$2 + 1 < 3$: Strategic voting in the Chilean Congressional Elections*

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Abstract

In 1989 Chile ended 17 years of dictatorship and resumed its democratic tradition with presidential and congressional elections. I take advantage of the peculiar features of the electoral system in congressional elections (commonly known as binominal) to test for the presence of strategic voting. I find that, under the preferred specification, having an election close to adoubling threshold (i.e. having both candidates elected from the same coalition) decreases the vote share of small coalitions by between 6.58% and 10.69%. The results are robust to different specifications, and to the inclusion of the lagged share of votes for small coalitions and the number of candidates as additional covariates. Most of the strategic behavior comes from supporters of small right-wing coalitions, that shift their vote to the center-right when the center-left is close to the doubling threshold.

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"Our political democracy is incomplete. As an example, electorally two-thirds are equal to one-third. A Chilean mathematical miracle."
—Marco Enríquez-Ominami, presidential candidate in the 2009 election

1 Introduction

Are voters strategic when they cast their vote? In most theoretical work on electoral competition it is assumed that individuals cast their vote either sincerely (according to their true preferences) or strategically (taking into account how other voters cast their vote). Since the work of Downs (1957) and Duverger (1964), strategic voting has played a fundamental role in understanding voters' behavior. Moreover, other theories trying to explain voters' behavior, such as expressive voting (Brennan and Hamlin, 1998), the existence of ethic voters (Feddersen and Sandroni, 2006), or voters that learn how to vote by trial and error (Bendor, Diermeier, and Ting, 2003) are not easy to reconcile with voters who behave strategically.

But whether strategic voting behavior is observed in elections is an empirical question. In fact, there is abundant evidence of strategic behavior in laboratory experiments. Palfrey (2009) surveys experimental work in three different settings (voter turnout, jury decision making, and elections with the swing voter's curse), and finds an overwhelming support for strategic behavior. But evidence from actual elections is more elusive. Cox (1997) analyzes different electoral systems — among those the Chilean electoral law — finding strong support for what he calls a generalization of Duverger's Law, the "M+1 rule" (that no more than M+1 candidates are viable in an M-seat district). Also in a comprehensive survey, Blais (2000) looks at how rational choice theory can explain the decision to vote. His conclusion is that election closeness matters, but the mechanism is not clear since people find hard to compute pivotal probabilities (the probability that their vote might decide the outcome of the election).

Alvarez and Nagler (2000) and Alvarez, Boehmke, and Nagler (2006) use individual data from surveys conducted during the British parliamentary elections to induce voters' preferences for candidates. They find that voters are less likely to vote for parties with little chance of winning, but find little evidence that voters vote more for their second-preferred candidate when the election is close. Fujiwara (2008) implements a regression discontinuity design to test for strategic voting using Brazilian municipal election data, finding support for the so called Duverger's Law, i.e. that in single-ballot elections voters' support for third placed candidates is significantly smaller than in double-ballot elections. Finally, a closely related paper is by Kawai and Watanabe (2010), who test for strategic voting using Japanese election data. This paper is interesting for a couple of features. First, the authors distinguish between strategic and misaligned voting, where the latter occurs when voters do not vote for their most preferred candidate (while a strategic voter could vote for his/her most preferred candidate while still taking into account other voters' decisions). Second, in order to find the share of strategic voters, they estimate a discrete choice model of voter's behavior, much in line with the literature following Berry, Levinsohn and Pakes (1995). They find that a large share of the electorate (75.3% to 80.3%) is strategic, while only a small fraction cast a misaligned vote (2.4% to 5.5%). Nonetheless, a key assumption in their identification strategy is that beliefs about race closeness do not change with sociodemographic covariates while preferences do.

The approach I follow in this paper aims to address the shortcomings encountered in the previous literature by taking advantage of three particular features of the Chilean electoral system. First, I make use of the fact that under the Chilean electoral system, commonly known as *binominal*, the two candidates with the largest share of votes might not necesarily be the two elected legislators in a given district. This outcome can only take place if the votes received by the most voted coalition more than doubles the number of votes of the following coalition, a phenomenon known as *doubling*. On the contrary, if the first coalition gets fewer than double the votes of the second coalition, a candidate from the latter coalition will be elected, regardless of the number of votes he or she receives. This mechanism makes the identification of strategic behavior for supporters of small coalitions much simpler: The hypothesis is that when the election is far from a *doubling* situation, there are no incentives for voters supporting small coalitions or independent candidates to vote strategically, since they are not going to alter the outcome of the election. But when the election is close to *doubling*, voters supporting small parties will find more useful to cast their vote either to help the largest coalition to get their second legislator or to help the second largest coalition not to lose its legislator. It is true that this empirical test will not capture other possible sources of strategic behavior, such as the decision to vote for a second-placed candidate in the largest coalition against the first placed candidate in the second largest.[1] However, finding support for strategic behavior with this exercise makes it hard to reject strategic behavior in elections overall.

The second feature of the Chilean elections that I take advantage of is the fact that pre-election polls are rare in congressional elections and even much so in Lower House elections.[2] This supports the use of previous elections as a reasonable proxy for voters' beliefs about the closeness of the following election. Finally, the third feature of the

[1]This source of strategic behavior is discussed in Cox (1997).

[2]This does not mean that candidates do not conduct their own pre-election polls, but rather that these results are not publicly available. Most polling companies follow presidential elections very close, but fail to provide public information on lower house elections. The most prestigious polling company, Centro de Estudios Publicos (CEP) only provides information on the presidential election and the popularity of "prominent" figures in the political arena.

Chilean system is mandatory voting, which makes abstention less of a concern.[3]

I estimate the model using a semi-parametric approach. Under the most preferred specification (controlling for year and district dummies), I find that having an election close to the doubling threshold reduces the share of votes going to small coalitions in between 6.58% and 10.69% (between 0.55% and 0.89% of the total number of votes). This reduction disappears when the election gets farther from the doubling threshold. One caveat of these findings is that I am not able to distinguish between strategic and misaligned voting, as in Kawai and Watanabe (2010), and therefore the results should be taken as lower-bound estimates of strategic voting. The results are robust to different specifications, and to the inclusion of the lagged share of votes for small coalitions and the number of candidates as additional covariates. Most of the strategic behavior comes from supporters of small right-wing coalitions, that shift their vote to the center-right when the center-left is close to the doubling threshold.

The Chilean binominal system and the literature fostered by its application are examined in detail in section 2, whereas section 3 discusses the empirical strategy formally, as well as data issues. Section 4 presents the results and robustness checks. Finally section 5 concludes.

