UNION RELATIVE WAGE EFFECTS: A CROSS-SECTION ANALYSIS USING ESTABLISHMENT DATA

DAVID BLANCHFLOWER*

THE question of the extent to which unions are able to raise the wages of their members over that of comparable non-union workers has received a great deal of attention in the past few years. The number of studies producing estimates of union relative wage effects in the United States in particular, has increased so dramatically that one leading commentator has likened the search for the 'true' differential, rather colourfully, to the quest for the Holy Grail (see Farber, 1983). The stimulus for much of this work has been the availability of large micro-data sets in the United States, most of which are at the level of the individual, e.g. the Current Population Surveys, although there are a few at the level of the establishment, e.g. the Employer Expenditure for Employee Compensation Survey and Area Wage Surveys. Some of the individual micro-data sets have the added advantage that the same set of individuals is observed at more than one date, e.g. the National Longitudinal Survey.[1]

The availability of large micro-data sets in the United States has enabled researchers to obtain estimates of the effects of union on relative wages, disaggregated by sex, colour, marital status, occupation, industry, region, city size, age, experience and a host of other worker and workplace characteristics. Moreover, a number of these studies have tried to take account of the possibility that union status is endogenous by estimating a system of equations including (at least) a union equation, and a wage equation. It does appear, however, that the estimates obtained from such studies are sensitive to the econometric specification of the equation system, the method of fitting the system to the data and the data set used (see Lewis, 1982b).

This wealth of studies for the United States is not repeated for Great Britain, where only a relatively few studies have produced estimates of union relative wage effects; Table 1 summarizes the results. As can be seen from the sixth column there is substantial variation in the estimates, which range in value from −6 per cent to +74 per cent. In four of the studies reported in Table 1, the unionisation variable was defined on the basis of union membership (Pencavel, 1974; McNabb and Demery, 1978; Minford, 1981; and Stewart, 1983a) whereas in the remaining studies it was defined on the basis of the coverage of collective agreements Coverage data were used in an attempt to control for wage spillovers between union members and those non-union workers who were covered by a collective agreement. In consequence, estimated differentials using coverage data tend to be larger than those obtained from membership data. The one major exception to this is in Minford (1981 and 1983) who obtained an estimate of 74 per cent for the union/non-union wage differential in Great Britain for 1979, from aggregate time series data. There is little empirical support elsewhere for an estimate of union relative wage effects of the magnitude of 70 per cent.[2] All of the other studies reported in Table 1 used cross-section data and attempted to control for differences in worker and workplace characteristics; the data were not ideal, however, as pointed out by Metcalf (1977):

> Data which indicate whether a person is a union member/covered (e.g raw 1973 New Earnings Survey data on coverage of collective agreements) do not simultaneously give personal characteristics such as age, skill, experience. Data which are good on personal characteristics (e g. the annual General Household Survey) do not indicate whether an individual is a union member/covered (p 158).

*Queen Mary College, University of London

TABLE 1

Survey of Estimates of the Impact of Unions on Relative Wages in Great Britain

Study	*Date*	*Observations*	*Data Set*	*Dependent Variable*	*Differential*
Pencavel[a] (1974)	1964	29 production industries	Various	Average hourly earnings of manual workers	0%–14.5%
Mulvey[b] (1976)	1973	77 MLH manufacturing industries	New Earnings Survey 1973 Census of Population, 1971.	Average hourly earnings of manual workers	26%–35%
Mulvey &[c] Foster (1976)	1973	99 occupational groups	New Earnings Survey, 1973	Mean gross weekly earnings of full-time adult males (manual and non-manual)	22%–36%
Nickell[d] (1977)	1972	121 MLH manufacturing industries	Various	Full-time male and female manual hourly earnings	18%–26%
McNabb &[e] Demery (1978)	1964	See Pencavel (1974)	Various	Average hourly earnings of manual workers	11.6%–16%
Wabe &[f] Leech (1978)	1973	101 MLH manufacturing industries	Various	Average hourly earnings for adult male manual workers	−6%–21.5%
Layard,[g] Metcalf & Nickell (1978)	1973	4300 individuals	General Household Survey, 1973 New Earnings Survey, 1973	Hourly wage rates of male (manual & non-manual) workers	25%
Mulvey &[h] Abowd (1980)	1974	26 two-digit industries	New Earnings Survey, 1974.	Average hourly earnings of male manual workers	3.8%–8.1%

Treble[j] (1981)	1973	See Mulvey (1976)	New Earnings Survey, 1973.	Average hourly earnings of adult male manual workers	8.3%–40%
Minford[k] (1981)	1964–1979	Aggregate time series data	Various	Real average earnings of industries	74%
Geroski &[l] Stewart (1982)	1973	See Mulvey (1976)	New Earnings Survey, 1973	Average hourly earnings of adult male manual workers.	10.5%–16 2%
Stewart[m] (1983a)	1975	5352 full-time manuals in manufacturing	National Training Survey, 1973	Weekly earnings of full-time male manual workers	7.7%
Nickell &[n] Andrews (forthcoming)	1951–1979	See Layard, Metcalf & Nickell (1978)	General Household Survey, 1973 New Earnings Survey, 1973	Hourly wage rate of male manual workers	7%–32%

Notes:

a Industry-wide agreements give a zero differential whereas plant bargains confer a differential of 14 5%.
b Workers covered by national agreements enjoy no earnings differential whereas those covered by company, district & local agreements enjoy at least as large a differential as those covered by national plus supplementary agreements
c Variations in the extent of occupational coverage of collective agreements are closely associated with variations in occupational earnings.
d Separate estimates of the differential for males of 20%–26% and for females of 17%–21%
e Allowance made for the presence of 'non-zero excess demand'.
f Used non-linear estimation techniques.
g The estimated differential is higher the lower the skill group
h Studies that do not contain data on both the union and non-union wages are likely to produce estimates of the differential biased by as much as 50 per cent
j It is possible to generate wide variations in the estimated differential by changing the proxy list
k Estimates derived from the Liverpool University Rational Expectations model
l Results indicate two types of industry; a small group of 'haves' dominated by closed shops with local bargaining. The others, the 'have nots' are the characteristic industries in U.K. manufacturing. Which counterfactual to conditionalise on makes a major difference
m First estimates (for the United Kingdom) based on individual data. Considerable variation in the differential across industries.
n Times series estimates of the differential derived using coverage data for 1973 based on the 1958 Standard Industrial Classification with independent variables relating to the mid-1960's.

