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The Effect of Unionisation on Wages in Great Britain: Estimates from the Labour Force Survey

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ABSTRACT

This paper provides benchmark estimates of the impact that trade unions have on the wage rates paid to workers in Great Britain using data from the Labour Force Survey. *This is done for a number of gender and occupational subgroups of the population using information on both union membership and union coverage.*

1. INTRODUCTION

UNION WAGE DIFFERENTIALS are a well-documented labour market phenomenon. Some of the earliest empirical work conducted on UK union/non-union earnings used aggregate industry data (see Metcalf, 1977 and Blanchflower, 1984 for summaries), with differentials estimated at between 70 per cent (by Minford, 1983, using a time-series estimation of a macro-model) and zero. More recent UK work has used micro-data to investigate the union issue, both at an establishment level (e.g. Stewart, 1987, 1991, 1995) and at an individual level (e.g. Shah, 1984; Blackaby *et al.*, 1991; Murphy *et al.*, 1992). Estimates obtained from these sources have typically been smaller than those derived from aggregate sources, but have shown a wide diversity across individuals and sectors. Added to this, some commentators have examined union membership wage differentials (e.g. Stewart, 1983; Green, 1988; Blanchflower, 1991), whilst others have used coverage (or an appropriate proxy) as their defining union condition (e.g. Blanchflower and Oswald, 1990, Stewart, 1991, 1995). However, general findings have been that wage differentials are greater for manual workers than for non-manuals, that union wage gaps are larger for women than they are for men, and that membership provides greater returns than coverage in individual level studies.

As highlighted by Andrews *et al.* (1998), methodological and data source differences have had a substantial impact upon the conclusions that can be drawn between such published studies. There is no single piece of work that provides a comprehensive analysis of the consistent effect of unionisation for male and female, manual and non-manual workers, looking at union membership and union coverage differentials. Invariably, this has been a result of data restrictions within the commonly used data sources in the UK and is closely linked to sample size considerations. Indeed, Andrews *et al.*'s own comprehensive study of union wage effects in the UK is hampered by their choice of data set (the *British Household Panel Survey*) and this means that they are unable to estimate separate relationships between unionisation and wages for female workers and for non-manual workers.[2]

By way of contrast, the *Labour Force Survey* does allow an analysis to be conducted over gender, occupation and union status subgroups. This important data set has been an under-utilised data source in the UK investigation of union wage effects.[3] Its chief advantage over other commonly employed data sources is its sheer size, which makes an examination of detailed population subgroups possible. Using a number of pooled cross-sections we will provide benchmark estimates for this data source of the hourly earnings premium unions are able to secure split by gender and a broad occupational grouping, with an emphasis placed upon isolating the influence of trade union membership and trade union coverage.[4] The framework within which this will be performed is set out in section 2 and the data employed is detailed in section 3. Section 4 presents the main results of this analysis before concluding comments are presented in section 5.

2. Methodology

The methodology employed in this study can be illustrated by considering a general framework measuring wage differentials between two identifiable groups of workers. The starting point is to define separate earnings functions for both sets of workers, who for ease of exposition we will refer to as being employed in sectors U and N:

$$\ln E_U = X_U \beta_U + \varepsilon_U \tag{1}$$

$$\ln E_N = X_N \beta_N + \varepsilon_N \tag{2}$$

E are hourly earnings, X is a vector of characteristics known to influence earnings, β is a conformable vector of estimated rewards to these characteristics, subscripts denote sector of employment, and the error terms ε_U and ε_N are assumed to be serially uncorrelated and normally distributed.[5]

The difference between equations (1) and (2) is the sectoral wage differential and following Blinder (1973) and Oaxaca (1973) this can be used to assess the importance of characteristic and coefficient (price) differences in accounting for mean wage differentials between the two sectors. Thus,

$$\overline{LnE_U} - \overline{LnE_N} = (\overline{X}_U - \overline{X}_N).w_\beta + w_x.(\beta_U - \beta_N) \tag{3}$$

where a bar indicates a mean value, $w_\beta = \lambda.\beta_U + (1-\lambda).\beta_N$, $w_X = (1-\lambda).\overline{X}_U + \lambda \overline{X}_N$ and ë $\in [0,1]$.

