Presidentialism and legislative fragmentation: Beyond coattail effects

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Abstract

Using data from a nearly comprehensive set of the world's electoral democracies, 1992–2014, this article empirically evaluates the impact of presidentialism upon legislative fragmentation. The analysis demonstrates that the impact is strong, consistent across a wide variety of political contexts, and conditioned by the type of presidential regime, the scope of presidential powers, electoral system effects, and essential party system properties. While much of the reasoning regarding the interplay between presidentialism and legislative fragmentation has been traditionally focused on short-term coattail effects of presidential elections, this study shows that these effects are real, but they are insufficient to make a significant impact upon the parameter of crucial importance for the functioning of presidential regimes: the number of parties in the legislature. The main impact of presidentialism is systemic, stemming from its tendency to restrict the number of parties to a limited set of viable competitors for the presidential prize.

Introduction

Legislative fragmentation—a concept that, in its simplest definition, refers to the number of parties in a legislative assembly or in legislative elections—has been found to be of crucial importance for executive–legislative relations in presidential regimes (Hicken and Stoll, 2011; Jones, 1995; Shugart and Carey, 1992) and thereby, for such regimes' short-term and long-term sustainability (Golder, 2006; Mainwaring, 1993; Stepan and Skach, 1993). There is a small but steadily growing body of literature on the effects of presidentialism on legislative fragmentation. In cross-national empirical research, it had been long recognized that presidential regimes had fewer legislative parties (Lijphart, 1994). Building on this finding, several scholars have advanced the argument according to which presidentialism shapes legislative fragmentation primarily by exerting coattail effects,

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with a party's prospects for success at the legislative level being contingent upon this party's prospects in the presidential election (Jones, 1994; Shugart, 1995). Early empirical studies tended to confirm the hypothesis that in this way, presidentialism reduced legislative fragmentation (Amorim Neto and Cox, 1997; Cox, 1997; Mozaffar et al., 2003), even though empirical support to this hypothesis was not entirely consistent, as other studies found that presidential elections either increased party system fragmentation (Filippov et al., 1999) or made no difference at all (Samuels, 2000). More recent research suggests that the reductive effects of presidentialism are real but contingent on the competitiveness of presidential elections (Clark and Golder, 2006; Golder, 2006; Williams-Wyche, 2014) and the scope of presidential powers (Hicken and Stoll, 2008, 2013; Stoll, 2015). Nevertheless, the effects estimated in the literature are rather weak, and some of them remain questionable (Elgie et al., 2014).

This study builds on the existing body of research but differs from it in several important respects. First, we argue that the effects of presidentialism upon legislative fragmentation cannot be reduced to the strategic calculations of elites seeking to solve their short-term coordination problems, as implied by the very notion of coattail effects. In our view, excessive concentration on coattail effects, with the resulting emphasis on the temporal proximity between presidential and legislative elections, tends to overshadow more fundamental causal relationships between presidentialism and party system properties. Second, while fully embracing the idea that these relationships are contingent upon the scope of presidential powers, we introduce the distinction between presidential and semi-presidential regimes as an additional factor. Third, in contrast to the bulk of previous research, this study draws a distinction between legislative fragmentation in the narrow sense of the term, that is, the number of parties in the legislature, and fragmentation in the electorate. Fourth, we extend the empirical scope of inquiry to a nearly comprehensive set of the world's electoral democracies embracing 130 country cases. On this basis, we empirically reassess the impact of presidentialism upon legislative fragmentation.

The study proceeds as follows. In the first section, we theoretically explain and refine the above-stated propositions and on this basis, develop a set of working hypotheses for empirical validation. The second section operationalizes the variables. The third section provides model specification and introduces the data for empirical analysis. In the final section, we report the findings.

Theory and working hypotheses

In his influential study of the relationship between presidentialism and legislative fragmentation, Golder (2006: 36) notices that there are 'two slightly different variants of this coattails story'. One of them, dubbed the 'short coattail hypothesis', reduces the impact of presidentialism to the temporal proximity of presidential and legislative elections, adding the effective number of presidential candidates as an important intermediary factor. The other, dubbed the 'long coattail hypothesis', argues that the effective number of electoral parties (ENEP) is conditioned by presidentialism irrespective of whether presidential and legislative elections are temporally proximate. When justifying this hypothesis, Golder (2006: 36) briefly notes that long presidential coattails 'indicate a highly institutionalized party system where the benefits a party receives from being associated with a particular presidential candidate persist long after a presidential election has taken place'. In the subsequent empirical analysis, he finds some evidence in favour of the long coattail hypothesis, but the short coattail hypothesis is better supported by his study.

There are, however, stronger theoretical reasons for expecting a negative relationship between presidentialism and legislative fragmentation. First and foremost, this expectation can be justified with reference to the famous proposition or 'law' of Duverger (1954: 127), according to which the plurality single-ballot electoral system favours the two-party system. Starting with the pioneering work of Rae (1967), followed by contributions of Shugart and Carey (1992), Lijphart (1994) and many others, this stream of research identifies district magnitude, defined as the number of seats to be allocated in the electoral district, as the principal determinant of party system fragmentation. By definition, the presidency is a single-member office, and therefore presidential elections, provided that they exert a systemic impact on the patterns of inter-party competition, can be expected to reduce the number of parties. It is true that many presidential elections are held by two-round majority systems that, consistent with the main tenets of the Duvergerian approach and as empirically shown by Golder (2006), mitigates the reductive effects of presidentialism. However, the available literature suggests that while allowing for some additional fragmentation at the electoral level (Birch, 2003), two-round majority systems, by setting an extremely high effective threshold of representation in the second round of voting, ultimately reduce legislative fragmentation (Lijphart, 1994). Second, presidentialism can be expected to have an impact on another fundamental property of party systems, nationalization, conventionally defined as the extent to which national parties receive similar vote shares across the territorial units of the state (Jones and Mainwaring, 2003). According to this line of reasoning, presidential elections naturally suppress local sentiment in the electorate by appealing to national political forces (Hicken and Stoll, 2011; Shugart and Carey, 1992). In this way, presidentialism undermines the electoral fortunes of small parties that owe their survival to the territorial bases of support. In the long run, this reduces the overall level of party system fragmentation. While the available empirical support to this theory is not entirely consistent (Brancati, 2008; Morgenstern et al., 2009), it is convincing enough to deserve one more empirical trial.

