Empirical Economics 31:113–142 (2006)
DOI 10.1007/s00181-005-0024-7

ORIGINAL PAPER

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Local flexibility in wage setting: evidence from the Norwegian local public sector

Received: 9 July 2002 / Accepted: 8 March 2005 / Published online: 20 January 2006

Abstract The public sectors in the Scandinavian countries have been prominent examples of centralized wage-setting systems. In Norway, more room for local flexibility was implemented by a wage frame system introduced in 1990 in which the national wage scale system merely works like a minimum wage system. We analyze the effect of this reform using a unique database where we can track employees and their local government over time and explore the consequences of controlling for fixed individual effects and fixed employer effects. We find that the wage dispersion increased across local governments after 1990, and that wages to some extent became more responsive to local government income, monopsony power and other local government characteristics after the reform. However, the numerical effects of the reform are estimated to be quite small.

Keywords Public sector labor market · Wage setting · Decentralization

1 Introduction

This study examines to what extent wage formation in the local public sector is responsive to changes in the wage setting rules. Many countries have seen a movement from rigid centralized pay systems towards more flexible systems with more responsibility delegated to the agency or local government level [Gregory and Borland (1999)]. The main motivation for this movement has been that centralized pay systems may lead to local labor market imbalances with excess

Comments from Kåre Johansen, Timo Teräsvirta, seminar participants at Geneva, Solstrand, Stavern and Trondheim, and two anonymous referees are greatly appreciated. We thank Jan Morten Dyrstad for invaluable help with the data. Some of the data is obtained from the Norwegian Social Science Data Services (NSD). However, NSD is not responsible for the analysis or the conclusion that are drawn.

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supply in some areas and excess demand in other areas. Moreover, centralized pay setting is likely to lead to systematic differences in average worker quality as employees sort themselves across employers due to differences in non-pecuniary workplace attributes. To the extent that local conditions differ, the delegation of pay responsibility to the local level is expected to give rise to different wage developments across units.

Within a simple competitive model of the labor market, decentralization implies that wages would react on the local labor market such that local excess demand or supply is eliminated. Hence, according to the competitive model, decentralization would generate an efficiency gain. However, it is relevant to consider alternatives to the competitive model when describing wage outcomes in the public sector. In a decentralized system, some establishments may have strong unions, other weak or nonexisting unions. Further, even with equally strong unions, the degree of fiscal competition among public employers may lead to different wages, and some employers may have monopsony power due to dominance in the local market or low employee mobility.[1] All these factors may lead to an inefficient allocation of labor and distorted wages. Equal employees will receive different wages dependent on the power of their local union and characteristics of the employer.

Public sector labor markets in most countries have strong unions and union wage models are therefore likely alternatives to the simple competitive model. Further, a long period of centralized pay setting with equal wage across establishments is likely to generate a norm of pay equality that may persist after decentralization. Thus, attempts to increase salary to eliminate excess demand for public employees in some areas may generate pay increases in other areas as local unions concerned with pay comparison try to implement the traditional pay equality norm. Similar effects may arise if wages are determined by efficiency wage mechanisms where worker productivity depends on own wages compared to the wages in other establishments (Akerlof and Yellen 1990). Accordingly, the actual relationship between local wage development and local conditions following decentralization may be smaller than predicted by the competitive model. To assess the possible efficiency gain from pay reforms, it is therefore of interest to study empirically the actual effect of decentralization. Such studies may also shed light on the likely effects of decentralization of pay bargaining within large private firms.

The existing empirical evidence on the outcomes from changing pay systems in the public sector is very limited. Elliott and Bender (1997) describe the attempts to decentralize pay bargaining and individualize pay in the central government in UK, Sweden and Australia and present preliminary evidence indicating that the reforms lead to increased earnings dispersion. But their study does not provide evidence to what extent decentralization changed the relationship between pay and local conditions. Further, some evidence exists on the relationship between wage outcomes and local labor market conditions under different pay systems operating at a certain point in time. Olson et al. (2000) compare wage responsiveness to local conditions between white- and blue-collar workers in the US federal government. They find that wages for white-collar workers, who are formally intended to follow national wage trends, are much less responsive to local labor market conditions than wages of blue-collar workers where the formal room for local wage setting is

[1] Several authors have argued that public sector employers have monopsony power [Ehrenberg and Schwarz (1986)] especially in areas with few competing communities.

much higher. While this evidence may reveal useful information on the performance of different wage setting rules, an obvious problem is to isolate the pure system effect when differences in wage setting rules are closely associated with differences in worker skills. Moreover, such cross-section evidence does not take into account that established norms on wage equality may continue to affect wage setting, irrespective of formal changes in wage setting rules.

The present paper contributes to the literature by using matched employee–employer data for local government employees in Norway 1986–1998 to investigate the effect of changing a centralized pay system to a system with more local pay discretion introduced in 1990. The problems of retaining and recruiting personnel within the centralized system lead to the expectation that wage shifting would occur as more flexibility were introduced. On the methodological side, the possibility to follow individual wage developments is of particular importance when investigating the impact of pay decentralization. Completely centralized systems imply absence of employer specific wage differentials conditional on individual worker characteristics [Edin and Zetterberg (1992)]. Following decentralization, the competitive model predicts employer wage differentials conditional on individual worker characteristics as local labor market imbalances are eliminated and local wages move towards their competitive levels. The panel nature of the individual data at our disposal enables us to control for both observed and unobserved individual and employer specific variables in a much more satisfactory way than in previous studies of public sector wage formation.

Complete rigidity and complete local flexibility should be considered as extreme cases, rarely observed in practice. Even in the old system, the local authorities could change the occupational classification in order to compensate for the centralized pay system. The evidence in Strøm (1995) indicates that such manipulation took place in the pre-reform phase, but that the local wage differences did not reflect differences in local labor market conditions. On the other hand, the pay reform did not remove the institution of national contracts. The national contracts continued to operate, and a national wage agreement existed within the whole period. But the important change was that after 1990 the central contracts stipulated for each occupation a wage frame available for the local governments, with considerable formal local freedom to place the employees within the wage frame. A central issue in this paper is therefore to analyze the extent to which the introduction of more formal flexibility changed the relationship between pay and local conditions. To allow for sluggishness in the adjustment to the reform, for instance due to slowly adjusting pay equality norms, we apply the Smooth Transition Regression Model (STR) as suggested by Teräsvirta (1998). This approach implies that the coefficients of the variables of interest in the wage model evolves according to a parametric transition function after the 1990-reform.

The paper is organized as follows. Section 2 contains a description of the changes that have taken place in the wage-setting system in the Norwegian local public sector during recent years. Section 3 presents our research expectations. Section 4 includes a descriptive data analysis of the changes in wage dispersion across local governments. Section 5 presents the econometric model, while the econometric results are given in Section 6. The effect of the wage-setting reform is investigated by the STR approach, and we distinguish between different types of workers. Section 7 contains concluding remarks.

2 The wage system

Until 1990, wage setting in the Norwegian local public sector could be characterized as the prototype of a centralized bargaining system. National bargaining between the Federation of Local Governments and national union federations determined the wage schedule for each single occupation, except for the leading managers. Given occupation and seniority, the wage level for each employee was almost exactly determined by the national wage contract. The major source of wage flexibility in the system was the discretion for local authorities to manipulate the occupational classification. For example, the authorities were relatively free to decide whether a worker was classified as a clerk, head clerk or head clerk with special grade. But this possibility was quite limited for some large groups of employees, as for example nurses and teachers. Strøm (1995), using data from 1985–1988, provides evidence that systematic differences in wage levels across local governments did exist within the centralized system. In particular this study found that wages for the lower skilled group (employees with basic commercial education) were affected by the wages of the higher skilled group (engineers) within the same local government and the political composition of the local council, while no effects of the external labor market was revealed. Systematic evidence on the existence of regional imbalances in the market for local public employees in the pre-reform period is not available. But recent evidence for Norwegian school teachers that were not covered by the reform documents significant sorting of teachers across schools and regions due to differences in student composition and other school characteristics [Bonesrønning et al. (2005)].

This evidence give some indication that the local adjustments taking place within the old system did not eliminate local imbalances and that the employees to some extent sorted themselves according to differences in job attributes. Thus, introducing more direct flexibility in wage setting at the local level may be seen as a response to the inadequacy of the old system.

During the 1980s the employer side represented by the Federation of Local Governments argued frequently that more local flexibility in pay setting was necessary in order to retain and recruit personnel. The largest trade union organizing low and medium skilled local government employees, the Norwegian Organization of Municipal Workers (Norsk Kommuneforbund), resisted reform arguing that decentralization would counter the goal of reducing income differentials and the principle of equal pay for equal work. However, there were substantial disagreement on this issue between this trade union and other smaller unions mainly organizing higher educated employees. As the number of employees with higher education increased, the opposition to decentralization weakened. Also in the central government sector the issue of introducing more wage flexibility at the establishment level was discussed. In November 1988, the government appointed a committee to consider changes in the wage system in the central government sector and the committee presented its report in December 1990.