This has meant that the majority of the studies using cross-section data have been at a high level of aggregation and, in general, relate to male manual workers in manufacturing. The one notable exception is Stewart (1983a), who, although still restricting his analysis to male manual workers in manufacturing, had information from the 1975 National Training Survey on an individual worker's union status and personal characteristics. A serious deficiency of previous cross-section studies using aggregate data was that the vector of explanatory variables used to control for labour quality and workplace characteristics was very limited. The model used also ignored the possibility of variation of the differential across industries. Stewart was able to show that there is *considerable* variation in the differential *across* industries, which will result in aggregation bias in industry-level studies. Moreover, the National Training Survey permitted a fuller standardisation for differences in 'labour quality', which resulted in a much lower estimated differential (of 7.7 per cent) than obtained in the majority of previous studies, which seem to have suffered from serious omitted variable bias.[3]

Similar results for the United States are provided in a recent paper surveying estimates of union relative wage effects derived from aggregate data in the United States by H. Gregg Lewis (1983). He observed that 'there is a noticeable, though not universal tendency' for estimates of the union/non-union differential obtained from equations using aggregate data (macro equations) to 'exceed by a considerable margin' estimates obtained from equations based upon individual micro data, the latter including a variable indicating an individual's union status (micro equations). Lewis goes on to argue that estimates of the union relative wage effect derived from 'macro' equations, which do not include a union status variable U, but only an extent of unionism variable (however measured) 'tend on the average to estimate the sum of two quite distinct coefficients...the union/non-union wage gap $\tilde{M}$ and the partial derivative a_y of W with respect to the extent-of-unionism variable Y' (p. 23).[4] Lewis shows that these two coefficients may be identified separately, *only* if a union status variable U *and* an extent of unionism variable Y are included on the right-hand side of a 'micro' wage equation. The wage gap or union/non-union wage differential may then be derived directly from the coefficient on U.[5]

In this paper estimates are produced of the union/non-union wage differential derived from a 'micro' wage equation at the level of the establishment, with data drawn from the 1980 Workplace Industrial Relations Survey. This is only the second study from Great Britain to derive such 'micro' estimates, and the first using establishment data. Moreover, it is much more comprehensive in its coverage of the economy than any previous study. The majority of studies surveyed in Table 1 restricted themselves to (male) manual workers in manufacturing; although the data set used does not permit separate estimates to be obtained by sex (it refers to 'typical' workers) it is possible to provide separate estimates for both semi-skilled and skilled manual workers in non-manufacturing, manufacturing and the British economy as a whole. Estimates have also been derived for clerical workers and middle managers disaggregated in a similar fashion: additionally I have obtained estimates for all groups disaggregated by industry both in the public and private sectors: these results will be reported elsewhere.

The paper is structured as follows. Section I outlines the methodology to be used to obtain empirical estimates. Section II describes the data set and variables used, whilst Section III provides estimates of the union/non-union wage differential for semi-skilled and skilled manual workers. Section IV compares the results obtained in this study with those obtained elsewhere. Section V summarises the main conclusions.

I Methodology

In earlier studies the method employed to estimate union relative wage effects from aggregate data, involved regressing the logarithm of the average wage on the proportion of workers who are union members/covered by collective bargaining agreements. In an attempt to overcome the absence of wage data in both the union and non-union sectors, following H. Gregg Lewis (1963), the following identity was used.

$$\ln W_i \equiv Y_i \ln W_i^u + (1 - Y_i) \ln W_i^n \tag{1}$$

where W_i = the weighted (geometric) average wage for group i

W_i^n = the weighted (geometric) average wage for the non-union sector

W_i^u = the weighted (geometric) average wage for the union sector

Y_i = the proportion of workers in group i who are union members/covered by collective agreements.

Rearranging (1) yields

$$\ln W_i \equiv Y_i (\ln W_i^u - \ln W_i^n) + \ln W_i^n \tag{2}$$

One would like to obtain a measure of the differential between W^u, the union wage and that existing in the absence of unions W^a.[6] Unfortunately W^a is unobservable. Although W^u and W^n are, in principle, observable, studies using aggregate data were limited to data for the (arithmetic) mean wage, W_i.[7] W^n may well differ from W^a due to spillover or threat effects which may arise because the existence of unionism in a certain group presents a threat to non-union employers that their employees will become unionised and gain wage increases. The existence of such a threat may induce the non-union employer to increase the wage of his (non-union) employees; if this occurs then W^n will be greater than W^a.[8] The best one is able to do is to estimate a wage gap M_i, which is likely to be an understatement of the wage gain if positive threat effects exist: there is every reason to believe that in Great Britain such threat effects are positive and non-negligible because many non-union workers are paid union rates. In this case the estimates of the wage gap will underestimate the true extent of the wage gain.

Let us define a 'wage gap' M_i, a 'wage gain' A_i and a 'threat differential' R_i.

$$A_i \equiv \frac{W^u - W^a}{W^a} \tag{3}$$

$$M_i \equiv \frac{W^u - W^n}{W^n} \tag{4}$$

$$R_i \equiv \frac{W^n - W^a}{W^a} \tag{5}$$

Studies using aggregate data obtained their estimates of union relative wage effects by rearranging (4) and substituting into (2) yielding

$$\ln W_i \equiv [\ln (1 + M_i)] Y_i + \ln W_i^n \tag{6}$$

Rearranging (5) and substituting into (6) yields

$$\ln W_i \equiv [\ln (1 + M_i)]\, Y_i + \ln (1 + R_i) + \ln W^a \tag{7}$$

As W^a is not directly observable it is replaced by a vector X of variables that are expected to be its major determinants. The term $\ln (1 + R_i)$ is assumed constant across groups and is treated as part of the stochastic disturbance term and ignored; if $\ln (1 + R_i)$ is not constant then the disturbance term is likely to be heteroscedastic.[9]

If data are available on wages in both the union and non-union sectors, then it is possible to avoid some of the difficulties encountered using the method described above. Following Lewis (1983) equations of the following form are estimated, at the level of the establishment.

$$\ln W_i = a_n + a_{nx} X_i + a_{ny} Y_i + (a_u - a_n)U_i + e \tag{8}$$

where the a's are estimated coefficients, the u's refer to union establishments and the n's to non-union establishments. U is the union status of the establishment (1 if unions are recognised, O otherwise). As before, Y is the extent of unionism variable (which in this case is measured at the level of the industry in which the establishment operates),[10] e is an error term. M_i is the coefficient on the union status variable U; to convert the wage gap to percentages, calculate $100\,(e^M - 1)$. Equation (8) was fitted by Stewart (1983) for example, as two separate equations of the following form.

$$\ln W_i = a_i + a_{ix} X + a_{iy} Y + e_i;\ i = u \text{ if } U = 1;\ i = n \text{ if } U = O \tag{9}$$

The regression coefficient on the union status variable U is unbiased and consistent as long as there is no covariance among the groups between union status and either the effects of unionism on the average relative wage of union workers, of the effects on non-union workers. There are grounds for suspecting that $\mathrm{Cov}(U_i\, e_i) \neq 0$: if $\ln (1 + R_i)$ is correlated with either U_i or any other explanatory variables, then OLS will produce biased and inconsistent estimators. In particular, one might expect union employers who are forced to pay higher wages, to respond by employing workers of higher quality. If the available data allowed for perfect control for worker quality this would not be a problem, but in practice the data are not perfect, which is likely to result in estimates of the wage gap that are upward biased. This should be borne in mind when reading the results in Section III.

II Data and Description of Variable

The data set used in this paper is the 1980 Workplace Industrial Relations Survey (henceforth WIRS) sponsored by the Department of Employment, Policy Studies Institute and the Social Science Research Council based on a representative sample of 3309 working establishments[11] distributed throughout England, Scotland and Wales. The sample was drawn from the 1977 Census of Employment; to be included in the sample an establishment had to have twenty-five or more employees (both full- and part-time) at the time of the 1977 Census and at the time of the survey (April–September 1980). Hence new establishments and those whose size increased from below twenty-five to above twenty-five and those that declined below twenty-five in the years 1977 to 1980 were excluded from the sample. Omitting these establishments, an extremely high response rate of 76.3 per cent was achieved, giving 2040 observations in the data set. The sample includes establishments drawn from both manufacturing and non-manufacturing, private as well as public sectors, although it does exclude agricultural and farming establishments and coalmining (see Daniel and Millward, 1983, p. 5).

