DISCUSSION PAPER SERIES

Downstream Effects of Voting on Turnout and Political Preferences: Long-Run Evidence from the UK

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Initiated by Deutsche Post Foundation

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ABSTRACT

Downstream Effects of Voting on Turnout and Political Preferences: Long-Run Evidence from the UK*

Does voting have downstream consequences for turnout and political preferences? While research initially showed strong support for the notion that the experience of voting fosters civic habits and political engagement, recent work has cast doubt on how universal these patterns are. We contribute to this debate by studying the short- and long-term impact of earlier voting eligibility on subsequent turnout and political preferences using rich panel data from the UK. Exploiting the eligibility cut-off for national elections within a regression discontinuity design, we document a short-run increase in party identification, political interest and democratic norms for those able to vote earlier. However, these short-term effects quickly fade away and do not translate into permanent changes in turnout propensity or political preferences. Our results imply that the transformative effects of voting are short-lived, at most, in a setting with low institutional barriers to vote.

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1 Introduction

The act of voting has long been portrayed as a transformative event in the political lives of citizens (De Kadt, 2017). A large amount of research, particularly looking at the United States, has shown that voting appears to be habit-forming, leading to increased participation in the subsequent election (Green and Shachar, 2000; Denny and Doyle, 2009; Meredith, 2009; Dinas, 2012; Fujiwara et al., 2016). However, recent research from a variety of other national contexts has cast doubt on the universality of this finding (Bhatti et al., 2016; Górecki, 2015; Bechtel et al., 2018; Hernæs, 2019). In addition, voting has recently been found to have little impact on political preferences either (Holbein and Rangel, 2020).

We contribute to this debate in four ways. First, we test for the effect of past voting eligibility on subsequent turnout and political preferences in a new context, the United Kingdom. The only previous work to examine the UK is Denny and Doyle (2009); however, their study relies on random effects estimations for individuals born in 1958 and only examines turnout for two elections. Second, we test for past voting effects not just on turnout but also on political interest and partisanship, building on Holbein and Rangel (2020). Third, we examine long-run effects by testing whether there is a transformative effect of voting up to three elections later. Finally, we examine the effect of past eligibility for a large set of elections in the UK, so we can show that our findings are not due to the characteristics of one election (Bhatti et al., 2016).

To identify the effect of past voting eligibility, we use a regression discontinuity (RD) design on individual-level survey data (BHPS/UKHLS) from the UK. This approach is appealing because of its simplicity and transparency: we compare the long-run outcomes of individuals who turned 18 just before an election (and thus had the opportunity to vote) with those who turned 18 just after an election (who were thus first able to vote one electoral cycle later). Our RD estimates cleanly identify the long-run effects of being able to vote at an earlier election.

Our rich data source provides new and credible causal evidence that past voting eligibility does not affect turnout at later elections. In addition, our findings show little variation based on election salience and outcomes as moderators of the impact of past vote. Overall, our study thus indicates that the positive effects found in work such as Coppock and Green (2016) may not apply to different contexts. Our analysis of political preferences provides insight into why this may be the case, as there is at most a short-term positive effect of past voting on political interest and partisanship, consistent with a 'first-time hype' effect (Bhatti et al., 2016). These findings cast further doubt on the hypothesised transformative effect of voting.

2 Voting as a transformative event

Extensive research has shown that voting in past elections makes people more likely to vote in a subsequent election (Green and Shachar, 2000; Meredith, 2009; Denny and Doyle, 2009; Dinas, 2012; Fujiwara et al., 2016), supporting the notion that voting is habit-forming (Plutzer, 2002; Fowler, 2006).

However, recent research questions this simple account. Outside the US, e.g., in Norway, Sweden, Finland, Switzerland and Brazil, evidence of the habit-forming effects of elections is much harder to find (Górecki, 2015; Bhatti et al., 2016; Bechtel et al., 2018; Hernæs, 2019; Holbein and Rangel, 2020). Perhaps other countries lack the institutional or other contextual characteristics that make voting habit-forming. Most importantly, there may be a moderating impact of institutional barriers (Bhatti et al., 2016): the less effort is required to register to vote, the smaller the effect of past eligibility. Moreover, turnout may naturally decline when people vote for the second time as the novelty of voting wears off, creating a countervailing, 'first-time hype' effect (Bhatti et al., 2016).

A key mechanism leading to habit formation builds on the hypothesis that voting is a transformative act (Holbein and Rangel, 2020). According to this hypothesis, voting strengthens individual attitudes such as political interest, their sense of civic duty or political efficacy (Bhatti et al., 2016). Voting then has an 'educative effect' (Mansbridge, 1999), increasing civic engagement more broadly (Braconnier et al., 2017; Holbein and Rangel, 2020). Going out to vote would therefore not just be important as a way of contributing to political decision-making, it would also make people better citizens in the long run (Mansbridge, 1999). Such downstream effects may be particularly likely when the target group is at a malleable age (Holbein and Rangel, 2020).[1]

Nevertheless, the evidence for downstream effects beyond voting is also mixed: some studies show clear effects of voting, for instance on political interest and knowledge (Braconnier et al., 2017; Shineman, 2018) and on non-electoral political participation (Khoban, 2019), while others present null results on outcomes such as political knowledge, information consumption, political memberships and social awareness (Loewen et al., 2008; de Leon and Rizzi, 2014; Rosenqvist, 2017; Holbein and Rangel, 2020). For example, de Leon and Rizzi (2014) and Rosenqvist (2017) find that earlier eligibility does not increase political knowledge of individuals in Brazil

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[1]Two other potential mechanisms underlying the habit to vote are not examined further here (Bhatti et al., 2016). First, voting may lead individuals to reflexively associate elections with going out to vote. Second, the habit-forming effect may be spurious to the extent that it emerges mainly because parties target previous voters more strongly when campaigning (Gerber et al., 2003).

and Sweden, respectively.

Further research has argued that the impact of past eligibility may vary depending on the electoral and institutional context. If the electoral context creates positive affect, such as at a first democratic vote, vote eligibility may have a particularly long-lasting effect (De Kadt, 2017; Kaplan et al., 2019). A more general role may be played by electoral competitiveness: if one's first election is close and exciting, this might strengthen habituation effects more than if the first election is uneventful (Franklin, 2004; Franklin and Hobolt, 2011). Our data allows us to check whether elections differ in their long-term effects; for instance, the 1997 election, which led to Tony Blair's premiership, may have been a different first-time voting experience than the more humdrum 2001 election.

