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Abstract

The statutory minimum wage in Japan has increased continuously for a few decades until the early 2000s even during a period of deflation. This paper examines the impact of the minimum wage on wage and employment outcomes under this unusual circumstance. We find that the minimum-wage increase resulted in the compression of the lower tail of the wage distribution among women and that the wage compression is only partially attributable to the loss of employment. The continuous increase in the minimum wage accounts for one half of the reduction in lower-tail inequality that occurred among women during the period between 1994 and 2003.

Keywords: minimum wage, wage inequality, employment loss, truncated distribution, deflation

JEL Classification Code: J23, J31, J38

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1 Introduction

Although many advanced industrialized countries share similar experiences of skill-biased technological change and increased exposure to international trade and outsourcing (Machin and Van Reenen, 1998), wage distributions have evolved differently among these countries. One representative example is the stark difference in changes in wage inequality between Japan and the United States. Wage inequality increased only moderately in Japan from the mid-1970s to the early 1990s, while it rose substantially in the United States (Katz, Loveman, and Blanchflower, 1995). Differences in inequality trends across countries are typically attributed to differences in wage-setting institutions, such as labor unions and minimum wages (Blau and Kahn, 1996; Freeman and Katz, 1996). When these institutions are compared between Japan and the United States, the real minimum wage deflated by the consumer price index rose by 60% in Japan from the early 1970s to the late 1990s, while it fell by 20% in the United States (OECD, 1998). DiNardo, Fortin, and Lemieux (1996), Lee (1999), and Teulings (2003) demonstrate that the erosion of the real value of the minimum wage accounts for a large part of the rise in wage inequality in the United States, but there has been a lack of formal analysis regarding the impact of the minimum wage on inequality trends in Japan.

The moderate increase in wage inequality occurred in Japan until the early 1990s for both men and women (Katz, Loveman, and Blanchflower, 1995). After the early 1990s, however, while male wage inequality increased, female wage inequality declined. To be precise, while the male wage distribution fanned out as the 10th percentile of the wage distribution declined after the late 1990s, the female wage distribution compressed as the 10th percentile of the wage distribution increased after the early 1990s (Figure 1). During the same period, the statutory minimum wage continued to increase in all prefectures, even after the median wage started to decline in the late 1990s. Since the rate of increase in the minimum wage was nearly uniform across all prefectures regardless of their differences in labor-market conditions, the 'bite' of the minimum wage rose more significantly in low-wage prefectures than in high-wage prefectures. In fact, the fraction of workers paid less than or equal to the minimum wage rose above 5% among women in some prefectures after the late 1990s, although it continued to be below 1% among men in most prefectures between 1994 and 2003 (Figure 2). While teenagers make up the majority of minimum-wage workers in the United

States (Flinn, 2010), women make up most of the minimum-wage workers in Japan. The aim of this paper is to investigate the extent to which a reduction in lower-tail inequality among women in Japan is attributable to the minimum-wage increase.

This paper's contribution to the literature is twofold. First, we examine the impact of the minimum wage on the wage distribution under wage deflation, in which the real value of the minimum wage increases even without a change in the statutory minimum wage. Since the revision of the statutory minimum wage inevitably lags behind inflation and deflation, the real value of the minimum wage typically varies with macroeconomic conditions. While the real value of the minimum wage fell under an inflationary economy in the 1980s in the United States, it rose under a deflationary economy in the 1990s in Japan. Moreover, for institutional reasons, the increase in the real value of the minimum wage was greater in low-wage prefectures than in high-wage prefectures in Japan. In this paper, following the approach developed by Lee (1999) and refined by Autor, Manning, and Smith (2010), we exploit variation in the minimum-wage bite across prefectures over time to estimate the impact of the minimum wage on the wage distribution.

Second, we quantitatively assess the impact of employment loss resulting from the minimum-wage increase on the wage distribution. Consider a simple competitive model, in which employment declines with a rise in the minimum wage as a result of a lower demand for labor. A rise in the minimum wage would then reduce employment for low-wage workers and compress the lower tail of the wage distribution. Loss of employment is widely recognized as one of the mechanisms that compresses the wage distribution, but Lee's (1999) approach is limited in its ability to determine the mechanism of wage compression. We thus introduce two nonparametric methods to measure the magnitude of the effect of employment loss on the wage distribution and deepen our understanding of how the labor market works. One method builds upon a version of the inverse probability weighting method (DiNardo, Fortin, and Lemieux, 1996), while the other builds upon a version of the trimming method (Lee, 2009). Although our approach relies heavily on these prior studies, we modify their methods to quantify the impact of employment loss resulting from the minimum-wage increase. For this purpose, we examine the adverse effects of the minimum wage on the labor market, and use the estimated employment elasticity with respect to the minimum wage to calculate the reweighting factor and the trimming threshold. Our analysis complements that of Autor, Manning, and Smith (2010), who quantify the effects of spike and spillover under the assumptions that the latent wage distribution is log-normal and that no loss of employment results from imposing the minimum wage.

This paper reveals that the increase in the minimum wage compressed the lower tail of the wage distribution among women in Japan and accounts for roughly one half of the reduction in lower-tail inequality that occurred during the 1994–2003 period. The 50–10 wage gap decreased by 7 log points during that period, but if there had been no change in the minimum wage, it would have decreased by only less than 4 log points. We also find that the increase in the minimum wage had adverse effects on new hires, hours worked, and employment for women. The loss of employment, however, only partially accounts for the compression of the lower tail of the wage distribution among women, indicating that the increase in the minimum wage resulted in an actual wage increase for low-skilled workers.

The remainder of this paper is organized as follows. Section 2 first describes data from two large-scale government surveys, one of which is an establishment survey (the Basic Survey on Wage Structure) used to examine the impact of the minimum wage on the wage distribution, new hires, and hours worked; the other is a household survey (the Employment Status Survey) used to estimate employment elasticity with respect to the minimum wage. It then introduces the minimum-wage system in Japan and finally discusses variation in the minimum-wage bite across prefectures over time. Section 3 first considers an empirical framework to examine the impact of the minimum wage on the wage distribution, then presents parameter estimates of the wage-compression effect, and finally assesses the quantitative contribution of the minimum-wage increase to changes in wage inequality. Section 4 provides an analysis of the effects on new hires, hours worked, and employment, followed by an analysis of the effect of employment loss on the wage distribution. The final section gives a summary and conclusions.

2 Evolution of the Wage Distribution

2.1 Data description

For the analysis of the wage distribution, new hires, and hours worked, we use repeated cross-sectional data from the Basic Survey on Wage Structure (BSWS) between 1994 and 2003, during which period the statutory minimum wage increased every year. The BSWS is compiled annually by the Ministry of Health, Labour and Welfare. The universe of the survey consists of private establishments with five 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, fishery, and the legislative, administrative, and judicial branches of local and national governments. Establishments are randomly selected in proportion to prefecture, industry, and establishment size according to the Establishment and Enterprise Census, which lists all establishments in Japan. Randomly selected establishments are asked to provide establishment information, such as the number of regular workers,[1] the number of new graduates hired, firm size, industry, and location, and to extract employee information, such as earnings, hours worked, employment status, age, sex, and educational attainment (only for full-time workers), from payroll records. Although board members, whose wages are set at a general meeting of shareholders, are not surveyed, all types of workers, including full- and part-time workers and regular and temporary workers, are surveyed when they are directly hired by randomly selected establishments.[2] There is neither bottom- nor top-coding. The sample of the analysis comprises approximately 800,000–890,000 male workers and 410,000–510,000 female workers from approximately 55,000 establishments for every year. We calculate hourly wages for all workers in the sample and weight them by the individual sampling weight in the analysis. Since the minimum-wage law in Japan applies to the straight wage rate, excluding allowances, we define hourly wages as scheduled earnings net of allowances divided by hours worked, where allowances include commutation allowance, perfect-attendance

[1]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. Part-time workers as well as full-time workers can be classified as regular workers if at least one of the above criteria is satisfied.

[2]A person in charge of personnel at each establishment is asked to randomly choose a certain 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.

allowance, and family allowance.[3] Although the survey is conducted between June 1 and June 30, the revised minimum wage becomes effective either September 30 or October 1. To maintain consistency between the survey date and the effective date, we merge the BSWS wage data in the current year with the statutory minimum wage data in the previous year. We thus analyze the effects of the minimum wage on the wage distribution, new hires, and hours worked eight months after the revision of the minimum wage.

For the analysis of female employment, we use repeated cross-sectional data from the Employment Status Survey (ESS) for the years 1997 and 2002. The ESS is compiled every five years by the Ministry of the Internal Affairs and Communications. The universe of the survey is all households in Japan, excluding foreign diplomats, foreign military personnel and their dependents, persons dwelling in Self Defense Force camps or ships, and persons serving sentences in correctional institutions. All household members 15 years or older are surveyed when their households are randomly selected. The ESS collects information on employment status, age, sex, educational attainment, and residential area as of October 1 of each survey year. We thus analyze the effect of the minimum wage on employment one year later. The sample of the analysis comprises approximately 500,000 and 470,000 women in 1997 and 2002, respectively. We weight all observations by the individual sampling weight in the analysis. The employment rates in the sample are 53.4% and 51.3% in 1997 and 2002, respectively.

2.2 Minimum-wage system

The minimum-wage law in Japan was enacted in 1959, under which most minimum wages were set under agreement among employers by region and industry, and a few were set under collective agreement.[4] The minimum-wage law was substantially revised in 1968 to ratify the International Labor Convention concerning the creation of minimum wage-fixing machinery. Since then, minimum wages have been decided by deliberations of the minimum-wage council. After prefectural minimum wages were set in all 47 prefectures in 1976, the *meyasu* (indication) system was introduced in 1978 as a response to the request by labor unions to reduce a large disparity in prefectural

[3] There is no custom of tipping in Japan.