This survey is the first of a new series designed to provide information about a broad range of topics within the industrial relations field. The survey incorporated interviews both with management and worker representatives, with the senior manager at the establishment who deals with industrial relations or in non-industrial establishments, staff or employee relations. In certain cases (165 cases out of 2040) the responsibility for industrial relations rests in more than one centre and in such cases a secondary management interview was also undertaken. In addition, at each establishment interviews were undertaken with worker representatives (who were nominated by management respondents) up to a maximum of three. In total 2439 interviews with worker representatives are included in the data set (for more details see Blanchflower, 1983). This paper restricts itself to data drawn from the management questionnaire (plus that from the 'Basic Workforce Data Sheet', a self-completion questionnaire received by respondents prior to the interview) because all questions used were asked of the management respondents, whereas only relatively few were asked of the worker representatives. The sample design incorporated the use of variable sampling fractions according to the number of employees at a Census Unit, which is essentially the number of employees working at the same address who are paid from the same place. In general there was a sufficient degree of correspondence between Census Units and establishments to serve as a frame from which the sample could be drawn. In order to ensure that large establishments were numerically well represented in the sample, variable sampling fractions were used: This necessitated a weighting of the data used in the following sections, to adjust for the inequalities of selection that were introduced because of the differential sampling of the size bands. The matrix of weights used are presented in Appendix A.[12]

The equation to be estimated is given in (10).

$$\begin{aligned}\ln W_i = {} & a_0 + a_1\,\mathrm{PARTE} + a_2\,\mathrm{MANE} + a_3\,\mathrm{REC15} + a_4\,\mathrm{FEMMANE} + \\ & a_5\,\mathrm{SHARES} + a_6\,\mathrm{SKILLED} + a_7\,\mathrm{TOTAL} + a_8\,\mathrm{FOREIGN} + \\ & a_9\,\mathrm{BLACK} + a_{10}\,\mathrm{EXTENT} + a_{11}\,\mathrm{MANUFAC} + a_{12}\,\mathrm{SHIFT} + \\ & a_{13}\,\mathrm{NMANPRES} + a_{14}\,\mathrm{OUTW} + a_{15}\,\mathrm{PJEVAL} + a_{16}\,\mathrm{PBR1} + a_{17}\,\mathrm{PBR2} + \\ & a_{18}\,\mathrm{LTDCO} + B_i\mathrm{IND} + e\end{aligned} \tag{10}$$

The variables included in (10) are as follows:

(PARTE)	the percentage of the labour force that work part-time (< 30 hrs/week)
(MANE)	the percentage of the labour force that are manual workers
(REC15)	a (1, 0) dummy for the existence of manual union(s) recognised for negotiating the pay and conditions of employment of their members
(FEMMANE)	the percentage of manual workers that are female
(SKILLED)	the percentage of those workers in the labour force who have received formal training (apprenticeship or equivalent)
(TOTAL)	the number of employees at the establishment
(FOREIGN)	a (1, 0) dummy if the establishment is foreign owned

(BLACK)	the percentage of the workforce of Asian, African, West Indian or similar origin
(EXTENT)	the percentage of manual workers in unions in the industry into which the establishment is classified (by principal product)
(MANUFAC)	a (1, 0) dummy if the establishment (plant) is in manufacturing
(SHIFT)	a (1, 0) dummy if there is any shift work at the establishment
(NMANPRES)	the number of unions that have members among manual employees at the establishment
(OUTW)	a (1, 0) dummy if the establishment uses outworkers
(LTDCO)	a (1, 0) dummy if the establishment is a private/public company
(SHARES)	a (1, 0) dummy if the establishment has a share ownership scheme
(PJEVAL)	the percentage of workers covered by job evaluation schemes
(PBR1)	a (1, 0) dummy if the establishment pays the majority of its male unskilled/semi-skilled employees by results
(PBR2)	a (1, 0) dummy if the establishment pays the majority of its female unskilled/semi-skilled employees by results
(IND)	Industry dummies. (Coefficients not reported)

Data on pay levels in WIRS are derived solely from the management questionnaire; there are no questions on pay levels in the worker representative questionnaires. Management respondents were asked for the gross (weekly) pay over the previous month of the 'typical' employee in four separate skill groups—semi-skilled and skilled manual workers, clerical workers and middle management. The question itself was designed to ensure a high response rate which it seems to have achieved; a major disadvantage, however, is that the wage data are broad-banded across 9 classes and open-ended at both ends. (The inability to include information on fringe benefits is likely to understate the true impact of unions, although the extent of this understatement is uncertain.)

The use of weekly earnings rather than hourly earnings has been criticised by Metcalf (1977) because

> it is hourly earnings which indicate the opportunity set (income and leisure) facing the individual: weekly earnings comprise the choice of hours given the opportunity set and the opportunity set itself... From the firm's point of view it is hourly earnings which determine its production costs: weekly earnings reflect the firm's choice between men and hours and do not measure the underlying labour costs (p. 161)

Metcalf, Nickell and Richardson (1976) estimated that unionised male workers supplied approximately two hours less labour per week than non-union workers, in 1966. Although there is no way of confirming these findings, because of the unavailability of hours data in WIRS, if this pattern has continued into the 1980's then an under-estimate of the impact of unions will be obtained.

In 'micro' wage equations at the level of the individual, it has become standard practice to include on the RHS a number of human capital variables such as experience, age, qualifications, training, etc., with the intention of controlling for labour quality. A group of variables to represent characteristics of the job in which the individual is currently employed (such as whether the worker works full- or part-time, has responsibility for the work of others, is involved in hazardous work, etc.) is also included in an attempt to equalise net advantages. Clearly one would expect that a higher quality workforce would result in higher gross weekly pay of the 'typical' employee. WIRS contains a number of variables that can be used as proxies for the quality of the labour force at the workplace.

The percentage of the workforce at the establishment that work part-time (PARTE) is a factor that is likely to affect the gross pay of a typical employee. As there is no information on the number of hours worked at the plant, this variable is likely to proxy lower average hours per employee and is thus expected to have a negative coefficient. The ratio of female manual workers to all manual workers (FEMMANE) is also expected to have a negative coefficient. Female manual workers tend to have relatively low human capital accumulation in anticipation of lower labour force participation; they also tend to be concentrated in smaller, low-paying establishments, which tend to operate in labour intensive, low-skill industries. Thus, a high proportion of female manual workers may indicate a low quality labour force.

A similar argument can also be applied to the percentage of the labour force of Asian, African, West Indian or similar origin (BLACK) which is included to pick up the effects of discrimination: the higher the proportion of black workers, the lower the (log) wage is expected to be.

The skill structure of the workforce is expected to be an important determinant of the gross wage of the typical employee at the establishment. A variable is included for the proportion of the workforce at the establishment who have received formal training, at apprenticeship level or above (SKILLED). A higher proportion of skilled workers is expected to indicate a high quality labour force and hence the coefficient is expected to have a positive sign. The proportion of manual workers in the establishment's labour force (MANE), is included to estimate the net effect of the establishment's relative demand for manual and non-manual workers; in certain cases a higher proportion of manual workers might improve those workers' position by strengthening their bargaining power, in other cases smallness in numbers may work to manual workers' benefit, hence this variable has an uncertain sign.

