The estimation of union wage differentials and the impact of methodological choices

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Received 10 November 1997; accepted 2 February 1998

Abstract

This paper demonstrates that methodological differences can matter a lot in the estimation of union/non-union wage differentials. Using individual-level data from the 1991 Wave of the British Household Panel Survey and a model evolved from replicating six existing British studies, we find that the model specification adopted has an important impact on the estimated differential and that the choice of which group means to use when evaluating the mean differential in multi-equation models is of considerable importance. There are also important differences between membership and coverage differentials and the earnings measure used and sample selected also make a difference. However, apart from firm size, the contents of the control vector used is not found to be of great importance.

JEL classification: J31; J51

Keywords: Wage differentials; Trade unions; Bargaining; Methodology; Replication

1. Introduction

Despite the recent decline in their coverage of the labour market, trade unions and collective bargaining are still of crucial importance to the wage setting process.

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PII: S0927-5371(98)00010-4
in Britain and most industrialized countries. The extent to which trade unions are able to drive a wedge between the pay of comparable union and non-union workers is therefore an important building block in understanding the determination of wages. There have now been many studies of such union/non-union wage differentials, particularly for the US, but also to a more limited extent for Britain and other countries. \(^1\) In order to be able to form judgments about the magnitude of the average differential in a particular country, how these differentials vary with characteristics of workers, firms or markets, and how they move through time, it is important to be able to compare the results from different studies. This is complicated by the many differences in methodology between the available studies. The prime focus of this paper is on these methodological differences between studies and on evaluating how important these differences are when comparing estimated differentials across studies.

We focus on the British labour market to make these methodological comparisons. There have been a number of empirical investigations of union wage differentials for Britain based on individual-level data. \(^2\) We contrast those by Stewart (1983), Green (1988), Blackaby et al. (1991), Blanchflower (1991), Lanot and Walker (1995) and Andrews et al. (1996). These six studies use very different datasets for different points in time and different modelling formulations. We use these to illustrate the methodological differences.

The formulations adopted by these studies differ in a number of dimensions. (i) They use different wage or earnings measures as the dependent variable. (ii) They use different measures of union status. (iii) Different control variables are included in the models. (iv) Different sample definitions are adopted. (v) The form of the econometric model specified differs across the studies. (vi) Where an average differential is evaluated at particular sample means, the group for whom the means are calculated differs. If we are to compare the results from these or other studies, we need to know the impact of these differences on the estimated differentials.

A seventh methodological difference concerns the estimation method used if union status is regarded as being potentially endogenous. Only one of the six studies listed above (Lanot and Walker, 1995) addresses this issue, and while methods for dealing with endogenous selection are interesting conceptually, in practice existing studies in the current context are unconvincing as they are unable to find credible identifying instruments. \(^3\) This remains a contentious issue, and so we prefer to concentrate on the other six methodological differences noted above.

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\(^1\) Lewis (1986) provides a comprehensive survey of the U.S. evidence up to that date.

\(^2\) See Booth (1995) for a recent review of the evidence as well as a discussion of the theoretical setting.

The British Household Panel Survey (BHPS) contains the requisite information to match the formulations of the previous studies in the most important dimensions and can thus be used to evaluate the influence on the resulting estimates of the differences listed above. Hourly, weekly and annual earnings measures can be constructed; both union membership and coverage are recorded; and most of the control variables used in previous studies are available in the dataset. This paper uses the 1991 Wave of the British Household Panel Survey (henceforth BHPS91) to estimate union wage differentials for that year and examines how these differentials vary with the methodological features listed above. 4

The analysis is conducted in two stages. In the first stage we replicate each of the specifications used in the six previous individual-level studies of union wage differentials in turn and estimate comparable models using our BHPS91 data. Even for a single dataset at a single point in time we find considerable variation in the estimates, due to differences in methodology. In the second stage we attempt to reconcile the findings from these various comparisons by adopting an over-simplified ‘base model’ and then investigating the impact on the estimated differentials of various deviations from this simple model in the direction of each of the methodological differences listed above.

The main findings of the paper are that the methodological differences between the studies compared matter a lot. There is an important difference between differentials with respect to union membership and differentials with respect to union presence or coverage at the workplace; it also matters to some degree whether we look at differentials in hourly, weekly or annual earnings. The formulation of the model adopted also matters. However, as long as some key controls are included, the inclusion or exclusion of the bulk of them does not seem to be important. In contrast the choice of which group’s means to use in the evaluation of the mean differential is found to be of considerable importance.

The paper is structured as follows. The next section discusses in more detail the six dimensions of methodological difference identified above, and presents a formal methodology for comparing union wage differentials from different model specifications. Section 3 presents the results of both our replications of the previous studies and our analysis of the impact of differences in methodology on the estimated differentials. Section 4 presents our conclusions.

2. A methodological framework

2.1. Methodological differences

The methodologies employed by the six studies we examine differ from one another in six main dimensions. These can be illustrated by considering a general

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4 We exclude individuals with imputed earnings.
framework for measuring ceteris paribus union/non-union wage differentials as follows. Given a vector of union characteristics, $u$, and a vector of other (control) factors, $x$, the conditional expectation of the log-wage for the $i$th individual in a selected group, $S$, is given by

$$E(\ln w_i | u_i, x_i) = f(u_i, x_i)$$

Taking $e_i$ to be the difference between this expectation and the observed log-wage, the wage determination model is given by

$$\ln w_i = f(u_i, x_i) + e_i$$

The function $f$, while linear in parameters, may be non-linear in variables if, for example, the model contains interaction terms. If the derivatives of $f$ with respect to the elements of $u$ vary with the characteristics of the individual, then a mean differential ($\Delta$) is calculated either by averaging the individual differentials ($\Delta_i$) or by evaluation at some specified point of means. Thus the mean differential is either estimated as

$$\Delta = \frac{1}{n_G} \sum_{i \in G} \{ \hat{f}(u_i^1, x_i) - \hat{f}(u_i^0, x_i) \} = \frac{1}{n_G} \sum_{i \in G} \Delta_i \quad \text{Method A}$$

or

$$\Delta = \hat{f}(n^G, x^G) - \hat{f}(n^{0G}, x^G) \quad \text{Method B}$$

where $n_G$ is the number of individuals in group $G$, $u_i^1$, $u_i^0$ are the ‘union’ and ‘non-union’ configurations of $u_i$, respectively, $\hat{f}$ denotes that the parameters of $f$ have been estimated and a superscript $G$ indicates an average over those in group $G$. The issues that the modeller must address and that lead to the main methodological differences between the papers can then be categorised as relating to the specification of $w$, $u$, $x$, $f$, $S$ and $G$. Between them the six studies that we compare differ in all six of these dimensions in addition, of course, to using different datasets referring to different years. Information on this and other details of the six studies is given in Table 1.

2.1.1. The earnings measure ($w$)

The most commonly used earnings measure in these studies is gross weekly earnings. However, Lanot and Walker (1995) use hourly earnings and Blanchflower (1991) uses annual earnings. Since we can write $\ln w = \ln(w/h) + \ln h$, where $h$ is the number of hours worked per week and $w$ is, for the moment, weekly earnings, then the difference between otherwise comparable estimates of effects on weekly and hourly (log) earnings will be the corresponding estimated union effects on the log of hours worked per week. Oswald and Walker (1993) using FES data and Stewart and Swaffield (1995) using BHPS data find that union
workers work fewer hours per week than non-union workers on average. Thus, other things equal, we might expect union wage differentials to be somewhat higher when measured in terms of hourly earnings than when measured in terms of weekly earnings.