Our approach does not differ from the short coattail hypothesis in identifying the fundamental causes of legislative fragmentation in presidential regimes. However, the classic formulations of the short coattail hypothesis place major emphasis on the immediate consequences of the reduction in fragmentation that occurs in presidential elections, whereas we theorize that these consequences are secondary to the long-term effects of presidentialism. In other words, presidentialism reduces party system fragmentation at the systemic level. This theory, however, has to be qualified in several ways. First, the proposition according to which the effects of presidentialism upon legislative fragmentation are conditioned by the scope of presidential powers has a strong intuitive appeal, as it is obvious that the size of the presidential prize is highly consequential for the competitive strategies of political actors, which has clear implications for party system format. From the side of the electorate, this line of reasoning is consistent with the well-grounded general theory that postulates the sensitivity of voters to patterns of political accountability (Duch and Stevenson, 2008). Several empirical tests (Hicken and Stoll, 2008, 2013; Stoll, 2015) have provided some support to this theory, albeit not without important refinements. In particular, Elgie et al. (2014) show that the way in which the number of presidential contenders shapes the legislative party system can only be predicted for an intermediate range of presidential powers, and propose a complex three-way interaction model of the related effects. Such refinements can and should be taken into account in empirical analysis. Nevertheless, there are solid theoretical reasons to hypothesize that as a general tendency,

H1. Stronger presidency is associated with greater reduction in legislative fragmentation.

The scope of presidential powers is a complex concept, and it can be operationally defined in several ways. When formulating the previous hypothesis, we referred to the conventional measures based on the coding of president's legislative and non-legislative powers along many dimensions, from veto/override to cabinet formation, as originally proposed by Shugart and Carey (1992). However, there is a related but different approach that makes a qualitative distinction between two types of presidential regimes, presidentialism in the narrow sense of the word and semi-presidentialism. First proposed by Duverger (1980) and later adopted and developed by several other scholars (Elgie, 2011; Shugart, 2005; Shugart and Carey, 1992), the concept of semi-presidentialism conventionally refers to a political regime in which there is a popularly elected fixed-term president and a prime minister and cabinet who are collectively responsible to the legislature. Further distinctions within this broad category can be made. The resulting sub-categories, such as president–parliamentary and premier–presidential regimes, have been productively employed in different streams of research. For the purposes of this analysis, however, it is necessary to make a primary distinction between presidentialism and semi-presidentialism on the basis of a parsimonious definition, as formulated above. While it is certainly true that to a certain extent, the standard measures of presidential powers do capture a distinction between the two types of presidential regimes, they also register the existence of rather weak presidential regimes, as well as semi-presidential regimes so strong that they are sometimes referred to as super-presidential (Ishiyama and Kennedy, 2001), on the one hand, and so weak that for most observers, they are nearly indistinguishable from parliamentary regimes, on the other hand. Thus, presidential power scores do not sufficiently discriminate between presidentialism and semi-presidentialism. More importantly, from the substantive perspective of this study, semi-presidential regimes differ from 'pure' presidentialism in terms of the structure of incentives for suppressing legislative fragmentation. While it is conventional to think of presidential powers with reference to the size of presidential prize, it is possible to argue that low legislative fragmentation is a prize in itself for the president, and the size of this prize varies depending on the type of presidential regimes.

In 'pure' presidential regimes, the loss of legislative majority is painful for the president, as he or she loses legislative support for the implementation of the stated policy objectives, but it does not prevent him or her from making cabinet appointment and thereby, from running the business of government in usual. In semi-presidential regimes, however, this situation is much more painful for the president, as the president either yields importance to the prime minister trusted by the legislative majority or engages in prolonged and conflict-ridden bargaining with the legislature regarding the candidacy of the prime minister and the composition of government. Thus, the prize of legislative majority under semi-presidentialism is much more valuable than in 'pure' presidential regimes. In the analysis below, we use the term 'institutionally defined presidentialism' to differentiate between 'pure' presidential government and semi-presidential government. The related hypothesis is, thus,

H2. Semi-presidentialism is conducive to lower levels of legislative fragmentation than 'pure' presidentialism. Therefore, if the scope of presidential powers is controlled for, there is a positive relationship between institutionally defined presidentialism and legislative fragmentation.

The remaining hypotheses of this study are complementary. Their purpose is to bring more precision into our empirical specifications. One of these hypotheses is based on the idea that the reductive impact of presidentialism can be realistically expected to occur only with modest numbers of competitors in presidential elections. This idea is very self-evident and therefore persistent in the literature. It has been found correct in several empirical studies (Clark and Golder, 2006; Golder, 2006; Williams-Wyche, 2014). It is important to note, however, that as this theory relates the level of fragmentation in presidential elections to the number of parties in legislative elections, the impact can be empirically established only if the latter aspect is also taken into account. We expect the discrepancy between the numbers of parties in presidential and legislative elections, further referred to as asymmetry, to be consequential for the resulting levels of legislative fragmentation. In particular, this effect is expected to happen because large discrepancies between the numbers of parties in presidential and legislative elections often serve as an indicator of under-institutionalized party systems, a factor that is important for the theory of coattail effects as presented in the recent literature (Golder, 2006), even though the impact of this factor has never been assessed empirically. Therefore, we hypothesize that

> *H3*. Asymmetry mitigates the reductive effect of presidential power upon legislative fragmentation.

The above-stated hypotheses are all built on our theoretical expectation that the impact of presidentialism upon legislative fragmentation is systemic. As such, it cannot be reduced to the short-term influences as epitomized in the notion of coattail effects. It has to be recognized, however, that the case for coattail effects is not only theoretically strong but also supported by some of the previous empirical research. Therefore, it has to be entered into the theoretical explanatory model and tested empirically. Since the topic is very well developed in the literature, building the related hypotheses is not problematic. It can be expected that temporal proximity between presidential and legislative elections affects legislative fragmentation. Hence,

> *H4*. If the number of important competitors in presidential elections is low, then the number of important parties in legislative elections goes down.

> *H5*. Concurrent or otherwise temporally close presidential elections contribute to lower levels of fragmentation.

