Estimating the Impact of the Minimum Wage Using Geographical Wage Variation *

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Abstract

This paper evaluates the impact on employment of the UK's introduction of a minimum wage in 1999 by exploiting the geographical variation in wages, which meant that the minimum wage's "bite" into an area's wage distribution differed considerably across the country. The results indicate that, although the minimum wage had differential wage-distribution effects across the 140 areas of the country, employment growth after its introduction was not significantly lower in areas of the country with a high proportion of low-wage workers, whose wages had to be raised to comply, from that in areas with a low proportion of such workers.

Keywords: Minimum wages, employment determination, geographical wage variation, difference-in-differences estimators.

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1. Introduction

There is considerable regional variation in wages across the UK. One consequence of this variation is that the extent to which the introduction of the National Minimum Wage impinged on an area's wage distribution also differed considerably across the country. In some areas (for example parts of the South East) it had relatively little impact on wages, since relatively few employees were being paid below the minimum wage prior to its introduction. In other areas many more employees needed their wages raising to comply with the newly introduced minimum. The National Minimum Wage reaches far further up the wage distribution in Northumberland or Lincolnshire, say, than in Surrey or Oxfordshire. This geographical variation can be exploited to assess the impact of the introduction of a minimum wage on the employment prospects of those whose wages were, or would have been, affected. This paper explores this line of reasoning for the 1999 minimum wage introduction in the UK.

On the basis of the standard textbook model of the labour market we would expect to see a relative decline in employment in low-wage areas where the minimum wage bit more deeply compared to higher-wage areas where relatively few employees' wages were affected, other things equal. This paper tests this prediction in a variety of ways using wage data from the New Earnings Survey and employment information from a variety of sources.

The approach being taken views the introduction of the minimum wage as a "quasi-experiment" with the "treatment effects" varying across local areas (Card, 1992a, Card and Krueger, 1995, chapter 4). The best known use of the "quasi-experiment" approach in the context of minimum wages is probably the evaluation by Card and Krueger (1994) of the impact of the 1992 New Jersey minimum wage increase on employment in fast-food restaurants. They use a sample of similar restaurants in neighbouring locations in eastern Pennsylvania, where the minimum wage rise did not apply, as a control group. This comparison exploits the US inter-state differences in minimum wage rates, and their increases, and does not translate to a country with a national minimum. However the "bite" of the national minimum varies across geographical areas and this can be used to evaluate its impact. As Card (1992a) puts it, "from an evaluation perspective, a uniform minimum wage is an under-appreciated asset."

The general "quasi-experiment" approach can be used with treatment and control groups defined in many ways. The idea at the core of the approach as applied here is an intuitively obvious one: other things being equal the largest effects of the introduction of the minimum wage on employment will be found where it has its largest effects on wages. Thus the impact can be estimated by comparing changes in employment rates between groups with high and low incidence of low-wage workers. Dolado et al. (1995) label it the "differential impact" approach.

The groups can be defined in various ways. The comparison can be between low-wage and higher-wage individual employees (Linnerman, 1982, and Abowd et al., 2000, for the US, Stewart, 2002, for the UK); or between low-wage and higher-wage companies (Card and Krueger, 1994), or as in this paper between low-wage and higher-wage geographical areas (Card, 1992a, and Deere et al., 1995, for the US, Dolado et al., 1995, for France and Spain).[1]

The employment effects of minimum wages remains one of the most contentious policy questions in economics and has generated a vast literature. (A useful review is provided by Brown, 1999). Research over the last decade has removed the earlier consensus on the subject. Much of this work has found employment effects to be absent or positive (e.g. Card and Krueger (1994, 1995, 2000), Abowd et al. (2000) for the US, Machin and Manning (1994), Dickens et al. (1999), Stewart (2002) for the UK). In contrast there are also many recent studies that find significant negative effects (e.g. Burkhauser et al. (2000), Neumark and Wascher (2000), Neumark et al. (2000) for the US, Machin et al. (2002) for the UK, Abowd et al. (2000) for France). The issue remains open for debate.

A national minimum wage was introduced in the UK on 1 April 1999, after a period of some years without wage floors (following the final abolition of the wages councils

[1] Other partitions have also been utilised in the literature, such as by industry (Dickens et al., 1999, and Machin and Manning, 1994, for the UK), by occupation (Dolado et al., 1995, for the Netherlands), by education, age and other demographic factors (Deere et al., 1995, for the US).

in 1993). The adult rate was set at £3.60 per hour, with a lower youth rate of £3.00 per hour for those aged 18-21 inclusive and a development rate of £3.20 per hour for adults in the first 6 months of a new job with accredited training. The youth rate subsequently rose in June 2000 and the adult rate in October 2000.

A number of different empirical strategies are pursued in this paper within this general approach of exploiting Britain's geographic wage variation to estimate the impact of the minimum wage's introduction on employment. Two main estimators are used. The first estimates the regression relationship between the change in the employment rate and the fraction of workers in an area initially below the minimum wage. The second is a difference-in-differences estimator based on comparing groups of low-wage and high-wage areas of the country. These two estimators are also both considered for individual-level data and for data aggregated to the area level. Sections 3 presents cross-sectional estimates of the impact on employment rates. Section 4 then uses panel data to incorporate a comparison with the earlier period without a minimum. Prior to that, in Section 2, the corresponding impact on the wage distribution is examined. For the methodology to be effective requires that there are significantly different wage movements in the two groups of local areas and a significant relationship with the fraction of workers in the area initially below the incoming minimum. Section 5 focuses the analysis more closely on those most likely to be affected within each area with a view to strengthening the test conducted. Section 6 provides conclusions.

2. Geographical wage variation

The cross-sectional area-level analysis conducted for this paper is based on the administrative geography of Great Britain as at April 1998 (i.e. after the local government reorganisation that took place between April 1995 and April 1998). The analysis is conducted at "local authority" level. These are defined to comprise the counties, Unitary Authorities and Metropolitan Councils in England and Wales and the Unitary Councils in Scotland.[2] The Greater London boroughs are aggregated into Inner London and Outer London. The resulting geography divides Great Britain into 140 "local areas".[3]

In the April 1998 New Earnings Survey, less than 1% of adult employees were paid below £3.60 per hour (the rate at which the minimum wage was subsequently introduced) in the Unitary Authorities of Bracknell Forest and West Berkshire (Newbury).[4] In contrast 12% of adult employees were paid below this level in both Redcar & Cleveland and Torbay. There is considerable variation across areas in the minimum wage's expected "bite" at the bottom of the wage distribution. Before using this geographic variation to estimate the minimum wage's impact on employment, it is necessary to demonstrate that there was indeed a differential impact of the minimum wage on the wage distribution across areas, which would then in turn lead us to expect differential employment growth.

