RIETI Discussion Paper Series 21-E-007

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https://www.rieti.go.jp/en/

RIETI Discussion Paper Series 21-E-007
January 2021

Estimating the Effects of the Minimum Wage Using the Introduction of Indexation*

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Abstract

We examine the impacts of the minimum wage on employment using the minimum wage hike induced by the introduction of indexation of the local minimum wage to the local cost of living. The revision of the Minimum Wage Act in 2007 of Japan essentially required the government to set the minimum wage indexed to the local cost of living with a five-year moratorium period. The government subsequently increased the minimum wage in areas where the cost of living was high relative to the local minimum wage. We find that minimum-wage hikes raised the wages of low-wage workers, but reduced the employment of young, less-educated men. A panel analysis based on matched Labour Force Survey data indicates that the minimum-wage hike decreased the job flows of prime-age men and women.

JEL Classification: J23, J38, J42, J64, J81
Keywords: Minimum Wage; Low Skill Workers; Employment; Japan

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* This study is conducted as a part of the Project "Empirical Analysis on Japanese Labor Market" undertaken at the Research Institute of Economy, Trade and Industry (RIETI). We are grateful for comments from David Card, Koichi Fujii, Dean Hyslop, Johnathan Meer, David Neumark, Fumio Ohtake, and seminar participants at the Asian Development Bank Institute, National Chengchi University, Motu Economic and Public Policy Research, the Victoria University of Wellington, Aoyamagakuin University, the University of Tokyo, and the annual meeting of the Asian and Australasian Society of Labor Economists. This study utilizes micro data of the Labour Force Survey conducted by the Ministry of Internal Affairs and Communications and the Basic Survey of Wage Structure conducted by the Ministry of Labour, Health and Welfare. This project has been financially supported by the Japan Society of Promotion of Science (JSPS), grant number 19H00589.

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1 Introduction

Numerous studies have assessed the effect of minimum wage on employment using the regional variation of the minimum wage, as reviewed by Card and Krueger (1995), Neumark and Wascher (2008), Meer (2018) and Clemens (2019). Key to the research is how to handle the endogeneity of the local minimum wage, because policymakers set the minimum wage by looking at local labor-market conditions. In the absence of an exogenous source of minimum-wage variation across states, US studies have struggled to control for the region- and time-specific unobserved shocks that could be correlated with minimum wages. One group of researchers controlled for region- and time-specific shocks by using the geographically proximate area as a control group (Dube et al., 2010). Another group of researchers claim that a traditional region- and time- fixed effects model performs better (Neumark et al., 2015; Meer and West, 2016; Clemens and Wither, 2019). The estimated impact of minimum wage on employment differs substantially by the construction of the counterfactual situation. In contrast with existing studies that attempt to construct counterfactual local labor-market conditions, we exploit the plausibly exogenous minimum-wage hike across regions.

The natural experiment we exploit is the change in the institutional setting of minimum-wage determination in Japan. The minimum wage in Japan is statutory, and local minimum wages are set for each of the 47 prefectures. Central and local minimum wage commissions are involved in determining the local minimum wage; in the process, the central minimum wage commission suggests the amount of minimum wage increase for each prefecture. Historically, the suggestion by the commission was linked to the average wage increase of micro scale establishments based on a survey for this specific purpose. The revision of the Minimum Wage Act in 2007 additionally required the government to set the minimum wage, referring to the amount of welfare benefits that is indexed to the local cost of living. Specifically, the revised Act aimed to close the gap between the amount of welfare benefit and the earnings of full-time minimum wage workers, which are more substantial in urban prefectures, where the cost of living is higher. Accordingly, to fill the gap, the minimum wage had been increased more significantly in urban prefectures than in rural prefectures. In other words, the revision of the Minimum Wage Act virtually introduced an indexation of the minimum wage to the local cost of living and induced heterogeneous minimum wage growth across prefectures. We exploit this heterogeneous growth of minimum wages across prefectures to estimate the impact of minimum-wage hikes on wages, employment, hours, and one-year transition of employment status.

We first examine the relationship between minimum wage and hourly wage using the Basic Survey of Wage Structure (BSWS), which allows us to precisely calculate the hourly wage based on the wage bill of establishments. We confirm that the minimum-wage hike increased the fraction of minimum-wage workers, defined as the workers whose wage is below $1.05 \times$ the local minimum wage. At the same time, we show that only about 3% of workers are minimum-wage workers, and they are concentrated among the less educated, as well as younger and older workers of both sexes.

We then analyze the effect of the minimum wage on employment, labor-force participation, and hours worked using the monthly Labour Force Survey (LFS). To pay attention to less-skilled workers, we focus only on those with a complete high-school education or less. Focusing on less-skilled workers, such as teenagers, is a typical practice in the literature, because only a small fraction of the workforce is directly affected by the minimum wage (Dube et al., 2010; Allegretto et al., 2011; Neumark et al., 2015). Accordingly, our analysis sample includes all less-educated people between 19 and 64 years old, dividing them into three age groups: 19-24, 25-59, and 60-64.

We implement the instrumental variable estimation exploiting the institutional feature that the minimum-wage hike after the 2007 revision of the Minimum Wage Act was largely motivated by a call to fill the gap between the welfare benefit amount and minimum wage earnings. Thus, we use the gap between the welfare benefit and the minimum-wage earnings in 2006 as the instrumental variable for subsequent minimum wage growth. We first examine if the gap is associated with the trends of the employment rate before 2007 and find that the employment growth of male youth was positively correlated with the gap between the welfare benefit and minimum-wage earnings. In contrast, the pre-trend of young women was negatively correlated with the gap. To address the potential bias of the instrumental variable estimation created by the correlation of the pre-trends and the IV, we allow for a flexible form of heterogeneous trends across prefectures before 2007.

The estimated impacts of minimum-wage hikes on employment are heterogeneous across workers' characteristics, according to the IV estimation allowing for a prefecture-specific linear time trend. We find that the minimum wage reduces employment among less-educated young men between 19 and 24 years old. The estimated elasticity is -1.2, which is larger than the estimates in the literature. We point to two reasons behind this discrepancy. First, the estimand of the IV estimator is the local average treatment effect of the minimum wage on employment identified off the increase of the minimum wage due to the introduction of the indexation of the minimum wage; employers can expect the future increase of the minimum wage according to the newly introduced rule and can substitute capital for labor, as found by Brummund and Strain (2020) from the evidence of state minimum-wage indexation. Second and more importantly, we argue that the OLS estimates in the literature suffer from an upward bias due to policy endogeneity, because policy makers are hesitant to increase the minimum wage in an adverse labor=market condition, as suggested by Baskaya and Rubinstein (2012) in the US context. This suggests the importance of seeking the exogeneous source of variation of local minimum-wage hikes in the difference-in-difference framework to address the policy endogeneity issue.

In addition, we find no precise estimates of the impact of minimum-wage hikes on employment among demographic groups other than less educated young men. We further find no significant impact of minimum-wage hikes on hours worked, conditional on being employed, among all demographic groups.

We further analyze the effect of the minimum-wage hike on the transition of employment status over a one-year period by constructing short panels. We do this by matching the outgoing rotation groups of adjacent years of the LFS to examine the impact of the minimum wage on the transition of employment status over a one-year period. We find that a minimum-wage hike reduces the transition probability from non-employment to employment among prime-age men, while it increases the probability that employed prime-age women stay employed. Thus, the minimum-wage hike lowers the transition between not being employed and being employed. This finding is consistent with theoretical predictions and empirical findings in previous studies indicating that a minimum-wage hike reduces job flow in the labor market (Portugal and Cardoso, 2006; Gittings and Schmutte, 2016; Dube et al., 2016; Brochu and Green, 2019).

1.1 Literature

This work contributes to the literature in two ways. First, it identifies the effects of the minimum wage on employment, drawing on a credible, exogenous source of minimum-wage variation. All recent US studies concur on the importance of controlling for region-specific shocks that can be correlated with the minimum wage. They disagree, however, on the proper construction of a counterfactual situation in terms of selecting control areas or modeling underlying region-specific time trends (Dube et al., 2010; Allegretto et al., 2011; Neumark et al., 2015; Meer and West, 2016; Cengiz et al., 2019; Clemens and Wither, 2019). Compared to the situation in the US, where state legislators vote for state minimum-wage hikes, which potentially creates policy endogeneity according to local labor-market conditions, the determination of the Japanese regional minimum wage is arguably more centralized and mechanical. Notably, the introduction of the indexation of the minimum wage to the local cost of living, due to the revision of the Minimum Wage Act in 2007, provides a credible exogenous source of variation of local minimum wages to identify the effect of the minimum wage cleanly. To the best of our knowledge, Baskaya and Rubinstein (2012) is the only recent study that instruments the state minimum wage by the interaction term of the federal minimum wage and states' tendency to let the federal minimum wage bind.

Second, we contribute to the Japanese policy debate. Political liberals traditionally support the minimum-wage hike as a policy tool to alleviate poverty in Japan. The political scene has changed dramatically in recent years; the minimum wage has also gained support from political conservatives as an income policy to boost wages to initiate wage-price dynamics in Japan, where wages have stagnated despite a very tight labor market since the early 2010s. At the Council on Economic and Fiscal Policy on November 24, 2015, Prime Minister Shinzo Abe committed to increasing the minimum wage by 3% annually until the average regional minimum wage reaches 1,000 yen. The International Monetary Fund (IMF) backs this policy stance,[1] and IMF staff economists Aoyagi et al. (2017) reported that minimum-wage hikes have a significant impact on wage increases. Some business owners and economists, however, express concern about job loss among unskilled workers due to minimum-wage hikes.

Numerous studies have examined the impact of minimum wage on employment in Japan. Early studies exploit the cross-sectional variation of the minimum wage to estimate the impact of minimum wage on employment, using cross-sectional data or panel data without prefecture fixed effects (Tachibanaki and Urakawa, 2006; Yugami, 2005; Ariga, 2007). Controlling for prefecture fixed effects is crucial, however, because the level of the minimum wage is systematically correlated with unobserved local labor-market conditions. Several studies have allowed for individual or prefecture fixed effects in the estimation of the minimum-wage impact on employment. Kawaguchi and Mori (2009) calculated the fraction affected by minimum wages, following Card (1992), using the Employment Status Survey from 1982 to 2002, and found that the stronger bite of the minimum wage reduces the employment of teens and married, middle-aged women. Kambayashi et al. (2013) reported that higher minimum wages reduce the female employment rate, based on data up to 2003. Kawaguchi and Yamada (2007) showed that a worker who is prone to be affected by a minimum-wage

[1] The Conclusion of the Mission for the 2018 Article IV Consultation with Japan published on October 4, 2018.

hike is more likely to lose her job, using the 1993-1999 panel surveys on women. Overall, minimum-wage studies that control for policy endogeneity, covering the sample period before the 2007 revision of the Minimum Wage Act, point to employment loss among less-skilled workers due to minimum-wage hikes.

