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Minimum Wage in a Deflationary Economy: 
The Japanese Experience, 1994–2003¹

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Abstract

The statutory minimum wage in Japan has steadily increased over the past few decades 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 increased bite of the minimum wage resulted in the compression of the lower tail of the wage distribution among women and that the wage compression is partially attributed to employment loss resulting from the minimum-wage increase. The increased bite of 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
1 Introduction

Although many advanced industrialized countries have shared 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 sharp difference in trends in wage inequality between Japan and the United States. Wage inequality increased only moderately in Japan from the middle 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 the 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). Indeed, 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. Nevertheless, there has been a lack of formal analysis regarding the impact of the minimum wage on inequality trends in Japan.1

The moderate increase in wage inequality in Japan occurred among both men and women until the early 1990s (Katz, Loveman, and Blanchflower, 1995). After the early 1990s, however, trends in male and female wage inequality in Japan started to diverge (Figure 1). While the male wage distribution has fanned out as the 10th percentile of the wage distribution has declined since the late 1990s, the female wage distribution has been compressed as the 10th percentile of the wage distribution has increased since the early 1990s. In contrast, the statutory minimum wage has continuously increased in all prefectures since the 1970s, even after the median wage started to decline in the late 1990s, in a way that does not eventually take into account regional differences in labor-market conditions. Consequently, the ‘bite’ of the minimum wage has increased significantly in low-wage prefectures. The proportion of workers paid less than or equal to the minimum wage has risen among women above 5% in some prefectures since the late 1990s, although it has been

1Some studies point to a spike in the wage distribution at the minimum wage, however. Hori and Sakaguchi (2005) illustrate the wage distribution in 2003 by prefecture and industry separately for full- and part-time workers using the Basic Survey on Wage Structure.
below 1% among men in most prefectures (Figure 2). The objective of this paper is to investigate the extent to which a reduction in lower-tail inequality among women in Japan can be attributed to the increased bite of the minimum wage.

The contribution of this paper to the literature is twofold. First, we examine the impact of an increased minimum wage on the wage distribution under wage deflation, in which the real value of the minimum wage would increase even without a change in the statutory minimum wage. The revision of the statutory minimum wage inevitably lags behind inflation or deflation. Therefore, the bite of the minimum wage rose under a deflationary economy in the 1990s in Japan, whereas it fell under an inflationary economy in the 1980s in the United States. Second, we quantitatively assess the impact of employment loss resulting from the minimum-wage increase on the wage distribution. In a simple competitive model, employment declines with a rise in the minimum wage as a result of lower demand for labor. The loss of employment for low-skilled workers will then cause a truncation of the wage distribution below the minimum wage. To shed light on the significance of the truncation effect, we develop two methods that do not require a distributional assumption, one of which builds upon a version of the inverse probability weighting method (DiNardo, Fortin, and Lemieux, 1996); the other builds upon a version of the trimming method (Lee, 2009). When employing these methods for quantifying the truncation effect, we modify the calculation of the reweighting factor and the trimming threshold by using estimates for the employment effect of the minimum wage.

This paper reveals that a steady increase in the bite of the minimum wage compressed inequality in the lower tail of the wage distribution among women in Japan. The increased bite of the minimum wage accounts for roughly one half of the reduction in lower-tail inequality during the period between 1994 and 2003. The 50–10 wage gap decreased by 7 log points during the period, but if there had been no change in the bite of the minimum wage, it would decrease by only less than 4 log points. Accordingly, the part-time wage penalty, which increased by 5 log points during the period, would have increased by 6 log points without any change in the minimum-wage bite. We additionally find that the increased bite of the minimum wage had adverse effects on new hires, which increased by 9 log points during the period, but if there had been no change in the bite of the minimum wage, it would increase by only less than 5 log points.

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2In monopsonistic models, the sign of the employment effect of the minimum wage depends on the degree of monopsony power (Bhaskar and To, 1999; Manning, 2003).
3Meyer and Wise (1983) estimate the effect of the minimum wage on a joint distribution of the wage and employment outcomes by the maximum likelihood method under a distributional assumption.
hours worked, and employment for women. The loss of employment resulting from the minimum-wage increase partially accounts for the minimum-wage effect on the 10th percentile of the wage distribution.

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 used to examine the effects of the minimum wage on the wage distribution, new hires, and hours worked; the other is a household survey used to estimate the effect on employment. It then introduces the minimum-wage system in Japan and finally discusses variation in the minimum-wage bite across regions and over time. Section 3 first considers an empirical framework to examine the impact of the minimum wage on the wage distribution, then presents the results of parameter estimates for the wage-compression effect, and finally assesses the quantitative contribution of the increased bite of the minimum wage to changes in wage inequality and the part-time penalty. Section 4 provides an analysis for the effects on new hires, hours worked, and employment, followed by an analysis for the effect of employment loss resulting from the minimum-wage increase 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. 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,\textsuperscript{4} 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.\textsuperscript{5} 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 550,000 establishments for every year. Each observation is weighted by a 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 allowance, and family allowance.\textsuperscript{6} 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 the 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

\textsuperscript{4}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 one of the above criteria is satisfied.