Figure 1 illustrates the relationship between the 1998-2000 change in different aspects of an area's (log-)wage distribution and the proportion of employees in that area who were paid below £3.60/hour for their basic wage in 1998. The figure shows smoothed (locally weighted) plots of these relationships. As we move from areas with a low proportion of employees paid below this level to areas with a high proportion of such employees, both the average wage of those at the bottom (defined as below £5 per hour) and the 5th. percentile of the distribution change sharply, while points further up

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[2] In England the local government reorganisation was implemented in four phases in April of each year from 1995 to 1998, with Unitary Authorities replacing the old two-tier system of County Councils and Local Authority District Councils. The counties and districts of Wales were replaced by Unitary Authorities in April 1996, and the Local Authority regions and districts of Scotland were replaced by Unitary Councils, also in April 1996. The reorganisation complicates comparisons across time prior to 1998, as will be seen below.

[3] Northern Ireland, while covered by the UK National Minimum Wage, is excluded from the analysis in this paper. In addition, the Scottish Unitary Councils of Eilean Siar (Western Isles), Orkney and Shetland are combined to enable the definition used here to be constructed from the variable "uacnty" in the LFS Local Area Data.

[4] Hourly wage data from the New Earnings Survey is used throughout this paper. The NES's great strengths are the likely accuracy of the wage data it provides, much of it direct from computerised payroll records, and the very large sample sizes. However since the sample is based on National Insurance numbers and taken from Inland Revenue PAYE records, it excludes most of those whose weekly earnings falls below the PAYE deduction threshold (particularly in small organisations). This threshold was £83.37/week at the time of the 1999 survey, meaning that someone on the minimum wage of £3.60/hour would fall below if they worked 23 hours/week or less.

the distribution (specifically the lower quartile and the median) show relatively little change.

Regression estimates of the relationship between changes in various features of the wage distribution of an area over this period and the proportion of employees in 1998 paid below the incoming minimum wage rates are given in Table 1. The proportion "low paid" in an area clearly significantly influenced the change in the bottom end of the area's wage distribution, but had only a relatively minor effect on the centre of the distribution.

A useful alternative way of evaluating the impact of these geographical differences is to define a group of local areas where the minimum wage is expected to have a high impact (due to their having a relatively high proportion of employees paid below the minimum wage rates prior to its introduction) and another group of areas where it is expected to have a low impact (due to having relatively few such employees). Comparison of these two groups will be one of the methods used below to evaluate the impact of the minimum wage on employment. So it is useful here to also examine the effect on the wage distribution for the two groups.

The two groups are constructed in the following way. The 140 "local areas" being used in the paper are classified according to the proportion of employees who were "low paid" in April 1998 (defined by the 1999 age-specific minimum wage rates: £3.60/hour for adults, £3.00/hour for youths). The "high impact" group is defined to be those local areas with the highest low-pay rates according to this definition and which between them constitute 10% of the employees in the 1998 NES. This requires 33 local areas. They are listed in the appendix. All have low-pay rates of 7.3% or above and overall 8.6% of employees in this group are low paid.

The "low impact" group is defined in an analogous fashion using the 10% of employees in the local areas with the lowest low-pay rates. This only requires 6 local areas, because one of them is Inner London, which is far larger in employment terms than the other areas. They are listed in the appendix also. All have low-pay rates of 2% or below and overall 1.9% of employees in this group are low paid. Some comparative statistics on the two groups are also given in the appendix. Kernel estimates of the wage densities for those in the two groups, based on the 1998 NES, are shown in Figure 2.[5] The individuals in the two groups clearly have very different wage distributions.

The two groups also exhibit significantly different changes at the bottom end of their wage distributions over the period 1998-2000. Figure 3 shows the shift in the kernel wage densities for the two groups. The "high impact" group shows a more marked shift. Table 2 presents difference-in-differences estimates for the impact of the minimum wage's introduction on various features of the wage distribution, treating the "high impact" group as the treatment group and the "low impact" group as the comparison group. Results from analyses at both the (unweighted) area level and the individual employee level (with suitable adjustment of the standard errors for the "clustering") are presented and paint a similar picture.[6] On the basis of this method of analysis the introduction of the minimum wage had a significant influence on the bottom end of the wage distribution, but relatively less impact on the centre of the distribution.

3. Cross-sectional estimation of the impact on employment rates

The first type of method for estimating the impact of the introduction of the minimum wage on employment considered is cross-sectional and based on the approach used by Card (1992a). The approach models the change in employment in a period straddling the introduction of the minimum. Section 4 below extends the Card approach by using panel data to construct a comparison with one or more earlier periods in which there was no minimum.

The impact of the introduction of the minimum wage on employment is estimated first by two similar methods, equivalent to those used on wages in the previous section. The first method estimates equations of the form

[5] The Kernel density estimates are constructed using the Epanechnikov kernel and an "optimal" kernel bandwidth. The "high impact" and "low impact" densities are based on 12,960 and 12,739 employees respectively.

(1) $$\Delta e_j = \alpha_0 + \beta p_j^{98} + \varepsilon_j$$

where e_j is the employment rate in area *j*, Δe_j is the change in this rate between a suitable point in late 1998 (prior to the introduction of the minimum wage) and the corresponding date in 1999 (after the introduction), and p_j^{98} is the proportion of those in area *j* who are low paid in April 1998 (with low pay being defined as below the appropriate age-specific minimum wage rate as subsequently introduced in April 1999).

The line of argument in the introduction suggests this as a suitable equation to estimate and test the hypothesis that $\beta < 0$. Card (1992a) interprets the equivalent of equation (1) as a reduced form equation from a very simple structural model that explains the change in the employment rate in area *j* as a movement along a labour demand schedule:

(2) $$\Delta e_j = \alpha_1 + \eta \Delta w_j + u_{1j}$$

and the wage increase in area *j* as a function of the proportion in the area who were "low paid" in 1998:

(3) $$\Delta w_j = \alpha_2 + \lambda p_j^{98} + u_{2j}$$

Then $\beta = \eta\lambda$, with λ assumed to be positive, implying that β has the same sign as η. All three equations can also incorporate additional regressors.

The proportion initially low paid is viewed in this system as a predetermined instrument for the endogenous wage change. The evidence of Section 2 suggests that if w_j is taken to be the overall average of the (log) wage in area *j*, the estimate of λ would be small and insignificantly different from zero. However if it is taken to be

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[6] The individual-level standard errors need to allow for possible correlation between the error terms within a local area (see Moulton, 1986). The t-ratios presented use standard errors which allow for this

the average across low paid workers (defined in some appropriate way), the estimate of λ can be large and highly significant.[7] Note however that estimation of equation (1) does not require specification of exactly which wage measure it is that appears in equations (2) and (3). Indeed equation (2) may contain multiple wage measures, each of which is related to p_j^{98} in an equivalent of equation (3), still producing equation (1) as the reduced form.