Only a few studies have exploited minimum-wage hikes after the revision of the Minimum Wage Act in 2007. Higuchi (2013) examined the impact on the employment of non-regular workers based on the Keio Household Panel Survey and concluded that minimum-wage hikes did not affect the employment of non-regular workers. A careful reading of the estimation results, however, reveals that the minimum-wage hikes indeed decreased the employment of non-regular workers, but the impact was not precisely estimated because of the small sample size, around 2,000 observations. Since the number of minimum-wage workers is as limited as 3% of all workers, we need to rely on a large sample of government surveys. Akesaka et al. (2018) estimated the impact of the recent minimum-wage hike on employment using the Employment Status Survey, which is a large government survey that takes place every five years. They found that the minimum-wage hike reduces the employment of teenage men and reduces the hours worked of all workers. Our study improves the granularity of the analysis exploiting the large sample of monthly or annual frequency. Okudaira et al. (2019) reported the impact of minimum-wage hikes after the 2007 revision negatively affected employment, with an estimated elasticity around -0.5 among all workers, relying on Census of Manufacture (Ministry of Economy, Trade and Industry). They also examine the heterogeneous impacts depending on the degree of labor market monopsony. They first identified the degree of monopsony by the gap between the observed labor share and the predicted labor share in the perfectly competitive labor market from the estimated production function. Then, they found a more substantial adverse impact in the region where the labor market is estimated to be competitive. In addition, they implemented the instrumental variable estimation using the gap between the welfare benefit and minimum-wage earnings as of 2006 as the instrumental variable. Our study complements their study by covering all industries.

2 Minimum-wage indexation to the local cost of living: Revision of the Minimum Wage Act in 2007

This section describes the institutional background of minimum-wage setting in Japan for introducing the identification strategy; more details are provided in Appendix A. Regional minimum wages, set for each of the 47 prefectures, were essentially indexed to the local average wage before 2007, because the increase in the local minimum wage was tightly linked to the average annual wage growth of workers in establishments with less than 30 workers, as documented by Tamada (2009). The partial revision of the Minimum Wage Act in 2007, which took effect on July 1, 2008, changed the policy process by adding a new requirement that the monthly earnings of full-time minimum-wage workers should not be lower than public assistance benefits. There was a gap between the monthly earnings of minimum-wage workers and the public assistance benefits in 12 prefectures as of fiscal year 2006.[2] The revised Act aimed to close the gap between the amount of welfare benefits and the earnings of full-time minimum-wage workers, which was more substantial in urban prefectures or prefectures with cold weather in the winter, where the cost of living is higher. Accordingly, to fill the gap, the minimum wage had been increased more significantly in urban prefectures than in rural prefectures. Between 2007 and 2016, for example, the increase of the minimum wage was 26% in Tokyo and Kanagawa prefectures, urban prefectures, and 12% in Tokushima prefecture, a rural prefecture, as indicated in Figure 1.

We use the gap between the amount of welfare benefits and the earnings of full-time minimum-wage workers as of 2006 as the instrumental variable to predict the amount of the minimum-wage hike in each prefecture. The left panel of Figure 2 shows the relationship

[2] A letter submitted to the Minister of Health, Labour and Welfare from the Central Minimum Wage Commission in 2008 reported that the monthly earnings of minimum-wage workers were lower than public assistance benefits in 12 prefectures, including Hokkaido, Aomori, Miyagi, Akita, Saitama, Chiba, Tokyo, Kanagawa, Kyoto, Osaka, Hyogo, and Hiroshima prefectures. The gap amount as of 2006 was 63 yen for Hokkaido, 20 yen for Aomori, 31 yen for Miyagi, 17 yen for Akita, 56 yen for Saitama, 35 yen for Chiba, 100 yen for Tokyo, 108 yen for Kanagawa, 47 yen for Kyoto, 53 yen for Osaka, 36 yen for Hyogo, and 37 yen for Hiroshima.

between the gap ($\ln(WB/MWE)_{j2006}$) and the minimum-wage growth between 2007 and 2016. There were 12 prefectures where the welfare benefit exceeded the minimum-wage earnings in 2006 ($\ln(WB/MWE)_{j2006} > 0$) and the initial gap and the subsequent minimum-wage growth was positively correlated. This positive correlation assures the validity of the instrumental variable. In contrast, the right panel of Figure 2 shows that the gap in 2006 is not correlated with the minimum-wage growth between 1998 and 2007. The absence of a correlation negates a concern that the correlation between the gap and the subsequent minimum-wage growth is due to the long-term trend of minimum-wage growth.

We attempt to estimate the causal impact of the minimum-wage hike on the change of labor-market outcomes using the gap between the welfare benefit and minimum-wage earnings as of 2006 as the instrumental variable. In this identification strategy, a probable threat to identification is that the initial gap between the minimum wage and welfare benefits is correlated with the underlying trends of labor-market outcomes. To address this concern, we estimate the relationship between the gap and the pre-trend of employment; we estimate the following one-year-interval first-difference model:

$$\Delta Emp_{jt} = \beta \ln(WB/MWE)_{j2006} + \tau_t + e_{jt}, \tag{1}$$

where ΔEmp_{jt} is the annual change in the employment rate in prefecture j in year t, using the period between 2002 and 2007; and $\ln(WB/MWE)_{i2006}$ is the gap between the welfare benefit and minimum-wage earnings in 2006. The estimation results using all the years are reported in Table 1. The regression coefficient is 0.10 (SE=0.03) for 19 to 24 year-old men and -0.20 (SE=0.05) for 19 to 24 year-old women. We do not find systematic trends for other gender and age groups.

Both the graphical and regression analyses indicate that the employment growth of the youth of both genders before the revision of the Minimum Wage Act is systematically correlated with the gap between welfare benefits and minimum-wage earnings. This heterogeneous pre-trend may bias the estimate of the effect of the minimum wage on employment; thus in the instrumental variable estimation, introduced in Section 5, we allow for a flexible form of heterogeneous trends across prefectures before 2007 to mitigate the effects of the pre-trends.

3 Data

This study uses the Labour Force Survey (LFS) to capture labor-market outcomes, including employment status and hours worked. The drawback of the LFS is its lack of accurate measurement of the hourly wage rate. To compensate for this drawback, we also use the Basic Survey on Wage Structure (BSWS). We briefly describe these two data sets below. Some additional details of these two data sets are presented in Appendix B.

The main data set used in this study is the Labour Force Survey (LFS). This monthly survey corresponds to the Current Population Survey in the US. The LFS generates repeated cross-section data, but its rotation sampling structure allows us to construct a short panel data set covering two years. We use the LFS as a source of repeated cross-section data, but we exploit the panel structure for the transition analysis of employment status. We restrict the analysis sample to non-college graduates between the ages of 19 and 64, because, as we will see below, the minimum wage is relevant to non-college-graduate workers.[3] We divide the analysis sample into three age groups: 19-24, 25-59, and 60-64. We pick the ages between 19 and 24 as indicating youth to avoid those who are still in high school, as almost all high-school graduates turn age 19 before graduation. The prime-age people are those between 25 and 59 years old, and the older people are those between 60 and 64 years old. The majority of employers in Japan set the mandatory retirement age at 60, and those workers who experience a mandatory retirement change in their employers or a transition to a temporary position with the same employer. The analysis sample covers the period between January 2002 and December 2016. A shortfall of the LFS is a lack of accurate

[3]The LFS asks respondents to identify their educational background by their enrollment status in school. We restrict the sample to those who are not enrolled, and highest education attainment is high school or less. Thus, the analysis sample does not include high-school graduates who are currently enrolled in higher educational institutions, such as a university or a professional training college.

measurement of the hourly wage rate.

To compensate for the lack of an accurate measurement of hourly wage in the LFS, we use worker-level micro data from the Basic Survey on Wage Structure (BSWS). The BSWS is an annual establishment survey that adopts a two-step random sampling procedure. In the first step, establishments are randomly selected from the population of all establishments that hire 5 or more employees. In the second step, the human resource department of the establishment randomly selects the workers from its payroll records. The BSWS is an ideal data set for analyzing the impact of the minimum wage on wages because of its accurate measurements of individual workers' monthly earnings and monthly hours worked transcribed from payroll records. The BSWS covers both full- and part-time workers, but it does not ask for the educational background of part-time workers. Since part-time workers are likely to be low-wage workers and more relevant to minimum wages, we basically do not restrict the analysis sample to less-educated workers. The sampling process of the BSWS uses prefecture $\times$ industry $\times$ establishment size as the strata. To keep the sampling error within a stratum below a certain level, larger establishments and establishments belonging to a rare industry are oversampled. To reflect this peculiar sampling structure of the survey, we use the sampling weight in the regression analysis to recover the population distribution throughout the analysis based on the BSWS. We also use the sampling weight so that the regression coefficients recover the population average treatment effects (Solon et al., 2015). Our analysis sample covers the years between 2002 and 2017.

The regional minimum wage set for each of the 47 prefectures is taken from the *Saitei Chingin Kettei Yoran* (Overview of Minimum Wage Determination) for each fiscal year published by the *Rodo Chosakai.* Because regional minimum wages are revised around October every year, we use the minimum wage that came into effect in October in year t to explain employment in October-December in year t and in January-September in year $t+1$.

