The Minimum Wage in a Deflationary Economy: The Japanese Experience, 1994–2003[1]

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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.

Keywords: 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

The real value of minimum wage increased by 50% between 1975 and 1995 in Japan (OECD, 1998, Chart 2.1). The statutory minimum wage has been revised every year and consistently increased in nominal terms in Japan. Due to a severe recession, the median wage has fallen nominally since 1999 (Figure 1). Nonetheless, the nominal minimum wage has steadily increased by inertia in late 1990s and early 2000s. Precisely, the real minimum wage increased by about 0.16 log points between 1994 and 2003, despite the economic downturn. Changes in the nominal minimum wage tend to lag behind general price inflation or deflation. During price and wage deflation, the real increment was higher than the nominal increment. The real value of 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.

Wage distribution has evolved differently among advanced industrialized countries, although these countries have shared the similar experience of rapid technological progress and increased exposure to international trade and outsourcing. In particular, the Japanese wage distribution has remained stable relative to Anglo-Saxon countries. The 90/50 and 10/50 log wage differentials do not demonstrate upward trends for male workers (Figure 2), which is 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). In contrast, the 10/50 and 90/50 log wage differentials have diverged for female workers. This trend implies dispersion at the upper tail and compression at the lower tail of the female wage distribution.

The goal of this paper is to assess 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. There is mixed evidence on the importance of the minimum wage as a determinant of the shape of the lower tail of the wage distribution in several countries. In the United States, DiNardo, Fortin, and Lemieux (1996) demonstrated that erosion of the real minimum-wage level during the 1980s contributed to the wage dispersion. 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) partially 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 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). 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 are set by the collective labor agreement between labor unions and firms, while no statutory minimum wage exists in Germany. Under such an institutional setting, deunionization leads to 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-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. 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.[1]

This study examined the evolution of Japanese wage distribution under conditions of wage deflation. 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.

Our identification strategy was basically to exploit regional variation in the effective minimum

[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 always bumpy.

wage over time, which is known as a difference-in-differences (or fixed-effects) approach. The 'effective' minimum wage can be measured by the difference between the log minimum wage and the log median wage. Because the statutory minimum wage and the nature of wage distribution evolve differently across prefectures, the minimum wage effect on the wage distribution can be isolated from unobserved prefectural heterogeneity and a common macroeconomic fluctuation. Our empirical model included prefecture-specific linear 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 a method to investigate to what extent truncation can explain the evolution of the wage distribution. Our method is a reweighting approach to recover the counterfactual distribution in the absence of disemployment and can be viewed as a variant of inverse probability method (Little and Rubin, 2002). To calculate a reweighting function, we analyzed how a binding minimum wage can affect employment. This proposed method does not require a distributional assumption about latent wage distribution and allows for a possible mechanical compression of the wage distribution associated with disemployment. 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. This study is the first to incorporate the employment effect of minimum wage into the analysis of wage compression.

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 stayed constant at around 0.51 between 1994 and 2003 for female workers but would have diverged by about 0.05 log points 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 0.05 log 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. 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: (1) a regional minimum wage based on collective agreement and (2) prefectural and industrial minimum wages 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.[2] The second system comprises the prefectural minimum wage covering all workers and the industrial minimum wage covering workers aged 18 to 65 in specific industries in specific

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

prefectures. This study focuses on the role of the prefectural minimum wage because the industrial minimum wage covers a small fraction of workers.[3]

Under the current system, prefectural minimum wages are revised every year in two steps. First, the central council classifies all Japanese prefectures into four ranks by actual wage levels and the standard cost of living and decides the minimum-wage increase (*meyasu*) for each rank. 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 central minimum-wage council is not an ad-hoc but standing institution under the administration of the Ministry of Welfare and Labor. The council consists of representatives of public interest (academics and a retired bureaucrat), employers, and employees. The statutory minimum wage has not been coordinated with any other policy such as a measure to promote small and medium enterprise. The *meyasu* system was introduced in 1978 to moderate regional disparity in the minimum wage. The council hardly ever alters their classification of the 47 prefectures into four ranks.

The regional minimum-wage councils' deliberations are largely influenced by decisions regarding the amounts of any minimum-wage increases, which are set annually by the central minimum-wage council. In effect, the current level of the prefectural minimum wage is thus approximately the 1977 level of the prefectural minimum wage plus the accumulated total of *meyasu* since 1978. However, the revised value of prefectural minimum wage is not exactly the same as that indicated by the central minimum-wage council. Thus, prefectural minimum wages could potentially be correlated with local economic conditions. We formally addressed concerns about policy endogeneity by using the 'indicated' (*meyasu*) minimum wage as an instrument. The 'indicated' minimum wage is calculated by the minimum wage in the previous year plus *meyasu*.

The political climate 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). Wage distributions are much more heterogeneous across pre-

[3] The report by Saitei Chingin no Arikata Kenkyukai (minimum wage study group) of the Ministry of Health, Labor and Welfare issued a recommendation for the revision of industrial minimum wage, including a possibility of its abolishment in 2005. In the fiscal year 2000, 4.5 million workers were covered by industry minimum wage while 52 million workers were covered by regional minimum wage according to the press release of the ministry of labor and welfare on January 25 in 2001.

fectures than minimum wages. The political process of the minimum-wage determination described above tends to be biased towards inertia. In effect, prefectural minimum wages have not been changed in response to economic shocks to local labor markets. Consequently, the minimum-wage bite became severe in rural areas during a recession.

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. This may be part of reasons that the central council consistently suggested an increase in the minimum wage. 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. The variations in effective minimum wage across prefectures over time are precisely measured in this study using a large micro-data set. We exploited them to identify how the minimum wage affected the wage distribution and employment.

The minimum wage is legally enforceable in the following manner. 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. 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.

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,[4] 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. 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*),[5] 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 hours – commutation allowance – perfect attendance allowance – family allowance)/contracted hours of work, which is consistent with the minimum-wage law.[6][7]

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.[8] The survey collects infor-

[4]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.

[5]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.

[6]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.

[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.

mation 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

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 expected wage for unemployed 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 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. Surprisingly, 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 main interest is in accounting for the reduction in the lower-tail inequality for female workers. Our estimation applied a model of the form

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

where $\widetilde{w}_{it}^{p} = \ln(w_{it}^{p}/w_{it}^{50})$ and $\ln\widetilde{mw}_{it} = \ln(mw_{it}/w_{it}^{50})$.[9] 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, 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

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

wage relative to the median wage caused by a one percent increase in the effective minimum wage in the linear specification. A quadratic term of the effective minimum wage is included as additional regressors to capture a nonlinear relationship 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 (FEs) were added to allow for unobserved heterogeneity in the dispersion of the latent wage distribution across prefectures.