The second initial estimator used is a difference-in-differences estimator, specifically the difference between 1999 and 1998 in the difference in the employment probability between the "high impact" areas group and the "low impact" areas group.[8] This can be viewed as the OLS estimator for an equation in which the proportion initially low paid, p_j^{98}, is replaced in equation (1) by an indicator variable for the "high impact" group:

(4) $$\Delta e_j = \alpha_3 + \theta d_j + \varepsilon_j^*$$

where $d_j = 1$ if area *j* is in the "high impact" group and $d_j = 0$ if in the "low impact" group. The OLS estimator of θ in equation (4) is $\hat{\theta} = \overline{\Delta e}^1 - \overline{\Delta e}^0$, where $\overline{\Delta e}^k$ is the average for group *k*. It is also useful for future reference to note that the estimate of θ can also be calculated from a regression, using two years of pooled data, of the employment rate on dummy variables for the high impact group, for the post minimum wage period and for the interaction between the two:

(5) $$e_{jt} = \gamma_0 + \gamma_1 y_t + \gamma_2 d_j + \theta y_t d_j + \varepsilon_{jt}^*$$

where $y_t = 1$ if *t* = 1999 and 0 if *t* = 1998. This will give the same estimate of θ. Again control variables can also be incorporated into equations (4) and (5). For the appropriate standard errors the variance of ε^* needs to be allowed to vary between years and groups.

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by a method which is a straightforward extension of that proposed by Huber (1967).

[7] This can be further increased if one focuses on more narrowly defined groups who have high proportions of low paid workers. This approach is taken in Section 5 below.

[8] See Meyer (1995) and Angrist and Krueger (1999) for discussion of this type of estimator.

Analogously to this, note that the same estimate of β in equation (1) would result from OLS estimation of an equation of the form

(6) $$e_{jt} = \gamma_0^* + \gamma_1^* y_t + \gamma_2^* p_j^{98} + \beta y_t p_j^{98} + \varepsilon_{jt}$$

A number of different data sources are used on both methods for the construction of employment rates. The first estimates presented use employment rates constructed from the Labour Force Survey Local Area Data files for September to November 1998 and the same months in 1999. These data provide the required area codes at the sub-regional level (whereas the main LFS release only identifies broad regions), but exclude income variables and band other variables (including age) to prevent identification of individuals with unusual combinations of characteristics.

Estimates based on these LFS employment rates are given in the first two rows of Table 3. The first employment rate considered is for those aged between 18 and the state retirement age (60 for women, 65 for men). The proportion low paid is constructed for the same age group from the 1998 NES. The estimate of β is -0.16 with a robust absolute t-ratio of 0.84. The hypothesis of no impact on employment rates cannot be rejected at any credible significance level.

This is illustrated in Figure 4, which gives a scatter plot of the data for the 140 local areas. There is not much evidence of a relationship between the change in the employment rate from autumn 1998 to autumn 1999 and the proportion in the area who were low paid in April 1998.

The difference-in-differences estimate, as described above and based on the comparison of the “high impact” and “low impact” groups of local areas defined in the appendix and used in Section 2 above, is given in the second column of Table 3 and is also insignificantly different from zero.

It is often suggested that young workers are one of the most vulnerable groups to any adverse employment effects of the minimum wage. However there is not much evidence to support this in this data. Restricting both employment and low pay rates to young workers gives an estimate of β that is insignificantly different from zero (and positive).[9] The difference-in-differences estimate is negative but also insignificantly different from zero.

One worry with these estimates is that the employment rates are based on area of residence while the proportion low paid is based on area of workplace. Might this be masking any relationship? An alternative to using the individual-based Labour Force Surveys is to use employer-based employment figures. The prime source of employment information for most purposes is the Annual Business Inquiry (ABI). Data from the ABI is available from 1998 onwards, with a survey date of December of each year.[10]

The total number of employees in employment was calculated from the ABI for the 140 local areas described above. These are then converted to employment rates (or more accurately employment-population ratios) by dividing by the ONS estimates of the population of working age in each local area.[11] However this is somewhat problematic and there's a trade-off here. The ABI provides the best available estimates of employment in a local area. However there's a well-known mismatch problem in the calculation of employment rates in this way, since population figures are residence-based, while ABI employment figures are workplace-based.

This problem is more serious for some areas than for others, being particularly serious for those areas where there is extensive in-commuting from surrounding areas. For some areas the calculated employment-population ratios can exceed 1. Using ABI employment and ONS population estimates the ratio exceeds 1 in 1998 or 1999 or both for 4 of the 140 areas: Inner London, Slough, Reading and Nottingham. These 4 areas are excluded from the analysis presented here for this construction of the employment rate.

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[9] The banding of the LFS Local Area Data age variable means that those aged 18-24 are used.

[10] The ABI replaced the Annual Employment Survey (AES) (conducted in September of each year). The AES methodology is now regarded as flawed and the AES data as giving a serious undercount. For 1998 the AES estimate of total employee jobs was around 900,000 lower than the ABI estimate (Partington, 2001).

Regression estimates for the change in the employment rate between December 1998 and December 1999 calculated in this way (on the sample of 136 local areas) are presented in the next row of Table 3. The estimated slope coefficient is similar to that for the equation using LFS-based rates, although slightly smaller. In particular it is again insignificantly different from zero.[12] The corresponding difference-in-differences estimate is also similar to that based on LFS-based rates and is insignificantly different from zero.[13]

A third possibility considered is to base the estimated employment rates on the same source as used for the calculation of the proportion in the local area below the minimum wage threshold, the New Earnings Survey. This matching of the sources for dependent and explanatory variables is obviously advantageous. However the NES has disadvantages for the calculation of employment rates. As pointed out above, it under-samples part-timers. In addition, other exclusions result from the fact that there is a lag (of about a month) between the drawing of the sample and the reference week in which the survey is conducted. This means that those not employed when the sample is drawn who enter employment before the survey date and those who move employer between the two dates and cannot be traced are excluded. Despite these drawbacks, it provides a useful alternative source of employment by local area. The NES employment numbers (multiplied by 100 to reflect the 1% sampling) are then converted to employment rates using the same ONS estimates of the population of working age in the local area as used above with the ABI.

Estimates for the equation for the change in the employment rate constructed in this way are given in the next row of Table 3. As with LFS and ABI-based rates, the estimated slope coefficient is insignificantly different from zero (and slightly closer to zero than both the others). The same comments also apply to the corresponding difference-in-differences estimate.

[11] See for example ONS (2001).

[12] The alternative of estimating an equation for the rate of growth of employment (from the ABI) with the rate of growth of working population in the area (from the ONS estimates) as an additional regressor gives a very similar (and insignificant) coefficient on the proportion low paid in 1998 and a coefficient on the population growth variable that is insignificantly different from 1.

