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National minimum wage and employment of young workers in the UK*

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) and many less developed countries do likewise (even Hong Kong, traditionally a bastion of the laissez-faire approach, introduced a minimum wage in 2010). Nevertheless, the minimum wage is a contentious measure, one that is often blamed for raising workers' earnings at the expense of worsening employment prospects for those out of work. Indeed, standard neoclassical economic theory predicts that, under 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 inconclusive. In a review, Neumark and Wascher (2007) argue that the bulk of the evidence points 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, 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 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) 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 lowers employment. Rohlin (2011) considers US firms' location choices and finds that increasing the state minimum wage discourages new firms from locating in the state. Hence, if anything characterizes the current state of the discourse on the employment effect of the minimum wage, it is a lack of consensus.

[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.

The UK introduced the current national minimum wage (NMW) framework relatively late, in April 1999.[2] Thereafter, the NMW has been subject to regular annual revisions, coming into effect every October from 2000 onwards. Since its introduction, the effect of the NMW on employment has been analyzed by a number of studies. Stewart (2004) and Dickens and Draca (2005) consider 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) utilize the fact that, unlike the NMW rages, average earnings vary considerably across the regions of the UK. They use 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. The main (and probably only) exception is the recent study by Dickens, Riley and Wilkinson (2012) who present evidence that the introduction and annual NMW increases reduce the employment of part-time women, a segment of the labor market that is especially exposed to the minimum wage.

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, the NMW was formulated with two distinct rates: the adult rate for workers aged 22 and over, and the so-called development rate for those between 18 and 21 years of age.[3] In 2004, an additional rate was introduced for those aged 16 and 17 who were not subject to the NMW until then. 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. This means that young workers earning the NMW rate relevant for their age are subjected to a sharp wage increase upon turning 18 and then again at 22. While productivity is likely to increase with age, workers who are 22 or 18 are at best only slightly more productive than those one year younger. Nevertheless, employers face a sharp increase in the wage bill for the workers attaining these ages. The objective of this paper is to determine whether this fact has any negative effect on the employmentw['bce tho,)5.4ffect discontinuity design (henceforth RDD; see Imbens and Lee, 2008; van der Klaauw, 2008; Lee and Lemieux, 2010). Arguably, the characteristics of workers on either side of the cutoff age are very similar and therefore the main difference between them is the applicable NMW rate.[4] The forcing variable, age, can be influenced neither by the workers nor by their employers (or anyone else, for that matter). Therefore, when comparing workers who are just above the cutoff age and those just below, the difference between them is as good as random. The 'treatment' category then consists of young workers older than the cutoff age while the rest constitute the 'control' group.

Our work extends the earlier research by Dickens, Riley and Wilkinson (2010, henceforth DRW) who consider the effect of age-related increases in the NMW on the employment of low-skilled young workers in the UK using also the regression discontinuity design. They find, somewhat surprisingly, that low-skilled young workers are significantly more likely to be employed and significantly less likely to be either unemployed or out of the labor force as they turn 22. They attribute this to an increase in their labor supply: if the development rate is below the reservation wage of some workers, the discontinuity effect can take two forms: besides the usual level (jump) effect, there can be a slope (kink) effect (see Dong, 2012, for theoretical formulation). Third, we recognize that while the case that we consider has some quasi-experimental properties (the fact that workers just below and just above the cutoff age are very similar in all characteristics other than the NMW rate that applies to them), it is also a case where the treatment occurs due to a deterministic rather than random process (aging).[7] Therefore, young workers' labor market outcomes can be affected by anticipation of the higher NMW rate already before they reach the cutoff age. To account for this, we estimate the discontinuity effect not only at the cutoff ages of 18 and 22 but for every month of age between 18 and 23. Finally, and rather trivially, our analysis is based on an extended data set relative to the one used by DRW.[8]

Our results are intriguing. In contrast to DRW (2010), we find that turning 22 has no effect on employment. Instead, and somewhat surprisingly, we find that male workers are less likely to be employed when they are around 21 years old. This finding is consistent with employers anticipating the wage hike that would occur at 22 and dismissing or not hiring workers approaching that threshold. In addition, we find also a negative effect of turning 18; moreover, the negative effect is found both for males and females at this age.

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. dat e4r8 what cio-.000-Tc8 Tc.0796 Tw[outcom)8-6smn

being employed, unemployed or inactive at the cutoff age. We estimate the following equation:

where y_i is equal to one if the individual is employed (unemployed, inactive), F is a standard normal cumulative distribution function, age_i is the age in months less the cutoff, d is a dummy variable equal to one when the individual's is at the cutoff age or older and θ again includes any remaining terms such as the constant and the covariates (qualifications, ethnic origin, apprenticeship, region of usual residence and being full time student). We allow for the effect of age to be different before and after the young workers attain the threshold age. This is standard in the regression discontinuity approach, reflecting the fact that the effect of the forcing variables may change after the cutoff. If we did not allow different slope coefficients, the pre-cutoff and post-cutoff relationships would be estimated using information contained in the both parts of the sample: those pertaining to the pre-treatment sub-sample would be estimated using information affected by the treatment and vice versa (see Lee and Lemieux, 2010). 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.

In expression (4), the jump in the probability of a particular employment status at the cutoff point (level effect) is measured as the marginal effect associated with the disctontinuity dummy, . However, because F is a non-linear (probit) function, computing the change in the slope is more complicated than merely comparing the coefficients of the age polynomial before and after the cutoff (and vs and). Norton *et al.* (2004) show how to evaluate the marginal effect for probit models and we adapt this procedure to our particular case:

Note that we evaluate this expression by double-differentiating the functional form at equal 0 and -1 and at equal 1 and 0. For robustness we also treat as a continuous variable and compute the slope change as the difference of the derivative of the response function at equal 1 and 0 but it does not change our findings (these results are available under request).