Estimated results were produced using ordinary least-squares (OLS) and two-stage least squares (2SLS), and standard errors were clustered at the prefecture level. Table 1 lists the results for female workers. 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 shows the increasing trend of the 10th percentile wage relative to the median wage. Next, we added 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 rose from 0.07 to 0.56 after the effective minimum wage was 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 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.

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. Moreover, a bias can also arise from the policy endogeneity. The direction of the bias is upward (downward) if the minimum-wage increase tends to be permitted by the local authority which has a small (large) lower-tail inequality. To work around the potential bias, we used an instrumental variable approach. 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. Therefore, unlike the analysis of Autor, Manning, and Smith (2008), our model allows 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.

Another concern is a spurious correlation arising from unobserved heterogeneity in the dispersion of the latent wage distribution across prefectures. In other words, a key 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 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 fixed-effects approach is suitable to solve such a problem. The 2SLS results suggest that measurement errors are negligible. Thus, the potential bias should not be exacerbated by the within transformation. 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 the 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 relationship between minimum wage and wage compression.

One way to assess the validity of our analysis is to examine the effect on the upper-tail inequality. 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. Importantly, the effect of the minimum wage plummeted and became statistically non-significant after we controlled for prefecture trends. Column 12 shows that a higher-order term is not significant, either. 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.

For the sake of completeness, Tables A1 lists the results for male workers. Column 1 shows that the 10/50 log wage differential was almost stable between 1994 and 2000 and started to erode slightly thereafter. 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. The coefficients for year dummies shrank when the effective minimum wage was held constant. Thus, the 10/50 log wage differential would have diverged if the real minimum-wage level were unchanged. The 2SLS estimates were identical to OLS estimates. The effective minimum wage had an even stronger effect after controlling for prefectural fixed effects and prefecture-specific linear trends. Nonetheless, the minimum wage played a less significant role in pushing up the lower tail of the wage distribution for male workers in terms of estimated coefficients and R^2.

Columns 7–12 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 a linear time trend. However, that an increase in the minimum wage would push up the 90th percentile wage is not very realistic. The results may indicate that the effective minimum wage was higher in prefectures where the male latent distribution is more dispersed. Thus, our estimate for the effect of the minimum wage on the lower-tail inequality for male workers may be biased downward. The rest of our analysis focuses on the female wage distribution.

4.3 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. In fact, these attributes are the key determinants of wages in the Japanese labor market. The 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 observed density of log hourly wages in year t is expressed as

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

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

$$\begin{aligned} f_{1994}\left(w|\,T=t\right) &= \int g\left(w|\,x,T=t\right)h\left(x|\,T=1994\right)dx \qquad (3)\\ &= \int g\left(w|\,x,T=t\right)\psi_t\left(x\right)h\left(x|\,T=t\right)dx. \qquad (4)\end{aligned}$$

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

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

The reweighting function can be estimated using a logit model applied to the pooled data from year 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 reexamined Lee's (1999) model of wage compression. Panel A in Table 2 displays the results in the selected columns of Table 1, while panel B in Table 2 shows the results using the counterfactual wage data. The minimum wage effect became greater for the female lower-tail (10/50) inequality and lesser

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

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.

4.4 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:

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

where $\widehat{\beta}_{1p}$ and $\widehat{\beta}_{2p}$ are the 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.

Figure 6 displays the actual and counterfactual wage distributions for female workers in 1994 and 2003. The horizontal axis is the log hourly wage. 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. The difference between the actual and counterfactual distribution illustrates the effect of the minimum-wage increase. 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 between 1994 and 2003. In fact, lower percentile wages increased considerably for female workers between 1993 and 2004. However, 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. It may appear that the minimum wage contributed to widening the upper-tail inequality slightly, but the effects were statistically non-significant, as seen in column 11 of Table 1.

4.5 The Part-time Pay Penalty

An increasing trend of part-time workers is a global phenomenon. For example, about 45% of female workers in Britain work part-time (Manning and Petrongolo, 2008). In keeping with this trend, the full-time/part-time wage differential has been of interest for 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. Reducing the full-time/part-time wage differential has been a heatedly debated policy issue.

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. 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 the British national minimum wage. Abe and Tanaka (2007) found that a minimum wage prevented wage erosion among part-time workers relative to full-time workers in Japan.

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 focused on female workers because the proportion of male part-time workers was very small.[11] Table 6 reports the actual and counterfactual pay gaps between full-time and part-time workers. The actual pay gap was 36.2 percent in 1994 and increased to 38.2 percent in 2003. However, the pay gap would have been 39.1 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 1 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.

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) did not control for either prefecture or year effects.

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

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

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

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

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 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. Prefecture dummies are included to allow for an unobserved prefecture effect. In light of the criticism by Card and Krueger (1995), we did not include the college enrollment rate as a regressor. Thus, the regressors include only exogenous variables in our preferred 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. 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 a fixed-effects approach. Table 3 reports how the minimum wage affected female employment by age group.[13] The estimated year effect was negative for all age groups, indicating a decline in the labor force attachment. The minimum wage effect was negative but nonsignificant except for females aged 31–59 years. However, column 3 reveals a moderate and

[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 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.

significant disemployment effect for females 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 was not statistically significant in any age group.

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.[14] Estimated employment elasticity was imprecise but greatest for those aged 22 or younger, as shown in column 1 of Tables 3 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 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 (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 time trends, and the residual, respectively. The model was estimated using a fixed-effects approach. The hat represents the estimate. 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 1.82-percent

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

decrease in the ratio of new graduates hired among female employees. 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 –0.31 with a standard error of 0.81, and R^2 was 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 compresses the lower tail of the wage distribution but reduces employment for low-skilled 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. 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 investigated to what extent truncation can explain the compression of lower-tail inequality.

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 is written as $f\left(w_t \middle| \widetilde{mw}_t, x_t\right)$ in year t. Then, the counterfactual density if the effective minimum wage had remained at the 1994 level is expressed as $f\left(w_t \middle| \widetilde{mw}_{1994}, x_t\right)$. The actual wage density is under the influence of truncation due to the minimum-wage increase between 1994 and 2003, whereas the counterfactual density is not. Denote an indicator variable for being employed by e_t. The following assumptions are required to quantify the effect of truncation.