There is some precedence in the literature for the inclusion in a wage equation of a size of establishment variable (see Sawyer, 1973; Mayhew, 1976, and Wabe and Leech, 1978). Hood and Rees (1974), for example, have argued that workers have a preference for work in smaller factories and any premium attributable to plant size, is to some degree, compensation for the disutility involved in working in large establishments (i.e. a compensating differential). Other explanations are that unions are more likely to have organised larger establishments due to their lower costs of organisation and be more militant because of the more impersonal nature of relationships. It has also been argued that large plants are more likely to operate payment by results schemes (PBR), and shift work than smaller plants, which thus produces higher wages. Large plants paying higher wages, argues Lester (1967) are likely to attract and hold higher quality employees. Hence he argues, the frequently observed (significant) positive relationships between wages and plant size arises due to the omission of certain labour quality variables. However, it may also be claimed that the disadvantages arising from the more impersonal aspects of large establishments necessitate higher payments to attract and hold a given quantity of labour. The number of workers at the workplace is included as the size of establishment variable (TOTAL); the coefficient on this variable is expected to have a positive sign, although it is not certain whether it will be significant in the presence of 'labour quality' variables. These include dummy variables for the presence of: i) shift working at the establishment level (SHIFT) which is expected to command a wage premium because of the inconvenience involved in such work; ii) payments by results schemes which are also expected to command a wage premium, as workers employed under incentive schemes tend to be more productive than those employed under time rates (PBR1 for whether the majority of male manual workers are paid by results; PBR2 for whether the majority of female manual workers are paid by results); iii) outworkers, many of whom are poorly paid (OUTW); iv) a share ownership scheme amongst employees could be a substitute for, or complement to, high wages (SHARES); v) the proportion of workers covered by job evaluation schemes (PJEVAL). Job evaluation schemes offer a systematic basis for establishing relative pay levels; they tend to exist in large establishments and are concentrated in the private and the manufacturing sectors in particular (see Daniel and Millward, 1983, pp. 203–8). Such formalised bargaining procedures are likely to carry compensating payments, hence the coefficient is expected to be positive; vi) Foreign ownership of the establishment (FOREIGN) which is often held to result in higher wages; however, evidence from a recent study suggests that this differential may not be as great as is generally believed. Indeed for the U.K., U.S. affiliates appear to pay lower wages than indigenous firms (see Dunning and Morgan, 1980); vii) the status of the organisation to which the establishment belongs, be it a limited company (LTDCO), a cooperative (COOP), a trust (TRUST), a nationalised industry (NATIND), a partnership (PARTNER), a quango (QUANGO) or government (GOVT)[13]; viii) A dummy variable to indicate whether the establishment is located in the manufacturing sector, classified by major activity. Layard, Metcalf and Nickell (1978) found that the estimates of the union relative wage effect they obtained from a macro equation are considerably smaller for manual workers in manufacturing than for the economy as a whole. Evidence provided by Thomson, Mulvey and Farbman (1977) also confirms this view,[14] hence the coefficient of this variable is expected to have a negative sign; ix) a total of 39 industry dummies are included, as rather crude proxies for the differences in product markets across industries (19 each in the manufacturing and non-manufacturing equations).

The final group of variables to be considered might be classified as 'union impact' variables. The first is the number of unions that have members amongst manual employees at the establishment (NMANPRES). There are two different types of multi-unionism: i) where each one of the main occupations at the establishment is organised by a different union; ii) where there is more than one union competing for membership within a group of workers at the establishment; the latter is less common than the first type. A multiplicity of unions at the workplace, of whatever type, creates problems because it 'makes difficult the effective use of manpower, promotes difficulties over job demarcation, has an in-built tendency to generate leap-frogging pay claims and imposes substantial demands upon management time when consultation, discussion and negotiation has to be duplicated over a number of bodies and conflicting demands have to be reconciled' (Daniel and Millward, 1983, p. 46).

Despite the fact that there are fewer unions than at the time the Donovan Commission reported (in 1966 there were 574 unions with a membership of 10,110,000, whereas in August 1983 there were 418 unions with 12,123,552 members),[15] multi-unionism still exists in 44 per cent of establishments that recognised manual unions. It is not certain, however, whether multi-unionism at the establishment strengthens or weakens the bargaining power of unions. On the one hand a pluralism of unions may strengthen the power of management who are able to play one union off against another. On the other hand, however, a plurality of unions may reinforce their overall bargaining power.

H. Gregg Lewis (1983) has argued for the inclusion in a 'micro' wage equation not only of a union status variable but also of an 'extent of unionism' variable. This extent of unionism variable Y is included because: 'There is a strong presumption, I think, that in the general equilibrium of the economy in the presence of unionism, the relative wage of each worker depends not only on his union status, sex, colour, schooling, experience, and like variables, but also on the extent of unionism in the whole work force and the distribution of workers by union status among work-force sectors' (p. 3). WIRS permits the derivation of an extent of industry unionism variable directly from the union membership data; the number of manual workers who are union members is summed over establishments and expressed as a proportion of the

TABLE 2

Means and Standard Deviations of Variables used in Section III

	Semi-skilled			*Skilled*		
	(1) *Great Britain*	*(2)* *Manu-facturing*	*(3)* *Non-manu-facturing*	*(4)* *Great Britain*	*(5)* *Manu-facturing*	*(6)* *Non-manu-facturing*
LMANWAGE	4.1919 (0 4001)	4.3633 (0.2760)	4.1083 (0.4234)			
LSKWAGE				4.5295 (0.2826)	4.6327 (0.2015)	4.4660 (0.3058)
MANE	59 8191 (27.201)	71 4761 (16.2835)	54.2377 (29.5311)	64.7824 (24.0819)	71 7643 (16 0893)	60.4883 (27.0117)
FEMMANE	21.3448 (23 3786)	20.6369 (24.3384)	21.6837 (25.866)	19.798 (24.392)	19.6928 (23.2503)	19.9009 (25.0819)
REC15*	0.6009 (0.4899)	0 7086 (0.4548)	0 5493 (0.4978)	0.6305 (0.4828)	0.7083 (0.455)	0.5826 (0.4934)
LTDCO*	0 6523 (0.4764)	0.9581 (0.2006)	0.5059 (0.5002)	0.6893 (0.4629)	0.9568 (0 2036)	0.5247 (0.4997)
SKILLED	17 3096 (21.9273)	27 0656 (23.6545)	12.6384 (19.344)	23.7358 (24.5386)	29 6864 (24.6961)	20.0717 (23.7328)
TOTAL	232.048 (662.009)	406.315 (994.0697)	148.6096 (392.56)	260.604 (703.82)	410.3144 (1002.659)	168.4184 (399.8123)
PARTE	18 7886 (22.724)	8 7655 (12 398)	23.5877 (24.8869)	15 417 (19 734)	8.174 (10 556)	19.8765 (22.5551)
BLACK	3.6492 (10.5846)	4.5905 (10.7229)	3.1985 (10.4929)	3.8676 (10.6777)	4.5458 (10 4106)	3.4500 (10.8239)
EXTENT	59.6985 (22.6788)	77 6995 (13 1384)	51.0798 (21.1804)	62.212 (22.262)	77.6586 (13 3109)	52.7005 (21.3178)
SHIFT*	0.4434 (0.4969)	0.4948 (0.5005)	0.4188 (0.4936)	0.4698 (0.4993)	0.4904 (0.5004)	0.4571 (0.4985)
NMANPRES	0.3889 (5.5571)	1 0893 (9.625)	0.0536 (1.0107)	0 4501 (6.0409)	1.0952 (9.6507)	0.0529 (1.1263)
OUTW	0.0793 (0.271)	0.1692 (0.3753)	0.0369 (0.1887)	0 0783 (0.2688)	0.1561 (0.3633)	0 0305 (0.1720)
SHARES*	0.0235 (0.1514)	0.0223 (0.1479)	0.024 (0.1531)	0.025 (0.1563)	0.0236 (0.1518)	0.0259 (0.159)
FOREIGN*	0.0619 (0.241)	0 1032 (0.3045)	0.0421 (0.2008)	0.0678 (0.2514)	0.0101 (0.3017)	0.0473 (0.2123)
PBR1*	0.2507 (0.4336)	0.3745 (0.4845)	0.1915 (0.3936)	0.2586 (0.438)	0.377 (0.4851)	0.1857 (0.3891)
PBR2*	0.1534 (0 3605)	0.3745 (0.4845)	0.0908 (0.2874)	0.154 (0.361)	0.2747 (0.4468)	0.0796 (0.0796)
MANUFAC*	0.3238 (0.4681)			0.3811 (0.4858)		