4 Analysis based on payroll records

4.1 Effect on wages

As a preparatory analysis, this section examines how minimum wages affect wages, using the BSWS. Figure 3 shows the relationship between the wage distributions of all workers ages 19-64 and the minimum wage by prefecture in 2002 and 2017 for representative prefectures: Tokyo, as an example of major urban agglomerations, and Tokushima, as an example of rural areas. In Tokushima prefecture, the wage distributions for men and women for both 2002 and 2017 are dented at the minimum wage, indicated by the vertical line. The figure thus illustrates that the minimum wage played an essential role in determining the shape of the wage distribution even before the partial revision of the Minimum Wage Act in 2007. In contrast, in Tokyo, only a tiny share of workers worked for the minimum wage in 2002, and the minimum-wage level hardly affected the shape of the wage distribution. In 2017, however, the shares of workers working for the minimum wage increased, and the shape of the wage distribution began to become dented at the minimum-wage level. This result is in contrast with the finding of Kambayashi et al. (2013), who, using data up to 2003, showed that minimum wages acted as a constraint only in rural areas and implied that as a result of the minimum-wage hike since 2007, minimum wages had started to act as a practical constraint in urban areas, too.

As a way to systematically examine the impact of the minimum wage on the lower end of the wage distribution, we first examine if the minimum-wage hike increases the fraction of workers who work near the minimum wage. To do so, we define minimum-wage workers as workers whose hourly wage is below 1.05 $\times$ minimum wage. We should note that in the analysis sample covering the years between 2002 and 2017, only 3.7% of workers had earnings in this range of hourly wage. This small fraction suggests that we need to narrowly restrict the analysis sample to those who were likely to have been affected by minimum-wage hikes. Figure 4 shows that the increase of the minimum wage between 2002 and 2017 and the increase in the fraction of workers who earned less than $1.05 \times$ MW in the corresponding period are positively associated.

We next examine the impact of the minimum wage on the wage distribution using the wage bins defined relative to the minimum wage. The wage bins are:

$$(0, MW), [MW, 1.05MW), [1.05MW, 1.10MW), ..., [1.25MW, 1.30MW].$$

We then define dummy variables for each bin that take 1 if the individual worker's wage falls in the bin. We then estimate the following linear probability model. For example, we estimate the following model for the wage bin $(0, MW)$:

$$1(W_{ijt} < MW_{jt}) = \beta \ln MW_{jt} + X_{ijt}\gamma + \theta_j + \tau_t + u_{ijt}, \tag{2}$$

where i is the index for individual, j is prefecture, and t is year. The indicator $1(W_{ijt} < MW_{jt})$ takes 1 if the wage is below MW, and takes 0 otherwise, MW_{jt} is the MW in prefecture j in year t, X_{ijt} is a set of dummy variables defined by age $\times$ gender. We estimate the models by weighted least squares using the sampling weight. Standard errors are robust against within-prefecture error correlation.

Table 2 reports the estimation results of the linear probability models, where the dependent variables are the dummy variables that take one if the wage falls within a specific range relative to the minimum wage. Column 1 shows that minimum-wage hikes increase the fraction of non-compliers. On average, 1% of workers work below the minimum wage, and a 10% increase in the minimum wage increases non-compliers by 0.9%. In contrast, Column 2 indicates that a 10% minimum-wage hike increases the fraction of workers whose wage belongs to the range $[MW, 1.05MW)$ by 1.7 %, whereas about 2% of workers belong to this category. Similarly, Column 3 indicates that minimum-wage hikes increase the fraction of workers whose wage belongs to the range $[1.05MW, 1.10MW)$, suggesting that minimum-wage hikes affect the wage of workers who earn above the minimum wage, called the spillover effect or ripple effect. This spillover reaches up to 15% above the minimum wage.

Minimum wages substantially affect the lower end of the wage distribution, but only 3.7% of workers earn below 1.05 × minimum wage. Thus, to assess the impact of minimum-wage hikes on employment, we need to focus on very low-skilled workers and identify the characteristics of minimum-wage workers. The caveat is that the BSWS does not record the educational background of part-time workers, who make up about 20% of the analysis sample and are more likely to work as minimum-wage workers; only around 1% of full-time workers are minimum-wage workers, compared to about 10% of part-time workers. With this drawback in mind, Figure 5 plots the fraction of workers that earn below $1.05 \times MW$ by the educational background and age of full-time workers. The fraction of minimum-wage workers is substantially higher among low-educated (high school or less) younger and older workers. Thus, to identify the impact of the minimum wage on employment, we need to focus on low-educated younger and older workers. Indeed, Clemens et al. (2020) demonstrate the need to focus on less-educated workers in minimum wage research in the US context.

To help depict the composition of minimum-wage workers, Appendix Tables C1 and C2 tabulate the industry and occupation distributions among minimum-wage workers as of 2017. A large fraction of minimum-wage workers worked in the manufacturing sector (23%), wholesale and retail trade (17%), and the hotel and restaurant industry (14%). In terms of occupation, a large fraction of minimum-wage workers worked as sales (35%), service (28%), and manufacturing process (25%) workers.

5 Analysis based on the Labour Force Survey

5.1 Econometric model

We have seen that minimum-wage hikes raise the wages of low-wage workers and that low-wage workers are concentrated among less educated and younger and older workers. We now ask whether these wage increases are associated with a decrease in their employment.

Since the number of workers who are affected by minimum-wage hikes is limited, we focus on the employment of less-educated workers, namely those with a high-school education or less. Note that the LFS records the educational background of all individuals, regardless of employment status, different from the BSWS.

The data we use are from the LFS, 2007 to 2016. Figures 6 shows on the horizontal axis the change in the natural log of the minimum wage between 2007 and 2016 and on the vertical axis the change in the employment rate of males and females by age groups 19-24, 25-59, and 60-64 in the same period. For young males, the increase in the minimum wage is associated with a decrease in employment; the negative association is particularly notable in two large prefectures that experienced a significant increase in the minimum wage, namely Tokyo and Kanagawa. For prime-age and older men, we find a slight negative association. In contrast, for women, the increase in the minimum wage is positively associated with a change in the employment of those who were young and prime-age. The association of minimum-wage change and employment rate change between 2007 and 2016, however, does not imply a causal impact of the minimum wage on employment because of potential policy endogeneity.

We now lay out the instrumental variable estimation method to attack the policy endogeneity issue exploiting the introduction of the indexation of the minimum wage to the local cost of living. As explained before, a part of the minimum-wage hike after the 2007 revision of the Minimum Wage Act was motivated to increase the earnings of minimum-wage workers so that earnings could catch up with the welfare benefit amount of a single-person household. At the revision of the Act, the government gave a five-year grace period until the earnings of full-time minimum wage workers could catch up with the amount of welfare benefit. We demonstrated that the gap between the welfare benefit and minimum-wage earnings calculated by the Ministry of Health, Labour and Welfare in 2006 indeed explains the level of the minimum wage after 2007. In the estimation, we need to take care of the fact that the trends of employment rates were systematically correlated with the gap between the welfare benefit and minimum wage earnings in 2006, as reported in Table 1.

We set up and estimate the following model by the two-stage least squares:

$$Y_{ijt} = \beta \ln MW_{jt} + X_{ijt}\gamma + \iota_j + \tau_t + \zeta_j \cdot Year + u_{ijt}, \tag{3}$$

$$\ln MW_{jt} = \sum_{y=2002}^{2016} \delta_y \ln(WB/MWE)_{j2006} \cdot 1(Year = y) + X_{ijt}\theta + \nu_j + \mu_t + \xi_j \cdot Year + v_{ijt}, \tag{4}$$

where Y_{ijt} is labor-market outcome of individual i living in prefecture j in year-month t, MW_{jt} is the minimum wage of prefecture j in year-month t, WB is the monthly welfare benefit amount of a single-headed household, MWE is the monthly earnings of minimum-wage workers, and thus $\ln(WB/MWE)_{j2006}$ is the log point difference between the welfare benefit and MW earnings in prefecture j in 2006. When the welfare benefit is lower than the earnings of MW workers $(WB/MWE)_{j2006}$ is defined as 1, thus $\ln(WB/MWE)_{j2006} = 0$. The vector X_{ijt} includes a set of dummy variables indicating each age; the population shares of males 19-24, 25-59, 60-64 and females 19-24, 25-59, and 60-64; and the prime-age unemployment rate of college graduates. The variables ι_j and ν_j are prefecture fixed effects and τ_t and μ_t are year×month fixed effects, ζ_j and ξ_j are prefecture-specific linear time trends.

The gap between the amount of welfare benefit and the earnings of minimum-wage workers, $\ln(WB/MWE)_{j2006}$, motivates central and local minimum-wage commissions to increase the minimum wage in prefecture j, particularly during the grace period between 2008 and 2012. To allow for the differential impact of the gap in 2006 on the minimum wage, we allow for different coefficients between 2002 and 2016. We would expect stable coefficients in the pre-period, between 2002 and 2007, an increase in coefficients between 2008 and 2012, and stable coefficients in 2013 and afterward. We impose restrictions $\delta_{2002} = \delta_{2007} = 0$ to identify the prefecture-specific linear time trends ξ_j, as articulated by Borusyak and Jaravel (2017); thus, the prefecture-specific state linear trend is estimated using the pre-treatment period.

Borusyak and Jaravel (2017) show that two time periods should be set to zero to identify the location-specific time trends, and we choose $\delta_{2002} = \delta_{2007} = 0$ so that the longest time span is used to identify the trends. Allowing for heterogeneous pre-trends across prefectures is essential, given that the trend of the employment rate was systematically correlated with the gap in 2006, as discussed in Section 2.

This instrumental variable estimation allows for policy endogeneity in the form of a correlation between u_{ijt} and v_{ijt} that would arise if minimum-wage commissions consider local labor-market conditions in their determination of minimum wages. The identification of the coefficients, however, requires that the error term of the structural equation u_{ijt} and the gap in 2006, $\ln(WB/MWE)_{j2006}$, not be correlated, conditional on prefecture and time fixed effects and the prefecture-specific linear time trend. As explained in Section 2, the level of the welfare benefit amount is determined by the cost of living of the region, and the gap between the welfare benefit amount and minimum-wage earnings is less likely to be correlated with subsequent temporary labor-market shocks, conditional on prefecture and year-month fixed effects and the prefecture-specific linear time trend. A caveat is that the instrumental variable estimation would estimate the impact of minimum wages on labor-market outcomes during the period between 2008 and 2012, because the variation of the minimum wage used is generated for the purpose of filling the gap between the welfare benefit and minimum-wage earnings. Thus, the instrumental variable estimator estimates the local average treatment effect.

