Wage Spillovers across Sectors: Evidence from a Localized Public-Sector Wage Cut*

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Abstract

We study how institutional wage reforms in one sector spill over to other sectors by analyzing the public sector. We leverage the Japanese policy reform that cut public-sector wages only in certain municipalities and the institutional setting in which only young workers are eligible for public-sector jobs. We find that a 1% public-sector wage cut reduces the private-sector wages of young workers by 0.3%, with larger spillovers in municipalities with a larger share of public workers. It also reduces the young population by 0.4%, suggesting a welfare decline based on spatial equilibrium and a decrease in private-sector labor demand.

JEL classification: H72, J31, J38, R13, R50.

Keywords: spillover, wage reform, public sector, local labor market, Japan.

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1 Introduction

Institutional rules in a labor market can play a crucial role in determining the wage rate (Boeri and Van Ours 2021). Often, some sectors are particularly influenced by such institutional rules. For example, some sectors or employers may have specific minimum wages (Derenoncourt, Noelke, Weil and Taska 2021; Demir 2023). Another example is the requirement of a uniform wage rate across locations, which many countries adopt in sectors such as healthcare and education (Staiger, Spetz and Phibbs 2010; Willén 2021). Reforms in these institutional rules can substantially change wages in some sectors, and they may even impact sectors not directly covered through spillover effects. Such spillovers should be accounted for when evaluating these reforms.

This paper studies the spillover effects of a public-sector wage cut on private-sector wages as a prominent example of institutional reforms that could induce wage spillovers across sectors. The public sector is an important employer in most countries, and its wage determination is heavily influenced by institutional settings. Among the OECD countries, the average employment share of the public sector is 18%, and the compensation of public employees accounts for, on average, 23% of the total government expenditure (OECD 2021). This strong presence of the sector may contribute to spillovers into the private sector. Moreover, it is widely documented that the public-sector wage rate is determined differently than the private-sector wage rate due to institutional rules (Katz and Krueger 1991; Gregory and Borland 1999; Morikawa 2016). Thus, a public-sector wage reform provides an interesting case for studying how institutional reforms in one sector affect other sectors, namely, the private sector, through spillover effects. Moreover, the public-sector wage itself is an important policy tool for achieving various objectives, such as gaining votes in the US federal elections (Borjas 1984) and achieving fiscal consolidation after the Global Financial Crisis (Forni and Novta 2014). Therefore, understanding the spillover effects of public-sector wage reforms is indispensable when evaluating policies regarding such reforms.

We leverage the geographic variation in the public-sector wages to estimate how public-sector wage reforms affect private-sector wages. A key empirical challenge for identifying the effects of public-sector wages is the endogeneity concern that the public-sector wage may be determined in reference to local private-sector wages, inducing reverse causality. We overcome this endogeneity concern by exploiting a Japanese policy reform that changed the formula for determining public wages in each municipality. Since this reform was introduced to address the long-standing pay gap between private and public wages, which varied across municipalities, the updates in the public-sector wages considered the *level* of the private-sector wages in a local labor market but did not account for its contemporaneous *trend*. Therefore, the policy reform induced exogenous variation in the public-sector wages after controlling for municipality fixed effects. Moreover, we exploit the presence of treatment and control groups within each municipality to further mitigate concerns that private-sector wage and population trends may be correlated with confounding region-specific trends. Specifically, we harness the institutional setting in which one must be under 30 years of age to become a public employee in many local governments, which creates a situation where the spillover effect of public-sector wage reforms is relevant only for workers younger than 30. This empirical setting leads us to adopt a triple-difference strategy (Olden and Møen 2022), which exploits both the exogenous local variation in the public-sector wages over time and the presence of treatment and control groups within each municipality.

We estimate that a 1% public-sector wage cut reduces the private wages of young workers by 0.3%. While we use a public-sector wage cut to estimate wage spillovers, this spillover estimate is comparable to wage spillover estimates across different employers in different contexts (e.g., Staiger et al. 2010; Bassier 2022). Importantly, we find that this wage spillover effect is stronger in regions with a larger share of public-sector workers. This confirms our intuition that there should be a greater impact on public-sector wage reform when there are more public-sector workers. This result supports the idea that our empirical strategy is likely to capture the effects of the public-sector wages rather than confounding factors.

We then analyze how the young population responds to the public-sector wage cut. We find that a 1% reduction in public-sector wages decreases the young population by 0.4%, and this negative effect is greater in municipalities with a larger share of public-sector workers. In addition to the importance of the young population itself as a local economic variable, there are two theoretical motivations for this analysis. First, it is suggestive of the welfare effect of the public wage cut. Intuitively, if the local public-sector wage increase improves workers' welfare in the affected area relative to other areas, then it should induce migration inflow. Therefore, the decline in population suggests that the public wage cut indeed harmed young workers' welfare, inclusive of its potential impacts on various aspects, such as the quality of public goods and job amenities. Second, given that population closely approximates the amount of workforce, it helps us distinguish whether the wage spillover is driven primarily by the increase in the private-sector labor supply to the private sector or the decrease in the private-sector labor demand. The decrease in population implies the decrease in the private-sector labor demand.

We also conduct several supplementary analyses. First, we estimate the effects of public-sector wages on youth unemployment rates and land prices. We find suggestive evidence that the unemployment rate increases and land prices decrease in response to public-sector wage cuts. The decrease in land prices is consistent with our spatial economic framework. Second, we explore the heterogeneous effects on private-sector wages with respect to gender, education attainment, industry, and firm size. We find evidence that the effects are greater for noncollege-educated workers. We also find that the spillover is greater in industries that receive young workers from the public sector. In contrast, we find little heterogeneity in terms of gender and firm size. Third, we investigate the sensitivity of our results to the largest cities and find that our results are robust by dropping them. Fourth, by using a spatial econometric model (Halleck Vega and Elhorst 2015), we find that our results are robust to considering the possibility that the public-sector wages of neighboring municipalities might also affect local labor market outcomes. Finally, we illustrate the aggregate impact of the policy reform by calculating the national-level impacts of the 2006-2010 public-sector wage cut based on our estimates, highlighting the importance of considering the spillover effects.

By analyzing the spillover effect of the public-sector wage cut on private wages, we provide novel evidence that institutional reforms in one sector can have sizable spillover effects on sectors not directly covered by these reforms. As a broad implication, our results highlight the importance of considering spillover effects when evaluating reforms in institutional wage rules, including minimum wages, anti-union laws, and equal-pay requirements across different geographical areas. More directly, our results shed light on how public-sector wages should be set. For example, public-sector wage cuts that are often adopted to achieve fiscal consolidation may significantly lower private-sector wages and reduce workers' welfare. This spillover effect should be counted as a cost of such an austerity measure. Our results are also suggestive of a macroeconomic policy to combat wage stagnation: Raising the public-sector wages may help resolve wage stagnation through spillover effects. In our context, the Japanese economy over the last 30 years has been characterized by the stagnation of both nominal and real wages (Ito and Hoshi 2020). Increasing the public-sector wages has been suggested to be an effective policy measure to combat this wage stagnation (e.g., Bernanke 2017), but its effectiveness has not been empirically investigated.[1] Our findings rationalize this proposal by providing causal evidence that a public-sector wage cut restrains private wages. Moreover, our results imply that welfare may improve for workers, at least for young workers.

This paper relates to four strands of literature. First, our paper contributes to a small but growing body of literature on wage spillover effects across sectors or employers. For instance, Staiger et al. (2010) and Willén (2021) examine the spillover effects of wage changes caused by the abolition of a nationally uniform wage schedule.[2] Derenoncourt et al. (2021) and Demir (2023) analyze the spillover effect of minimum wages that apply only to a subset of employers. Bassier (2022) studies the spillover effects of collective wage bargaining on sectors not covered by these bargaining rules. In addition to its own importance, the public sector is an interesting setting for investigating such spillovers because of its large employment share and the crucial role of institutional rules in determining wages. We contribute to this literature by providing novel evidence that public-sector

[1]Despite the lack of formal evidence, this possibility has also been noted by some politicians. For example, Goshi Hosono, a member of the House of Representatives and a former Minister of the Environment, stated that "To achieve a wage increase, especially in a rural area, we should increase the public-sector wages to promote the private-sector pay raise" (https://twitter.com/hosono_54/status/1585775663314702336, last accessed on April 15, 2024. The original quote in Japanese has been translated by the authors).

[2]Berger, Herkenhoff and Mongey (2022) shows that the wage spillover result of Staiger et al. (2010) is consistent with their oligopsonistic labor market model.

wage cuts reduce private-sector wages and population in the local labor market.[3] Interestingly, despite our focus on the wage cut and numerous other contextual differences, our preferred estimate of private wage elasticity with respect to public-sector wages is 0.34, which is close to several estimates in this literature (e.g., Staiger et al. 2010; Bassier 2022).

Second, our paper contributes to reduced-form studies on the relationship between public and private wages. The majority of studies have documented a positive relationship between public-sector and private-sector wages (e.g., Ehrenberg and Goldstein 1975; Lacroix and Dussault 1984; Gregory and Borland 1999; Lamo, Pérez and Schuknecht 2012; Afonso and Gomes 2014; Abdallah, Coady and Fah Jirasavetakul 2023).[4] While these correlations are suggestive, most studies in this literature do not address the endogeneity of public-sector wages, presumably because their exogenous variation is rare. An important exception is Telegdy (2018), who uses a national wage reform in Hungary that uniformly raised the public-sector wages across locations. Consistent with our findings, he finds that private-sector workers who were more strongly hit by the public wage shock, especially young workers, experienced a wage increase. Our contribution is to exploit the novel exogenous variation in the local public wage itself and causally estimate the elasticity of private wages with respect to public wages. Moreover, our local variation in the public wage allows us to estimate the causal effects on population and conduct a welfare evaluation of the policy reform based on a spatial economic framework.

Third, our study also relates to studies analyzing the impact of reforms in public wages and employment by fully specifying and calibrating an economic model of the labor market (e.g., Burdett 2012; Gomes 2015, 2018; Bradley, Postel-Vinay and Turon 2017; Bermperoglou, Pappa and Vella 2017; Albrecht, Robayo-Abril and Vroman 2019; Chang, Lin, Traum and Yang 2021; Lu and Kameda 2024). Since we exploit a quasi-experiment, we can estimate the causal impact of public wages without committing to a particular labor market structure. Moreover, our spatial equilibrium framework allows us to infer the qualitative welfare implications of public wage reform from population responses. Therefore, by estimating the effect on the population, we can analyze welfare implications without fully specifying the labor market structure.

While we focus on the spillover effects of public-sector wages, a growing number of studies have examined the impact of the public-sector *employment* in a local labor market, such as the increase in the private-sector jobs in response to an increase in the public-sector jobs. (Faggio and Overman 2014; Zou 2018; Faggio 2019; Auricchio, Ciani, Dalmazzo and de Blasio 2020; Jofre-Monseny, Silva and Vázquez-Grenno 2020; Becker, Heblich and Sturm 2021; Guillouzouic, Henry and Monras 2024; Lee, Ko and Kim 2024; Chirakijja forthcoming; Franklin, Imbert, Abebe and

[3] The spillover of the public-sector wage cut on private-sector wages and the population is consistent with the monopsony power of the public sector as an employer, which is sensible given the significant employment share of the public sector and consistent with experimental evidence by Dal Bó, Finan and Rossi (2013).

[4] However, some studies do not find a positive relationship (e.g., Auld, Christofides, Swidinsky and Wilton 1980).

Mejia-Mantilla 2024).[5] Although this burgeoning literature exemplifies substantial academic and policy interest in the public sector, these papers do not analyze the spillovers of public-sector *wages*. In stark contrast, we analyze the spillovers of public-sector *wages* on private-sector wages using a local labor market approach, which we achieve by leveraging a distinctive Japanese public-sector wage cut and institutional setting.

This paper is organized as follows. Section 2 explains the relevant institutional background. Section 3 describes the data. Section 4 presents our main results on the spillover effects on private-sector wages. Section 5 analyzes the response of the young population to public-sector wages and discusses the welfare implication. Section 6 presents additional results. Section 7 concludes.

2 Empirical setting

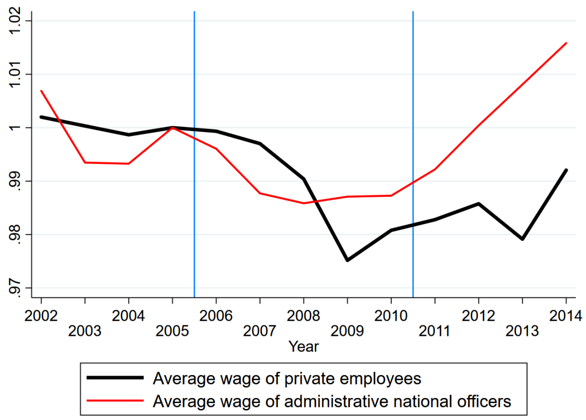
This section describes the institutional contexts relevant for our analysis. We start with general background information about the Japanese public sector. We then describe the public-sector wage reform during 2006-2010, which cut public-sector wages in some municipalities but not in other municipalities. Finally, we describe the situation in which the private sector and the public sector compete with each other for young workers due to institutional settings, while such competition essentially does not exist for older workers. These institutional contexts motivate our empirical strategy in Section 4.

Background information on the Japanese public sector We provide relevant background information about the Japanese public sector. The total number of public officials in Japan was approximately 3.5 million in 2010, constituting approximately 6% of the total employment, and the compensation of public employees constitutes 14% of the total government expenditure. Both are among the lowest in OECD countries (the OECD averages are 18% and 23%, respectively; OECD 2021). Therefore, finding a significant impact of public-sector wages in the Japanese context would also suggest a significant impact in other countries in which the public sector is more sizable. Among the public officials, 2.85 million, or more than 80%, are local public officials. The remaining 20% are national public officials.

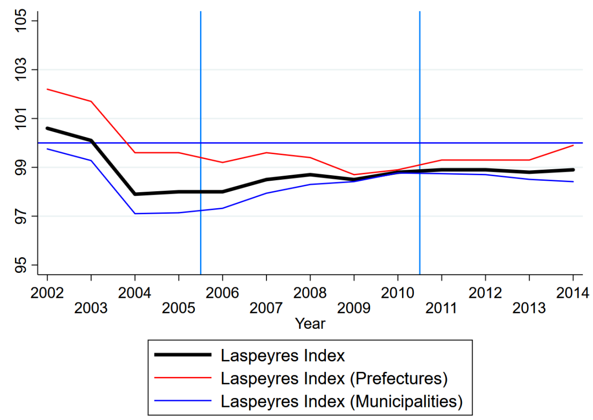
Salary levels for national public-sector workers are balanced with those of private-sector employees by the National Personnel Authority, the central administrative agency in charge of human resource management in public administration. In Japan, public wages are not determined through negotiations between employers and unions because collective bargaining rights and the right to

[5]More generally, a growing number of studies adopt the local labor market approach that exploits a cross-sectional variation to study policy impacts (e.g., Rosen 1979; Roback 1982; Kline and Moretti 2014; Suárez Serrato and Zidar 2016; Nakamura and Steinsson 2018; Monras 2019; Yamagishi 2021).

Figure 1: Public-sector wage trends in Japan



(a) National average wages relative to the 2005 basis



(b) The average Laspeyres index of local governments

Source: Basic Survey on Wage Structure and White Paper on Public Employees. Note: Figure 1a shows the national average wages of private employees and national officers relative to the baseline in 2005, which are shown by black and red lines, respectively. Following the definition of wages in the Basic Survey on Wage Structure, the wage here includes the base salary and all allowances except overtime pay and bonus. Figure 1b shows the trends of the Laspeyres index. The Laspeyres index shows the previous year's wage level of each local government relative to the central government. The index represents the base salary level of local government employees when the base salary level of national employees is set at 100. To eliminate the influence of academic background, years of experience, and the composition of staff in local governments, the index is calculated by applying the average salary of local officials by education and years of experience to the national staff composition.

strike are severely restricted for most public workers (Shimoi 2017). Figure 1a shows the national average private wages and the wages of national public officials. While both wages tend to exhibit a similar trend, the wage levels of national public-sector workers are not necessarily equal to those of private-sector workers. This implies that in addition to labor market conditions, the administration's discretionary policies affect the average wages of national public-sector workers.[6] Later in this section, we introduce a policy reform that suddenly changed the public-sector wages irrespective of contemporaneous trends in private wages.

According to the Local Public Service Act, the wage level of local public-sector workers must be balanced with the wage level of national public-sector workers, local public-sector workers in other jurisdictions, and local private-sector workers. This rule is called the "equal pay principle", and following this rule implies that the wage level of local public-sector workers depends on the wage level of local private-sector workers. However, as shown by Kawasaki and Nagashima (2007), Aoki

[6] For example, to raise funds for reconstruction following the Great East Japan Earthquake, wages for national public servants were reduced by an average of approximately 7.8% in 2012 and 2013, and some local governments followed this measure. Since this measure is temporal and our results are robust even after excluding data after 2012, we ignore this measure in the figures and analyses of this paper.

(2021), and Marumi (2023), local governments actually focus on the wage gap between the local and national governments because of the strict guidance of the national government to minimize this gap. As an example of such guidance, the national government annually publicizes the "Laspeyres index", which shows the previous year's salary level for each local government relative to the central government. Local governments with relatively high wages are frequently reported in newspapers and magazines by referring to the index (Morikawa, 2016), pressuring them to lower local public-sector wages. Figure 1b shows the Laspeyres index of base salaries for local governments, in total and separately for prefectural and municipal governments. Figure 1b shows that throughout our sample period, the Laspeyres index is stable near 100 and ranges from 97 to 102, implying that the gap between the wages of nationally employed and locally employed public officials is only 3%, at most.

Regardless of their job duties, national government employees receive regional allowances of a certain percentage of their base salary depending on their place of residence. The regional allowance is a place-based wage premium adjusted for the price level in each region, which is independent of individual characteristics, such as age, education, and job. Specifically, the effective wage rate of individual i in municipality j in year t is

$$\text{Effective wage}_{ijt} = (1 + \text{Regional allowances rate}_{jt}) \times \text{Base wage}_{it}. \tag{1}$$

As such, regional allowances act as a multiplicative wage premium of the nationally uniform base wage that depends on individual characteristics.

The regional allowance rate for national government employees also dictates regional allowance rates for local officials. Although local governments have discretion over the level of their regional allowances, they are effectively required to closely follow the national level as they do for the base salary (Aoki 2021).[7] Therefore, the national regional allowance rate in municipality j represents the overall public-sector wage level in this municipality.

Note that the Japanese public-sector wage system is not an outlier in the sense that similar systems can be found in other countries. For example, France has a regional allowances (indemnité de résidence) system, where the local civil service pay is expected to follow the national one.[8] However, the Japanese system is distinctive in that the variation in regional allowances is large (0-20%), while variation in the French system ranges from 0% to 3% of the base salary (Ministry of Public Transformation and Service of France, 2023). The large variation in regional allowances in Japan

[7] The wage gap between national and local governments, accounting for regional allowances, is also stable and is published annually as the "modified Laspeyres index".

[8] While in some countries, either trade unions are involved in the salary determination process or the federal system dictates that local public officials be paid differently in different regions, Japan and France are similar in that trade unions have a relatively limited role in salary determination and local governments are expected to follow the central government salary determination process.

is helpful for empirically identifying the effects of public-sector wages.

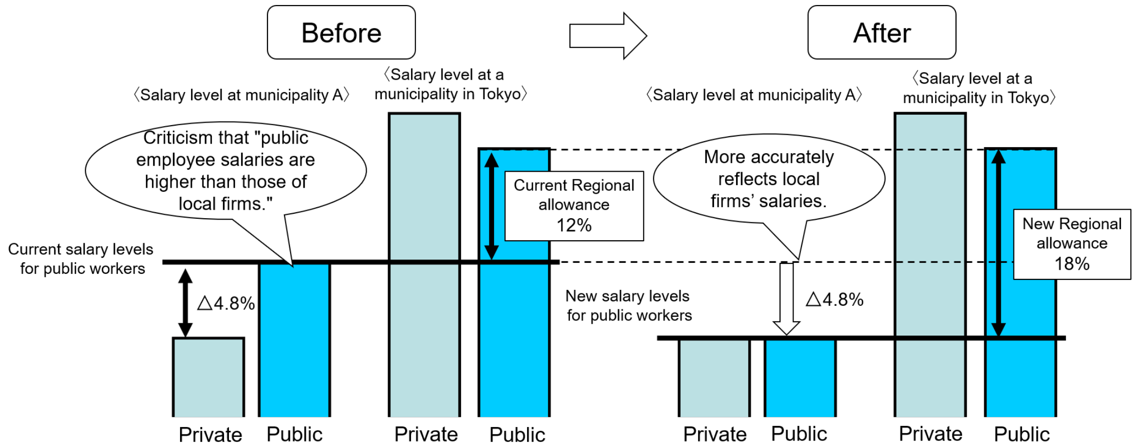
Public-sector wage reform during 2006-2010. During 2006-2010, the Japanese government reformed national public-sector wages by reducing the base wage and introducing a new regional allowance schedule with greater regional variation. The reform is illustrated in Figure 2. The main goal of the reform was to improve the fiscal balance by reducing the wage rate of national public-sector workers in nonurban areas, which was criticized as being too high relative to private wages. This was achieved by combining the following two reforms. First, the base salary, which was nationally uniform, gradually decreased each year over five years. The total wage cut was 4.8% of the original base wage. This reform reduced the public wage of nonurban areas ("municipality A" in Figure 2), but it also reduced the public wage rate in urban areas ("a municipality in Tokyo" in Figure 2), which seems undesirable as the public wage becomes lower than the private wage. The second reform corrected this by introducing new regional allowance rates. Although the old regional allowances were paid in limited locations and in limited amounts, the new regional allowances expanded the coverage and amounts. As a result, urban areas, such as Tokyo, did not experience a public wage decrease, and some even experienced a wage increase due to increased regional allowances despite the reduction in the base wage.

We make three further remarks about the implementation of the policy reform. First, note that the new regional allowance system increased the wage premium in both municipalities that were already covered in the older regional allowance system (e.g., "a municipality in Tokyo" in Figure 2) and newly covered municipalities that did not receive the wage premium in the older system. Second, the new regional allowance system was introduced in 2006, after which the regional allowance rates gradually increased over five years until the completion of the policy reform in 2010.[9] Finally, the amount of the new regional allowances was calculated based on each municipality's 10-year average wage index, which means that the regional allowances do not account for contemporaneous trends in the labor market.