2 Chilean congressional elections: the binominal system and its performance

In 1989 Chile ended 17 years of dictatorship and resumed its democratic tradition with presidential and congressional elections. Nonetheless, the electoral system implemented for that election was not the one which was operative in 1970, the previous congressional election. A new constitution enacted in 1980[4] together with an electoral bill laid down in 1988[5] changed the electoral system from a proportional one to the binominal system that currently operates.[6]

The Congress therefore consists of two chambers: The *Senado* (Senate or upper house) and the *Cámara de Diputados* (lower house, equivalent to the U.S. House of Representatives). The former has 36 members that represent 18 two-member circumscriptions, while the latter has 120 members representing 60 two-member districts. Each circumscription groups between 1 and 8 districts. In what follows I use the terms circumscription and district interchangeably. Congressmen are elected every 4 years, while

[3] Cerda and Vergara (2009) analyze voters' turnout in Chile using both aggregate and individual data, and conclude that the observed decline in turnout is mainly due to low participation of the youth. This in turn is due to under-registration of this group (registration is voluntary), and not due to a lower level of participation once registered to vote.

[4] Constitución Política de Chile (1980).

[5] Ley 18700 Orgánica Constitucional sobre Votaciones Populares y Escrutinios.

[6] See Navia (2004) for a complete description of the origins of the binominal system in Chile.

senators' tenure is 8 years, with half of the chamber renewed each 4 years.

All parties and independent candidates have to join a *pacto* or *lista* (from now on, coalition). These coalitions have to be the same in all circumscriptions and districts, but a coalition can abstain of presenting candidates in a given district. A coalition cannot present more than 2 candidates per district. If independent candidates do not want to join a coalition, they will be labeled in the ballot as "Independent". Each list is open, so the electorate cast their vote directly for the candidate (one vote for a single candidate). Figure 1 shows a typical ballot for the lower house elections in 2005. In this case 3 coalitions presented 2 candidates each.

To determine the winners in each district, it is necessary to pool the votes by coalition and allocate the seats by the D'Hondt method: As explained before, if the coalition with the largest share of votes more than doubles the number of votes of its follower coalition, a second legislator of the former coalition is elected. Otherwise the largest coalition elects one candidate, as well as the second-largest coalition, regardless of the number of votes that the latter list receives.

To better understand this rule, consider the following setting. There are three coalitions, A, B and C, with $\#A \geq \#B > \#C$ and $\#X$ means "number of votes for coalition X". Note that I am restricting the number of votes of coalition C to be strictly smaller than those of coalition B. If this is not the case, there could be another source of strategic behavior for small parties' supporters, since it would be possible to dispute the second seat. Even though this does not seem to occur in the Chilean context[7], it will be consider as a robustness check in the results. Finally, denote as $X1$ and $X2$ the candidates that coalition X puts up. Then, there are two possible outcomes of this election:

$$A1 \text{ and } B1 \text{ are elected if } \#A < 2 \cdot \#B \tag{C.1}$$

$$A1 \text{ and } A2 \text{ are elected if } \#A \geq 2 \cdot \#B \tag{C.2}$$

The set of examples presented in Table 1 illustrate all possible election outcomes. In both cases 1 and 2 condition (C.1) is satisfied, so one legislator in each of the largest coalitions is elected. In case 2 coalition A "fails to double", since even though its two candidates obtain the first and second larger shares of votes, it obtains less than double the votes of coalition B. Therefore candidates A1 and B1 are elected, even though in Case 2 candidate A2 has the second-largest share of votes. Note also that coalition C will not have representation in the Congress, despite obtaining one-fifth of the votes. On the other hand, Cases 3 and 4 satisfy condition (C.2). Coalition A "doubles" coalition

[7]Only 6.94% of the elections have the number of votes of the second-largest coalition below 1.5 times the votes of the third-largest coalition, or when $\#B < 1.5 \cdot \#C$. This number reduces to 1.94% when $\#B < 1.1 \cdot \#C$.

B and gets candidates A1 and A2 elected. Case 4 is a peculiar one, since candidate B1 has a larger share of votes than candidate A2. All of these cases are of interest, since it is in the neighborhood of "doubling" that the discontinuity is produced.

As in the previous example, these twenty years of democracy have been characterized by two large and relatively stable coalitions: the center-left *Concertación de Partidos por la Democracia* —from now on, *Concertación*[8]—, and the center-right coalition, *Alianza*.[9] Among the smaller coalitions, the one grouping the Comunist party and other small parties is the one with the highest vote share. Since 1989 there have been 6 elections, and Case 1 has been the most frequent outcome with 73% of the total sample. Cases 2, 3 and 4 follow, representing 14.32%, 8.45% and 4.23% of the elections, respectively.[10]

The special features of this electoral system have made it worth studying by political scientists, but they have focused mainly on the system's innate bias towards the Right.[11] In fact, most studies describe the Chilean binominal system as engineered to favor the political allies of the dictatorship. Rabkin (1996) argues that the system, though controversial, has provided political stability by encouraging the creation of broad coalitions. But Rahat and Sznajder (1998) show, using data from the first two elections (1989 and 1993), that the system indeed favors the second largest coalition, *Alianza*. They also provide anecdotal evidence that the system was engineered: the electoral law was only completed after the 1988's Plebiscite[12] so that information about voters' behavior was available, and that districts borders were gerrymandered to artificially equalize representation.[13] Zucco (2007) argues that the binominal system did not only benefit the second-largest coalition, but also the largest one. But even though this could be true, he is not able to refute that the binominal system was engineered to favor the Right. Finally, Carey and Siavelis (2005) provide another consequence of having the binominal electoral rule: To get a chamber majority, coalitions have to put strong candidates in risky districts. Therefore, to keep the coalition united, the *Concertación* has developed the strategy of offering appointed post to unsuccessful congressional candidates as a way to insure them against a potential electoral loss. They find that being

[8]The *Concertación has also won the last four presidential elections.*

[9]The center-right coalition has changed its name several times, being *Alianza* the one used in the last legislative election.

[10]This numbers consider lower house elections only, though the numbers for the upper chamber are very similar.

[11]The only notable exception is Cox (1997) that looks at the incentives to behave strategically when only two coalitions are competing. This is not going to be the focus in this paper.

[12]This referendum was held to determine the continuity of Mr. Pinochet for another eight years in office. The options were "Yes" and "No". Option "No" won with a 55.99% of the votes.

[13]When looking at the 1988's Plebiscite outcome per *Cámara de Diputados*' district, Rahat and Sznajder, 1998 do not find a district were the "No" vote is 10% higher than the national average, while there were 11 districts featuring the opposite case.

in a senatorial race and getting a higher share of votes increases the probability of getting an appointment in government.

None of these studies has put too much emphasis on the strategic behavior of voters, namely that when faced with an election close to "doubling", extreme voters will vote for a candidate of one of the two largest coalition instead of their most preferred candidate.