Of course, it would be utterly unrealistic to assume that presidentialism is the primary determinant of legislative fragmentation. Therefore, any empirical model based on the related theory should include variables accountable for different, presumably very strong, factors. There is a large stream of research that explains party system fragmentation as a product of complex interactions between two groups of factors, sociological and institutional (Amorim Neto and Cox, 1997; Coppedge, 1997; Ferrara, 2011; Ordeshook and Shvetsova, 1994; Powell, 1982). Within the latter group, a prominent role is traditionally ascribed to electoral systems. As already discussed above, this literature identifies district magnitude as the strongest institutional determinant of legislative fragmentation. Hence,

> *H6*. The greater the district magnitude, greater the number of legislative parties.

The mainstream literature on the political consequences of electoral rules, while finding that the 'law' of Duverger, and therefore district magnitude as an explanatory factor, worked very well at the level of individual districts, at the same time revealed that it did not necessarily account for variations in the fragmentation of party systems at the national level. When explaining the exceptional cases, many scholars, including Rae (1967), Katz (1980), Powell (1982), Riker (1982) and Cox (1997) identified the cross-district heterogeneity of the vote as a primary limitation of the validity of the 'law' of Duverger. Otherwise, single-member plurality systems can be compatible with very high levels of national party system fragmentation, which happens primarily due to the ability of small regionally based parties to obtain representation (Meguid, 2008; Norris, 2004). Therefore, the level of territorial homogeneity of the vote has to be accounted for when testing the impact of presidentialism upon legislative fragmentation:

> *H7*. The greater the level of territorial homogeneity of the vote, smaller the number of legislative parties.

Finally, there is one societal determinant of legislative fragmentation that has been, traditionally, in line with the foundational argument of Lipset and Rokkan (1967), employed as a major explanatory variable both in general inquiries (Amorim Neto and Cox, 1997; Ordeshook and Shvetsova, 1994) and in studies focused on the effects of presidentialism (Golder, 2006). These studies convincingly demonstrate that ethnic diversity is positively related to legislative fragmentation. It is also traditional to link this explanatory variable to one of the factors discussed above, district magnitude, normally by entering an interactive term into multiple regression analysis. Based on the massive empirical evidence presented in this literature, we hypothesize that

> *H8*. The positive impact of ethno-linguistic diversity upon legislative fragmentation will be particularly strong if the territorial homogeneity of the vote is low, and it will be mitigated otherwise.

Of course, the voluminous literature on the determinants of multipartism identifies many other causal relationships. In this study, however, we intend merely to test the impact of presidentialism on legislative fragmentation under a set of controls that can be expected to bear particularly heavily on the dependent variable, not to provide a comprehensive evaluation of the causes of cross-national variation in legislative fragmentation.

The variables

This study operationalizes its key dependent variable, legislative fragmentation, as the effective number of legislative parties (*ENLP*). We employ the mathematical definition of the effective number of parties as proposed by Golosov (2010)

$$ENLP = \sum_{1}^{x} \frac{p_i}{p_i + p_1^2 - p_i^2}$$

where sigma stands for summation; x, for the raw number of parties, and p_i and p_1, for the fractional shares of the national votes or seats received by the ith and the largest parties,

respectively. The same formula was used for determining the *ENEP* and the effective number of presidential parties (*ENPP*). The reason why this definition is preferred to a more conventional formula of Laakso and Taagepera (1979) is that, while being very close to it in the conceptual sense, the new index poses an advantage that is important for this study: its values register greater distances between the cases of very low and very high fragmentation, which is particularly useful if the empirical sample includes such cases in sufficient numbers. It has to be mentioned that an analysis that used the Laakso–Taagepera index, not reported here, produced results that were different mostly in the numerical values of the coefficients. When computing the indices of electoral and legislative fragmentation, we systematically counted independent candidates as individual parties. All votes cast for different candidates belonging to the same parties were aggregated. The presence of right skewness in the distribution of our both dependent variables, ENEP and ENLP, makes it difficult to meet the ordinary least squares (OLS) basic assumptions and weakens our subsequent findings. Therefore, we decided to resort to natural log transformation of both dependent variables.

For the operationalization of Presidential Power (PP), we employ the composite index of presidential powers developed by Doyle and Elgie (2016) on the basis of many measures previously proposed in the literature. The values of the index are normalized Prespow1 scores as published at the Presidential Power website of Elgie (2015). Presidentialism (PS) is operationally defined as a dichotomous variable, with 1 indicating the presence of 'pure' presidentialism and 0 indicating absence. The concept of asymmetry (*AS*) between the structures of competition in presidential and legislative elections is operationalized as absolute difference

$$AS = |\, ENEP - ENPP \,|$$

where *ENEP* and *ENPP* stand for the effective numbers of parties in legislative elections and the most recent or concurrent presidential elections, respectively. The reason for our preference of absolute difference over raw difference is that while the number of competitors in presidential elections is often smaller than the number of parties in legislative contests, this is not necessarily the case, and we expect that any kind of discrepancy, by pointing to the under-institutionalization of a party system in accordance with the above-cited argument of Golder (2006), will be associated with additional incentives for party proliferation.

In our operationalization of short-term coattail effects (*CE*), we fully rely on the previous literature and simply follow Golder (2006) by defining it as the standard index of temporal proximity between presidential and legislative elections, as developed by Amorim Neto and Cox (1997)

$$CE = \left(2 \times \left| \frac{L_t - P_{t-1}}{P_{t+1} - P_{t-1}} \right| - \frac{1}{2} \right)$$

where L_t 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 operationalization of district magnitude is scarcely problematic. It is conventional to define it as the mean over all district magnitudes in a logged form, $\log(DM) = \log(N_s/N_d)$, where *DM* is average district magnitude, N_s is the number of seats and N_d is the number

of electoral districts (Lowery et al., 2010). In most electoral systems, the procedure of averaging does not pose any difficulty. However, some of them deserve special attention. For mixed superposition systems (Massicotte and Blais, 1999), we single out their proportional sections and apply the above formula only to the respective numbers of seats and electoral districts. This operational definition is extended to the Hungarian electoral system of 1990–2010 with its very partial compensation (Ferrara et al., 2005). For mixed systems with complete or nearly complete compensation, such as in Germany and New Zealand, N_s is defined as the overall number of seats, and N_d as the number of districts in the proportional section of elections. For multitier proportional systems with double vote, such as in Guatemala and Nicaragua, only the sub-national district tiers are taken into account. Note that our operational definitions of legislative fragmentation and district magnitude are consistent in the sense that the values of these variables have been established on the basis of the same sets of seats.