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

[15] 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, $f_*(\omega) = f(\omega \,|\omega \geq mw) = \frac{f(\omega)}{\Pr(\omega \geq mw)} \geq f(\omega)$. Thus, $\omega^{50} - \omega^{10} \geq \omega_*^{50} - \omega_*^{10}$.

on the distribution of workers' attributes, i.e., $f\left(x_t \mid e_t=1, \widetilde{mw}_t\right) = f\left(x_t \mid e_t=1, \widetilde{mw}_{1994}\right)$; (c) selection on observables, i.e., $\Pr\left(e_t=1 \mid w_t, \widetilde{mw}_{1994}, x_t\right) = \Pr\left(e_t=1 \mid \widetilde{mw}_{1994}, x_t\right)$.

Under the first two assumptions, the minimum wage affects the wage distribution solely through the disemployment effect. These assumptions are extreme but suitable for our purpose to remove the truncation effect. The counterfactual density can be nonparametrically constructed as follows:

$$f\left(w_t \mid e_t=1, \widetilde{mw}_{1994}\right) = \int f\left(w_t \mid e_t=1, \widetilde{mw}_t\right) \theta\left(w_t, \widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right) \eta\left(\widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right) dx, \tag{8}$$

where

$$\begin{aligned}
\theta\left(w_t, \widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right) &= \frac{\Pr\left(e_t=1 \mid w_t, \widetilde{mw}_{1994}, x_t\right)}{\Pr\left(e_t=1 \mid w_t, \widetilde{mw}_t, x_t\right)} \cdot \frac{\Pr\left(e_t=1 \mid \widetilde{mw}_t, x_t\right)}{\Pr\left(e_t=1 \mid \widetilde{mw}_{1994}, x_t\right)}, \\
\eta\left(\widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right) &= \frac{\Pr\left(e_t=1 \mid \widetilde{mw}_{1994}, x_t\right)}{\Pr\left(e_t=1 \mid \widetilde{mw}_t, x_t\right)} \cdot \frac{\Pr\left(e_t=1 \mid \widetilde{mw}_t\right)}{\Pr\left(e_t=1 \mid \widetilde{mw}_{1994}\right)}.
\end{aligned}$$

In general, the weight, $\theta\left(w_t, \widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right)$, is not defined well for non-employed individuals because their offered wages are not observed. Nonetheless, under the selection-on-observables assumption that $\Pr\left(e_t=1 \mid w_t, \widetilde{mw}_{1994}, x_t\right) = \Pr\left(e_t=1 \mid \widetilde{mw}_{1994}, x_t\right)$ and $\Pr\left(e_t=1 \mid w_t, \widetilde{mw}_t, x_t\right) = \Pr\left(e_t=1 \mid \widetilde{mw}_{1994}, x_t\right)$, the weight, $\theta\left(w_t, \widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right)$, is one for all observations. The weighting function, $\eta\left(\widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right)$, is the ratio of the employment rate at the 1994 minimum-wage level to the employment rate at the current minimum-wage level. Thus, in the weighting procedure, more weight is put on workers who would have a higher propensity to be employed if the minimum wage were unchanged. An estimate of the response probability can be obtained from a probit model:

$$\Pr\left(e_{kt}=1 \mid \widetilde{mw}_{it}, x_{jt}\right) = \Phi\left(\rho_{j0} + \rho_{j1}\widetilde{mw}_{it} + \zeta_j t + d_i \delta_j\right). \tag{9}$$

where k is an index for individual, j is an index for worker's attributes x. Specifically, workers are classified into four age groups ($\leq$22, 23-30, 31-59, $\leq$60). All variables are allowed to vary with demographic group j. The model is estimated using the individual-level micro data from ESS. The minimum wage effects were negative and statistically significant for all age groups. The estimated coefficients ρ_{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). Given the estimated coefficients, we can calculate the weight for every year over the sample periods by virtue of linear time trends, although the ESS were collected only in the years 1997 and 2002, during the sample period of 1994–2003. Higher-order terms of time trends cannot be identified from two-period data.

Figure 9 illustrates the 2003 counterfactual distribution in the absence of disemployment effect along with the actual distribution. The counterfactual distribution fully overlaps with the actual distribution. Panel C in Table 2 shows the results that are reproduced using the counterfactual sample. The results are identical to those using the actual sample. Therefore, the truncation effect is negligible in the analysis of wage compression.

5.4 Upper Bound of the Truncation Effect

We proposed a general approach to examine the effect of truncation on the wage distribution. The approach can be viewed as a variant of inverse probability weighting method and works under the selection-on-observables assumption. However, the analysis conducted above was limited by data constraints. Education is not included in workers' attributes, and time effects were not flexibly specified. A possible misspecification could potentially reduce the effect of truncation. In addition, the employment loss may occur from the bottom end of the wage distribution within a group defined by observed characteristics, whereas it is assumed to occur randomly conditional on observed characteristics in our proposed method. If that is the case, our previous analysis might underestimate the effect of truncation.

Thus, we developed an alternative approach to quantify the upper bound of the mechanical effect. This approach does not rely on the selection-on-observables assumption. The estimated disemployment effect reported in Panel A in Table 3 was used to recover the counterfactual wage distribution if employment loss did not occur. The change in the log employment rate caused by changes in the minimum wage 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. Assume that the population size is unchanged. Then, the change in the number of employed can be expressed as

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

Using this equation, 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 job by group and prefecture between years $t-1$ and t. Then, calculate the total number of unemployed workers between 1995 and 2003, $N_{it}^{add} = -\sum_j \min\{\Delta emp_{jit}, 0\}$.

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.

We recovered the wage distribution in the absence of disemployment for all demographic groups. In the process of creating the counterfactual sample, we lost the observations from the first year of the sample period. In the procedure, a value of zero was imputed to the log hourly wage for unemployed individuals. This imputation is extreme but suitable for our purpose to examine the upper bound of the truncation effect. If the imputed wage is lower than the 10 percentile wage, the 10/50 log wage differential will be unchanged. If it is higher, then the truncation effect should be even smaller.

The results of the 10/50 log wage differential in Table 1B were reproduced for the counterfactual sample in panel D of Table 2. Estimation results differed only marginally. Assuming 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.

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 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. 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 in 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. However, we obtained similar results for wage compression 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 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.