* denotes a dummy variable

total manual workers in each industry. The coefficient of this variable is expected to take a positive sign, although whether it will be significant in the presence of the union status variable remains uncertain.[16]

The final variable to be defined is the 'union status' variable which is crucial for our estimation of the wage gap. There has been some discussion in the U.K. literature in the context of 'macro wage' equations about the correct specification of the 'unionisation' variable, Pencavel (1974) used union membership whereas most subsequent authors used coverage by collective agreements. WIRS includes both the number of manual workers who are members of unions and whether or not the union(s) at the workplace were recognised for collective bargaining. The difficulty in using the union membership data to determine the union status of the establishment involves the use of some inevitably ad hoc criteria such as whether more than 50 per cent of manual workers were union members. Even though a large percentage of workers could be members of unions this does not mean that such unions actually *do* bargain on these workers' behalf, at the establishment. If unions are recognised for bargaining at the establishment, even if total membership is relatively small, it is probably prudent to class these establishments as 'union'. Hence the presence (or not) of recognised trade unions at the establishment is used as the indicator of the union status of establishments. Obviously this variable is expected to have a positive influence on wages.

Table 2 shows the means and standard deviations of these variables included in our empirical estimation, at the level of the economy and disaggregated into manufacturing and non-manufacturing sectors. Columns 1–3 are for semi-skilled manual, columns 4–6 for skilled manuals. The mean wage of semi-skilled workers is £66.08 (antilogarithm of 4·1909) whereas for skilled workers it is £92.71 (antilogarithm of 4.5295); although the dispersion of wages is rather larger in the former case than the latter. The most interesting differences are not between skilled and semi-skilled workers, however, but between the manufacturing and non-manufacturing sectors. Establishments in the manufacturing sector appear to be considerably larger, employ a higher proportion of skilled workers as well as a higher proportion of manual workers and a lower proportion of part-time workers than establishments in the non-manufacturing sector. A considerably larger proportion of establishments recognise unions for purposes of collective bargaining (71 per cent and 55 per cent in the semi-skilled equations and 71 per cent and 59 per cent for skilled respectively). The extent of industry unionism is similarly much larger in manufacturing than in non-manufacturing (78 per cent and 51 per cent for the semi-skilled and 78 per cent and 53 per cent for the skilled respectively). These marked differences between the manufacturing and non-manufacturing sectors are discussed in the following section.

III Results

The results of estimating equation (10) using the technique of OLS, with data drawn from the 1980 Workplace Industrial Relations Survey for semi-skilled manuals and skilled manual workers, are presented in Table 3. For both skill groups, equations are presented at the level of the British economy (equations 1 and 5) as well as disaggregated into manufacturing (equations 2 and 5) and non-manufacturing sectors (3 and 7). The total sample sizes available are 1604 for semi-skilled workers, and 1355 for skilled workers.

As was noted earlier, a major weakness of the data set is the measurement of the dependent variables; the typical gross pay of both skilled and semi-skilled workers is grouped and open-ended. Somewhat arbitrarily, arithmetic midpoints are allocated to the internal groups; £30 per week is used to close the open-ended group with weekly earnings less than £50; and £170 is imposed on the open-ended group with weekly

TABLE 3

Regression Equations Explaining Gross Weekly Wages of 'typical' Semi-skilled Manual Employees

Semi-skilled

	(1) *Great Britain*	*(2)* *Manufacturing*	*(3)* *Non-Manufacturing*	*(4)* *Great Britain*
PARTE	−0.0052[c]	−0.0050[c]	−0.0050[c]	−0.00332[c]
	(10.36)	(5.32)	(8.76)	(8.58)
MANE	0.0035[c]	0 0023[c]	0.0036[c]	0.03116[c]
	(8.45)	(3.26)	(6 90)	(9.80)
FEMMANE	−0.0045[c]	−0.005[c]	−0.004[c]	−0 00420[c]
	(8.80)	(7.95)	(6.09)	(10.68)
REC15	0.0967[c]	0.0195	0.129[c]	0 07341[c]
	(5.59)	(0.87)	(5.53)	(5 50)
LTDCO	0.0600[b]	0.065	0.082[c]	0 05088[c]
	(2.18)	(1.25)	(2.60)	(2.39)
SKILLED	−0.0015[c]	−0.0010[b]	−0.0018[c]	−0 00133[c]
	(3.54)	(2.39)	(2.90)	(4.12)
TOTAL	0.00003[c]	0.00001	0.00001[c]	0.00003[c]
	(2 76)	(1.07)	(3.46)	(3.47)
FOREIGN	0.1010[c]	0.061[a]	0.142[c]	0 08533[c]
	(3.28)	(1.90)	(2 93)	(3 58)
BLACK	−0.0023[c]	−0 0013	0.0034[c]	−0 00167[c]
	(3.23)	(1.50)	(3.59)	(3.08)
EXTENT	0 0030[a]	0.0089[c]	0.0043	0.0027[b]
	(1.81)	(5.24)	(1.63)	(2.13)
SHIFT	0.033[b]	0.068[c]	0.0187	0 02510[b]
	(2.02)	(3.00)	(0 85)	(1.97)
NMANPRES	0.0022[a]	0 0025[b]	−0 0010	0 00215[b]
	(1.68)	(2 57)	(0.11)	(2.15)
OUTW	−0 034	−0 056[b]	−0.010	−0 0298
	(1.15)	(1.96)	(0.22)	(1.29)
SHARES	−0.068	−0.130[b]	−0.050	−0.0593
	(1.44)	(2 08)	(0 86)	(1.43)
MANUFAC	−0.0600[a]	n/a	n/a	−0.03892
	(1.90)			(1.56)
PBR1	0 0670[c]	0.0110	0.109[c]	0.05850[c]
	(2.87)	(0.55)	(3.19)	(4 19)
PBR2	0.013	0.0110	0.020	0 0120
	(0.50)	(0.55)	(0.50)	(0.58)
PJEVAL	−0.0001	0.00001	−0.00001	−0.000025
	(1 56)	(0 85)	(1 32)	(1 07)
Constant	3.903	3.637	3 821	3.954
$\bar{R}^2$	0.5203	0.4570	0.4920	0.5334
N	1604	519	1084	1604
Wage gap	10.15%[c]	1.97%	13 77%[c]	7 62%[c]
S D. of wage gap	0.01906	0.02286	0.02655	0.01437

a Significant at the 10% level on a two-tailed test
b Significant at the 5% level on a two-tailed test
c Significant at the 1% level on a two-tailed test

