Policy-based Voting and the Type of Democracy

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Policy-based Voting and the Type of Democracy

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It has often been argued that various features of what Lijphart (1999) calls consensus democracy, such as proportional representation, multiparty systems, coalition governments, and significant opposition influence on legislation promote party-voter linkages based on policy and ideology, while probably undermine accountability for performance in office. The latter, in its turn, is often thought to be promoted by features like majoritarian electoral rules, two-party systems, single-party governments, and executive dominance over the legislature. In this paper we examine these propositions using the extensive empirical evidence provided Module 2 of the Comparative Study of Electoral Systems (CSES) and a new set of measures for Lijphart’s political system variables provided by Adrian Vatter and Julian Bernauer for the purposes of this volume and described in chapter X. We depart from an important thread of the prior literature that found that executives and legislatures are slightly more representative of citizen preferences under proportional than majoritarian systems (see e.g. Powell 2000; Golder and Stramski 2010). Instead, we focus on a causally prior factor, namely the incidence of policy-based voting, or, to put it in a way that is conceptually both more accurate and better linked to macro phenomena, the responsiveness of aggregate election returns to shifts of policy preferences among citizens. Sections one to three discuss our dependent variable, theoretical expectations, and statistical models, respectively. Section four presents the empirical analysis and section five concludes.

1. Policy-based voting among citizens

At least some correspondence between voters’ policy positions and their vote is widely considered essential for creating a normatively desirable feature of democracy, namely policy-based linkages between citizens and their political representatives. This is obviously not the only desirable type of such linkages, and holding politicians accountable for performance (a.k.a. ‘valence’ issues, or competence and integrity) may even have a tradeoff with the degree to which citizen preferences on policy (a.k.a. ‘positional’, i.e. divisive, issues) impact the vote (Powell 2000, 165). Yet, a stronger impact of citizens’ policy preferences on the vote should be, and is widely considered, desirable for democracy as a sign of those preferences exercising a stronger prospective control over policy choices by elected elites.

Comparatively evaluating the total impact of citizens’ policy preferences on vote choices across a large number of democracies is not an easy task, though. We approximate it by looking at how much a sizeable shift in ideological self-placements among citizens on an abstract ideological scale (left vs. right) might have changed the aggregate distribution of votes across parties. This focus on a single ideological dimension is justified mainly by previous demonstrations in the extant literature of its usefulness for cross-contextual comparisons regarding the impact of policy-oriented voting (see e.g. van der Brug 1997; van der Eijk et al. 2005), which in its turn provided by the ability of the left-right semantics to absorb whatever the relevant dimensions of policy and ideological differentiation are among the parties in a given party system (see Fuchs and Klingemann 1989; Kitschelt and Hellemans 1990; van der Brug et al. 2009). At the same time, this focus on a single ideological dimension allows us to root our theoretical expectations in the prior literature on how political institutions impact ideological differentiation among competing political parties. Golder and Stramski’s (2010) analysis, which also used CSES data, already established that the distribution of left-right positions in legislatures of ‘proportional’ systems is somewhat more congruent with the distribution of left-right positions in the electorate than the same distribution in the legislatures of ‘majoritarian systems’. We seek to add to their test by focusing on responsiveness of aggregate election outcomes to shifts in citizens’ preferences, i.e. by examining whether such differences in elite-
mass congruence between consensus and majoritarian democracies can conceivably come about as a result of supposed institutional influences on voting behaviour.

2. Polity-level determinants
Following Duch and Stevenson’s (2008) classification of structural factors conducive for retrospective economic voting and Knutsen and Kumlin’s (2005) discussion of the mechanisms through which ideological polarization among parties may impact voting behaviour, we expect that political institutions impact policy-based voting among citizens via three key routes: through (a) the political supply – i.e., smaller or bigger political differences among the parties, which is clearly a major influence on the incidence of ideological voting among citizens (Lachat 2008) –; (b) the political communication of these differences from elite actors to voters; and (c) voters’ motivation to act on the ideological differences that they see.