The first term on the right-hand side of equation (3) is the difference in average earnings arising from differences in average levels of characteristics between sectors. The second term is that part of the average earnings difference attributable to differences in the prices (or returns) to these characteristics. The choice of λ in equation (3) gives rise to the familiar index-number problem. When λ=0 we use the coefficients from sector N as base in the framework, and when $\lambda=1$ we use those of sector U. Thus, the decomposition framework will be different for each possible choice of λ.

Other possible choices of λ have also been suggested by Reimers (1983) and Cotton (1988), conforming to the idea that the appropriate value of λ in the decomposition framework should lie somewhere between those currently observed in both sectors. Whilst providing unique decompositions and thus avoiding the index-number problem, the assumptions made by both Reimers and Cotton are somewhat arbitrary and have no theoretical foundations. Alternatively, the approach of Neumark (1988) and Oaxaxa and Ransom (1994) advocates a value of λ that would be seen to arise if wages were determined competitively between sectors. Thus, the competitive wage structure ($\beta*$) is given by

$$\beta^* = \Omega\beta_U + (1-\Omega)\beta_N \tag{4}$$

where the weighting matrix (Ω) is defined such that $\Omega = (X_U' X_U + X_N' X_N)^{-1}(X_U' X_U)$. This provides the basis for a unique decomposition framework where the difference in mean hourly earnings may then be decomposed as

$$\overline{LnE_U} - \overline{LnE_N} = (\overline{X}_U - \overline{X}_N)\beta^* + [\overline{X}_U(\beta_U - \beta^*) + \overline{X}_N(\beta^* - \beta_N)] \tag{5}$$

which can be separated into a characteristics component (i.e. the first term on the right-hand side) and a price component (i.e. the term in square parentheses). From equation (5), the size of the sectoral hourly earnings markup (D) can be shown to be equal to

$$D = (\exp[\overline{X}_U(\beta_U - \beta^*) + \overline{X}_N(\beta^* - \beta_N)] - 1).100 \tag{6}$$

3. Data

The data used in this analysis come from the *Labour Force Survey* (LFS), a large-scale survey conducted by the Office for National Statistics (ONS). Switched from an annual to a quarterly basis in 1992, it aims to produce a sample of approximately 60,000 responding households in Great Britain every quarter. Over the course of the survey respondents are interviewed on five separate occasions, commencing in the quarter they enter the survey and then once more in each of the next subsequent four quarters. Following their fifth interview respondents are replaced by a new cohort. This rotating sample design means that within any one quarter approximately one-fifth of all respondents are being interviewed for the first time, one-fifth for the second time etc., all the way up to the fifth who are being interviewed for the final time. There is, therefore, an eighty per cent overlap of respondents from any one quarter to the next. To avoid any possible double-counting we ensure that individuals are only picked up once during their participation within the LFS. This is done by selecting only those respondents who have completed the fifth and final wave of their interviews, which is also the only occasion that earnings-related questions are asked in the sample used here.[6]

Information contained within the LFS allows the definition of union membership and union coverage and both of these measures are utilised in the analysis that follows. Specifically, coverage is proxied by union recognition at place of work, whereby a series of questions are asked to determine if trade unions are present at a respondent's workplace that they are entitled to join and which are recognised by the management of the establishment. Membership details are gained directly from a question asking whether a respondent is a trade union or staff association member.

