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The Wage Impact of Trade Unions in the UK Public and Private Sectors

By David G. Blanchflower† and Alex Bryson‡

†*Dartmouth College, University of Stirling, IZA, CESifo, NBER and Bank of England*
‡*Policy Studies Institute and Centre for Economic Performance*

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There is a growing gap in the union membership wage premium between public and private-sector workers in the United Kingdom. Using the Labour Force Surveys of 1993–2006, the gap between the membership premium in the public and private sectors closes with the addition of three-digit occupational controls, although significant wage premia remain in both sectors. However, using the Workplace Employment Relations Survey of 2004, the public-sector union membership wage premium remains roughly twice the size of the private-sector membership premium, having accounted for workplace fixed effects and workers' occupations, job characteristics, qualifications and demographics.

Introduction

This paper draws attention to a trend hitherto unnoticed, namely an increase in the size of the union membership wage premium in the UK public sector relative to the private sector. This finding is surprising because it implies that union bargaining power has held up reasonably well in the public sector, despite the fact that union density has declined at a similar rate to the decline in the private sector. It is also surprising because it seems implausible that there should be a substantial union wage premium in the public sector. Unlike in the private sector, the majority of public-sector workers have pay set directly through collective bargaining or through pay review bodies, whether or not they are union members. In seeking to account for these findings, we explore a number of possible explanations.

We first establish the link between union membership and collective bargaining coverage in the two sectors to see whether membership is more closely aligned with coverage in the public sector, perhaps resulting in a higher premium. We consider the possibility that some workers misclassify themselves as public-sector workers when, in fact, they work in public-sector workplaces under contract to private-sector firms. If, as is likely, these 'misclassified' workers are more likely to be union non-members, this could account for a seemingly large public-sector union membership wage premium. We assess how much of the gap in the membership premium between the two sectors is accounted for by worker heterogeneity which is usually unobserved in more parsimonious wage models.

We focus on two potential sources of unobserved heterogeneity which may bias the union wage premium estimates. The first is the nature of employment contracts and, in particular, permanent versus temporary contracts and full-time versus part-time working. It is conceivable that, since such a high percentage of public-sector workers are union members, those that are not may be the most marginalized of workers in the sector—in contrast to the private sector, where most workers are non-members. Non-membership may therefore proxy particularly poor working conditions in the public sector relative to the private sector.

The second potential source of heterogeneity is the occupational distribution of workers by union status in the two sectors. We also use linked employer–employee data, to address the possibility that the union membership premium in both sectors reflects unobserved workplace heterogeneity rather than unionization *per se*. This investigation is prompted by the possibility that in the more highly unionized public sector non-members may be disproportionately located in workplaces with lower union density and/or poorer terms and conditions, relative to their member counterparts, leading to a larger union membership wage gap than in the private sector.

We find that the public and private-sector membership wage premia persist having controlled for a full range of individual, job and workplace characteristics. Using data from the Labour Force Survey, the gap between the membership premium in the public and private sectors closes with the addition of occupational controls. However, using data from the Workplace Employment Relations Survey 2004, the public-sector union membership wage premium remains roughly twice the size of the private-sector membership premium, having accounted for workplace fixed effects and workers' occupations, job characteristics, qualifications and demographics. Furthermore, the membership wage premium among workers covered by collective bargaining is apparent only in the public sector. We argue that these findings are worthy of further investigation, since they have public policy implications for wage setting.

I. Union Effects on Wages

A primary goal of trade unions is to maintain and improve workers' terms and conditions, particularly workers who are members of the union, through collective bargaining with employers. Whether unions are successful depends, in large part, on their bargaining strength—which is based on their ability to restrict the supply of labour to the employer—and on the ability of employers to concede above-market wages (Freeman and Medoff 1984; Blanchflower *et al.* 1990).

Unions' bargaining strength is enhanced by the percentage of all workers they represent, and leads to a higher union wage premium (Freeman and Medoff 1981; Lewis 1986; Forth and Millward 2002). Where the vast majority of workers in a given industry are covered by collective bargaining, union-negotiated wages have less impact on the employer's cost competitiveness than in instances in which competing employers have ready access to non-union labour. This is because above-market wage costs are faced by all competitors. Unions' success in raising wages is further enhanced if the price elasticity of demand for products or services in the industry is low, as might be the case where there is a monopoly or oligopolistic production, since employers are able to meet additional costs from above-normal profits or to pass the additional costs on to consumers without undue fear of being undercut by other producers.

It is normally assumed that the mechanism by which unions create a union wage premium is through their direct impact on covered workers' wages through pay bargaining. However, there is a variety of ways in which a union–non-union wage differential can emerge. The first is unions' ability to limit downward wage flexibility in times of hardship relative to their uncovered counterparts: this shows up as a countercyclical rise in the premium (Blanchflower and Bryson 2004a). A second is the possibility that union-induced wage hikes limit worker entry to the union sector, or result in job cuts that increase the supply of labour to the non-union sector, thereby lowering wages relative to those paid in the covered sector. A third union wage effect, which may compress the union wage differential, is the 'threat' effect, whereby non-union employers raise their wages to avoid the threat of unionization (Rosen 1969; Freeman and Medoff 1981; Farber 2003). Unions may also have more indirect effects on wages. For instance, their 'voice' face lengthens job tenure, which itself is often correlated with higher wages, and alters the incentives employers and workers face when investing in their human capital.[1]

II. The Union Wage Premium in the Public Sector

The empirical literature on the union wage premium is one of the largest in labour economics and dates back to the seminal work of H. Gregg Lewis (1963). Economists' interest lies in the fact that wages set through collective bargaining may differ from wages that employers set unilaterally. The literature indicates that, in general, union bargaining results in wages above the market rate and in a wage distribution that is more compressed than the distribution in the non-union sector (Freeman and Medoff 1984; Blanchflower and Bryson 2003). Traditionally the literature has focused on the private sector, where unions' ability to appropriate rents can have direct implications for employers' competitiveness, their demand for labour, and price setting. Less attention has been paid to the size of the union wage premium in the public sector. This is because it is much more heavily unionized, and it is often assumed that most workers will receive collectively bargained rates of pay, even if they are not themselves union members, through practices that extend collectively bargained rates either as a matter of public policy, or because the practice is regarded as commensurate with the public sector's role as a 'good employer' and a 'fair employer'. Furthermore, although union pay setting in the public sector has implications for public–private-sector wage differentials, the absence of a financial maximand in the public sector has meant that it has a less direct bearing on employment growth and workplace survival in the public sector.