[13] Age disaggregation of the ABI employment figures, to enable consideration of youths only, is not available.

Another possible variable construction is to focus on those initially (i.e. before the introduction of the minimum wage) in employment and use the matching of individuals across years in the NESPD to construct employment "continuation rates". More specifically for each local area, the constructed variable is the proportion of those in employment in April 1998 who are also in employment in April 2000. (An additional advantage of this construction is that it produces an employment "rate" without the need for matching employment and population figures.) The next rows of Table 3 give estimates based on this construction. Both the coefficient on the proportion in the area initially low paid and the difference-in-differences estimator are positive and significantly different from zero. The estimates for youths are also positive, but are insignificant. Again there is no evidence of the introduction of the minimum wage reducing employment.

4. Comparison with the period without a minimum using panel data

The above analysis looks at the cross-sectional relationship between the change in the employment rate over a period straddling the introduction of the minimum wage and the pre-minimum wage proportion low paid in the area. Although the equation estimated can be specified in terms of panel data (for two dates) on the employment rate, as in equation (6), in terms of the change in the employment rate, the relationship examined is cross-sectional.

There is an implicit assumption that there would be no relationship between them in the absence of the minimum wage, so that any relationship found for the straddling period can be attributed to the minimum wage. This is a strong assumption. One way to address the issue is to compare the estimates for the 1998-2000 period with equivalent ones for an earlier period that did not straddle the introduction of a minimum wage, but rather was a period when there was no wage floor in place throughout.

The period used initially for this comparison is 1996-1998. The effect of the introduction of the minimum wage can then be estimated as the difference between the slope coefficient for the period 1998-2000, which straddles the introduction, and the equivalent coefficient for the period 1996-1998, which is entirely before the introduction. In the case where d_j is used in place of p_j this is equivalent to a difference-in-differences estimator for the change in the employment rate (the equivalent of equation (5) for $\Delta_2 e_{jt}$ instead of e_{jt}). When the variable p_j is used, the estimator can be produced by pooling the two cross-sections and estimating the equation

$$(7) \qquad \Delta_2 e_{jt} = \mu_0 + \mu_1 p_{j(t-2)} + \mu_2 y_t + \beta p_{j(t-2)} y_t + \varepsilon_{jt}$$

for t = 1998, 2000, where p_{jt} is the proportion low paid in local area j in year t and y_t is the time period dummy with $y_t = 1$ if t = 2000. The estimate of β then measures the impact of the minimum wage. Estimation of equation (7) by OLS is equivalent to estimation of two separate equations of the form of equation (1) for $y_t = 1$ and $y_t = 0$, i.e. for t = 2000 and t = 1998, with the estimate of β then calculated by subtraction.

The geography used above in Section 3 cannot be used for this estimation, due to changes in coding reflecting the local government reorganisation of the mid-1990s. The administrative geography of Britain underwent extensive changes over the period April 1995 to April 1998 as part of the reorganisation of local government. The area coding included on the NES also changed to reflect this. This requires the use of a slightly more aggregated geography than used above – to 97 local areas – to give a consistent definition of areas across the years. This more aggregated version corresponds to the coding used on the NES up to 1996. The different codings for 1997 and for 1998 onwards are mapped into the old geography.

For the 97 local areas in this panel-consistent geography, estimation of equation (7) gives an estimate of β of 0.177 with a robust t-ratio of 0.66 (Table 3). Thus the differencing (or equivalently, allowance for existence of a relationship in the absence of a minimum wage) produces an estimate of the effect of the minimum wage introduction that is positive, but insignificantly different from zero.

The corresponding difference-in-differences estimation requires a re-definition of the "high impact" and "low impact" groups in terms of this panel-consistent geography. The "treatment" group is constructed on this geography in a similar way to that used above. It contains areas with a high proportion of employees who are "low paid" in April 1998 (defined by the 1999 age-specific minimum wage rates (£3.60/hour for adults, £3.00/hour for youths)). The 13 areas comprise the top 10% of the employees in the 1998 NES in terms of these "low pay" rates.[14] The "comparison" group used contains the local areas in the matching bottom 10% of the NES 1998 sample in terms of "low pay" rates.[15]

The corresponding difference-in-differences estimate is given in the second column of Table 3 and has an even lower t-ratio than that when the proportion low paid is used. Restricting attention to youths raises both the estimate of β and its t-ratio (the latter to 1.14), but both it and the difference-in-differences estimate are still insignificantly different from zero.

The final part of the analysis for this section extends the approach and strengthens the test of the impact of the minimum wage in a number of ways. First it uses the entire panel from 1994 onwards (1994 being chosen so that the entire pre-minimum wage period is a period without wage floors, since the wages councils were abolished in 1993), rather than considering just one inter-temporal comparison of Δe. Second, it moves the analysis to the level of the individual employee. Third, it incorporates individual-level control variables for relevant characteristics. Fourth, it considers conditional employment probabilities for one year ahead. For the difference-in-differences estimator, the following model is estimated:

$$(8) \quad \Pr[E_{it+1}=1 \mid E_{it}=1] = \Lambda\{x_{it}'\beta + \pi_1 d_{it}^{(1)} + \pi_2 d_{it}^{(2)} + \pi_3 m_{t+1} + \theta^* d_{it}^{(1)} m_{t+1} + \phi d_{it}^{(2)} m_{t+1} + \delta_t\}$$

[14] The local areas in the group are: Northumberland, Cleveland, Dyfed, Humberside, Cornwall, Gwynedd, Dumfries, Lincolnshire, Hereford & Worcester, Norfolk, Devon, Isle of Wight and Durham (listed in low-pay rate order, highest first). All have low pay rates of 6% and above and the overall low-pay rate for the group is 7.4%.

[15] The group contains 16 local areas, consisting of the counties of Buckinghamshire, Oxfordshire and Surrey and 13 of the London boroughs. All have low pay rates of 1.7% and below and the overall low-pay rate is 1.5%.

where $d_{it}^{(1)} = 1$ if individual i was working in a "high impact" local area, $d_{it}^{(2)} = 1$ if working in neither a "high impact" nor a "low impact" local area, m_{t+1} is a binary indicator taking the value 1 if the new minimum wage was in place at time $t+1$, δ_t are time effects for the remaining time periods and Λ is the logit transformation. Thus group 1 (the "high impact" group) acts as the "treatment" group and the base group (the "low impact" group) acts as the "comparison" group. The parameter θ^* leads to the difference-in-differences estimator of θ after the usual scaling to take account of the non-linear functional form. This model is similar to that used in Stewart (2002), where the difference-in-differences estimator used is based on "treatment" and "control" groups defined in terms of segments of the individual-level wage distribution.

