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Minimum Wages and Schooling: Evidence from the UK's Introduction of a National Minimum Wage

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Abstract
This paper uses the introduction of the national minimum wage in the UK in April 1999 as a 'natural experiment' to analyse the impact of minimum wages on enrolment in schooling. At the time of its introduction, only workers aged 18 years or more were covered by the legislation. The paper uses panel data for a sample of young people in a given school-year cohort, some of whom were aged 18 years in spring 1999 and therefore eligible to receive the national minimum wage, and others who were aged only 17 years. We compare participation in post-compulsory schooling for the two groups, both before and after the enactment of the legislation and find robust evidence that eligibility for the national minimum wage significantly reduces the probability of participation in post-compulsory schooling for young people living in areas where the national minimum is high relative to local earnings.

Keywords: minimum wages; enrolment in schooling; natural experiment approach
JEL classification: J22, J24, J38

1. Introduction

Few issues in economics are as contentious as the effects of statutory minimum wages on labour market outcomes. While much of the debate focuses on employment, the potential impact on enrolment in schooling is also a matter of some dispute. On the one hand, by raising wages for the unskilled and reducing the wage differentials between skilled and unskilled labour, wage floors reduce the incentives to invest in further education or training. It has been suggested that "reduced training opportunities or lowered educational attainment could be much more widespread than disemployment effects, and could – by lowering skill acquisition at young ages – have longer lasting consequences for the affected individuals" (Neumark and Wascher, 2003). Others argue that by increasing the relative demand for higher skilled workers, minimum wages may actually increase the incentives to invest in education and training in order to compete effectively for more skilled jobs (Cahuc and Michel, 1996). In contrast to the very large empirical literature examining the employment effects of minimum wages, empirical studies of the effects on school enrolment are relatively few in number and almost all relate to North America. In this paper, we use the introduction of a national minimum wage in the UK in Spring 1999 as a 'natural experiment' to investigate the effect of minimum wages on enrolment in post-compulsory schooling among a cohort of young people aged 17/18 years.

In April 1999, following the recommendations of the independent Low Pay Commission, the UK government introduced a national minimum wage. This followed a period of many years in which there was no statutory minimum wage for most sectors of the UK economy[1]. The new legislation specified a minimum hourly rate for all employees aged 22 years or more, and a lower 'youth development' rate for those aged 18 to 21 years.

[1] The exception was the agricultural sector where the Agricultural Wages Board (AWB) had a statutory obligation to fix minimum wages for employees in England and Wales. The Board also had discretionary powers to decide other terms and conditions of employment, e.g. holidays and sick pay.

Individuals below the aged of 18 years were exempt from the legislation until 2003, when a third "youth" rate was introduced for those aged 16 and 17 years. This paper makes use of the fact that assignment to the treatment – eligibility to receive the national minimum wage – is determined by being on either side of a fixed age threshold of 18 years. Our empirical analysis relates to a sample of young people drawn from a given school-year cohort, some of whom were aged 18 years in April 1999 and hence belong to the treatment group, and others who were below the age of 18 years and form the control group. Thus, we avoid the necessity of making assumptions about the comparability of a treatment and control group whose members are drawn from different parts of the wage distribution or from different geographical areas as in previous studies (e.g. Stewart, 2004).

Uncovering the causal effect of eligibility for the minimum wage is complicated by the possibility that other unobservable factors associated with the individual's relative age may affect a young person's participation in post-compulsory schooling. Those who are relatively old in their school year appear to have significant educational and social advantages over their younger counterparts and such advantages have been shown to persist into adulthood, through differences in human capital accumulation and in the development of softer skills such as maturity and leadership (Bedard and Dhuey, 2006; Cunha et al, 2006). For our sample, we have detailed information on their participation in education and other activities during the spring of 1998, before the introduction of the national minimum wage, and again during early summer 1999, once the legislation was in place. The availability of panel data allows us to control for any unobserved heterogeneity that may be associated with the relative age of an individual and hence correlated with their eligibility for the minimum wage.

In what follows, we use a conditional logit framework to estimate the treatment effect of eligibility for the national minimum wage on enrolment in post-compulsory schooling.

With the effect of eligibility for the minimum wage restricted to be homogenous across members of the school-year cohort, we find no evidence of a statistically significant impact on the probability that a young person is enrolled in full-time post-compulsory schooling. However, the 'bite' of the national minimum - the extent to which it affected the distribution of earnings in the local labour market - varies considerably across Great Britain as shown in Stewart (2002). It would be surprising if the national minimum wage had the same impact on young people's behaviour in the relatively well-paid areas of south-east England as in the less prosperous local labour markets of the north-east. If the treatment effect is allowed to vary with the 'bite' of the national minimum wage then eligibility to receive the national minimum wage is found to significantly reduce the probability of enrolment in post-compulsory schooling for young people living in relatively low paid areas. Taking the ratio of the national minimum wage to the 20th percentile of the pre-legislation distribution of hourly earnings as the measure of 'bite', our results suggest that the national minimum wage reduced enrolment in schooling in those local areas where the 'bite' exceeded some 64 percent. This was the case in around 60 percent of local areas in England and Wales, covering nearly 70 percent of the school-year cohort.

The paper is organised as follows. Section 2 reviews the existing empirical evidence on the effects of minimum wages on school enrolment. Section 3 explains the model set-up and estimation strategy, and the data are described in some detail in section 4. A premise of the estimation strategy is that the introduction of the national minimum wage in 1999 significantly increased the wages of those workers covered by the legislation relative to those of their younger counterparts and we examine the evidence for this in section 5. Section 6 presents our main results on the effects of minimum wages on school enrolment, and the robustness of these findings is assessed in section 7.

2. Related Literature

The theoretical implications of minimum wages for enrolment in schooling are far from clear cut. In a simple static setting, introducing a minimum wage, like any price change, has income and substitution effects on an individual's choices. The introduction of a wage floor, assuming that it is effective, raises wage rates for some young unskilled workers and in so doing increases their opportunity costs of schooling. At the same time, by increasing their expected income, the minimum wage may induce some individuals to reduce their hours of work and consume more schooling. So for example a young person may choose to switch from full-time working to part-time working combined with enrolment in schooling. This simple static analysis becomes more complicated when we allow for the possibility that by reducing the relative demand for unskilled labour, a minimum wage may reduce the probability of employment for young school leavers, offsetting the effects of higher wages on expected income.

From the perspective of an investment decision, if the introduction of a minimum wage reduces wage differentials between skilled and unskilled workers then the expected returns to human capital investment are reduced. Others have argued that by increasing the relative demand for more skilled labour, minimum wages increase the incentives to invest in education in order to compete for higher skilled jobs (Cahuc and Michel, 1996), or to increase the probability of gaining minimum wage employment (Agell and Lommerud, 1997).

The empirical evidence on this question is mixed. Early influential studies by Neumark and Wascher (1995a,b,c) using US data report that increases in minimum wages lead to lower rates of school enrolment among 16 to 19 year olds, coupled with higher rates of inactivity (i.e. not in school and not employed), particularly among individuals in the youngest age category (16 and 17 year olds) and ethnic minorities. They conclude that

higher minimum wages result in individuals leaving full-time education in order to 'queue' for better-paid jobs. Employers tend to substitute in favour of these higher quality young workers, resulting in higher rates of unemployment among their lower quality counterparts. A more recent study by the same authors updating the data to 1998 confirms these earlier findings (Neumark and Waschter, 2003). Turner and Demiralp (2001) also report evidence that Black and Hispanic teenagers and teenagers based in inner-cities areas are more likely to become inactive as a result of a minimum wage increase.

Card (1992) focuses on the effect of the 27% increase in the California state minimum wage in 1988 and finds evidence of decreases in school enrolment in California relative to other comparable states that did not experience an increase in the minimum. More recently, Chaplin, Turner and Paper (2003) analyse US Department of Education data covering the entire population of public school students in the US. They find evidence of lower state-level continuation ratios, particularly between grade 9 and grade 10 (corresponding roughly to ages 16 to 17 years), for states with higher minimum wages.

In addition to the US evidence, there have been a series of studies based on Canadian data, exploiting the fact that minimum wage rates vary by province as well as by time. Here, the evidence on the effects on school enrolment rates is more mixed. Landon (1997) found evidence that higher minimum wages are associated with lower school enrolment rates among 16 and 17 year olds. By contrast, Baker (2003) found no evidence of an effect on the enrolment of those subject to compulsory schooling laws (ages 15 to 16 years) and a modest positive effect on older age groups (17 to 19 years and 20 to 24 years). Campolieti, Fang and Gunderson (2003) report no significant effects on school enrolment rates or rates of employment.

