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Departamento de Estadística
Universidad Carlos III de Madrid
Calle Madrid, 126
28903 Getafe (Spain)
Fax (34) 91 624-98-49

NATIONAL MINIMUM WAGE AND LABOUR MARKET OUTCOMES OF YOUNG WORKERS

Jan Fidrmuc and J.D. Tena

Abstract

We analyze the impact of the national minimum wage (NMW) in the UK on the employment of young workers. We utilize the regression discontinuity approach to assess the impact of age-related increases in the NMW when workers turn 18 and 22. The previous literature has found little evidence of an adverse impact of the NMW on the UK labour market, both when considering the age-related increases or the regular annual increases that apply to all NMW rates. We fail to find any effect of turning 22 on employment. However, we find a significant and negative effect of male workers turning 21. We also find a negative effect for both genders upon turning 18. The age-related NMW rates may have an adverse effect on employment of young workers, with this effect possibly occurring already well in advance of reaching the threshold age.

Keywords: minimum wage; employment; unemployment; young workers

Fidrmuc, Brunel University, and CESifo Munich. Contact information: Department of Economics and Finance, Brunel University, Uxbridge, UB8 3PH, United Kingdom. Email: Jan.Fidrmuc@brunel.ac.uk; Tena, Departamento de Estadística, Universidad Carlos III, C/Madrid 126. 28903 Getafe (Madrid), Spain. Email: jtena@est-econ.uc3m.es.
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1 Introduction

The imposition of a mandatory minimum wage, whether at national, regional or industry level, is a common instrument of economic policy: most OECD countries impose some form of a minimum wage (Dolton and Rosazza-Bondibene, 2011). Even Hong Kong, which traditionally espoused a laissez-faire approach to regulation, recently introduced a minimum wage. Nevertheless, the minimum wage is a contentious measure, one that is often blamed for raising workers' earnings at the expense of lowering employment. Indeed, standard neoclassical economic theory predicts that, under the assumption of competitive markets, a wage floor should either have no effect on employment (if set at a sufficiently low rate) or it should lower employment (by preventing the least productive workers from finding work at market-clearing wages).[1]

To date, the empirical evidence on the employment effect of the minimum wage is equally contentious. In a review, Neumark and Wascher (2007) argue that the bulk of studies point to a negative employment effect of introducing (or increasing) the minimum wage both in the US and in other countries. Workers who are most likely to be affected by the minimum wage, such as young workers and the low-skilled, are said to experience especially large disemployment effects (nevertheless, they find that the negative effect is mitigated somewhat when young workers are subject to a lower minimum wage rate). The range of estimated elasticities, however, is very broad: from significantly negative to significantly positive. This resonates with the findings of an earlier overview study by Dolado et al. (1996) who consider the employment effect of minimum wage rules in France, the Netherlands, Spain and the UK. Their results are inconclusive, with the estimated effects ranging from negative (especially for young workers again) to positive. The meta studies by Card and Krueger (1995b) and Doucouliagos and Stanley (2009), likewise, conclude that there is little evidence that the minimum wage causes lower employment. Hence, if anything characterizes the state of the current discourse on the employment effect of the minimum wage, it is lack of consensus.

Despite the lack of conclusive evidence of adverse employment effect of minimum wages, policy makers seem concerned about it, especially during the current economic and financial crisis. Ireland took a largely unprecedented step in February 2011 when it reduced the minimum wage by €1 (from €8.65 to €7.65, corresponding to a 12% cut); this move, however, was subsequently reversed in April 2011. When the reduction was announced, it was justified by the need to increase the labour market flexibility and ensure that Irish firms remain competitive during the crisis period. The bail-out package offered to Greece by the *troika* of the European Commission, ECB and IMF, similarly, stipulated a reduction in the minimum wage. In February 2012, Greece undertook to lower its minimum wage by 22% (the rate for under-25s was lowered even more

[1] Once we relax the assumption of competitive markets, however, the theoretical predictions can change dramatically. Assuming monopsony in the labor market, in particular, can result in a positive employment effect of the minimum wage (Dolado et al., 1996): monopsony employer can push wages below the marginal product of labor, thereby maximizing profits while depressing employment. Imposing a wage floor, correspondingly, reduces the employer's profits and increases employment.

dramatically, by 35%). Hence, it appears that, at least during the times of extraordinary economic hardship, policy makers assume that the relationship between employment and minimum wage is negative.[2]

The UK introduced the current national minimum wage (NMW) framework relatively late, in April 1999.[3] Thereafter, the NMW has been subject to regular annual revisions, coming into effect every October from 2000 onwards. Since shortly after its introduction, the effect of the NMW on employment has been analyzed by a number of studies. Stewart (2004) and Dickens and Draca (2005) considered the effect of the NMW's introduction and the impact of the annual minimum-wage increases in the subsequent years. Dolton, Rosazza-Bondibene and Wadsworth (2009) utilized the fact that, unlike the NMW rages, average earnings vary considerably across the regions of the UK. They used the resulting variation in the 'bite' of the NMW at the regional level to assess its impact on employment. Invariably, these studies (as well as others not cited here) find little evidence that the UK NMW has had an adverse effect on employment.

In this paper, we seek to contribute further to this discussion. We focus on a particular institutional feature of the UK minimum wage regulation: the existence of separate (lower) rates for young workers. At its introduction in 1999, besides the regular (adult) rate, the NMW specified a separate rate for those between 18 and 21 years of age (the so-called development rate).[4] In 2004, a separate rate was introduced for those aged 16 and 17 (who were exempt from the NMW until then). This allows employers to hire young workers at a discount relative to the adult rate. Specifically, the ratio between the adult rate and the development rate has remained approximately 1.2 since 1999. The ratio between the development rate and the 16-17 rate has been approximately 1.35. In turn, this means that young workers who earn only the NMW rate relevant for their age experience a sharp wage increase upon turning 18 and then again at 22.[5] While productivity is likely to increase with age, workers who are 22 are at best only marginally more productive than those who are 21 years old. Therefore, if the NMW affects employment, this should be especially apparent in the case of young workers.

The impact of age-related increases in the NMW on the employment of young workers in the UK has been investigated in earlier research by Dickens, Riley and Wilkinson (2010, henceforth DRW). They apply a regression-discontinuity approach to investigate how the employment status of low-skilled young workers

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[2] The lowering of the minimum wage is part of a broader effort aimed at facilitating *real* exchange rate depreciation at a time when the *nominal* exchange rate is irrevocably fixed. At a time when other prices are falling, keeping the nominal minimum wage frozen would be equivalent to increasing it in real terms.

[3] Up until 1993, the Wages Councils had the power to set minimum wages for specific industries (not all industries had a Wages Council). No minimum wage was in place in the period between 1993 and 1999.

[4] From October 2010, the upper limit for the development rate has been lowered to 20. The data used in our analysis, however, pertain to the period before this change.