List PR should facilitate greater ideological differentiation among parties than majoritarian electoral rules (Cox 1990), as well as a more ideological and less personality- and performance-based electoral competition (Katz 1980), i.e. a stronger communication of existing differences to voters and the stronger priming of the latter on ideology. The same could probably be expected from multipartyism and power sharing among partisan actors at the national level, i.e. oversized multiparty coalitions, minority government dependent on some cooperative opposition parties in the legislature, and a legislative process that gives significant veto powers to the opposition. Compared to single party governments, twopartism, and executive dominance of the legislative process, all these institutions ought to reduce the incentives for ideological convergence among the parties, help to highlight existing policy differences, and probably prime citizens more on ideological and policy differences than on credit and blame for performance. In line with an extensive literature going back to Powell and Whitten (1993), we would expect that the reverse applies to performance-based voting: the more majoritarian a country’s institutions, the stronger the impact of voters’ performance evaluations on the vote will be.

It is less clear whether a country’s location along the ‘federal-unitary’ dimension of Lijphart’s typology can affect our dependent variable. Federalism, bicameralism, judicial review and decentralization should promote horizontal accountability and power-sharing between political actors, which might be instrumental for clarifying existing policy differences and undermining voters’ motivation for performance-(as opposed to policy-) based voting. Yet these expectations are not very strongly motivated because the institutional features associated with this dimension of differentiation among democracies concern relations between territorial units and functionally separated actors (such as judicial and executive power), rather than electorally competing political parties. Hence one could equally expect them to create divisions within parties and policy convergence across them, thus undermining policy differentiation between parties as well as its communication to voters and voters’ motivation to act on perceived ideological differences.

All in all, we hypothesize a tradeoff between the impacts of performance-based and policy-based voting across countries. We further hypothesize that Lijphart’s (1999) ‘executive-parties’ dimension of consensus democracy promotes policy-based voting partly through ideological polarization between the parties, but also independently of that. Finally and less clearly, country locations along Lijphart’s federal-unitary dimension of democratic systems should probably not affect policy-based voting either through ideological polarization or independently of that.
3. Research design, data and measures

Our analysis aims at a better understanding of how various characteristics of the political system explain differences across elections in the extent to which policy preferences – as captured by citizens’ left-right self-placements – influence the vote. Given our interest in macro-determinants of micro-level behaviour, we need observations that are comparable both across different individuals acting in the same election, and across a significant number of elections that took place in different political institutional contexts. The comparative dataset that best meets the above criteria while also featuring suitable indicators of citizens’ left-right placements and performance evaluations is Module 2 of the Comparative Study of Electoral Systems project. This module was administered as part of national post-election surveys to probability samples of the voting population following national elections that took place on five continents between 2001 and 2006, yielding a total of 42 political contexts and 61,018 voters for our aggregate- and individual-level analyses, respectively.¹

The individual-level analysis consists in deriving comparable measures of the impact of policy preferences on the vote choice for each of our 42 political contexts, while the aggregate level analysis simply regresses these measures on quantitative indicators of political system characteristics, taking into account the statistical error with which these measures were calculated on the basis of effect parameters in the individual-level models.

We could only do this via a single multilevel statistical model if the dependent variable in our individual-level analysis had been identically coded across all political contexts, for instance into “left” and “right”, or into “pro-government parties/candidates” and “all other parties/candidates”. Such standardized measures would, however, hardly do any justice to the complex differences in the nature of electoral choices between consensus and majoritarian democracies.

