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THE ROLE OF MAJORITY STATUS IN CLOSE ELECTIONS STUDIES

by Matteo Alpino* and Marta Crispino**

Abstract

Many studies exploit close elections in a regression discontinuity framework to identify partisan effects, i.e. the effect of having a given party in office on the outcome. We argue that, when conducted on single-member districts, such analysis may identify a compound effect: the partisan effect, plus the majority status effect, i.e. the effect of being represented by a member of the legislative majority. We provide a simple strategy to disentangle the two effects, and test it with simulations. Finally, we show the empirical relevance of this issue using real data.

JEL Classification: C21, D72.
Keywords: partisan effect, single-member districts, regression discontinuity design.
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Contents

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1 Introduction

Since Lee (2008), Lee et al. (2004) and Pettersson-Lidbom (2008), many papers use a regression discontinuity design (RDD) that exploits close elections (CE) to estimate the effect of a given party being in office on some outcome (e.g. public spending).

We argue that when the data is made of first-past-the-post districts to elect members of a parliament, the treatment effect cannot be interpreted as a pure *partisan effect* (PE), because it is potentially compounded with the effect of being represented by a member of the majority, i.e. a *majority status effect*. Consider one term when the democrats have conquered the majority of seats. In this case, all districts are either won by a democrat in the majority or by a republican in the opposition. Instead, if republicans have won the majority of seats, all districts are either won by a democrat in the opposition or by a republican in the majority. In other words, representatives differ not only in their party affiliation, but also in their majority status. Since most applications combine data pooled from several election-years, the estimated effect is a weighted average of these two different joint effects, making its interpretation complicated.

Note that the bundling of these two effects naturally occurs in this electoral system due to institutional features that make party affiliation mechanically correlated with majority status. This issue is therefore distinct from the fact that party identity is sometimes correlated with politician's characteristics such as gender or ethnicity due to complicated patterns of representation, a problem analyzed by Marshall (2022).

Majority status is a characterising feature of all members of parliament, and has the potential to have an effect on the outcome in many applications that aim at estimating the PE: pork barrel spending, party incumbency advantage, roll call voting, campaign financing, etc. In fact, majority members are likely to have greater agenda setting power and to serve in key positions in legislative committees, or in the cabinet; together they can pass legislation without relying on the support of members of different parties; in some countries, the majority in the parliament elects the executive. Finally, there is evidence that majority status matters for the ability to secure federal transfers and campaign contributions (Albouy, 2013; Cox and Magar, 1999).

2 Related literature

Our paper fits in the large literature in economics and political science that applies RD design to close-elections, first initiated by Lee (2008), Pettersson-Lidbom (2008) and Lee et al. (2004) in order to estimate different types of *partisan effects*.

Pettersson-Lidbom (2008) studies whether left-wing municipal governments implement different fiscal policies than right-wing ones in Sweden. The importance of the left-right dimension in determining economic policies has been investigated by many others in different settings (see e.g. Ferreira and Gyourko, 2009; Solé-Ollé and Viladecans-Marsal, 2013). In a similar vein, Meyersson (2014) investigates the effect of Islamic party rule on female education attainments in Turkey, while Brollo and Nannicini (2012) focus on the effect of party alignment between different layers of government on transfers to local jurisdictions in Brazil.

A relevant stream of literature, originated by Lee (2008), aimed at estimating the *incumbent party advantage*, that is at answering the question: "From the party's perspective, what is the electoral gain to being the incumbent party in a district, relative to not being the incumbent party?" (Lee, 2008, page 692).[1] In his original specification, the running variable is the reference party vote share in t, and the outcome is the reference party vote share in $t+1$, or an indicator for the victory of the reference party in $t+1$. Many papers have applied this design to other

[1]This estimate is different from the incumbency advantage previously studied in political science, which focused on the effect of running with an incumbent *candidate* (Gelman and King, 1990).

settings, in particular in single-member districts elections (for instance, Lee et al., 2004; Lee, 2008; Cattaneo et al., 2015; Uppal, 2010). Moreover, Fouirnaies and Hall (2014) provide evidence that party incumbency in the U.S. has a positive effect on campaign contributions to the party from lobbies. Kendall and Rekkas (2012) estimate positive incumbency advantage in the Canadian House, and Uppal (2009) negative ones in India state legislatures. Eggers and Spirling (2017) provide evidence from the UK House, where more than two parties field candidates, and show that the party incumbency effect after a Conservative-Liberals race is much larger than the one after a Conservative-Labor race.

3 The compounded effect

Consider an electoral system, where representatives are elected in n single-member first-past-the-post districts. Each of two parties fields one candidate in every district. Define D_{it} as a dummy equal to one if the democratic party (D) wins the election in district i, in election-year t, and M_{it} as a dummy equal to one if the district i in year t belongs to the majority, i.e. to the party whose candidates won in the majority of districts. Thus, D_{it} captures the party affiliation and M_{it} the majority status. Note that, by definition, D_{it} and M_{it} are mechanically related: when party D holds the majority then $D_{it} = M_{it}$; when D is in the opposition, then $D_{it} = 1 - M_{it}$.

We are interested in estimating the PE, i.e. the causal effect of party D being in office on some outcome Y_{it}. Assume that in true data generating process Y_{it} is a function of both D_{it} and M_{it} (e.g. the level of federal funding of a district may depend on the party affiliation of its representative, and on its majority status)[2] and that electoral outcomes in all districts are randomized.[3]

Consider regressing Y_{it} on D_{it} using cross-sectional data from one election-year t

[2]See Albouy (2013) for evidence in this respect.

[3]Indeed the issue under discussion is not limited to RDD CE, but to all research designs.

when the democrats have the majority. Using this dataset the coefficient on D_{it} corresponds to the compound effect of being represented by a democrat, and by a majority member, because $D_{it} = M_{it}, \forall i$. If instead at t republicans have won the majority, the same coefficient would capture the compound effect of being represented by a democrat, and by an opposition member, because $D_{it} = 1 - M_{it}, \forall i$. Finally, when data include several election-years, the estimated coefficient is a weighted average of these two joint effects. In particular, it identifies the pure PE only if majority status has no effect on the outcome (ruled out by assumption), or if the covariance between D_{it} and M_{it} is zero,[4] which is not true in general. In fact, such covariance crucially depends on the relative number of democratic-controlled (when $D_{it} = M_{it}$, positive covariance) versus republican-controlled (when $D_{it} = 1 - M_{it}$, negative covariance) years. Specifically, it decreases (in absolute value) as the dataset is more balanced in terms of democratic-controlled and republican-controlled years; it becomes negligible in case of perfect balance, because for each observation such that $D_{it} = M_{it}$, there is one such that $D_{it} = 1 - M_{it}$. Starting from perfect balance, the covariance increases (decreases) as the fraction of democratic-controlled years increases (decreases).[5] Note that typically studies that estimate a (local) regression of Y_{it} on D_{it} use datasets with an unbalanced number of republican-controlled and democratic-controlled years, and thus they do not necessarily identify the pure PE.

3.1 Identification of the PE

To identify the PE, formally defined in the supplementary material B, the data must include more than one election-year and exhibit variation in the party who controls the assembly.[6] Assume that D_{it} is randomized; our main strategy is to simply control for M_{it} in the regression of Y_{it} on D_{it}. Note that M_{it} depends

[4]This follows from the omitted variable bias formula.

[5]See supplementary material A for a proof.

[6]Note that it is not possible to identify heterogeneous effects, such as the PE on majority members. In fact, we cannot credibly compare democratic districts in years when democrats have the majority to republican districts when republicans have the majority due to year-level confounders. See the supplementary material B.

only on D_{it} and on which party has the majority in the assembly. It is therefore sufficient to assume that the overall majority is determined at the national level (and not at the district level) and to control for time fixed effects to safely include M_{it} in the regression without introducing a selection bias. The assumption is more likely to hold (i) when the number of districts n is large, and thus small is the probability that the outcome in one district determines the overall majority, and (ii) the smaller the fraction of districts that never changes political color, because in that case the control of the assembly would be determined only by the outcome in the few contestable districts. Both i) and ii) are testable. Finally, note that Albouy (2013) already makes the same assumption with the aim to identify M_{it}, but he does not discuss the importance of controlling for M_{it} in order to identify the PE, which is our focus.