Similarly, since we can write \( \ln(we) = \ln w + \ln e \) where \( e \) is the number of weeks worked per year, the difference between otherwise comparable estimates of effects on annual and weekly (log) earnings will be the corresponding estimated union effects on the log of weeks worked per year. Since union workers typically have longer job tenure and lower turnover rates, we might expect them to have fewer weeks out of work in the previous year. In this case, other things equal, union differentials in annual earnings will be higher than those in weekly earnings.

2.1.2. Union status variables (\( u \))

Two types of union variable are included in these studies, those that indicate membership and those taken to indicate union presence at the workplace in a form potentially affecting the pay of the individual (whether they are a member or not). The latter type we will label ‘coverage’, although in some datasets the measure available is only a proxy for this. Green (1988), Blackaby et al. (1991) and Blanchflower (1991) have measures of both individual membership and coverage, Stewart (1983) and Lanot and Walker (1995) have only membership and Andrews et al. (1996) only coverage.

If all workers in receipt of the union negotiated rate get a fixed mark-up over non-union pay, then we would expect the coverage differential to be the important one, with membership not having an additional influence on pay. In fact where both types of information are available in a dataset, it is found that covered members get paid more than covered non-members. One possible explanation of this is provided by the establishment-level evidence. The presence of a closed shop, or more generally high union density, at the establishment results in an above-average union wage differential and coverage on its own is not enough (Stewart, 1987). At the individual level, while members and non-members doing the same job in the same establishment will earn the same, when comparing across establishments, membership is a closer indicator of a differential than coverage. (The conditional probability of a closed shop or high density given membership is higher than that given coverage.)

The studies that employ a coverage variable differ considerably in its definition. The GHS asks ‘Is there a trade union or staff association where you work, which

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\footnote{The explanation of Oswald and Walker (1993) of this is that the contract curve in wage–hours space, which lies to the left of the labour supply schedule for non-union workers, is downward-sloping. However, Stewart and Swaffield (1995) present evidence that union workers are no more likely to be rationed below desired hours than non-union workers.}
Table 1
Comparison of the six studies and their replications for BHPS91

<table>
<thead>
<tr>
<th>Author</th>
<th>Data* Year</th>
<th>Dependent variable</th>
<th>Sample selection and size</th>
<th>Method usedb Union info.c</th>
<th>Original differentialsâ</th>
<th>1991 BHPS differentialsâ</th>
</tr>
</thead>
<tbody>
<tr>
<td>Stewart (1983)</td>
<td>NTS 1975</td>
<td>Log gross weekly earnings</td>
<td>Full-time manual male employees in manufacturing; $N = 5352$</td>
<td>(4) with $\theta_0 - \theta_1 = 0$</td>
<td>M</td>
<td>0.074 (0.011)</td>
</tr>
<tr>
<td>Green (1988)</td>
<td>GHS 1983</td>
<td>Log gross weekly earnings</td>
<td>Full-time manual employees; $N = 2411$</td>
<td>(4)</td>
<td>$C^c$</td>
<td>−0.007</td>
</tr>
<tr>
<td>Blackaby et al.</td>
<td>GHS 1983</td>
<td>Log gross weekly earnings deflated by a monthly index of average earnings</td>
<td>Full-time manual male employees; $N = 2355$</td>
<td>(6)</td>
<td>$M$</td>
<td>0.116</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>$U$</td>
<td>0.109</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>0.115</td>
<td>0.133 (0.039)</td>
</tr>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>0.198</td>
<td>0.222 (0.038)</td>
</tr>
<tr>
<td>Source</td>
<td>Survey</td>
<td>Period</td>
<td>Variable Description</td>
<td>Sample Size</td>
<td>Estimate 1</td>
<td>Estimate 2</td>
</tr>
<tr>
<td>-------------------------------</td>
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<td>---------------------------------------------------------------------------------------</td>
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</tr>
<tr>
<td>Blanchflower (1991)</td>
<td>BSAS (R)</td>
<td>1983–1989</td>
<td>Log annual earnings, based on mid-points of grouped data</td>
<td>N = 5270</td>
<td>0.041</td>
<td>0.017</td>
</tr>
<tr>
<td>Lanot and Walker (1995)</td>
<td>FES (LW)</td>
<td>1978–1986</td>
<td>Log real hourly gross reported earnings</td>
<td>N = 20,727</td>
<td>0.078</td>
<td>0.003</td>
</tr>
<tr>
<td>Andrews et al. (1996)</td>
<td>NES (ABU)</td>
<td>1978</td>
<td>Log real gross weekly earnings</td>
<td>N = 46,790 (1978) and 38,609 (1985)</td>
<td>0.041</td>
<td>(0.017)</td>
</tr>
</tbody>
</table>


See text for equation numbers.

Membership, Coverage, and total union (see text for full explanation).

Standard errors in parentheses.

Constructed from model 4 with $p = 0.064$. 
people in your type of job can join if they want?'', with no mention as to whether
the union is recognised for bargaining purposes. The BSAS asks about trade
unions or staff associations at the place of work recognised by the management for
negotiating pay and conditions, but does not ask whether they cover the type of
job that the individual does. The NES indicates pay covered by a negotiated
collective agreement. The BHPS data used here asks ‘‘Is there a trade union, or
similar body such as a staff association, recognised by your management for
negotiating pay or conditions for people doing your sort of job in your workplace?’’

There is less, but still some, difference in the membership questions. In most
cases individuals are asked whether they are members of a trade union or staff
association. The NTS question only refers to trade unions without mention of staff
associations. The FES provides rather different information, indicating only
whether trade union dues are deducted from the respondent’s earnings at source or
not. This will under-report membership, but will have moved closer to it over time
as the incidence of ‘check-off’ has increased.

2.1.3. The sample selected (S)
The most commonly used sample is manual male full-time employees. We take
this as the ‘baseline’ sample in the discussion here. Relative to this, Stewart
includes full-time female employees, with an additive dummy, as well as males
and estimates a slightly lower differential for female workers. Yaron (1990) finds
this in a more general specification (also estimated on the GHS). This suggests
that inclusion of female workers in the sample will reduce the overall mean.
Blanchflower (1991), however, finds higher differentials for women. Blanchflower
(1991) also includes non-manuals and part-timers. The available evidence suggests
that union differentials are much smaller for non-manual than manual workers.
The impact of this exclusion is uncertain a priori.

2.1.4. The control vector (x)
All six studies use different control vectors from one another and it is not
possible to predict a priori the combined effect of the differences between any two
studies. However, two factors are worthy of particular mention in this context.
Existing evidence suggests that firm size is strongly correlated with both unionisa-
tion and pay. Larger firms pay more and are more likely to be unionised. Thus

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6 The sample used by Blanchflower (1991) gives a mean of 0.469 for this variable compared with a
mean of 0.464 for the membership variable. These seem far too close together relative to other data
sources and cast some doubt on the interpretation of this variable.
exclusion of controls for firm size may be expected to result in higher differentials. Blanchflower (1991), Lanot and Walker (1995) and Andrews et al. (1996) do not include controls for firm size and might be expected to be affected by this. 7

A second potentially important variable not included by some of the studies is job tenure. Existing evidence suggests that this too is (positively) correlated with both unionisation and pay. Thus exclusion of controls for job tenure may also be expected to result in higher differentials. The studies by Stewart (1983), Blanchflower (1991) and Lanot and Walker (1995) do not include controls for job tenure and would be affected by this.

2.1.5. The form of the econometric model (f)

Four of the six studies use separate equations by union status, equivalent to including a complete set of interactions with the union status variable(s). Blanchflower (1991) and Andrews et al. (1996) use single equations with only additive union status dummies. Andrews et al., who only have coverage information, use a single coverage dummy. Their specification thus assumes that the coverage differential is the same for all individuals. Blanchflower uses additive dummies for membership and coverage. This specification, in addition to the constancy of these differentials, assumes the membership effect to be the same for those who are covered and those who are not.