Nevertheless, it came as a surprise to most observers when the parties in the national negotiations in the local sector already in the spring of 1990 agreed on a new wage system in the local sector, the so-called “wage frame system”. This system implies that each occupation is placed in a “wage frame”. In the 1990–1992 contract, each wage frame consisted of a basic wage schedule and an alternative wage schedule, both with stipulated wages dependent on seniority. The actual wage schedule used for individual employees is to be decided at the local level. As an alternative to these two seniority-dependent schedules, the municipality could also choose to give the worker a wage independent of seniority. This wage was one step higher on the wage scale than the highest wage in the two seniority-dependent schedules. In later contracts, the number of alternative schedules increased somewhat, but the basic features of the system have been more or less unchanged after 1990.[2] According to the national contract, the lowest wage schedule within a wage frame should be considered as a basic wage. Thus, the new wage system looks very much like a minimum wage system.

Table 1 illustrates the changes in the wage system during the recent years exemplified by the wages for engineers, nurses and pre-school teachers with zero seniority. According to Table 1, the reform in 1990 introduced substantial formal flexibility into the wage system. The difference between the lowest and highest possible wage increased from 0 in the 1988–1990 contract to 24% in the 1990–1992 contract, and from 1998 there is no maximum wage.

The 1990 reform did introduce a substantial amount of local flexibility in the local wage determination, but the implementation of the new system has been constrained by relatively restrictive limits on wage increases that can be given in local negotiations. National contracts have stipulated the total amount that can be negotiated locally. For example, in 1990, the amount available for local negotiations was 1.12% of the average total wage costs in 1989. In addition, the national contracts have sometimes contained explicit descriptions of the distribution of wage increases across groups. For instance, the central bargained contract in 1990 stated that at least 60% of the amount available for local negotiations should be given to low-wage workers, and the local negotiators were advised to favor women and workers in the health care sector. While the national contracts put constraints on the formal local negotiations, the full flexibility of the new wage system was available for new hires. In some respect this resembles the academic

Table 1 Wage schedules for Engineers (Occupational code 7084), Nurses (Occupational code 7174) and Pre-school teachers (Occupational code 6709). Zero seniority. Yearly wage in NOK

Contract period	1986–1988	1988–1990	1990–1992	1992–1994	1994–1996	1996–1998	1998–2000
Minimum wage	120,134	124,500	145,451	147,636	158,100	174,600	194,100
Maximum wage	120,134	124,500	179,873	190,133	201,000	224,300	–
Percentage difference	0	0	23.7	28.8	27.1	28.5	–

[2] There are no formal restrictions on which of these three alternative methods that can be used. For instance, one nurse can be put in the basic wage schedule, another in the alternative schedule, and a third on the wage independent of seniority. To illustrate, in the 1996–1998 contract the basic wage schedule for a nurse with zero seniority implied a yearly wage of NOK 174600, while the two alternative seniority dependent wage schedules implied NOK 178000 and NOK 183000, respectively. The wage schedule independent of seniority implied a yearly wage of NOK 224300 as seen from Table 2. Further, in principle, one individual can be put on the different schedules at any point in their career.

labor market in the US as described by Ransom (1993), with market wages paid to new hires, but not to others. Below we therefore investigate the extent to which the reform effect is different for newly hired workers versus incumbent workers.

3 Research expectations

In this section we lay out in more detail our research expectations. The determinants of pay in the local public sector is likely to be a complicated mix of the formal constraints given in the national wage contracts, characteristics of the fiscal system, and the political decision making process. We take as point of departure an individual earnings equation based on human capital theory where wages depend on a set of individual characteristics such as education, gender and age. Within the centralized pay system with little opportunity to make local adjustments except for manipulation of the occupational classification, traditional human capital variables should to a large extent describe the dispersion of wages between workers. When the system allows for local adjustment, wages should respond to local conditions as the state of the local labor market.

For example, a local government located in an area with an initial shortage of qualified pre-school teachers could begin to bid up the wage of these workers in order to retain and recruit personnel of required quality. Thus, while the wage distribution in the Norwegian local public sector should be almost fully described by a traditional Mincerian wage equation in the pre-1990-period, we expect that human capital variables are less important in the post-1990 period. The evidence in Elliott and Bender (1997) and Olson et al. (2000) lead us to expect higher wage differentiation both across and within local governments after the reform.

While supply and demand factors obviously may play a role in local government wage formation, trade unions are present and have bargaining rights. As is well known, the predicted reduced form effect of most variables does not differ between competitive and imperfect competition models, implying that formal tests of underlying theoretical models are difficult to conduct based on reduced form models. Nevertheless, we may draw some tentative conclusions based on the effect of explicit measures of fiscal competition and bargaining power variables.

3.1 Local government budget

In the centralized fiscal system in Norway, the largest part of municipal revenue is determined by the national government. A centralized financing system of the local government sector has been the counterpart to the centralized wage-setting system. The main local tax rates (income tax and wealth tax) are regulated by the central government and are equal across the country. It is a tax revenue sharing system with extensive use of central government grants.

We expect a positive relationship between local wages and the size of the local government budget. In a competitive model a rise in the local government budget is expected to increase the wage via higher demand. In a wage bargaining model, higher government budget increases the bargained wage in the same way as higher productivity or product prices in private firms as demonstrated by Strøm (1995).

Thus, we expect that the effect of budget size on local government wages increases as wage setting is decentralized.

3.2 Local unemployment rate

We expect that local wages are negatively linked to the local unemployment rate. If higher local unemployment rate increases the number of workers offering themselves for work in the local government sector, the competitive model predicts an inverse relationship between local public wages and local unemployment. Within a union model, the probability to become unemployed for a trade union member increases as the unemployment rate increases, also generating an inverse relationship between wages and unemployment.

3.3 Monopsony power

Several authors have argued that public employers have monopsony power in the labor market. Empirical evidence suggests that wages in the US is lower in municipalities located in areas with few competing communities, see for example Ehrenberg and Schwarz (1986). However, such an effect can also be rationalized within a trade union bargaining model. The higher the number of surrounding communities, the higher is the probability for unionized local government workers to find work elsewhere in the case of an employment reduction or a conflict in their own community, increasing the bargained wage.

3.4 Political strength

Within a wage bargaining model, the actual wage outcomes will depend on the bargaining power of the parties involved in local negotiations, and variables influencing the amount of money available at the bargaining table, conditional on the total local budget.

Recent Norwegian empirical evidence suggests a negative relationship between public sector spending and measures of the strength of the political leadership, see Kalseth and Rattsø (1998) and Falch and Rattsø (1999). Their interpretation of this evidence is that weak local governments create more room for lobbying by interest groups. To what extent will political strength influence wages in the local public sector within such a setting? A possible story, presented in Falch and Strøm (2003), is the following. Consider a local government consisting of two or more service producing sectors and an exogenously given total budget. Interest groups representing the users of the public services want a high service level in their sector. The political leadership has to balance the interest groups' pressure for high employment against other spending items. When strong political leaderships are able to withstand the interest group pressure, a larger “cake” that can be used on wages is created, and a larger cake available in the wage bargaining will increase the wage, all else equal. Thus, even though the direct effect of political strength is increased local government bargaining power in the wage bargaining, political strength may increase the wage under some plausible circumstances. In the case of efficiency wages where a higher wage increases the productivity of the employees, more resources available after the interaction with the interest groups will under some highly plausible conditions also increase the wage level. Thus, the effect on public sector wages of a strong political leadership is ambiguous. Nevertheless, as wage setting is decentralized and the potential effects of political strength are allowed to play out in wage setting, we expect the importance of political strength to increase.

3.5 Fiscal competition

Empirical evidence and theories of wage bargaining and efficiency wages suggest that private sector wages are negatively related to the degree of product market competition [Nickell (1999)]. Similar mechanisms may exist in the public sector. In the literature on public finance it is argued that taxpayer mobility may put constraints on the ability of the local public sector unions to extract part of the local budget in a decentralized fiscal system. Higher local wages lead to increased local tax rates which tend to erode the tax base and the local budget due to dissatisfied taxpayers leaving the community, see Courant et al. (1979). According to this view, the wages should be lowest in areas where mobility across communities is easy, all else equal.

The mechanism of fiscal competition is relevant also in a centralized fiscal system as in Norway since higher wages must be financed by lower level of public services. In such a system, decisions on where to locate may partly be based on comparison of local service levels. Below we construct a measure of fiscal competition, and expect to find that local wages became more responsive to fiscal competition when the wage setting was decentralized.