In reality D_{it} is not randomized and thus researchers rely on the RDD CE. In this design Calonico et al. (2019) recommend including controls, which is crucial in our identification strategy, only to improve precision and after checking that such controls are balanced at the threshold. This recommendation is based on the presumption that covariates imbalance might suggest that the potential outcome function is not continuous at the threshold, so that the crucial identifying assumption is violated. Furthermore, the authors add that covariates can be included to restore identification if the researchers are willing to impose additional assumptions. In our case, we are aware that M_{it} might not be balanced at the threshold, and that the outcome might be a function of it. In fact, as elaborated above, we propose to include M_{it} in the regression under the additional assumption that assembly control is determined at the national level.

Finally, note that if our argument does not convince the reader on the viability of controlling for M_{it}, it is always possible to balance the sample in terms of years with democratic/republican control, so that the correlation between M_{it} and D_{it} is negligible and is not necessary to include majority status. In practice, one may selectively drop years or, more efficiently, use post-stratification (Miratrix et al., 2013), i.e. re-weight the sample such that observations under the two types of years have equal weight.

4 Simulations

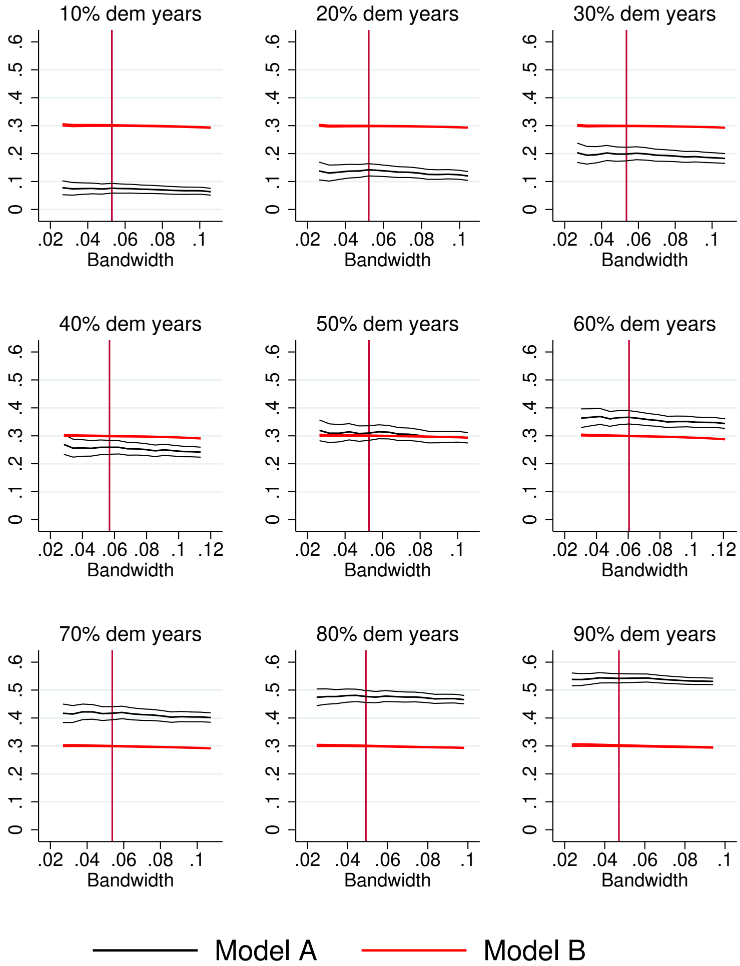
We simulate elections in 601 single-member districts to elect representatives of a parliament in a two-party system for 100 election years. The outcome Y_{it} is a function of majority status, party identity, the vote share X_{it} for the democratic party, and random components at the year and district level.[7]

We estimate two models: A) the standard one with a constant and D_{it}, and B) our specification augmented with M_{it} and year fixed effects. Both include a linear function in the margin of victory estimated separately on each side of the threshold. Figure 1 plots the point estimate of the coefficient on D_{it} for the two models together with the 95% confidence intervals (CI), as a function of the bandwidth. Crucially, the estimates are performed separately in nine different samples of 50 election years, each characterized by a different ratio of democratic to republican years, corresponding to the panels of Figure 1.

Model A (black) provides an unbiased estimate of the PE (i.e., 0.3) only when the sample is composed by the same number of democratic and republican years (central panel). In all other cases, the estimate is either upward biased (with more democratic years) or downward biased (with more republican years). The sign and size of the bias is thus consistent with what predicted in Section 3. On the contrary, model B (red) always estimates a coefficient centered on the true effect.

[7]See supplementary material C for details.

Figure 1: Estimates of PE in simulated data. True PE is 0.3.



Note: Vertical red lines indicate the optimal bandwidth by Calonico et al. (2014). Linear model estimated with OLS with standard errors adjusted for heteroskedasticity.

5 Evidence from real data on the U.S. House

We perform similar analyses on real data, aiming at showing that controlling for majority status can affect estimates of the PE in the predicted direction. Throughout the section, we present results from models A and B, as well as a third specification with both D_{it} and M_{it} but without fixed effects. For more details on data and estimation see supplementary materials, E, F and G.

5.1 Roll-call voting and incumbency advantage 1946-1994

We replicate the analysis in Lee et al. (2004) using the original dataset, which includes results for the U.S. House in the period 1946-1994, and voting scores of representatives on a right-left scale 0-100 based on roll-call votes. In this sample there is only one republican-controlled year. The authors use a RDD CE to estimate the PE on three outcomes: contemporaneous policy stance RC_{it}, policy stance in the next term RC_{it+1}, and the treatment in the next term D_{it+1} (incumbency advantage). Results, reported in Table 1, show that including majority status considerably affects the estimate of the coefficient on D_{it} for all outcomes.

Table 1: Replication of Lee et al. (2004)

	Outcome : RC_{it+1}			Outcome : RC_{it}			Outcome : D_{it}		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
D_{it}	20.75	13.15	17.63	48.28	60.99	57.91	0.530	0.337	0.389
	(1.98)	(2.84)	(2.94)	(1.30)	(1.87)	(1.93)	(0.058)	(0.069)	(0.064)
M_{it}		10.31	7.17		-14.18	-11.36		0.262	0.182
		(2.84)	(2.94)		(1.82)	(1.93)		(0.050)	(0.048)
Time-FE	No	No	Yes	No	No	Yes	No	No	Yes
Observations	887	887	887	955	955	955	887	887	887

Note: Linear model estimated with OLS without controlling for the margin of victory. Robust standard errors in parenthesis. Bandwidth = 2 percentage points.

Despite some differences, the qualitative conclusion in Lee et al. (2004) is robust to this replication. Nevertheless this exercise shows that the PE changes more than one would expect in a valid RDD CE when we control for majority status.

5.2 Roll-call voting 1947-2008

We extend the dataset in the previous section until 2008, obtaining a sample with 23 terms under democratic control, and 8 under republican control. The estimation is conducted separately on subsamples that feature a different ratio of observations from democratic- and republican-controlled years, resulting in different covariance between D_{it} and M_{it}. For simplicity, we only focus on the PE on contemporaneous roll-call voting RC_{it}. Table 2 reports the results. In the most balanced period

1982-2004 the correlation between D_{it} and M_{it} is close to zero. As expected, the coefficient on D_{it} is the same (approximately 56) irrespective of whether we control for majority status. The coefficient on M_{it} is approximately -5, suggesting that majority members have on average a less liberal stance compared to opposition members, holding party constant. Results from the other subsamples are broadly consistent with what predicted theoretically in Section 3: relative to 1982-2004, the coefficient on D_{it} in the model without M_{it} is lower the more democratic years (positive covariance), and higher the more republican years (negative covariance). Furthermore, in all partially unbalanced subsamples controlling for majority status yields a coefficient on D_{it} closer to 56, relative to the model without M_{it}. Introducing time fixed effects makes little difference. The results confirm our theoretical insights which, however, has a limited quantitative relevance in this application, due to the moderate effect of majority status on roll-call voting.

Table 2: Roll-call voting.