[4] See Hamaguchi (2009) for minimum-wage legislation in Japan.

minimum wages. Industrial minimum wages have remained as a system to complement collective bargaining in some industries and prefectures, where industrial minimum wages were set higher than prefectural minimum wages and the unionization rate was lower than 30% in the mid-1980s.

At present, according to Article 1 of the Minimum Wage Act, the purpose of minimum-wage legislation is to guarantee a minimum amount of wages and improve working conditions for low-wage workers, thereby contributing to the stability of workers' lives, improving the quality of the labor force, protecting fair competition in business, and developing a sound national economy. Since the minimum-wage policy's goal is not limited to the welfare of minimum-wage workers, but extends to the welfare of low-wage workers in general, we do not restrict our analysis to the effects on minimum-wage workers, and primarily examine the effects on the lower tail of the wage distribution.

Under the current indication system, the central minimum-wage council classifies all 47 prefectures into four ranks and indicates the increased amount of prefectural minimum wages by rank every year. Prefectural minimum wages are then decided by deliberations of prefectural minimum-wage councils. Since the central minimum-wage council is not an ad-hoc but a standing institution, its decisions have been insusceptible to politics. In fact, Japan's minimum-wage policy has not been coordinated with any other policy, such as unemployment insurance and measures to promote small- and medium-sized enterprises. The central minimum-wage council consists of representatives of public interest (academics and retired bureaucrats), employers, and employees. Since consensus has never been reached between representatives of employers and employees, the increased amount of prefectural minimum wages each year has been presented as the view of representatives of the public interest. Prefectural minimum-wage councils almost always approve the indication of the central minimum-wage council, as evidenced by the fact that the actual increased amount of prefectural minimum wages is typically identical to the one indicated by the central minimum-wage council. Therefore, the continuous increase in prefectural minimum wages is a consequence of the current system, in which the central minimum-wage council indicates the increased amount of prefectural minimum wages not by prefecture but by rank every year, and prefectural minimum-wage councils effectively follow the central minimum-wage council's indication. The increase in prefectural minimum wages continued until the early 2000s, even after the median wage started to decline in the late 1990s.

2.3 Regional variation in the minimum-wage bite

As of 2000, four minimum wage ranks A, B, C, and D, respectively, comprised 3, 10, 18, and 16 prefectures. The three prefectures in rank A are Tokyo, Osaka, and Kanagawa (a prefecture located next to Tokyo); the 10 prefectures in rank B are typically located around Tokyo and Osaka; the 16 prefectures in rank D are located either in Tohoku (the north end district of the main island), Chugoku (the west end district of the main island), Shikoku (the fourth largest island in the southwest region of Japan), or Kyushu (the third largest island in the southwest end of Japan); and 18 prefectures in rank C are the rest (Figure 3). A total of 10 prefectures experienced a change in rank, though only by one level, during the 1994–2003 period. Since the intended purpose of the indication system was to moderate disparity in prefectural minimum wages, the rate of increase in prefectural minimum wages was greatest for the lowest-ranked prefectures. The difference between the lowest rank and the other three ranks was only modest, however, and the difference between the three ranks excluding the lowest rank was negligible. To be precise, the average increasing rates of prefectural minimum wages between 1994 and 2003 were 13.9, 14.0, 13.9, and 14.4% in ranks A, B, C, and D, respectively. Since the rate of increase in prefectural minimum wages was nearly uniform regardless of local labor-market conditions, prefectural minimum wages still differ greatly across prefectures, even a few decades after the introduction of the indication system, and moreover, the real value of the minimum wage increased more substantially in low-wage prefectures than in high-wage prefectures.

Figure 4 illustrates the log wage distributions for women in two prefectures selected by rank in 1994 and 2003. We select two prefectures that have either one of the highest or one of the lowest proportion of minimum-wage workers in each rank, in order to highlight regional and time variation in the minimum-wage bite within and between ranks. For both rank A prefectures, such as Tokyo and Osaka, and rank B prefectures, such as Shizuoka and Aichi, the log wage distributions were bell-shaped in 1994 but lost their shape in 2003 in the prolonged recession. Nonetheless, the proportion of workers paid the minimum wage was very small in these higher-ranked prefectures. In contrast, the proportion of workers paid the minimum wage was relatively high in some lower-ranked prefectures. While in one rank-C prefecture, Miyagi, there was no concentration of workers around the minimum wage in either 1994 or 2003, in another rank-C prefecture, Hokkaido, the wage density had already spiked moderately in 1994 and was highest in 2003 at the minimum wage. The wage density spiked even more evidently in a couple of rank-D prefectures, such as Okinawa (and also Aomori), where many workers were already working at the minimum wage in 1994. In another rank-D prefecture, Yamagata, however, the minimum wage did not bind in either 1994 or 2003. To summarize, the minimum-wage bite was fairly high in some, if not all, lower-ranked prefectures, especially after 2000, while it was relatively negligible in higher-ranked prefectures both in the 1990s and after 2000. Most employers appear to have complied with the minimum-wage law. If an employer's noncompliance is detected by the prefectural labor standards inspection office, the employer must pay a fine of up to 20,000 yen and compensate employees for the difference between the minimum wage and the actual wage. Even without detection, the employer's reputation would be severely damaged if the public becomes aware that employees are paid less than the minimum wage.

Some studies point to spikes in the wage distribution at the minimum wage in Japan. Hori and Sakaguchi (2005) illustrate the wage distribution by prefecture and industry separately for full- and part-time workers using the 2003 BSWS. Nonetheless, such a large regional difference in the change of the wage distribution associated with the minimum wage has not been documented. Furthermore, none of the studies estimate the impact of the minimum wage on the entire wage distribution in Japan.

3 Impact of the Minimum Wage on the Wage Distribution

3.1 Model specification and identification

The minimum-wage hike affects the wage distribution through three channels: employment loss (truncation), spike (censoring), and spillover (see Card and Krueger 1995; Brown, 1999; Neumark and Wascher, 2008 for the survey). First, consider a competitive labor market, in which homogeneous workers are paid the value of their marginal product of labor. The minimum-wage hike will then lead to a loss of employment for workers paid less than the new minimum wage and a truncation of the wage distribution below the minimum wage.[5] Second, suppose that firms retain workers paid less than the minimum wage by cutting fringe benefits and training programs, requiring a higher level of work effort, reducing profits, or passing on the cost to consumers. The minimum-wage hike will then raise the wages of these low-skilled workers up to the minimum wage, and the wage density will spike at the minimum wage. Finally, suppose that there is a certain degree of substitutability between workers with different skills. As the minimum wage raises the costs of hiring minimum-wage workers, there will be a higher demand for workers who are more skilled than minimum-wage workers. The effect of the minimum-wage hike will then spill over to the wages of workers earning more than the minimum wage. In all three cases, the minimum-wage increase pushes up the wages of workers at the bottom end of the wage distribution and compresses the lower tail of the wage distribution.[6]

We now set up the model relating the minimum wage to the wage distribution. Let w_{it}^{p} (w_{it}^{q}) denote the pth (qth) percentile of the wage distribution in prefecture i in year t, mw_{it} the minimum wage, and x_{it} other determinants of the wage distribution. Following Lee (1999), we describe the theoretical relation between the minimum-wage bite and the wage differentials as follows:

$$\ln w_{it}^{p} - \ln w_{it}^{q} = f\left(\ln mw_{it} - \ln w_{it}^{q}, x_{it}\right), \tag{1}$$

where f is a function that maps the minimum-wage bite onto the p–q log wage differential. Since the minimum-wage bite varies not only by the statutory minimum wage but also by the wage distribution across prefectures over time, the minimum-wage bite is measured by the log differential between the minimum wage and the qth percentile wage, which is referred to as the ‘effective’ minimum wage. In this paper, we look at the effect of the minimum wage on all percentiles of the wage distribution from $p = 5$ to $p = 95$, in order to examine the presence and size of the spillover effect and determine a valid specification. The effect of the minimum wage may appear not only in the 10th percentile but also in the 20th or higher percentiles though spillover effects, though it is

[5] In monopsonistic models, whether the minimum-wage hike will decrease employment depends on the degree of monopsony power (Bhaskar and To, 1999; Manning, 2003).

[6] The truncation, censoring, and spillover effects described here are more in line with a competitive model. These effects, however, could be explained not only by a competitive model with imperfect substitution between workers with different skills (Teuling, 2003; Aaronson and French, 2007), but also by monopsonistic models (Bhaskar and To, 2003; Manning, 2003; Aaronson and French, 2007), a search model with bargaining (Flinn, 2010), and a tournament model (Lazear and Rosen, 1981).

unlikely to appear in the 80th or higher percentiles. Since spillover effects are considered to occur by the change of the demand for workers who are substitutes for minimum-wage workers (Teulings, 2003) and the outside option in wage bargaining between firms and workers (Flinn, 2010), the higher the percentile the weaker is the spillover effect. When we specify the regression model, we must set the qth percentile to be high enough so as not to be subject to the influence of the minimum wage. Since the qth percentile is used to normalize the minimum wage, if the minimum wage affects the qth percentile, the estimated coefficient on the effective minimum wage cannot be interpreted as the effect of the minimum wage on the pth percentile. Lee (1999) sets the qth percentile to be the median, but then finds a significant effect of the minimum wage on upper-tail wage inequality in the United States in some specifications. To circumvent this problem, Bosch and Manacorda (2010) adopt the 70th percentile as the qth percentile. Since we find similar results in Japan when we set the qth percentile to be the median, we adopt the 70th percentile as the qth percentile.