[5] In most of our analysis, we focus on those subject to the 18-21 rate. The workers aged 16-17 differ from their older counterparts in several important ways: they are more likely to be in full-time education, their employability is lowered by restrictions such as not being allowed to sell alcoholic beverages and their eligibility to benefits is more limited. Therefore, it is difficult to discern whether any employment effects that may occur upon turning 18 are due to becoming eligible to the higher NMW rate or whether they are entirely attributable to the age effect.

changes when they turn 22. They find, somewhat surprisingly, that low-skilled young workers who turn 22 are significantly more likely to be employed and significantly less likely to be either unemployed or out of the labour force. Such a positive effect of becoming eligible for the higher NMW rate is counter intuitive. Nevertheless, they attribute this to an increase in labour supply by young workers: if the lower development rate is below the reservation wage of some workers, such workers postpone their labor market entry until after they can be certain of earning a sufficiently high wage. The positive employment effect for low-skilled workers, moreover, appears rather robust: they find no significant effect when workers reach 21 or 23 years of age (falsification tests) and neither do they find any effect at the age of 22 before the minimum wage was introduced in the UK. However, the result disappears when they consider all workers rather than only the low-skilled ones.[6] This is especially peculiar as they find the NMW to have a positive effect for the types of workers who are most likely to be paid the minimum wage and should therefore be more adversely affected than workers overall.

In this paper, we revisit and explore further the result of DRW (2010) with a somewhat longer data series. Our analysis differs from theirs in a number of potentially important aspects. First, we consider all workers rather than only low-skilled ones. Young workers are generally more likely to be subject to the minimum wage, more or less independently of their skill level. DRW indeed report that the shares of low and high skilled workers paid the minimum wage are only marginally different from one another: 10% of high skilled vs 11% of low skilled workers aged 21 earn less than the adult rate.[7] Second, while we follow DRW in implementing the discontinuity approach, we interpret the discontinuity effect differently. Specifically, since the effect of age on employment can be different before and after the discontinuity, we account for this when evaluating the effect of reaching the relevant age threshold. Finally, our analysis is based on extended data set relative to the one used by DRW.

While we replicate some of DRW's results, we also find important differences. In line with their result for all workers, we find that the effect of turning 22 on employment is not significant. Somewhat surprisingly, however, we find that male workers are less likely to remain employed after they turn 21. While reaching the age of 21 has no effect on the minimum wage, this finding may be consistent with employers anticipating the wage hike that would occur at 22 and shedding workers approaching that threshold well in advance of them reaching it. When considering the effect of turning 18, we find also a negative effect of turning 18; moreover, the negative effect is found both for males and females.

Our finding for males turning 21 reflects the specific nature of the framework that we are considering. The regression discontinuity approach is a *quasi-experimental* method of evaluating the effect of treatment that is assigned randomly (at least in approximation, see Lee and Lemiux, 2010, p. 283). In this case, the treatment and its timing is entirely deterministic: all young workers age (except Peter Pan) and they turn 22 in an amount of time that is perfectly known in advance to them and to their employers (as well as to their employers accountants

[6] Low skilled workers are defined as those whose qualifications are no higher than the GCSE exams (equivalent to incomplete high school).

[7] Table 3 (p. 26), Dickens, Riley and Wilkinson (2010).

and lawyers). It is therefore not surprising to see the effect of age take place even before the workers reach the age threshold.[8]

The next Section presents the data used in our analysis. The results of the discontinuity analysis are in Section 3. Section 4 concludes the paper by summarizing the results and suggesting some tentative avenues for further work.

2 Data

Our analysis is based on the UK Labour Force Survey (LFS). The LFS is a quarterly nationally-representative survey of households across the UK. Each quarterly file contains information on approximately 60 thousand households and over 100 thousand individuals aged 16 and above. Each household is retained in the survey for five consecutive quarters, with one-fifth of households replaced in each wave. The survey contains detailed demographic and socio-economic information on the respondents, including, importantly, their labour market outcomes. Since the NMW was introduced in April 1999, we use all quarterly datasets available since then. Our data thus span the period from April-June 1999 to October-December 2009, pooling all available LFS waves during this period.

The LFS contains information on the precise date of birth of every respondent.[9] We use this information to compute the age of each individual in months. Crucially, we also have the date the survey was carried out. By comparing these two dates, we can determine the precise age of each respondents in months, on the day when the survey was carried out. We thus know exactly whether a particular individual is 21 or 22 at the time of the survey, even when their birthday falls within the month in which they were interviewed. As is common in the regression-discontinuity literature, we redefine age so that it takes the value of 0 in the month when the individual reaches the threshold age of 22 (or 18) years.

Our treatment of age differs slightly from that implemented by DRW who only consider the year and month in which the respondent was born and compare this with the year/month when the survey was carried out. As a result, for each discrete age in months, some respondents are in fact falling short of that age according to their approach while all respondents are correctly aged in our analysis.[10] DRW therefore use the information on age in

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[8] In essence, this is similar to models of speculative crises (Salant and Henderson, 1978, and Krugman, 1979) which predict exchange rate crises rationally occur *in anticipation* of foreign exchange reserves being depleted and not when the reserves actually reach zero.

[9] This information is not available in the publicly released LFS datasets. We are grateful to the Low Pay Commission and the Office for National Statistics for making the restricted release of the LFS available to us.

[10] For example, consider the case when a group of respondents, all born in April (of any year) are interviewed on 15th April. When considering only the month and year of birth, it would appear that all of them have already passed their birthday. One needs to therefore use also the date of birth to determine the true age of each individual.

years, also contained in the LFS, to correctly classify those respondents who appear to have reached the threshold ages of 18 and 22 years, without similarly correcting the age in months of the remaining individuals.[11]

3 Employment Effect of NMW on Young Workers

To assess the impact of age-related MNW increases, we start by looking at individuals on either side of 22 years of age (corresponding to 264 months). For our baseline model, we consider individuals who are within 15 months of their 22nd birthday. Since, as explained above, we redefined age so that it takes the value of 0 when the respondent reaches the relevant threshold, our analysis considers young workers whose ages in months fall between -15 and 15 months. While each LFS quarterly data set contains information on around 100 thousand individuals, only a relatively small fraction of those are workers in the relevant ages. By considering a wider window (+/-15 months), we are effectively maximizing the number of observations that we can utilize in our analysis. As a robustness check, we replicate the analysis also for 12 and 6 month intervals.