These new regional allowances created a larger cross-sectional variation in public-sector wages at the municipality level, which we exploit to identify their effects on local economic outcomes. Figure 3 shows the impacts of the policy reform on the effective wages of public workers in municipalities in Japan as a whole and in the Kanto region of Japan. Panel 3a clarifies that the increase in effective wages due to increased regional allowance rates was concentrated in urban areas, most notably in the three largest metropolitan areas (Tokyo, Osaka, and Nagoya). However, Panel 3b clarifies that there was substantial variation in the change in the effective wage across municipalities within a smaller region or a metropolitan area. In particular, while urban municipalities tended

[9]For instance, Tokyo experienced a six percentage-point increase in the rate of regional allowances, with an average annual increase of 1.2 percentage points from 2005 to 2010.

Figure 2: Graphical illustration of the national public-sector wage reform

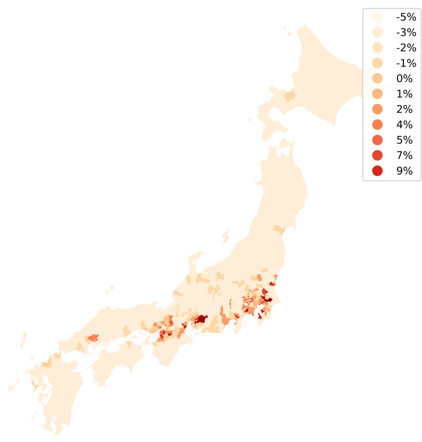


Note: This figure illustrates the national public-sector wage reform, which reduced the base salary and introduced a new regional allowances system. The figure is taken from https://www.mext.go.jp/b_menu/shingi/chukyo/chukyo3/041/siryo/__icsFiles/afieldfile/2019/05/15/1416840_001.pdf (in Japanese, last accessed on June 5, 2023), where comments in the figure are partly modified and translated into English by the authors.

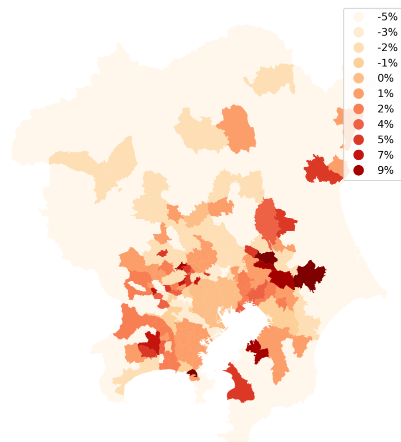
to experience an increase in regional allowances, most urban municipalities did not experience the largest increase, and the increase in the effective wage was limited (e.g., Tokyo's 23 wards experienced approximately 0% change in the effective wage). This is because our policy variation has both intensive and extensive margins, providing a large increase in regional allowance rates for municipalities that were newly eligible for regional allowances. Taken together, the results show that the base salary of the public sector was reduced in most of Japan due to the reform, while some areas even experienced an increase in the public-sector wage due to the new regional allowance schedule. Thus, the new regional allowance policy exogenously created greater variation in the public-sector wage across local labor markets.

As shown in Kawasaki and Nagashima (2007), Aoki (2021), and Marumi (2023), the wage rates of local public officials closely track those of national public officials working in the same area through administrative guidance by the national government, although local governments can, in principle, determine the wage rates of their public officials. Therefore, although the policy reform during 2006-2010 concerned national public officials, it accompanied an almost parallel change in the wages of the local public sector. Consequently, what we identify as the effect of national public-sector wages in a given municipality can be approximately interpreted as the effects of overall public-sector wages in the municipality, including both locally employed and nationally employed public officials.

Figure 3: Changes in the effective wage



(a) Japan



(b) Kanto area

Note: These figures show the changes in the effective wage level before and after the introduction of the new regional allowances in Japan (panel 3a) and the Kanto region (panel 3b). The Kanto region is one of the regions of Japan centered on the Tokyo metropolitan area. The darker red areas indicate a higher percentage change in the level of regional allowances payments. In both panels, some small islands (e.g., Okinawa) are omitted for visibility. From the equation (1), the effective wage in 2010 can be shown as $(1 + RA_{i,2010}) \times \text{Base wage}_{i,2010}$, where RA is the regional allowances rate. The change in effective wage can be calculated as $(1 + RA_{i,2010}) \times \text{Base wage}_{i,2010} - (1 + RA_{i,2005}) \times \text{Base wage}_{i,2005} = (\text{Base wage}_{i,2010} - \text{Base wage}_{i,2005})(1 + RA_{i,2010}) + (RA_{i,2010} - RA_{i,2005}) \times \text{Base wage}_{i,2005} = -0.048 \times (1 + RA_{i,2010}) + (RA_{i,2010} - RA_{i,2005})$, where we normalize $\text{Base Wage}_{i,2005}$ to 1 in the last equation. The wage changes shown in the figures are based on this calculation, where we round the values to the nearest decimal place.

Workers' age and public-sector jobs in the labor market. In addition to examining the policy reform described above, we harness the Japanese institutional setting in which only young workers are primarily affected by public-sector jobs in the labor market. There are two key institutional features behind this situation.[10] First, the lifelong employment system is conventional in the Japanese labor market, which substantially limits labor mobility for older workers (e.g., Genda, Kondo and Ohta 2010; Ito and Hoshi 2020). This implies that it is unusual for older workers to leave their current job. Consistent with this, the turnover rate for local government employees in their 20s is approximately twice as high as the turnover rate for other age groups, as shown in Figure A.1. Second, most public-sector jobs mandate a recruitment exam for screening, and an upper age limit is typically imposed for screening.[11] For national public officials, the upper age limit is usually set at 30.[12] There are similar upper age limits for exams for local public officials.[13] Overall, in our Japanese setting, both entering and leaving public-sector jobs are restricted among workers aged 30 and above.

In Figure 4, we present the private wage time series for each 10-year age category, separately for municipalities that did or did not experience the public-sector wage cut. Prior to the reform in 2006, workers in all age groups faced a similar wage trend in all municipalities. However, after the public-sector wage cut in 2006, the wage trend of young workers in municipalities that experienced the public wage cut fell below that in municipalities that did not experience the public-sector wage cut. In contrast, no such divergence is observed for workers older than 30. These wage trends in Figure 4 are consistent with our assumption that only workers younger than 30 were affected by the public-sector wage cut.

Overall, both the institutional setting and the empirical evidence suggest that primarily young workers were affected by public-sector wages, while little effect is expected for older workers. In particular, only young workers are directly influenced by the decrease in the public-sector wages because only young workers can effectively take up a public-sector job. Moreover, as private-sector firms employing young workers engage in competition with the public sector, their wage policies are also expected to be affected by public-sector wages. In contrast, older workers cannot directly enjoy the public-sector wage increase by obtaining public-sector jobs, and private firms employing them will not change their wage policies because they do not compete with the public sector in the

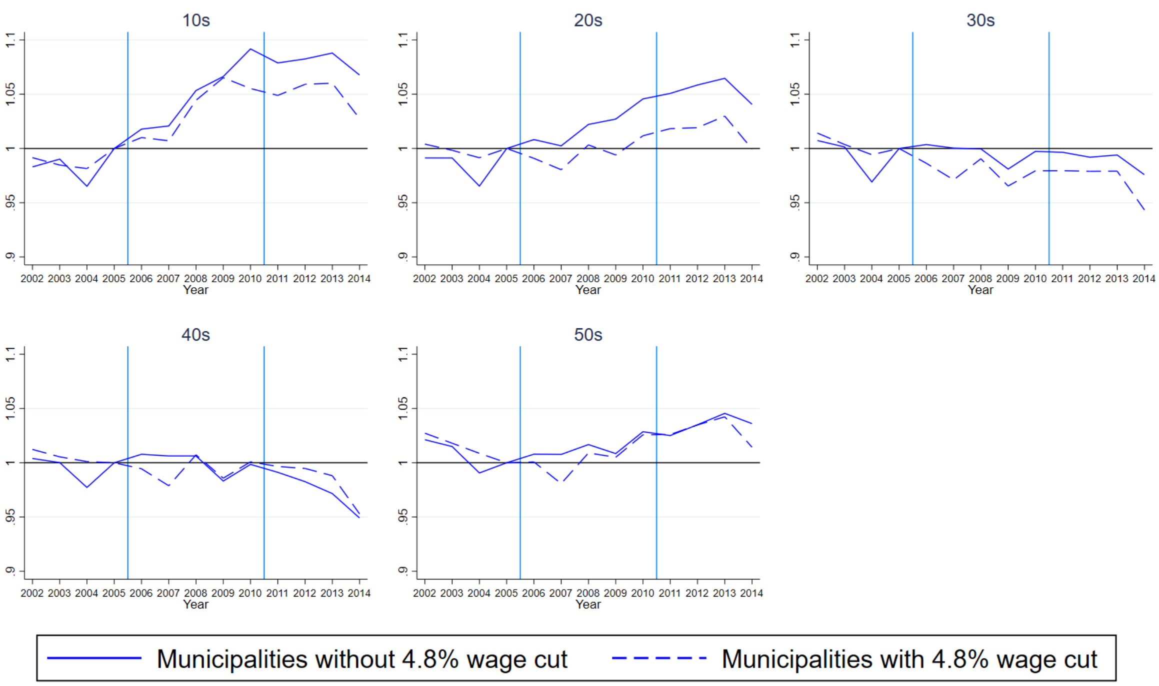
[10]In addition to the Japanese institutional features, two general reasons support the expectation that the impact is concentrated among young workers. First, Telegdy (2018) finds that public-sector wage reform in Hungary affected primarily young workers. Second, there may be information asymmetry in workers' productivity. In particular, firms may easily observe older worker's productivity from past task performance, but the productivity of young workers is harder to observe, and firms may use the public-sector wages as a yardstick to determine their wages.

[11]There is exceptional mid-career recruitment that is free of the upper age limit, but it constitutes only a negligible fraction of total public-sector employment.

[12]For example, see https://jp.stanby.com/magazine/entry/220961 (in Japanese, last accessed on May 13, 2023).

[13]The upper age limit of 30 is most prevalent in the local public sector, but some local governments are slightly more tolerant than the national government (https://90r.jp/nenrei.html, last accessed on March 15, 2024. In Japanese).

Figure 4: The trend of average wages by region following the 4.8% wage reduction by age groups



Note: Each panel shows average private wages by whether the region experienced the 4.8% reduction in the public-sector wage for each age group. The solid and dashed lines show, respectively, the average private wage in areas where the effective wage in the public sector was reduced by 4.8% and in areas where it was not reduced. To control for municipality-specific trends, all wages are shown relative to those of workers in their 60s. For each wage time series, we normalize the 2005 wage rates to one.

labor market. In this paper, motivated by the upper age limit of recruitment exams in the public sector, we define young workers as workers younger than 30 and the rest as older workers.

3 Data

We combine various datasets to examine the effect of the reform of national officials' regional allowances from 2006 to 2010. As outcome variables, we first analyze the wage rate of the private sector and then local population. We investigate unemployment rates and land prices as additional analyses. We construct two datasets: one is based on individual private workers' data for the analysis of the private sector's wage levels, and the other is based on municipality-level data for the analysis of population, unemployment rates, and land prices. Our data cover the period from 2002 to 2014, covering four years before and after the policy reform period (2006-2010).

We use the definition of municipalities as of 2015 throughout our sample period.[14] Because of the large-scale mergers of municipalities in Japan in the 2000s, we need to suitably aggregate merged municipalities when constructing the municipal-level variables. We follow Kondo (2023) to identify merged municipalities. In constructing municipality-level variables, we take the sum of merged municipalities for aggregate values (e.g., population), and we take a weighted average for per capita values (e.g., the tax income per capita).[15]

Public-sector wages. The policy variable we focus on is the regional allowance rate for national public workers.[16] Since the regional allowance payment level is stipulated in rule 9-49 of the National Personnel Authority for the year in question, the payment level for each year was obtained using D1-law.com, a database of Japanese laws and rules. The revised regional allowance levels are also included in the Fact-finding Survey on Compensation of Local Government Employees, which was used in compiling the data.

We also consider the number of local officials and their salary levels as potential control variables. We obtained these data from the Fact-finding Survey on the Compensation of Local Government Employees. The local officials' average salary is available based on education and years of experience. The average salary in our definition contains a base salary and allowances (except the regional allowances) but not overtime pay and bonuses since overtime pay and bonuses depend heavily on firm-specific idiosyncratic shocks.[17]

[14]The last municipal merger was in 2014, so we use the 2015 criteria.

[15]For example, the tax income per capita is calculated by taking the weighted average of the tax income per capita of merged municipalities, where we use the premerger population of each municipality as weights.

[16]In studying the effect of national public-sector wages, we do not need data on the base wage of national public officials because the year fixed effects in our regression models absorb changes in the nationally uniform base wage over time (see also equation 1 and footnote 24).

[17]We exclude the regional allowance from the local officials' salaries because it is determined with close reference to

Private-sector wages. We use individual-level microdata from the Basic Survey on Wage Structure. The target population for this survey is randomly selected employees from randomly selected establishments from regional and industry strata, and we weight observations by sampling weights included in the data to ensure representativeness. The survey takes the data of different employees every year and records the location, scale, and industry of each employee's workplace, as well as sampling rates and information regarding each employee, such as wage, gender, age, and education. While the individual earnings information includes the base salary, allowances, overtime pay, and bonuses, we construct the salary as the base salary and the allowances since the overtime pay and bonuses are substantially influenced by idiosyncratic shocks at the firm level.[18] Moreover, we focus on full-time workers aged between 15 and 64 years in our analysis since the Basic Survey on Wage Structure does not contain temporal workers' data before 2005.[19] For the analysis, we use the wage as a dependent variable, which is calculated by dividing the salary by the scheduled working hours. The summary statistics of these data are shown in Table C.1.

Population. We take each municipality's demographic information from the population, demographic and household surveys based on the Basic Resident Registration System. This allows us to observe the municipal population annually, separately for five-year age groups. For the population analysis, municipal data are needed, and we obtain balanced panel data for 1731 municipalities.[20] The summary statistics of this dataset are shown in Table C.2.

Unemployment rates. The unemployment rate is taken from the Population Census. The unemployment rate is based on self-reports and is defined as the number of people searching for a new job divided by the working population. Since census data are available only every five years (2000, 2005, 2010, 2015), we use linear interpolation to obtain unemployment rates in intermediate years. As in the population analysis, 1731 municipalities are included in the sample.

Land prices. We take the land price information from the Land Market Value Publication (*kouji chika*). Our municipal land price data are based on the changes in repeated appraisal prices for the

the regional allowances rate of national public officials (see Section 2). Given that the prescribed working time in local governments is 38 hours and 45 minutes, the salary level divided by this number of hours equals the wage.

[18]The salary here is defined as the amount of cash paid in June, at the time of the Basic Survey on Wage Structure, according to the payment conditions and calculation method predetermined by the labor contract, collective labor agreement, or employment regulations of the business office, including base salary and allowances but excluding overtime pay and bonuses. It is not take-home pay but rather the amount before the deduction of income tax, social insurance premiums, etc.

[19]This is likely a modest limitation for our purpose because part-time jobs in the Japanese public sector were relatively rare in the data period (Goto 2021).

[20]While Japan had 1741 municipalities as of 2015, ten municipalities with missing data due to the Great East Japan Earthquake are omitted. These municipalities are also omitted from the analysis of the unemployment rate and land price.

same land plot. In the Japanese context, land appraisal prices are based on transaction prices and well reflect market conditions. Indeed, studies have found a strong correlation between appraisal and transaction prices (LaPoint 2021; Yamagishi and Sato 2023).

We construct our municipal land price index as follows. We first calculate the ratio of the official land price of the current year to that of the previous year for all land plots that are appraised in two consecutive years.[21] For each year, we then calculate the average ratio within each municipality. Finally, we use this ratio to construct a municipality-level land price index, normalizing the 2002 price to 1. Importantly, our land price data account for any fixed characteristics of each land plot, even if they are unobservable.[22] The number of municipalities included in the land price analysis is 1355, which is somewhat less than the number included in our population and unemployment analysis because only municipalities with a land price survey point are included. The summary statistics of this dataset are shown in Table C.3.

Municipal fiscal data As potential control variables, we obtain each municipality's fiscal information from the Local Government Finance Survey. We collect local tax revenues per capita, lump-sum transfers (called local allocation taxes, LAT) per capita, and earmarked subsidies (called national treasury disbursements, NTD) per capita.

4 Spillover effects on private-sector wages

This section analyzes the spillover effect of the public-sector wage cut on private-sector wages. Section 4.1 briefly discusses the theoretical mechanisms behind wage spillover effects, and this section focuses on identifying the changes in equilibrium private-sector wages in response to the public-sector wage cut. Section 4.2 introduces our empirical strategy that exploits the wage cut and stronger exposure of young workers to the public-sector wage cut. Section 4.3 presents and discusses our main empirical results.

4.1 Potential mechanisms behind spillover effects

Before we estimate the wage spillover effect of the public-sector wage cut, we briefly discuss the underlying mechanisms behind the spillovers. First, note that the public-sector wage cut may impact both the labor supply and demand in the private sector. For the labor supply, more workers are

[21] We use land price data for any land use in our main analysis. Focusing on residential land plots hardly changes our conclusion.

[22] This feature is analogous to the repeat sales index of Case and Shiller (1987, 1989) in that it utilizes multiple observations of prices for the same land plot, but our price data are based on multiple appraisals rather than multiple transactions. LaPoint (2021) uses a similar land price index to ours in a different context, calling it a "repeat appraisal index."

likely to supply their labor in the private sector as public sector jobs become less attractive due to the wage cut (e.g., Burdett 2012; Gomes 2015; Bradley et al. 2017; Albrecht et al. 2019). The labor demand may decline if the public-sector wage reduces the quality of public goods, which contributes to firms' production (Borjas 1984). The labor demand also decreases if the public-sector wage cut exacerbates the monopsony power of private firms, which might occur if the labor supply to the private sector becomes more inelastic as public sector jobs become less attractive than outside options (e.g., Caldwell and Danieli 2024). Alternatively, labor demand may decrease if a "demonstration effect" of the public-sector wage exists (Afonso and Gomes 2014). For example, firms may not observe the productivity of young workers but may learn about it by observing the public-sector wages. The public-sector wage cut would then decrease the labor demand by reducing the perceived marginal productivity of young workers.

Both the decline in labor demand and the increase in labor supply predict a wage decline in the private sector. In this section, we do not distinguish between these two scenarios but rather focus on identifying the effect of the public-sector wage cut on the equilibrium wage rates in the private-sector labor market. In Section 5, we obtain suggestive evidence that the labor demand is reduced by documenting a decrease in private sector employment. In the next subsection, we describe our empirical strategy for estimating the response of equilibrium private-sector wage rates to the public-sector wage cut.

4.2 Empirical strategy

We estimate the effect of public workers' wages on private workers' wages considering intergenerational heterogeneity by the following triple-difference model:

$$\ln w_{i,j,t,private} = \beta RA_{j,t} \times Young_i + \mu_{j,t} + \sum_{k=\text{young or old}} (\iota_j^k + \tau_t^k) + \gamma X_{i,j,t} + \epsilon_{i,t} \tag{2}$$

$$\begin{aligned} \ln w_{i,j,t,private} = & \sum_{t \neq 2005} \beta_t \{\tau_t \times (RA_{j,2010} - RA_{j,2005}) \times Young_i\} \\ & + \mu_{j,t} + \sum_{k=\text{young or old}} (\iota_j^k + \tau_t^k) + \gamma X_{i,j,t} + \epsilon_{i,t}, \end{aligned} \tag{3}$$

where i is an individual worker, j is the municipality where worker i lives, and t is the year. $Young_i$ is a dummy variable taking the value of 1 if i's age is less than 30 and 0 otherwise. $X_{i,j,t}$ is the vector of control variables (i.e., individual characteristics[23]), and $\epsilon_{i,t}$ is the error term. $\mu_{j,t}$, ι_j^k, and τ_t^k capture the municipal-year, municipal, and year fixed effects, respectively. $w_{i,j,t,private}$ is the

[23] Note that our specification includes municipal-year-fixed effects, which absorb all municipal characteristics.

salary level of private worker i in municipality j at t. $RA_{j,t}$ is the regional allowance rate in the national public sector, and this can be interpreted as the index of the public-sector wage level.[24] We use the clustered standard errors at the municipality level.

The coefficient of interest in equation (4) is β, which shows the elasticity of private-sector wages relative to public-sector wages. In equation (5), the event study specification, the coefficient of interest is β_t, which is the elasticity of the private-sector wage in year t with respect to the total change in the regional allowance rate from this policy change (i.e., $RA_{j,2010} - RA_{j,2005}$). Note that both approaches are complementary. The first specification (4) summarizes all the information in the single elasticity β, while the second specification (5) permits the policy reform to have a different impact on private wages in different years.

Our identification assumption behind equations (2) and (3) is a triple-difference strategy that combines a quasi-experiment in regional allowances and an institutional setting in which only young workers are primarily exposed to the policy change. Note that simply regressing the public workers' salary level on the corresponding private workers' salary level may suffer from endogeneity in that the public workers' base wage is annually determined in reference to the previous year's base wage of the corresponding private workers. In particular, we expect a positive bias in β if public-sector wages tend to increase when private-sector wages are high. To address this endogeneity issue, we first utilize the quasi-experimental variations in national officials' regional allowance levels in each municipality j, $RA_{j,t}$. Given that the change in regional allowances resulting from the 2006 reform was unexpectedly introduced and that the previous 10-year average wage index determined the regional allowance rate, $RA_{j,t}$ is less likely to suffer from endogeneity.

To further address the potential endogeneity that $RA_{j,t}$ is systematically correlated with municipality-specific time trends, we take older workers as the control group and compare the evolution of the wage difference between young and older workers across municipalities. To implement such a triple-difference strategy, we introduce municipality-year fixed effects $\mu_{j,t}$, which can flexibly capture municipality-specific trends. Importantly, these include the municipality-specific wage trend that may be correlated with the regional allowance rate $RA_{j,t}$. We also control for municipality and year fixed effects separately for young and old workers (ι_j^k and τ_t^k). As a result, we compare the differential private wage gap between young and older workers in municipalities with different regional allowance rates to estimate the elasticities β and β_t. Note that although this is a two-way fixed effects specification, we use the ordinary least squares estimation since the so-called "nega-

[24] To see why, let $Base_{t,k}$ be the index of the base salary level of national public-sector workers in year t for worker type k (= young or old). From equation (1), the log wage of national public-sector workers of type k is written as $\ln(1 + RA_{j,t})Base_{k,t}$. We write its associated regression coefficient for the young as β, while we assume that the regression coefficient is zero for older workers because they are assumed to be unaffected by the public-sector wages (see Section 2). Here, $\ln(1 + RA_{j,t})Base_{k,t} = \ln(1 + RA_{j,t}) + \ln Base_{k,t} \simeq RA_{j,t} + \ln Base_{k,t}$, but we do not need to explicitly control for $\ln Base_{k,t}$, as it is absorbed by the year fixed effects τ_t^k. As a result, the coefficient of $RA_{j,t} \times Young_i$ equals β in equation (2).

tive weight" problem does not arise because the timing of the policy change is the same across all municipalities (Roth, Sant'Anna, Bilinski and Poe 2023). Our triple-difference strategy relaxes the common trend assumption required in a difference-in-differences strategy by accommodating any municipality-specific time trend common to young workers and others, where the municipality is the level at which the public-sector wage schedule varies (Olden and Møen 2022).