3   Case description, data and empirical approach

Case   Elections to the House of Commons of the United Kingdom ('General Elections') must take place within five years. Turnout at General Elections was between 70 and 80% from the 1950s to 1997 and has hovered around 65% since the early 2000s. To be eligible to vote in the General Elections, one must be British citizens and have turned 18 on election day. Additionally, eligible voters must be on the electoral register; this was traditionally done by post, but can nowadays also quickly be done online. Once registered, individuals are eligible to vote in all upcoming elections. Thus, our focus is on a country with comparatively low institutional barriers to vote. In contrast to studies on the US, our focus on the UK makes null findings more likely.

Data   We use BHPS/UKHLS data for the empirical analysis (University of Essex, 2019). The surveys provide information on political variables such as voting behaviour, political interest, party preferences and whether voting is seen as a social norm. We use the survey waves covering 1991-2018, thus we can examine the General Elections of 1992, 1997, 2001, 2005, 2010, 2015 and 2017.[2] A special licence version of the data contains information on the birth month and year, allowing us to determine voting eligibility at each election.

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[2]In the period we study, UK voters could also vote in subnational, European and local elections. To ensure that our results are not confounded by European or subnational elections, in a robustness check we exclude individuals within a six-months window on each side of the eligibility cut-off date for European elections and, second, restrict the sample to individuals residing in England. We find no evidence that subnational and EP elections affect our results (see Appendix Figure A.5).

Empirical approach We use an RD design to estimate the long-run effects of past eligibility. Our empirical approach exploits that eligibility is a deterministic function of an individual's date of birth and the election date. The age cut-off generates exogenous variation in the age at which individuals receive the first opportunity to vote. Intuitively, we compare the political outcomes of individuals who turned 18 just before an election with those of individuals turning 18 just after the same election (Meredith, 2009; Coppock and Green, 2016; Dahlgaard, 2018). The idea underlying this approach is that individuals around the eligibility cut-off should be identical apart from the difference in the first eligibility to vote. Importantly, our estimates are not confounded by life cycle effects, which are often drawn on to explain low turnout of young voters. The main identifying assumption is that individuals cannot manipulate the 'running variable', i.e. an individual's age on election day (Imbens and Lemieux, 2008). Given the strict enforcement rules and that eligibility is the product of two random events—date of birth and the date of election—this is unlikely. We provide supplementary evidence that supports this uncontroversial assumption: we show that pre-determined covariates are smoothly distributed around the cut-off (Figure A.1) and that there is no bunching of observations on either side of the cut-off (McCrary density test in Figure A.2).

We estimate the following regression model for voting:

$$y_{i,t_0+k} = \beta_0 + \beta_1 I(Z_{i,t_0} > 0) + \beta_2 Z_{i,t_0} + \beta_3 Z_{i,t_0} \cdot I(Z_{i,t_0} > 0) + \gamma_{t_0+k} + \epsilon_{i,t_0+k} \tag{1}$$

where y_{i,t_0+k} denotes the voting outcome of individual i at election t_{0+k}, where $k \in \{0, 1, 2, 3\}$. Thus, we examine the effect of earlier eligibility on voting for the initial and three subsequent elections. Z_{i,t_0} denotes the running variable, i.e., the relative age at the initial election t_0. To maximise precision of the estimates, we use a symmetric 48-months observation window on either side of the cut-off date. $I(Z_{i,t_0} > 0)$ takes the value of 1 for individuals eligible to vote at election t_0, the initial eligibility-determining election, zero otherwise. We exclude individuals who turn 18 in the month of an election as we cannot determine eligibility for those. In our baseline specification we control for the running variable linearly, which captures age-related linear time trends in the outcomes, and allow for different slopes on both sides of the cut-off. Our coefficient of interest, β_1, captures any discontinuous jumps in voting outcomes between just-eligibles and just-ineligibles. We include fixed effects for the election for which we analyse voting, γ_{t_0+k}, to capture differences between elections with respect to average turnout, salience, and media attention. For estimation, we use the *rdrobust* package for Stata by Calonico et al. (2017) with the bias-corrected standard errors. We provide several specification checks to assess

the robustness of our results.

As political preferences are collected annually, we slightly modify the regression specification and include survey year fixed effects instead of election fixed effects to absorb time-specific differences in these outcomes.

4 Results

Voting   We begin our empirical analysis by examining how just becoming eligible to vote affects first time voting, and how past eligibility affects average turnout for three subsequent elections. Figure 1 presents the results, with each panel showing the reduced form relationship between the running variable, relative age at election t_0, and turnout at these different elections. We plot the average voting shares for each monthly bin of relative age and fit linear trends separately for each side of the cut-off. We also report the RD estimates in the figure for our baseline specification with linear age trends.

Panel A clearly shows that having passed the age threshold of 18 affects turnout for the newly eligibles: We observe a discontinuous jump in the probability to vote of around 53 percentage points (pp) between individuals who turned 18 just before the election date (to the right of the vertical axis) and individuals who turned 18 just after the election. To the left of the cut-off, 1.1% of just-ineligibles report having voted in this election. We attribute this small share to reporting errors as these individuals were legally not eligible to vote. When we manually set the voting outcomes to 0 for the ineligibles, we estimate a jump in turnout of 56.4 pp. Figure 2 shows that the RD estimates are stable and robust to including control variables and specifying the running variable linearly or quadratically.[3]

Does past voting eligibility then have an effect on subsequent turnout? Panels B to D in Figure 1 display the results for three subsequent elections. If voting is habit-forming, we would expect that individuals who had the opportunity to vote earlier would exhibit higher turnout at subsequent elections. However, we find that past eligibility does not generate any long-run effects on voting. The outcomes trend smoothly on each side of the cut-off and, consistent with prior studies, the figures also indicate an initial increasing propensity to vote as individuals get older (e.g., see Phelps, 2004; Bhatti et al., 2012). The point estimates at the cut-off are all close to zero and confidence intervals mostly include zero, overall providing little support for large effects on habit formation. Our estimates are fairly precise as we can rule out with 95%

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[3]Appendix Figure A.3 plots the first *observed* voting age against birth months. Reassuringly, we observe a pronounced discontinuity in the first observed voting age at each election cut-off date.