We estimate the model separately by gender and age group (19-24, 25-59, 60-64) to allow for different fixed effects and prefecture-specific time trends, because both the baseline labor-market outcomes and the effect of the minimum wage on outcomes presumably differ across groups. The sample period includes the period when the Act on the Employment of Elderly Persons was revised, and it promoted the employment of older people between the ages of 60 and 64 (Kondo and Shigeoka, 2017). The impact of this Act is presumably absorbed by time fixed effects, because it was a national policy without regional variation.

5.2 The first stage

We first report the estimation results of the first-stage equation, Equation (4). Figure 7 reports the estimated coefficients, γ_y, and their 95% confidence intervals. The coefficients had hovered around 0 by 2007, and started to creep up from 2008 to 2013, reaching 0.6, and then hovered around 0.6.[4] This result is consistent with the notion that minimum wage councils had increased regional minimum wages to fill the gap between WB and MWE in 2006 during the period between 2008 and 2013. A reason why the difference of the estimated coefficients before and after the grace period is about 0.6, instead of 1.0, is that a significant part of the initial gap was resolved by the uniform percentage increases in regional minimum wages captured by the year-month fixed effects. F-statistics for the null hypothesis that all the coefficients for the interaction terms between the gap in 2006 and year dummy variables between 2003 and 2006 equal zero is 1.60. The F-statistic for the null hypothesis that all the coefficients for the interaction terms between the gap in 2006 and year dummy variables between 2008 and 2016 equal zero is 203.40, however, and thus we do not face the weak instrumental variable problem.

5.3 Employment

Next, we report the estimation results of the following reduced-form equation:

$$Emp_{ijt} = \sum_{y=2002}^{2016} \theta_y^E \ln(WB/MWE)_{j2006} \cdot 1(Year = y) + X_{ijt}\pi^E + \omega_j^E + \psi_t^E + \lambda_j^E \cdot Year + e_{ijt}^E, \quad (5)$$

where Emp_{ijt} is the dummy variable indicating the employment status; $\ln(WB/MWE)_{j2006}$ is the gap between the welfare benefit and the monthly earnings of minimum wage workers as of 2006; and the vector X_{ijt} includes the set of age dummy variables, the fraction of each age and sex group in the population, and the unemployment rate of college graduates. As mentioned before, when the welfare benefit is lower than the earnings of minimum-wage

[4] The estimation results are unchanged when the prefecture-specific linear time trend and the restriction $\delta_{2002} = 0$ are dropped.

workers, $\ln(WB/MWE)_{j2006} = 0$. To identify the prefecture-specific linear time trend λ_j, we impose the assumptions that $\theta_{2002} = \theta_{2007} = 0$. Figure 8 reports the estimated θ_y and its 95% confidence intervals. The gap between the welfare benefit and minimum-wage earnings as of 2006 is negatively associated with the subsequent employment rate among men in general, and it is statistically significant among those between the ages of 19 and 24. In contrast, among all age groups of women, the gap in 2006 and subsequent employment is not systematically correlated. This result implies that the minimum-wage hike induced by the initial gap between the welfare benefit and minimum-wage earnings reduced the employment of young men between 19 and 24 years old.

Table 3 Panel A reports the OLS and IV estimates for men and women by age groups. The OLS estimate for men ages 19-24 is -0.78, where the average employment rate of this group is 0.80; thus, the implied elasticity is -0.98. The IV estimate is -0.93, implying the elasticity -1.2. Both OLS and IV estimates are precisely estimated and statistically significant. These negative estimates indicate that minimum-wage hikes significantly reduce the employment of high-school-graduate youth, presumably the group with the lowest skill.

We need an explanation for this large negative impact of minimum wage on employment. As a baseline, Okudaira et al. (2019) reported an estimated elasticity around -0.5 among all workers based on OLS, using the minimum-wage hikes after the 2007 revision for identification. First, our estimated elasticity based on OLS estimate, -0.98, is larger than the baseline estimate, but the difference is not odd given that the estimate was obtained for high-school graduates between 19 and 24 years old, a group of low-skilled workers. Second, the larger negative impact of the minimum wage on employment based on the IV estimate than the OLS estimate is consistent with policy endogeneity; when the local economy booms, the government impose minimum-wage hike, thus the OLS estimate is upward biased. Third, the estimand of our IV estimator is the local average treatment effect of the minimum wage on employment, exploiting the variation of the minimum wage induced by the introduction of the indexation to the local cost of living. Since the government aimed to eliminate the gap between the minimum wage and the local cost of living in the five-year period, employers foresaw the future increase of the minimum wage in 2007. Employers could have adjusted capital investment with a long-term perspective and that may well make the estimate larger than the usual estimate. Indeed, Brummund and Strain (2020) report that the disemployment effect of the minimum-wage hike in a state that indexes its minimum wages to inflation is around 3 times larger than the disemployment effect in a state without indexation. They argue that their finding is due to the employers' forward-looking capital investment behaviors. Compared with the indexation to the inflation rate, Japan's situation makes the future increase in the minimum wage even more foreseeable, because the government announced its intent to eliminate the gap between the minimum wage and the welfare benefit in 5 years from 2007, and the gap was public information as of 2007.

For prime age men between 25 and 59 years old, neither the OLS estimate nor the IV estimate is statistically significant. The OLS and IV estimates for those who are ages 60-64 are negative but imprecisely estimated. These results are consistent with the reduced-form estimates reported in Figure 8. We find some suggestive evidence that minimum-wage hike reduced the employment of older men, but the estimate is not precise enough to conclude anything definitive.

For women, the minimum-wage hike did not affect the employment of any age group in a statistically significant way. The absence of the impact of minimum wage on women's employment rate across all age groups is consistent with the findings from Figure 8. In sum, we cannot conclude anything definitive about the effect of the minimum-wage hike on employment among women.

Overall, we find that the minimum wage reduces employment among less-educated men between 19 and 24 years old. The estimated elasticity based on the IV estimate is -1.2, which is larger than estimates in the literature. We argue, however, that the upward bias of the OLS estimate due to policy endogeneity and the predictable increase of the minimum wage in our context makes our estimate larger than the previous estimates. We find no precise estimates, however, of the impact of minimum-wage hikes on other demographic groups.

5.4 Labor-force participation

We next examine the impact of the minimum wage on labor-force participation. A higher minimum wage could reduce employment, but it could simultaneously encourage labor-force participation (LFP) because of high offered wage under the search friction in the labor market. At the same time, however, high unemployment rate potentially induced by the minimum wage hike may discourage LFP. Thus, the impact of a minimum-wage hike on LFP cannot be signed *a priori*. Thus, to analyze the impact of the minimum wage on LFP, we estimate the following reduced-form equation:

$$LFP_{ijt} = \sum_{y=2002}^{2016} \theta_y^L \ln(WB/MWE)_{j2006} \cdot 1(Year = y) + X_{ijt}\pi^L + \omega_j^L + \psi_t^L + \lambda_j^L \cdot Year + e_{ijt}^L, \quad (6)$$

where LFP_{ijt} is the dummy variable indicating LFP, including being unemployed. Other variables are the same as in the employment equation, and the same identification assumption is imposed.

Figure 9 shows the estimated θ_y in the above equation. We find negative impacts of the initial gap on LFP among younger and older men, though these are imprecisely estimated. For other demographic groups, we do not find significant impacts. Table 3 Panel B tabulates both OLS and IV estimates by demographic groups. Focusing on IV estimates, the estimated negative impact on LFP attenuates from the estimated negative impact on employment for younger and older men. Due to the decrease in the negative impact, the estimated impact becomes statistically insignificant for younger men between 19 and 24 years old. For example, for younger men, the IV estimate for employment was -0.93, but the IV estimate for LFP was -0.56. The smaller negative impact implies that some young men are discouraged from participating in the labor force because of deteriorated job prospects due to the minimum-wage hike. While not precisely estimated and only suggestive, the same discussion may apply to older men. As we would expect from the reduced-form estimates, we do not find that the minimum wage has a significant impact on the LFP among women.

5.5 Hours worked

The analysis heretofore has focused on the effect of the minimum wage on the extensive margin of employment status. The LFS asks for the hours worked in the last week of the previous month for those who were in employment. Using this variable, we estimate the impact of the minimum wage on hours worked, conditional on being employed, to shed light on the intensive margin of the potential adjustment. The estimated model is:

$$\ln(hour)_{ijt} = \sum_{y=2002}^{2016} \theta_y^H \ln(WB/MWE)_{j2006} \cdot 1(Year = y) + X_{ijt}\pi^H + \omega_j^H + \psi_t^H + \lambda_j^H \cdot Year + e_{ijt}^H, \tag{7}$$

where $\ln(hours)_{ijt}$ is the natural logarithm of hours worked in the last week of the previous month. Other variables are the same as the those in the employment equation, and the same identification assumption is imposed.

Figure 10 reports the estimated θ_y of the reduced-form model. The estimates show that the initial gap does not systematically affect the weekly hours worked for all demographic groups and that most of the estimated coefficients are statistically insignificant. Table 3 Panel C reports the OLS and IV regression results of the natural logarithm of hours worked based on the same specification as before, using high-school graduate (or less) men and women as the analysis sample. OLS estimates indicate that the minimum-wage hike reduces the hours worked of younger and older men and women. Consistent with the reduced-form results, however, the IV estimates show that the minimum-wage hike does not affect the hours worked of almost all demographic groups except for 25-59 years old male. For 25-59 years old male, 10 percent increase in MW increases hours worked by 1.4 percent, which is quantitatively limited. Considering the policy endogeneity of minimum-wage hikes, it is fair to conclude that minimum-wage hikes do not affect the hours worked. Thus, the minimum-wage hikes do not affect the intensive margin of employment.

5.6 Event study method

Recent studies on the minimum wage, such as Meer and West (2016) and Cengiz et al. (2019), adopt an event study approach that includes lags and leads of $\ln MW_{jt}$ to validate the common trend assumption. This approach works well when the minimum wage in some regions increases significantly at different times, as in the case of the US state minimum wage, because such a discrete change in the minimum wage creates a significant variation of minimum wages across time in the same region. In contrast, as explained in the previous section, the 2007 revision of the Minimum Wage Act sets a background for the trend heterogeneity of minimum-wage hikes across prefectures and creates a strong serial correlation of the minimum-wage hikes. To give a concrete image, suppose that we observe a decline of employment in a particular region where the minimum wage continuously rises; in this case, attributing an observation to current, past, or future minimum-wage increases is fundamentally difficult. Thus our study does not adopt an event study approach; instead, we address the policy endogeneity issue by estimating the model by using an instrumental variable method that exploits the gap between the full-time monthly earnings of minimum wage workers and the amount of public assistance in 2006 as the instrumental variable in this section.

