

THE NATIONAL MINIMUM WAGE AND THE SUBSTITUTABILITY BETWEEN YOUNG AND OLD WORKERS IN LOW PAID OCCUPATIONS*

by

GAUTHIER LANOT

Umeå School of Business and Economics, Economics, Umeå University
and

PANOS SOUSOUNIS[†]

Keele Management School, Economics, Keele University

We study the effect of the National Minimum Wage (NMW) on the workforce age composition within the low paying sectors of the British economy. Our interest is in the degree of substitutability between labour inputs (young and old employees). We find evidence that both the introduction and the regular upratings of the NMW have a significant effect on the observed changes to the relative wages and to the relative wage bills but not to relative employment. We estimate the elasticity of substitution, between ‘young’ age groups and older workers (55+) to be zero, while that of ‘prime’ (22+) age and older workers, to be around 0.79. Our estimates therefore imply significant complementarity between younger and old employees.

1 INTRODUCTION

Since its introduction in April 1999, the UK National Minimum Wage (NMW) has set distinct rates for different age groups: at first covering separately workers aged between 18 and 22 with the Youth Development Rate (YDR), and workers older than 22 with the Adult Rate. Two sub-minimum rates were subsequently introduced in 2004 and 2010 for workers between 16 and 17 and for apprentices, respectively, and since October 2010 the adult rate applies to all workers aged 21 or above. In April 2016, the UK government has introduced a ‘National Living Wage’ (NLW) set at £7.20 (with the view of increasing it to £9 by 2020, the equivalent to 60 per cent of median earnings given current growth forecasts) which applies to all workers aged 25 or over, while the NMW rates set in October 2015 apply for younger workers (£6.70 for workers aged 21–24, £5.30 for the YDR, and £3.87 for workers between 16 and 17).

* Manuscript received 14.11.14; final version received 27.5.16.

[†] We would like to thank the editors and four anonymous referees for constructive comments, and Tim Butcher and the participants of the UK Low Pay Commission 2013 workshops for helpful comments and suggestions during the early stages of the manuscript. Financial support from the UK Low Pay Commission is gratefully acknowledged.

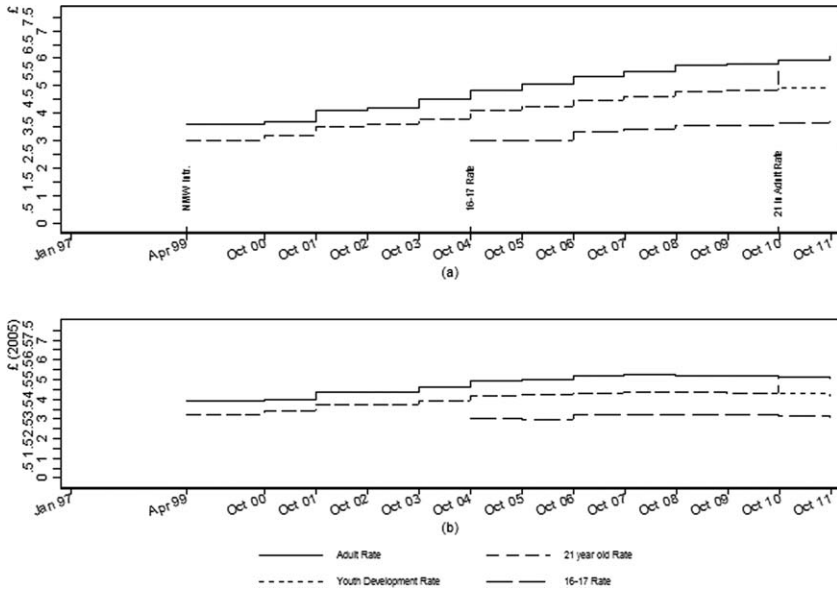


FIG. 1. UK National Minimum Wage, Nominal and Real value, 1998 to 2011.

Note: Panel (a) describes the nominal value of the UK National Minimum Wage Rates, Pound Sterling. Panel (b) presents the value of the UK NMW deflated by the Consumer Price Index (100 in 2005).

Our aim is to measure the elasticity of substitution between ‘young’ and old workers, which rationalizes the effect of the introduction and the regular up rating of the different rates of the NMW on the composition of the workforce. Panel (a) of Fig. 1 documents the evolution of the age contingent rates over the sample period as well as their distinct levels in nominal terms (from £3.60 in April 1999 to £5.93 in October 2010 for the Adult Rate, and from £3 to £4.92 for the YDR, while the rate for workers between 16 and 17 ranges from £3 in October 2004 to £3.64 in October 2010). We observe that these increase steadily over the sample period. The lower part of the figure documents the evolution of the same rates in real terms (deflated by the consumer price index, CPI, for the UK, CPI = 100 in 2005), and shows that the real value of most rates peaked in 2008 and decreased slowly thereafter.

While the consequences of operating minimum wage laws on overall employment are debated on both theoretical and empirical grounds, (Hamermesh, 1993; Card and Krueger, 1997; Manning, 2005; Metcalf, 2006; Neumark and Wascher, 2008) the arguments in favour or against a minimum wage are well understood. (Neumark and Wascher, 2008, p. 95) claim that “the evidence points toward disemployment effects”, but Butcher (2012) reaches the opposing conclusion that minimum wages have not had a significant adverse effect on employment (see Dolton *et al.*, (2015) for another

recent contribution). Other studies (see Holmlund, 2014; Metcalf, 2006) find no or even positive employment effects. Stewart (2002) reports that while the NMW has affected low-paid workers' wage distribution, it has had insignificant area- and individual-level effects on employment. Similarly, Stewart (2004a, 2004b) finds no evidence of negative effects on employment. Stewart and Swaffield (2006) find that adverse employment effects of the NMW are discernible in the (basic) hours of work and at a lag. The authors report a 1.5 hours reduction in weekly basic hours among men, and a smaller 0.6 hours reduction for women. Low Pay Commission (2014) find a 3–4 hours reduction among young workers (18–20 year olds), noting that their estimates are based on small samples, but find no systematic effect among adults.

Assuming that all types of labour are perfect substitutes (the homogeneity hypothesis), the theoretical prediction of the standard (competitive) model of the labour market in the presence of a minimum wage floor is clear: if the minimum wage is set above the market clearing wage, employment declines. The magnitude of the decrease in employment is determined by the distance between the competitive wage and the legal minimum and the elasticity of the demand for labour. Explicitly rejecting the homogeneity hypothesis, thus assuming that there are substantial complementarities in production between age groups, the first UK Low Pay Commission (1998) report suggests that the design of the distinct NMW rates was based on the comparison of the average productivity of low skill workers, hence, the relative NMW rate between two age groups reflects the differences in the marginal product of labour between the two groups.

Over time, as we show in Fig. 2, the relative minimum wages (the ratio of the minimum wage rates for the younger age groups relative to the adult rate) were kept relatively constant. The most significant changes are associated with the introduction of a new rate (e.g. in 2004 for the 16–17 rate) or a modification of the definition of the groups themselves (as happens later in 2010 when the 21 year olds became part of the adult group). However, the requirement to satisfy the minimum wage regulation does not bind the relative market wages above or below a given value.¹ While the minimum wage differentials are kept relatively constant, market wages for each labour type can still change in response to demand or

¹Consider two types of labour, young and old, and their wages, w_y and w_o , respectively, while \underline{w} denotes the minimum wage for the age group. Requiring that $w_y > \underline{w}_y$ and $w_o > \underline{w}_o$, is obviously equivalent to $\frac{w_y}{w_o} > \frac{\underline{w}_y}{\underline{w}_o}$ and $w_o > \underline{w}_o$ for positive wages. Wages for the different labour types can satisfy the constraints while their ratio can be either greater or smaller than the ratio of the minimum wages $\frac{\underline{w}_y}{\underline{w}_o}$. Imagine for example that the minimum wages for both type of labour are identical, the market wages are both greater than the minimum wage, but w_o is twice w_y . In that case, the ratio of market wages is equal to 1/2 while the ratio of minimum wages is equal to 1.

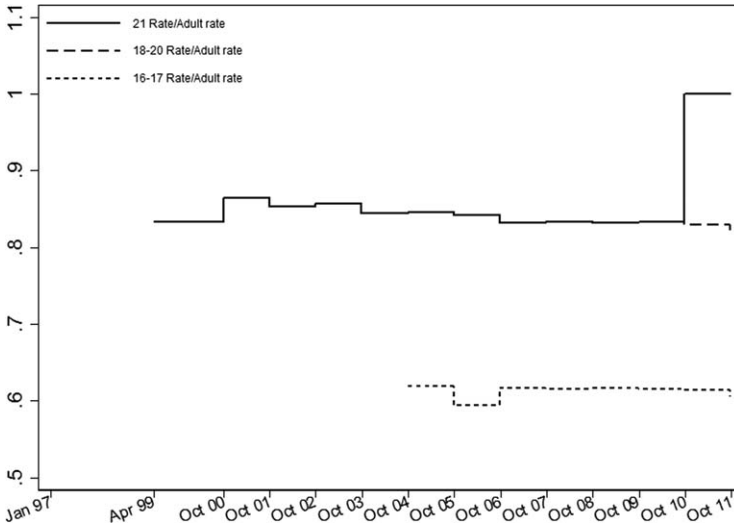


FIG. 2. Relative Minimum Wage Rates, 1998–2011.

Note: UK, National Minimum Wage Rates relative to the Adult Rate.

supply shocks, or in response to changes in the value of the NMW rates, so that the market wage differentials affect the age composition of labour demand.

If the hypothesis of homogeneity of labour is rejected, wage differentials between age groups, and the effects of the NMW rates on the wage differentials will determine the composition of the employed workforce. Metcalf (2006), in his review of the effect of the minimum wage, suggests that the homogeneity hypothesis is questionable in principle. However, studies which look directly at the pattern of substitution between young (younger than 22) and old (older than 55) labour are rare. The OECD (OECD, 2013) fails to find any substantial reduction in the employment of young workers when the employment of the old increases (in this case following the 2008 recession), and in a few cases it even reports moderate increases. The OECD concludes that this observation is consistent with some complementarity between the two types of labour.

As part of her study of a reform to the minimum wage in Portugal, Pereira (2003) reports evidence of sizeable substitution between teenage workers and workers in their early 20. Similarly, Hyslop and Stillman (2007) find that the effect of a reform of the rules that govern the eligibility of teenage workers to the minimum in New Zealand, in the early 2000s, has had sizeable effects on their wage and employment but no discernible effects on the wage or the employment of older (20 and over) workers, which argues against perfect substitution. While the framework of analysis proposed by Hyslop and Stillman (2007) does not allow for the straightforward derivation of an implied elasticity of substitution between types of labour, their findings do not support the hypothesis of perfect substitutability.

In this paper, we study how the structure of the UK minimum wage has, over time, affected the relative share of old (55+) and young (16–22) workers in the labour force. In particular, we estimate the elasticities of substitution between young (16–22) and old (55+) employees in the low paid occupations (as defined by the Low Pay Commission) in Great Britain. We assume that younger and older workers are distinct inputs in a Constant Elasticity of Substitution (CES) labour aggregate. Given this assumption we can easily characterize the relationship between the wage differentials between younger and older workers on the one hand and the composition of the workforce on the other. In this case it is the elasticity of substitution between types of labour, which determines the demand response, in terms of relative employment size, to a change in the relative wages. This approach was suggested by Katz and Murphy (1992), and employed by Card (2001), and Card and Lemieux (2001). Focusing on the effect of immigration, Borjas (2003) studies the relationship between employment composition and the wage structure in the US. The task is made difficult in general, since the employment composition and the wage structure are determined endogenously on the labour market and reflect the interactions between the supply and the demand of the distinct types of labour. To deal with the simultaneity, Card (2001) and Borjas (2003) argue that immigration into the US provides an exogenous source of variation of the supply of labour. In our context, the introduction of the UK NMW, its regular up-rating and the modifications of its design over time provide us with similar (arguably) exogenous variations to the relative price structure between younger and old workers. This exogenous source of variation (which we could understand as a change to the relative supply curve) allows us to identify the elasticity of substitution between young and old labour.