	1946-2006			Dem. control: 1978-1992		Rep. control: 1994-2004		1978-1994			1990-2004		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
D	49.48	53.72	53.59	48.32	48.43	61.80	61.79	50.35	53.49	53.43	58.93	57.54	57.47
	(1.54)	(1.51)	(1.51)	(3.19)	(3.25)	(2.90)	(2.97)	(2.83)	(2.66)	(2.71)	(2.48)	(2.51)	(2.53)
M		-6.73	-6.29						-4.83	-4.74		-3.80	-3.62
		(0.79)	(0.79)						(1.42)	(1.38)		(1.31)	(1.25)
Electoral cycle FE	No	No	Yes	No	Yes	No	Yes	No	No	Yes	No	No	Yes
Mean Y if D=1	67	67	67	68	68	75	75	69	69	69	73	73	73
Mean Y if D=0	15	15	15	16	16	11	11	15	15	15	12	12	12
Obs. in dem years (%)	78	78	78	100	100	0	0	86	86	86	32	32	32
Corr(D,M)	0.56	0.56	0.56	1.00	1.00	-1.00	-1.00	0.72	0.72	0.72	-0.36	-0.36	-0.36
Observations	3699	3682	3682	843	843	531	531	980	969	969	781	777	777

	1982-2004			Dem. control: 1954-1976		Rep. control: 1946+1952		1946-1976			1946-1958		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
D	55.90	56.36	56.31	44.17	44.55	57.08	56.95	47.70	52.48	51.63	49.22	52.02	51.28
	(2.53)	(2.46)	(2.46)	(2.44)	(2.40)	(3.44)	(3.31)	(1.89)	(2.02)	(1.96)	(2.31)	(2.37)	(2.31)
M		-5.01	-4.82						-5.88	-4.53		-4.82	-3.46
		(1.20)	(1.18)						(1.05)	(1.07)		(1.00)	(1.05)
Electoral cycle FE	No	No	Yes	No	Yes	No	Yes	No	No	Yes	No	No	Yes
Mean Y if D=1	71	71	71	63	63	63	63	64	64	64	64	64	64
Mean Y if D=0	13	13	13	19	19	7	7	16	16	16	14	14	14
Obs. in dem years (%)	54	53	53	100	100	0	0	88	88	88	74	74	74
Corr(D,M)	0.07	0.07	0.07	1.00	1.00	-1.00	-1.00	0.75	0.75	0.75	0.47	0.47	0.47
Observations	1145	1135	1135	1677	1677	279	279	2269	2264	2264	1067	1063	1063

Note: Linear model estimated with OLS controlling linearly for the margin of victory on each side of the threshold. Standard errors clustered at the electoral district. Bandwidth = 0.183 selected using the method by Calonico et al. (2014).

5.3 Electoral financing 1979-2006

We estimate the effect of a victory of the democratic party in a district on the campaign funds raised by the incumbent party in the next election.[8] Since most incumbents seek re-election, this is almost equivalent to testing whether democratic members raise more funds than their republican colleagues to finance their re-election campaign. This could happen if members of one party are on average more able to attract funds, or if donors have a partisan bias. The analysis is interesting in light of Cox and Magar (1999), who find that majority status yields an advantage in terms of campaign financing. The outcome is the amount of campaign funds (in thousands of 1990 dollars) raised in a district from non-investor donors by the party that won the previous election.

As before, in the balanced subsample (1978-2004) the coefficient on D_{it} is the same (approximately -133) irrespective of whether we control for majority status (see Table 3). Moreover, here the coefficient on M_{it} is sizable (80), and thus its omission makes for very large difference in the estimate of the coefficient on D_{it} in unbalanced subsamples: -51 in 1978-1992 versus -205 in 1994-2004. As before, controlling for majority status makes the estimate of the coefficient on D_{it} more similar across subsamples.

[8] Data is from Fouirnaies and Hall (2014) but our analysis is different and it is not a replication.

Table 3: Campaign financing.

	1978-2004			Dem. control: 1978-1992		1978-1994			Rep. control: 1994-2004		1990-2004		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
D	-132.83	-145.06	-133.02	-51.13	-33.76	-95.17	-127.84	-114.25	-205.34	-219.08	-165.61	-95.01	-108.75
	(45.92)	(47.76)	(41.57)	(37.92)	(30.51)	(38.75)	(43.57)	(31.63)	(84.42)	(79.65)	(75.32)	(73.49)	(66.90)
M		82.80	78.61				57.34	64.63				109.01	120.66
		(27.20)	(22.49)				(27.74)	(21.19)				(39.45)	(35.41)
Electoral cycle FE	No	No	Yes	No	Yes	No	No	Yes	No	Yes	No	No	Yes
Mean Y if D=1	327	327	327	220	220	256	256	256	461	461	442	442	442
Mean Y if D=0	467	467	467	258	258	324	324	324	669	669	622	622	622
Obs. in dem years (%)	52	52	52	100	100	80	80	80	0	0	16	16	16
Corr(D,M)	0.05	0.05	0.05	1.00	1.00	0.60	0.60	0.60	-1.00	-1.00	-0.68	-0.68	-0.68
Observations	1056	1056	1056	554	554	690	690	690	502	502	599	599	599

Note: Linear model estimated with OLS controlling linearly for the margin of victory on each side of the threshold. Standard errors clustered at the electoral district in parenthesis. Bandwidth = 0.09 selected using the method by Calonico et al. (2014).

6 Conclusion

We show how and when majority status can affect the interpretation of the PE in RDD CE studies. We propose an identification strategy based on controlling for majority status and validate it with simulated and real data, including those used in Lee et al. (2004). In the latter case, our specification does not alter the qualitative conclusion of the study, but in other applications the empirical relevance of our point is significant.

Despite our focus on first-past-the-post systems, where party and majority status are realized simultaneously, our argument is more broadly relevant to contexts where the alignment between different layers (local versus national) or branches (president versus parliament) of government is expected to matter. Furthermore, our paper is relevant not only for RDD CE studies, but also for other research designs aimed at estimating the PE, since our argument is not about failure of specific identification assumptions.

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Supplementary material

A Sample covariance between D_{it} and M_{it}

Denote by $m(\cdot)$, $s(\cdot)$, and $c(\cdot, \cdot)$ the sample mean, sample variance, sample covariance respectively. Notice that we have variables varying both within years ($t = 1, ..., T$) and districts ($i = 1, ..., n$). Let $\mathbf{D} = (\mathbf{D}_1, \mathbf{D}_2, ..., \mathbf{D}_T)$, where $\mathbf{D}_t = (D_{1t}, D_{2t}, ..., D_{nt})$ (define $\mathbf{M}$ and $\mathbf{M}_t$ similarly). The sample covariance between $\mathbf{D}_t$ and $\mathbf{M}_t$ is

$$\begin{aligned} c(\mathbf{M}_t, \mathbf{D}_t) =& \frac{1}{n-1} \sum_{i=1}^{n} [M_{it} - m(\mathbf{M}_t)][D_{it} - m(\mathbf{D}_t)] = \\ =& \frac{n}{n-1} \left[m(\mathbf{M}_t \mathbf{D}_t) - m(\mathbf{M}_t) m(\mathbf{D}_t) \right], \end{aligned} \tag{A1}$$

where $m(\mathbf{M}_t \mathbf{D}_t) = \frac{1}{n} \sum_{i=1}^{n} M_{it} D_{it} := \frac{1}{n} \mathbf{M}_t \cdot \mathbf{D}_t$, the operator "$\cdot$" being the inner product.