The regression discontinuity approach allows us to estimate how labor-market outcomes are affected by whether a worker is older than the age from which the adult rate of NMW applies. Here, we use the econometric methodology proposed by Lee and Card (2008) to account for the uncertainty in the choice of functional forms in regression discontinuity designs in the case of discrete covariates. In this setting, it is no longer possible to estimate the impact of a covariate on the dependent variable by simply computing averages within arbitrarily small neighborhoods of the cutoff point, even with an infinite amount of data. Instead, it is necessary to choose a particular functional form for the model relating the outcomes of interest to the treatment-determining variable. Of course, it has to be tested whether the specification error of the proposed functional form is not significantly different form a fully flexible functional form that allows for different impacts of the discrete values of the covariate for each different age.

Our econometric specification is initially similar to that of DRW. We consider the effect of the discontinuity at 22 years of age on the probability of being employed, unemployed or inactive. In particular, let y_i be a variable equal to one if the individual is employed (unemployed, inactive). Then, we estimate the following equation:

$$E[emp \backslash age, dum] = \mathrm{F}(\theta + \alpha_0 * age_i + \alpha_1 * age_i^2 + \alpha_0^* * age_i * dum + \alpha_1^* * age_i^2 * dum + \beta * dum) = \mathrm{F}(u) \qquad (1)$$

where F is a normal distribution function, age_i is the age in months less the threshold (264 months for 22 years), *dum* is a dummy taking value of 1 when the individual is 22 or older and θ includes additional terms such as the constant and the selected covariates. More specifically, as in DRW we includes as explanatory variables the individual's qualifications, ethnic origin, apprenticeship, region of usual residence and whether the individual is a full time student. For our baseline results, age takes the form of a quadratic polynomial which we test against an alternatives fully-flexible specification with each age in months captured by a separate dummy. As is standard

[11] While in principle this methodological difference should be rather innocuous, it may be one of the reasons for some of our results differing from those of Dickens, Riley and Wilkinson

when using the regression discontinuity approach, we allow for the effect of age to be different before and after the young workers attain the threshold age. It is important to note that in this specification any change in employment probability associated with the discontinuity can stem either from the level effect (coefficient of the discontinuity dummy variable) or the slope effect (change in the effect of the age polynomial). Therefore, the effect of reaching the threshold age can be evaluating by combining the coefficient estimated for the discontinuity dummy with the different effects of age before and after the threshold[12]:

$$F(\theta+\beta)-F(\theta-\alpha_0^*+\alpha_1^*+\beta)-F(\theta)+F(\theta-\alpha_0+\alpha_1)$$

Table 1 reports regression results for the probability of being employed. We estimate separate regressions for males and females as well as for both genders together. The top panel presents results obtained while controlling for additional socio-economic characteristics while the bottom panel contain those obtained without additional controls. Unlike DRW, we consider all individuals, regardless of their skill level: as we argued above, young workers, skilled and unskilled, have very similar propensities to be paid the NMW. Specification (1) is tested against a fully flexible functional form. For men we cannot rejected that both specifications are significantly different at the conventional levels while for women the quadratic specification was rejected, in which case we also considered a cubic specification with no material change in the results. The row denoted *discontinuity* reports the combined effect of the discontinuity dummy and the change in the coefficient estimates for the age polynomial. *Dummy*, in contrast, stands for the coefficient estimated for the discontinuity dummy alone. DRW only consider the sign and significance of this latter coefficient, which we believe ignores a potentially important part of the discontinuity effect. However, neither the full discontinuity effect nor the discontinuity dummy on its own are significant. This is in line with the findings of DRW who also report an insignificant result when they include all individuals rather than only the low skilled ones. Hence, we find no significant effect, whether positive or negative, of turning 22 on young workers' employment.

For the sake of comparability, we replicate DRW's analysis of low-skilled workers, defined as those who left school at the age of 16 after completing their GCSEs and those who report having no qualifications. DRW found a significant positive effect of turning 22 for low-skilled workers, suggesting that becoming eligible for the adult NMW rate increases rather than reduced employment. Our results replicating their analysis are summarized in Table 2. They are broadly in line with those of DRW but weaker.[13] In particular, while the discontinuity dummy is always positive, it is never significant for females; it is significant for males and for all

[12] In non linear models, the marginal effect of a change in two interactive variables (age and dum) is not equal to the marginal effect of changing just the interaction term. Moreover, the sign may be different for different observations. Norton *et al.* (2004) explain how to compute interactive effects for probit models and we adapt this procedure to our particular case.

[13] Note that while we attempt to replicate DRW's results, there are some potentially important differences between their analysis and ours. In particular, we consider a 15-month window before/after the individual's 22nd birthday while they only consider 12 months, we compute the age in months slightly differently as discussed above, our data include three additional quarters in 2009, and, finally, although we sought to include the same covariates as them, it is possible that some of the covariates may be different or are formatted differently.

workers but only in the 5-10% range. More importantly, the combined effect of the discontinuity dummy and age polynomial is never even close to being significant. We are therefore unable to confirm their finding of a positive employment effect of turning 22 and becoming eligible for the adult NMW rate.

Next, Tables 3 and 4 present the regression results for unemployment and inactivity, considering once more all workers regardless of their skill level. Again, the full effect of the discontinuity is never significant. Note however that the dummy alone is significant and negative in the regressions for unemployment with all individuals: this mirrors the similar finding of DRW; as we argue above, accepting this as the true effect of the discontinuity would be wrong as it ignores the fact that the effect of the age polynomial is also allowed to change upon surpassing the age threshold.

In summary, we find thus no evidence that the approximately 20% increase in the rate of the NMW at the age of 22 has any effect – whether positive or negative – on young workers' employment, unemployment or inactivity. This conclusion does not depend on whether we consider all young workers or only the unskilled ones.

To probe the NMW effect on young workers further, we undertake a number of extensions. In Table 5, we consider the effect of turning 22 on employment conditional on the individual's employment status (employed, unemployed or inactive) in the previous quarter. It may well be that the increase in the NMW rate that applies to workers as they reach their 22nd birthday affects employed and unemployed workers differently: while some of those who were employed at 21 may lose their jobs, others may only enter the labour market or intensify their job search attracted by the higher wage. If this were the case, then the overall effect, presented in Table 1, could be insignificant because these two kinds of effects cancel each other out. The analysis is again presented separately for males and females (to save on space, we are omitting the results for both genders). In the first two columns of Table 5, we present the estimates for the probability of remaining employed, conditional on being employed already. The estimated effect of turning 22 is negative, especially for men, but it is not even close to being significant at conventionally accepted levels. Hence, young workers who were employed at the age of 21 are no more or less likely to be employed after their 22nd birthday. The next two columns present the estimates of the probability of being employed at 22, conditional on being unemployed before. The last two columns, in turn, present the corresponding estimates for those who were inactive before the quarter in which they turned 22. Again, none of these coefficients are significant, suggesting that controlling for the labour market status of young workers just before they turn 22 makes little difference to our findings.