As covariates, $X_{i,j,t}$, we consider both individual i's characteristics and municipal fiscal characteristics. The individual characteristics consist of the dummies for gender and university degree, age categories with 5-year intervals from 15 to 64, and their interaction terms. In addition, we include the logarithm of scheduled working hours. The inclusion of these control variables is analogous to the earnings functions à la Mincer, but we use dummy variables to relax functional form restrictions (Kawaguchi 2011).[25] As fiscal characteristics, we use the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita.[26] In addition, although the salary level of municipal public workers corresponding to private worker i in municipality j at t, $w_{i,j,t,public}$, could suffer from endogeneity, we include it in some specifications to investigate the sensitivity of our results to the potential discretionary wage changes by the municipal government.[27]

4.3 Spillover estimates of private-sector wages

Figure 5 shows the effect of public-sector wages on private wages from the event study specification (3). Given that the public-sector wage reform led to a reduction in wages, note that the positive coefficients indicate a reduction in private sector wages. In each panel, β_t is estimated based on specifications with different sets of control variables. In all panels, there is little indication that the parallel trend assumption is violated in the pre-treatment period. Because the reform of regional allowances was conducted gradually from 2006 to 2010, the estimated β_ts in each panel gradually increased from 2006 to 2010 and reached approximately 0.4 after 2011. The positive impact on private wages is confirmed by Table 1, which reports an elasticity estimate of approximately 0.35 from the specification (2). In particular, our preferred estimate in column (2), which controls for individual characteristics, suggests that a 1% public-sector wage cut induces a decrease in private

[25] Our control variables are meant to compare workers with similar observable characteristics because our aim is to identify the elasticity with respect to public-sector wages β. The coefficients on our covariates should not necessarily be taken as causal due to endogeneity, such as self-selection of schooling (Heckman, Lochner and Todd 2006).

[26] In the logarithmic transformation, one is added to the original number to prevent data omission.

[27] Because the average salary level of local public employees by education and years of service are available, we match each private worker's corresponding salary level of municipal public workers by the education and years of service estimated from age and education to construct this variable. Workers lacking a corresponding municipal public worker are omitted from the sample.

wages of 0.32%.[28] Given that the municipal-year fixed effect is controlled in our triple-difference strategy, municipality-specific shocks would not explain this result. Therefore, this result suggests that the policy change in regional allowances induced a spillover from the wages of public workers to the wages of private workers.

The identified effect of the public-sector wage on the private wage is not only statistically significant but also economically meaningful. We illustrate this in two ways. First, we compare the effect size to the gender wage gap and college premium, two of the most salient wage disparities in the data. The regional allowance rate $RA_{j,t}$ ranges from 0-0.18 in our data, implying that moving from the lowest regional allowance rate to the highest increases wages by approximately 6% ($0.32 \times 0.18 \simeq 0.057$). This amounts to approximately one-fifth of the gender wage gap and one-third of the college premium in our sample.[29] Second, the estimated elasticity implies a substantial aggregate impact on private wages. In particular, we calculate in Section 6.4 that at the national level, the private-sector wages decreased by 111.6 billion yen due to the policy reform. Moreover, despite our focus on the wage cut, the public sector, and numerous other contextual differences, our preferred elasticity estimate of 0.32 is relatively close to several spillover estimates in the literature (e.g., Staiger et al. 2010; Bassier 2022).

The spillover effect on the private-sector wages suggests the monopsony power of the public sector because the public sector can affect market-based wages. If the public-sector internalizes this spillover effect, it acts as a price maker in the labor market. It is sensible to consider the public sector monopsonic given its significant employment share. The monopsony power of the public sector in the labor market is also consistent with Dal Bó et al. (2013), who experimentally find a finite labor supply elasticity to the public sector.[30]

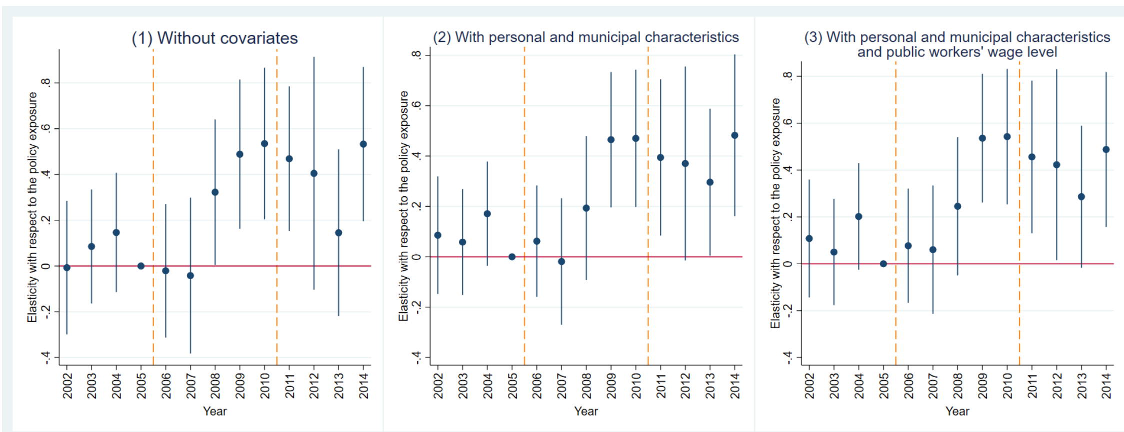
Larger wage spillovers in municipalities with a greater share of public workers. We may naturally hypothesize that the wage spillover effect of the public sector would be stronger in areas with a greater share of public-sector workers. To understand this heterogeneity, we repeat the same analysis as in Sections 4.3 and 5.3 for the two subsamples, where the sample is split by the share of public workers in the region: one subsample is regions with a higher-than-median number of municipal public workers per capita, and the other is regions with a lower-than-median number of municipal public workers per capita. In our data, the median number of municipal public workers

[28]The estimated coefficients of the control variables are consistent with the results of Kawaguchi (2011), who apply the Mincer wage equation to Japanese data.

[29]We calculate the gender wage gap and college premium in our sample as follows. First, our regression results in column (2) of Table 1 include gender and college dummies, which interact with five-year age categories. We then construct the weighted average of gender dummies and college dummies, where the weight is the frequency of each age category in the sample.

[30]Note that the labor supply elasticity is also finite in our context because we find no evidence that the public sector could not hire workers after the wage cut.

Figure 5: The elasticity of the private-sector wage in year t with respect to the regional allowance change during 2006-2010



Note: This figure shows the estimated β_t for each year in the eq.(3). This corresponds to the elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010, when the reform of regional allowances was conducted. 95% confidence intervals based on standard errors clustered at the municipality level are also shown. In all panels, we control municipal, year, and municipal-year fixed effects. In panel (1), we do not control any other variables. In panels (2)-(3) we control for individual (age dummies interacted with the college education dummy and gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in panel (3).

Table 1: Regression results on private-sector wages based on (2)

	(1)	(2)	(3)
	log(wage rate of private workers)		
Regional allowances × Young dummy	0.3791***	0.3233***	0.3583***
	(0.1120)	(0.0984)	(0.0942)
log(base wage of local municipal workers)			0.0736***
			(0.0105)
Year fixed effect	Yes	Yes	Yes
Municipality fixed effect	Yes	Yes	Yes
Municipality-year fixed effect	Yes	Yes	Yes
Individual and municipal fiscal characteristics	No	Yes	Yes
N	12194536	12194536	11668764
R^2	0.264	0.514	0.508

Standard errors clustered at a municipal level in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: The regression results of estimating equation (2) are presented. In columns (2)-(3), we control for individual (age dummies interacted with college education dummy and gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in column (3). Note that the sample size of column (3) is limited because samples lacking public workers corresponding to the municipal public workers in terms of education and experience are omitted.

per capita is approximately 0.01, i.e., one municipal public worker for every 100 residents.[31] We expect the effects of public-sector wages on private wages and the young population to be stronger in the sample with a greater share of public-sector workers.

Consistent with this hypothesis, our results show that the impact of public wages is greater in municipalities with a larger share of public employees. Figure 6 shows that the spillover of the public wage on the private wage is larger in the regions with more public workers, especially in the post-treatment period 2011–2014. Moreover, while the coefficients are statistically significant after the policy reform in regions with a higher share of public-sector workers, the significance is more limited in regions with a lower share. Consistent with this, Table B.1 shows that while the size of the coefficients in the specification (2) for the regions with a higher-than-median number of municipal public workers per capita is 0.35 in panel (a), the corresponding estimates are approximately 0.25 for the regions with a lower-than-median number of municipal public workers per capita.

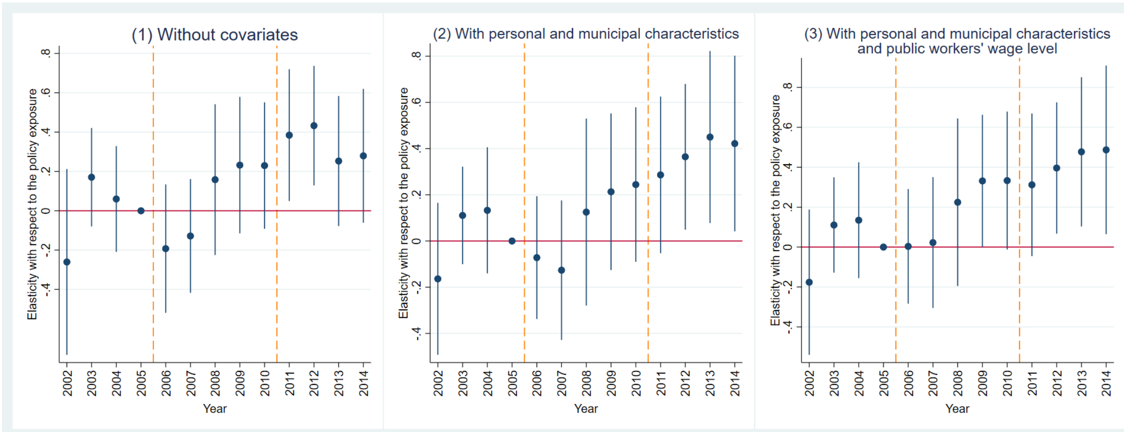
Overall, the effects of public-sector wages on private wages are larger in areas with a greater share of public workers. Given the small share of public-sector workers in Japan compared to other OECD countries (OECD 2021), the results in this section suggest that our wage spillover estimate may serve as a lower bound for the effects of public-sector wages in other countries. Moreover,

[31]Strictly speaking, it is 0.0098 for the data of private wage, 0.0120 for the data of population, and 0.0106 for the data of land price in FY2005.

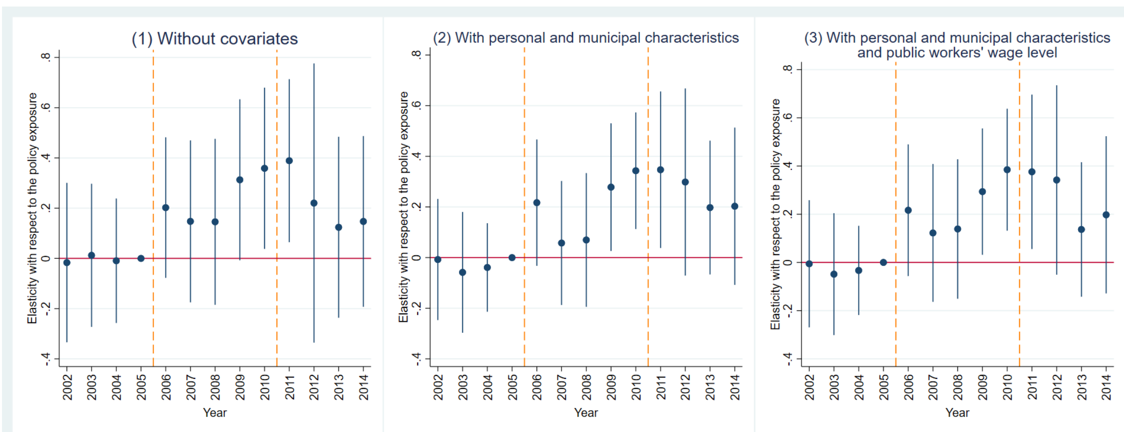
the larger effects in areas with a greater share of public workers suggest that our empirical analysis indeed captures the effects of public-sector wages but not other confounding factors. Indeed, it would be difficult to explain the heterogeneous effects with respect to the public worker share if our results were spuriously driven by confounding factors unrelated to the public sector.

Figure 6: The elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010 in the regions with a higher- and lower-than-median share of public employees

(a) Analysis for the regions with a higher-than-median share of public employees



(b) Analysis for the regions with a lower-than-median share of public employees



Note: These figures show the estimated β_t for each year in the eq.(3). This corresponds to the elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010, when the reform of regional allowances was conducted. We restrict the sample to regions with a higher-than-median (lower-than-median) share of public employees in 2005 in figures (a) and (b). 95% confidence intervals based on standard errors clustered at the municipality level are also shown. In all panels, we control municipal, year, and municipal-year fixed effects. In panel (1), we do not control any other variables. We control covariates regarding individual (age dummies interacted with college education dummy and gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita) in panels (2) and (3). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in column (3) although it is not controlled in the other columns. Note that the sample size of column (3) is limited because samples lacking public workers corresponding to the municipal public workers in terms of education and experience are omitted.

5 Population response and welfare implication

The previous section reveals that the 1% public-sector wage cut induced an approximately 0.3% reduction in the private-sector wages of young workers. This section analyzes the response of the young population to the public-sector wage cut. We discuss the theoretical motivations behind this analysis in Section 5.1. We describe our empirical strategy in Section 5.2 and present our empirical results in Section 5.3.

5.1 Theoretical motivations

In addition to the importance of population itself as a local economic outcome, there are two theoretical motivations for analyzing the population response to the public-sector wage cut. First, it is suggestive of the welfare impact of the public-sector wage cut. Although the decline in private-sector wages among young workers is suggestive of their welfare decline, this finding is still inconclusive because the public-sector wage cut may also have other effects. For example, it may affect the quality of public goods (Borjas 1984) and job amenities in the private sector. To address this, we analyze the population response of young workers motivated by a simple spatial equilibrium logic á la Rosen (1979) and Roback (1982) (see Appendix D for details of the model). Intuitively, if young workers' welfare declines due to the public-sector wage cut in their local labor market, more of them are likely to move out of the market. Notably, the welfare decline is inclusive of various potential effects caused by the public-sector wage cut, including effects on public goods quality and job amenities.

Second, the population response is suggestive in distinguishing whether the wage spillover is caused by the demand or supply shift in the private-sector labor market. As discussed in Section 4.1, the equilibrium wage decline in the private sector can be induced by either an increase in the labor supply or a decrease in demand. However, they make contrasting predictions about employment: an increase in supply is associated with more employment, while a decrease in demand is associated with less employment.[32] Overall, analyzing the population response helps us better understand welfare implications and the underlying mechanisms of wage spillovers.

5.2 Empirical strategy

Using municipality-level panel data, we estimate the following regression equations for estimating the effect of public workers' wages on the young population:

[32] To check whether the population response reflects the workforce response, we also analyze the unemployment rate response in Section 6.1.

$$\ln(\text{YO ratio of Pop})_{j,t} = \beta RA_{j,t} + \iota_j + \eta_{t,p} + \gamma X_{j,t} + \epsilon_{j,t} \tag{4}$$

$$\ln(\text{YO ratio of Pop})_{j,t} = \sum_{t \neq 2005} \beta_t \{\tau_t \times (RA_{j,2010} - RA_{j,2005})\} + \iota_j + \eta_{t,p} + \gamma X_{j,t} + \epsilon_{j,t}, \tag{5}$$

where j is the municipality in prefecture p and t is the year. The outcome variable, $\ln(\text{YO ratio of Pop})_{j,t}$, is defined as the logarithm of the ratio of the young population to the older population (Young $\text{Pop}_{j,t}$ / Older $\text{Pop}_{j,t}$). $X_{j,t}$ is the vector of control variables (i.e., municipal characteristics), and $\epsilon_{j,t}$ is the error term. ι_j, $\eta_{t,p}$, and τ_t capture the municipality, prefecture-year, and year fixed effects, respectively. We weight the observations by the population as of 2000. We use the clustered standard errors at the municipality level.

The coefficients of interest, β and β_t in equations (4) and (5), are interpreted as the elasticity of the young population with respect to the regional allowances. This is because, as in the triple-difference strategy for analyzing private wages, using the young-to-older population ratio allows us to flexibly control for municipality-specific trends by taking the older population as the control group. To illustrate this point, suppose the following model of the young population: $\ln(\text{Young Pop})_{j,t} = \beta RA_{j,t} + \iota_j^{'} + \eta_{t,p}^{'} + \Xi_{j,t}^{'} + \gamma^{'} X_{j,t} + \epsilon_{j,t}^{'}$, where $\iota_j^{'}$, $\eta_{t,p}^{'}$, and $\Xi_{j,t}^{'}$ are municipality, prefecture-year, and municipality-year fixed effects, respectively. Since the municipality-year specific trend $\Xi_{j,t}^{'}$ is perfectly collinear with the regional allowances $RA_{j,t}$, this model does not identify β. However, now suppose an analogous model for the older population: $\ln(\text{Older Pop})_{j,t} = \iota_j^{''} + \eta_{t,p}^{''} + \Xi_{j,t}^{'} + \gamma^{''} X_{j,t} + \epsilon_{j,t}^{''}$. Here, the regional allowance ($RA_{j,t}$) does not appear because it is assumed to have no effect on the older population.[33] Note also that the municipality-year fixed effect ($\Xi_{j,t}^{'}$) is assumed to be age-independent. Then, by subtracting the former from the latter, we obtain the estimation equation (4).[34] The coefficient β is therefore interpreted as the effect of regional allowances on the young population.[35] Similarly, β_t in equation (5) is also interpreted as the effect on the young population.

As control variables, we consider the same set of fiscal characteristics as in the private-sector wage analysis: the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita. We also consider the average salary level of municipal public workers in some specifications.

[33]The analysis using the logarithm of the old population as the dependent variable shows that the regional allowance has no effect on the old population, which is consistent with this assumption. The results are available from the authors upon request.

[34]To see this, we can define $\iota_j \equiv \iota_j^{'} - \iota_j^{''}$. Other variables are defined analogously.

[35]Note that the result is qualitatively the same even when we use the logarithm of the young population as the dependent variable.

5.3 Estimated responses of the young population and welfare implications

Figure 7 shows the impact of public-sector wages on the young population from the event study specification (5). In all specifications, the downward trend in the young population (relative to the older population) in municipalities that experienced an increase in public wages stopped after 2006. This indicates that the increase in the public wage stopped the decline in the young population. Therefore, this result supports the expectation that a public-sector wage cut induces the outflow of the young population.

Table 2 reports the elasticity of the young population with respect to public wages from the regression specification (4). In Columns (1)-(3), we find that the elasticity is approximately 0 for all specifications. However, this regression estimate of public wage elasticity is misleading because it does not account for the preexisting trends apparent in Figure 7, leading to substantial underestimation of the true elasticity of the public-sector wages. To correct for this underestimation, in Columns (4)-(6), we account for preexisting trends by following a common detrending procedure of the outcome variable (e.g., Kleven, Landais, Saez and Schultz 2013; Monras 2019; Garcia-López, Jofre-Monseny, Martínez-Mazza and Segú 2020).[36] Columns (4)-(6) show that the young population becomes approximately 0.4% smaller in response to a 1% reduction in public-sector wages after controlling for preexisting trends. In particular, our preference in column (5), which accounts for observable municipality-year control variables and preexisting trends, suggests that the young population decreases by 0.4% in response to a 1% reduction in public-sector wages.

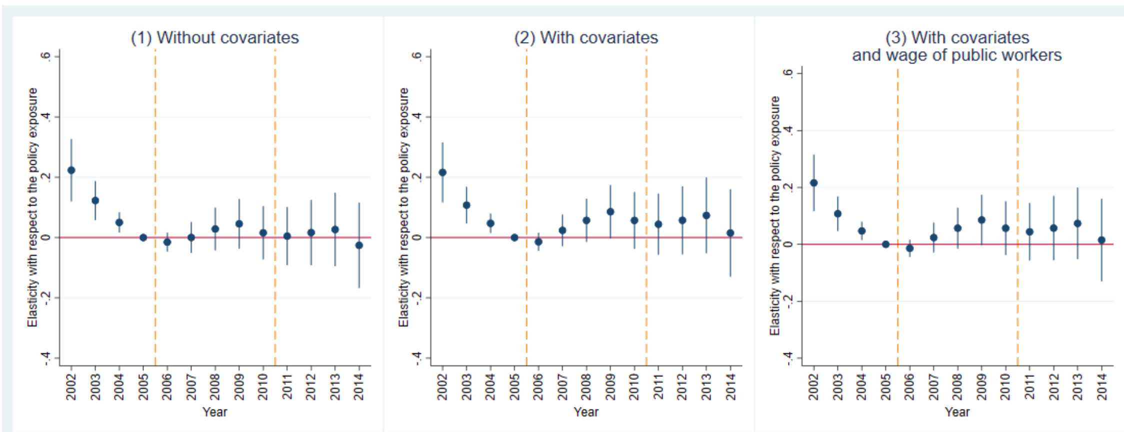
Larger impacts in municipalities with a greater share of public workers. Similar to the case for private-sector wages, we find evidence that the impact on the young population is greater in municipalities with a greater share of public-sector workers. Figure 8 shows that a larger preexisting trend can be seen for the region with many public workers compared with the region with few public workers, while such trends changed after 2006 in both regions. This suggests that the policy effect of public wage reduction on the young population is greater for regions with many public workers when the preexisting trend is controlled. Table B.2 indeed shows that after the preexisting trend is removed, the magnitude of the coefficient is approximately 0.66 in the regions with many public workers and 0.37 in the regions with few public workers. This result supports the idea that the young

[36] Specifically, to construct the detrended outcome variable $\widetilde{\ln(\text{YO ratio of Pop})}_{ijt}$, we first estimate the municipality-specific pretrend by the following equation:

$$\ln(\text{YO ratio of Pop})_{jt} = \iota_j + \tau_t + \iota_j \times \tau_t + \epsilon_{jt} \quad \text{for } \ t < 2006.$$

We predict the values of $\ln(\text{YO ratio of Pop})_{jt}$ for entire samples, including $t \geq 2006$, from this result and compute the residuals corresponding to $\widetilde{\ln(\text{YO ratio of Pop})}_{jt}$.