6 Impacts on transitions of employment status

Most existing studies on minimum wage have analyzed the effect of minimum-wage hikes on the level of employment, as in this study so far, yet relatively few studies have examined the impact of minimum wage on job turnover. Portugal and Cardoso (2006) is arguably the first influential study to examine the impact of minimum-wage hikes on job flow, using Portuguese data, and found that higher minimum wages reduce labor-market turnover. Gittings and

Schmutte (2016) found a similar pattern using US data. Dube et al. (2016) also found that a higher minimum wage reduces job flow but does not affect the level of employment using US data. To explain their findings, they set up a search model allowing for on-the-job search. In a version of the theoretical model, they showed that having a higher minimum wage makes the distribution of outside offers less attractive and reduces job to job transitions. Brochu and Green (2019) also set up a search theoretic model, in which employers learn their match quality with workers after the probation period. In their model, having a higher minimum wage increases the cost of employment during the probation period and makes employers reluctant to fire existing workers. Thus, they predicted that having a higher minimum wage reduces the employment to unemployment transition, and they found consistent evidence of this based on Canadian job flow data.

In light of the development of literature regarding the minimum-wage effect on job flow, we now examine the effect of the minimum wage on the transition between non-employment and employment over one year, using short panel data constructed by matching the second- and fourth-month surveys of the LFS, as explained in Appendix B. Here we note a caveat that our matched panel data are not as rich as the job flow data used in previous studies, because we can only measure the employment status in $t-1$ and t and cannot distinguish between those who stay on a job and those who change jobs in between. Thus the analysis here is not free from time-aggregation bias (Shimer, 2005; Nordmeier, 2014).

We examine the level of the minimum wage faced by an individual between year $t-1$ and t on the employment status in year t, conditional on the employment status in year $t-1$. Since the minimum wage is revised every year, the effective minimum wages an individual faces between $t-1$ and t are MW_{t-1} and MW_t. Given that LFS is a monthly survey and the minimum wage is generally revised in October, the weight given to MW_{t-1} or MW_t should depend on the month of the survey.

For example, individuals in October surveys that ask about their employment status for the last week of September face MW_{t-1} for 12 months. Similarly, individuals in November surveys face MW_{t-1} for 11 months and MW_t for 1 month. In this fashion, we define the weighted average of minimum wages faced by an individual between year $t-1$ and t as $\widetilde{MW}_{jt}$ for the purpose of panel transition analysis. We examine if this effective minimum wage $\widetilde{MW}_{jt}$ affects the transition between non-employment and employment in the duration between the previous year and the current year by estimating the following linear probability model:

$$\begin{aligned} E_{ijt} = \gamma_0 + \gamma_1 \ln \widetilde{MW}_{jt} + X_{ijt}\gamma_2 + \theta_j + \tau_t + u_{ijt}, \\ \text{conditional on } E_{ijt-1} = \{0, 1\}. \end{aligned} \tag{8}$$

The dependent variable E_{ijt} is the dummy variable indicating if individual i in prefecture j is employed in year t. The vector X_{ijt} includes the set of age dummy variables, the fraction of each age and sex group in the population, and the unemployment rate of college graduates. We divide the sample by the employment status of the previous year, denoted as E_{ijt-1} to allow for the asymmetric effect of the minimum wage on the transitions from non-employment to employment or employment to non-employment. Because this is fundamentally a transition analysis, we did not include the prefecture-specific time-trend term, $\xi_j \cdot Year$. Instead, the prefecture fixed effects, θ_j, capture the prefecture-level unobserved heterogeneity in the transition.

The upper panel of Table 4 tabulates the estimation results for men by employment status in the previous year and age groups. The first three columns report the regression results of current employment status on the effective minimum wage, conditional on being employed in the previous year. All the estimated coefficients are negative, suggesting that the higher minimum wage reduces the transition probability from non-employed to employed, but the estimate is not statistically significant for youth, perhaps because of the small sample size. Among prime-age men, a 10% increase of the effective minimum wage reduces the transition probability by about 2.5%, whereas the average transition probability is 28%. The negative impact is even more significant for younger and older workers. In contrast, a higher minimum wage does not reduce the probability of staying employed in statistically significant ways among those who were employed in the previous year, as reported in columns 4 to 6 in Table 4. Overall, this transition analysis reveals that the negative impact of the effective minimum wage on employment among prime-age men reported in Table 3 was due to the decreased probability of transitioning from non-employment to employment.

The lower panel of Table 4 reports the estimation result of the transition analysis for women. None of the estimated impacts on the transition from non-employment to employment are statistically significant, as reported in Columns 1 - 3. A minimum-wage hike increases $P(E_t|E_{t-1})$ for prime-age women, however, by 0.07 in a statistically significant way. A 10% increase in the minimum wage increases the probability of staying employed by 0.7 percentage point relative to its mean 92%. In contrast, the increase of MW reduces $P(E_t|E_{t-1})$ significantly among older women. A 10 percentage-point increase reduces the current employment probability by 2.4 percentage points.

Overall, the transition analysis based on short-panel data implies that a minimum-wage hike reduces the transition between non-employment status and employment status among prime-age men and women. Thus, behind the insignificant impact of a minimum-wage hike on the stock of employment among prime age men and women, it reduces the job flow among them.

7 Conclusion

This paper examined the impact of minimum wages on the labor-market outcomes of various demographic groups, exploiting the heterogeneous increase in regional minimum wages in Japan from 2007 that essentially required minimum-wage commissions to index local minimum wages to the local cost of living.

An analysis of payroll records showed that the minimum wage indeed increased the probability that a worker worked near the minimum wage, suggesting that minimum wages indeed formed the wage floor of low-wage workers in Japan between 2002 and 2017. The estimated impacts of minimum wages on labor-market outcomes were heterogeneous across demographic groups. The increase of the minimum wage reduced the employment of young, less-educated men (19 - 24 years old), with an elasticity around -1.2. In contrast, minimum-wage hikes did not affect the employment of other demographic groups in a statistically significant way. An increase of the minimum wage did not significantly affect the hours worked of any of the demographic groups that we examined.

The estimated elasticity of the minimum wage on employment among less-educated young men, -1.2, is larger than the US estimates for teenagers in the literature, which roughly ranges from 0 to -0.3. This estimate, however, is not far from the estimate, -0.5, using all workers in the manufacturing sector in Japan by Okudaira et al. (2019). Furthermore, even in the US context, Baskaya and Rubinstein (2012) report that the short-run implicit elasticity of teenage employment to the minimum wage ranges between -0.75 and -1.20 using the plausibly exogenous minimum-wage increase for identification. Clemens and Wither (2019) also found the implied elasticity of low-wage workers' employment to the minimum wage to be -1, using the federal minimum-wage hike around 2008 in the middle of the Great Recession. Both studies suggest that using the minimum-wage hike independent of local labor-market conditions for identification tends to render significant negative estimates. Furthermore, Brummund and Strain (2020) reports that the disemployment effect of the minimum wage is three times larger when the state minimum wage is indexed to the cost of living than when it is not indexed. Thus, the relatively large estimate of this study is attributed to the exogenous variation of the local minimum wage caused by the introduction of the indexation of the minimum wage in Japan induced by the 2007 revision of the Minimum Wage Act. This result suggests the importance of seeking the exogeneous source of variation of local minimum wages in the research design to exploit the heterogeneous timing of minimum-wage hikes across geographic areas to address the issue of policy endogeneity of minimum-wage hikes.

Overall, the negative impacts of the minimum-wage hikes on employment is found to be limited among less-educated young men. Thus, we do not attempt to claim that minimum-wage hikes have a detrimental impact on overall employment. Instead, we point out that minimum-wage hikes increase the wage of low-wage workers at the cost of the employment of less-educated young men. Policymakers should pay particular attention to the negative impacts among this demographic group in designing minimum-wage policies, because literature has found that the employment status of youth has persistent impacts on subsequent labor-market outcomes, and the impact is particularly large in Japan (Genda et al., 2010).

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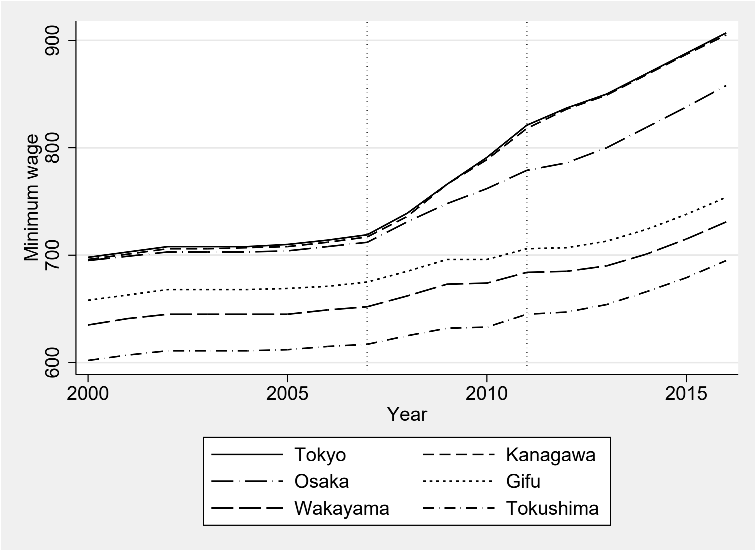
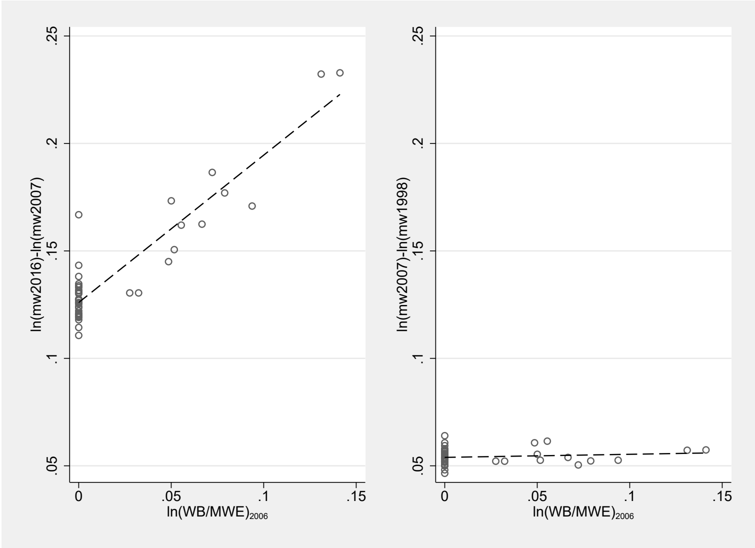
Figure 1: Trends of prefecture minimum wage