3 Data and empirical strategy

3.1 Data and descriptive statistics

The data comes from the Chilean Electoral Service (Servicio Electoral) and the Election Qualifying Court (Tribunal Calificador de Elecciones), and is publicly available.[14] It contains the total number of votes per candidate in each of the 6 elections held since 1989. The data is available for both chambers, but the focus is on Lower House elections because, as mentioned before, individuals do not have polls to ascertain the closeness of the race. Table 2 shows the summary statistics.

Election closeness is defined in this context as the ratio of votes for the largest coalition over the votes of the second-largest, and has an average of 1.513. This number is one standard deviation below 2, the ratio needed to have a double (both candidates elected from the same coalition). This explains why there are so few doubles in the sample (12%, and only 7% in the Upper House). On the other hand, the fact that there are only a few doubles implies that the distribution of seats in the Lower House is generally 50% center-left and 50% center-right, and therefore when a voter is pivotal in a given district, it is very likely that she is pivotal at the national level as well.

The average number of candidates is 6.66, and this means that on average there are 2.66 candidates coming from small coalitions, since the 2 largest coalitions always put up 2 candidates each, with only one exception in an Upper House district. The average number of coalitions per election is 3.8, but there are some Lower House districts with up to 7 coalitions in a given year (note that each independent candidate that does not join a coalition is considered as a different coalition). Finally, the share of votes going to small coalitions is 9.5% and 8.8% in the Lower and Upper Houses, respectively. Even though these numbers seem small, there is plenty of variation both within and between election years: The standard deviation in Lower Chamber elections is 7%, with a maximum support for small coalitions of 35%. On the other hand, both 1989 and 2001 elections denote the extremes regarding support for small parties, with 13.2% and 6.9%, respectively.

[14] www.servel.cl and www.tricel.cl, respectively.

3.2 Testing for strategic voting

In this section I formally test for strategic behavior. The hypothesis is that when the previous election was close to a *doubling* situation, the incentive for small-coalition supporters to vote for one of the two main coalitions should be larger, and therefore a decrease in the support for small coalitions should be observed. Figure 2 presents a scatterplot of the average share of votes of small coalitions in election $t+1$ as a function of the ratio of votes of the two largest coalitions in election t, for Lower House elections. The solid vertical line marks the threshold for having a doubling, i.e. when the share of votes of the largest to the second-largest coalition is equal to 2. For observations located to the left of this line, one candidate from the largest coalition and one candidate from the second largest are elected. For observations located to the right of the threshold the two candidates from the largest coalition are elected.

Figure 2 also shows two sets of non-parametric fits computed at each side of the threshold, a local linear polynomial and a locally weighted scatterplot smoothing estimator (lowess). The estimates show a reduction in the share of votes of small coalitions both for the doubling and the no-doubling case, providing evidence that some supporters of small coalitions do behave strategically, at least for elections close to the doubling threshold. Figures 3 and 4 show the same scatterplot but using a different bin size to compute the average of the dependent variable: 0.05 and 0.1 (the bin size is 0.025 in Figure 2). The observed pattern is the same, i.e. a decrease in the support for small coalitions when close to the doubling threshold, but the magnitude seems to be smaller in figures 3 and 4.

The non-parametric estimates in the previous figures do not take into account other controls, as well as interactions between the controls and the variables of interest. In order to address these issues I therefore adopt the following econometric model:

$$
\begin{aligned}
v^3_{i,t+1} = {} & \beta_0 + \sum_{k=1}^{4} \beta_k \, (distance_{i,t})^k + \beta_5 \, around(\theta)_{i,t} + \beta_6 \, double_{i,t} + \\
& \beta_7 \, around(\theta)_{i,t} * distance_{i,t} + \beta_8 \, distance_{i,t} * double_{i,t} + \beta_9 \, around(\theta)_{i,t} * double_{i,t} + \\
& \beta_{10} \, around(\theta)_{i,t} * distance_{i,t} * double_{i,t} + \beta_{11} X_{i,t} + \xi_t + \mu_i + \epsilon_{i,t+1} \qquad (1)
\end{aligned}
$$

where

$$distance_{i,t} = \mid \frac{v^1_{i,t}}{v^2_{i,t}} - 2 \mid \tag{2}$$

$$around(\theta)_{i,t} = \mathit{1} * [distance_{i,t} \leq \theta] \, , \theta = \{.1, .2, .3\} \tag{3}$$

$$double_{i,t} = \mathit{1} * \left[\frac{v^1_{i,t}}{v^2_{i,t}} \geq 2 \right] \tag{4}$$

$v^j_{i,t}$ indicates the vote share of the j-th largest coalition in discrict i in election t, $j = \{1, 2, 3\}$. In a slight abuse of notation, $v^3_{i,t}$ also includes the vote share of the third and smaller coalitions, as well as the vote of independent candidates. $distance_{i,t}$ measures the distance to the doubling threshold, re-centered at 2. It enters in the regression as a fourth degree polynomial, to allow for a flexible interaction with the dependent variable, as the one seen in figure 1. $around(\theta)_{i,t+1}$ is a dummy variable that takes the value of 1 when the ratio of votes of the two largest coalitions is in a neighborhood of size θ of the doubling threshold. I try with three different values for θ (.1, .2 and .3). $double_{i,t}$ is a dummy variable that takes the value of 1 when a double has occured. Finally $X_{i,t}$ is a vector of variables that affect voters' choices in the election at year t (number of candidates and the lagged share of votes for small coalitions).

There is no clear prediction about the sign of β_1 to β_4, since there should not be any strategic incentive for small-coalition supporters outside the neighborhood of the doubling threshold (see figure 2). If there is strategic voting inside the doubling neighborhood, β_5 should be negative, and β_7 should be positive, since the effect of strategic voting on small coalitions should go down when going further away from the doubling threshold. The strategic incentive might be asymmetric, and therefore the interactions with *double* are of interest as well.

4 Results

4.1 OLS estimates

Table 3 shows the OLS estimates for equation (1) considering different values of θ, but not considering fixed effects. Odd columns show the coefficients for *around* and its interaction with *distance*, while even columns include the other interactions. For $\theta = .1$ both coefficients on *around* and $around * distance$ have the expected signs and are statistically significant. The coefficient on $around * double$ is very small and not significant, suggesting that the strategic effect is symmetric in the neighborhood of the doubling threshold. In addition, the small F test (p-value of 0.082) of *around* and the other interactions in column (2), indicates that these variables are barely significant when taken together. Therefore in the following specifications only *around* and its interaction with *distance* are included.

When $\theta = .2$ the coefficients on *around* and *around∗distance* preserve the signs, but the magnitudes are smaller and significance is lost. In fact, when $\theta = .3$ the coefficient show the opposite signs. This suggests that the strategic effect is quite narrow: the share of votes for small coalitions is reduced only when the previous election was very close to doubling, below 20% of the threshold. Therefore in all the following specifications $\theta = .1$ will be used.