3.6 Sluggish adjustment

We take into account that adjustment to reforms usually takes time. In our context three arguments are of particular relevance. First, if established pay equality norms across local governments are maintained by local unions also in the decentralized regime, the actual impact from the change in wage setting rules may be quite small at least in the first years as norms change slowly. Second, the formal flexibility defined in the central wage agreements increased over time. Third, much of the flexibility is related to new hires. Thus, as more of the employees are hired after the reform, we expect increased variability across local governments. To allow for a gradual adjustment towards a new equilibrium, we estimate a smooth transition regression model where the effect of local conditions changes gradually through time according to a parametric transition function [Teräsvirta (1998)].

3.7 Newly hired versus incumbent workers

As commented on above, the flexibility of the new wage system was fully available only for newly hired workers. We therefore expect that the effect of the reform was stronger for newly hired workers than for incumbent workers. Thus, in addition to the baseline model, we will also estimate models that allow the reform response to differ between these two groups of workers.

4 Descriptive data analysis

In this section we present the development in the wage distribution for local government employees during the period 1986–1998. A primary goal is to investigate whether the distribution across local governments changed in any significant way after the introduction of the wage frame system in 1990. The data include employees working above 20% of full time job in the welfare services provided by the Norwegian municipalities.[3] The first part of Table 2 presents the number of observations each year and the development in the aggregate real wage during the empirical period. After a high wage growth from 1985 to 1986, the mean real wage dropped markedly from 1986 to 1987. Afterwards, the real wage varied little before it started to rise in the late 1990s.

Following the discussion in Section 2 and 3, the changes in the wage system in 1990 are expected to increase the wage dispersion. Table 2 shows that the standard deviation of the log wage decreased up to 1990, and stayed fairly constant afterwards. The compression occurred at the right tail of the distribution. The median wage in relation to the wage at the 10 percentile actually increased over time, while the wage at the 90 percentile compared to the median wage decreased. Looking at percentile gaps may give an inadequate picture, however, since the wage distribution is not continuous but stepwise due to the use of a common wage scale. A drop in the median wage in 1987 explains the special result that year.

This evidence is at odd with the prior believes. However, a major weakness of the above measures of wage dispersion is that they do not control for differences in individual characteristics of the employees. The compression of the wage distribution may be a result of reduced variation over time in individual characteristics. For example, the share of workers without college and university education decreased during the period from about 50% to below 40%. To control for individual characteristics, we regress for each year individual wages against a number of individual characteristics,

$$\mathrm{w_{ij}} = \mathrm{x_{ij}}\beta + \eta_{\mathrm{ij}} \tag{1}$$

where w_{ij} is the log wage of individual i in municipality j, x_{ij} is a vector of individual characteristics, β is a vector of estimated parameters and η_{ij} is the residual. Individual characteristics included are dummy variables for education (six categories), age (seven categories) and working hours (three categories). By adding terms that interact the variables with the gender dummy, all variables are allowed to have gender specific effects. The variables used are defined in the Appendix.

The variation of η is the wage variation conditional on measured individual characteristics, and is consequently either a result of unmeasured individual characteristics

[3] Even though primary and lower secondary education are the responsibility of the municipalities, the teacher union bargain with the central government over wages and other issues. Teachers were the only group of workers not covered by the 1990-reform, and are not included in our sample.

Table 2 Monthly wage and wage dispersion

			Unconditional log wage dispersion			Log wage dispersion conditional on individual characteristics[2]		
Year	Observations	Mean real wage[1]	Standard deviation	Percentile gap: 50/10	Percentile gap: 90/50	Unweighted std. dev.	Weighted std. dev. within municipalities	Weighted std. dev. across municipalities
1986	111,569	16,719	0.159	0.075	0.263	0.103	0.102	0.014
1987	116,479	15,626	0.168	0.036	0.311	0.110	0.109	0.015
1988	132,298	15,357	0.157	0.074	0.262	0.105	0.104	0.014
1989	141,930	15,432	0.151	0.070	0.250	0.101	0.100	0.013
1990	149,084	15,886	0.144	0.079	0.214	0.096	0.095	0.014
1991	162,965	15,577	0.141	0.078	0.211	0.095	0.094	0.013
1992	173,781	15,563	0.146	0.085	0.202	0.100	0.099	0.016
1993	185,581	15,701	0.149	0.084	0.232	0.101	0.100	0.015
1994	192,063	15,725	0.148	0.083	0.229	0.102	0.100	0.015
1995	200,723	15,752	0.147	0.081	0.225	0.103	0.101	0.015
1996	206,390	16,489	0.142	0.078	0.215	0.101	0.100	0.015
1997	203,207	16,439	0.143	0.075	0.227	0.103	0.102	0.016
1998	208,727	17,572	0.133	0.108	0.179	0.095	0.094	0.015

[1] 1998-NOK, deflated by the consumer price index
[2] Individual characteristics included are two dummy variables for part time work, age categories and educational categories, all interacted with gender. The variables are defined in the Appendix

or local government wage policy.[4] According to the second part of Table 2, controlling for individual characteristics reveals a temporary drop in wage dispersion in 1990, the reform year, and 1991. However, overall, wage inequality is remarkably stable over the period, in contrast to the reduced inequality observed in the unconditional wage. Thus, the reduction in the overall wage inequality found in the raw data can be explained by reduced variation in measured individual characteristics.[5]

If the municipalities used the increased flexibility in local wage policy introduced by the 1990 reform to design their own pay policy based on local conditions as hypothesized in Section 3, we should observe increased variation in wages across the local governments in the post-1990 period. A simple way to test this hypothesis is to decompose the variation into two parts, variation within municipalities and variation across municipalities. The two last columns in Table 2 present the standard deviations within and between municipalities, weighted by the number of employees. There is a tendency of increased variation across municipalities over time, and an increasing share of the variance of the conditional log wage is due to variation across municipalities. The within municipality standard deviation is seven and six times higher than across municipality standard deviation in 1986 and 1998, respectively. Notice also that the temporary drop in the overall standard deviation in 1990 and 1991 is due to a drop in the within municipality variation, while the variation across municipalities in fact increased in 1990.

There are two weaknesses by using the residuals from (1) to calculate variation across municipalities. First, we are not able to test whether the variation is significant in statistical terms. A standard deviation of 1.5% as we find for the population of employees may not seem overwhelming in economic terms. Second, if this variation is significant, Eq. (1) may be misspecified because it does not take into account local conditions forming the local wage policy. Thus, we run the following regression for each year

$$\mathrm{w_{ij}} = \mathrm{x_{ij}}\beta + \Omega_\mathrm{j} + \eta'_{\mathrm{ij}} \tag{2}$$

where Ω_j is a dummy variable for municipality j, and η'_{ij} is the residual.

Table 3 reports the F-statistics for the null hypothesis of no municipal specific effects. The municipal specific effects are clearly jointly significant in all years, implying that the model in Eq. (1) is rejected in favor of Eq. (2). Most interestingly, however, is the finding that the F-value makes a jump in 1990, the reform year, and in 1993. The economic importance of the municipal specific effects also seems to have increased over time. The standard deviation of the municipal specific effects is clearly higher after the 1990 reform than before the reform with the exception of the year 1987.

[4] The effect of the included variables in this regression and the regressions below are as expected. The wage increases with educational level and age, and the (full-time equivalent) wage is reduced by working part-time. The wage is lower for old full time working women than men, but women loose less from working part-time and their age profile is flatter. The differences between men and women probably reflect differences in job types. Women is relatively numerous in health care and care for the elderly, while men are to a large extent working with infrastructure and administration.

[5] The wage regression explain 50–60 percent of the wage variation, and the coefficient of determination is falling over time. Investigating the 50/10 and 90/50 percentile gaps, the conditional wage inequality at the lower end of the log wage distribution is, interestingly, about 1 percent higher than the unconditional distribution. At the upper end of the distribution, however, the conditional distribution is clearly lower than the unconditional distribution, and it is stable over time.

Table 3 Estimated municipal specific effects

Year	Observations[1]	F-test for no municipal specific effects, $DF \approx (440,\infty)$	Std. dev., weighted	Corrected std. dev., weighted[2]
1986	445	4.4	0.0136	0.0122
1987	452	5.0	0.0153	0.0139
1988	446	5.4	0.0141	0.0129
1989	446	5.3	0.0130	0.0119
1990	446	9.3	0.0158	0.0151
1991	446	8.2	0.0142	0.0134
1992	437	9.3	0.0152	0.0144
1993	438	13.5	0.0180	0.0174
1994	434	13.8	0.0179	0.0173
1995	434	12.0	0.0165	0.0159
1996	434	13.0	0.0167	0.0161
1997	434	13.9	0.0176	0.0170
1998	434	15.0	0.0166	0.0161

[1] In 1986–1992, one municipality is missing, and in 1986, seven additional municipalities are missing. The reduced number of municipalities in 1988, 1992 and 1994 are due to merging
[2] Standard deviation of the municipal specific effects corrected by the method of Haisken–DeNew and Schmidt (1997)

Because the municipal specific effects are estimated, the weighted standard deviation across the groups is an upwardly biased estimate of the true standard deviation as pointed out by Krueger and Summers (1988). Group-specific coefficients have a sampling error because they are estimated. Haisken-DeNew and Schmidt (1997) shows that by renormalizing the coefficients such that their weighted average is equal to zero, the true standard deviation can be straightforwardly calculated. The standard deviations of the municipal specific effects using their method are presented in Table 3. This measure makes a jump in the reform year, and is higher after the reform than before the reform. It also increases during the post-reform period. Thus, there is some support for the hypothesis that the reform in the wage system induced more wage variation between municipalities. In the following we investigate in a systematic way to what extent this increased variation is due to local labor market factors and variables describing the fiscal and political situation of the local governments as discussed in Section 3.