However, these circumstances have begun to change in the last two decades in a way that makes estimation of the public-sector union wage premium of greater interest and concern. First, the Workplace Industrial Relations Surveys show that public-sector union density dropped in the second half of the 1980s and continued to fall into the 1990s (Table 1). The *Labour Force Survey* shows union density in the public sector has continued to fall since then, albeit at a slow rate (Table 2).

Second, pay determination in the public sector began to change. As Table 1 shows, the percentage of workplaces recognizing unions for pay determination fell from 99% in 1984 to 87% in 1990, a figure that has remained stable since. By 2004, 17% of public-sector workplaces with 10 or more employees had no workers covered by collective bargaining, and a further 27% used other methods of pay setting for at least some of their employees (Kersley *et al.* 2006: 183–4). In one-third (32%) of public-sector workplaces some employees' pay was set through Pay Review Bodies, and in 28% some employees' pay was set unilaterally by management (Kersley *et al.* 2006).

The third change has been the increase in competition among providers of public services. Some public services are now offered by private-sector organizations, while many non-core activities supporting public services have been contracted out. Table 3 shows that subcontracting has risen substantially in both the public and private sectors since 1990. In the public sector this is apparent in six of the seven activities for which we have data. These changes have been occasioned by increased concerns regarding the size of the public-sector pay bill and the cost efficiency of private versus public providers of services. Managers in the public sector are thus focused on cost efficiencies, and there is

TABLE 1
UNIONIZATION IN THE PUBLIC AND PRIVATE SECTORS, 1980–2004

	1980	1984	1990	1998	2004
% workplaces with recognized union					
Public	94	99	87	87	87
Private	50	48	38	24	22
% workplaces with members but no union recognition					
Public	5	1	12	11	10
Private	11	10	11	12	15
Union density, workplace-weighted					
Public	85	80	73	61	64
Private	41	32	25	13	13
Union density, employee-weighted					
Public	84	81	72	58	58
Private	57	45	36	25	20

Note: All workplaces with 25 + employees.
Source: Workplace Industrial Relations Surveys, 1980–2004; authors' calculations.

TABLE 2
LABOUR FORCE SURVEY UNION DENSITY AND RAW TRADE UNION MEMBERSHIP WAGE PREMIUM, UK EMPLOYEES, 1995–2006

	Union membership density (%)		Raw membership wage premium (%)	
	Private	Public	Private	Public
1995	21.6	61.5	13.2	26.9
1996	20.9	61.1	13.0	26.7
1997	20.2	61.3	12.1	23.3
1998	19.5	61.0	9.7	22.4
1999	19.3	59.9	8.9	27.4
2000	18.8	60.2	7.5	27.5
2001	18.6	59.3	6.3	29.9
2002	18.2	59.7	6.1	28.9
2003	18.2	59.1	8.3	26.5
2004	17.2	58.8	6.4	23.4
2005	17.2	58.6	8.1	22.3
2006	16.6	58.8	9.1	22.5

Note: Sample consists of all employees.
Source: Grainger and Crowther (2007).

generally more careful scrutiny of pay-setting in the public sector than there was a few decades ago.

Together, these changes in the public sector mean that it is both more meaningful to estimate the union membership premium in the public sector than it used to be and that the findings are perhaps of more policy relevance than they once were.

TABLE 3
PERCENTAGE OF WORKPLACES SUBCONTRACTING SERVICES, 1990–2004

	Private sector			Public sector		
	1990	1998	2004	1990	1998	2004
Cleaning	55	71	72	56	61	66
Security	33	46	45	13	29	31
Catering	20	22	19	36	52	50
Building maintenance	66	71	72	54	67	73
Printing/photocopying	29	25	18	10	14	15
Payroll	13	17	25	5	25	41
Transport documents/goods	47	52	43	21	29	23

Note: Workplaces with 25 + employees; authors' calculations.
Source: Workplace Employment Relations Surveys, 1990–2004.

III. IDENTIFYING THE UNION WAGE PREMIUM

The above discussion highlights the potential causal effect that unions may have on wages in both the covered and uncovered sectors. However, there are serious difficulties in isolating the causal impact of unions on wages because of the difficulties in identifying the counterfactual—that is, what wages would look like in the absence of unions. The presence of unions in the economy can change the level and distribution of wages generally. In theory, these general equilibrium effects may both raise and reduce the level of aggregate wages in the economy (Farber 2001). Since it is not possible to observe wages in the absence of unions the effect is very difficult to estimate. Instead, estimates of union wage effects assume a partial equilibrium framework.[2] The union–non-union wage differential (the wage gap), defined as

$$(1) \qquad \Delta = \frac{W_u - W_n}{W_n},$$

is estimable because we observe the wages of members (W_u) and non-members (W_n). The union wage gap in equation (1) can be usefully approximated by the difference in log wages, implying that

$$(2) \qquad \Delta \approx \ln(W_u) - \ln(W_n).$$

The union wage gap may reflect the direct effect of unions on the wages of unionized workers, and the offsetting effects on non-union workers.