Little empirical evidence exists for countries outside of North America. In 2001, New Zealand reformed its minimum wage legislation, reducing the age of eligibility for the adult minimum to 18 years and at the same time increasing the youth minimum wage rate from 60 percent to 80 percent of the adult rate. Hyslop and Stillman (2007) examine the effects of these reforms on labour market outcomes for 16 to17 year olds and 18 to 19 year olds. They find evidence that these reforms had negative effects on participation in full-time education for both age categories. In addition, there was some evidence that the reforms led to higher rates of unemployment and inactivity among the younger age group, but this finding was less robust.

Previous studies of the impact of the introduction of a national minimum wage in the UK have focused on its effects on employment, hours of work and the distribution of earnings. There is a large body of empirical work, much of it reviewed in Metcalf (2008). Drawing together the results of these studies, Metcalf concludes that the national minimum wage contributed to higher levels of real and relative earnings for low paid workers, and to a significant decline in inequality in the lower half of the earnings distribution. There is little evidence of significant adverse effects on employment, although there is some evidence of a reduction in hours of work among those whose pay was raised as a result of the legislation (Stewart and Swaffield, 2008). None of the studies undertaken to date have considered the question addressed in the present paper; namely what has been the impact of the national minimum wage of on enrolment in schooling?

3. Model set-up and estimation strategy

The UK Labour Party came into government in May 1997 with a manifesto commitment to introduce a national minimum wage. The details of the legislation were not yet decided and an independent Low Pay Commission was tasked with producing recommendations on the coverage and the level of the proposed wage floor. The Low Pay Commission produced its

report in the summer of 1998, and shortly after, the details of the legislation were published. With effect from April 1st 1999, a national minimum wage (NMW) of £3.60 per hour would apply for all those aged 22 years or more, with a lower 'development' rate of £3.00 per hour for those aged 18 to 21 years. Those below the age of 18 years were exempt from the legislation.

Against this background, the school-year cohort of young people used in our empirical analysis completed their compulsory schooling, and proceeded to further education, work-based training, employment or in some cases, inactivity/unemployment. Our data relates to a sample of young people who completed their compulsory schooling during the summer of 1997, shortly after the new Labour government took power. The first sweep of data on this group was collected in the spring of 1998, prior to the publication of the Low Pay Commission's recommendations. The second sweep of data was collected in the early summer of 1999, after the introduction of the national minimum wage. At the time of the second sweep, some members of cohort were aged 18 years and hence eligible to receive the lower 'development' rate of £3 per hour; others were still only 17 years of age and hence not covered by the legislation.

The question we wish to address in this paper is what would have been the rate of enrolment in full-time education of those young people eligible to receive the national minimum if the legislation had not been introduced, and do their observed participation rates differ significantly from these. The approach adopted is to compare changes in the enrolment in full-time education between spring 1998 and early summer 1999 for the 'treatment' group of those aged 18 years at the time of the second sweep, with the experience of the 'control' group of those aged only 17 years.

To be more precise, let S_{ij} denote the school enrolment status of individual i at sweep j (S_{ij} =1 if i is enrolled in full-time education at sweep j, S_{ij} =0 otherwise, for j=1 or 2). Suppose that in the absence of the minimum wage legislation, the probability of enrolment in post-compulsory schooling evolves over time according to some function of the individual's age, which itself is determined by the individual's date of birth and the survey date. Further assume that the introduction of the national minimum wage has a constant effect, θ, on the enrolment rate for those treated and no effect on enrolment rates for members of the control group. Under these assumptions, the enrolment status of individual i in sweep j may be modelled as

$$\begin{aligned} &S_{ij} = 1 \quad \text{if} \quad \alpha_i + \delta.T_j + \theta.NMW_{ij} + \varepsilon_{ij} > 0 \\ &S_{ij} = 0 \quad \text{otherwise.} \end{aligned} \tag{1}$$

Where α_i is an unobserved time-invariant individual-specific effect which subsumes any date-of-birth effects; T_j is an indicator variable that takes the value of one for j=2 and is equal to zero otherwise. NMW_{ij} is the treatment indicator; so $NMW_{ij} = 1$ if individual i is aged 18 years and j=2, and is equal to zero otherwise. ε_{ij} is an unobservable error term. The simple specification in (1) may be extended by adding a vector of additional control variables, $\mathbf{x}_{ij}$, that are thought to affect the probability that an individual is enrolled in school at a given age.

There are a number of alternative approaches to estimating a model (1). One is to parameterize the distribution of the α_i conditional on T_j and NMW_{ij} making the model fully parametric. The main drawback of this so-called random effects approach is that if the distributional assumptions do not hold then in general all the parameter estimates are inconsistent. The alternative is to treat the α_i as parameters and thereby avoid making any

assumptions regarding their distribution. With a large number of individuals and a small fixed number of time periods, as in the present case, the number of parameters increases with sample size giving rise to the 'incidental parameters' problem which leads also to inconsistent parameter estimates. This problem can be avoided by identifying a feature of the model that depends on the parameter(s) of interest, in this case the treatment effect, θ, but not on the α_i. An example of this approach is the conditional logit model (Chamberlain, 1984). If we assume that the error ε_{ij} is logistically distributed independent of α_i, T_j, NMW_{ij} then conditioning on $S_{i1} + S_{i2} = 1$

$$\Pr[S_{i1} = 1 \mid S_{i1} + S_{i2} = 1, \alpha_i, D_{i1}, NMW_{i1}, D_{i2}, NMW_{i2}] = \frac{1}{1 + \exp(\delta + \theta . NMW_{i2})} \qquad (2)$$

In other words, for those individuals whose enrolment status changes between sweep 1 and sweep 2 of the survey, the probability that it changes from 1 to 0, as opposed to changing from 0 to 1, is described by a logit model with explanatory variables equal to the first difference of the variables in (1) and does not depend on the α_i. The treatment effect θ can be estimated from (2) without making any assumptions on the individual-specific effects, α_i.

As noted by Honore (2002), it is intuitively appealing that the individuals who do not switch enrolment status are not used to estimate the treatment effect, θ, since their behaviour can be rationalized by an extremely large or an extremely small values of α_i for any value of θ. However, there are costs to this approach. Most notably, by estimating θ in (2) we can assess whether or not the treatment – in this case eligibility for the national minimum wage - has a significant impact on individual behaviour. We also can estimate the effect of the treatment on the probability that the individual is enrolled in full-time schooling conditional

on a particular value for α_i, but it is not possible to calculate the average effect of the treatment across the distribution of α_i in the population.

The key identifying assumption of this model is that eligibility for the national minimum wage is the only source of discontinuity in behaviour at age 18 years. Clearly this is questionable. It is possible that due to custom and practise, young workers move from juvenile to adult pay scales on attaining the age of 18 years, and this produces a discontinuity in behaviour. A second identifying assumption is that the introduction of the national minimum wage has no impact on the younger age group who are not directly covered by the legislation. However, there may be wage spillovers. Firms may choose to pay the minimum wage to all their younger workers, irrespective of whether they are 18 years of age or younger. Alternatively, firms may increase their demand for workers in the younger age category who are not covered by the legislation, leading to increases in the wages of this group. The robustness of our results to possible violations of these assumptions are investigated later in the paper.

4. Data

Our data comes from the ninth Youth Cohort Study for England and Wales (YC9). The Youth Cohort Study is a longitudinal study of young people between the ages of 16 and 20 years focusing on their education, training and employment. The sample is selected from pupils attending eligible schools in the maintained and independent sectors (excluding special schools) by taking those who were born on the 5th, 15th and 25th of each month. In the case of YC9, this provided a total sample of 22,498, of which 21,105 were in England. The first sweep of data was collected by postal questionnaire and telephone interview between March and May 1998. The number of legible responses received by the cut-off date in early June 1998 was 14,662, a response rate of 65.6%. The second sweep of YC9 data was collected a

little over a year later, between April and July 1999, with questionnaires sent out to all those who had responded in sweep 1. In this case, a total of 9,662 legible responses were obtained by the cut-off date in late July. It is this sample of 9,662 individuals who provided data at both sweeps that is used in the analysis that follows.

The Youth Cohort Study provides detailed information on the young person's current activities – schooling, training and work - together with a monthly calendar of their main activity in each of the preceding twelve months. At the first sweep, data on a wide range of background characteristics also is collected; this includes previous schooling and qualifications gained, ethnicity, parent's education and occupation. The individual-level data from YC9 is supplemented by data relating to the local labour market including measures of earnings, employment composition and unemployment[2].

The YC9 sample is designed to be representative of the population who reached the minimum school-leaving age in 1996/97. However, there is ample evidence of differential response rates by gender and by school attainment level. To avoid potential biases from this source, sample weights are used designed to match the responding sample at sweep 2 to the population of England and Wales with respect to number of known characteristics including gender, region, school type, GCSE attainment levels[3].