\textsuperscript{5}A person in charge of personnel at each establishment is asked to randomly choose a number of workers from the pool of employees using specific instructions for random sampling, including the sampling probability, which is dependent on the industry and establishment size.

\textsuperscript{6}There is no custom of tipping in Japan.
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. Each observation is weighted by a 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. 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 minimum wages. Industrial minimum wages have remained as a system to complement collective bargaining in some industries and prefectures, where the unionization rate was lower than 30% and industrial minimum wages were set higher than prefectural minimum wages in the middle 1980s.

Under the 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, the minimum-wage policy in Japan 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. Consensus has never been reached between representatives of employers and employees, however. The increased amount of prefectural minimum wages thus has been presented as the view of representatives of public interest every year. Prefectural minimum-wage councils virtually always conform to the indication of the central minimum-wage council, as evidenced by the fact that the actual

\footnote{See Hamaguchi (2009) for minimum-wage legislation in Japan.}
increased amount of prefectural minimum wages is typically identical to the one indicated by the central minimum-wage council. In sum, prefectural minimum wages increased every year in all prefectures as 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 wages are set in conformance with the central minimum-wage council’s indication. The steady increase 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. Three prefectures in rank A are Tokyo, Osaka, and Kanagawa (a prefecture located next to Tokyo); 10 prefectures in rank B are typically located around Tokyo and Osaka; 16 prefectures in rank D are located either in Tohoku (the north end district of main island), Chugoku (the west end district of main island), Shikoku (fourth largest island in the southwest of the main island), or Kyushu (third largest island in the southwest end of the main island); and 18 prefectures in the rank C are the rest (Figure 3). A total of 10 prefectures experienced a change in rank, though only by one level, during the period between 1994 and 2003. Since the intended purpose of the indication system was to moderate disparity in prefectural minimum wages, the average percentage increase of minimum wages was greatest among the lowest-ranked prefectures. Nevertheless, the difference between the lowest rank and other three ranks was only modest, and the difference between the three ranks excluding the lowest rank was negligible. The average percentage increases of prefectural minimum wages across prefectures in ranks A, B, C, and D between 1994 and 2003 were 13.9%, 14.0%, 13.9%, and 14.4%, respectively. Thus, disparity in prefectural minimum wages has remained largely unchanged, even since the indication system was introduced in 1978. The uniform increase in prefectural minimum wages implies that the bite of the minimum wage would have 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. To highlight regional and time variation in the minimum-wage bite within and between ranks, we select two prefectures that have either one of the highest or one of the lowest
proportion of minimum-wage workers in each rank. Both for 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. The proportion of workers who earned the minimum wage was, however, very small in these higher-ranked prefectures. The minimum wage was more binding in some lower-ranked prefectures. While in a rank C prefecture, Miyagi, low-wage workers were not concentrated around the minimum wage either in 1994 or 2003, in another rank C prefecture, Hokkaido, the wage density had already spiked moderately as of 1994, and it was highest at the minimum wage in 2003. 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. In sum, the bite of the minimum wage 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.

3 Impact of the Minimum Wage on the Wage Distribution

3.1 Model specification and identification

The increased bite of the minimum wage affects the wage distribution through three channels: truncation, censoring (spike), and spillover (see Card and Krueger 1995; Brown, 1999; Neumark and Wascher, 2008 for the survey). First, the minimum-wage hike leads to a loss of employment for workers who were paid less than the minimum wage in a competitive labor market, where the wage equals the value of the marginal product of labor. In this case, the wage distribution will be truncated below the minimum wage. Second, if firms retain workers who were 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 wages of these low-skilled workers can rise to the minimum wage. The wage distribution then will be censored at the minimum wage, and the wage density will spike at the minimum wage. Finally, as the minimum wage raises the costs of hiring minimum-wage workers, the demand for workers who are more skilled than minimum-wage workers will relatively increase. The effect of the minimum wage will then spill over to the wages of workers who earn 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 inequality in the lower tail of the wage distribution.8