In Table 6, we consider only those young workers who they earn less than the adult rate when they are 21. Such workers are bound to be affected by the age-mandated increase in the NMW upon turning 22. The previous analyses, in contrast, included all workers, regardless of whether their wages had to be raised. As before, we are unable to find any significant discontinuity effect on employment probability. One drawback of this analysis, however, is the rather small sample size.

As the last robustness check, we perform falsification tests for workers turning 21 and 23 (Table 7). The finding of no significant effect at 22 years of age may be either attributed to the NMW having no impact on employment, or it may indicate that the employment effect does not coincide with the workers' 22nd birthdays. In particular, employers may seek to dismiss workers in a way that could not be easily construed as motivated by the age-related NMW increase. If this is the case, then we might expect the employment effect to take place at some point before or after the workers turn 22. This is indeed what appears to happen: male workers are significantly less likely to remain employed after turning 21; in contrast, reaching their 23rd birthday has no significant impact on employment of males or females. The fall in employment probability at 21 years for men may be an anticipation effect: employers are aware of the age-related NMW increase that young workers are entitled to after their 22nd birthday and dismiss them well in advance of the relevant date. Note that this negative result only appears when we consider the combined effect of the discontinuity dummy and the coefficients for the age polynomial: the dummy alone is not significant. This again highlights the importance of assessing the full effect of the discontinuity and the changed effect of age rather than considering only the coefficient of the discontinuity dummy.

Finally, we replicate the discontinuity analysis at 21st, 22nd and 23rd birthday with 6 and 12 month estimation windows instead of 15 months used to generate the results discussed so far. In general, the regressions generate weaker results (see the Appendix). Those obtained with the 6 month window are never significant. This may be due to the lower number of observations with the shorter estimation window. Moreover, the discontinuity effect may take time to become sufficiently pronounced. The regressions utilizing the 12 month window generally paint the same picture as those discussed above. In particular, the discontinuity effect is negative both at the age of 21 and 22 for males; the former is marginally significant at 10% while the latter is not significant.

The finding of a significant negative effect for males at the age of 21 is interesting and perplexing at the same time. Therefore, we pursue it further and consider the discontinuity for every age in one-month increments between 20 and 23 years. Since we estimate dozens of coefficients, it is more instructive to depict the results graphically. Figure 1 presents the results for males. The solid line captures the employment effect while the dotted lines correspond to the 95% confidence interval. An interesting pattern emerges: the probability of being employed dips in the neighborhood of both the 21st and 22nd birthdays (252 and 264 months, respectively). Only the first dip is significant, however, suggesting that young male workers are significantly less likely to be employed as they approach their 21st birthday. The employment probability rebounds subsequently: just after the 20th birthday, then again at approximately at 21.5 years of age and at around 22.5. The first two of those peaks are significantly different from zero whereas being between 22 and 23 is not associated with a significant increase in the employment probability. The estimates for women are strikingly different, as Figure 2 shows: none of the age effects between 21 and 23 years are significant. Moreover, as is also the case in Table 1, the quadratic age polynomial is rejected by the model, as is also the cubic alternative (the latter results, presented in Figure 3, also yield insignificant effects).[14]

We can only speculate what drives these results. The age-related NMW rates apply equally to men and women yet we only observe age-related effects for the former. This may reflect the fact that the labour market positions of men and women are substantially different from each other. As we argued above, the negative effect around men's 21st birthday may be due to anticipation effects whereby employers choose to dismiss workers well in advance of the age-related NMW increase. Moreover, as the age-related NMW increases follows an entirely deterministic process that all young workers are subject to, the effect on employment can indeed occur at

[14] Because of the insignificant results obtained for ages between 21 and 23 and also in the light of the quadratic/cubic polynomial being rejected, we did not extend the analysis for women to their 20th birthday.

any time before the discontinuity: an employer seeking to avoid employing 22 years olds likewise has little incentive to hire someone who is 21 or 21.5 years old.

An alternative explanation could be that the negative effect around the 21st birthday is due to an influx of university graduates into the job market which increases the competition for jobs. However, while it is true that university students graduate when they are 21 (assuming they went to university immediately after completing secondary education), the bulk of them enter the job market in the summer or autumn after graduation. They would therefore reach 21 years of age during their final year in university and only a small fraction of them would be turning 21 exactly at the time when they graduate.

Finally, we also consider the NMW threshold at 18 years of age. Recall that those turning 18 become eligible for the development rate which historically has been some 35% above the 16-17 rate. As before, we consider all workers, irrespective of skills (although the differences in skill levels at this age are not particularly large). Table 8 reports the results. The effect of turning 18 appears significantly negative for both genders: becoming eligible for the higher NMW rate is associated with lower employment probability. Note that again this negative effect becomes apparent only when we consider both the coefficient estimated for the discontinuity dummy and the changed effect of the age polynomial: the dummy itself is not significantly different from zero (except for females). The insignificant coefficient for the discontinuity dummy is in line with the finding of DRW. The differences in the conclusions reached when considering the discontinuity dummy only and when looking also at the changed effects of the age polynomial again underscores the importance of assessing the full effect of the discontinuity.

As we argued before, turning 18 is associated with a host of other important changes besides becoming eligible for a higher NMW rate. For example, UK law requires anyone selling or serving alcohol to be 18 or older, which makes those under 18 ineligible to work in bars, restaurants and many shops. This makes the negative effect that we found all the more remarkable. Again, an alternative explanation would link the effect that we observe to the end of full-time secondary education. In the UK, education is currently compulsory until the age of 16 but many students stay enrolled for another two years to complete their secondary education. Those who do so without enrolling in higher education upon graduating then generally enter the job market when aged 18. Nevertheless, as with university graduates, few would leave full time education close to their 18th birthday. Rather, the students in their final year of secondary education turn 18 over the course of their last year. Therefore, the negative employment effect is unlikely to be attributable to changes in participation in education.

Note that our analysis is based on the estimation of the functional form presented in expression (1). However, it is also relevant to study if main conclusions in the paper still are upheld when we adopt an specification similar to (1) but imposing the restrictions $\alpha_0 = \alpha_0^*$ and $\alpha_1 = \alpha_1^*$. Although we prefer specification (1) because it already encompassed this restricted case and also it allows for comparison with DRW, the constrained version of the model has the advantage that restrictions are accepted at the usual confidence levels and that the discontinuity impact can be easily computed as the marginal effect of the dummy variable dum without taking consideration of the interaction effect between age and dum. Under the restricted model for all workers, we also find strong evidence of the negative effect of NMW on the probability of employment at 18 (-0.02 with a p-value of 0.001). Moreover, the falsification impact at 21 is negative but only marginally significant

(-0.009 with a p-value of 0.097). Finally, we find no significant effect at 22 (0.002 with a p-value of 0.39) and a significant and positive impact at 23 (0.01 with p-value:0.009).

