Keeping the rascals in: Anti-political-establishment parties and their cost of governing in established democracies

JOOST VAN SPANJE

European University Institute, Florence, Italy and Amsterdam School of Communication Research, The Netherlands

Abstract. Coalition governments in established democracies incur, on average, an electoral ‘cost of governing’. This cost varies across coalition partners, and is higher for anti-political-establishment parties. This is because, if such a party participates in a coalition, it loses the purity of its message by being seen to cooperate with the political establishment. In order to demonstrate that anti-political-establishment parties suffer an additional cost of governing, this article builds on the work by Van der Brug et al. and refines the standard cost of governing theory by ‘bringing the party back in’. The results of the analyses, based on 594 observations concerning 51 parties in seven Western European countries, cast doubt on the conventional concept of a cost of governing that pertains to all parties equally. The findings call for a major revision of the standard cost of governing literature, while adding a significant contribution to the debate on strategies against parties that may constitute a danger to democracy.

Keywords: far right parties; communist parties; elections; government; Western Europe

Coalition governments in established democracies incur, on average, an electoral ‘cost of governing’ (e.g., Strøm 1990; Paldam 1986; Nannestad & Paldam 2002; Powell & Whitten 1993; Rose & Mackie 1983). In the literature, this effect has been assumed to apply to all parties equally. In multiparty systems, however, variation may exist in the extent to which coalition partners are vulnerable to this cost (see, e.g., Buelens & Hino 2008: 159). The French Communist Party, for instance, lost more than six percentage points following its participation in government in the early 1980s. A more recent case is the downfall of the Freedom Party of Austria (FPÖ) after holding office. Do some parties systematically incur a larger cost of governing than others? If so, why?

These questions typically have been ignored in the relevant literature. By taking the government as the unit of analysis, the standard cost of governing literature has implicitly assumed that characteristics of the individual parties are irrelevant. In this article, however, I argue that anti-political-establishment parties (Abedi 2002, 2004; Schedler 1996) lose the purity of their message by...
being seen to cooperate with the political establishment, and suffer an additional cost of governing as a result. In order to demonstrate this, I go beyond the existing literature to examine the cost of governing by party instead of by government. On the basis of 594 observations concerning 51 parties in seven Western European countries, I demonstrate that the cost of ruling is higher for anti-political-establishment parties than for others. In doing so, I refine the standard cost of governing theory by ‘bringing the party back in’.

My conclusions are also important beyond their scientific relevance. In showing that anti-political-establishment parties incur an additional cost of governing when they are brought in from the cold, I contribute to the debate on how to protect (the quality of) democracy from parties or movements that may constitute a danger to it. My findings imply that, from the perspective of ‘defending democracy’ (e.g., Pedahzur 2004; Capoccia 2005), it may in some circumstances be an effective strategy to invite an anti-political-establishment party to join a government coalition once it seems to have obtained ‘coalition potential’ (Sartori 1976: 122). Where anti-political-establishment parties, like other parties, seek office, policy and votes (Strøm & Muller 1999), such an invitation would pose a strategic dilemma for the anti-political-establishment party. If it rejects the opportunity, it would neither be able to share in the spoils of office nor have its policies enacted. It would also forego the legitimacy and visibility associated with government status. If, on the other hand, the offer is accepted, this would be likely to generate a considerable electoral cost to the anti-political-establishment party, as demonstrated in this study.

In this article, I begin with an overview of previous work on the topic, including three general theories explaining the cost of governing. After this, I present a new, party-specific cost of governing theory. I formulate two hypotheses on the basis of this theoretical framework. I describe the data (and case selection) and the method that I use to test the hypotheses – including control variables used. I then present the results of my analyses. The article concludes by summarising my findings and putting them into a wider perspective.

**Previous work**

Governments in multiparty systems have been found to incur a loss estimated to average between 1.0 and 3.15 percentage points per government (e.g., Paldam 1986: 19; Nannestad & Paldam 2002: 21; Powell & Whitten 1993: 410; Rose & Mackie 1983; Strøm 1990: 124). As Nannestad and Paldam (2002: 28) point out, the cost of ruling effect is ‘unusually constant’ in established democracies, bearing in mind the differences, for example, in electoral system and
party system between countries. They therefore refer to this effect as a ‘basic fact’, which should build on some ‘deep parameter in human behavior’ (Nannestad & Paldam 2002: 22, 28).

What would this deep parameter be? In the existing literature at least three theories have been put forward that would explain the cost of ruling: the coalition of minorities idea, the median gap theory and the grievance asymmetry theory (see Nannestad & Paldam 2002: 28–33; see also Vermeir & Heyndels (2007) for an overview).

Let us start with the coalition of minorities idea. In his landmark study on democracy, Anthony Downs (1957) suggested that an administration, even if it always represents a majority on each issue, will always put off some minority of voters when it acts. The voters belonging to the minority are likely to vary from issue to issue. As a result, the opposition will be able to mobilise a coalition of minorities, which may add up to a majority of voters at some point (Downs 1957: 55–60). Mueller (1970: 20) links this idea to the decrease of American presidents’ popularity over time. In addition, he alludes to the idea that disillusionment of groups of voters will increase over time due to differences between promises made by the incumbent political actor and his actual choices. In order to be elected, a political actor ‘invariably says or implies he will do more than he can do and some disaffection of once bemused supporters is all but inevitable’ (Mueller 1970: 20). Once in government, the actor’s true preferences will be gradually revealed, and groups of voters will defect as a consequence. Nannestad and Paldam (2002: 28–29) elaborate upon this idea, and acknowledge the possibility that this mechanism explains the cost of ruling in multiparty systems.²

An alternative theory builds on work by Hotelling (1929). In a setting where two parties compete for voters whose preferences are (close to standard normally) distributed along a one-dimensional ideological axis, the parties will end up competing for the median voter. As a result, the parties tend to position themselves near the centre of the distribution (cf. Hotelling, 1929). Yet, they do not completely converge, according to Paldam and Skott (1997), so as to avoid being completely indistinguishable. Under these circumstances, the voters in the gap between the two parties can do no better than to vote for the opposition party at each election. This is because these centrist voters will find the government policy too leftist when the left party is in power, and too right-wing when the right party is in government. Paldam and Skott argue that these centrist swing votes equal the cost of governing. Stevenson (2002) extends this model so that it also explains why this cost goes up as government duration increases.

