Impacts of Minimum-Wage Hikes on Wages and Employment in Japan*

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Abstract

We examine the impacts of heterogeneous increases in regional minimum wages following the revision of the Minimum Wage Act in 2007 in Japan using payroll records and Labor Force Survey. We find minimum-wage hikes raised the wages of less-skilled workers but reduced the employment of less-educated young-adult and prime-age men. It also significantly reduced the working hours of both young and older workers of both genders. At the same time, the minimum-wage hikes encouraged the labor-force participation of middle-aged women, whose labor-force attachment had been historically low. A panel analysis based on matched Labor Force Survey indicates that minimum wage hike decreased job flows of prime-age men and women.

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1 Introduction

The minimum-wage hike is widely supported as a policy tool to alleviate poverty in many countries. In addition, the minimum wage has gained support as an income policy to boost wages to initiate wage-price dynamics in Japan, where wages have apparently stagnated despite a very tight labor market since the early 2010s. In the Council on Economic and Fiscal Policy on November 24, 2015, Prime Minister Shinzo Abe committed to increasing the minimum wage 3% annually until the average of regional minimum wage reaches 1,000 yen. This policy stance is backed by the International Monetary Fund (IMF) (the Conclusion of the Mission for the 2018 Article IV Consultation with Japan published on October 4, 2018), and IMF staff economists Aoyagi et al. (2017) reported a significant impact of minimum-wage hikes on wage increase. Some business owners and economists, however, express their concern for job loss among unskilled workers due to the minimum-wage hikes.

Several studies have examined the impact of minimum wage on employment in Japan, but most of them analyzed the experience before the revision of the Minimum Wage Act in 2007. A challenge for researchers analyzing the minimum-wage impact on employment in Japan has been the fact that regional minimum wages are routinely raised every year by similar percentages across regions. This institutional feature has made it difficult to disentangle the minimum-wage impact from the macroeconomic shock common across prefectures. The situation was somewhat similar to the situation of the US before each state started to set its own minimum wage above the federal minimum wage in the late 1980s. The revision of the Minimum Wage Act in 2007 changed the situation, however. Japan's minimum wage is set individually for each of the 47 prefectures reflecting local labor-market conditions, particularly the wage growth of workers working in micro firms that hire 30 or less employees. The revision of the Act required the government to set the minimum wage, referring to the amount of welfare benefits that is determined by the local cost of living; specifically, the revised Act aimed to close the income gap between those working for the minimum wage and those receiving welfare benefits. The size of the initial gap between the earnings of full-time minimum wage workers and the amount of welfare benefit was larger in urban prefectures, where the cost of living is higher. Accordingly, to fill the gap between the minimum wage and the welfare benefit, 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 ranged from 12% in Tokushima prefectures to 26% in Tokyo and Kanagawa 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, using large-scale micro data of government statistics, namely the Basic Survey of Wage Structure and the Labor Force Survey.

We first document how minimum wage affects the wage distribution using the Basic Survey of Wage Structure, an ideal data set for analyzing the impact of 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, containing around 1.2 million workers annually. The analysis demonstrates that the minimum wage had become binding in high-wage prefectures, such as Tokyo prefecture, by 2017, whereas it had not been binding in 2002. This change is in contrast to the experiences in low-wage prefectures, such as Tokushima prefecture, where the minimum wage had been binding ever since 2002. The regression analysis allowing for prefecture and year fixed effects also confirms that the minimum-wage hike increased the fraction of 'minimum wage workers' who are paid just above the minimum wage. At the same time, we show that only a fraction of workers are minimum wage workers, namely those who work below the minimum wage + 5%, and they are concentrated among less educated, younger or older workers of both sexes. This finding motivates us to focus on workers without higher education to analyze the impact of the minimum wage on employment.

The accurate record of hours worked and monthly earnings in the BSWS further allows us to assess the impact of minimum wage hike on hours worked and monthly earnings. We find minimum wage hike uniformly reduces hours worked of all types of workers and particularly negative impact on younger and older men and women. For these groups of workers, the estimated elasticity range from -0.26 to -0.56. Because of this significant reduction of hours worked, minimum wage hike reduces monthly salary of younger workers of both sexes with the estimated elasticity of -0.47 and -0.62. On the other hand, minimum wage hike do not affect the the monthly salary of prime-age and older workers of both sexes in statistically significant ways.

We then analyze the effect of the minimum wage on employment, labor force participation and hours worked using the monthly Labor Force Survey, which randomly samples around 50,000 households each month. To pay particular attention to less-skilled workers, we focus only on those with a complete high-school education or less. In addition, as a background development of the Japanese labor market, we observe an increase in labor-force participation among demographic groups whose labor-force participation had been traditionally low, such as women and older adults. The trend of labor-force participation may well be different across regions, and thus we employ a model allowing for prefecture, year-month fixed effects and prefecture-specific year trend to allow for prefecture-specific trends.

The estimated impacts of minimum-wage hikes on employment are heterogeneous across workers' characteristics. Among less-educated young men between the ages of 19 and 24, a 10% minimum-wage hike reduces the employment rate by 4.8 points where the average employment rate is 80%, which implies that the elasticity is around -0.6. Similarly, a 10% minimum-wage hike reduces the employment rate of prime-age men by 1.2 points where the average employment rate is 89%. In contrast, a 10% minimum-wage hike increases the employment rate of older men ages 65-69 by 3 points where the average employment rate is 45%. We find similar impacts of minimum-wage hikes on labor-force participation. Looking at the impact on women, a 10% minimum-wage hike reduces the employment rate of prime-age women by 1.3 percentage points where the average employment rate is 65%. It also increases that of older women ages 65-69 by 2.7 percentage points where the average employment rate is 26%. In addition, we find that a minimum-wage hike has a larger impact on labor-force participation than on employment for prime-age and older less-educated women. Overall, a minimum-wage hike decreases the employment rate of less-educated young and prime-age men. In contrast, a minimum-wage hike moderately increases the employment of less-educated prime-age women and significantly increases the employment of older people of both genders.

We further look into the effect of minimum-wage hikes on the hours worked among those who work based on the Labor Force Survey. We find that a minimum-wage hike reduces the hours worked among young and older workers of both genders, with the elasticity between -0.17 and -0.40. These estimated negative impacts are almost in line with the estimated impacts based on payroll records while the size of estimated impacts are attenuated perhaps due to the non-classical self-report measurement error of the hours worked.

We checked the robustness of our results from two aspects. First, we consider the possible impact of minimum wage hike on employment growth instead of employment level as pointed out by previous literature. If minimum wage affects the employment growth of a region, controlling for region-specific linear time trends suffers from over-control bias. We confirmed that our estimation results do not change almost at all between specifications with and without region-specific liner time trends. Second, to address potential policy endogeneity bias that minimum wage is determined by temporal shocks to local labor markets, we implement the instrumental variable estimation exploiting the institutional feature that minimum wage hike after 2007 revision of the minimum wage act was partially motivated to fill the gap between welfare benefit amount and minimum wage earnings. Essentially using the gap between welfare benefit and minimum wage earnings in 2007 as instrumental variable for minimum wage, we implement instrumental variable estimation. Due to the smaller variation of minimum wages explained by the instrumental variable, the estimates are less precise than the OLS estimates but we could still confirm the negative impact of minimum wage hike on employment of young men and positive impact of it on prime-age women. More importantly, the changes of coefficients between OLS and IV estimates are not consistent with the typical policy endogeity story. In the end, we take the OLS estimates with prefecture-specific linear time trends as the preferred estimates.

We construct the short panels by matching the outgoing rotation groups of adjacent years of LFS to examine the impact of minimum wage on the transition of employment status in one year period. Among men, the analysis indicates that minimum wage hike reduces the transition probability of prime-age and older men out of employment in the previous year to be employed in the current year. Among women, minimum wage hike increases the probability of prime-age women who are employed in the previous year staying employed in this year. Thus, minimum wage hike generally reduces the transition probability between being not employed and employed.