Figure 1: Relationship between relative age at election t_0 and turnout at subsequent elections

Panel A: Election t_0

Panel B: Election t_{0+1}

Panel C: Election t_{0+2}

Panel D: Election t_{0+3}

RD estimate: .533 (0.012)

RD estimate: -.032 (0.016)

RD estimate: -.02 (0.016)

RD estimate: -.015 (0.015)
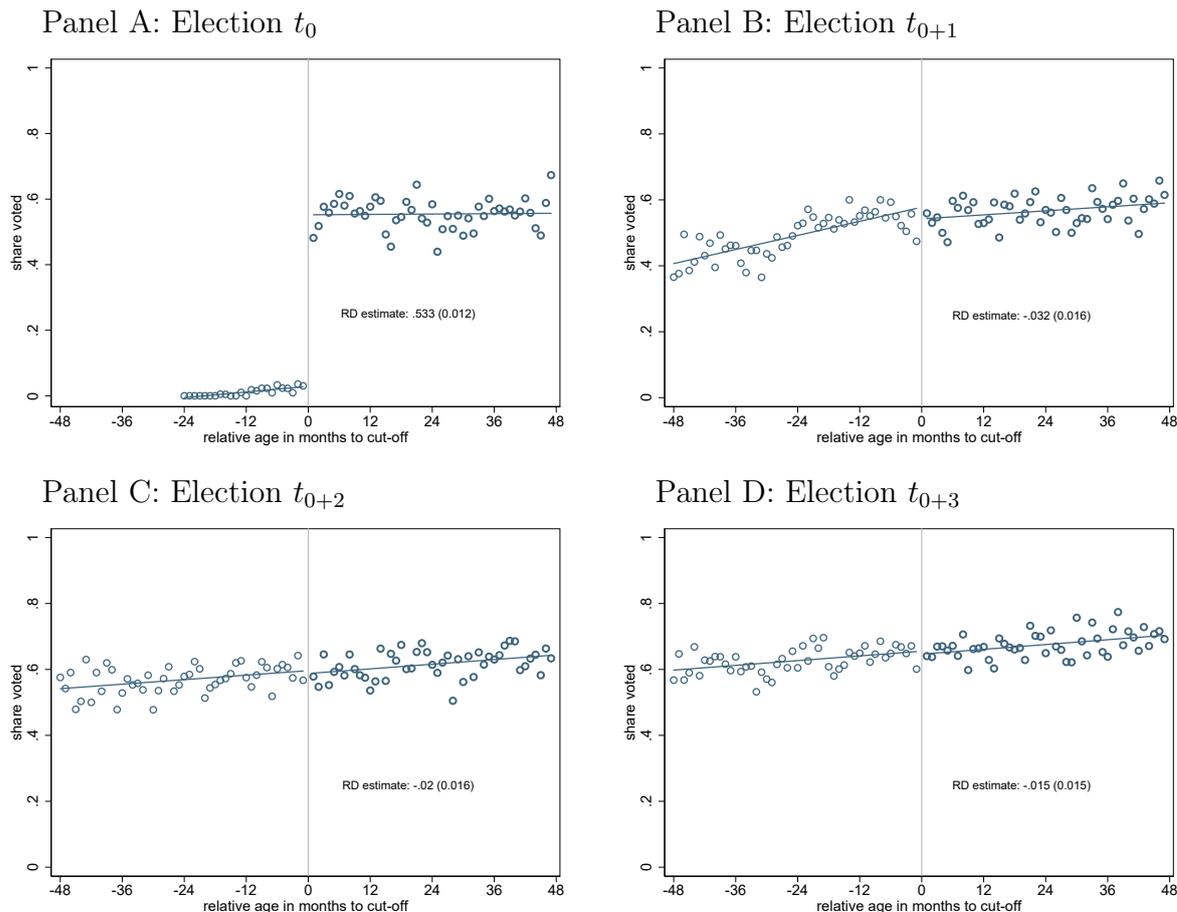
Note: Dots represent average voting shares for individuals born in a certain month relative to the cut-off date; at election t_0, individuals born to the right of the cut-off have the right to vote. We pool information on all available elections RD estimates include election fixed effects. Standard errors shown in parentheses. Respective $N = 13987, 18017, 18525, 20034$.
Source: BHPS / UKHLS, own calculations.

confidence that the effect in any of three elections later is larger than 1.5 pp. We do observe a small, statistically significant negative effect of past eligibility on voting at the subsequent election (t_{0+1}) in our baseline specification. Although this is consistent with a 'first-time-hype' effect for the previously ineligibles, our robustness checks presented in Figure 2 show that this coefficient is not robust and becomes insignificant when using quadratic time trends.

The final specification in Figure 2 pools the data for all subsequent elections to gain precision. The estimates show a precisely estimated zero indicating that past eligibility does not affect aggregate long-run voting behaviour. Figure 2 confirms the stability of the estimated coefficients. Appendix Figure A.4 shows that these results persist when considering up to six subsequent elections. We therefore conclude that past voting eligibility does not increase future turnout; if anything, we estimate a slightly negative, though statistically insignificant, effect.

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Figure 2: RD specifications - effect of earlier eligibility on subsequent elections
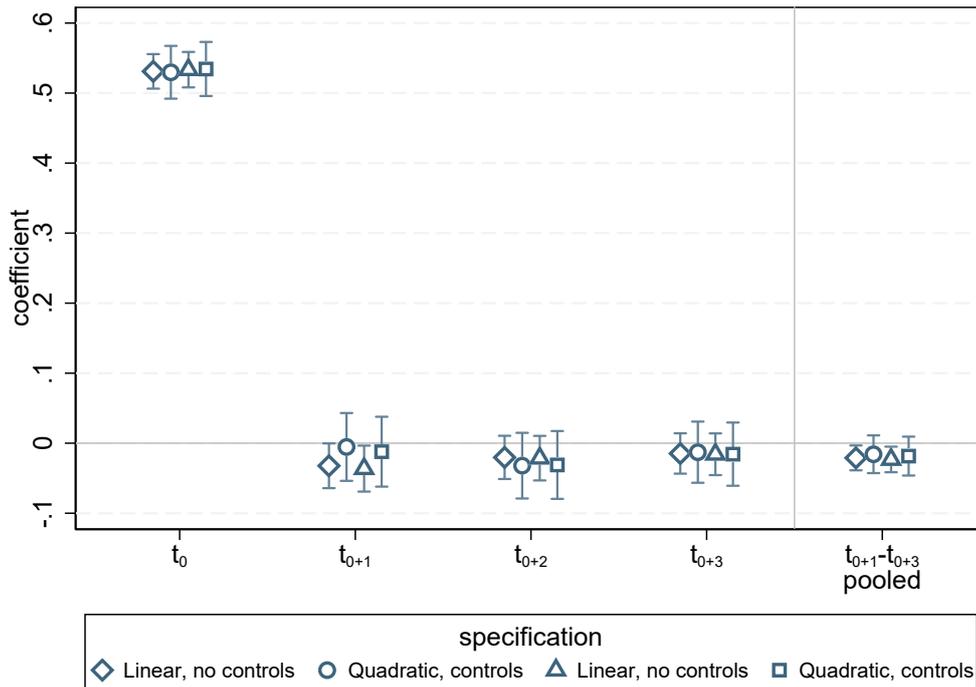
Note: Figure shows coefficients and 95% confidence intervals for different specifications. The first specification (linear, no controls) corresponds to the RD estimate shown in the respective Panel of Figure 1. All estimates include election fixed effects. Respective $N = 13987, 18017, 18525, 20034, 56411$.