A third theory explaining the cost of ruling – that of grievance asymmetry (see also Mueller 1970: 23; Bloom & Price 1975) – is based on the idea that
poor performance of the country’s economy has a larger influence on the incumbent government’s popularity than good performance. An increase in inflation, for example, will hurt a government more than it would benefit from a decrease. Mueller (1970: 23; emphasis in original) even finds that ‘an economy in slump harms a President’s popularity, but an economy which is improving does not seem to help his rating’. As economic performance is bound to average out over time, governments will, on average, incur a cost of governing as a result. Notwithstanding the lack of scholarly consensus about this effect (Lewis-Beck 1988; Van der Brug et al. 2007), several studies (e.g., Nannestad & Paldam 1997; Soroka 2006) have demonstrated that there is such ‘grievance asymmetry’, which Nannestad and Paldam (2002: 32, 33–36) link to the recurrent finding that losses have larger effects on people’s preferences than gains (e.g., Tversky & Kahneman 1991).

A party-specific cost of governing theory

These three theories have in common that they are based on two-party systems such as that of the United States. Because the winning party usually forms a government in two-party systems, no distinction between party and government is needed in these settings. This is obviously different in multi-party contexts. Yet the government has almost always been the unit of analysis in the existing literature. When applied to multiparty systems, then, the cost of ruling has implicitly been supposed to pertain to all the coalition partners equally.

In this article, by contrast, I suggest that the cost of governing is different for different kinds of parties. This is not an entirely new argument, as Buelens and Hino (2008: 159) have casually hinted that ‘Communists/Extreme left’ and ‘Right-wing populists’ seem to lose more than members of other party families (for party-specific effects resulting from a country’s economic performance, see Feld & Kirchgässner 2000; Geys & Vermeir 2008; Van der Brug et al. 2007). What is new, however, are the two theories I present here, both of which aim to explain why these parties incur higher losses than other parties – and my empirical test of these theories.

Why would some parties lose more votes after government participation than others? Clearly, such an effect is not expected on the basis of the median gap theory as this theory applies to two-party systems with a one-dimensional ideological space. The grievance asymmetry theory does not make any predictions on this point either as it concerns entire government coalitions. In doing so, it does not explain additional losses suffered by one of a government coalition’s parties. The coalition of minorities idea, by contrast, can readily be
modified to apply to multiparty systems, and to base predictions on concerning party-specific costs of governing. Thus, I build my theory on this idea, originating in Downs’ seminal work.

I begin with the notion that government parties lose out because they cannot deliver on their promises – or, more precisely, once in government they are unable to live up to voters’ expectations. These expectations can be about particular policies – as argued in the literature mentioned above – but also about other considerations upon which voters may base their vote choice. If anything, parties that suffered greatly after government participation, such as the French Communists (PCF) and the Austrian FPÖ, have in common that they (aim to) mobilise against the political establishment. Thus, their additional cost of governing may be due to their attitude towards the establishment, which raises expectations among the electorate.

Willingly or unwillingly, these parties are perceived as opposing the political establishment. As such, they belong to a specific party type – that of the anti-political-establishment party (Schedler 1996; Abedi 2002, 2004). As a result of this image, they attract voters with anti-establishment feelings. Once one of these parties joins a government coalition, however, it will have to cooperate with the establishment. In doing so, it will most likely disappoint voters who expected it to ‘throw the rascals out’ (Miller & Wattenberg 1985) instead of keeping the rascals in by joining them. Insofar as the party was previously receiving votes on the basis of its anti-political-establishment stance, some of those votes may be lost. Both in expressive accounts of rational voting (e.g., Greene & Nelson 2002; Brennan & Hamlin 1998) and in instrumental accounts (e.g., Austen-Smith 1983; Enelow & Hinich 1984; Shepsle 1991), a party that mobilises on anti-political-establishment feelings and ideologies is seen as likely to lose electoral support if it joins a governing coalition. This is a cost that other parties do not face.

An anti-political-establishment party that has participated in a government coalition with established parties can no longer credibly make the claim that it aims to kick the establishment out. Voters will therefore defect from these parties after government participation (cf. Sani 1978: 94, 1979: 73–82). I thus expect an anti-political-establishment party to lose electoral support after taking up government responsibility through the very act of accepting to share government responsibility with the establishment. I formulate my first hypothesis below.

**H1:** An anti-political-establishment party that joins a government coalition will lose votes over and above what other parties would expect to suffer in the next election.

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A second reason why parties incur an extra cost of governing compared to others is associated with not delivering on their promises in terms of policy. Because they generally take up radical policy positions, far left and far right parties that forge coalitions with other parties (such as the French PCF and the Austrian FPÖ) have to compromise on many issues. As a result, they are likely to face more disaffection among their supporters after government participation than more moderate parties, so that relatively many of their voters might switch to other parties, or abstain. The defecting radical voters will presumably not immediately be replaced by more moderate voters since such parties’ radical images will not change right away, even if they have participated in government. If the far left or far right party governs with parties that are ideologically distinct — for instance, if communists enter a coalition that includes right-of-centre parties, or if a radical right party shares power with social democrats — it may have to make more compromises, and therefore lose more votes, than if it strikes deals with ideologically similar parties. From this perspective, the worst-case scenario for a far left or far right party is a coalition with ideologically distinct partners — resulting, in practice, in a centrist government. The additional cost of governing related to this should be observed each time such a coalition is formed. Note that this rule would hold for any party, and not just for the far left and far right. The more radical the party and the more centrist the government coalition it enters, the higher the predicted electoral cost.

\[ H2: \text{The more radical a party’s ideological profile and the more centrist the ideological profile of the government coalition it joins, the more votes the party will lose in the next election.} \]

Control variables

The literature suggests that several variables be controlled for when assessing hypotheses regarding the electoral cost of governing in established democracies. I explicitly control for all the variables that have been shown to affect changes in electoral success. First of all, three characteristics of the government formed in the period since the previous election will be added — whether it is a minority government or a majority government, how many coalition partners it has and its duration in months. Minority governments are expected to incur a smaller cost of governing than majority governments because they can credibly make the claim that their efforts were blocked by other parties (e.g., Paldam 1991: 123; Powell & Whitten 1993: 401, 408; Rose & Mackie 1983; Strøm 1990: 124). The cost of governing is also expected to be attenuated by a
government’s number of coalition partners. The more parties in government, the less clear it is for voters which party can be held accountable for the policies enacted, so the smaller the cost of governing becomes (e.g., Powell & Whitten 1993: 401–402, 403–404) – at least, the cost by party. The government duration variable captures the expected effect that the longer an incumbent government has been in power, the larger its cost of governing becomes (e.g., Nannestad & Paldam 2002: 27; Stevenson 2002).