Figure 7: The elasticity of the young population in year t with respect to the change in regional allowances during 2006-2010



Note: This figure shows the estimated β_t for each year in the eq.(5) that the logarithm of the ratio of young-to-older population is the dependent variable. As discussed in Section 4.2, this corresponds to the elasticity of the young population in year t with respect to the change in regional allowances during 2006-2010, when the reform of regional allowances was conducted. In all panels, we control municipality, year, and prefecture-year fixed effects. In panel (1), we do not control any other variables. We control covariates (the logarithms of local tax revenue per capita, LAT per capita, NTD per capita, and the number of local officials per capita) in panels (2) and (3). We add the average salary level of public workers in municipality j at t, $w_{j,t,public}$ in panel (3). 95% confidence intervals based on standard errors clustered at the municipality level are also shown.

Table 2: Regression results on the young population based on (4)

	(1)	(2)	(3)	(4)	(5)	(6)
	log(YO ratio of Pop)			$\widetilde{\text{log(YO ratio of Pop)}}$		
Regional allowances	-0.0748	-0.0246	-0.0269	0.4609***	0.4065***	0.3730***
	(0.0550)	(0.0556)	(0.0560)	(0.1069)	(0.1043)	(0.1029)
log(tax revenue per capita)		0.0534***	0.0544***		-0.0897***	-0.0748***
		(0.0176)	(0.0178)		(0.0243)	(0.0234)
log(LAT per capita)		0.0033	0.0032		0.0065**	0.0065*
		(0.0022)	(0.0022)		(0.0032)	(0.0034)
log(NTD per capita)		0.0012	0.0013		-0.0019	-0.0005
		(0.0021)	(0.0021)		(0.0037)	(0.0036)
log(municipal public workers per capita)		0.0173*	0.0168*		-0.0444***	-0.0517***
		(0.0093)	(0.0093)		(0.0160)	(0.0158)
log(base wage of local municipal workers)			-0.0224			-0.3305***
			(0.0311)			(0.0453)
Year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipal fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Detrended outcome variable	No	No	No	Yes	Yes	Yes
N	22503	22503	22503	22503	22503	22503
R^2	0.967	0.967	0.967	1.000	1.000	1.000

Standard errors in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

The regression results of estimating equation (4) in which the logarithm of the ratio of the young to the older population, log(YO ratio of Pop), and its detrended variable, $\widetilde{\text{log(YO ratio of Pop)}}$, are the dependent variables are presented. As discussed in Section 4.2, this allows us to estimate the elasticity of the young population with respect to public-sector wages. Young and older populations correspond to the 15-29 and 30-64 age groups, respectively. $\widetilde{\text{log(YO ratio of Pop)}}$ is the value that detrends the 2002-2005 region-specific linear trends from log(YO ratio of Pop). Specifically, to construct the detrended outcome variable $\widetilde{\text{log(YO ratio of Pop)}}_{jt}$, we first estimate the municipality-specific pretrend via the following equation

$$\ln(\text{YO ratio of Pop})_{jt} = \iota_j + \tau_t + \iota_j \times \tau_t + \epsilon_{jt} \quad \text{for} \quad t < 2006.$$

We make predicted values of $\ln(\text{YO ratio of Pop})_{jt}$ for entire samples, including $t \geq 2006$, from this result and compute the residuals corresponding to $\widetilde{\ln(\text{YO ratio of Pop})}_{jt}$.

The average salary level of public workers in municipality j at t, $w_{j,t,public}$, is controlled in columns (3) and (6), although it is not controlled in the other columns because it is likely to suffer from endogeneity.

population indeed responds to changes in public-sector wages.[37]

Implications for welfare and the underlying mechanism of wage spillovers. In light of our spatial equilibrium model following Rosen (1979) and Roback (1982) (see Appendix D for details), the decrease in the young population in response to the public-sector wage cut implies that the wage cut harmed young workers' welfare. Intuitively, if young workers' welfare declines due to the public-sector wage cut in their local labor market, more of them are likely to move out of the region. Notably, the welfare decline is inclusive of various potential effects caused by the public-sector wage cut, including a decrease in the public-sector wage, wage spillover in the private sector, changes in the quality of public goods, and changes in job amenities.

This result has several policy implications. For instance, public-sector wage cuts to achieve fiscal consolidation, which were widely adopted in many countries in the aftermath of the Global Financial Crisis (Forni and Novta 2014), may have harmed young workers. Another implication is that increasing public-sector wages may be an effective policy tool for achieving wage increases and improving welfare, which may be particularly useful when wages stagnate due to prolonged recessions, as in the Japanese context (as suggested by, for example, Bernanke 2017).

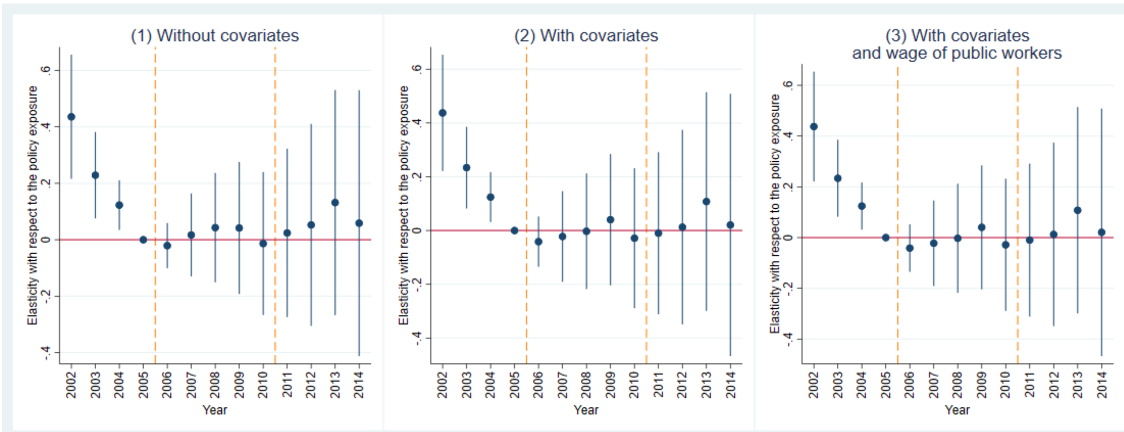
However, it should also be noted that the public-sector wage cut may be beneficial for older workers. First, wage spillovers and the opportunity to obtain a better job in the public sector may be irrelevant for older workers, which is likely to be the case in our context because there is little opportunity for older workers to obtain a public-sector job. Second, a decrease in the young population may lower land prices, which decreases the housing costs of older workers. In contrast to the case for young workers, who are mobile, the lower housing cost may actually increase older workers' welfare (see Appendix D). Consistent with the lower housing cost due to the out-migration of young workers, Section 6.2 provides suggestive evidence that the public-sector wage cut had a negative impact on land prices. Overall, the welfare effect of the public-sector wage cut would be heterogeneous across workers of different ages: young workers would likely suffer from the public-sector wage cut, but older workers might benefit from it.

As discussed in Sections 4.1 and 5.1, the decrease in private-sector wages and workforce size together imply that labor demand shifted downward in the private-sector labor market. The decrease in employment is further reinforced by our finding that the public-sector wage cut, if any, induces an increase in unemployment rates (Section 6.1). While our data do not allow us to formally unpack how the public sector wage cut caused the decline in the labor demand, we point out that the increase in monopsony power is consistent with our results. First, monopsony power implies that the wage

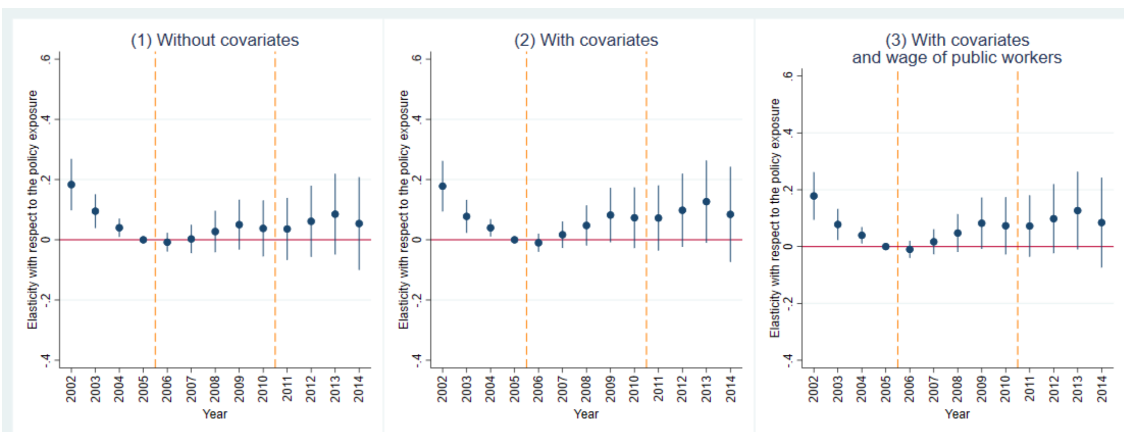
[37] In addition to conducting the wage and young population analyses, we confirm that the similar analyses of young unemployment and land prices show larger effects in areas with a greater share of public workers. The results are available from the authors upon request.

Figure 8: The elasticity of the young population in year t with respect to the change in regional allowances during 2006-2010 for the regions with a higher- and lower-than-median share of public employees

(a) Analysis for the regions with a higher-than-median share of public employees



(b) Analysis for the regions with a lower-than-median share of public employees



Note: This figure shows the estimated β_t for each year in the eq.(5) that the ratio of the young to the older population is the dependent variable. As discussed in Section 4.2, this corresponds to the elasticity of the young population in year t with respect to the change in regional allowances during 2006-2010, when the reform of regional allowances was conducted. We restrict the sample to regions with a higher-than-median (lower-than-median) share of public employees in 2005 in panels (a) and (b). In all panels, we control municipal, year, and municipal-year fixed effects. In panel (1), we do not control any other variables. We control covariates (the logarithms of local tax revenue per capita, LAT per capita, NTD per capita, and the number of local officials per capita) in panels (2) and (3). We add the average salary level of public workers in municipality j at t, $w_{j,t,public}$ in panel (3). 95% confidence intervals based on standard errors clustered at the municipality level are also shown.

decline in response to the public-sector wage cut is observed only for young workers, which is suggested by the different wage trends by age (Figure 4). As competition between the private and public sectors occurs only for young workers in our context (Section 2), the labor supply elasticity of the private sector should change only among young workers, inducing a rise in monopsony power (i.e., the markdown rate). Second, we find that the wage and population responses are greater in areas with a larger share of public-sector workers. This is consistent with Azar, Huet-Vaughn, Marinescu, Taska and Von Wachter (forthcoming) showing that monopsony power tends to be greater in areas with lower population density. We expect greater monopsony power in municipalities with a greater share of public-sector workers because these municipalities tend to be more rural. Overall, while we cannot rule out other explanations behind the downward shift in private-sector labor demand, monopsony power is one possible explanation.

6 Additional analysis

We conduct several supplementary analyses. Section 6.1 estimates the effect of public-sector wages on youth unemployment rates, finding some evidence that the public-sector wage cut might increase youth unemployment rates. Section 6.2 estimates the effect on land prices. We find that the public-sector wage cut decreases land prices, which is consistent with the population decline found in Section 5.3. Section 6.3 considers the heterogeneous effects on private-sector wages with respect to gender, educational attainment, industry, and firm size. We find that the wage elasticity is greater for less-educated workers. The spillover is greater in industries that receive more young workers from the public sector. We find little heterogeneity in terms of gender and firm size. Section 6.4 presents further additional results. We first investigate the sensitivity of our results to the largest cities and public-sector wages of neighboring municipalities and find that our results are robust to these considerations. We also calculate the national-level impacts of the 2006–2010 public-sector wage cut based on our estimates to illustrate the quantitative relevance of wage spillovers in evaluating the public-sector wage cut policy.

6.1 The effects on youth unemployment rates

Thus far, we have focused on the private wages of young workers and the young population as the outcomes of our analyses. In addition to these outcomes, the unemployment rate may depend on the public-sector wage rate by changing workers' job search behavior (e.g., Gomes 2015; Bradley et al. 2017; Albrecht et al. 2019). The analysis of unemployment is also interesting for examining whether the decrease in population we found in Section 5.3 translates into a decrease in employment.

We analyze the effect of the change in regional allowances on the unemployment rate through regression based on the specifications of (4) and (5). The outcome variable is the log ratio of the unemployment rate of young workers to that of older workers. By the same argument as in Section 4.2, the estimate corresponds to the elasticity of the unemployment rate of young workers with respect to public-sector wages. We note that our results on unemployment should be interpreted with caution because of the data frequency: unemployment rate data at the municipality level are available only for census years (2000, 2005, 2010, 2015). We use linear interpolation for intermediate years.

Figure A.2 and Table B.3 present the results. Figure A.2 indicates that although there is an overall positive trend, the increase in the unemployment rate became smaller after 2005 in municipalities that did not experience the public-sector wage cut. In terms of the magnitude of the effect, Columns 4–6 of Table B.3, which account for the positive preexisting trend, show that a 1% increase in regional allowances reduces the unemployment rate for young workers by approximately 0.49%. Although the estimate is somewhat noisy and statistically insignificant, it indicates the possibility that public-sector wage cuts might exacerbate youth unemployment rates. Overall, we find suggestive evidence that the unemployment rate of young workers increased in response to the public-sector wage cut.

6.2 The effects on land prices

We have shown in Section 5.3 that the young population decreased in response to the decrease in the public-sector wages. The decreased population is, in turn, predicted to lower land prices due to weaker housing demand (see Appendix D for more details).

Motivated by this, this subsection analyzes land prices. We use municipality-level panel data for analyzing land prices. We use the specifications (4) and (5), where the outcome variables are replaced with the log land price index.

Figure A.3 and Table B.4 show the results for the effect on land prices. Figure A.3 shows the effect of the public wage on the land price after the reform, where the elasticity is approximately 0.5. Table B.4 shows that a 1% decrease in public-sector wages induces a 0.54% decrease in land prices.[38]

Overall, we find evidence that land prices decreased in response to the public-sector wage cut, which is consistent with the decrease in population found in Section 5.3. Based on our theoretical prediction in Appendix D, the decreased land prices suggest that the welfare of young workers decreased. However, the decline in land prices may have benefited older workers, who are unlikely to be affected by the public-sector wage cut in the labor market but benefit from the decreased housing cost (see the discussion in Section 5.3). As such, the land price responses are indicative

[38]We also find that the elasticity is greater for regions with a greater share of public workers: 1.03 for regions with a greater share and 0.46 for regions with a lower share. This is consistent with Section 5.3 that the population decreases more in regions with a higher share of public workers.

that the public-sector wage cut may have harmed young workers but not older workers.

6.3 Heterogeneity in the effects on private wages

We investigate the heterogeneity in the effect on private wages with respect to firm and worker characteristics. Overall, we find that the effects of public-sector wages on private wages tend to be greater for noncollege-educated workers.The spillover is also larger in industries that receive more young workers from the public sector. In contrast, we find little heterogeneity by gender and firm size.

Gender. Figure A.4 and Table B.5 present the estimated elasticity of private wages for men and women, respectively. We do not find a significant difference in the elasticity estimates for men and women. Although women are more likely to work in the public sector for reasons such as a smaller gender wage gap and other job amenities (Gomes and Kuehn 2019), such heterogeneity does not seem to translate into gender differences in spillover elasticity.

Education. A subsample analysis of college graduates and nongraduates revealed that nongraduates are more affected by the public wages of the local municipality. Figure A.5 and Table B.6 show that the elasticity of private wages is approximately 0.32 and is statistically significantly positive for nonuniversity graduates, but it is near zero and nonsignificant for university graduates after including control variables. Given that our research design exploits the variation in the public-sector wages across municipalities, this result implies that private wages for university graduates refer less to the public wage in the local municipality, while private wages for noncollege graduates are heavily influenced by it.

We view this result as stemming from the locality of job searches of noncollege graduates. Studies have shown that the geographical scope of job searches is generally narrower for noncollege graduates (Kaplan and Schulhofer-Wohl 2017; Marinescu and Rathelot 2018). Moreover, the Japanese institutional setting reinforces the location of job searches in the labor market for high school graduates. In Japan, firms first send job postings to high schools, and high schools allocate job openings to their students (Genda et al. 2010). Since it is rare for students to apply for jobs and firms are likely to send their job postings to nearby high schools, this institutional system limits the geographical scope of job searches for high school graduates.[39] In contrast, college graduates are generally more mobile, and they search for more distant jobs. As a result, in hiring a worker, firms recruiting noncollege graduates are more likely to recruit within the municipality, and the labor market competition for workers is more likely influenced by the public-sector wage rate in

[39] See, for instance, https://lab.jinjib.co.jp/archives/1016/ for more description of the localized nature of the Japanese job market for high school graduates (in Japanese. Last accessed on April 25, 2024).

the same municipality, at the level of our identifying variation. Overall, the heterogeneity resulting from education can stem from the local nature of our research design, which may not detect wage spillover effects for workers who search for more distant jobs.

Industry. We conduct a subsample analysis by industries with high and low exposure to the labor flow from the public sector, measured by the flow of workers under 29 from the public sector into the industry (normalized by the number of workers under 29 in the industry). The labor flow from the public sector and the number of workers in each industry are obtained from the Survey on Employment Trends and the Labor Force Survey, respectively.[40] Because the definition of industry classification in the Survey on Employment Trends changed several times before 2009, it is challenging to define the level of exposure consistently before and after 2009.[41] Therefore, the analysis is based on the top three industries with the highest exposure since 2009 (Electricity, Gas, Heat Supply and Water; Transport and Postal Services; Real Estate and Goods Rental and Leasing), when comparisons are possible.[42]

Table B.7 shows that the effect of the public wage on private wages is approximately 0.74 for industries with high exposure to labor flow from the public sector, although the effect is approximately 0.21 for other industries. Figure A.7 confirms a similar tendency that the effect is larger for the industries with high exposure to labor flow from the public sector[43] This result is in line with Bassier (2022), who find larger wage spillover effects of collective bargaining reforms in sectors that have more worker flows with the treated sectors.

Firm size. We analyze whether large and small firms react differently to public-sector wages. Figure A.6 and Table B.8 show the results of the subsample analysis by whether the worker belongs to a company with more than 100 employees or not.[44] Table B.8 shows that the effect of public wages

[40]Since data of the Labor Force Survey are absent for 2011 due to the Great East Japan Earthquake, note that the inflows of workers from the public sector compared to the number of workers in the industry cannot not be defined for 2011.

[41]Note that we omit the data for 2011 due to missing data (see footnote 40). Since the individual private-sector wage data include data up to the three-digits classification (*chu bunrui*) of the industry in which the worker is employed, it is possible to perform subsample analysis by adjusting the industry classification using the three-digits classification data. However, we cannot define the level of exposure before and after 2009 because the Employment Trends Survey includes only data for the two-digits classification.

[42]We confirm that even if we use the top two or top four industries (the fourth industry is "Services, N.E.C."), the result does not change. Note that Electricity, Gas, Heat Supply and Water and Transport and Postal Services were always among the three industries with the largest labor flows from the public sector compared to the number of workers in the industries; in addition, they were not affected by changes in industry classification, even before 2009.

[43]One caveat here is that the standard error in the analysis of the high-exposure industries is large compared to the corresponding figures for the low-exposure industries. This is because the sample size of the high-exposure analysis is approximately one-sixth that of the low-exposure analysis.

[44]The threshold of 100 is chosen because the government sets public-sector wages referring to the private wage rates of firms with more than 100 employees, arguing that the working conditions are similar to those in the public sector

on private wages is approximately 0.27 for workers in firms with more than 100 employees and 0.22 for those in firms with fewer than 100 employees. Figure A.6 confirms a similar tendency and shows that the effect of public wages on private wages is similar for both small and large companies. Therefore, both large and small firms might be similarly influenced by public-sector wages.

6.4 Other additional analyses

This section discusses three additional analyses. First, we consider the potential impact of public-sector wages of neighboring municipalities. Second, we explore the sensitivity of our results to exceptionally large cities, such as Tokyo. Third, we illustrate the quantitative importance of wage spillovers by considering the national-level aggregate economic impact of the public-sector wage cut from 2006–2010.

Public-sector wages of neighboring municipalities. We analyze the possibility that not only the public-sector wage rate of a municipality but also the wage rates of neighboring municipalities may affect outcomes. Theoretically, incorporating commuting and general equilibrium effects on the welfare level of the marginal worker leads to such an effect by creating intermunicipal dependence (see Appendix D for details).[45] For this purpose, we use the so-called "SLX" model in spatial econometrics (Halleck Vega and Elhorst 2015), which adds to our main regression the spatial lag that summarizes the regional allowance rates of neighboring municipalities. Adding such a spatial lag term, Figure A.8 and Table B.9 repeat the analysis of Figure 5 and Table 1 for private wages. Similarly, Figure A.9 and Table B.10 repeat the analysis of Figure 7 and Table 2 for the young population. The results of this analysis show that the regional allowance has a slight spillover effect on neighboring municipalities, especially for the young population, but such an effect is limited. Reassuringly, the inclusion of the spatial lag does not change the main result, which implies that our baseline result is robust.

Sensitivity of the results to the largest cities. We explore whether our empirical results are driven by the fact that the increased regional allowance rates are concentrated in the three major metropolitan areas centered on Tokyo, Osaka and Nagoya. Excluding these areas from the sample in the analysis, Table B.12 replicates Table 1 for the private wages of young workers, and Table B.11 replicates Table 2 for the young population. In this analysis, the coefficients for the policy effect

(Aoki 2021).