What does this mean in terms of the closeness of the election? For $\theta = .1$ it means that the largest coalition wins the second seat to the second-largest coalition by at most 61.30% to 29.19%, or it loses the seat by at most 59.28% to 31.20%.[15]

How large is the strategic effect? According to the results in Column (1), being at the threshold reduces the share of votes for small coalitions in 10.69%. Since the share of votes for small coalitions is 8.33% around the threshold, this implies that 0.89% of the electorate shifts their support from smaller coalitions to one of the two largest ones. Nonetheless, it is not straightforward to compare this result with the findings in previous literature, because this estimation is aimed at testing for one possible strategic behavior, the one of supporters of small coalitions. In addition, this reduced-form approach is only able to capture misaligned voting, which is a subset of strategic voting. Kawai and Watanabe (2010) find that between 2.4% and 5.5% of voters in their sample engage in misaligned voting, which seems to be consistent with my findings, given that I only look at small coalition supporters.

4.2 Fixed effects estimates

Table 4 shows the estimates when time and districts fixed effects are included. In Column (1) I include a set of Upper House district dummies to control for unobserved Upper House district characteristics. As mentioned before, there are 19 Upper House districts that group between 1 and 8 Lower House districts. Even though the number of parameters estimated increases considerably, the coefficient on *around* is still negative and very precise. The coefficient on *around* ∗ *distance* stays positive but not significant. Including a set of year dummies improves the results, shown in Column (2). *around* is even more precisely estimated and *around* ∗ *distance* is still non-significant but with a p-value of 0.16. This suggests that the year dummies play an important role, and even much so when a lower house district fixed effects regression is performed in columns (3) and (4). Column (3) does not include year dummies, and *around* loses significance and

[15] This calculation is done as follows: Let v^i be the vote share of the i-th coalition (with v^3 including third and smaller coalitions). From Table 1 take the average figure for v^3 in Lower House elections (9.516%). Therefore with $v^1 + v^2 = 90.484\%$ and $\frac{v^1}{v^2} \in [1.9, 2.1]$ (given by the 10% bandwidth) I can compute the extreme outcomes for v^1 and v^2.

the interaction with *distance* changes sign. In column (4) year dummies are included again, and *around* is again significant at the 5%. $around * distance$ has the expected sign, but its estimate is very imprecise. Overall Columns (2) and (4) provide strong support for strategic behavior, even after controlling for all possible heterogeneity accross districts.

To further test the robustness of the results presented above, I ran similar regressions to column (2) in Table 4, but using different values for the threshold, ranging from 1.5 to 3.0. It would be hard to justify the previous results as evidence of strategic behavior if the coeficient on *around* or its interaction with *distance* have the correct signs and are significant around other thresholds. The results are in Appendix A, and show that this is not the case.

4.3 Including lagged vote share of small coalitions and number of candidates as covariates

To analyze persistence in the share of votes for small coalitions the lagged share of votes for small coalitions ($v^3_{i,t}$) is included as a covariate in Column (1) of Table 5. Nonetheless, this variable is correlated with the unobserved fixed effect characteristics, leading an inconsist estimate. The solution is to use the GMM estimator developed by Arellano and Bond (1991). Column (2) shows these estimates, but in this case the Sargan test rejects the null hyphotesis that the overidentifying restrictions are valid. Furthermore there is evidence of second order serial correlation in the first differences, invalidating the use of lagged variables as instruments. In Column (3) an additional covariate, the number of candidates in $t+1$, is included. The coefficient on this regressor is positive and significant, i.e. more candidates for small coalitions increase their share of votes.[16] Nonetheless it is very likely that parties in small coalitions are also acting strategically: they decide whether to put up candidates by taking into account the previous share of votes for small candidates, and therefore the number of candidates is an endogenous variable. The estimates presented in Column (4) address this problem by specifying the number of candidates as an endogenous regressor, and instrumenting it with its lagged variable. Here both the lagged share of votes for small coalitions and the number of candidates seem to have a positive effect on the dependent variable, though statistically insignificant. The two regressors of interest, *around* and $around*distance$, have the expected signs (*around* is significant at the 10%). Finally, there is no evidence of second order autocorrelation and that the overidentifying restrictions are not valid.

[16] Including the total number of candidates is almost equivalent to include the number of small-coalition candidates, since the two largest coalitions almost always put up two candidates each.

4.4 Left-wing and right-wing small coalitions

The share of votes for small coalitions, $v_{i,t}^3$, includes both left and right wing small coalitions. Therefore it is possible that if the election closeness affects supporters of left and right coalitions differently, then both effects could be netted out. To investigate this possibility I ran the model in Column (4) of Table 4, but now distinguishing between the vote share of left and right small coalitions. I also split the sample according to the leading coalition, i.e. whether the coalition with the largest vote share is the center-left (Concertación) or center-right (Alianza). Table 6 shows the results. The first three columns have the vote share of left-wing small coalitions as the dependent variable. In column (1) I consider those cases where the leading coalition ($v_{i,t}^1$) is the center-left and when the leading coalition is the center-right. Here the results do not show evidence of strategic voting. Columns (2) and (3) consider those cases where the leading coalition is the center-left and center-right, respectively. Again here both coefficients of interest are not significant and most do not have the expected signs.

The rest of table 6 looks at the determinants of the vote share of right-wing small coalitions. When there is no distinction regarding the leading coalition (column (4)), the coefficient on *around* is negative and significant, while the coefficient on its interaction with *distance* is positive but very imprecise. Nonetheless, both are jointly significant at the 5%. This strategic behavior comes mainly from those cases where the center-left is the leading coalition (column (5)), where the coefficient on *around* is even larger. There seems to be no strategic behavior when the leading coalition is the center-right, as shown in column 6. Therefore, it is possible to conclude that supporters of right small coalitions, when faced with a center-left coalition that it close to the doubling threshold, decide to cast their vote for the center-right instead of giving their support to their sincere preference.

As the reader may have noticed, the coefficients for *around* in columns (3) and (6) are extremely large, not even comparable to the previous results. The reason for this lies behind the small number of observations considered in these specifications, and the very low predictive power of the regressions.[17] Therefore although informative, these results cannot be taken as evidence of strategic voting.

[17] A standard procedure to measure the quality of the fit is to run a regression of the predicted vote shares for small coalitions on the actual vote share and a constant. If the fit is good, the coefficient on the actual vote share should be close to one, and the one on the constant close to zero. For the case of the specification in column (3), this regression gives a coefficient on the intercept of 6.15 and 0.014 on the slope, with an R^2 of 0.0001. Compare these values with the ones obtained when specification (2) of Table 4: 5.79 on the intercept, 0.339 on the slope, and an R^2 of 0.339. Even though the fit in the case of the latter is far from perfect, its performance is more than acceptable.