Second, government parties have turned out to be more severely punished on average in elections since the 1990s (Nannestad & Paldam 2002: 26) as a result of increasing voter volatility in Western democracies. Hence, a dummy variable will be included, distinguishing elections held since 1989 from those held earlier.9

Third, an economic indicator, the GDP growth in the election year,10 will be used as a control variable (see also Van der Brug et al. 2007). This variable will be added in interaction with Powell and Whitten’s ‘clarity of responsibility’ variable, which indicates how easy it is for voters in a given political context to identify which (government) party or parties can be held responsible for the country’s economic policies (Powell & Whitten 1993; Whitten & Palmer 1999; Royed et al. 2000; Van der Brug et al. 2007; Stevenson & Vavreck 2000).

Case selection

In this article, I use the concept of the ‘anti-political-establishment party’ as developed in the literature (Schedler 1996; Abedi 2002, 2004). Anti-political-establishment parties ‘order the political world in a particular way. They draw up a triangular symbolic space by (simultaneously) constructing three actors and their relationships: the political class, the people and themselves. . . . While citizens and anti-political-establishment actors live in peace and harmony, their relationship with the political establishment is deeply antagonistic’ (Schedler 1996: 293).

A party is classified as anti-political-establishment on the basis of the authoritative categorisation by Abedi (2004: 32–75). A party qualifies as anti-political-establishment if it is ‘a party that challenges the status quo in terms of major policy issues and political system issues’, ‘perceives itself as a challenger to the parties that make up the political establishment’ and ‘asserts that there exists a fundamental divide between the political establishment and the people. It thereby implies that all establishment parties be they in government or in opposition are essentially the same’ (Abedi 2004: 12). The concept of the ‘anti-political-establishment party’ is based on that of the ‘anti-system party’ (Sartori 1966, 1970, 1976), which has been used frequently in recent research.
However, a definitive and up-to-date categorisation of ‘anti-system parties’ does not exist in the literature. In this study I use Abedi’s classification of parties as anti-political-establishment or not based on the three criteria mentioned above (see also Abedi 2002: 573–575).

Anti-political-establishment parties have shared in government responsibility in only a few established democracies since World War II – all of them in Western Europe. These parties have done so on 19 occasions in Austria, Belgium, Finland, France, Ireland, Italy, Luxembourg and the Netherlands.

In the countries under study, all the parties are included for which at least six successive national-level electoral results are recorded since the Second World War – both governing and opposition parties. All these parties’ election results since 1945 are included, adding up to 594 observations concerning 51 parties. Table 1 gives an overview of the electoral performance of the anti-political-establishment parties and their coalition partners before and after government participation.

These cases comprise the entire universe of observations for which comparable electoral scores are available regarding government participation of anti-political-establishment parties in established democracies. Although most often negative, the change in electoral performance after government participation of anti-political-establishment parties varies considerably across the observations, ranging from the Freedom Party of Austria (FPÖ) in 2002 (−16.9 per cent), to the Italian National Alliance (AN) in 1996 (+2.2 per cent).

Data and method

The two hypotheses are tested on the basis of a dataset compiled especially for this study. These data are derived from an EJPR political data collection publication (Woldendorp et al. 1998), as well as from updates of this database, also published in EJPR (Katz & Koole 1997, 1999, 2002; Koole & Katz, 1998, 2000, 2001; Katz 2003; Van Biezen & Katz 2004, 2005, 2006; Bale & Van Biezen 2007, 2008). See Table 2 for descriptive statistics of the data.

As I am interested in testing an effect that is assumed to be equal across space and time and in short-term adjustments rather than long-term effects, it makes sense to pool the data (e.g., Kittel & Winner 2005: 289) into a ‘time-series cross-sectional’ structure. In this article I deal with the many complications associated with this type of data structure in the standard way (Stimson 1985: 945) – that is, assessing the effects on the basis of OLS regression analysis with panel-corrected standard errors (PCSEs), as well as party fixed-effects where appropriate (Beck & Katz 1995, 1996; Beck 2001).
Table 1. Governments including anti-political-establishment (APE) parties in established democracies, 1945–2008

<table>
<thead>
<tr>
<th>Country</th>
<th>Governments including APE party</th>
<th>APE party in government</th>
<th>Result APE party, ante [% of vote]</th>
<th>Result APE party, post [% of vote]</th>
<th>Change in result APE party [% of vote]</th>
<th>Change in result coalition partners [% of vote]</th>
</tr>
</thead>
<tbody>
<tr>
<td>Belgium</td>
<td>1977–1978</td>
<td>People’s Union (VU)</td>
<td>10.0 (1977)</td>
<td>7.0 (1978)</td>
<td>-3.0</td>
<td>+4.9</td>
</tr>
<tr>
<td>Finland</td>
<td>1995–1999</td>
<td>Green League (VIHR)</td>
<td>6.5 (1995)</td>
<td>7.3 (1999)</td>
<td>+0.8</td>
<td>-8.8</td>
</tr>
<tr>
<td>Ireland</td>
<td>1948–1951</td>
<td>Family of the Republic (CnP)</td>
<td>13.2 (1948)</td>
<td>4.1 (1951)</td>
<td>-9.1</td>
<td>+3.5</td>
</tr>
<tr>
<td>Ireland</td>
<td>1948–1951</td>
<td>Offspring of the Land (CnT)</td>
<td>5.5 (1948)</td>
<td>2.9 (1951)</td>
<td>-2.6</td>
<td>+3.5</td>
</tr>
<tr>
<td>Italy</td>
<td>1994</td>
<td>National Alliance (AN)</td>
<td>13.5 (1994)</td>
<td>15.7 (1996)</td>
<td>+2.2</td>
<td>+1.8</td>
</tr>
<tr>
<td>Italy</td>
<td>1993</td>
<td>Federation of Greens (FdV)</td>
<td>2.8 (1992)</td>
<td>2.7 (1994)</td>
<td>-0.1</td>
<td>N/A[^a]</td>
</tr>
<tr>
<td>Country</td>
<td>Governments including APE party</td>
<td>APE party in government</td>
<td>Result APE party, ante [% of vote]</td>
<td>Result APE party, post [% of vote]</td>
<td>Change in result APE party [% of vote]</td>
<td>Change in result coalition partners [% of vote]</td>
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</tr>
<tr>
<td>Italy</td>
<td>1994 Go Italy (FI)</td>
<td>21.0 (1994)</td>
<td>20.6 (1996)</td>
<td>-0.4</td>
<td>+3.9</td>
<td></td>
</tr>
<tr>
<td>Italy</td>
<td>1994 Northern League (LN)</td>
<td>8.4 (1994)</td>
<td>10.1 (1996)</td>
<td>+1.7</td>
<td>+1.8</td>
<td></td>
</tr>
<tr>
<td>Luxembourg</td>
<td>1945–1946; 1946–1947 Communist Party (PCL)</td>
<td>13.5 (1945)</td>
<td>2.5(^b) (1948)</td>
<td>-11.0</td>
<td>+13.4</td>
<td></td>
</tr>
<tr>
<td>Netherlands</td>
<td>2002 List Pim Fortuyn (LPF)</td>
<td>17.0 (2002)</td>
<td>5.7 (2003)</td>
<td>-11.3</td>
<td>+3.2</td>
<td></td>
</tr>
</tbody>
</table>