Eligibility for the national minimum wage is determined from the information provided on the individual's month and year of birth, and on the month that the sweep 2 survey is returned. In the absence of information on the actual date of birth and of survey return, the estimate of the individual's age is correct only to within ±1 month. Given this measurement error, all those whose estimated age at the date of the sweep 2 return is 217

[2] The YCS contains information on the local education authority of the individual. LEAs are matched to local authority level data on labour market variables available through the National Online Manpower Information Service (NOMIS).

[3] For details of the construction of the sampling weights see Finch et al (2004)

months or more are 18 years of age and eligible to receive the national minimum wage [$NMW_{i2} = 1$]. Those with an estimated aged of 215 months or less are only 17 years-old and therefore not covered by the legislation [$NMW_{i2} = 0$]. The group with an estimated age of 216 months includes both eligible and ineligible individuals. One way to handle this group is simply to exclude them from the estimation sample altogether, and this is the approach adopted in much of the analysis that follows. An alternative is to assign them an average 'treatment' value on the assumption that their true age is uniformly distributed across the interval 215 to 217 months, in which case $NMW_{i2} = 0.5$ for those with an estimated age of 216 months. We present some results based on this approach as a further check on the robustness of our findings.

Before proceeding with the analysis, we take a look at some descriptive statistics for the treatment and control groups in order to identify any possible systematic differences in relevant characteristics. Aside from age, the only sample characteristic of those listed in Table 1 that differs significantly across the two groups is gender, with a four percentage point difference in the proportion of females in the treatment group and in the control group. This is a reflection of the tendency of girls to complete and return the survey more promptly than boys, with the result that they are younger on average at the date of survey return. Controlling for the month of survey return, the difference in the gender composition of the two groups is small, around 1 percentage point.

5. The impact of the introduction of a national minimum wage on the wages of young workers

A premise of our empirical analysis is that the introduction of the national minimum wage significantly increased the wages of those workers covered by the legislation relative to their younger counterparts. The Low Pay Commission in their second report concluded that the

introduction of the development rate had had a substantial impact on the pay of young workers, affecting "the earnings of a larger proportion of 18-21 year olds than those aged 22 and above who have benefited from the full minimum wage".(Low Pay Commission (2000), p 79). However, this does not rule out the possibility that those below the age of 18 years also benefited significantly from the legislation.

Some information on the earnings of young people before and after the introduction of the national minimum is available from the YC9. For those employed in a full-time or part-time job or on a government-supported training programme at the time of the survey, information is collected on their usual take-home pay, after deductions but including overtime and bonuses, on a weekly or a monthly basis as appropriate, and their usual weekly hours of work.[4] Clearly there are a number of drawbacks to this data for current purposes. The data relates to take-home pay rather than gross wages. An estimate of hourly pay is computed by converting reported earnings to a weekly basis and then dividing by reported weekly hours of work. This introduces two possible sources of measurement error which combine multiplicatively and in order to reduce their possible effects, we exclude the 1 percent tails of the sample distribution in both reported earnings and reported hours of work.

Summary statistics for the distribution of hourly earnings for those in the YC9 sample who reported earnings from employment in both sweep 1 and sweep 2 are shown in Table 2. These suggest that differences in the earnings distribution of the two groups were small in sweep 1, but by sweep 2, there is a significant differential in favour of the older age group, particularly in the lower half of the earnings distribution. These findings can be seen more clearly in the kernel density estimates of (ln) hourly earnings depicted in Figure 1. Here we can see a significant shift to the right in the distribution of earnings for the older age

[4] Those with more than one current job are asked to provide this information for the job with the most hours of work For further details see Finch et al (2004)

group relative to their younger counterparts between sweep 1 and sweep 2. The shift is even more pronounced if we exclude from the sample those enrolled on government-supported training schemes, some of whom would have been exempt from the national minimum wage under the terms of the legislation (see Figure A1 of the Appendix).

While there is a clear evidence of an earnings differential in favour of the older age group opening up between sweep 1 and sweep 2 of YC9, it does not follow that this is a direct result of the introduction of the national minimum wage. As already observed, it could be that young workers tend to transfer to higher adult rates of pay when they reach the age of 18 years. If this is the case then the key identifying assumption of the analysis – namely that the NMW legislation is the only source of a discontinuity in the relationship between the individual's age and their behaviour – would not hold.

One way to investigate this further would be to undertake a comparable analysis of the earnings distribution of the two groups – those aged 18 years at sweep 2 and those still aged 17 years at sweep 2 – for an earlier school-year cohort, pre-dating the introduction of the national minimum wage. Unfortunately this is less straightforward than it sounds. For the two preceding Youth Cohort Studies, the successive sweeps were carried out at two-yearly intervals rather than annually, and so the second sweeps took place when the participants were aged 18/19 years rather than aged 17/18 years. We have to go back to Youth Cohort Study 6, the first sweep of which was conducted in the spring of 1992, to obtain panel data with the same age structure as YC9. However there are substantial differences in the design of the questionnaire for this earlier study and that for YC9.

Moreover, there are major changes in further education and training provision over the period 1992 to 1998 which limits further the comparability of the two surveys.[5]

With these limitations in mind, the kernel density estimates of (ln) hourly earnings for those who reported earnings from employment in both sweep 1 and 2 of YC6 are shown in Figure A2 of the Appendix. The kernel density estimates look very different from those for YCS9. They display strong bi-modality and there is evidence of significant earnings differential in favour of the older age group in both sweeps. That said, we see little evidence that this differential increased substantially between sweep 1 and sweep 2, particularly in the lower half of the distribution.

6. National minimum wage and enrolment in full-time education

The focus of our empirical analysis is the effects of eligibility for the national minimum wage on enrolment in full-time post-compulsory schooling[6]. The raw data on enrolment rates at sweep 1 and sweep 2 by eligibility for the national minimum wage are reported in Table 3. These raw differences suggest a small negative impact on participation in full-time education with the enrolment rate for the treated group falling by 11.88 percentage points between sweeps 1 and 2, compared with a decline of just 10.73 percentage points for the younger control group. However, these differences evaporate if we control for the timing of the return of the sweep 2 survey. In local education authorities in England and Wales, the school year formally ends in the third week of July. For some of those who returned the sweep 2 survey in July 1999, their school year had effectively ended, and as a consequence, they are less likely to report participation in full-time education as their main activity. At the same time, those returning the sweep 2 survey late are more likely to have reached the age of 18 years by

[5] The introduction of General National Vocational qualifications (GNVQ) based on two years of full-time study were introduced nationally in 1993. The Modern Apprenticeship scheme was established in 1995, replacing previous government subsidized training provision such as theYouth Training Scheme.

[6] Enrolled and attending as a full-time student in a school or college of further education in the state-maintained or independent sector.

the date of the survey return. To eliminate this source of spurious correlation between eligibility for the national minimum wage and enrolment in full-time education, we include dummy variables for the month in which the survey is returned at each sweep, as well as a dummy variable for the sweep itself, when estimating the conditional logit model (2).

Of the sample of 8,823 young persons, 1,214 changed school enrolment status between sweep 1 and sweep 2 of the survey and these provide the estimation sample for the conditional logit model (2). The estimated coefficients of the logit model, together with their standard errors clustered at the local area level, are reported in Table 4[7]. Column 1 of table 4 shows the results for the most basic specification of the logit model, with unobserved time-invariant individual-specific effects, a dummy variable for the survey sweep, dummy variables for the month of survey return, and the treatment variable, NMW_{ij}. In this case, the estimated effect of the national minimum wage on the probability of enrolling in full-time schooling is small, positive and statistically insignificant. The results are largely unchanged when we include additional controls for conditions in the local area labour market including measures of youth unemployment and the sectoral composition of local employment.

Up until this point, the response of young people to eligibility for the national minimum wage is assumed to be homogenous. However, the 'bite' of the national minimum wage – the extent to which it affected the distribution of earnings in a local area – varies considerably across the UK as shown in Stewart (2002). At its introduction in April 1999, the national minimum wage was set at £3.60 per hour (£3 per hour for workers aged 18 to 21years). As a proportion of hourly earnings at the lowest quintile of the distribution in each of the 171 local authority areas of England and Wales, the minimum wage varied between 86 percent and 37 percent (44 percent if the City of London is excluded). Given this

[7] Estimated by pseudo-maximum likelihood using the clogit procedures with sampling weights in Stata 10.

considerable spatial variation, we would expect that the effect of the minimum wage on young people's behaviour to vary spatially also.

To capture potential spatial variation in the treatment effect, we interact eligibility for the national minimum wage with the local 'bite' of the national minimum as measured by the ratio of the value of the national minimum wage to the value of hourly earnings at the lowest quintile of the distribution for full-time workers in the local area in 1998[8]. The results are reported in the third column of table 4. Allowing for this form of heterogeneity, we find that eligibility for the national minimum wage has a well-determined, statistically significant effect on the probability of enrolment in schooling. The effect is negative for young people residing in areas where the local 'bite' of the minimum wage is large, and it increases as the local 'bite' declines, becoming positive for those living in areas where hourly earnings are relatively high.