Appendix

Derivation of equation (8). The counterfactual wage densitiy is

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

The wage density conditional on x is

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

Note that

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

Thus,

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

where

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

The density of observed characteristics is

$$\begin{aligned}
f\left(x_t|\, e_t = 1, \widetilde{mw}_{1994}\right) &= \frac{\Pr\left(e_t = 1|\, \widetilde{mw}_{1994}, x_t\right) f\left(x_t|\, \widetilde{mw}_{1994}\right)}{\Pr\left(e_t = 1|\, \widetilde{mw}_{1994}\right)} \\
&= \frac{\Pr\left(e_t = 1|\, \widetilde{mw}_{1994}, x_t\right) f\left(x_t|\, \widetilde{mw}_t\right)}{\Pr\left(e_t = 1|\, \widetilde{mw}_{1994}\right)} \qquad (11)
\end{aligned}$$

Note that

$$
\begin{aligned}
f\left(x_t \mid e_t=1, \widetilde{mw}_t\right) &= \frac{f\left(x_t, e_t=1 \mid \widetilde{mw}_t\right)}{\Pr\left(e_t=1 \mid \widetilde{mw}_t\right)} \\
&= \frac{\Pr\left(e_t=1 \mid \widetilde{mw}_t, x_t\right) f\left(x_t \mid \widetilde{mw}_t\right)}{\Pr\left(e_t=1 \mid \widetilde{mw}_t\right)} \\
f\left(x_t \mid \widetilde{mw}_t\right) &= f\left(x_t \mid e_t=1, \widetilde{mw}_t\right) \frac{\Pr\left(e_t=1 \mid \widetilde{mw}_t\right)}{\Pr\left(e_t=1 \mid \widetilde{mw}_t, x_t\right)}.
\end{aligned}
$$

Thus,

$$
f\left(x_t \mid e_t=1, \widetilde{mw}_{1994}\right) = f\left(x_t \mid e_t=1, \widetilde{mw}_t\right) \eta\left(\widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right)
$$

where

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

Therefore, the counterfactual wage densitiy can be written as

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

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Table 1: How 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	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 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					
ln(MW/W50)	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					
	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: 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)	(0.01)	(0.028)
R^2	0.582	0.12	0.716	0.786

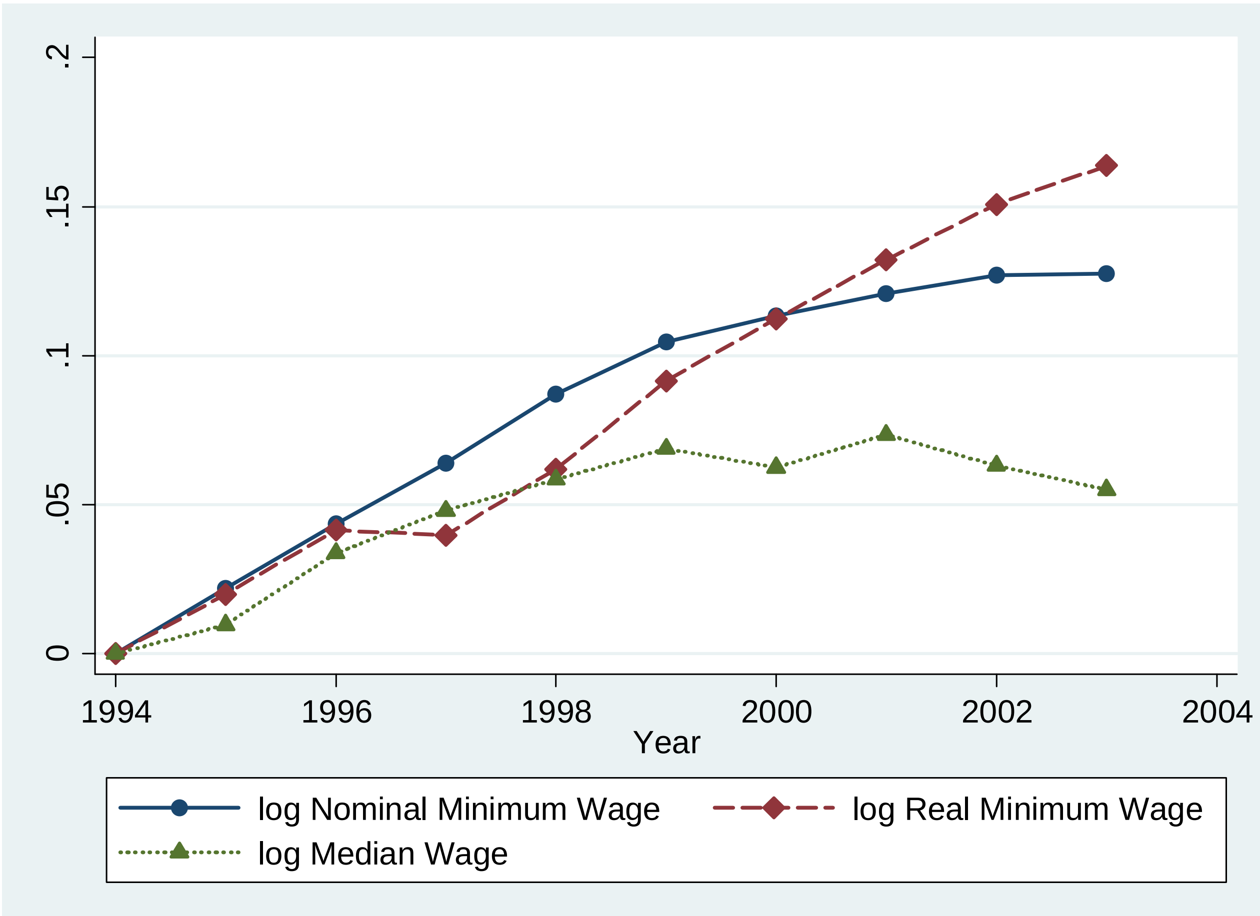
Notes: A total of 94 observations are included. Standard errors in parentheses are clustered at the prefecture level. MW and W50 represent minimum wage and median wage, respectively.

Table 4: 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	7.053	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.

Figure 1: 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 (base year) and 2006 when the substitution bias and quality upgrading are taken into account.

Figure 2: Trends in lower and upper tail wage inequality.