Notes: \(^a\) The main government party, Christian Democracy, did not contest any elections after the formation of this government. The other parties’ results were: -11.4 per cent (Italian Socialist Party) and +4.3 per cent (Democratic Party of the Left). \(^b\) Assessing the electoral result of the PCL before and after its government participation comes with problems of comparability as the 1948 elections were held in only half of the country. There is, however, no reason to believe that the result would be completely different if the elections had been held in the other half of the country as well. After all, the electoral support for communist parties is relatively evenly spread in most countries. In any case, leaving out the observations pertaining to Luxembourg does not substantially change our findings (see also the sensitivity analyses in the penultimate section of this article).

Table 2. Descriptive statistics of the data used (seven countries, 1945–2008)

<table>
<thead>
<tr>
<th>Variable</th>
<th>Observations</th>
<th>Mean</th>
<th>Standard deviation</th>
<th>Minimum</th>
<th>Maximum</th>
</tr>
</thead>
<tbody>
<tr>
<td>Performance change</td>
<td>594</td>
<td>-0.10</td>
<td>4.06</td>
<td>-19.40</td>
<td>18.60</td>
</tr>
<tr>
<td>Incumbent</td>
<td>594</td>
<td>0.46</td>
<td>0.50</td>
<td>0.00</td>
<td>1.00</td>
</tr>
<tr>
<td>Minority government</td>
<td>594</td>
<td>0.14</td>
<td>0.35</td>
<td>0.00</td>
<td>1.00</td>
</tr>
<tr>
<td>Number of government parties</td>
<td>594</td>
<td>2.88</td>
<td>1.39</td>
<td>1.00</td>
<td>7.00</td>
</tr>
<tr>
<td>Duration of government</td>
<td>594</td>
<td>29.75</td>
<td>16.88</td>
<td>2.99</td>
<td>64.77</td>
</tr>
<tr>
<td>After 1989</td>
<td>594</td>
<td>0.33</td>
<td>0.47</td>
<td>0.00</td>
<td>1.00</td>
</tr>
<tr>
<td>GDP growth</td>
<td>594</td>
<td>3.63</td>
<td>2.85</td>
<td>-6.40</td>
<td>18.9</td>
</tr>
<tr>
<td>Clarity of responsibility</td>
<td>594</td>
<td>0.39</td>
<td>0.49</td>
<td>0.00</td>
<td>1.00</td>
</tr>
<tr>
<td>APE party</td>
<td>594</td>
<td>0.28</td>
<td>0.45</td>
<td>0.00</td>
<td>1.00</td>
</tr>
<tr>
<td>Centrism government</td>
<td>594</td>
<td>2.05</td>
<td>0.90</td>
<td>1.00</td>
<td>3.00</td>
</tr>
<tr>
<td>Radicalism party</td>
<td>594</td>
<td>0.70</td>
<td>0.15</td>
<td>0.38</td>
<td>1.00</td>
</tr>
</tbody>
</table>

Note: APE = anti-political establishment.
The non-stationarity of the data may affect the results of the type of regression analysis that I perform. Because some of the parties in my dataset have been in decline since the Second World War (e.g., the French Communist Party) and other parties have performed increasingly well over time (e.g., the Finnish Greens), some of the series of election results are non-stationary. The results of a Fisher test for panel unit roots indicate that seven out of the 51 series are non-stationary (using a 5 per cent critical value). In the penultimate section, I present results of alternative specifications of my models. I use the change in electoral result as the dependent variable in order to avoid problems due to this non-stationarity. Compared to estimating level models, estimating change models has the additional benefit that it generally removes two other major potential problems related to panel data analysis: cross-sectional heterogeneity and autocorrelation.

The models are tested for any remaining autocorrelation using Breusch-Godfrey tests (e.g., Kennedy 2003: 149; Greene 2003: 269–271), and for heteroskedasticity using ARCH tests (e.g., Greene 2003: 244–245; Kennedy 2003: 147). I include party-fixed effects in my models, as suggested by two statistical indicators: the Breusch-Pagan Lagrangian multiplier test (Breusch & Pagan 1980) and the Hausman test (Hausman 1978). Supplementary analyses show that my results are robust.

The dependent variable in my analysis is change in a party’s national-level vote share. A dichotomous variable identifying parties that have been in government since the previous election is included in each of the models. In addition, a dummy variable distinguishing between anti-political-establishment and other parties is added to the analysis. The interaction of this variable with the incumbency identifier is the main independent variable used to test the first hypothesis. This interaction effect is expected to indicate an ‘extra cost of governing’ for anti-political-establishment parties (H1), and is therefore hypothesised to be negative.

In order to test the second hypothesis, a three-way interaction variable is included in the model. This is an interaction between the incumbency dummy mentioned before, an indicator of the political complexion of the government and one of the ideological profile of each party. The government’s ideological profile is measured based on a five-point scale in the dataset by Woldendorp et al. (1998), recoded into three categories: centrist governments (3), governments of either centre-right or centre-left complexion (2) and governments with either left-wing or right-wing dominance (1). The party’s ideological profile is indicated on a scale ranging from 0 to 1. The smaller the distance from the relevant end of the political spectrum, the higher the value. Thus, if a communist party has a position of 0.91, this means that the party is positioned at 1 – 0.91 = 0.09 points from the left end of a 0–1 scale, and 0.91 points
removed from the opposite (right-wing) pole. The parties’ ideological positions are derived from various expert surveys conducted in recent decades (see Carter 2005: 216–226; Laver & Schofield 1998: 252–265). The three-way interaction variable should have a negative impact \((H_2)\). After all, the more centrist the government (positive) and the more radical the party (positive), the more compromises the government party has to make, and the more votes it is expected to lose as a consequence (negative). Unlike the \(H_1\) effect mentioned earlier, the \(H_2\) effect concerns all parties rather than anti-political-establishment parties only.

If the far left and far right parties’ extra cost of governing can be entirely attributed to the compromises that they have to make \((H_2)\), the additional cost of governing for incumbent anti-political-establishment parties – if any – should disappear when I add the three-way interaction variable, which should have a negative effect. If, by contrast, the impact of the three-way interaction does not hold when controlling for the anti-political-establishment interaction variable, and the latter variable yields a negative effect, then the additional cost of ruling is completely due to parties’ anti-political-establishment profiles \((H_1)\). It is also possible that neither effect occurs, or that both simultaneously occur.