Literature

Economists have studied the employment effect of the minimum wage for many years without reaching a consensus. (Neumark and Wascher, 2008, Chapter 3) covered much of the research on the topic up until 2008 and concluded that reliable studies point to a disemployment effect of minimum wage on low-skilled workers. More recent developments of the literature in the US emphasize the importance of controlling for the state-specific time trend (Allegretto et al., 2011) or detailed geographic area-specific shock (Dube et al., 2015), because these unobserved economic conditions may well be correlated with state-specific decisions on the minimum wage. Both Allegretto et al. (2011) and Dube et al. (2015) found that minimum-wage hikes had a small impact on teenage employment. Neumark et al. (2015), however, criticized their approach, claiming that the linear state-specific trend is not a proper way to capture the long-term macroeconomic condition and that choosing control regions from a narrowly defined geographic area is not a proper way to select control regions. In the end, Neumark et al. (2015) pointed out that controlling for region $\times$ period fixed effects removes a significant fraction of informative identification variation. Regarding the region-specific unobserved time trend, Meer and West (2015) pointed out that a minimum-wage hike can affect the growth rate of employment of the affected region. In this case, conditioning on the region-specific employment trend causes a bad (an over) control problem. To address the problem, Meer and West (2015) proposed estimating models that use the employment growth rate as the dependent variable.

All recent US studies concur on the importance of controlling for region-specific shock that can be correlated with the minimum wage, but they disagree on the proper construction of a counterfactual situation in terms of selecting control areas or modeling underlying region-specific time trends. 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 more centralized and mechanical. This institutional difference arguably makes policy endogeneity less of a concern, but it is not perfectly resolved. Thus, in light of recent developments of the literature in the US, we check the robustness of our results to the inclusion of period-specific regional effects or region-specific time trends. Overall, we show that our basic results are quite robust against the choice of the counterfactual situation.

This paper also aims to add knowledge to a growing body of literature on the effect of minimum wages on employment in Japan. Using data from the 2002 Employment Status Survey, Tachibanaki and Urakawa (2006) report that a high Kaitz index raises the wages of women in their 20s, but does not reduce their employment. In contrast, using prefecture-level data from the 2000 Population Census of Japan, Yugami (2005) finds a positive correlation between the minimum wage and the unemployment rate. Ariga (2007) uses the School Basic Survey and The Labour Market for New Graduates to construct prefecture-level panel data from 1977 to 2002 and reports that a higher minimum wage raises the initial salary of high-school graduates, but lowers the number of job openings. None of these studies allowed for prefecture-level unobserved characteristics that could be correlated with regional minimum wages, but controlling for prefecture fixed effects is crucial to studying the impact of the regional minimum wage on employment, because minimum wages are systematically lower in rural prefectures than in urban prefectures and the rural/urban differences may well affect employment outcomes.

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 hike is more likely to lose her job, using the 1993-1999 panel surveys on women conducted by the Institute for Research on Household Economics. Overall, the 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. Looking at his estimation results closely, we notice that the minimum-wage hikes indeed decreased the employment of non-regular workers, but the impact was not precisely estimated because of a 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 large sample of government surveys. Relying on a large sample of the census of manufacturers, Okudaira et al. (2019) reported the impact of minimum-wage hikes after the 2007 revision negatively affected the employment, with an estimated elasticity around -0.5 among all workers. They also identified the degree of monopsony captured by the gap between the predicted labor share from the estimated production function and found a larger negative impact in the region where the labor market is estimated to be competitive.

We are the first to estimate the effect of the minimum wage on employment using the policy change after 2007 in Japan, based on large-scale survey data covering all sectors and paying attention to the potential endogeneity of regional minimum wages.

2 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, Labor 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," for each of the four regional groups, based primarily on the average annual wage increase of workers in establishments with 29 or fewer workers. The Central Minimum Wage Council generally starts in June and devises the Criterion by the end of July. Upon the receipt of 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. The Regional Minimum Wage Council decides the minimum wage of the prefecture in August, and the new minimum wage comes into effect from October 1st. This two-step process using Criteria has taken place every year since 1978, and the regional 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). The aim of this provision is 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. 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 are expected, by raising minimum wages (Nakakubo, 2009).

In practice, the revision of the Minimum Wage Act made the Central Minimum Wage Council indicate the amount of the minimum-wage hike required to fill the gap between public assistance benefits and the minimum wage for prefectures where public assistance benefits exceed the minimum wage, in addition to the traditional Criterion for the minimum-wage hike. For prefectures where the gap amount is indicated, the Central Minimum Wage Council recommends adopting the Criterion or the gap amount, whichever is larger. The Regional Minimum Wage Councils generally follow the recommendations to eliminate the gap.

Eliminating the gap between public assistance benefits and the minimum wage effectively became the major determinant of the regional minimum wage after 2007. This change in the minimum wage determination structure is clearly seen in Figure 1. The figure shows that the growth rate of the minimum wage kinked from 2007 in urban prefectures, such as Tokyo, Kanagawa, and Osaka, while there was no such kink for rural prefectures, such as Gifu, Wakayama, and Tokushima. The divergence of the minimum wage in urban prefectures from that in rural prefectures is largely due to the larger gap between the minimum wage and the welfare benefit in urban prefectures, reflecting the high housing cost in urban prefectures. As evidence, the left panel of Figure 2 shows the relationship between the initial gap, obtained from the report by the Central Minimum Wage Council, and the minimum wage increase between 2007 and 2016. [1] There is a clear upward-sloping relationship, indicating that

[1] Total public assistance benefits of a person aged 12-19 is the sum of public assistance payments (Category 1 benefits + Category 2 benefits + End of year assistance + Winter allowance) and housing assistance.

the increase in the minimum wage was indeed greater in prefectures with a greater public assistance-minimum wage gap. In contrast, the right panel of Figure 2 shows that the gap and the minimum wage growth between 1999 and 2008 are hardly correlated. In sum, regional minimum wage growth after 2007 is mainly determined by the gap between the public assistance benefit and the minimum wage.

A probable threat to identification is that the initial gap between the minimum wage and welfare benefits, which determines the subsequent minimum wage growth, was correlated with the underlying trends of labor-market outcomes. To examine this possibility, we examine the employment rate growth between 2002 and 2007 and the gap in 2007. Figure 3 show that the gap in 2007 is not systematically correlated with the employment growth of either gender between 2002 and 2007.

3 Data

We draw on two government survey data sets to measure wages and employment.

For wages, we use micro data from the Basic Survey on Wage Structure (BSWS). This annual survey chooses establishments based on stratified sampling by prefecture, industry, and size of 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 the Economic Census conducted in 2014, the 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

Among the components, housing assistance makes up a significant part of the total benefits.

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 69. The BSWS is designed to calculate the wages by the cell of prefecture × industry $\times establishmentsize$ with certain precision. To keep the standard error / mean ratio below a certain level, larger establishments are over sampled. To reflect this peculiar sampling structure of the survey, we use the sampling weight to recover the population distribution throughout the analysis. We also use the sampling weight so that the regression coefficients recover the population average treatment effects (Solon et al., 2015).

For employment, we use micro data from the Labor Force Survey, which randomly selects about 40,000 households by using a geographically stratified sampling structure; we use data from 2002 to 2017 for our analysis, as the sampling structure was substantially revised in 2002.[2] The survey has a rotating sample structure; a selected household responds to the first 2 months of the surveys, with a 10-month break, and 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 only used 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

[2]The survey was not conducted in the area severely affected by the natural disaster. This 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 includes Wakayama prefecture for July 2009 and June 2014 and Tottori prefecture for July 2015.

as not working. We used 2002-2016 surveys to construct the analysis sample. The analysis sample is restricted mainly to non-college graduates between the ages of 19 and 69 because, as we will see below, the minimum wage is relevant to less-educated workers.

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 Impacts on wages

This section examines the extent to which minimum wages act as a binding constraint by looking at the relationship between the minimum wage and the wage distribution in each year and each prefecture. Figure 4 shows the relationship between the wage distributions of all workers ages 19-69 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. As can be seen in the panels for Tokushima, 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 important role in determining the shape of the wage distribution even before the partial revision of the Minimum Wage Act in 2007. In contrast, the panel for Tokyo for 2002 shows that only a very small share of workers worked for the minimum wage and that the minimum-wage level hardly affected the shape of the wage distribution at all. By 2010, however, the shares of both men and women working for the minimum wage had increased, and the shape of the wage distribution had begun to be dented by 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 an effective 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 between 2002 and 2017, only 3.7% of workers earned this range of hourly wage. This suggests that we need to narrowly restrict the analysis sample to those who are likely to be affected by minimum-wage hikes. Figure 6 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 are positively associated.