Notice that, from the definition of majority status it follows that

$$c(\mathbf{M}_t, \mathbf{D}_t) = \begin{cases} s(\mathbf{D}_t), & \text{if } m(\mathbf{D}_t) > 0.5 \\ -s(\mathbf{D}_t), & \text{if } m(\mathbf{D}_t) < 0.5. \end{cases} \tag{A2}$$

The average of the covariances across electoral years can be written as:

$$\frac{1}{T} \sum_{t=1}^{T} c(\mathbf{M}_t, \mathbf{D}_t) = \frac{n}{n-1} \left[\frac{1}{T} \sum_{t=1}^{T} m(\mathbf{M}_t \mathbf{D}_t) - \frac{1}{T} \sum_{t=1}^{T} m(\mathbf{M}_t) m(\mathbf{D}_t) \right]. \tag{A3}$$

Using (A1) and (A3), we can write the overall sample covariance as:

$$
\begin{aligned}
c(\mathbf{M},\mathbf{D}) =& \frac{nT}{nT-1}\left[m(\mathbf{MD}) - m(\mathbf{M})m(\mathbf{D})\right] = \\
=& \frac{nT}{nT-1}\left[\frac{1}{T}\sum_{t=1}^{T} m(\mathbf{M}_t\mathbf{D}_t) - \frac{m(\mathbf{M})}{T}\sum_{t=1}^{T} m(\mathbf{D}_t)\right] = \\
=& \frac{n}{nT-1}\left[\frac{n-1}{n}\sum_{t=1}^{T} c(\mathbf{M}_t,\mathbf{D}_t) + \sum_{t=1}^{T} m(\mathbf{M}_t)m(\mathbf{D}_t) - m(\mathbf{M})\sum_{t=1}^{T} m(\mathbf{D}_t)\right] = \\
=& \frac{n}{nT-1}\left[\underbrace{\frac{n-1}{n}\sum_{t=1}^{T} c(\mathbf{M}_t,\mathbf{D}_t)}_{A} + \underbrace{\sum_{t=1}^{T} m(\mathbf{D}_t)\left[m(\mathbf{M}_t) - m(\mathbf{M})\right]}_{B}\right].
\end{aligned}
\tag{A4}
$$

Now, the first element in the square parenthesis in (A4) (labeled as A) is not equal to zero in general. Using (A2) we can write, with a slight abuse of notation:

$$
\sum_{t=1}^{T} c(\mathbf{M}_t,\mathbf{D}_t) = \sum_{t\in DemYears} s(\mathbf{D}_t) - \sum_{t\in RepYears} s(\mathbf{D}_t). \tag{A5}
$$

The summation in equation (A5) is equal to zero if the sample features the same number of democratic-controlled years and republican-controlled years, and the variance of the treatment dummy is constant across years. It is important to notice that: a) the absolute value of the term A decreases as the dataset is more balanced in terms of democratic-controlled years and republican-controlled years; b) the term A increases as the fraction of democratic-controlled years increases; c) the term A decreases as the fraction of republican years increases.

The second element in the square parenthesis in (A4) (labeled as B) is never exactly equal to zero. In fact, we can write:

$$
m(\mathbf{M}_t) - m(\mathbf{M}) = \frac{1}{n}\sum_{i=1}^{n}\left[M_{it} - m(\mathbf{M}_t)\right], \tag{A6}
$$

which, in practice, is never equal to zero because $M_{it} \neq m(\mathbf{M}_t)$, unless all the districts are conquered by one party.[1] Nevertheless, the term B is likely to be often negligible,

[1] Majority status is a dummy, so its mean can not be equal to any value taken by the variable unless

as it involves differences between two numbers both between 0.5 and 1, than multiplied times a number between 0 and 1. As such, $A + B$ is in general different from zero.

B Saturated models and heterogeneous effects

The data generating process (DGP)

$$Y_{it} = \gamma_0 + \gamma_1 D_{it} + \gamma_2 M_{it} + \varepsilon_{it}. \tag{A7}$$

restricts the functional form of the conditional expectation function. In other words, it has only three parameters compared to the four groups of districts in the data: democratic districts that belong to majority, democratic districts that belong to opposition, republican districts that belong to majority, and republican districts that belong to opposition.[2] Let us assume instead the more general DGP that not only includes D_{it} and M_{it}, but also their interaction:

$$Y_{it} = \gamma_0 + \gamma_1 D_{it} + \gamma_2 M_{it} + \gamma_3 D_{it} \cdot M_{it} + \varepsilon_{it}. \tag{A8}$$

The model in (A8) is fully saturated, because it has one different parameter for each of the values taken by the conditional expectation function.[3] However, this model, even if saturated, does not allow to identify heterogeneous effects of D_{it} conditional on different value of M_{it}. To see why, consider that the quantity

$$\mathbb{E}[Y_{it}|D_{it} = 1, M_{it} = 1] - \mathbb{E}[Y_{it}|D_{it} = 0, M_{it} = 1] = \gamma_1 + \gamma_3$$

actually compares democratic districts in years when democrats have control of the house, to republican districts when republicans have control of the house. This opens the possibility that the estimate is biased by a *partisan effect* at the house level, or more generally

they are all zero (impossible), or all ones (one party wins all the seats).

[2]These groups can be described as: democratic districts in years when democrats hold control of the house, democratic districts when republicans hold control, republican districts when republicans hold control, republican districts when democrats hold control.

[3]The four values are: $\mathbb{E}[Y_{it}|D_{it} = 0, M_{it} = 0] = \gamma_0$; $\mathbb{E}[Y_{it}|D_{it} = 1, M_{it} = 0] = \gamma_0 + \gamma_1$; $\mathbb{E}[Y_{it}|D_{it} = 0, M_{it} = 1] = \gamma_0 + \gamma_2$; $\mathbb{E}[Y_{it}|D_{it} = 1, M_{it} = 1] = \gamma_0 + \gamma_1 + \gamma_2 + \gamma_3$.

by year-level confounders. Augmenting the specification in (A8) with an indicator variable for democratic control of the house, that is $\mathbb{1}(\overline{D}_t > 0.5)$, $\overline{D}_t = \sum_{i=1}^{n} D_{it}/n$, does not help. It actually results in perfect collinearity because districts represented by the democratic party, that belong to the majority, in years when the republicans hold control of the house do not exist by construction.[4] This fact is reflected in the possibility to rewrite (A8), as:

$$Y_{it} = \beta_0 + \beta_1 D_{it} + \beta_2 \mathbb{1}(\overline{D}_t > 0.5) + \beta_3 D_{it} \cdot \mathbb{1}(\overline{D}_t > 0.5) + \varepsilon_{it}, \tag{A9}$$

by using the definition

$$M_{it} = D_{it} \cdot \mathbb{1}(\overline{D}_t > 0.5) + (1 - D_{it}) \cdot [1 - \mathbb{1}(\overline{D}_t > 0.5)]. \tag{A10}$$

The coefficients in (A9) are such that $\gamma_0 = \beta_0 + \beta_2$, $\gamma_1 = \beta_1 - \beta_2$, $\gamma_2 = -\beta_2$, and $\gamma_3 = \beta_3 + 2\beta_2$. Yet a different way to write the exact same model is the following:

$$Y_{it} = \alpha_0 + \alpha_1 D_{it} + \alpha_2 M_{it} + \alpha_3 \mathbb{1}(\overline{D}_t > 0.5) + \varepsilon_{it}, \tag{A11}$$

where $\beta_0 = \alpha_0 + \alpha_2$, $\beta_1 = \alpha_1 - \alpha_2$, $\beta_2 = \alpha_3 - \alpha_2$ and $\beta_3 = 2\alpha_2$. Use the definition of M_{it} in (A10) into (A11) to obtain (A9). This model is analogous to the reduced-form model in Albouy (2013), that includes D_{it} and M_{it}, and year fixed effects.

To sum up the models in (A8), (A9) and (A11) are equivalent and even if they do not restrict the functional form of the DGP, they do not allow to identify heterogeneous effects of D_{it} with respect to M_{it}. However, it is possible to identify the arithmetic average between the effect of D_{it} when democrats have majority status and the effect of D_{it} when democrats have opposition status. We define this as the average *partisan effect* (PE)[5]. The PE can be estimated by either one of equations (A8), (A9) and (A11):

$$\text{PE} = \alpha_1 = \beta_1 + \beta_3/2 = \gamma_1 + \gamma_3/2. \tag{A12}$$

[4] In other words, there would be five parameters for the same four values of the conditional expectation function.

[5] Of course in a RD setting the PE will be local in the sense that it applies only to observations in the neighborhood of the threshold.