5 Econometric model

This section presents the measurement of the explanatory variables at the local government level and discusses estimation issues.

5.1 Measurement of additional explanatory variables

Definitions of the all variables included in the empirical model and descriptive statistics are reported in Appendix Table A1. In addition to the variables discussed below, the model includes the same individual characteristics as in Section 4.

5.1.1 Local government budget

The sum of the exogenous income tax, wealth tax, and unconditional grant on per capita form is included as the *local government income* variable. The real value of the income depends on the payroll tax rate, which to some extent vary across the local governments and over time. In the case where the exogenous income is the only income source and employment is the only input in the production, the budget constraint of a municipality can be written $R=w(1+t)L$, where R is income per capita, t is the payroll tax rate and L is employment per capita. The budget constraint can also be written $R^*=R/(1+t)=wL$, where R^* is the after tax income. To simplify the model, this income measure is used in the following. There is clearly a scope for an economically important budget effect because the standard deviation of the budget measure used in the empirical investigation below is about 1/3 of the mean value.

5.1.2 Local unemployment rate

Two measures of local labor market conditions are considered. First, we include the municipal *unemployment* rate. The municipal unemployment rate varies in the data from zero to 0.26. It is an open question whether wages respond stronger to the unemployment rate in a wider geographic area then the municipality, and in sensitivity analyses we consider the unemployment rate in the labor market regions defined by Statistics Norway based on the commuting statistics described below, including at average 4.8 municipalities.

5.1.3 Monopsony power

To capture monopsony effects, we need a measure of the possibility to commute to another workplace, that is, another municipality. Our idea is that the attractiveness and availability of jobs in other municipalities depends on the commuting costs. With high commuting costs, the labor supply schedule will be relatively steep, and accordingly there is a relatively high degree of monopsony power. We measure commuting possibilities using census data of residence and working place in 1990. Actual commuting across local government borders are expected to be closely related to the possibilities of changing job without changing residence. The commuting possibilities in municipality j can be approximated either by the share of the total number of employees (both in the public and private sectors) *living* in municipality j and working in another municipality or alternatively the share of the employees *working* in municipality j and living in another municipality. We use the mean of these separate measures of commuting out of and into the municipality.

$$(\textit{Index of monopsony power})_{\mathrm{k}} = \frac{1}{2}\left(\mathrm{E}^{\mathrm{r}}_{\mathrm{kk}} \Big/ \sum_{\mathrm{j}=1}^{\mathrm{J}} \mathrm{E}^{\mathrm{r}}_{\mathrm{jk}} + \mathrm{E}^{\mathrm{w}}_{\mathrm{kk}} \Big/ \sum_{\mathrm{j}=1}^{\mathrm{J}} \mathrm{E}^{\mathrm{w}}_{\mathrm{jk}}\right) \qquad (3)$$

where E^r_{jk} $\left(E^w_{jk}\right)$ is the number of employees in municipality j residing (working) in municipality j and working (residing) in municipality k. This index is positively related to monopsony power. Since this information is available for one year only, there is no time variation in the variable, and consequently the variable cannot be included in the empirical model together with fixed employer effects. Thus, an important caveat is that our index of monopsony power may pick up some general time invariant attractiveness of the municipality, which may be negatively related to the index.

5.1.4 Fiscal competition

Similar to the case of monopsony, the degree of fiscal competition will depend on mobility possibilities, i.e., the possibility to change resident without changing job. Thus, it may be difficult to disentangle the effects of fiscal competition and monopsony power empirically. But while monopsony power is independent of whether the mobility is easy to one large or many small municipalities, the degree of fiscal competition depends in addition on the number of competing local governments. Thus, for fiscal competition, we use an inverse Herfindahl index calculated as

$$(\mathit{Index\ of\ fiscal\ competition})_{\mathrm{k}} = 2\left(\sum_{\mathrm{j=1}}^{\mathrm{J}}\left(\mathrm{E}_{\mathrm{jk}}^{\mathrm{r}}\Big/\sum_{\mathrm{j=1}}^{\mathrm{J}}\mathrm{E}_{\mathrm{jk}}^{\mathrm{r}}\right)^{2}+\sum_{\mathrm{j=1}}^{\mathrm{J}}\left(\mathrm{E}_{\mathrm{jk}}^{\mathrm{w}}\Big/\sum_{\mathrm{j=1}}^{\mathrm{J}}\mathrm{E}_{\mathrm{jk}}^{\mathrm{w}}\right)^{2}\right)^{-1} \tag{4}$$

This index takes into account both the size of the commuting out of and into municipality k, and the number of municipalities for which there is “major” commuting. The theory suggests that the wage is negatively related to the index as a higher index means more fiscal competition. Notice that even though this measure also is based on data from the 1990 census, there is some variation over time due to merging of municipalities. When two municipalities merge, the fiscal competition for the surrounding municipalities is reduced.

5.1.5 Strength of political leadership

Our measure of political strength is an Herfindahl index of party fragmentation in the local council.

$$(\mathit{Index\ of\ political\ strength})_{\mathrm{k}} = \sum_{\mathrm{p=1}}^{\mathrm{P}}\left(\mathrm{R}_{\mathrm{pk}}\Big/\sum_{\mathrm{p=1}}^{\mathrm{P}}\mathrm{R}_{\mathrm{pk}}\right)^{2} \tag{5}$$

where R_{pk} is the number of representatives of party p in the local council in municipality k. Increased party fragmentation reduces the index, and is expected to reduce the possibilities for the political leadership to oppose pressure from interest groups. Accordingly, this variable can be interpreted as measuring the strength of the political leadership.[6]

[6] Other variables measuring political structure are tested in Falch and Strom (2003). They find similar effects of several variables intended to measure political strength.

5.1.6 Sociodemographic variables

Finally, in addition to the variables discussed above, the model includes the population size, variables describing population composition and settlement pattern, and indexes for centrality, pay roll tax zones and industry structure as defined in the Appendix. The former variables are intended to pick up variation in the relative size of the different sectors within local governments, while the latter variables are included to control for some characteristics of the external labor market. Notice that it may be important to include all relevant characteristics of the municipalities in models without fixed municipal effects as discussed below.

5.2 Estimation issues

Assume that the logarithm of the wage of individual *i* in municipality *j* in year *t* is given by

$$\mathrm{w_{ijt}} = \beta_{\mathrm{t}}\mathrm{X_{ijt}} + \rho_{\mathrm{t}}\mathrm{V_i} + \theta_{\mathrm{i}} + \varepsilon_{\mathrm{ijt}} + \alpha\mathrm{Y_{jt}} + \mu\mathrm{Z_j} + \phi_{\mathrm{j}} + \nu_{\mathrm{jt}} + \varphi_{\mathrm{t}} \tag{6}$$

X_{ijt} and V_i are observable individual characteristics that vary and are constant over time, respectively, θ_i and ε_{ijt} are unobservable individual characteristics that are constant and vary over time, Y_{jt} and Z_j are time varying and time independent observable municipal characteristics, and ϕ_j and ν_{jt} are unobservable municipal characteristics that are constant and vary over time, respectively. ϕ_t captures all factors that only vary over time. β_t, ρ_t, θ_i, α, μ, ϕ_j, and φ_t are parameters to be estimated.

We are able to take into account unobservable individual effects because the data track employees and their local government over time. In order to estimate Eq. (6), we take deviation from individual mean to reduce the dimensionality of the model.

$$\begin{aligned}(\mathrm{w_{ijt}} - \overline{\mathrm{w}}_{\mathrm{i}}) = \beta_{\mathrm{t}}(\mathrm{X_{ijt}} - \overline{\mathrm{X}}_{\mathrm{i}}) + \alpha(\mathrm{Y_{jt}} - \overline{\mathrm{Y}}_{\mathrm{i}}) + \mu(\mathrm{Z_j} - \overline{\mathrm{Z}}_{\mathrm{i}}) \\ + (\phi_{\mathrm{j}} - \overline{\phi}_{\mathrm{i}}) + (\varphi_{\mathrm{t}} - \overline{\varphi}_{\mathrm{i}}) + \eta''_{\mathrm{ijt}}\end{aligned} \tag{7}$$

$\overline{w}_i$ is the mean wage of individual *i* during the period he is represented in the data, the mean of the other variables are defined in the similar way, and $\eta''_{ijt} = \varepsilon_{ijt} + \nu_{jt} - \overline{\varepsilon}_i - \overline{\nu}_i$ is the residual.

