Transaction costs in electoral coordination: How turnout shapes changes in the number of parties*

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ABSTRACT

This article examines the forces shaping changes in the number of parties between consecutive elections. We argue that the transaction costs in electoral coordination depend on the turnout level in the previous election. The greater the number of peripheral voters entering the electorate, the less likely a substantial change in the distribution of partisan support in the subsequent election. The argument is tested using data for 313 parliamentary elections in 63 countries from 1990 to 2011, and two cases studies of countries using compulsory voting (the Netherlands and Australia).

1. Introduction

Variation in the number of parties across districts or countries has attracted considerable attention in political science. Cross-sectional variation in party system fragmentation is explained by (a weak version of) the interactive hypothesis: electoral rules interact with social diversity to determine the number of parties competing in elections (Milazzo et al., 2018: 939). The number of parties has been shown to drop after several elections using the same set of rules—four according to evidence from Crisp et al. (2012). Learning progressively reduces inexperience-induced coordination failures, so votes wasted on hopeless parties and candidates decrease. As electoral systems and social diversity rarely change, the expectation is that the number of parties will remain stable across elections in established democracies.

Surprisingly, the number of parties substantially changes between consecutive elections. In a sample of 329 elections in 66 countries assembled by Lublin (2017), the correlation between the effective number of electoral parties in the election held at time \( t \) and the election held at time \( t-1 \) in a given country is only 0.86. Interestingly, there are substantial differences in the stability of party system fragmentation across established democracies. For instance, the effective number of electoral parties in national elections was much more variable in Israel (variation coefficient of 25.09 percent from 1992 to 2009) than in the Netherlands (variation coefficient of 12.25 percent from 1994 to 2010) when both countries had employed a single national district for decades.

This article sheds light on the conditions under which electoral coordination on a set of parties over repeated elections remains (un)stable, a crucial dimension of party system change (Sartori, 2005; Mair, 1997). A large body of research has examined the emergence and success of new parties in recent years (Bolleyer, 2013), with new party entry associated with increases in turnout (see Heath and Ziegfeld, 2018). However, while party votes substantially change between consecutive elections, new party entry is quite exceptional. In a sample of 324 national elections in 19 West European countries from 1946 to 2015 assembled by Emanuele and Chiaramonte (2018: 478), in 184 elections there had not been any new party reaching 1 percent of the votes; in 107 elections new parties got between 1 and 5 percent of votes; and in 33 elections new parties obtained more than 5 percent of the votes.

We argue that the number and type of voters in the election influences how stable the distribution of partisan support will be in future elections. The transaction costs for a change in the number of parties (i. e., the costs of researching, negotiating, and publicizing the emergence of new parties and the costs if they fail (Cox, 1997: 252)) increase with turnout level and, in particular, with the number of peripheral voters entering the electorate in the election held at \( t-1 \). Core voters consistently vote in every election and have strong partisan attachments that cause them to vote for the same party across elections, whereas peripheral voters have weaker partisan leanings and are more likely to support new entrants or simply switch their vote. Only core voters go to the polls when turnout is low. As turnout is likely to increase in future elections due to the mobilization of peripheral voters, the probability of a substantial change in party votes increases. By contrast, fragmentation

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generated in high-turnout elections is less likely to change, given that it already captures the preferences of peripheral voters. Additionally, political elites might sense an opportunity to enter the party system when the turnout level in previous elections is low. We call this the stability hypothesis.

This article contributes to a better understanding of the conditions under which the number of parties is more prone to change. Our research has significant political and policy implications. First, our findings allow us to increase the predictability of the future behavior of elites and voters, a crucial parameter for foreign investors. For example, unstable party systems are less attractive to foreign direct investment inflows due to the higher expropriation risks (Bellinger and Son, 2019). Second, the stability of party systems crucially affects democratic politics (Mainwaring, 2018).

Our theoretical expectations are tested using a large comparative analysis of 313 parliamentary elections in 63 countries from 1990 to 2011. We then provide evidence from an in-depth analysis of two countries (the Netherlands and Australia) using compulsory voting under different electoral systems.

2. Arguments

Party system change encompasses three different dimensions: fragmentation (i.e., the number of relevant parties), polarization (i.e., the ideological distance covered by these parties), and power alternation (i.e., the degree of closure of the governmental arena) (Casal Bétoa and Enyedi, 2016; Nwokora and Pelizzo, 2018).

The conventional wisdom in political science suggests that the number of parties in democracies is a function of the number of cleavages, the permissiveness of electoral systems, and the number of elections held under the same electoral rules. According to the strong version of the interactive argument (Cox, 1997), cleavages create demand for parties. Social heterogeneity increases the number of parties when there is a permissive electoral system (i.e., proportional representation [PR] with high district magnitudes). A weak version of the interactive argument proposes that “even under first-past-the-post (FPTP) electoral rules, there is a relationship between social diversity and the number of parties, thus suggesting that restrictive rules are not as powerful a constraint on electoral behavior and outcomes as is usually supposed” (Milazzo et al., 2018: 962). Since coordination is impeded during early elections under new rules, the number of parties remains unstable in the first years of electoral democracies (Tavits and Annus, 2006).