We now set up the model relating the bite of the minimum wage to the shape of the wage distribution, building upon Lee (1999). Let \( w_{it}^p \) (\( w_{it}^q \)) denote the \( p \)th (\( q \)th) percentile of the wage distribution, \( mw_{it} \) the minimum wage, and \( x_{it} \) other determinants of the wage distribution in prefecture \( i \) in year \( t \). The theoretical relation between the minimum-wage bite and the wage distribution can be stated as

\[
\ln w_{it}^p - \ln w_{it}^q = f (\ln mw_{it} - \ln w_{it}^q, x_{it}),
\]

where \( f \) is a function that maps the minimum-wage bite onto the \( p - q \) log wage differential. Since the degree of the minimum-wage bite varies not only by the statutory minimum wage but also by the wage distribution across prefectures and time, the minimum-wage bite is measured by the ‘effective’ minimum wage, i.e., the log differential between the minimum wage and the \( q \)th percentile wage. We examine the effect of the minimum wage on all percentiles of the wage distribution from \( p = 5 \) to \( p = 95 \). The minimum wage might affect not only the 10th percentile but also the 20th or higher percentiles though the spillover effect, but it would not affect the 80th or higher percentiles, because the higher the percentile the weaker is the degree of substitutability for minimum-wage workers. We can thus see the presence and size of the spillover effect and determine a valid specification by looking at the minimum wage’s effect on the entire distribution.

When estimating equation (1), however, the \( q \)th percentile must be set high enough so as not to be

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8The 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, 2001; Manning, 2003; Aaronson and French, 2007), a search model with bargaining (Flinn, 2010), and a tournament model (Lazear and Rosen, 1981).
subject to the influence of the minimum wage. Since the $q$th percentile is used to normalize the minimum wage, if the minimum wage affects the $q$th percentile, we cannot interpret the estimated coefficient on the effective minimum wage as the effect of the minimum wage on the $p$th percentile. Although Lee (1999) and Autor, Manning, and Smith (2010) use the median as the $q$th percentile for their studies in the United States, Bosch and Manacorda (2010) adopt the 70th percentile as the $q$th percentile for their study in urban Mexico. As we have seen in Figure 4, there are large spikes around the minimum wage in the wage distribution among women in some low-wage prefectures in Japan, where the minimum wage conceivably has an influence on the median wage for women. We thus set the $q$th percentile to be the 70th percentile here.

Figure 5 plots the $p$–70 log wage differential along with the effective minimum wage for $p = 10, 20, 80, 90$ 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. In addition, the 10–70 log wage differential is more strongly associated with the effective minimum wage than the 20–70 log wage differential. The figure implies the validity of the specifications and the presence of spillover effect.

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

$$\ln \tilde{w}_{it}^p = \beta_1^p \ln \tilde{m} \tilde{w}_{it} + \beta_2^p (\ln \tilde{m} \tilde{w}_{it})^2 + x_{it} \gamma^p + u_{it}$$

for $i = 1, 2, \ldots, 46, 47$ and $t = 1994, 1995, \ldots, 2002, 2003$, where $\tilde{w}_{it}^p = w_{it}^p / w_{70}^p$, $\tilde{m} \tilde{w}_{it} = mw_{it} / w_{70}^p$, and $u_{it}$ 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. In the regression model above, we exploit differences in the statutory minimum wage and the wage distribution across prefectures and time to identify the effect of the minimum wage on the wage gap. On the one hand, there is a large cross-sectional variation in prefectural minimum wages originating from a difference in prefectural minimum wages set in the middle 1970s but no
significant 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. The calculation of the minimum wage in each prefecture and year also takes into account a variation in industrial minimum wages across prefectures and time, but the proportion of workers to whom industrial minimum wages apply is so small that industrial minimum wages do not create a significant variation additionally. On the other hand, 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 a variation in the wage distribution across prefectures and 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 a bias because of heterogeneity in the dispersion of the latent wage distribution in the absence of the minimum-wage system. Suppose that the minimum wage is more binding, i.e., the log differential between the minimum wage and the $q$th percentile wage is less, in prefectures with higher wage dispersion in the absence of the minimum-wage system. The direction of the bias will be downward for $p < q$ and upward for $p > q$, since the $p - q$ wage gap would be greater in absolute value in prefectures with higher wage dispersion, and its sign is negative for $p < q$ and positive for $p > q$. We then additionally control for prefecture effects, prefecture-specific trends, and the age, education, and industry composition of the workforce to allow for heterogeneity in the latent wage distribution across prefectures and years. Prefecture effects are unobserved but can be cancelled out by taking a first difference between year $t$ and year $t - 1$.

$$\Delta \ln \bar{w}_{it}^p = \beta_1^p \Delta \ln \bar{w}_{it} + \beta_2^p (\ln \bar{w}_{it})^2 + \Delta x_{it}^p + \Delta u_{it}^p,$$  \hspace{1cm} (3)$$

where $\Delta x_{it}$ 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 the division bias because of sampling error (measurement error) in the percentile of the wage distribution (Lee 1999; Autor, Manning, and