At the outset, the model in Eq. (7) has a particular attractive feature. Controlling for fixed individual effects is appropriate in order to distinguish between pure wage policy and sorting mechanisms. A local government may have a high wage because it is attractive, and thereby attract employees with high unobservable skills. Our model completely control for sorting to the extent that employee sorting is independent of individual time varying variables omitted from the model. In addition, in the case of fixed employer effects o_j, the bias in the standard errors of the estimators discussed by Moulton (1987) is absent. The Moulton bias arises if the residuals are correlated within the municipalities, which are taken into account by the employer specific effects.

Notice that in Eq. (7) ϕ_j is identified by employees moving between municipalities. Thus, a necessary assumption for consistent estimators with ordinary least square is that the term $\left(\phi_j - \overline{\phi}_i\right)$ is exogenous with respect to the wage.[7] Abowd et al. (1999) use a similar model for the French private sector under the assumption of exogenous mobility, and their data include thousands of firms. In that case, Eq. (7) cannot be estimated by OLS because of the high dimensionality. In our case of about 440 local governments, however, the number of parameters in Eq. (7) is reasonable finite, and the preferred model can be estimated.

To the extent that mobility is endogenous, the municipal fixed effects will be correlated with the error term and hence produce biased estimates. If this is the case the reduced form wage equation should not include the municipal specific effects f_j. On the other hand, in this case, it is important that all relevant exogenous characteristics of the municipalities are included in the model in order to avoid inconsistent estimates due to omitted variables. Thus, whether or not to include municipal fixed effects is not easy to decide a priori. Below, we will present the results from different specifications of Eq. (7) along with a discussion of the shortcomings of different estimation methods. Thereafter, we investigate whether the wage formation changed after the reform in the wage-setting system in 1990.

A simple way to investigate reform effects is to use a dummy variable approach, where the municipal characteristics are interacted with a dummy variable equal to zero before 1990 and equal to unity thereafter. However, a serious shortcoming of this method is that it ignores that adjustments to reforms usually are smooth processes towards a new equilibrium. The effect of a reform is likely to increase the first years after the reform. For the Norwegian local government wage formation, we expect increased response to the reform over time, following the arguments in Section 3 above. Our second method to test whether the response to the variables of interest changed after the reform is inspired by the nonlinear model formulation of Teräsvirta (1998). A flexible approach is to interact the variables of interest with a parametric transition function, $\overline{S}$, defined as

$$\overline{\mathrm{S}} = \frac{1}{1 + \mathrm{e}^{-\gamma\left(\overline{\mathrm{T}} - \mathrm{c}\right)}} \tag{8}$$

This functional form assumes convergence to a new equilibrium and allows the adjustment to take place from the start of the estimation period. The solid line in Fig. 1 illustrates the logistic smooth transition function defined by Eq. (8). One advantage of such a function is that it can be described by only two parameters γ and c. The transition variable $\overline{T}$ is a trend, c is the switching point defined by the year in which the transition to the new equilibrium is fastest, and γ determines the slope of the transition function at the switching point (γ>0). $\overline{S}$ is bounded between zero and unity.

[7] In the data there are 383,912 employees, for which about 80,000 are present only one single year and therefore yield no information when we control for fixed individual effects. 13.5% of the employees change municipality at least once during the empirical period, which implies that the employer fixed effects are not perfectly collinear with the individual fixed effects. As the employment duration increases, the number of employees in the data decreases (to about 18,000 for duration of 12 years). As the number of years an employee is in the data increases, the share of the employees changing employer at least once increases, with a maximum for a duration of 9 years (17%).

However, due to our institutional knowledge, we restrict the transition function such that the transition to a new equilibrium cannot start prior to the reform year 1990. Thus, the transition variable T is set equal to zero before the reform and is a linear trend thereafter. The transition function used is scaled such that it is bounded between zero and unity,

$$S = \frac{1+k}{k}\left(\frac{1}{1+e^{-\gamma(T-c)}} - \frac{1}{1+k}\right) \tag{9}$$

where $k = e^{\gamma c}$. The transition function S given by Eq. (9) is illustrated by the dotted line in Fig. 1. Notice that the dummy variable approach is a special case of Eq. (9), namely $\gamma \rightarrow \infty$ and $c=0$. For $\gamma \rightarrow 0$, the transition function is approximately linear.

The model estimated to investigate reform effects is

$$w_{ijt} = \beta_t X_{ijt} + \rho_t V_i + \theta_i + \alpha_1 Y_{jt} + \mu_1 Z_j + (\alpha_2 Y_{jt} + \mu_2 Z_j)S + \phi_j + \varphi_t + \eta_{ijt} \tag{10}$$

The effect of a variable a particular year, say the effect of Y, is equal to $(\alpha_1+\alpha_2*S)$. α_2 measures the reform effect, that is the change in the effect of Y when the transition to the new equilibrium is completed (S=1). Thus, the model formulation allows the reform effect to differ between the variables, but the model restricts the transition path to the new equilibrium to be equal for all variables.

6 Estimation results

Before analyzing the effect of the wage-setting reform, we present the results from wage equations with constant parameters during the empirical period in order to check the sensitivity of model specification. Results from different specifications of Eqs. (6) and (7) are presented in Table 4. In addition to the reported variables, all models include the age composition of the population, settlement pattern, and dummy variables for centrality, pay roll tax zone, and industry structure. Column

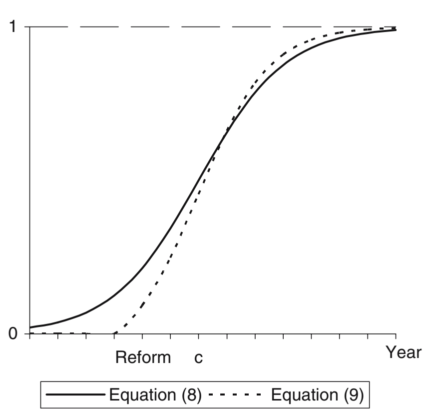


Fig. 1 Smooth transition function

Table 4 Estimation results

	(1)	(2)	(3)	(4)
log(local government income/ (1+payroll tax rate))	0.016 (22.0)	0.017 (33.1)	0.003 (8.67)	−0.000 (0.63)
Unemployment	−0.079 (11.2)	−0.068 (14.1)	0.003 (1.18)	0.001 (0.35)
Index of political strength	0.015 (7.89)	0.025 (19.0)	0.014 (13.8)	0.010 (9.17)
log(index of monopsony power)	0.008 (4.11)	−0.021 (14.7)	−0.012 (5.98)	–
log(index of fiscal competition)	0.019 (12.9)	−0.002 (2.46)	−0.014 (8.93)	−0.052 (6.30)
log(population)	0.003 (17.5)	0.005 (34.8)	−0.002 (7.69)	0.017 (13.9)
Time specific effects	Yes	Yes	Yes	Yes
Observable individual characteristics	No	Yes[1]	Yes[1]	Yes[1]
Individual specific effects [No. of individuals]	No	No	Yes [302,779]	Yes [302,779]
Municipal specific effects [No. of municipalities]	No[2]	No[2]	No[2]	Yes[3] [452]
Number of observations	2,174,062	2,174,057	2,093,598	2,093,598
R^2	0.072	0.562	0.947	0.947
Equation standard error	0.14656	0.10072	0.03775	0.03765

Empirical period is 1986–1998, *t*-values in parentheses
[1] Individual characteristics included are two dummy variables for part time work, age categories and educational categories. The variables are defined in the Appendix. The coefficients are restricted to be equal in two subsequent years (1986 and 1987, 1988 and 1989, etc.), and the effects depend on gender in a time invariant way
[2] Municipal control variables included are the share of children, youth and elderly in the population, the share of the population living rural, and dummy variables for centrality (seven categories), pay roll tax zone (six categories) and industry structure (22 categories)
[3] Municipal control variables included are the share of children, youth and elderly in the population. The additional municipal control variables used in the other model specifications have no time variation

(1) includes only observable municipal characteristics and time specific effects. The income elasticity is 0.016, and the unemployment elasticity at mean value of unemployment is −0.003.[8] Both effects are numerically small, but highly significant. Both population size and the indexes of political strength, fiscal competition and monopsony power have a positive effect. The effects of the latter two variables change sign in more comprehensive specifications, which indicates that the unexpected sign in column (1) is a result of misspecification.

