IZA DP No. 4949

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May 2010

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ABSTRACT

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The statutory minimum wage has steadily increased for decades in Japan, while the median wage has fallen nominally since 1999 because of a severe recession. We use large micro-data sets from two government surveys to investigate how the minimum wage has affected the wage distribution under unusual circumstances of deflation. The compression of the lower tail of the female wage distribution is largely explained by an increased real value of the minimum wage. 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.

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1 Introduction

Wage distributions have evolved differently among advanced industrialized countries, although these countries have shared similar experiences of rapid technological progress and increased exposure to international trade and outsourcing. In particular, the Japanese wage distribution has remained stable relative to those of Anglo-Saxon countries. The 90/50 and 10/50 log-wage differentials exhibit no particular trend for male workers from 1994 to 2003 (Figure 1, Panel A), which may be at odds with a recent polarization in the U.K. and U.S. labor markets, in which employment in high-skilled and low-skilled jobs has expanded at the expense of medium-skilled jobs (Goos and Manning, 2007; Autor, Katz, Kearney, 2008). Nonetheless, both the 10th and 90th percentile wages have increased compared to the 50th percentile wage for female workers (Figure 1, Panel B). This trend implies dispersion at the upper tail and compression at the lower tail of the female wage distribution.

While compression occurred at the lower tail of the female wage distribution, the nominal minimum wage has steadily increased for institutional reasons, despite economic downturns (Figure 2). The median wage has fallen nominally since 1999 because of a severe recession. During price and wage deflation, the rise in the minimum wage was higher in real terms than in nominal terms. Because revisions of the statutory minimum wage tend to lag behind general price inflation or deflation, the real value of the 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 fact, the real minimum wage increased by a weighted average of 0.16 log points across 47 prefectures between 1994 and 2003.

Our hypothesis is that an increase in the real value of the minimum wage contributed to the compression of the wage distribution among low-skilled workers. There is mixed evidence on the importance of the minimum wage among labor-market institutions as a determinant of the evolution of the wage distribution. In the United States, DiNardo, Fortin, and Lemieux (1996) demonstrated that erosion of the real minimum-wage level during the 1980s contributed to wage dispersion. Lee (1999) found that erosion of the real value of the minimum wage caused by general price inflation almost completely explained wage dispersion over the corresponding period. Autor, Manning, and Smith (2008) confirmed that the minimum wage plays a certain 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. Studies conducted in the United Kingdom have reported that the introduction of the British national minimum wage in 1999 did not contribute much 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 small (Dickens and Manning, 2004a, 2004b). 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. German minimum wages have been set by collective labor agreements between labor unions and firms, although no statutory minimum wage exists in Germany. Under such an institutional setting, deunionization leads to an erosion of minimum wages. Some relevant evidence has also been found in Japan. Abe and Tanaka (2007) pointed out that the prefectural minimum wage contributed to a reduction in the wage gap between full- 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. These studies examined only the effect on the mean wage, however. Hori and Sakaguchi (2005) illustrated the wage distribution in 2003 by prefecture and industry separately for full- and part-time workers, but did not conduct a formal regression analysis for the relation between minimum wage and wage distribution.[1]

This study examines the evolution of the Japanese wage distribution under conditions of wage deflation, with a central focus on the compression in lower-tail inequality for female workers. We analyze an econometric model that explains standard measures of wage inequality in terms of the effective minimum wage. The 'effective' minimum wage can be measured by the distance between minimum wage and median wage in the distribution of log hourly wage. Our identification strategy is basically to exploit regional variation in the effective minimum wage over time, which is known as a difference-in-differences (or fixed-effects) approach. Because the statutory minimum wage and the nature of the wage distribution evolve differently across prefectures, the minimum wage effect on the wage distribution can be isolated from unobserved prefectural heterogeneity and common macroeconomic fluctuations. Our empirical model includes prefecture-specific linear time trends, as well as prefecture and time effects, to allow for possible changes in the dispersion of the latent wage

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[1]A spike around the minimum-wage level is somewhat obscure in their illustration for a couple of reasons. First, the sample was split into small subgroups. Second, the bin width chosen was so narrow that the distribution was ragged everywhere.

distribution. Instrumental variable methods are used to assess robustness against measurement errors and policy endogeneity. The kernel reweighting approach proposed by DiNardo, Fortin, and Lemieux (1996) is also employed to examine the alternative hypothesis that changes in workforce composition mechanically lowered lower-tail inequality. Using the regression coefficients, the counterfactual wage distribution without an increase in the effective minimum wage is created to quantify the contribution of the increased real minimum-wage level on the evolution of the wage distribution between 1994 and 2003.

Wage compression can occur via three channels: censoring, spillover, and truncation. Lee (1999) revealed the presence of a spillover effect, but his approach cannot exactly differentiate among the three effects. Autor, Manning, and Smith (2008) decomposed the total effect into censoring and spillover effects in the absence of the truncation effect. Their approach assumes that the latent wage distribution has a lognormal distribution in the absence of minimum wages and that introducing a minimum wage has no adverse effect on employment. If the minimum wage has a disemployment effect, however, as documented by Neumark and Wascher (2008), a possible truncation of the wage distribution because of the minimum-wage increase can mechanically reduce lower-tail inequality. In this study, we develop two methods to examine to what extent truncation can explain wage compression. Our first approach is a variant of the inverse probability method (Little and Rubin, 2002). The proposed methods do not require a distributional assumption about the latent wage distribution.

Our analysis reveals that an increase in the minimum wage significantly pushed up wages for low-skilled female workers. Compression of the lower tail of the female wage distribution is largely explained by the minimum wage's increased real value. The actual 10/50 log-wage differential stayed constant at around 0.51 between 1994 and 2003 for female workers, but would have diverged by about 0.05 log points if there had been no increase in the effective minimum wage. Composition effects might be the main driving force for widening female upper-tail inequality but not for reducing female low-tail wage inequality. Moreover, we find that the minimum-wage hike contributed to a reduction in the pay gap between full- and part-time workers by about 0.05 log points at the lower end of the wage distribution. Furthermore, the compression of the wage distribution's lower tail is not attributable to the mechanical effect associated with truncation of the wage distribution, although

a moderate adverse effect of minimum wage on employment is observed among middle-aged female workers.

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 quantify 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 relation between minimum wage and wage compression, using a counterfactual sample in the absence of disemployment. The last section presents our conclusions.

2   Statutory Minimum Wage in Japan

The Minimum Wages Law in Japan was enacted in 1959 and substantially revised in 1967. The current law defines two types of minimum wages: (1) regional minimum wages based on collective agreement and (2) prefectural and industrial minimum wages based on the research and deliberations of minimum-wage councils. The first type of minimum wages is premised on craft-wide or industry-wide bargaining by which minimum wages are agreed upon and extended to nonunionized workers within the same sector. Such bargaining, however, does not really exist under the Japanese enterprise union system; in practice, all minimum wages in Japan are currently of the second type.[2] The second type of minimum wages comprises prefectural minimum wages covering all workers and industrial minimum wages covering workers aged 18 to 65 in specific industries in particular prefectures. This study focuses on prefectural minimum wages because industrial minimum wages cover a small fraction of workers.[3]

Under the current system, prefectural minimum wages are revised every year in two steps. First, the central minimum-wage council classifies all 47 Japanese prefectures into four ranks by actual wage levels and the standard cost of living and suggests how much the minimum wage should be

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[2]Moreover, the first type was formally abolished in 2008.

[3]The report by Saitei Chingin no Arikata Kenkyukai (minimum wage study group) of the Ministry of Health, Labor and Welfare recommended revising the industrial minimum wage in 2005, including possibly abolishing it. In fiscal year 2000, 4.5 million workers were covered by the industry minimum wage, while 52 million workers were covered by the regional minimum wage, according to a press release of the Ministry of Health, Labor and Welfare on January 25 in 2001.

increased ("meyasu") for each rank. The central minimum-wage council is not an ad-hoc but a *standing* institution administered by the Ministry of Health, Labor and Welfare. The council consists of representatives of public interest (academics and retired bureaucrats), employers, and employees, and its decisions have been insusceptible to politics. The minimum-wage policy has not been coordinated with any other policy, such as unemployment insurance and a measure to promote small and medium enterprises, which is different from minimum-wage systems in other countries. Second, the chief of the prefectural labor bureau determines the level of the prefectural minimum wage based on the regional minimum-wage councils' deliberations. The regional minimum-wage councils' deliberations are strongly influenced by the central minimum-wage council's suggestions for the increased amount of minimum wage.

The *meyasu* system was introduced in 1978 to moderate regional disparities in minimum wages. The classification of 47 prefectures into four ranks has hardly ever changed, however. Prefectural minimum wages have not been revised in a way that reflects local labor-market conditions for any particular prefecture. Prefectural minimum wages vary within the same rank because of the difference in the minimum-wage level before the *meyasu* system was introduced.[4] The current system in which the central minimum-wage council suggests the minimum-wage increase every year induced *inertia*. In fact, the real value of minimum wage increased by 50% between 1975 and 1995 (OECD, 1998, Chart 2.1). This trend continued even during a recession.

The minimum wage is legally enforceable in the following manner. The prefectural labor standards inspection office 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. Moreover, the minimum wage seems to be enforced largely 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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[4]The current minimum-wage level is approximately the same as the 1977 level plus the accumulated total of annual minimum-wage increases since 1978.

3 Data

Our analysis uses 1994–2003 micro data from the Basic Survey of Wage Structure (BSWS), which is compiled annually by the Ministry of Health, Labour and Welfare. 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*,[5] and establishments and individual files are then merged using an establishment identification number. Both full- 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. The available information includes each worker's wages, age, sex, educational attainment (only for full-time workers), full-/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*),[6] the number of new graduates hired, firm size, industry, and location.