Figure 5 plots the p–70 log wage differential along with the effective minimum wage for $p =$ 10, 20, 80, and 90 in 1994 and 2003. Both the 10–70 and the 20–70 log wage differentials are positively associated with the effective minimum wage, whereas neither the 80–70 nor the 90–70 log wage differentials vary with the effective minimum wage. Moreover, the 10–70 log wage differential is more strongly associated with the effective minimum wage than the 20–70 log wage differential. The figure confirms the presence of spillover effects and the validity of the specifications.

Based on the observations in Figure 5 that the minimum-wage effect on lower-tail inequality is well approximated by a quadratic function, we estimate a model of the form:

$$\ln \widetilde{w}_{it}^{p} = \beta_1^p \ln \widetilde{mw}_{it} + \beta_2^p \left(\ln \widetilde{mw}_{it}\right)^2 + x_{it}\gamma^p + u_{it}^p \tag{2}$$

for $i = 1, 2, \ldots, 46, 47$ and $t = 1994, 1995, \ldots, 2002, 2003$, where $\widetilde{w}_{it}^{p} = w_{it}^{p} / w_{it}^{70}$, $\widetilde{mw}_{it} = mw_{it} / w_{it}^{70}$, and u_{it}^p is the error term. The vector of controls x_{it} includes year effects, prefecture effects, prefecture-specific linear trends, and the age, education, and industry composition of the workforce. The age, education, and industry composition of the workforce is measured by the share of workers for seven 10-year band age groups, four education groups, and 16 industry groups. We calculate mw_{it} by first assigning industrial minimum wages to workers in industries covered by industrial minimum wages and prefectural minimum wages to all other workers and then averaging their minimum wages by prefecture and year. As described above, there is large cross-sectional variation in prefectural minimum wages originating from their difference at the time of the legislation in the mid-1970s but not much time variation in prefectural minimum wages, except that the lowest-ranked prefectures experienced a slightly greater increase in prefectural minimum wages than prefectures in the other three ranks. Since the proportion of workers to whom industrial minimum wages apply is very small, industrial minimum wages do not create any additional significant variation. In contrast, as we have seen in Figure 4, there is a great deal of variation in the wage distribution across prefectures and time. Thus, after conditioning prefecture fixed effects, a large part of the variation in the effective minimum wage is attributable to variation in the wage distribution across prefectures over time.

We begin our analysis by estimating the benchmark model, in which we control only for year effects. The ordinary least squares (OLS) estimation presumably suffers from bias because of heterogeneity in the dispersion of the latent wage distribution. Suppose that the minimum wage is more binding in prefectures, in which the wage distribution would be more dispersed if there were no minimum wage. The direction of the bias would then be downward for $p < q$ and upward for $p > q$, since the p–q wage gap is greater in absolute value in prefectures with higher wage dispersion. We thus include prefecture effects, prefecture-specific trends, and the age, education, and industry composition of the workforce to account for heterogeneity in the latent wage distribution. Prefecture effects are unobserved but can be cancelled out by taking a first difference between year t and year $t-1$.

$$\triangle \ln \widetilde{w}_{it}^{p} = \beta_1^p \triangle \ln \widetilde{mw}_{it} + \beta_2^p \triangle \left(\ln \widetilde{mw}_{it}\right)^2 + \triangle x_{it}\gamma^p + \triangle u_{it}^p, \tag{3}$$

where $\triangle x_{it}$ includes year effects, prefecture effects (as a difference in prefecture-specific linear trends), and changes in the age, education, and industry composition of the workforce. The first-difference (FD) estimation, however, can exacerbate bias because of measurement error in the percentile of the wage distribution caused by sampling (Lee 1999; Autor, Manning, and Smith, 2008). As seen in equations (2) and (3), the qth percentile appears as a denominator in both the dependent and independent variables. In this case, measurement error in the qth percentile can cause an upward bias. To circumvent this measurement error problem, we perform an instrumental variable (IV) estimation where we use the lagged value of the effective minimum wage as an instrument for $\triangle \ln \widetilde{mw}_{it}$. We make two assumptions to ensure that the instrument satisfies the exclusion restriction. First, the effect of the minimum wage is (approximately) linear, i.e., $\beta_2^p = 0$.[7][8] Second, the measurement error takes the multiplicative form, such that $w_{it}^r = w_{it}^{r*} c_i^r e_{it}^r$ for $r = p, q$, where the observed rth percentile consists of the actual rth percentile (w_{it}^{r*}), the persistent measurement error (c_i^r), and the i.i.d. measurement error (e_{it}^r). Under this assumption, the measurement error ($c_i^r e_{it}^r$) is serially correlated, but the change in its log is i.i.d., and thus we can control for the persistent measurement error caused by sampling. After substituting the restrictions imposed by the two assumptions into equation (4), we can rewrite the estimating equation as follows:

$$\triangle \ln \widetilde{w}_{it}^p = \beta_1^p \triangle \ln \widetilde{mw}_{it} + \triangle x_{it} \gamma^p + \triangle v_{it}^p, \tag{4}$$

where $\triangle v_{it}^p = \triangle u_{it}^p + \triangle \ln e_{it}^p - (1 + \beta_1^p) \triangle \ln e_{it}^{70}$, and the error term comprises only i.i.d. errors that are not correlated with the lagged value of the effective minimum wage. More specifically, the instrument we use is $\left(\ln \overline{mw}_{i,t-2} - \ln w_{i,t-2}^{70}\right)$, where $\overline{mw}_{it}$ is prefectural minimum wages calculated based on the increased amount of prefectural minimum wages that the central minimum-wage council indicates by rank every year. Since we do not use variation in prefectural minimum wages that were set by prefectural minimum-wage councils, the IV estimates are consistent even in the presence of endogeneity of prefectural minimum-wage councils' decisions. The first-stage *F* statistic is 7.26, indicating that the instrument also satisfies the rank condition.

[7] The instrumental-variable estimator is generally inconsistent in nonlinear models with measurement error (Chen, Hong, and Nekipelov, 2011).

[8] Since there is no substantial difference in marginal effects and counterfactual wage changes regardless of whether the quadratic term is included or excluded in the OLS and FD estimations, the potential bias because of the linear specification is expected to be quantitatively marginal in our analysis.

3.2 Parameter estimates

Table 1 reports the marginal effects of the minimum wage on the p–70 wage gap for $p = 10$, 20, 30, 40, 50, 60, 80, 90, when the marginal effects are evaluated at the sample mean of the effective minimum wage across all prefectures and years.[9] Since both the wage gap and the effective minimum wage are measured in log terms, the marginal effects indicate the percentage changes in the wage gap in response to a one-percent increase in the minimum wage. The estimated minimum-wage effects are greatest for the 10th percentile and shrink gradually as the percentile becomes higher in the OLS estimation (column 1). R^2 is also largest at 0.74 for the 10th percentile and monotonically declines to 0.04 as it approaches the 90th percentile. The estimated minimum-wage effects are statistically significant in the lower tail but not in the upper tail of the wage distribution. Given the fact that the fraction of minimum-wage workers is at most 10% in Okinawa, these results confirm the presence of spillover effects. After controlling for prefecture effects and prefecture-specific trends, the estimated minimum-wage effects become larger in the FD estimation (column 2). This result remains the same regardless of including or excluding prefecture-specific trends. After further controlling for the workforce's age, education, and industry composition, the estimated minimum-wage effects increase slightly more (column 3). The results described thus far are consistent with those of Lee (1999), who estimates the impact of the minimum wage on the entire wage distribution in the United States with and without state fixed effects. The increase in the estimated minimum-wage effects, however, may be attributable to the upward bias arising from the measurement error. Indeed, when the IV estimation is performed to circumvent the measurement-error problem, the minimum-wage effects turn out to be smaller. The IV estimates lie between the OLS estimates and the FD estimates for the effects on lower-tail inequality, although they are imprecise for the effects on upper-tail inequality. The IV estimates obtained in this paper are somewhat larger than those in Autor, Manning, and Smith (2010) but comparable with those in Bosch and Manacorda (2010), who estimate the impact of the erosion of the minimum wage in Mexico. As it is documented that the minimum wage traditionally serves as a scale to determine wages as well as welfare benefits in Mexico, there is anecdotal evidence suggesting that the minimum wage

[9]The marginal effect is $\widehat{\beta}_1^p + 2\widehat{\beta}_2^p \overline{\ln \widetilde{mw}}$, where $\overline{\ln \widetilde{mw}}$ is the sample mean of the effective minimum wage across all prefectures and years.

has been used as a scale to determine the wages of part-time workers and temporary workers in Japan.[10]

When we repeat the same analysis for men, we obtain similar results in that the estimated minimum-wage effects are greatest for the 10th percentile and shrink gradually as the percentile becomes higher. The results are less compelling, however, in that the estimated minimum-wage effects are rather imprecise in the IV estimation and sometimes statistically significant for the effects on upper-tail inequality in the FD estimation. Nonetheless, given the fact that women make up most of the minimum-wage workers in Japan, the minimum wage is obviously much more important for the female wage distribution. Therefore, we evaluate the impact of the minimum wage according to the analysis for women in this paper.

The minimum-wage increase may cause firms to reduce the level of allowances. We examine the offsetting effect on allowances for women by calculating the p–70 wage gap from the wage rate, including commutation allowance, perfect-attendance allowance, and family allowance. We then find that the OLS, FD, and IV estimates, respectively, decrease by 3.0, 8.5, and 14.0 percentage points for the minimum-wage effect on the 10–70 wage gap and by 2.8, 9.0, and –3.0 percentage points for the minimum-wage effect on the 20–70 wage gap. These results indicate that the wage-compression effect is slightly attributable to this offsetting effect.