According to estimates provided by Crisp et al. in 21 countries (2012: 153) and the evidence shown by Lago and Lago (2000) in a new democracy, Spain, where the preferences of voters were largely unknown, the first couple of elections held under a new set of electoral rules show strong evidence of pronounced coordination issues that nevertheless subside quickly, until, by the fifth election, coordination becomes indistinguishable from that in more experienced settings.

Electoral systems and social diversity are largely constant, however, so the interactive hypothesis cannot account for short-term changes in the distribution of partisan support in established democracies. Filling this gap, Chhibber and Kollmann (1998) show that the degree of political and economic (de)centralization affects the number of parties. In particular, “as national governments exert more political or economic control over local areas, candidates have greater incentives to associate themselves with national organizations, and voters have greater incentives to abandon locally competitive but nationally noncompetitive parties” (Chhibber and Kollmann, 1998: 329). Thus, under increasing centralization, there will be fewer parties, although it is not clear how many years or elections are necessary to observe this effect (see also Lublin, 2012).

Change in the number of parties has also been addressed by realignment theory (Mayhew, 2000; Sundquist, 1983) and theories of new party entry and volatility (Bolleyer, 2013; Lago and Torcal, 2020; Tavits, 2008). Although both approaches rightly argue that not all party systems are equally stable, they do not indicate the conditions under which the number of parties changes. On the one hand, realignment theory is a macro-level description of voting patterns with unclear micro-level foundations. It does not present an explicit account based on individual behavior for why realignment occurs and when existing equilibria are challenged. As Mayhew (2000: 471) summarizes, “the claims of the realignments genre do not hold up well, and the genre’s illuminative power has not proven to be great.” On the other hand, Tavits (2008: 114) has argued that most research on new entrants has focused on identifying variables that facilitate or hinder new party emergence and success, but no theoretical model has been articulated. The few existing theoretical arguments for the fluctuations in the support of new entrants (e.g., Tavits, 2008) are derived from the conventional coordination model embedded in the interactive argument, but offer no insights to reformulate it.

The existing research only identifies the electoral system as a determinant of the transaction costs for changes in the number of parties. According to Cox (1997: 264–265), “in strong electoral systems [i.e., majoritarian systems], the costs of failed coordination are particularly high.... Only really important issues can force realignments in stronger systems, whereas smaller issues can motivate realignment in proportional systems.” Even if the argument is more nuanced—as it is not just majoritarian systems compared to PR but also the fact that low district magnitudes and high legal thresholds play a role—it is, however, underspecified. The conditions under which the number of parties will change in a particular election or why the distribution of partisan support is more stable in some countries than others when using a similar electoral system are not indicated. In short, the explanation of the change or stability of the number of parties is not compelling.

We argue that the transaction costs for a change in the number of parties derive from turnout. The lower the turnout level in the election held at time $t-1$, the more likely a change in the number of parties in the election held at time $t$ due to the new party entry and/or the redistribution of partisan support among existing parties. Stability Hypothesis: The lower the turnout level in the election held at time $t-1$, the greater the probability of a change in the number of parties in the subsequent election. More specifically, the correlation between the number of parties at time $t$ and time $t+1$ should increase with turnout at time $t-1$.

Two mechanisms at the voter and elite level account for this hypothesis.

2.1. Voters

Sundquist (1983: 37) argues that changes in the configuration of party systems are composed of the conversion of established voters to new party allegiances and the mobilization of citizens who previously had not voted. Using the classic distinction between concentric circles in the American political universe (Campbell, 1960), electorates are made up of those voters who vote consistently in every election and have strong partisan attachments (‘core voters’); those who enter and exit the electorate, who have weaker partisan leanings and are more vulnerable to short-term forces (‘peripheral voters’); and those who never vote (‘nonvoters’). Turnout in a given election is the result of adding a fixed amount of (core) voters who always vote and a variable amount of (peripheral) voters who sometimes vote.

When examining the electoral consequences of variation in voter turnout in the United States, Hansford and Gomez (2016: 270, 286) argue that peripheral voters are less likely to support the electoral status quo and, compared with core voters, bring both change and noise when they go to the polls because they played a less active role in establishing the status quo in previous elections. This assumption is strongly supported by empirical, comparative evidence. In Denmark (Bhatti et al., 2019) or the United States (Nawara, 2016), peripheral voters share more characteristics with abstainers than with core voters. Core voters are
more likely to be older, better educated, more interested in politics, earn higher incomes than peripheral voters, and have stronger partisan attachments. In sum, peripheral voters are the segment of the electorate available to change partisan preference (i.e., the availability condition of competition in Bartolini’s terms (2002: 93–94)).

In low-turnout elections, peripheral voters are not significantly mobilized. The election results mainly reflect core voters’ preferences. If turnout increases in future elections, it means that peripheral voters are going to the polls and the probability of a substantial change in partisan support increases. In contrast, fragmentation generated in high-turnout elections is less likely to change as it is already capturing peripheral voters’ preferences. In future elections turnout will hardly increase, no more peripheral voters will be mobilized, and the probability of a change in the number of parties decreases.

Sanctions for not voting are not exactly core voters. However, they should be on average better informed and have a stronger political party preference than peripheral voters in voluntary voting systems given that voting is to some extent a process of habit formation (Dinas, 2012). When including compulsory voting, our assumption is that there are fewer peripheral voters in compulsory voting systems than in voluntary voting systems.