Data about wages and hours include individuals' contracted pay, overtime pay, allowances (e.g., for family and transportation), contracted hours of work, and overtime hours between June 1 and June 30 over the corresponding period. Because the effective date of change in prefectural minimum wages is either September 30 or October 1 of each year, we merge statutory minimum-wage data in year $t-1$ with wage and hour data from the BSWS in year t. Japanese minimum wage laws apply to the straight wage rate, excluding allowances. We define hourly wage as (wages for contracted hours – commutation allowance – perfect attendance allowance – family allowance)/contracted hours of work,

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[5]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.

[6]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.

which is consistent with the minimum-wage law.[7] A change in the minimum wage conceivably may affect the level of allowances. Results obtained in our analysis are unchanged, however, even when hourly wage is defined as wages (including allowances) for contracted hours divided by contracted hours of work.

The analysis on how the minimum wage affects employment also includes data from a household survey that covers non-employed as well as employed individuals. We use 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.[8] 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. This study draws on micro data about employment status, educational attainment, age, sex, and residential area. Overall, the sample includes approximately 1 million individuals, with a half-million males and a half-million females for each year that the survey was conducted. The sample is restricted to data with valid age, educational background, and employment status.

4 Role of the Minimum Wage

4.1 Evolution of the Wage Distribution

The steady increase in the statutory minimum wage occurred in a political process that induces nearly automatic increases in minimum-wage levels. All prefectural minimum wages rose in the political climate for the equalization of minimum-wage levels across prefectures, despite serious slumps in rural areas. The minimum-wage bite indeed became severe in low-wage prefectures in the prolonged recession. Figures 3A and 3B illustrate the wage distribution in Aomori and Tokyo in 1994 and 2003 for male and female workers with hourly wages between 400 and 3,500 Japanese yen. Aomori, which is a prefecture located at the north end of Japan's main island, is classified as Rank D (with the lowest minimum wage), while Tokyo is classified as Rank A (with the highest minimum wage). The wage

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

[8]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.

distribution moved dramatically toward the lower end in Aomori from 1994 to 2003. A moderate 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 in Aomori. Surprisingly, the wage density is highest at the minimum-wage level for female workers in 2003. As indicated by the density of hourly wages below 1,000 Japanese yen, the proportion of low-wage workers also increased for both males and females in Tokyo, although not to the same extent as in Aomori. The minimum wage did not seem to bind either male or female workers in Tokyo over the sample period.

Figure 4 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 these prefectures 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 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. Variation in the effective minimum wage is exploited to identify how the minimum wage affected the wage distribution and employment in the subsequent regression analysis.

Figure 5 illustrates the relation 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 in both panels A and B, and it is greater in panel A for male workers than in 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 for 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. An 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

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, 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 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 expected wage for unemployed workers (Manning, 2003). Finally, an increased minimum wage may result in a truncation of the wage distribution associated with disemployment. The disappearance of the bottom end of the wage distribution can change the distance between its 10th and 50th percentiles.

We conduct a regression analysis to investigate the cause for wage compression. Our regression framework is based on Lee (1999).

$$\widetilde{w}_{it}^{p} = \beta_{1p}\widetilde{mw}_{it} + \beta_{2p}\widetilde{mw}_{it}^{2} + d_t\gamma + d_i\delta + t_i\zeta + u_{it}, \tag{1}$$

where $\widetilde{w}_{it}^{p} = \ln\left(w_{it}^{p}/w_{it}^{50}\right)$ and $\ln\widetilde{mw}_{it} = \ln\left(mw_{it}/w_{it}^{50}\right)$.[9] The variable w_{it}^{p} is the pth percentile of the wage distribution, mw_{it} is the minimum wage, d_t is a vector of year dummies, d_i is a vector of prefecture dummies, t_i is a vector of prefecture-specific linear time trends, i is the index for prefecture, and t is the index for year. The minimum-wage bite is measured by the log of the minimum wage relative to the median wage. Parameter β_{1p} 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, if no higher-order term of the effective minimum wage is included. The quadratic term is included as an additional regressor to capture a nonlinear relation between the minimum wage and wage compression, as illustrated by Lee (1999). Year effects represent the evolution of the wage distribution over time, the real minimum-wage level being constant. Prefectural fixed effects (FE) are added to allow for unobserved heterogeneity in the dispersion of the latent wage distribution across prefectures. Prefecture trends are added to control for macroeconomic trends more flexibly.

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

Our main interest is in accounting for the reduction in the lower-tail inequality for female workers. Table 1 lists the results for the 10/50 log-wage differential among female workers. Estimated results are produced using ordinary least-squares (OLS) and two-stage least squares (2SLS), and standard errors are clustered at the prefecture level. The analysis begins 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 shows the increasing trend of the 10th percentile wage relative to the median wage. Next, we include the effective minimum wage as an additional regressor. Column 2 indicates that the coefficient for the effective minimum wage is positive and significant. R^2 rises from 0.07 to 0.56 after the effective minimum wage is added. 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. Moreover, the estimated year effects are virtually zero or sometimes negative, conditional on the effective minimum wage. These results suggest that wage compression can be explained entirely by increases in the effective minimum wage over the sample period.

OLS may suffer from an upward bias associated with measurement errors, however, as Autor, Manning, and Smith (2008) emphasized. Because the median wage appears on 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. Moreover, a bias can also arise from policy endogeneity. Although the prefectural authorities generally confirm the increased amount of minimum-wage suggested by the central minimum-wage council, the revised value of prefectural minimum wage is sometimes not exactly the same as that indicated by this council. The direction of the bias is upward (downward) if a large increase in the minimum wage is permitted by the prefectural authority that has a small (large) lower-tail inequality. Existing literature tends to ignore policy endogeneity. Nonetheless, we can address econometric concerns about policy endogeneity, as well as measurement errors, by using the Japanese minimum-wage system. Specifically, as the instrument for the effective minimum wage, we use the *meyasu* minimum wage, which is calculated by the minimum wage in the previous year plus "meyasu". The median of the log wage within a prefecture over the sample period is also used as the instrument, as in Autor, Manning, and Smith (2008). As shown in column 3, the

2SLS estimates are 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 resulting from sampling errors or policy endogeneity is negligible in our analysis.

A key concern is a spurious correlation arising from unobserved heterogeneity in the dispersion of the latent wage distribution across prefectures. In other words, the important assumption required to identify the minimum wage effect is no correlation between the effective minimum wage and the dispersion of the latent wage distribution across prefectures. In the absence of a minimum wage, the 10/50 log-wage differential should be negative and lower in a prefecture with a higher level of wage dispersion. If the effective minimum wage is positively correlated with the dispersion of the latent wage distribution, the effect of the minimum wage on the 10/50 log-wage differential will be biased downward. In contrast, the effect on the 90/50 log-wage differential will be biased upward, because the sign of the correlation with the wage dispersion is reversed. A standard solution to this problem is to control for prefecture fixed effects. Although the fixed-effects approach is generally vulnerable to measurement errors, the 2SLS results suggest that such a concern is unwarranted here. Column 4 shows that the minimum wage had a greater effect on wage compression when we controlled for prefecture effects. This result can be interpreted as correcting for omitted-variable bias. Prefecture-specific linear trends are added to allow for possible changes in the dispersion of the latent wage distribution. Column 5 shows a slightly larger minimum wage effect. Furthermore, column 6 exhibits a non-linear relation between minimum wage and wage compression.

One way to assess the validity of our analysis is to examine the effect on upper-tail inequality. Columns 7–12 report the results of the 90/50 log-wage differential. Column 7 suggests that the 90/50 log-wage differential was almost stable during the sample period. Column 8 shows a positive and significant effect of minimum wage, but R^2 increases by only 6 percentage points. The effective minimum wage does not significantly explain the 90/50 wage differential. OLS estimates are identical to 2SLS estimates and similar to fixed-effects estimates. Importantly, the effect of the minimum wage plummets and becomes statistically nonsignificant after controlling for prefecture trends. Column 12 shows that a higher-order term is not significant, either. Overall, our findings suggest that 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.

For the sake of completeness, Table A1 lists the results for the 90/50 log-wage differential among male workers. Column 1 suggests that the 10/50 log-wage differential was almost stable between 1994 and 2000 and started to erode slightly thereafter. Column 2 shows a positive and significant effect of minimum wage and R^2 rises from 0.06 to 0.38. Coefficients for year dummies shrink when the effective minimum wage is held constant. The results imply that the 10/50 log-wage differential would have diverged if the real minimum-wage level remained unchanged. 2SLS estimates are identical to OLS estimates. The effective minimum wage has a stronger effect after controlling for prefectural fixed effects and prefecture-specific linear trends. The minimum wage, however, played a less significant role in pushing up the lower tail of the wage distribution for male workers in terms of the magnitude of estimated coefficients and R^2. Columns 7–12 report the results of the 90/50 log-wage differential. Column 7 suggests a slight decline in the male upper-tail inequality during the sample period. Column 8 shows that, contrary to our expectation, the effective minimum wage had a positive and significant effect on the 90/50 log-wage differential and R^2 rises from 0.04 to 0.40. Again, 2SLS estimates are identical to OLS estimates. The minimum wage effect becomes smaller, but still remains after controlling for prefecture effects and its interaction terms with a linear time trend. That an increase in the minimum wage would push up the 90th percentile wage is not very compelling. These results indicate that the effective minimum wage might be higher in prefectures where the male latent distribution is more dispersed. Thus, the estimate for the effect of the minimum wage on male lower-tail inequality may be biased downward.

4.3 Minimum Wage Effects or Composition Effects?

In Japan, the labor force has been aging, and job tenure increased for female workers in the period between 1994 and 2003. In fact, these attributes are key determinants of wages in the Japanese labor market. Shifts in workforce composition may have mechanically raised or lowered wage inequality. To isolate the minimum wage effect from the composition effect, we employ the kernel reweighting approach proposed by DiNardo, Fortin, and Lemieux (1996, hereafter DFL).