The possibility that eligibility for the national minimum wage is confounded with other factors related to the individual's relative age within the school-year cohort remains a concern. The analysis to date captures the potential effects of the individual's relative age through a time-invariant individual-specific effect. However, it is possible that relative age effects are not time-invariant but rather cause the probability of enrolment in schooling to evolve differently over time. For example, individuals who are more mature may be less likely to drop-out of schooling. To allow for this, we include a low-order polynomial of the individual's relative age interacted with the sweep 2 dummy variable in the logit model. Column 4 of table 4, reports the results obtained with a quadratic function of relative age, while column 5 shows the results for a linear spline function with a knot at the equivalent of

[8] For England and Wales as a whole, the 10th percentile of the hourly earnings distribution in Spring 1998 was £4.60 and the 20th percentile was £5.53, compared with a national minimum wage of £3.60. The 20th percentile is used in preference to the 10th percentile at the local area level because the sample estimates are more reliable and there are no missing values.

age 18 years at sweep 2. The estimates of the effects of eligibility for the national minimum wage are largely unaffected by the inclusion of these additional terms in the relative age of the individual.

As a final exercise, we consider the possibility that other time-invariant characteristics may cause the probability of enrolment in schooling to evolve differently over the two sweeps of the survey. Young persons with higher levels of prior academic attainment or with more highly-educated parents may be more likely to remain enrolled in post-compulsory schooling. To control for these potential effects, we include a number of individual characteristics interacted with the sweep 2 dummy variable. The characteristics considered are: the individual's age relative to the cohort average; gender; ethnicity; parent's education; and the number of higher grade (grades A* to C) GCSEs achieved by the completion of compulsory schooling. As can be seen in columns 6 and 7 of table 4, our results with respect to the treatment effects of eligibility for the national minimum are robust to the inclusion of these additional controls. In this case, the absolute magnitudes of the coefficients increase somewhat and the corresponding standard errors tend to increase also, but estimated coefficients for the treatment effect remain statistically significant at the 5 percent level.

Figure 3 shows the estimated logit coefficients, $\hat{\theta}(E_a^{LQ}) = \hat{\theta}_1 + \hat{\theta}_2 . \overline{W}_{\min} / E_{a,}^{LQ}$, for different values of lowest quintile earnings E_a^{LQ}, using the parameter estimates reported in column 6 of table 4. Also depicted are the 90 and 95 percent confidence interval bands for the estimates. Eligibility for the national minimum wage reduces the probability of enrolment in schooling in local authority areas with lowest quintile earnings of less than £5.60 per hour. However, the confidence intervals of the estimates are relatively wide, and the effects are significantly negative at the 10 percent level only for areas where lowest quintile earnings are below £4.10 per hour. Of the 171 local authority areas in England and

Wales, 106 had lowest quintile earnings of less that £5.60 in 1998, prior to the introduction of the national minimum, only two had lowest quintile earnings of less than £4.10.

As an alternative to interacting eligibility for the national minimum wage with a continuous measure of the minimum wage's 'bite' into the local area's earnings distribution as in table 4, we could allow the effects to vary with local area earnings according to a step function. Table A4 of the appendix shows a set of results using this type of specification and the pattern of treatment effects is consistent with that depicted in figure 3. For young people residing in local authorities in the lower tail of the distribution with respect to lowest quintile hourly earnings, eligibility for the NMW has a significant negative effect on enrolment in full-time education. For those residing in local authorities in the middle range of the distribution, eligibility for the NMW has no discernible effect; while for those residing in local authorities in the upper tail of the distribution, the estimated effects are generally positive, but small and not statistically significant.

As already noted, it is not possible in a conditional logit framework to calculate the average effect of the treatment on the probability of enrolling in full-time schooling across the distribution of time-invariant individual-specific effects in the population. However, we can assess the effect of the treatment on the probability of an individual quitting full-time schooling between sweeps 1 and 2 (conditional on a change in enrolment status) since this does not depend on the unobserved individual-specific effects (see equation (2)). In the upper part of figure 4, we show the estimated effect of the national minimum wage on the probability of quitting full-time schooling for a representative 18 year old.[9] (The broken lines

[9] Representative individual is a 18 year-old white male with 5 GCSE qualifications grades A* to C; mother educated to the level of A-levels or equivalent; resident in the East Midlands region; no change in local labour market conditions between sweeps.

depict the corresponding 90 percent confidence bands for the estimate.[10]) In this particular example, eligibility for the national minimum wage, more than doubles the probability of dropping out of schooling in the case of 18 years-olds living in local areas where the lowest quintile earnings are below £4.40. For this exercise, we assume that there is no change in local labour market conditions between the two sweeps in order to highlight the relationship between the magnitude of the treatment effect - as measured by the difference in the two probabilities – and hourly earnings in the local area. Using the actual sample data, as in the lower part of figure 4, produces 'noise' about the underlying negative relationship between the magnitude of the average treatment effect and the level of earnings in the local area.

7. Robustness

Before concluding, we review the results of a number of estimation exercises undertaken to investigate the robustness of our finding. The first of these involves narrowing the age range of young people included in our estimation sample. The full sample is unbalanced in the sense that it includes individuals within -3 months and +9 months of their 18th birthday at sweep 2. Columns 2 to 4 of table 6 show the results obtained if the age interval around the 18th birthday is narrowed to provide a more balanced sample. For ease of comparison, column 1 of table 6 reports the results obtained with the full sample (i.e as in column 5 of table 4). As one would expect, the smaller samples result in larger standard errors, but generally speaking, our findings are robust to narrowing the age interval around the 18th birthday.

The second exercise is to re-estimate the model including the observations on those whose estimated age at the time the sweep 2 survey return is equal to 216 months. Up until this point, this group has been excluded from the analysis on the grounds that we are unable

[10] The confidence limits are computed using the 'delta' method and consequently the values are not restricted to lie in the interval [0,1].

to distinguish a priori between those members who are covered by the legislation and those who are not. An alternative approach is to include them in the estimation sample with NMW_i set equal to 0.5, the average treatment for members of this group on the assumption that birth dates and survey return dates are uniformly distributed across the month. The results of this exercise for the quadratic specification are reported in column 5 of table 6 and are very similar to those for the sample with the 216 months age group excluded.

The question remains whether our results are being driven by the effects on enrolment in schooling of attaining the age of 18 years per se, rather than a consequence of becoming eligible to receive the national minimum wage. To investigate this further, we re-estimate the model as specified in table 4, but with the dependent variable defined in terms of the individual's enrolment status in the previous November, rather than at the time the survey was completed. To be precise, the dependent variable is now S_{ij}^N =1 if i is enrolled in full-time education in the November prior to sweep j, and S_{ij}^N =0 otherwise. Recall that the national minimum wage legislation did not come into effect until 1st April 1999, and hence the minimum wage was not in place in either November 1997 or November 1998. We construct a new treatment variable $Nov18_i$ based on the individual's age in the November prior to sweep 2 of the survey (i.e. November 1998). For those whose estimated age at November 1998 is 217 months or more, $Nov18_i = 1$; for those aged 215 months or less at November 1998, $Nov18_i = 0$. If our results are being driven by factors related to being 18 years-old, rather than a consequence of eligibility for the national minimum wage, then we should expect to find a similar pattern of treatment effects in the new specification. A can be seen in table 7, this is evidently not the case. The estimated coefficients associated with the treatment effect in this case are much smaller in magnitude and very poorly determined.

8. Concluding remarks

This paper uses the introduction of a national minimum wage in the UK in April 1999 as a 'quasi experiment' to assess the impact of statutory minimum wages on participation in further education by young people. This empirical analysis has a number of distinct advantages over many previous studies of the effects of minimum wages. First, we are able to exploit the fact that assignment to the treatment is determined by being either side of a fixed age threshold. Second, our panel data allows us to control for unobserved time-invariant heterogeneity using a difference-in-differences type estimator.

Across the cohort as a whole, the average effect of the national minimum wage on enrolment in post-compulsory schooling appears to be negligible. However, for young people living in areas where earnings are relatively low, eligibility for the national minimum wage is associated with a significant reduction in the probability of enrolling in full-time schooling. It may seem surprising that a relatively modest increase in their average expected earnings from employment should have a significant impact on a young person's decision to invest in human capital. However in low-paid local labour markets areas, the introduction of the national minimum wage significantly compressed differentials between unskilled and skilled workers, reducing the expected returns on human capital investment. This, as much as any short-term increase in the income, lies behind the observed reduction in enrolment rates in further education.