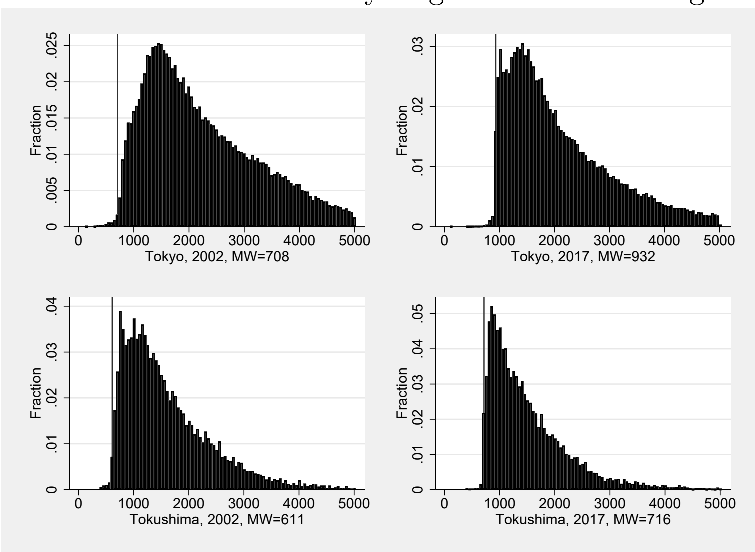
Note: Tokyo, Kanagawa and Osaka are prefectures where the welfare benefit amount exceeded the monthly earnings of minimum wage workers as of 2006. Gifu, Wakayama, and Tokushima are prefectures where the reversal did not occur.

Figure 2: Gap between MW and welfare benefit in 2006 vs. MW growth



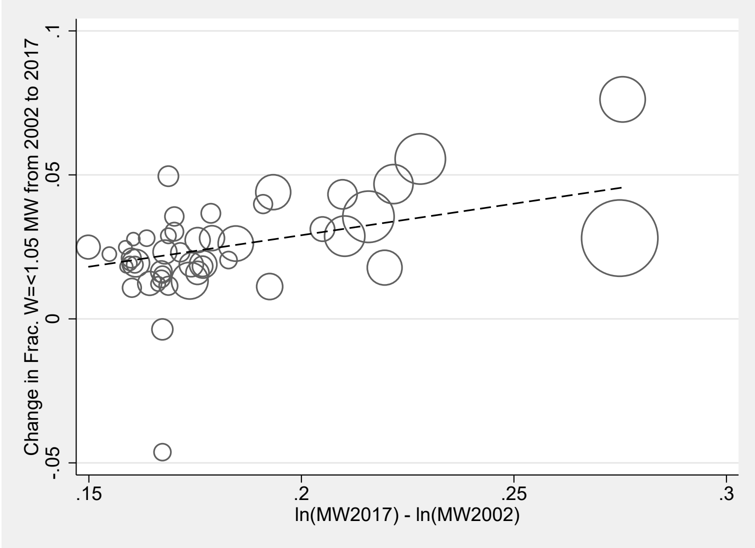
Note: Left panel shows the minimum wage growth after the revision of minimum wage act in 2007. Right panel shows that the minimum wage growth before the revision.

Figure 3: Distribution of hourly wages of all workers ages 19-64



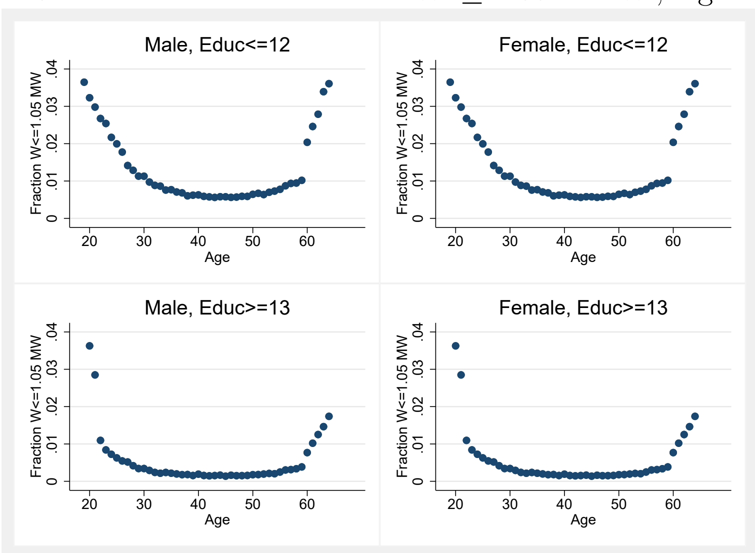
Note: The vertical lines are the minimum wages.

Figure 4: MW impact on the bottom end of the wage distribution



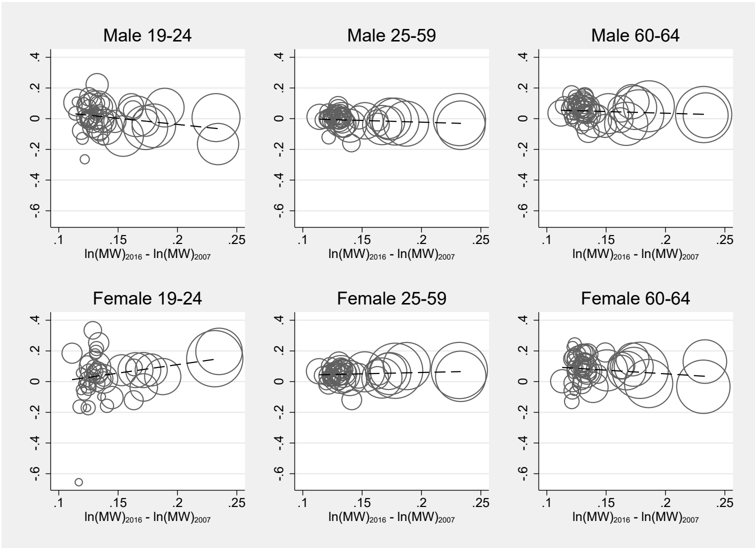
Note: The size of the bubble corresponds to the size of employment in each prefecture in 2002.

Figure 5: Fraction of workers with $W \leq 1.05 \times MW$, Ages 19-64



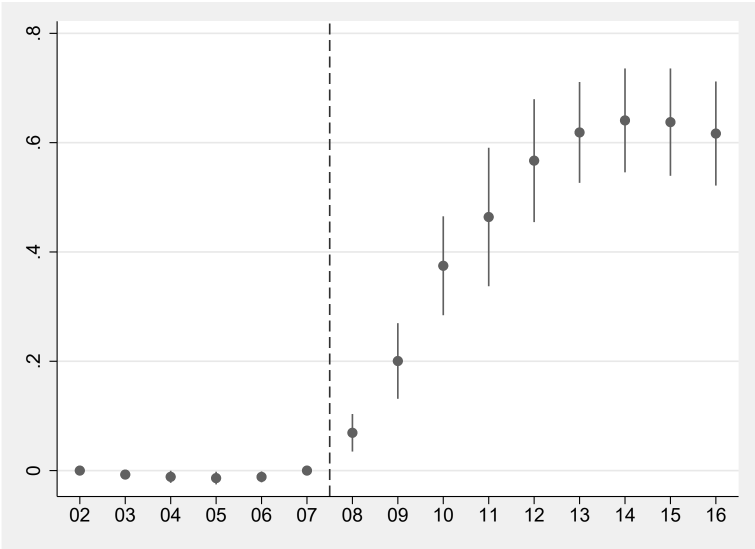
Note: Basic Survey of Wage Structure, 2002-2017 waves are all pooled. The sample is restricted to full-time workers because educational background is not available for part-time workers.

Figure 6: Change in minimum wage 2007-2016 and change in employment 2007-2016



Note: The size of the circle corresponds to the number of employments in each prefecture in 2007.

Figure 7: The gap between the welfare benefit and MW earnings in 2006 and MW

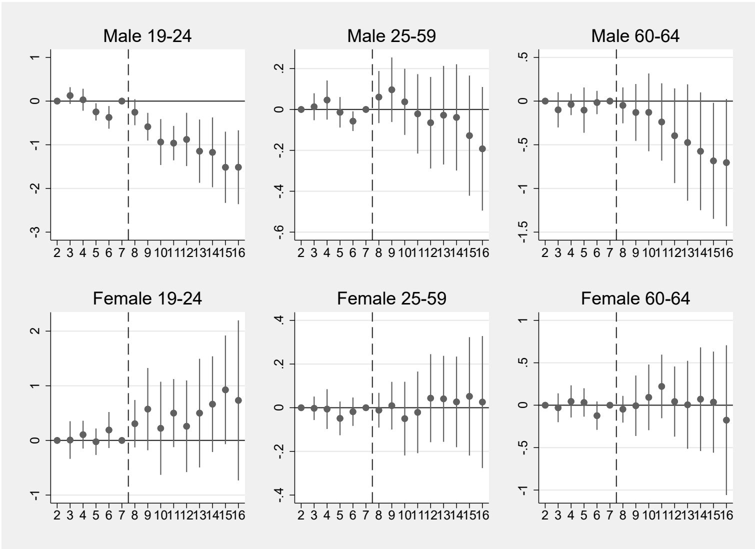


Regression coefficients and 95% confidence intervals of $\hat{\delta}_y$ of the estimated model:

$$\ln MW_{jt} = \sum_{y=2002}^{2016} \delta_y \ln(WB/MWE)_{j2006} \cdot 1(Year = y) + X_{ijt}\zeta + \nu_j + \mu_t + \xi_j \cdot Year + v_{jt},$$

where $(WB/MWE)_{j2006}$ is the ratio of welfare benefit and minimum-wage earnings in prefecture j in 2006. When the welfare benefit amount exceeds the minimum-wage earnings, $(WB/MWE)_{j2006}$ is defined to be 1. The restrictions $\delta_{2002} = \delta_{2007} = 0$ are imposed to identify the prefecture-specific linear time trends ξ_j. The 95% confidence intervals are calculated based on the standard errors robust against clustering within prefectures denoted by j.