The other four studies use separate equations by union status. Stewart (1983) and Lanot and Walker (1995) only have membership information and partition their samples accordingly. Green (1988) estimates separate equations for members and non-members and includes a coverage dummy in each equation. Thus the effect of membership is allowed to vary with all other characteristics, but the effect of coverage only between members and non-members. Blackaby et al. (1991) estimate separate equations for covered members, covered non-members and uncovered non-members. Uncovered members are excluded from the sample. The impact of these alternative formulations is difficult to see a priori and is examined in more depth in Section 2.2.

2.1.6. Definition of the mean differential (G)

When separate equations are estimated, an additional issue to be addressed is which means to use to evaluate the average differential. This is a familiar index number problem. Which mean is appropriate depends on the question that one wants to ask. Stewart (1983) argues that the interesting question is how much more does the unionised individual earn than would be the case if he or she were

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7 In the original version of Andrews et al. (1996) that we replicated, they did not include firm size. In the version referenced they have since added firm size to the control vector, and find that it roughly halves the differential for most specifications.
non-union, but everything else remained the same. Hence the appropriate average is over those who are unionised, i.e., those currently in receipt of such a differential. If non-union means are used, a similar question is asked, but of a non-union individual. The use of the overall mean gives a weighted average of the two. One disadvantage of the use of the overall mean is that the answers to the two questions posed above can remain the same, but the calculated average differential can change if the proportion unionised does.

Given that the characteristics of union and non-union workers differ in a number of regards, the estimated average differential may well differ according to whether union or non-union means (or a combination of the two such as the overall means) are used. Green (1988) evaluates the average differential at the overall sample means. Blackaby et al. (1991) calculate two types of differential and in each case evaluate at the non-union means. They calculate a differential for covered members over covered non-members, evaluating at the means for covered non-members, and they calculate a differential for covered members over uncovered non-members, evaluating at the means for uncovered non-members. Lanot and Walker (1995) present average differentials calculated for both the member and non-member samples and find the two to be very similar for their particular measure of union membership. Evidence given below indicates that this is not always the case.

2.2. Comparing differentials from different model specifications

The interpretation of an estimated union wage differential depends crucially on the specification of the model from which it is derived. Here we examine in more detail how the interpretation depends on the specification of the function \( f \) and the variables used in the vector \( u \). We present a classification system to assist in the comparison of differentials from studies based on different specifications.

We take as our starting point that we are interested in the differentials between three groups of workers defined as follows:

\[
\begin{align*}
C = 1, & \quad M = 1; \\
C = 1, & \quad M = 0; \\
C = 0 & \quad (M = 1 \text{ or } M = 0);
\end{align*}
\]

where \( M = 1 \) indicates union membership and \( C = 1 \) indicates union ‘coverage’ (more specifically that there is a union at the place of work which is recognised for bargaining). The main reason for choosing these three categories, rather than the four that could be derived from the variables \( M \) and \( C \), is that these are the three groups of prime economic interest. In addition, on practical grounds, there are typically not enough non-covered union members to treat them as a separate group and estimate differentials with any degree of precision. We assume that being a union member does not influence the level of pay of uncovered workers.
We regard this as a more reasonable assumption than simply discarding this group. To emphasise, where \( C = 0 \) we ignore the distinction between members and non-members.

Using this categorisation, we can define the following differentials: 9

\[
\Delta_y(C; M = 1) = E(\ln w_i | C_i = 1, M_i = 1; x_i) - E(\ln w_i | C_i = 0; x_i)
\]

\[
\Delta_y(C; M = 0) = E(\ln w_i | C_i = 1, M_i = 0; x_i) - E(\ln w_i | C_i = 0; x_i)
\]

The union membership differential is defined as follows:

\[
\Delta_y(M; C = 1) = E(\ln w_i | C_i = 1, M_i = 1; x_i) - E(\ln w_i | C_i = 1, M_i = 0; x_i)
\]

and for a given individual is related to the other two by:

\[
\Delta_y(M; C = 1) = \Delta_y(C; M = 1) - \Delta_y(C; M = 0). \tag{1}
\]

\( \Delta_y(C; M = 0) \) can be termed the union coverage differential and \( \Delta_y(C; M = 1) \) the total union differential. Average differentials can then be calculated by averaging these individual differentials across the individuals in some group \( G \) (Method A) or by evaluation at the group \( G \) sample means of the \( x \)'s (Method B).

If the same group is used for the calculation of the average differentials corresponding to each of the three individual differentials defined above, then Eq. (1) will hold for the average differentials as well as for those for a particular individual. However, while it might be appropriate to construct \( \Delta_y(C; M = 1) \) by averaging across either the \( (C = 1, M = 1) \) or \( C = 0 \) groups, it is not appropriate for the \( (C = 1, M = 0) \) group. A similar argument holds for the other differentials. While Eq. (1) holds for each individual, it only holds for average differentials if they are evaluated at the same means. This might be viewed as inappropriate since it always includes a group for whom the comparison is irrelevant. Adding-up makes less sense at the aggregate level than for an individual worker.

We now turn to the parameterisation of these differentials. When there is a single binary union status variable, typically estimates of the differential are derived either from a single equation with the dummy variable included, or from separate equations on ‘union’ and ‘non-union’ sub-samples. Generalising to three groups of worker generates a large number of models, with an even larger number of implied estimates of the three differentials defined above. Two types of model can be distinguished: single equation models, where the function \( f \) is taken to be additively separable in \( u \) and \( x \) (type I models), and separate equation models, where \( f \) allows the effects of the elements of \( u \) on \( \ln w \) to vary with \( x \) (type II models).

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8 Of the BHPS sample used for the ‘base’ model in Section 3, 24 out of 906 individuals fall into this category. This represents 2.6% of the sample and 6.4% of those not covered.

9 Strictly the differentials in each case are given by \( d_y = \exp(\Delta_y) - 1 \). However, for simplicity of terminology we will also refer to the \( \Delta_y \)’s as differentials throughout. To aid the distinction between the two we adopt the convention that \( \Delta_y \) is written as a decimal (e.g., 0.226) whereas \( d_y \) is written and discussed in percentage terms (e.g., 25.3% when \( \Delta_y = 0.226 \)).
2.2.1. Type I: single equation models

Given the above focus, the appropriate parameterisation of a single equation regression model, when information on both $M$ and $C$ is available, is

$$\ln w_i = x_i' \beta + \alpha_i M_i C_i + \alpha_2 (1 - M_i) C_i + \epsilon_i.$$  \hspace{1cm} (2)

Expected log–pay in each of the $C, M$ groups relative to that in the $C = 0, M = 0$ group is given by:

<table>
<thead>
<tr>
<th>$C = 0$</th>
<th>$M = 0$</th>
<th>$M = 1$</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\alpha_2$</td>
<td>0</td>
<td>0</td>
</tr>
<tr>
<td>$\alpha_1$</td>
<td>0</td>
<td>1</td>
</tr>
</tbody>
</table>

The three differentials defined above can be calculated as $\Delta(C; M = 1) = \alpha_1$, $\Delta(C; M = 0) = \alpha_2$, and $\Delta(M; C = 1) = \alpha_1 - \alpha_2$. (Corresponding average and individual differentials are the same since the latter do not vary with $i$.) $\alpha_2$ is interpreted as the union coverage differential, $\alpha_1$ as the total union differential. The union membership differential is given by subtraction.