The observed density of log hourly wages in year t is expressed as

$$f\left(w\middle|\tau=t\right)=\int g\left(w\middle|x,\tau=t\right)h\left(x\middle|\tau=t\right)dx, \tag{2}$$

where $g\left(w\middle|x,\tau=t\right)$ is the density of wages for workers' observed attributes x in year t, and $h\left(x\middle|\tau=t\right)$ is the density of attributes x in year t. As shown by DFL, the counterfactual density in year t if the observed attributes are fixed at 1994 levels can be written as

$$f_{1994}\left(w\middle|\tau=t\right)=\int g\left(w\middle|x,\tau=t\right)h\left(x\middle|\tau=1994\right)dx \tag{3}$$

$$=\int g\left(w\middle|x,\tau=t\right)\psi_t\left(x\right)h\left(x\middle|\tau=t\right)dx. \tag{4}$$

where $\psi_t\left(x\right)=h\left(x\middle|\tau=1994\right)/h\left(x\middle|\tau=t\right).$ Thus, calculating the counterfactual density requires reweighting a price function $g\left(w\middle|x,\tau=t\right)$ by the ratio of the two composite functions $h\left(x\middle|\tau=1994\right)$ and $h\left(x\middle|\tau=t\right).$ It is difficult to estimate the composite function, however, because x is high-dimensional. By applying Bayes's rule, the reweighting function can be written as

$$\psi_t\left(x\right)=\frac{\Pr\left(\tau=1994\middle|x\right)}{\Pr\left(\tau=t\middle|x\right)}\cdot\frac{\Pr\left(\tau=t\right)}{\Pr\left(\tau=1994\right)}. \tag{5}$$

The reweighting function can be estimated using a logit model applied to the pooled data from years 1994 and t. The attributes x include a full set of dummy variables for age and job tenure.[10]

After reweighting the data in a way that holds the distribution of skills constant over time, we reexamine Lee's (1999) model of wage compression. Panel A in Table 2 displays results in selected columns of Table 1, while panel B in Table 2 shows the results using the counterfactual wage data. The minimum wage effect decreases slightly, but remains significant for female lower-tail (10/50) inequality. The effect on female upper-tail (90/50) inequality becomes smaller and nonsignificant in column 4 and negative and significant in column 5. The former suggests that wage compression can be largely explained by the minimum-wage hike. The latter implies that the expansion of female upper-tail inequality can be attributed to aging and lengthening job tenure in the labor force. In effect,

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[10]Educational information is not available for part-time workers in BSWS.

composition effects may be large at the upper tail of the wage distribution, but they cannot account for most of the reduction in lower-tail inequality.

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

Increases in the real value of the minimum wage contributed to a compression between the 10th and 50th percentiles of the wage distribution from 1994 to 2003, especially among female workers, as described thus far. If the effective minimum wage had remained unchanged over the 10-year period, wage compression might not have occurred. Following Lee (1999), we construct a counterfactual wage distribution without any increase in the effective minimum wage to quantify the relation between the minimum wage and wage compression in more detail. The counterfactual wage in 2003 is 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 is simulated as follows:

$$\widehat{w}^p_{k,i,2003} = w^p_{k,i,2003} - \widehat{\beta}_{1p} \left(\widetilde{mw}_{i,2003} - \widetilde{mw}_{i,1994} \right) - \widehat{\beta}_{2p} \left(\widetilde{mw}^2_{i,2003} - \widetilde{mw}^2_{i,1994} \right), \qquad (6)$$

where $\widehat{\beta}_{1p}$ and $\widehat{\beta}_{2p}$ are estimated coefficients obtained from the regression of the percentile wage differential on the effective minimum wage, its square, year dummies, prefecture dummies, and prefecture-specific time trends. The pth percentile varies with prefecture, year, and sex.

Our aim here is to examine the extent to which the minimum-wage bite can explain compression at the lower tail of the wage distribution. Figure 6 displays the actual and counterfactual wage distributions for female workers in 1994 and 2003. The lower tail of the actual wage distribution in 2003 is compressed for female workers. The compression is displayed by the spike at the lower tail of the wage distribution. The difference between the actual and counterfactual distributions in 2003 illustrates the effect of the minimum-wage increase. We find that the minimum-wage increase pushed up wages at the lower-tail of the distribution. 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 attributed mostly to the minimum-wage increase.

Figure 7 displays the actual and counterfactual changes in the log hourly wage by percentile between 1994 and 2003. Lower percentile wages actually increased by 0.02 to 0.08 log points for female workers between 1993 and 2004, whereas the counterfactual changes without an increase in the effective minimum wage are close to 0.01 log point from the 10th to 35th wage percentiles. Thus, the rise in actual lower percentile wages can be largely attributed to the minimum-wage hike. The difference between the actual and counterfactual wage changes indicates a significant spillover effect on workers who earn more than the minimum wage. It may appear that the minimum wage contributed to widening the upper-tail inequality slightly, but the effects are statistically nonsignificant, as seen in column 12 of Table 1.

4.5   Part-time Pay Penalty

An increasing trend of hiring part-time workers is a global phenomenon. For example, about 45% of female workers in Britain work part-time (Manning and Petrongolo, 2008). Along with this trend, the full-time/part-time wage differential has been of interest in recent years. In Japan, the fraction of part-time workers in the workforce increased from 21.5 to 32.0 percent among female workers between 1994 and 2003. Indeed, the full-time/part-time wage differential has been a hotly debated policy issue.

The minimum wage's effect on wage compression has an implication for the part-time penalty, i.e., the pay gap between full- and part-time workers. Employees who are paid the minimum wage are typically part-time workers. Thus, the minimum-wage hike can cause a reduction in 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 British national minimum wage was introduced. Abe and Tanaka (2007) found that having a minimum wage prevented wage erosion among part-time workers relative to full-time workers in Japan.

We directly quantify how the minimum wage affected the pay gap between full- and part-time

workers using the counterfactual wage distribution without an increase in the effective minimum wage. The analysis focuses on female workers, because the proportion of male part-time workers was very small.[11] Table 3 reports the actual and counterfactual pay gaps between full- and part-time workers. The actual pay gap was 36.2 percent in 1994 and increased to 38.2 percent in 2003. The pay gap would have been 39.1 percent in 2003, however, 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 1 percentage point at the mean.

Figure 8 illustrates the full-time/part-time log-wage differential by wage percentile. The pay gap between full- 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. The pay gap would have expanded, however, if no increase in the minimum wage had occurred. The minimum wage had a greater effect at the lower tail of the wage distribution. The pay gap without the minimum-wage increase, for example, would have expanded by about 5 percentage points at the 25th percentile.

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. 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; Kawaguchi and Mori, 2009) but no consensus has been reached, and the cross-sectional analyses conducted by Tachibanaki and Urakawa (2007) and Abe and Tamada (2008) controlled for neither prefecture nor year effects.

To examine how the minimum wage affected employment, we conduct a standard pseudo-panel

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

data analysis, as put 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

$$\ln\left(\frac{emp_{jit}}{pop_{jit}}\right) = \rho_j \widetilde{mw}_{it} + d_t\gamma_j + d_i\delta_j + u_{jit}, \tag{7}$$

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.[12] Parameter ρ represents the wage elasticity of labor demand for minimum-wage workers. If parameter ρ_j is negative, an increase in the minimum wage reduces the employment rate for group j. A common macroeconomic shock is flexibly captured by year dummies. The results obtained in this paper change only marginally, when the employment rate for male college graduates aged between 31 and 59 is included as an additional regressor to enable further control of aggregate fluctuations in employment. Prefecture dummies are included to allow for an unobserved prefecture effect. In light of Card and Krueger's (1995) criticism, we do not include the college enrollment rate as a regressor. Thus, regressors include only exogenous variables in our preferred specification of labor demand.

The employment rate is calculated from ESS data; these surveys were conducted only in years 1997 and 2002 during the sample period of 1994–2003. The effective minimum wage is calculated from the BSWS, as in the previous analysis.

Our analysis focuses 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. The model is estimated using the fixed-effects approach. Table 4 reports how the minimum wage affected female employment by age group.[13] The estimated year effect is negative for all age groups, indicating a decline in the labor-market attachment. The minimum wage effect is negative but nonsignificant, except for females aged 31–59 years. However, column 3 reveals a moderate and significant disemployment effect for females

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

[13]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 nonsignificant for all demographic groups after controlling for prefectural fixed effects. These results are not surprising, because decreasing the number of employees can reduce firms' labor costs more effectively than reducing the number of hours of work when fixed costs arise in employment. Most previous studies have focused on the effect on employment.

aged 31–59 years. 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. For the sake of completeness, Table A2 reports the results of the disemployment effect by age group among male workers. The disemployment effect is not statistically significant for any age group.

5.2 Effect on New Hires

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.[14] Estimated employment elasticity is imprecise but greatest for those aged 22 or younger, as shown in column 1 of Tables 4 and A2. 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. Nonetheless, very few studies to date, with the exception of Portugal and Cardoso (2006), have explored the effect on worker flow.

We examine 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 (from schools, colleges, and universities) hired by each firm is available from the BSWS for every year between 1994 and 2003. The results for female new graduates are outlined below.

$$\ln\left(\frac{newhire_{it}}{employee_{it}}\right) = \underset{(0.58)}{-1.82}\widetilde{mw}_{it} + d_t\widehat{\gamma} + d_i\widehat{\delta} + t_i\widehat{\zeta} + \widehat{u}_{it}, \qquad R^2 = 0.91,$$

where $newhire$ is the number of new graduates hired, $employee$ is the number of regular workers, and the last four terms are prefecture dummies, year dummies, prefecture-specific linear time trends, and the residual, respectively. The model is estimated using the fixed-effects approach. The hat represents the estimate. A total of 470 observations are included. Standard errors in parentheses are clustered at the prefecture level.