The data used run from the Autumn of 1993 to the Winter of 1995, a total of ten quarters in all. Data availability dictated the starting point, as the questions used to establish union coverage were not included in the LFS until the third quarter of 1993. Meanwhile, the end point was chosen so as to provide a sufficiently large sample for the gender and occupational splits outlined earlier whilst being mindful of aggregation problems over time. By pooling the separate quarters and after selecting only non-agricultural, non-personal service employees for whom there was no missing information, we were left with somewhere near 28,000 (full-time) males and 27,000 females of working age who had completed their final interview and had hourly earnings data available.[7]

4. Results

Some preliminary descriptives of employee trade union affiliation are set out in Table 1. The figures in the final column show that nearly one half of all workers in the sample used here (44.1 per cent) have no connection with trade unions in the course of their jobs. Such a figure is greater than the proportion of all workers reporting themselves as being a union member and is in stark contrast to the experience of the 1970s. Indeed, at the end of 1995 there were only 8.1 million trade union members in the UK (DfEE, 1997), the lowest number since 1945 and 39 per cent down on the peak achieved in 1979.

Table 1: Union Representation (%) by Gender and Occupation: LFS 1993q3-1995q4

Trade Union Affiliation	*Males*		*Females*		*All*
	Non-Manual	*Manual*	*Non-Manual*	*Manual*	
Uncovered non-member	45.1	40.1	43.7	52.1	44.1
Uncovered member	3.1	3.7	3.5	1.8	3.3
Covered non-member	15.8	9.0	17.4	14.5	14.8
Covered member	36.0	47.2	35.4	31.6	37.8
Sample size	15,925	12,312	21,768	5,509	55,514

Disaggregating these trends along the lines of gender and manual/non-manual status highlights a wide divergence in trade union attachment across subgroups of the population. For males, 50.9 per cent of manual workers are trade union members as compared with only 39.1 per cent of non-manuals. A substantially larger proportion of non-manual workers, though, enjoy union coverage at work without actually being a member of a union themselves (15.8 per cent compared with only 9.0 per cent for manuals). There is also a larger part of the non-manual workforce (45.1 per cent) with neither membership nor coverage than there is for manual workers (40.1 per cent).

For females, union membership is proportionally higher for non-manual workers (38.9 per cent) than it is for manual workers (33.4 per cent), and a greater fraction of non-manuals are covered by a trade union at work without having joined than is found for manuals (17.4 per cent compares with 14.5 per cent). The largest proportion of workers in any of the four groups with no trade union affiliation whatsoever is found for manual females, where over half of all workers are neither a union member nor covered by a union at work. The comparable figure for non-manual females is more than eight percentage points lower at 43.7per cent.[8]

Earnings equations were fitted for each of the four gender/occupation groups outlined in Table 1, and a full list of the variables used and their definitions may be found in the Appendix.[9] F-tests were performed to indicate whether it was appropriate to combine the four categories of union status in a single earnings equations. Such tests clearly demonstrated that the pooling of union states was inappropriate and so a unionisation dimension was also introduced into the analysis.[10] However, the relatively small number of individuals reported as being a union member but not covered by a union at their place of work made a separate analysis of this category untenable. Moreover, union members in non-covered establishments are typically members of professional organisations, which have rather different characteristics from typical trade unions. It was therefore decided to exclude this group of workers from the analysis.[11]

While it is impractical to give a blow-by-blow account of the individual OLS estimates for each of the population subgroups by union status, they did nonetheless conform to a familiar pattern: hourly earnings increase with years of formal schooling, experience (though at a decreasing rate),[12] job tenure and generally with establishment size; married (and cohabitating) individuals enjoy a wage premium over other marital states, as do healthy individuals relative to those with reported health problems; a lack of formal educational qualifications significantly reduces employee hourly earnings; there are large regional variations in wage rates, with the highest rates being found in the South East of England and London; large hourly earnings differentials are evident across both industry and occupation dimensions; and finally, being of an ethnic origin other than 'white' substantially reduces hourly wage rates.[13]