5 Conclusions

The Chilean Congressional elections and the features of its binominal system constitute an excellent setup to test for strategic voting. In particular, the way in which seats are allocated in each district and the lack of pre-election polls for the Lower Chamber allow me to look at the relationship between the share of votes going to small coalitions and the outcome of the previous election as an appropiate test for strategic voting. I perform this test in a semi-parametric framework and find that having the last election at the doubling threshold decreases the vote share for small coalitions in between 6.58% and 10.69% in the following election. The results are robust to including a set of year dummies and a set of Upper House district dummies, though when Lower House district dummies are included the coefficient on $around * distance$ becomes non significanct. The results are also robust to including the lagged share of votes for small coalitions and the number of candidates in each election. Since the latter is endogenous to the closeness of the election, the model is estimated using the Arellano and Bond (1991) estimator. Finally, I distinguish between left and right wing coalitions, and find support for a one-sided strategic behavior in the case of supporters of small right-wing coalitions, who behave strategically when the center-left is close to the doubling threshold.

Why don't supporters of small left-wing coalitions behave strategically? One possibility is that these are *die-hard* voters that will not "sell" their vote to get more seats to the center-left, even if this in turn favors the center-right. Another possibility is that small coalitions are able to negotiate the center-left support for some of their preferred policies, in return for having weak candidates in all districts (or maybe a couple of strong candidates in a few districts). This might explain why the strategic behavior is smaller when fixed effects are included (and therefore only the within-variation is exploited), in addition to the non-existent strategic behavior of supporters of small left-wing coalitions.

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Figure 1: Ballot for district 34 in the 2005 lower house election. The figure was generously provided by the Chilean Electoral Service.

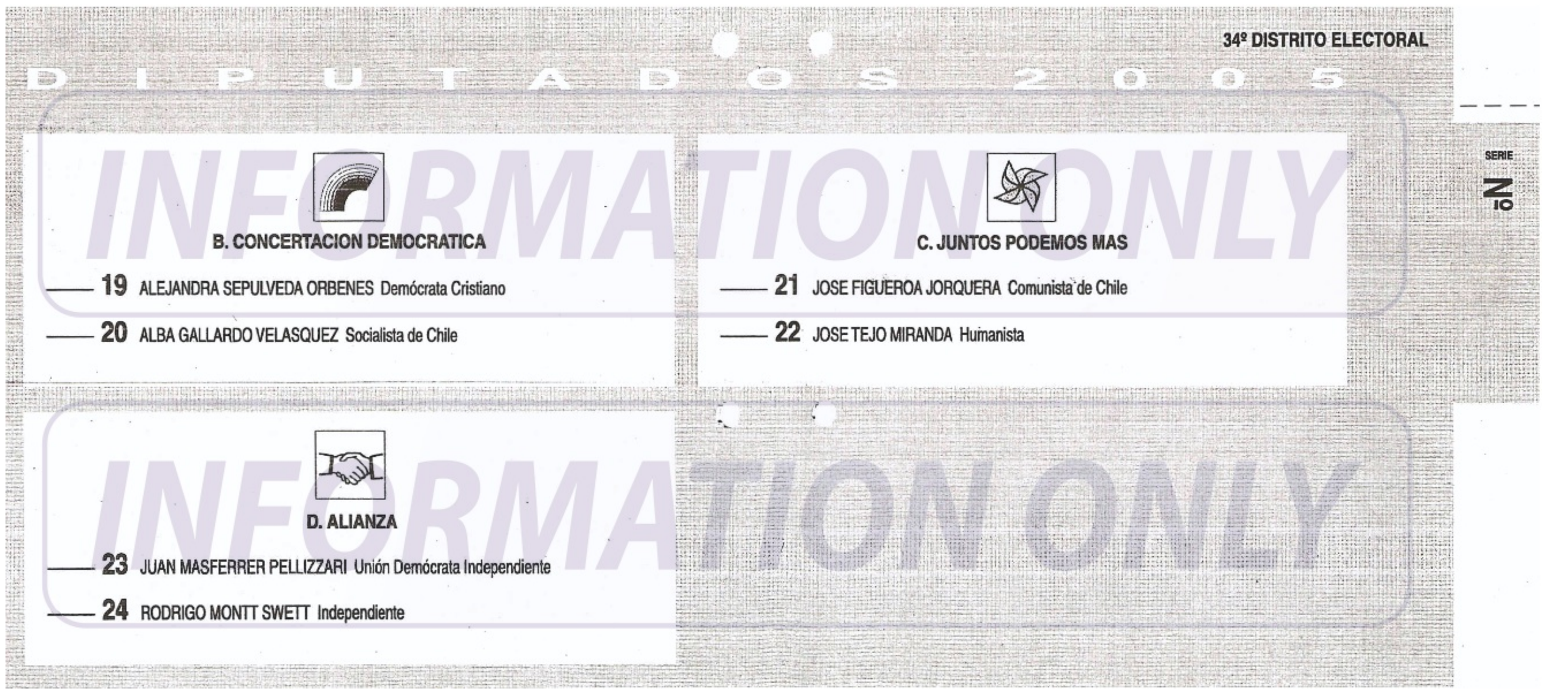

Figure 2: Share of votes of small coalitions in election $t + 1$, by likelihood of doubling in election t. (Bin size = 0.025)