Finally, our findings identify a short-term negative effect on enrolment in post-compulsory schooling from minimum wages. The analysis in this paper is unable to assess the long-term consequences of minimum wages for this cohort of young people. It may be - as Neumark and Waschter argue – that lower rates of school enrolment and reduced skill acquisition have long-lasting consequences for the affected individuals. It may be that over the longer term, minimum wages lead firms to upgrade the general skills of its workers and

the quality of the jobs that they offer as suggested by Acemoglu and Pischke (2003). The robust identification of the effects of statutory minimum wages on labour market outcomes over the longer term remains the challenge for future empirical work.

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Table 1: Sample characteristics

	Group 1: Eligible for NMW Age ≥ 217 months at sweep 2	**Group 2: Not eligible for NMW** Age ≤ 215 months at sweep 2
Sample size	6988	1857
Average age at sweep 2 (months)	220.88	214.31
	(2.64)	(0.74)
Gender (% female)	48.94	52.86
Ethnicity (% white)	87.98	88.79
Parent's education (% at least one parent has a degree level qualification	21.66	22.77
Type of school attended (% private)	7.39	6.62
Average number of GCSEs grade A-C at completion of compulsory schooling	4.53	4.18
	(3.85)	(3.91)
Local area: average unemployment rate, Spring 1998	3.74	3.62
	(1.71)	(1.65)
Local area: average hourly earnings of full-time workers, Spring 1998	8.06	8.02
	(1.01)	(1.02)

* Summary statistics based on weighted data

Table 2: The Introduction of the NMW and the Earnings of Young Workers

	Group 1: Eligible for NMW		**Group 2: Not eligible for NMW**	
	Sweep 1: April-June 1998	**Sweep 2: May-July 1999**	**Sweep 1: April-June 1998**	**Sweep 2: May-July 1999**
Number reporting hourly take-home pay in sweeps 1 & 2	2647	2647	672	672
Lowest decile	1.33	2.42	1.35	2
Lower quartile	2.10	3	2	2.73
Median	2.78	3.6	2.78	3.25
Upper quartile	3.33	4.17	3.4	4
Highest decile	4	5	4.1	5

* Summary statistics based on weighted data

Table 3 : The Introduction of the NMW and Enrolment in Full-time Education

	Group 1: Eligible for NMW Age ≥ 217 mths at sweep 2	**Group 2: Not eligible for NMW** Age ≤ 215 months at sweep 2
Sample size	6970	1853
Sweep 2- enrolment in full time education (%)	57.53	61.03
Sweep 1- enrolment in full time education (%)	69.41	71.76
	-11.88	-10.73
May 1999 returns - sample size	2369	1002
Sweep 2- enrolment in full time education (%)	67.54	66.19
Sweep 1- enrolment in full time education (%)	74.76	73.70
	-7.22	-7.51
June 1999 returns - sample size	2790	656
Sweep 2- enrolment in full time education (%)	56.68	57.16
Sweep 1- enrolment in full time education (%)	67.67	68.56
	-10.99	-11.40
July 1999 returns - sample size	1811	195
Sweep 2- enrolment in full time education (%)	47.53	51.35
Sweep 1- enrolment in full time education (%)	69.39	73.87
	-21.86	-22.52

Table 4: The Effects of Eligibility for NMW on Enrolment in Full-time Education

Dependent variable: Individual currently enrolled in full-time education (s_{ij})

	1	2	3	4	5	6	7
NMW_i	0.2615 (0.2698)	0.2742 (0.2647)	2.9763** (1.4573)	2.9284** (1.3166)	2.9763** (1.4573)	3.3858** (1.4641)	3.2164** (1.6060)
$[W_{min}/E^{LQ}_{a,t-1}]$ * NMW_i			-4.2298** (1.7881)	-4.2106** (1.7584)	-4.2298** (1.7881)	-5.2320** (2.0497)	-5.3023** (2.0356)
$[W_{min}/E^{LQ}_{a,t-1}]$			-1.8794 (4.1391)	-2.0915 (4.1504)	-1.8794 (4.1391)	-7.0056 (4.9933)	-6.7010 (5.0171)
Sweep	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Month of survey return	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Individual fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Local labour market conditions (time-varying)		Yes	Yes	Yes	Yes	Yes	Yes
Birth date *sweep				quadratic	linear spline	quadratic	linear spline
Time invariant controls*sweep						Yes	Yes
Pseudo R-squared	0.4495	0.4564	0.4615	0.4651	0.4646	0.4960	0.4961
Log pseudo-likelihood	-531.91	-522.13	-517.19	-513.69	-514.22	-484.10	-483.94

Notes: There are 1220 observations at each sweep. Robust standard errors are reported in parentheses, clustered at the local authority area level throughout.
****** denotes significance of coefficient at the 5% level and * denotes significance at the 10% level.
Labour market conditions: number of unemployed aged less than 18 years in local authority area; proportion of local authority employment in the service sector and in public administration.
Time invariant controls: gender; ethnicity; parent's education; number of GCSE qualification grades A-C gained by end of compulsory schooling; region of residence.
$W_{min}/E^{LQ}_{a,j-1}$ is the ratio of the value of the national minimum wage to the lowest quintile of the distribution of hourly earnings for full-time workers in the local authority in year prior to sweep.

Table 5: The Effects of Eligibility for NMW on Enrolment in Full-time Education: Robustness checks

Dependent variable: Individual currently enrolled in full-time education (s_{ij})

	1 **Full sample** **3/+9 mths**	**2** **-3/+5 mths**	**3** **-3/+4 mths**	**4** **-3/+3 mths**	**5** **Including those aged 216 mths**
NMW_i	3.3858** (1.4641)	2.9516* (1.8378)	2.6814 (1.8904)	3.6449* (2.2666)	3.4539** (1.5091)
$NMW_i * W_{\min} / E^{LQ}_{a,j-1}$	-5.2320** (2.0497)	-4.7617** (2.3630)	-4.7592* (2.6562)	-6.4547** (3.3005)	-5.3586** (2.1001)
$W_{\min} / E^{LQ}_{a,j-1}$	-7.0056 (4.9933)	-5.5680 (6.1751)	-5.5149 (6.7013)	-7.4548 (6.9085)	-7.3167 (4.8836)
Sweep	Yes	Yes	Yes	Yes	Yes
Month of survey return (dv)	Yes	Yes	Yes	Yes	Yes
Individual fixed effects	Yes	Yes	Yes	Yes	Yes
Local labour market conditions (time-varying)	Yes	Yes	Yes	Yes	Yes
Birth date *sweep	quadratic	quadratic	quadratic	quadratic	quadratic
Time-invariant controls*sweep	Yes	Yes	Yes	Yes	Yes
No. of observations in each sweep	1220	780	663	558	1318
Pseudo R-squared	0.4960	0.4789	0.4757	0.5040	0.5091

Notes: see notes to table 4

Table 6: The Effects of Eligibility for NMW on Enrolment in Full-time Education – Falsification Check

Dependent variable: Individual enrolled in full-time education in November prior to survey (s_{ij}^{N})

	1	2	3	4	5	6	7
Nov18$_i$	-0.3300 (0.2904)	-0.3514 (0.2958)	0.5805 (2.1861)	0.4449 (2.3513)	0.5925 (2.1402)	0.7687 (2.6958)	0.5093 (2.4787)
$[W_{\min}/E_{a,t-1}^{LQ}]$ * *Nov18*$_i$			-1.4442 (3.3837)	-1.4060 93.3569)	-1.4676 (3.3434)	-1.4580 (3.9330)	-1.4820 (3.9225)
$[W_{\min}/E_{a,t-1}^{LQ}]$			1.4723 (5.3461)	1.7638 (5.3578)	1.4305 (5.3578)	1.2637 (5.8418)	1.2187 (5.8779)
Sweep	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Month of survey return	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Individual fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Local labour market conditions (time-varying)		Yes	Yes	Yes	Yes	Yes	Yes
Birth date *sweep				quadratic	linear spline	quadratic	linear spline
Time invariant controls*sweep						Yes	Yes
Pseudo R-squared	0.4878	0.4945	0.4946	0.4953	0.4946	0.5232	0.5231
Log pseudo-likelihood	-448.19	-441.35	-441.25	-440.67	-441.21	-416.29	-416.36

Notes: There are 883 observations at each sweep. Robust standard errors are reported in parentheses, clustered at the local authority area level throughout. ****** denotes significance of coefficient at the 5% level and * denotes significance at the 10% level.
Time-varying controls: number of unemployed aged less than 18 years in local authority area; proportion of local authority employment in the service sector and in public administration. **Time invariant controls**: gender; ethnicity; parent's education; number of GCSE qualification grades A-C gained by end of compulsory schooling; region of residence.
$W_{\min}/E_{a,j-1}^{LQ}$ is the ratio of the value of the national minimum wage to the lowest quintile of the distribution of hourly earnings for full-time workers in the local authority in year prior to sweep j