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9We calculate $mw_{it}$ by taking a weighted average of minimum wages across workers by prefecture and year, after assigning industrial minimum wages to workers in industries covered by industrial minimum wages and prefectural minimum wages to all other workers. Industrial minimum wages are always higher than prefectural minimum wages.
Smith, 2008). Since the \( q \)th percentile appears as a denominator of both the wage gap and the effective minimum wage, the sampling error in the \( q \)th percentile can generate a spurious positive correlation between the two variables. To circumvent this problem, we perform an instrumental variable (IV) estimation. Applying this method to equation (3) generally requires two assumptions. First, the effect of the minimum wage is (approximately) linear, i.e., \( \beta_2 = 0 \). Second, the measurement error takes the multiplicative form, such that \( w_{it}^r = w_{it}^{r*}e_{it}^r \), where \( w_{it}^r \) and \( w_{it}^{r*} \) are the observed and actual \( r \)th percentile, respectively, and \( e_{it}^r \) and \( e_{it}^{r*} \) are the persistent and i.i.d. components in the measurement error, respectively. The estimating equation can then be rewritten as

\[
\Delta \ln \tilde{w}_{it}^p = \beta_1 \Delta \ln \tilde{mw}_{it} + \Delta x_{it}\gamma_p + \Delta v_{it}^p, \tag{4}
\]

where \( \Delta v_{it}^p = \Delta u_{it}^p + \Delta \ln e_{it}^p - (1 + \beta_1^P) \Delta \ln e_{it}^{70} \). Although the measurement error is allowed to be serially correlated, the change in the log is i.i.d, and thus, the error term in equation (4) is i.i.d. Therefore, the lagged value of the effective minimum wage is a legitimate instrument. We thus use \( (\ln \tilde{mw}_{i,t-2} - \ln w_{i,t-2}^{70}) \) as an instrument for \( \Delta \ln \tilde{w}_{it} \), where \( \tilde{mw}_{it} \) 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. The first stage \( F \) statistic is 7.26. The reason for using \( \tilde{mw}_{it} \) is to avoid a potential endogeneity of prefectural minimum wages that prefectural minimum-wage councils can ultimately decide. In this specification, we do not use a variation in prefectural minimum wages that were set at the prefecture level for identification. The results remain unchanged regardless of whether using \( mw_{it} \) or \( \tilde{mw}_{it} \), however, as can be expected from the fact that prefectural minimum-wage councils effectively conform to the indication of the central minimum-wage council.

Equation (4) can be derived by substituting the restrictions imposed by the two assumptions into the regression model:

\[
\Delta \ln \tilde{w}_{it}^p = \beta_1^P \Delta \ln \tilde{mw}_{it}^* + \beta_2^P \Delta (\ln \tilde{mw}_{it}^*)^2 + \Delta x_{it}\gamma_p + \Delta u_{it}^p, \quad \text{where} \quad \tilde{w}_{it}^p = w_{it}^{p*}/w_{it}^{70*} \quad \text{and} \quad \tilde{mw}_{it}^* = mw_{it}/w_{it}^{70*}. \quad \text{Without the linearity assumption, the error term would be} \quad \Delta v_{it}^p = \Delta u_{it}^p + \Delta \ln e_{it}^p - (1 + \beta_1^P) \Delta \ln e_{it}^{70} + \beta_2^P \Delta (\ln (e_{it}^{70}/e_{it}^{70}))^2 - 2\beta^P \Delta \ln (\tilde{mw}_{it}) \Delta \ln (e_{it}^{70}/e_{it}^{70}). \quad \text{The reason to impose the linearity is that the instrumental-variable estimator is generally inconsistent in nonlinear models (Chen, Hong, and Nekipelov, 2011). We expect that the potential bias resulting from the linear specification would be quantitatively marginal in the analysis here, since the OLS and FD estimates for marginal effects and counterfactual wage changes remain largely unchanged regardless of whether including or excluding the quadratic term. Under the assumptions here, we can control for the persistent component in the measurement error caused by sampling.}
3.2 Parameter estimates

Table 1 reports the marginal effects of the effective 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 weighted average of the effective minimum wage across all prefectures and years. 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 effective minimum wage. As expected, the estimated marginal effect is greatest for the 10th percentile and shrinks gradually as the percentile becomes higher. The effects of the minimum wage are statistically significant in the lower tail but not in the upper tail of the wage distribution. The $R^2$ when controlling only for year effects is large at 0.74 for the 10th percentile, but it monotonically declines to 0.04 as the percentile approaches the 90th percentile. After controlling for prefecture fixed effects and prefecture-specific trends, the estimated marginal effects of the minimum wage become larger. After further controlling for the age, education, and industry composition of the workforce, the estimated marginal effects increase slightly more. The increase in the estimated marginal effects, however, may be attributed to the division bias arising from the sampling error. When the IV estimation is performed to circumvent this problem, the estimated marginal effects are indeed smaller. The IV estimates lie between the OLS estimates and the FD estimates for the marginal effects of the minimum wage on the lower percentiles of the wage distribution.