The effect of local government income is in accordance with our hypothesis. The income effect can be a result of union power in local wage bargains, and it is therefore relevant to compare the estimated effect with the amount of rent-sharing in the private sector. The estimated effect is of the same magnitude as the effect of profit found by Christofides and Oswald (1992), and somewhat lower than the effect found by Blanchflower et al. (1996). According to our results in the specifications without fixed individual effects, a rise of one standard deviation in income (that is 32%) increases the wage by 0.5%. The unemployment effect is well

[8] Unemployment is not included at logarithmic form because it is equal to zero in some observations.

below the usual findings for the private sector, supporting the view that public sector wages are less responsive to unemployment than private sector wages [Gregory and Borland (1999)]. The small effect may also reflect that the wages respond to unemployment in a larger geographical area. We have replaced municipal unemployment with unemployment in the labor market regions defined by Statistics Norway. With this alternative unemployment measure, the effect increases by about 75%, but is still very small.

Column (2) shows the results when observable individual characteristics are included. With this specification, the wage tends to be highest in municipalities with low monopsony power and a low degree of fiscal competition. In column (3) individual fixed effects are added to the model. Now the size of the income effect drops, and the unemployment effect disappears.[9] This may imply that the coefficients of these variables in column (1) and (2) are biased because of omitted unobservable individual characteristics correlated with unemployment and income. But it also illustrates the problem of identifying the effects of employer characteristics within this class of models. Local government income and unemployment are likely to be autocorrelated and evolving in a more or less common way over time across municipalities. When individual fixed effects and time dummies are included there may simply not be enough variation left in the data to identify the effects of these variables. If the remaining variation to some extent is measurement error, this may bias the coefficients towards zero [Hamermesh (2000)]. In addition, with individual fixed effects, the remaining variation across municipalities is due to workers moving between municipalities. Thus, unbiased estimators require that the change in the workforce in a municipality is uncorrelated with the explanatory variables, which is a quite restrictive assumption.

The full model in column (4), including both individual and municipal fixed effects, assumes that mobility is exogenous with respect to the wage. Then the only variation left to identify the effects of local characteristics is municipal specific development in these variables, conditional on observed and unobserved individual effects. Overall, it is quite possible that the results in columns (1) and (2) overstate the true responses to local characteristics, while the specifications with fixed individual effects understate the responses.

The negative effect of the index of monopsony power found in column (2) and (3) indicates that wages to some extent depend on labor market conditions as expected. It is not possible, however, to reveal whether the effect is a pure monopsony effect or whether there is an effect via local wage bargaining. The results regarding the monopsony power variable must, however, be interpreted with care because the index does not vary over time. Accordingly we are not able to identify the effect of this variable in the full model with municipal fixed effects in column (4). Further, in this model specification, the effect of our index of fiscal competition is identified only through mergers of municipalities during the period. Given these caveats, however, it is interesting that both fiscal competition and monopsony power has the expected effects in all models except the naïve specification in column (1). According to the results in column (3), a rise of one standard deviation in these indexes, that is 19 percent for monopsony power and 34% for fiscal competition, reduces the wages by 0.23 and 0.48%, respectively.

[9] Using the unemployment in the labor market region, the effect becomes negative but is still insignificant.

The most robust result across the specifications is the positive effect of the index of political strength of the local council. The result should be compared with studies of local government spending. Such studies typically find that a strong political leadership is able to hold down spending. Our results indicate that this must be due to reduced employment or non-wage spending, and that some of the reduced spending on these items is used to pay higher wages as discussed in Section 3 above.

In our view, the above discussion also indicates that there is no easy and obvious way to estimate the effects of employer specific variables in panel data models of wage formation. In the econometric literature, fixed effects specifications are usually recommended, but the amount of unit specific time variation in the data required to identify the effects of employer specific variables in such specifications may not to be available in many real world applications. Thus, when investigating reform effects below, we choose to present the results from different specifications.

6.1 Reform effects

Did the response to local characteristics increase after the 1990-reform? Initially, we focus on the fixed individual effect model, column (3) in Table 4. A nonlinear estimation method is not possible to perform directly in a fixed effect model. Thus, γ and c in Eq. (10) are estimated by a two-dimensional grid search method.[10] The parameters minimizing the standard error of the equation are γ=2.5 and c=3.0. [11] The behavioral change estimated by the smooth transition approach is illustrated for two of the variables in Fig. 2. The transition to the new equilibrium mainly occurred between 1991 and 1993, and according to the model, 99% of the adjustment towards the new equilibrium was completed within 1994. Notice that the modeling approach restricts the transition period to be equal for each variable in the model, but the reform effects (the difference between the pre- and post-reform effects) are freely estimated for each variable.

The results from the smooth transition approach are reported in column (2) in Table 5. For comparison purposes, column (1) presents the results from the dummy variable approach. The two approaches yield very similar results for all variables except unemployment. The income elasticity increased from 0.02 in the pre-reform period to 0.05 or 0.07 in the new equilibrium according to the dummy variable approach and smooth transition approach, respectively. The income effect in Table 4, assuming constant effect during the whole empirical period, can loosely be seen as a weighted mean of the pre- and post-reform effects. The same pattern of the reform effect on the wage response holds for most of the other local characteristics in the model. Both the effects of political strength, monopsony power and fiscal competition increased after the reform, from a level below the "average" responses estimated in Table 4 to a level above the "average" responses. This evidence supports our initial expectations that the wage system reform would make wages more

[10] The values used for γ in the grid are 0.2, 0.5, 1, 1,5, ..., 10, and the values used for c are 0, 1, ..., 10.

[11] This implies that S = 0.01, 0.07, 0.50, 0.93, and 0.99 in 1990, 1991, 1992, 1993, and 1994, respectively.

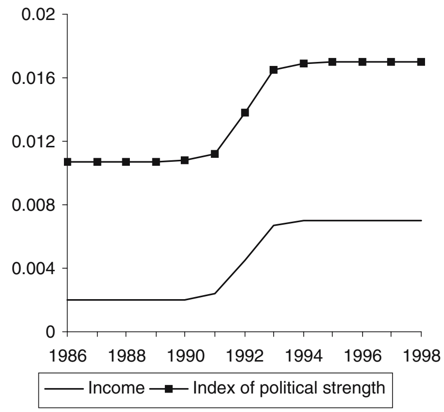


Fig. 2 The transition of the effects of income and political strength

dependent on local characteristics. As to the effect of unemployment, the dummy variable approach and the smooth transition approach give a mixed picture. In the former approach, the responsiveness to unemployment is weak and insignificant, while in the latter approach, there is a shift from a negatively significant effect prior to the reform to a positive effect after the reform. Although the reform effect is significantly positive, the unemployment effect in the post-reform period is not significantly different from zero at 5% level.

Is this result robust with respect to model specification? Column (3) Table 5 presents the results from the specification with only measured individual and municipal variables included in the model, using the same transition function as in column (2). Consistent with the earlier results, the effects of unemployment and local government income are much higher in this specification. More important, however, the response to all variables increased after the reform in this specification, and the signs of the reform effects are equal to the former model specification for all variables except unemployment. Column (4) presents the results when the smooth transition approach is applied to the full model including both individual and employer specific effects. This model also gives the same qualitative effect of the reform. The sizes of the reform effects differ somewhat across the specifications, but the conclusion that the reform increased the responsiveness to most local characteristics is independent of model specification.

6.2 Newly hired workers versus incumbent workers

As pointed out in Sections 2 and 3, the new system introduced in 1990 has merely been in full effect only when new employees are hired. Accordingly, in a given time period, the wage response to local characteristics may differ for employees hired during the period and employees that have stayed in the same local government for several time periods. If the wage formation is equal for the two groups, this indicates that the relationship between wages and local characteristics is equal in the hiring process and in the promotion and local bargaining processes. If, however, the response to local characteristics is strongest for newly hired employees, there is a kind of overshooting. To illustrate this point, consider a rich

Table 5 Reform effects

Regression	(1)[1]		(2)[1]		(3)[2]		(4)[3]	
Method	Dummy variable approach		Smooth transition approach		Smooth transition approach		Smooth transition approach	
Effects	Pre-reform effects	Reform effects	Pre-reform effects	Reform effects	Pre-reform effects	Reform effects	Pre-reform effects	Reform effects
log(local government income/(1+payroll tax rate))	0.002 (6.50)	0.003 (6.86)	0.002 (5.93)	0.005 (9.95)	0.009 (14.8)	0.018 (18.4)	−0.001 (3.55)	0.002 (3.34)
Unemployment	−0.001 (0.17)	−0.005 (1.15)	−0.012 (3.31)	0.019 (3.95)	−0.043 (6.24)	−0.070 (6.60)	−0.004 (1.22)	0.013 (2.63)
Index of party fragmentation	0.008 (6.04)	0.010 (7.59)	0.011 (9.29)	0.006 (4.61)	0.007 (3.44)	0.027 (9.85)	0.010 (7.76)	0.001 (0.40)
log(index of monopsony power)	−0.008 (3.66)	−0.006 (4.11)	−0.009 (4.18)	−0.006 (4.05)	−0.015 (7.17)	−0.008 (2.70)	–	–
log(index of fiscal competition)	−0.012 (7.28)	−0.002 (1.91)	−0.012 (7.48)	−0.004 (3.96)	−0.001 (0.94)	−0.003 (1.20)	−0.078 (9.32)	−0.001 (4.04)
log(population)	−0.003 (14.1)	0.002 (16.8)	−0.004 (17.9)	0.002 (23.5)	0.001 (7.56)	0.006 (23.1)	−0.009 (5.51)	0.003 (24.6)
γ	∞		2.5		2.5		2.5	
c	0		3.0		3.0		3.0	
Time specific effects	Yes		Yes		Yes		Yes	
Observable individual characteristics	Yes		Yes[1]		Yes		Yes	
Individual specific effects [No. of individuals]	Yes [302,779]		Yes [302,779]		No		Yes [302,779]	
Municipal specific effects [No. of municipalities]	No		No		No		Yes [452]	