Table 1 reports the estimation results of linear probability models where the dependent variables are the dummy variables that takes one if the wage falls within a certain range relative to the minimum wage. The models are estimated with prefecture and year fixed effects by weighted least squares using the sampling weight. Standard errors are robust against within prefecture error correlation. Column 1 shows that minimum-wage hikes increase the fraction of non-compilers. On average, 1% of workers work below the minimum wage and a 10% increase of the minimum wage increases the non-compilers by 0.9%. In contrast, Column 2 indicates that a 10% minimum-wage hike increases the fraction of workers earning above the minimum wage but below 1.05$\times$ minimum wage by 1.7 %, whereas about 2% of workers belong to this category. Similarly, Column 3 indicates that minimum-wage hike increases the fraction of workers earning above 1.05$\times$minimum wage and below 1.10 $\times$ minimum wage, suggesting that minimum-wage hikes affect wage of workers who earn above minimum wage, called the spillover effect or ripple effect. This spillover reaches up to 20% above the minimum wages.

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. To identify the characteristics of workers who earn around the minimum wage, but the caveat is that the BSWS does not record the educational background of part-time workers, which consists of about 20% of the analysis sample and are more likely to work as the minimum wage workers; only around 1% of full-time workers are minimum-wage workers, while around 10% of part-time workers are minimum-wage workers. With this drawback in mind, Figure 6 plots the fraction of workers that earns below $1.05 \times MW$ by educational background and age. It is apparent that the fraction of minimum-wage workers is substantially high among low-educated (high -school or less) younger and older workers. Thus to identify the impact of minimum wage on employment, we need to focus on low-educated younger and older workers.

4.2 Impacts on monthly hours worked and salary

The payroll record, BSWS, includes monthly hours worked, including overtime, and monthly salary payment, including overtime pay. We now examine the impact of minimum-wage hikes on monthly hours worked and salary payment. To repeat, the BSWS does not record the educational background of part-time workers, who are more likely to work as minimum-wage workers. Thus, we do not divide the sample by educational background. Table 7 tabulates the regression results of the natural log of monthly hours worked on the natural log of minimum wage by sub-samples. The first column shows that a 10% increase in the minimum wage reduces hours worked by 1.3%. The sub-group analysis further indicates that the impacts are larger for younger and older workers of both genders.

Regarding monthly salary, Table 3 tabulates the regression result of the natural log of monthly salary on the natural log of minimum wage. Column 1 shows that minimum-wage hikes do not affect the monthly salary, on average; however, the sub-sample analysis shows quite heterogeneous impacts. On one hand, a minimum-wage hike reduces the monthly salary of younger workers of both genders; a 10% increase in minimum wage reduces younger workers' salary by about 5 - 6%. On the other hand, a 10% increase of the minimum wage increases the monthly salary of prime age men by about 2%. Because the number of prime age men is about five times greater than the number of younger men and women, in total, a minimum-wage hike does not affect the monthly salary, on average.

Overall, the minimum-wage hike negatively affects the hours worked of younger workers, and it in turn negatively affects their monthly salary.

5 Analysis based on Labor Force Survey

5.1 Employment

We have seen that minimum-wage increases since 2007 raised the wages of low-skilled workers. We now ask whether these wage increases have been associated with a decrease in their employment. Since the number of workers who are affected by minimum hikes is limited, we focus on the employment of less-educated workers, namely those with a high-school education or less.

The data we use are from the Labor Force Survey, 2002 to 2016. Figure 7 shows on the horizontal axis the change in the natural log of the minimum wage between 2002 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 minimum wage, namely Tokyo and Kanagawa. For prime-age and older men, we find slight negative association. In contrast, for females, the increase of minimum wage is positively associated with the change of the employment of young and prime-age women.

The association of minimum wage change and employment rate change between 2007 and 2016 of course does not imply the causal impact of minimum wage on employment.

A particular concern is the association of pre-trend of employment rate and the change of minimum wage; if employment rate of certain group is on the downward trend in prefectures experiencing significant minimum wage hike, we mistakenly attribute the decrease in employment rate to the increase in minimum wage. To examine how the minimum wage hikes are associated with the pre-trends of employment, Figure 8 plots the change of employment rate between 2002 and 2007 against the change in MW between 2007 and 2016. The figure suggests that employment rate of young men were growing before 2007 in prefectures that experienced significant minimum wage hike afterward. We find opposite result for younger women. These results suggest the importance of allowing for prefecture specific time trends in the estimation.

To examine the impact of the minimum wage on various labor-market outcomes, we estimate the following model using lower-educated people, namely those with a high-school education or less, as the analysis sample:

$$Y_{ijt} = \beta \ln MW_{jt} + \theta_j + \tau_t + \xi_j \times Year + u_{ijt} \tag{1}$$

where Y_{ijt} is the indicators for being employed or in the labor force, and hours worked of individual i living in prefecture j in year t. The explanatory variable $\ln MW_{jt}$ is the natural logarithm of minimum wage in prefecture j in year-month t. This specification allows for a prefecture-specific linear time trend in terms of calendar year.

Recent literature on the minimum wage often adopts an event study approach that includes lags and leads of $\ln MW_{it}$ to validate the common trend assumption. This approach works well when the minimum wage in some regions increases significantly, 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 hike. To give a concrete image, suppose that we observe a decline of employment in a certain region where the minimum wage continuously rises; in this case, we would have a hard time separately attributing the observation to the current, past, or future minimum-wage increases. Thus our study does not adopt an event study approach, but we already have confirmed that the labor-market trends before 2007 are not systematically correlated with the trend increase of the minimum wage after 2007.

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 are presumably different across groups. We pick the ages between 19 and 24 as youth to avoid those who are still in high school, as almost all high-school graduates turn age 19 before graduation. Those between the ages of 25 and 59 are considered to be prime age people, and those between the ages of 60 and 64 are considered as older people. The majority of employers in Japan set a mandatory retirement age at 60, and those workers who experience the mandatory retirement change their employers or transition to temporary positions with the same employer.

Table 4 reports the estimation results for men and women by age groups. The estimation result for men ages 19-24 shows that a 10% increase of the minimum wage decreases the employment probability by 4.8 percentage points, where the average employment rate of this group is 80%; thus, the implied elasticity is 0.6, which is precisely estimated and statistically significant. This negative estimate indicates that minimum-wage hikes significantly reduce the employment of high-school-graduate youth, a group with the lowest skill. The statistically significant estimate for ages 25-59 is -0.12, implying -0.13 as the elasticity. Thus, we find a moderate negative impact of minimum-wage hikes on the employment of prime-age, less-educated men. The estimate for those who are ages 60-64 is close to zero and not statistically significant.

For women, minimum-wage hikes do not affect the employment of youth and older workers in statistically significant ways. For prime-age women ages 25-59, a minimum-wage hike even increases the employment rate in a statistically significant way; a 10% minimum-wage hike increases the employment rate by 1.3 percentage points, where the average employment rate is 65%.

As a way to check if the minimum wage hike coincided with the temporary negative shock to the local labor market, we estimate the same model using college graduate workers as the analysis sample who are less likely to be subject to minimum wage. The estimation results are reported in Table 5. None of the estimated coefficients is statistically significant. Note that the those who are in school is not included in the analysis sample and youth (19-24) in the analysis sample becomes about one-third compared to the high school graduates. Thus, one might wonder if the negative coefficient is statistically insignificant because of the inflated standard errors due to the small sample size. Without considering the clustered sampling, the standard error would be $1/\sqrt{3}$, thus the negative coefficient for young men would be statistically significant at margin, but the size of estimated coefficient is about 2/5. Overall, the analysis of employment responses among college graduates suggest that the employment results for high school graduates or less are not mere artifact caused by the temporary shock correlated with minimum wage hike.

Overall, we find that the minimum wage has a negative impact on employment among less-skilled workers with a relatively high employment rate, such as young workers and prime-age men. In contrast, we find a negligible or even slightly positive impact of minimum-wage hikes among the groups of workers with lower employment rates, such as women or older men. Using the number of observations of each group as the weight, the weighted average of the impacts is close to zero. Thus, our overall estimate is smaller than the around -0.5 found by Okudaira et al. (2019) from all workers in the manufacturing sector.