B.1 The average *partisan effect*

Assume that each district has four potential outcomes: $Y_{it}^{D,M}, Y_{it}^{D,O}, Y_{it}^{R,M}, Y_{it}^{R,O}$, where the first apex refers to the party (democrat or republican) and the second to the majority status (majority or opposition). Let δ_t be a dummy for D having the majority at t: $\delta_t = \mathbb{1}(\overline{D}_t > 0.5)$. The observed outcome is thus:

$$\begin{aligned} Y_{it} =& D_{it} \cdot \delta_t \cdot Y_{it}^{D,M} + \\ & D_{it} \cdot (1 - \delta_t) \cdot Y_{it}^{D,O} + \\ & (1 - D_{it}) \cdot \delta_t \cdot Y_{it}^{R,O} + \\ & (1 - D_{it}) \cdot (1 - \delta_t) \cdot Y_{it}^{R,M}. \end{aligned} \tag{A13}$$

We are interested in identifying the *partisan effect* (PE), defined as:

$$\begin{aligned} \beta = \text{PE} &= 1/2 \cdot \left[Y_{it}^{D,M} + Y_{it}^{D,O} - Y_{it}^{R,M} - Y_{it}^{R,O} \right] \\ &= \underbrace{1/2 \cdot [Y_{it}^{D,M} + Y_{it}^{D,O}]}_{\textit{average potential outcome if democrat}} - \underbrace{1/2 \cdot [Y_{it}^{R,M} + Y_{it}^{R,O}]}_{\textit{average potential outcome if republican}} \\ &= 1/2 \cdot [\underbrace{Y_{it}^{D,M} - Y_{it}^{R,M}}_{\textit{PE on the majority members}} + \underbrace{Y_{it}^{D,O} - Y_{it}^{R,O}}_{\textit{PE on the opposition members}}]. \end{aligned} \tag{A14}$$

The PE has an intuitive interpretation: it can be written as the difference between the average potential outcome when the district is democrat and the average potential outcome when the district is republican (second line of (A14)) or, equivalently, as the average between the PE on the majority members and the PE on the opposition members (third line of (A14)).

C Main simulation

Here we provide additional details on the simulation used in the paper. We take the number of districts n equal to 601, and the number of election-years T equal to 100.[6] For each election-year t we proceed as follows: first, we draw the identity of the party who holds control of the assembly, with probability 0.5 each. The vote share for the democratic

[6]The number of districts is of the same order of magnitude of real-world lower houses.

party in each district i is then drawn from a beta distribution:

$$X_{it} \sim Beta(\vartheta_t, 10 - \vartheta_t), \tag{A15}$$

where ϑ_t depends on which party holds control of the assembly. In particular, ϑ_t is drawn from a uniform $\mathcal{U}[5.1, 5.5]$ if the democrats hold control of the assembly, and from $\mathcal{U}[4.5, 4.9]$ if republicans hold control, to make sure that $E[X_{it}] > 0.5$ in case of democratic control, and $E[X_{it}] < 0.5$ in case of republican control.[7] The variables D_{it} and M_{it} follow from X_{it}.

We assume the following DGP for the outcome:

$$\begin{aligned} Y_{it} = \; & 0.5 + 0.3D_{it} + 0.3M_{it} + 0.5\mathbb{1}(\overline{D}_t > 0.5) + \\ & + 20X_{it}^3 - 20X_{it}^2 + 2X_{it} + 0.5 + \\ & + \theta_t \sim \mathcal{N}(0, 0.05) + \varepsilon_{it} \sim \mathcal{N}(0, 0.03). \end{aligned} \tag{A16}$$

The PE is thus equal to 0.3.

Table B1 reports: the summary statistics of key variables, separately for years with democratic majority and republican majority (upper panel); the correlation coefficients between some of the key variables (central panel); the correlation coefficients and covariances between D_{it} and M_{it} in sub-samples with different ratios of democratic to republican years (lower panel). Note that when the balance between democratic-controlled years and republican-controlled years is perfect, the covariance between M_{it} and D_{it} is zero. Instead, when we consider different sub-samples, the covariance increases as the fraction of democratic years increases, while it decreases as the fraction decreases.

[7] Note that $E[X_{it}] = \vartheta_t/10$, so in years of democratic control the mean of the distribution is between 0.51 and 0.55, and in years of republican control is between 0.45 and 0.49.

Table B1: Summary statistics - simulated data.

	Republican majority		Democratic majority	
	mean	sd	mean	sd
Democrats' vote share, X_{it}	0.468	0.150	0.531	0.151
Democratic seat (0/1), D_{it}	0.416	0.493	0.579	0.494
Majority status (0/1), M_{it}	0.584	0.493	0.579	0.494
Interaction term (0/1), $D_{it} \times M_{it}$	0.000	0.000	0.579	0.494
Democratic majority (0/1), $\mathbb{1}(\overline{D}_t > 0.5)$	0.000	0.000	1.000	0.000
Outcome variable, Y_{it}	0.063	0.443	0.500	0.327
$n \times T$	28247		31853	

$\mathbb{C}$or(., .)	X_{it}	D_{it}	M_{it}	$M_{it} \times D_{it}$	$\mathbb{1}(\overline{D}_t > 0.5)$
Democratic seat, D_{it}	0.82	1.00			
Majority status, M_{it}	0.05	0.06	1.00		
Interaction, $M_{it} \times D_{it}$	0.58	0.66	0.56	1.00	
Democratic majority, $\mathbb{1}(\overline{D}_t > 0.5)$	0.21	0.16	-0.00	0.63	1.00
Outcome variable, Y_{it}	-0.21	-0.15	0.40	0.40	0.49

% dem. maj. years	10	20	30	40	50	60	70	80	90
$\mathbb{C}$or(D_{it}, M_{it})	-0.80	-0.59	-0.39	-0.19	0.00	0.20	0.40	0.60	0.80
$\mathbb{C}$ov(D_{it}, M_{it})	-0.19	-0.15	-0.10	-0.05	0.00	0.05	0.10	0.15	0.20

D Alternative simulation

In this alternative simulation, we attempt to produce a distribution of vote share across districts that is more similar to the actual distribution in the U.S. House. We take again the number of districts n equal to 601; here we assume that 51 districts are highly competitive, 275 are democratic-leaning and 275 republican-leaning. We take the the number of election-years T equal to 100. For each election-year t we proceed as follows: first, we draw the identity of the party who holds control of the assembly, with probability 50% each. The vote share for the democratic party in each of the 51 competitive districts is then drawn from a beta distribution:

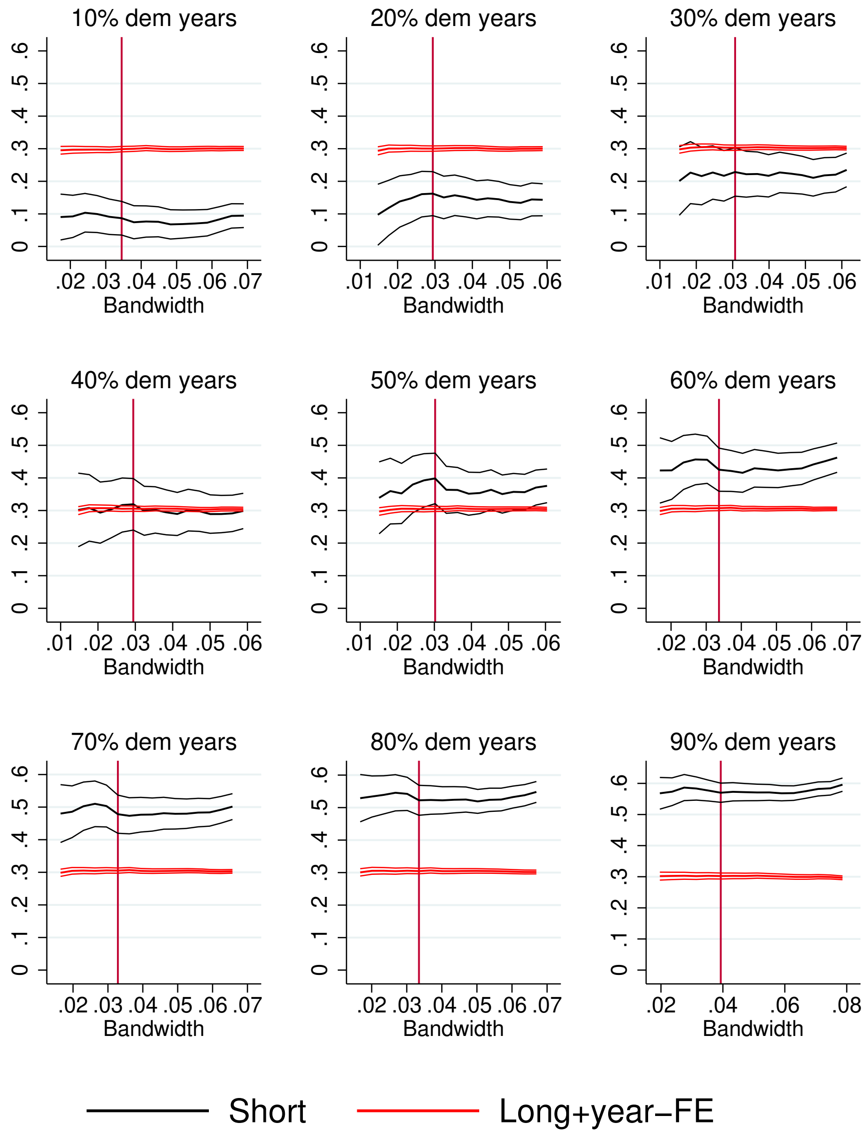
$$X_{it} \sim Beta(100\vartheta_t, 100(1 - \vartheta_t)), \tag{A17}$$