2.2. Elites

As peripheral voters are more willing to consider modifying their party choice, they are the “sought-after prize of the parallel efforts by competitors and the incentive for party competition” (Bartolini, 2002: 93). When examining mobilization efforts in national, regional, and European elections in France, Germany, and Spain, Golder et al. (2017: chapter 4) find that as turnout decreases, campaigning is more focused on core voters.

On the other hand, party systems are dramatically affected by strategic entry decisions (Cox, 1997: chapter 8). New party entry depends on the turnout level in the previous election and peripheral voter mobilization. Tavits (2008: 129) finds strong empirical evidence in East European countries supporting the hypothesis that “an increase in turnout is associated with higher support for (new parties) . . . due to dissatisfied voters who would have stayed home were they satisfied with the status quo.” More recently, Heath and Ziegfeld (2018), relying on Indian data, show that the entry of newly founded parties and successor parties enhances turnout. The mechanism driving this correlation is that “when new parties enter an area, they do not have core supporters to target, so their next best strategy is to mobilize unattached voters” (Heath and Ziegfeld, 2018: 572).

The logical follow-up to this argument is that political fragmentation is more likely to increase when turnout at time $t$ is higher than at $t-1$. This increase in turnout would imply that political entrepreneurs have seized the opportunity to profitably enter the party system through the mobilization of previously apathetic voters. This is, for instance, what Lublin (2004) shows when explaining realignment in the US South as due largely to the entry of many new voters.

3. Data and methods

To test our hypothesis, two analyses are undertaken. First, we perform a large-N comparative cross-national analysis composed of 313 lower-house elections in 63 countries. We then perform a more detailed district-level analysis of the Netherlands and Australia.

3.1. The cross-national analysis

3.1.1. Sample

The data and methods used here closely follow those of the recent and comprehensive analysis of party system fragmentation in Lublin (2017). The cases are all minimally democratic legislative elections according to Freedom House from 1990 to 2011, yielding a total of 313 elections in 63 countries. The number of elections per country varied from a minimum of one in Greece, Cape Verde, and Italy to a maximum of 11 in the United States, with an average of five elections per country. See the appendix for the elections included in the analysis.

The most complete model in Lublin’s analysis contains 349 observations in 65 countries. Our most complete model contains 313 observations and 63 countries. We have excluded two countries (Monaco and Samoa) due to the availability of data. Founding elections in new democracies have also been dropped because it is not possible to use lagged variables. Similarly, electoral reforms were undertaken in some countries from 1990 to 2011. As we are predicting the effective number of electoral parties in the election held at time $t$ using the fragmentation in the election held at time $t-1$, noise is added to the analysis when the electoral system changes. There is no rule of thumb to determine which kinds of electoral reforms create a new electoral system. Following Lipphart (1994), Renwick concentrates on major electoral reforms, that is, “any change in the electoral formula or a change of at least 20 per cent in legal threshold, district magnitude, or assembly size” (Renwick, 2011: 464). We follow Katz (2005: 58) and limit the meaning of major electoral reforms of national electoral systems to the wholesale replacement of the electoral formula through which the chamber of parliament is elected. Accordingly, we have excluded from the sample the first election held after a major electoral reform: New Zealand in 1996, Lesotho in 2002, Romania in 2008, and Bulgaria in 2009. Data are taken directly from Lublin, with the exception of our key independent variables (turnout and whether there is compulsory voting in the country), two control variables (the length of democracy and economic performance), and some lagged values for the dependent variable (the effective number of electoral parties).

3.1.2. Dependent variable

The dependent variable is measured in two different ways to show the robustness of our results. First is the effective number of electoral parties, ENEP, from Laakso and Taagepera (1979) at time $t$. For ENEP values before 1990, the sources are Bormann and Golder (2013) and Michael Gallagher (https://www.tcd.ie/Political_Science/people/michael_gallagher/ElSystems/index.php). The second measurement is the absolute difference in ENEP between two consecutive elections. As we are interested in the stability of fragmentation, and therefore the direction (positive or negative) of the change in fragmentation does not matter, our dependent variable has been measured as $ABS(\text{ENEPT}-\text{ENEPT}_{t-1})$.

3.1.3. Models and independent variables

To examine the conventional wisdom, we estimate three models:

$$
\text{ENEPT} = \beta_0 + \beta_1 \text{ENEPT}_{t-1} + \epsilon_t
$$

$$
\text{ENEPT} = \beta_0 + \beta_1 \text{ENEPT}_{t-1} + \beta_2 \ln(\text{exmag}) + \beta_3 \ln(\text{Length of democracy})
+ \beta_4 \text{ENEP}_{t-1} \times \ln(\text{exmag}) + \beta_5 \text{ENEP}_{t-1} \times \ln(\text{Length of democracy})
+ \beta_6 (\text{Controls}) + \epsilon_t
$$

$$
ABS(\text{ENEPT} - \text{ENEPT}_{t-1}) = \beta_0 + \beta_1 \text{ENEPT}_{t-1} + \beta_2 \ln(\text{exmag})
+ \beta_3 \ln(\text{Length of democracy}) + \beta_4 \text{ENEP}_{t-1} \times \ln(\text{exmag}) + \beta_5 \text{ENEP}_{t-1}
\times \ln(\text{Length of democracy}) + \beta_6 (\text{Controls}) + \epsilon_t
$$

1 The lagged value of the effective number of parties for Monaco and Samoa is not available, as only their founding elections are included in Lublin’s sample.