An important issue to point out is that in our particular model this interaction effect could be nonzero even if and . This is because of non-linearity which implies that the marginal effect of age depends also on the parameter . Therefore, expression (5) provides a more complete picture of the discontinuity effect than that provided simply by in the traditional level effect. The traditional approach, instead, focuses only on the jump effect and thus ignores the fact that the slope coefficient can change at the cutoff as well. The traditional approach, instead, focuses only on the jump effect and thus ignores the fact that the slope coefficient can change at the cutoff as well.

3 NMW and 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). Table 1 reports regression results for the probability of being employed. We present estimates for males and females separately as well as for both genders together, and with and without additional covariates. Unlike DRW (2010), we consider all individuals, regardless of their skill level: as we argued above, both skilled and unskilled young workers have very similar propensities to be paid the NMW. Specification (4) is tested against a fully flexible f

2. They are broadly in line with those of DRW but somewhat weaker.[10] In particular, while the discontinuity dummy is always positive, it is never significant for females, and for males and for all workers it is significant only in the 5-10% range. More importantly, the combined level and slope effect 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.

at conventionally accepted levels. Hence, young workers who were employed at the age of 21 are not 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 labor 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 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 or not. As before, we are unable to find any significant discontinuity effect (level or slope) on employment probability. One drawback of this analysis, however, is the rather small sample size, which may be responsible for the lack of significant results.

As the last robustness check, we repeat the discontinuity analysis for workers turning 21 and 23 years of age (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, age is a deterministic process and employers can take action motivated by workers reaching a particular age before or after they actually attain that age. This is indeed what appears to happen: the slope effect suggests that 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. Note that this negative result only appears when we consider the slope effect; the level effect is not significant. This again highlights the importance of assessing both effects of the discontinuity rather than considering only the coefficient of the discontinuity dummy.[11]

The fall in employment probability at 21 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 and/or they refrain from hiring

[11] We replicate the discontinuity analysis at 21st, 22nd and 23rd birthday with 6 and 12 month estimation windows instead of 15 months (see the Appendix). The results obtained with the 6 month window are never significant. This may be due to the lower number of observations when using the shorter estimation window. Moreover, the discontinuity effect may take time to become sufficiently pronounced. The regressions with 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 significant at 10% while the latter is not significant.

workers between the ages of 21 and 22. We pursue this possibility further and repeat our analysis for every age in one-month increments between 18 and 23 years. Since we estimate dozens of coefficients, it is more instructive to depict the results graphically. Figure 1 presents the slope effect for males, Figures 2 and 3 summarize the findings for females (using quadratic and cubic age polynomial, respectively) and Figure 4 features those for both genders combined. The solid line captures the employment probability while the dotted lines correspond to the 95% confidence interval. An interesting pattern emerges. The employment probability goes up and down, occasionally being significant positive or negative. Most of these upsurges and dips are not very pronounced and tend to be observed only for a very short period. This is to be expected, given that we estimate a relatively large number of coefficients. We observe, nevertheless, a significantly negative employment probability for both males and females when they are 18 (we return to this below). Thereafter, the effect appears consistently positive for both males and females (the latter when age is accounted for with a quadratic polynomial) for several months when they are between 18 and 19 years old: this is likely attributable to the end of full-time secondary education. Then, the employment probability is negative for young males for some five months around their 21st birthday; no such effect is observed for females at this or any other age.[12]

We can only speculate what drives these results. The age-related NMW rates apply equally to men and women yet we only observe negative employment effect for the former. This may reflect the fact that the labor market positions of men and women are substantially different. As we argued above, the negative eted fnm4 the f vf m40er ag3406032bse4(oes du03 TcTw[ity 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. Turning 18 is associated with a significantly negative slope effect for both genders (as is already apparent in the Figures): becoming eligible for the higher NMW rate is associated with lower employment probability. Note that again this negative effect is observed only when we consider the slope effect: 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. 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. This may explain why the employment probability first dips around the 18^{th} birthday and then rises, both for males and females.

Note that our analysis is based on estimating the functional form in expression (4). However, it is also relevant to study if the main conclusions in the paper are upheld when we adopt a specification similar to (4) but imposing the restrictions and . Although we prefer specification (4) because it already encompassed this restricted case and also it allows for comparison with DRW, the constrained version of the model is interesting since it allows us to test the contribution of allowing slope parameters to change at the threshold age on the estimation of the jump effect. Under the restricted model for all workers, we also find strong evidence of a negative jump effect of NMW on the probability of employment at 18 (-0.02 with a p-value of 0.001). Moreover, the impact at 21 is negative but

neighborhood of the cutoff age, whether before or after. The fact that we find a negative effect approximately one year before it should 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 22 to 21. Future research will show how this has affected 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 diminished.

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 slope effect.

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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

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	.00107	.00190	.00175	.00037	.000200
	(.00126)	(.00135)	(.00195)	(.00212)	(.00160)	(.00170)
Dum[(2)]	-.008830	-.00919	-.01013	-.01104	-.00844	-.00819
	(.00425)*	(.00452)*	(.00659)	(.0071)	(.00535)	(.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)

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	-.00004	-.01189	.01636	.00030	-.00500
	(.00158)	(.00181)	(.00936)	(.01102)	(.00663)	(.00518)

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

	Males	Females

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

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

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)

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	-.00136	-.00665	.01457	-.01932	-.00362
	(.01389)	(.00508)	(.01375)	(.01321)	(.01955)	(.00484)
Dum(2)	-.00170	-.00589	.02335	.00031	-.02845	-.01020
	(.02049)	(.01353)	(.02011)	(.00506)	(.01264807)	(.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.