When the evaluation is done using the continuous proportion low paid variable in place of the dummies for the two groups of local areas, the equation takes the form:

$$\text{(9)} \qquad \Pr[E_{it+1} = 1 \mid E_{it} = 1] = \Lambda\{x_{it}'\beta + \pi_3 m_{t+1} + \pi_4 p_{it} + \beta^* p_{it} m_{t+1} + \delta_t\}$$

where p_{it} = the proportion in individual i's local area who are "low paid" in real terms (after adjustment by the RPI to April 1999 terms). Again the estimate of β^* is scaled to take account of the non-linear functional form to give an estimate of β.

The results for equation (9) are given in the first column of Table 4. The table gives "marginal effects" of the logit coefficients. The estimate of β is positive and just significant at the 5% level for the all ages sample without area fixed effects, but less significant when area fixed effects are included. The difference-in-differences estimates based on the "treatment" and "comparison" groups defined above are given in the second column and are insignificant, with or without the area fixed effects. Indeed the area fixed effects seem to make very little difference to the estimate of θ.

Because the "comparison" group used is dominated by London, a second comparison group is also examined. It consists of the bottom 10% of the NES 1998 sample in terms of "low pay" rates, but excluding London boroughs.[16] The difference-in-differences estimator based on this "comparison" group is also insignificantly different from zero for the all-ages sample, both with and without area fixed effects.

For youths the estimate of β is insignificantly different from zero, positive without area fixed effects, negative with them. The difference-in-differences estimates for this group are small and insignificant for both comparison groups.

5. Focusing on those most likely to be affected

A potential problem with the results in the previous two sections may be that, since only a minority of those in any local area have their wages raised as a result of the introduction of the minimum wage (see Section 2), any employment effects for this minority may get swamped by the lack of an effect for the majority in the remainder of the wage distribution whose wages are not affected. This can be addressed by restricting the analysis to those most likely to be affected. There is an illustration of this above when looking at youths only. Alternative definitions of those at high risk of being affected are potentially useful and are examined in this section.

Thus this section repeats the later stages of the analysis of the previous two sections restricting the analysis to high risk groups, defined in various ways. The first set of estimates presented are of equation (7) using constructed area-level panel data. (Estimates of this equation using the full sample are given in the final block of Table 3.) For each definition of the high risk group considered, employment and low-pay rates are calculated for each local area in each year for members of that high risk group only. Estimates of β in equation (7) and of θ in the equivalent equation using the high and low impact groups of local areas in place of the continuous proportion variable are given in the first two columns of Table 5.

[16] The group contains the 7 counties of Buckinghamshire, Oxfordshire, Surrey, Berkshire, Hertfordshire, Cambridgeshire and Bedfordshire (listed in low-pay rate order, lowest first). All have low pay rates of 2.6% and below and the overall low-pay rate is 2.3%.

The second set of estimates presented in Table 5 are for equations (8) and (9) using individual-level panel data. (Estimates of these equations using the full sample are given in Table 4.) For each definition of the high-risk group considered, the proportion low paid is calculated as described above. The sample for the estimated logit models is then also restricted to those in the high risk group. Estimates of β and θ from the equivalents of equations (8) and (9) are given in the last two columns of Table 5.

The first high-risk group considered are women, who are more than twice as likely as men to have been low paid. The estimates for women of both β and θ at both the area and individual levels are positive but insignificantly different from zero. A similar conclusion results if attention is restricted to those who have been in their jobs less than 12 months (who are more likely to be low paid than those who have been in their jobs longer than this).

The next two high risk groups are both defined in terms of (major group) occupations. Those in manual, personal service and clerical occupations are much more likely to be low paid than those in other occupations, with those in unskilled occupations most likely of all. The next row considers estimates for those in Standard Occupational Classification (SOC) major groups 4-9. This consists of those in: clerical & secretarial occupations (SOC major group 4), craft & related occupations (5), personal & protective service occupations (6), sales occupations (7), plant & machine operatives (8) and other occupations (9). It therefore excludes: managers & administrators (SOC major group 1), professional occupations (2) and associate professional & technical occupations (3). Again all four estimates are insignificantly different from zero. The area-level difference-in-differences specification gives a very small negative estimate.

The next row restricts attention all the way down to unskilled occupations (SOC major group 9) only, with the consequent considerable reduction in cell sizes. Again all four estimates are insignificant, although two are now negative. The individual-level estimate of β is the most so, but still only has an absolute t-ratio of 1.13.

Looking next at industrial sectors, again two specifications of "high risk" are considered, one fairly broad and one very tight. The broad one uses the 6 sectors (out of 15) that have above average low-pay probabilities. These are: hotels and restaurants (1992 SIC main division H), agriculture, hunting and forestry (A), wholesale and retail trade: repair of motor vehicles, motorcycles and personal household goods (G), health and social work (N), other community, social and personal service activities (O) and real estate, renting and business activities (K). This last includes for example industrial cleaning and security activities. Using this sample, again all four estimates in Table 5 are positive but insignificant (although the significance of the individual-level estimate of θ is marginal). When only the hotels and restaurants industry (which is the SIC main division with the highest proportion of low paid workers) is used, with accompanying reduction in cell sizes, the individual-level estimate of θ is negative, although insignificant. The other three estimates are positive and insignificant.

The estimates so far in this section all use a single variable to define the "high risk" criterion. An alternative is to combine these criteria (and others) and model the probability of being low paid (i.e. being paid below the subsequent minimum), following part of the analysis in Card and Krueger (1995). The model fitted includes as explanatory variables: age (a quartic), part-time, in job less than 12 months, 8 occupation dummies (for SOC major groups), 14 industry dummies (for SIC main divisions) plus gender and its interaction with all the other variables listed. This model is then used to predict the probability of an individual needing their wage rate raised to comply with the incoming minimum. Those with predicted probabilities in the top quartile are then taken to be the "high risk" group and analysed as above.

A further enhancement to provide additional controls for the area-level analysis, again following Card and Krueger (1995), is that those with predicted probabilities in the lower half of all workers are used to construct area-level measures of the change in the employment probability and the change in the average wage for this group to capture area-specific movements in the part of the distribution unaffected by the minimum wage. The area-level and individual-level estimates of β and θ for this definition of the "high risk" group are given in the final row of Table 5. All four estimates are positive and insignificant.

Overall the range of estimates presented in Table 5 provide very little evidence to suggest an adverse employment effect of the introduction of the minimum wage. All the estimates, of both β and θ at both the area-level and the individual-level, are insignificant and the great majority of them are positive.

6. Conclusions

This paper evaluates the impact on employment of the UK's introduction of a minimum wage in 1999 by exploiting the geographical variation in wages, which meant that the minimum wage's "bite" into an area's wage distribution differed considerably across the country. The paper tests the prediction of the standard textbook model of the labour market that there should be a relative decline in employment in areas with high proportions of low-wage workers compared to areas where relatively few employees' wages were affected.