[45]For example, Monte, Redding and Rossi-Hansberg (2018) and Borusyak, Dix-Carneiro and Kovak (2022) highlight the importance of incorporating such geographical interdependence in a reduced-form analysis. In addition to commuting and the general equilibrium welfare effect, another reason for the interdependence is that in setting their wage rates, private firms might use not only the public wage rates in their own municipality but also those in neighboring municipalities as yardsticks (Besley and Case 1995; Kishishita and Yamagishi 2021).

tend to be larger than those in the baseline analyses. This may be because areas other than the three largest metropolitan areas tend to have a higher share of public employees, and the spillover effect of public employees is more significant, as discussed in Sections 4.3 and 5.3.

National-level aggregate economic impacts. Based on the elasticity estimates of the effects of public-sector wages on the local economy presented thus far, we quantify the national-level aggregate economic impacts of the public wage cut during 2006–2010, the policy reform analyzed in this paper. As discussed in Section 2, the public wage reform was intended to improve the fiscal balance by reducing the base wage of public workers, although the reform increased local public wages in some municipalities due to the increase in regional allowances. However, as our analyses show, the effect of this reform was not limited to reductions in public wages but was also linked to declines in private-sector wages, land prices, and increases in the youth unemployment rate.[46] While such potential side effects were recognized by some policymakers (see, for example, House of Representatives 2015), we provide suggestive quantification of them. We briefly present the results in the main text, and Appendix E provides the technical details.

We first calculate the fiscal benefit of the policy reform. In line with the Japanese government's estimate (House of Representatives 2015), we estimate that other things being equal, the public wage reform was estimated to reduce local public wages by approximately 200 billion yen in 2010.

However, as our analyses show, the effect of this reform was not limited to reductions in public wages but was also linked to declines in private-sector wages and land prices and increases in the youth unemployment rate. Considering the differences in public wage changes by region and the results of our estimates, we calculate that at the national level, the public wage reform would have reduced the wages of young full-time workers by approximately 111.6 billion yen, reduced the taxable assessed value of land by approximately 193.9 billion yen, and increased youth unemployment by approximately 10,470 in 2010. Taken together, when evaluating the public-sector wage cut as a policy, it is important to recognize the possibility that the fiscal benefit of 200 billion yen might accompany substantial negative effects on the private wages and unemployment of young workers and on land prices.[47]

[46] Note that the national-level impact on the population was nearly zero since international migration is negligible in our context. For example, approximately 0.0001% of the total Japanese population moved out of Japan from 2009 to 2010.

[47] The presence of the personal income tax and residential tax on labor income and the property tax on land implies that the actual fiscal benefit could be significantly less than 200 billion yen. In particular, assuming the marginal income tax rate of 10%, the residential tax rate of 10%, and the property tax rate of 1.4% based on the Japanese tax system, the income tax revenue decreases by approximately 40 billion yen and the property tax revenue by 2.7 billion yen, implying that the actual fiscal benefit might be approximately 157.3 billion yen. While this back-of-the-envelope calculation is incomplete because it misses other taxes (e.g., corporate income taxes), it illustrates the importance of accounting for the side effects of the public wage reform when estimating its fiscal benefits.

7 Conclusion

How do institutional wage reforms in one sector spill over to other sectors? This paper studies the spillover effects of a public-sector wage cut on private-sector wages as a prominent example of institutional reforms that could induce wage spillovers across sectors. We leverage the Japanese policy reform that cut public-sector wages only in certain municipalities and the institutional setting in which only young workers are eligible for public-sector jobs. We find that a 1% decrease in public-sector wages reduces the wages of young workers by 0.3%, with larger spillovers in municipalities with a greater share of public workers. A 1% decrease in public-sector wages also reduced the young population in affected municipalities by 0.4%, suggesting a negative welfare effect on young workers in spatial equilibrium and a downward shift in the labor demand for young workers. Overall, we have uncovered an unintended side effect of the public-sector wage cut: the decline in the private-sector wages and welfare for young workers. Such spillover effects should be taken into account when evaluating institutional wage reforms. More broadly, this highlights the importance of considering spillover effects in evaluating reforms in institutional wage rules, including minimum wages, anti-union laws, and equal-pay requirements across different geographical areas.

We close this paper with the reminder that since our identification is based on the local variation in public-sector wages within a country, the effect size may be different when the policy reform is at the national level. This is an inherent problem in recent empirical studies that exploit cross-sectional variation to shed light on macroeconomic issues. A carefully designed structural model may be useful in linking the effects of local policy changes to the effects of national changes (Nakamura and Steinsson 2018). Using quasi-experimental variation to infer national-level public wage reform in the context of public-sector wages is an important next step.

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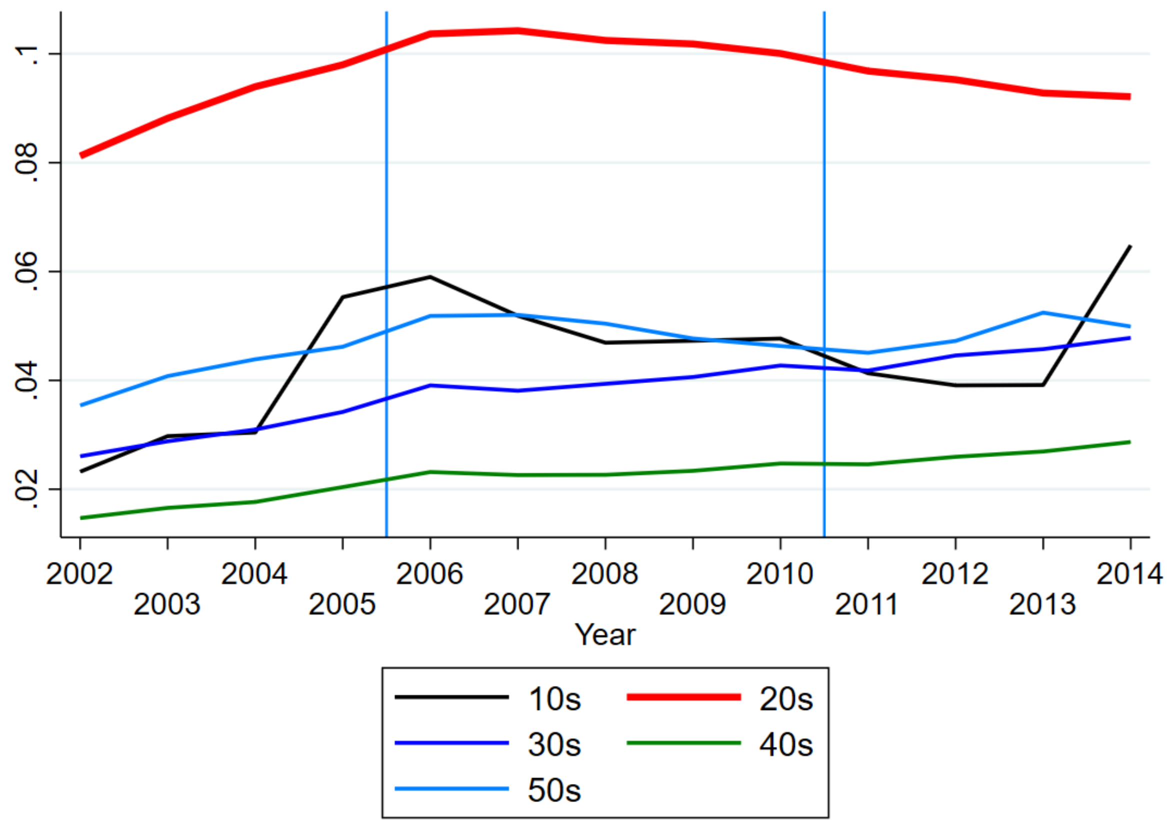
Appendix to "Wage Spillovers across Sectors: Evidence from a Localized Public-sector wage Cut" (Not for Publication)

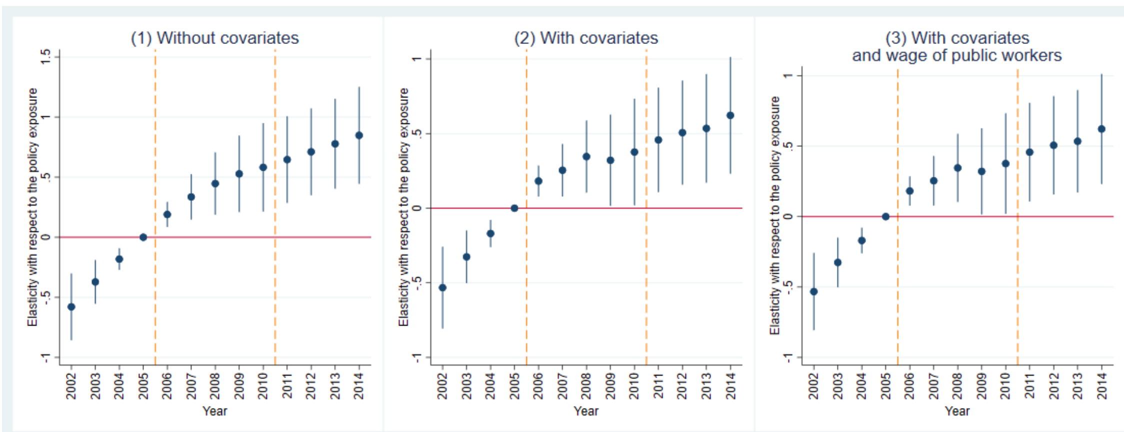
A Omitted Figures

Figure A.1: Trends in the turnover rate of local government employees by age group



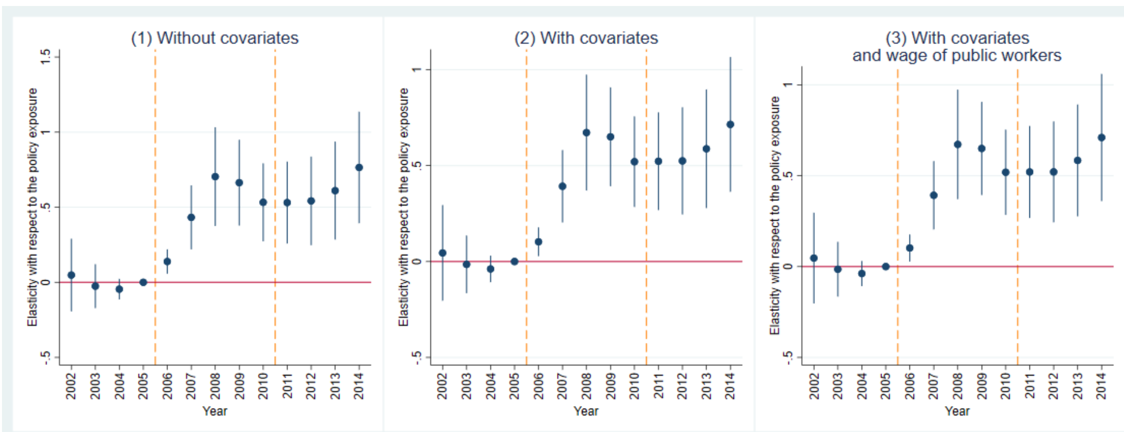
Note: This figure shows the turnover rate of local public-sector workers by age group. The turnover rate is calculated as the number of employees leaving the local government divided by the number of local government employees in each age group. The turnover rate for workers in their 60s is omitted from the figure because it exceeds 2, which would collapse the figure.

Figure A.2: The elasticity of the ratio of the young and old unemployment rates in year t with respect to the regional allowance change during 2006-2010



Note: These figures show the estimated β_t for each year in the eq.(5), where the dependent variable is the ratio of the young and old unemployment rates. This corresponds to the elasticity of the ratio of the unemployment rate for 15-29-year-olds and the unemployment rate for 30-64 year olds in year t with respect to the change in regional allowances during 2006-2010, when the reform of regional allowances was conducted. The unemployment rate is a linear interpolation of five-year census data to one-year data. 95% confidence intervals based on standard errors clustered at the municipality level are also shown. In all panels, we control municipality, year, and prefecture-year fixed effects. In panel (1), we do not control any other variables. We control covariates (the logarithms of local tax revenue per capita, LAT per capita, NTD per capita, and the number of local officials per capita) in panels (2) and (3). We add the average salary level of public workers in municipality j at t, $w_{j,t,public}$ in panel (3).

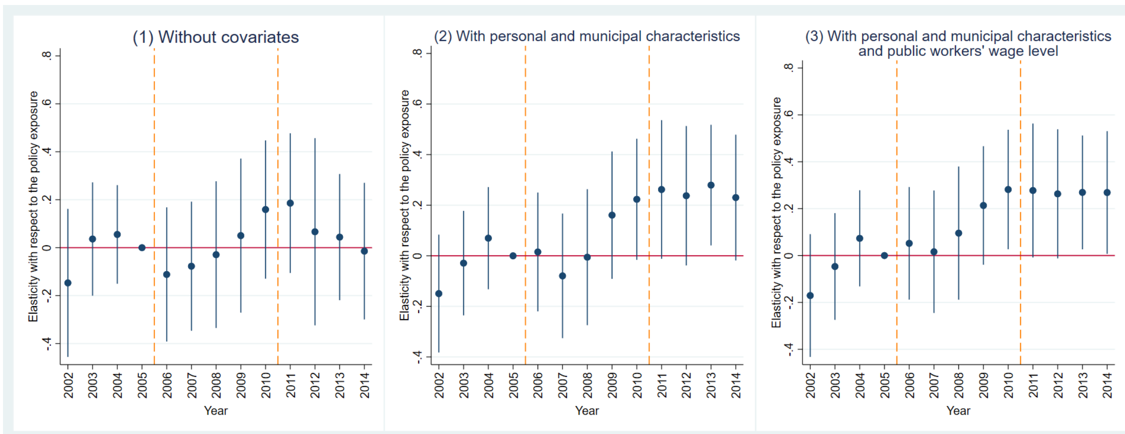
Figure A.3: The elasticity of land prices in year t with respect to the change in regional allowances during 2006-2010



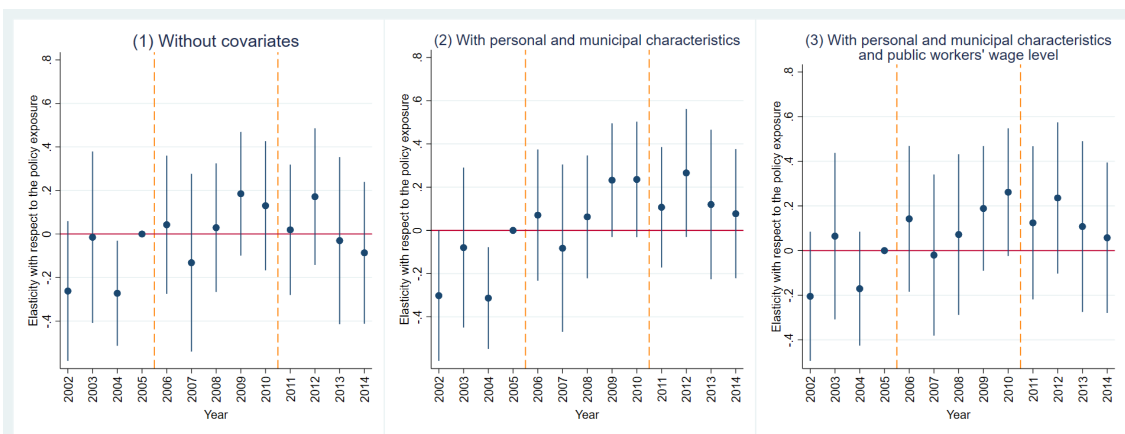
Note: These figures show the estimated β_t for each year in the eq.(5), with the land price as the dependent variable. This corresponds to the elasticity of the land price in year t with respect to the change in regional allowances during 2006-2010, when the reform of regional allowances was conducted. In all panels, we control municipal, year, and municipal-year fixed effects. In panel (1), we do not control any other variables. We control covariates (the logarithms of local tax revenue per capita, LAT per capita, NTD per capita, and the number of local officials per capita) in panels (2) and (3). We add the average salary level of public workers in municipality j at t, $w_{j,t,public}$ in panel (3). 95% confidence intervals based on standard errors clustered at the municipality level are also shown.

Figure A.4: The elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010 for men and women

(a) Analysis for men



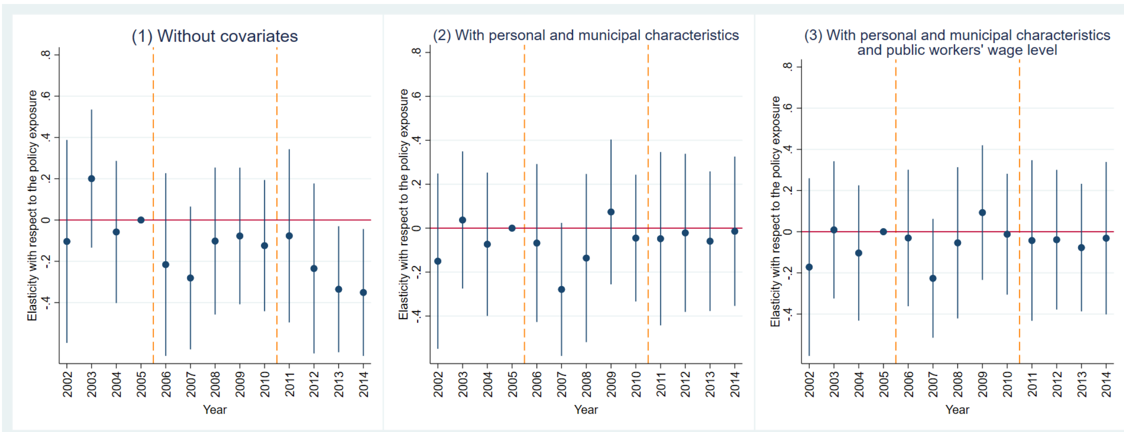
(b) Analysis for women



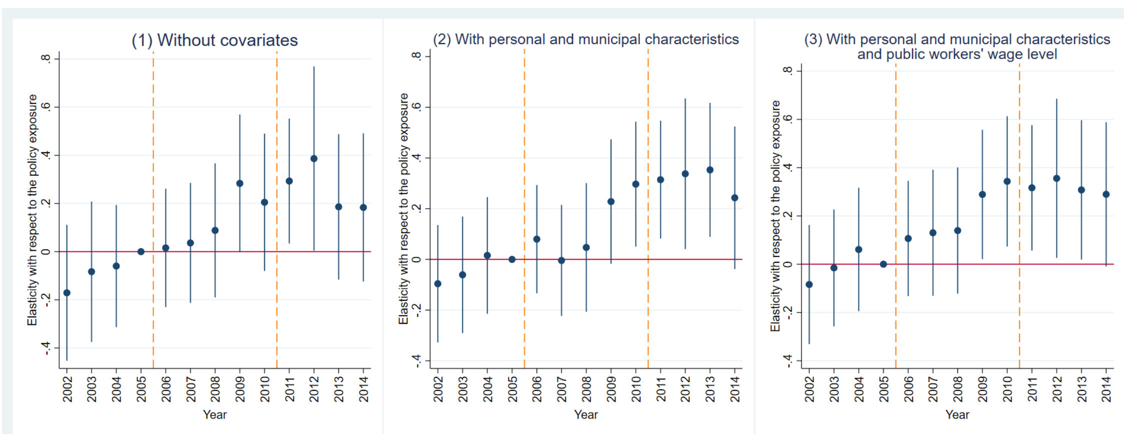
Note: These figures show the estimated β_t for each year in the eq.(3). This corresponds to the elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010, when the reform of regional allowances was conducted. We restrict the sample to men (women) in figure (a) ((b)). 95% confidence intervals based on standard errors clustered at the municipality level are also shown. In all panels, we control municipal, year, and municipal-year fixed effects. In panel (1), we do not control any other variables. In panels (2)-(3), we control for individual (age dummies interacted with college education dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita.). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in panel (3).

Figure A.5: The elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010

(a) Analysis for university graduates



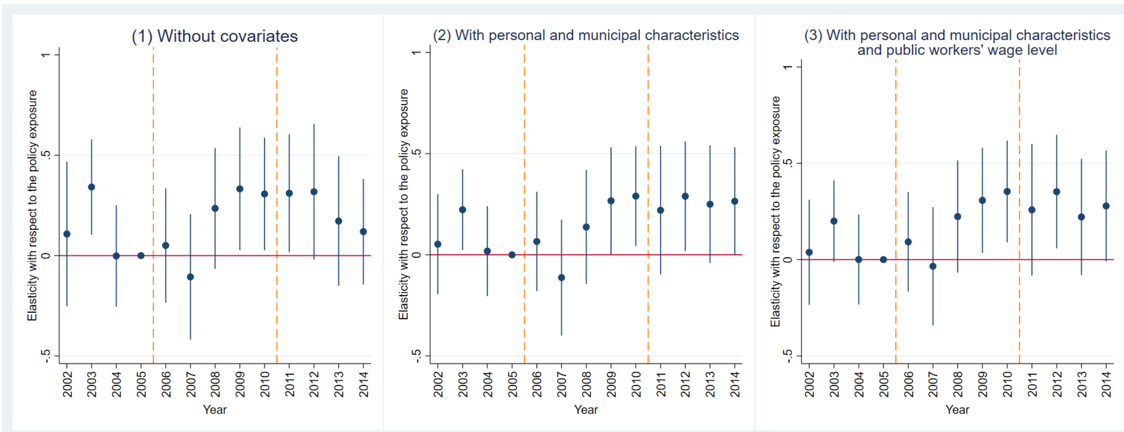
(b) Analysis for non-university graduates



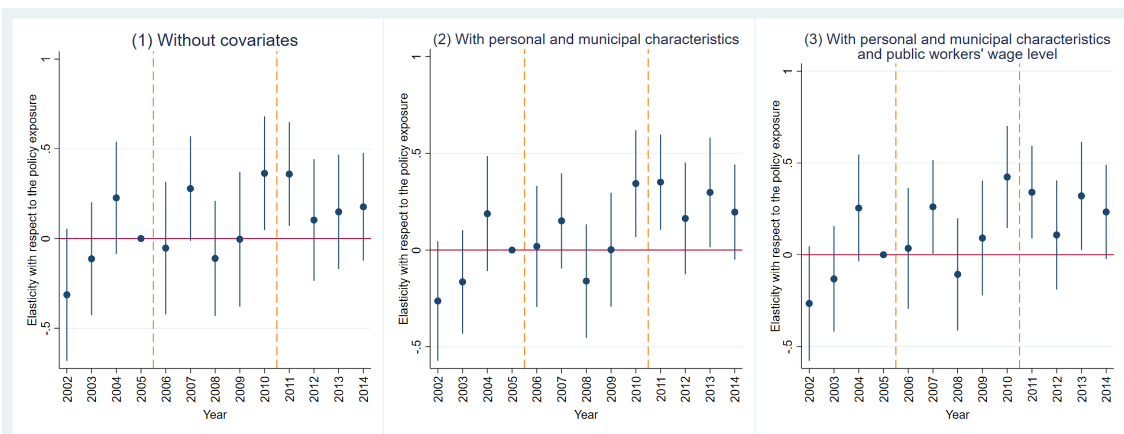
Note: These figures show the estimated β_t for each year in the eq.(3). This corresponds to the elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010, when the reform of regional allowances was conducted. We restrict the sample to (non)university graduates in figure (a) ((b)). 95% confidence intervals based on standard errors clustered at the municipality level are also shown. In all panels, we control municipal, year, and municipal-year fixed effects. In panel (1), we do not control any other variables. In panels (2)-(3), we control for individual (age dummies interacted with the gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in panel (3).