Eq. (2) can be contrasted with

$$\ln w_i = x_i' \beta + \gamma_1 M_i + \gamma_2 C_i + \epsilon_i.$$  \hspace{1cm} (3)

Expected log–pay in each of the $C, M$ groups relative to that in the $C = 0, M = 0$ group in this case is given by:

<table>
<thead>
<tr>
<th>$C = 0$</th>
<th>$M = 0$</th>
<th>$M = 1$</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>$\gamma_1$</td>
<td>$\gamma_1 + \gamma_2$</td>
</tr>
<tr>
<td>$\gamma_2$</td>
<td>0</td>
<td>1</td>
</tr>
</tbody>
</table>

The contrast between the two equations is clear. Eq. (2) takes $E(\ln w_i | C_i = 0, M_i = 1; x_i) - E(\ln w_i | C_i = 0, M_i = 0; x_i)$ to be zero. Eq. (3) takes it to be equal to $E(\ln w_i | C_i = 1, M_i = 1; x_i) - E(\ln w_i | C_i = 1, M_i = 0; x_i)$. We argued above that the first of these is the more plausible a priori assumption to make.

On the basis of Eq. (3) the three differentials are $\Delta(M; C = 1) = \gamma_1$, $\Delta(C; M = 0) = -M \gamma_1 + \gamma_2$, and $\Delta(C; M = 1) = (1 - M) \gamma_1 + \gamma_2$. Common practice is to regard $\gamma_1$ as the membership differential and $\gamma_2$ as the coverage differential. The former is obviously appropriate whatever the choice of $G$. The latter depends on the choice of $G$. With $G$ as the whole sample it is $-\bar{M} \gamma_1 + \gamma_2$, where $\bar{M}$ is the proportion of the whole sample who are union members; when $G$ is non-covered workers it is $-\bar{p} \gamma_1 + \gamma_2$, where $\bar{p}$ is the proportion of non-covered workers who are union members. The coverage differential is only $\gamma_2$ when averaging over a group for which $M = 0$, that is covered non-members.
While Eq. (3) might be viewed as a natural formulation of the effects and has been used in the literature, in the context of the structure laid out above it would only be strictly appropriate in the case where \( \bar{p} \) was zero. In this case Eqs. (2) and (3) are reparameterisations of one another. \(^{10}\) The bigger \( \bar{p} \), the bigger the gap between the two structures. \(^{11}\)

Both models are nested within a more general one, which adds a term \( \alpha_3 M(1 - C) \) to Eq. (2). This, of course, estimates a separate conditional mean for all four cells in the \( C \) by \( M \) crosstabulation. Testing Eq. (2) against this joint alternative tests whether \( \alpha_3 = 0 \), that is it tests whether it is valid to group the two types of non-covered worker together as suggested. However, testing Eq. (3) against the joint alternative tests \( \alpha_3 = \alpha_1 - \alpha_2 \). It is a test of independence, as formulation (3) implies \( \Delta(M) = E(\ln w_i | M = 1; x_i) - E(\ln w_i | M = 0; x_i) = \gamma_1 \) is independent of \( C_i \), and \( \Delta(C) = E(\ln w_i | C = 1; x_i) - E(\ln w_i | C = 0; x_i) = \gamma_2 \) is independent of \( M_i \). When the number of uncovered members is relatively small (as indicated by \( \bar{p} \) for example), as is typically the case, it would not be surprising to find that neither approach can be rejected against the joint alternative. \(^{12}\)

Two of the six studies listed in Table 1 use single equation models. Blanchflower has information on both \( M \) and \( C \) and estimates Eq. (3). Andrews et al. only have coverage information and estimate Eq. (3) with \( \gamma_1 = 0 \).

2.2.2. Type II: separate equation models

If we generalise the specification to allow the impact of the \( x \)-vector to vary with membership, but not coverage, the regression model can be parameterised as model (4) in Table 2, which also shows associated individual-level differentials. Eq. (1) will hold for each individual \( i \), but will only hold for average differentials if the same group \( G \) is used for all three differentials. As discussed above, this only makes sense if it is the overall (whole sample) mean that is used.

In evaluating average differentials, one can apply either Method A or Method B for any one of four groups: the whole sample and our three groups of interest. For the membership differential both methods give the same outcome, noting that the

---

\(^{10}\) This suggests another problem with using the overall mean. As the number of covered workers goes to zero the three differentials should converge to \( \Delta(M; C = 1) - \gamma_1 \), \( \Delta(C; M = 0) - \gamma_2 \), and \( \Delta(C; M = 1) - \gamma_1 + \gamma_2 \). This does not happen when \( G \) is the whole sample, as \( \bar{M} \) does not converge to zero.

\(^{11}\) Since \( \bar{p} \) is typically small (see footnote 8), the difference should not be great in practice.

\(^{12}\) To illustrate, for the raw BHPS data, the raw coverage differential is 0.069 and the raw membership differential is 0.213, giving a raw total union differential of 0.282. When independence is imposed, i.e., Eq. (3) is estimated without controls, the three differentials are 0.097, 0.194, and 0.291 respectively. Using Eq. (3) rather than Eq. (2) does not misrepresent the effects by a large amount (due to the fact that \( \bar{p} \) is relatively small). However, there is some distortion of the relative magnitudes of the (raw) coverage and membership differentials. The approach laid out above has the virtue of not imposing independence.
uncovered group is inappropriate in this case. For the coverage differential, $\Delta(C; M = 0)$, the two methods give the following:

<table>
<thead>
<tr>
<th></th>
<th>Model</th>
<th>Individual differentials</th>
</tr>
</thead>
<tbody>
<tr>
<td>(4) $\ln w_i = \begin{cases} x'_i \beta_1 + \theta_0 C_i + \epsilon_i, &amp; \text{if } M_i = 0 \ x'_i \beta_1 + \theta_0 C_i + \epsilon_i, &amp; \text{if } M_i = 1 \end{cases}$</td>
<td>$\Delta(M; C = 1) = \frac{x'_i (\beta_1 - \beta_0)}{\theta_1 - \theta_0}$</td>
<td>$\Delta(C; M = 0) = -M \frac{x'_i (\beta_1 - \beta_0)}{\theta_1 - \theta_0}$</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$\Delta(C; M = 1) = (1 - M) \frac{x'_i (\beta_1 - \beta_0)}{\theta_1 - \theta_0}$</td>
</tr>
<tr>
<td>(5) $\ln w_i = \begin{cases} x'_i \beta_0 + \phi_0 M_i + \epsilon_i, &amp; \text{if } C_i = 0 \ x'_i \beta_1 + \phi_0 M_i + \epsilon_i, &amp; \text{if } C_i = 1 \end{cases}$</td>
<td>$\Delta(M; C = 1) = \phi_1$</td>
<td>$\Delta(C; M = 0) = x'_i (\beta_1 - \beta_0) - M \phi_0$</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$\Delta(C; M = 1) = x'_i (\beta_1 - \beta_0) + (\theta_1 - M \phi_0)$</td>
</tr>
<tr>
<td>(6) $\ln w_i = \begin{cases} x'_i \beta_0 + \epsilon_i, &amp; \text{if } C_i = 0 \ x'_i \beta_0 + \epsilon_i, &amp; \text{if } C_i = 1, M_i = 0 \ x'_i \beta_1 + \epsilon_i, &amp; \text{if } C_i = 1, M_i = 1 \end{cases}$</td>
<td>$\Delta(M; C = 1) = x'_i (\beta_1 - \beta_0)$</td>
<td>$\Delta(C; M = 0) = x'_i (\beta_1 - \beta_0)$</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$\Delta(C; M = 1) = x'_i (\beta_1 - \beta_0)$</td>
</tr>
<tr>
<td>(7) $\ln w_i = \begin{cases} x'_i \beta_0 + \epsilon_i, &amp; \text{if } C_i = 0, M_i = 0 \ x'_i \beta_0 + \epsilon_i, &amp; \text{if } C_i = 1, M_i = 0 \ x'_i \beta_1 + \epsilon_i, &amp; \text{if } C_i = 1, M_i = 1 \end{cases}$</td>
<td>$\Delta(M; C = 1) = x'_i (\beta_1 - \beta_0)$</td>
<td>$\Delta(C; M = 0) = x'_i (\beta_1 - \beta_0)$</td>
</tr>
<tr>
<td></td>
<td></td>
<td>$\Delta(C; M = 1) = x'_i (\beta_1 - \beta_0)$</td>
</tr>
</tbody>
</table>