The fact that unionization is not randomly assigned means that it is very difficult to isolate the true causal impact of unions on wages. Biased estimates are likely to occur because factors unobserved by the analyst that affect wages may also affect worker and employer selection into the covered sector. Thus, union status is endogenous with respect to wages. Selection into union status is likely to be a function of both worker and employer choices (Abowd and Farber 1982). The wage standardization policy of trades unions is well known to workers and will be most appealing to those workers with low underlying earnings potential, since they have most to gain through unionization. However, not all workers who desire union employment can find union jobs (Bryson and Freeman 2006). This affords employers the opportunity to pick workers from the queue, and since, unlike the analyst, they are able to observe the quality of workers in the queue they will choose the best in the queue. As Farber (2001, p. 19) notes, the two selection processes appear to have offsetting effects on the estimated wage gap, with worker selection implying negative bias and employer selection implying positive bias. However, the effect of double selection on ordinary least squares estimates of the union wage gap is uncertain *a priori*, since it depends on the relative size of the two biases.

This is not the only selection issue that may affect estimates of union wage effects. A second is a union's choice of employer for organizing, a choice that is likely to be influenced by the cost of organizing, the benefits of organizing and, in particular, the availability of surplus profits, as well as by working conditions conducive to workers' desire for union solutions (such as low or unfair pay). Third, employers may have some choice as to whether they are in the covered or uncovered sector, or the type of collective agreement they adhere to.

For these reasons analysts have experimented with alternative methods in identifying the effect of unions on wages. Ever since Lewis's pioneering research (1963, 1986), in which he argued that ordinary least squares (OLS) estimates were the least biased estimator of union wage effects, most analysts have contented themselves with estimates of a union membership wage premium based on OLS. However, OLS returns an unbiased union impact only where all factors influencing both unionization and wages are accounted for. This 'selection on observables' assumption, known as the conditional independence assumption in the treatment literature, requires a very rich set of covariates.

However, even with rich cross-sectional data, there are likely to be factors determining both unionization and wages that are not observable to the researcher. These might include worker motivation, which may lead workers to become union members—if, for instance, they wish to have a voice in workplace organization or job design—as well as affecting their wages (e.g. through the effort they devote to their job). This has led researchers to explore methods of tackling selection on unobservables. With cross-sectional data, this entails the simultaneous estimation of union status and earnings to account for the simultaneity. The approach relies on assumptions regarding functional form and the use of instrumental variables that affect the probability of union status but do not have a direct bearing on wages. These instruments are hard to find, and it is generally difficult to design them into surveys. Furthermore, they often lead to unstable estimates which are frequently much larger than those obtained through other methods (Lewis 1986).[3]

Our data do not contain suitable instruments. We therefore adopt an OLS approach to estimation in this paper. We do not seek to control for the potential endogeneity of union membership: rather, we adopt the standard approach to estimation of the union–non-union wage gap using individual-level data and estimating by OLS. That is, at the individual worker i level, at time t,

$$\ln W_{it} = aX_{it} + \delta_{it} U_{it} + \varepsilon_{it}, \tag{3}$$

where subscript it indexes individuals over time, X_{it} is a vector of worker, job and workplace characteristics, U_{it} is a dummy variable indicating union membership and ε_{it} is a random component. (The t subscript is omitted from the Workplace Employment Relations Survey analyses, since we use a single cross-section.) The parameter δ represents the average proportional difference in wages between union and non-union workers adjusted for worker and workplace characteristics.

IV. Data and Estimation

We use the Labour Force Surveys from 1993 to 2006 to estimate the union wage premium over time and the Workplace Employment Relations Survey 2004 (WERS) to estimate the premium with linked employer–employee data. The Labour Force Survey (LFS) is a quarterly sample survey of households living at private addresses. Its purpose is to provide information on the UK labour market that can then be used to develop, manage, evaluate and report on labour market policies. The survey seeks information on respondents' personal circumstances and their labour market status during a specific reference period, normally a period of one week or four weeks (depending on the topic) immediately prior to the interview. Each quarter's LFS sample of 53,000 UK households is made up of five waves, each of approximately 11,000 private households. Each wave is interviewed in five successive quarters, such that in any one quarter one wave will be receiving its first interview, one wave its second and so on, with one wave receiving its fifth and final interview. Thus, there is an 80% overlap in the samples for each successive quarter. Average gross hourly pay data for employees and those on a government scheme have been available from winter 1992/93 onwards.[4] Data on union status is available only in autumn quarters, so our data-set consists solely of autumn quarters from 1993 to 2006. The LFS has the advantage that it is publicly available and has a large number of control variables.[5]

WERS is a cross-sectional survey of workplace managers responsible for employment relations linked to employees working in those workplaces.[6] Union wage premia have been estimated previously using various sweeps of the WERS data (Blanchflower 1984, 1986; Booth and Bryan 2004). The 2004 survey is confined to workplaces with at least five employees. Both manager and employee surveys have high response rates. The dependent variable is hourly wages derived from banded wage data and continuous hours' data. We take the mid-point in the wage bands and divide by hours.[7]

In addition to the high quality of the data, WERS has a number of advantages in estimating the union wage premium. First, it contains rich workplace covariates that are usually lacking in the individual or household-level data used in many studies. Recent empirical research indicates that the paucity of employer controls tends to result in an upward bias in union wage effects. This is because unionized workplaces tend to be better paying than non-union workplaces for reasons that are not directly attributable to union membership (Blanchflower and Bryson 2004b).[8] The linkage of employees to employer data is thus likely to reduce the bias in estimating union wage effects.