2 It is calculated as follows. For $n$ parties receiving votes,

$$
N = \frac{\sum_i p_i}{\sum_i p_i}
$$

The element $p_i$ is the proportion of votes obtained by party $i$ in the election.
Model 0, our baseline specification, captures fragmentation change using the effective number of electoral parties in the election held at time $t-1$ as a predictor of the effective number of parties in the election held at time $t$. If $\beta_0 < 1$, electoral fragmentation is lower in the current election than in the previous one; if $\beta_0 > 1$, electoral fragmentation has increased between the two consecutive elections.\(^3\) Models 1 and 2 examine the conventional wisdom. On the one hand, the natural logarithm of the exclusion magnitude, $\ln(\text{exmag})$, gauges electoral system permissiveness. The exclusion threshold is the maximum percentage of the vote required to gain a seat in the assembly based on district magnitude, legal threshold, or other factors: $\text{Exclusion Magnitude} = \frac{100}{\text{Number of seats}} - 1$. Larger values indicate more permissive systems. The data come from Lublin (2017). On the other hand, $\ln(\text{Length of democracy})$ is the age in years of the current regime classified by democracy; when applicable, ages were extended to 1870. The source is Cheibub et al. (2010). The relationship is expected to have the functional form of an exponential decay, so logged rather than raw values of the variable are entered in the models.

We are particularly interested in the interactive terms $\text{ENEP}_{t-1} \times \ln(\text{exmag})$, and $\text{ENEP}_{t-1} \times \text{Length of democracy}$. As the transaction costs increase in stronger electoral systems and in established democracies, the coefficient for $\text{ENEP}_{t-1}$ should be closer to 1 the less permissive the electoral system and the older the democracy. We also control for economic performance, in particular the rate of change in the (constant dollar) annual gross domestic product (GDP) from the World Bank. When accounting for the determinants of party system equilibrium, Mainwaring et al. (2017) and Lago and Torcal (2020) have shown that economic growth is a crucial variable. Given previous studies (Golder, 2006), we estimate two models for each election: $\text{ENEP}_{t-1}$ has been measured according to the number of registered voters.\(^3\) The source is the Voter Turnout Database from the International Institute for Democracy and Electoral Assistance (IDEA) and Adam Carr’s website, when the data are not provided in IDEA.

To examine the stability hypothesis, we estimate two models for each dependent variable. When $\text{ENEP}_{t-1}$ is the outcome to be explained, the models are as follows:

$$\text{ENEP}_{t} = \text{Model}(1) + \beta_1 \text{Turnout}_{t-1} + \beta_2 \text{ENEP}_{t-1} \times \text{Turnout}_{t-1} + \varepsilon$$

$$\text{ENEP}_{t} = \text{Model}(1) + \beta_1 \text{Compulsory Voting} + \beta_2 \text{ENEP}_{t-1} \times \text{Compulsory voting} + \varepsilon$$

Model 3 adds to Model 1 $\text{Turnout}_{t-1}$ and the interaction $\text{ENEP}_{t-1} \times \text{Turnout}_{t-1}$. Turnout has been measured according to the number of registered voters.\(^4\) The source is the Voter Turnout Database from the International Institute for Democracy and Electoral Assistance (IDEA) and Adam Carr’s website, when the data are not provided in IDEA.

Our hypothesis is that turnout is a stabilizer that dampens fluctuation in political fragmentation between consecutive elections. In order to test this hypothesis, we have to show that, all else being equal, the correlation between the number of parties at time $t-1$ and time $t$ increases with turnout at time $t-1$. A direct effect of turnout at $t-1$ on the number of parties at $t$ does not capture this stabilizing effect of turnout, but whether turnout is correlated with the number of parties. When running our baseline model (0), the coefficient on $\text{ENEP}_{t-1}$ captures the correlation between fragmentation in two consecutive elections; the closer to 1 the coefficient is, the more stable the fragmentation is. When including in the model the interaction $\text{ENEP}_{t-1} \times \text{Turnout}_{t-1}$ we are testing the stabilizing role of previous turnout. In order to support our hypothesis we should find that the coefficient on $\text{ENEP}_{t-1}$ tends to 1 the greater the turnout at time $t-1$.

Model 4 is a robustness check and replicates Model 3 using Compulsory Voting (1 if there is formal compulsory —whether enforced or not— voting; 0, otherwise) instead of $\text{Turnout}_{t-1}$. As compulsory voting increases turnout, our expectation is that the coefficient for $\text{ENEP}_{t-1} \times \text{Compulsory Voting}_{t-1}$ should be closer to 1 in those countries using compulsory voting. The data source is IDEA.