In the following I test four models: a null model with only controls (Model 1); a model with the \(H_1\) variable (Model 2); one with the \(H_1\) as well as the \(H_2\) variable (Model 3); and a reduced-form model that includes only those variables that yield a significant impact (Model 4).

**Results**

The hypotheses are tested on the basis of several similar models explaining the electoral performance of government parties in the seven selected countries since the Second World War. I first estimate a model including all the control variables listed above (see Model 1 in Table 3).

Model 1 is able to explain 7 per cent of the variance in change in the electoral performance of the 51 parties under study since the Second World War. The results of the Breusch-Godfrey and ARCH tests indicate that the null hypothesis of no autocorrelation (Breusch-Godfrey) and no heteroskedasticity (ARCH) cannot be rejected. As the continuous variables are centred on their means, Model 1 predicts that parties in political contexts with a low clarity of responsibility incur a cost of governing of 1.90 percentage points when all other continuous variables are held constant at their means and the dummies are zero. This effect is statistically significant at the \(p = 0.01\) level (one-tailed). Parties in high clarity contexts lose significantly more votes whereas parties in broader coalitions lose significantly fewer votes.
Table 3. Analysis of the data used (change in electoral results of 51 parties in seven countries, 1945–2008)

<table>
<thead>
<tr>
<th></th>
<th>Model 1</th>
<th>Model 2</th>
<th>Model 3</th>
<th>Model 4</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>b (PCSE)</td>
<td>b (PCSE)</td>
<td>b (PCSE)</td>
<td>b (PCSE)</td>
</tr>
<tr>
<td>Constant</td>
<td>-0.45 (3.02)</td>
<td>-2.13 (2.70)</td>
<td>0.99 (0.92)</td>
<td>-1.44 (2.15)</td>
</tr>
<tr>
<td>Incumbent</td>
<td>-1.90*** (0.65)</td>
<td>-1.60* (0.71)</td>
<td>-1.59* (0.75)</td>
<td>-3.14*** (0.50)</td>
</tr>
<tr>
<td>Minority government</td>
<td>0.41 (1.06)</td>
<td>0.52 (1.07)</td>
<td>0.61 (1.08)</td>
<td></td>
</tr>
<tr>
<td>Minority government*Incumbent</td>
<td>-0.13 (1.13)</td>
<td>-0.14 (1.15)</td>
<td>-0.26 (1.21)</td>
<td></td>
</tr>
<tr>
<td>Number of government parties</td>
<td>-0.15 (0.20)</td>
<td>-0.15 (0.21)</td>
<td>-0.21 (0.21)</td>
<td>-0.15 (0.17)</td>
</tr>
<tr>
<td>Number of government parties*Incumbent</td>
<td>0.79** (0.31)</td>
<td>0.83** (0.31)</td>
<td>0.91** (0.33)</td>
<td>0.68** (0.25)</td>
</tr>
<tr>
<td>Duration government</td>
<td>-0.00 (0.01)</td>
<td>-0.00 (0.01)</td>
<td>-0.01 (0.01)</td>
<td></td>
</tr>
<tr>
<td>Duration government*Incumbent</td>
<td>-0.01 (0.03)</td>
<td>-0.02 (0.03)</td>
<td>-0.02 (0.03)</td>
<td></td>
</tr>
<tr>
<td>After1989</td>
<td>0.08 (0.48)</td>
<td>0.08 (0.52)</td>
<td>0.00 (0.54)</td>
<td></td>
</tr>
<tr>
<td>After1989*Incumbent</td>
<td>-0.97 (0.91)</td>
<td>-0.93 (0.92)</td>
<td>-0.91 (0.96)</td>
<td></td>
</tr>
<tr>
<td>GDP growth</td>
<td>-0.07 (0.14)</td>
<td>-0.06 (0.14)</td>
<td>-0.08 (0.14)</td>
<td></td>
</tr>
<tr>
<td>GDP growth*Incumbent</td>
<td>0.04 (0.19)</td>
<td>0.01 (0.18)</td>
<td>0.04 (0.19)</td>
<td></td>
</tr>
<tr>
<td>Clarity of responsibility</td>
<td>1.12 (1.28)</td>
<td>1.22 (1.26)</td>
<td>1.27 (1.27)</td>
<td></td>
</tr>
<tr>
<td>Clarity of responsibility*Incumbent</td>
<td>-2.64*** (0.79)</td>
<td>-2.63*** (0.79)</td>
<td>-2.59*** (0.79)</td>
<td></td>
</tr>
<tr>
<td>Interaction</td>
<td>Weight 1</td>
<td>Weight 2</td>
<td>Weight 3</td>
<td>Weight 4</td>
</tr>
<tr>
<td>-------------------------------------------------</td>
<td>----------</td>
<td>----------</td>
<td>----------</td>
<td>----------</td>
</tr>
<tr>
<td>GDP growth*Clarity of responsibility</td>
<td>0.17 (0.18)</td>
<td>0.16 (0.18)</td>
<td>0.20 (0.18)</td>
<td></td>
</tr>
<tr>
<td>GDP growth<em>Clarity of responsibility</em>Incumbent</td>
<td>-0.22 (0.23)</td>
<td>-0.18 (0.23)</td>
<td>-0.21 (0.23)</td>
<td></td>
</tr>
<tr>
<td>APE party</td>
<td>-0.03 (0.70)</td>
<td>-0.05 (0.74)</td>
<td>-0.09 (0.67)</td>
<td></td>
</tr>
<tr>
<td>APE party*Incumbent (H1)</td>
<td>-3.26** (1.35)</td>
<td>-3.39** (1.38)</td>
<td>-2.93* (1.35)</td>
<td></td>
</tr>
<tr>
<td>Centrism government</td>
<td>0.34 (0.34)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Centrism government*Incumbent</td>
<td>-1.07 (2.11)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Radicalism party</td>
<td>-0.47 (0.60)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Radicalism party*Incumbent</td>
<td>2.20 (2.89)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Centrism government*Radicalism party</td>
<td>-1.12 (1.55)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Centrism government<em>Radicalism party</em>Incumbent</td>
<td>0.19 (3.46)</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

N 594 594 594 594
Adjusted R² 0.07 0.08 0.08 0.07
Breusch-Godfrey test (2 lags) 0.58 0.63 0.66 0.62
ARCH test (2 lags) 0.68 0.74 0.71 0.74