Figure 8: The gap between the welfare benefit and minimum-wage earnings in 2006 and employment among the less educated ($Educ \leq 12$)

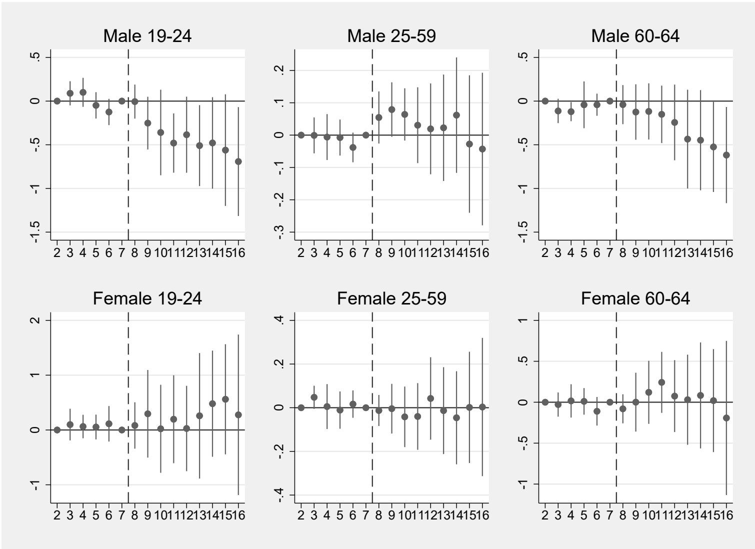


Note: Regression coefficients and 95% confidence intervals of $\hat{\theta}_y^E$ of the estimated model:

$$Emp_{ijt} = \sum_{y=2002}^{2016} \theta_y^E \ln(WB/MWE)_{j2006} \cdot 1(Year = y) + X_{ijt}\pi^E + \omega_j^E + \psi_t^E + \lambda_j^E \cdot Year + e_{ijt}^E,$$

where $(WB/MWE)_{j2006}$ is the ratio of the welfare benefit and minimum-wage earnings in prefecture j in 2006. When the welfare benefit amount exceeds the minimum-wage earnings, $(WB/MWE)_{j2006}$ is defined to be 1. The restrictions $\theta_{2002}^E = \theta_{2007}^E = 0$ are imposed to identify the prefecture-specific linear time trends λ_j^E. The 95% confidence intervals are calculated based on the standard errors robust against clustering within prefectures, denoted by j.

Figure 9: The gap between welfare benefits and minimum-wage earnings in 2006 and LFP among the less educated ($Educ \leq 12$)

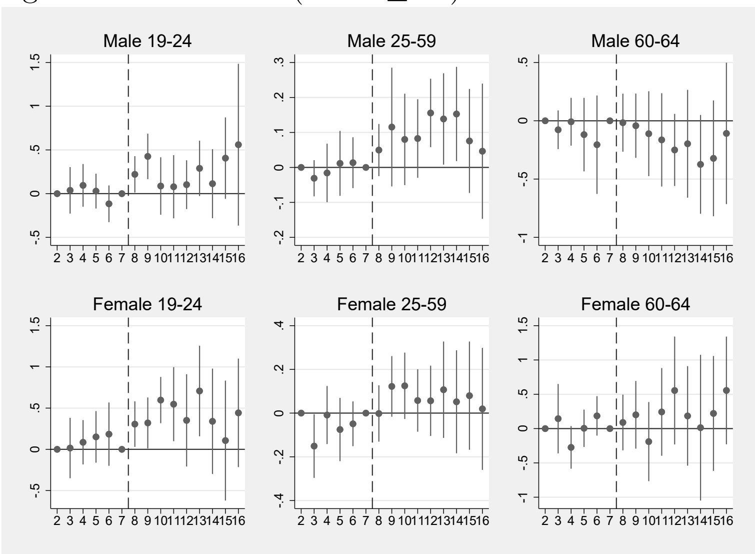


Note: Regression coefficients and 95% confidence intervals of $\hat{\theta}_y^L$ of the estimated model:

$$LFP_{ijt} = \sum_{y=2002}^{2016} \theta_y^L \ln(WB/MWE)_{j2006} \cdot 1(Year = y) + X_{ijt}\pi^L + \omega_j^L + \psi_t^L + \lambda_j^L \cdot Year + e_{ijt}^L,$$

where $(WB/MWE)_{j2006}$ is the ratio of the welfare benefit and minimum-wage earnings in prefecture j in 2006. When the welfare benefit amount exceeds the minimum-wage earnings, $(WB/MWE)_{j2006}$ is defined to be 1. The restrictions $\theta_{2002}^L = \theta_{2007}^L = 0$ are imposed to identify the prefecture-specific linear time trends λ_j^L. The 95% confidence intervals are calculated based on the standard errors robust against clustering within prefectures, denoted by j.

Figure 10: The gap between the welfare benefit and minimum-wage earnings in 2006 and ln hours worked among the less educated ($Educ \leq 12$)



Note: Regression coefficients and 95% confidence intervals of $\hat{\theta}_y^H$ of the estimated model:

$$\ln(hour)_{ijt} = \sum_{y=2002}^{2016} \theta_y^H \ln(WB/MWE)_{j2006} \cdot 1(Year = y) + X_{ijt}\pi^H + \omega_j^H + \psi_t^H + \lambda_j^H \cdot Year + e_{ijt}^H,$$

where $(WB/MWE)_{j2006}$ is the ratio of welfare benefit and MW earnings in prefecture j in 2006. The restrictions $\theta_{2002}^H = \theta_{2007}^H = 0$ are imposed to identify the prefecture-specific linear time trends λ_j^H. The 95% confidence intervals are calculated based on the standard errors robust against clustering within prefectures, denoted by j.

Table 1: Annual change of the employment rate between 2002 and 2007

	(1)	(2)	(3)	(4)	(5)	(6)
Sex	Male	Male	Male	Female	Female	Female
Age	19-24	25-59	60-64	19-24	25-59	60-64
$\ln(WB/MWE)_{j2006}$	0.10	-0.01	0.01	-0.20	-0.01	0.00
	(0.03)	(0.01)	(0.02)	(0.05)	(0.01)	(0.02)
Observations	235	235	235	235	235	235

Note: Weighted least squares using the prefecture-level sum of sampling weight in 2002 as the weight. Prefecture-level clustering robust standard errors in parentheses. Year fixed effects are included.

Table 2: Regression of the wage bin indicator on $\ln(MW)$

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$W_{ijt} \in$	(0, MW)	[MW, 1.05MW)	[1.05MW, 1.10MW)	[1.10MW, 1.15MW)	[1.15MW, 1.20MW)	[1.20MW, 1.25MW)	[1.25MW, 1.30MW]
$\ln(MW)_{jt}$	0.09	0.17	0.12	0.04	0.03	0.01	-0.03
	(0.03)	(0.09)	(0.04)	(0.01)	(0.02)	(0.02)	(0.02)
Frequency	0.01	0.02	0.03	0.03	0.03	0.03	0.03

Note: Prefecture-level clustering robust standard errors are in parentheses. Weighted least squares use sampling weights. Prefecture and year fixed effects are included. N=18,979,132

Table 3: Effect on labor-market outcomes among the less educated ($Educ \leq 12$)

	(1)	(2)	(3)	(4)	(5)	(6)
Demographic group	M19-24	M25-59	M60-64	F19-24	F25-59	F60-64
Panel A: Employment						
OLS	-0.78	-0.06	-0.53	0.02	0.02	0.13
	(0.32)	(0.08)	(0.31)	(0.38)	(0.12)	(0.25)
IV	-0.93	-0.08	-0.63	0.07	0.08	0.37
	(0.43)	(0.13)	(0.39)	(0.41)	(0.16)	(0.37)
Mean	0.80	0.89	0.67	0.70	0.65	0.42
Observations	152484	1810572	403582	133423	1907371	482288
P-value of Hausman-Wu	0.77	0.34	0.66	0.74	0.27	0.58
Panel B: LFP						
OLS	-0.77	0.00	-0.24	-0.07	0.04	0.23
	(0.23)	(0.07)	(0.25)	(0.38)	(0.11)	(0.25)
IV	-0.56	0.02	-0.46	0.09	0.02	0.47
	(0.32)	(0.08)	(0.29)	(0.46)	(0.13)	(0.37)
Mean	0.91	0.95	0.74	0.79	0.69	0.44
Observations	152484	1810572	403582	133423	1907371	482288
P-value of Hausman-Wu	0.59	0.21	0.57	0.87	0.25	0.66
Panel C: ln($Hours$)						
OLS	-0.43	-0.16	-0.18	0.45	-0.11	0.07
	(0.30)	(0.13)	(0.22)	(0.27)	(0.13)	(0.56)
IV	-0.06	0.14	-0.31	0.31	0.09	0.15
	(0.26)	(0.07)	(0.30)	(0.31)	(0.16)	(0.53)
Mean of Hours	44.14	47.24	41.39	38.62	34.36	32.20
Observations	120857	1604541	268544	92613	1236848	201236
P-value of Hausman-Wu	0.59	0.21	0.57	0.87	0.25	0.66

Note: Prefecture-level clustering robust standard errors are in parentheses. Prefecture and year fixed effects and the prefecture-specific year trend are included.

Table 4: Effect on current employment among men by employment status of the previous year among the less educated ($Educ \leq 12$)

	(1) NE19-24	(2) NE25-59	(3) NE60-64	(4) E19-24	(5) E25-59	(6) E60-64
Panel A: Men						
OLS	-0.53	-0.24	-0.41	-0.08	-0.03	-0.09
	(0.47)	(0.10)	(0.19)	(0.16)	(0.03)	(0.10)
IV	-0.86	-0.33	-0.32	0.07	0.00	0.03
	(0.65)	(0.11)	(0.27)	(0.20)	(0.02)	(0.12)
Mean	0.41	0.28	0.14	0.94	0.97	0.89
Observations	5785	42371	26294	20761	358100	66816
P-value for Hausman-Wu	0.256	0.184	0.401	0.288	0.194	0.394
Panel B: Women						
OLS	0.56	-0.08	-0.05	0.19	0.05	-0.24
	(0.40)	(0.05)	(0.08)	(0.24)	(0.03)	(0.09)
IV	0.80	0.00	-0.01	0.09	0.08	-0.19
	(0.35)	(0.08)	(0.08)	(0.30)	(0.04)	(0.10)
Mean	0.37	0.17	0.07	0.90	0.92	0.86
Observations	7129	144944	60516	16008	277031	50382
P-value for Hausman-Wu	0.151	0.184	0.407	0.240	0.193	0.368

Note: Prefecture-level clustering robust standard errors are in parentheses. Prefecture and year fixed effects are included. $\ln \widetilde{MW}$ is the effective minimum wage between $t-1$ and t.