As shown above, a one-percent increase in the effective minimum wage leads to a 1.82-percentage decrease in the ratio of new graduates hired among female employees. For male new graduates,

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[14]In most cases, Japanese employment regulations were not put into statutory form, but were established by court precedents (Sugeno 2002). The "abuse of dismissal rights" doctrine, however, was legislated in Article 16 of the Labor Contract Act in 2008.

however, the effect on new hires is statistically nonsignificant and smaller than that for female new graduates. The estimated coefficient on the effective minimum wage is –0.31 with a standard error of 0.81, and R^2 is 0.84. 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.

5.3 Removing the Truncation Effect

Up to this point, the results have confirmed that an increase in the effective minimum wage compressed the lower tail of the wage distribution but reduced employment for low-skilled female workers. We are now concerned about a mechanism of wage compression. Our analysis suggests that the minimum-wage hike caused employment loss among low-skilled female workers. The lower tail of the wage distribution might be mechanically compressed by the truncation of the bottom end of the wage distribution associated with disemployment. Indeed, 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.[15] Thus, we investigate to what extent wage compression can be explained by employment loss resulting from the minimum-wage increase.

The problem can be described as a sample selection problem. The observed density of log hourly wages conditional on the effective minimum wage $\widetilde{mw}$ and workers' observed attributes x can be written as $f\left(w_t | \widetilde{mw}_t, x_t\right)$ in year t, while the counterfactual density if the effective minimum wage had remained at the 1994 level can be expressed as $f\left(w_t | \widetilde{mw}_{1994}, x_t\right)$. The actual wage density can be influenced by the truncation of the wage distribution caused by the minimum-wage increase between 1994 and 2003, whereas the counterfactual density cannot. Let e_t denote an indicator variable for being employed. The following assumptions are required to quantify the truncation effect.

Assumptions. (a) The minimum wage has neither a spillover effect nor a censoring effect, i.e., $g\left(w_t | e_t = 1, \widetilde{mw}_t, x_t\right) = g\left(w_t | e_t = 1, \widetilde{mw}_{1994}, x_t\right)$; (b) The minimum wage has no direct effect on

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

the distribution of individual attributes, i.e., $h\left(\left. x_t \right| \widetilde{mw}_t\right) = h\left(\left. x_t \right| \widetilde{mw}_{1994}\right)$.

Under the two assumptions, the minimum wage affects the wage distribution solely through the disemployment effect. These assumptions are extreme, but suitable for our purpose, which is to examine the extent to which truncation can explain the compression of lower-tail inequality. As derived in the Appendix, the counterfactual density can be nonparametrically constructed as follows:

$$f\left(\left. w_t \right| e_t = 1, \widetilde{mw}_{1994}\right) = \int g\left(\left. w_t \right| e_t = 1, \widetilde{mw}_t, x_t\right) \theta\left(\widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right) h\left(\left. x_t \right| e_t = 1, \widetilde{mw}_t\right) dx,$$

(8)

where

$$\theta\left(\widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right) = \frac{\Pr\left(\left. e_t = 1 \right| \widetilde{mw}_{1994}, x_t\right)}{\Pr\left(\left. e_t = 1 \right| \widetilde{mw}_t, x_t\right)} \cdot \frac{\Pr\left(\left. e_t = 1 \right| \widetilde{mw}_t\right)}{\Pr\left(\left. e_t = 1 \right| \widetilde{mw}_{1994}\right)}.$$

The weighting function is essentially the ratio of the employment rate at the 1994 minimum-wage level to the employment rate at the current minimum-wage level. The weighting procedure is quite intuitive. To remove the truncation effect on the wage distribution, more weight is given to workers who would have a higher propensity to be employed without an increase in the effective minimum wage, i.e., to those who are more likely to disappear from the labor market because of the minimum-wage hike. An estimate of the response probability can be obtained from a probit model:

$$\Pr\left(\left. e_{kt} = 1 \right| \widetilde{mw}_{it}, x_{jt}\right) = \Phi\left(\pi_{j0} + \pi_{j1}\widetilde{mw}_{it} + d_i\delta_j + \zeta_j t\right).$$

(9)

where d_i is a vector of prefecture dummies, t is a linear time trend, k is an index for individual, j is an index for worker's attributes x. Specifically, workers are classified into four age groups (≤ 22, 23-30, 31-59, ≤ 60). All parameters are allowed to vary with group j. The model is estimated using the individual-level micro data from the ESS. Minimum-wage effects are negative and statistically significant for all age groups. The estimated coefficients π_{j1} (standard errors clustered at the prefecture level) are, in order of age, -0.133 (0.210), -0.127 (0.124), -0.283 (0.111), -0.115 (0.110). Using the regression coefficients, we can calculate the weight for every year over the sample period by virtue of linear time trends, although the ESS was collected only in years 1997 and 2002 during the sample period of 1994–2003. Higher-order terms of time trends cannot be identified from two-period data.

We omit an illustration of the counterfactual distribution in the absence of the disemployment

effect, because it almost fully overlaps with the actual distribution. Panel C in Table 2 shows the results that are reproduced using counterfactual data. Results are identical to those using the actual sample. Therefore, the truncation effect is negligible in our analysis of wage compression.

5.4   Upper Bound of the Truncation Effect

Employment loss may occur exclusively from the bottom end of the wage distribution for each skill group. In that case, our proposed method could underestimate the truncation effect. Thus, we alternatively propose a more conservative approach. Our idea is to add workers who are unemployed because of the minimum-wage increase at the lower end of the distribution. The alternative method is developed to quantify the upper bound of the mechanical effect.

The standard pseudo-panel model of employment implies that the change in the log employment rate caused by the minimum-wage increase between years $t-1$ and t can be expressed as $\Delta \ln \left(\frac{emp_{jit}}{pop_{jit}} \right) = \widehat{\rho}_j \Delta \widetilde{mw}_{it}$, where $\widehat{\rho}_j$ is the estimated coefficient for the effective minimum wage in the fixed-effect estimates of the employment equation for group j. Assuming that the population size is unchanged, the decrease in the number of employed can be expressed as

$$\Delta emp_{jit} = \widehat{\rho}_j \Delta \widetilde{mw}_{it} \cdot emp_{jit}. \tag{10}$$

The counterfactual wage distribution for group j in prefecture i in year t can be constructed in the following steps.

1. Substituting the actual change in the effective minimum wage yields the number of workers who lost their jobs by group and prefecture between years $t-1$ and t. Then, calculate the total number of unemployed workers between 1995 and 2003, $N_{it}^{add} = -\sum_j \min \{\Delta emp_{jit}, 0\}$.

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

21

The counterfactual scenario considered here is that workers who are laid off because of the minimum-wage increase could stay in the lowest-wage job. In the first step, we lose observations from the first year of the sample period in creating the counterfactual sample. In the second step, a value of zero is imputed to the log hourly wage for unemployed individuals to add them from the bottom end of the distribution. This imputation procedure is extreme but suitable for our purpose to remove the upper-bound truncation effect.

The results of the 10/50 log-wage differential in Table 1 are reproduced for the counterfactual sample in panel D of Table 2. Estimation results differ only marginally. Eliminating the upper bound of the truncation effect, our analysis provides the lower bound of the censoring and spillover effects. Indeed, the lower-bound estimates are smaller than baseline estimates, but the difference is minimal. The results suggest that a change in the effective minimum wage affects the wage distribution mostly through censoring and spillover. An alternative imputation procedure will not change the results if all imputed wages are lower than the 10th percentile wage. If some imputed wages are greater than 10th percentile wage, the truncation effect will be even smaller and the censoring and spillover effects will be larger.

6   Conclusions

This study has examined how the minimum wage affected the wage distribution between 1994 and 2003 in Japan, the world's second-largest economy. Japan's experience since the 1990s mirrors the U.S. experience in the 1980s and 1990s. The real value of the minimum wage substantially increased because of a fall in the median wage in a deflationary economy and because of a steady increase in the statutory minimum wage. As a consequence, the minimum-wage hike compressed the lower tail of the wage distribution in Japan, whereas a fall in the effective minimum wage resulted in 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. These results held even after controlling for composition effects. 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. We also found that the increase in the effective minimum wage decreased the full-time/part-time wage differential by 5 percentage points at the lower tail of the wage distribution among female workers. The minimum-wage hike reduced employment for low-skilled, middle-aged female workers. The disemployment effect was –0.31 in elasticity terms. We obtained similar results for wage compression, however, after recovering the wage distribution in the absence of disemployment. The reduction in the lower-tail inequality from female workers cannot be attributed to the effect of truncation.

To conclude, 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 worker flow for each skill group across industries. Second, the minimum-wage hike may affect college enrollment and occupational choices. Constructing a model of educational and occupational choices would be helpful to examine how the minimum wage affects complex individual choices.

Appendix

Derivation of equation (8).

The counterfactual wage density is

$$f\left(w_t|\, e_t = 1, \widetilde{mw}_{1994}\right) = \int g\left(w_t|\, e_t = 1, \widetilde{mw}_{1994}, x_t\right) h\left(x_t|\, e_t = 1, \widetilde{mw}_{1994}\right) dx \qquad (11)$$

where $g\left(w_t|\, e_t = 1, \widetilde{mw}_{1994}, x_t\right)$ is the conditional density of hourly wages, and $h\left(x_t|\, e_t = 1, \widetilde{mw}_{1994}\right)$ is the conditional density of workers' attributes.