Figure A.6: The elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010

(a) Analysis for workers belonging to large companies (with over 100 employees)



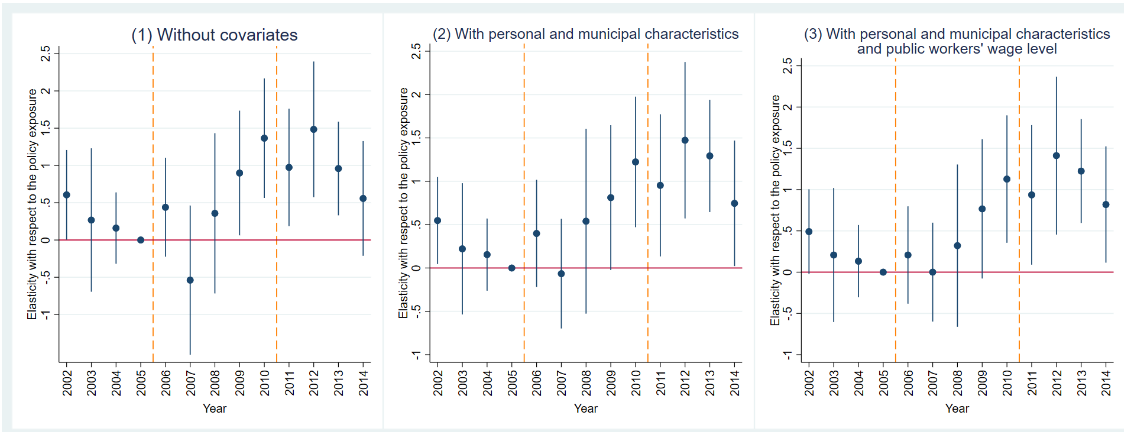
(b) Analysis for workers belonging to small companies (with fewer than 100 employees)



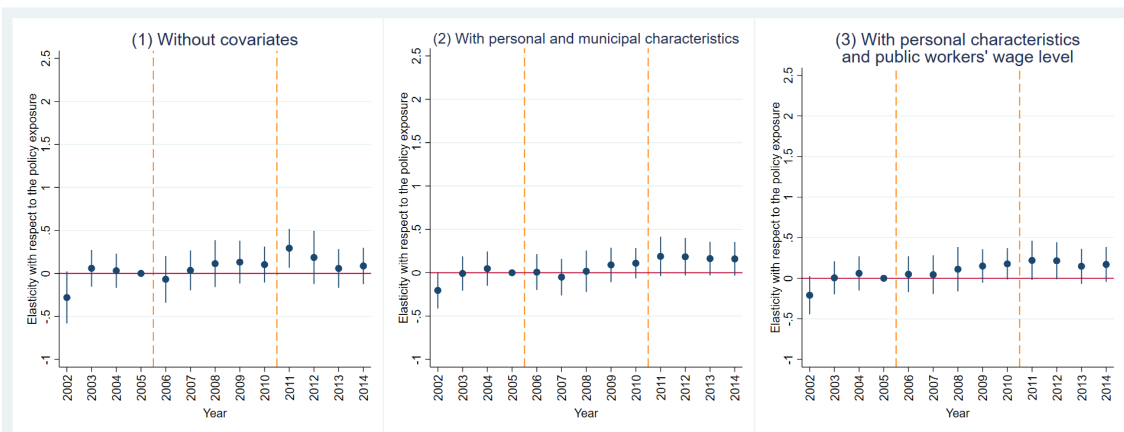
Note: These figures show the estimated β_t for each year in the eq.(3). This corresponds to the elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010, when the reform of regional allowances was conducted. We restrict the sample to workers belonging to companies with more (less) than 100 employees in figure (a) ((b)). 95% confidence intervals based on standard errors clustered at the municipality level are also shown. In all panels, we control municipal, year, and municipal-year fixed effects. In panel (1), we do not control any other variables. In panels (2)-(3), we control for individual characteristics (age dummies interacted with college education dummy and gender dummy and the logarithm of prescribed working hours). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in panel (3).

Figure A.7: The elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010

(a) Analysis for industries with a large inflow of young labor from the public sector compared to the number of workers



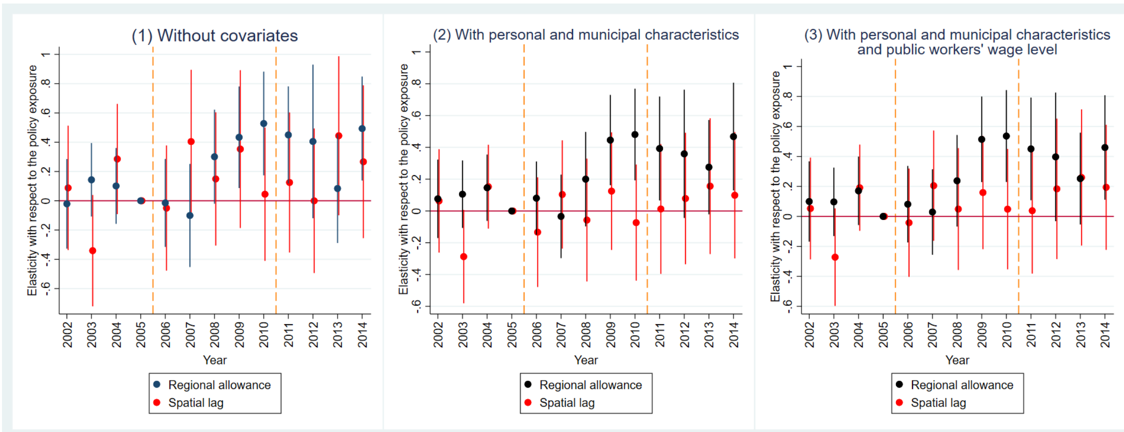
(b) Analysis for industries with a small inflow of young labor from the public sector compared to the number of workers



Note: These figures show the estimated β_t for each year in the eq.(3). This corresponds to the elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010, when the reform of regional allowances was conducted. For analysis regarding the industries with high labor inflow from the public sector, we restrict the sample to the workers belongs to the top three industries (Electricity, Gas, Heat Supply and Water; Transport and Postal Services; Real Estate and Goods Rental and Leasing) with relatively high inflows of workers from the public sector from 2009 to 2014[a]. For the analysis for the industries with low labor inflow from the public sector, we restrict the sample to workers belonging to industries other than the top three industries. 95% confidence intervals based on standard errors clustered at the municipality level are also shown. In all panels, we control municipal, year, and municipal-year fixed effects. In panel (1), we do not control any other variables. In panels (2)-(3), we control for individual (age dummies interacted with college education dummy and gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in panel (3).

[a]Note that due to the Great East Japan Earthquake, the data of 2011 are lacking here.

Figure A.8: The elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010 estimated based on the SLX model



Note: These figures show the estimated β_t (the black one) and β_t^{SLX} (the red one) for each year in the equation

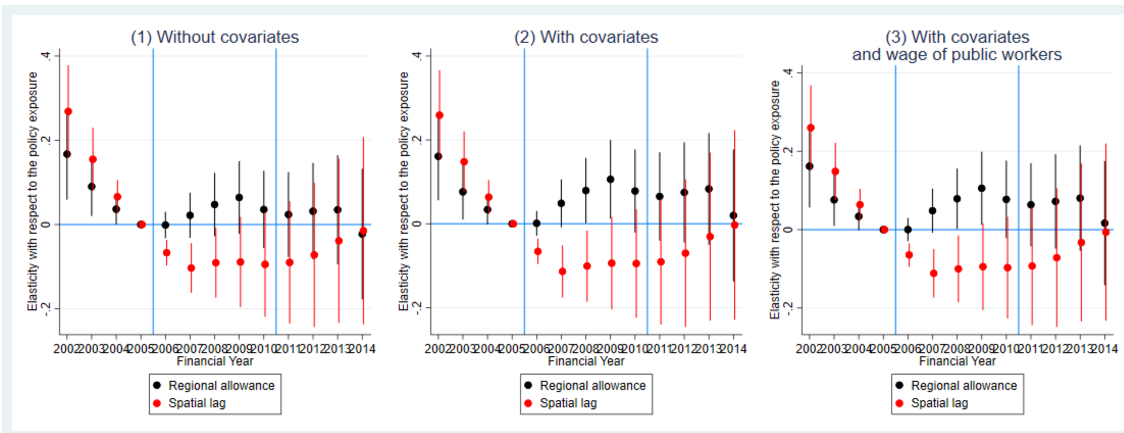
$$
\begin{aligned}
\ln w_{i,j,t,private} = & \sum_{t \neq 2005} [\beta_t \{\tau_t \times (RA_{j,2010} - RA_{j,2005})\} + \beta_t^{SLX} \{\tau_t \times (\boldsymbol{W}RA_{j,2010} - \boldsymbol{W}RA_{j,2005})\}] \times Young_i \\
& + \mu_{j,t} + \sum_{k=\text{young or old}} (\iota_j^k + \tau_t^k) + \gamma X_{i,j,t} + \epsilon_{i,t}.
\end{aligned} \tag{A.1}
$$

β_t and β_t^{SLX}, respectively, correspond to the elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010 and with respect to the spatial lag of the change in regional allowances, when the reform of regional allowances was conducted. The spatial lag is made based on the exponential type of spatial weight matrix whose element (i, j) is

$$
\omega_{i,j} = \begin{cases} \frac{\exp(-\delta d_{i,j})}{\sum_{j=1}^{n} \exp(-\delta d_{i,j})}, & \text{if } d_{i,j} < d, i \neq j, \delta > 0 \\ 0 & \text{otherwise,} \end{cases} \tag{A.2}
$$

where we set decay parameter δ as 1.2 kilometers by using the `spgen` command in Stata (See Kondo 2016 for details). 95% confidence intervals based on standard errors clustered at the municipality level are also shown. In panel (1), we do not control any other variables. In panels (2)-(3), we control for individual (age dummies interacted with college education dummy and gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in panel (3).

Figure A.9: The elasticity of the ratio of the young to the older population in year t with respect to the change in regional allowances during 2006-2010 estimated based on the SLX model



Note: These figures show the estimated β_t (in black) and β_t^{SLX} (in red) for each year in the equation

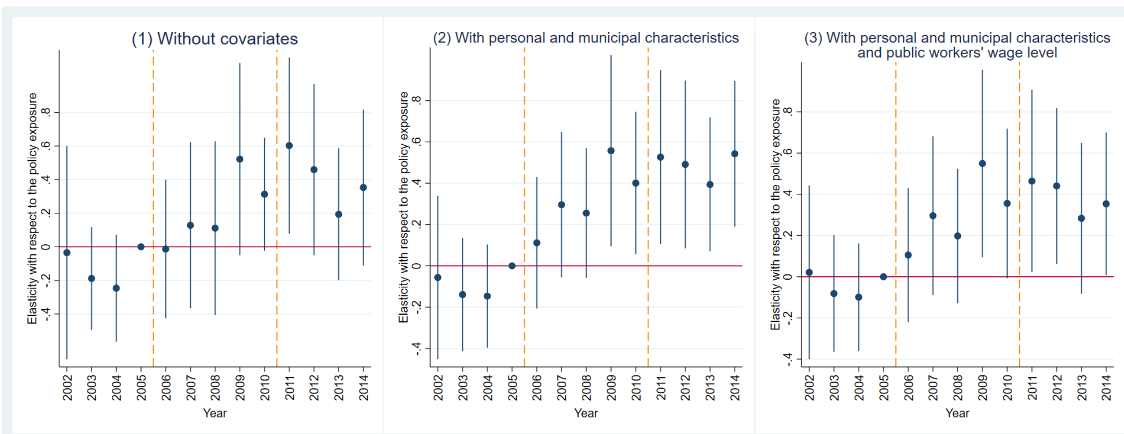
$$\ln(YOratio)_{j,t} = \sum_{t \neq 2005} [\beta_t\{\tau_t \times (RA_{j,2010} - RA_{j,2005})\} + \beta_t^{SLX}\{\tau_t \times (\boldsymbol{W}RA_{j,2010} - \boldsymbol{W}RA_{j,2005})\}] + \iota_j + \eta_{t,p} + \gamma X_{j,t} + \epsilon_{j,t}. \tag{A.3}$$

β_t and β_t^{SLX}, respectively, correspond to the elasticity of the ratio of the young to the older population in year t with respect to the change in regional allowances during 2006-2010 and with respect to the spatial lag of the change in regional allowances, when the reform of regional allowances was conducted. The spatial lag is made based on the exponential type of spatial weight matrix whose element (i, j) is

$$\omega_{i,j} = \begin{cases} \frac{\exp(-\delta d_{i,j})}{\sum_{j=1}^{n} \exp(-\delta d_{i,j})}, & \text{if } d_{i,j} < d, i \neq j, \delta > 0 \\ 0 & \text{otherwise,} \end{cases} \tag{A.4}$$

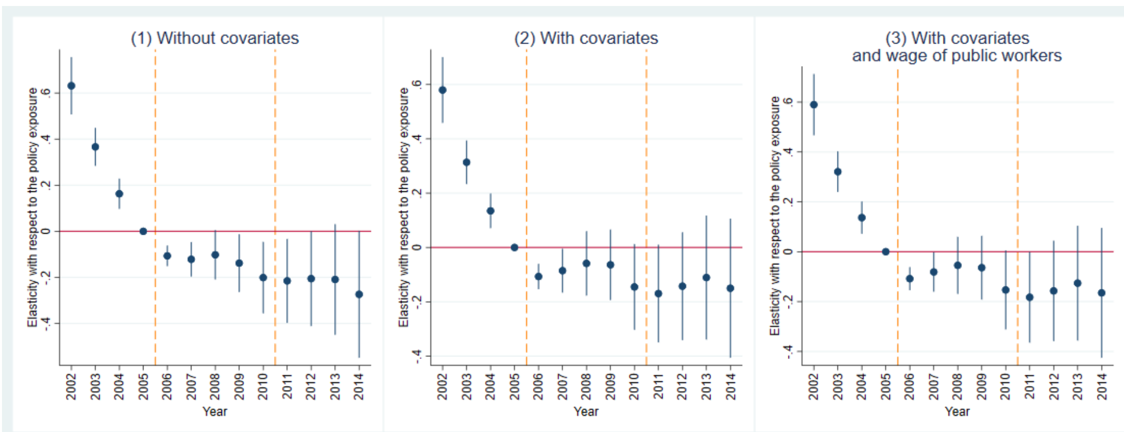
where we set decay parameter δ as 1.2 kilometers by using the `spgen` command in Stata (See Kondo 2016 for details). 95% confidence intervals based on standard errors clustered at the municipality level are also shown. In panel (1), we do not control any other variables. In panels (2)-(3), we control for the covariates regarding municipalities (the logarithms of local tax revenue per capita, LAT per capita, NTD per capita, and the number of local officials per capita). We add the average salary level of public workers in municipality j at t, $w_{j,t,public}$ in panel (3). 95% confidence intervals based on standard errors clustered at the municipality level are also shown.

Figure A.10: The elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010 for regions other than the three largest metropolitan areas



Note: These figures show the estimated β_t for each year in the eq.(3). This corresponds to the elasticity of the private-sector wage in year t with respect to the change in regional allowances during 2006-2010, when the reform of regional allowances was conducted. We restrict the sample to regions other than three metropolitan areas (Tokyo, Kanagawa, Chiba, Ibaraki, Saitama, Aichi, Mie, Osaka, Kyoto, Nara, and Hyogo prefectures), which the laws for the respective regions designate as the three metropolitan areas, in figure. 95% confidence intervals based on standard errors clustered at the municipality level are also shown. In all panels, we control municipal, year, and municipal-year fixed effects. In panel (1), we do not control any other variables. In panels (2)-(3), we control for individual (age dummies interacted with college education dummy and gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to the private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in panel (3).

Figure A.11: The elasticity of the ratio of the young to the older population in year t with respect to the change in regional allowances during 2006-2010 for regions other than the three largest metropolitan areas



Note: These figures show the estimated β_t for each year in the eq.(5), where the ratio of the young to the older population is the dependent variable. We restrict the sample to regions other than three metropolitan areas (Tokyo, Kanagawa, Chiba, Ibaraki, Saitama, Aichi, Mie, Osaka, Kyoto, Nara, and Hyogo prefectures), which the laws for the respective regions designate as the three metropolitan areas, to make figure. In all panels, we control municipal, year, and municipal-year fixed effects. In panel (1), we do not control any other variables. We control covariates (the logarithms of local tax revenue per capita, LAT per capita, NTD per capita, and the number of local officials per capita) in panels (2) and (3). We add the average salary level of public workers in municipality j at t, $w_{j,t,public}$ in panel (3). 95% confidence intervals based on standard errors clustered at the municipality level are also shown.

B Omitted Tables

Table B.1: Regression results on private-sector wages based on (2) for the regions with a higher- and lower-than-median share of public employees

	(1)	(2)	(3)	(4)	(5)	(6)
	(a) Region with many public workers			(b) Region with few public workers		
	log(wage rate of private workers)					
Regional allowances × Young dummy	0.3401***	0.3515***	0.4116***	0.1858**	0.2534***	0.2671***
	(0.0983)	(0.1206)	(0.1113)	(0.0925)	(0.0679)	(0.0664)
log(base wage of local municipal workers)			0.1069***			0.0505***
			(0.0175)			(0.0088)
Year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipality fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipality-year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Individual and municipal fiscal characteristics	No	Yes	Yes	No	Yes	Yes
N	6851803	6851803	6540347	7248851	7248851	6947525
R^2	0.307	0.546	0.540	0.213	0.472	0.466

Standard errors clustered at a municipal level in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: The regression results of estimating equation (2) are presented. We restrict the sample to regions with a higher-than-median (lower-than-median) share of public employees in 2005 in panel (a) ((b)). In columns (2), (3), (5), and (6), we control for individual (age dummies interacted with college education dummy and gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in columns (3) and (6) although it is not controlled in the other columns. Note that the sample sizes of columns (3) and (6) are limited because samples lacking public workers corresponding to the municipal public workers in terms of education and experience are omitted.

Table B.2: Regression results on the ratio of the young to the older population based on (4) for the regions with a higher- and lower-than-median share of public employees

	(1)	(2)	(3)	(4)	(5)	(6)
	Region with many public workers					
	log(YO ratio of Pop)			$\widetilde{\text{log(YO ratio of Pop)}}$		
Regional allowances	-0.1303	-0.1368	-0.1461	0.7781***	0.6597**	0.6081**
	(0.1628)	(0.1666)	(0.1656)	(0.2467)	(0.2660)	(0.2575)
log(tax revenue per capita)		-0.0113	-0.0099		-0.1086***	-0.1007***
		(0.0148)	(0.0147)		(0.0257)	(0.0254)
log(LAT per capita)		-0.0016	-0.0013		0.0033	0.0049
		(0.0060)	(0.0059)		(0.0079)	(0.0079)
log(NTD per capita)		0.0058**	0.0059**		0.0045	0.0046
		(0.0027)	(0.0027)		(0.0045)	(0.0044)
log(municipal public workers per capita)		0.0062	0.0060		-0.0364	-0.0374
		(0.0116)	(0.0117)		(0.0231)	(0.0240)
log(base wage of local municipal workers)			-0.0372			-0.2076***
			(0.0372)			(0.0694)
Year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipal fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Detrended outcome variable	No	No	No	Yes	Yes	Yes
N	11245	11245	11245	11245	11245	11245
R^2	0.968	0.968	0.968	1.000	1.000	1.000

Standard errors in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

	(1)	(2)	(3)	(4)	(5)	(6)
	Region with few public workers					
	log(YO ratio of Pop)			$\widetilde{\text{log(YO ratio of Pop)}}$		
Regional allowances	-0.0276	0.0168	0.0154	0.4071***	0.3718***	0.3552***
	(0.0559)	(0.0574)	(0.0575)	(0.0987)	(0.0994)	(0.0985)
log(tax revenue per capita)		0.0793***	0.0800***		-0.0376	-0.0292
		(0.0296)	(0.0298)		(0.0339)	(0.0332)
log(LAT per capita)		0.0049**	0.0049**		0.0054	0.0050
		(0.0023)	(0.0023)		(0.0035)	(0.0036)
log(NTD per capita)		-0.0028	-0.0026		-0.0114***	-0.0091**
		(0.0032)	(0.0032)		(0.0043)	(0.0041)
log(municipal public workers per capita)		0.0132	0.0123		-0.0542***	-0.0644***
		(0.0126)	(0.0126)		(0.0189)	(0.0180)
log(base wage of local municipal workers)			-0.0240			-0.2785***
			(0.0403)			(0.0583)
Year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipal fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture-year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Detrended outcome variable	No	No	No	Yes	Yes	Yes
N	11258	11258	11258	11258	11258	11258
R^2	0.970	0.970	0.970	1.000	1.000	1.000

Standard errors in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

The regression results of estimating equation (4) in which the logarithm of the ratio of the young to the older population, log(YO ratio of Pop), and its detrended variable, $\widetilde{\text{log(YO ratio of Pop)}}$, are the dependent variables are presented. Young and older populations correspond to the 15-29 and 30-64 age groups, respectively. $\widetilde{\text{log(YO ratio of Pop)}}$ is the value that detrends the 2002-2005 region-specific linear trends from log(YO ratio of Pop). Specifically, to construct the detrended outcome variable $\widetilde{\text{log(YO ratio of Pop)}}_{jt}$, we first estimate the municipality-specific pretrend by the following equation

$$\ln(\text{YO ratio of Pop})_{jt} = \iota_j + \tau_t + \iota_j \times \tau_t + \epsilon_{j,t} \quad \text{for} \quad t < 2006.$$

We make predicted values of $\ln(\text{YO ratio of Pop})_{jt}$ for entire samples, including $t \geq 2006$, from this result and compute the residuals corresponding to $\widetilde{\ln(\text{YO ratio of Pop})}_{jt}$.