5.2 Labor-force participation

To further investigate the reason why the estimated effects on employment are different, we examine the impact of the minimum wage on labor-force participation, because a higher minimum wage could encourage labor-force participation under labor-market-search friction. Table 6, reporting the regression results for less-educated men and women, shows that the estimated negative impact on labor-force participation is slightly attenuated from the estimated negative impact on employment for young and prime-age men. For example, a 10% minimum-wage hike decreases employment by 1.2 percentage points and decreases labor-force participation by 0.9 percentage point. The gap implies that some of those who lost their jobs due to minimum-wage hikes remain in the labor force to look for another job. Thus, a minimum-wage hike increases the unemployment rate. This finding is understandable, given the strong labor-force attachment of prime-age men represented by the mean labor-force participation rate of 95%.

For women, the last three columns in Table 6 report the effect of minimum-wage hikes on the labor-force participation of less-educated women by age groups. The minimum-wage hike increases the labor-force participation of prime-age women; a 10% increase of the minimum wage increases their labor-force participation by 1.5 percentage points. This impact is larger than the impact on employment, which is 1.3 percentage points. The combination of these findings implies that minimum-wage hikes induce the labor-force participation of prime-age women, whose labor-force participation rate is traditionally low, and increases their employment, but about 14% of the increased labor force is not absorbed into employment and remains unemployed.

Regardless whether minimum-wage hikes increase or decrease employment, they increase unemployment, presumably because job losers and new entrants to the labor market have a stronger incentive to stay on labor market because of higher wages induced by higher minimum wages.

5.3 Hours worked

The Labor Force Survey asks the hours worked in the last week of the previous month for those who were in employment. Using this variable, we estimated the impact of the minimum wage on hours worked. Table 7 reports the regression result of 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. The estimation result for young men, ages 19-24, shows that a 10% increase of the minimum wage decreases the hours worked by 33%. The minimum wage does not affect the hours worked of prime-age less-educated men in a statistically significant way. Meanwhile, a 10% minimum-wage hike reduces the hours worked of older men, ages 60-64, by 1.7%, which is marginally significant. For women, the results are qualitatively similar to those of men, indicating that the minimum-wage hike does not affect the hours worked of prime-age less-educated women, but it reduces the hours worked of young and older less-educated women in statistically significant ways. Quantitatively, a 10% increase in the minimum wage decreases the work hours by 2% of young workers, ages 19-24, and by 4% of older workers, ages 60-64. Overall, minimum-wage hikes reduce the hours worked of young and older workers of both sexes. These results are consistent with results based on payroll records.

6 Robustness checks

6.1 Minimum wage effect on prefecture-specific trends

Meer and West (2015) criticized the model specification with region-specific time trends that is routinely estimated, claiming that the model fails to capture the impact of minimum wage on the region-specific time trend. In other words, they pointed out that minimum wages affect not only employment level but also employment growth. In our model notation, their claim is written as

$$Y_{ijt} = \beta \ln MW_{jt} + \theta_j + \tau_t + \xi(\ln MW_{jt}) \times Year + u_{ijt}. \tag{2}$$

In this model specification, controlling for the region-specific linear time trend, $\xi_j \times Year$, causes a over control problem because the treatment of interest affects the explanatory variable. To address this miss-specification issue, they proposed to estimate the first difference model, which regress ΔY_{ijt} on $\ln MW_{jt}$ along with time-fixed effects, but estimating the first difference model instead of fixed effects model could render different results due to the time series property of the idiosyncratic error term u_{ijt} even if the model is correctly specified (Wooldridge, 2010, 10.7). Thus, instead of pursuing the suggested avenue, we simply drop the prefecture fixed effects to check the robustness of the results to avoid the issue of over control.

The estimation results are reported in Tables 13, 14 and 15 for employment, labor force participation and natural logarithm of hours worked. The estimated coefficients only change up to 2 digits after the decimal point. From this constancy of the results based on specifications with and without prefecture-specific time trends, we would conclude that the effects of minimum wage hike on the region-specific trends in labor market outcomes are not very important in Japanese context.

6.2 Instrumental variable estimation

A part of the minimum wage hike after the 2007 revision of minimum wage act was motivated to increase the earnings of minimum wage workers so that the earnings catch up with the welfare benefit amount of a single person household. At the revision of the act, the government gives five-years grace period until the earnings of full-time minimum wage workers catch up with the amount of welfare benefit. Thus the gap between welfare benefit and minimum wage earnings in 2007 should explain the level of minimum wage between 2007 and 2012 in a way that larger the initial gap, the higher the minimum wage. Based on this institutional setting, we set up and estimate the following model by the two stage least squares:

$$Y_{ijt} = \beta \ln MW_{jt} + \theta_j + \tau_t + u_{ijt}, \tag{3}$$

$$\ln MW_{jt} = \sum_{y=2002}^{2016} \gamma_y \ln(WB/MWE)_{j2007} \times 1(Year = y) + \nu_j + \mu_t + v_{ijt}, \tag{4}$$

where Y_{ijt} is labor market outcomes of individual i living in prefecture j in year t, MW_{jt} is the minimum wage of prefecture j in year t, WB is the monthly welfare benefit amount of a single household headed household, MWE is the monthly earnings of minimum wage workers, $1(\cdot)$ is the indicator function that takes one if the statement in the parenthesis is true and zero otherwise, θ_j and ν_j are prefecture fixed effects and τ_t and μ_t are year×month fixed effects.

The gap between the amount of welfare benefit and the earnings of minimum wage workers, $\ln(WB/MWE)_{j2007}$, motivate the 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 2007 on minimum wage, we allow for different coefficients between 2002 and 2016. We would expect stable coefficient in the pre-period, 2002 and 2007, the increase in coefficients between 2008 and 2012, stable coefficient in 2013 and afterward.

This instrumental variable estimation allows for the policy endogeneity in the form of correlation between u_{ijt} and v_{ijt} that would arise if the minimum wage commissions consider the local labor market conditions in its determination of minimum wages. On the other hand, the identification of the coefficients requires the error term of the structural equation u_{ijt} and the gap in 2007, $\ln(WB/MWE)_{j2007}$, are not correlated conditional on prefecture and time fixed effects. As explained in Section 2, the level of 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 the subsequent temporary labor market shocks conditional on prefecture and year-month fixed effects.

There are several caveats apply to the instrumental variable estimation. The first is that we do not allow of the prefecture-specific linear time trends because the prefecture-specific linear time trends capture the substantial part of the increase of minimum wage in each prefecture after the 2007 revision of the minimum wage act depending on the gap of welfare benefit amount and minimum wage earnings in 2007. The analysis in the previous subsection suggests that prefecture-specific linear time trends do not play important role though. The second 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 minimum wage used is generated for the purpose of filling the gap between welfare benefit and minimum wage earnings.

Figure 9 reports the estimated coefficients, γ, and its 95% confidence intervals in Equation (4). The coefficients had hovered around 0.7 by 2007, and started to creep up from 2008 to 2013 reaching to 1.1, then again had hovered around since then. This result is consistent with the notion that the minimum wage councils had increased regional minimum wages to fill the gap between WB and MWE in 2007 from 2008 to 2012. A reason why the difference of the estimated coefficients before and after the grace period is about 0.4, instead of 1.0, is that a significant part of the initial gap is resolved by the uniform percentage increases in regional minimum wages captured by the year-month fixed effects; the gaps were larger in the prefectures where the level of minimum wages were higher, thus the same percentage increase in minimum wages means larger amount of minimum wage increase. F-statistics for the null hypothesis that all the coefficients for the interaction terms between the gap in 2007 and year dummy variables are zero is XXX, thus we do not face the weak instrumental variable problem.

Table 8 reports the IV estimates of the employment equation. The standard errors become roughly twice as large as the OLS estimates reported in Table 4. Regardless of inflated standard errors, the coefficients for younger men is -0.67 with standard error 0.26. This coefficient is slightly larger than the OLS estimate -0.48 in its absolute value, but with the large standard error of the IV estimate, the difference is within the range of sampling error. The IV estimate for prime age women is 0.31 with standard error 0.15, which is statistically significant. Again this estimate is larger than the OLS estimate, 0.13, reported in Table 4, but with the large standard error of IV estimate, we argue that the difference is within the range of sampling error. Given the opposite direction of the changes in coefficients for younger men and prime-age women, it is difficult to attribute the changes of estimated coefficients to a simple policy endogeneity that positive labor market shock is positively correlated with minimum wage hike, which results in positive biases of OLS estimates.