The first assumption states that the conditional wage density can be written as

$$g\left(w_t|\, e_t = 1, \widetilde{mw}_{1994}, x_t\right) = g\left(w_t|\, e_t = 1, \widetilde{mw}_t, x_t\right). \qquad (12)$$

The second assumption implies that the conditional density of workers' attributes can be written as

$$\begin{aligned}
h\left(x_t|\, e_t = 1, \widetilde{mw}_{1994}\right) &= \frac{\Pr\left(e_t = 1|\, \widetilde{mw}_{1994}, x_t\right) h\left(x_t|\, \widetilde{mw}_{1994}\right)}{\Pr\left(e_t = 1|\, \widetilde{mw}_{1994}\right)} \\
&= \frac{\Pr\left(e_t = 1|\, \widetilde{mw}_{1994}, x_t\right) h\left(x_t|\, \widetilde{mw}_t\right)}{\Pr\left(e_t = 1|\, \widetilde{mw}_{1994}\right)}. \qquad (13)
\end{aligned}$$

By applying Bayes's rule,

$$\begin{aligned}
h\left(x_t|\, e_t = 1, \widetilde{mw}_t\right) &= \frac{h\left(x_t, e_t = 1|\, \widetilde{mw}_t\right)}{\Pr\left(e_t = 1|\, \widetilde{mw}_t\right)} \\
&= \frac{\Pr\left(e_t = 1|\, \widetilde{mw}_t, x_t\right) h\left(x_t|\, \widetilde{mw}_t\right)}{\Pr\left(e_t = 1|\, \widetilde{mw}_t\right)} \\
h\left(x_t|\, \widetilde{mw}_t\right) &= \frac{\Pr\left(e_t = 1|\, \widetilde{mw}_t\right)}{\Pr\left(e_t = 1|\, \widetilde{mw}_t, x_t\right)} \cdot h\left(x_t|\, e_t = 1, \widetilde{mw}_t\right). \qquad (14)
\end{aligned}$$

Substituting (14) into (13) yields

$$h\left(x_t|\, e_t = 1, \widetilde{mw}_{1994}\right) = \theta\left(\widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right) h\left(x_t|\, e_t = 1, \widetilde{mw}_t\right). \qquad (15)$$

where

$$\theta\left(\widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right) = \frac{\Pr\left(e_t = 1 \middle| \widetilde{mw}_{1994}, x_t\right)}{\Pr\left(e_t = 1 \middle| \widetilde{mw}_t, x_t\right)} \cdot \frac{\Pr\left(e_t = 1 \middle| \widetilde{mw}_t\right)}{\Pr\left(e_t = 1 \middle| \widetilde{mw}_{1994}\right)}.$$

After collecting the conditional wage density (12) and the conditional density of workers' attributes (15), the counterfactual wage density (11) can be written as

$$f\left(w_t \middle| e_t = 1, \widetilde{mw}_{1994}\right) = \int g\left(w_t \middle| e_t = 1, \widetilde{mw}_t, x_t\right) \theta\left(\widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right) h\left(x_t \middle| e_t = 1, \widetilde{mw}_t\right) dx.$$

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

	(1) OLS	(2) OLS	(3) 2SLS	(4) FE	(5) FE	(6) FE	(7) OLS	(8) OLS	(9) 2SLS	(10) FE	(11) FE	(12) FE
Estimation Methods	OLS	OLS	2SLS	FE	FE	FE	OLS	OLS	2SLS	FE	FE	FE
Dependent Variables	10/50 log Wage Differential						90/50 log Wage Differential					
$\ln(MW/W50)$	–	0.39	0.38	0.54	0.61	1.34	–	0.21	0.21	0.27	0.11	0.62
		(0.04)	(0.04)	(0.09)	(0.05)	(0.38)		(0.11)	(0.12)	(0.13)	(0.10)	(0.73)
$[\ln(MW/W50)]^2$	–	–	–	–	–	0.77		–	–	–	–	0.53
						(0.39)						(0.74)
Year 1995	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00	-0.01	-0.01	-0.01	-0.01	-0.00	-0.00
	(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	-0.01
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Year 1997	0.00	-0.01	-0.01	-0.01	-0.02	-0.02	-0.02	-0.03	-0.03	-0.03	-0.01	-0.01
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Year 1998	0.01	-0.01	-0.01	-0.02	-0.02	-0.02	-0.03	-0.04	-0.04	-0.04	-0.01	-0.01
	(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 1999	0.00	-0.02	-0.02	-0.03	-0.03	-0.03	-0.02	-0.03	-0.03	-0.04	0.01	0.01
	(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 2000	0.01	-0.01	-0.01	-0.02	-0.03	-0.03	-0.02	-0.04	-0.04	-0.04	0.01	0.01
	(0.00)	(0.00)	(0.00)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Year 2001	0.02	-0.01	-0.01	-0.02	-0.03	-0.03	-0.01	-0.03	-0.03	-0.03	0.03	0.03
	(0.00)	(0.01)	(0.01)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Year 2002	0.02	-0.01	-0.01	-0.03	-0.03	-0.04	-0.00	-0.02	-0.02	-0.03	0.04	0.04
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Year 2003	0.03	-0.01	-0.01	-0.02	-0.03	-0.03	-0.01	-0.03	-0.03	-0.03	0.04	0.04
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)
Constant	-0.35	-0.14	-0.15	-0.06	0.56	-0.70	0.60	0.71	0.71	0.74	13.94	13.07
	(0.00)	(0.02)	(0.02)	(0.05)	(0.21)	(0.69)	(0.01)	(0.06)	(0.07)	(0.07)	(0.39)	(1.34)
Prefecture trends	No	No	No	No	Yes	Yes	No	No	No	No	Yes	Yes
R^2	0.07	0.56	–	0.55	0.79	0.80	0.03	0.09	–	0.18	0.54	0.54

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

Table 2: Controlling for changes in workforce composition and the disemployment effect. Sample: Females, high school education or less, 1997 and 2002

	(1)	(2)	(3)	(4)	(5)	(6)
Estimation Methods	OLS	FE	FE	OLS	FE	FE
Dependent Variables	10/50 log Wage Differential			90/50 log Wage Differential		
	Panel A: Baseline					
	0.39	0.54	0.61	0.21	0.27	0.11
	(0.04)	(0.10)	(0.06)	(0.11)	(0.13)	(0.10)
	Panel B: No Change in Workforce Composition					
ln(MW/W50)	0.32	0.41	0.56	0.16	-0.15	0.23
	(0.03)	(0.02)	(0.05)	(0.10)	(0.05)	(0.09)
	Panel C: No Truncation Effect					
	0.39	0.54	0.61	0.21	0.28	0.13
	(0.04)	(0.09)	(0.05)	(0.11)	(0.13)	(0.09)
	Panel D: No Employment Loss					
	0.40	0.56	0.59	0.18	0.25	0.10
	(0.04)	(0.08)	(0.05)	(0.11)	(0.11)	(0.09)
Prefecture trends	No	No	Yes	No	No	Yes

Notes: A total of 470 observations are included in Panels A to C, and a total of 423 observations are included in Panel D. Standard errors in parentheses are clustered at the prefecture level. MW and W50 represent minimum wage and median wage, respectively. Other covariates include year dummies.

Table 3: Actual and counterfactual pay gaps between full-time and part-time female workers.

Sample	(1)	(2)	(3)	(4)	(5)	(6)
	1994 Actual		2003 Actual		2003 Counterfactual	
	Full-time	Part-time	Full-time	Part-time	Full-time	Part-time
log Wage	7.053	6.691	7.141	6.759	7.115	6.723
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
log Wage Differentials	0.362		0.382		0.391	
	(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.

Table 4: How the minimum wage affected the employment rate.
Dependent variable: log employment rate
Sample: Females, High school education or less, 1997 and 2002

	(1)	(2)	(3)	(4)
Estimation Methods		FE		
Age Groups	≤22	23–30	31–59	≥60
ln(MW/W50)	-0.768	-0.367	-0.313	-0.373
	(0.569)	(0.315)	(0.169)	(0.489)
Year 2002	-0.044	0.033	-0.017	-0.099
	(0.032)	(0.018)	(0.010)	(0.028)
R^2	0.582	0.120	0.716	0.786

Notes: A total of 94 observations are included. Standard errors in parentheses are clustered at the prefecture level. MW and W50 represent minimum wage and median wage, respectively. The base year is 1997.