3.3 Inequality trends

The parameter estimates presented above consistently indicate that the minimum-wage increase resulted in a reduction in lower-tail inequality among women under various specifications. To assess the quantitative contribution of the minimum-wage increase to inequality trends, we calculate counterfactual wages as if there had been no change in the effective minimum wage during the 1994–2003 period, based on the OLS, FD, and IV estimations, whose results are reported in columns 1, 2, and 4 of Table 1. For a worker k whose hourly wages rank at the pth percentile in prefecture i in year t, the counterfactual log wages in 2003 can be calculated by subtracting the

[10] According to the interviews the author conducted with people in charge of personnel management in restaurant chains, convenient store chains, supermarkets, security companies, and building maintenance companies in Sapporo, Hokkaido in July 2010, the wages of part-time workers start at 680 yen per hour, with reference to the minimum wage of 678 yen, and increase with job tenure in all the companies. Employers say that they have changed their wage scale according to the revision of the minimum wage.

change in the pth percentile wage resulting from the minimum-wage increase from the actual log wages in 2003 as follows:

$$\widehat{\ln w}^{p}_{k,i,2003} = \ln w^{p}_{k,i,2003} - \widehat{\beta}^{p}_{1} \left(\ln \widetilde{mw}_{i,2003} - \ln \widetilde{mw}_{i,1994}\right) - \widehat{\beta}^{p}_{2} \left[\left(\ln \widetilde{mw}_{i,2003}\right)^2 - \left(\ln \widetilde{mw}_{i,1994}\right)^2\right] \tag{5}$$

for $p = 5, 6, \ldots, 94, 95$.

Figure 6 illustrates the actual and counterfactual changes in log wages between 1994 and 2003 by percentile of the wage distribution. In the figure, the effect of the minimum wage is represented as the difference between the actual and counterfactual changes for each percentile. All three estimates of the counterfactual wage changes are significantly smaller than the actual wage changes in the lower tail of the wage distribution but not in the upper tail of the wage distribution, indicating that the minimum wage contributed to the increase in low-wage workers' wages. At least one half of the increase in the wages of workers below the 20th percentile is attributable to the minimum-wage increase.

Table 2 reports the actual and counterfactual changes in wage inequality measured by the p–50 wage gap for $p =$ 10, 20, 30, 40, 60, 70, 80, 90. Based on the IV estimates, the minimum-wage increase accounts for 43% (63%) of the reduction in the gap between the 50th percentile wage and the 10th (20th) percentile wage but for none of the change in the gap between the 50th percentile wage and the 30th or higher percentile wage. Based on the OLS and FD estimates, the minimum-wage increase accounts for an even greater part of the reduction in the 50–10 and the 50–20 wage gaps.

The results of the counterfactual analysis in this paper are more moderate than those in Lee (1999), who argues that a rise in wage inequality in the United States is explained mostly by the erosion of the minimum wage. His argument is drawn from an analysis of counterfactual wages calculated based on the OLS estimation when state effects are not included. In light of our earlier results, we consider the difference in the results to be largely attributable to the difference in econometric specifications, i.e., the choice of qth percentile and the use of an instrumental variable.

4 Employment Loss and Its Impact on the Wage Distribution

4.1 New hires, hours worked, and employment

The minimum wage provided a wage floor for female workers during the period of recession after the 1990s. As a consequence, wages of workers at the bottom end of the wage distribution increased. This brings up the question of how firms reacted to such an increase in labor costs. Many studies examine employment adjustments in response to the minimum-wage increase among the possible responses, in order to gauge the degree of competition in labor markets and reveal an unintended adverse effect of minimum-wage legislation.[11] One reason for mixed results on the employment effect of the minimum wage in the literature is that employment-adjustment costs are asymmetric between hiring and firing because of government regulations that levy high firing costs on firms. Basically, the employment effect of the minimum wage should be less evident when examining how much employment declines in response to the minimum-wage increase than when examining how much employment expands in response to the erosion of the minimum wage. Given the fact that protection for regular employees is stringent in Japan (OECD, 2004),[12] and the minimum wage has continuously increased, firms may adjust the quantity of labor input by reducing the number of new hires and the number of hours worked for existing employees. We thus begin the analysis for the (dis-)employment effect of the minimum wage by looking at the effects of the minimum wage on the number of new graduate hires ($hire$) and the average hours worked ($hour$) for women in an establishment j in year t using the BSWS.

$$hire_{jt} = \alpha_1 \ln \widetilde{mw}_{it} + x_{jt}\pi_1 + v_{1jt},$$

[11] Brown, Gilroy, and Kohen (1982) summarize time-series studies as suggesting that the estimated elasticities of teenage employment with respect to the minimum wage range from –0.1 to –0.3 in the United States. Neumark and Wascher (2007) review panel data studies and case studies beginning in the early 1990s and document a wide range of estimates of employment elasticity from near minus one to above zero. Also, a few studies estimate the employment effect of the minimum wage in Japan. Kawaguchi and Yamada (2006) employ an approach similar to that of Currie and Fallick (1996) and find a negative employment effect of the minimum wage for women. Kawaguchi and Mori (2009) use a similar approach to that developed by Card (1992) and show that the minimum-wage hike reduced the employment of male teenagers and middle-aged married women.

[12] 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.

$$\ln hour_{jt} = \alpha_2 \ln \widetilde{mw}_{it} + x_{jt}\pi_2 + v_{2jt},$$

where $\widetilde{mw}_{it} = mw_{it} / w_{it}^{50}$, x_{jt} is a vector of controls, which include the number of regular employees, fourth-order polynomials in the average employee age, prefecture effects, year effects, and prefecture-specific linear time trends, and v_{1jt} and v_{2jt} are the error terms. Here, we normalize the minimum wage by the 50th percentile wage instead of the 70th percentile wage, since $\widetilde{mw}_{it}$ comes closer to the conventional measure of the minimum-wage bite, known as the Kaitz index in the analysis of the effect on employment. We are not concerned about the choice of percentile, since it appears only on the right-hand side. Table 3 indicates that the adverse effects of the minimum wage are not negligible. The new-hires elasticity is large at –1.84, and the hours elasticity is moderate at –0.115.

We then look at the effect of the minimum wage on the probability of being employed for a woman k of age group a in year t, using the ESS.

$$\Pr\left(e_{kt} = 1 | \widetilde{mw}_{ait}, x_{kt}\right) = \Phi\left(\alpha_3 \widetilde{mw}_{ait} + x_{kt}\pi_3\right), \tag{6}$$

where $\widetilde{mw}_{ait} = mw_{it} / w_{ait}^{50}$, e_{kt} is an indicator for being employed, x_{kt} is a vector of controls, which include fourth-order age polynomials, prefecture effects, year effects, and prefecture-specific linear time trends, and Φ represents the standard normal cumulative distribution function. As discussed in Card and Krueger (1995), the reduced-form equation (6) can be viewed as an approximation of the function of the demand for minimum-wage workers relative to median-wage workers. We exploit variation in the effective minimum wage across 10-year band age groups to identify the disemployment effect of the minimum wage after flexibly controlling for age effects. Since the minimum wage potentially affects school enrollment (Card and Krueger, 1995; Neumark and Wascher, 2008), we neither include education as a control nor exploit variation in the effective minimum wage across education groups. We find that employment elasticity with respect to the minimum wage is –0.340 when evaluated at the sample mean of the effective minimum wage across all prefectures and years, and it ranges from –0.407 in Okinawa to –0.267 in Tokyo when evaluated at the sample means by prefecture. The results remain unchanged regardless of including or excluding prefecture-specific trends and quadratic trends.

4.2 Reweighting method

The analysis thus far has demonstrated that three effects (truncation, censoring, and spillover effects) underlie the impact of the minimum wage on the wage distribution. We now turn to the analysis for quantifying the magnitude of the truncation effect.[13] By doing so, we can better understand how the labor market works. Recognizing that the problem we consider here is a sample-selection problem, we can infer the extent to which wage compression is attributable to the loss of employment by correcting for the selection bias arising from the truncation effect. To do so, we consider the counterfactual wage density that is not subject to the influence of employment loss resulting from the minimum-wage increase. Let $f\left(w_t|\,\widetilde{mw}_t, x_t\right)$ denote the actual wage density conditional on the effective minimum wage $\widetilde{mw}$ and observed attributes x in year t, $f\left(w_t|\,\widetilde{mw}_{1994}, x_t\right)$ the counterfactual wage density if there were no change in the effective minimum wage since 1994, and e_t an indicator for being employed. We make two assumptions to examine the extent to which wage compression is attributable solely to the truncation effect. First, there is neither a spillover effect nor a censoring effect: $g\left(w_t|\,e_t = 1, \widetilde{mw}_t, x_t\right) = g\left(w_t|\,e_t = 1, \widetilde{mw}_{1994}, x_t\right)$. Second, the minimum wage has no direct effect on the distribution of observed attributes: $h\left(x_t|\,\widetilde{mw}_t\right) = h\left(x_t|\,\widetilde{mw}_{1994}\right)$. As derived in the Appendix, the counterfactual density can be nonparametrically constructed as follows:

$$f\left(w_t|\,e_t = 1, \widetilde{mw}_{1994}\right) = \int g\left(w_t|\,e_t = 1, \widetilde{mw}_t, x_t\right) h\left(x_t|\,e_t = 1, \widetilde{mw}_t\right) \theta\left(\widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right) dx, \tag{7}$$

where

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

The reweighting function is essentially the ratio of the counterfactual employment rate without a change in the effective minimum wage to the actual employment rate. In this reweighting procedure, we give more weight to workers who would have a higher propensity to be employed without

[13]It can be shown that truncation mechanically reduces lower-tail inequality as follows. For an arbitrary continuous distribution, it must be that $\int_{w^p}^{w^q} f\left(w\right) dw = \left(q-p\right)/100$ for $p < q$, where w is the log hourly wages, w^p and w^q are the pth and qth percentiles of the wage distribution, and $f\left(\cdot\right)$ is the probability density function. For the distribution truncated below the minimum wage mw, it must be that $\int_{w_*^p}^{w_*^q} f_*\left(w\right) dw = \left(q-p\right)/100$, where w_*^p and w_*^q are the pth and qth percentiles of the truncated distribution. Then, $f_*\left(w\right) = f\left(w\,|w \geq mw\right) = f\left(w\right)/\Pr\left(w \geq mw\right) \geq f\left(w\right)$. Thus, $w^q - w^p \geq w_*^q - w_*^p$.

the minimum-wage increase (i.e., to those who are more likely to disappear from the labor market as a result of the minimum-wage increase) to control for the truncation effect. The estimate of each response probability can be obtained from probit model (6).