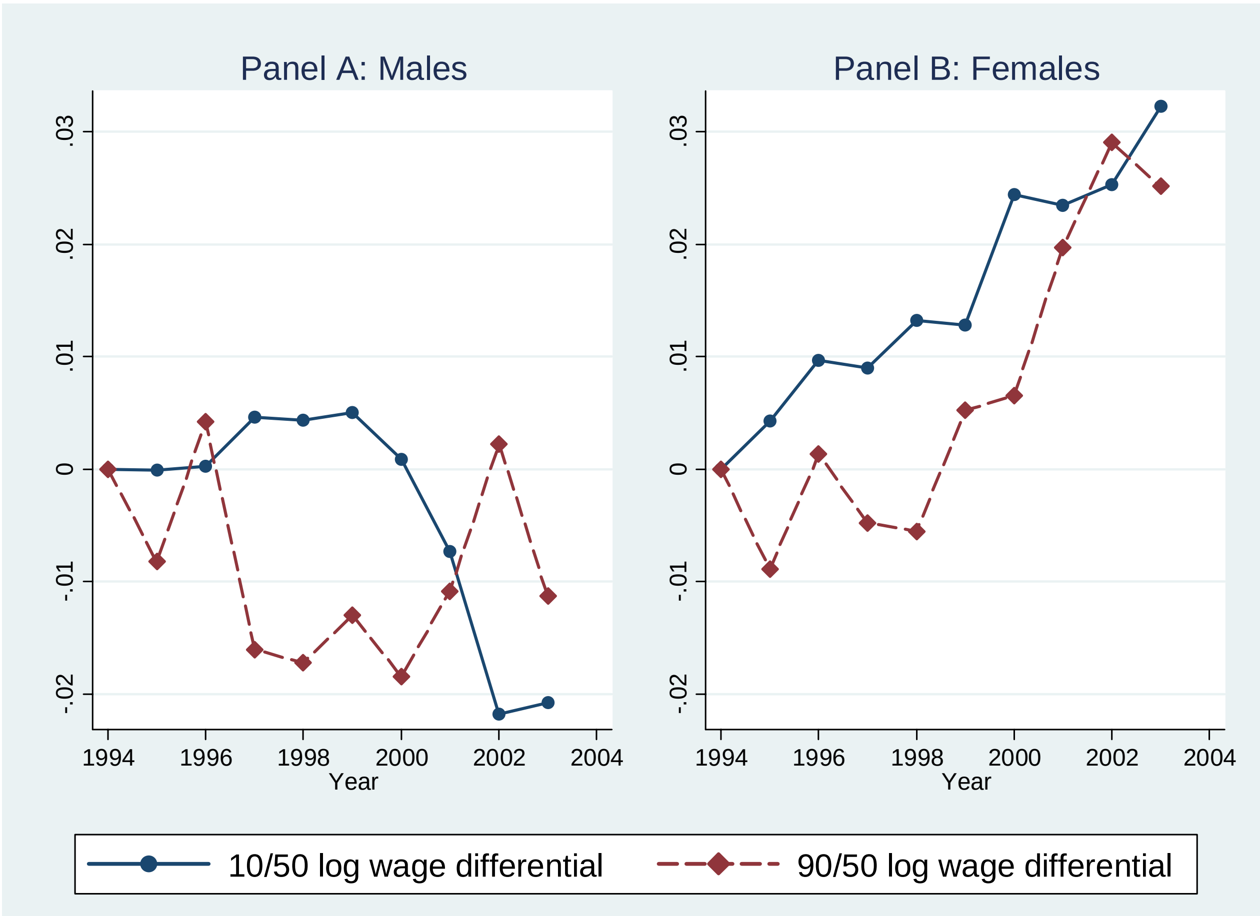


Figure 3: The ratio of the minimum wage to the median wage by prefecture in 1994 and 2003.

Figure 4: Wage distribution by selected prefecture and year.

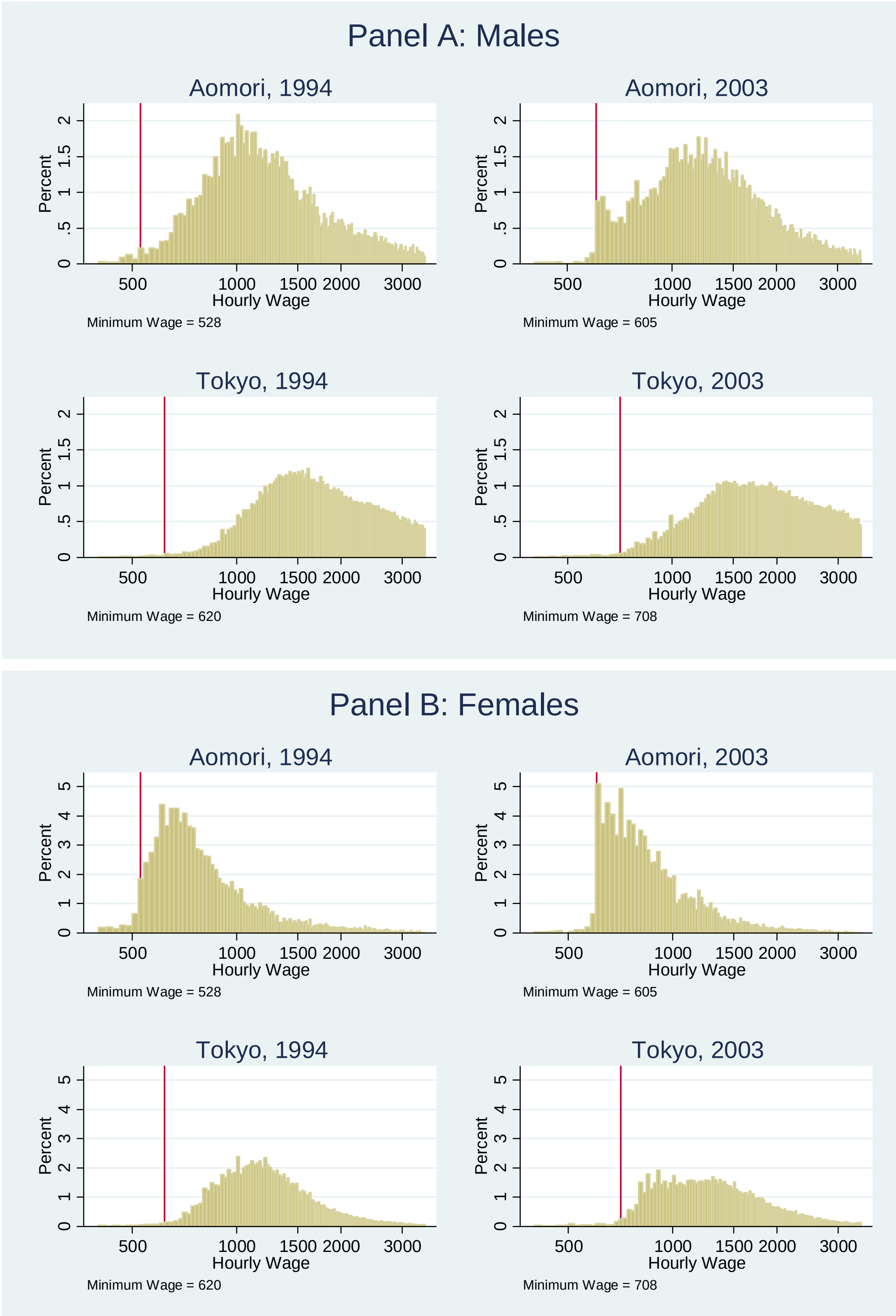


Figure 5: Wage compression and minimum wage.