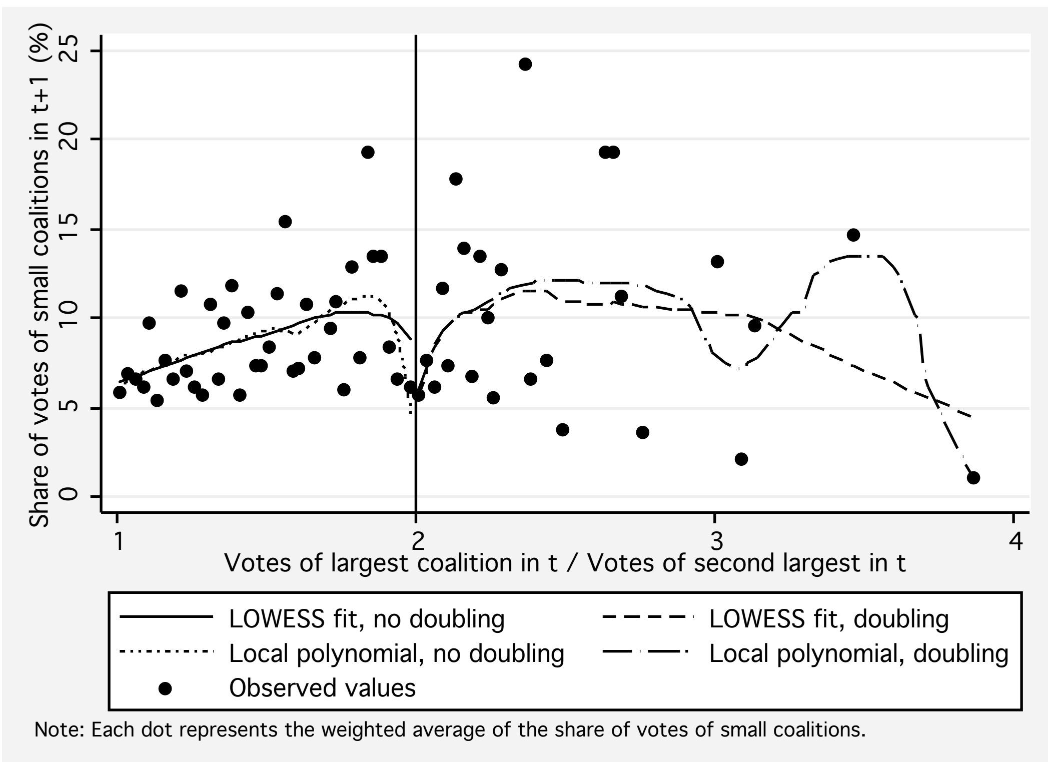

Figure 3: Share of votes of small coalitions in election $t+1$, by likelihood of doubling in election t. (Bin size = 0.05)

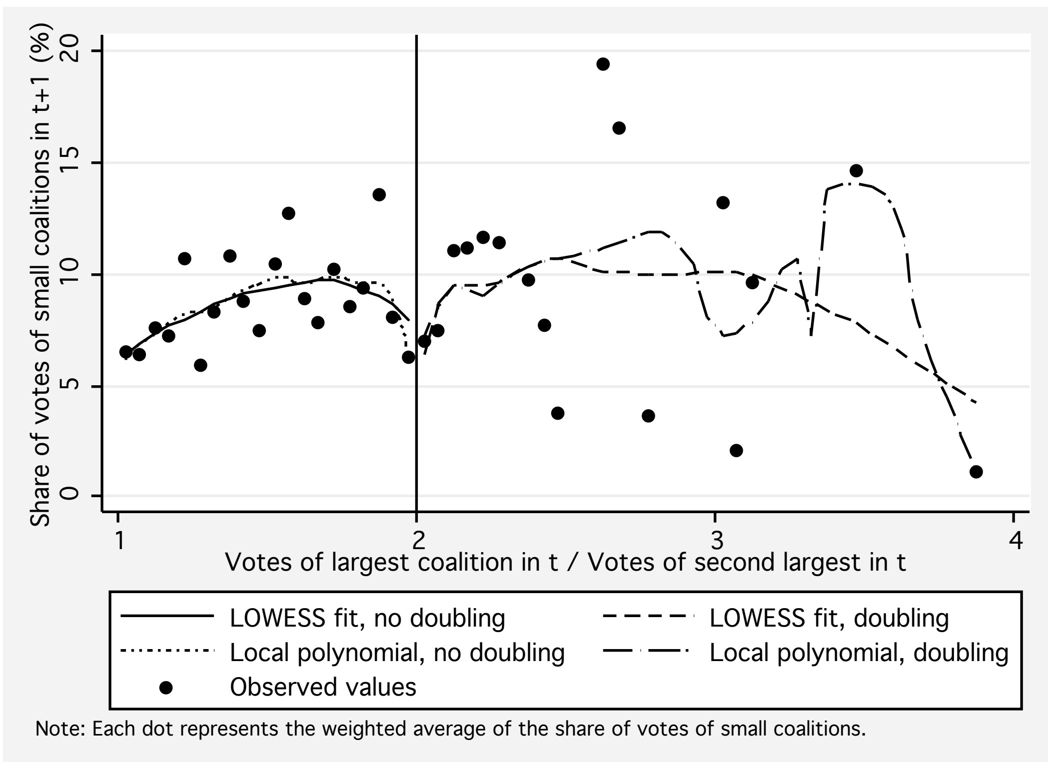

Figure 4: Share of votes of small coalitions in election $t+1$, by likelihood of doubling in election t. (Bin size = 0.1)

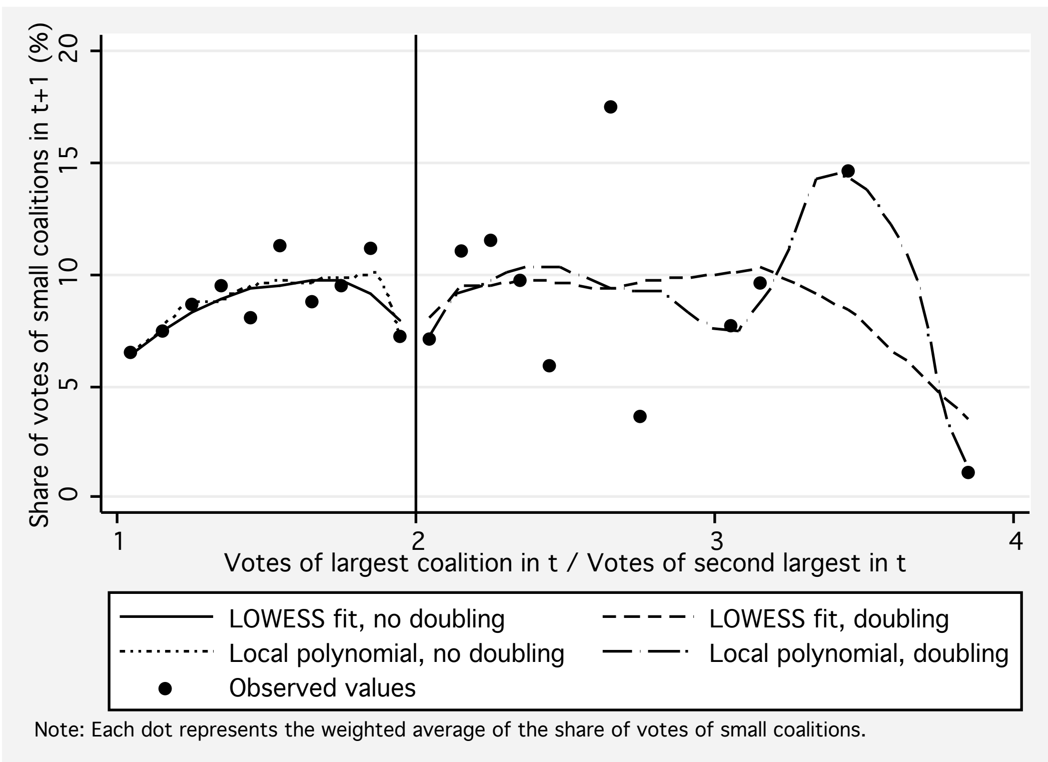

Table 1: Four examples of election outcomes.