Hence, our results suggest that the age related NMW rates are indeed having an adverse effect on the employment of young workers. However, the nature of the effect is not entirely straightforward. In particular, it is possible that, due to anticipation effects, employers dismiss workers in advance of them reaching the age when the higher wage is supposed to take effect.

5 Conclusions

The received wisdom concerning the effect of the UK national minimum wage is that it has had little adverse impact on employment. In this paper, we revisit this issue. We consider young workers and investigate whether their employment prospects are affected by the fact that different rates apply to different age groups: the minimum wage rates are different for those who are 16-17, 18-21 and above 22 years old. Using the regression-discontinuity approach, we find that although the effect of turning 22 is negative, it is not statistically significant. This contrasts with an earlier finding by Dickens, Riley and Wilkinson (2010) who argued that becoming eligible for the higher adult rate from the age of 22 increases the employment of unskilled young workers. We believe their finding is potentially flawed because they do not take into account the fact that the effect of age on employment probability may also change at discontinuity. Specifically, their analysis (as ours) allows the discontinuity to affect the dependent variable through the coefficient of the discontinuity dummy as well as by allowing age to have a different effects before and after the discontinuity. Dickens, Riley and Wilkinson only consider the former effect. When we account for the combined effect of the dummy and the changed effect of age, we find that turning 22 has no effect on the employment of young workers, whether they are unskilled or skilled.

In contrast, we do find evidence of a negative employment effect for males at the age of 21. While in the period we have studied the NMW does not change at this age, we believe this result may be driven by the anticipation of the later minimum wage increase at 22. This reflects the specific nature of the case that we, and Dickens, Riley and Wilkinson, consider. While the regression discontinuity approach is usually used to study the effects of outcomes that are assigned (approximately) randomly, there is nothing random about the outcome in this case: all young workers eventually turn 22. The effect associated with the discontinuity (higher NMW rate applying to those aged 22 and above) therefore can occur anywhere in the neighborhood of the discontinuity age, whether before or after. The fact that we find a negative effect approximately one year before we initially expected it to occur intuitively makes sense. The cost of hiring a 21-year old is substantially lower only for employers seeking short-term staff; those wishing to retain this worker in the long term would enjoy only a temporary cost advantage.

Our findings thus suggest that the age specific minimum wage rates do affect employment. This is confirmed also by our finding that both genders experience a negative employment effect at the age of 18, when they become eligible for the 18-21 NMW rate (35% higher than the 16-17 rate).

The UK NMW rules concerning young workers were modified in October 2010 in that the threshold age for the adult rate has been lowered from 22to 21. Future research will show whether this has hurt the employment prospects of young workers. Our findings would suggest that the age at which this effect occurs may shift further so that even workers younger than 21 may see their employment prospects affected.

Finally, our work has two important methodological implications. First, it underscores that when applying the regression discontinuity approach to non-random deterministic processes through time, the effect need not coincide with the discontinuity. Instead, it can occur either before or after the discontinuity is reached. Second, it is important to correctly account for the effect of the regression discontinuity in cases when it can entail both level and slope effects. In particular, the negative employment effects that we find at 18 and 21 are only apparent when we consider both the coefficient estimated for the discontinuity dummy and the change in the coefficients for age after the discontinuity.

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Table 1 Discontinuity Effect on Employment: All Young Workers. Marginal effects at mean values and standard deviations between brackets.

	All		Males		Females	
	with covariates	without covariates	with covariates	without covariates	with covariates	without covariates
Discontinuity[(1)]	.00122 (.00244)	.00227 (.00236)	-.00228 (.00331)	.00055 (.00328)	.00368 (.00353)	.00356 (.00336)
Dum[(2)]	.00482 (.00800)	.00480 (.00772)	.00567 (.01097)	.00502 (.0107)	.00589 (.01154)	.00348 (.01103)
No. observations	136,591	136,591	66,582	66,582	70,009	70,009
Chi-statistic for Whole regression	26345.97	638.70	15412.56	480.74	12942.46	218.54
Pr>Chi	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
R2	0.1524	0.0037	0.1918	0.0060	0.1411	0.0024
Chi-statistic for quadratic	27.11	29.11	27.55	. 34.08	44.13	53.25
Pr>Chi	0.3503	0.2539	0.3292	0.1063	0.0105	0.0008

Notes: (1) estimated discontinuity effect taking into account the impact of age and the threshold dummy variable; (2) estimated impact of the threshold dummy variable. Significance levels denoted as * 5% and ** 1%. Source: Labour Force Survey.

Table 2 Discontinuity Effect on Employment: Low Skilled Young Workers. Marginal effects at mean values and standard deviations between brackets.

	All		Males		Females	
	with covariates	without covariates	with covariates	without covariates	with covariates	without covariates
Discontinuity[(1)]	.00211 (.00418)	.00224 (.00415)	.00214 (.00555)	.00270 (.00561)	.00061 (.00595)	.00193 (.00589)
Dum[(2)]	.02940 (.01402)*	.02241 (01386)	.03380 (.01852)	.02807 (.01859)	.02486 (.02002)	.01822 (.01971)
No. observations	43809	43809	20457	20457	23352	23352
Chi-statistic for Whole regression	2686.26	3.24	1621.56	42.32	1174.80	14.47
Pr>Chi	0.0000	0.6633	0.0000	0.0000	0.0000	0.0129
R2	0.0478	0.0001	0.0705	0.0018	0.0370	0.0005
Chi-statistic for quadratic	45.31	43.99	24.89	30.52	61.38	58.20
Pr>Chi	0.0077	0.0109	0.4683	0.2054	0.0001	0.0002

Notes: (1) estimated discontinuity effect taking into account the impact of age and the threshold dummy variable; (2) estimated impact of the threshold dummy variable. Significance levels denoted as * 5% and ** 1%. Source: Labour Force Survey.

Table 3 Discontinuity Effect on Unemployment. Marginal effects at mean values and standard deviations between brackets.

	All		Males		Females	
	with covariates	without covariates	with covariates	without covariates	with covariates	without covariates
Discontinuity[1]	.00118 (.00126)	.00107 (.00135)	.00190 (.00195)	.00175 (.00212)	.00037 (.00160)	.000200 (.00170)
Dum[2]	-.008830 (.00425)*	-.00919 (.00452)*	-.01013 (.00659)	-.01104 (.0071)	-.00844 (.00535)	-.00819 (.00565)
No. observations	136,591	136,591	66,582	66,582	70,009	70,009
Chi-statistic for Whole regression	3489.80	61.34	2721.18	44.54	1170.22	15.95
Pr>Chi	0.0000	0.0000	0.0000	0.0000	0.0000	0.0070
R2	0.0446	0.0008	0.0621	0.0010	0.0347	0.0005
Chi-statistic for quadratic	19.40	15.69	26.00	23.85	23.16	20.95
Pr>Chi	0.7776	0.9237	0.4078	0.5278	0.5682	0.6955

Notes: (1) estimated discontinuity effect taking into account the impact of age and the threshold dummy variable; (2) estimated impact of the threshold dummy variable. Significance levels denoted as * 5% and ** 1%. Source: Labour Force Survey.