Second, WERS contains multiple observations on employees per workplace, thereby allowing us to estimate workplace fixed-effects models which compare the wages of union members and non-members within the same workplace. Third, WERS identifies which occupations within the workplace have their pay set by collective bargaining. Since the employee data contain occupational classifications, we are able to identify which employees are covered by collective bargaining. Assuming that workplace managers responsible for employment relations know more about methods of pay determination than their employees, these coverage data are likely to be less prone to measurement error than surveys relying on employee responses. Where covered and uncovered workers are present in the same workplace, multiple employee observations linked to workplace data make it possible to estimate the within-workplace effect of coverage. However, when interpreting such estimates one needs to bear in mind the 'spillover' effect of coverage in pay-setting for non-covered workers. Forth and Millward (2002) present evidence of this spillover effect. Booth and Bryan (2004) use multiple worker observations per workplace to estimate the union membership wage premium among covered workers.

A fourth reason why WERS is a valuable data source for investigating the union membership wage premium is that it helps us overcome an important measurement issue that might result in an upward bias of the union wage premium in the public sector. The measurement problem arises from the potential for private-sector employees working in public-sector establishments to misclassify themselves as public-sector employees. Hicks *et al.* (2005) point to a discrepancy of more than 1.2 million workers between estimates of public-sector employment derived from the Labour Force Survey, based on households, and the numbers derived from administrative and survey data from public sector organizations.[9] In 2004 Q1 there were 5,746,000 public-sector employees in the United Kingdom using data from public-sector organizations, compared with 6,907,000 using data from the LFS (Hicks *et al.* 2005, Table 8). They argue that the numbers in the LFS are considerably higher because

> respondents can unknowingly report themselves in the public sector when really they are in the private sector according to National Account definitions. . . . Employees working for agencies and for contractors can also classify themselves as working in the public sector in the LFS when in reality, because their employer is a private sector organization, they should be allocated to the private sector according to the National Accounts definitions. (Hicks *et al.* 2005, pp. 9–10)

WERS does not face this problem, since the public-sector status of employees is ascribed to them by their employer. Employers identify which sector they are in and then provide a listing of their employees from which employee respondents are drawn. Thus, in the case of WERS, an employee working on a public-sector site for a private-sector contractor would be correctly classified as a private-sector employee.

There is some dispute in the literature as to what is deemed the appropriate set of conditioning variables in identifying the union membership wage premium. In the first place, some argue that the membership premium is not relevant: what matters is coverage by collective bargaining, since this is what determines pay levels. In practice, many datasets do not contain a coverage variable, or if they do there are concerns about measurement error, leading analysts to prefer the membership indicator. In our analysis we focus on the membership premium but test its sensitivity to the inclusion and exclusion of a coverage dummy in our workplace data. In these workplace data we also consider the membership premium among covered employees.

Second, there is some discussion as to whether to incorporate occupational codes, since one may argue that they simply represent segments of the wage distribution. In our case it is important to test the sensitivity of our results to the inclusion of occupational dummies to establish whether any difference in the size of the membership premium between the public and private sectors is simply due to different occupational distributions of members and non-members in the two sectors.

Third, analysts vary in their choice of job-related conditioning variables: some argue that job tenure and other variables are themselves a function of unionization and, as such, remove some of the union-based differences in wages across members and non-members. Again, in our case we test the sensitivity of results to more and less parsimonious models to establish the degree to which any membership premium can be explained by these job-related differences across members and non-members.

V. Results

Over the last twenty years or so there have been declines in union membership across most OECD countries (Blanchflower 2007). The decline in unionization rate that has occurred in the United Kingdom does not appear to have been accompanied by a large fall in compensation of union workers relative to non-union workers. Table 2 shows some decline in the raw union membership wage premium in both the public and private sectors. However, research tracking the regression-adjusted premium over a longer time frame suggests that some of this recent decline may be countercyclical rather than secular (Blanchflower and Bryson 2003, 2004a).

Table 4 reports estimates of the union membership wage premium for the United Kingdom using the LFS over the period 1993–2006. Part (a) shows the raw premium in the public and private sectors for all workers and for men and women separately. The raw premium is 26 log points or 29.7% (calculated as the antilog of 0.2602–1). The raw premium is lower in the second half of the period in the private sector than in the first half of the period for both men and women. The public-sector raw union membership wage premium is higher than in the private sector throughout the period, the gap being much more pronounced in the case of women. In contrast to the private sector, the raw premium is roughly constant over time in the public sector and even rises a little for men.

Part (b) of the table shows regression-adjusted estimates of the union membership wage premium, controlling for variables identified in the notes to the table. These controls account for a sizeable part of the raw membership premium in both the private and public sectors, but a significant premium remains for all 27 estimates presented. In all but one case, the public-sector union wage gap is higher than in the private sector.

Table 5(a) shows the distribution of union members and non-members in the public and private sectors across dimensions of their jobs. In the public sector, Professional Occupations and Associate Professional/Technical Workers make up 57% of union members, compared with only 33% of non-members; in the private sector the figures are 21% and 18%, respectively. Thus, union members are heavily concentrated in professional and technical occupations in the public sector, unlike their non-member counterparts. This is not the case in the private sector, where union members are more evenly distributed across occupations, although they are 'over-represented' among Process Plant and Machine Operatives. One might therefore expect union members to be relatively highly paid among public-sector workers simply because of their occupational status.

Part (b) of Table 5 shows union density within occupations across the public and private sectors. Public-sector workers are more heavily unionized than private-sector workers in all occupations. However, the rank order of union density by occupation differs in the two sectors. In the public sector the occupation with the highest union density is Professionals, followed by Associate Professionals/Technical Workers. So, not only do they account for a high percentage of all unionized workers in the public sector, these occupations also have very high union density and, as such, might be expected to have substantial bargaining power. In the private sector the occupation with the highest union density, and thus the potential for high bargaining power, is Process, Plant and Machine Operatives.