Notes: * p < 0.05; ** p < 0.01; *** p < 0.001 (one-tailed). PCSE = panel-corrected standard error. Party-fixed effects are included in the analyses but not shown in the table. The continuous independent variables are centred on their means. APE = anti-political establishment.
The value of the coefficient may be somewhat misleading when it comes to interpreting this cost as it pertains, for instance, to the impossible average value of 2.88 coalition partners. I run simulations using the Clarify software (King et al. 2000; Tomz et al. 2001), which facilitate interpretation of the effects found. The results of the simulations indicate that the expected vote loss incurred by a party that shares government responsibility with one other party in a low clarity context, holding all other parameters constant at zero (which is the mean of the centred continuous variables), is 1.65 percentage points (standard error = 0.67). If it does not join a government coalition, by contrast, the party is expected to obtain 0.96 percentage points more than at the previous election (standard error = 0.65), which falls comfortably outside of the 95 per cent confidence interval of the expected result for incumbents. This suggests a cost of governing of 1.65 + 0.96 = 2.61 percentage points per party in low clarity contexts with two government parties. The effects yielded by the interaction variable identifying minority governments, by the interaction variable that captures the duration of the government and by the post-1989 interaction variable do not reach conventional levels of statistical significance.

The impact of the variable related to the economy is not significant, which may seem surprising in light of empirical evidence of economic voting (e.g., Powell & Whitten 1993; Stevenson & Vavreck 2000; Whitten & Palmer 1999; Royed et al. 2000). Van der Brug et al. (2007), however, show that electoral effects based on objective economic indicators are usually small. In addition, they argue that these effects are difficult to detect at the party level when one does not take into account that some parties cannot win many more votes because they have already mobilised almost their full potential, and others cannot lose many votes because their vote has been reduced to their bedrock supporters (Van der Brug et al. 2007: 15–18). Note that when I also include inflation figures, I find a significant effect that is much the same as one of the effects found by Van der Brug et al., and my conclusions do not substantively change. In sum, economic voting does not appear to affect the findings of interest in this study.

In order to test the first hypothesis, the $H1$ variable – the interaction of anti-political-establishment party and incumbency – is added to the model, as well as the anti-political-establishment party identifier. The resulting Model 2 explains 8 per cent of the variance. The results of my second model show that anti-political-establishment parties suffer an extra cost of governing. The effect of the $H1$ variable is in the expected direction ($b = -3.26$) and is significant at the $p < 0.01$ level (one-tailed) in the presence of all the other variables, which confirms my first hypothesis. As expected, the inclusion of the cost of governing pertaining to anti-political-establishment parties attenuates the
effect of the general cost of governing. The last-mentioned effect changes from $b = -1.90$ in Model 1 to $b = -1.60$ in Model 2.

Based on simulations using Clarify, I estimate that anti-political-establishment parties receive 4.74 percentage points (standard error = 1.45) less of the vote on average after government participation in low clarity contexts with one other coalition partner, which equals 57 per cent of the average size of such parties at the previous election (8.26 per cent). If they do not join, they obtain a vote share that is 0.80 percentage points (standard error = 0.74) larger than their previous result. This means that the cost of governing for anti-political-establishment parties in these cases is $4.74 + 0.80 = 5.54$ percentage points.20

$H2$, by contrast, is not supported. When I add the six relevant variables in order to test this hypothesis (see Model 3 in Table 3), the explained variance does not increase. I do not find the predicted negative impact of the three-way interaction variable that would indicate that parties lose more when they enter a centrist government while having a radical ideological profile itself. The impact is not in the predicted direction. The effect yielded by the incumbent anti-political-establishment identifier ($H1$) retains its significance at the $p < 0.01$ level (one-tailed) and its strength ($b = -3.39$) in the presence of this variable.

The finding that the $H2$ variable does not have the hypothesised impact can be illustrated on the basis of Clarify simulations. A maximally radical party faces an expected loss of 1.83 percentage points (standard error = 1.36) after participation in a two-party centrist government coalition. An incumbent centrist party is not expected to lose less than this in elections after having shared power in such a two-party centrist coalition, holding the other variables constant. By contrast, it is expected to incur a loss of 2.06 percentage points (standard error = 1.29). To sum up, the empirical evidence supports my first hypothesis, but not my second one.

My final model, Model 4, only includes variables that both have a significant impact and are of theoretical interest, as well as the corresponding first-order effects. The model fit is only slightly lower ($R^2 = 0.07$) than that of Models 2 and 3. This parsimonious model demonstrates the hypothesised result even when non-significant effects are excluded as the $H1$ variable still has a significant negative impact ($b = -2.93$, $p < 0.05$, one-tailed).

**Sensitivity analyses**

Four additional kinds of analyses are performed in order to test the robustness of my results, as advocated in the literature on times-series cross-sectional data
(e.g., Wilson & Butler 2007: 119). First, the findings might result from country-specific effects rather than general mechanisms, or, alternatively, they might be attributed to outliers or errors in the dataset. Thus, Model 3 is re-estimated seven times, excluding one political system from the analysis each time. The results appear to be stable, as the coefficient of the $H1$ variable varies from $b = -4.65$ (without the Italian cases) to $b = -1.96$ (if Austria is left out), with a mean of $-3.49$ and a standard deviation of 1.08. The effect induced by the $H2$ variable, by contrast, does not reach commonly applied levels of statistical significance in any of the analyses.

Second, compared to models with the level of electoral success as a dependent variable, change models suffer from the drawback that they accumulate measurement error in the dependent variable. I therefore rerun the analysis shown in Model 3 once more, this time with the electoral performance as the dependent variable instead of the change in electoral result and with a lagged dependent variable in order to account for serial correlation in the data. Substituting the dependent variable by the electoral performance level variable, and adding the previous vote share as an independent variable, causes the explained variance of Model 3 to increase to over 90 per cent. Apart from this, the results are very similar, with the $H1$ variable ($b = -2.15$) reaching statistical significance at the $p = 0.05$ level (one-tailed). The fact that its impact is slightly smaller than in Model 3 is due to the inclusion of a lagged dependent variable, which tends to suppress the effect of other independent variables (Achen 2000).

Third, the non-stationarity of seven of the 51 panels may affect the results. If I therefore re-estimate Model 3 on the basis of a reduced dataset only including the 44 stationary panel series, the $H1$ variable yields an effect that is similar ($b = -4.21$) to that in Model 3 and still significant at the $p < 0.01$ (one-tailed) level even though the number of observations is restricted ($N = 516$). At the same time, the $H2$ variable fails to show a substantial impact.