The territorial homogeneity of the vote (*TH*) is operationally defined as the index of party system nationalization (Golosov, 2016). The index is derived from the Herfindahl–Hirschman index of concentration and thereby, from the standard deviation (Feld and Grofman, 2007), a measure with very solid theoretical credentials. The formula is thus

$$TH = \sum_{1}^{n} \left(\left(1 - \frac{n - \left(\left(\sum_{1}^{n} s_i \right)^2 \Big/ \sum_{1}^{n} s_i^2 \right)}{n-1} \right) p_i \right)$$

where sigma stands for summation, *n* stands for the number of electoral districts, s_i stands for the percentage or fractional share of the vote received by the *i*th party in each of the electoral districts, and p_i, for the fractional share of the vote received by the *i*th party nationally. For nationwide districts, the index is always defined as 1. Of course, this bears no substantive implications for our understanding of the levels of party system nationalization in the respective countries, indicating only the fact that in nationwide districts, the territorial homogeneity of the vote is inconsequential for seat allocation. Indeed, the reason why we prefer the term 'territorial homogeneity of the vote' to the term 'party system nationalization' is that we seek to avoid unnecessary references to the territorial penetration of political parties, an aspect certainly beyond the scope of this analysis. The operational definition of ethnic fractionalization used in this study is derived from Alesina et al. (2003). In the recent years, the Alesina index has become conventional in research on the effects of presidentialism.

Model specification and the data

As explained above, we view presidentialism from two important perspectives: the narrow perspective treating presidentialism as a certain type of political regime, and the broad perspective seeing it as a combination of the main and interaction effects of our two key variables: presidential power *(PP)* and institutional presidentialism (*PS*). Specifically, we are not merely interested in presidentialism in a narrow 'institutional' sense, but rather in a broader sense regarding it also as a function of the presidential power. As a result, our key construct of interest can be defined using the combination of main and interaction

terms of our key variables: $b_1 PP + b_2 PS + b_9 PP \times PS$. We construct our model with inclusion of this central to our analysis combination of main and interactive effects.

In the majority of previous studies, the notions of fragmentation in the electorate and in the legislature are implicitly assumed to be identical. It is, however, quite obvious that they are not. In a perfectly proportional electoral system, there would be no discrepancy between the levels of electoral fragmentation and legislative fragmentation. But no real-life electoral system is perfectly proportional, which means that there is always a discrepancy. Thus, the impact of presidentialism upon party system fragmentation can be assessed by placing it into two different analytical frameworks. However, both approaches are consistent with the model that, using the same notation as introduced in the previous section, writes as follows

$$\log(ENP) = b_0 + b_1 PP + b_2 PS + b_3 AS + b_4 CE + b_5 ENPP + b_6 \log(DM) + b_7 EF +$$
$$b_8 TH + b_9 PP \times PS + b_{10} EF \times TH + b_{11} PP \times AS + b_{12} ENPP \times CE + e$$

where $\log(ENP)$ is the natural logarithm of effective number of electoral/legislative parties, PP is the presidential power, PS is the institutional presidentialism, AS is the asymmetry, CE is the coattail effects, $ENPP$ is the effective number of presidential parties/candidates, $\log(DM)$ is the natural logarithm of district magnitude, EF is the ethnic fractionalization index, TH is the territorial homogeneity, and $PP \times PS$, $EF \times TH$, $PP \times AS$, $ENPP \times CE$ are the various interaction effects between the variables of interest, serving the purpose of hypothesis testing.

When building the set of observations for statistical analysis, we sought primarily to maximize its diversity by including as many country cases as possible. At the first step, the pool of potentially available cases was provided by those 144 independent countries that were assigned by Freedom House (2015) to the category of electoral democracies for at least 1 year within the chronological limits of our inquiry, 1992 through 2014. Due to the unavailability of data on earlier elections, the Comoros are represented by elections held in 2015. The lower chronological limit is defined by the collapse of the Soviet Union, an event that was highly consequential for many regions of the world. The chronological limits of inquiry are rather wide, but this choice can be justified with reference to the fact that with more narrow limits—confined, for instance, to the recent decade—it would be necessary to exclude a large group of countries that ceased to be electoral democracies in the 2000s or later.

In fact, even among the pre-selected 144 cases, not all are fit for this study. Four countries did not hold any legislative elections during their democratic periods, normally very brief. Six countries, all of them in the Pacific, conducted their elections on a non-partisan basis. Upon the exclusion of these cases, we had to additionally exclude four more countries, all of them in Francophone Africa, due to the lack of sufficient electoral data. We considered the completeness of the electoral data—including the lack of the 'others' category and the availability of sub-national electoral returns for the estimation of the territorial homogeneity of the vote—as essential criteria for inclusion. One country case, Maldives, was excluded from some of the models because of the lack of information on ethnic fractionalization. It has to be mentioned that our experimentation with incomplete data on thus excluded cases, not reported below, demonstrated that with them, the statistical models would have become even stronger because of the increased explanatory power of the semi-presidentialism variable. All in all, with 129 of the theoretically available 134

country cases included, it would be fair to say that the set of observations is nearly identical to the general population of the world's electoral democracies.

Each of the 129 countries enters the empirical sample with one election to the single or lower chamber of its national legislature, invariably held during the respective countries' democratic periods. In our view, many difficulties in the previous research on party system properties stemmed from the disproportionate presence of long-established democracies in empirical samples, which is unavoidable if some of the countries enter the dataset with more than one election. Of course, one way to deal with this problem is to employ complex multilevel modelling. However, the lack of variation in the institutional variables due to institutional rigidity can prevent us from building models with complex variance structures. An alternative solution is simply to avoid a multilevel structure in the data, which explains our strong preference for having each country represented by one observation only. It is however clear that this methodological choice invites concerns about the possible influence of idiosyncrasies in particular elections upon the overall results of statistical analysis. Thus, in those cases when sufficient information on more than one elections was available, we selected those of them that could be viewed as typical, which was established by comparing the ENLP in the given election with the 'systemic effective number of parties' (Golosov, 2013). Maximum correspondence between the two parameters qualified for inclusion. The list of country cases and data sources is provided in Online Appendix A.