Therefore the dependent variable in our individual level analyses will be Vote choice, a multinomial variable with a country-specific coding scheme that simply lists the major alternatives on the ballot in the given election (e.g., Labour, Conservative and Liberal Democrat in the 2005 British general election). For obvious reasons, the individual-level dependence of such a dependent variable on policy preferences and other factors can only be estimated via vote

¹ We kept elections in the analysis even with a lower level of democracy as long as they appeared to have been sufficiently competitive and influential for institutional variables related to the distinction between consensus and majoritarian democracies to have any impact on the incidence of performance- or policy-oriented voting behaviour. Hence, we only dropped one of the CSES2 elections – the 2005 presidential election in Kyrgyzstan – from the analysis, because of the insufficient variation on the vote choice variable caused by a candidate who collected about ninety percent of all votes and doubts about whether vote-counting was sufficiently fair to award voters real influence on government formation. We treat as two separate cases – though occurring in identical context as far as consensual vs. majoritarian traits of democracy are concerned – the elections for the French- and Dutch-speaking lists in Belgium, Eastern and Western Germany, and in Quebec and the English-speaking provinces of Canada, because the major within-country differences in their party systems would have made the estimation of identical vote functions meaningless for these regions. While these cases are hardly as independent of each other as elections in two neighbouring countries, they are substantially different with respect to the key dependent variables in our aggregate level analysis (see Figure 3). Introducing a correction for clustering for just three pairs of cases out of a total of 42 seemed too much ado for too little and was thus avoided in the analysis.
functions estimated separately for each of the 42 contexts in the analysis. It would be both technically impossible and substantively meaningless to estimate macro-micro interactions between the individual-level parameters of such a model on the one hand, and political system characteristics on the other. Therefore the individual- and aggregate-level analyses are carried out in separate steps. The dependent variable in the second step is not a particular parameter in the micro-level multinomial regressions but a quantity derived with post-estimation data manipulation. Otherwise, however, the logic of multilevel analysis remains intact (cf. Achen 2005 and Long Jusko and Shively 2005 regarding methodological considerations; and Duch and Stevenson 2008 for a closely related example).

The individual-level analysis involves estimating a vote function separately for each of the 42 contexts. The model posits a multinomial logit link function $fn$ between the Vote choice variable on the left-hand side and its presumed determinants, including Left-right self-placement, on the right-hand side. For ease of reading, Equation 1 greatly simplifies the tedious notation for such a model and replaces the maths of the logit link function with a simple and generic reference to the $fn$ link function:

\[
\text{VoteChoice} = fn(B_0 + B_1(\text{Strength of party identification}) + \\
+ B_2(\text{Left - right self - placement}) + B_3(\text{Left - right self - placement}^2) + \\
+ B_4(\text{Government performance evaluation}) + \\
+ B_5(\text{Strength of party identification})(\text{Left - right self - placement}) + \\
+ B_6(\text{Strength of party identification})(\text{Left - right self - placement}^2) + \\
+ B_7(\text{Strength of party identification})(\text{Government performance evaluation}) + \\
+ B_8(\text{Female}) + B_9(\text{Age}) + B_{10}(\text{Age}^2) + B_{11}(\text{Education low}) + B_{12}(\text{Education high}) + \\
+ B_{13}(\text{Farm job}) + B_{14}(\text{Manual work}) + B_{15}(\text{Rural residence}) + B_{16}(\text{Income}) + \\
+ B_{17}(\text{Devout}) + B_{18}(\text{Minority 1}) + B_{19}(\text{Minority 2})
\]

The first problem that arises with the justification of this model is the notorious dependence of both ideological self-placements and performance evaluations on party preference itself, i.e. that our vote function may suffer from an endogeneity problem (see e.g., Knutsen 1997 and Evans and Anderson 2006). Since our derivation of the dependent variable for the aggregate-level analysis crucially depends on the twenty B parameter vectors estimated for each of the 42 political contexts with the above equation, this is an important objection. We answer to it in three ways.