To account for these differences in the occupational profiles of workers in the public and private sectors, part (c) of Table 4 adds over three hundred and fifty detailed 3-digit occupational controls to the LFS specifications discussed above. Comparing results in part (c) with those in part (b), the addition of detailed occupational controls results in a larger union membership wage premium in the private sector. The separate analyses by sex reveal that this result is driven by what happens to the male union membership wage premium. Among men, the addition of occupational controls doubles the premium in the private sector and lowers it slightly among women. In the public sector, on the other hand, the introduction of occupational controls lowers the membership premium for

TABLE 4
UNION WAGE DIFFERENTIALS, 1993–2006, LFS

	All	Private	Public
(*a*) *No controls*			
1993–2006	0.2602 (n = 196,069)	0.1864 (n = 139,374)	0.2677 (n = 56,249)
1993–1999	0.2870 (n = 89,790)	0.2190 (n = 64,176)	0.2665 (n = 25,365)
2000–2006	0.2368 (n = 106,279)	0.1559 (n = 75,198)	0.2687 (n = 30,884)
Men			
1993–2006	0.1661 (n = 95,388)	0.1116 (n = 75,394)	0.1366 (n = 19,768)
1993–1999	0.1869 (n = 44,077)	0.1397 (n = 34,585)	0.1275 (n = 9363)
2000–2006	0.1471 (n = 51,311)	0.0853 (n = 40,809)	0.1445 (n = 10,405)
Women			
1993–2006	0.3402 (n = 100,681)	0.1955 (n = 63,980)	0.3140 (n = 36,481)
1993–1999	0.3656 (n = 45,713)	0.2189 (n = 29,591)	0.3166 (n = 16,002)
2000–2006	0.3190 (n = 54,968)	0.1739 (n = 34,389)	0.3120 (n = 20,479)
(*b*) *With controls*			
1993–2006	0.0939 (n = 185,778)	0.0662 (n = 132,315)	0.1280 (n = 53,463)
1993–1999	0.1227 (n = 88,793)	0.0882 (n = 63,417)	0.1289 (n = 25,156)
2000–2006	0.0868 (n = 105,491)	0.0521 (n = 74,709)	0.1271 (n = 30,782)
Men			
1993–2006	0.0443 (n = 90,077)	0.0407 (n = 71,391)	0.0568 (n = 18,686)
1993–1999	0.0629 (n = 43,450)	0.0602 (n = 34,075)	0.0515 (n = 9265)
2000–2006	0.0397 (n = 50,805)	0.0312 (n = 40,448)	0.0611 (n = 10,357)
Women			
1993–2006	0.1374 (n = 95,701)	0.1005 (n = 60,924)	0.1538 (n = 34,777)
1993–1999	0.1752 (n = 45,343)	0.1206 (n = 29,342)	0.1583 (n = 15,891)
2000–2006	0.1290 (n = 54,686)	0.0842 (n = 34,261)	0.1498 (n = 20,425)
(*c*) *Plus narrow occupation controls*			
1993–2000	0.1163 (n = 105,067)	0.1138 (n = 75,022)	0.0814 (n = 29,798)
2001–2006	0.0789 (n = 89,183)	0.0823 (n = 63,059)	0.0681 (n = 26,124)
Men			
1993–2000	0.0913 (n = 51,409)	0.1009 (n = 40,387)	0.0453 (n = 10,900)
2001–2006	0.0714 (n = 42,821)	0.0814 (n = 34,106)	0.0444 (n = 8715)
Women			
1993–2000	0.1338 (n = 53,658)	0.1154 (n = 34,635)	0.0966 (n = 18,898)
2001–2006	0.0826 (n = 46,362)	0.0759 (n = 28,953)	0.0788 (n = 17,409)

Notes: Equations in parts (*b*) and (*c*) all include the following controls: year dummies, age and its square, gender, four race dummies, 47 highest qualifications dummies, 61 industry dummies and 21 region of work dummies. Dependent variable is the log of hourly earnings—gross weekly pay divided by basic usual hours. In parts (*a*)–(*c*) year dummies are included. In part (*c*) there are 370 and 353 occupation dummies respectively for years 1993–2000 and 2001–2006. Samples include all employees. All estimates are significantly different from zero at the 0.01 level.
Source: Labor Force Surveys, Autumn quarters.

both men and women, the effect being most pronounced for women. As a result, the union membership wage premium for men is lower in the public sector than it is in the private sector, whereas among women the membership premium is roughly comparable across the two sectors.

We explore the link between the union membership wage premium in the two sectors and employees' occupational status by estimating the membership premium within each

TABLE 5(a)
DISTRIBUTION OF UNION AND NON-UNION WORKERS IN THE PUBLIC AND PRIVATE SECTORS (%)

	Public sector		Private sector	
	Non-union	Union	Non-union	Union
Managers and senior officials	8%	6%	18%	11%
Professional occupations	16%	30%	8%	9%
Associate professional/technical	17%	27%	10%	12%
Administrative and secretarial	24%	14%	13%	9%
Skilled trade occupations	2%	2%	10%	15%
Personal service occupations	17%	12%	6%	4%
Sales and customer service occupations	1%	1%	12%	8%
Process plant and machine operatives	1%	1%	8%	19%
Elementary occupations	13%	7%	14%	12%
	Public sector		Private sector	
	Non-union	Union	Non-union	Union
% all workers	11	17	58	14
% female	67	61	42	32
Male	33	39	53	66
Non-white	6	5	5	5
London & South East	34	27	35	25
Job tenure	7	11	5	11
Ever do overtime	38	56	47	62
Fulltime	58	78	73	86
Permanent job	85	95	94	97
Seasonal	*	*	1	*
Fixed contract	10	4	2	2
Temp agency	1	*	1	*
Casual work	2	*	2	*
Other temporary	1	1	1	*
Basic work hours	29	34	34	36
$\leqslant$10 employees at workplace	10	6	27	9
$\geqslant$ 50 employees at workplace	63	68	43	71
Degree or higher degree	23	33	15	11
Tenure 1 year or less	31	12	38	15
Tenure $\geqslant$ 10 years	26	50	19	46

Note: Sample based on all employees. *<1%.

major occupational classification. Table 6 does this using pooled LFS data for the period 2001–06. The membership premium is positive and statistically significant in both sectors for all occupational groups, with one exception, i.e. managers and senior officials in the private sector, where the premium is weakly negative. In general, it seems that the membership wage premium for higher occupations is larger in the public sector than in the private sector, whereas the premium for lower occupations is higher in the private sector. The membership premium is greater in the public sector than the private sector for Managers and Senior Officials, Professional Occupations and Associate Professionals and Technical Staff; it is roughly the same among Administrative and Secretarial Staff. However, the union membership premium among both Personal Service Occupations and Elementary Occupations is much larger in the private sector than in the public sector.