We draw our evidence from the UK Quarterly Labour Force Survey (LFS) and the Annual Survey of Hours and Earnings (ASHE) over the period 1997–2010. These two data sources are commonly used to assess the effect of the minimum wage on the labour market (see Low Pay Commission (2010) and later reports for examples). We construct two pseudo panels, one from each data source, where we aggregate employment, wage and wage bill data at the level of an occupation in a region in a given year. Over time, and within each occupation \times region cell we are able to measure the fraction of workers in a given age group who are liable to be affected by the future uprating of the NMW. We argue that these quantities are valid instruments which we can use to estimate the elasticity of substitution despite the simultaneous determination of the relative size of the workforce and of the wage differentials between age groups. Our findings suggest substantial complementarities in production between young and older workers. The separate analyses of the two data sources provide results that are comparable. Hence, the mechanism which supports our instrumental variable methodology is robust to the origin of the information we use.

The remainder of the paper is organized as follows, the next two sections describe the construction of the datasets and the ratios of interest, and

our empirical strategy, respectively. The following section discusses estimates of both the reduced form and structural equations, and the final section concludes.

2 DATA

We base our analysis on the two surveys, the Labour Force Survey (LFS) and the Annual Survey of Hours and Earnings (ASHE),² which are most commonly used to assess the effect of the minimum wage on the labour market. The LFS is a quarterly survey of households in the UK with around 140,000 respondents (earlier surveys are larger) in about 60,000 households. Respondents are followed for five quarters, and earnings data is collected in the first and final wave from 1998 (only the fifth wave collects earnings data before 1998). We pool all respondents who report earnings data³ within a 12-month period from 1st October in year $t - 1$ to 30th September in year t . Fortuitously, this matches the regular timing of changes to NMW rates over the sample period.

The ASHE is an administrative survey based on a sample of employees taken from HM Revenue and Customs PAYE records, which collects information from all employees with a particular final two digits to their National Insurance numbers. Data is collected directly from payroll records and should provide reliable information on earnings and hours. ASHE records information on gross pay, that is before tax, national insurance or other deduction and does not include payments in kind.⁴

We assume that the relevant labour markets for low paid workers are well described by the occupation (low-pay occupations are those where the minimum wage matters the most), region as well as the age of any such worker. Hence, we sort individual observations according to (low-pay) occupation,⁵ region of

²Stewart (2011) discusses in detail the relationship between ASHE and the New Earnings Survey (NES).

³From the LFS we lose 9.4 per cent of the sample in year 2001, while from the ASHE, we lose 0.1 per cent in 1998, 1.8 per cent in 1999, 1.9 per cent in 2000, 2 per cent in 2001 and 1.7 per cent in 2002. We relax this restriction as a robustness check: the results remain unaltered (available from the authors upon request).

⁴There are two discontinuities in the ASHE dataset within our sample period. In 2004, information (additional surveys) on employees starting a new job between January and April (survey reference date) was included. In 2006, large businesses that return their data electronically ('special arrangements') to the Office for National Statistics (ONS) were treated as a separate stratum in the ASHE—we include the new stratum in year 2004 and exclude the new stratum in year 2006. This means that we maintain a steady number of homogeneous individuals in the yearly data with the least possible loss of information.

⁵We assign individuals to cells using the definitions of low-pay and major occupational groups according to the Standard Occupation Classification frameworks of 1990 and 2000. To be consistent with other studies, we follow the definitions used by the UK Low Pay Commission and the Office of National Statistics. Table A1 in the appendix summarizes the low-pay occupation definitions/groups: Retail, Hospitality, Social Care, Food Processing, Leisure, Travel and Sport, Cleaning, Agriculture, Security, Childcare, Textiles and Clothing, Hairdressing, Office Work.

residence⁶ and age group for each survey year. Since the age of a worker determines the National Minimum Wage Rate, we group workers into five age bands: 16–17 year olds ('very young'), 18–20 year olds ('young'), 21 year olds, 22–54 year olds ('middle aged') and 55+ year olds ('older'). We divide the adult group (older than 22) into two groups since (1) skills sets and incentives may differ over the age distribution of this group, thus aggregation may mask important differences, and (2) most low-pay occupations are manual or physically demanding (e.g. fruit-picking, night watches etc.) and 'older' workers may not be perceived (and indeed not be, if health status deteriorates with age) as productive as their 'younger' counterparts.⁷ Finally, we create the 21 year olds group to be consistent with the change of the definition to the age group eligible for the youth development rate in October 2010.⁸

Within each cell we focus on a few variables: l_y and l_o measure the number (count) of young and old employees,⁹ respectively, and the ratio of young to old workers within each occupation \times region cell in a given time period is then $l_{y/o} = \frac{l_y}{l_o}$. We measure the wage rate using the variables HOURPAY (gross hourly pay) for the LFS, and HE (average hourly earnings) for the ASHE. The wage ratio is derived directly from the within cell averages as $w_{y/o} = \frac{w_y}{w_o}$. For the wage bill ratio, we consider the ratio of the sum of weekly earnings (hourly wage, w , \times usual hours of work, h) among workers, hence in some cell c we measure the relative wage bill as:

$$w_{y/o}^* = \frac{\sum_{\text{all young workers in } c} w_y \cdot h_y}{\sum_{\text{all older workers in } c} w_o \cdot h_o}$$

⁶Our regional classification uses the Government Offices for the Regions (GORs) classification: North East, North West, Yorkshire and the Humber, East midlands, West midlands, East of England, London, South East, South West, Wales, Scotland, Northern Ireland (LFS sample only) and Outside the UK. Northern Ireland data is owned by the Department for Trade and Investment (www.detini.gov.uk) and is not included in ASHE. Data for employees outside the UK not available due to ASHE's sampling design. We include Merseyside with the North West in 1997 as well, for consistency. To make the information drawn from the LFS comparable to the information derived from the AHSE, we exclude observations from Northern Ireland and outside the UK, which are not surveyed by the ASHE.

⁷We estimate our models with 'prime' workers as the comparison group. Table A2 in the appendix presents these results.

⁸From 1st October 2010, workers who are employed as apprentices are covered by the national minimum wage at a lower rate. Since 2010, the apprentice rate is set about two pounds forty pence below the youth development rate. However, apprentices are identified in the LFS only. We do not pursue the effect of the apprentice rate here.

⁹We have also constructed the employment ratio using the total number of usual hours worked by young and old workers in any given cell. The main conclusions remain unaltered.

The number of survey observations per cell, n_c , does vary over the sample period and between surveys from 10 to more than 250 individuals.¹⁰ All cell averages are obtained using the original dataset sampling weights.

By construction, our synthetic panel data should contain 10 regions \times 13 occupations \times 4 age groups¹¹ \times 14 time periods = 7280 synthetic (cell or pseudo) observations for either dataset. By grouping individuals in cells that are sufficiently ‘large’ we potentially bypass problems of (non-classical) measurement errors. Verbeek and Nijman (1993) argue for cell sizes in excess of 100 individuals even though this number could be lower, if observations are sufficiently homogenous within any cell at any point in time. We obtain 5720 pseudo-observations in the LFS (78.5 per cent of maximum coverage) and 6564 observations in the ASHE (90.2 per cent of the maximum coverage) where both the wage ratios and the employment ratios are observed; similarly we obtain 5735 observations in the LFS (78.8 per cent of the maximum coverage) and 6573 observations in the ASHE (90.3 per cent of the maximum coverage) where both the relative wage bill and the relative employment are observed.¹² Figs. 3, 4 and 5 present the average (log) ratios of employment, wages and wage bills over the sample period, respectively.

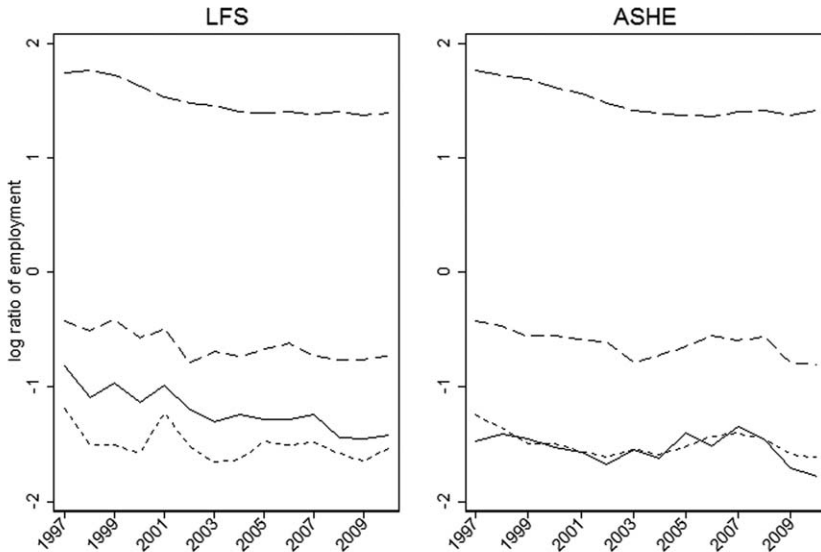
Starting from Fig. 3 note first that the patterns for the older age group (22–54 relative to the 55+) and the 18–20 group are very similar across LFS and ASHE. This provides some reassurance that the two surveys, arising from different data sources, do not provide conflicting evidence. It is for the younger group that the differences are the largest. It appears that, at least at the start of the sample period, the LFS over-estimates the relative size of the employment of the very young (16–17) to old workers. The ASHE instead suggests that relative to older workers, the employment sizes for the very young (16–17) and for the 21 year olds are more or less equal over the sample period.¹³

¹⁰In theory, n_c would not vary over survey year but in practice, as Browning *et al.* (1985) also note, socio-economic and/or other factors such as emigration, could affect sample cell sizes. Understandably, cells for the very young (16–17 year old) and young (18–20 year old) workers in certain occupations such as Security, Agriculture, or Child Care are expected to, and indeed do, have small sizes, while occupations such as Retail, Hospitality and Cleaning have large sizes (>100). In small cells the cell’s sample average may not be a precise estimate of the cell’s population mean. We note this potential caveat for the very young and young cohorts but given the available data we are unable to address it any further. We do not employ a formal threshold cell size for inclusion in the sample other than that imposed by the UK Data Service Secure Lab as per the data usage agreement, namely cells with <10 observations are treated as ‘empty’. The exclusion of cells of few individuals, namely <40 has produced qualitatively unaltered results.

¹¹We are interested in ratios relative to the older age group, hence, we focus on the observations of these ratios for four groups only.

¹²For a given cell we must be able to measure the employment ratio between young and old workers, and either the relative wage or the relative wage bill. If one or more of these is missing in a given cell the information becomes missing.

¹³The divergence between the LFS and ASHE samples could be partly due to sampling error.