Table 4 Discontinuity Effect on Inactivity. Marginal effects at mean values and standard deviations between brackets.

	All		Males		Females	
	with covariates	without covariates	with covariates	without covariates	with covariates	without covariates
Discontinuity[(1)]	-.00151	-.00347	.00038	-.00252	-.00451	-.00389
	(.00160)	(.00220)	(.00249)	(.00291)	(.00334)	(.00323)
Dum[(2)]	.00539	.00444	.00695	.00615	.00287	.00474
	(.00698)	(.00705)	(.00819)	(.00919)	(.01072)	(.01047)
No. observations	136,591	136,591	66,582	66,582	70,009	70,009
Chi-statistic for Whole regression	29973.84	541.74	20380.64	446.08	13752.84	189.13
Pr>Chi	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
R2	0.1971	0.0036	0.3135	0.0069	0.1614	0.0022
Chi-statistic for quadratic	21.83	25.18	27.69	24.00	30.59	46.73
Pr>Chi	0.6455	0.4521	0.3225	0.5194	0.2030	0.0053

Notes: (1) estimated discontinuity effect taking into account the impact of age and the threshold dummy variable; (2) estimated impact of the threshold dummy variable. Significance levels denoted as * 5% and ** 1%. Source: Labour Force Survey.

Table 5 Probability of Employment Conditional on Employment Status in Previous Quarter. Marginal effects at mean values and standard deviations between brackets.

	All		Males		Females	
	with covariates	without covariates	with covariates	without covariates	with covariates	without covariates
Discontinuity[(1)]	-.00184 (.00158)	-.00004 (.00181)	-.01189 (.00936)	.01636 (.01102)	.00030 (.00663)	-.00500 (.00518)
Dum[(2)]	.00483 (.00822)	.00114 (.00843)	-.01864 (.04345)	.01636 (.05514)	.03364 (.02418)	.02886 (.01552)
No. observations	27921	26030	3956	2671	6795	11815
Chi-statistic for Whole regression	42.09	30.76	7.89	11.21	7.48	10.13
Pr>Chi	0.0000	0.0000	0.1625	0.0473	0.1876	0.0716
R2	0.0037	0.0029	0.0017	0.0033	0.0016	0.0014

Notes: (1) estimated discontinuity effect taking into account the impact of age and the threshold dummy variable; (2) estimated impact of the threshold dummy variable. Significance levels denoted as * 5% and ** 1%. Source: Labour Force Survey.

Table 6 Probability of Employment for Workers Earning Less than Adult Rate. Marginal effects at mean values and standard deviations between brackets.

	Males	Females
Discontinuity[(1)]	.000242 (.01783)	-.00684 (.01279)
Dum[(2)]	.014173 (.04104)	.008331 (.03483)
No. observations	1365	1931
Chi-statistic for Whole regression	4.06	7.96
Pr>Chi	0.5404	0.1582
R2	0.0047	0.0066

Notes: None of the estimations include covariates. (1) estimated discontinuity effect taking into account the impact of age and the threshold dummy variable; (2) estimated impact of the threshold dummy variable.

Significance levels denoted as * 5% and ** 1%. Source: Labour Force Survey. The regressions do not contain additional control variables due to low number of observations.

Table 7 Falsification Tests: Discontinuity Effects at 21 and 23. Marginal effects at mean values and standard deviations between brackets.

	21 years		23 years	
	Males	Females	Males	Females
Discontinuity[(1)]	-.00994 (.00326)**	-.001039 (.00349)	.00435 (.00318)	-.00179 (.00336)
Dum[(2)]	-.00764 (.01150)	-.00186 (.01184)	.01043 (.01023)	-.01325 (.01138)
No. observations	68324	70647	65206	70622
Chi-statistic for Whole regression	17001.14	12155.02	13443.49	14310.83
Pr>Chi	0.0000	0.0000	0.0000	0.0000
R2	0.1947	0.11285	0.1879	0.1602

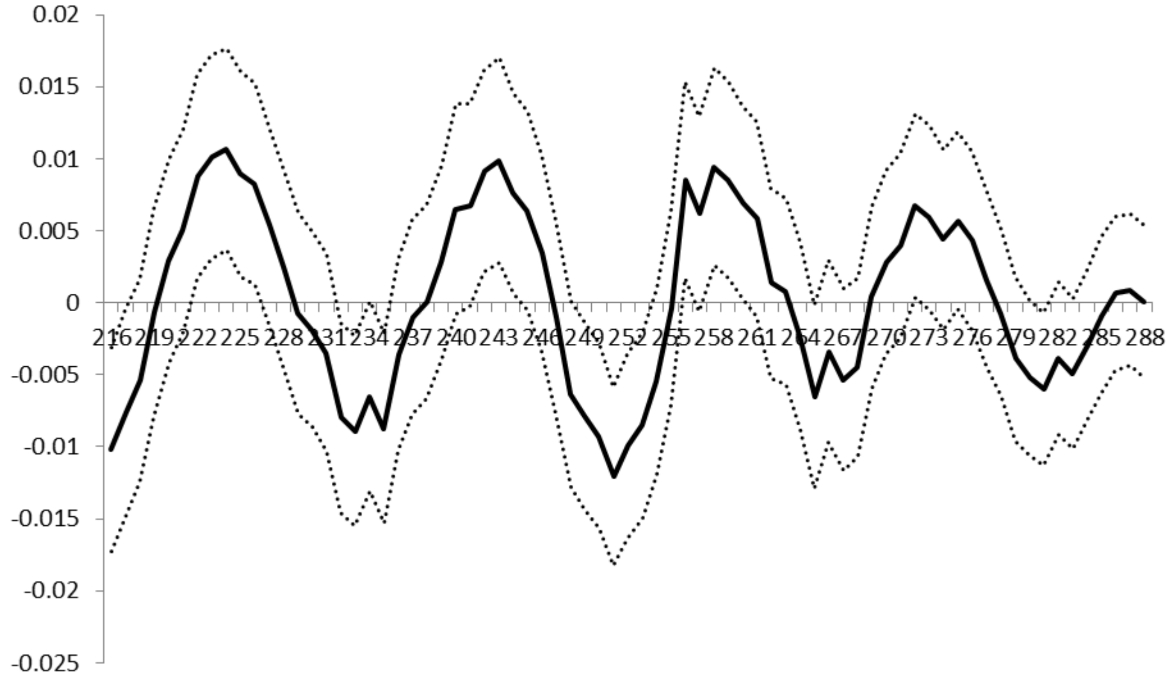
Notes: All the estimations include covariates. (1) estimated discontinuity effect taking into account the impact of age and the threshold dummy variable; (2) estimated impact of the threshold dummy variable. Significance levels denoted as * 5% and ** 1%. Source: Labour Force Survey.

