actual and counterfactual changes in the log hourly wage by percentile, 1994–2003.

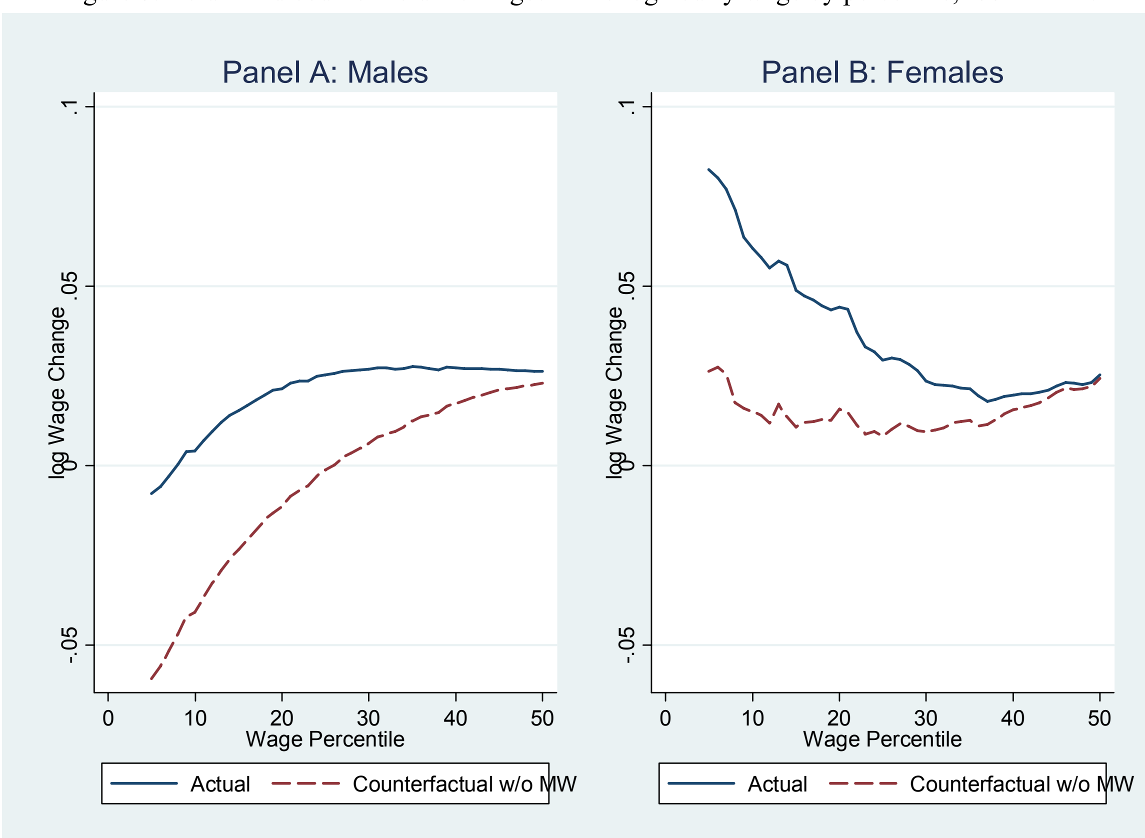


Figure 8: Female full-time/part-time log wage differential.