where $\bar{x}_M$ contains means for $M = 1$ observations, $\bar{x}_{01}$ contains means for the $C = 0, M = 1$ observations, $\bar{x}$ contains means for the whole sample and $\bar{x}_0$ for the $C = 0$ observations. In this context, $\bar{p}$ is the sample estimate of $\bar{E}(M_i = 1 | C_i = 0)$ and $\bar{M}$ is the sample estimate of $\bar{E}(M_i = 1)$. The objection that might be raised to the two expressions which contain $\bar{M}$ is given in footnote 10, and so $\bar{M}$ can be replaced by $\bar{p}$. Similar expressions to the above six can be derived for the union differential.

Among the studies we are considering, Green estimates model (4) but only discusses membership differentials. Stewart (1983) and Lanot and Walker (1995) only have access to membership data, and both estimate model (4) with $\theta_0 = \theta_1 = 0$.

An alternative formulation to model (4) which allows the impact of the $x$-vector to vary with coverage, but not membership is model (5) of Table 2. As above, estimated differentials could replace $M_i$ by $\bar{p}$, the proportion of non-covered workers who are union members or could adopt an alternative approach equivalent to one of those discussed for model (4). Blanchflower presents estimates of model (5), but uses the estimates of Eq. (3) to calculate his estimated wage differentials.

Note that both models (4) and (5) have a dummy variable with separate coefficients in the two equations which capture the difference in expected (log)
wages between uncovered members and non-members. Given the arguments above, this might be viewed as inappropriate and clearly gives rise to some problems with regard to the calculation of differentials.

A more general strategy is to split the sample into three categories, as given by model (6). Again Eq. (1) holds if and only if the same mean is used for the three differentials. While Eq. (2) is nested within model (6), models (3) and (4) are not and model (5) only is under the restriction that $\phi_0 = 0$. Note also that model (6) is more heavily parameterised than these other models, resulting in some loss of precision. An alternative three equation model, used by Blackaby et al., specifies model (7) in Table 2, and excludes those with $C = 0$ and $M = 1$ from the sample. However this exclusion seems hard to justify. The differentials are as for model (6) but with $\beta_{\text{w}}$ in place of $\beta_0$, although the coverage and total union differentials will not be completely comparable in terms of interpretation.

3. Results

3.1. Replication of six original studies on BHPS91

The differentials estimated in our first stage replications of the six original studies using BHPS91 are summarised in the final column of Table 1. These are set along side the differentials obtained from the original study. Most of the existing evidence suggests a total union differential in the region of 8–12%, with Blackaby et al.’s 22% somewhat of an outlier. Notice that both Green (1988) and Blackaby et al. (1991) use the 1983 GHS, but their estimates are not directly comparable due to the different model constructions and different samples selected. While they give very similar membership differentials, the total union differential of Blackaby et al. is roughly twice that of Green.

The ‘1991 BHPS’ column of Table 1 indicates that differentials based on the same dataset can vary widely. Contrast the three estimates for the Green specification with those of the Blanchflower specification. Leaving aside differences in sampling and measurement, variations in the estimates reported in the column headed ‘differential’ are due to variations in the date of the dataset used and variations in methodology; variations in the estimates reported in the ‘1991 BHPS’ column are due to variations in methodology alone. We seek to understand why there is such wide variation in the ‘1991 BHPS’ column of Table 1. For example, why does the Green specification deliver a total

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13 The full versions of these replications are available from the corresponding author on request.
14 22% corresponds to $D = 0.198$ (see footnote 9).
15 By comparing the two columns, in principal one should be able to piece together movements in the differential through time, although, in fact, no sensible pattern to the time-series movement of the differentials can be detected.
union differential of around 8%, while the Blanchflower specification generates one of 34%? Section 3.2 therefore examines how the differentials calculated from the BHPS91 replications vary with differences in the sample selected and definition of the dependent variable, with differences in the specification of the model and structure of the union effects in the model adopted, with differences in the specification of the control vector and with differences in the group over which the average differentials are calculated.

3.2. The impact of variations in specification and methodology

This stage of the analysis starts by reporting results for a ‘base model’ specification which can then be compared with the formulations used in the existing literature. We select full-time manual male employees in all industries and consider their gross weekly pay. Our base model is Eq. (2), and we specify a simple set of control variables containing the most important ones from various studies, including (2-digit) density, but excluding industry dummies and other industry-level variables. In this base model the total union differential of 0.076 is made up of a membership differential of 0.056 and a coverage differential of 0.020, the latter being insignificantly different from zero.

3.2.1. Model specification and the structure of union effects

We first consider the various ways it is possible to construct differentials, as outlined in Section 2 above. Instead of estimating Eq. (2) we estimate Eqs. (3–6) using the same specification of $w$, $x$ and $S$. Table 3, column ‘All’ reports differentials evaluated at overall sample means for all such specifications. The traditional approach (e.g., Blanchflower, 1991), which imposes independence, reduces the union differential by 2%. It increases the coverage differential by a similar amount when evaluated at $M = 0$. Formally, as expected, the two specifications are observationally equivalent, with encompassing $t$-statistics of 0.45 and 1.11 for Eqs. (2) and (3), respectively.

The most unrestricted specification is model (6), which splits the sample three ways. For direct comparisons with single equation methods, differentials are calculated at the average characteristics of the whole sample. Again the union effect is dominated by the membership differential, which is now 0.121. The coverage differential is insignificantly different from zero.

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16 More precisely, those (i) who either ‘worked last week’ or ‘normally work’, and (ii) who work 30 h a week or more. Manual workers are those whose SEG definition includes the word ‘manual’ plus agricultural workers. Their pay is usual gross pay in current job adjusted to a weekly basis. The choice of full-time manual male employees matches four of the six studies. The use of all industries matches five of the six studies.

17 Full definitions, regression results and variable means are contained in the working paper version of this paper and are available on request.

18 Standard errors are calculated as indicated by Stewart (1987).
This and the base model are the more appealing specifications (a Chow test of the latter against the former delivers an $F$-statistic of 1.55, $p$-value 0.007) as the other specifications contain the equivalent of a dummy variable for uncovered members. Recall that there are only 24 employees in this group, and our methodology seeks to group them with uncovered non-members, rather than isolate them as in models (4) and (5). Generally the estimates of the differentials are closer to those from the three-way split than those from single-equation methods.

Finally, in two specifications membership information is missing. If we compare the third panel of Table 3 with the second it would appear that the coverage effect picks up most of the membership effect in this case. This is not true when comparing the seventh panel with the sixth, but does suggest that studies which estimate Eq. (3) (e.g., Andrews et al., 1996) where only coverage is available may not be too misspecified relative to Eq. (3) with membership included.

To conclude, there is variation in the estimated effects across the specifications. Estimates of the membership differential (evaluated at overall means) vary between 4% and 13% depending on the formulation used. In all cases the membership differential dominates the coverage one. This is sometimes viewed as a counterintuitive result as coverage is viewed as the more appropriate measure of unionism for collective bargaining. However, stronger membership than coverage effects can also be found in the three original studies which used both variables (Green, 1988; Blanchflower, 1991; Blackaby et al., 1991) (see Table 1). As discussed in Section 2, this is what would be predicted in a model in which there is variation in the differential with the ‘strength’ of the union and this strength is related to the presence of a closed shop and/or high union density.