The results presented indicate that, although the introduction of the National Minimum Wage had differential effects on the wage distribution across the areas of the country, employment growth after its introduction was not significantly different in areas of the country with a high proportion of low-wage workers, whose wages had to be raised to comply, from that in areas with a low proportion of such workers. The findings in the paper are consistent with the view that the minimum wage's introduction had no systematic adverse effect on employment.

References

Abowd, John M., Francis Kramarz, David N. Margolis and Thomas Philippon (2000), "A Tail of Two Countries: Minimum Wages and Employment in France and the United States", mimeo, CREST, Paris, September.

Angrist, Joshua and Alan Krueger (1999), "Empirical Strategies in Labor Economics", in O. Ashenfelter and D. Card (eds.), *Handbook of Labor Economics*, Volume 3 (Elsevier).

Brown, Charles (1999), "Minimum Wages, Employment and the Distribution of Income", in O. Ashenfelter and D. Card (eds.), *Handbook of Labor Economics*, Volume 3 (Elsevier).

Burkhauser, Richard V., Kenneth A. Couch and David C. Wittenburg (2000), "A Reassessment of the New Economics of the Minimum Wage Literature with Monthly Data from the Current Population Survey", *Journal of Labor Economics*, 18, 653-680.

Card, David (1992a), "Using Regional Variation in Wages to Measure the Effects of the Federal Minimum Wage", *Industrial and Labor Relations Review*, 46, 22-37.

Card, David (1992b), "Do Minimum Wages Reduce Employment? A Case Study of California, 1987-1989", *Industrial and Labor Relations Review*, 46, 38-54.

Card, David and Alan B. Krueger (1994), "Minimum Wages and Employment: A Case Study of the Fast-food Industry in New Jersey and Pennsylvania", *American Economic Review*, 772-793.

Card, David and Alan B. Krueger (1995), *Myth and Measurement: The New Economics of the Minimum Wage*, Princeton NJ: Princeton University Press.

Deere, Donald, Kevin Murphy and Finis Welch (1995), "Employment and the 1990-1991 Minimum Wage Hike", *American Economic Review*, 85, 232-237.

Dickens, Richard, Stephen Machin and Alan Manning (1999), "The Effects of Minimum Wages on Employment: Theory and Evidence from Britain", *Journal of Labor Economics*; 17, 1-22.

Dolado, Juan, Francis Kramarz, Stephen Machin, Alan Manning, David Margolis and Coen Teulings (1995) The Economic Impact of Minimum Wages in Europe, *Economic Policy*, 23, 317-72.

Huber, P. J. (1967), "The Behaviour of Maximum Likelihood Estimates Under Non-standard Conditions", *Proceedings of the Fifth Berkeley Symposium on Mathematical Statistics and Probability*, 1, 221-233.

Kim, Taeil and Lowell J. Taylor (1995), "The Employment Effect in Retail Trade of California's 1988 Minimum Wage Increase", *Journal of Business and Economic Statistics*, 13, 175-182.

Linnerman, Peter (1982), "The Economic Impacts of Minimum Wage Laws: A New Look at an Old Question", *Journal of Political Economy*, 443-469.

Machin, Stephen and Alan Manning (1994), "The Effects of Minimum Wages on Wage Dispersion and Employment: Evidence from UK Wages Councils", *Industrial and Labor Relations Review*, 47, 319-329.

Machin, Stephen, Alan Manning and Lupin Rahman (2002), "Care Home Workers and the Introduction of the UK National Minimum Wage", mimeo, LSE.

Meyer, Bruce (1995), "Natural and Quasi-experiments in Economics", *Journal of Business and Economic Statistics*, 13, 151-161.

Moulton, B. R. (1986), "Random Group Effects and the Precision of Regression Estimates", *Journal of Econometrics*, 32, 385-397.

Neumark, David, Mark Schweitzer and William Wascher (2000), "The Effects of Minimum Wages Throughout the Wage Distribution", NBER Working Paper 7519.

Neumark, David and William Wascher (2000), "Minimum Wages and Employment: A Case Study of the Fast-food Industry in New Jersey and Pennsylvania: Comment", *American Economic Review*, 90 (December), 1362-1396.

Office for National Statistics (2001), *Regional Trends, No. 36* (editors: Jon McGinty and Tricia Williams). Also available at http://www.statistics.gov.uk.

Partington, James (2001), "The Launch of the Annual Business Inquiry", *Labour Market Trends*, May, 1-7.

Stewart, Mark B. (2002), "The Impact of the Introduction of the UK Minimum Wage on the Employment Probabilities of Low Wage Workers", Economics Department, University of Warwick, Research Paper 630.

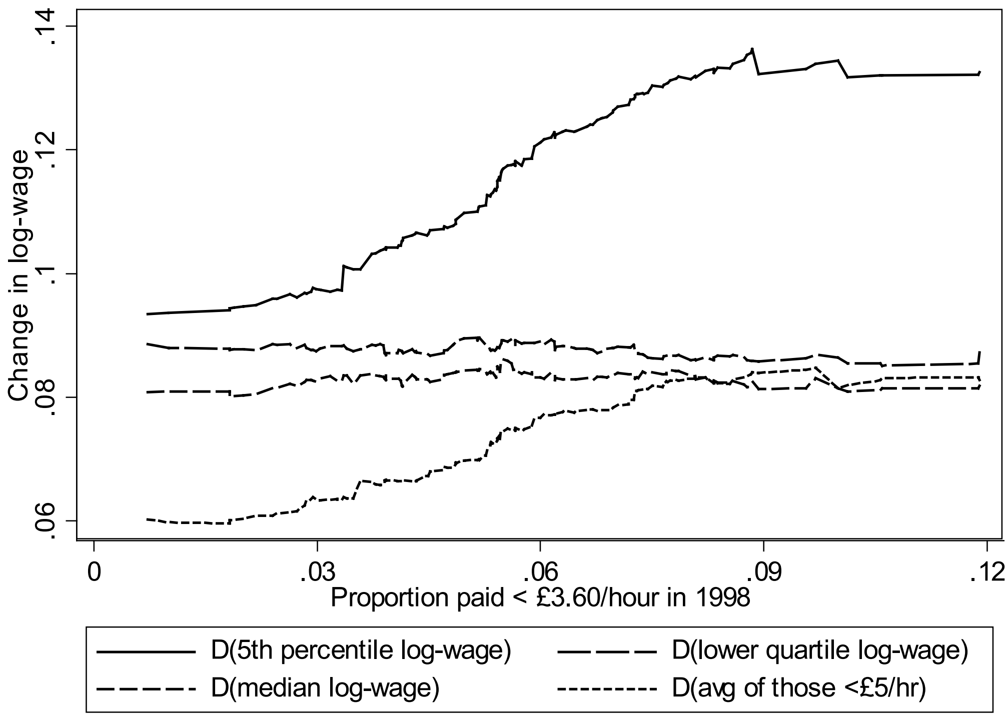

Figure 1: Changes in Area Wage Distributions, 1998-2000

Figure 2: Kernel Wage Density Estimates (1998 NES)