TABLE 5(b)
UNION DENSITY BY OCCUPATION IN PUBLIC AND PRIVATE SECTORS (%)

	Public	Private
Managers and senior officials	52.9	12.2
Professional occupations	72.2	21.0
Associate professional and technical	69.2	21.2
Administrative and secretarial	44.7	13.0
Skilled trades occupations	62.9	25.0
Personal service occupations	49.9	12.7
Sales and customer service occupation	48.2	13.0
Process, plant and machine operatives	64.2	34.8
Elementary occupations	42.2	16.0

Notes: $n = 196{,}366$; all employees.
Source: Labour Force Surveys Autumn quarters (September, October and November), 2001–06 plus August 2003 and December 2006.

TABLE 6
UNION MEMBERSHIP WAGE PREMIUM BY OCCUPATION, 2001–2006

	Wage gap	*t*-statistic	*n*
Public sector			
Managers and senior officials	0.0420	2.39	1885
Professional occupations	0.0728	6.29	6293
Associate professional and technical	0.0690	6.79	5915
Administrative and secretarial	0.0723	7.45	4794
Personal service occupations	0.1113	9.17	3632
Elementary occupations	0.0753	4.89	2531
Private sector			
Managers and senior officials	–0.0247	1.65	10,802
Professional occupations	0.0319	1.92	4917
Associate professional and technical	0.0456	3.36	6578
Administrative and secretarial	0.0701	5.11	8110
Skilled trades	0.1303	11.36	6878
Personal service occupations	0.1965	9.16	3544
Sales and customer service	0.0381	3.14	7259
Process, plant & machine operatives	0.1568	16.68	6317
Elementary occupations	0.1321	11.85	8654

Notes: Equations include the following controls: year dummies, age and its square, gender, four race dummies, 47 highest qualifications dummies, 61 industry dummies and 21 region of work dummies. Dependent variable is the log of hourly earnings – gross weekly pay divided by basic usual hours.
Source: Labor Force Surveys, Autumn quarters.

The other point emerging from part (a) in Table 5 is that the gap in job quality between union members and non-members, as measured by being on fixed-term contracts and in part-time jobs, is larger in the public sector than the private sector. That is to say, union non-members are more likely to be on poorer contracts than union members in the public sector than they are in the private sector, confirming our conjecture mentioned in the Introduction. This may reduce non-members' wages relative to members' wages and, since the discrepancies are greater in the public than in the private sector, they may increase the wages of members relative to non-members to a greater degree in the public than in the private sector. Of course, these job features are endogenous with respect to union membership. Nevertheless, we decided to test the sensitivity of our results to the inclusion of these controls, together with job tenure, which is lower for non-members than members in both sectors. We re-estimated the three equations in row 1 of part (b) of Table 4 for 1993–2006 for the whole economy and then for the private and public sectors, simply adding controls for years of job tenure with the current employer and the nature of the employment contract distinguishing between seasonal, casual, fixed-term and 'other' non-permanent contracts as well as between full-time and part-time employees. The results were as follows:

	Table 4 estimates	New estimates
All	0.0939 ($n = 185{,}778$)	0.0485 ($n = 185{,}623$)
Private sector	0.0662 ($n = 132{,}315$)	0.0226 ($n = 132{,}204$)
Public sector	0.1280 ($n = 53{,}463$)	0.0824 ($n = 53{,}419$)

Including these controls lowers the estimated union membership premium in both sectors, the effect being particularly pronounced in the private sector. Thus, contrary to expectations, rather than reducing the gap in the union membership wage premium between the two sectors, introducing controls for job quality further accentuates the gap.

As noted earlier, the way employees are sampled in WERS means we are unlikely to encounter the misclassification problem that might plague the LFS. Furthermore, the coverage variable is also less likely to be prone to error, since it is collected from workplace managers of employment relations rather than relying on employees' self-reporting.

Table 7 presents OLS estimates of the union membership wage premium using WERS. The raw premium is substantially larger in the public sector than in the private sector (columns (1) and (2)).[10] The introduction of controls for demographic characteristics, qualifications, 3-digit occupational classification, region and workplace characteristics reduces the premium in both sectors very substantially and goes a long way towards closing the gap in the membership premium between the two sectors (columns (3) and (4)). However, the membership premium in the public sector remains a little higher than the premium in the private sector—in contrast to the findings in part (c) of Table 4. The introduction of potentially endogenous job controls reduces the membership premium still further in both sectors, but it remains statistically significant in both sectors and is a little larger in the public sector (columns (5) and (6)).