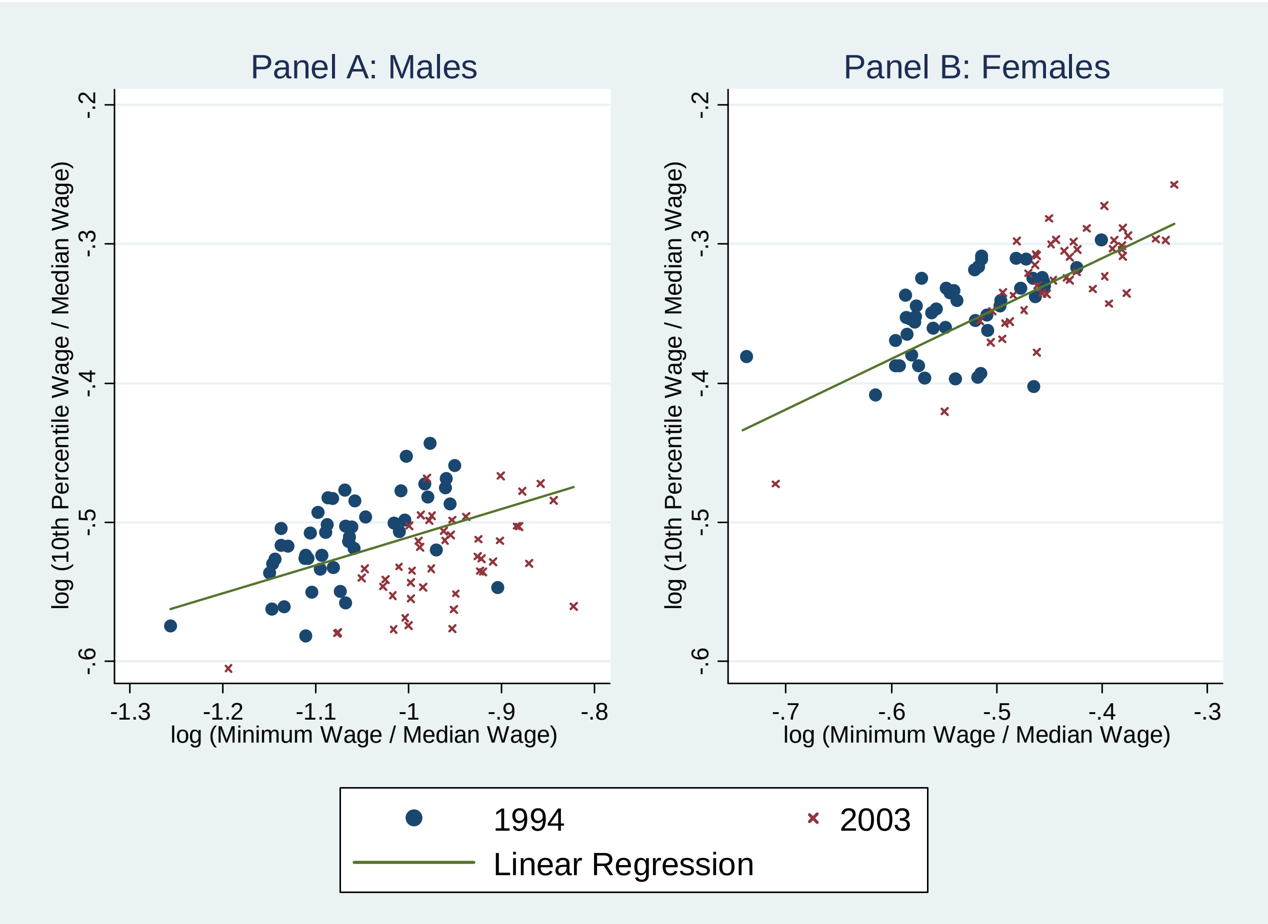


Figure 6: Actual and counterfactual female log wage distributions.