The first three columns of Table 4 present the estimated minimum-wage effects reproduced after reweighting. The results are almost identical to those reported in Table 1, indicating that the truncation effect is minimal if employment loss occurs with the same probability for workers with the same attributes.

4.3 Trimming method

The results reproduced after reweighting suggest that employment loss has a minimal impact on the wage distribution, but the reweighting method would understate the truncation effect if employment loss occurs exclusively from the bottom end of the wage distribution. We thus employ a more conservative approach that enables us to see the upper bound of the truncation effect. Suppose that workers lose their jobs as a result of the minimum-wage increase in the current year if they were paid less than a certain percentile of the wage distribution in the previous year. We can then control for the truncation effect to the maximum extent by excluding workers at the bottom end of the wage distribution from the sample in the previous year. While the reweighting method assumes that employment loss occurs with the same probability for workers with the same attributes, the trimming method considers the case when employment loss occurs in the order of wage percentile from the bottom end of the distribution. We calculate the threshold for trimming the wage distribution by prefecture and year by multiplying the year-by-year percentage change in the effective minimum wage by the estimated elasticities with respect to the minimum wage, ranging from –0.407 in Okinawa to –0.267 in Tokyo. We assign the trimming threshold to zero if the year-by-year percentage change in the effective minimum wage is not positive. The estimated trimming threshold then ranges from the zero to the fourth percentiles. The percentage of prefectures where the wage distribution is trimmed at any percentile is 19, 23, 34, 70, 32, 47, 28, 55, 19, and 0% in respective years from 1991 to 2003. The trimming threshold is zero in the last year of the sample period, because it is the base year with the highest minimum wage. In most cases of trimming, the trimming threshold is the first percentile. The percentage of prefectures where the wage distribution is trimmed at the first percentile is 17, 19, 34, 60, 26, 38, 28, 40, 17, and 0% in respective years from 1991 to 2003.

The last three columns of Table 4 present the estimated minimum-wage effects reproduced after trimming. Although the OLS estimates are almost unchanged, the FD and IV estimates of the minimum-wage effect on the 10th percentile wage decrease by 28% and 6%, respectively. After controlling for the truncation effect, there is less difference in the minimum-wage effect on the 10th percentile wage among the three estimates. While the FD estimates indicate that employment loss accounts for a significant part of the minimum-wage effect on the 20th and higher percentiles, the IV estimates indicate that employment loss does not account for the minimum-wage effect on the 20th and higher percentiles. In sum, we confirm that, while the truncation effect is present, its magnitude is limited, even if employment loss occurs exclusively from the bottom end of the wage distribution. Therefore, the wage compression is not just the result of a mechanical effect arising from the loss of employment.

5 Summary and Conclusions

In this paper, we have examined the impact of the minimum wage on the wage distribution and the mechanisms for the wage compression that occurred during the 1994–2003 period among women in Japan, one of the world's largest economies. For institutional reasons, the statutory minimum wage continuously increased in all prefectures, even during a period of deflation, and consequently, the minimum-wage bite substantially increased among women in low-wage prefectures. Japan's experience since the 1990s mirrors the U.S. experience in the 1980s and 1990s. The continuous increase in the minimum wage resulted in a reduction in wage inequality in Japan, while the erosion of the minimum wage led to a rise in wage inequality in the United States. Our analysis revealed that the increase in the minimum wage accounts for approximately one half of the reduction in 50–10 wage gap among women and that a large part of the wage compression is not attributable to the loss of employment resulting from the minimum-wage increase. We also found that the increase in the minimum wage had adverse effects on new hires, hours worked, and employment. To conclude, the minimum wage provided a wage floor for women in Japan's deflationary economy. This benefit of the minimum-wage system, however, came at the expense of employment opportunities. The findings of this paper confirm the policy trade-off between employment opportunities and wage differentials.

Appendix

Derivation of equation (7).

The counterfactual wage density without any change in the effective minimum wage is written as

$$f\left(w_t \middle| e_t = 1, \widetilde{mw}_{1994}\right) = \int g\left(w_t \middle| e_t = 1, \widetilde{mw}_{1994}, x_t\right) h\left(x_t \middle| e_t = 1, \widetilde{mw}_{1994}\right) dx, \tag{8}$$

where $g\left(w_t \middle| e_t = 1, \widetilde{mw}_{1994}, x_t\right)$ is the counterfactual wage density conditional on attributes x, and $h\left(x_t \middle| e_t = 1, \widetilde{mw}_{1994}\right)$ is the counterfactual density of attributes. Under the first assumption,

$$f\left(w_t \middle| e_t = 1, \widetilde{mw}_{1994}\right) = \int g\left(w_t \middle| e_t = 1, \widetilde{mw}_t, x_t\right) h\left(x_t \middle| e_t = 1, \widetilde{mw}_{1994}\right) dx. \tag{9}$$

The second assumption implies

$$h\left(x_t \middle| e_t = 1, \widetilde{mw}_{1994}\right) = \frac{\Pr\left(e_t = 1 \middle| \widetilde{mw}_{1994}, x_t\right) h\left(x_t \middle| \widetilde{mw}_t\right)}{\Pr\left(e_t = 1 \middle| \widetilde{mw}_{1994}\right)}. \tag{10}$$

By applying Bayes's rule,

$$\begin{aligned} h\left(x_t \middle| e_t = 1, \widetilde{mw}_t\right) &= \frac{\Pr\left(e_t = 1 \middle| \widetilde{mw}_t, x_t\right) h\left(x_t \middle| \widetilde{mw}_t\right)}{\Pr\left(e_t = 1 \middle| \widetilde{mw}_t\right)} \\ h\left(x_t \middle| \widetilde{mw}_t\right) &= h\left(x_t \middle| e_t = 1, \widetilde{mw}_t\right) \cdot \frac{\Pr\left(e_t = 1 \middle| \widetilde{mw}_t\right)}{\Pr\left(e_t = 1 \middle| \widetilde{mw}_t, x_t\right)}. \end{aligned} \tag{11}$$

Substituting (11) into (10) yields

$$h\left(x_t \middle| e_t = 1, \widetilde{mw}_{1994}\right) = h\left(x_t \middle| e_t = 1, \widetilde{mw}_t\right) \theta\left(\widetilde{mw}_{1994}, \widetilde{mw}_t, x_t\right), \tag{12}$$

where

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

Substituting (12) into (9) yields

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

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Table 1: Marginal effects of the minimum wage on the wage distribution

log wage differentials	OLS	FD	FD	IV
	(1)	(2)	(3)	(4)
10–70	0.552	0.887	0.928	0.642
	(0.047)	(0.067)	(0.049)	(0.199)
20–70	0.379	0.751	0.804	0.437
	(0.049)	(0.060)	(0.053)	(0.232)
30–70	0.234	0.530	0.602	0.495
	(0.048)	(0.034)	(0.046)	(0.137)
40–70	0.127	0.362	0.464	0.356
	(0.045)	(0.038)	(0.039)	(0.182)
50–70	0.056	0.260	0.360	0.357
	(0.036)	(0.049)	(0.049)	(0.208)
60–70	0.013	0.104	0.177	0.216
	(0.020)	(0.045)	(0.042)	(0.157)
80–70	0.015	0.023	0.052	–0.132
	(0.022)	(0.036)	(0.038)	(0.185)
90–70	0.026	0.070	0.080	–0.117
	(0.039)	(0.070)	(0.085)	(0.310)
year effects	yes	yes	yes	yes
prefecture effects		yes	yes	yes
prefecture-specific trends		yes	yes	yes
other controls			yes	

Notes: Standard errors in parentheses are clustered at the prefecture level. The sample includes 470 prefecture-year observations. Each observation is weighted by the sum of individual sampling weights by prefecture and year. Other controls include the shares of workers by age, education, and industry. In the last column, the IV estimation is performed after first-differencing using the log differential between two years lags of indicated prefectural minimum wages and the 70th percentile wage as instrument.
Data source: BSWS.