While not statistically significant, it is worth drawing attention to the positive IV estimates for older workers of both genders. As mentioned before, the IV estimation hinges on the variation of minimum wages between 2008 and 2012. This period overlaps with the period that the employment rate of older workers aged 60-64 increased because of the implementation of Employment Stability Act for Older Workers that requires firms to offer employment opportunity up to age 65. Kondo and Shigeoka (2017) document that the impact of the Act was larger for larger firms where the mandatory retirement at age 60 and back-loaded wage payment is more prevalent. Since the gap between WB and MWE were larger in urban prefectures in 2007 and the larger firms are concentrated in the urban prefectures, the instrumental variable and the error term in the structural equation may be positively correlated, resulting in the positive asymptotic bias in the IV estimates. Thus, IV estimates for older workers should be interpreted with a grain of salt.

Table 9 reports the IV estimates for labor force participation equation. The estimates are qualitatively similar to the IV estimates for employment equation reported in Table 8. Thus the same caveats apply here. Finally, Table 10 reports the IV estimates of ln hours equation. The estimates are similar to the OLS estimates reported in Table 7 but with inflated standard errors. Thus all the estimated coefficients lost statistical significance expect for the negative impact among younger men.

Overall, it is rather difficult to derive definitive conclusions based on the IV estimation due to the inflated standard errors because of the limited variation of minimum wage between 2008 and 2012. Regardless of imprecise estimates, the result indicates that minimum wage hike reduced the employment rate of less educated young men while it encouraged employment of less educated prime-age women. More importantly, the heterogeneity in the difference of OLS and IV estimates across demographic groups is not consistent with a potential policy endogeneity concern. In the end, in the absence of strong evidence for the biases of OLS estimates, we take OLS estimates as the preferred estimates because the OLS estimates are more precise and identified off the minimum wage variation of the whole sample period.

7 Impacts on transitions of employment status

We now examine the effect of the minimum wage on the transition between non-employment and employment over a one-year period, using the short panel data constructed by matching the second month and fourth month surveys of the Labor Force Survey, as explained in the data section. In particular, 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 minimum wages are 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 the employment status of the last week of September face the MW_{t-1} for 12 months. At the other extreme, individuals in November surveys face MW_{t-1} for 11 months and MW_t for 1 month. We define the weighted average of minimum wages faced by an individual between year $t-1$ and t as $\widetilde{MW}_{jt}$. 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:

$$E_{ijt} = \gamma_0 + \gamma_1 \ln \widetilde{MW}_{jt} + \theta_j + \tau_t + u_{ijt}, \quad \text{conditional on } E_{ijt-1} = \{0, 1\}. \tag{5}$$

We divide the sample by the employment status of the previous year 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, $\xi_j \times Year$. Instead, the prefecture fixed effects, θ_j, presumably capture the prefecture-level unobserved heterogeneity in the transition.

Table 11 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%, whereas the average transition probability is 28%. The negative impact is even larger for 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 11. Overall, this transition analysis reveals that the negative impact of the effective minimum wage on employment among prime-age men reported in Table 4 was due to the decreased transition probability from non-employment to employment.

We now attempt to relate the estimates from the stock analysis reported in Table 4 and the estimates from the flow analysis reported in Table 11. We focus our attention to the cases where the impacts are estimated precisely for both stock and flow models. The key equation to relate the stock with the flow is the transition equation:

$$E_t = P(E_t|NE_{t-1})NE_{t-1} + P(E_t|E_{t-1})E_{t-1}, \tag{6}$$

where E_t is the stock of employment in year t, NE_t is the stock of non-employment in year t, $P(E_t|NE_{t-1})$ is the transition probability from non-employment to employment from year $t-1$ to year t, $P(E_t|E_{t-1})$ is the transition probability from employment to employment from year $t-1$ to year t. We consider the effect at the steady state by dropping the time subscript and dividing the equation by $NE + E$ to obtain

$$\frac{E}{NE+E} = P(E_t|NE_{t-1})\frac{NE}{NE+E} + P(E_t|E_{t-1})\frac{E}{NE+E}. \tag{7}$$

For prime-age men, the effect of the minimum-wage hike on $P(E_t|NE_{t-1})$ is -0.21 with statistical significance and on $P(E_t|E_{t-1})$ is -0.02 without statistical significance, according to Columns 2 and 5 of Table 11; and $\frac{NE}{NE+E}$ is 0.89 and $\frac{NE}{NE+E}$ is 0.11, according to Column 2 of Table 4. Neglecting the impact that is not statistically significant, the expected equilibrium impact is -0.18, which is even larger than the estimated impact of the stock model, -0.12, reported in Column 2 of Table 4. There are several reasons for this discrepancy. First, the flow model is not precisely estimated relative to the stock model, Second, the transition of employment status may not be a simple Markov process, and thus the employment status in the previous year does not fully capture the dynamic process. The discrepancy makes the quantitative argument difficult, but qualitatively, the flow analysis for men suggests that the negative impact of the minimum-wage hike on prime-age employment is partially attributable to the reduced transition from non-employment to employment.

Table 12 reports the estimation result of the transition analysis for women. None of the estimated impacts on the transition from non-employment to employment is statistically significant, as reported in the Columns 1 - 3. In contrast, a minimum-wage hike increases $P(E_t|E_{t-1})$ of prime-age women by 0.08 in statistically significant ways. A 10% increase in the minimum wage increases the probability of staying employed by 0.8 percentage point relative to its mean 92%. Applying the formula of (7) neglecting statistically insignificant coefficient, the implied impact of a minimum-wage hike on a steady-state employment rate is about 0.05, which is smaller than the estimate of 0.13 reported in Column 5 of Table 4. Thus, again, the quantitative result of the flow analysis does not match up with the result of the stock analysis. Qualitatively, however, the results suggest that the positive impact of the minimum wage on prime-age women is partly due to its positive impact on the probability that employed women stay employed.

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. By suppressing the transition probability from non-employed to employed, minimum wage hike reduces the stock employment rate of prime-age men; On the other hand, by enhancing the probability of employed prime-age women to stay employed, it increases their stock employment rate.

8 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, aiming to fill the gap between the amount of public assistance benefits and the earnings of full-time minimum-wage workers.

The analysis of payroll records shows that the minimum wage indeed increased the probability that a worker works near the minimum wage, suggesting that minimum wages certainly formed the wage floor of low-wage workers in Japan between 2002 and 2017. The fraction of workers who work around the minimum wage, namely below 1.05 $\times$ minimum wage, however, is just around 3% of all workers. This fact implies the importance of paying particular attention to low-skilled workers, who are presumably more susceptible to minimum-wage hikes.

The estimated impacts of minimum wages on labor-market outcomes are heterogeneous across demographic groups. The increase of the minimum wage reduces the employment of young, less-educated men with an elasticity around -0.6. Similarly it reduces the employment of prime-age, less-educated men with an elasticity around -0.13. In contrast, minimum-wage hikes do not affect the employment of young and older less-educated women and less-educated older men. It even slightly increases the employment of less-educated prime-age women. Generally speaking, minimum-wage hikes reduce the employment of demographic groups whose employment rate was originally high; in contrast, they do not affect or slightly increase the employment of demographic groups whose employment rate was originally low though increasing the labor-force participation rate.

An increase of minimum wage decreases the hours worked of young and older less-educated workers of both genders. The estimated elasticity ranges from -0.17 to -0.40. In contrast, minimum-wage hikes do not affect the hours worked by prime-age workers of either gender.

Overall, the impacts of minimum-wage hikes are heterogeneous across demographic groups; One on hand, minimum-wage hikes weakly increase the employment of sub-groups through encouraging labor-force participation, while on the other hand, they significantly decrease the employment of low-skilled workers, especially among less-educated young men. Literature has found that the employment status of youth has persistent impacts on subsequent labor-market outcomes in Japan (Kondo, 2007; Genda et al., 2015). Thus, policy makers should pay particular attention to the negative impacts on young less-educated men.