The increased bite of the minimum wage may cause firms to reduce the level of allowances. We thus examine the offsetting effect on allowances by calculating the $p$–70 wage gap from the wage rate including commutation allowance, perfect-attendance allowance, and family allowance. We find that the effect of the minimum wage on the 10–70 wage gap decreases by 3 to 14 percentage points, indicating that the wage-compression effect can be partially attributed to this offsetting effect.

\[ \text{The marginal effect is } \hat{\beta}_1 + 2\hat{\beta}_2 \ln \text{mw}, \text{ where } \ln \text{mw} \text{ is the weighted average of the effective minimum wage across all prefectures and years.} \]
3.3 Inequality trends

The parameter estimates under various specifications consistently indicate that the increased bite of the minimum wage resulted in a reduction in lower-tail inequality among women. To further examine the quantitative contribution of the increased bite of the minimum wage to inequality trends, we calculate counterfactual wages if there were no change in the effective minimum wage during the period between 1994 and 2003 based on the OLS, FD, and IV estimation, whose results are reported in columns 1, 2, and 4 of Table 1. For a worker $k$ whose hourly wages rank at the $p$th percentile in prefecture $i$ in year $t$, the counterfactual log wages in 2003 are calculated by subtracting the change in the $p$th percentile wage resulting from the minimum-wage increase from the actual log wages in 2003 as follows:

$$
\ln \hat{w}_{k,i;2003}^p = \ln w_{k,i;2003}^p - \beta_1^p (\ln \bar{w}_{i;2003}^p - \ln \bar{w}_{i;1994}^p) - \beta_2^p \left[ (\ln \bar{w}_{i;2003}^p)^2 - (\ln \bar{w}_{i;1994}^p)^2 \right]
$$

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. The effect of the minimum wage is represented as a 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 increased bite of the minimum wage resulted in an increase in wages of workers in the lower-tail of the wage distribution. At least one half of the wage increase below the 20th percentile can be attributed to the increased bite of the minimum wage.

Table 2 reports the actual and counterfactual changes in the $p$–50 wage gap for $p = 10, 20, 30, 40, 60, 70, 80, 90$. Based on the IV estimates, the increased bite of the minimum wage 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. The OLS and FD estimates indicate that the increased bite of the minimum wage accounts for an even greater part of the reduction in the 50–10 or the 50–20 wage gap.
3.4 Part-time wage penalty

Part-time employment has been increasingly common in recent years in many countries. Although most OECD countries have enacted new legislation for ‘equal treatment provisions’ since the early 1990s, a significant full-/part-time wage differential has remained in some countries. A cross-country comparison suggests that the part-time penalty is associated with the in-work poverty rate (OECD, 2010). The part-time penalty has been a hot policy issue in recent years in Japan, where the part-time penalty is highest among the OECD countries. Looking at our data set, while the proportion of part-time workers increased from 24 to 36% between 1994 and 2003, the full-time/part-time wage differential increased from 33 to 38 log points.

The steady increase in the minimum wage has an implication for a change in the part-time penalty. Given that part-time workers are paid less than full-time workers, the minimum-wage increase would push up the wages of part-time workers more significantly than the wages of full-time workers. If there had been no change in the minimum-wage bite, the part-time penalty may have increased even more. We use the counterfactual log wages calculated from equation (5) to more directly examine the effect of the increased bite of the minimum wage on the part-time wage penalty.

Figure 7 illustrates changes in the full-/part-time log-wage differential between 1994 and 2003 by percentile of the wage distribution. The increase in the part-time penalty is greater toward the upper percentiles. All three estimates of the counterfactual changes are significantly greater than the actual changes from the 15th percentile to the 60th percentile, indicating that the minimum-wage increase contributed to preventing the part-time penalty from increasing further. The bottom row of Table 2 reports that the actual and counterfactual changes in the log wage differential between full- and part-time workers. Based on the IV estimates, the part-time penalty would have increased by 27% more at the mean if there had been no increase in the minimum-wage bite.

\[\text{Manning and Petrongolo (2008) find that, after the national minimum wage was introduced in the United Kingdom, the percentage increase in wages at the bottom end of the wage distribution was greater for part-time workers than for full-time workers. Abe and Tanaka (2007) point out that the prefectural minimum wage affects the wage gap between full- and part-time workers in Japan.}\]
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, there was an increase in the wages of workers at the bottom end of the wage distribution. This brings up the question of how firms reacted to such an increase in labor costs. Enormous studies have examined 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 the minimum-wage legislation.\(^{13}\) 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 declined in response to the minimum-wage increase than when examining how much employment expanded in response to the erosion of the minimum-wage bite. Given that protection for regular employees is very stringent in Japan (OECD, 2004),\(^{14}\) and the minimum wage has steadily increased, firms might adjust the quantity of labor input by reducing the number of new hires or the number of hours worked for existing employees.\(^{15}\) 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 \((\text{newhire})\) and the average hours worked \((\text{hour})\) for women in an establishment \(j\) in year \(t\) using the BSWS.

\[
\text{newhire}_{jt} = -1.94 \ln \bar{mw}_{it} + x_{jt}, \tag{1.11}
\]

\(^{13}\)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. There are also a few studies that 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) take a similar approach to that used by Card (1992) and show that the minimum-wage hike reduced the employment of male teenagers and middle-aged married women.