TABLE 4

Regression Equations Explaining Gross Weekly Wages of 'typical' Skilled Manual Employees

	Skilled			
	(5) Great Britain	*(6) Manufacturing*	*(7) Non-Manufacturing*	*(8) Great Britain*
PARTE	−0.00167[c]	−0.0004	−0.00185[c]	−0.00140[c]
	(3.28)	(0 37)	(2.89)	(2.89)
MANE	0.00180[c]	0.00102[a]	0.00244[c]	0.00169[c]
	(4.70)	(1.67)	(5.04)	(4 64)
FEMMANE	−0.004[c]	−0.0033[c]	−0.0048[c]	−0.00379[c]
	(8.68)	(6.27)	(6.96)	(8.62)
REC15	−0.0042	−0.0067	−0.0091	−0 0030
	(0.29)	(0.35)	(0.45)	(0.21)
LTDCO	0.0337	−0.0304	0.0609[b]	0 0367[a]
	(1 54)	(0.58)	(2.53)	(1 76)
SKILLED	−0.0039	0.00027	−0.0007	−0.0003
	(1.18)	(0.63)	(1.48)	(1.07)
TOTAL	0 00003[c]	0.00001	0.00009[c]	0.00003[c]
	(3 50)	(1.34)	(3.94)	(3.60)
FOREIGN	0 0367	0.05137[a]	0.0202	0.0414[a]
	(1.45)	(1.85)	(0.48)	(1.72)
BLACK	0 00076	−0.00025	0.00128	0.00066
	(1 27)	(0.10)	(1.48)	(1.17)
EXTENT	0 00227	−0 00062	−0.0010	0 00173
	(1 53)	(0 46)	(1.06)	(1.38)
SHIFT	0 0082	0.0763[c]	−0 0193	0.0127
	(0.57)	(3.83)	(0.99)	(0.93)
NMANPRES	0.00233[b]	0.00254[c]	−0.00258	0.00313[c]
	(2.28)	(3.06)	(0.35)	(3.27)
OUTW	0.0200	0 042[a]	−0 050	0 0192
	(0.78)	(1.68)	(1.00)	(0.79)
SHARES	0.0292	0.0300	0 0162	−0.0329
	(0 75)	(0.57)	(0.29)	(0.89)
MANUFAC	−0.02465	n/a	n/a	−0.0232
	(0 93)			(0.87)
PBR1	0.0211	0.00243	0.0309	0 0236
	(1.10)	(0.10)	(1 10)	(1.30)
PBR2	0.05426[b]	0.0259	0 0934[b]	0.0455[b]
	(2 36)	(1.32)	(2.44)	(2.08)
PJEVAL	−0.00003[c]	−0.000013	−0.000025[c]	−0.00002[c]
	(5.79)	(0.16)	(5.3)	(3 81)
Constant	4.343	4.630	4.403	4.366
$\bar{R}^2$	0.4076	0.2620	0.4039	0.4087
N	1355	516	839	1355
Wage gap	−0.42%	−0.67%	−0 91%	−0.3%
S D of wage gap	0.01442	0.01902	0.02004	0.0142

a Significant at the 10% level on a two-tailed test
b Significant at the 5% level on a two-tailed test
c Significant at the 1% level on a two-tailed test

earnings or more than £160. To test the sensitivity of the results to these assumptions, the equations were re-estimated using data drawn from the 1980 New Earnings Survey.[17] The mean of the weekly earnings of a comparable sample of full-time manual employees whose earnings fell in that range are allocated to each establishment. The results of this experiment, for the economy as a whole, are reported in equation (4) for semi-skilled manuals and equation (8) for skilled manuals.

The distribution of workers within the wage groups used is reported in Table 5 with the first two columns derived from WIRS data and the third from the NES data on full-time employees. The one striking difference between the data from the two surveys is the proportion of workers earning over £140 per week (0.99 per cent for the semi-skilled, 4.98 per cent for the skilled and 13.3 per cent for full-time employees from the N.E.S.). It is encouraging, however, to find that there is considerable agreement between the findings, regardless of the procedure used to allocate wage levels to establishments,[18] although the N.E.S. data, in all cases, produces slightly lower estimates of the wage gap. The results for both skilled and semi-skilled manual workers by sector are discussed in the remainder of this section, concentrating on the results obtained using the WIRS data.

TABLE 5

Wage Distributions by Skill Group

Typical pay	*(1) % of semi-skilled*	*(2) % skilled*	*(3) % of full-time manual employees*
Less than £50	14.91	1.53	6.36
£50 but under £60	14 08	5.27	6.70
£60 but under £70	19.70	7.76	9.14
£70 but under £80	17.46	10.05	10.43
£80 but under £90	14.20	16.70	11.70
£90 but under £100	9 35	17.33	11.27
£100 but under £110	5 09	17.39	10.54
£110 but under £120	2.66	11.37	8 69
£120 but under £140	1.60	7 62	11.87
£140 but under £160	0.83	3.46	6.47
£160 and over	0.12	1.52	6.83
	100	100	100

Source: Columns 1 and 2: 1980 Workplace Industrial Relations Survey
Column 3: 1980 New Earnings Survey

The British Economy

Equation (1) explains 52 per cent of the variation of the wages of semi-skilled manual workers between establishments, which for a cross-section study using micro-economic data, is relatively high. The signs of the coefficients are very much as expected, with 14 out of 18 of the coefficients significant at the 10 per cent level or better on a two-tailed test and with 9 significant at the 1 per cent level. The coefficients take highly plausible values broadly consistent with prior hypotheses and are extremely stable as the list of regressors is varied. Moreover these variables performed very similarly in both equations (1) and (4) as the assumptions were changed regarding the measurement of the dependent variable.

In all specifications tried, the percentage of the workforce who worked part-time was highly significant and on its own it accounted for a remarkable 35.3 per cent of the variation of the dependent variable. Although there are considerable differences in the size of the coefficient in equations (1) and (4), in both cases the coefficients are highly stable. (It is not absolutely clear why this difference should occur, but a possible explanation is that the gross pay of the 'typical' employee has actually been interpreted as the typical gross pay of employees, which is much more readily influenced by the presence of part-time workers.) Although not accounting for anything like as large a proportion of the variation of the dependent variable as PARTE, the variables FEMMANE, MANE and SKILLED all have coefficients that were highly significant and very stable. The variable FEMMANE performed according to prior expectations: the higher the proportion of manual workers who were female the lower tends to be the gross wage of the typical semi-skilled manual. At the same time, the higher the proportion of manual workers in the workforce the higher is the gross wage of the typical semi-skilled worker. The variable SKILLED was included to indicate the level of human capital at the establishment; the negative sign on the coefficient is therefore, inconsistent with our priors.