Figure 1: Impact of the Introduction of the NMW on Earnings of Young Workers – Kernel Density Estimates of (Ln) Hourly Earnings

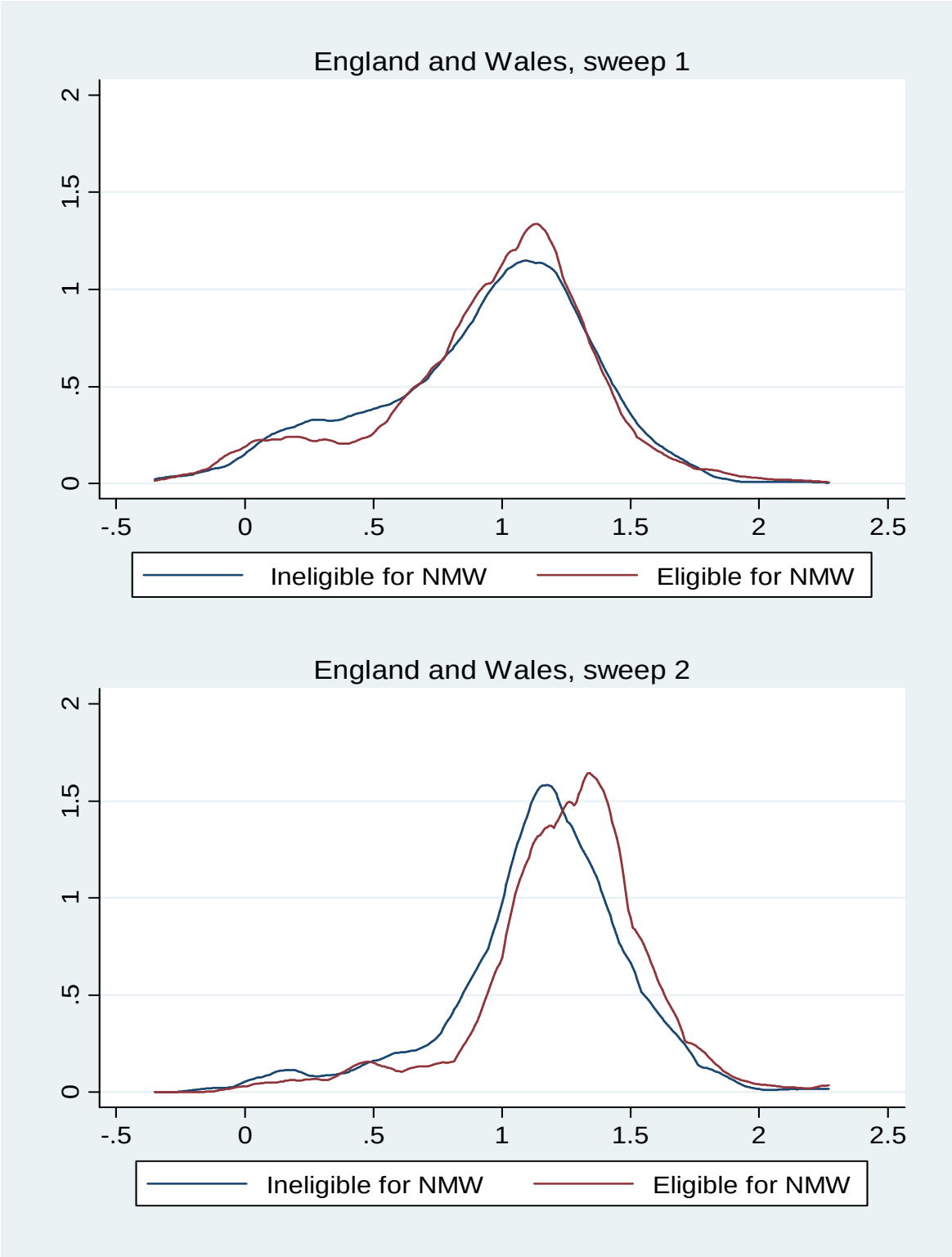


Figure 2: Marginal Effect of Eligibility for the National Minimum Wage (logit coefficients).