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

\(^{15}\)Few studies to date, with the exception of Portugal and Cardoso (2006), have explored the effect of the minimum wage on worker flow.
\[ \ln \text{hour}_{jt} = -0.115 \ln \text{mw}_{it} + x_{jt} \gamma, \]

where \( \text{mw}_{it} = mw_{it} / w_{50}^{it} \), and \( x_{jt} \) is a vector of controls that include the number of regular employees, fourth order polynomials in the average employee age, prefecture effects, year effects, and prefecture-specific linear time trends. The sample includes 518,502 establishment-year observations. The regressions are weighted by the sum of sampling weights by establishment and year. Standard errors in parentheses are clustered at the prefecture level. The implied new hires’ elasticity at the sample mean is –1.84, indicating that the adverse effect of the minimum wage is not negligible. 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.

\[ \hat{\Pr}(e_{kt} | \text{mw}_{it}, x_{kt}) = \Phi \left( -0.654 \text{mw}_{ait} + x_{3kt} \gamma \right), \]

where \( \text{mw}_{ait} = mw_{it} / w_{50}^{ait} \), \( e \) is an indicator for being employed, and \( x \) is a vector of controls that include fourth-order age polynomials, prefecture effects, year effects, and prefecture-specific linear time trends. We exploit a 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. The sample includes 972,479 individual-year observations. The regression is weighted by a sampling weight. Standard errors in parentheses are clustered at the prefecture level. The implied employment elasticities with respect to the minimum wage range from –0.407 in Okinawa to –0.267 in Tokyo, when evaluated at the sample means of the effective minimum wage by prefecture.

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 understanding the size of the truncation effect.\(^{16}\) Our analysis here is complementary

\(^{16}\)We can show that truncation mechanically reduces lower-tail inequality as follows. For an arbitrary continuous distribution, it must be that \( \int_{w^{10}}^{w^{50}} f(w) dw = 0.4 \), where \( w \) is the log hourly wages, \( w^{10} \) and \( w^{50} \) are the 10th and 50th percentiles of the wage distribution, and \( f(\cdot) \) is the probability density function. For the distribution truncated below
to Autor, Manning, and Smith (2010), who estimate the size of censoring and spillover effects in the case of no truncation effect under a distributional assumption. Recognizing that the problem considered here is a sample-selection problem, we can infer the extent to which wage compression is attributable to the loss of employment resulting from the minimum wage by correcting for the selection bias arising from the truncation effect. To do so, we consider the actual wage density and the counterfactual wage density that is not subject to the influence of employment loss resulting from the minimum-wage increase. Let \( f(w_t | \tilde{mw}_t, x_t) \) denote the actual wage density conditional on the effective minimum wage \( g_{\tilde{mw}} \) and observed attributes \( x \) in year \( t \), \( f(w_t | \tilde{mw}_{1994}, x_t) \) the counterfactual wage density if there had been 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 the wage compression can be attributed solely to the truncation effect. First, there is neither a spillover effect nor a censoring effect:

\[
g(w_t | e_t = 1, \tilde{mw}_t, x_t) = g(w_t | e_t = 1, \tilde{mw}_{1994}, x_t).
\]

Second, the minimum wage has no direct effect on the distribution of observed attributes:

\[
h(x_t | \tilde{mw}_t) = h(x_t | \tilde{mw}_{1994}).
\]

As derived in the appendix, the counterfactual density can be nonparametrically constructed as follows:

\[
f(w_t | e_t = 1, \tilde{mw}_{1994}) = \int g(w_t | e_t = 1, \tilde{mw}_t, x_t) h(x_t | e_t = 1, \tilde{mw}_t) \theta(\tilde{mw}_{1994}, \tilde{mw}_t, x_t) \, dx,
\]

where

\[
\theta(\tilde{mw}_{1994}, \tilde{mw}_t, x_t) = \frac{Pr(e_t = 1 | \tilde{mw}_{1994}, x_t) \cdot Pr(e_t = 1 | \tilde{mw}_t)}{Pr(e_t = 1 | \tilde{mw}_t) \cdot Pr(e_t = 1 | \tilde{mw}_{1994})}.
\]

The reweighting function is essentially the ratio of the counterfactual employment rate without a change in the minimum-wage bite to the actual employment rate. In this reweighting procedure to control for the truncation effect, we give more weight to workers who would have a higher propensity to be employed without an increase in the effective minimum wage, i.e., to those who are more likely to disappear from the labor market as a result of the minimum-wage increase. An estimate of each response probability can be obtained from probit model (6).