The findings

Before turning to discussion of our findings from the multivariate regression analysis, we apply a non-parametric regression between the measures of presidential power and effective number of electoral and legislative parties with exclusion of only one case that strongly affected our regression results, Comoros 2015 (Figure 1, see country codes in Online Appendix A). The reference model has been specified as linear. All our findings with respect to presidential regimes are well predicted by the theory, yielding statistically significant effect of presidential power on fragmentation for both graphs at $\alpha = 0.05$ and $\alpha = 0.1$ levels. Since the sample size is small and there is no substantive justification for which outliers ought to be excluded, in our further analysis we decided to keep all of the observations in our multivariate regression analysis and account for any possible outliers by reporting cluster robust standard errors. However, due to substantive considerations, we did perform several robustness tests. First, we replicated the analysis for a reduced set of countries that qualified as electoral democracies for a minimum of five consecutive years between 1992 and 2014, thus isolating the possible impact of short-lived and uncertain democracies in general and of those cases where either presidential or legislative elections had been held in non-democratic conditions. Second, we replicated the analysis for a reduced set of countries that does not include micro-states, operationally defined as countries with populations of less than 100,000, because such countries have been omitted from several previous studies. The results of the tests are reported in Online Appendix B (Tables B1 to B4). They are not significantly different from those reported below.

Let us now proceed to the multivariate regression analysis. For each presented model with different definitions of the dependent variable, we report cluster robust standard errors. The results of the analysis that uses the effective number of parties in the electorate and legislature as the dependent variable are reported in Tables 1 and 2, respectively. To evaluate the robustness of our findings, we also resort to estimation of models with

Figure 1. The effect of presidential power on effective number of parties.

different sets of included covariates as well as perform our analysis on the different data subsets: the whole dataset with the parliamentary regimes included and the restricted one containing only presidential/semi-presidential regimes. As a result, we estimate models M(01)–M(03) on the whole dataset and models M(04)–M(06) for the data containing only presidential/semi-presidential regimes.

In Table 1, all the models illustrate remarkably high percentage of explained variance reaching in its maximum 85% for the whole dataset and 87% for the subsample of our data. The central hypothesis of interest with regard to effects of presidential strength is supported by our findings: a strong and statistically significant negative effect of the presidential power on the effective number of parties in the electorate holds across different specifications of the models. For instance, according to M(01), one point increase in

Table 1. Factors of effective number of parties in the electorate.

	M(01)	M(02)	M(03)	M(04)	M(05)	M(06)
Constant	1.55***	1.05***	0.71***	1.12***	0.59***	0.52***
	(0.15)	(0.13)	(0.04)	(0.07)	(0.14)	(0.08)
Presidential power (PP)	−0.36*	−0.4***	−0.49***	−0.37***	−0.39**	−0.3*
	(0.21)	(0.15)	(0.11)	(0.18)	(0.16)	(0.17)
Presidentialism (PS)	0.12**	0.1	−0.06	0.15***	0.08	0.00
	(0.06)	(0.08)	(0.07)	(0.05)	(0.07)	(0.09)
Asymmetry (AS)	0.15***	0.15***	0.18***	0.15***	0.16***	0.19***
	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)	(0.02)
Coattail effects (CE)	−0.02	0.00	−0.15	−0.09***	−0.05**	0.00
	(0.08)	(0.07)	(0.21)	(0.04)	(0.02)	(0.19)
Presidential Parties (ENPP)	0.07***	0.08***	0.09***	0.14***	0.14***	0.15***
	(0.01)	(0.01)	(0.01)	(0.05)	(0.05)	(0.00)
District magnitude (DM)	0.11***	0.1***	0.06***	0.1***	0.09***	0.05***
	(0.01)	(0.01)	(0.02)	(0.01)	(0.01)	(0.01)
Ethnic fractionalization (EF)	−0.2**	0.86***		−0.01	1.01***	
	(0.09)	(0.19)		(0.05)	(0.21)	
Territ. homogeneity (TH)	−0.99***	−0.34**		−0.68***	−0.02	
	(0.16)	(0.17)		(0.08)	(0.11)	
PP × PS		0.00	0.54***		0.11	0.33
		(0.23)	(0.12)		(0.13)	(0.22)
EF × TH		−1.39***			−1.41***	
		(0.36)			(0.23)	
PP × AS		0.01	−0.02		−0.01	−0.04
		(0.01)	(0.04)		(0.03)	(0.04)
ENPP × CE			0.06			0.00
			(0.08)			(0.07)
R^2	0.84	0.85	0.79	0.86	0.87	0.84
Obs.	129	129	130	69	69	70

Source: See Online Appendix A.
ENPP: effective number of presidential parties.
Cluster robust standard errors in parentheses. Models: M(01)–M(03)—the models estimated for the whole dataset; M(04)–M(06)—the models estimated only for the presidential regimes.
Significance levels: *$p \leqslant 0.1$; **$p \leqslant 0.05$; ***$p \leqslant 0.01$.

the index of presidential power contributes to about 43% ($e^{0.36}$) decrease in the effective number of parties as suggested by our theory. However, contrary to our theoretical expectations, the main effect of the presidentialism, understood as a type of government, while failing to show statistically significant results, remains consistently positive and statistically significant in the models M(01) and M(04). For instance, the mere country's belonging to presidentialism compared to semi-presidentialism increases its electoral fragmentation by about 16% ($e^{0.15}$), *ceteris paribus*. Interestingly though, after inclusion of interaction effects between presidentialism and presidential power, the significance of the presidentialism's main effect disappears raising, well-grounded suspicions that these effects can be partly moderated by presidential strength: one interaction effect between presidential power and presidentialism is statistically significant in model M(03). According to model M(03), while the presidential power in the states with 'pure'

Table 2. Factors of effective number of parties in the legislature.