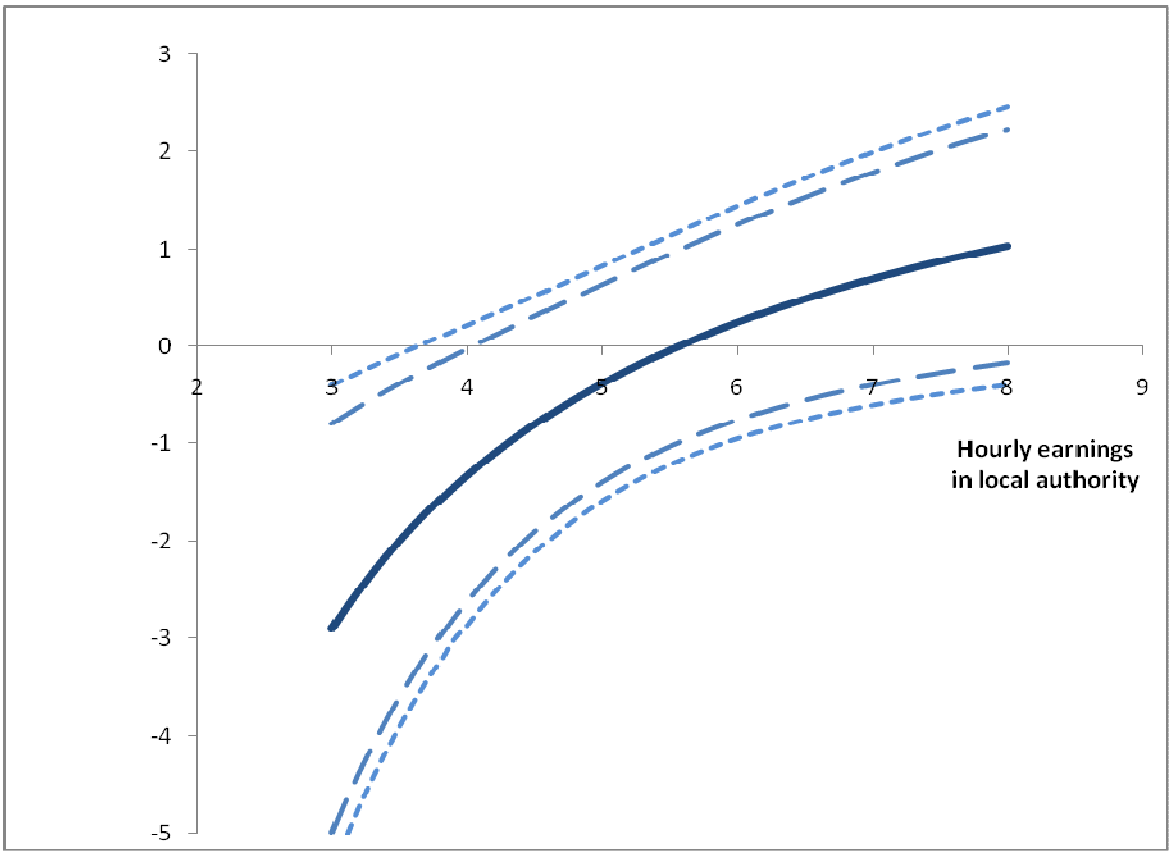
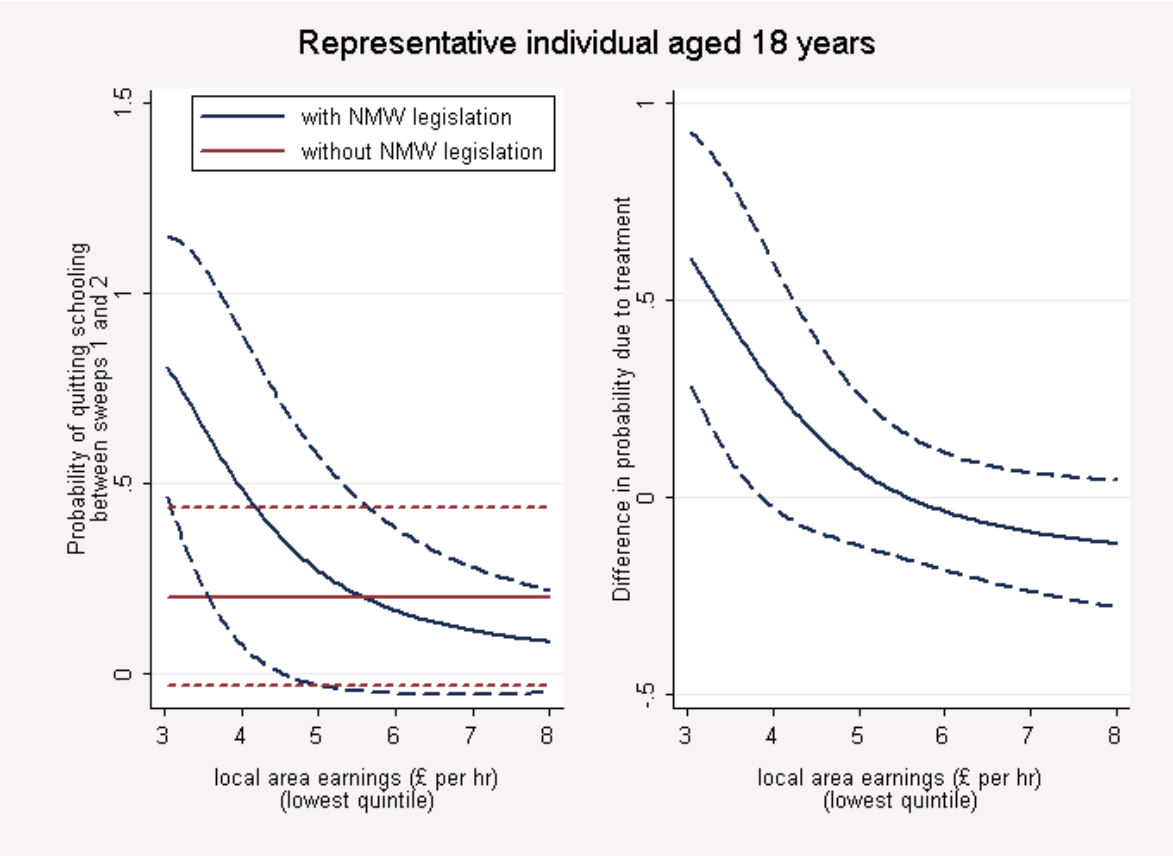
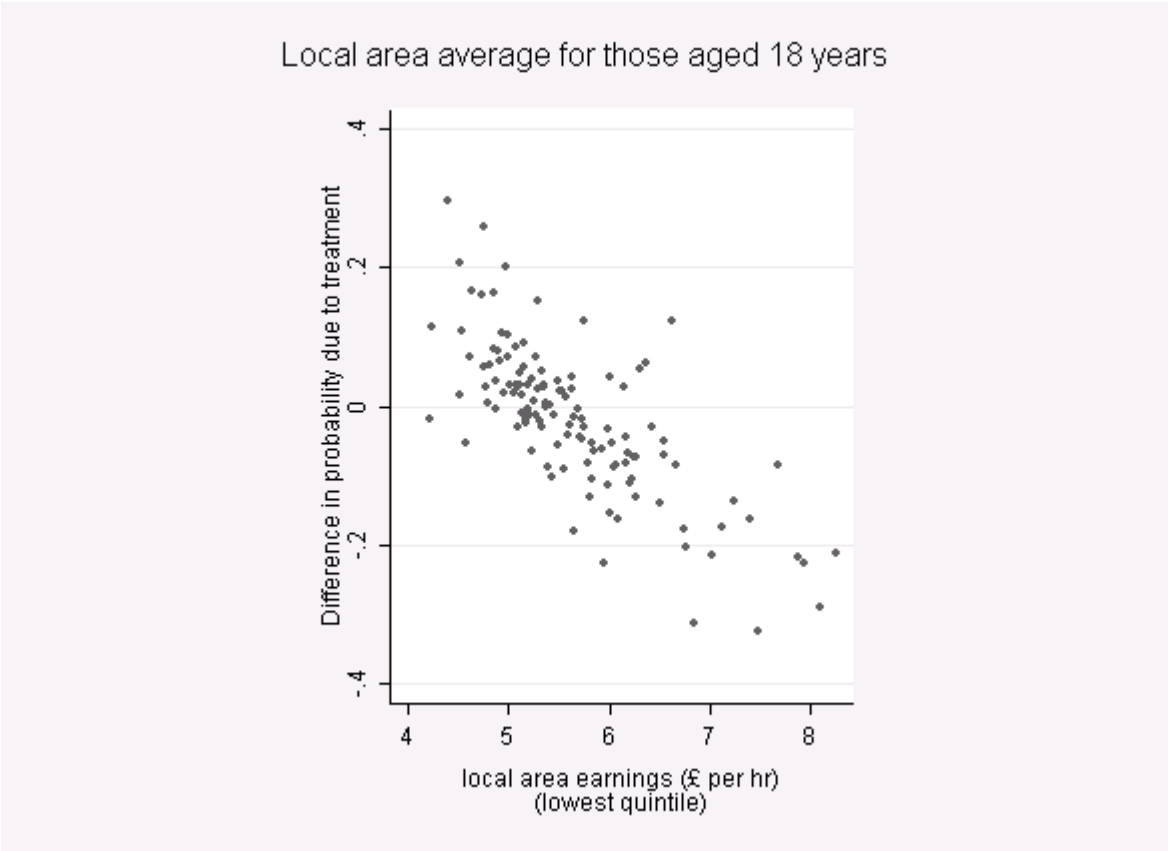


Figure 3: Impact of Eligibility for the National Minimum Wage.

Estimated effect on the probability of quitting full-time schooling for those aged 18 years.





Appendix

Table A1: Age Composition of YC9 sample

Estimated age at sweep 2 in months	Number in sample	Estimated age at sweep 2 in months	Number in sample
213	344	**220**	786
214	662	**221**	747
215	851	**222**	736
216	817	**223**	789
217	815	**224**	742
218	843	**225**	539
219	775	**226**	196

Table A2: Impact of the Introduction of the NMW on Earnings of Young Workers (excluding those on government supported training schemes)

	Group 1: Eligible for NMW		Group 2: Not eligible for NMW	
	Sweep 1: April-June 1998	Sweep 2: May-July 1999	Sweep 1: April-June 1998	Sweep 2: May-July 1999
Number reporting hourly take-home pay in sweeps 1 & 2	2337	2337	586	586
Lowest decile	2	2.75	1.96	2.42
Lower quartile	2.5	3.17	2.5	2.95
Median	3	3.75	3	3.33
Upper quartile	3.5	4.33	3.59	4
Highest decile	4.14	5	4.28	5

* Summary statistics based on weighted data

Table A3: Earnings Distribution of Young Workers, Youth Cohort 6 (216)

	Group 1: Aged 18 years at sweep 2		Group 2: Aged 17 years at sweep 2	
	Sweep 1: April- June 1992	**Sweep 2: April-June 1993**	**Sweep 1: April- June 1992**	**Sweep 2: April-June 1993**
Number reporting hourly take-home pay in sweeps 1 & 2	2161	2161	1266	1266
Lowest decile	0.88	0.96	0.78	0.90
Lower quartile	1.13	1.5	0.99	1.28
Median	1.97	2.41	1.75	2.13
Upper quartile	2.5	3	2.25	2.7
Highest	2.96	3.51	2.77	3.18

* Summary statistics based on weighted data

Figure A1: Impact of the Introduction of the NMW on Earnings of Young Workers (excluding those on Government Supported Training schemes) – Kernel Density Estimates of (Ln) Hourly Earnings