LFS and ASHE data supplied by the UK Data Service Secure Lab

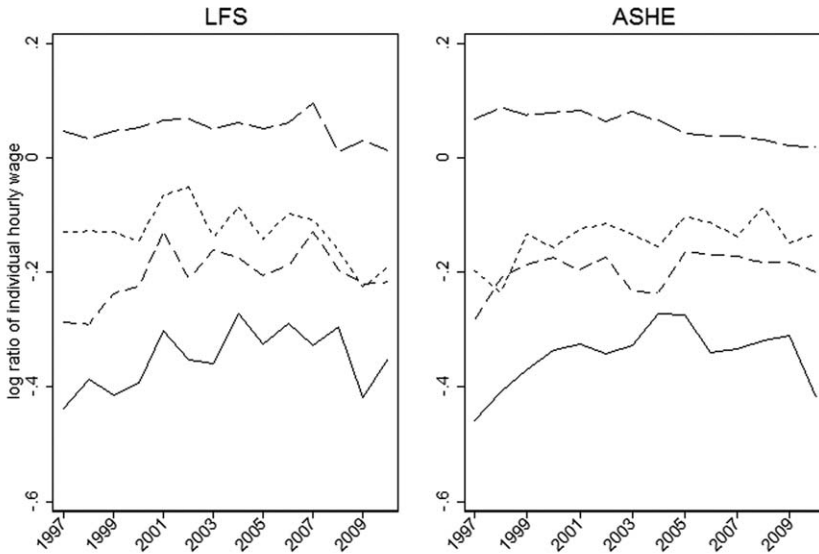
FIG. 3. Average (Over the Low Pay Occupations and Regions) of the Log Ratio of Employment.

Note: Ratio of 16–17 (solid), 18–20 (dashed), 21 (short dashed) and 22–54 (long dashed) to 55+ year old individuals in employment over time.

The mild decline in the ratios suggested by the LFS and, to a lesser extent, the ASHE sample can be attributed to the increased participation rates in higher and further education among the young over the sample period, the smaller size of the younger cohort, the increased participation of old workers in the labour force, or a combination of the three.¹⁴

Figure 4 presents the evolution of the wage ratios. The evidence from the LFS is more volatile but consistent with what we can observe from the ASHE. Both samples suggest an initial moderate upward trend from the late 1990s until the early 2000s, after when the ratio of the average wage of the ‘young’ relative to that of the old appears to level out for all age groups. Either survey allows a comparable ranking of the relative wages over the sample period: 22–54 year olds are better paid than older workers (the

¹⁴We explore the evolution of labour inputs further in Figs. A1 and A2 in the appendix. We observe that while numbers of ‘old’ workers remain stable over the sample period, those of ‘younger’ workers slightly decline in both low-pay occupation and overall. This is more pronounced in the LFS rather than the ASHE sample, where numbers appear more stable. Nonetheless, a very moderate decrease in the employment of the 16–17 year olds can be argued for in the ASHE sample as well. This is suggesting that it may in fact be a reduction in the supply of labour among the young, rather than an increase in the supply of labour among the old, that drives the observed trajectory of the ratios. The ‘dip’ in the LFS in 2001 is attributed to missing values, see footnote 3.



LFS and ASHE data supplied by the UK Data Service Secure Lab

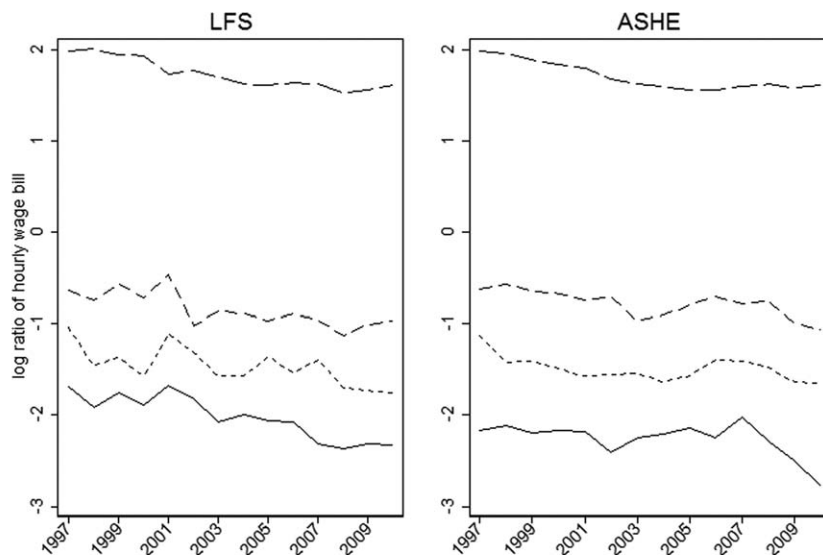
FIG. 4. Average (Over the Low Pay Occupations and Regions) of the Wage Differential Between Young and Old Workers.

Note: Ratio of 16–17 (solid), 18–20 (dashed), 21 (short dashed) and 22–54 (long dashed) to 55+ year old individuals in employment over time.

differential is between 5 per cent and 10 per cent), 21 year olds are paid less than older workers and relatively more than the 18–20 year olds. Finally, younger workers are paid the least with a differential relative to older workers ranging from 30 per cent to 40 per cent.

We believe that ASHE reports higher quality wage data since the information it contains is drawn from payroll records and is not self-reported as in the LFS. It is noteworthy that very young workers (16–17 year olds) appear to have experienced a reduction in their hourly wage relative to that of old workers at the same time as (or close after) the introduction of the 16–17 year old NMW rate (Oct 2004). Moreover, for the 18–20 and 21 year olds, the 2001 increase in the NMW development rate pushed wages up by almost 10 per cent.

The average ratios of the wage bills (Fig. 5) exhibit patterns that are similar to the employment ratios and wage ratios. As we would expect from the evidence presented in Figs. 3 and 4, the wage bill of middle aged workers in a low pay occupation is more than 15 per cent larger than the wage bill of older workers in a similar occupation. The wage bill of the very young is about 20 per cent less than that of older worker, and between 5 per cent and 10 per cent of that of the 18–21 year olds. Because of its small size, the wage bill of the 21 year olds is somewhere in between. The relative labour cost of



LFS and ASHE data supplied by the UK Data Service Secure Lab

Fig. 5. Average (Over the Low Pay Occupations and Regions) of the Log Ratio of Relative Wage Bill Paid to Young and Old Workers.

Note: Ratio of 16–17 (solid), 18–20 (dashed), 21 (short dashed) and 22–54 (long dashed) to 55+ year old individuals in employment over time.

‘young’ age groups appears to fall over time in the LFS, but has remained relatively constant in the ASHE.

Finally, Fig. 6 presents some evidence concerning the measured proportion affected by the introduction or the uprating of the minimum wage. Within a particular cell and an age group g , we measure the proportion of workers in period $t - 1$ who are paid a wage between the NMW current at time $t - 1$, $w_g(t-1)$, and the NMW that will apply at time t , $w_g(t)$. We denote these quantities $z_y(s, r, t-1)$ and $z_o(s, r, t-1)$ for young¹⁵ and old workers, respectively, where s indexes occupations and r regions.^{16,17} For the

¹⁵ $z_y(s, r, t-1)$ refers to a generic ‘young’ age group. A more accurate notation would be $z_{y_i,t-1}(s, r, t-1)$ where $i=(1, 2, 3, 4)$ for each of the 16–17, 18–20, 21 and 22–54 year old groups, we omit the i subscript for simplicity.

¹⁶We make the plausible assumption that the NMW upratings lead to a wage change of the directly affected workers only and thus have no further effect higher up the wage distribution.

¹⁷The ‘fraction affected among the young’, $z_y(s, r, t-1)$, in the ASHE sample has an average value, over the period, of 4 per cent ($s.d.=15$ per cent) for the 16–17, 15 per cent ($s.d.=19$ per cent) for the 18–20, and 15 per cent ($s.d.=24$ per cent) for the 21 year old, while for the 22–54 year old workers the average is 19 per cent ($s.d.=12$ per cent). The ‘fraction affected among older workers’, $z_o(s, r, t)$, has a mean of 21 per cent ($s.d.=16$ per cent). In the LFS sample the corresponding means (standard deviation figures in parentheses) of $z_y(s, r, t-1)$ are, 12 per cent ($s.d.=25$ per cent), 25 per cent ($s.d.=26$ per cent), 16 per cent ($s.d.=27$ per cent)

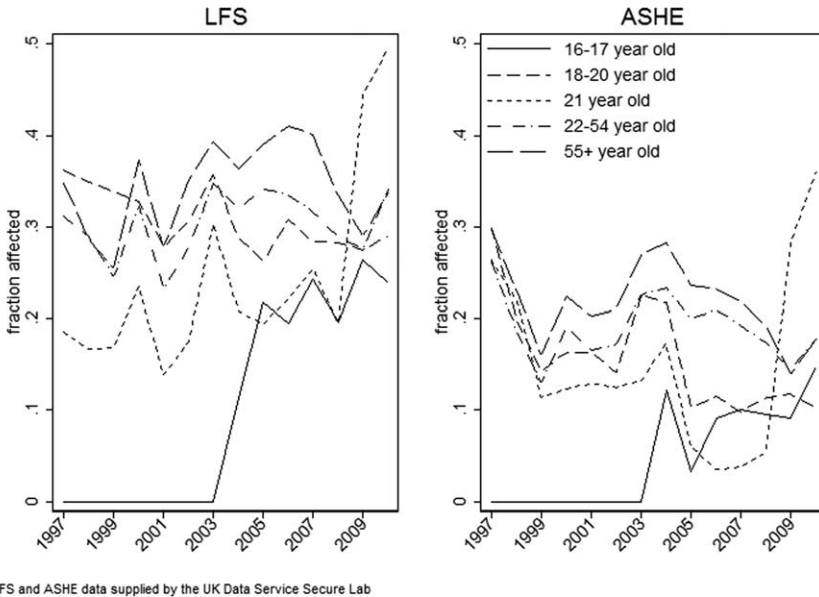


FIG. 6. Average (Over Low Pay Occupations and Regions) of the ‘fractions affected’, $z_y(s, r, t)$ and $z_o(s, r, t)$ Over Time.

Note: Proportions affected by uprating of NMW among 16–17 (solid), 18–20 (dashed), 21 (short dashed), 22–54 (dash and dot) and 55+ (dashed, long) year old individuals.

period before the NMW was introduced, 1997–99, we calculate the proportion using the introductory NMW rates. This is similar to the ‘fraction affected’ variable (Card, 1991), which is used in the literature as an instrument¹⁸ for the effect of the minimum wage. We present the evidence from both data sources split into the relevant age groups¹⁹ as well as for older (55+) workers. In the absence of demand side reactions these proportions are directly correlated with increases in total cost (measured by the wage bill) or increases in the average/marginal cost (measured by the average wage), which follow the introduction or the regular uprating of the minimum wage.

The two data sources provide significantly different measurements for the size of the fraction affected by the next uprating of the minimum wage, especially for the younger age groups. The Low Pay Commission (2010),

cent) and 20 per cent (*s.d.* = 17 per cent), respectively, while that of $z_o(s, r, t-1)$ is 22 per cent (*s.d.* = 22 per cent).

¹⁸For a critical discussion of other minimum wage variables used in the literature see Lemos (2005).

¹⁹For the young, we use the corresponding national minimum wage rate for the age group each individual belongs to, in each period.

Fig. 4.1, reports differences in these estimates between the two sources of data. At the start of the sample period, both datasets have an almost equal number of ‘affected’ workers, however, in later years, the proportions diverge in absolute terms, with a smaller affected fraction measured in the ASHE than the LFS sample.