The size of plant, the presence of shift working and payments by results amongst the majority of males and foreign ownership of the establishment all appear to raise the gross pay of the 'typical' employee, as does working for a limited company. A higher proportion of black workers has the opposite effect, however. The manufacturing dummy is significant at the 10 per cent level on a two-tailed test in equation 1, although it is not significant in (4). In none of the specifications tried did the coefficients on PJEVAL, SHARES, PBR2 or OUTW ever achieve significance.

Finally, turning to the 'union impact' variables, starting first with the NMANPRES variable, which indicates that the number of unions that have members among manual employees has a small significant effect upon wages, independent of plant size. This variable works in a very similar way to EXTENT which also has a small significant effect upon wages; both variables are significant at the 10 per cent level in (1) and at the 5 per cent level in (4). These results provide limited confirmation of Lewis' (1983) finding that the higher the extent of unionism in an industry the higher the wage, (although it must be said, that the size of this effect is relatively small).

The union status dummy works well, with a 't'-statistic of over 5 and a coefficient of 0.0967, which gives an estimated wage gap of 10.15 per cent in equation (1) with a standard deviation of 0.01906 and a wage gap of 7.62 per cent and a standard deviation of 0.01437 in equation (4). (This standard deviation is calculated using a procedure suggested by Treble.)[19]

Equation (5) presents the results obtained from estimating equation (10) for skilled manual workers making the same assumptions about the distribution of the dependent variable as were made for the semi-skilled. (Equation (8) uses NES data rather than midpoints.) These equations do not work as well as those for the semi-skilled with only seven out of eighteen variables significantly different from zero at the 10 per cent level or better (five of these seven were significant at 1 per cent). MANE, PARTE, FEMMANE and TOTAL all work in a very similar way for skilled manuals as they did for the semi-skilled. The coefficients of the LTDCO, SKILLED, FOREIGN, BLACK, SHIFT, EXTENT, PBR1 and MANUFAC variables change from significance in the semi-skilled equations to insignificance in the skilled equations. The significance of NMANPRES, PBR2 and PJEVAL change in the opposite direction however.

The most surprising result is the finding of a very small (negative) coefficient of the union status variable REC15; this produces a wage gap of −0.4 per cent which is not significantly different from zero. Sherwin Rosen (1970) has argued that the union/non-union differential varies inversely with the degree of substitutability with non-unionised resources. Hence the demand for skilled workers is less elastic than that for the unskilled, which suggests that skilled workers will have a higher wage gap than will semi-skilled workers. Precisely the reverse is found, however; the wage gap for the skilled is *less* than for the semi-skilled, confirming the findings of Ashenfelter (1978) for the United States that unions narrow skill differentials. This may occur if trade unions deliberately used policies to advance the cause of the lower paid, particularly when skilled and semi-skilled workers are in the same union (or bargain together) as frequently occurs in British Industry.

The Manufacturing and non-manufacturing sectors

When the results for the economy as a whole are disaggregated into two sectors, manufacturing and non-manufacturing a number of important differences emerge. The results by skill group are examined separately, starting with the semi-skilled; equation (2) reports results for manufacturing and (3) for non-manufacturing. As might be expected, given the preponderance of establishments in non-manufacturing, equation (3) performs very similarly to equation (1). Given that the vast majority of studies in Great Britain have confined themselves to analysing the manufacturing sector, it is interesting to note how different equation (2) is from equations (1) and (3). (These differences are rather less apparent in the case of the skilled manual worker equations.) The existence of shift systems, share, ownership schemes, and the presence of outworkers all appear to be significant determinants of the earnings of the typical employee in manufacturing, but not in non-manufacturing. The size of establishment, the presence of payments by results schemes amongst males and black workers are significant determinants of the earnings of the typical employee in non-manufacturing but not in manufacturing.

There are considerable differences between the two sectors in the performance of the 'union impact' variables. In the manufacturing sector, the weekly earnings of a typical semi-skilled manual employee are higher, the greater the extent of industry unionism and the more unions there are that represent manual employees at the establishment. Union recognition at the establishment is not a significant determinant of wages in manufacturing; it is very important in non-manufacturing, however, whereas the EXTENT and NMANPRES variables do not have significant effects. When estimates of the wage gap are derived from the union status (recognition) variable, that for non-manufacturing (13.77 per cent) is considerably larger than that for manufacturing (1.97 per cent).

This pattern does disappear, however, in the skilled manual worker equations; the estimated wage gaps not only for the economy as a whole but also for manufacturing and non-manufacturing are all very close to zero. The NMANPRES variable has much the same effect as for the semi-skilled but the EXTENT variable performs very poorly. It is interesting to note that many of the differences between the two sectors that emerged in the semi-skilled equations are repeated here; the LTDCO, NMANPRES, TOTAL, SHIFT and OUTW variables are good examples of this. In addition PJEVAL and PBR2 work in equation (7) but not in (6).

IV Comparisons with Previous Findings

There are a number of similarities between the findings of this study and other studies, both in the United Kingdom and in the United States. For example, a number of studies for the United States have found that the wage gap for non-manufacturing is larger than that for manufacturing and by approximately the same order of magnitude as that reported here for semi-skilled manual workers (eg. Ashenfelter, 1978; Freeman, 1982). There are no directly comparable studies for Great Britain but Thomson, Mulvey and Farbman (1977) who examined crude earnings differentials and

Layard, Metcalf and Nickell (1978) who estimated a coverage differential found that the value of their estimates for manufacturing was lower than that for the economy as a whole.

Stewart (1983a) presented the first estimates of the union/non-union wage differential or wage gap for the United Kingdom, using micro-data at the level of the individual. He estimated an average differential of 7.7 per cent for male manual workers in manufacturing in 1975; this was considerably lower than most previous estimates which, without exception, were derived from aggregate data at either the industry or occupation level. This study not only provides much more recent estimates of the wage gap (for Great Britain in 1980) but also extends the analysis to semi-skilled and skilled manual workers in manufacturing, as well as in non-manufacturing. Moreover, this is the first time that micro-data at the level of the establishment (drawn from the 1980 Workplace Industrial Relations Survey) has been used to estimate wage gaps in Great Britain. This is not the first time that such data have been used, however. A number of recent studies in the United States such as Alpert (1983), Antos (1983) and Freeman (1982) have used establishment data to estimate wage gaps. Such estimates tend to be rather smaller than those derived from individual data (see Freeman and Medhoff, 1981, p. 53).

The wage gaps reported in this paper are consistently smaller than estimates obtained in previous studies that used aggregate data, confirming the work of Stewart (1983a) in Great Britain and Lewis (1983) in the United States. The wage gap in the British economy (excluding agriculture and coal mining) for 'typical' semi-skilled manual workers, estimated at 10.15 per cent and in the non-manufacturing sector at 13.77 per cent, are both significant at the one per cent level. The remaining estimates, including *all* those for 'typical' skilled manual workers are found to be insignificantly different from zero. The wage gaps in manufacturing for both skill groups are even smaller than those obtained by Stewart (1983a), although they refer to different groups and different time periods. Over the five years that separate the two studies, the manufacturing sector experienced a decline of nearly three quarters of a million employees in employment, or approximately 10 per cent of the total workforce (source: Employment Gazette, Feb. 1979; Jan. 1981). In these circumstances it would be unsurprising if unions did not moderate pay claims, because of the possible employment consequences that might accompany high wage demands.