First, prior cross-national research by Duch and Stevenson (2008: 123ff) showed that relative differences between elections in the amount of performance-oriented voting remain virtually identical whether or not the estimates are “purged” of endogeneity in performance evaluations. We expect that the same applies to our measure of ideological voting, which closely parallels Duch and Stevenson’s measure of the economic vote in that they are both based on the net association between vote choice and a single introspective question about the respondent – how s/he finds the economy and where s/he is on a left-right scale, respectively – without using any information about the ‘objective’ responsibility of parties for the economy or, what would be conceptually equivalent to that, the ‘objective’ left-right position of the parties. Second, we see no theoretical reason to suspect that relative cross-national differences regarding performance vs.
policy influences on vote choices could be severely obscured by cross-national variation in the extent to which our indicators of performance evaluations and policy preferences are endogenous to vote choice itself. If anything, our intuition is that endogeneity (i.e. the tendency to rationalize voting preferences in performance-oriented or in policy-oriented ways) increases with the true effect of performance- and policy considerations on vote choices. Should that be the case, cross-contextual variation in the endogeneity of our measures of performance- and policy-oriented considerations do not alter relative cross-national differences in the influence of these factors on vote choices. Third, the endogeneity of these measures to party preferences must logically be a largely non-existent or at least greatly diminished problem among non-partisan respondents. Therefore, our actual measures of the impact of policy preferences and performance-evaluations on the vote will, in the analysis below, be based on estimates that we are making for such respondents. For this last reason, the vote functions that we estimate with individual level data for each of the 42 electoral contexts include the interactions of performance evaluations and policy preferences with Strength of party identification as shown in Equation (1) above.

A further question is due about whether Left-right self-placement and Government performance evaluation really achieve what our model may seem to expect from them, namely to capture all relevant performance- and policy-related considerations. We do not think so. To be sure, the left-right semantics is largely avoided in favour of other terminologies (or political dimensions) in some countries like the US, and only in the Japanese data set we have a supposed local functional equivalent (a “progressive-conservative” scale) to avoid the probable Euro-centric bias in the measurement of the respondent’s ideological position. Hence our Left-right self-placement variable may have an unequal ability to capture policy preferences in the different countries in the analysis.

While this line of thinking is worth to explore further, we note that the validity of our analysis does not in fact depend on the excessively strong assumption that Left-right self-placement and Government performance evaluation capture all relevant performance- and policy-related considerations in every country; not even on the slightly weaker assumption that they do so to an equal extent in all political contexts. Rather, the right question about the validity of our analysis is whether the undeniable and inescapable measurement problems of our proxies are systematically correlated with political system characteristics. Only if that is the case can our findings about the impact of the latter on the former be misguided. At this point we see neither theoretical reasons to expect this, nor a feasible empirical way to study whether such a distortion could emerge. Therefore we leave these important issues for further study and proceed to probe the existing data.

Another question regarding Equation (1) concerns the question of whether an omitted variable bias may impact our estimates about policy and performance-oriented voting in the given systems. We respond to this concern by including as many socio-demographic control variables in the model as possible given the CSES 2 data set. We expect these variables to capture such shared causes of vote choice on the one hand, and policy preferences or performance evaluations on the other that, if left uncontrolled in the vote function, may create spurious correlations between vote choice and the given political attitudes. The reader may wonder why we do not add similar controls for political attitudes. The reason is that our theoretical interest here is in the total effect of policy preferences and performance evaluations on the vote. What any other political attitude variable (for instance, issue attitudes, satisfaction with democracy, or leader evaluations) would add is nothing else but some rather more specific policy and/or performance related considerations. Thus, these controls would influence the
estimated impact of Left-right self-placement and/or Government performance evaluation on vote choice for an entirely wrong reason. Therefore the only additional political attitude variable that appears next to Left-right self-placement and Government performance evaluation in our model is the Strength of party identification, which appears there only because of our intention to estimate the impact of the theoretically relevant variables on vote for non-partisan respondents. As discussed above, this restriction seems advisable given the inevitable endogeneity of both policy positions and performance evaluations to vote choice among partisan voters.