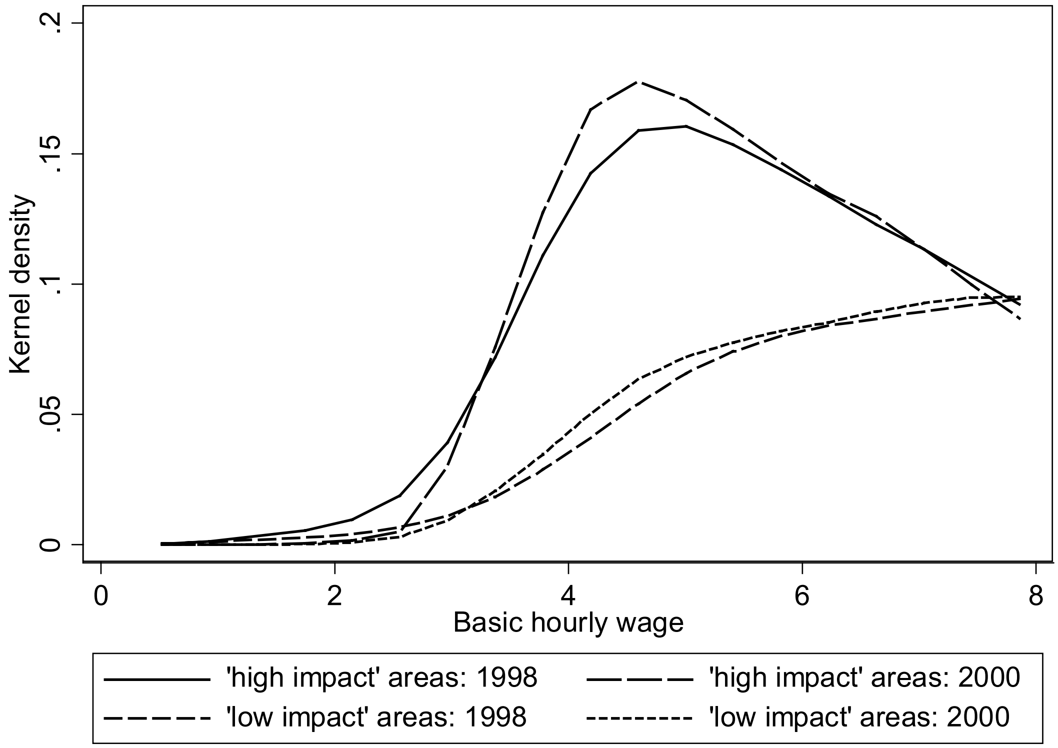

Figure 3: Lower Ends of Kernel Wage Densities for 1998 and 2000

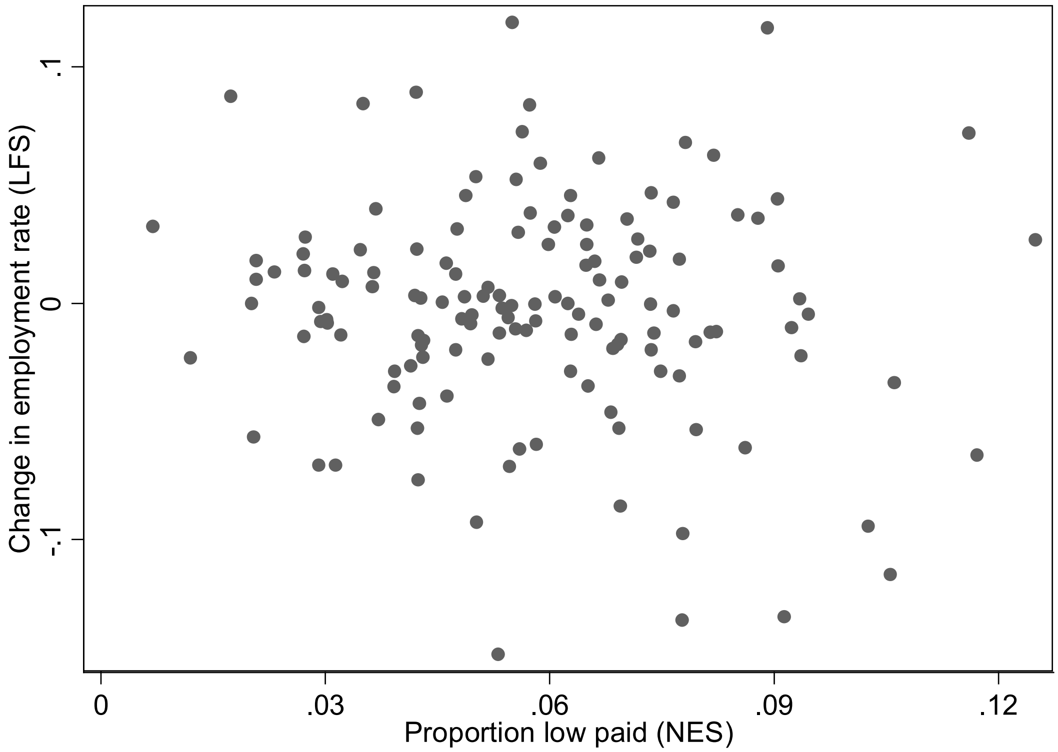

Figure 4: Area-level Scatter Plot of Employment Change against Proportion Low Paid

Table 1

Area-level Regression Estimates of Relationship Between Wage Growth and the Proportion Low-paid in 1998

Dependent variable	Coefficient on proportion low paid in 1998
Δ average log-wage	.175 (1.42)
Δ average log-wage for those with wage < £5/hour	.520 (4.11)
Δ median log-wage	.143 (0.74)
Δ 10th. percentile of log-wage	.778 (4.46)
Δ 5th. percentile of log-wage	1.094 (5.25)
Δ lower quartile of log-wage	.382 (2.93)
Δ proportion low-paid	-.717 (19.28)

Notes:

1. Source: New Earnings Survey.

2. Each row gives an area-level regression using 140 “local areas”.

3. Robust absolute t-ratios in parentheses.

Table 2

Difference-in-differences estimates of minimum wage impact on various aspects of the wage distribution

Dependent variable	Difference-in-differences estimate
<u>Area-level difference-in-differences:</u>	
Average log-wage	-.001 (0.01)
Proportion low-paid	-.053 (12.40)
Average log-wage for those with wage < £5/hour	.044 (2.66)
5th. percentile of log-wage	.079 (2.27)
10th. percentile of log-wage	.054 (1.28)
Lower quartile of log-wage	.009 (0.18)
<u>Individual-level difference-in-differences:</u>	
Log(wage)	-.009 (0.95)
Low-paid indicator	-.048 (15.21)
Quantile regression (5th. percentile) for log-wage	.054 (4.97)
Quantile regression (10th. percentile) for log-wage	.031 (3.05)
Quantile regression (lower quartile) for log-wage	.005 (0.48)

Notes:

1. Source: New Earnings Survey.
2. Robust absolute t-ratios in parentheses.
3. The t-ratios for the individual-level equations are based on standard errors that allow for "clustering" by local area.
4. The t-ratios for the quantile regressions use bootstrap standard errors.
5. Individual-level estimates based on sample of size 51,709.