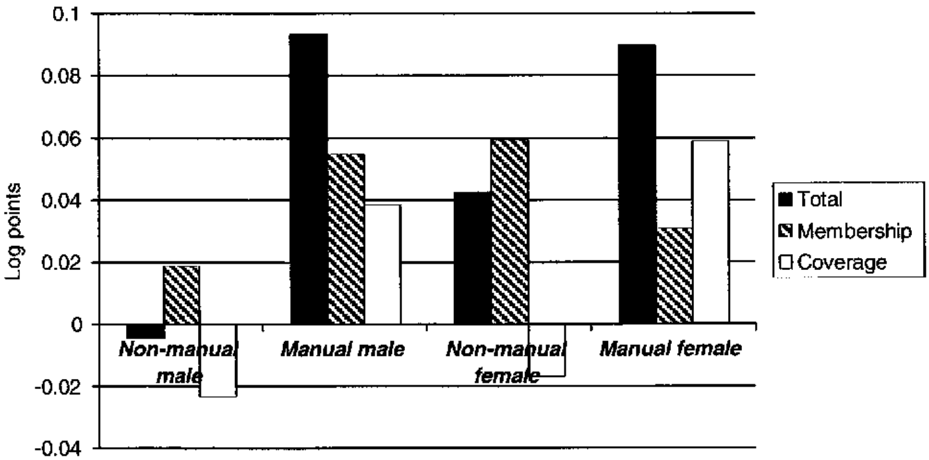
Comparisons of the above equations that had been estimated separately across union states were carried out and it was found that a sizeable number of coefficients were statistically different from one another between equations, in particular those relating to plant size and work experience. These results would tie in generally with those found by Blackaby *et al.* (1991) in the UK and with the general evidence cited in Hirsch and Addison (1986) for the US that earnings profiles are flatter in education and experience in the union sector.

By adopting pairwise comparisons of the previous wage equation estimates it is possible to separate out the union influence upon wages into an effect arising due to union membership and an effect arising due to union coverage. The total union effect is derived by comparing employees with no trade union affiliation whatsoever (i.e. non-members who are not covered by a union at work) against those workers who are not only covered by a union at their place of work but who are also union members. To isolate the effect of membership from this, union members with coverage at work are compared against non-members who are covered at work. The effect of coverage is gauged by comparing non-union members who are covered by a union with those non-members who are not covered.

The results of the decomposition analysis, as set out earlier in equation (5) and using the above pairwise comparisons, can be found in the Appendix. The figures show how much of the difference in average earnings between any two comparison groups can be attributed to either a 'characteristic' effect or a 'price' effect (see section 2) and the 'price' component from these decompositions is then used along the lines set out in equation (6) to calculate union wage premia. Whilst it is not our intention to concentrate upon the overall decomposition results themselves, it should be noted that the nature of the breakdown of earnings is entirely consistent with previous research of this nature. For example, the difference in mean hourly earnings between manual male workers who are covered trade union members and comparable uncovered non-unionists is 0.2741 log points. Of this difference, nearly two thirds (or 0.1806 log points) is attributable to the superior average workforce characteristics of covered unionists — the characteristic component. The remainder of the difference (0.0935 log points) is due to a more favourable labour market return on the average characteristics of covered unionists — the price component. More generally, though, the results would indicate that characteristic asymmetries are the major factor behind the observed union wage differences for male and female, manual and non-manual employees alike. The importance of a differing workforce composition, though, is consistently higher (in terms of its percentage contribution) for non-manual workers. For males, the chief contributors to this 'characteristic' effect are job tenure and size of workplace, whilst the importance of mean education levels and public sector employment are emphasised in the female analysis in addition to job tenure.[14]

A visual representation of the 'price' component from the decompositions is given in Figure 1 and the union premia implied by such figures are shown in Table 2.[15] The visual representation of the log point differences given in Figure 1 confirms three general trends known from the existing literature. The first is that returns to unionisation are greater for manual workers (both male and female alike) than they are for non-manuals. Secondly, the returns to membership are generally higher than are the returns to union coverage. This is true for all categories except for manual females, for whom coverage appears to impart greater pecuniary rewards. And thirdly, the benefits to unionisation appear to be greater for women than they are for men. Whilst such a pattern is clearly evident for non-manual workers, it is not so apparent between manuals. We should remember, though, that the occupational structure of female employment is highly skewed towards the non-manual sector. Indeed, over eighty per cent of the sample of women used here are in non-manual occupations.