The differences may arise from the differences in the way the hourly wage is measured in each data source (the LFS self-reported wage measure may not be as accurate as that of the ASHE).²⁰ However, the correlations between $z_y(s, r, t-1)$ and $z_o(s, r, t-1)$, from the LFS and ASHE, respectively, appear similar: -0.0371 and -0.0458 for the 16–17 year old, 0.1192 and 0.1255 for the 18–20 year old, 0.1819 and 0.0704 for the 21 year old, and 0.5467 and 0.5352 for the 22–54 year old.

We complete this section by exploring the variability of the log ratios, wage, wage bills and employment, by running OLS regressions on the complete set of occupation, region and year dummies. Table 1 presents the F -statistics on the joint significance of each set of dummies in turn. Overall, the results suggest that for all age groups, there exists significant variation across occupations and regions over time (in both samples).²¹ It is striking, however, that the year dummies do not seem to have much joint explanatory power for the wage bill ratios and the employment ratios for the ‘young’ age groups excluding the 22–54 year old (the evidence is more marginal for the wage ratios). Finally, we note that the regional dummies do not systematically explain the variability of the relative wage ratios for the 21 year old age group (LFS), or the relative employment ratios for the 18 to 20 year old (LFS and ASHE).

3 EMPIRICAL STRATEGY

The purpose of the modeling framework we present in this section is to describe the relationship between the relative employment size of ‘young’ and older workers on the one hand, and the relative wage of young and older workers or the relative wage bill paid to young and older workers on the other. We then show how this simple theoretical framework leads to a specification which can be brought to the data.

We divide the workforce in a given occupation into ‘young’, middle aged and older workers. Besides the different types of labour, the technology involves other inputs for example capital, K . We further assume that the production function depends on young and old labour through a labour

²⁰In the LFS the hourly wage is derived from self-reported weekly earnings and usual hours worked and thus likely to measure these proportions imprecisely, if there are recall errors. Hayes *et al.* (2007) find that ASHE consistently reports higher earnings (and by extension wage) data than the LFS.

²¹We note, however, that this could occur if the dummy for one category only in a particular set of dummies explains a sizeable part of the variance.

TABLE 1
OLS REGRESSIONS OF LOG RATIOS ON OCCUPATION, REGION AND YEAR DUMMIES

Variable	Age group	LFS						ASHE									
		All		Occupation		Region		All		Occupation		Region					
		F-stat	p val	F-stat	p val	F-stat	p val	F-stat	p val	F-stat	p val	F-stat	p val				
ln $w_{y/o}$	16-17 year olds	5.384	0.000	13.32	0.000	2.060	0.025	1.384	0.160	5.169	0.000	9.946	0.000	2.775	0.002	2.223	0.007
	18-20 year olds	4.391	0.000	8.300	0.000	2.759	0.002	2.227	0.007	5.601	0.000	9.709	0.000	4.360	0.000	2.118	0.011
	21 year olds	2.170	0.000	4.107	0.000	0.824	0.606	1.289	0.213	3.292	0.000	4.252	0.000	3.768	0.000	1.907	0.025
	22-54 year olds	2.835	0.000	5.280	0.000	2.157	0.018	1.076	0.376	5.613	0.000	11.08	0.000	2.180	0.017	3.060	0.000
	16-17 year olds	31.99	0.000	75.73	0.000	2.760	0.002	2.228	0.007	41.00	0.000	104.2	0.000	5.101	0.000	1.639	0.069
ln $w_{y/o}^*$	18-20 year olds	49.73	0.000	136.4	0.000	2.127	0.020	2.153	0.010	92.29	0.000	259.7	0.000	1.864	0.046	1.923	0.024
	21 year olds	26.01	0.000	67.84	0.000	2.436	0.007	2.778	0.001	50.72	0.000	145.1	0.000	1.497	0.134	1.570	0.087
	22-54 year olds	27.37	0.000	66.26	0.000	2.723	0.003	5.813	0.000	52.39	0.000	114.3	0.000	4.562	0.000	8.240	0.000
	16-17 year olds	87.30	0.000	219.9	0.000	2.395	0.008	2.104	0.012	86.54	0.000	238.1	0.000	3.135	0.001	0.929	0.522
	18-20 year olds	95.31	0.000	277.1	0.000	0.981	0.458	1.597	0.079	127.9	0.000	370.7	0.000	0.645	0.776	1.468	0.122
ln $l_{y/o}$	21 year olds	52.14	0.000	136.7	0.000	3.554	0.000	1.668	0.062	67.26	0.000	188.7	0.000	1.723	0.071	1.281	0.218
	22-54 year olds	40.29	0.000	88.62	0.000	5.705	0.000	7.437	0.000	61.68	0.000	118.6	0.000	9.255	0.000	9.244	0.000

Note: The entries in the table present the results from the regression of the endogenous variables, (respectively: log wage ratio, log wage bill ratio, and log employment ratio) on occupation, region and year dummies. The entries present the F -statistics that correspond to the test of joint significance of alternative sets of dummy variables: occupations, region, year dummies separately or all dummies together. The first two columns in each case present the F -statistics and the p value for the null hypothesis that the parameters for all dummies are jointly all 0. The left hand side of the table presents the results based on the pseudo panel derived from the LFS and the right hand side of the table presents the results based on the pseudo panel derived from the ASHE.

aggregate, L_{yo} , as well as the size of the rest of the labour force (the ‘middle aged’ group), L_{ma} . The production function takes the form:

$$Q = F(K, L_{yo}, L_{ma}), \tag{1}$$

where the young/old aggregate is defined as a Constant Elasticity of Substitution (CES) aggregate and satisfies:

$$L_{yo}^\rho = \Theta l_y^\rho + l_o^\rho, \tag{2}$$

where Θ captures the differences in productivity between young and old workers. The parameter ρ , $-\infty \leq \rho \leq 1$, is a function of the elasticity of substitution σ between the two types of labour²² and describes the substitution between young and old workers keeping the size of the young/old aggregate constant. This parameter summarizes the role that technological choices play in evaluating the effect of a labour market intervention which affects the relative price of young and old labour such as, e.g. age contingent minimum wages.

Assuming that wages are given, the first-order conditions for the minimization of costs of producing a level of output in terms of l_y and l_o are:

$$\begin{aligned} v \frac{\partial F}{\partial L_{yo}} \frac{\partial L_{yo}}{\partial l_y} &= w_y, \\ v \frac{\partial F}{\partial L_{yo}} \frac{\partial L_{yo}}{\partial l_o} &= w_o, \end{aligned}$$

where v is the Lagrange multiplier associated with the production requirements constraint. These two expressions together imply the following relationship between the relative wages and the relative employment:

$$\ln w_{y/o} = \ln \Theta + (\rho - 1) \ln l_{y/o}, \tag{3}$$

where $w_{y/o} = \frac{w_y}{w_o}$ and $l_{y/o} = \frac{l_y}{l_o}$ as already defined.

This condition suggests simply that the young/old log wage differential determines the (logarithm of the) relative utilization of the two age groups. Hence, this is a model of the relative demand of young and old workers. If the two age groups are paid the same wage, i.e. $\ln w_{y/o} = 0$, the quantity $\sigma \ln \Theta$ determines the optimal relative employment size. Hence, at equal pay the employment of young workers will be larger than the employment of older workers, if the former are more productive than the latter i.e. if $\Theta > 1$. Given the relative productivity of the young, the employment differences will be larger the larger σ is.

The model is identical to the Constant Elasticity of Substitution (CES) model used in other contexts, in particular when studying the effect of

²² $\rho = 1 - \frac{1}{\sigma}$ or equivalently $\sigma = \frac{1}{1-\rho}$

demographic changes or the effect of immigration on relative employment and wages, but with similar objectives, by Katz and Murphy (1992), Card (2001) and Borjas (2003). The parameter of the (logarithm of the) relative utilization of the two inputs, i.e. $\rho - 1 = -1/\sigma$ identifies directly the elasticity of substitution between the two kinds of labour (within the young/old aggregate).

Equation (3) can be restated as a relationship between the relative quantities, $\ln l_{y/o}$, and the relative pay, $\ln w_{y/o}$, or as a relationship between the relative quantities, $\ln l_{y/o}$, and the relative wage bill $\ln w_{y/o}^*$:

$$\ln l_{y/o} = \theta - \sigma \ln w_{y/o}, \quad (4)$$

and

$$\ln l_{y/o} = \frac{\theta}{1-\sigma} + \frac{1}{\rho} \ln w_{y/o}^*, \quad (5)$$

where $\theta \equiv \sigma \ln \Theta$. In principle, either relationship can be estimated from available data, since we observe both relative wages, $w_{y/o}$, and relative wage bills, $w_{y/o}^*$, and either specification will provide an estimate of the elasticity of substitution σ , or equivalently ρ . We base our empirical work on relationships (4) and (5).

To capture the variability between groups we describe how the productivity parameter θ captures the differences in productivity between occupations, s , regions, r and periods, t . Furthermore, we allow for other differences that cannot be explained in terms of specific effects and which we assume are random. More precisely, we specify the productivity component as a general heterogeneous trend specification:

$$\theta(s, r, t) = [\mu(s) + \lambda(r) + \kappa(t)]t + u(s, r, t), \quad (6)$$

where the parameters μ , λ and κ are the effects specific to an occupation, a region, and a time period (respectively) on the trend of the relative productivity differences between young and older workers. $u(s, r, t)$ captures the random unexplained productivity differences between groups. For simplicity, we assume that this random component behaves like a random walk such that the first difference (in time) $u(s, r, t) - u(s, r, t-1) \equiv v(s, r, t)$ is an idiosyncratic innovation. Therefore, the difference $\Delta\theta(s, r, t)$, within the same (occupation \times region) group, contains region, occupation, and time specific components, as well as an idiosyncratic random component:

$$\begin{aligned} \Delta\theta(s, r, t) &\equiv \theta(s, r, t) - \theta(s, r, t-1) \\ &= \mu(s) + \lambda(r) + \tilde{\kappa}(t) + v(s, r, t), \end{aligned} \quad (7)$$

where $\tilde{\kappa}(t) \equiv t\Delta\kappa(t) + \kappa(t-1)$. Despite its complicated form, $\tilde{\kappa}(t)$ is specific to period t , and therefore can be thought of as a period specific effect that explains relative changes between period t and $t-1$ that are common to all observations occupations and regions.

Equations (4) and (5) expressed in terms of first differences become:

$$\Delta \ln l_{y/o}(s, r, t) = -\sigma \Delta \ln w_{y/o}(s, r, t) + \mu(s) + \lambda(r) + \tilde{\kappa}(t) + v(s, r, t), \quad (8)$$

and

$$\begin{aligned} \Delta \ln l_{y/o}(s, r, t) &= \frac{1}{\rho} \Delta \ln w_{y/o}^*(s, r, t) \\ &+ \frac{1}{1-\sigma} (\mu(s) + \lambda(r) + \tilde{\kappa}(t) + v(s, r, t)). \end{aligned} \quad (9)$$

Equations (8) and (9) specify a structural model linking changes in relative employment to changes in relative wages, or to changes in the relative wage bills between age groups. This is a natural specification since the differential reform to the minimum wage across age groups are likely to be reflected in changes in the relative quantities. The heterogenous trend specification rationalizes the presence of occupation, regional and time specific effects in the formulation in terms of first differences. The exact parameter identified depends on the variable on the RHS: equation (8) identifies σ while equation (9) identifies $1/\rho$. Given the restrictions on the technology, σ must be positive and equivalently, $1/\rho$ must be outside the (0, 1) interval.