	M(01)	M(02)	M(03)	M(04)	M(05)	M(06)
Constant	1.29***	0.67***	0.48***	1.28***	0.59***	0.44***
	(0.22)	(0.16)	(0.03)	(0.24)	(0.17)	(0.07)
Presidential power (PP)	−0.6***	−0.56***	−0.71***	−0.67***	−0.6**	−0.56**
	(0.18)	(0.06)	(0.04)	(0.19)	(0.29)	(0.28)
Presidentialism (PS)	0.25***	0.19***	0.02	0.24***	0.17**	0.07
	(0.02)	(0.06)	(0.04)	(0.05)	(0.07)	(0.05)
Asymmetry (AS)	0.11***	0.13***	0.16***	0.11***	0.13***	0.16***
	(0.01)	(0.01)	(0.02)	(0.02)	(0.01)	(0.03)
Coattail effects (CE)	−0.02	−0.02	0.04	−0.06	−0.02	0.12
	(0.12)	(0.08)	(0.19)	(0.06)	(0.03)	(0.32)
Presidential Parties (ENPP)	0.05***	0.05***	0.07***	0.05	0.05	0.1
	(0.01)	(0.01)	(0.01)	(0.05)	(0.05)	(0.07)
District magnitude (DM)	0.16***	0.15***	0.11***	0.14***	0.13***	0.08***
	(0.00)	(0.01)	(0.02)	(0.01)	(0.02)	(0.01)
Ethnic fractionalization (EF)	−0.13	1.18***		−0.07	1.23***	
	(0.13)	(0.24)		(0.14)	(0.1)	
Territ. homogeneity (TH)	−0.95***	−0.16		−0.85***	0.00	
	(0.22)	(0.15)		(0.28)	(0.24)	
PP × PS		0.07	0.6***		0.07	0.35*
		(0.16)	(0.1)		(0.13)	(0.19)
EF × TH		−1.74***			−1.81***	
		(0.36)			(0.25)	
PP × AS		−0.03	−0.05		−0.02	−0.05
		(0.02)	(0.06)		(0.05)	(0.07)
ENPP × CE			−0.01			−0.04
			(0.08)			(0.15)
R^2	0.72	0.75	0.68	0.7	0.75	0.68
Obs.	129	129	130	69	69	70

Source: See Online Appendix A.
ENPP: effective number of presidential parties.
Cluster robust standard errors in parentheses. Models: M(01)–M(03)—the models estimated for the whole dataset; M(04)–M(06)—the models estimated only for the presidential regimes.
Significance levels: *$p \leqslant 0.1$; **$p \leqslant 0.05$; ***$p \leqslant 0.01$.

presidentialism contributes to a roughly 5% increase in electoral fragmentation compared to other political systems, which is statistically insignificant, in the semi-presidential regimes the effect is negative and statistically significant contributing to 63% ($e^{0.49}$) decrease in electoral fragmentation. Since presidentialism as a broad category accounts for both institutional arrangements and degree of presidential power, we need to apply the Wald test, which enables us to test the joint null hypothesis, in which none of the composing coefficients can be simultaneously equal to zero. Since the probability of the null hypothesis of Wald test is $p < 0.05$, our findings for model M(03) are conclusive about the presence of a strong statistically significant effect of presidentialism on electoral fragmentation.

In Table 1, M(02), one can also observe a strong and statistically significant effect of asymmetry on electoral fragmentation: the district's increase by 10% contributes to about

1.4% ($1.10^{0.15}$) rise in the level of electoral fragmentation. In similar vein, all the models show both the ENPP and logged average district magnitude being positive and statistically significant in their main effects. Likewise, TH poses a consistently negative effect on our dependent variable, for instance, contributing to an almost two and a half times drop ($e^{0.99}$) in electoral fragmentation in model M(01). Moreover, in this model, the effect of ethnic fractionalization index on electoral fragmentation is negative: an increase of the ethnic fractionalization by one unit contributes to a decrease in the electoral number of parties by 0.22% ($e^{0.20}$). As expected in theory, the effect of ethnic fractionalization on the effective number of parties in model M(05) seems to be moderated by territorial homogeneity: in the regimes with small territorial homogeneity, we observe a positive effect of ethnic fractionalization on the number of parties, whereas in the regimes with high territorial homogeneity the opposite effect is observed (see Figure B1 in Online Appendix B).

Importantly, in models M(04) and M(05) the main effects for short-term coattail effects are statistically significant and negative: as coattail effects increase by one unit, electoral fragmentation drops down by 10% ($e^{0.09}$), keeping the other variables constant.

Table 2 illustrates our findings with regard to effective number of parties in the legislature, being generally similar to our results discussed in Table 1. Again, the effect of our key variable, presidential power, is consistently strong in magnitude, negative in its sign and statistically significant across all the models. Similarly to Table 1, the effect of 'pure' presidentialism on the effective number of parties in the legislature is positive in models M(01) and M(04). Also, in M(03) similarly to M(02) the marginal effect of semi-presidential compared to 'pure' presidential regimes on the effective number of parties in the legislature is much more negative in magnitude, that is, reducing effective numbers of electoral/legislative parties by half ($e^{0.71}$) and 12% ($e^{0.11}$), respectively.

Our findings with regard to other explanatory variables show patterns that are almost similar to Table 1 for asymmetry, logged average district magnitude, territorial homogeneity and ethnic fractionalization.

And here again, the importance of coattail effects seems to be overly exaggerated with none of the models yielding statistically significant results for this term. In substantive terms, this means that while short-term coattail effects can exert some impact upon the effective number of parties in electoral dimension, they are not strong enough to serve as a significant determinant of fragmentation in the elected legislature, especially, if compared to the effects of presidential power. At this level, the reductive effect of broadly defined presidentialism is exerted primarily through its systemic impact upon the structure of party competition.