The minimum wage \( mw \), it must be that \( f_{w_{10}^{50}} w_{10} \, dw = 0.4 \), where \( w_{10} \) and \( w_{50} \) are the 10th and 50th percentiles of the truncated distribution. Then, \( f_{*} (w) = f (w | w \geq mw) = \frac{f(w)}{Pr(w \geq mw)} \geq f (w) \). Thus, \( w_{50} - w_{10} \geq w_{50}^{10} - w_{10}^{10} \).

This result applies to the 70–10 wage gap.
The first four columns of Table 3 present estimates for the marginal effects of the minimum wage reproduced after reweighting. The results are almost identical to those reported in Table 1, indicating that the truncation effect is minimal.

4.3 Trimming method

The results reproduced after reweighting suggest that employment loss has a minimal impact on the wage distribution. Nevertheless, the truncation effect should be underestimated if employment loss occurs exclusively from the bottom end of the wage distribution. We thus employ a more conservative approach that enables us to quantify the upper bound of the truncation effect. Consider the scenario when jobs disappear in the order of wage percentile from the bottom end of the distribution as a result of the minimum-wage increase in the current year. This means that workers would lose their jobs 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. The trimming threshold can be calculated by prefecture and year as follows. We first calculate the year-by-year percentage change in the effective minimum wage by prefecture and year and then multiply it by the estimated elasticities with respect to the minimum wage evaluated at the sample means of the effective minimum wage by prefecture. The trimming threshold is zero in the last year of the sample period, since it is the base year with the highest minimum wage. We also 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. 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 four columns of Table 3 present estimates for the marginal effects of the minimum wage reproduced after trimming. Although the OLS estimates are almost unchanged, the FD and IV estimates for the minimum-wage effect on the 10th percentile wage decrease by 31% and 8%,
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. The IV estimates indicate that employment loss does not account for the minimum-wage effect on the 20th and higher percentiles, while the FD estimates indicate that employment loss accounts for a significant part of the minimum-wage effect on the 20th and higher percentiles.

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 period between 1994 and 2003 among women in Japan, one of the world’s largest economies. For institutional reasons, the statutory minimum wage steadily increased in all prefectures, even during a period of deflation, and consequently, the bite of the minimum wage substantially increased, especially among women in low-wage prefectures. Japan’s experience since the 1990s mirrors the U.S. experience in the 1980s and 1990s. The increased bite of the minimum wage compressed inequality in the lower tail of the wage distribution in Japan, whereas a fall in the minimum-wage bite resulted in a sizable increase in wage inequality in the United States. Our analysis revealed that the increased bite of the minimum wage accounts for approximately one half of the reduction in the 50–10 wage gap among women and that the compression of the lower-tail of the wage distribution is partially attributable to the loss of employment resulting from the minimum-wage increase. We additionally find that the increased bite of the minimum wage had adverse effects on new hires, hours worked, and employment for women. 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 cost of employment loss among women. The findings of this paper imply a policy trade-off between the reduction in wage inequality and the loss of employment.
Appendix

Derivation of equation (7).

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

\[ f(w_t|e_t=1, \bar{m}w_{1994}) = \int g(w_t|e_t=1, \bar{m}w_{1994}, x_t) h(x_t|e_t=1, \bar{m}w_{1994}) \, dx, \tag{8} \]

where \( g(w_t|e_t=1, \bar{m}w_{1994}, x_t) \) is the counterfactual wage density conditional on attributes \( x \), and \( h(x_t|e_t=1, \bar{m}w_{1994}) \) is the counterfactual density of attributes. Under the first assumption,

\[ f(w_t|e_t=1, \bar{m}w_{1994}) = \int g(w_t|e_t=1, \bar{m}w_t, x_t) h(x_t|e_t=1, \bar{m}w_{1994}) \, dx. \tag{9} \]

The second assumption implies

\[ h(x_t|e_t=1, \bar{m}w_{1994}) = \frac{\Pr(e_t=1|\bar{m}w_{1994}, x_t) h(x_t|\bar{m}w_t)}{\Pr(e_t=1|\bar{m}w_{1994})}. \tag{10} \]

By applying Bayes’s rule,

\[ h(x_t|e_t=1, \bar{m}w_t) = \frac{\Pr(e_t=1|\bar{m}w_t, x_t) h(x_t|\bar{m}w_t)}{\Pr(e_t=1|\bar{m}w_t)} \]

\[ h(x_t|\bar{m}w_t) = h(x_t|e_t=1, \bar{m}w_t) \cdot \frac{\Pr(e_t=1|\bar{m}w_t)}{\Pr(e_t=1|\bar{m}w_t, x_t)}. \tag{11} \]