Source: BHPS and UKHLS, own calculations.

As elections differ in their salience, which may affect both election turnout and long-run effects, we also examine the downstream voting effects of each individual election in Table A.1. Three key findings emerge. First, being newly eligible to vote always affects turnout at the first election, with the effect ranging from 37 pp (2005) to 69 pp (1992 and 1997). Second, we find no consistent evidence for a first-time hype effect, as evidenced by the opposite signs for t_{0+1} for the elections 1992 and 1997 and the zero coefficients for the remaining elections. Third, we find no long-run effect of past eligibility, in particular not for elections with larger first-time turnout.

Political preferences Next, we examine whether the younger age at first vote affects political preferences and interest over the next three election cycles. We focus on the following outcomes: party identification, defined as supporting or feeling closer to one political party; political interest, coded as whether individuals have high or very high interest in politics; whether voting is seen as a social norm—a proxy for civic duty (Blais, 2000; Gerber and Rogers, 2009); and party preference, measured as whether an individual states a party she will vote for. We

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code all dependent variables as indicator variables for ease of interpretation. We only use observations from the election year if the interview date took place after the election.

Table 1: Relationship between relative age at election t_0 and political preferences

Election cycle	t_0 (1)	t_{0+1} (2)	t_{0+2} (3)	t_{0+3} (4)
Supporting a political party or close to one party				
RD estimate	0.037**	-0.010	-0.001	-0.001
	(0.015)	(0.013)	(0.012)	(0.012)
N	20,418	25,538	26,530	29,744
Interested in politics				
RD estimate	0.025	-0.004	0.013	-0.001
	(0.016)	(0.014)	(0.014)	(0.014)
N	15,564	19,503	19,090	20,884
Agree that voting is a social norm				
RD estimate	0.095	-0.032	-0.044	-0.025
	(0.058)	(0.050)	(0.053)	(0.049)
N	1,433	1,659	1,407	1,719
States a party to vote for				
RD estimate	0.052**	0.034**	-0.002	-0.007
	(0.021)	(0.017)	(0.017)	(0.017)
N	9,594	12,205	11,773	12,384

Note: Table shows the relationship between relative age at election t_0 and a range of political preferences. Election cycle t_0 denotes the period between the first election where individuals are around the eligibility cut-off and the next election, t_{t+1} the election cycle thereafter etc. All estimates include survey year fixed effects. Dependent variables are covered over the period 1991-2019 with the exception of voting as a social norm which has only been included since 2010. Bias corrected standard errors in parentheses, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: BHPS and UKHLS, own calculations.

Table 1 provides the corresponding estimation coefficients for each outcome over time.[4] The table provides evidence that past vote eligibility affect political preferences in the short-run, that is at election t_0, leading to higher levels of party identification, political interest (though statistically insignificant), perceptions of voting as a social norm, and party preference.[5] These differences would be consistent with the notion that voting has downstream effects. However, these short-term effects fade away quickly, becoming insignificant and hovering around zero at t_{0+3}. We therefore conclude that past eligibility does increase individuals' initial connection to and interest in politics, but that these effects are altogether short-lived.

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[4] For the corresponding reduced form graphs, see Figures A.6 to A.9
[5] We do not observe strong evidence for an effect in favour of one particular party, see Appendix Table A.2.

5 Conclusion

Our research shows that voting is not universally habit-forming. We find no evidence of a difference in turnout based on past vote eligibility. This null effect largely holds across six elections from 1992 to 2017. As such, our results confirm recent work (Bhatti et al., 2016; Bechtel et al., 2018; Hernæs, 2019) on the non-existence of a habit of voting outside the US. One explanation for the different results could potentially be institutional differences, as the process of voter registration and voting itself is less restrictive in the UK compared to the US where some states follow restrictive policies with respect to state voter ID laws, automatic voter registration, and early and mail-in voting. Finally, we note that our findings contradict earlier results for the UK presented in Denny and Doyle (2009), who find evidence of a habit of turnout. However, our study uses a stronger causal research design, covers more elections and includes several birth cohorts, thus providing more general results.

Our findings also provide evidence that voting only has a short-term impact on key political orientations. Specifically, there is no long-term effect of voting on political interest, party identification or seeing voting as a social norm. Like Holbein and Rangel (2020), these results question whether voting in fact makes people better citizens (Mansbridge, 1999). Voting may generate a first-time hype effect (Bhatti et al., 2016), but this wears off by the time of the next election.

Our study also has implications for the debate on the effect of lowering the voting age to 16. Supporters sometimes argue that doing so could increase turnout in the long-run, based on two assumptions. First, voters under 18 would vote at higher rates than those over 18. This expectation is reasonable, for example as results from Austria show high levels of voting among voters under 18 (Aichholzer and Kritzinger, 2020). However, the second key assumption is more problematic, as increased turnout among those under 18 would have to establish habits of voting among this cohort. Our results show that instilling habits may not be so straightforward after all, so lowering the voting age may need to be justified using other arguments.

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Online Appendix A: Supplementary material