where ϑ_t depends on which party holds control of the assembly. In particular, ϑ_t is drawn from a uniform $\mathcal{U}[0.51, 0.55]$ if the democrats hold control of the assembly, and from $\mathcal{U}[0.45, 0.49]$ if republicans hold control.[8] The vote share for the other districts is drawn from a beta distributions with parameters 250 and 150 in case of democratic-leaning districts, and 150 and 250 in case of republican-leaning districts[9]; the seat in these districts can be only occasionally won by the underdog party. The final distribution of X_{it} is thus trimodal, and in the RD design the estimating sample will be made mainly by highly competitive districts, as happens in real applications. The rest of the exercise is the same as in the baseline simulation. The results are in line with those obtained using the baseline simulation. The model that controls for majority status and time fixed effects performs well in all subsamples; the standard model is more biased the more unbalanced is the sample.

[8]In this way, in years of democratic control the expected value of the distribution in the competitive districts is between 0.51 and 0.55, and in years of republican control is between 0.45 and 0.49.

[9]This corresponds, to an expected value of 5/8 for democratic-leaning districts and of 3/8 for republican-leaning districts.

Figure B1: Estimates of *partisan effect* in simulated data. True effect=0.3.



Note: each panel reports estimates of the *partisan effect* α_1 from a different sub-sample of 50 election-years, with different ratios of democratic years reported below. Estimates and 95% confidence interval are plotted against the bandwidth used. Vertical red lines indicate the optimal bandwidth by Calonico et al. (2014). Estimation by OLS, and standard errors adjusted for heteroskedasticity. The "short" model include as regressors: D_{it}, the margin of victory, and its interaction with an indicator for observations to the right of the threshold. The "long+year-FE" model control for both majority status and year fixed effects. The true *partisan effect* is equal to 0.3 .

E Data

E.1 Replication of Lee et al. (2004)

The dataset in Lee et al. (2004) includes electoral results for the U.S. House in the period 1946-1994, and voting scores of House representatives on a right-left scale 0-100 based on high-profile roll-call votes.[10] The unit of analysis is the district-year. The timing notation is as follows: t denotes electoral terms, so $t = 1984$ denotes the election in November 1984, and congressional voting in years 1985 and 1986 (U.S. House representatives are elected every two years.). The authors drop the years that ends with two because they correspond to the time when the boundaries of the district change. They also drop observations for which either D_{it} or D_{it-1} are missing. The final sample is thus composed by electoral terms $t =$1948, 1950, 1954, 1956, 1968, 1960, 1964, 1966, 1968, 1970, 1974, 1976, 1978, 1980, 1984, 1986, 1988, 1990. However, in all these terms the House was under democratic control. Therefore, we introduce back in the sample $t = 1946$ to break the perfect correlation between D_{it} and M_{it}. Summary statistics of the key variables are reported in this Online appendix.

E.2 Roll-call voting in U.S. House 1947-2008

We download data on U.S. House elections held between 1946 and 2006 from the Constituency-level election archive (Kollman et al., 2016) maintained by the University of Michigan.[11] We follow Lee et al. (2004) in measuring roll-call voting on the liberal-conservative scale using the ADA scores adjusted according to the methodology by Groseclose et al. (1999). In particular, we download the dataset by Anderson and Habel (2009), who make available this measure until 2008.[12] We match the two datasets by name, surname, state and election year, collapse the data at the electoral term-district level, and use as outcome the adjusted ADA score averaged across the term.

[10]The measure used is the voting score constructed by Americans for Democratic Action (ADA). It is based on about twenty high-profile roll-call votes per Congress, and ranges from 0 to 100, where lower score represents more conservative voting record. The measure is adjusted to ensure comparability over time following Groseclose et al. (1999).

[11]`http://www.electiondataarchive.org/`

[12]`dataverse.harvard.edu/dataset.xhtml?persistentId=hdl:1902.1/12339`

E.3 Electoral financing in U.S. House 1979-2006

We download the replication data of the paper by Fouirnaies and Hall (2014). They estimate the incumbency advantage in campaign financing in the U.S. House. They find that the incumbent party raises more funds than the other party. Data available at: `stanforddpl.org/papers/fouirnaies_hall_financial_incumbency_2014`. The dataset includes information on campaign financing for U.S. House elections held between 1980 and 2006, and electoral results for U.S. House elections held between 1978 and 2004. The original source of the data on campaign financing is the U.S. Federal Election Commission. The time coverage includes both democratic-controlled years and republican-controlled years. We take as outcome variable the campaign funds raised for the election at $t+1$ in district i by the party that won the election at t in i. We exclude from the outcome variable funds from "investor" donors. Investor donors include the categories of donors that finance candidates in exchange for policy favors, and not on ideological grounds (Snyder, 1990; Fouirnaies and Hall, 2014). These include Political Action Committees (PACs) connected with corporations, cooperatives, and Trade, Health and Membership PACs. The categories included in our outcome variables are mainly "consumer" donors (individuals and non-connected PACs), and party contributions.[13] The outcome variable is measured in thousands of 1990 U.S. dollars. The running variable is the margin of victory which is calculated, slightly differently from what used elsewhere in this paper, using the democratic party's share of the total votes received by Democrats and Republicans in i at t.[14]

F Additional empirical results

F.1 Replication of Lee et al. (2004)

The reader may wonder if the changes in the coefficients are due to a general violation of the assumption of quasi-random assignments, rather than due to the relationship between D_{it} and M_{it}. To test this, we augment the specification with a vector of representative's

[13] We exclude "investor" donors because the estimates obtained using those categories as outcome variable are small and not significant, and so not very useful to illustrate the confounding role of majority status. These estimates are available upon request.

[14] We use the same running variable as in Fouirnaies and Hall (2014).

	Republican majority		Democratic majority	
	mean	s.d.	mean	s.d.
Democrats' margin of victory	0.047	0.246	0.082	0.230
Democratic seat (0/1), D_{it}	0.420	0.494	0.596	0.491
Majority status (0/1), M_{it}	0.580	0.494	0.596	0.491
Interaction term (0/1), $D_{it} \times M_{it}$	0.000	0.000	0.596	0.491
Democratic majority (0/1), $\mathbb{1}(\overline{D}_t > 0.5)$	0.000	0.000	1.000	0.000
ADA score, RC_{it}	22.326	29.411	41.914	32.633
Observations	791		13577	

Table B2: Summary statistics of key variables in Lee et al. (2004)

characteristics available in the replication data: age, gender, education, occupation, military service and an indicator for having a relative in politics.[15] If the assumption of quasi-random assignments is violated, the introduction of controls that have predictive power on the outcome would potentially affect the coefficient on D_{it}. This is not the case as shown in Table B3: for all three outcomes the coefficient on D_{it} barely changes when we add controls, even if a joint test of significance of these variables rejects the null at conventional significance levels (columns 4 to 6).

Finally, we test the robustness of our results to the choices of bandwidth and estimator. We focus on two models: the model with only D_{it}, and our preferred specification which controls for majority status and time fixed effects. Here control for a linear function in the margin of victory on each side of the threshold and we report estimates obtained using bandwidths between 3.25 and 12 percentage points.[16] The estimates, reported in Figure B2 along with 95% confidence intervals, draw a similar picture as those in Table B3. Our preferred specification (in red) delivers an higher estimate than the model with only D_{it} (in black) for RC_{it}, and a lower one for RC_{it+1} and D_{it+1}.