Figure 1: Log Point Union Wage Effects by Gender and Occupation: LFS 1993q3-1995q4



A more detailed description of these results can be gained from Table 2. For non-manual males, the overall union markup (i.e. including the effects of membership and coverage) is marginally negative at -0.45 per cent (Table 2, column 1) but when the effect of union membership is taken out this disadvantage increases in magnitude to -2.3 per cent (column 3). The effect of membership alone, in comparison, is to increase wages by just under two per cent (column 2). Previous researchers have generally found small positive hourly earnings markups for non-manual male workers of the order of about 4-5 per cent (Green, 1988; and Symons and Walker, 1990), although Murphy *et al.'s* (1992) estimate is somewhat higher at 10 per cent. This latter figure is more in line with the annual markups found by Blanchflower and Oswald (1990) and Blanchflower (1991).[16]

Table 2: Percentage Mean Wage Markups by Gender and Occupation: LFS 1993q3-1995q4

		TOTAL *covered member vs. uncovered non-member*	*MEMBERSHIP* *covered member vs. covered non-member*	*COVERAGE* *covered non-member vs. uncovered non-member*
Males	Non-Manual	-0.45	1.90	-2.30
	Manual	9.80	5.66	3.93
Females	Non-Manual	4.35	6.12	-1.67
	Manual	9.41	3.14	6.07

For manual males, unions exert a substantial influence upon the hourly earnings of workers (the total markup shown in Table 2, column 1 is 9.8 per cent), with both membership (5.7 per cent) and coverage (3.9 per cent) found to be important factors (columns 2 and 3 respectively). Other studies have generally found a 10-13 per cent hourly wage differential in favour of union members (Shah, 1984; Symons and Walker 1990; Murphy *et al.*, 1992), although the figure of 18 per cent found by Yaron (1990) appears out of line with these.

Turning now to females, the results show that there are again undoubtedly gains to be had from unionisation. For non-manual workers, the overall markup of 4.4 per cent (Table 2, column 1) is driven by the effect that union membership has on wages. Whereas the membership markup is well over six per cent (column 2), the effect associated with coverage is actually negative (-1.7 per cent, column 3). These estimates fall between the widely differing membership markups of 2.7 and 14.6 per cent presented by Green (1988) and Main and Reilly (1992) respectively.[17]

For manual females, the overall union markup of 9.4 per cent (Table 3, column 1) is on a par with the 8.6 and 10 per cent estimates given by Green (1988) and Yaron (1990). However, both of these studies examined membership differentials, which we find to be somewhat smaller at only 3.1 per cent (column 2). Much larger at 6.1 per cent, though, is the coverage markup (column 3). Whilst undoubtedly substantial, the size of the overall estimate would lead us to conclude that whilst manual women have much to be gained from unionisation, the benefits arising appear to be no greater than they are for men.[18]

5. Concluding comments

Using data from the *Labour Force Survey*, substantial wage premia are found to exist for unionised workers, both male and female. These premia are greater for manual workers than they are for non-manual workers, and are larger for female non-manuals than they are for male non-manuals. Even in spite of the array of legislation implemented during the 1980s intent upon curtailing union influence over rent extraction,[19] it appears that they are still able to command considerable pecuniary benefits for their members and those whose interests they represent.

In line with other commentators, we would also suggest that women have more to gain from unionisation than do men. This is more apparent for non-manual workers; as such, the mean total markup is -0.5 per cent for males and 4.4 per cent for females. Meanwhile for manual workers, for whom estimated wage premia are that much greater, the pattern of female advantage is not so obvious. There is, however, a more clear-cut result when the focus is restricted to union coverage differentials.