Lewis (1986) reviewing earlier U.S. work (including Jones (1982) and Mincer (1983)) as well as his own estimates concludes that membership differentials are slightly bigger than coverage differentials, but regards the distinction as relatively unimportant for the U.S. More recent studies investigate the relationship between membership and coverage rather than viewing them as alternatives as in the earlier work. Schumacher (1997), for example, seeks to explain why there is a penalty associated with free-riding (i.e., why there is a positive membership differential in our terminology). His basic framework is the same as ours, Eq. (2), and he finds coverage and membership differentials both of the order of ten percent. In a similar vein, Barth et al. (1998), using Norwegian data, and Reilly (1996), using Canadian data, find that membership differentials disappear once establishment-level union density is included, supporting our conjecture in Section 2.

While variation in the way the union effect is constructed can explain some of the variation in estimated differentials from the same dataset, it does not explain all the differences found in Table 1. In Table 3 we note where the methodology behind a differential coincides with a published study, which can then be compared with Table 1. For example, if we consider Andrews et al., Table 3 reports a coverage differential of 0.057 using the other features of the base model,
Table 3
Variation in union effects

<table>
<thead>
<tr>
<th>Model</th>
<th>$\Delta_i$</th>
<th>Evaluated at the mean characteristics of:</th>
<th>$C = 0$</th>
<th>$C = 1$, $M = 0$</th>
<th>$C = 1$, $M = 1$</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Base model (2)</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Coverage</td>
<td>$\alpha_2$</td>
<td>0.020 $(0.038)$</td>
<td>0.020 $(0.038)$</td>
<td>0.020 $(0.038)$</td>
<td>0.020 $(0.038)$</td>
</tr>
<tr>
<td>Membership</td>
<td>$\alpha_1 - \alpha_2$</td>
<td>0.056 $(0.039)$</td>
<td>0.056 $(0.039)$</td>
<td>0.056 $(0.039)$</td>
<td>0.056 $(0.039)$</td>
</tr>
<tr>
<td>Union</td>
<td>$\alpha_1$</td>
<td>0.076 $(0.031)$</td>
<td>0.076 $(0.031)$</td>
<td>0.076 $(0.031)$</td>
<td>0.076 $(0.031)$</td>
</tr>
<tr>
<td><strong>Traditional (3)</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Coverage $^b$</td>
<td>$\gamma_2$</td>
<td>0.036 $(0.034)^b$</td>
<td>0.036 $(0.034)^b$</td>
<td>0.036 $(0.034)^b$</td>
<td>0.036 $(0.034)^b$</td>
</tr>
<tr>
<td>Coverage</td>
<td>$\gamma_2 - M \gamma_1$</td>
<td>0.019 $(0.046)$</td>
<td>0.034 $(0.035)$</td>
<td>0.036 $(0.034)$</td>
<td>$-$</td>
</tr>
<tr>
<td>Membership</td>
<td>$\gamma_1$</td>
<td>0.036 $(0.035)^b$</td>
<td>$-$</td>
<td>0.036 $(0.035)$</td>
<td>0.036 $(0.035)$</td>
</tr>
<tr>
<td>Union</td>
<td>$(1 - M) \gamma_1 + \gamma_2$</td>
<td>0.055 $(0.028)$</td>
<td>0.070 $(0.031)$</td>
<td>$-$</td>
<td>0.036 $(0.034)$</td>
</tr>
<tr>
<td><strong>Restricted traditional (3)</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Membership only</td>
<td>$\gamma_1$ ($\gamma_2 = 0$)</td>
<td>0.058 $(0.028)$</td>
<td>$-$</td>
<td>0.058 $(0.028)$</td>
<td>$-$</td>
</tr>
<tr>
<td>Coverage only</td>
<td>$\gamma_2$ ($\gamma_1 = 0$)</td>
<td>0.057 $(0.028)^{ABU}$</td>
<td>$-$</td>
<td>0.057 $(0.028)^{ABU}$</td>
<td>$-$</td>
</tr>
<tr>
<td><strong>Parameter vector varies with $M$, no coverage (4)</strong></td>
<td>$\lambda_i^j (\beta_1 - \beta_3)$</td>
<td>0.074 $(0.029)$</td>
<td>$-$</td>
<td>0.102 $(0.033)^{UW}$</td>
<td>0.045 $(0.038)^{UW}$</td>
</tr>
</tbody>
</table>

$^*$ Parameter vector varies with $M$, no coverage (4)
<table>
<thead>
<tr>
<th>Parameter vector varies with ( M ) (4)</th>
<th>( C = 0 )</th>
<th>( C = 1, M = 0 )</th>
<th>( C = 1, M = 1 )</th>
</tr>
</thead>
<tbody>
<tr>
<td>Coverage (A) (- M_i \phi (\beta_1 - \beta_0) + \theta_0)</td>
<td>0.055 (0.041)</td>
<td>0.011 (0.043)</td>
<td>0.009 (0.044)</td>
</tr>
<tr>
<td>Coverage (B) (0.038 (0.041))</td>
<td>0.010 (0.044)</td>
<td>0.009 (0.044)</td>
<td>–</td>
</tr>
<tr>
<td>Membership (x_i'(\beta_1 - \beta_0) + (\theta_1 - \theta_0))</td>
<td>0.083 (0.042)</td>
<td>–</td>
<td>0.080 (0.043)</td>
</tr>
<tr>
<td>Union (A) ((1 - M_i) x_i'(\beta_1 - \beta_0) + \theta_1)</td>
<td>0.138 (0.040)</td>
<td>0.138 (0.036)</td>
<td>–</td>
</tr>
<tr>
<td>Union (B) (0.120 (0.039))</td>
<td>0.136 (0.035)</td>
<td>–</td>
<td>0.150 (0.064)</td>
</tr>
</tbody>
</table>

<table>
<thead>
<tr>
<th>Parameter vector varies with ( C ), no membership (5)</th>
<th>( C = 0 )</th>
<th>( C = 1 )</th>
</tr>
</thead>
<tbody>
<tr>
<td>Coverage (x_i'(\delta_1 - \delta_0))</td>
<td>0.069 (0.038)</td>
<td>0.098 (0.030)</td>
</tr>
<tr>
<td>Parameter vector varies with ( C ) (5)</td>
<td>( C = 0 )</td>
<td>( C = 1, M = 0 )</td>
</tr>
<tr>
<td>Coverage (x_i'(\delta_1 - \delta_0) - M_i \phi_i)</td>
<td>0.023 (0.054)</td>
<td>0.048 (0.036)</td>
</tr>
<tr>
<td>Membership (\phi_i)</td>
<td>0.090 (0.036)</td>
<td>–</td>
</tr>
<tr>
<td>Union (x_i'(\delta_1 - \delta_0) + (\phi_i - M_i \phi_0))</td>
<td>0.114 (0.050)</td>
<td>0.138 (0.034)</td>
</tr>
</tbody>
</table>

| Parameter vector varies by all three groups (6) | \( C = 0 \) | \( C = 1, M = 0 \) | \( C = 1, M = 1 \) |
| Coverage \(x_i'(\beta_{10} - \beta_0)\) | 0.013 (0.052) | 0.026 (0.047) | 0.025 (0.052) |
| Membership \(x_i'(\beta_{11} - \beta_{10})\) | 0.121 (0.043) | – | 0.092 (0.040) |
| Union \(x_i'(\beta_{11} - \beta_0)\) | 0.107 (0.040) | 0.164 (0.036) | – |

\[ \text{BMS}^+ \]