As a policy mix, some policy makers suggest combining a minimum-wage hike with an investment subsidy to promote the mechanization of production processes to mitigate the impacts on the performance of small- and medium-scale firms. This investment subsidy would further decrease the employment of low-skilled workers through the substitution of labor for capital. Policy makers should instead consider conditional cash transfer programs targeting low-income earners as an income policy.

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Figure 1: Trends of prefecture minimum wage

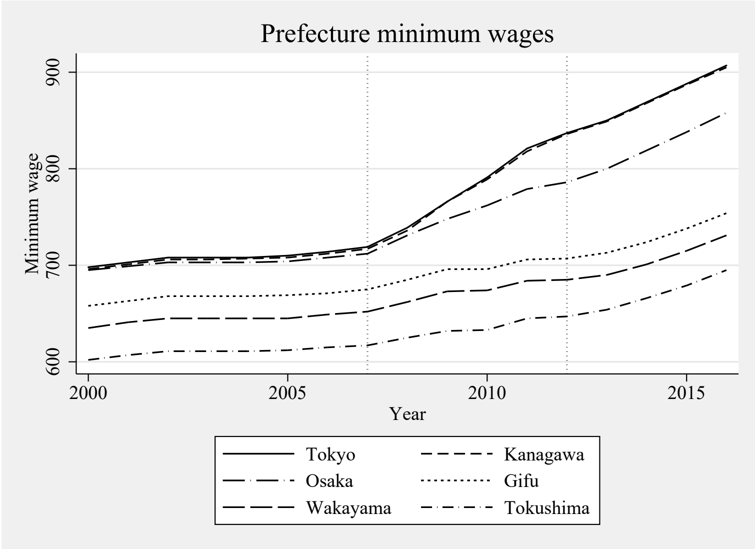


Figure 2: Gap between MW and welfare benefit in 2007 vs. MW growth

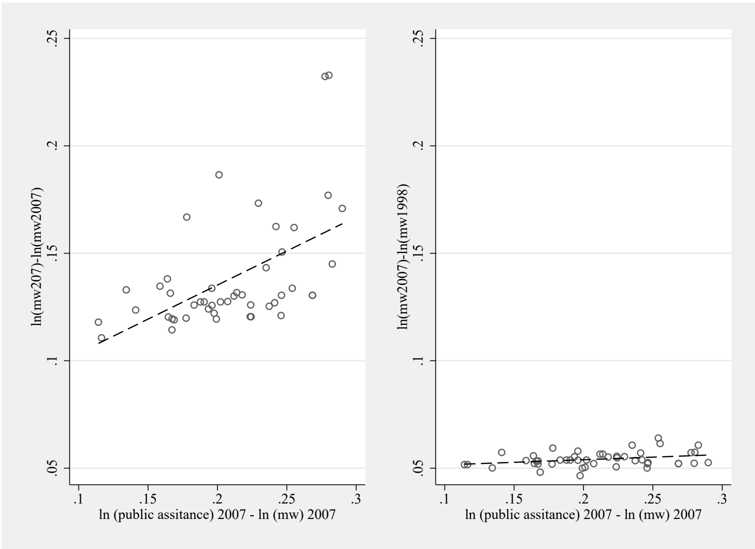


Figure 3: Gap between MW and welfare benefit in 2007 vs. employment pre-trend

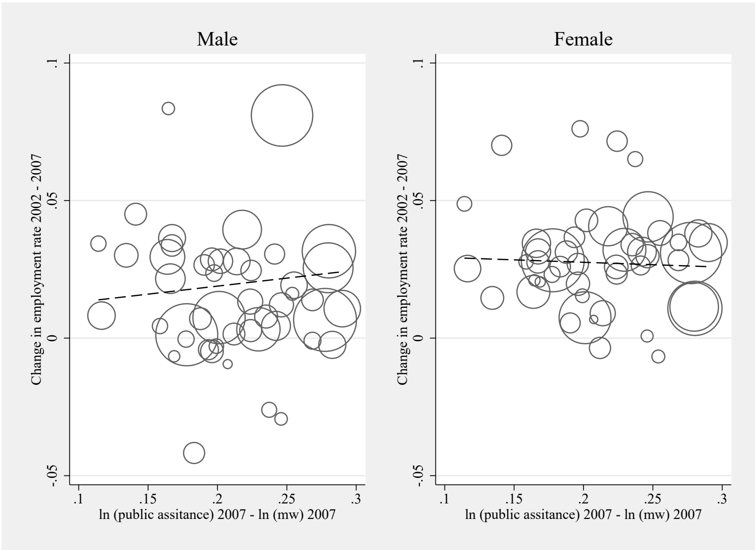


Figure 4: Distribution of hourly wages of full-time workers age 19-64

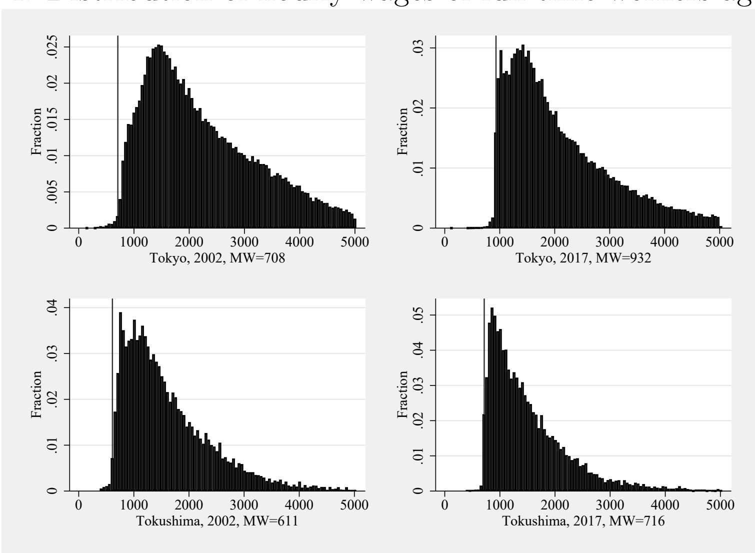