The average salary level of public workers in municipality j at t, $w_{j,t,public}$, is controlled in columns (3) and (6), although it is not controlled in the other columns because it is likely to suffer from endogeneity.

Table B.3: Regression results on the ratio of the young unemployment rate and the old unemployment rate based on (4)

	(1)	(2)	(3)	(4)	(5)	(6)
	YO ratio of unemployment rate			$\widetilde{\text{YO ratio of unemployment rate}}$		
Regional allowances	0.8607***	0.6609***	0.6297***	-0.5124	-0.4945	-0.4914
	(0.1467)	(0.1426)	(0.1410)	(0.3701)	(0.3564)	(0.3561)
log(tax revenue per capita)		-0.1870***	-0.1731***		0.1516*	0.1502*
		(0.0395)	(0.0390)		(0.0860)	(0.0858)
log(LAT per capita)		-0.0091**	-0.0091**		-0.0260**	-0.0260**
		(0.0044)	(0.0043)		(0.0102)	(0.0102)
log(NTD per capita)		-0.0257***	-0.0244***		-0.0043	-0.0044
		(0.0056)	(0.0055)		(0.0131)	(0.0131)
log(municipal public workers per capita)		-0.0549*	-0.0616**		0.0387	0.0394
		(0.0287)	(0.0280)		(0.0677)	(0.0684)
log(base wage of local municipal workers)			-0.3077***			0.0305
			(0.0729)			(0.1741)
N	22479	22479	22479	22479	22479	22479
R^2	0.904	0.905	0.906	1.000	1.000	1.000

Standard errors in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

The table presents the regression results of estimating equation (4), in which the ratio of the unemployment rate among those aged 15-29 and among those aged 30-64, YO ratio of Pop, and its detrended variable, $\widetilde{\text{YO ratio of unemployment rate}}$, are the dependent variables. The young and older populations correspond to the 15-29 and 30-64 age groups, respectively. $\widetilde{\text{YO ratio of Unemployment rate}}$ is the value that detrends the 2002-2005 region-specific linear trends from the YO ratio of Unemployment rate. Specifically, to construct the detrended outcome variable $\widetilde{\text{YO ratio of Unemployment rate}}_{jt}$, we first estimate the municipality-specific pretrend by the following equation

$$\text{YO ratio of Unemployment rate}_{jt} = \iota_j + \tau_t + \iota_j \times \tau_t + \epsilon_{jt} \quad \text{for } t < 2006.$$

We make predicted values of YO ratio of Unemployment rate$_{jt}$ for entire samples, including $t \geq 2006$, from this result and compute the residuals corresponding to $\widetilde{\text{YO ratio of Unemployment rate}}_{jt}$.

The average salary level of public workers in municipality j at t, $w_{j,t,public}$, is controlled in columns (3) and (6), although it is not controlled in the other columns because it is likely to suffer from endogeneity.

Table B.4: Regression results on the land price based on (4)

	(1)	(2)	(3)
		log(land prices)	
Regional allowances	0.5677***	0.5423***	0.5395***
	(0.1573)	(0.1421)	(0.1413)
log(tax revenue per capita)		0.0909**	0.0924**
		(0.0370)	(0.0367)
log(LAT per capita)		-0.0177***	-0.0177***
		(0.0052)	(0.0052)
log(NTD per capita)		-0.0048	-0.0047
		(0.0046)	(0.0046)
log(municipal public workers per capita)		-0.0321	-0.0329
		(0.0225)	(0.0227)
log(base wage of local municipal workers)			-0.0322
			(0.0679)
Year fixed effect	Yes	Yes	Yes
Municipal fixed effect	Yes	Yes	Yes
Prefecture-year fixed effect	Yes	Yes	Yes
N	17615	17615	17615
R^2	0.922	0.924	0.924

Standard errors clustered at the municipal level in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

The table presents the regression results of estimating equation (4), in which the land price is the dependent variable. The average salary level of public workers in municipality j at t, $w_{j,t,public}$, is controlled in column (3), although it is not controlled in the other columns because it is likely to suffer from endogeneity.

Table B.5: Regression results on private-sector wages based on (2) for men and women

	(1)	(2)	(3)	(4)	(5)	(6)
		(a) Men			(b) Women	
			log(wage rate of private workers)			
Regional allowances × Young dummy	0.0806	0.2498***	0.2863***	0.1706*	0.2419**	0.2400**
	(0.0870)	(0.0682)	(0.0655)	(0.0928)	(0.0987)	(0.0994)
log(base wage of local municipal workers)			0.1603***			-0.1191***
			(0.0077)			(0.0186)
Year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipality fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipality-year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Individual and municipal fiscal characteristics	No	Yes	Yes	No	Yes	Yes
N	9788905	9788905	9408084	4311241	4311241	4079192
R^2	0.284	0.498	0.491	0.249	0.354	0.34

Standard errors clustered at the municipal level in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: The regression results of estimating equation (2) are presented. We restrict the sample to regions with a higher-than-median (lower-than-median) share of public employees in 2005 in panel (a) ((b)). In columns (2), (3), (5), and (6), we control for individual (age dummies interacted with the college education dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in columns (3) and (6), although it is not controlled in the other columns. Note that the sample size of columns (3) and (6) is limited because samples lacking public workers corresponding to the municipal public workers in terms of education and experience are omitted.

Table B.6: Regression results on private-sector wages based on (2) for university graduates and non-university graduates

	(1)	(2)	(3)	(4)	(5)	(6)
	(a) University graduates			(b) Non-university graduates		
	log(wage rate of private workers)					
Regional allowances × Young dummy	-0.2414**	-0.0015	0.0380	0.3313***	0.3255***	0.3251***
	(0.0956)	(0.0807)	(0.0768)	(0.0754)	(0.0695)	(0.0727)
log(base wage of local municipal workers)			0.3370***			-0.0297***
			(0.0055)			(0.0073)
Year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipality fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipality-year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Individual and municipal fiscal characteristics	No	Yes	Yes	No	Yes	Yes
N	4286262	4286262	4279137	9813339	9813339	9207675
R^2	0.313	0.525	0.528	0.221	0.411	0.404

Standard errors clustered at the municipal level in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: The regression results of estimating equation (2) are presented. We restrict the sample to (non)university graduates in panel (a) ((b)). In columns (2), (3), (5), and (6), we control for individual (age dummies interacted with the gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in columns (3) and (6), although it is not controlled in the other columns. Note that the sample size of columns (3) and (6) is limited because samples lacking public workers corresponding to the municipal public workers in terms of education and experience are omitted.

Table B.7: Regression results on private-sector wages based on (2) for workers in the industries with high and low exposure to labor flow from the public sector

	(1)	(2)	(3)	(4)	(5)	(6)
	(a) Workers belonging to industries with high exposure			(b) Workers belonging to industries with low exposure		
	log(wage rate of private workers)					
Regional allowances × Young dummy	0.6274**	0.7478**	0.7491**	0.1943***	0.2100***	0.2399***
	(0.2684)	(0.3127)	(0.3042)	(0.0596)	(0.0434)	(0.0435)
log(base wage of local municipal workers)			0.1115***			0.0779***
			(0.0105)			(0.0110)
Year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipality fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipality-year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Individual and municipal fiscal characteristics	No	Yes	Yes	No	Yes	Yes
N	1930518	1930518	1863507	12170022	12170022	11624202
R^2	0.282	0.456	0.454	0.276	0.533	0.527

Standard errors clustered at the municipal level in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: The regression results of estimating equation (2) are presented. We restrict the sample to the workers in industries with high (low) exposure to the labor flow from the public sector in panel (a) ((b)), which is determined by whether the industry is one of the top three industries with the highest exposure (Electricity, Gas, Heat Supply and Water; Transport and Postal Services; Real Estate and Goods Rental and Leasing) since 2009. In columns (2), (3), (5), and (6), we control for individual (age dummies interacted with the college education dummy and gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in columns (3) and (6), although it is not controlled in the other columns. Note that the sample size of columns (3) and (6) is limited because samples lacking public workers corresponding to the municipal public workers in terms of education and experience are omitted.

Table B.8: Regression results on private-sector wages based on (2) for workers belonging to large and small companies

	(1)	(2)	(3)	(4)	(5)	(6)
	(a) Workers belonging to large companies			(b) Workers belonging to small companies		
	log(wage rate of private workers)					
Regional allowances × Young dummy	0.2352***	0.2453***	0.2237***	0.1399	0.2288***	0.2813***
	(0.0697)	(0.0753)	(0.0737)	(0.0799)	(0.0679)	(0.0652)
log(base wage of local municipal workers)			0.0397***			0.1106***
			(0.0077)			(0.0113)
Year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipality fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Municipality-year fixed effect	Yes	Yes	Yes	Yes	Yes	Yes
Individual and municipal fiscal characteristics	No	Yes	Yes	No	Yes	Yes
N	4914582	4914582	4635093	9185985	9185985	8852654
R^2	0.260	0.441	0.435	0.271	0.541	0.534

Standard errors clustered at the municipal level in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: The regression results of estimating equation (2) are presented. We restrict the sample to workers belonging to companies with more (fewer) than 100 employees in panel (a) ((b)). In columns (2), (3), (5), and (6), we control for individual (age dummies interacted with the college education dummy and gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in columns (3) and (6), although it is not controlled in the other columns. Note that the sample size of columns (3) and (6) is limited because samples lacking public workers corresponding to the municipal public workers in terms of education and experience are omitted.

Table B.9: Regression results on private-sector wages based on the SLX model

	(1)	(2)	(3)
	log(wage rate of private workers)		
Regional allowances × Young dummy	0.3544***	0.3110***	0.3396***
	(0.1209)	(0.1069)	(0.1029)
Spatial lag of Regional allowances × Young dummy	0.1796	0.0918	0.1406
	(0.1321)	(0.1160)	(0.1182)
log(base wage of local municipal workers)			0.0736***
			(0.0105)
Year fixed effect	Yes	Yes	Yes
Municipality fixed effect	Yes	Yes	Yes
Municipality-year fixed effect	Yes	Yes	Yes
Individual and municipal fiscal characteristics	No	Yes	Yes
N	12194536	12194536	11668764
R^2	0.264	0.514	0.508

Standard errors clustered at the municipal level in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: These tables show the estimated β and β^{SLX} in the equation

$$\ln w_{i,j,t,private} = [\beta RA_{j,t} + \beta^{SLX}\boldsymbol{W}RA_{j,t}] \times Young_i + \mu_{j,t} + \sum_{k=\text{young or old}} (\iota_j^k + \tau_t^k) + \gamma X_{i,j,t} + \epsilon_{i,t}. \tag{B.1}$$

The results of β and β^{SLX} are shown

The spatial lag is made based on the exponential type of spatial weight matrix whose (i, j) element is

$$\omega_{i,j} = \begin{cases} \frac{\exp(-\delta d_{i,j})}{\sum_{j=1}^{n}\exp(-\delta d_{i,j})}, & \text{if } d_{i,j} < d, i \neq j, \delta > 0 \\ 0 & \text{otherwise,} \end{cases} \tag{B.2}$$

where we set decay parameter δ as 1.2 kilometers by using the `spgen` command in Stata (See Kondo 2016 for details). In columns (2) and (3), we control for individual (age dummies interacted with the college education dummy and gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in column (3), although it is not controlled in the other columns. Note that the sample size of column (3) is limited because samples lacking public workers corresponding to the municipal public workers in terms of education and experience are omitted.

Table B.10: The elasticity of the ratio of the young to the older population in year t with respect to the change in regional allowances during 2006-2010 estimated based on the SLX model

	(1)	(2)	(3)	(4)	(5)	(6)
	log(YO ratio of Pop)			$\widetilde{\text{log(YO ratio of Pop)}}$		
Regional allowances	-0.0441	0.0039	0.0019	0.3590***	0.3151***	0.2890***
	(0.0573)	(0.0590)	(0.0593)	(0.1112)	(0.1114)	(0.1108)
Spatial lag of regional allowances	-0.1608*	-0.1516*	-0.1543*	0.5351***	0.4878***	0.4529***
	(0.0858)	(0.0869)	(0.0874)	(0.1330)	(0.1346)	(0.1362)
log(tax revenue per capita)		0.0517***	0.0528***		-0.0843***	-0.0701***
		(0.0175)	(0.0177)		(0.0239)	(0.0231)
log(LAT per capita)		0.0037*	0.0037*		0.0051	0.0052
		(0.0021)	(0.0021)		(0.0033)	(0.0035)
log(NTD per capita)		0.0011	0.0012		-0.0015	-0.0001
		(0.0021)	(0.0021)		(0.0036)	(0.0035)
log(municipal public workers per capita)		0.0159*	0.0153*		-0.0399**	-0.0473***
		(0.0092)	(0.0092)		(0.0162)	(0.0159)
log(base wage of local municipal workers)			-0.0252			-0.3226***
			(0.0313)			(0.0451)
N	22490	22490	22490	22490	22490	22490
R^2	0.967	0.967	0.967	1.000	1.000	1.000

Standard errors clustered at the municipal level in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: These table show the estimated β and β^{SLX} in the equation

$$Y_{j,t} = \beta RA_{j,t} + \beta^{SLX} \boldsymbol{W} RA_{j,t} + \iota_j + \eta_{t,p} + \gamma X_{j,t} + \epsilon_{j,t}, \tag{B.3}$$

where the logarithm of the ratio of the young to the older population, log(YO ratio of Pop), and its detrended variable, $\widetilde{\text{log(YO ratio of Pop)}}$, are the dependent variables are presented. Young and older populations correspond to the 15-29 and 30-64 age groups, respectively.

$\widetilde{\text{log(YO ratio of Pop)}}$ is the value that detrends the 2002-2005 region-specific linear trends from log(YO ratio of Pop). Specifically, to construct the detrended outcome variable $\widetilde{\text{log(YO ratio of Pop)}}_{jt}$, we first estimate the municipality-specific pretrend by the following equation

$$\ln(\text{YO ratio of Pop})_{jt} = \iota_j + \tau_t + \iota_j \times \tau_t + \epsilon_{j,t} \quad \text{for} \;\; t < 2006.$$

We make predicted values of $\ln(\text{YO ratio of Pop})_{jt}$ for entire samples, including $t \geq 2006$, from this result and compute the residuals corresponding to $\widetilde{\ln(\text{YO ratio of Pop})}_{jt}$.

The spatial lag is made based on the exponential type of spatial weight matrix whose (i, j) element is

$$\omega_{i,j} = \begin{cases} \frac{\exp(-\delta d_{i,j})}{\sum_{j=1}^{n} \exp(-\delta d_{i,j})}, & \text{if} \;\; d_{i,j} < d, i \neq j, \delta > 0 \\ 0 \;\; \text{otherwise}, \end{cases} \tag{B.4}$$

where we set decay parameter δ as 1.2 kilometers by using the `spgen` command in Stata (See Kondo 2016 for details). 95% confidence intervals based on standard errors clustered at the municipality level are also shown.

The average salary level of public workers in municipality j at t, $w_{j,t,public}$, is controlled in columns (3) and (6), although it is not controlled in the other columns because it is likely to suffer from endogeneity.

Table B.11: Regression results on the ratio of the young to the older population based on (4) excluding three metropolitan areas

	(1)	(2)	(3)	(4)	(5)	(6)
	log(YO ratio of Pop)			$\widetilde{\text{log(YO ratio of Pop)}}$		
Regional allowances	-0.4755***	-0.3710***	-0.3850***	1.0637***	0.7348***	0.6270***
	(0.0916)	(0.0828)	(0.0832)	(0.1394)	(0.1369)	(0.1293)
log(tax revenue per capita)		0.0150	0.0172		-0.1832***	-0.1662***
		(0.0123)	(0.0128)		(0.0262)	(0.0254)
log(LAT per capita)		0.0167***	0.0171***		-0.0165***	-0.0140**
		(0.0039)	(0.0039)		(0.0055)	(0.0056)
log(NTD per capita)		0.0033	0.0033		-0.0025	-0.0025
		(0.0020)	(0.0020)		(0.0037)	(0.0036)
log(municipal public workers per capita)		0.0178*	0.0179*		-0.0559***	-0.0545***
		(0.0102)	(0.0102)		(0.0166)	(0.0162)
log(base wage of local municipal workers)			-0.0385			-0.2945***
			(0.0325)			(0.0508)
N	16159	16159	16159	16159	16159	16159
R^2	0.973	0.973	0.973	1.000	1.000	1.000

Standard errors clustered at the municipal level in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

The table presents the regression results of estimating equation (4) in which the logarithm of the ratio of the young to the older population, log(YO ratio of Pop), and its detrended variable, $\widetilde{\text{log(YO ratio of Pop)}}$, are the dependent variables. Young and older populations correspond to the 15-29 and 30-64 age groups, respectively. $\widetilde{\text{log(YO ratio of Pop)}}$ is the value that detrends the 2002-2005 region-specific linear trends from log(YO ratio of Pop). Specifically, to construct the detrended outcome variable $\widetilde{\text{log(YO ratio of Pop)}}_{jt}$, we first estimate the municipality-specific pretrend by the following equation

$$\ln(\text{YO ratio of Pop})_{jt} = \iota_j + \tau_t + \iota_j \times \tau_t + \epsilon_{jt} \quad \text{for } \ t < 2006.$$

We make predicted values of $\ln(\text{YO ratio of Pop})_{jt}$ for entire samples, including $t \geq 2006$, from this result and compute the residuals corresponding to $\widetilde{\ln(\text{YO ratio of Pop})}_{jt}$. We restrict the sample to regions other than three metropolitan areas (Tokyo, Kanagawa, Chiba, Ibaraki, Saitama, Aichi, Mie, Osaka, Kyoto, Nara, and Hyogo prefectures), which the laws for the respective regions designate as the three metropolitan areas. The average salary level of public workers in municipality j at t, $w_{j,t,public}$, is controlled in columns (3) and (6), although it is not controlled in the other columns because it is likely to suffer from endogeneity.

Table B.12: Regression results on private-sector wages for regions excluding three metropolitan areas

	(1)	(2)	(3)
	log(wage rate of private workers)		
Regional allowances × Young dummy	0.5131***	0.4237***	0.4364***
	(0.1253)	(0.1139)	(0.1197)
log(base wage of local municipal workers)			0.0204**
			(0.0084)
Year fixed effect	Yes	Yes	Yes
Municipality fixed effect	Yes	Yes	Yes
Municipality-year fixed effect	Yes	Yes	Yes
Individual and municipal fiscal characteristics	No	Yes	Yes
N	8903248	8903248	8503467
R^2	0.176	0.440	0.433

Standard errors clustered at the municipal level in parentheses
* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Note: The regression results of estimating equation (2) are presented. We restrict the sample to regions other than three metropolitan areas (Tokyo, Kanagawa, Chiba, Ibaraki, Saitama, Aichi, Mie, Osaka, Kyoto, Nara, and Hyogo prefectures), which the laws for the respective regions designate as the three metropolitan areas. In columns (2) and (3), we control for individual (age dummies interacted with the college education dummy and gender dummy and the logarithm of prescribed working hours) and municipal fiscal characteristics (the logarithms of local tax revenue per capita, lump-sum transfer per capita, earmarked subsidies per capita, and the number of local officials per capita). The average salary level of public workers corresponding in terms of education level and experience to private worker i in municipality j at t, $w_{i,j,t,public}$, is controlled in column (3), although it is not controlled in the other columns. Note that the sample size of column (3) is limited because samples lacking public workers corresponding to the municipal public workers in terms of education and experience are omitted.

C Summary Statistics

Table C.1: Summary statistics of the dataset for wage analysis

	Areas without the regional allowances		Areas receiving the regional allowances	
	(1) -2005	(2) 2006-	(3) -2005	(4) 2006-
Panel A. Regional allowances				
Regional allowances	0.00	0.00	0.04	0.08
	(0.00)	(0.00)	(0.05)	(0.05)
Panel B. Individual private workers' characteristics				
Wage of private workers	16.58	16.09	21.08	20.14
	(9.67)	(9.19)	(12.58)	(12.38)
Gender	0.70	0.68	0.74	0.69
	(0.46)	(0.47)	(0.44)	(0.46)
University	0.18	0.22	0.35	0.40
	(0.38)	(0.41)	(0.48)	(0.49)
Age	40.47	41.22	39.74	40.42
	(11.77)	(12.03)	(11.65)	(11.83)
Prescribed working hours	165.71	165.05	161.97	161.74
	(20.23)	(20.25)	(19.61)	(20.69)
Percentage of workers aged less than 30	0.23	0.21	0.24	0.22
	(0.42)	(0.40)	(0.43)	(0.42)
Percentage of workers in companies with over 100 employees	0.44	0.41	0.29	0.29
	(0.50)	(0.49)	(0.45)	(0.45)
Panel C. Municipal characteristics				
Population	1.6e+05	1.7e+05	7.6e+05	7.7e+05
	(1.8e+05)	(1.8e+05)	(8.5e+05)	(8.6e+05)
Local tax revenue per capita	119.97	130.02	170.58	175.80
	(43.83)	(46.30)	(52.69)	(47.59)
LAT per capita	104.27	113.79	27.73	26.45
	(75.24)	(90.11)	(26.61)	(26.47)
NTD per capita	38.29	60.79	45.98	59.77
	(21.48)	(95.29)	(24.37)	(26.88)
Number of municipal public workers	1,631.05	1,542.87	9,212.98	7,737.79
	(1,677.01)	(1,588.27)	(12613.63)	(10022.95)
Average income of corresponding municipal public workers	3,292.17	3,085.41	3,373.80	3,237.96
	(1,105.72)	(1,083.46)	(1,059.09)	(992.97)
N	1806222	4044021	2223787	4120531

The summary statistics of the areas where regional allowances have been provided (at least once) and the areas where regional allowances have not been provided. Means are shown as coefficients, and standard deviations are given in parentheses. Note that the characteristics of urban municipalities are more representative, reflecting the larger sample size in urban areas.