The variables $\Delta \ln l_{y/o}(s, r, t)$, $\Delta \ln w_{y/o}(s, r, t)$ and $\Delta \ln w_{y/o}^*(s, r, t)$ on the right hand side and the left hand side of equations (8) or (9) are endogenously determined in equilibrium. However, the introduction of the NMW and its yearly uprating, as well as the definition of youth sub-minimum wage rates, provide exogenous²³ sources of variability to the relative wages or to the relative wage bills. We identify the variation of the minimum wage to a variation to the supply of labour (i.e. the wage that must be paid to secure some positive level of labour supply).

We assume that the change of the relative wage or of the relative wage bills from one period to the next depends on the proportion of workers in period $t - 1$ who are paid a wage between the NMW current at time $t - 1$ and the NMW that will apply at time t . Therefore, we argue that the fractions affected by future upratings, which we described in the previous section, are good instrumental variable candidates. In the absence of demand side reactions these proportions are directly correlated with increases in total cost (measured by the wage bill) or increases in the average/marginal cost (measured by the average wage), which follow the introduction or the regular uprating of the minimum wage.

Hence, we assume that the reduced form model which describes the determination of the wage differential in response to a change in the minimum wage for either age group is:

²³As long as minimum wages do not change as a response to changes in labour demand (current or expected), NMW upratings are most likely to be exogenous. Even though this cannot be guaranteed, (Neumark and Wascher, 2008, p. 93) note, '[...] we suspect that this problem [endogeneity bias] is more likely to arise in the context of the UK Wage Councils than in cases where minimum wages are enacted by legislatures, for which there often seems to be much more regard for political than economic timing'.

$$\begin{aligned} \Delta \ln w_{y/o}(s, r, t) &= \gamma_y z_y(s, r, t-1) + \gamma_o z_o(s, r, t-1) \\ &+ \phi_w(s) + \chi_w(r) + \psi_w(t) + \zeta_w(s, r, t), \end{aligned} \quad (10)$$

We assume that the unobserved components in this wage equation share the same error structure as the structural equations (8) and (9), and $\phi_w(s)$, $\chi_w(r)$ and $\psi_w(t)$ capture the occupation, region and time specific effects while $\zeta_w(s, r, t)$ captures all other idiosyncratic deviations from the conditional mean. This implies that the reduced form equation for the relative wage bills is:

$$\begin{aligned} \Delta \ln w_{y/o}^*(s, r, t) &= \delta_y z_y(s, r, t-1) + \delta_o z_o(s, r, t-1) \\ &+ \phi_{w^*}(s) + \chi_{w^*}(r) + \psi_{w^*}(t) + \zeta_{w^*}(s, r, t), \end{aligned} \quad (11)$$

where $\delta_y \equiv (1-\sigma)\gamma_y$ and $\delta_o \equiv (1-\sigma)\gamma_o$. The specific effects $\phi_{w^*}(s)$, $\chi_{w^*}(r)$ and $\psi_{w^*}(t)$ and the idiosyncratic error term are related to the structural parameters in equation (8) and the reduced form parameters in equation (10) in the following way:

$$\begin{aligned} \phi_{w^*}(s) &= \mu(s) + (1-\sigma)\phi_w(s), \\ \chi_{w^*}(r) &= \lambda(r) + (1-\sigma)\chi_w(r), \\ \psi_{w^*}(t) &= \tilde{\kappa}(t) + (1-\sigma)\psi(t), \\ \zeta_{w^*}(s, r, t) &= v(s, r, t) + (1-\sigma)\zeta_w(s, r, t). \end{aligned}$$

The reduced form equation for the change in the relative employment size has a similar expression:

$$\begin{aligned} \Delta \ln l_{y/o}(s, r, t) &= \beta_y z_y(s, r, t-1) + \beta_o z_o(s, r, t-1) \\ &+ \phi_l(s) + \chi_l(r) + \psi_l(t) + \zeta_l(s, r, t), \end{aligned} \quad (12)$$

where $\beta_y \equiv -\sigma\gamma_y$ and $\beta_o \equiv -\sigma\gamma_o$. The specific components of the reduced form are related to the specific components of the structural form (8) and to the components of the reduced form equation for the changes in the wage such that:

$$\begin{aligned} \phi_l(s) &= \mu(s) - \sigma\phi_w(s), \\ \chi_l(r) &= \lambda(r) - \sigma\chi_w(r), \\ \psi_l(t) &= \tilde{\kappa}(t) - \sigma\psi(t), \\ \zeta_l(s, r, t) &= v(s, r, t) - \sigma\zeta_w(s, r, t). \end{aligned}$$

We observe that when the two inputs are complements in production, i.e. whenever $\sigma = 0$ or equivalently when $\rho = -\infty$, equations (10) and (11)

suggest that the relative employment size is determined independently of the relative wages or the relative wage bills. When $\sigma = 0$ the reduced form equations suggest that the change in relative employment depends only on the specific effects which describe the evolution of productivity and not the components which describe how the relative wages vary or the fraction affected by the uprating to the NMW. Furthermore, in that case, the wage bill varies with the fraction affected and the sum of the specific effects that determine the productivity changes and those that determine the wage increases. The relative wage and the relative wage bill depend on the changes to the NMW, while the reduced form for the relative employment size does not respond to a change of the NMW.

If instead $\sigma = 1$, the Cobb-Douglas case, the observed wage bill is determined entirely by the components of the demand for labour: the components that determine how the wage changes for a particular (occupation \times region) group are irrelevant. The observed relative employment changes given by equation (12), on the other hand, depends only on the difference between the specific effects from the demand and supply side. Furthermore, the fractions affected by the NMW, $z_y(s, r, t-1)$ and $z_o(s, r, t-1)$, determine the changes to the relative employment in exactly the opposite way the fractions affected determine the changes to the relative wage, see equation (10).

The structural equations impose exclusion restrictions. These require that the proportions of young and old workers affected by NMW upratings have an effect on the log young/old ratio of employment only through their effect on the relative young/old average wage or wage bill. This excludes the possibility of a direct effect of the instruments on the relative employment size, i.e. $l_{y/o}$. Since the number of instruments is larger than the number of endogenous regressors on the right hand side of the structural equations (8) or (9) we can test whether or not aspects of the exclusion restrictions are supported by the data.

4 RESULTS

Recent theoretical and applied developments in empirical econometrics (as they are discussed for example in Angrist and Pischke (2008) argue that for the effects of economic policy to be rationalized in our modelling framework, the effect of the exogenous variables must be ‘sufficiently’ significant across all reduced form equations, if subsequent IV estimates of the structural relationships are to be informative. In practical terms, this means that we can conclude that the policy has had sizeable effects on the outcome of interest, if the test statistics summarizing the explanatory power of the proportion affected by the change in the NMW i.e. the

F -statistics for the test of the hypothesis that the instruments can be excluded from the reduced form equations, take a 'large enough' value. The literature suggests a larger threshold value for the F -statistic than the critical value of conventional tests at the 5 per cent level. (Stock and Watson, 2012) suggest as a rule of thumb that the F -statistic should be larger than 10 at least, and Angrist and Pischke (2008) provide a discussion of this rule in an empirical applied context and a theoretical justification.

We apply the approach we describe in the previous section to several definitions for the young workers group. Furthermore, we perform the same analysis on the middle aged workers (i.e. those aged between 22 and 54) relative to the older workers (those aged 55 or more). The last set of results provides a base line against which we can compare the results we obtain for the younger age groups. Over the sample period, the middle aged group share the same NMW rate as the older workers. The instrumental variable strategy is still effective in this case since in general the proportions affected by the uprating will be different between age group and between occupation and region at any point in time.

Table 2 presents parameter estimates of the reduced form equations (10), (11) and (12) for each age group relative to the old workers group within a particular low-pay occupation, region and year. The set of regressors contains the instruments (the fractions affected) as well as occupation, region and year dummies. Starting from the top panel, we observe that the proportion of young workers receiving a wage between the current NMW and next period's NMW, has a negative and statistically significant effect on the change in the log ratio of hourly wages. This is measured consistently and significantly across datasets. This means that the wage differentials between young and older workers are reduced in regions where the proportion affected is larger than average. Given the estimates, this effect is substantial (an implied semi-elasticity greater than 0.5 in absolute value, in almost all cases). Similarly, the effect of the proportion of old workers affected has a significant positive effect on the change in the log ratio of hourly wages. The F -statistics suggest that our instruments contribute significantly to the explanation of the observed change of the relative wages. The computed F -statistics are 'large' for all age groups and well above the suggested rule-of-thumb threshold of 10. Hence the uprating of the NMW, as it is measured using the fractions affected, has a significant effect on the relative wages between younger and older workers.

The middle panel of Table 2 reports the estimated effects on changes of the relative wage bill. Estimates from both the LFS and ASHE produce almost the same pattern of significance, a larger than average proportion of a younger group paid at an hourly wage between the current NMW

TABLE 2
REDUCED FORM EQUATION ESTIMATES (OLS)

	LFS				ASHE			
	16-17 year olds	18-20 year olds	21 year olds	22-54 year olds	16-17 year olds	18-20 year olds	21 year olds	22-54 year olds
$\Delta \ln(w_{t/o}^*)$	-0.623*** (0.079)	-0.648*** (0.044)	-0.699*** (0.054)	-0.387*** (0.060)	-0.695*** (0.110)	-0.563*** (0.046)	-0.656*** (0.055)	-0.307*** (0.068)
γ_o	0.421***	0.397***	0.286***	0.489***	0.260*	0.221***	0.151	0.282***
R^2	(0.090)	(0.066)	(0.105)	(0.034)	(0.138)	(0.062)	(0.104)	(0.036)
N	0.136	0.195	0.219	0.152	0.072	0.120	0.136	0.055
$F(H_o : \gamma_y = \gamma_o = 0)$	824	1153	731	1482	853	1396	1147	1584
p value	42.729	123.064	88.956	119.168	21.245	79.046	72.369	36.156
$F(H_o : \gamma_y + \gamma_o = 0)$	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
p value	62.579	217.300	167.488	41.012	39.933	148.517	143.903	20.568
	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
$\Delta \ln(w_{t/o}^*)$	-0.737*** (0.234)	-0.664*** (0.135)	-0.682*** (0.140)	-0.492*** (0.175)	-0.539** (0.270)	-0.264** (0.121)	-0.693*** (0.139)	0.036
δ_o	0.704***	0.716***	0.493*	0.833***	0.675*	0.596***	-0.040	0.038
R^2	(0.276)	(0.204)	(0.274)	(0.098)	(0.350)	(0.166)	(0.266)	(0.072)
N	0.044	0.055	0.074	0.059	0.044	0.040	0.043	0.027
$F(H_o : \delta_y = \delta_o = 0)$	829	1160	734	1484	856	1398	1151	1584
p value	8.455	17.465	13.842	38.675	3.702	8.570	12.503	0.201
$F(H_o : \delta_y + \delta_o = 0)$	0.000	0.000	0.000	0.000	0.025	0.000	0.000	0.818
p value	9.918	24.200	23.655	7.858	4.000	4.732	24.874	0.072
	0.002	0.000	0.000	0.005	0.046	0.030	0.000	0.788

TABLE 2 (Continued)

	LFS			ASHE				
	16-17 year olds	18-20 year olds	21 year olds	22-54 year olds	16-17 year olds	18-20 year olds	21 year olds	22-54 year olds
$\Delta \ln(I_{it0})$								
β_y	-0.191 (0.122)	-0.008 (0.077)	-0.134 (0.088)	-0.128 (0.117)	-0.165 (0.171)	0.099 (0.095)	-0.207** (0.101)	0.206* (0.113)
β_o	-0.285** (0.140)	0.096 (0.113)	-0.203 (0.177)	0.127** (0.064)	0.240 (0.222)	0.216 (0.135)	-0.027 (0.215)	-0.234*** (0.060)
R^2	0.040	0.031	0.047	0.017	0.072	0.038	0.023	0.046
N	882	1225	819	1509	863	1405	1176	1584
$F(H_o : \beta_y = \beta_o = 0)$	3.326	0.363	1.739	2.366	1.004	1.902	2.100	8.340
p value	0.036	0.696	0.176	0.094	0.367	0.150	0.123	0.000
$F(H_o : \beta_y + \beta_o = 0)$	2.455	0.010	2.315	1.197	0.924	1.089	4.162	3.354
p value	0.118	0.920	0.128	0.274	0.337	0.297	0.042	0.067

Source: LFS and ASHE data supplied by the Secure Data Service.