Technical details about the meaning and coding of the variables entering Equation (1) are presented in our online appendix.\(^2\) Obviously, many of these variables (particularly income) came with missing values for many respondents and we were concerned about the impact that dropping these cases may have on the accuracy and efficiency of the estimates. Therefore we used the Amelia 2 package of Honaker \textit{et al.} (2007) to multiply impute all missing values in our individual-level data sets.\(^3\) During all individual-level statistical analyses we made use of the country-specific socio-demographic weight variables deposited with the CSES data set, and consistently excluded everyone who did not give a valid answer (other than did not vote, spoiled the ballot, or voted blank/invalid) to the question about vote choice.

The first yield of our individual level statistical analyses was a vast number of multinomial logistic coefficients for each of 42 political contexts that are of no theoretical interest here and cannot be presented for reasons of space. To estimate what impact our two key theoretical variables have on vote choices in the given context, we derived four quantities for each party/candidate that make use of these initial regression coefficients. These four quantities reflected the expected fractional share of the party/candidate among all voters in the analysis under the given model and the observed characteristics of these respondents except that they all obtain the minimum score on the Strength of party identification variable and, in the first case, their Left-right self-placement moves one sample standard deviation to the left of its observed value; in the second, their Left-right self-placement moves one sample standard deviation to the right of its observed value; in the third, their Government performance evaluation drops one sample standard deviation below its observed value; and, in the fourth, their Government performance evaluation rises one sample standard deviation above its actual value.

The difference between the second and first of these quantities gives our estimate about the Impact of policy preferences on the vote share of the given party/candidate in the given electoral context; the difference between the fourth and the third provides our estimate about the Impact of performance evaluations on the vote share of the given party/candidate. To estimate the statistical error regarding these coefficients that derive from the fact that we base our estimates on random samples of the relevant population, we bootstrapped the estimation of these two differences (see Efron and Tibshirani 1993 for a discussion of this method). Specifically, we took 200 random subsamples (with replacement) from each of the 42 samples in the analysis, and re-estimated the parameters of Equation (1), the relevant sample standard deviations and the resulting estimates about the Impact of policy preferences and the Impact of performance evaluations for each resampling. The bootstrapping process provided us with 200 estimates for each of the 230 parties/candidates that we could distinguish between in our analysis.

\(^2\) See http://www.personal.ceu.hu/departs/personal/Gabor_Toka/Policy/

\(^3\) We created five imputed data sets for each political context separately. The imputations were based on a slightly larger set of variables than those listed in the online appendix. Technical details are available from the authors.
Figures 1 and 2 about here

Figure 1 and 2 give a sense for how these estimates look like using the example of the countries that are the best (albeit in the majoritarian case imperfect) examples of what Lijphart meant by majoritarian and consensus democracy, respectively. The horizontal axis shows the estimated impact of policy considerations, and the vertical axis the impact of performance evaluations on the vote share of each party. Light dots indicate the location of the 200 bootstrap estimates for each party. The party acronym is printed exactly at the mean value of the 200 estimates for the given party. What the chart signals above all is that very distinctive numerical estimates are obtained for the different parties in both democracies, and that some parties would greatly benefit or lose (by up to as much as 30 percent of the total vote) if a two standard deviations change occurred in everyone’s left-right placement and/or satisfaction with government. The two figures also illustrate the difficulty of making comparisons across political systems using these party-level estimates. Probably against our theoretical expectations, the expected impact of such a vast change in policy preferences in the electorate is slightly bigger (a roughly 30 percent swing in the case of both Labour and the Conservatives) on the main ideological antagonists in the UK than in Switzerland. However, it could be that the UK estimates are higher only because of a method artefact, namely because the individual parties tend to be smaller in multiparty Switzerland than in the two-and-a-half party system of the UK. If so, the Swiss parties, taken individually, could not lose as much from any swing of public opinion as their UK counterparts. However, since more parties’ vote is affected by ideological change in Switzerland than in the UK, the combined shift of the vote may actually be higher in the first than in the second. Therefore comparisons about effects size may be misleading at the level of individual parties. Furthermore, since the gains and losses of the individual parties cancel out within a system, the party-level estimates cannot be treated as independent observations in an aggregate level analysis.