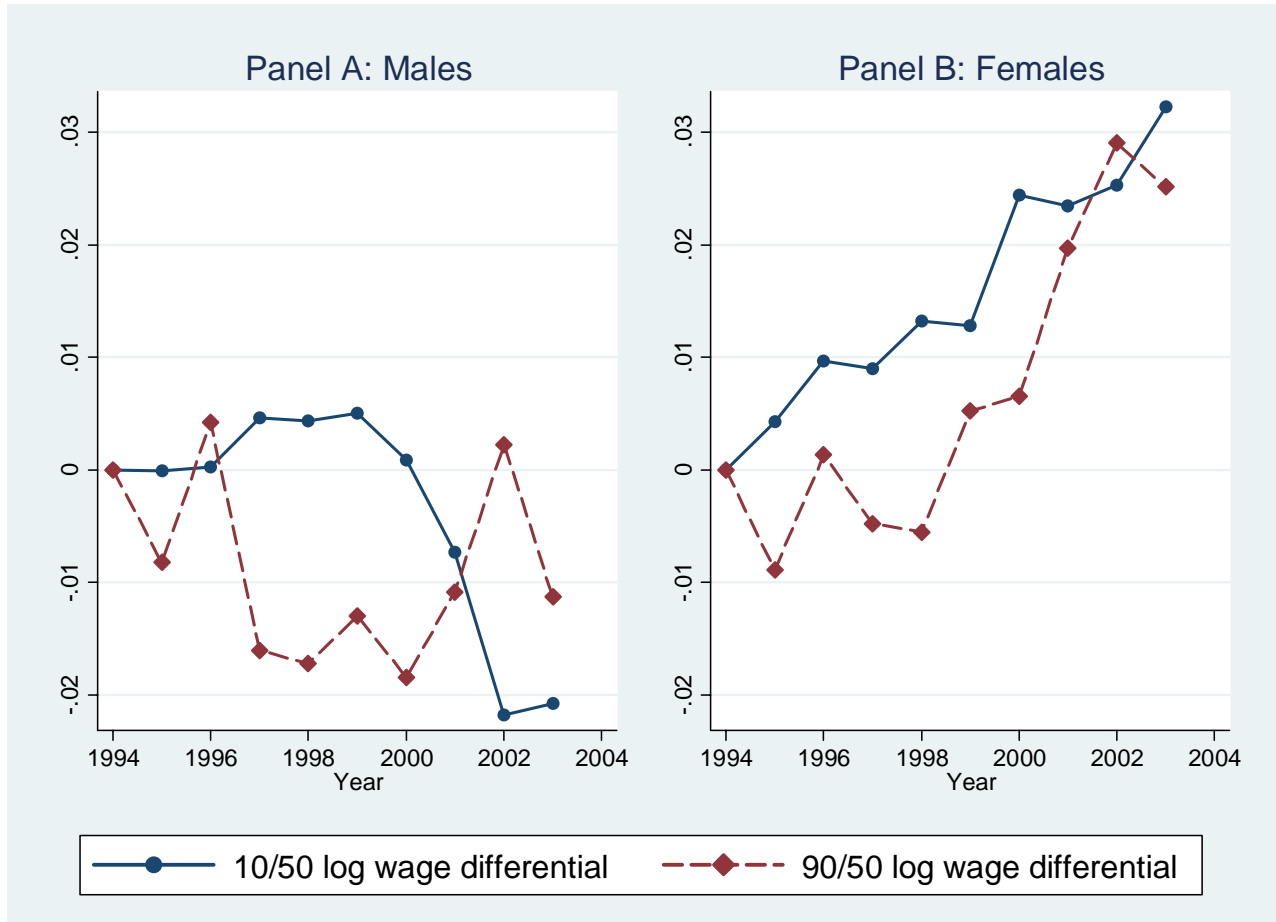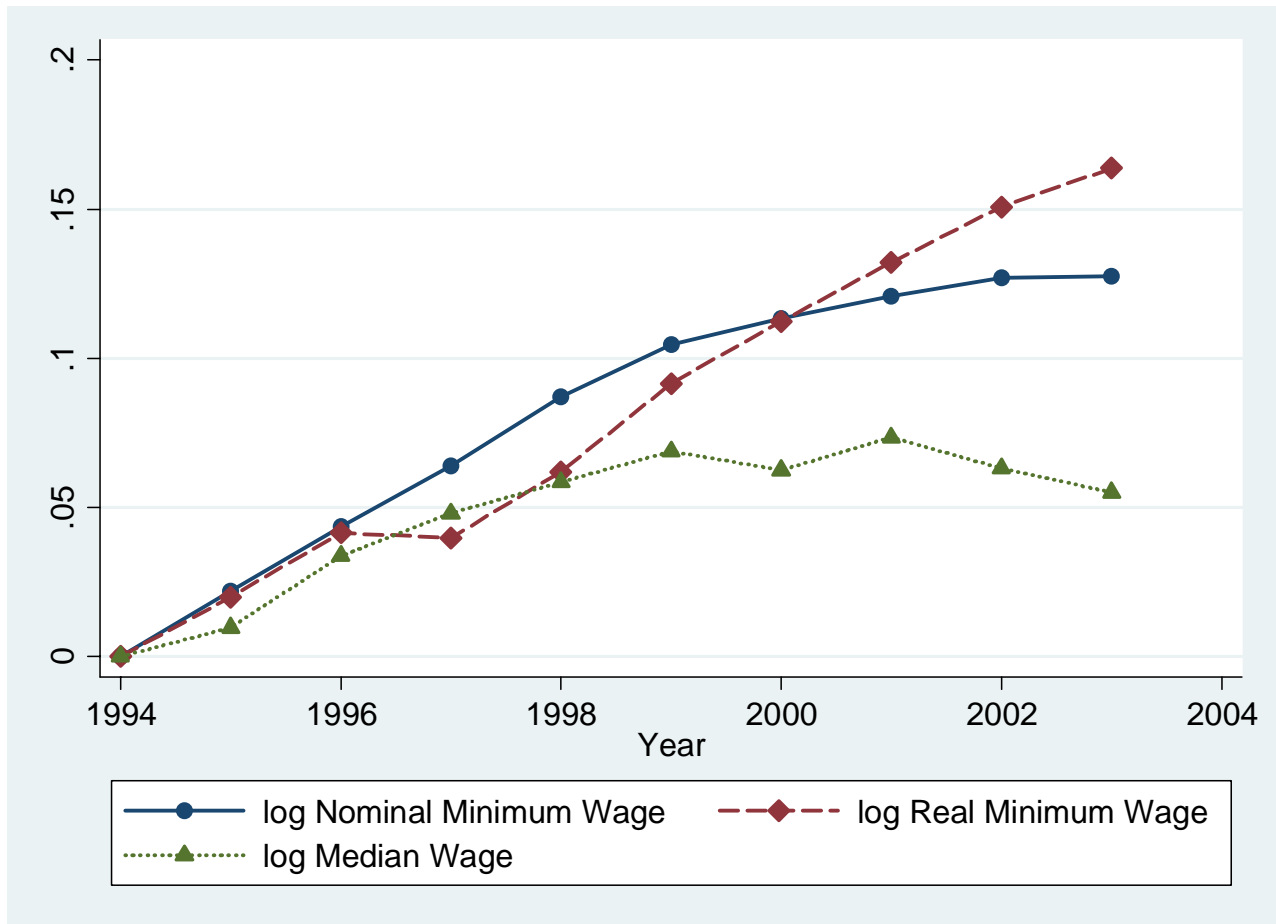Figure 1: Trends in lower and upper tail wage inequality.

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 consumer price index (CPI). The CPI is adjusted according to the results of Broda and Weinstein (2007) that the average rate of deflation was 1.2 percent per year between 1998 and 2006 when the substitution bias and quality upgrading are taken into account.

Figure 3: Wage distribution by selected prefecture and year.