Table 8 Discontinuity Effects at 18. Marginal effects at mean values and standard deviations between brackets.

	Males	Females	All
Discontinuity[(1)]	-0.01018	-.01009	-0.00984
	(0.00361)**	(.00362)**	(0.00255)**
Dum[(2)]	-0.00238	-.0253495	-.012706
	(0.01253)	(.01263)*	(0.00888)
No. observations	67641	65023	132664
Chi-statistic for Whole regression	16587.27	9896.45	25665.83
Pr>Chi	0.0000	0.0000	0.000
R2	0.1788	0.1110	0.1410

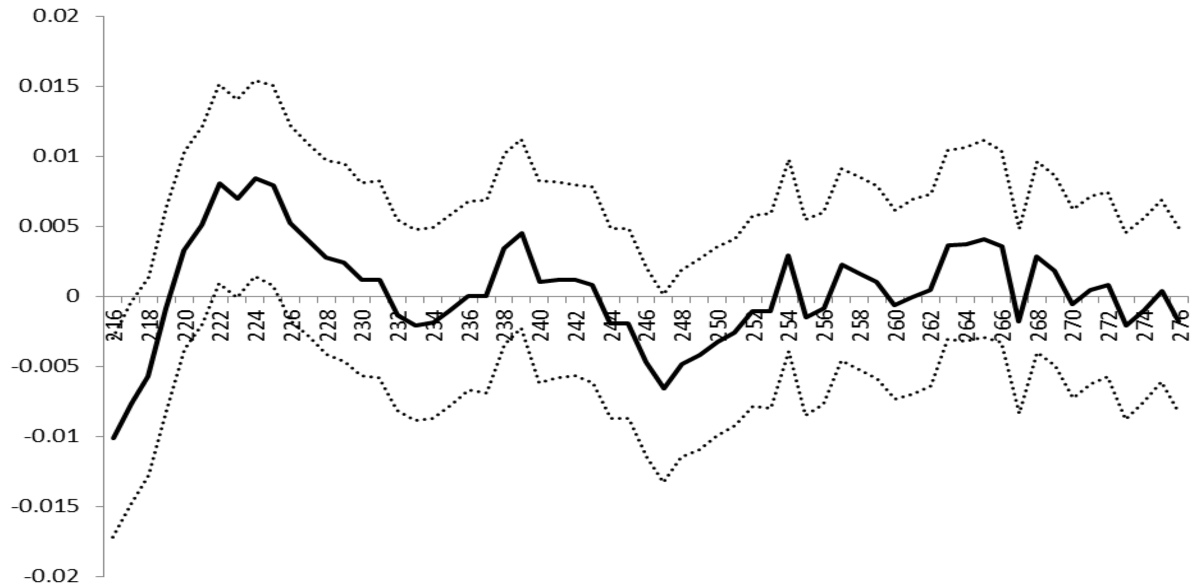
Notes: All the estimations include covariates. (1) estimated discontinuity effect taking into account the impact of age and the threshold dummy variable; (2) estimated impact of the threshold dummy variable. Significance levels denoted as * 5% and ** 1%. Source: Labour Force Survey.

Figure 1 Discontinuity Effects by Month, Ages 18 to 24, Males



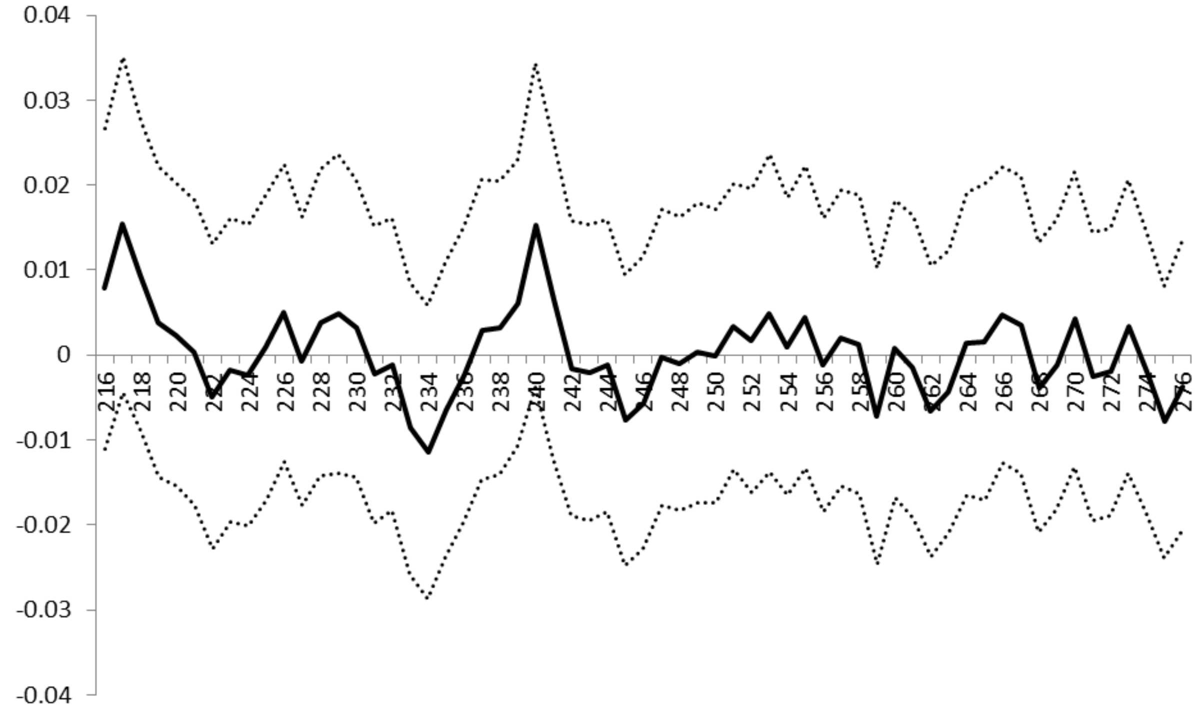
Notes: The points at which birthdays occur are: 18 years (216 months), 19 (228), 20 (240), 21 (252), 22 (264), 23 (276) and 24 years (288 months). Dotted lines represent the 95% confidence interval.