A Minimum-wage determination in Japan and the revision of the Minimum Wage Act in 2007

This section introduces the policy process of how regional minimum wages, set for each of the 47 prefectures, are determined. The Ministry of Health, Labour and Welfare assembles the Central Minimum Wage Council, consisting of the representatives of labor, management, and public welfare, based on the tripartite principle. The Council divides all 47 prefectures into four regional groups, based primarily on regional income level, and provides a recommendation for the amount of the hourly rate of the minimum-wage hike, called a "Criterion (*Meyasu*)," for each of the four regional groups, based primarily on the average annual wage growth of workers in establishments with less than 30 workers, captured by the Fact-finding Survey on Minimum Wages (*Saitei Chingin ni Kansuru Jittai Chosa*). The Central Minimum Wage Council generally starts in June and devises the Criterion by the end of July. Upon receiving its Criterion, the Regional Minimum Wage Council, assembled in each of the 47 prefectures based on the tripartite principle, decides the amount of the minimum-wage hike for that prefecture in August, and the new minimum wage comes into effect from October 1. This two-step process using the Criteria has taken place every year since 1978, and local minimum wages have increased every year with just a few exceptions.

The partial revision of the Minimum Wage Act in 2007, which took effect on July 1, 2008, did not principally change the policy process, but the revision added a new requirement that must be considered when determining regional minimum wages. The new requirement is manifested in Article 9, Paragraph 3, which governs regional minimum wages: "When considering [the] living expenses [mentioned in the previous Paragraph of the Act], consistency with policies concerning public assistance benefits shall be considered, so that workers can lead a life with a minimum standard of health and culture" (authors' translation). This provision aims to resolve the so-called reversal phenomenon (*gyakuten gensho*), which refers to the fact that the total amount of benefits one can receive while being on public assistance has come to exceed the monthly income one can earn when working full-time on minimum wages. The reversal occurred in 12 prefectures as of fiscal year 2006, according to a letter submitted to the Minister of Health, Labour and Welfare from the Central Minimum Wage Commission in 2008.[5] In practice, the aim was to close the gap between the amount of welfare benefits and the earnings of full-time minimum-wage workers in the subsequent two years, and up to five years at maximum in prefectures where significant adverse impacts were expected, by raising minimum wages (Nakakubo, 2009).

Public assistance is a national income transfer program for the poor. Those who pass asset and income tests are eligible to receive the benefit; the tests are intrusive, such that the caseworker investigates all the assets and the family members' ability to support the welfare applicants. Because of this rigorous approval process, the government estimated that the take-up rate among eligible is 32.1% as of 2007 based on Comprehensive Survey of Living Conditions.[6] The benefit amount is the difference between the standard welfare

[5] The 12 prefectures include Hokkaido, Aomori, Miyagi, Akita, Saitama, Chiba, Tokyo, Kanagawa, Kyoto, Osaka, Hyogo, and Hiroshima. The gap amount as of 2006 was 63 yen for Hokkaido, 20 yen for Aomori, 31 yen for Miyagi, 17 yen for Akita, 56 yen for Saitama, 35 yen for Chiba, 100 yen for Tokyo, 108 yen for Kanagawa, 47 yen for Kyoto, 53 yen for Osaka, 36 yen for Hyogo, and 37 yen for Hiroshima.

[6] Estimation of the number of low-income households below the welfare standard amount by the public

amount (*Seikatsu Hogo Gaku*) and the total income, including labor earnings and other social security transfers, such as pension benefits. The standard welfare amount depends on the number of family members, the earnings capacity of each member, and the local cost of living. As examples, Ministry of Health, Labour, and Welfare documents that the standard welfare amount for a 68-year-old single-person household is 80,820 yen (about US$800) in metropolitan districts, whereas it is 62,640 yen (about US$620) in rural districts.[7] The difference in welfare benefits largely reflects the difference in housing costs. The regional difference in the standard welfare amount became the source of heterogeneous growth of minimum wages when the revision of the Minimum Wage Act required the indexation of the minimum wage to the standard welfare amount.

B Data sets

This study uses the Labour Force Survey (LFS) to capture such labor-market outcomes as employment status and hours worked. The drawback of the LFS is the lack of accurate measurement of the hourly wage rate. To supplement this drawback, we also use the Basic Survey on Wage Structure (BSWS).

The LFS is a household survey that randomly selects about 40,000 households by using a geographically stratified sampling structure; we use data from 2002 to 2016 for our analysis, as the sampling structure was substantially revised from 2002. The survey was not conducted in the area severely affected by the natural disaster. This case includes Iwate, Miyagi, and Fukushima prefectures between March and August 2011 and Kumamoto prefecture during May and June 2016. Since the geographic stratum is the region, a collection of prefectures, no survey district may be sampled from a small prefecture. This case includes Wakayama prefecture for July 2009 and June 2014 and Tottori prefecture for July 2015.

The survey has a rotating sample structure; a selected household responds to the first 2 months of the surveys, followed by a 10-month break, and then the last 2 months of the surveys. The surveys of different months can be matched using survey district ID, birth year, birth month, and sex. The fourth-month survey is called the special survey and records educational background, as well as annual income, in addition to the labor-force status recorded in other months. To obtain the educational background of individuals, we used only the households that could be matched to its fourth-month surveys. Given that the fourth month is available, 98% is matched for the third month, 86% is matched for the second month, and 85% is matched for the first month. The labor-force status of each household member selected for the survey is classified as belonging to one of the following nine categories: (1) Mainly working, (2) In school but also working, (3) Homemaker but also working, (4) Working but temporarily out of work, (5) Unemployed, (6) In full-time education, (7) Homemaker, (8) Other (the elderly, etc.), and (9) Unspecified. We define categories 1-3 as working and 4-8 as not working. The analysis sample is restricted mainly

assistance section of the social and assistance bureau of the Ministry of Health, Labour and Welfare released on April 9th, 2010. `https://www.mhlw.go.jp/stf/houdou/2r98520000005olm-img/2r98520000005oof.pdf` (Accessed on November 9, 2020)

[7] The document titled Public Assistance System by the Ministry of Health, Labour, and Welfare summarizes the program. `https://www.mhlw.go.jp/english/topics/social_welfare/dl/outline_of_the_public_assistance_system_20101004.pdf` (Accessed on November 9, 2020).

to non-college graduates between the ages of 19 and 64, because, as we will see below, the minimum wage is relevant to non-college-graduate workers.

A shortfall of the LFS is a lack of an accurate measurement of the hourly wage rate. The only individual earnings information available in the survey is the range of annual income, which is asked in the last month of the survey (the second month in the second year). From this annual earnings information and the hours worked in the last week of the previous month, we can construct hourly wage by imposing assumptions. However, the constructed hourly wage is not precise enough to define those who are treated by the minimum-wage hike and those who are not, as was done by Cengiz et al. (2019) and Clemens and Wither (2019). Thus to analyze the impact of the minimum wage on hourly wages, we rely on the BSWS instead. Also, we do not restrict the analysis sample to "low wage" workers in the employment analysis using the LFS, as Cengiz et al. (2019) did using the Current Population Survey in the US.

The BSWS is an ideal data set for analyzing the impact of the minimum wage on wages because of its accurate measurements of individual workers' monthly earnings and monthly hours worked, transcribed from payroll records, and the large sample size, in that it contains around 1.2 million workers annually. This annual survey chooses establishments based on stratified sampling by prefecture, industry, and size of the establishment; it focuses on private establishments with 5-9 employees that belong to firms that hire 9 or fewer employees, and all private and public corporations' establishments with 10 or more employees. According to 2014 Economic Census for Business Frame (Ministry of Internal Affairs and Communications) conducted in 2014, private establishments with 5 or more employees hire 88% of total employees. The survey asks establishments to randomly sample its employees from payroll records and provide information on their employees' scheduled cash earnings and scheduled work hours in June. The survey contains information on about 1.2 million employees at approximately 45,000 establishments each year. Using the information from this survey, we calculate the hourly wage rate for our analysis as [(Scheduled cash earnings - Commuting allowance - Perfect attendance allowance - Family allowance) / Scheduled work hours], following the definition of hourly wage in the Minimum Wage Act. We used 2002-2017 surveys to construct the analysis sample that includes all individuals between ages 19 and 64. The BSWS is designed to calculate the wages by the cell of prefecture $\times$ industry $\times$ establishment size with substantial precision. To keep the sampling error below a certain level, larger establishments, which are relatively rare and more heterogeneous in the population, are oversampled. To reflect this peculiar sampling structure of the survey, we use the sampling weight in the regression analysis to recover the population distribution throughout the analysis based on the BSWS. We also use the sampling weight so that the regression coefficients recover the population average treatment effects (Solon et al., 2015).

C Industrial and occupational composition of minimum-wage workers

Table C1: Industry composition of MW workers in 2007

	(1)
	pct
Mining	0.07
Construction	0.70
Manufacturing	22.74
Electlicity gas and water	0.24
Information and communications	0.93
Transport and postal services	4.12
Wholesale and retail Trade	16.69
Finance and insurance	2.91
Real estate	6.94
Scientific research and technics	1.07
Hotel and restaurant	13.94
Life and amusement services	9.96
Education	3.96
Medical and health care	3.50
Compound services	1.14
Other services	11.09
Total	100.00

Table C2: Occupation composition of MW workers in 2007

	(1)
	pct
Administrative workers	0.18
Specialist professionals	6.98
Clerical workers	0.46
Sales workers	34.90
Service workers	27.60
Security workers	2.09
Transport operation workers	2.72
Manufacturing process workers	25.07
Total	100.00