Note: The entries in this table present the OLS parameter estimates for the instruments $z_{y,t}(s, r, t-1)$ and $z_{o,t}(s, r, t-1)$ in the reduced form equations (10), (11) and (12). The estimated equations include occupation, region and year dummies. The left hand side of the table presents the results based on the pseudo panel derived from the LFS and the right hand side of the table presents the results based on the pseudo panel derived from the ASHE. Standard errors in parentheses * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

and next period's NMW is associated with a smaller change in the relative wage bill, while a larger than average proportion of old workers affected is associated with a larger change in the relative wage bill in both samples. The F -statistics, which describes the 'strength' of the association, are in half the cases smaller than 10. In particular, in the ASHE the only 'strong' enough association arises for the 21 year olds. In the LFS's case the groups 18–20, 21 exactly, and between 22 and 54 show a similarly strong association. In this last case, the difference between the two data sources arises because the proportion affected among older workers appears to explain the relative wage bill paid to the 22–54 year olds using the data derived from the LFS, while this is not the case when the data is derived from the ASHE.

Concerning the relative employment size (bottom panel of Table 2), we only get significant estimates using data derived from the ASHE, among the 21 and the 22 to 54 year olds, while using data from the LFS, among the very young. The F -statistics are systematically smaller than the threshold value of 10 suggesting that the effect of minimum wage upratings, as captured by the proportion of 'workers affected', has not had any sizeable effects on employment composition.

We also test in each case the null hypotheses that the sum of the effects of the proportion affected by the uprating of the NMW in the young and old group is equal to zero. For the relative wage rate and the relative wage bill, we reject the null hypothesis, at conventional levels, in all cases across datasets, except for the 22–54 year olds for the wage bill equation in the ASHE. In the former case (wage rate) the test suggests that the uprating of the NMW does change the relative wage of young and old workers. When it comes to the relative employment size, however, the effect of the uprating is only precisely estimated among the 21 year olds and the 22–54 year olds in the ASHE. In these two cases, we reject the hypothesis that the uprating does not change the relative size of employment.

Our results so far suggest that the introduction and uprating of the NMW has had a significant effect on the determination of the relative wages and relative wage bills, while the NMW had no systematic effect on the evolution of relative employment (i.e. in terms of the employment size of younger workers relative to the employment size of old workers). Hence, one of the usual requirements for the application of Instrumental Variables (IV) methodology is not satisfied systematically for all age groups: the candidate instruments should be significant in the reduced form equations for all endogenous variables whether they are on the RHS or the LHS of the structural equation of interest (see Angrist and Pischke, 2008, for a discussion). This conclusion is not sensitive to the presence or absence of outliers in the

TABLE 3
2SLS IV ESTIMATES

	LFS				ASHE			
	16–17 year old	18–20 year old	21 old	22–54 year old	16–17 year old	18–20 year old	21 old	22–54 year old
$\Delta \ln(l_{i/o})$ on $\Delta \ln(w_{i/o})$								
$-\sigma$	0.028 (0.181)	0.009 (0.112)	0.148 (0.128)	0.205 (0.125)	0.190 (0.244)	-0.236 (0.169)	-0.009 (0.170)	-0.796*** (0.211)
N	824	1153	731	1482	853	1396	1147	1584
J -statistic	9.188	0.140	0.853	0.202	1.186	5.415	0.002	0.136
p value	0.002	0.708	0.356	0.653	0.276	0.020	0.960	0.713
$\Delta \ln(l_{i/o})$ on $\Delta \ln(w_{i/o}^*)$								
$1/\rho$	0.017 (0.130)	-0.018 (0.100)	0.133 (0.112)	0.125* (0.067)	0.387** (0.170)	0.094 (0.176)	0.045 (0.152)	-3.351 (6.324)
N	829	1160	734	1484	856	1398	1151	1584
J -statistic	9.604	0.239	1.286	0.490	0.034	8.223	0.001	0.451
p value	0.001	0.625	0.257	0.484	0.852	0.004	0.969	0.502

Source: LFS and ASHE data supplied by the Secure Data Service.

Note: The entries in the table present the 2SLS parameter estimates of interest in the structural form equations (8), σ in the top half of the table, and (9), $1/\rho$ in the bottom half of the table. Each substantive column consider a distinct age group relative to the older workers. The set of instrumental variables includes: $z_y(s, r, t-1)$ and $z_o(s, r, t-1)$ as well as occupation, region and year dummies, overall 36 variables, while the set of variables included on the RHS of the structural equations include $\Delta \ln(w_{i/o})$ (top half), $\Delta \ln(w_{i/o}^*)$ (bottom half) as well as occupation, region and year dummies. The left hand side of the table presents the results based on the pseudo panel derived from the LFS and the right hand side of the table presents the results based on the pseudo panel derived from the ASHE.

Robust standard errors in parentheses * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

reduced form equation.²⁴ This is, however, consistent with complementarity between inputs, i.e. whenever σ is equal or close to 0 or, equivalently, whenever ρ approaches $-\infty$. This was clear from the reduced form equations (10), (11) and (12), which suggest that the relative employment size should be determined independently from the relative wages or the relative wage bills whenever $\sigma = 0$.

Table 3, presents the estimated structural parameters from Instrumental Variables (IV) estimation. The specification always controls for occupation, region and year effects although these estimated parameters are not presented. We report over-identification test statistics (J -statistics), which assess whether the exclusion of one of our two instruments from the structural equation is supported by the data. When more instruments than endogenous variables are available, a test of over-identifying restrictions is possible. The test assumes that one instrument

²⁴We investigate the sensitivity of the estimates to the presence of extreme observations in the samples. We run the first stage regressions by excluding extreme values from our samples by calculating the Cook's Distance and using the conventional $4/n$ threshold as a cutoff point. Even though some of the estimated effects are slightly larger in magnitude, the significance and signs of the coefficients are unaltered. These estimates are available from the authors upon request.

TABLE 4
LIML ESTIMATES

	LFS				ASHE			
	16-17 year old	18-20 year old	21 year old	22-54 year old	16-17 year old	18-20 year old	21 year old	22-54 year old
$\Delta \ln(I_{i/o})$ on $\Delta \ln(w_{i/o})$								
$-\sigma$	0.019 (0.191)	0.009 (0.112)	0.148 (0.128)	0.205* (0.125)	0.195 (0.247)	-0.238 (0.172)	-0.009 (0.170)	-0.797*** (0.212)
N	824	1153	731	1482	853	1396	1147	1584
J -statistic	9.186	0.140	0.853	0.202	1.185	5.415	0.002	0.136
p value	0.002	0.708	0.356	0.653	0.276	0.020	0.960	0.713
$\Delta \ln(I_{i/o}^*)$ on $\Delta \ln(w_{i/o}^*)$								
$1/\rho$	-0.225 (0.227)	-0.021 (0.100)	0.121 (0.116)	0.123* (0.067)	0.386** (0.170)	-0.214 (0.295)	0.045 (0.152)	-7.803 (19.228)
N	829	1160	734	1484	856	1398	1151	1584
J -statistic	7.690	0.238	1.274	0.489	0.034	6.261	0.001	0.216
p value	0.005	0.626	0.259	0.485	0.852	0.012	0.969	0.642

Source: LFS and ASHE data supplied by the Secure Data Service.
 Note: The entries in the table present the LIML parameter estimates of interest in the structural form equations (8), σ in the top half of the table, and (9), $1/\rho$ in the bottom half of the table. Each substantive column consider a distinct age group relative to the older workers. The set of instrumental variables includes: $z_y(s, r, t-1)$ and $z_o(s, r, t-1)$ as well as occupation, region and year dummies, overall 36 variables, while the set of variables included on the RHS of the structural equations include $\Delta \ln(w_{i/o})$ (top half), $\Delta \ln(w_{i/o}^*)$ (bottom half) as well as occupation, region and year dummies. The left hand side of the table presents the results based on the pseudo panel derived from the LFS and the right hand side of the table presents the results based on the pseudo panel derived from the ASHE.
 Robust standard errors in parentheses * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$.

is valid and then tests for the validity of all other instruments i.e. whether the instruments are uncorrelated with the error term of the structural equation of interest and are legitimately excluded from that equation. We accept the over-identification hypothesis, that is that both instruments can be excluded from the structural equation of interest, in almost all models across the two data sources (the exception is for the 16-17 year olds in the LFS sample, and the 18-20 year olds in the ASHE sample).

Just-identified 2SLS is approximately unbiased, however, even then, when the instruments are weak (the first stage is in actuality zero) the estimates may suggest a causal relationship when one is in fact not present. For over-identified models, the estimates are biased with the bias being an increasing function of the number of instruments. Given our concerns about the validity of the instruments, based on the first stage results, we also produce limited information maximum likelihood estimates (LIML) which are approximately median-unbiased and robust in the presence of weak instruments (Angrist and Pischke, 2008). Table 4 presents LIML estimates, comparable to those from 2SLS estimation of Table 3.

For the model linking relative employment to the relative average wage, equation (8), the 2SLS and LIML results are similar, which lends more

support to our estimates. Given our model specification, the absolute value of the parameter estimate of (the first difference of) the log ratio of average wages between young and old workers provides a measure of the elasticity of substitution, σ . The original parameter of interest is negative and statistically significant for the 22–54 year olds (at the 1 per cent level) in the ASHE sample only. This implies that the elasticity of substitution is around 0.79 (LIML estimate) for the 22–54 year olds, while for the same age group the estimate from the LFS sample is insignificant. For the 18–20, and 21 year olds the ASHE estimates are also negative but insignificant, while the LFS estimates are positive and insignificant. For the very young, 16–17 year old, both the LFS and ASHE estimates are insignificant and positive. We conclude that the value of the estimated elasticity of substitution argues in favour of some significant complementarity between the 22–54 year old and the 55+ age group.

Looking at the estimates of model (9) which describes the relationship between relative employment and the wage bill measure, we find that, again, the 2SLS and LIML estimates in both the LFS and ASHE samples are similar. The coefficient estimates for the 16–17 year olds in ASHE, and for the 22–54 year olds in the LFS, imply that $\sigma < 0$, which is not admissible (in our specification values of $1/\rho$ between 0 and 1 are not consistent with a positive value of σ). These findings raise concerns over the validity of these estimates. For the rest of the age groups, we fail to obtain any statistically significant result from either dataset or estimator, however, the estimated effects have the ‘correct’ (negative) sign in ASHE.