Table 2: Actual and counterfactual changes in log wage differentials from 1994 to 2003

log wage differentials	Actual	Counterfactual		
		OLS	FD	IV
	(1)	(2)	(3)	(4)
10–50	0.032	0.010	0.004	0.018
20–50	0.017	–0.001	–0.008	0.006
30–50	–0.003	–0.011	–0.017	–0.008
40–50	–0.007	–0.009	–0.013	–0.008
60–50	0.018	0.019	0.023	0.023
70–50	0.033	0.033	0.039	0.041
80–50	0.040	0.035	0.044	0.049
90–50	0.042	0.021	0.029	0.042

Notes: Counterfactual changes in the log wage differentials in columns 2, 3, and 4 are calculated based on the specifications used to produce the results in columns 1, 2, and 4 of Table 1, respectively.
Data source: BSWS.

Table 3: Elasticities of new hires, hours, and employment with respect to the minimum wage

	New Hires	Hours	Employment
	(1)	(2)	(3)
Elasticity	–1.84	–0.115	–0.340
	(1.01)	(0.044)	(0.121)

Notes: Minimum wage elasticities are evaluated at the sample means. Standard errors are clustered at the prefecture level. The sample includes 518,502 establishment-year observations in columns 1 and 2 and 972,479 individual-year observations in column 3. The regressions are weighted by the sum of individual sampling weights by establishment and year in columns 1 and 2 and by the individual sampling weight in column 3.
Data Sources: BSWS and ESS.

Table 4: Marginal effects of the minimum wage after controlling for the truncation effect

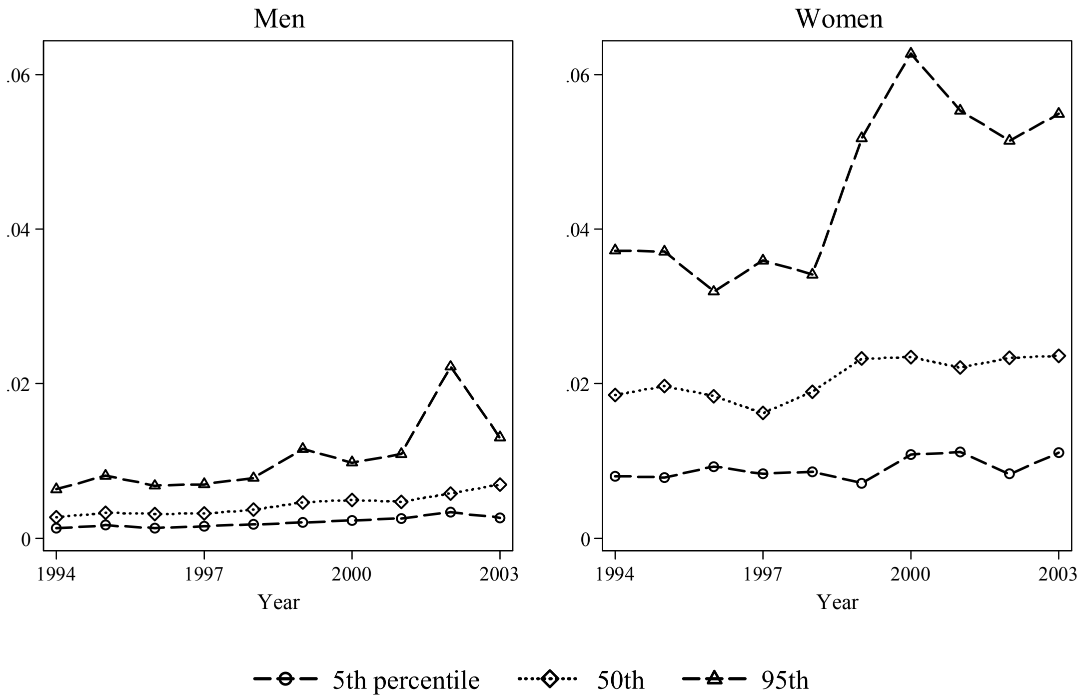
	Reweighting			Trimming		
log wage differentials	OLS	FD	IV	OLS	FD	IV
	(1)	(2)	(3)	(4)	(5)	(6)
10–70	0.553	0.888	0.636	0.539	0.641	0.604
	(0.045)	(0.066)	(0.193)	(0.046)	(0.066)	(0.307)
20–70	0.380	0.750	0.435	0.374	0.575	0.532
	(0.047)	(0.059)	(0.225)	(0.047)	(0.059)	(0.339)
30–70	0.236	0.530	0.485	0.231	0.386	0.681
	(0.045)	(0.033)	(0.132)	(0.046)	(0.040)	(0.340)
40–70	0.129	0.364	0.344	0.124	0.249	0.428
	(0.042)	(0.037)	(0.172)	(0.043)	(0.039)	(0.334)
50–70	0.059	0.263	0.338	0.055	0.176	0.506
	(0.034)	(0.048)	(0.193)	(0.035)	(0.049)	(0.408)
60–70	0.014	0.108	0.199	0.013	0.068	0.296
	(0.019)	(0.045)	(0.147)	(0.020)	(0.042)	(0.290)
80–70	0.013	0.022	–0.110	0.016	0.048	–0.217
	(0.021)	(0.035)	(0.170)	(0.021)	(0.032)	(0.331)
90–70	0.024	0.068	–0.087	0.025	0.087	–0.291
	(0.037)	(0.069)	(0.284)	(0.038)	(0.063)	(0.534)
year effects	yes	yes	yes	yes	yes	yes
prefecture effects		yes	yes		yes	yes
prefecture-specific trends		yes	yes		yes	yes

Notes: Standard errors in parentheses are clustered at the prefecture level. The sample includes 470 prefecture-year observations. Each observation is weighted by the sum of individual sampling weights by prefecture and year. In the regressions whose results are reported in columns 1 to 4, sampling weights are modified according to the reweighting factor described in Section 4.2.
Data source: BSWS.

Figure 1: Changes in the wage distribution and the average minimum wage

Notes: We calculate the 10th, 50th, and 90th percentiles of the wage distribution and the average minimum wage by year and plot their logarithmic values after normalizing them to zero in 1994 both for men and women.
Data source: BSWS.

Figure 2: Changes in the fraction of minimum-wage workers



Notes: We calculate the fraction of workers paid less than or equal to the minimum wage by prefecture and year and plot the fraction in prefectures at the 5th, 50th, and 95th percentiles by year both for men and women.
Data source: BSWS.