Because of these concerns we decided to aggregate the party-level estimates into a single figure (for each of the 200 resamplings) for each political context with the help of a modified Pedersen-index. Proposed for the measurement of aggregate electoral volatility in a country k, the original Pedersen-index summed up half the sum of the absolute value of the (positive or negative) change in each party i’s share of the vote. This index is useful for our present purposes but needs to be adjusted when applied to bootstrap estimates. In Taiwan, for instance, the impact of policy preferences is quite close to zero and some resamplings thus suggest positive while other suggest negative effects of the same change on the same party’s share of the vote. If we simply summed up the absolute values of the changes across parties disregarding differences between estimates that point in totally different partisan directions, we would obtain exaggerated figures about the total impact of policy preferences on Taiwanese election outcomes. Indeed, we would inevitably obtain a positive value of the Pedersen-index for all resamplings, which would mistakenly suggest statistically significant effects even where the effects are really not significant at all and are entirely inconsistent in direction across the resamplings.

The key dependent variables of our aggregate level analysis are thus defined by Equations (2) and (3), which state that the impact of policy preferences/performance evaluations in election k is calculated, for each resampling p using the original Pedersen index (as half the sum of the absolute value of the expected impact on each party i within the same k context), except that the absolute value of the estimated impact is multiplied by an \( a_{kip} \) factor that can only be either plus one (when the estimated impact for party/candidate i in context k in resampling p
has the same sign as the mean estimate for the same party across the 200 resamplings) or minus one (when the same two figures have the opposite sign):

\[
\text{(Impact of policy preferences)}_{k_{ip}} = \frac{1}{200} \sum_{i=1}^{200} \left( a_{k_{ip}} \text{[Impact of policy preferences]}_{k_{ip}} \right),
\]

where

\[
a_{k_{ip}} = \frac{1}{200} \sum_{j=1}^{200} \text{[Impact of policy preferences]}_{k_{ip}} \times \text{[Impact of policy preferences]}_{k_{ip}}
\]

\[
\text{(Impact of performance evaluations)}_{k_{ip}} = \frac{1}{200} \sum_{i=1}^{200} \left( a_{k_{ip}} \text{[Impact of performance evaluations]}_{k_{ip}} \right),
\]

where

\[
a_{k_{ip}} = \frac{1}{200} \sum_{j=1}^{200} \text{[Impact of performance evaluations]}_{k_{ip}} \times \text{[Impact of performance evaluations]}_{k_{ip}}
\]

Figure 3 displays the 200 bootstrap estimates about the location of the 42 political contexts in the two-dimensional space formed by the impact of policy preferences and performance evaluations on the vote. Table 1 gives mean estimates and their confidence intervals for each context. There are a quite a few pairs and even triads of contexts for which the cloud of bootstrap estimates clearly overlap, suggesting no statistically significant differences between these pairs and triads of cases. In the most extreme instance of such similarities, we even had to display the acronyms for South Korea (KR) and Portugal 2005 (PT-05) at some distance from the actual mean value of the estimates for these contexts because otherwise their acronyms would get mixed up in the chart with the one for Quebec (CA-Q). Yet by and large the estimates are very distinctive regarding individual contexts, with each of them appearing to be significantly different, at least in one of the two dimensions, from a 90+ percent majority of all other cases in our sample of 42 contexts.