Finally, the extra cost of governing for anti-political-establishment parties could be due to their lack of government experience. An anti-political-establishment party has, almost by definition, not previously participated in a governing coalition. As soon as such a party gains access to power, this lack of experience may show, resulting in disappointment for party supporters. As a consequence, the party is likely to lose more votes after government participation than the established parties do. An additional cost of governing for an anti-political-establishment party could thus be attributed to the fact that it has generally less governing experience than the other parties rather than to its anti-political-establishment stance.

This rival hypothesis is assessed on the basis of a variable that measures the number of times (thus, a continuous variable) that the party had already gained access to power before the relevant election period. For example, the Northern
League in Italy, that was incumbent in 2006 and had participated in a government coalition one time before (in 1994), is attributed a ‘1’ concerning the election period 2001–2006. This effect is expected to be positive, as the more experienced a party is, the smaller its cost of governing should be. If the H1 variable’s impact disappeared if government experience were accounted for, the alternative hypothesis would be confirmed. This is because my theory implies that an anti-political-establishment party loses more votes on average than other parties, notwithstanding any general extra costs for inexperienced government parties. As it turns out, however, the additional cost of governing is robust to the number of times that a party has held office before. I do not find the predicted positive impact that would indicate that a party’s cost of governing is higher the first time it holds office than the second, and the second time higher than the third time, and so on. More importantly, the impact of the H1 variable is not affected. On the contrary, it produces a stronger effect (b = −3.64), which still is significant at the p < 0.01 level (one-tailed). Thus, the empirical evidence does not support the claim that this effect is due to a lack of government experience. In sum, the results of these sensitivity tests indicate that my findings are robust.

**Discussion**

This article offers a fresh approach to the topic of the cost of governing by focusing on individual parties instead of on governments as a whole. In doing so, I am able to detect differences between different kinds of parties. The article’s findings contribute to the cost of governing literature in three ways. First, in accordance with the existing literature, the empirical evidence supports the idea of a cost of governing (see also, e.g., Strøm 1990; Paldam 1986; Nannestad & Paldam 2002; Powell & Whitten 1993; Rose & Mackie 1983). Second, it has been demonstrated that the cost of ruling varies across party types. Third, it is shown that *anti-political-establishment parties* suffer an additional cost of governing (H1). The intuition behind this is that these parties lose votes insofar as they are perceived to have transformed from being ‘truly’ anti-political-establishment to being ‘part of the political establishment’. The evidence does not support the claim that this effect is due to a lack of government experience. Nor is the additional cost of governing related to the policy compromises that these parties have to make when sharing power (H2). Note, however, that my findings pertain to only anti-political-establishment parties in postwar Western democracies.

The evidently critical role of party traits in these results calls into question many findings in the existing cost of governing literature. Earlier research,
which did not account for the differences in party characteristics, may have led to imprecise estimates of the cost of governing. At least, my findings call for a revision of results reported in earlier studies.

My findings also have implications for coalition-building theory. Coalitions with anti-political-establishment parties are less likely to form than coalitions without such parties (Martin & Stevenson 2001: 46). It is suggested that the reason for this is that established parties lose votes if they enter a government with an anti-political-establishment party. Martin and Stevenson (2001: 37), for example, refer to the theory that ‘because strong social norms exist against admitting parties to government that are not committed to the maintenance of the democratic system, the electoral costs of forming coalitions with such parties are prohibitive’ (see also Geys et al. 2006: 966). My results, however, suggest that the established parties do not have any particular reason to worry about this. It is the anti-political-establishment parties that, on average, suffer greatly from such coalitions and that therefore have an interest in avoiding them. The established parties that bring an anti-political-establishment party in from the cold, by contrast, on average make electoral gains from doing so.

Finally, my conclusions have important implications for the defence of democracy. In light of potential consequences in terms of enacted policies and government style, it might be undesirable to allow anti-political-establishment parties to enter government coalitions. Yet, it is an important finding that, without running any additional electoral risk themselves, established parties can pose a major dilemma for an anti-political-establishment party by inviting it to share government power. Its expected additional cost of governing renders this choice particularly difficult for an anti-political-establishment party. If it turns down the offer, the anti-political-establishment party denies itself a rare opportunity to have its share in the spoils of office and escape its isolated position. If it accepts the invitation, the party is likely to have its teeth pulled by a large electoral cost of governing. So, if the aim is to prevent the anti-political-establishment party from becoming a major electoral threat, the other parties might want to consider allowing it to share government responsibility once it has acquired coalition potential. Rather than trying to defeat it, it seems to be a better strategy to force upon the challenger party a choice between irrelevance in terms of office, and irrelevance in terms of votes.

Acknowledgements

I would like to thank Jørgen Bølstad, Elias Dinas, Cees van der Eijk, Mik Laver, Peter Mair, Sergi Pardos, Steven Poelhekke, Till Weber and three anonymous reviewers as well as the editors of EJPR for their very useful
suggestions. Special thanks to Mark Franklin for his help throughout the entire project. All errors and omissions in the manuscript remain my sole responsibility.

Notes

1. Buelens and Hino (2008: 159) show that the cost of governing varies by party type. However, their analyses do not reveal any clear link between party type and cost of governing among ‘newly governing parties’ – that is, parties that have shared government power between 1950 and 2004 (Buelens & Hino 2008: 163–164).

2. Nannestad and Paldam (2002: 28–29) are sceptical of the coalition of minorities ideas. Because of its link to campaign promises in Mueller’s second idea, they maintain that this idea requires the assumption of a ‘constant sucker fraction’ over time and space – in other words, that a constant share of the electorate can be, and is, fooled in each election wherever, and whenever. Clearly, this assumption is unlikely to be met. However, this does not apply to Mueller’s first-mentioned idea, which is based on Downs (1957: 55–60). It is certainly possible that the government lets down a certain share of the electorate each time it acts, thereby potentially organising its own opposition. This does not require a constant sucker fraction, and is therefore not problematic in this sense.

3. Lewis-Beck (1988: 78–79, 107, 156) does not find strong empirical evidence in support of the existence of grievance asymmetry in Western European countries. Furthermore, whether or not the cost of ruling can be attributed to economic performance remains open to question as it explains only a tiny fraction of the variation in government popularity (Van der Brug et al. 2007). For these and other reasons (see Van der Brug et al. 2007: 16–18), the empirical evidence for the existence of such grievance asymmetry is not beyond doubt.

4. In a similar vein, several studies have found different effects of economic indicators on support for particular German government and opposition parties without explaining this variation (Feld & Kirchgässner 2000; Geys & Vermeir 2008).

5. Stevenson (2002: 177), however, argues that this theory is also applicable to multiparty systems such as in Belgium, Germany and the Netherlands. This is because, although the Christian Democrats dominate these countries’ party systems, they have to ally with either the left or the right to govern. In these countries, a similar mechanism may therefore occur, as centrist voters would want to see alternation between centre-left and centre-right coalitions.