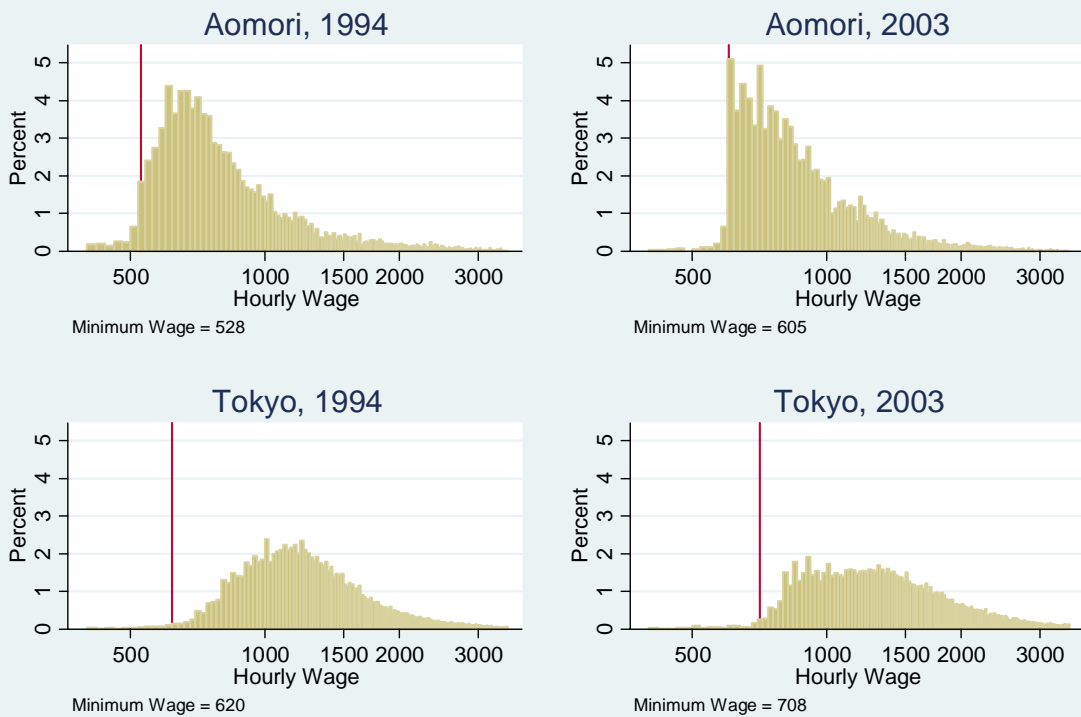Figure 4: The ratio of the minimum wage to the median wage by prefecture in 1994 and 2003.
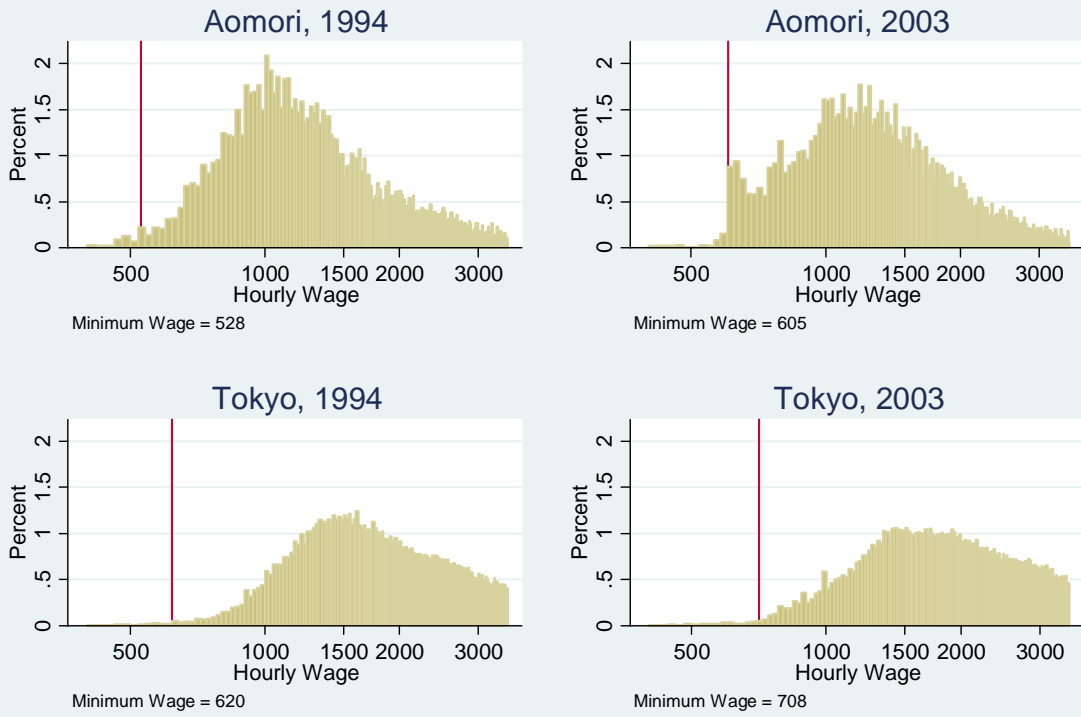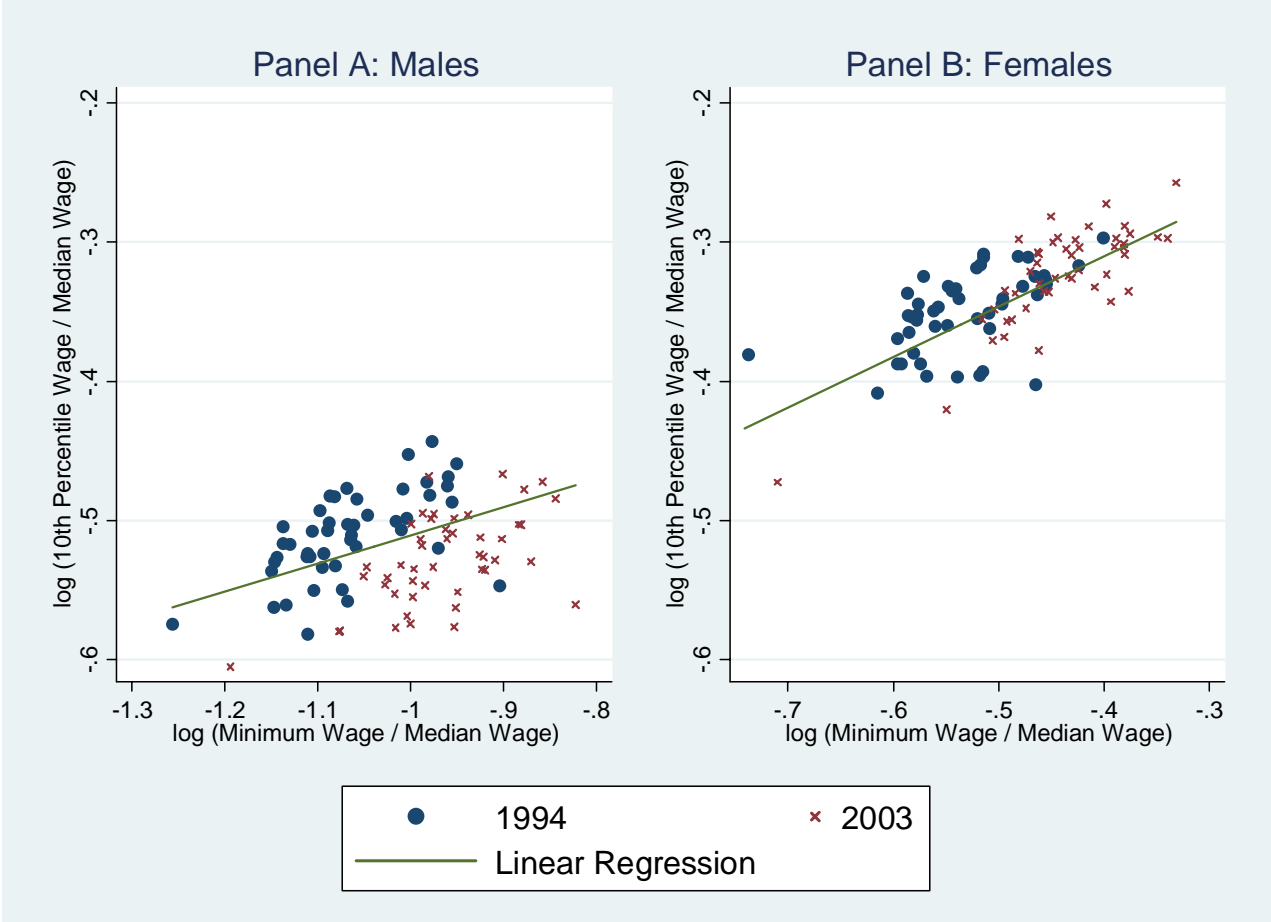
Figure 5: Wage compression and minimum wage.