Figure 3 and Table 1 about here

Our key substantive question is whether political system characteristics influence the incidence of policy-based voting. The aggregate level analysis of this question estimates the simple OLS regression model shown in Equation (4) below. The impact in question is the function of a \( \beta_0 \) constant, the weighted sum of the given elections k’s score on an arbitrarily selected n number of political system characteristics, and an election specific component (or in other words residual error of the fitted values from the model). The constant of the model and the election-specific components of the impact have no theoretical relevance in our case. Instead, it is the weighting of the political system characteristics by a set of regression coefficients called \( \beta_m \) that reveals what features of the political institutional context make the impact of policy preferences bigger or smaller in elections. The model parameters can be identified and their margin of error can be empirically estimated under the relatively weak assumption that the
\( \varepsilon_k \) election-specific error term is normally distributed around a mean of zero. The validity of the results of course depends on the significantly more demanding assumption that our sample of \( k \) elections was selected at random and is composed of independent events.

\[
(Impact \ of \ policy \ preferences)_k = \beta_0 + \sum_{m=1}^{o} \beta_m (Political \ system \ characteristics)_{mk} + \varepsilon_k \tag{4}
\]

Two variables on system characteristics – Executive_parties and Federalism_Unitarism, respectively – were provided by Adrian Vatter and Julian Bernauer for this volume to locate a large number of contemporary democracies on Lijphart’s two conceptual and empirical dimensions of majoritarian vs. consensus democracy. Annual values on the elementary variables making up the two indices were standardized and averaged across the years from 1997 to the year of the election covered by the CSES2 survey data in the given country. We estimate Equations (4) with 39 of our 42 macro-level cases.\(^4\) In addition, our analysis makes use of a measure of Polarization, which shows the standard deviation of the expert-estimated left-right positions of each country’s relevant parties, with the parties weighted by their number of voters in the CSES 2 survey. The source of the expert estimates about party locations was the Macro Data Set accompanying the CSES2 survey.\(^5\) Relevant OLS regression estimates are displayed in Table 2 and will be discussed below.

Table 2 about here

4. Empirical findings

The online appendix provides test results regarding the statistical significance of the various effects in Equation 1 that involve left-right self-placement as an indicator of policy preferences. It suffices here to summarize these results just briefly. A simple direct effect of left-right self-placement on vote choice is clearly significant in all but a few non-European electoral contexts, which are the English-speaking provinces of Canada, Hong Kong, the Philippines, and Taiwan (where the effect is of borderline significance). The additional effects of the squared value of left-right self-placements are also statistically significant in well over half the electoral contexts, especially often where the number of parties/candidates in the analyses is relatively large. This makes good theoretical sense as in conventional spatial models based on policy-based voting some parties in the more complicated multiparty systems may have the highest probability of support not on the extremes of the ideological spectrum but somewhere closer to the centre. Therefore we feel that the inclusion of this squared term in the common model for all political contexts is well justified.

Taken together, the two interaction effects of left-right self-placement and its squared term on the one hand, and strength of partisanship on the other are statistically insignificant in a majority of the 42 contexts. However, they appear to register significant effects in a lot more

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\(^4\) Two elections in Taiwan and one in Hong Kong had to be dropped from this analysis for lack of data on some institutional features.

\(^5\) Missing expert judgements were single-imputed using information about each party’s left-right position in all CSES2 countries as estimated via the mean value of all voters’ placements of these parties on a left-right scale in the CSES2 survey, or, if that was also missing, via the mean self-placement of each party’s voters in the survey.
cases than we would expect this to occur just by chance. In nearly all these latter contexts, the effects of policy preferences tend to increase with the strength of partisanship, which is consistent with our expectation that we see here a spuriously inflated effect among partisans due to a greater endogeneity of ideology to party preference in this group. This seems to support our decision that the estimates for non-partisans should provide the best guide to the actual effect of policy preferences on vote choice. Last but not least, the comparison of the likelihood ratio statistics presented in the online appendix regarding models 1 and 4 suggests that at least some of the above effects of policy preferences are statistically significant in every single context except Taiwan in 2001 and, on a closer call, in the Philippines.