\(^a\)Numbers in brackets refer to equation in main text.
\(^b\)ABU, B, BMS, G, LW and S are acronyms which refer to the six replication studies defined in Table 1.
\(^c\)All differentials calculated by methods A and B (see text), which coincide except where indicated. Notice that if method B uses \( \overline{M} \) instead of \( \overline{M} \) for ‘All’, the coverage and union differentials are 0.013 and 0.095, respectively.
\(^d\)Differentials for model (7) were also calculated, but the numbers were very close (small movements in third decimal place) to those for model (6).
\(^e\)The BMS+ in the final row indicates that this differential differs from BMS due to the inclusion of uncovered union members.
whereas in Table 1, on their own specification, this is estimated as 0.133. Because
the dependent variable and reference group coincide, this difference must be
attributable to the specification of the control vector: for example, Andrews et al.
(1996) do not include variables for education, firm size or industry-level member-
ship density. Our replication of Blanchflower’s model on BHPS91 also differs
from the base model considerably (total union differentials of 0.293 and 0.072,
respectively). Here the difference is more likely due to comparing manual males
with all employees. These issues are investigated further below.

3.2.2. Evaluation at different means

In the remaining columns of Table 3 we look at how the estimated differentials
from type II models differ with G, the sample over which the means are defined.
Differentials are calculated by both Methods A and B, which usually coincide
except where indicated. In model (4) without a coverage variable (the model used
by Stewart, 1991, 1995; Lanot and Walker, 1995) the membership differential that
is 0.074 when the overall means are used falls to 0.045 when the means for
members are used and rises to 0.102 when those for non-members are used.
Similarly with model (5) without a membership variable, the coverage differential
of 0.069 for the overall means is a weighted average of 0.098 (means if not
covered) and 0.048 (means if covered). In other words, the differential is lower for
the ‘more unionised’ group. Of the remaining type II models, ignoring coverage

<table>
<thead>
<tr>
<th>Replication</th>
<th>Whole sample</th>
<th>Mean used</th>
<th>Non-union members</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>Union members</td>
<td>Non-union members</td>
</tr>
<tr>
<td>Stewart: membership</td>
<td>0.020 (0.039)</td>
<td>0.070* (0.047)</td>
<td>−0.035 (0.048)</td>
</tr>
<tr>
<td>differential</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Green: membership</td>
<td>0.036* (0.037)</td>
<td>0.017 (0.042)</td>
<td>0.053 (0.051)</td>
</tr>
<tr>
<td>differential</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Lanot and Walker: membership</td>
<td>0.153 (0.026)</td>
<td>0.156* (0.029)</td>
<td>0.151* (0.027)</td>
</tr>
<tr>
<td>differential</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Blackaby et al.:</td>
<td>All covered</td>
<td>Covered members</td>
<td>Covered</td>
</tr>
<tr>
<td>(A) Membership</td>
<td>0.070 (0.058)</td>
<td>0.055 (0.066)</td>
<td>non-members</td>
</tr>
<tr>
<td>differential</td>
<td>Covered members and uncovered non-members</td>
<td>Covered Members</td>
<td>Uncovered</td>
</tr>
<tr>
<td>(B) Total union</td>
<td>0.150 (0.043)</td>
<td>0.089 (0.065)</td>
<td>0.222* (0.038)</td>
</tr>
<tr>
<td>differential</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

*Mean used in original study.
differentials which are usually insignificant, the same effect occurs about half the time. If one takes the view that the appropriate way to model type II models is the three-way split (model 6) then the conclusion is that which means are used matters a lot—the total union differential is three times higher when means for uncovered workers are used than when means for covered members are used—but does not vary in any obvious systematic way.

In Table 4 we return to the original replications of the type II models, but look at the differentials based on use of various means. There is considerable variation both in the membership differentials for a given model at different means and for a given mean across models. Which mean is used clearly matters, but again the direction and extent of the difference in differentials depends on which type of model is being used.

3.2.3. Dependent variable and sample effects

In this subsection we consider deviations in the dependent variable (w) and sample definitions (S) in the directions of the six studies while retaining other aspects of our base model. These are considered one at a time, and are listed in Table 5 along with the effects on the estimates of $\alpha_1$ and $\alpha_2$.

When the definition of the dependent variable changes from weekly earnings to hourly earnings, the coverage and total differentials both rise. This is expected (as discussed in Section 2) since the weekly earnings differential can be decomposed into differentials in hourly earnings and hours, and there is considerable evidence

<table>
<thead>
<tr>
<th>Deviation</th>
<th>$\alpha_2$</th>
<th>$\alpha_1$</th>
<th>N</th>
<th>$R^2$</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Base model</strong></td>
<td>0.020 (0.038)</td>
<td>0.076 (0.031)</td>
<td>906</td>
<td>0.484</td>
</tr>
<tr>
<td><strong>Changes to dependent variable</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Log hourly earnings (LW)</td>
<td>0.057 (0.036)</td>
<td>0.104 (0.030)</td>
<td>898</td>
<td>0.488</td>
</tr>
<tr>
<td>Log annual earnings (B)</td>
<td>-0.066 (0.050)</td>
<td>0.049 (0.041)</td>
<td>823</td>
<td>0.541</td>
</tr>
<tr>
<td><strong>Changes to sample</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Manufacturing sector only (S)</td>
<td>0.015 (0.050)</td>
<td>0.072 (0.039)</td>
<td>423</td>
<td>0.458</td>
</tr>
<tr>
<td>Manual females added (G)</td>
<td>0.051 (0.035)</td>
<td>0.095 (0.028)</td>
<td>1100</td>
<td>0.511</td>
</tr>
<tr>
<td>Uncovered members excluded (BMS)</td>
<td>0.021 (0.037)</td>
<td>0.077 (0.031)</td>
<td>882</td>
<td>0.490</td>
</tr>
<tr>
<td>Agricultural workers excluded (BMS)</td>
<td>0.007 (0.039)</td>
<td>0.076 (0.031)</td>
<td>883</td>
<td>0.479</td>
</tr>
<tr>
<td>Non-manual males added (B)</td>
<td>-0.042 (0.026)</td>
<td>0.022 (0.023)</td>
<td>1874</td>
<td>0.553</td>
</tr>
<tr>
<td>Non-manual males and females added (B)</td>
<td>-0.019 (0.022)</td>
<td>0.075 (0.019)</td>
<td>3142</td>
<td>0.530</td>
</tr>
<tr>
<td>Part-time males added (B)</td>
<td>0.017 (0.038)</td>
<td>0.082 (0.032)</td>
<td>926</td>
<td>0.527</td>
</tr>
<tr>
<td>Married males only (LW)</td>
<td>-0.018 (0.043)</td>
<td>0.057 (0.032)</td>
<td>654</td>
<td>0.294</td>
</tr>
<tr>
<td>Part-time males added, no dummy (LW)</td>
<td>0.029 (0.041)</td>
<td>0.119 (0.034)</td>
<td>926</td>
<td>0.459</td>
</tr>
<tr>
<td>Age restricted to 21–65 (LW)</td>
<td>-0.005 (0.037)</td>
<td>0.078 (0.030)</td>
<td>811</td>
<td>0.304</td>
</tr>
</tbody>
</table>

The acronyms B, BMS, G, LW and S refer to the original models (see Table 1).
that union workers work shorter hours than their non-union counterparts. This partly explains why the differential from replicating the Lanot and Walker model on the BHPS (17%) is one of the higher ones. When the dependent variable is specified in terms of annual earnings (as used by Blanchflower, 1991), the total differential falls (contrary to the prediction in Section 2) and the coverage differential is negative.\footnote{Lewis (1986) estimates for the U.S. that the differential in weekly earnings is lower than that in hourly earnings by 0.018 on average and higher than that in annual earnings by 0.012 on average.}