WERS 2004 confirms our expectation that coverage and membership are more highly correlated in the public sector than the private sector. In the public sector 63% of union members belong to an occupation that their workplace manager says has pay set through collective bargaining, compared with 51% in the private sector. However, the addition of a coverage dummy has little impact on the pattern of these results.[11]

We tackle the issue of potential unobserved workplace heterogeneity by running estimates of the union membership wage premium with workplace fixed effects using a robust estimator (Table 8). Workplace heterogeneity plays an important role in reducing the size of the union membership wage premium in the private sector—cutting it by around a half (compare models (1) in Tables 7 and 8). However, the premium actually rises in the public sector when comparing the wages of members versus non-members in the same workplace (compare models (2) in Tables 7 and 8). The introduction of

TABLE 7
OLS ESTIMATES OF THE UNION MEMBERSHIP WAGE PREMIUM

	(1)	(2)	(3)	(4)	(5)	(6)
	Private	Public	Private	Public	Private	Public
Union member	0.120	0.202	0.051	0.064	0.035	0.052
	(10.72)	(15.73)	(5.83)	(6.28)	(3.91)	(5.01)
Constant	2.113	2.115	2.351	2.516	2.414	2.654
	(383.28)	(206.75)	(31.76)	(13.95)	(32.16)	(14.67)
Observations	12,818	5934	12,818	5934	12,818	5934
R-squared	0.01	0.04	0.53	0.52	0.55	0.53

Notes:
t-statistics in parentheses.
Models (1) and (2): raw premium without controls.
Models (3) and (4): includes following controls: male, age (9 dummies), academic qualifications (8 dummies), vocational qualifications (3 dummies), health problem, white British, household composition (4 dummies); 3-digit SOC (81 dummies); industry (12 dummies); region (10 dummies); establishment size (7 dummies); single workplace organization; workplace age (4 dummies); foreign owned (private sector only).
Models (5) and (6): (3)/(4) plus the following controls: permanent contract; full-time contract; workplace tenure (5 dummies); training days (4 dummies); gender segregation of the job at the workplace (6 dummies).
Sample consists of all employees.
Source: WERS 2004.

TABLE 8
ESTIMATES OF THE UNION MEMBERSHIP WAGE PREMIUM WITH WORKPLACE FIXED EFFECTS

	(1)	(2)	(3)	(4)	(5)	(6)
	Private	Public	Private	Public	Private	Public
Union member	0.054	0.237	0.037	0.060	0.024	0.049
	(4.54)	(18.10)	(3.61)	(5.64)	(2.35)	(4.59)
Constant	2.129	2.093	2.389	2.519	2.523	2.701
	(451.44)	(205.62)	(40.06)	(24.04)	(38.40)	(24.48)
Observations	12818	5934	12818	5934	12818	5934
R-squared	0.00	0.06	0.32	0.45	0.34	0.47

Notes
t-statistics in parentheses.
Private sector models contain fixed effects for 1237 workplaces. Public sector models contain fixed effects for 484 workplaces.
Models (1) and (2): raw premium without controls.
Models (3) and (4) contain following controls: male, age (9 dummies), academic qualifications (8 dummies), vocational qualifications (3 dummies), health problem, white British, household composition (4 dummies); 3-digit SOC (81 dummies).
Models (5) and (6): (3)/(4) plus the following controls: permanent contract; full-time contract; workplace tenure (5 dummies); training days (4 dummies); gender segregation of the job at the workplace (6 dummies).
Sample consists of all employees.

demographic controls and 3-digit occupational classification reduces the size of the estimated premium in both sectors, particularly in the public sector. Nevertheless, the membership premium remains larger in the public sector than in the private sector (models (3) and (4)). The inclusion of potentially endogenous job controls in models (5) and (6) reduces the premium still further but it remains statistically significant in both

TABLE 9

UNION MEMBERSHIP WAGE PREMIUM AMONG COVERED WORKERS

	Private	Public
OLS		
1. No controls	0.083 (4.70)	0.151 (9.51)
2. (1) + demographics, 3-digit SOC, workplace controls	0.053 (3.71)	0.062 (5.07)
3. As 2 + job controls	0.041 (2.86)	0.046 (3.64)
Workplace fixed effects		
4. No controls	0.030 (1.75)	0.157 (9.74)
5. (1) + demographics, 3-digit SOC	0.015 (.97)	0.054 (4.08)
6. As 5 + job controls	0.008 (.50)	0.037 (2.77)

Notes:
t-statistics in parentheses
Model coefficients with t-statistics in parentheses.
Demographic controls: male, age (9 dummies), academic qualifications (8 dummies), vocational qualifications (3 dummies), health problem, white British, household composition (4 dummies).
Workplace controls: industry (12 dummies); region (10 dummies); establishment size (7 dummies); single workplace organization; workplace age (4 dummies); foreign owned (private sector only).
Job controls: permanent contract; full-time contract; workplace tenure (5 dummies); training days (4 dummies); gender segregation of the job at the workplace (6 dummies).
Fixed effects models contain 303 workplaces in the private sector and 384 in the public sector.
Sample consists of employees (full + part-time).
Source: WERS.

sectors and is roughly twice as large in the public sector as it is in the private sector. These within-workplace estimations of the union membership wage premium which also include detailed occupation controls are strong evidence of a union effect on wage setting in both sectors, and of a larger effect in the public sector.

We further attempted to determine whether there was a membership premium among those workers who, according to their workplace manager, had their pay set via collective bargaining. We estimated both OLS and fixed-effects models using the same controls as in the estimates presented above. The raw membership premium among covered workers is roughly twice as large in the public sector as it is in the private sector, but is only marginally larger once regression-adjusted (Table 9, rows 1 and 2). Within-workplace estimates of the membership premium among covered workers were estimated using fixed-effects estimation (rows 4–6). With workplace fixed effects there appears to be no statistically significant union membership wage premium among covered workers in the private sector, confirming Booth and Bryan's (2004) previous analyses using WERS 1998. However, in contrast, members in the public sector continue to be paid a premium over their non-member counterparts in the same workplaces, even when controlling for detailed occupational classification. That premium is in the order of around 5% (4% when potentially endogenous job controls are included in row 6).[12]

VI. CONCLUSIONS

Research on trade unions' wage effects is confined largely to the private sector. However, the growth of the non-union public sector has made it more meaningful to consider union effects in that sector, while increasing competition in the provision of public services has meant that union effects are attracting greater policy interest. The paper is motivated by two initial observations. First, the raw membership premium continues to be large in both sectors, although it has declined a little. Second, the raw membership wage premium is much larger in the public sector than in the private sector. We investigated potential reasons for this sectoral difference in the membership wage premium using the Labour Force Survey and the Workplace Employment Relations Survey. Our starting point was the observation that there are substantial differences in the distribution of union members in the public and private sectors according to job characteristics and job quality. Aspects of this worker and workplace heterogeneity are not usually captured in standard wage regression models, and so may account for the differences in the premium across the two sectors.