Table C.2: Summary statistics of the dataset for population and unemployment rate analysis

	Areas without the regional allowances		Areas receiving the regional allowances	
	(1) -2005	(2) 2006-	(3) -2005	(4) 2006-
	Panel A. Regional allowances			
Regional allowances	0.00	0.00	0.02	0.06
	(0.00)	(0.00)	(0.04)	(0.05)
	Panel B. Demographic characteristics			
Population	36932.27	35813.29	2.1e+05	2.1e+05
	(60949.07)	(60672.33)	(3.3e+05)	(3.4e+05)
Population aged 15-29	6,439.55	5,410.02	40638.13	35151.25
	(11423.97)	(9,779.60)	(64970.65)	(56893.69)
Population aged 30-64	16958.74	16490.95	1.0e+05	1.1e+05
	(28776.33)	(28785.59)	(1.7e+05)	(1.7e+05)
Unemployment rate for 15-29	0.09	0.10	0.09	0.09
	(0.04)	(0.04)	(0.02)	(0.02)
Unemployment rate for 30-64	0.04	0.05	0.05	0.05
	(0.02)	(0.02)	(0.01)	(0.01)
	Panel C. Administrative characteristics			
Local tax revenue per capita	107.38	120.25	144.26	151.14
	(66.63)	(85.19)	(47.31)	(43.18)
LAT per capita	230.33	264.15	34.68	34.08
	(221.37)	(263.55)	(35.07)	(37.02)
NTD per capita	47.97	75.31	29.17	44.15
	(106.16)	(167.75)	(16.13)	(23.17)
Number of municipal public workers	442.39	398.78	2,019.13	1,804.60
	(613.37)	(568.93)	(4,047.27)	(3,453.39)
Average income of municipal public workers	3,247.07	3,173.34	3,431.02	3,287.83
	(212.19)	(191.96)	(218.37)	(183.57)
N	5464	12294	1460	3285

The summary statistics of the areas where regional allowances have been provided (at least once) and the areas where regional allowances have not been provided. Means are shown as coefficients, and standard deviations are given in parentheses.

Table C.3: Summary statistics of the dataset for land price analysis

	Areas without the regional allowances		Areas receiving the regional allowances	
	(1) -2005	(2) 2006-	(3) -2005	(4) 2006-
	Panel A. Regional allowances			
Regional allowances	0.00	0.00	0.02	0.06
	(0.00)	(0.00)	(0.04)	(0.05)
	Panel B. Demographic characteristics			
Land price	0.93	0.74	0.90	0.78
	(0.07)	(0.12)	(0.08)	(0.14)
	Panel C. Administrative characteristics			
Local tax revenue per capita	110.11	120.49	144.36	151.26
	(57.34)	(57.18)	(47.33)	(43.18)
LAT per capita	151.26	171.21	34.67	34.07
	(103.63)	(125.95)	(35.10)	(37.06)
NTD per capita	35.51	61.62	29.20	44.20
	(27.13)	(139.92)	(16.13)	(23.18)
Number of municipal public workers	560.92	507.58	2,023.93	1,808.81
	(681.25)	(632.58)	(4,051.79)	(3,457.19)
Average income of municipal public workers	3,284.95	3,202.02	3,432.45	3,288.63
	(190.12)	(173.82)	(216.93)	(183.17)
N	3964	8919	1456	3276

The summary statistics of the areas where regional allowances have been provided (at least once) and the areas where regional allowances have not been provided. Means are shown as coefficients, and standard deviations are given in parentheses.

D A Formal Exposition of the Theoretical Framework

Setup. A worker can choose to live in a small region or the outside world. Workers consume local residential amenities, numeraire goods, and land. Workers are heterogeneous in their idiosyncratic tastes to live in the small region. L denotes the population of the small region, and the total population in this economy is normalized to one. Since the region is small, we take the utility of living in the outside world, $\bar{V}$, as exogenous. We relax this small region assumption later in this section.

Upon choosing to live in the small region, a worker obtains a random job offer.[D.1] We denote the size of the public sector by L_g. Letting u denote the unemployment rate in this region, the number of private-sector workers is $(1-u)L - L_g$. Let p_g be the probability of obtaining a public-sector job and w_g be the public-sector wage. Then, the assumption of a random job offer implies that $p_g = L_g/L$. With the probability $\left((1-u)L - L_g\right)/L = 1 - u - p_g$, the worker works for the private sector and obtains wage w. Finally, with probability $uL/L = u$, the worker remains unemployed and obtains home production income b. We take (w_g, L_g) as exogenous because the government can determine this value by setting the public-sector wage and size. Prior to drawing a random job offer, we suppose that workers consider the expected income by evaluating the labor market attractiveness

[D.1] While we assume the random job offer for simplicity, our main arguments do not rely on this assumption.

of the small region: $p_g w_g + ub + (1 - p_g - u)w$.

For our purpose, we simply posit that the private-sector wages and the unemployment rate follow some functions of public sector variables (w_g, L_g). Specifically, we assume that the equilibrium private-sector wage in the private labor market (w) and the unemployment rate (u) follow

$$w = W(w_g, L_G, L), \quad u = U(w_g, L_g, L). \tag{D.1}$$

We do not specify a particular microfoundation of the private-sector labor market to derive the functions W and U to accommodate various labor market structures considered in the fully structural literature on the effects of public-sector wages. Such microfoundations include a search-theoretic model in which workers choose between private and public jobs.

A worker also derives utility from amenity A, which again depends on (w_g, L_g). For instance, a larger public sector may improve public services. A higher public-sector wage may also improve the quality of public services by, for example, inducing more effort from public-sector workers (c.f., Borjas 1984; Shapiro and Stiglitz 1984).[D.2] An alternative interpretation is that $A()$ summarizes the job amenities in this region, including in-kind benefits and unpaid overtime working hours.[D.3] As in the private-sector wages and unemployment rates in (D.1), we choose not to specify a particular microfoundation of $A()$ to accommodate various situations.

Finally, there is a land market. A worker consumes one unit of land at price r, regardless of employment status. Letting $r(L)$ be the inverse land supply function, which is increasing because of the increasing marginal cost of land supply for landlords. Then, the housing market equilibrium requires the land price to equal $r(L)$. For simplicity, we suppose that the land is owned by absentee landlords who spend their revenue outside the economy.

Worker i chooses to live in the small region if it brings utility higher than the outside world:

$$A + \left(p_g w_g + ub + (1 - p_g - u)w\right) - r + \epsilon_i \geq \bar{V}, \tag{D.2}$$

where ϵ_i is her idiosyncratic taste for the small region. The left-hand side is the expected utility of worker i of living in the small region, which we specify as the sum of the amenities A, the consumption of the numeraire good $\left(p_g w_g + ub + (1 - p_g - u)w\right) - r$, and the idiosyncratic taste. Letting $\bar{\epsilon}$ be the marginal worker such that she is indifferent between the small region and the outside world, the population of the small region is given by $L = 1 - F(\bar{\epsilon})$, where $F()$ is the cumulative distribution function of ϵ_i. In particular, this implies that $\bar{\epsilon} = F^{-1}(1 - L)$. To simplify the notation,

[D.2] The public service quality can be considered as net of the local tax payment for financing them, which may increase with public sector wages (Brueckner 1982).

[D.3] For instance, we may write the expected job amenities as $A = p_g A_g + (1 - u - p_g)A_p$, where A_g and A_p are the job amenities in the public and private sectors, respectively.

let $G(L) = F^{-1}(1 - L)$. $G()$ is decreasing in L.

Equilibrium Conditions The spatial equilibrium condition, which determines the total population L given (w, u), is as follows:

$$\underbrace{A(w_g, L_g)}_{\text{Amenities}} + \underbrace{\left(p_g w_g + ub + (1 - p_g - u)w\right)}_{\text{Expected labor income}} - \underbrace{r(L)}_{\text{Land cost}} + \underbrace{G(L)}_{\text{Idiosyncratic taste}} = \underbrace{\bar{V}}_{\text{Outside utility}} . \tag{D.3}$$

(w, u) are determined by the private-sector labor market equilibrium conditions (D.1). Finally, land market equilibrium is implicit in the condition that land cost equals the inverse land supply curve $r(L)$. These equilibrium conditions imply that the private-sector wage w, unemployment rate u, population level L, and land prices r all depend on the public-sector wage w_g. We empirically estimate the effects of w_g on these four variables.

Our model is concerned with a single type of mobile worker. In our empirical analysis, we posit that young workers are affected by the public-sector wage reform, while older workers are not. To address this situation in our theoretical framework, we may interpret the mobile workers in our model as young workers. The simplest way of extending our model to introduce older workers is to assume that they are immobile and obtain the exogenous wage rate w_o, implying that it is independent of the public-sector wages and employment (w_g, L_g). The independence of the public sector variables is motivated by our argument that the public-sector labor market for older workers is thin in our Japanese context (see Section 2), and the assumption of geographical immobility is consistent with the common empirical finding that the intercity mobility rate sharply decreases with age (e.g., Ishikawa 2016; Kaplan and Schulhofer-Wohl 2017). Then, after slightly modifying the land market equilibrium condition such that the land price is equal to $r(L + L_o)$, where L_o is the number of immobile older workers in the small region and $L + L_o$ is the total number of workers, the remaining equilibrium conditions are the same as the equilibrium conditions (D.1) and (D.3).[D.4]

Welfare Implications of Changing Public-sector wages. Our model yields simple welfare implications for the public wage cut: it harms workers' welfare if it decreases the population and lowers land prices.[D.5] Intuitively, holding land prices fixed, a decrease in the public-sector wage

[D.4] Yamagishi (2021) formally shows in a similar model that when there are two groups of workers with different levels of geographical mobility, land prices more strongly reflect the willingness of more mobile workers to pay to live in the small region. The case of immobile older workers can be thought of as the limit case of this result.

[D.5] To see this, let $A + \left(p_g w_g + ub + (1 - p_g - u)w\right) - r + \epsilon_i$ be worker i's welfare. Suppose that the population L decreases after the reduction in the public-sector wage, which is equivalent to the decrease in r because of the land market equilibrium condition. The decrease in the population implies that $\bar{\epsilon}$ is higher since $F^{-1}(1 - L) = \bar{\epsilon}$. The spatial equilibrium condition (D.3) then implies that $A + \left(p_g w_g + ub + (1 - p_g - u)w\right) - r$ is smaller. Therefore, for a given ϵ_i, welfare decreases if worker i lives in the small region after the change in the public-sector wage. The utility of workers in the outside economy remains constant, so that no worker in the economy benefits from the public-sector wage cut.

affects workers' welfare by changing the wage when employed in the public sector, the wage when employed in the private sector, the unemployment rate, and amenities. However, without discerning the underlying mechanisms behind welfare, the welfare decrease lowers land prices by inducing out-migration so that the spatial equilibrium condition (D.3) holds. Moreover, the decreased population lowers land prices by reducing the demand for land. Therefore, in light of our model, our empirical analysis of population and land prices provides sufficient statistics of the qualitative welfare implications of policy reform regarding public-sector wages. This result motivates us to examine population and land prices as outcome variables.

Note, however, that this result applies only to mobile workers in the model. If there are additional immobile workers, such as the older workers discussed above, then such workers may experience a welfare gain even if the population and land prices decrease. For instance, the public-sector wage cut may lower the land price without affecting the wages, unemployment rates, or amenities of older workers, which would increase their utility through lower housing costs.[D.6] As we see in Sections 5.3 and 6.2, this situation may be empirically relevant in our context.

Spillover Effects from Other Regions. The above equilibrium conditions (D.1)–(D.3) imply that the local economic outcomes (w, u, L, r) are independent of the public-sector wages in other regions. This is a consequence of two implicit assumptions. First, there is no commuting across regions. For instance, a worker in region i may commute to work at a public sector job in region j, but then the public-sector wage rate of region j should be considered a relevant aspect of local labor market i. Second, we assume that the welfare level of the marginal worker is exogenously determined at outside utility level $\bar{V}$. In a general equilibrium model that explicitly models the outside option, however, such an outside utility level is endogenously determined, which in turn depends on the public sector wage rate of all regions. By relaxing these two assumptions, the model predicts that public-sector wages in all regions affect local economic outcomes in the region i: (w_i, u_i, L_i, r_i).

To illustrate this point, suppose that there are N regions in this economy, indexed by $i = 1, ..., N$. The economy is closed in the sense that the total population is fixed, and we normalize it to one. Note that the economy no longer has the "outside region" associated with exogenous utility $\bar{V}$.

We introduce commuting as follows. The attractiveness of local labor market j is denoted as the expected utility from available job opportunities: $\left(p_{gj} w_{gj} + u_j b_j + (1 - p_{gj} - u_j) w_j\right)$. Workers living in i choose workplace j maximizing the expected utility while incurring τ_{ij}, the bilateral commuting cost between i and j. The attractiveness of workplace j for workers in residence i inclusive of commuting cost is written as $\tau_{ij}^{-1} \left(p_{gj} w_{gj} + u_j b_j + (1 - p_{gj} - u_j) w_j\right)$.[D.7] Workers also

[D.6]See Yamagishi (2021) for more discussion on workers' heterogeneity and welfare implications.

[D.7]This implicitly assumes that workers choose their workplace prior to knowing whether they will obtain a private-sector job, public-sector job, or no job offer.

face an idiosyncratic taste shock for workplace j that follows the Type-I extreme value distribution. When workers choose their residence and workplace to maximize their utility, the probability of a worker living in i working in j is written as follows:

$$\frac{\exp\left(\tau_{ij}^{-1}\left(p_{gj}w_{gj}+u_jb_j+(1-p_{gj}-u_j)w_j\right)\right)}{\sum_{j=1,\ldots,N}\exp\left(\tau_{ij}^{-1}\left(p_{gj}w_{gj}+u_jb_j+(1-p_{gj}-u_j)w_j\right)\right)}$$

Prior to the realization of the idiosyncratic shock, the expected value of workplace options by living in region i is written as the well-known log-sum formula (see Train 2009):

$$\ln\left(\sum_{j=1,\ldots,N}\exp\left(\tau_{ij}^{-1}\left(p_{gj}w_{gj}+u_jb_j+(1-p_{gj}-u_j)w_j\right)\right)\right)$$

Note that in the special case of prohibitive commuting costs (i.e., $\tau_{ij}=\infty$ for $j\neq i$ and $\tau_{ii}=1$), this expression decreases to $p_{gi}w_{gi}+u_ib+(1-p_{gi}-u_i)w_i$, which is the attractiveness of the local labor market i in our baseline model without commuting.

In this setup, the equilibrium conditions are that for all $i=1,\ldots,N$ and $j\neq i$:

$$A_i+\ln\left(\sum_{k=1,\ldots,N}\exp\left(\tau_{ik}^{-1}\left(p_{gk}w_{gk}+u_kb_k+(1-p_{gk}-u_k)w_k\right)\right)\right)-r_i(L_i)=$$

$$A_j+\ln\left(\sum_{k=1,\ldots,N}\exp\left(\tau_{jk}^{-1}\left(p_{gk}w_{gk}+u_kb_k+(1-p_{gk}-u_k)w_k\right)\right)\right)-r_j(L_j) \tag{D.4}$$

$$w_i=W(w_{gi},L_{gi},L_i) \tag{D.5}$$

$$u_i=U(w_{gi},L_{gi},L_i), \tag{D.6}$$

where we suppose that the land market clearing condition in each location i makes the land cost equal to $r_i(L_i)$. These equilibrium conditions, together with the total population constraint $\sum_{i=1,\ldots,N}L_i=1$, determine the endogenous variables of the model: $(w_1,\ldots,w_N)$, $(u_1,\ldots,u_N)$, $(L_1,\ldots,L_N)$, and $(r_1,\ldots,r_N)$.

Equations (D.4)–(D.6) naturally extend (D.1)–(D.3) to include commuting and general equilibrium effects. First, the spatial equilibrium condition (D.4) includes the workplace access term of location i:$\ln\left(\sum_{k=1,\ldots,N}\exp\left(\tau_{ik}^{-1}\left(p_{gk}w_{gk}+u_kb_k+(1-p_{gk}-u_k)w_k\right)\right)\right)$. Second, it no longer features the exogenous utility of living in an outside region, but it explicitly requires the spatial equilibrium condition between region i and all other regions $j(\neq i)$. On the other hand, we simplify the spatial equilibrium condition by assuming that workers do not have an idiosyncratic taste for each residence

i. Introducing it does not change our argument.

The equilibrium conditions (D.4)–(D.6) imply that the wage rate, population, unemployment rate, and land price of region i depend on the public-sector wages of *all* regions $\boldsymbol{w_g} = (w_{g1}, ..., w_{gN})$. With commuting and general equilibrium changes in the welfare of outside utility, there are two channels through which an increase in w_{gj} affects private sector wages in region $i(\neq j)$. First, it improves the workplace access of region i through commuting. This increases the demand for living in region i. In particular, the public-sector wages of nearby regions can have a large effect because of relatively low commuting costs. On the other hand, improvement in the workplace access of region $k(\neq i)$ also occurs, which reduces the demand for living in region i. Which effect dominates is an empirical question. Since w_i, u_i, and r_i are all functions of population L_i, we cannot theoretically determine how commuting and general equilibrium welfare changes affect these variables. As a result, since our main empirical specification ignores changes in w_{gj} in considering the effect of public-sector wages in region i, this could induce a bias whose sign is unknown a priori. In Section 6.4, we investigate the sensitivity of our results by using a spatial econometric model to explicitly account for the public-sector wages in other regions (Halleck Vega and Elhorst 2015). We find that our results are robust in this alternative model.

E Details on Calculating the Aggregate Economic Impact

In this appendix, we explain how we calculate the economic impact of the public wage reform in Section 6.4.

Our estimation results show the extent to which the outcomes change with the change in the regional public wage level. Therefore, by multiplying the coefficients of the outcomes, the changes in the effective public wage, and the actual amount of the outcome variables in each municipality, we can estimate how much the outcomes in each municipality changed due to the public wage reform.

For the changes in the effective public wage in each region, since the effective wage in region i in 2010 can be shown as $(1 + RA_{i,2010}) \times \text{Base wage}_{i,2010}$, where RA is the regional allowance rate, the change in the effective wage in region i can be calculated as $(1 + RA_{i,2010}) \times \text{Base wage}_{i,2010} - (1 + RA_{i,2005}) \times \text{Base wage}_{i,2005} = (\text{Base wage}_{i,2010} - \text{Base wage}_{i,2005})(1 + RA_{i,2010}) + (RA_{i,2010} - RA_{i,2005}) \times \text{Base wage}_{i,2005} = -0.048 \times (1 + RA_{i,2010}) + (RA_{i,2010} - RA_{i,2005})$, where we normalize $\text{Base Wage}_{i,2005}$ to 1 in the last equation.

We now explain each outcome variable in more detail. For each calculation, we use data from 2010, when the reform ended, to prevent differences due to the base year of each data point.

Impact on private earnings Since our study shows that a 1% decrease in public wages leads to an approximately 0.35% decrease in private wages for young full-time workers aged 15 to 29, we can infer the impact of the public wage reform on private earnings if the total amount of earnings for young full-time workers in each municipality is available.

However, since such data are not available, we approximate them using alternative data. In the absence of data regarding the number of young full-time workers by municipality, the number of young full-time workers is reproduced using the number of full-time workers in private sectors[E.1] and the percentage of the working-age population aged 15-29 years from the census in each municipality. In addition, since data on the average wage of young full-time workers are not available for some municipalities, we use the average wage of the prefecture from the Basic Survey on Wage Structure to infer the average wage of each municipality. We then construct the approximate total amount of earnings for young full-time workers in each municipality by multiplying the approximate number of young full-time workers by the average wage of young full-time workers.

By multiplying the estimated elasticity of the private wage of young full-time workers, the changes in the effective public wage, and the approximate total amount of earnings for young full-time workers in each municipality, the reduction in earnings for young full-time workers due to the reduction in public wages is estimated as 111.6 billion yen for 2010.

Impact on the taxable assessed value of land Given that our analysis of land price shows that a 1% decrease in public wages leads to an approximately 0.54% decrease in the land price, we can infer the impact of public wage reform on the land price if the total value of land in each municipality is available.

However, since the total value of land in each municipality is not available, we alternatively use data on the taxable assessed value of land, which is from the Local Government Finance Survey. This variable is measured when property taxes on land are levied. The assessed value of land may be reduced to a portion of the transaction value of land, depending on its use to reduce property taxes, and is often said to average approximately 70% of the transaction value of land.[E.2] Therefore, our inference might be a lower bound of the impact of the reform on land transaction prices.

By multiplying the estimated elasticity of land price, the changes in the effective public wage, and the taxable assessed value of land in each municipality, the reduction in the taxable assessed value of land due to the reduction in public wages is estimated as 193.9 billion yen for 2010.

Impact on unemployment among the young population The result of our analysis on unemployment shows that a 1% decrease in public wages leads to an approximately 0.5% increase in the ratio of young and old unemployment rates, controlling for the preexisting trend. Although this

[E.1]The sample is limited to workers in the secondary and tertiary industries, which are the subject of our analysis.

[E.2]Consistent with this, Yamagishi and Sato (2023) compares the land assessment values for property taxation and transacted land prices, finding that a 1% increase in the former is associated with a 1% increase in the latter.

result is nonsignificant, it offers suggestive evidence that public wage cuts increase the youth unemployment rate. Therefore, we use this result to infer the impact of public wage reform on youth unemployment.

Consistent with our argument in Section 2 that only the young are affected by the public-sector wage reform, we suppose that the public-sector wage reform affects the unemployment rate of the young. This change in the ratio of young and old unemployment rates due to public wage reform can be derived by multiplying the changes in public wages and the estimation result, 0.5%. This is denoted as "ΔYO ratio of Unemployment". Since the unemployment rate is the ratio of unemployed individuals to the labor force, the change in the number of young unemployed individuals can be calculated as "ΔYoung unemployment = ΔYO ratio of Unemployment $\times$ Old unemployment rate $\times$ Young labor force" in each municipality. Since the census reports the labor force and unemployment by age group in each municipality, we use these data to infer the impact.

Applying the calculation, we can infer that the estimated increase in the number of young unemployed individuals is approximately 10,471 in 2010.

The total amount of public wage reduction in local governments In addition to estimating the economic impact on the various outcomes, we infer the total amount of public wage reduction in local governments, considering that the changes in the effective public wage were different for each region. Since the Fact-Finding Survey on the Compensation of Local Government Employees includes the number and average wage of local government employees by age and education level for each locality, we calculate that the total amount of public wage reduction for all municipalities due to the reform can be estimated at 104.6 billion yen in 2010 based on these data.

However, the public wage reduction for prefectures cannot be exactly estimated because different regional allowance rates are applied in different municipalities within the same prefecture, and it is not known how many prefectural workers are located in each municipality. Since the number of prefectural workers is approximately 1.1 times greater than the number of municipal workers and prefectural workers tend to be located in the prefectural capital, which tends to have higher regional allowance rates than other regions, we suppose that the total amount of public wage reduction for all prefectures due to the reform is similar to that for all municipalities, i.e., approximately 100 billion yen in 2010.

Taken together, the estimated total amount of local public wage reduction is approximately 200 billion yen in 2010.

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