F.2 Roll-call voting in U.S. House 1947-2008

[15] We pick these control variables because they are readily available in the replication dataset. There is evidence that some of these politicians' characteristics affect policy in other contexts (Clots-Figueras, 2011; Lahoti and Sahoo, 2020; Alesina et al., 2019).

[16] The optimal bandwidth by Calonico et al. (2014) is 6 percentage points when the outcome is D_{it+1} or RC_{it+1}, and 7.5 percentage points when the outcome is RC_{it}.

Table B3: Replication of Lee et al. (2004): additional controls

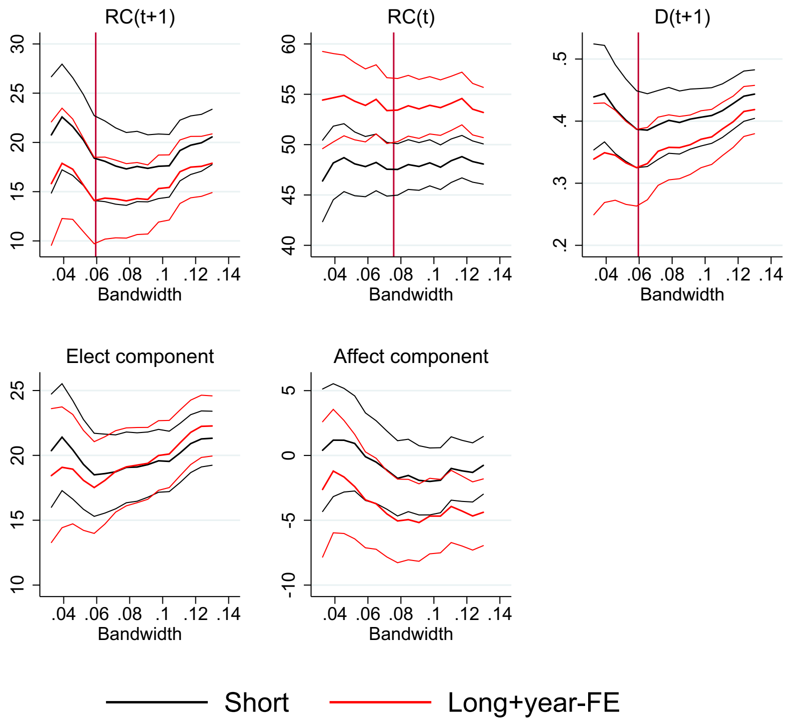
	Outcome variable: RC_{it+1}					
	(1)	(2)	(3)	(4)	(5)	(6)
D_{it}	20.75	13.15	17.63	18.84	12.20	16.41
	(1.98)	(2.84)	(2.94)	(2.06)	(2.97)	(3.06)
M_{it}		10.31	7.17		9.10	5.62
		(2.84)	(2.94)		(2.94)	(3.07)
Time-FE	No	No	Yes	No	No	Yes
Controls	No	No	No	Yes	Yes	Yes
P-value controls				0.09	0.16	0.00
Observations	887	887	887	887	887	887

	Outcome variable: RC_{it}					
	(1)	(2)	(3)	(4)	(5)	(6)
D_{it}	48.28	60.99	57.91	45.83	59.57	58.33
	(1.30)	(1.87)	(1.93)	(1.36)	(1.88)	(2.08)
M_{it}		-14.18	-11.36		-15.25	-14.45
		(1.82)	(1.93)		(1.78)	(2.11)
Time-FE	No	No	Yes	No	No	Yes
Controls	No	No	No	Yes	Yes	Yes
P-value controls				0.00	0.00	0.00
Observations	955	955	955	955	955	955

	Outcome variable: D_{it+1}					
	(1)	(2)	(3)	(4)	(5)	(6)
D_{it}	0.530	0.337	0.389	0.540	0.350	0.388
	(0.058)	(0.069)	(0.064)	(0.059)	(0.069)	(0.064)
M_{it}		0.262	0.182		0.267	0.186
		(0.050)	(0.048)		(0.049)	(0.048)
Time-FE	No	No	Yes	No	No	Yes
Controls	No	No	No	Yes	Yes	Yes
P-value controls				0.04	0.03	0.00
Observations	887	887	887	887	887	887

Note: OLS regressions without controlling for the margin of victory. Robust standard errors in parenthesis. Observations included only if the margin of victory is between ± 2 percentage points. Controls include dummies for age, gender, relative who served, secondary education, college, last occupation and military service.

Figure B2: Replication of Lee et al. (2004): bandwidth robustness

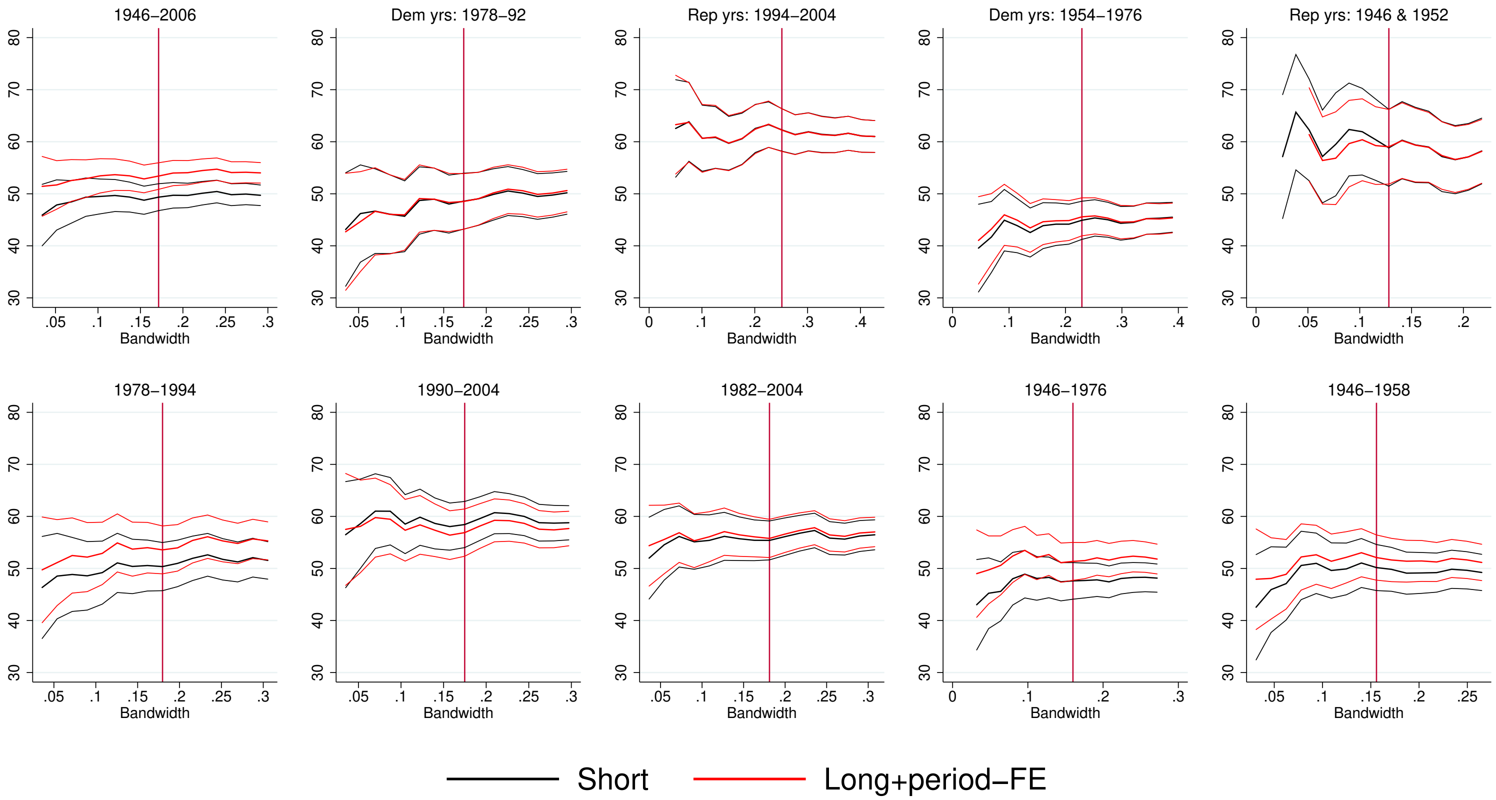


Note: The three upper panel report RD estimates of the *partisan effect* and 95% confidence interval plotted against the bandwidth used. Vertical red lines indicate the optimal bandwidth by Calonico et al. (2014). Estimation by OLS, and standard errors adjusted for heteroskedasticity. The "short" model includes: D_{it}, the margin of victory, and its interaction with an indicator for observations to the right of the threshold. The "long+year-FE" model also controls for majority status and year fixed effects. The elect component is the product of the estimates in the central and right upper panels. The *affect* component is the difference between the estimate in the upper left panel and the *elect* component.