	Case 1	Case 2	Case 3	Case 4
Coalition A	40%	50%	60%	70%
Candidate A1	**30%**	**30%**	**35%**	**60%**
Candidate A2	10%	20%	**25%**	**10%**
Coalition B	40%	30%	30%	20%
Candidate B1	**22%**	**18%**	18%	18%
Candidate B2	18%	12%	12%	2%
Coalition C	20%	20%	10%	10%
Candidate C1	11%	11%	6%	6%
Candidate C2	9%	9%	4%	4%

Note: Vote shares of elected candidates are in boldface.

Table 2: Chilean congressional elections, summary statistics.

	N	Mean	St. Dev.	Min	Max
Election Closeness	426	1.513	0.493	1.002	5.226
Lower House	360	1.518	0.517	1.003	5.226
Upper House	66	1.486	0.336	1.002	2.430
Share of Doubles	426	0.120	0.325	0	1
Lower House	360	0.128	0.334	0	1
Upper House	66	0.076	0.267	0	1
No. Candidates	426	6.660	1.310	3	11
Lower House	360	6.781	1.288	4	11
Upper House	66	6.000	1.240	3	9
No. Coalitions	426	3.775	0.768	2	7
Lower House	360	3.831	0.766	2	7
Upper House	66	3.470	0.707	2	5
Share of votes for small coalitions (%)	426	9.402	7.055	0	35.11
Lower House	360	9.516	6.933	0	35.11
Upper House	66	8.783	7.714	0	29.58
1989 (Lower House)	60	13.233	8.537	0	35.11
1993 (Lower House)	60	7.158	4.388	0	22.04
1997 (Lower House)	60	10.411	6.147	2.72	28.85
2001 (Lower House)	60	6.857	5.235	0	27.06
2005 (Lower House)	60	8.978	5.874	2.58	28.75
2009 (Lower House)	60	10.460	8.471	0	32.16

Note: Election Closeness is defined as the ratio of votes for the largest coalition over the votes of the second-largest.

Table 3: Testing for strategic voting, results without fixed effects.

Dependent variable: % of votes for small coalitions, $t+1$						
	$\theta = .1$		$\theta = .2$		$\theta = .3$	
	(1)	(2)	(3)	(4)	(5)	(6)
$around(\theta)$	-10.69***	-10.68**	-7.798	-6.138	10.19	11.63
	(3.799)	(4.506)	(4.930)	(5.947)	(9.027)	(9.219)
$around(\theta) * distance$	52.20*	57.78*	53.97**	42.34	-22.15	-29.34
	(26.70)	(34.72)	(26.25)	(34.28)	(26.90)	(28.24)
$around(\theta) * double$		-0.0561		-3.229		-4.063
		(3.604)		(4.439)		(6.873)
$around(\theta) * distance * double$		-11.02		22.93		17.08
		(42.71)		(38.05)		(16.73)
$distance * double$		4.091		3.215		1.931
		(5.168)		(5.216)		(7.699)
$distance$	-25.52	-33.55	-12.70	-19.75	61.49	54.29
	(19.48)	(22.80)	(23.76)	(26.31)	(39.97)	(39.56)
$distance^2$	39.15	56.82	18.08	33.71	-91.07	-76.30
	(31.98)	(39.54)	(37.01)	(42.70)	(57.84)	(58.07)
$distance^3$	-27.65	-42.23	-14.35	-27.21	48.99	37.45
	(20.72)	(27.23)	(22.86)	(27.81)	(33.93)	(35.08)
$distance^4$	6.267	9.596	3.498	6.445	-8.912	-6.297
	(4.515)	(5.910)	(4.809)	(5.854)	(6.883)	(7.140)
$double$	2.697*	1.234	2.637*	1.739	2.447*	2.466
	(1.412)	(2.772)	(1.430)	(3.050)	(1.446)	(6.195)
Constant	14.77***	15.96***	12.32**	13.32**	-4.343	-3.238
	(3.837)	(4.388)	(4.929)	(5.389)	(9.246)	(9.146)
Observations	300	300	300	300	300	300
Obs. where $around(\theta) = 1$	30	30	70	70	100	100
F test	15.21	5.657	11.23	10.70	3.511	2.911
F test for *around* and interactions	3.994**	2.093*	2.129	1.438	1.633	1.181
adjusted R^2	0.0659	0.0606	0.0507	0.0470	0.0427	0.0401

Notes: Robust standard errors in parentheses. (*** p<0.01, ** p<0.05, * p<0.1)

Table 4: Testing for strategic voting: Including fixed effects.

Dependent variable: Share of votes for small coalitions, $t+1$				
	(1)	(2)	(3)	(4)
$around(.1)$	-8.887**	-9.109***	-6.351	-6.577*
	(3.632)	(3.297)	(4.037)	(3.566)
$around(.1) * distance$	32.05	43.14	-8.403	4.231
	(33.84)	(30.93)	(42.02)	(38.57)
$distance$	-20.50	-19.89	-22.39	-21.62
	(16.93)	(15.83)	(17.74)	(16.02)
$distance^2$	31.54	27.35	36.77	32.16
	(28.27)	(26.41)	(30.17)	(27.35)
$distance^3$	-22.31	-18.62	-25.05	-21.06
	(18.81)	(17.26)	(20.37)	(18.27)
$distance^4$	5.009	4.211	5.311	4.462
	(4.184)	(3.764)	(4.574)	(4.057)
$double$	1.944	1.812	1.430	1.418
	(1.361)	(1.263)	(1.455)	(1.355)
Constant	22.42***	20.79***	13.47***	11.99***
	(4.261)	(3.982)	(3.385)	(3.028)
Upper House districts dummies	Yes	Yes	No	No
Year dummies	No	Yes	No	Yes
Lower House districts fixed effects	No	No	Yes	Yes
Observations	300	300	300	300
F test	5.727	5.420	5.075	5.414
F test for around and interactions	3.431**	4.002**	2.686*	2.838*
adjusted R^2	0.212	0.267	0.035	0.113

Notes: Robust standard errors in parentheses, adjusted for clustering on Lower House districts in (3) and (4). (*** p<0.01, ** p<0.05, * p<0.1)

Table 5: Testing for strategic voting: Adding lagged share of votes.