Thus, our central hypotheses are partly confirmed. While a stronger presidency is associated with a decreasing level of fragmentation (H1), a regime's institutional affiliation with presidentialism is, on the contrary, conducive to an increased level of electoral and legislative fragmentation (H2). Our minor hypotheses are also partly supported by our analysis. Specifically, Hypothesis H3 is only partially confirmed: the asymmetry's main effect is positive, but in a majority of the models its interaction with the presidential power fails to show any statistical significance. Our findings are also in line with expectations for H4: the number of important competitors in presidential elections, indeed, is positively correlates with the number of important parties in legislative election. Unsurprisingly, the positive effect of district magnitude on fragmentation, as stated in Hypothesis H6, is also supported by our empirical findings. Hypothesis H7 predicting the negative effect of territorial homogeneity on fragmentation is confirmed
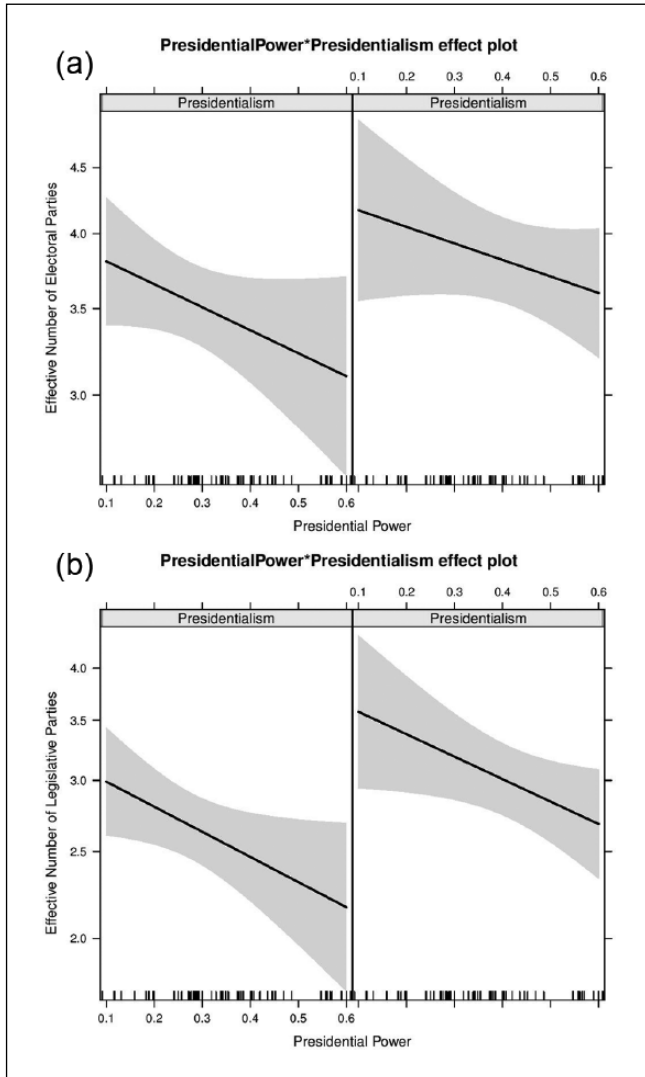
Figure 2. Effects of Presidential Power on Effective Number of Parties for Semi-presidential (left) and 'Pure' presidential (right) regimes.

by our data analysis. Moreover, our claim in H8 about how the effect of ethnic fractionalization on ENP has to be conditioned on territorial homogeneity is strongly supported by our findings.

In the next stage, we resort to visualization of effects of presidential power on electoral fragmentation for both presidential and semi-presidential regimes using *effects* package in R. The primary motivation for this choice is the difficulty of interpreting results with included interaction term, $PP \times PS$, and the importance of disentangling the effects between both types of regimes. The panel at the top left of Figure 2 displays a strong negative impact of presidential power on electoral fragmentation for semi-presidential regimes, while at the top right figure the milder negative effect is detected for 'pure' presidential

regimes. Our Hypothesis H2 agrees with our findings: the average electoral fragmentation in semi-presidential regimes is smaller compared to 'pure' presidential regimes. The mean effects in both graphs are accompanied with 90% confidence intervals, which only partially overlap across the graphs, therefore demonstrating that the differences between both types of regimes can be statistically significant for the medium-range values of presidential power. Both panels at the bottom of Figure 2 illustrate similarly strong negative effects of presidential strength on legislative fragmentation. However, both graphs without observed overlapping confidence intervals assert the presence of differences between presidential and semi-presidential regimes for all values of presidential strength. Therefore, our Hypothesis H2 is also confirmed: semi-presidentialism most likely leads to lower levels of fragmentation than it is the case with 'pure' presidentialism.

Panel data analysis

We admit that since our data are characterized by a degenerate temporal dimension, this restricts us from controlling for country-specific heterogeneity and can seriously bias our estimates of interest. This problem could be easily solved by the panel data analysis. Therefore, in order to check for robustness of our findings, we additionally build a small panel dataset including the previously described presidential power index, ENEP from Constituency-Level Elections Archive (CLEA, 2014), the presidentialism variable derived from the Database of Political Institutions (2012) as well as ethnic fractionalization index and district magnitude. The ENEP in a country's party system is calculated at the national level following Laakso and Taagepera's (1979) specification. For simplicity reasons and various data-building issues, unfortunately, we were not able to include other covariates in our models. However, we think that even this limited set of alternative empirical data would be enough to implement an additional robustness check of our findings. As a result of merging all our data together, we obtained unbalanced panel data for 226 unique observations, 43 countries and time period ranging from 1977 to 2011. All panel data analysis has been implemented using the *plm* package in *R*.

Table 3 presents the results from three separate models: pooled time-series cross-sectional, fixed-effects and random-effects models with only additive and both additive/interactive effects included in each case. We also estimate robust standard errors for all our models. Since the results of the pooled OLS analysis can be biased, to account for this possibility we also utilize two alternative approaches. On the one hand, we use the fixed-effect models, which enable us to control for time-invariant unobservables for each of the countries. On the other hand, we use a random-effects model resting on the normality assumption, and implying that the effects are drawn from the normal distribution of some larger population. The results from the Hausman test allow us to differentiate between both approaches and find out which method is most appropriate to use. Due to the lack of any statistical significance of a chi-squared statistic computed for the Hausman test, the random-effects models seem to stand out as a more optimal solution compared to the fixed-effects models. However, since the existing studies often report country fixed-effects models, we do report these results. Table 4 presents the models with panel corrected standard errors.

Our findings from the panel data regression analysis are remarkably consistent with the findings derived from our original data: presidential power yields a statistically significant negative effect on the ENEP for both our OLS models and random-effects model M(05). For instance, in model M(05) the increase in presidential power by 10% results in

Table 3. Panel data analysis with cluster robust standard errors.