Table 5 (continued)

Regression	(1)[1]		(2)[1]		(3)[2]		(4)[3]	
Method	Dummy variable approach		Smooth transition approach		Smooth transition approach		Smooth transition approach	
Effects	Pre-reform effects	Reform effects	Pre-reform effects	Reform effects	Pre-reform effects	Reform effects	Pre-reform effects	Reform effects
Number of observations	2,093,598		2,093,598		2,174,062		2,093,598	
R^2	0.947		0.947		0.562		0.947	
Equation standard error	0.03775		0.03774		0.10069		0.03763	

Empirical period is 1986–1998, *t*-values in parentheses

[1] Specification of column (3) Table 5 extended with the reform effects

[2] Specification of column (2) Table 5 extended with the reform effects

[3] Specification of column (4) Table 5 extended with the reform effects

local government hiring a worker in a given time period. Because governments with high income have more ability to pay, this employee will receive a high wage relative to other workers in that particular time period. Under overshooting, this income effect will decline once the employee becomes an incumbent worker, reducing the wage premium over time. The next question is why such an overshooting may take place. A possible explanation is that local unions dominated by incumbent workers with high seniority may pursue a wage equalization policy with respect to newly hired workers and incumbents. To implement such an equalization policy the union may simply prevent new employees being hired at a relatively high wage from receiving wage increases through local bargaining in the following years.

To investigate empirically these questions we divide the sample into two subsamples; newly hired workers and incumbent workers. We define incumbent workers as employees working in the same local government in two consecutive years. Newly hired workers are defined as employees recruited during the last year.[12] We will focus on a model specification without fixed effects because the vast majority of the employees in the newly hired workers subsample are observed in this subsample only once. A grid search is performed for each sample to allow the transition to a new equilibrium to differ across the groups, but the transition path is restricted to be equal for each variable within the subsamples.[13]

The results are shown in Fig. 3, and details are delegated to Appendix Table A2. The model for incumbent workers is very similar to the model using the whole sample. This is as expected because the major part of the employees is incumbents. Newly hired workers differ from incumbents in two ways in their response to local characteristics. First, government income and political strength have strongest effect on the wages of newly hired employees both before and after the 1990-reform. This evidence suggests some kind of overshooting in the wage setting process as discussed above. As the newly hired workers become incumbents, the responses to these factors are reduced. Political strength seems in particular to have a large effect in the hiring process, indicating that local governments with a strong political leadership use wages more actively in the recruitment policy than municipalities with a weak political leadership.

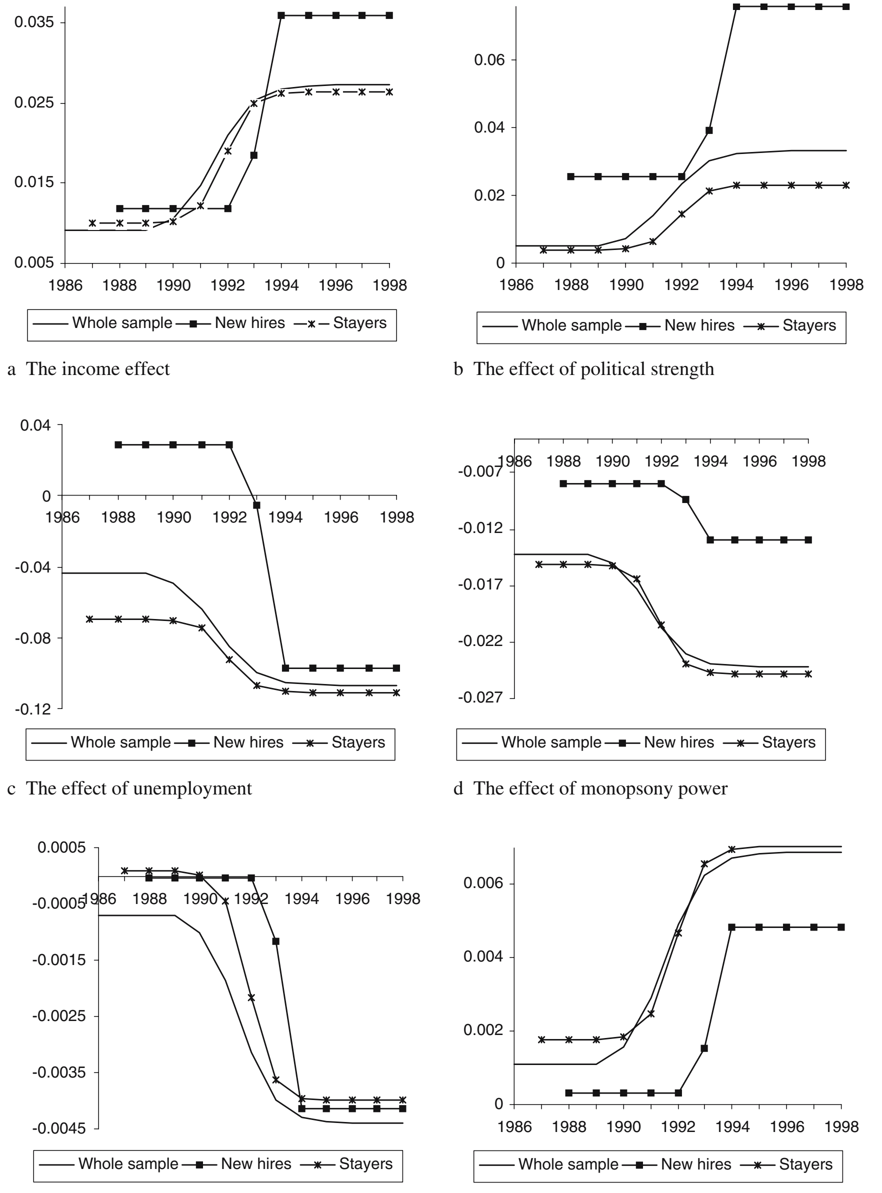
Second, the response to local labor market conditions is weak for newly hired employees. This probably reflects that new employees are to some extent hired from a wider region than the local labor market, making the state of the local labor market less important in the hiring process. When the new employees become incumbents, the local labor market seems to grow in importance.

7 Discussion and concluding comments

The purpose of this paper was to examine the effects of increased formal flexibility in wage setting in the local public sector using a change in the wage setting system

[12] Because we do not know the employment history of the employees prior to our empirical period, the first year of the data (1986) must be excluded from both subsamples. For the newly hired workers subsample, we have also excluded 1987 to avoid mistaking employees on leave in 1986 as hired in 1987.

[13] Because the model in this section does not include fixed effects, a denser grid is used in the grid search. The values used for γ in the grid is 0.1, 0.2, 0.3, ..., 10, and the values used for c is 0, 0.1, 0.2, ..., 10.



a The income effect

b The effect of political strength

c The effect of unemployment

d The effect of monopsony power

e The effect of fiscal competition

f The effect of population

Fig. 3 Different responses on the wages of newly hired workers and incumbent workers: (**a**) The income effect; (**b**) the effect of political strength; (**c**) the effect of unemployment; (**d**) the effect of monopsony power; (**e**) the effect of fiscal competition; (**f**) the effect of population

in Norway in 1990. Decisions on public sector pay have traditionally been highly centralized. Delegation of decisions on salary issues to lower level institutions has been an important element in recent public sector reforms in many countries. Standard textbook reasoning suggest that decentralization would increase pay differences and be efficiency enhancing as decentralized decisions on pay issues moves the local wages toward their competitive levels. However, both common observation and previous research suggest that public sector labor markets are not well described by the competitive model. Union influence and monopsony power are frequently considered as important in the public sector. Thus, the actual outcome from decentralization is an empirical question. Yet, existing evidence on the actual outcomes from such shifts in wage setting responsibility is scarce.

Since budget size is an important determinant of labor demand for local government employees, we expected a stronger positive relationship between local wages and the size of the local government budget as wage setting were decentralized. Further, we expected that decentralization would make wages more responsive to local labor market conditions represented by the local unemployment rate and a measure of local monopsony power.