Using the Labour Force Survey we show just how heterogeneous the union wage premium is across workers. In particular, it differs systematically in the two sectors across occupations: in higher occupations the membership premium is larger in the public sector than in the private sector, whereas in lower occupations it is larger in the private sector than it is in the public sector. If all occupations are pooled in models that contain detailed occupational controls, there is no differential across sectors in the membership premium among women, and the male membership premium is actually lower in the public sector than in the private sector. On the other hand, contrary to expectations, controlling for potentially endogenous job quality variables actually accentuates the gap between the membership premium in the public and private sectors.

Using the Workplace Employment Relations Survey confirms that the gap between the membership wage premium in the public and private sectors closes when controlling for a more extensive set of variables. However, in contrast to the LFS, the membership premium remains a little larger in the public sector than in the private sector, whatever one controls for. This is even the case within workplaces. The public-sector union membership wage premium remains roughly twice the size of the private-sector membership premium having accounted for workplace fixed effects and workers' occupations, job characteristics, qualifications and demographics. Furthermore, the membership wage premium *among* workers covered by collective bargaining is apparent only in the public sector.

We conclude that the membership wage premium in the public sector is not merely an artefact of private-sector workers misclassifying themselves as working in the public sector. Rather, the union membership wage premium in the public sector is very robust to controlling for a wide range of detailed workplace, job and demographic controls. The same is true for the private-sector membership wage premium. Thus, despite the weakening of unions in the United Kingdom over the last quarter-century, unions continue to play an important role in wage formation. For the first time, we have shown that this is also the case in the public sector, even within workplaces and controlling for detailed occupational classification. Given the importance that government attaches to pay equity in general, and to the role of the public sector in achieving equitable pay outcomes, these findings are worthy of further investigation.

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NOTES

1. Unions are also known to have large effects on fringe benefits and non-wage labour costs, as well as on methods of payment.

2. Where union wage-setting affects a large percentage of the working population and union effects are sizeable, general equilibrium effects are likely to be substantial. For an example of such a study see Sanner (2003).
3. Reasons why IV impacts differ so much have recently been discussed in Heckman *et al*. (2006). They show that IV estimates vary where returns to treatment are heterogeneous and individuals select into treatment with partial knowledge of idiosyncratic returns.
4. The quarterly LFS launched in 1992 in Great Britain and in 1994 in Northern Ireland operated on a seasonal quarter basis: March–May (Spring), June–August (Summer), September–November (Autumn) and December–February (Winter). The reasons for this were as follows. (i) Many activities associated with the labour market occur seasonally and follow the pattern of the school year. This was more the case when the LFS first started, at which point more young people left school at Easter than in the summer. (ii) Easter can cause difficulty as it varies in timing between March and April—so ensuring that Easter is always covered by the same quarterly survey period avoids this problem. In May 2006 the LFS moved to calendar quarters which means the microdata are no longer available on a seasonal basis (spring–winter). The main reason why the Office for National Statistics ONS is moving to calendar quarters for the LFS is that it is an EU requirement.
5. This is in contrast with the Annual Survey of Hours and Earnings (ASHE), formerly known as the New Earnings Survey (NES), which is a sample 1% of workers and is based on employer records. ASHE does not contain crucial details, for example of workers' education level, race or union membership. Also, since 2000 the microdata have not been publicly available. For a discussion of the differences in the earnings data in the LFS and ASHE, see Ormerod and Ritchie (2007).
6. For full details of the survey, see Kersley *et al*. (2006).
7. The top wage band is top-coded using a value of 1.5 times the lower band. We tested the sensitivity of our results to the use of interval regression techniques. They make no substantive difference to the results. We also tested the sensitivity of our results to the use of survey weights. The unweighted estimates produce slightly larger estimates of the membership premium but also make no substantial difference to our overall results.
8. Other studies (e.g. Farber 2001) have shown that biases in OLS cross-sectional estimates arising from unobserved heterogeneity may either upwardly or downwardly bias the 'true' impact.
9. Downloadable at www.statistics.gov.uk
10. The public-sector raw union membership premium in WERS is similar to that for the LFS in 2004, but the private-sector premium is a little bigger in WERS than it is in the LFS (see Table 2).
11. These estimates are available from the authors on request.
12. We also examined data on wage changes from the longitudinal element of the LFS. We pooled together data from 1997–2004 and 2006 for individuals aged 16–64; no data are available for 2005 because of the switch to calendar years. Approximately half of the individuals in the quarterly LFS files attrited for one reason or another. In total that gave us 36,113 observations, of which 3.6% of workers switched from union in year 1 to non-union in year 2 while 4.0% switched from non-union in year 1 to union in year 2. We were unable to find any statistically significant results using these data (results not reported). The reason for this finding is likely due in part to measurement error as well as to attrition bias. It is also likely to be driven by the fact that there are only small numbers of workers changing their union status: the decline in density has largely been driven by old (union) firms dying and new (non-union) firms being born. At the same time, there has been a rise in the proportion of individuals who have never been union members even in unionized firms (Bryson and Gomez 2005).

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