We consider 0 as an estimate of the elasticity of substitution which the data does not reject, and we conclude that the evidence presented in Tables 3 and 4 is consistent with complementarity between younger age groups and the old age group. This is also a conclusion we reach from the analysis of the reduced form equations i.e. the NMW has a significant, consistent effect on the changes of the relative average wage within low-pay occupation, region and year but it has no effect on the labour force age composition.

Card and Lemieux (2001) find the elasticity of substitution between age groups for the UK to range between 2 and 4, depending on model specification, which is markedly different to our findings. Borjas *et al.* (2012)²⁵ note that the CES framework can be sensitive to the definition of the qualification groups and indeed (Borjas *et al.*, 2012, p. 209) argue that “there is no convincing empirical evidence that indicates how best to pool”. In the UK, we expect most individuals employed in the low-pay sectors to hold qualifications at or below A-levels (broadly comparable to a High School diploma in the US). Since the majority of pupils achieving A-levels are likely to continue to tertiary education, any attempt at splitting the sample into finer

²⁵The authors’s argument considers distinct skills groups but a similar argument could be argued in terms of age.

education groups may suffer from the problem of (dis)aggregation across groups of very different sizes. Moreover, Borjas *et al.* (2012) argue that the CES framework could further be sensitive to (untestable) assumptions about the evolution of demand conditions for specific groups of workers, and suggest interpreting results with caution. Furthermore, the literature also provides evidence which points to complementarity rather than substitutability in production between young and older labour (Gruber and Wise, 2010; Kalwij *et al.*, 2010).

We further check the validity of our results by estimating confidence sets robust to weak instruments following the suggestion of Moreira (2003), who argue that in IV regression with a single endogenous variable and potentially weak instruments, the conventional (Wald-type) inferential statistics are unreliable. Table 5 presents statistics for each age group and dataset, which are valid whether or not the instruments are weak. We report the Anderson-Rubin (AR) and conditional likelihood ratio (CLR) confidence sets, as well as p values for the hypotheses of the coefficient of the relative average wage and the relative cost of labour being equal to zero and one, see Mikusheva and Poi (2006) for implementation details. The authors note that if the instruments are weak, the Anderson-Rubin confidence set could be uninformative, i.e. contains the whole real line, or even be empty, (\emptyset). In contrast the CLR confidence set always contains the LIML estimated value.

In our model linking the change in relative average employment to the change in relative average wage, the coefficient of the change in the relative average wage is a direct measure of the elasticity of substitution between ‘young’ and ‘old’ workers, hence, accepting the hypothesis of $\sigma = 0$ ($\rho = -\infty$) means we have to accept that the two types of labour are perfect complements in production. Accordingly, testing for $\sigma = 1$ (equivalently, $\rho = 0$) is testing whether our underlying production function is of the Cobb-Douglas type.

The results from the LFS sample suggest that we cannot reject the hypothesis of perfect complementarity of ‘young’ and ‘old’ labour ($\sigma = 0$) across age groups and tests except for the 22–54 year olds in the ASHE sample (based on both tests). This finding is in line with our IV estimates, which use conventional Wald-type confidence intervals, thus suggesting that the potential weakness of the instruments does not affect inference much. On the other hand, the hypothesis of $\sigma = 1$ (the Cobb-Douglas case) is rejected for all age groups in both the LFS and ASHE samples except for the 16–17 year olds. This implies that the age composition of the wage bill in any given occupation adjusts in response to wage changes with respect to ‘young’ and ‘old’ workers.

Looking at the wage bill measure, we get similar results, accepting the null hypothesis of $\sigma = 0$ for all age groups apart from the 22–54 year

TABLE 5
COVERAGE CORRECTED CONFIDENCE INTERVALS AND CONDITIONAL P VALUES

$\Delta \ln(w_{it})$ on $\Delta \ln(w_{it}^*)$	LFS				ASHE				
	Statistic	CI_L	CI_U	p value for $H_0 : \sigma=1$	p value for $H_0 : \sigma=0$	CI_L	CI_U	p value for $H_0 : \sigma=1$	p value for $H_0 : \sigma=0$
16-17 year old	Conditional LR	-0.361	0.452	0.000	0.807	-0.177	1.002	0.050	0.180
	Anderson-Rubin	-0.265	0.359	0.000	0.156	-0.194	1.025	0.057	0.162
	CLR	-0.198	0.289	0.000	0.718	-0.594	0.172	0.000	0.276
18-20 year old	AR	-0.259	0.351	0.000	0.936	∅		0.000	0.012
	CLR	-0.130	0.443	0.000	0.286	-0.261	0.484	0.000	0.563
	AR	-0.176	0.491	0.000	0.375	-0.303	0.528	0.000	0.462
22-54 year old	CLR	-0.003	0.509	0.000	0.053	-1.395	-0.415	0.000	0.000
	AR	-0.064	0.575	0.000	0.150	-1.558	-0.316	0.000	0.000
$\Delta \ln(w_{it})$ on $\Delta \ln(w_{it}^*)$	LFS				ASHE				
Statistic	CI_L	CI_U	p value for $H_0 : \sigma=\infty$	p value for $H_0 : \sigma=0$	CI_L	CI_U	p value for $H_0 : \sigma=\infty$	p value for $H_0 : \sigma=0$	
16-17 year old	CLR	-5.687	0.462	0.014	0.960	0.069	1.287	0.103	0.038
	AR	-0.677	0.300	0.004	0.132	-0.045	1.633	0.187	0.060
	CLR	-0.262	0.204	0.000	0.825	-2.123	0.236	0.000	0.479
18-20 year old	AR	-0.367	0.240	0.000	0.954	∅		0.000	0.012
	CLR	-0.260	0.406	0.000	0.370	-0.330	0.344	0.000	0.502
	AR	-0.314	0.427	0.000	0.345	-0.404	0.363	0.000	0.426
22-54 year old	CLR	0.004	0.282	0.000	0.045	-0.801	2.502	0.000	0.000
	AR	-0.042	0.310	0.000	0.131	-0.572	2.176	0.000	0.000

Source: LFS and ASHE data supplied by the Secure Data Service.

Note: The entries in the table show the Confidence Intervals around the parameter value 0 given by the Conditional Likelihood Ratio Test proposed by Moreira (2003) as well as the Confidence Intervals based on the Anderson and Rubin statistic. These are computed using the conddiverg command in Stata described in Mikusheva and Poi (2006). These confidence intervals are robust to the use of weak instruments. It is possible for the confidence intervals based on the AR statistic to be empty or unbounded intervals. Confidence intervals based on the CLR statistics are never empty. The table shows the p values for the test of different null hypothesis for the parameters of interest expressed in terms of the value of σ . The left hand side of the table presents the results based on the pseudo panel derived from the LFS and the right hand side of the table presents the results based on the pseudo panel derived from the ASHE.

old in the ASHE sample. Further, we reject the hypothesis of $\sigma=\infty$ across samples and age groups except for the very young, again from ASHE. All our estimates suggest considerable, if not perfect, complementarity. We, therefore, reach similar conclusions from both data sources, namely, ‘young’ and ‘old’ workers exhibit substantial complementarities.

5 CONCLUSION

We analyze data from 1997–2010 drawn from the LFS and the ASHE with a view to characterizing the effect of the different NMW age-based rates and their uprating on the relative wages and the age related employment structure among low-pay occupations. Our analysis suggests that, if anything, the introduction and uprating of the NMW has a significant effect on the determination of wages and wage bills, while the NMW has no systematic effect on the evolution of relative employment. The evidence points in the direction of substantial, if not perfect, complementarity between the young age groups (18–20 year old and 21 year old) and old workers (more than 55 year old). This in turn, suggests that the differences of the NMW between age groups may not matter much when it comes to determining the labour force composition. In that sense, the current structure of the minimum wage appears innocuous. However, the evidence we report shows that the regular upratings of the NMW has a significant effect on the relative wages between younger and old workers.

In the future as the UK National Living Wage is introduced for workers aged 25 or more, the question of the substitutability between age groups, this time between younger and adult groups, is set to become more central to the question of the design of the minimum wage rates. Our current study is not designed to provide guidance in this case but its methodology would apply.

DISCLAIMER

This work was based on data from the Quarterly Labour Force Survey and the Annual Survey of Hours and Earnings, produced by the Office for National Statistics (ONS) and supplied by the Secure Data Service at the UK Data Archive. The data are Crown Copyright and reproduced with the permission of the controller of HMSO and Queen’s Printer for Scotland. The use of the data in this work does not imply the endorsement of ONS or the Secure Data Service at the UK Data Archive in relation to the interpretation or analysis of the data. This work uses research datasets which may not exactly reproduce National Statistics aggregates.

APPENDIX

TABLE A1
DEFINITIONS OF LOW-PAYING OCCUPATIONS BY SOC2000 AND SOC1990 CODES

<i>Low-paying occupation</i>	<i>SOC2000</i> ⁽²⁾	<i>SOC1990</i> ⁽³⁾⁽⁴⁾
Retail	1234, 5496, 711, 7125, 721, 925	178, 720, 721, 722, 730, 731, 732, 790, 791, 792, 954, 959
Hospitality	5434, 9222-9225	620, 621, 622, 951, 952, 953
Social care	6115	644
Employment agencies	n/a	n/a
Food processing	5431-5433, 8111	580, 581, 582, 800, 801, 802, 809
Leisure, travel and sport	6211, 6213, 6219, 9226, 9229	630, 699, 875, 999
Cleaning	6231, 9132, 923	670, 671, 956, 957, 958
Agriculture	5119, 9111, 9119	900, 902, 903
Security	9241, 9245, 9249	615, 619, 955
Childcare	6121-6123, 9243, 9244	650, 651, 659
Textiles and clothing	5414, 5419, 8113, 8137	553, 556, 559
Hairdressing	622	660, 661
Office work	4141, 4216, 9219	460, 461, 462

Source: LFS and ASHE data supplied by the Secure Data Service.

Notes: (1) n/a is not applicable. (2) Low-paying occupation definitions (SOC 2000) provided by the UK Low Pay Commission. Low Pay Commission report 2010, Appendix 4: Review of the Low-paying sectors, Table A4.1, p. 243. (3) Adapted from data from the Office for National Statistics licensed under the Open Government Licence v.1.0: OOSS User Guide 2000: 22, Occupational Information Unit, Office for National Statistics. (4) Some relationships were adapted from: Elias, P., and Purcell, K. (2004) 'SOC(HE) A classification of occupations for studying the graduate labour market', Researching Graduate Careers Seven Years On; Research Paper No. 6, Warwick Institute for Employment Research, Table A3, p. 40.