Since local governments are ruled by politically elected representatives, the strength of the political leadership relative to employee organizations and interest groups may affect the wage outcomes in a decentralized system. Accordingly, we expected that the relationship between local wages and measures of political strength of the political leadership became more important after decentralization. In addition, wage bargaining models suggest that product market competition exerts a downward pressure on wages. Fiscal competition between local governments may serve a similar function in the local public sector. Therefore we expected that wages became more responsive to measures of fiscal competition after the decentralization of wage setting.

Adjustments to reforms generally take time, and in a case with norms of pay equality across local governments established within the centralized system we expected that the response to local characteristics increased gradually through time. We also expected that the new wage system would be most important for the wages of new hires as national limits on the size of wage increases to be given in formal negotiations were in place throughout the period investigated.

This paper explores a large matched individual-municipal data set covering 1986-1998 to investigate these hypotheses of the effect of increased formal flexibility in local public sector wage setting in Norway from 1990. Consistent with our expectations, we find evidence that the wage differences increased somewhat as the wage setting was decentralized, and that local wages become more responsive to the local budget. But we also find evidence consistent with the hypothesis that introduction of more flexibility in wage setting increased the effect of monopsony power, fiscal competition and variables characterizing the strength of the local political leadership. However, contrary to our expectations, wages seem to be largely unaffected by regional labor market pressure represented by the local unemployment rate both before and after the reform. Our results support the hypothesis that adjustment to the system reform took place gradually through time. We also find evidence that the response to the local budget and the political strength of the local council is higher for newly hired employees than for incumbents. However, the numerical effect of these variables decreased significantly when the new workers became incumbents.

Taken together, our empirical evidence indicates that decentralization of wage setting in the public sector produces outcomes that in some respects deviate from the competitive model. Moreover, as the changes in wage distribution across local governments are fairly small, it suggests that established pay equality norms across government workplaces continue to play an important role also in the decentralized regime.

Thus, our evidence suggests that changing the formal system may not itself be sufficient to induce the local actors to react to local market conditions in the desired way since the actual wage setting process in the local governments seems to be influenced by local unions, local interest groups, local monopsony power and pay equality norms, which change only slowly through time. The results should be of interest to other countries that already have, or intend to implement reforms in public sector pay systems. Although institutional settings clearly differ across countries, our results suggest that common arguments on the expected efficiency gains from increased local flexibility in wage setting in public institutions may be exaggerated. On the methodological side this paper seems to be one of the first to study how the introduction of more flexibility in public sector wage setting affects the relationship between individual wages and local characteristics using panel data that makes it possible to control for both observed and unobserved individual and employer effects. More research using this approach in different fiscal systems is warranted to provide more definitive empirical evidence on the likely effects of pay system reforms in the public sector.

Appendix

The sample is restricted to employees working above 20% of full time in the welfare services provided by the municipalities. This covers about 55% of the employees in the local public sector. The county authorities employ almost all the excluded employees. Seven municipalities are missing in 1986 and one municipality is missing in 1986–1992. In addition, the capital area (Oslo) is excluded due to a special institutional setup. Table A1 defines the variables used in the analysis.

Table A1: Data definition and sources

Definition and sources	Mean value	Standard deviation
Dependent variable. Source is the Federation of Local Governments		
Real wage, calculated as the monthly wage excess of supplement due to for example overtime and night work, divided by the share of full time job. Deflated by CPI.	16,032	2,720
Individual characteristics. Source is the Federation of Local Governments		
Master degree	0.02	
Engineer	0.02	
Education in business and administration	0.01	
College education	0.45	
High school education or less	0.42	
Education missing	0.08	
Age up to 19 years	0.002	
Age 20–24 years	0.02	
Age 25–29 years	0.08	
Age 30–39 years	0.26	

Definition and sources	Mean value	Standard deviation
Age 40–49 years	0.31	
Age 50–59 years	0.22	
Age above 60 years	0.10	
Man	0.23	
Working less than 50% of full time	0.13	
Working between 50 and 99% of full time	0.42	
Working full time	0.45	
Municipal characteristics. Source is the Norwegian Social Science Service.		
Local government income/(1+payroll tax rate). Local government income is calculated as the sum of income taxes, wealth taxes and unconditional grants on per capita form. Deflated by CPI.	15,521	4,905
Unemployment, measured in the spring. For 1986–1992, the variable is calculated as the number of registered unemployed in relation to calculated work force (the number of employees living in the municipality plus the number of registered unemployed). For 1993–1998, the unemployment rate is available.	0.04	0.02
Index of political strength, see the main text for definition.	0.26	0.07
Index of monopsony power, see the main text for definition.	0.77	0.15
Index of tax competition, see the main text for definition.	1.68	0.57
Population.	34,390	52,150
Children, the share of the population between 0 and 6 years of age.	0.09	0.01
Youth, the share of the population between 7 and 15 years of age.	0.12	0.02
Elderly, the share of the population above 80 years of age.	0.04	0.01
Share of population living rural. Rural is defined as outside towns and city centers with more than 2,000 inhabitants.	0.34	0.28
Centrality, an index including seven categories defined by commuting time to small, medium sized and big cities.		
Payroll tax zone. The payroll tax rate is equal within zones, and the number of zones varies over time between four and six.		
Industry structure, an index including 22 categories defined by the relative distribution on a five-digit industry measure.		

Table A2: Different response on new hires and incumbent workers

Column	(1)	(2)		(3)	(4)		(5)	(6)	
Sample	All workers			Newly hired workers[1]			Incumbent workers[2]		
Method	OLS	Smooth transition approach		OLS	Smooth transition approach		OLS	Smooth transition approach	
Effects	1986–1998 constant effects	Pre-reform effects	Reform effects	1988–1998 constant effects	Pre-reform effects	Reform effects	1987–1998 constant effects	Pre-reform effects	Reform effects
Log(local government income/ (1+payroll tax rate))	0.017 (33.1)	0.009 (13.8)	0.018 (18.0)	0.021 (13.0)	0.012 (6.18)	0.024 (8.18)	0.017 (29.8)	0.010 (13.4)	0.016 (14.9)
Unemployment	−0.068 (14.1)	−0.044 (5.95)	−0.063 (5.70)	−0.021 (1.56)	0.029 (1.78)	−0.126 (4.22)	−0.081 (15.1)	−0.070 (8.69)	−0.041 (3.51)
Index of political strength	0.025 (19.0)	0.005 (2.55)	0.028 (9.88)	0.054 (13.7)	0.026 (4.94)	0.050 (6.29)	0.019 (12.8)	0.004 (1.73)	0.019 (6.07)

Column	(1)	(2)		(3)	(4)		(5)	(6)	
Sample	All workers			Newly hired workers[1]			Incumbent workers[2]		
Method	OLS	Smooth transition approach		OLS	Smooth transition approach		OLS	Smooth transition approach	
Effects	1986–1998 constant effects	Pre-reform effects	Reform effects	1988–1998 constant effects	Pre-reform effects	Reform effects	1987–1998 constant effects	Pre-reform effects	Reform effects
Log(index of monopsony power)	−0.021 (14.7)	−0.014 (6.23)	−0.010 (3.23)	−0.011 (2.80)	−0.008 (1.47)	−0.005 (0.62)	−0.022 (13.7)	−0.015 (5.89)	−0.010 (2.85)
Log(index of fiscal competition)	−0.002 (2.46)	−0.001 (0.44)	−0.004 (1.70)	−0.001 (0.51)	−0.000 (0.01)	−0.004 (0.74)	−0.002 (1.69)	0.000 (0.05)	−0.004 (1.69)
Log(population)	0.005 (34.8)	0.001 (5.68)	0.006 (23.6)	0.003 (7.21)	0.000 (0.68)	0.005 (6.57)	0.005 (33.0)	0.002 (7.39)	0.005 (19.4)
γ	–	1.4		–	9.8		–	2.1	
c	–	2.5		–	4.1		–	2.9	
Time specific effects	Yes	Yes		Yes	Yes		Yes	Yes	
Observable individual characteristics	Yes	Yes		Yes	Yes		Yes	Yes	
Individual specific effects	No	No		No	No		No	No	
Municipal specific effects	No	No		No	No		No	No	
R^2	0.562	0.562		0.557	0.557		0.556	0.556	
Equation standard error	0.10072	0.10069		0.10807	0.10803		0.09867	0.09864	
Number of observations	2,174,057			293,299			1,691,053		

t-values in parentheses. The regression in column (1) is the same regression as column (2) Table 5

[1] Newly hired workers are defined as: (1) Employees that are new in the data; (2) Employees that have moved to a new municipality; (3) Employees that have been out of the data in more than three consecutive years. The two first years of the empirical period (1986 and 1987) are dropped. Category (1), (2) and (3) include 84, 12 and 4% of the sample, respectively

[2] Incumbent workers are defined as employees working in the same municipality in two consecutive years

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