Figure A.1: Covariate balancing checks
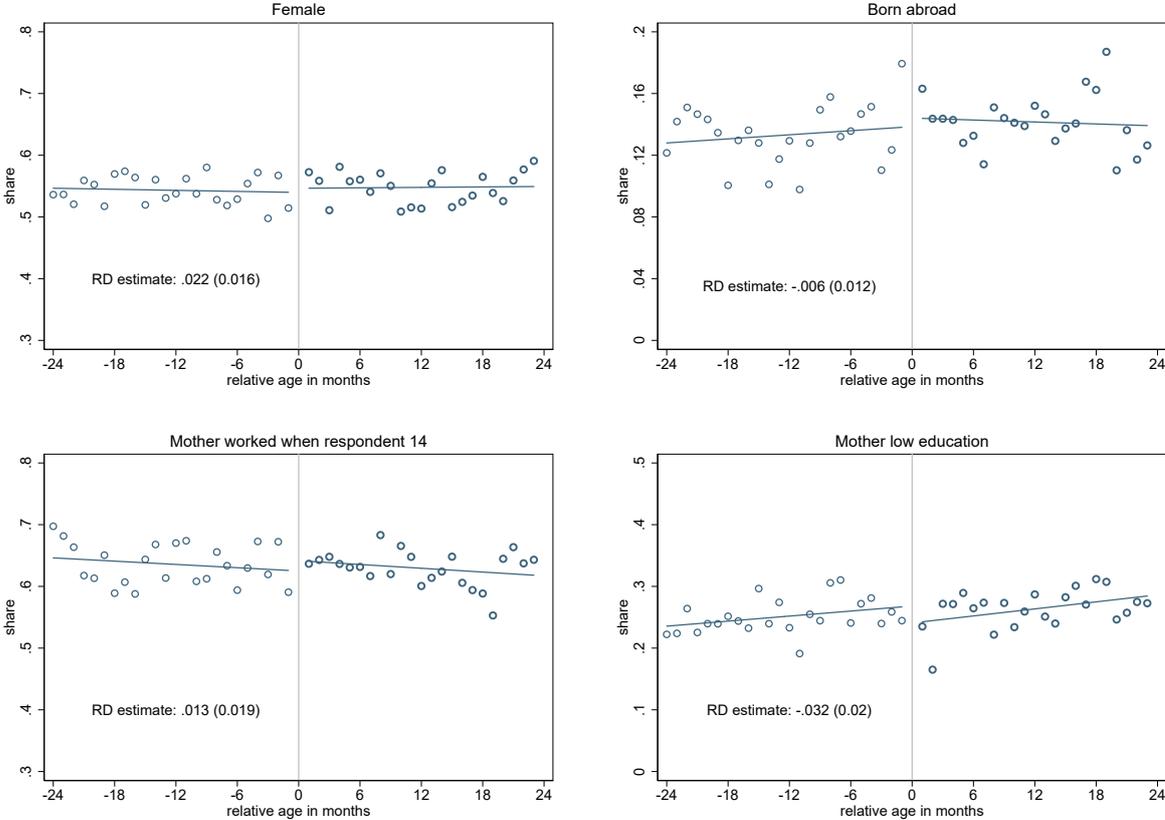
Figure A.2: Manipulation test of running variable
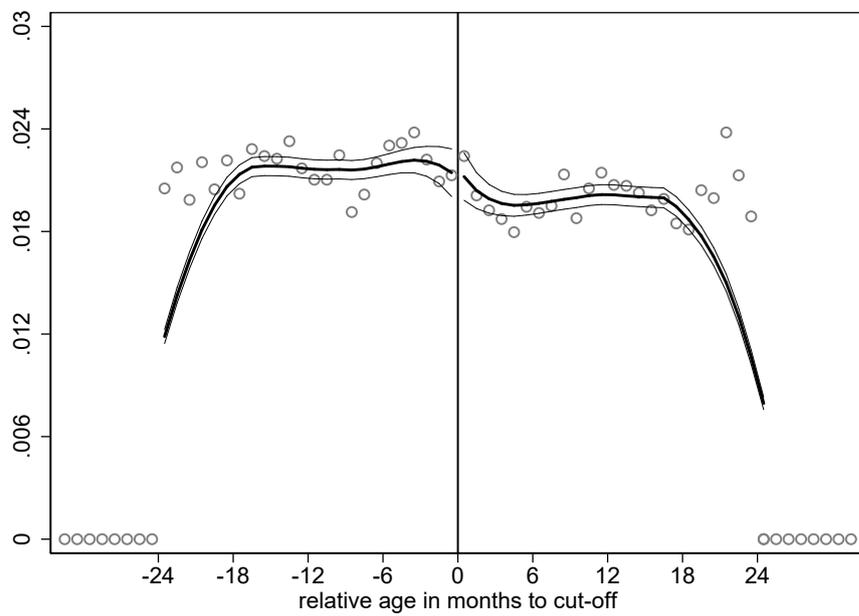
Note: Figure shows a McCrary (2008) density test of observations around the cut-off of the running variable (relative age in months). T-statistic of density test is 0.423. *Source:* BHPS and UKHLS, own calculations.

Figure A.3: Month of birth and age at first observed vote
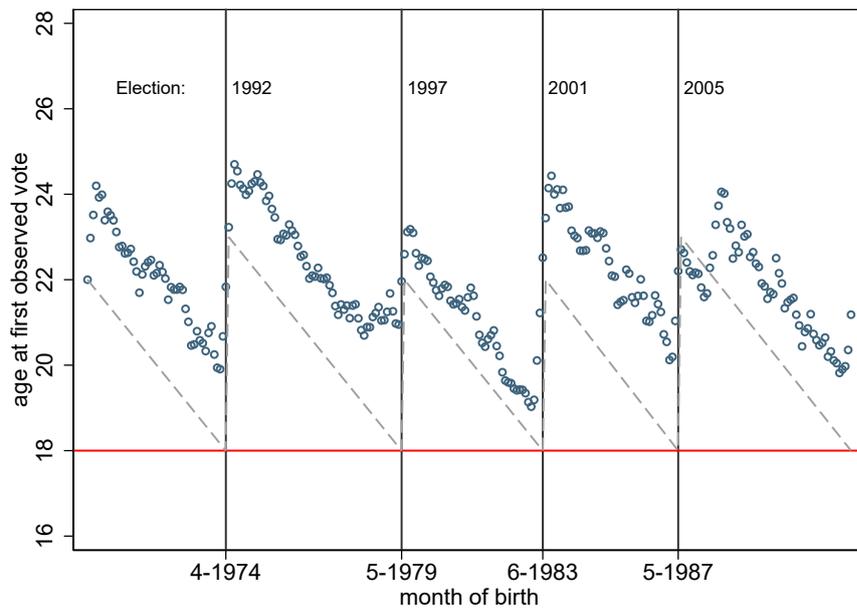
Note: Figure shows age at first observed vote by month of birth. For this analysis, we have to make two additional sampling restrictions to calculate the first observed voting age more reliably. We reduce the observation window to two years around each cut-off to avoid using the same individual twice (once on each side of the cut-off). Second, we restrict the sample for each election to individuals aged up to 27 years, which corresponds to two electoral cycles (4+5 years). This restriction reduces the problem of not observing the first vote in the data if individuals join the sample after their first vote. $N = 4912$.
Source: BHPS and UKHLS, own calculations.