Figure 6: Actual and counterfactual female log wage distributions.
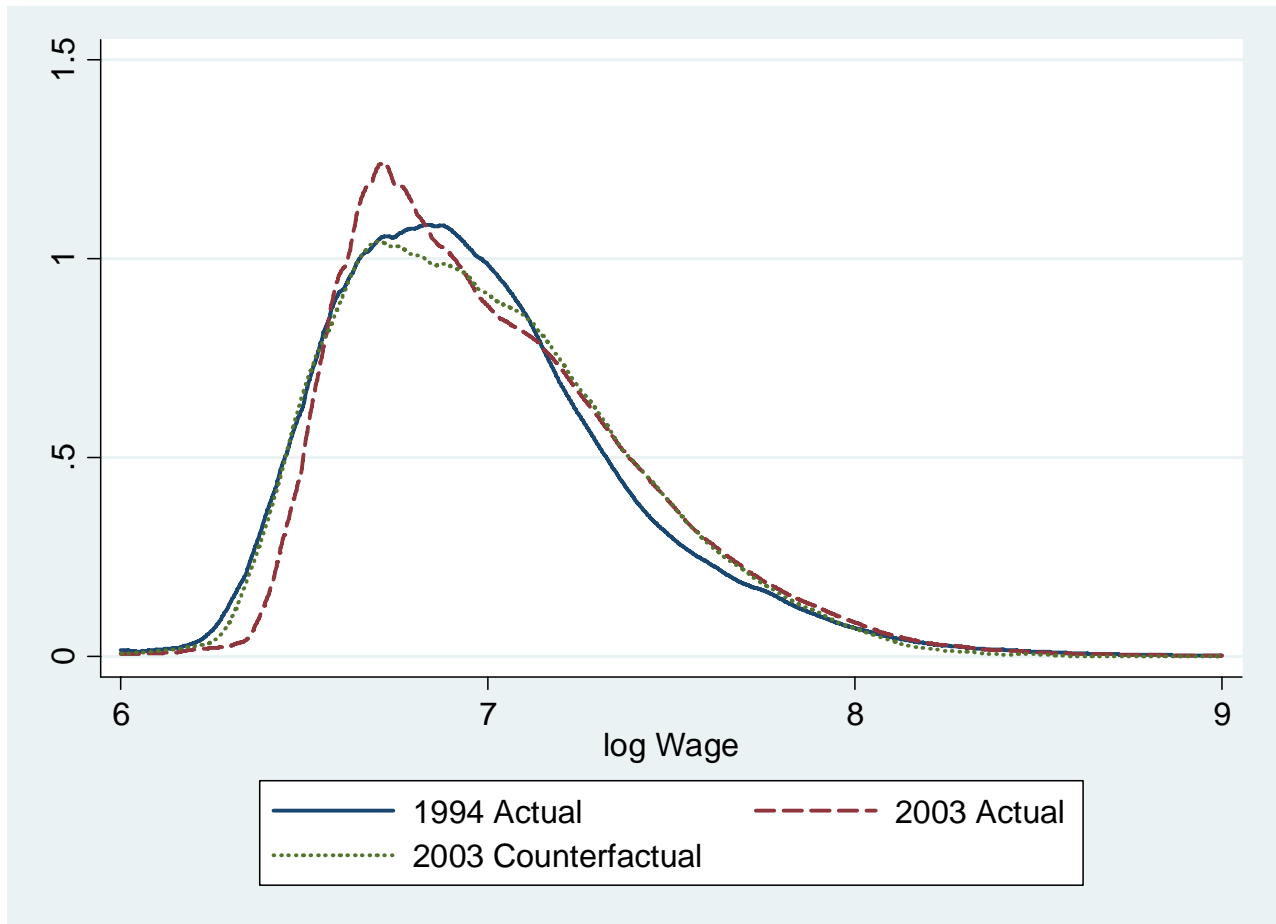
Figure 7: Actual and counterfactual changes in female log hourly wage by percentile, 1994–2003.
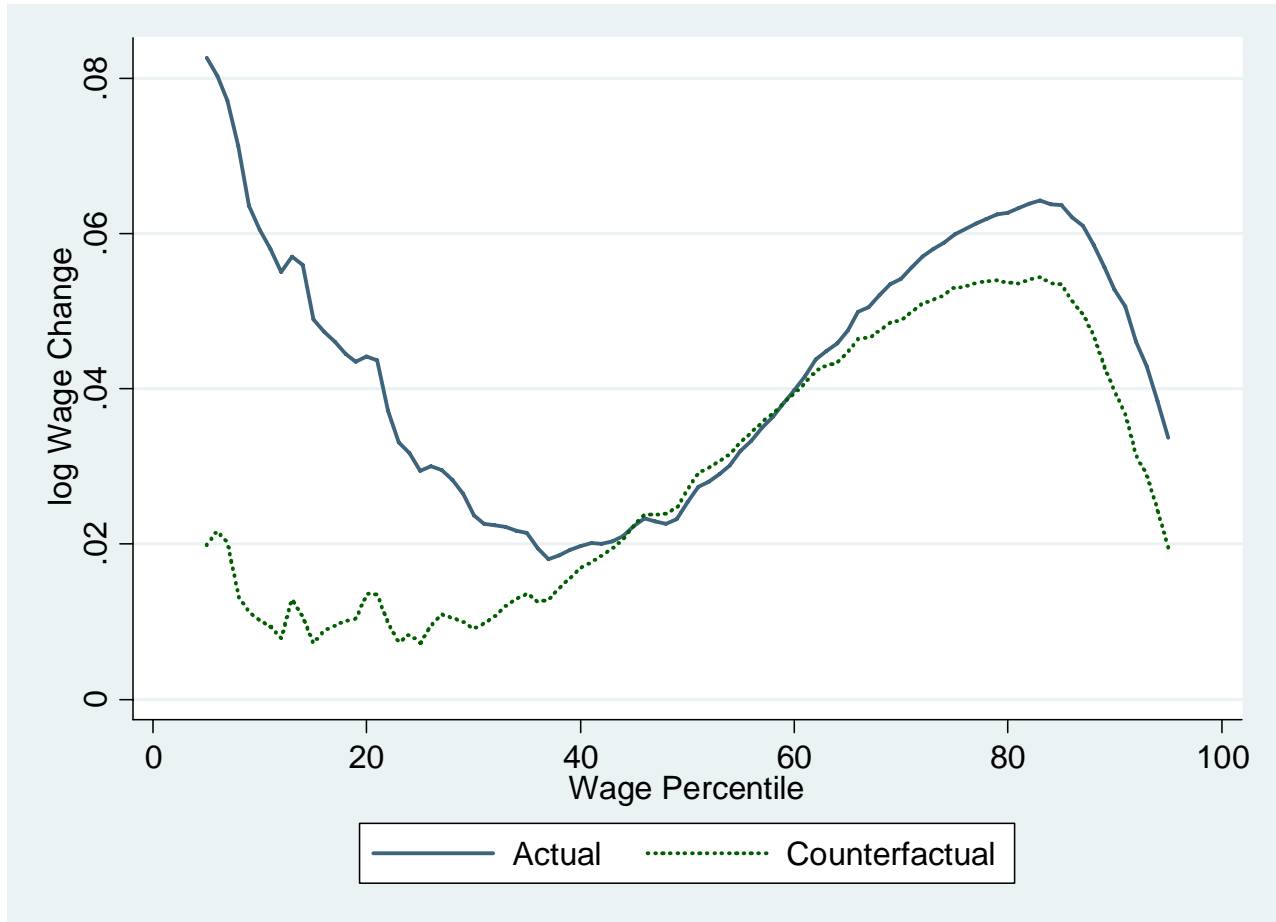
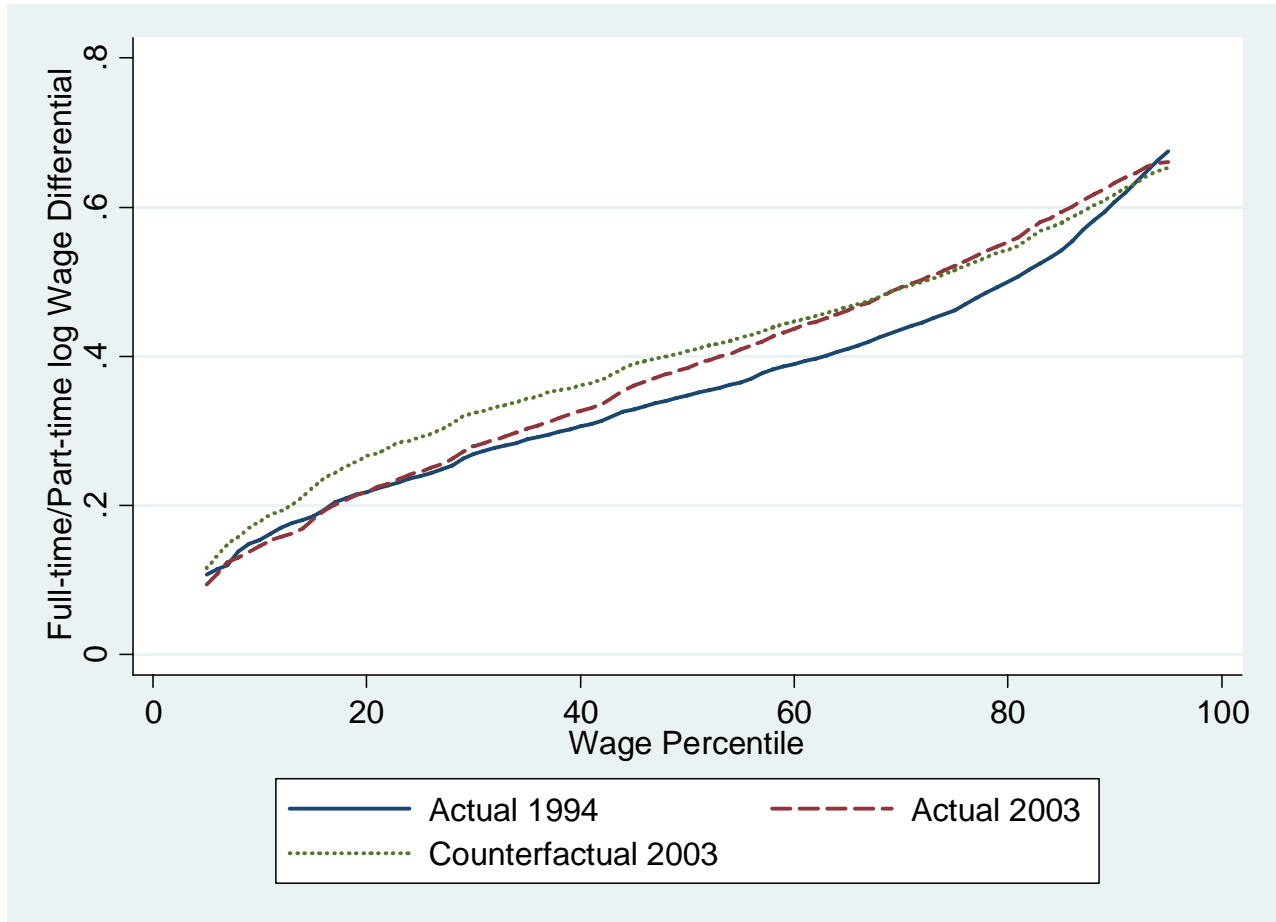Figure 8: Female full-time/part-time log wage differential.

Table A1: How 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	FE	OLS	OLS	2SLS	FE	FE	FE
Dependent Variables	10/50 log Wage Differential						90/50 log Wage Differential					
$\ln(MW/W50)$	–	0.29	0.28	0.49	0.50	0.70	–	0.42	0.42	0.32	0.27	0.31
		(0.07)	(0.07)	(0.06)	(0.06)	(0.31)		(0.12)	(0.13)	(0.06)	(0.06)	(0.82)
$[\ln(MW/W50)]^2$	–	–	–	–	–	0.10	–	–	–	–	–	0.02
						(0.15)						(0.40)
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	-0.00
	(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	-0.02
	(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 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	-0.02	-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 2000	-0.00	-0.02	-0.02	-0.03	-0.04	-0.03	-0.02	-0.05	-0.05	-0.05	-0.02	-0.02
	(0.00)	(0.01)	(0.01)	(0.00)	(0.01)	(0.00)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Year 2001	-0.01	-0.03	-0.03	-0.04	-0.05	-0.04	-0.01	-0.04	-0.05	-0.04	-0.01	-0.01
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Year 2002	-0.02	-0.04	-0.04	-0.06	-0.07	-0.06	-0.01	-0.04	-0.04	-0.03	-0.00	-0.00
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Year 2003	-0.02	-0.05	-0.05	-0.07	-0.07	-0.07	-0.02	-0.06	-0.06	-0.05	-0.02	-0.02
	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Constant	-0.51	-0.20	-0.21	0.01	-0.96	0.12	0.62	1.06	1.07	0.96	7.10	7.11
	(0.00)	(0.08)	(0.08)	(0.06)	(0.18)	(0.17)	(0.01)	(0.13)	(0.13)	(0.06)	(0.19)	(0.23)
Prefecture trends	No	No	No	No	Yes	No	No	No	No	No	Yes	No
R^2	0.06	0.38	–	0.52	0.68	0.53	0.04	0.40	–	0.40	0.64	0.64

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

Table A2: How the minimum wage affected the employment rate.
Dependent variable: log employment rate
Sample: Males, High school education or less, 1997 and 2002

	(1)	(2)	(3)	(4)
Estimation Methods		FE		
Age Groups	≤ 22	23–30	31–59	≥ 60
ln(MW/W50)	-0.428	0.072	-0.067	-0.251
	(0.460)	(0.162)	(0.086)	(0.377)
Year 2002	-0.059	-0.057	-0.036	-0.112
	(0.029)	(0.010)	(0.005)	(0.024)
R^2	0.650	0.853	0.922	0.861

Notes: A total of 94 observations are included. Standard errors in parentheses are clustered at the prefecture level. MW and W50 represent minimum wage and median wage, respectively. The base year is 1997.