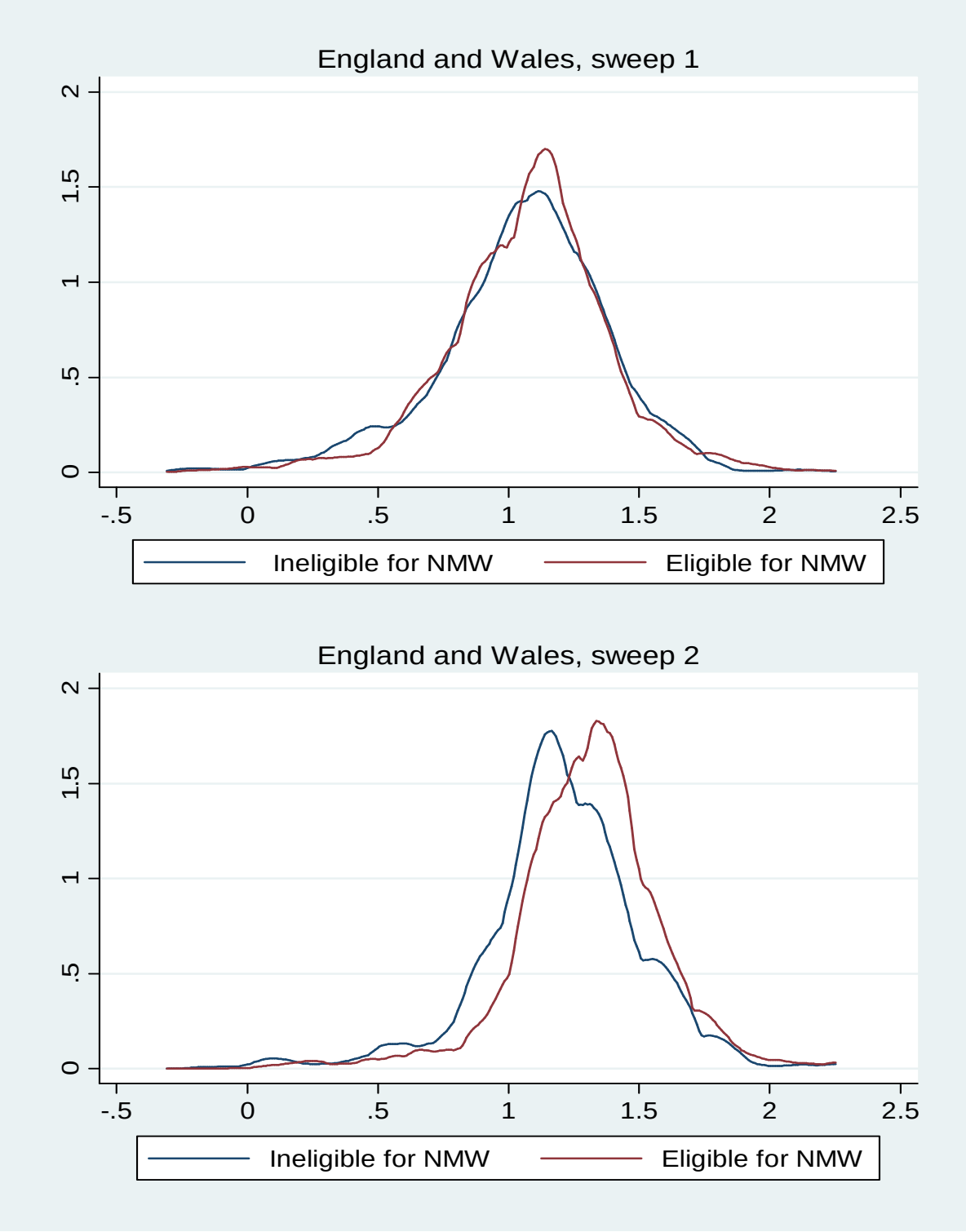


Figure A2: Kernel Density Estimates of (Ln) Hourly Earnings for YC6 (216)

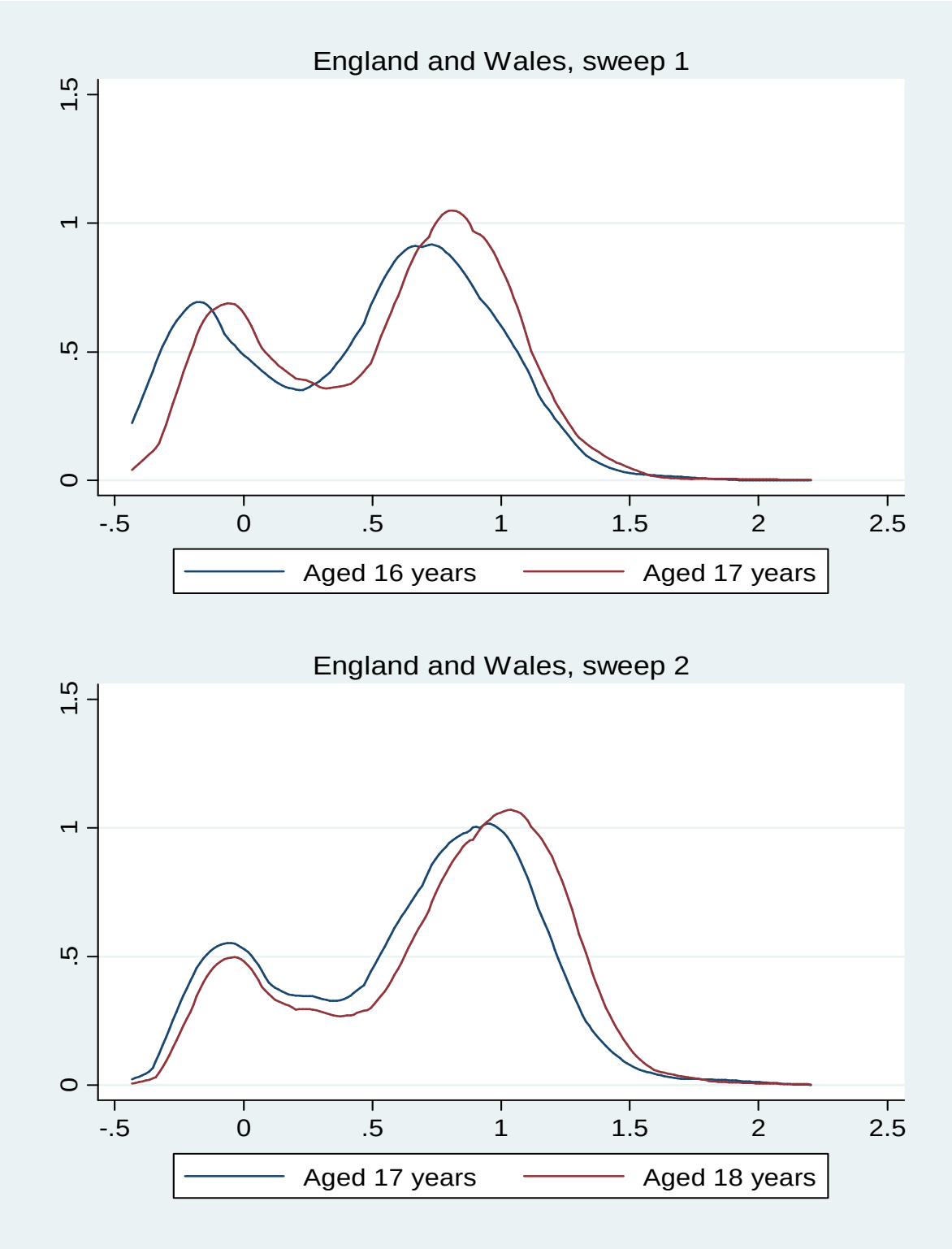


Table A4: The Effects of Eligibility for NMW on Enrolment in Full-time Education – Step Functions

The intervals of the step function are defined such that λ percent of school-year cohort reside in local authorities with lowest quintile earnings of less than E_a^L; and of greater than E_a^U

Dependent variable: Individual currently enrolled in full-time education (s_{ij})

	$\lambda = 5$	$\lambda = 10$	$\lambda = 15$	$\lambda = 20$	$\lambda = 25$
	$E_a^L = 4.76; E_a^U = 6.54$	$E_a^L = 4.87; E_a^U = 6.25$	$E_a^L = 4.96; E_a^U = 6.09$	$E_a^L = 5.03; E_a^U = 5.82$	$E_a^L = 5.10; E_a^U = 5.72$
NMW_i $(E_{a,98}^{LQ} \le E_a^L)$	-1.1316*	-0.9323	-1.0677*	-0.4199	-0.4005
	(0.6398)	(0.6441)	(0.6120)	(0.5906)	(0.5580)
NMW_i $(E_a^L < E_{a,98}^{LQ} \le E_a^U)$	0.0191	0.0091	-0.0475	-0.0284	0.0937
	(0.5883)	(0.5895)	(0.6024)	(0.5689)	(0.5606)
NMW_i $(E_{a,98}^{LQ} > E_a^U)$	0.0316	0.2698	0.5296	0.3713	0.1176
	(0.6770)	(0.6760)	(0.6528)	(0.7010)	(0.6937)
Pseudo R-squared	0.4934	0.4952	0.4984	0.4939	0.4936
Log pseudo-likelihood	-486.51	-484.79	-481.78	-486.02	-486.31
NMW_i $(E_{a,98}^{LQ} \le E_a^L)$	-1.1504**	-0.9410**	-1.0222**	-0.3937	-0.4881
	(0.4143)	(0.4692)	(0.3886)	(0.3502)	(0.3127)
NMW_i $(E_a^L < E_{a,98}^{LQ} \le E_a^U)$	-	-	-	-	-
NMW_i $(E_{a,98}^{LQ} > E_a^U)$	0.0493	0.2618	0.5730	0.3973	0.0322
	(0.4732)	(0.4954)	(0.4671)	(0.4336)	(0.4181)
Pseudo R-squared	0.4934	0.4952	0.4984	0.4939	0.4936
Log pseudo-likelihood	-486.51	-484.79	-481.79	-486.02	-486.33

Notes: There are 1220 observations at each sweep. Robust standard errors are reported in parentheses, clustered at the local authority area level throughout. ****** denotes significance of coefficient at the 5% level and * denotes significance at the 10% level. **Additional controls:** as in column 5 of table 4.

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