Figure A.4: Relationship between relative age at election t_0 and turnout at subsequent elections

Panel A: Election t_{0+4}

Panel B: Election t_{0+5}

Panel C: Election t_{0+6}
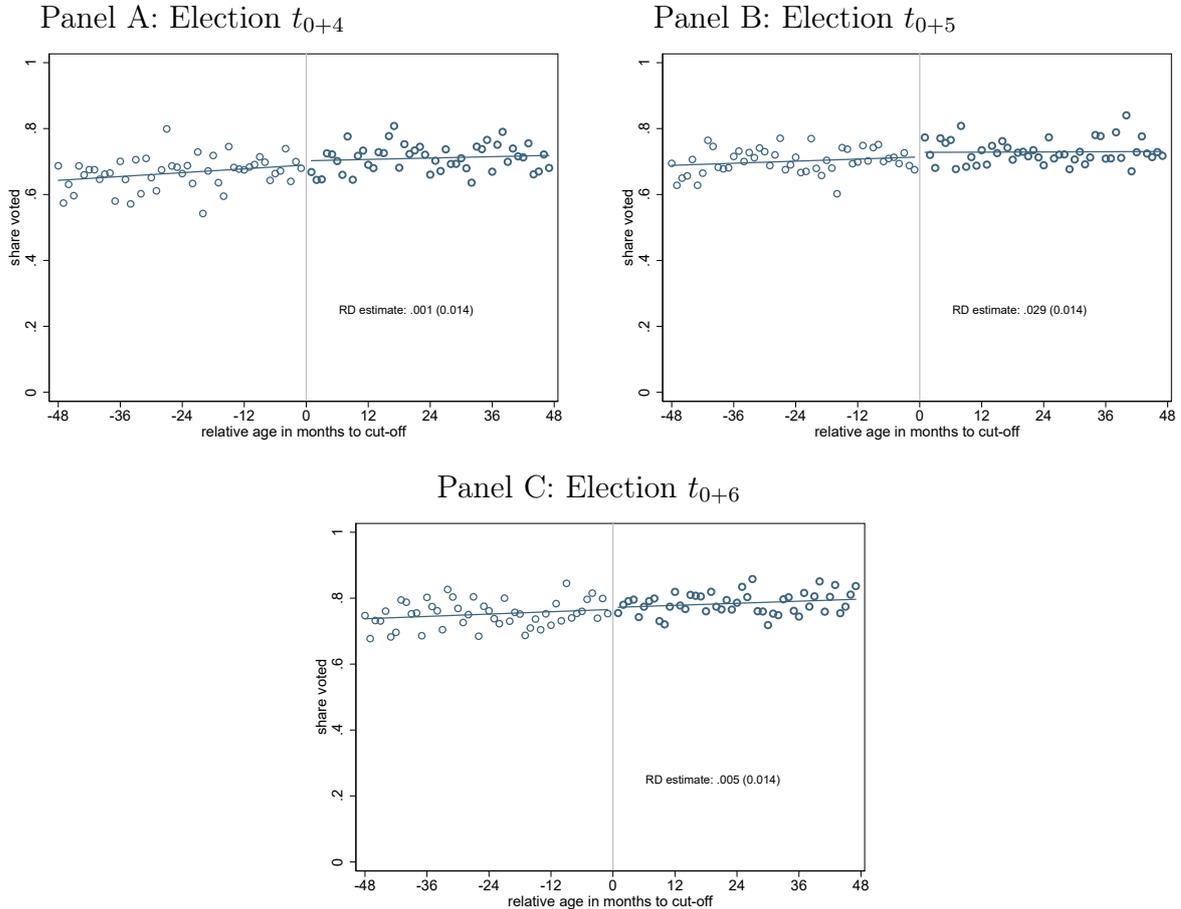
Note: This figure shows the relationship between the running variable, relative age at election t_0, and average turnout at different subsequent elections. The dots represents the average voting shares for individuals born in a certain month relative to the cut-off date; at election t_0, individuals born to the right of the cut-off have the right to vote. We exclude observations who turn 18 in the month of each election as we cannot assign eligibility for those individuals cleanly. We pool information on all available elections (1992, 1997, 2001, 2005, 2010, 2015 and 2017). As two elections took place in 1974, the cut-off date cannot be cleanly assigned in our framework, thus we exclude elections for which t_0 is 1974; for voting in t_{0+4} this is the 1992 election, for t_{0+5} 1997 and for t_{0+6} the 2001 election. RD estimates include election fixed effects. Standard errors shown in parentheses. Respective $N = 19862, 18405, 16126$.
Source: BHPS and UKHLS, own calculations.

Figure A.5: Robustness of voting coefficients to European and subnational elections
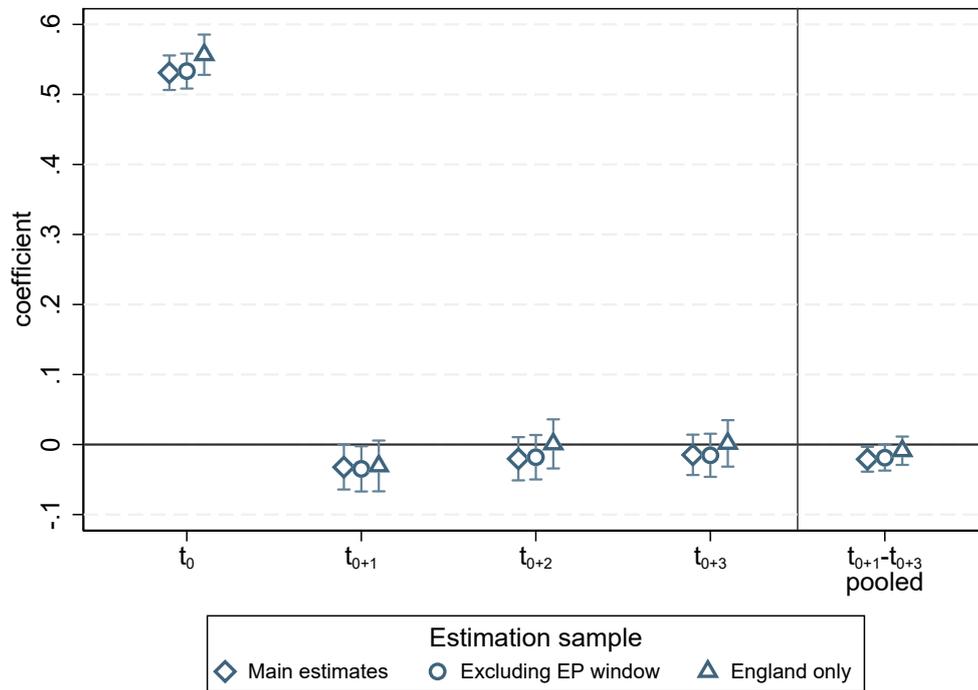
Note: Figure shows robustness checks for voting at the 'initial' and subsequent elections. Main estimates shown in Figure 2. The first robustness check excludes all individuals who were within six months of the age cut-off relative to an election for the European Parliament (EP). The second set of estimates is restricted to individuals residing in England to avoid subnational elections in Wales, Scotland and Northern Ireland biasing the results. *Source:* BHPS and UKHLS, own calculations.