Figure 3: Minimum wage ranks in 2000

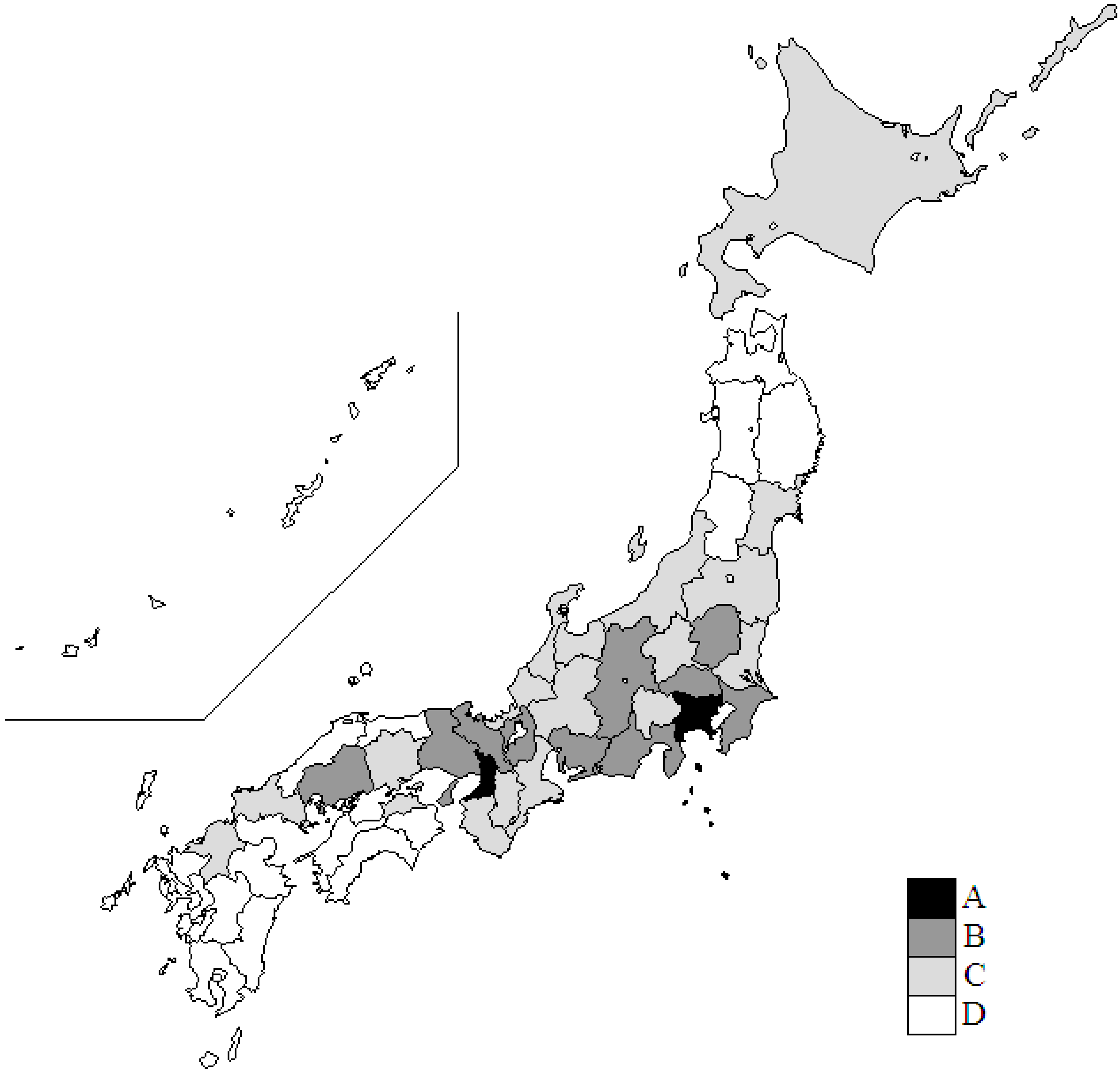
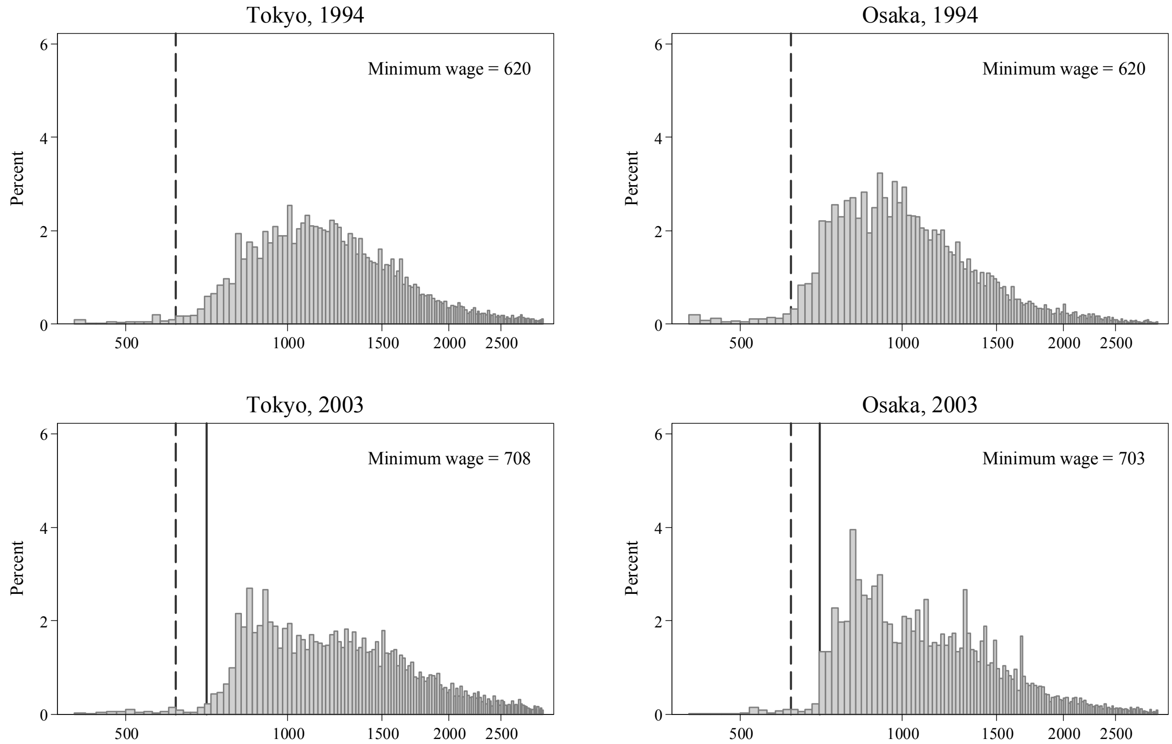
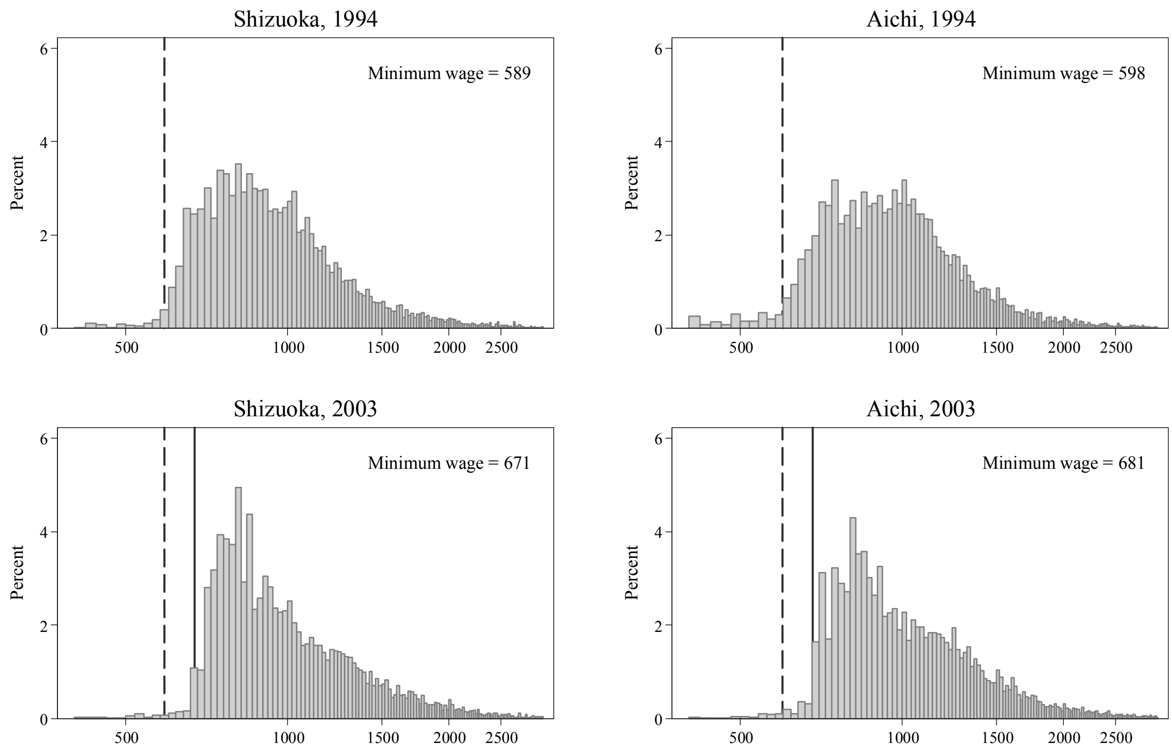


Figure 4: Changes in the log wage distributions from 1994 to 2003 by selected prefecture