6. As a case in point, the two anti-political-establishment parties for which I have reliable data on left-right placement before and after government participation – the Freedom Party of Austria (FPO) and the Northern League (LN) in Italy – did not become less radical after government participation. On the contrary, data derived from the European Election Studies show that the FPO was positioned at 7.54 on a 1–10 left-right scale before entering a government coalition (in 1999) and at 8.11 after (in 2004), and the LN also radicalised after it entered a government coalition (from 5.32 in 1999 to 7.81 in 2004).

7. Neither right-leaning governments that include radical left parties nor left-leaning coalitions that include radical right parties are likely or empirically observed. Therefore, anti-political-establishment parties are expected to make the most far-reaching compromises in centrist governments.
8. If several governments are formed during one election period, I select the longest serving government.

9. Although the findings by Nannestad and Paldam (2002: 26) do not point in this direction, it could be the case that the average loss government parties suffer has gradually increased since 1945 rather than in an abrupt fashion at the end of the 1980s. Employing an alternative variable, the time elapsed since 1945, does not change any of the conclusions of this article, however. When I re-estimate Model 3 (see Table 3) with this variable instead of the post-1989 one, it does not have a substantial impact. The $H1$ variable yields an effect ($b = -3.25; p = 0.05$, one-tailed) that is highly similar to the one in Model 3, and the $H2$ variable does not have a substantial impact, as in Model 3.

10. The models were rerun with GDP growth in the year before the elections (e.g., Whitten & Palmer 1999) and the average GDP growth since the previous election. In addition, the analyses were performed including unemployment change and inflation data (e.g., Powell & Whitten 1993; Van der Brug et al. 2007), both regarding the election year, the year before, and the entire election period. They were estimated in interaction with the ‘clarity of responsibility’ variable (Powell & Whitten 1993), as well as not in interaction with it. Either way, none of these variables yielded any impact that substantially affected the conclusions presented in this article.

11. The only Dutch case, the List Pim Fortuyn (LPF)’s government access in 2002, has to be left out of the analysis because of the short string of this flash party’s election results (it contested only three elections to the national parliament). As a result, the Netherlands does not constitute a case in my study, which is based on the other seven countries mentioned. As the LPF lost heavily (11.3 percentage points) in the elections immediately following its government participation, excluding this case may have rendered the test of my argument more conservative and therefore increases the confidence in my findings.

12. The number of successive observations for each party should be at least six in order to obtain reliable estimates with the time-series cross-sectional analysis that I use in this article. Estimating the models on the basis of data concerning only the parties for which I have at least ten successive observations, as recommended by Beck (2001: 274), my conclusions do not substantially change (31 parties, $N = 456$). Rerunning Model 3 on these 456 observations, the $H1$ variable has an effect in the predicted direction ($b = -4.85$) significant at the $p = 0.001$ level (one-tailed). The $H2$ variable does not yield any significant impact.

13. Note that I include all parties: those in government as well as those in opposition, anti-political-establishment as well as other parties.

14. The six variables constituting this three-way interaction variable are also included so as to avoid misspecification of the model. The incumbency variable has already been included in the models shown. Thus, I add five variables along with the three-way interaction variable, as is shown in Table 3.

15. For the operationalisation of the original five-point political complexion of the government scale, see Woldendorp et al. (1998: 128). This measure is used in the absence of any more reliable alternative indicator of the ideological positions of all the governments since 1945 that is comparable over time across the seven countries under study.

16. The party-level analyses that included inflation figures were performed on the basis of a reduced data set ($N = 404$) as inflation data were not available for all of my observations.

17. When I perform analyses at the government level, however, I do find significant effects of GDP growth, notwithstanding the much smaller number of observations ($N = 92$). Yet,
including GDP growth, inflation or unemployment change figures, regardless of how they were measured, does not affect my substantive results (see also note 10).

18. It is possible that an anti-political-establishment party joins a government coalition only if it expects to lose at the next election (cf. Strøm 1990). In that case, the effect found would not reveal any mechanism of anti-political-establishment parties losing extra votes, but just capture their anticipated losses at the next election. However, it is not only up to an anti-political-establishment party whether or not it joins a government coalition. Moreover, there have been many instances in the countries under study where anti-political-establishment parties could have entered government coalitions but did not. I can therefore safely attribute the effect found to the voters’ disillusionment with the anti-political-establishment party they voted for before it entered a coalition with one or several established parties.

19. Note that, unlike the cost of ruling for the anti-political-establishment parties, the cost of governing for their coalition partners is not substantially larger than for other parties. A variable identifying these coalition partners yields a positive effect ($b = 1.94$, not shown), which reaches statistical significance at the $p < 0.05$ level (two-tailed). This indicates that, in contrast to a common assumption in the relevant literature (e.g., Martin & Stevenson 2001: 37; Geys et al. 2006: 966), including an anti-political-establishment party in a government coalition does not involve any additional risk for the established parties. In fact, it is, on average, profitable for them to do so.

20. It is possible that voters punish governing parties if their coalition breaks down during the election period, resulting in instability and inefficiency. The additional cost of governing for anti-political-establishment parties could then (partly) be explained by the relatively many early breakdowns of the governments in which these parties participate. However, adding a variable identifying governments that survived the entire election period to Model 3, based on data from Woldendorp et al. (1998) and updates (Katz & Koole 1997, 1999, 2002; Koole & Katz 1998, 2000, 2001; Katz 2003; Van Biezen & Katz, 2004, 2005, 2006; Bale & Van Biezen, 2007, 2008), along with its interaction with the incumbency dummy, neither yields an effect nor substantially affects the impact of the variables of interest. The $H1$ variable retains its effect ($-3.43$, significant for $p = 0.01$, one-tailed), while the $H2$ variable still fails to have an impact.

21. As an exception to this rule, it is possible that an anti-political-establishment party was ousted from such a coalition and became anti-political-establishment after this (again), such as Western European communist parties in 1947–1948 (see Abedi 2004: 32–75).

22. Martin and Stevenson refer to ‘anti-system parties’ (Sartori 1966, 1970, 1976) – a concept upon which the ‘anti-political-establishment party’ concept was based.

23. See note 19.

References


Address for correspondence: Joost van Spanje, Amsterdam School of Communication Research (ASCoR), University of Amsterdam, Kloveniersburgwal 48, 1012 CX Amsterdam, The Netherlands. Tel.: +31 (0) 20 525 4827; Email: j.vanspanje@uva.nl; Webpage: http://home.medewerker.uva.nl/j.h.p.vanspanje