Figure A.6: Supporting a political party or close to one party
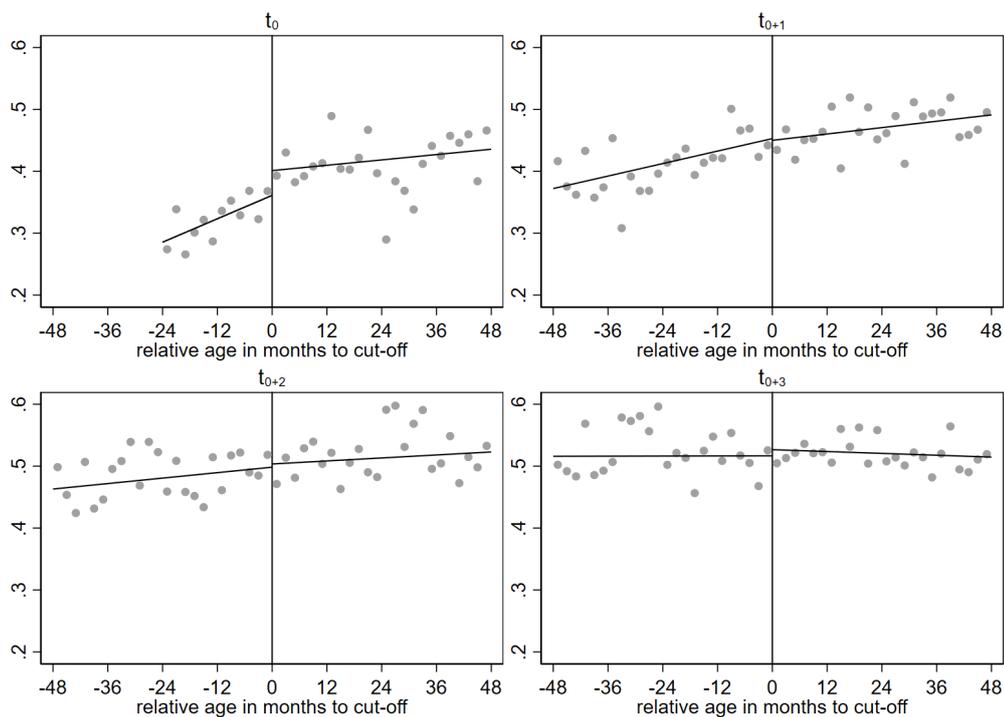
Note: Figure shows the relationship between relative age at election t_0 and and indicator for support for a political party or being closer to one party. Election cycle t_0 denotes the period between the first election where individuals are around the eligibility cut-off and the next election, t_{t+1} the election cycle thereafter etc. Figure shows raw data, i.e. survey year fixed effects as in Table 1 are not included.

Source: BHPS and UKHLS, own calculations.

Figure A.7: Interested in politics
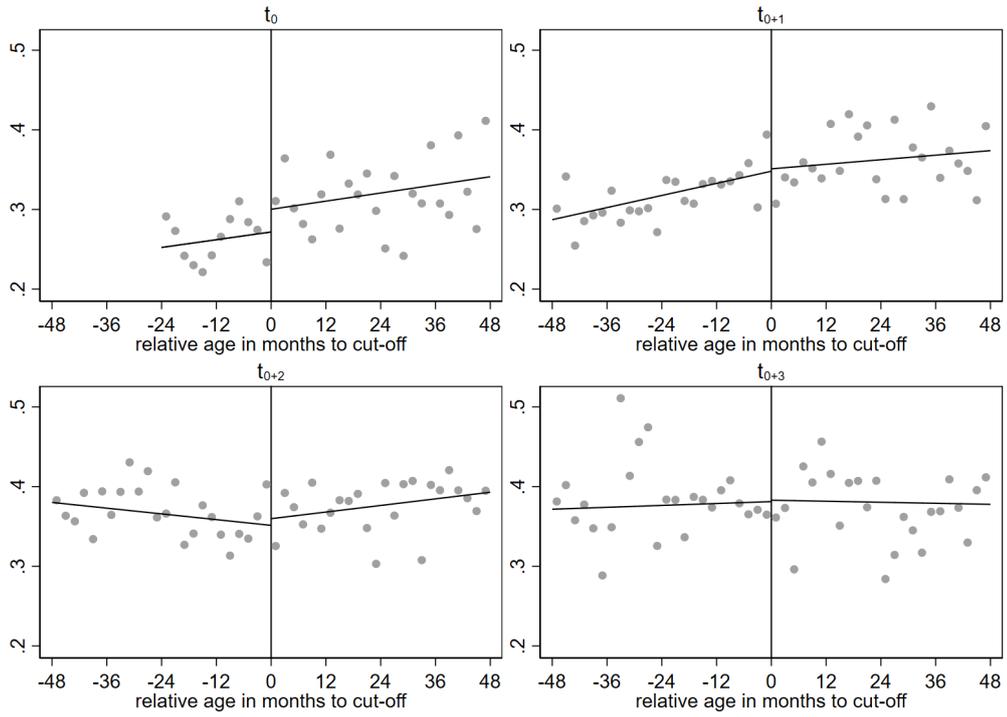
See notes for Figure A.6.