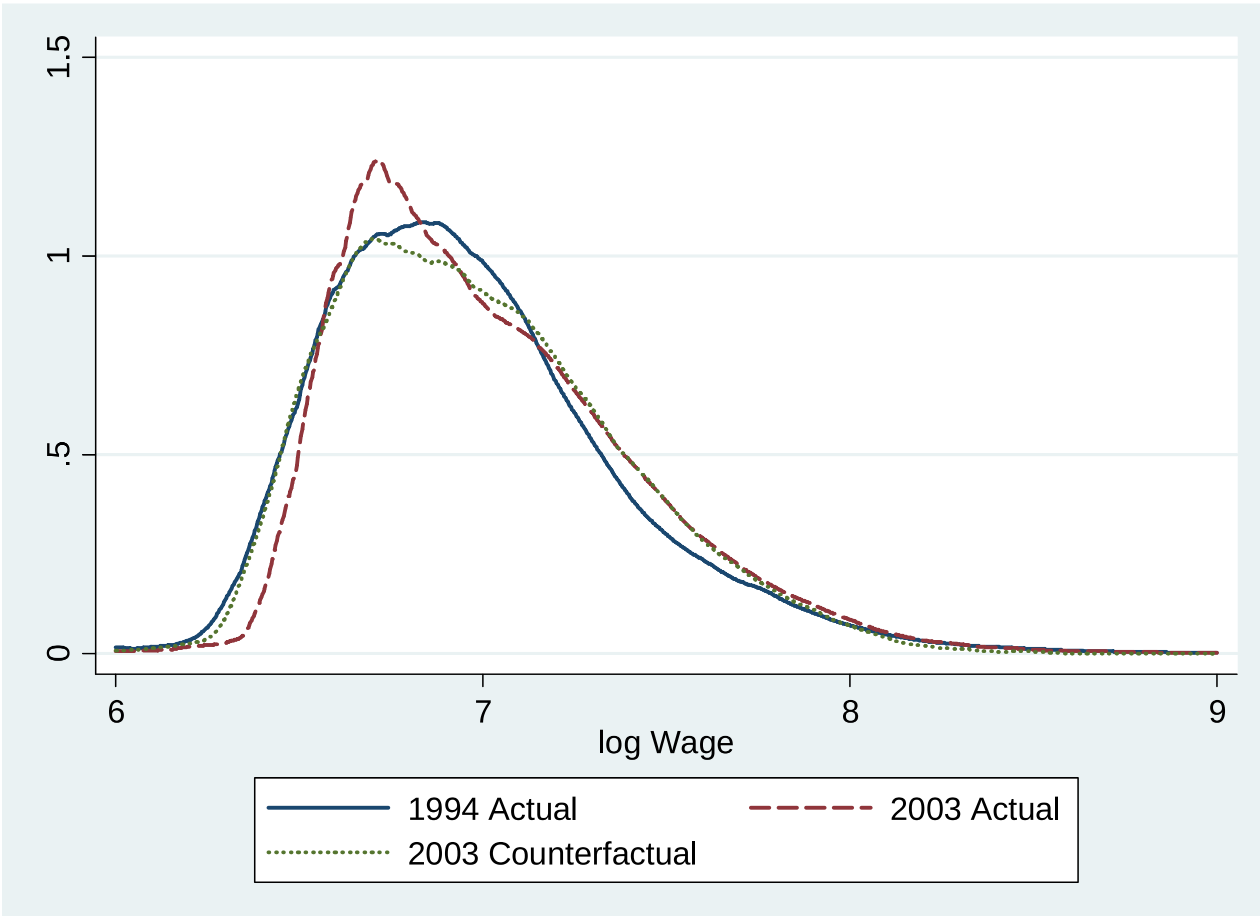


Figure 7: Actual and counterfactual changes in female log hourly wage by percentile, 1994–2003.

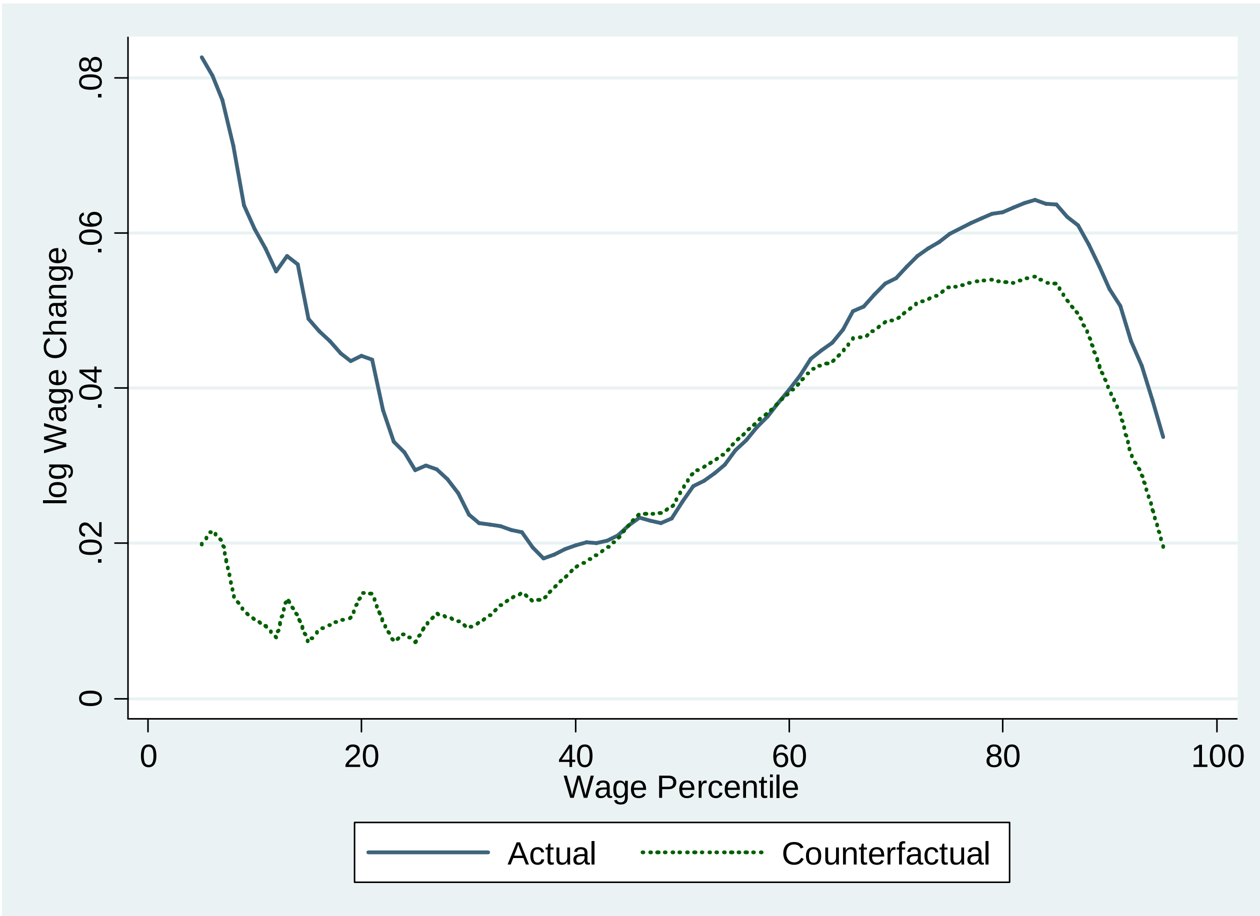


Figure 8: Female full-time/part-time log wage differential.