When $\text{ABS} (\text{ENEP}_{t-1} - \text{ENEP}_{t-1})$ is the dependent variable, the models are as follows:

$$\text{ABS} (\text{ENEP}_{t-1} - \text{ENEP}_{t-1}) = \text{Model}(2) + \beta_1 \text{Turnout}_{t-1} + \beta_2 \text{ENEP}_{t-1} \times \text{Turnout}_{t-1} + \varepsilon$$

$$\text{ABS} (\text{ENEP}_{t-1} - \text{ENEP}_{t-1}) = \text{Model}(2) + \beta_1 \text{Compulsory Voting} + \beta_2 \text{ENEP}_{t-1} \times \text{Compulsory voting} + \varepsilon$$

High fragmentation in the previous election should affect the extent to which fragmentation is likely to change in the subsequent election: the greater the number of parties at $t-1$, the less stable the political fragmentation is. As the coefficient on $\text{ENEP}_{t-1}$ is expected to be positive (and this is what our results show), if turnout at $t-1$ is a stabilizer of political fragmentation, the coefficient on the interactions $\text{ENEP}_{t-1} \times \text{Turnout}_{t-1}$ and $\text{ENEP}_{t-1} \times \text{Compulsory Voting}_{t-1}$ should be negative. That is, the greater the turnout at $t-1$, the more stable the fragmentation will be (i.e., the coefficient on $\text{ENEP}_{t-1}$ should be less than 0).

In all of the models we included three controls that might change between two consecutive elections: the effective number of presidential candidates, $\text{ENPRES}$,\(^5\) the temporal proximity of presidential and legislative elections, $\text{proximity}$,\(^6\) and the interaction between them. Strong evidence suggests that presidential elections have a significant coattails effect on legislative fragmentation (Golder, 2006). In Lublin’s original estimates, the effective number of electorally relevant ethnoregional groups and the percentage of seats allocated in upper tiers are included. However, as those are time-invariant, such variables cannot capture change and are therefore excluded in the models. The descriptive statistics of the variables are shown in Table 1.

### Table 1

<table>
<thead>
<tr>
<th>Variables</th>
<th>Obs</th>
<th>Mean</th>
<th>Std. Dev.</th>
<th>Min</th>
<th>Max</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\text{ENEP}$</td>
<td>313</td>
<td>4.15</td>
<td>1.96</td>
<td>1.67</td>
<td>11.21</td>
</tr>
<tr>
<td>$\text{Turnout}_{t-1}$</td>
<td>313</td>
<td>73.6</td>
<td>13.62</td>
<td>28.07</td>
<td>97.16</td>
</tr>
<tr>
<td>Compulsory voting</td>
<td>316</td>
<td>0.19</td>
<td>0.39</td>
<td>0</td>
<td>1</td>
</tr>
<tr>
<td>$\ln(\text{exmag})$</td>
<td>313</td>
<td>1.83</td>
<td>1.53</td>
<td>0</td>
<td>6</td>
</tr>
<tr>
<td>$\ln(\text{Length of democracy})$</td>
<td>313</td>
<td>3.53</td>
<td>0.93</td>
<td>0.00</td>
<td>4.97</td>
</tr>
<tr>
<td>GDP$^{-1}$</td>
<td>313</td>
<td>2.87</td>
<td>3.48</td>
<td>-14.4</td>
<td>12.3</td>
</tr>
<tr>
<td>$\text{ENPRES}$</td>
<td>313</td>
<td>1.17</td>
<td>1.57</td>
<td>0</td>
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<tr>
<td>Proximity</td>
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<td>0.24</td>
<td>0.38</td>
<td>0</td>
<td>1.4</td>
</tr>
</tbody>
</table>

The descriptive statistics correspond to Model 3 in Table 2, except the compulsory voting variable that corresponds to Model 4.

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\(^3\) Strictly speaking, this interpretation corresponds to a model with no constant.

\(^4\) It could be argued that for some countries, obtaining and relying on the figure of registered voters is tricky. For example, for the United States, McDonald and Popkin (2001) have recommended relying on voting eligible population (VEP) rather than registered voters. Obtaining VEP figures in a comparative perspective is not an easy task, though. The only effort of which we are currently aware is Stockemer’s (2017). We prefer relying on registered voters, as this maximizes the number of observations in our sample, and Stockemer’s VEP figures are based on the assumption that in each country there is the same distribution of immigrants and emigrants of voting age.

\(^5\) It is calculated as follows. For $n$ parties receiving votes, $N = \sum_{i=1}^{n} p_i$. The element $p_i$ is the proportion of votes obtained by candidate $i$ in the presidential election.

\(^6\) Proximity is calculated as $2 \times (L_1 - P_{t-1} - P_{t-1} + 1/2)$ where $L_i$ is the year of the legislative election; $P_{t-1}$ is the year of the previous presidential election; and $P_{t-1}$ is the year of the next presidential election. The variable takes the maximum value of 1 when the two elections are concurrent and the minimum value of 0 when the presidential election is held halfway between legislative elections and in parliamentary regimes.
In our view, there are no reasons to expect that turnout in the previous election was endogenous to ENEP. In a meta-analytic assessment of the determinants of turnout, Cancela and Geys (2016: 267) show that the effect of political fragmentation (i.e., the number of parties running in the election) on turnout has been extensively researched: the conclusion is that the number of parties in the election has “little direct, independent relation to voter turnout.”

Following Clark and Matt (2006: 690) and Lublin (2017: 379) and considering the small number of elections per country available in our data, our models are estimated using pooled ordinary least squares (OLS) with robust standard errors clustered by country. The small number of elections per country prevents us from using methods for standard time-series cross-section data.