Figure 5: MW impact on bottom end of wage distribution

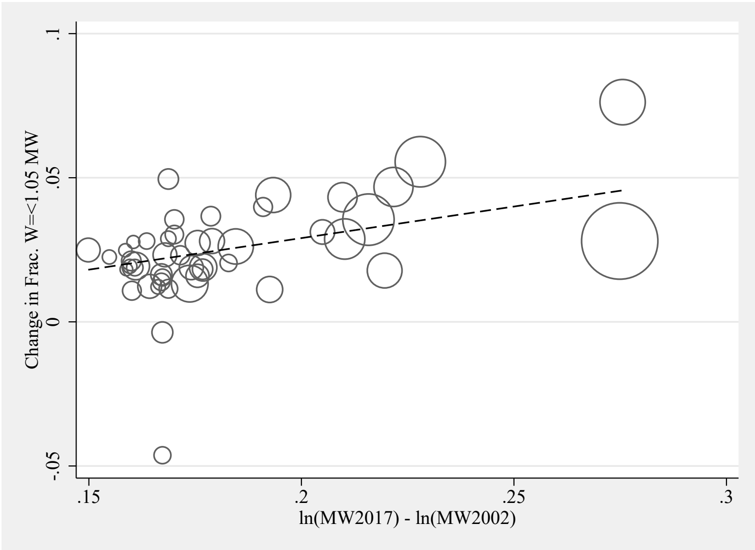


Figure 6: Fraction of workers with $W \leq 1.05 \times MW$, Age: 19-69

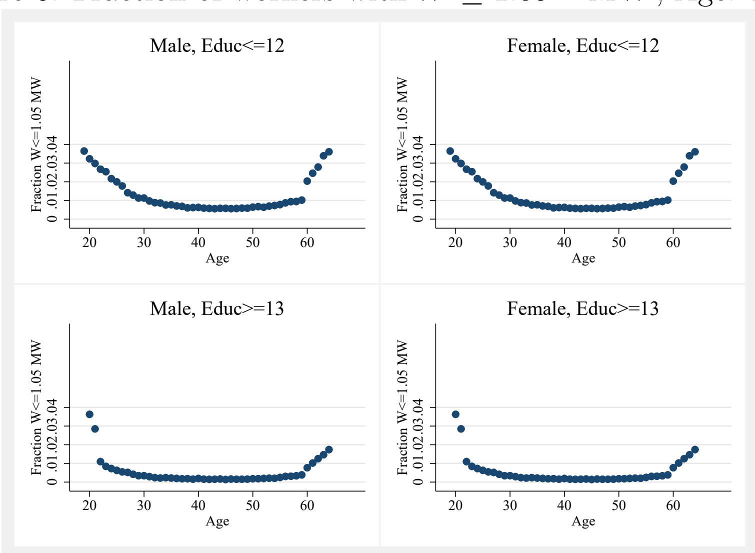


Figure 7: Change in MW 2007-2016 and change in employment 2007-2016

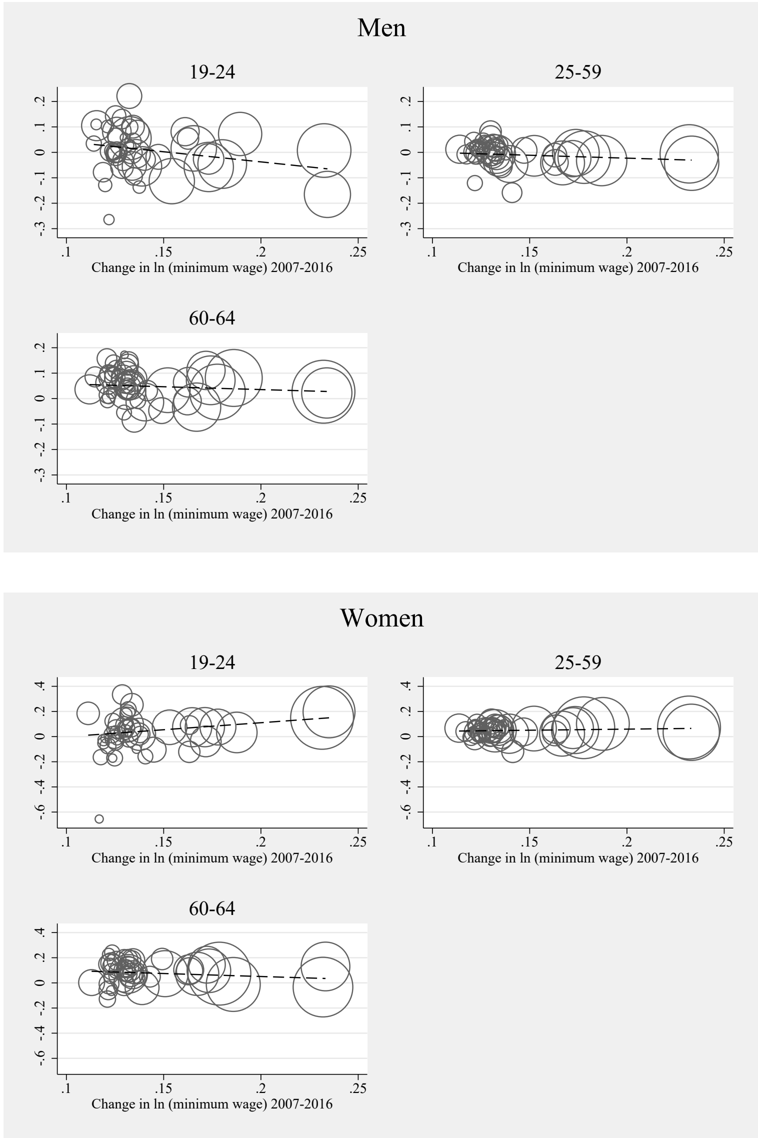


Figure 8: Change in MW 2007-2016 and change in employment 2002-2007

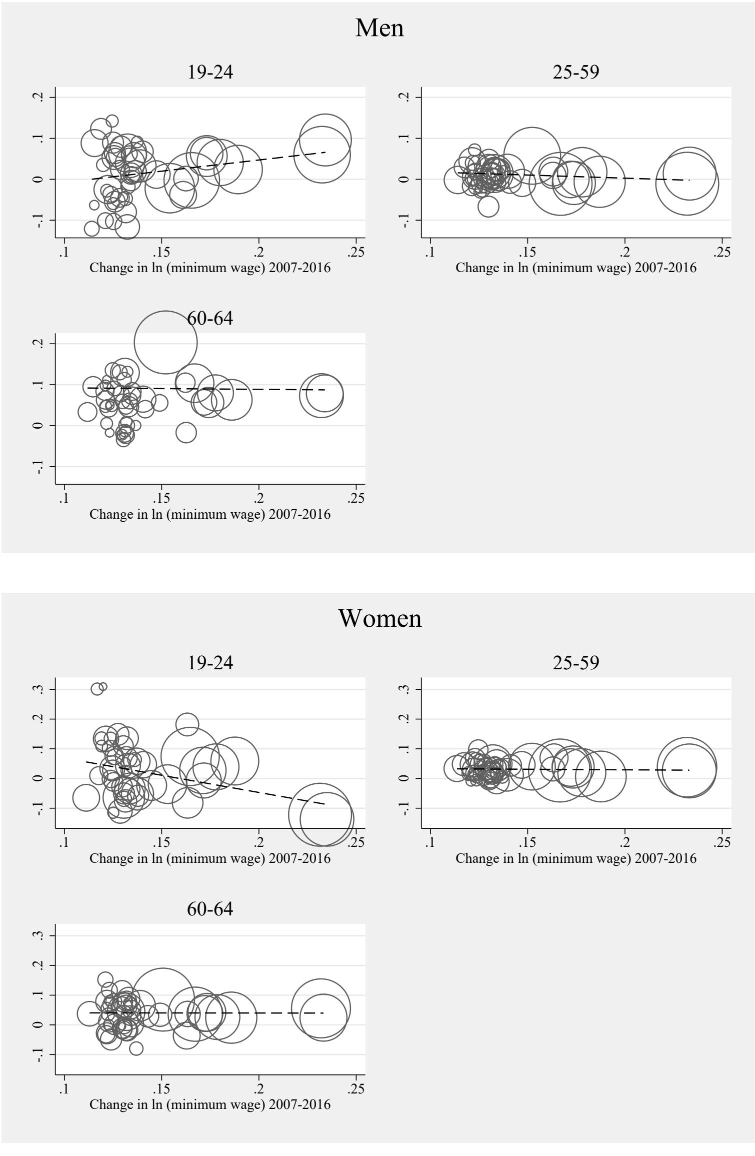
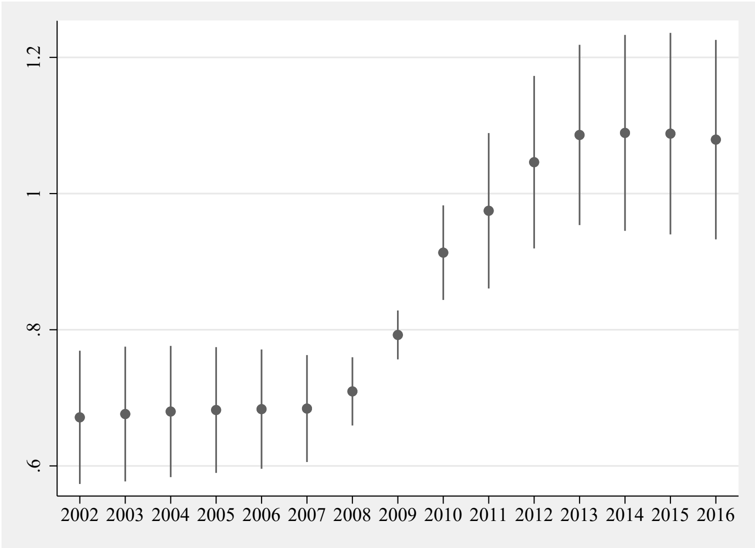


Figure 9: MW and the gap between welfare benefit and MW earnings in 2007



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

$$\ln MW_{jt} = \sum_{y=2002}^{2016} \gamma_y \ln(WB/MWE)_{j2007} \times 1(Year = y) + \nu_j + \mu_t + \upsilon_{ijt},$$

where $(WB/MWE)_{j2007}$ is the ratio of welfare benefit and MW earnings in prefecture j in 2007. Confidence intervals are calculated based on the standard errors robust against clustering within prefectures denoted by j.