Dependent variable: Share of votes for small coalitions, $t+1$	(1) Fixed Effects	(2) Arellano-Bond	(3) Arellano-Bond	(4) Arellano-Bond[a]
$around(.1)$	-6.825*	-4.906	-3.834	-7.674*
	(3.780)	(3.710)	(3.844)	(3.919)
$around(.1) * distance$	6.958	24.81	48.12	46.78
	(40.13)	(30.22)	(35.53)	(37.57)
Lagged share of votes for small coalitions	0.0273	-0.173	0.0461	0.114
	(0.0752)	(0.108)	(0.0570)	(0.0721)
Number of candidates			2.003***	0.654
			(0.309)	(0.692)
Constant	14.38***	11.79***	-5.121	6.277
	(3.111)	(4.122)	(3.864)	(5.952)
Observations	300	240	240	240
Adjusted R^2	0.1106			
p-value from Sargan test		0.0011	0.2909	0.2090
p-value from test for 1st order autocorrelation		0.0233	0.0001	0.0004
p-value from test for 2nd order autocorrelation		0.0580	0.3995	0.9496

Notes: Robust standard errors in parentheses, adjusted for clustering on Lower House districts in (1) and computed using the WC-robust estimator of Windmeijer in (2), (3) and (4). The Sargan test was computed without using the robust estimator. All regressions include year dummies and lower house districts fixed effects, as well as a fourth-order polynomial on *distance*. (*** p<0.01, ** p<0.05, * p<0.1)

[a] The lagged number of candidates was specified as endogenous.

Table 6: Looking at left and right small coalitions.

Dependent variable: Share of votes for small coalitions, election $t+1$						
	Left small coalitions			Right small coalitions		
	(1)	(2)	(3)	(4)	(5)	(6)
Largest coalition	Either	Center Left	Center Right	Either	Center Left	Center Right
around(.1)	1.502	-0.703	258.3	-5.801**	-6.594**	113.0
	(2.929)	(3.169)	(170.6)	(2.799)	(3.162)	(221.3)
around(.1) × *distance*	-27.39	-21.05	-1421	2.222	11.61	-698.8
	(29.00)	(28.95)	(889.3)	(27.66)	(29.70)	(1140)
distance	6.723	-8.334	1624	-24.27*	-26.28	705.4
	(13.70)	(14.79)	(1110)	(14.11)	(16.25)	(1441)
$distance^2$	-13.64	8.908	-3666	40.86*	45.08	-1626
	(23.23)	(25.48)	(2503)	(23.94)	(27.99)	(3233)
$distance^3$	9.154	-3.582	3545	-27.50*	-29.89	1587
	(15.32)	(16.59)	(2411)	(15.76)	(18.26)	(3091)
$distance^4$	-2.171	0.240	-1248	6.075*	6.498*	-559.3
	(3.349)	(3.570)	(843.7)	(3.418)	(3.909)	(1071)
double	1.098	1.403	-27.28	1.147	1.017	-9.445
	(1.075)	(1.138)	(20.08)	(1.171)	(1.301)	(25.79)
Constant	5.112*	7.445**	-247.8	4.606*	6.699**	-106.3
	(2.609)	(2.895)	(173.6)	(2.692)	(2.848)	(225.6)
Observations	300	225	74	300	225	74
F test	3.789***	3.366***	913.8***	2.018**	1.675*	0.453
F test for *around* and interactions	0.4702	0.8719	1.6362	3.0379**	2.6441*	1.1178
# districts	60	57	35	60	57	35
adjusted R^2	0.0167	0.0639	0.0699	0.0521	0.0488	0.103

Notes: Robust standard errors in parentheses, adjusted for clustering on lower house districts. All regressions include year dummies and lower house districts fixed effects. (*** p<0.01, ** p<0.05, * p<0.1)

APPENDIX

A Robustness check: Changing the threshold.

The value of the threshold is represented by the parameter η, and therefore $around_\eta(.1)$ is a dummy variable that takes the value of 1 *close* is in the 10%-neighborhood of η. Column (3) in this table is the exact same regression analyzed in Column (2) of Table 4, and it is the only one that has the correct signs and significant coefficients. Column (5) has the correct signs but non-significant coefficients. Furthermore, the the F-test for *around* and $around * distance$ reject joint significance (p-value of 0.17).

Table 7: Robustness check: Changing the threshold.

	Dependent variable: Share of votes for small coalitions, election $t+1$						
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
	$\eta = 1.5$	$\eta = 1.75$	$\eta = 2$	$\eta = 2.25$	$\eta = 2.5$	$\eta = 2.75$	$\eta = 3$
$around_\eta(.1)$	6.701	8.762	-9.109***	6.482	-34.39	129.1**	105.1
	(7.197)	(6.611)	(3.297)	(7.605)	(44.01)	(53.88)	(148.7)
$around_\eta(.1)\times$	-13.86	-34.65	43.14	-17.89	60.82	-182.0**	-98.56
distance	(13.87)	(25.44)	(30.93)	(31.27)	(95.01)	(73.74)	(141.4)
distance	16.57*	17.58*	-19.89	14.02	18.11	15.41*	16.27
	(10.03)	(9.225)	(15.83)	(9.512)	(11.76)	(9.167)	(11.76)
$distance^2$	-34.02*	-35.27*	27.35	-29.53	-36.52	-34.50*	-34.57
	(19.82)	(18.96)	(26.41)	(18.96)	(26.83)	(18.94)	(27.28)
$distance^3$	20.38	21.17	-18.62	18.00	21.97	22.26	21.46
	(15.11)	(15.02)	(17.26)	(14.71)	(23.36)	(15.09)	(23.92)
$distance^4$	-3.922	-4.100	4.211	-3.501	-4.261	-4.551	-4.261
	(3.652)	(3.688)	(3.764)	(3.592)	(6.118)	(3.723)	(6.284)
double	1.745	1.898	1.812	1.002	2.114	1.304	1.670
	(1.284)	(1.327)	(1.263)	(1.512)	(1.438)	(1.307)	(1.408)
Constant	14.67***	14.11***	20.79***	15.03***	17.93***	15.10***	18.44***
	(3.542)	(3.491)	(3.982)	(3.493)	(3.808)	(3.397)	(3.757)
Observations	300	300	300	300	300	300	300
Obs. *around*=1	55	30	30	13	2	4	2
F test	4.875	4.830	5.420	4.886	4.168	5.030	4.161
F test for *around*							
and interactions	0.5966	0.9292	4.0018	0.7603	1.7799	4.8290	0.2751
adjused R^2	0.245	0.248	0.267	0.248	0.249	0.263	0.245

Notes: Robust standard errors in parentheses in (1), (2), (3), (4), and (6). Standard errors in (5) and (7) where computed using the jackknife procedure. All regressions include year and upper house dummies. The parameter η indicates the location of the threshold ($\eta = 2$ in all previous tables).
(*** p<0.01, ** p<0.05, * p<0.1)