Figure A.8: Agree that voting is a social norm
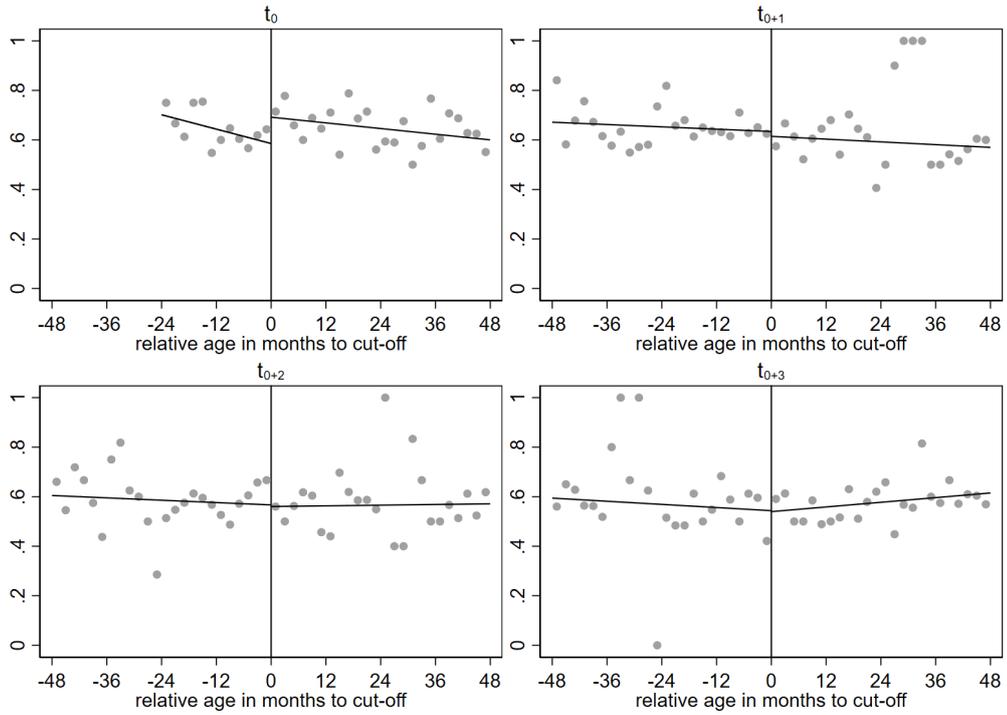
See notes for Figure A.6.

Figure A.9: States a party to vote for
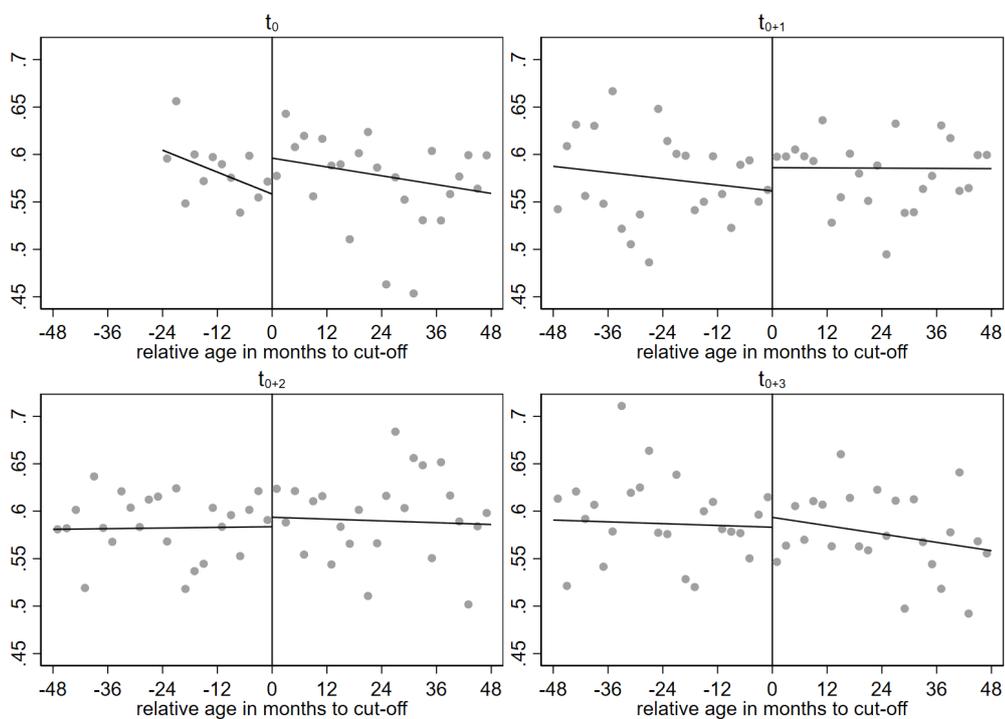
See notes for Figure A.6.

Table A.1: Relationship between relative age at election t_0 and turnout by election

Election	t_0 (1)	t_{0+1} (2)	t_{0+2} (3)	t_{0+3} (4)
RD estimate election 1992	0.694***	-0.122**	0.028	-0.010
	(0.041)	(0.056)	(0.049)	(0.039)
N	964	1,376	1,514	1,574
RD estimate election 1997	0.692***	0.098**	0.055	0.059*
	(0.031)	(0.041)	(0.037)	(0.034)
N	2,082	2,895	3,108	3,299
RD estimate election 2001	0.490***	-0.031	-0.002	-0.037
	(0.026)	(0.035)	(0.034)	(0.032)
N	3,162	4,063	4,409	4,993
RD estimate election 2005	0.371***	-0.059	-0.141***	-0.045
	(0.027)	(0.037)	(0.037)	(0.035)
N	2,871	3,470	3,394	3,726
RD estimate election 2010	0.505***	0.014	0.000	-0.038
	(0.035)	(0.049)	(0.047)	(0.044)
N	1,692	2,068	2,222	2,330
RD estimate election 2015	0.562***	-0.005	-0.088	0.047
	(0.042)	(0.057)	(0.057)	(0.054)
N	1,175	1,439	1,423	1,566
RD estimate election 2017	0.593***	0.029	0.017	-0.044
	(0.033)	(0.047)	(0.041)	(0.040)
N	2,041	1,944	2,455	2,546

Note: Table show RD estimates show in Figure 1 by election.

Source: BHPS and UKHLS, own calculations.

Table A.2: Relationship between relative age at election t_0 and party preferences

Election cycle	t_0 (1)	t_{0+1} (2)	t_{0+2} (3)	t_{0+3} (4)
Would vote for Labour				
RD estimate	0.025	-0.004	-0.007	-0.005
	(0.021)	(0.016)	(0.016)	(0.015)
N	8,039	10,110	9,624	10,436
Would vote for Conservatives				
RD estimate	0.001	0.001	-0.007	-0.009
	(0.013)	(0.012)	(0.012)	(0.012)
N	8,039	10,110	9,624	10,436
Would vote for Liberal Democrats				
RD estimate	0.022	0.026**	0.000	-0.008
	(0.013)	(0.011)	(0.012)	(0.012)
N	8,039	10,110	9,624	10,436
Would vote for other party				
RD estimate	0.009	0.017	0.012	0.012
	(0.017)	(0.014)	(0.015)	(0.014)
N	8,039	10,110	9,624	10,436

Note: Table shows the relationship between relative age at election t_0 and party preference. Election cycle t_0 denotes the period between the first election where individuals are around the eligibility cut-off and the next election, t_{t+1} the election cycle thereafter etc. All estimates include survey year fixed effects. Dependent variables are covered over the period 1991-2019. Bias corrected standard errors in parentheses, *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.
Source: BHPS and UKHLS, own calculations.