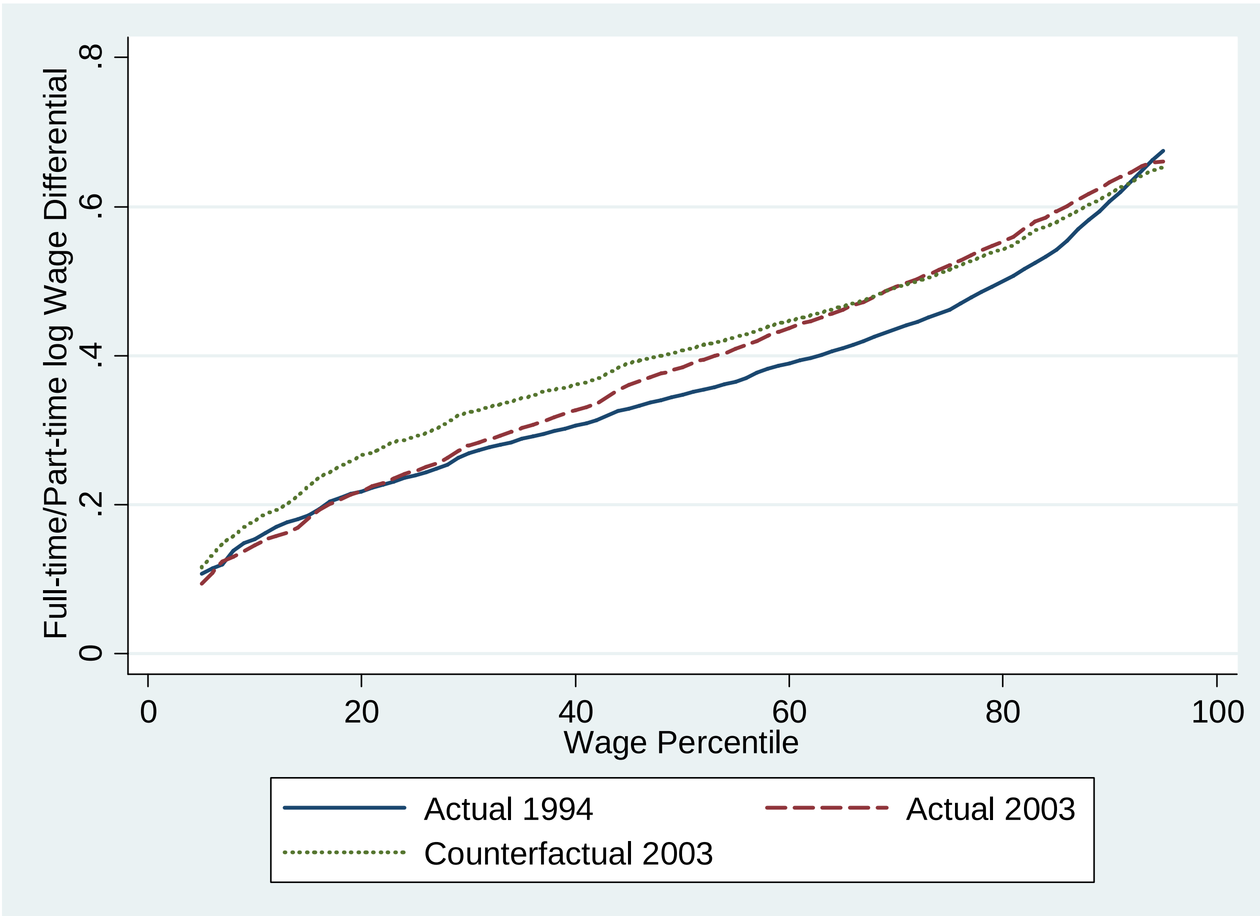


Figure 9: Counterfactual distributions in the absence of disemployment effect.

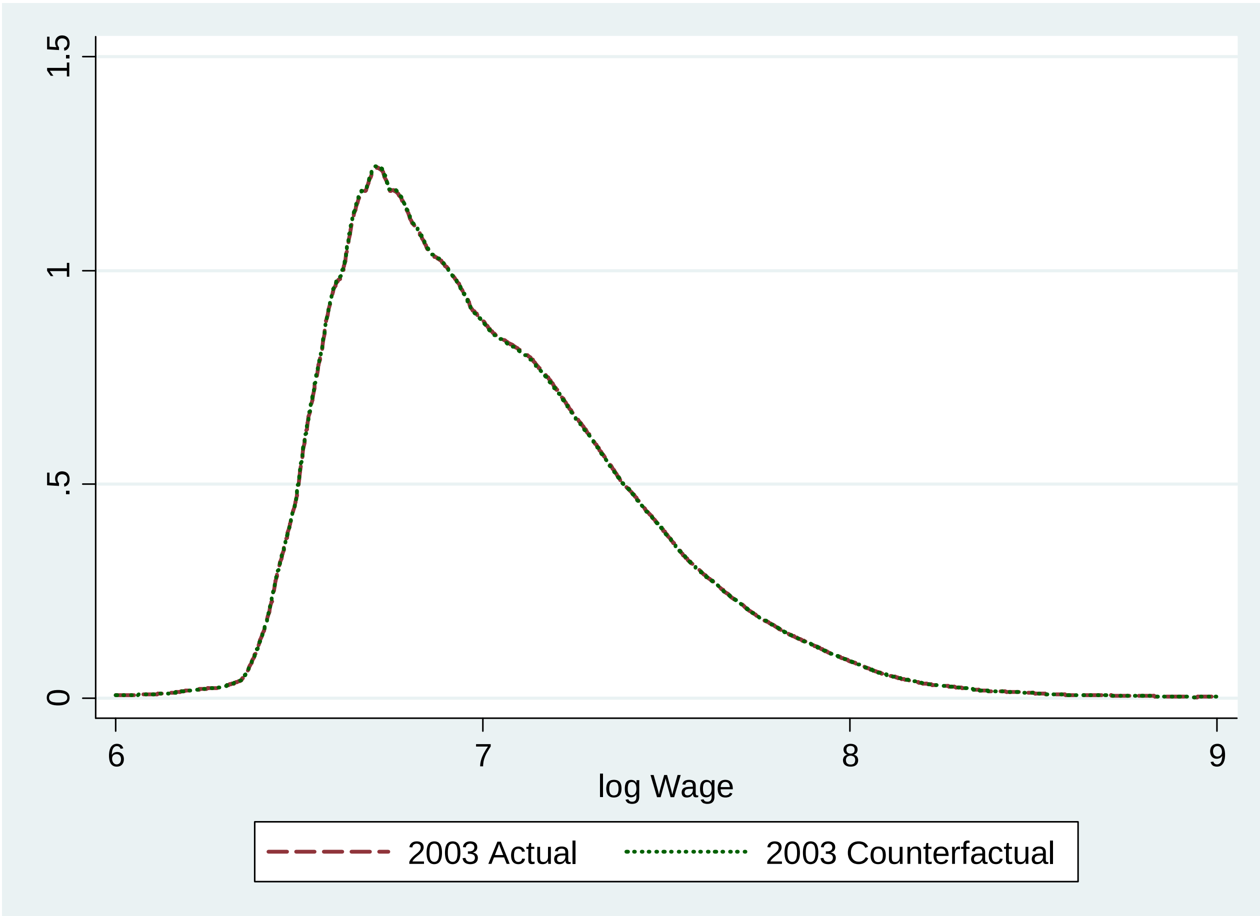


Table A1: How the minimum wage affected the wage distribution.
Sample: Males, 1994–2003

	(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.