	M(01)	M(02)	M(03)	M(04)	M(05)	M(06)	M(07)	M(08)
Constant	1.06***	1.08***			1.21***	1.36***	1.22***	1.32***
	(0.18)	(0.18)			(0.15)	(0.17)	(0.18)	(0.20)
Logged Pres. Power (PP) (logPP)	−0.09**	−0.09**	−0.004	0.17	−0.09*	−0.03	−0.08	−0.03
	(0.04)	(0.04)	(0.01)	(0.22)	(0.05)	(0.07)	(0.05)	(0.07)
Presidentialism (PS)	−0.01	−0.05	0.003	−0.48**	0.04	−0.21	0.06	−0.21
	(0.15)	(0.26)	(0.21)	(0.20)	(0.11)	(0.16)	(0.15)	(0.18)
Semi-Presidential. (SP)	−0.01	−0.28	−0.17	−0.72**	−0.08	−0.32	−0.07	−0.34
	(0.21)	(0.35)	(0.21)	(0.36)	(0.18)	(0.32)	(0.16)	(0.23)
Ethnic fractionaliz.	0.08	0.09					−0.11	0.01
	(0.31)	(0.31)					(0.28)	(0.3)
District magnitude	0.13	0.13					0.04	0.04
	(0.14)	(0.14)					(0.12)	(0.12)
PS × PP	−0.02	−0.02		−0.19***		−0.13***		−0.13**
	(0.15)	(0.15)		(0.05)		(0.04)		(0.05)
SP × PP	−0.09	−0.09		−0.25		−0.11		−0.11
	(0.11)	(0.11)		(0.22)		(0.16)		(0.08)
Hausman χ^2					1.26 (df=3)	4.53 (df=5)		
Sig.					0.74	0.48		
R^2	0.1	0.1	0.003	0.04	0.15	0.16	0.15	0.16
Obs.	226	226	43	43	43	43	43	43

Sources: Constituency-Level Elections Archive (CLEA, 2014); Database of Political Institutions (2012).
Cluster robust standard errors in parentheses. Listed models: M(01)–M(02)—the pooled ordinary least squares (OLS) models; M(03)–M(04)—the fixed-effects models; M(05)–M(08)—the random-effects models.
Significance levels: *$p \leqslant 0.1$; **$p \leqslant 0.05$; ***$p \leqslant 0.01$.

Table 4. Panel data analysis with panel corrected standard errors.

	M(03)	M(04)	M(05)	M(06)	M(07)	M(08)
Constant			1.21***	1.36***	1.22***	1.32***
			(0.15)	(0.17)	(0.19)	(0.21)
Logged Pres.	0.004	0.17	−0.09*	−0.03	−0.08	−0.03
Power (PP) (logPP)	(0.13)	(0.19)	(0.05)	(0.07)	(0.06)	(0.07)
Presidentialism	0.00	−0.48*	0.04	−0.21	0.06	−0.21
(PS)	(0.17)	(0.24)	(0.11)	(0.16)	(0.12)	(0.18)
Semi-Presidential.	−0.16	−0.72	0.04	−0.32	−0.07	−0.34
(SP)	(0.32)	(0.46)	(0.18)	(0.33)	(0.2)	(0.34)
Ethnic fractionaliz.					−0.11	0.01
					(0.28)	(0.29)
District magnitude					0.04	0.04
					(0.1)	(0.1)
PS × PP		−0.19***		−0.13***		−0.13***
		(0.05)		(0.04)		(0.04)
SP × PP		−0.25		−0.11		−0.11
		(0.26)		(0.16)		(0.16)
Hausman χ^2			1.26 (*df*=3)	4.53 (*df*=5)		
Sig.			0.74	0.48		
R^2	0.003	0.04	0.15	0.16	0.15	0.16
Obs.	43	43	43	43	43	43

Standard errors for a panel model according to the Beck and Katz (1995) method, a.k.a. Panel Corrected Standard Errors (PCSE) in parentheses. Listed models: M(03)–M(04)—the fixed-effects models; M(05)–M(08)—the random-effects models.
Significance levels: *$p \leqslant 0.1$; **$p \leqslant 0.05$; ***$p \leqslant 0.01$.

a decrease in electoral fragmentation by roughly 1%, holding all other variables constant. All other models, except model M(04), demonstrate consistently negative, but statistically insignificant effects of presidential power on electoral fragmentation. As far as interactive effects are concerned, both findings from the fixed-effects and random-effects models demonstrate that if the state is characterized by a 'pure' form of presidentialism, its ENEP drops down by 0.19% compared to parliamentary regimes in M(04), and by 0.13% in M(06), respectively, with both coefficients being statistically significant at $p<0.01$ and $p<0.1$ levels. The inclusion of the ethnic fractionalization index and district magnitude did not improve the goodness-of-fit of random-effects models M(07) and M(08) in Table 3. Moreover, since both of these variables are time invariant, these were dropped out of our fixed-effects model, leaving us with the reduced version of the model. Therefore, our models M(03) and M(04) seem to be the most appropriate for our purposes and seem to be generally supportive of our earlier findings.

This being said, it should be stressed that more research has to be done in order to establish the validity of the reported models with a greater degree of empirical certainty. Of course, it would be worth exploring the hypotheses with more data. While our priority in building the dataset for testing working hypotheses was to maximize the breadth of countries included into the study, the narrowness of time period weakens the argument. However, while recognizing that our findings should be viewed as preliminary in the sense that more research is needed for them to be additionally validated and refined, we also find that the reported model is robust enough to serve as a direction for further empirical investigation.

Conclusion

Much of the reasoning regarding the interplay between presidentialism and legislative fragmentation has been traditionally focused on short-term coattail effects. The results of this study suggest that these effects are real, they do influence fragmentation in the electorate, but they are simply insufficient to make a significant impact upon the parameter of crucial importance for the functioning of presidential regimes: fragmentation in the legislature. The main impact of presidentialism is systemic, stemming from its tendency to restrict the number of parties in the system to a limited set of viable competitors for the presidential prize, and from its ability to appeal to national political forces at the expense of local political sentiment. Consistent with previous research, we find that the reductive influence of presidentialism upon the number of legislative parties is conditioned by the scope of presidential powers, the level of party system fragmentation, and electoral system effects. In addition, this study demonstrates that the level of the territorial homogeneity of the vote is highly consequential for legislative fragmentation, that party system under-institutionalization mitigates the impact of presidentialism, and that the type of presidential regime—'purely' presidential or semi-presidential—is highly consequential for the number of legislative parties.

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