Substituting (11) into (10) yields

\[ h(x_t|e_t=1, \bar{m}w_{1994}) = h(x_t|e_t=1, \bar{m}w_t) \theta(\bar{m}w_{1994}, \bar{m}w_t, x_t), \tag{12} \]

where

\[ \theta(\bar{m}w_{1994}, \bar{m}w_t, x_t) = \frac{\Pr(e_t=1|\bar{m}w_{1994}, x_t)}{\Pr(e_t=1|\bar{m}w_t, x_t)} \cdot \frac{\Pr(e_t=1|\bar{m}w_t)}{\Pr(e_t=1|\bar{m}w_{1994})}. \]

Substituting (12) into (9) yields

\[ f(w_t|e_t=1, \bar{m}w_{1994}) = \int g(w_t|e_t=1, \bar{m}w_t, x_t) h(x_t|e_t=1, \bar{m}w_t) \theta(\bar{m}w_{1994}, \bar{m}w_t, x_t) \, dx. \]
References


Table 1: The marginal effects of the minimum wage on the wage distribution

<table>
<thead>
<tr>
<th>log wage differentials</th>
<th>OLS (1)</th>
<th>FD (2)</th>
<th>FD (3)</th>
<th>IV (4)</th>
</tr>
</thead>
<tbody>
<tr>
<td>10–70</td>
<td>0.552</td>
<td>0.887</td>
<td>0.928</td>
<td>0.642</td>
</tr>
<tr>
<td></td>
<td>(0.047)</td>
<td>(0.067)</td>
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<td>(0.199)</td>
</tr>
<tr>
<td>20–70</td>
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<td>0.751</td>
<td>0.804</td>
<td>0.437</td>
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<td>(0.060)</td>
<td>(0.053)</td>
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</tr>
<tr>
<td>30–70</td>
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<td>0.602</td>
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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 sampling weights by prefecture and year. Other controls include the share of workers by age, education, and industry. In the third 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.
Table 2: Actual and counterfactual changes in the log wage differentials, 1994–2003

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Notes: The 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.
Table 3: The minimum-wage effect after controlling for the truncation effect

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<td>0.022</td>
<td>−0.110</td>
<td>0.016</td>
<td>0.048</td>
<td>−0.217</td>
<td>(0.021)</td>
<td>(0.035)</td>
</tr>
<tr>
<td></td>
<td>(0.021)</td>
<td>(0.035)</td>
<td>(0.170)</td>
<td>(0.021)</td>
<td>(0.032)</td>
<td>(0.331)</td>
<td>(0.021)</td>
<td>(0.032)</td>
</tr>
<tr>
<td>90–70</td>
<td>0.024</td>
<td>0.068</td>
<td>−0.087</td>
<td>0.025</td>
<td>0.087</td>
<td>−0.291</td>
<td>(0.037)</td>
<td>(0.069)</td>
</tr>
<tr>
<td></td>
<td>(0.037)</td>
<td>(0.069)</td>
<td>(0.284)</td>
<td>(0.038)</td>
<td>(0.063)</td>
<td>(0.534)</td>
<td>(0.038)</td>
<td>(0.063)</td>
</tr>
</tbody>
</table>

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 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 4.3.
Figure 1: Trends in the 10th, 50th, and 90th percentile wages and the minimum wage

![Graph showing trends in wages](image1)

**Notes:** These are measured in log terms and normalized to zero in 1994.

Figure 2: Trends in the proportion of minimum wage workers by prefecture

![Graph showing proportions](image2)

**Notes:** The statutory minimum wage varies by prefecture, industry, and year.
Figure 3: Minimum wage ranks in 2000
Figure 4: Distributions of log wages in 1994 and 2003 by selected prefecture

Rank A

Tokyo, 1994
Minimum wage = 620

Tokyo, 2003
Minimum wage = 708

Osaka, 1994
Minimum wage = 620

Osaka, 2003
Minimum wage = 703

Rank B

Shizuoka, 1994
Minimum wage = 589

Aichi, 1994
Minimum wage = 598

Shizuoka, 2003
Minimum wage = 671

Aichi, 2003
Minimum wage = 681
Notes: The dashed and solid vertical lines indicate the minimum wage levels in 1994 and 2003, respectively.
Figure 5: Wage-compression effect

Notes: The regression lines are drawn based on the results in column 1 of Table 1. The size of symbol is proportional to the sum of sampling weights by prefecture and year.
Figure 6: Changes in log wages by percentile of the wage distribution, 1994–2003

Notes: The shaded area represents the 95% confidence interval.
Figure 7: Changes in the full-/part-time log wage differential, 1994–2003

Notes: The shaded area represents the 95% confidence interval.