Figure 2 Discontinuity Effects by Month, Ages 18 to 24, Females (quadratic age polynomial)



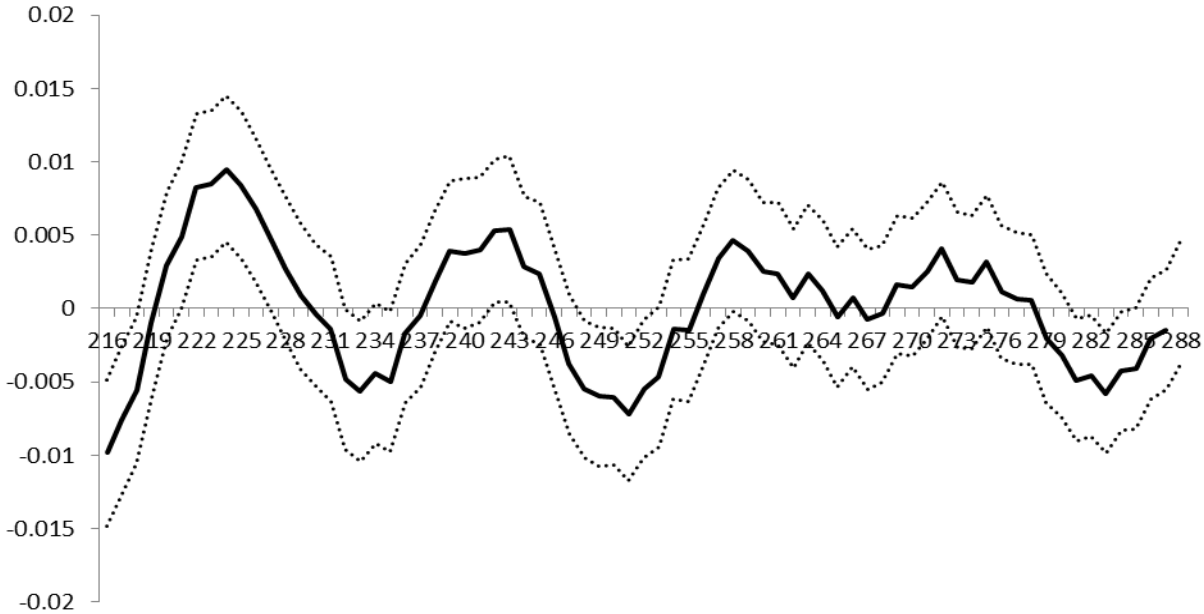
Notes: The points at which birthdays occur are: 18 years (216 months), 19 (228), 20 (240), 21 (252), 22 (264), 23 (276) and 24 years (288 months). Dotted lines represent the 95% confidence interval.

Figure 3 Discontinuity Effects by Month, Ages 18 to 24, Females (cubic age polynomial)



Notes: The points at which birthdays occur are: 18 years (216 months), 19 (228), 20 (240), 21 (252), 22 (264), 23 (276) and 24 years (288 months). Dotted lines represent the 95% confidence interval.

Figure 4 Discontinuity Effects by Month, Ages 18 to 24, Both Genders



Notes: The points at which birthdays occur are: 18 years (216 months), 19 (228), 20 (240), 21 (252), 22 (264), 23 (276) and 24 years (288 months). Dotted lines represent the 95% confidence interval.

Appendix

Regression-discontinuity analysis: Alternative time windows

Total workers. Discontinuity Effects at 21, 22 and 23

	21 years		22 years		23 years	
	6 months	12 months	6 months	12 months	6 months	12 months
Discontinuity[1]	.00092	-.00461	.00116	-.00045	-.00961	.00096
	(.00969)	(.00350)	(.00965)	(.00350)	(.00891)	(.00334)
Dum[2]	.01341	-.00430	.01026	.01483	-.01239	-.00188
	(.01425)	(.00945)	(.01395)	(.02617)	(.01323)	(.00876)
No. observations	57797	109453	57513	108102	56417	107005
Chi-statistic for Whole regression	11048.03	21478.97	11245.37	20836.73	10430.78	19855.19
Pr>Chi	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
R2	0.1458	0.1496	0.1536	0.1520	0.1563	0.1562

Notes: All the estimations include covariates. (1) estimated discontinuity effect taking into account the impact of age and the threshold dummy variable; (2) estimated impact of the threshold dummy variable. Significance levels denoted as * 5% and ** 1%. Source: Labour Force Survey.

Male workers. Discontinuity Effects at 21, 22 and 23

	21 years		22 years		23 years	
	6 months	12 months	6 months	12 months	6 months	12 months
Discontinuity[1]	.01042	-.00883	-.00024	-.00239	.01077	.00532
	(.01352)	(.00476)	(.00793)	(.00479)	(.01269)	(.00459)
Dum[2]	.02918	-.00307	.00052	-.00303	-.00365	.00668
	(.01976)	(.01316)	(.01919)	(.01260)	(.01750)	(.01159)
No. observations	28583	53899	27978	52724	27086	51396
Chi-statistic for Whole regression	6610.71	13098.40	6656.79	12248.60	5547.02	10567.76
Pr>Chi	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
R2	0.1812	0.1900	0.1955	0.1919	0.1885	0.1888

Notes: All the estimations include covariates. (1) estimated discontinuity effect taking into account the impact of age and the threshold dummy variable; (2) estimated impact of the threshold dummy variable. Significance levels denoted as * 5% and ** 1%. Source: Labour Force Survey.

Female workers. Discontinuity Effects at 21, 22 and 23

	21 years		22 years		23 years	
	6 months	12 months	6 months	12 months	6 months	12 months
Discontinuity(1)	-.00925 (.01389)	-.00136 (.00508)	-.00665 (.01375)	.01457 (.01321)	-.01932 (.01955)	-.00362 (.00484)
Dum(2)	-.00170 (.02049)	-.00589 (.01353)	.02335 (.02011)	.00031 (.00506)	-.02845 (.01264807)	-.01020 (.01295)
No. observations	29214	55554	29535	55378	29331	55609
Chi-statistic for Whole regression	5040.66	9529.44	5505.22	10287.81	5987.72	11228.77
Pr>Chi	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000
R2	0.1290	0.1282	0.1417	0.1417	0.1628	0.1602

Notes: All the estimations include covariates. (1) estimated discontinuity effect taking into account the impact of age and the threshold dummy variable; (2) estimated impact of the threshold dummy variable. Significance levels denoted as * 5% and ** 1%. Source: Labour Force Survey.

Not-for-publication Appendix. Computation of Marginal effects.

There is a discontinuity in equation (1) when $age_i = 0$ and the derivative $F'(u)$ is not defined for that point. In this case, following Norton et al. (2004), the discontinuity can be computed as

$$\frac{\Delta \frac{\Delta F(.)}{\Delta age}}{\Delta dum} = \frac{\Delta[F(\theta + \beta * dum) - F(\theta - \alpha_0 * (1 - dum) + \alpha_1 * (1 - dum) - \alpha_0^* * dum + \alpha_1^* * dum + \beta * dum)]}{\Delta dum}$$

$$= F(\theta + \beta) - F(\theta - \alpha_0 + \alpha_1 - \alpha_0^* + \alpha_1^* + \beta) - F(\theta) + F(\theta - \alpha_0 + \alpha_1)$$