In the other deviations in Table 5, it is the sample definition which changes. The total differential falls dramatically when non-manuals are added (as in the work of Blanchflower, 1991): from 8% down to 2%.\footnote{For the U.S., Lewis (1986) finds white-collar union differentials to be smaller than those for blue-collar workers by about 10 percentage points.} When a non-manual dummy and interactions with the two union variables are included, the estimated membership, coverage and total differentials for non-manuals are all negative. Addition of female workers (Green) or male part-timers (particularly if no dummy is included to control for them) (Lanot and Walker) raises the total union differential: in the last case to 13%.\footnote{Reviewing the U.S. evidence, Lewis (1986) concludes that “(1) the sign of the mean male-minus-female wage gap difference is ambiguous, and (2) the numerical magnitude of the difference is close to zero.” (p. 118).}

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3.2.4. Control vector effects

We look finally at the impact of changes in the control vector. The potential modifications that one could examine run into the millions. To keep things within bounds we group potential control factors into six groups (see the rows of Table 6). We then look, for each of the six groups of variables and for each of the six studies, at the effect of replacing the variables in the base model control vector in that group by those in the same group used in the particular study. The resultant estimates are presented in Table 6. To illustrate, if the group of firm/industry variables used in the base model (which are industry-level membership density, a public sector dummy, and firm-size dummies) are replaced by the industry dummies used by Lanot and Walker, the estimate of $\alpha_2$ rises from 0.020 to 0.034 and the estimate of $\alpha_1 - \alpha_2$ rises from 0.056 to 0.080. Across the variations considered, $\alpha_2$ varies between zero and 0.047 (compared to 0.020 in the base model) and $\alpha_1 - \alpha_2$ varies between 0.044 and 0.083 (compared to 0.056 in the base model). The three instances of 0.080 and above result from the adding of industry dummies by Blanchflower, Lanot and Walker and Andrews et al. However, none of the changes is particularly dramatic. Interestingly, when the education variables are taken out of the base model (Andrews et al., row 1) the differentials change very little.
Table 6: Effects of changes to the control vector

<table>
<thead>
<tr>
<th>Education subgroups</th>
<th>$a_1 - a_2$</th>
<th>$a_1 - a_3$</th>
<th>$a_1 - a_4$</th>
<th>$a_1 - a_5$</th>
<th>$a_1 - a_6$</th>
<th>$a_1 - a_7$</th>
<th>$a_1 - a_8$</th>
<th>$a_1 - a_9$</th>
<th>$a_1 - a_{10}$</th>
<th>$a_1 - a_{11}$</th>
<th>$a_1 - a_{12}$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Education</td>
<td>0.055</td>
<td>0.053</td>
<td>0.052</td>
<td>0.050</td>
<td>0.048</td>
<td>0.046</td>
<td>0.045</td>
<td>0.044</td>
<td>0.043</td>
<td>0.042</td>
<td>0.041</td>
</tr>
<tr>
<td>Experience/other</td>
<td>0.047</td>
<td>0.046</td>
<td>0.045</td>
<td>0.044</td>
<td>0.043</td>
<td>0.042</td>
<td>0.042</td>
<td>0.041</td>
<td>0.040</td>
<td>0.039</td>
<td>0.038</td>
</tr>
<tr>
<td>Regional personal</td>
<td>0.038</td>
<td>0.037</td>
<td>0.036</td>
<td>0.035</td>
<td>0.034</td>
<td>0.033</td>
<td>0.033</td>
<td>0.032</td>
<td>0.031</td>
<td>0.030</td>
<td>0.029</td>
</tr>
<tr>
<td>Industry</td>
<td>0.037</td>
<td>0.036</td>
<td>0.035</td>
<td>0.034</td>
<td>0.033</td>
<td>0.032</td>
<td>0.032</td>
<td>0.031</td>
<td>0.030</td>
<td>0.029</td>
<td>0.028</td>
</tr>
<tr>
<td>Other</td>
<td>0.035</td>
<td>0.034</td>
<td>0.032</td>
<td>0.031</td>
<td>0.030</td>
<td>0.029</td>
<td>0.028</td>
<td>0.027</td>
<td>0.026</td>
<td>0.025</td>
<td>0.024</td>
</tr>
</tbody>
</table>

Using the base sample, the base model's $x$-vector is split into sub-groups and then singularly varied toward each of the replicated models (see text). Base model estimates for comparison: 0.056, 0.039, 0.020, 0.038.
Another set of experiments conducted involves dropping variables from the base model (singly or in groups as appropriate). As expected from the discussion in Section 2, differentials rise if firm size is not included among the controls. The coverage differential rises to 0.053, the membership differential to 0.062 and the total union differential to 0.115. Excluding industry union membership density without replacing it by industry dummies (which were included in the modification discussed earlier in the section) increases the membership differential to 0.096, but does not affect the coverage differential. These are the two main instances where the estimated differentials are sensitive to the dropping of elements of the control vector. Exclusion of years of experience raises the coverage differential slightly, but less than these other two changes and exclusion of any of the other variables has only even more minor effects.

4. Conclusions

The main finding of the paper is that methodological differences can matter a lot. The ballpark 10% figure often quoted may be somewhat of an illusion, with variations in methodology and data tending to offset each other. The estimated union wage differentials produced vary considerably with various aspects of the specification adopted. Some specific findings are as follows.

1. There is an important difference between differentials with respect to union membership and differentials with respect to union presence or coverage at the workplace. Except in specifications without a membership variable, coverage effects are typically small.

2. Even among studies with the same sample selected, the same dependent variable, the same controls and both membership and coverage variables included, there is considerable variation in the estimated membership differential (from 4% to 13%) with the model specification used.

3. It matters to some degree whether we look at differentials in hourly, weekly or annual earnings. Hourly earnings differentials are bigger than those in weekly earnings, which in turn are bigger than those in annual earnings.

4. As regards sample definition, differentials are lower when non-manuals are added and the inclusion of female workers seems to raise estimated differentials.

5. It is important to control for firm size. However the inclusion or exclusion of most of the other control variables typically used does not seem to be very important.

This is a much larger effect on the differentials than Lewis (1986) finds for the U.S. He finds a relatively modest effect of excluding firm size controls and does not include it in the dozen or so main adjustment factors for omissions that he calculates. He does, however, find that differentials decline as establishment or firm size increases.
(6) The choice of which group means to use when evaluating the mean differential in multi-equation models is found to be of considerable importance. Even with all else held constant the differentials vary with the group chosen to average over and not always in the same direction. This difference interacts with that resulting from the form of model chosen.

In sum, methodological differences can cause considerable differences in estimated union/non-union wage differentials. Considerable care should therefore be taken when making comparisons across studies which differ in the methodological aspects discussed in this paper.

Acknowledgements

Mark Stewart and Joanna Swaffield thank the ESRC (under grant number R000234448) for financial assistance, and Martyn Andrews and Richard Upward thank the Faculty of Economic and Social Studies, University of Manchester for financial assistance. The data used in this paper were made available through the ESRC archive. The data were originally collected by the ESRC Research Centre on Micro-Social Change at the University of Essex. Neither the original collectors of the data nor the Archive bear any responsibility for the analyses or interpretations presented here. We are indebted to David Blackaby, David Blanchflower, Francis Green, Gauthier Lanot and Ian Walker for answering all our questions about their own papers so quickly and to Robin Bladen-Hovell for his comments at a preliminary stage of the work. Comments from presentations at Manchester University, the University of North Wales at Bangor and the EMRU Labour Economics Conference in Dundee are gratefully acknowledged.

References