There are a number of similarities between the results obtained in this study and those of Ashenfelter (1978) for the United States. In both studies the wage gaps for semi-skilled manuals (operatives) are larger in non-manufacturing than manufacturing; in Ashenfelter's study this pattern was also repeated for craftsmen but is not replicated in this study for skilled manual workers. Additionally, both studies find evidence that trade unions have the effect of narrowing skill differentials; this contrasts with the findings of Stewart (1983a) who obtained a higher differential the higher the skill level. (It should be noted, however, that the degree of narrowing is slower in the manufacturing sector than it is in the non-manufacturing sector which was not included in Stewart's sample.)

V CONCLUSIONS

In this paper estimates of the union/non-union wage differential were derived from an important new data set, the 1980 Workplace Industrial Relations Survey. This data set used the establishment as its basic unit and included detailed information on both workers and workplace characteristics including whether or not unions were recognised for purposes of bargaining at the workplace. Using this recognition data to determine the union status of the establishment and controlling for differences in worker and workplace characteristics, estimates of union relation wage effects were obtained that were considerably smaller than those obtained in previous studies that had used aggregate data.

Estimated effects were found to be greater in the case of semi-skilled manual workers than for skilled manual workers; estimates derived in the non-manufacturing sector (which had been excluded in virtually all previous studies) tended to be larger than those for the manufacturing sector, with the greatest difference occurring for the semi-skilled. If omitted quality variables are positively correlated with union status then even these estimates will be *upward* biased. Possible solutions to this problem include the use of longitudinal data (to remove 'fixed-effect' error terms) and the estimation of a simultaneous equation system (where the wage equation(s) include a selectivity correction factor such as the Inverse Mills' Ratio).[20] Despite the fact that a number of studies in the United States have used such techniques to control for this selectivity bias,[21] no such adjustments have been made in British studies. Hence the estimates reported here should be regarded as preliminary; the extent of this selectivity bias is the subject of current research.

The inability to control for differences in labour quality and an absence of data on the union status of an establishment in studies using aggregate data led H. Gregg Lewis to argue that estimates of the union/non-union wage differential derived in such studies 'should be ignored in estimating the mean wage gap in the U.S. workforce as a whole and in its parts' (1983, p. 24). I see no reason for altering this view in the case of the British manual workforce.

Appendix A

The matrix of weights applied were as follows:-

	Interviews achieved	*Weighting factors*
A) Single Census Units		
25– 49 employees	359	2.4857
50– 99 employees	395	1.3377
100– 199 employees	374	0.8182
200– 499 employees	356	0.4721
500– 999 employees	264	0.1890
1000–1999 employees	128	0.1735
2000+ employees	88	0.0941
B) Multiple Census Units	76	1.251–0.056
Total	2040	

Individual weights were calculated for each of the 76 Multiple Census Units within the range given above. (These occurred because some establishments were listed in the Census of Employment more than once. Since a Census Unit consists of employees at the same establishment who are paid from the same place, in some instances a single establishment was represented by more than one Census Unit.)

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NOTES

1. For a discussion of the advantages of using 'panel' data the interested reader is referred to Lewis, 1982a and Freeman and Medoff, 1981.

2. It is not clear to me which 'differential' Minford is actually measuring. It appears to bear little resemblance to the others reported in Table 1 (or to those reported here) due to his failure to control for differences in worker and workplace characteristics or time Moreover Nickell (1984) argues that the empirical model from which Minford's estimates are derived is 'fundamentally mis-specified even if one accepts the basic theory expounded by the author' (p 9).
3. Metcalf (1977) surveyed these studies and concluded that 'Unions do appear to influence relative wages. The absolute magnitude of the wage differential they achieve for those covered by collective agreements is uncertain but a figure of 20 per cent may not be dreadfully in error' (p 169) He had warned earlier, however, that 'It is possible that after controlling for labour quality, M (the wage differential) will be found to be small' (p. 164)
4. W is the natural logarithm of the worker's wage $\tilde{M}$ is the partial derivative of W with respect to the individual's union status.
5. Lewis (1983) defines the wage gap for an individual as the excess of his real wage if unionised (covered by a collective agreement) over his real wage if non-union (not so covered) given his working conditions. The term was first coined by Mincer (1981).
6. Lewis calls the differential a 'wage gain', defined as follows:

$$A_i \equiv \frac{W^u - W^a}{W^a}$$

It is the 'excess of an individual's real wage in the presence of unionism over his real wage in the absence of unionism' (1983, p. 24).
7. See Treble and McGrady (1983) for a discussion of the (minor) consequences of using arithmetic rather than geometric means in the measurement of union relative wage effects from aggregate data.
8. The possibility does exist of course, that the non-union wage W^n, could be *lower* than W^a, this could occur, for example, if unions are able to raise wages for their members by reducing employment in that sector, increasing the supply of workers in the non-union sector and hence the equilibrium wage would fall below that existing in the absence of unions
9. See Treble (forthcoming) for a discussion of the problems of correcting for heteroscedasticity where the dependent variable is a logarithm.
10. Defined by principal product.
11. Defined as 'places of employment at a single address or site'.
12. For more details of the weighting procedure used see Daniel and Millward (1983, Appendix B).
13. In all specifications of equation (12) tried, only LTDCO ever reached significance, hence the other 'status of organization' variables are excluded from the results reported here.
14. They argue that 'the differential in non-manufacturing is almost double that in manufacturing, probably reflecting the low wages in some wages council industries in the uncovered sector of non-manufacturing, since it is the no-agreement wage which seems low by comparison with the average rather than the covered wage being high' (p. 181)
15. Data kindly supplied by the Department of Employment.
16. It was also possible to derive an 'extent of industry unionism' variable based upon the concept of recognition, which was used to derive the 'union status' variable. This variable performed almost exactly the same as that based on union membership. The variable based on membership was chosen as it fits in more closely with variables used in previous studies.
17. Data kindly supplied by D. Capron of the Department of Employment.
18. Stewart (1983b) has criticised this method of allocating values to observations and argues that it results in inconsistent estimates. He advocates the use of a 'two-step estimator' to assign each observation the conditional expectation $E[y_i \mid A_{k-1} < y_i \leqq A_k]$ where y_i is the unobserved dependent variable and A_k and A_{k-1} are the upper and lower bounds of the kth range of the logarithm of earnings respectively. Stewart found that utilising such *ad hoc* procedures as the ones adopted here tend to result in an upward bias in the coefficient on the union status variable of up to 20 per cent. The direction and the extent of this bias will be investigated at a later date.
19. Treble (1981) shows that if one accepts the normality of the coefficient on the union status variable the estimate of the wage gap will be log normally distributed. One can then compute its standard deviation from the following formula

$$\text{Var } M = e^{2\mu+2\sigma^2} - e^{2\mu+\sigma^2}$$

where μ = value of coefficient
σ = standard deviation of the coefficient
M = the union/non-union wage differential (or wage gap)

20 See Lewis (1982a,b) for a discussion of these methods.

21. Examples of studies that obtain estimates of the wage gap using longitudinal data are Mellow (1981) and Mincer (1981). Examples of studies estimating simultaneous equation systems are Lee (1979) and Schmidt (1978).

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