4. Results

4.1. The cross-national analysis

The results of our cross-national estimates are shown in Table 2. The baseline model (Model 0) shows that 84 percent of the effective number of electoral parties in the previous election is transferred to the next election. The variable is statistically significant at the 0.01 percent level. Interestingly, when running the model with no constant (not shown), the coefficient is below 1 (0.98); that is, the average fragmentation of party systems tended to decrease. This result is not surprising. The number of parties mainly changes in the first elections when coordination failures are reduced, and many new democracies are included in the sample. As electoral fragmentation decreases over time, the additional transaction costs for the coordination of a different set of parties in strong electoral systems, established democracies, and high-turnout elections make the expected signs of \( \beta_{\text{ENEP}, \text{t}, \text{1} \times \text{Turnout}, \text{t}, \text{1}} \times \text{Compulsory voting} \) positive. That is, the coefficient on \( \text{ENEP}, \text{t}, \text{1} \) should be greater (i.e., it should be closer to 1) when turnout increases. Of course, a graphical simulation is crucial to see the effect.

Models 1 and 2 test the conventional wisdom. In Model 1 the interaction between \( \text{ENEP}, \text{t}, \text{1} \) and the length of democracy is positive and statistically significant at the 0.01 percent level. The reduction in the number of parties is therefore greater in new democracies than in established democracies. This result is clearly in line with the conventional wisdom: the older a democracy, the more stable the party system. The left-hand side in Fig. 1 illustrates this effect. The coefficient for \( \text{ENEP}, \text{t}, \text{1} \) moves toward 1, and in a monotonically increasing way the older the democracy. The interaction between \( \text{ENEP}, \text{t}, \text{1} \) and the permissiveness of the electoral system has the expected negative sign, but it is not statistically significant. The right-hand side in Fig. 1 shows the effect of the electoral systems’ permissiveness on fragmentation change. None of the controls can account for the short-run change in the number of electoral parties. In Model 2 all the coefficients are insignificant, although \( \text{ENEP}, \text{t}, \text{1} \) has the expected positive sign.

Models 3 and 4 test the stability hypothesis using \( \text{ENEP}, \text{t}, \text{1} \) as dependent variable. As expected, the two interactions between \( \text{ENEP}, \text{t}, \text{1} \) and the turnout level in the previous election (M3) and whether the election is held under compulsory voting (M4) are positive and statistically significant at the 0.01 percent level. The higher the turnout level, the more effective the number of electoral parties that are transferred to the next election. In Fig. 2 the effect of the two variables on fragmentation change is shown. As can be seen in the left-hand side of Fig. 2, when the turnout level in the previous election is very high (i.e., greater than 85 percent, for instance in Mongolia), the coefficient for \( \text{ENEP}, \text{t}, \text{1} \) approaches 1. However, when turnout in the previous election is approximately 50 percent (for instance in Poland), only 70 percent of \( \text{ENEP} \) in the previous election is transferred to the next election. The coefficients and statistical significance of the remaining variables show little change in comparison with Model 1. The interaction between \( \text{ENEP}, \text{t}, \text{1} \) and the length of democracy remains positive and statistically significant at the 0.01 percent level, whereas all other variables remain at no statistical significance. The fit of the models testing the stability hypothesis, 0.780 and 0.785, is larger than in Models 1 and 2.

Finally, Models 5 and 6 test the stability hypothesis using \( \text{ABS}(\text{ENEP}, \text{t}, \text{1} - \text{ENEP}, \text{t}, \text{1}) \) as dependent variable. The most relevant result is that the two interaction terms, \( \text{ENEP}, \text{t}, \text{1} \times \text{Turnout}, \text{t}, \text{1} \) and \( \text{ENEP}, \text{t}, \text{1} \times \text{Compulsory voting} \), have the expected negative sign and are statistically significant at the 0.01 percent level: in two countries with an unusually high fragmentation in the election at \( t-1 \), fragmentation will drop more in the subsequent election in the country with the highest turnout at \( t-1 \).

4.1.1. Robustness checks

To show that our results are not driven by unusual observations or outliers, we have followed two strategies. First, we ran the models testing the stability hypothesis (Models 3 and 4) after sequentially deleting those observations whose Studentized residuals were greater than 3 and 2 in absolute value. The number of observations dropped from 313 to 306 and 296 when using the lagged turnout as independent variable and from 316 to 311 and 297 when using compulsory voting as a regressor. As can be seen in Tables A1–A2 in the appendix, the results do not change appreciably.

Second, we used a jackknife procedure, in which we sequentially dropped each country one at a time and then re-estimated Models 3 and 4 (i.e., those testing the stability hypothesis) for each of the reduced data sets. In Fig. 3 we plot the t-statistics for the two interactive terms, \( \text{ENEP}, \text{t}, \text{1} \times \text{Turnout}, \text{t}, \text{1} \) and \( \beta_{\text{ENEP}, \text{t}, \text{1} \times \text{Compulsory voting}} \), after separately estimating the 63 models. In the left-hand side, we can see that all of the t-statistics for the interactive terms \( \text{ENEP}, \text{t}, \text{1} \times \text{Turnout}, \text{t}, \text{1} \) are above 2.5, except when Poland is dropped, when the t-statistic is 1.68 and statistically significant at the 0.1 percent level. In the right-hand side of Fig. 3 we can see that all of the t-statistics for the \( \text{ENEP}, \text{t}, \text{1} \times \text{Compulsory voting} \) term are greater than 2 (i.e., the coefficient is statistically significant at the 0.05 percent level or better), with the exception of the model in which Brazil is dropped, where the t-statistic is 1.84 (statistically significant at the 0.07 percent level). This result is not surprising. There are 12 countries using compulsory voting, comprising 61 observations, and six observations correspond to Brazil (9.84 percent).