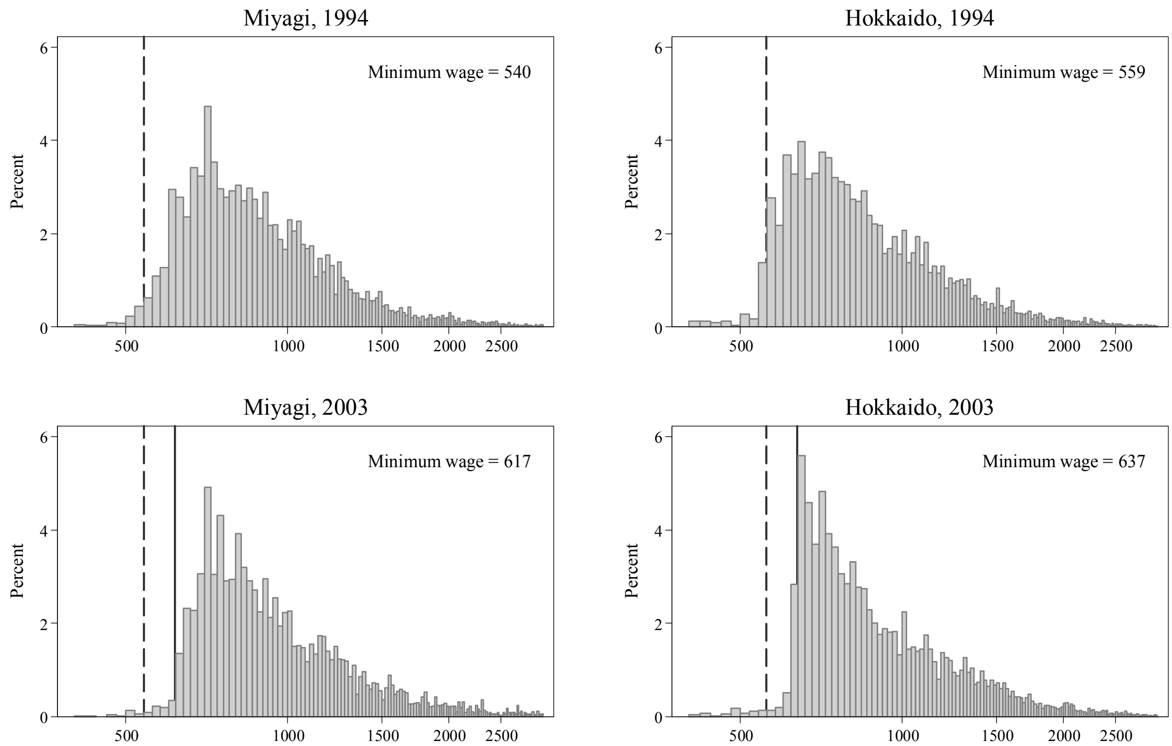
Rank A



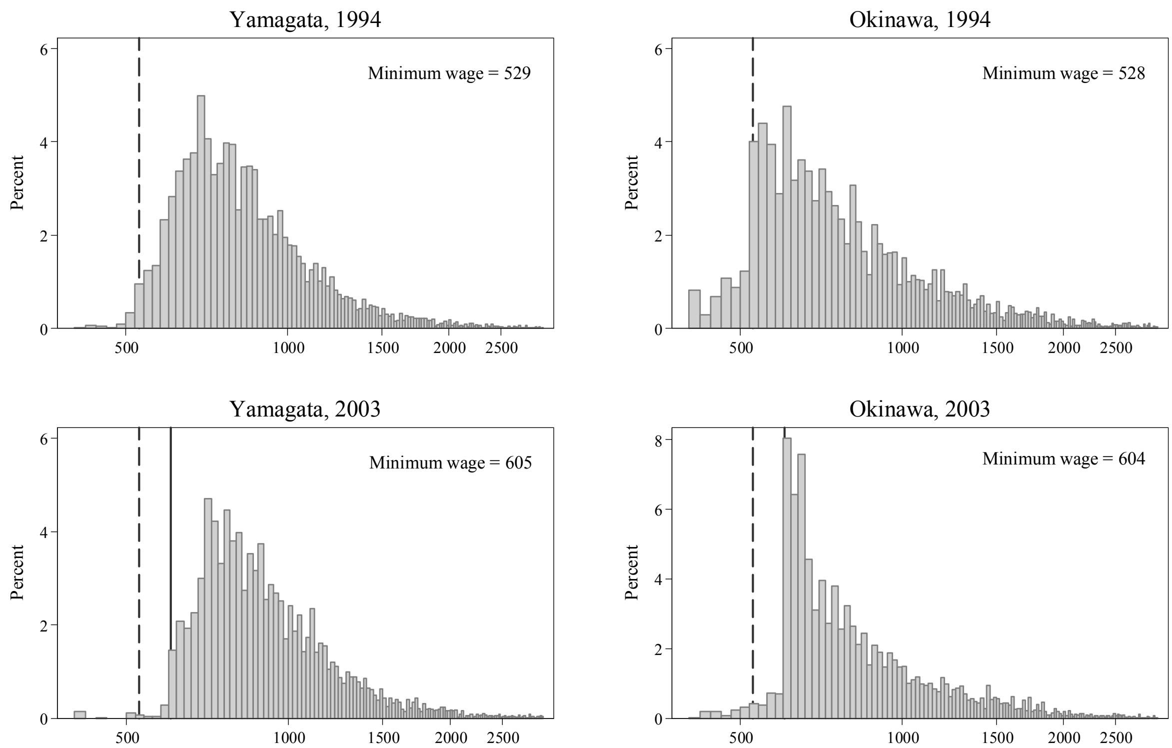
Rank B



Rank C

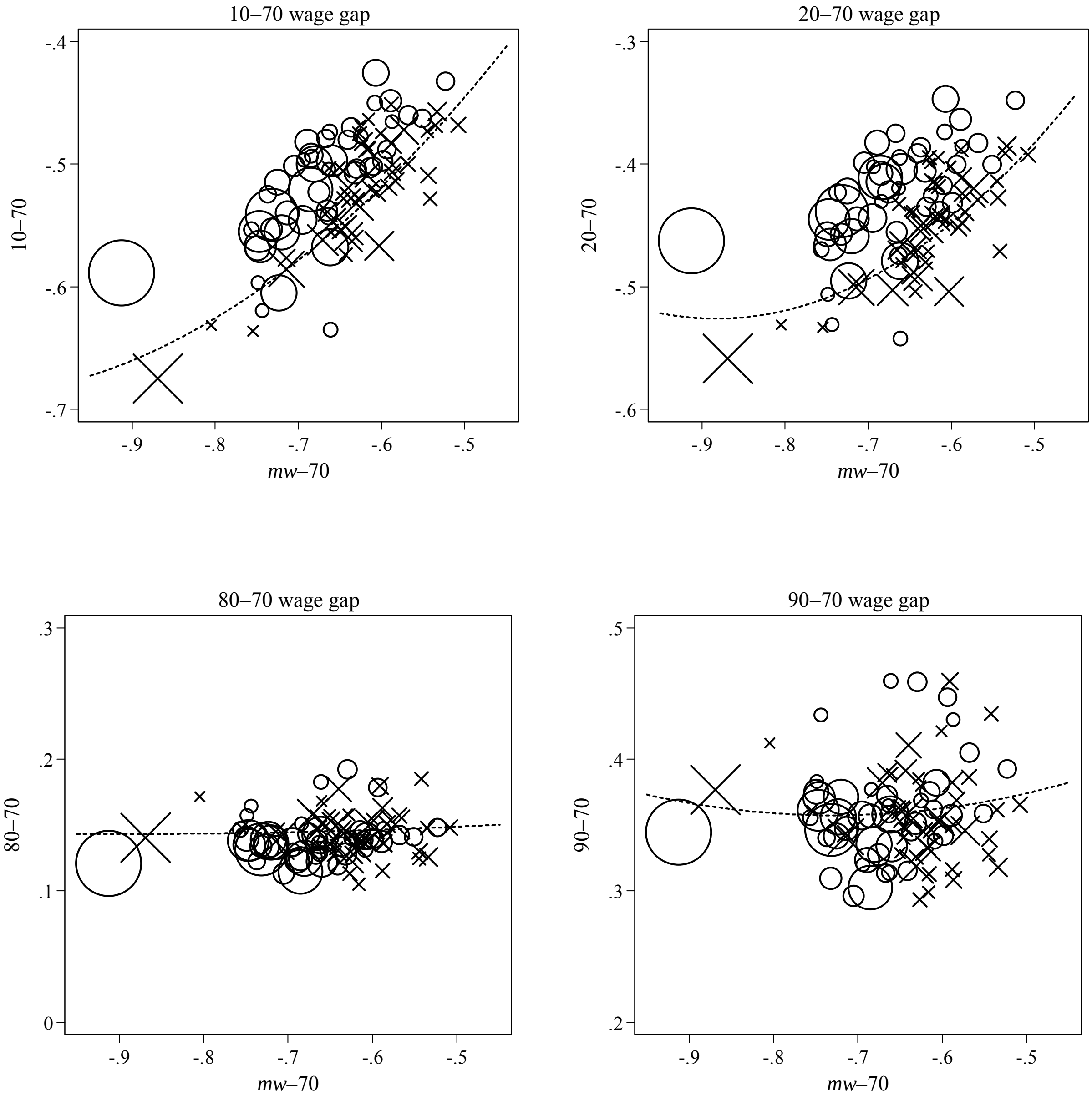


Rank D



Notes: The dash and solid vertical lines indicate the minimum wage levels in 1994 and 2003, respectively.
Data source: BSWS.

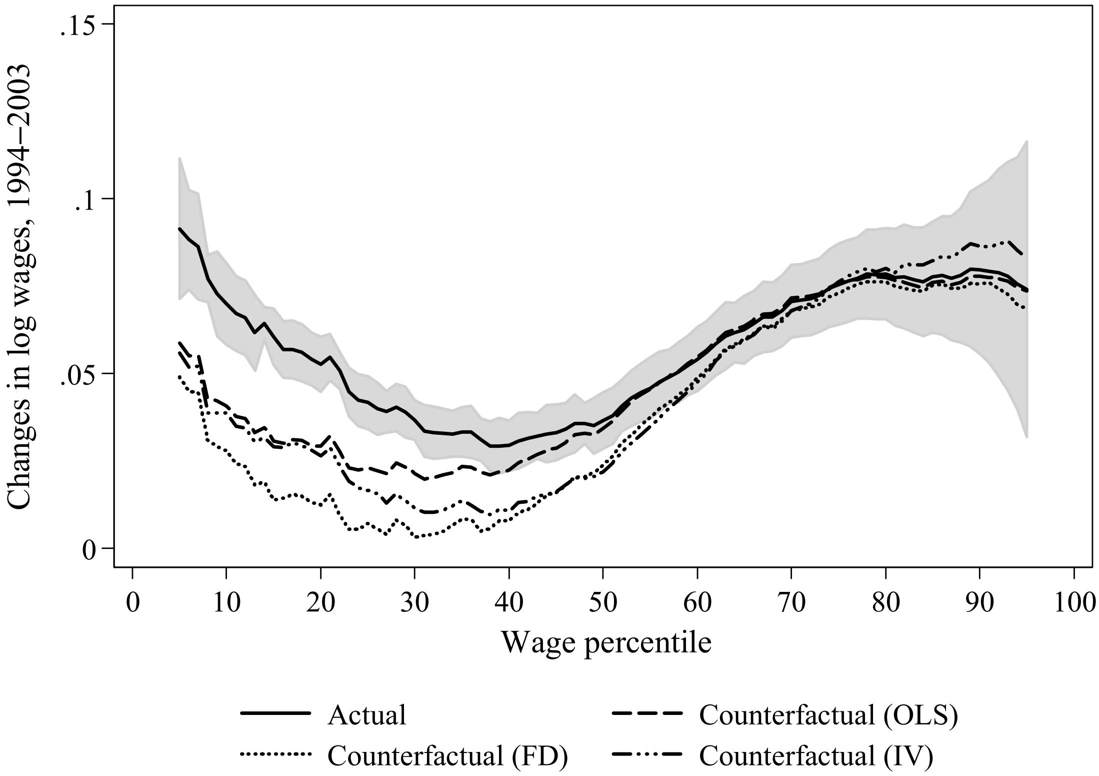
Figure 5: Wage-compression effect



○ 1994 × 2003 ------- Regression line

Notes: The vertical axis is the *p*–70 log wage differential for *p* = 10, 20, 80, and 90, and the horizontal axis is the log differential between the minimum wage and the 70th percentile wage. The regression lines are drawn based on the results in column 1 of Table 1. The size of the symbol is proportional to the sum of individual sampling weights by prefecture and year.
Data source: BSWS.

Figure 6: Actual and Counterfactual changes in log wages by percentile of the wage distribution from 1994 to 2003



Notes: The solid line represents the actual wage changes, while the dash and dotted lines represent the counterfactual wage changes if there was no change in the effective minimum wage. The shaded area represents the 95% confidence interval.
Data source: BSWS.