Table 1: Fraction of workers wages relative to MW

	(1) Below	(2) 0-5	(3) 5-10	(4) 10-15	(5) 15-20	(6) 20-25	(7) 25-30
$\ln(MW)$	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)
Mean	0.01	0.02	0.03	0.03	0.03	0.03	0.03

Prefecture-level clustering robust standard errors in parentheses.
Weighted least squares using sampling weight.
Prefecture and year fixed effects are included.

Table 2: Impact of MW on ln hours worked based on payroll record

	(1) All	(2) M19-24	(3) M25-59	(4) M60-64	(5) F19-24	(6) F25-59	(7) F60-64
$\ln(MW)$	-0.13	-0.56	-0.12	-0.38	-0.56	-0.07	-0.26
	(0.04)	(0.07)	(0.02)	(0.07)	(0.12)	(0.04)	(0.16)
Observations	18979132	1056623	9524884	743411	974318	6191637	488259

Prefecture-level clustering robust standard errors in parentheses.
Weighted least squares using sampling weight.
Prefecture and year fixed effects are included.

Table 3: Impact of MW on ln monthly salary based on payroll record

	(1) All	(2) M19-24	(3) M25-59	(4) M60-64	(5) F19-24	(6) F25-59	(7) F60-64
$\ln(MW)$	0.02	-0.47	0.19	-0.21	-0.62	0.03	-0.13
	(0.08)	(0.08)	(0.09)	(0.18)	(0.14)	(0.09)	(0.22)
Observations	18979132	1056623	9524884	743411	974318	6191637	488259

Prefecture level clustering robust standard errors in parentheses.
Weighted least squares using sampling weight.
Prefecture and year fixed effects are included.

Table 4: Effect on employment among less educated ($Educ \leq 12$)

	(1) M19-24	(2) M25-59	(3) M60-64	(4) F19-24	(5) F25-59	(6) F60-64
$\ln(MW)$	-0.48	-0.12	0.04	-0.09	0.13	-0.01
	(0.13)	(0.03)	(0.15)	(0.09)	(0.06)	(0.11)
Mean	0.80	0.89	0.67	0.70	0.65	0.42
Observations	152536	1811170	403689	133470	1907972	482427

Prefecture level clustering robust standard errors in parentheses
Prefecture and year fixed effects, prefecture specific year trend are included

Table 5: Effect on employment among college graduates ($Educ = 16$)

	(1) M19-24	(2) M25-59	(3) M60-64	(4) F19-24	(5) F25-59	(6) F60-64
$\ln(MW)$	-0.20	-0.02	0.12	0.08	0.04	-0.37
	(0.19)	(0.02)	(0.16)	(0.16)	(0.08)	(0.26)
Mean	0.75	0.95	0.72	0.78	0.68	0.41
Observations	49677	1006358	117401	46504	469183	31770

Prefecture level clustering robust standard errors in parentheses
Prefecture and year fixed effects, prefecture specific year trend are included

Table 6: Effect on labor-force participation among less educated

	(1) M19-24	(2) M25-59	(3) M60-64	(4) F19-24	(5) F25-59	(6) F60-64
$\ln(MW)$	-0.47	-0.09	0.01	-0.02	0.15	-0.00
	(0.07)	(0.04)	(0.12)	(0.09)	(0.05)	(0.10)
Mean	0.91	0.95	0.74	0.79	0.69	0.44
Observations	152536	1811170	403689	133470	1907972	482427

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

Table 7: Effect on $\ln(hoursworked)$ among less-educated working men

	(1) M19-24	(2) M25-59	(3) M60-64	(4) F19-24	(5) F25-59	(6) F60-64
$\ln(MW)$	-0.33	-0.07	-0.17	-0.20	0.06	-0.40
	(0.15)	(0.05)	(0.10)	(0.10)	(0.06)	(0.14)
Observations	120901	1605057	268612	92640	1237290	201314

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

Table 8: Effect on employment among less educated: IV estimates

	(1) M19-24	(2) M25-59	(3) M60-64	(4) F19-24	(5) F25-59	(6) F60-64
$\ln(MW)$	-0.67	-0.04	0.33	0.04	0.31	0.41
	(0.26)	(0.08)	(0.32)	(0.17)	(0.15)	(0.31)
Mean	0.80	0.89	0.67	0.70	0.65	0.42
Observations	152536	1811170	403689	133470	1907972	482427

Prefecture level clustering robust standard errors in parentheses
Prefecture and year fixed effects are included

Table 9: Effect on labor-force participation among less educated: IV estimates

	(1) M19-24	(2) M25-59	(3) M60-64	(4) F19-24	(5) F25-59	(6) F60-64
$\ln(MW)$	-0.58	-0.11	0.22	0.10	0.30	0.45
	(0.15)	(0.07)	(0.28)	(0.16)	(0.13)	(0.31)
Mean	0.91	0.95	0.74	0.79	0.69	0.44
Observations	152536	1811170	403689	133470	1907972	482427

Prefecture-level clustering robust standard errors in parentheses.
Prefecture and year fixed effects are included

Table 10: Effect on $\ln(hoursworked)$ among working men: IV estimates

	(1) M19-24	(2) M25-59	(3) M60-64	(4) F19-24	(5) F25-59	(6) F60-64
$\ln(MW)$	-0.55	-0.08	-0.16	-0.33	0.15	-0.20
	(0.26)	(0.06)	(0.16)	(0.30)	(0.11)	(0.23)
Observations	120901	1605057	268612	92640	1237290	201314

Prefecture-level clustering robust standard errors in parentheses.
Prefecture and year fixed effects are included

Table 11: Effect on current employment among men by employment status of previous year and age

	(1) NE19-24	(2) NE25-59	(3) NE60-64	(4) E19-24	(5) E25-59	(6) E65-69
$\ln \widetilde{MW}$	-0.62	-0.21	-0.40	-0.07	-0.02	-0.08
	(0.51)	(0.10)	(0.19)	(0.15)	(0.03)	(0.10)
Mean	0.41	0.28	0.14	0.94	0.97	0.89
Observations	5786	42393	26304	20769	358222	66833

Prefecture-level clustering robust standard errors in parentheses.
Prefecture and year fixed effects are included

Table 12: Effect on current employment among men by employment status of previous year and age

	(1) NE19-24	(2) NE25-59	(3) NE60-64	(4) E19-24	(5) E25-59	(6) E65-69
$\ln \widetilde{MW}$	0.70	-0.07	-0.05	0.14	0.08	-0.21
	(0.39)	(0.05)	(0.08)	(0.24)	(0.03)	(0.08)
Mean	0.37	0.17	0.07	0.90	0.92	0.86
Observations	7133	144982	60534	16013	277137	50398

Prefecture level clustering robust standard errors in parentheses.
Prefecture and year fixed effects are included

A Specifications without prefecture-specific linear time trends

Table 13: Effects on employment among high school graduates

	(1) M19-24	(2) M25-59	(3) M60-64	(4) F19-24	(5) F25-59	(6) F60-64
$\ln(MW)$	-0.48	-0.12	0.05	-0.09	0.12	-0.01
	(0.13)	(0.03)	(0.15)	(0.09)	(0.06)	(0.11)
Mean	0.80	0.89	0.67	0.70	0.65	0.42
Observations	152536	1811170	403689	133470	1907972	482427

Prefecture level clustering robust standard errors in parentheses.
Prefecture and year fixed effects are included

Table 14: Effects on labor force participation among high school graduates

	(1) M19-24	(2) M25-59	(3) M60-64	(4) F19-24	(5) F25-59	(6) F60-64
$\ln(MW)$	-0.47	-0.08	0.01	-0.02	0.14	-0.00
	(0.07)	(0.04)	(0.12)	(0.08)	(0.05)	(0.10)
Mean	0.91	0.95	0.74	0.79	0.69	0.44
Observations	152536	1811170	403689	133470	1907972	482427

Prefecture level clustering robust standard errors in parentheses.
Prefecture and year fixed effects are included

Table 15: Effects on ln hours worked among high school graduates

	(1) M19-24	(2) M25-59	(3) M60-64	(4) F19-24	(5) F25-59	(6) F60-64
$\ln(MW)$	-0.33	-0.07	-0.17	-0.21	0.05	-0.40
	(0.15)	(0.05)	(0.10)	(0.10)	(0.06)	(0.14)
Observations	120901	1605057	268612	92640	1237290	201314

Prefecture level clustering robust standard errors in parentheses.
Prefecture and year fixed effects are included