5. A closer look at The Netherlands and Australia

In our second empirical analysis, we examine the stability hypothesis with district-level data from two countries (the Netherlands and Australia) using compulsory voting at some other time in the last century. This different identification strategy will serve as a robustness check of our results, in particular when the two countries employ different electoral systems.

5.1. The Netherlands

The Netherlands adopted compulsory voting at the same time as universal suffrage and proportional representation with a single national district allocating at least 100 seats. The sanctions for not voting were not severe. According to Irwin (1974: 292), “the fine for not voting was never high. . . and it was seldom levied (only 577 of a potential 400,000 non-voters were even brought to court in 1966). Instead, obedience was never high. . . and it was seldom levied (only 577 of a potential 400,000 non-voters were even brought to court in 1966). Instead, obedience was simply recognition that that was the law and the law should be obeyed.”

In the 13 elections held with compulsory voting between 1918 and 1967, the average turnout was 93.31 percent and the variation coefficient was 2.55 percent, and in the 15 elections held after the repeal of compulsory voting, from 1970 to 2017, the average turnout was 80.55 percent and the variation coefficient was 5.33 percent. In other words, turnout was much more variable when there was no compulsory voting.

\[ \text{To save space, we exclusively focus on those models using } \text{ENEP} \text{ as dependent variable.} \]
According to the stability hypothesis, the effective number of electoral parties should be more stable from 1918 to 1967 than from 1970 to 2017. Using the threshold suggested by Emanuele and Chiaramonte (2018), in five elections the overall vote share of new parties was above 5 percent (in 1922 and 1937 when using compulsory voting and 1994, 2002, and 2017 when not using compulsory voting). In order to have enough observations to run our models (i.e., we have one observation per election given that a single national district is employed), we use data from the 28 elections in the 1918–2017 period. To test the stability hypothesis and to maximize the degrees of freedom, we estimated the following models:

\[ \text{ABS}(\text{ENEP}_t - \text{ENEP}_{t-1}) = \beta_0 + \beta_1 \text{Turnout}_{t-1} + \epsilon_t \]  \hspace{1cm} (8)

\[ \text{ABS}(\text{ENEP}_t - \text{ENEP}_{t-1}) = \beta_0 + \beta_1 \text{Compulsory voting} + \epsilon_t \]  \hspace{1cm} (9)

Our expectation is that \( \beta_1 \text{Turnout}_{t-1} \) and \( \beta_1 \text{Compulsory voting} \) should depress the change in party system fragmentation between elections.

### 5.2. Australia

In Australia we focus on the last two elections with voluntary voting.
(1919 and 1922) and the first two with compulsory voting (1925 and 1928). No new relevant parties emerged in the four elections. Because in Australia a majority preferential system in 75 single-member districts is employed and most of the districts are the same before and after the adoption of compulsory voting, we have enough observations for running our models with only four elections. We focus only on those districts existing in the four elections and whose electoral boundaries did not change. Given that we are examining elections held in a short period of time (from December 1919 to November 1925), societal changes affecting our results are expected to be negligible or at least much less important than in the Netherlands.

The adoption of compulsory voting in 1924 immediately raised turnout, almost 91 percent in 1925 and 93 percent in 1928, and reduced differences across districts: the variation coefficient was almost 20 percent in 1922, but only 3.5 percent in 1925. The stability hypothesis will be tested through the estimation of Models 5 and 6 using district-level data from 1919 to 1922 elections and 1925 and 1928 elections.

### 5.3 Results

#### 5.3.1 The Netherlands

As we already know, turnout in the Netherlands was less variable across elections under compulsory voting (1917–1967) than in elections under voluntary voting (1971–2017). Therefore, our expectation is that the effective number of electoral parties should be more stable in the former period than in the latter. The first piece of evidence supporting the stability hypothesis is displayed in Fig. 4. The effective number of electoral parties in every election is plotted separately for the two periods.

![Stability Hypothesis](image1)

**Fig. 2.** Simulating the effect of turnout and compulsory voting on the change in the number of parties. The dashed lines represent the 95 percent confidence interval.

![Jackknife Test. Dropping Countries](image2)

**Fig. 3.** Jackknife test.
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periods using a 95 percent confidence interval. As can be seen, the range of ENEP within the confidence interval is substantially smaller when using compulsory voting (about one effective electoral party) than when using voluntary voting (about 2.5 effective electoral parties). Despite the narrower margin in the compulsory voting period, the fit of the linear model is much better than in the voluntary voting period.

This graphical pattern is strongly supported by the regression results displayed in Table 3. The absolute difference in ENEP between two consecutive elections is lower when using compulsory voting (first model) and the greater the turnout level in the previous election (second model). The two variables are statistically significant at the 0.05 percent level.