The general direction of these results is consistent with previous estimates that have been derived using a variety of alternative data sources. The results presented here, though, calculated across gender, occupation and union status subgroups, provide unique estimates from an important UK data source and will establish benchmark estimates for future analysis with this data set.

Appendix

Variable definitions for analysis

Hourly earnings	The natural logarithm of gross hourly earnings from employment in January 1996 prices. Gross hourly earnings are defined as actual gross weekly earnings deflated by usual weekly hours worked excluding unpaid overtime.
Education	The number of years spent in full-time education. Dummy variable (10) indicating highest equivalent educational qualification.
Work experience	The number of years of potential work experience (males) or imputed labour market experience (females). Entered in linear and quadratic form.Dummy variable (3) indicating job tenure with current employer.
Marital status	Dummy variable (3) indicating marital status.
Health	Dummy variable indicating a health problem or disability that limits the kind of paid work that can be undertaken.
Ethnic origin	Dummy variable indicating a white ethnic background.
Location	Dummy variable (12) indicating region of residence.
Type of employment	Dummy variable indicating part-time employment (females only). Dummy variable (9) indicating industry of employment. Dummy variable (8) indicating occupation of employment. Dummy variable indicating public sector employment.
Workplace size	Dummy variable (5) indicating size of establishment.
Employment selectivity	Inverse Mills ratio from reduced form employment probit.

Table 1: Decomposition of Union Wage Differential by Gender and Occupation: LFS 1993q3-1995q4

Males					
Non-manual			*Manual*		
Covered Member vs Uncovered Non-Member					
Earnings difference	Characteristic component	Price component	Earnings difference	Characteristic component	Price component
0.1465	0.1510	-0.0045	0.2741	0.1806	0.0935
Covered Member vs Covered Non-Member					
Earnings difference	Characteristic component	Price component	Earnings difference	Characteristic component	Price component
0.0532	0.0344	0.0188	0.1473	0.0923	0.0550
Covered Non-Member vs Uncovered Non-Member					
Earnings difference	Characteristic component	Price component	Earnings difference	Characteristic component	Price component
0.0933	0.1166	-0.0233	0.1268	0.0883	0.0385

Females					
Non-manual			*Manual*		
Covered Member vs Uncovered Non-Member					
Earnings difference	Characteristic component	Price component	Earnings difference	Characteristic component	Price component
0.3745	0.3319	0.0426	0.2353	0.1454	0.0899
Covered Member vs Covered Non-Member					
Earnings difference	Characteristic component	Price component	Earnings difference	Characteristic component	Price component
0.2092	0.1499	0.0594	0.0511	0.0202	0.0309
Covered Non-Member vs Uncovered Non-Member					
Earnings difference	Characteristic component	Price component	Earnings difference	Characteristic component	Price component
0.1652	0.1820	-0.0168	0.1842	0.1253	0.0590

ENDNOTES

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2. Likewise, Hildreth (1999) also reports poorly defined coefficient estimates because of limited degrees of freedom when using the *British Household Panel Survey* to examine movements in the union wage differential in the 1990s.

3. Andrews, Bell and Upward (1998) do present estimates from the *Labour Force Survey* when they look at movements in coverage differentials for manual males, but we are unaware of any other published work that makes use of this data source.

4. Given the lack of convincing identifying instruments in the current data (particularly when we want to condition upon union membership and union coverage), the endogeneity of union status has been unaddressed. Evidence from Robinson (1989) and Lanot and Walker (1998) would lead us to believe, though, that the mean-based OLS estimates presented here will provide a lower bound upon the true influence that unions are able to exert over wages. This should be borne in mind when interpreting the results presented later.

5. As the data used in the analysis combines ten quarters of the *Labour Force Survey*, log hourly earnings are re-based to the same time period. This is achieved by subtracting the influence of time dummies from a pooled OLS regression from individual log hourly earnings. This is a similar procedure to that adopted by Blau and Kahn (1992).