	Republican majority		Democratic majority	
	mean	s.d.	mean	s.d.
Democrats' margin of victory	0.050	0.380	0.117	0.377
Democratic seat (0/1), D_{it}	0.496	0.500	0.580	0.494
Majority status (0/1), M_{it}	0.504	0.500	0.580	0.494
Interaction term (0/1), $D_{it} \times M_{it}$	0.000	0.000	0.580	0.494
Democratic majority (0/1), $\mathbb{1}(\overline{D}_t > 0.5)$	0.000	0.000	1.000	0.000
ADA score, RC_{it}	41.552	36.340	43.730	32.436
Observations	2785		8468	

Table B4: Summary statistics, U.S. House electoral terms 1947-2008

Figure B3: *Partisan effect* and *majority status effect* on conservativeness in roll-call voting.



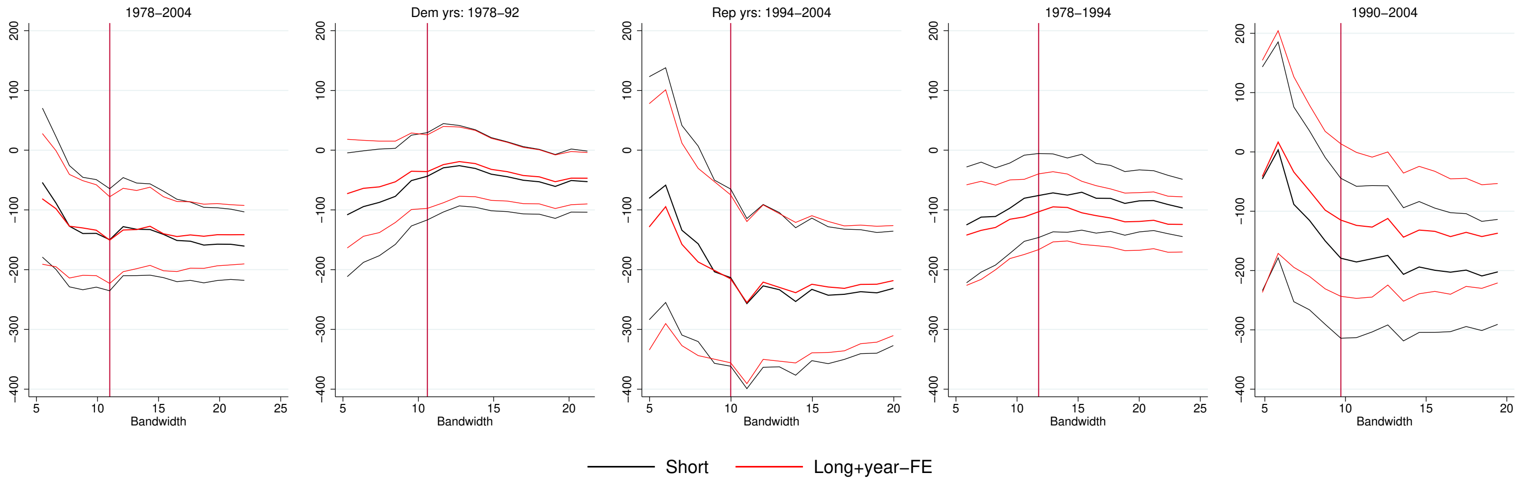
Note: RD estimates of the *partisan effect* and 95% confidence intervals plotted against the bandwidth used. Outcome variable: adjusted ADA score (Groseclose et al., 1999; Anderson and Habel, 2009). Lower values of the ADA score represents more conservative roll-call voting; higher values, more liberal roll-call voting. Vertical red lines indicate the optimal bandwidth by Calonico et al. (2014). Estimation by OLS, and standard errors clustered at the district level. The "short" model includes: D_{it}, the margin of victory, and its interaction with D_{it}. The "long+year-FE" model also controls for majority status and electoral term fixed effects.

F.3 Electoral financing in U.S. House 1979-2006

	Republican majority		Democratic majority	
	mean	s.d.	mean	s.d.
Margin of victory	-0.034	22.325	4.946	24.596
Democratic seat (0/1), D_{it}	0.479	0.500	0.588	0.492
Majority status (0/1), M_{it}	0.521	0.500	0.588	0.492
Interaction term (0/1), $D_{it} \times M_{it}$	0.000	0.000	0.588	0.492
Democratic majority (0/1), $\mathbb{1}(\overline{D}_t > 0.5)$	0.000	0.000	1.000	0.000
Funds at $t + 1$ for incumbent party	423.950	385.118	191.163	230.149
Observations	1945		2383	

Table B5: Summary statistics, U.S. House electoral terms 1979-2006 from Fouirnaies and Hall (2014)

Figure B4: *partisan effect* and majority status con campaign financing



Note: Outcome variable are campaign funds from non "investor" donors in thousands of 1990 U.S. dollars (Fouirnaies and Hall, 2014). RD estimates of the *partisan effect* and 95% confidence intervals plotted against the bandwidth used. Vertical red lines indicate the optimal bandwidth by Calonico et al. (2014). Estimation by OLS, and standard errors clustered at the district level. The "short" model includes: D_{it}, the margin of victory, and its interaction with D_{it}. The "long+year-FE" model also controls for majority status and electoral term fixed effects.

G Additional details on the replication of Lee et al. (2004)

The research question in Lee et al. (2004) is the following: do voters affect or merely elect policies? To answer, the authors rely on U.S. House district-level election data. They use a RD design to estimate the causal effect of having the democratic party in office (the treatment is D_{it}) on three outcome variables: a measure of policy stance on a right-left scale, RC_{it}, the same measure in the subsequent term, RC_{it+1}, and the treatment variable itself in the next election D_{it+1}. Their test is inspired by the model in Alesina (1988), and its logic can be explained as follows. The effect of D_{it} on RC_{it+1} can be decomposed into two components: on the one hand, D_{it} affects the equilibrium probability that democrats will be in office next term as well, and therefore will implement their preferred policy: the *elect* component; on the other hand, D_{it} affects the underlying popularity of the democratic party, and therefore the extent to which the democrats must compromise on their policy stance to please the electorate: the *affect* component. The *elect* component can be estimated separately as the product between the effect of D_{it} on D_{it+1}, and the effect of D_{it} on RC_{it}. Finally, the *affect* component is obtained by subtracting the *elect* component from the joint effect. The strategy is formalized in the following equations:

$$RC_{it+1} = \quad constant + \pi_1 D_{it} + \varepsilon_{it} \tag{A18}$$

$$RC_{it} = \quad constant + \pi_2 D_{it} + \varepsilon_{it} \tag{A19}$$

$$D_{it+1} = \quad constant + \pi_3 D_{it} + \varepsilon_{it} \tag{A20}$$

$$\pi_1 = \quad elect\ component + affect\ component \tag{A21}$$

$$\pi_2 \cdot \pi_3 = \quad elect\ component \tag{A22}$$

Despite the differences in some of the RD estimates, the main qualitative conclusion in Lee et al. (2004) is robust to our replication exercise. The estimates of the *elect* component are large and positive with or without controlling for majority status and time fixed effects. To see why, recall that the *elect* component is the product between π_2, whose estimate is higher using our specification, and π_3, whose estimate is lower using our specification. Our preferred specification delivers a lower estimate of the *affect* component, but still not significantly different from zero for many bandwidth choices. The overall conclusion is that the *elect* component largely dominates the *affect* component in elections to the U.S. House, as in the original paper.[17]

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