TABLE A2
2SLS ESTIMATES WITH 'ADULT' (22+) WORKERS AS THE COMPARISON GROUP

	<i>LFS</i>			<i>ASHE</i>		
	<i>16-17 year old</i>	<i>18-20 year old</i>	<i>21 year old</i>	<i>16-17 year old</i>	<i>18-20 year old</i>	<i>21 year old</i>
$\Delta \ln(I_{i/o})$ on $\Delta \ln(w_{i/o})$						
$-\sigma$	0.275 (0.195)	0.070 (0.104)	0.296** (0.119)	0.317 (0.234)	-0.367** (0.146)	0.029 (0.159)
<i>N</i>	860	1234	749	856	1455	1171
<i>J</i> -statistic	1.519	0.568	0.511	0.060	0.030	0.000
<i>p</i> value	0.218	0.451	0.475	0.806	0.864	0.990
$\Delta \ln(I_{i/o})$ on $\Delta \ln(w_{i/o}^*)$						
$1/\rho$	0.321** (0.150)	0.047 (0.101)	0.279*** (0.089)	0.417** (0.197)	-0.769 (0.555)	0.069 (0.141)
<i>N</i>	864	1243	752	859	1457	1175
<i>J</i> -statistic	0.717	0.727	0.009	0.001	0.202	0.012
<i>p</i> value	0.397	0.394	0.926	0.973	0.653	0.914

Notes: (1) Robust standard errors in parentheses. (2) For the 21 year old, the sample extends to 2009 since from 2010, 21 year old are paid the adult NMW rate. (3) * $p < 0.10$; ** $p < 0.05$; *** $p < 0.01$. (4) Instrumented: $\Delta \ln(w_{i/o})$. (5) Number of Instruments: 36. (6) Included Instruments: Region, Occupation, and Year dummies. (7) Excluded Instruments: z_o , z_y . (8) Source: Labour Force Survey data supplied by the Secure Data Service.

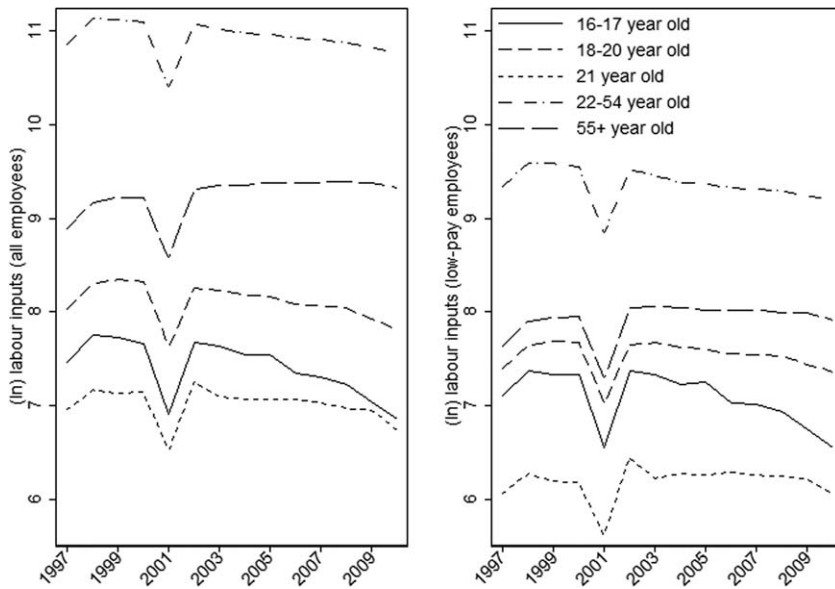


FIG. A1. Evolution of Labour Inputs, $I_{y/o}$, by Age Group from the LFS Sample
 The left panel depicts employees in all major occupational groups, while the right panel those employees in low-pay occupations. The pronounced dip in observation in 2001 is due to missing values, see main text footnote 3.

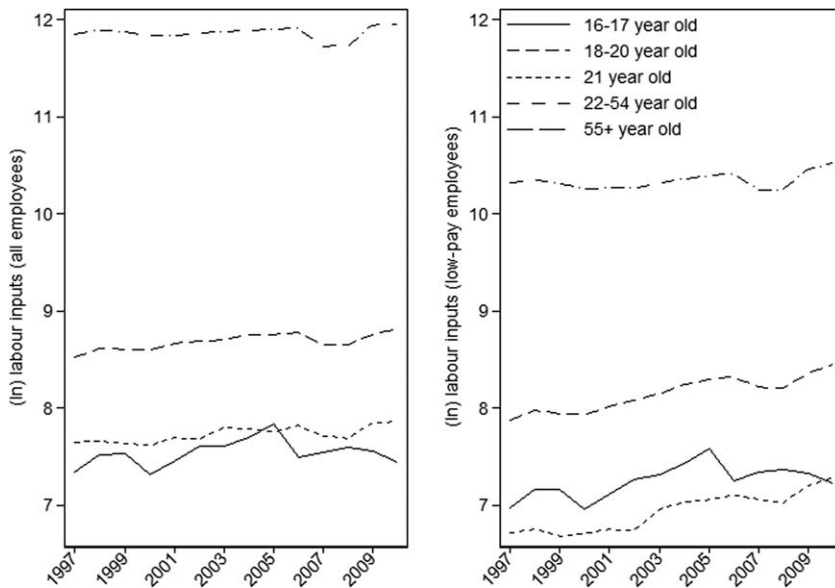


FIG. A2. Evolution of Labour Inputs, $I_{y/o}$, by Age Group from the ASHE Sample
 The left panel depicts employees in all major occupational groups, while the right panel those employees in low-pay occupations.

REFERENCES

- Angrist, J. D. and Pischke, J.-S. (2008). *Mostly Harmless Econometrics: An Empiricist's Companion*, Princeton University Press, Princeton, New Jersey, USA.
- Borjas, G. J. (2003). 'The Labor Demand Curve is Downward Sloping: Reexamining the Impact of Immigration on the Labor Market', *The Quarterly Journal of Economics*, Vol. 118, No. 4, pp. 1335–1374.
- Borjas, G. J., Grogger, J. and Hanson, G. H. (2012, 02). 'Comment: On Estimating Elasticities of Substitution', *Journal of the European Economic Association*, Vol. 10, No. 1, pp. 198–210.
- Browning, M., Deaton, A. and Irish, M. (1985, May). 'A Profitable Approach to Labor Supply and Commodity Demands over the Life-Cycle', *Econometrica*, Vol. 53, No. 3, pp. 503–543.
- Butcher, T. (2012). 'Still Evidence-Based? The Role of Policy Evaluation in Recession and Beyond: The Case of the National Minimum Wage', *National Institute Economic Review*, Vol. 219, pp. R26–R40.
- Card, D. (1991, May). 'Do Minimum Wages Reduce Employment? A Case Study of California, 1987–89', *Working Paper 3710*, National Bureau of Economic Research.
- Card, D. (2001). 'Immigrant Inflows, Native Outflows, and the Local Market Impacts of Higher Immigration', *Journal of Labor Economics*, Vol. 19, No. 1, pp. 22–64.
- Card, D. and Krueger, A. B. (1997). *Myth and Measurement: The New Economics of the Minimum Wage*, Princeton University Press, Princeton, New Jersey, USA.
- Card, D. and Lemieux, T. (2001). 'Can Falling Supply Explain the Rising Return to College for Younger Men? A Cohort-Based Analysis', *The Quarterly Journal of Economics*, Vol. 116, No. 2, pp. 705–746.
- Dolton, P., Bondibene, C. R. and Stops, M. (2015). 'Identifying the Employment Effect of Invoking and Changing the Minimum Wage: A Spatial Analysis of the UK', *Labour Economics*, Vol. 37, pp. 54–76.
- Gruber, J. and Wise, D. (2010). *Social Security Programs and Retirement Around the World: The Relationship to Youth Employment*, Illinois and London, University of Chicago Press.
- Hamermesh, D. S. (1993). *Labor Demand*, Princeton University Press, Princeton, New Jersey, USA.
- Hayes, R., Omerod, C. and Ritchie, F. (2007, 01). 'Earnings: Summary of Sources and Developments', *Economic Labour Market Review*, Vol. 1, No. 1, pp. 42–47.
- Holmlund, B. (2014). 'What do Labor Market Institutions Do?', *Labour Economics*, Vol. 30, pp. 62–69.
- Hyslop, D. and Stillman, S. (2007). 'Youth Minimum Wage Reform and the Labour Market in New Zealand', *Labour Economics*, Vol. 14, No. 2, pp. 201–230.
- Kalwij, A., Kapteyn, A. and de Vos, K. (2010). 'Retirement of Older Workers and Employment of the Young', *De Economist*, Vol. 158, No. 4, pp. 341–359.
- Katz, L. F. and Murphy, K. M. (1992, February). 'Changes in Relative Wages, 1963–1987: Supply and Demand Factors', *The Quarterly Journal of Economics*, Vol. 107, No. 1, pp. 35–78.
- Lemos, S. (2005). 'How Do Alternative Minimum Wage Variables Compare?', *Discussion Papers in Economics 05/6*, University of Leicester. Department of Economics.
- Low Pay Commission. (1998). *The National Minimum Wage: First Report of the Low Pay Commission 1998: presented to Parliament by the president of the Board of Trade by Command of Her Majesty June 1998*. [London]: Stationery Office.

- Low Pay Commission. (2010). National Minimum Wage: Low Pay Commission Report 2010: presented to Parliament by the Secretary of State for Business, Innovation and Skills by Command of Her Majesty March 2010. [London]: Stationery Office.
- Low Pay Commission. (2014). National Minimum Wage: Low Pay Commission Report 2014: presented to Parliament by the Secretary of State for Business, Innovation and Skills by Command of Her Majesty March 2014. [London]: Stationery Office.
- Manning, A. (2005). *Monopsony in Motion: Imperfect Competition in labor Markets*, Princeton University Press, Princeton, New Jersey, USA.
- Metcalfe, D. (2006). On the impact of the British National Minimum Wage on pay and employment (Working Paper 1481a). London: London School of Economics, Centre for Economic Performance.
- Mikusheva, A. and Poi, B. P. (2006). 'Tests and Confidence Sets with Correct Size When Instruments are Potentially Weak', *The Stata Journal*, Vol. 6, No. 3, pp. 335–347.
- Moreira, M. J. (2003). 'A Conditional Likelihood Ratio Test for Structural Models', *Econometrica*, Vol. 71, No. 4, pp. 1027–1048.
- Neumark, D. and Wascher, W. L. (2008). *Minimum Wages*, The MIT Press, Massachusetts, USA.
- OECD. (2013). *Employment Outlook*, Paris, OECD Publishing, Paris.
- Pereira, S. C. (2003, April). 'The Impact of Minimum Wages on Youth Employment in Portugal', *European Economic Review*, Vol. 47, No. 2, pp. 229–244.
- Stewart, M. B. (2002). 'Estimating the Impact of the Minimum Wage Using Geographical Wage Variation', *Oxford Bulletin of Economics and Statistics*, Vol. 64, pp. 583–605.
- Stewart, M. B. (2004a). 'The Employment Effects of the National Minimum Wage', *The Economic Journal*, Vol. 114, No. 494, pp. C110–C116.
- Stewart, M. B. (2004b). 'The Impact of the Introduction of the U.K. Minimum Wage on the Employment Probabilities of Low-Wage Workers', *Journal of the European Economic Association*, Vol. 2, No. 1, pp. 67–97.
- Stewart, M. B. (2011). 'The Changing Picture of Earnings Inequality in Britain and the Role of Regional and Sectoral Differences', *National Institute Economic Review*, Vol. 218, No. 1, R20–R32.
- Stewart, M. B. and Swaffield, J. K. (2006). 'The Other Margin: Do Minimum Wages Cause Working Hours Adjustments for Low-Wage Workers?', *The Warwick Economics Research Paper Series (TWERPS) 746*, University of Warwick. Department of Economics.
- Stock, J. H. and Watson, M. W. (2012). *Introduction to Econometrics: International Edition*, (3rd ed.), Pearson Education Limited, Harlow, England.
- Verbeek, M. and Nijman, T. (1993). 'Minimum MSE Estimation of a Regression Model with Fixed Effects from a Series of Cross-Sections', *Journal of Econometrics*, Vol. 59, Nos 1–2, 125–136.