According to our argument, the mechanism accounting for this greater stability of the party system when elections are held under compulsory voting is the lower number of peripheral voters in comparison with those elections held under voluntary voting. When voting is voluntary, peripheral voters emerge, and they are less supportive of the electoral status quo. The evidence from a survey conducted immediately after the provincial election in 1970, the first one after the repeal of compulsory voting, strongly supports our argument. Respondents who recalled voting in the 1967 national election under compulsory voting but abstained in the 1970 provincial election under voluntary voting were younger, less interested in politics, and had weaker party identification and lower political efficacy than those who recalled voting in both elections (Irwin, 1974).

5.3.2. Australia

The relationship at the district level between the (absolute) change in ENEP and turnout under two situations –(1) last two elections with voluntary voting (1919 and 1922) and turnout in 1919; and (2) first two elections (1925 and 1928) with compulsory voting and turnout in 1925–using a 95 percent confidence interval is plotted in Fig. 5. As can be seen, while there is a strong negative correlation between the two variables when using voluntary voting, it virtually disappears when compulsory voting is adopted.

The regression results in Table 4 are in line with the graphical evidence. The districts before and after the adoption of compulsory voting have been pooled. As in the Netherlands, the absolute difference in ENEP between two consecutive elections is lower with compulsory voting than with voluntary voting (M7) and the greater the turnout level in the previous election (M8).

6. Conclusions

The exploration of the forces that shape changes in the number of parties shows that the greater the turnout level in a given election, the less likely a substantial change in the distribution of partisan support in the subsequent election. Using the classical distinction between core and peripheral voters, we argue that transaction costs for electoral coordination depend on the number of peripheral voters mobilized in the previous election. In low-turnout elections, only core voters go to the polls. A massive mobilization of peripheral voters, who are less likely to support the electoral status quo, might challenge the existing party system. By contrast, after a high-turnout election, in which many peripheral voters are mobilized, the existing distribution of partisan support is not at risk. At the same time, for political entrepreneurs, the turnout level in the previous election is a proxy capturing the opportunity for new, successful entries. The argument is also strongly supported with cross-national evidence and two case studies in the Netherlands and Australia.

We derive two substantive and institutional implications from our analysis of how turnout affects the transaction costs in electoral

### Table 3

<table>
<thead>
<tr>
<th></th>
<th>M7</th>
<th>M8</th>
</tr>
</thead>
<tbody>
<tr>
<td>Compulsory voting (0 = no)</td>
<td>-0.574** (0.247)</td>
<td>-0.0385** (0.0173)</td>
</tr>
<tr>
<td>Turnout, t</td>
<td>1.031*** (0.212)</td>
<td>4.115** (1.555)</td>
</tr>
<tr>
<td>Constant</td>
<td>0.161</td>
<td>0.156</td>
</tr>
<tr>
<td>Observations</td>
<td>27</td>
<td>27</td>
</tr>
<tr>
<td>R-squared</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

**p < 0.05; ***p < 0.01. Estimation is by OLS. Robust standard errors in parentheses.

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8 The data are from Bormann and Golder (2013) and Gallagher. For those elections held before 1946, we have calculated ENEP using the elections results from http://www.nlverkiezingen.com/TK1837P.html (retrieved March 14, 2020).

9 The data are from CLEA (Kollman et al., 2019).
coordination. First, the link between electoral systems and turnout is puzzling within existing research (Grofman and Selb, 2011). Countries using proportional representation have higher national-level turnout than countries using single-member single plurality, yet turnout does not increase with the (effective) number of parties. According to our argument, turnout and the number of parties are not correlated over time. When fragmentation is generated in high-turnout elections, turnout should drop in future elections due to the demobilization of peripheral voters and the strategic withdrawal of nonviable parties, whereas the effective number of parties should not change significantly.

Second, party system institutionalization has traditionally been considered a necessary condition for the consolidation of democracy (Mainwaring, 2018). As the transaction costs for the coordination of a different set of parties increase with the turnout level, the adoption of compulsory voting in new democracies should be seriously considered in institutional engineering.

Finally, the introduction of universal male suffrage and women’s suffrage meant a massive entry of inexperienced voters and led to a radical new electoral arena. It is worthwhile exploring whether the impact of moving from limited to universal suffrage on political fragmentation was modulated by turnout changes.

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Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.electstud.2021.102349.

References


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Table 4
The determinants of the change in the number of parties in Australia.

<table>
<thead>
<tr>
<th>Model</th>
<th>Turnout, -1</th>
<th>Constant</th>
<th>R-squared</th>
</tr>
</thead>
<tbody>
<tr>
<td>7</td>
<td>-0.296*** (0.064)</td>
<td>1.396*** (0.305)</td>
<td>0.167</td>
</tr>
<tr>
<td>8</td>
<td>-0.013*** (0.003)</td>
<td>0.462*** (0.056)</td>
<td>0.151</td>
</tr>
</tbody>
</table>

**p < 0.01. Estimation is by OLS. Robust standard errors in parentheses.

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Fig. 5. Turnout and the effective number of electoral parties in Australia. The shaded areas represent the 95 percent confidence interval.

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Appendix A. Supplementary data

Supplementary data to this article can be found online at https://doi.org/10.1016/j.electstud.2021.102349.