6. The design of the LFS means that not all questions are asked in every quarter. Thus, it is sometimes necessary to 'match in' a question response from a wave that does not correspond with a respondent's final interview. For example, all questions relating to union status are asked only in the Autumn quarter of each year. In this instance, a number of observations were therefore dropped where workers had changed jobs between the collection of union status and earnings data. However, these exclusions, imposed to avoid potentially spurious correlations, had no bearing upon the results reported later.

7. A large part of the female workforce is employed on a part-time basis. Indeed, part-time employment accounts for over one third of non-manual workers and over one half of manual workers in the sample used here. It was decided to retain both groups of women in the following analysis to allow a fuller picture to be painted of the effects of unionisation upon the wages of women.

8. Whilst the above figures refer to all female workers, the same general trends are exhibited for both full-timers and part-timers. It is noticeable, though, that fewer part-time employees have no trade union affiliation in their jobs than do their full-time counterparts. Over one half of all part-time workers, be they employed in manual or non-manual occupations, are neither union members nor enjoy the benefits of union coverage at their place of work.

9. The choice of variables entering the model was largely motivated by theoretical considerations and the existing literature. In a comprehensive analysis of individual-level studies in the UK, Andrews *et al.* (1998) concluded that so long as some key controls were included, the inclusion or exclusion of the majority of them did not appear to be important. Of particular importance was the inclusion of controls for firm size, of which there are five categories in this analysis.

10. Tests for parameter equality gave F-values of 3.11, 4.85, 5.10 and 1.87 for non-manual males, manual males, non-manual females and manual females respectively. This is against a critical value of approximately 1.3 at the 1 per cent level.

11. In their work, Andrews *et al.* (1998) retained this group of workers and made no distinction between members and non-members who were not covered by a union, assuming that being a union member did not influence the level of pay of uncovered workers. We find that F-tests for such a restriction are failed.

12. Whilst the experience variable included for males is the conventional measure of potential labour mar-

ket experience, a value of labour market experience for females has been imputed along the lines suggested by Zabalza and Arrufat (1985) and Wright and Ermisch (1991). This involves predicting the probability of workforce participation for a woman of a given age in previous years, a procedure which has also been used by Blackaby *et al.* (1997). As a practical issue, the decompositions presented later are hardly affected by whether or not this experience adjustment is made.

13. Pre-selection into employment has been controlled for by including the inverse Mills ratio, estimated separately for males and female from reduced form probits, as a regressor variable in the wage equations. The variables that entered these participation probits were the standard set of family background variables, housing tenure, marital status etc. For females, a full sample of participants and non-participants was used in the probit equation, whilst for males only labour market participants were included. The decomposition results reported later, though, were neither contingent upon the inclusion of the constructed selectivity term nor its precise treatment. For example, netting out the influence of selectivity from the dependent variable and performing the decompositions around a framework of wage offer differentials gave premiums of the same order of magnitude as those presented later.

14. Due to space constraints these results are not shown in the Appendix Table, but they are available upon request from the authors.

15. The price difference component from the 'membership' and 'coverage' decompositions will sum exactly to the price component in the 'total' decomposition. However, because wage markups are calculated as exponential transformations of these figures, there is a slight discrepancy between the total wage markup and the sum of the membership and coverage markups.

16. All of these mentioned studies used membership as their measure of union status.

17. The figure by Green was derived from the coefficient on a gender dummy in a pooled equation rather than from a separately estimated equation for female workers. Meanwhile, Main and Reilly's estimate is for all full-time females.

18. By way of contrast, though, evidence from the coverage markups would indicate that the premiums experienced by women do outweigh those paid to male workers.

19. Machin and Stewart (1996) provide a discussion of this legislation, ranging from the *1980 Employment Act* to the *1993 Trade Union Reform and Employment Rights Act.*

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