Table 3

Estimates of minimum wage impact on employment: area-level analysis

Source for employment rates	Employment growth equation: slope coefficient	Difference-in-differences estimate
LFS Local Area Data files:		
All aged 18+	-.165 (0.84)	-.020 (0.44)
Youths only (18-21)	.077 (0.33)	-.034 (0.39)
Ratio of total employment from ABI to ONS estimate of population of working age:	-.121 (0.78)	-.024 (0.64)
Ratio of NES frequency x 100 to ONS estimate of population of working age:	-.052 (0.36)	-.010 (0.22)
NES continuation probability:		
All aged 18+	.484 (2.54)	.068 (4.77)
Youths only (18-21)	.165 (1.25)	.041 (0.93)
NES continuation probability: Difference relative to 1996-98 change:		
All aged 18+	.177 (0.66)	.003 (0.11)
Youths only (18-21)	.353 (1.14)	-.006 (0.12)

Notes:

1. Information on proportion low paid in 1998 is taken from the NES for that year.

2. See text for sources for employment rates.

3. Robust absolute t-ratios in parentheses.

Table 4

Estimates of minimum wage impact on employment: individual-level analysis

		Difference-in-differences Estimates	
Specification	*Coefficient on proportion low paid*	*Comparison group A*	*Comparison group B*
All aged 18+:			
without area fixed effects	.390 (2.01)	.011 (1.42)	.012 (1.55)
with area fixed effects	.285 (1.66)	.010 (1.35)	.012 (1.54)
Youths only (18-21):			
without area fixed effects	.111 (0.31)	-.002 (0.17)	.001 (0.14)
with area fixed effects	-.033 (0.10)	-.003 (0.31)	.0004 (0.04)

Notes:

1. Estimated by Logit models using NES data with controls for year (dummies), age (4th. order polynomial), gender, part-time status, and an indicator for less than 12 months in job.
2. Robust absolute t-ratios that allow for "clustering" in parentheses.
3. Sample size = 789,141.

Table 5

Estimates of minimum wage impact on employment: individual-level analysis for certain "high risk" groups

	Area-level		Individual-level	
High risk group:	β	θ	β	θ
Women only	.169 (0.86)	.024 (0.74)	.198 (1.32)	.001 (0.10)
In job less than 12 months	.201 (0.89)	.047 (1.33)	.067 (0.66)	.012 (1.34)
In an occupation in SOC major groups 4 – 9	.053 (0.25)	-.008 (0.34)	.129 (0.90)	.009 (1.11)
In an unskilled occupation (major group 9)	.135 (0.73)	-.030 (0.66)	-.122 (1.13)	.016 (1.49)
In one of the industries with above-average proportions low paid	.013 (0.09)	.001 (0.03)	.127 (1.29)	.014 (1.76)
In the hotels and restaurants industry	.136 (0.78)	.097 (1.29)	.089 (0.90)	-.035 (1.50)
Predicted probability in top quartile	.157 (1.17)	.009 (0.29)	.0003 (0.01)	.007 (0.53)

Notes:

1. Estimated by Logit models using NES data with controls for year (dummies), age (4th. order polynomial), gender, part-time status, and an indicator for less than 12 months in job.
2. Robust absolute t-ratios that allow for "clustering" in parentheses.
3. Standard Occupational Classification (SOC) major groups 4-9 (for the estimates in the 3rd. row) consists of : clerical & secretarial occupations (SOC major group 4), craft & related occupations (5), personal & protective service occupations (6), sales occupations (7), plant & machine operatives (8) and other occupations (9), which are unskilled.
4. The industries with above average proportions of low-paid employees (for the estimates in the 5th. Row) are: hotels and restaurants (1992 SIC main division H), agriculture, hunting and forestry (A), wholesale and retail trade: repair of motor vehicles, motorcycles and personal household goods (G), health and social work (N), other community, social and personal service activities (O) and real estate, renting and business activities (K), this last including for example industrial cleaning and security activities.
5. Sample sizes vary across the rows. The smallest are 16,365 for hotels and restaurants (row 6) and 64,129 for unskilled occupations (row 4).

Appendix: Definitions of high- and low-impact "local area" groups

High-impact group:
Areas with the highest proportion of employees who are "low paid" in April 1998 (defined by the 1999 age-specific minimum wage rates (£3.60/hour for adults, £3.00/hour for youths)). These areas all have 7.3% or more who are low paid by this definition. Between them the 33 areas comprise 10% of the employees in the 1998 New Earnings Survey.

> Torbay, Redcar & Cleveland, North Ayrshire, Northumberland, Blackpool, Darlington, East Riding of Yorkshire, Middlesborough, Pembrokeshire, East Lothian, Argyll & Bute, Lincolnshire, Carmarthenshire, Gwynedd, Hartlepool, Blaenau Gwent, Shropshire, Worcestershire, Conwy, Aberdeenshire, Kingston upon Hull, Cornwall & Scilly Isles, Inverclyde, Monmouthshire, Clackmannanshire, East Dunbartonshire, Perth & Kinross, Wrexham, Powys, Dumfries & Galloway, Devon, Isle of Anglesey, Stockton-on-Tees.

Low-impact group:
Areas with the lowest proportion of employees who are "low paid" in April 1998 (defined by the 1999 age-specific minimum wage rates (£3.60/hour for adults, £3.00/hour for youths)). These areas all have less than 2% who are low paid by this definition. Between them the 6 areas comprise 10% of the employees in the 1998 New Earnings Survey.

> Bracknell Forest, West Berkshire (Newbury), Milton Keynes, Inner London, Reading, Oxfordshire.

Table A.1

Comparison of the "high impact" and "low impact" groups

	"High Impact" group	"Low Impact" group
Sample size	12,960	12,739
Mean basic hourly wage	£7.46	£12.13
Median basic hourly wage	£6.10	£9.66
10th. percentile	£3.62	£4.98
5th. percentile	£3.25	£4.24
Proportion "low paid"	.086	.019

Source: New Earnings Survey, 1998.