Another and probably more interesting way of looking at the significance of these effects is look at the bootstrapped estimates about the expected impact of a two standard deviation shift in everyone’s policy preferences on the vote share of individual parties. Since the Impact of policy preferences may be either underestimated (due to the imperfect measurement of policy preferences) or overestimated (if left-right self-placement is endogenous to vote choice even among the non-partisan), it is instructive to consider this issue in a relative perspective, i.e. whether policy preferences or performance evaluations exercise a bigger effect on the vote share of parties/candidates. Considering all 230 parties/candidates in the analysis, we find that a two sample standard shift in performance evaluations would change (positively or negatively) the average party’s vote share by 9.7 percentage point; while a similarly large shift in policy preferences would change the same share by 10.4 percentage point (data not shown). Considering this evidence, one may want to conclude that policy preferences and performance evaluations are equally well (or equally poorly) reflected in election outcomes.

Of course, the picture varies considerably across individual parties/candidates. Figure 1 and 2 illustrate the general trends that emerge. Apparently there are rather big positive effects of the simulated two standard deviation shift to the right among citizens on the vote share of right-wing parties like the Conservatives in Britain or the Swiss People’s Party (SVP/UDC). This is mirrored by the large negative effect of such changes on the vote share of left-wing parties like Labour in Britain and the Social Democratic Party (SP/PS) in Switzerland. More centrist parties like the Christian Democrats (CVP) in Switzerland or the Liberal Democrats in Britain are little affected by this change: they may get a more or less different set of individuals voting for them when such massive ideological shift occurs in society, but their overall share of the vote changes little in the process. In this sense the vote share of these centrist parties says rather little about policy preferences in society compared to what is revealed by the ups and downs of the distinctly left- and right-wing parties.

Similar trends emerge regarding the impact of performance evaluations. The simulated increase in satisfaction boosts the vote shares of the UK government (provided by Labour at the time of the survey), and undermine the vote of the main opposition party. Meanwhile the vote share of the Lib Dems would barely be affected by this change, presumably because they attract medium-satisfied voters, whose proportion does not change much under the given scenario. Most interesting is that in Switzerland, where all the parties in the chart save the Greens (GPS/PES) and the ‘others’ have been part of the federal government all the time since 1943, the vote share

6 Of course, the negative and positive signs of the changes in Figures 1 and 2 merely indicate which parties would be differently affected by homogeneous movements of the electorate in a particular dimension. The size of the effect would be the same but the sign reversed if we simulated the impact of a similar movement in a leftward (or, in terms of performance evaluations, dissatisfied) direction.
of parties is still quite sensitive to shifts in performance evaluations, even if not quite as much as in Britain. Apparently not all Swiss government parties were equally willing and able to claim credit from voters satisfied with the government, and at least one government party (the SVP/UDC) managed to attract a relatively dissatisfied electorate in the 2003 election. Hence even under such a vastly oversized coalition government as the one in office in Switzerland, elections can still remain a barometer of public opinion with respect to satisfaction with performance. Conversely, even in the relatively majoritarian democracy of the UK, election results show more than just citizens’ satisfaction with government. Since the parties stake out distinct ideological positions and the voters apparently respond to that accordingly, their left-right position substantially influences the distribution of the vote over and above whatever influence is exercised by (dis)satisfaction with government performance.

For reasons explained in section three, we believe that the impact of political system characteristics is better examined at a higher level of aggregation than those of the individual parties, and for this reason we cumulate the party-level effects with our slightly modified Pedersen-index shown in Equations (2) and (3) into the 200 times 42 context-level estimates displayed in Figures 3 and Table 1. The first remarkable finding here is the apparent negative relationship between the location of electoral contexts along the two dimensions of Figure 3 (significant at the p=.03 level). The negative correlation ($r=-0.34$) is consistent with theoretical expectations and suggests a weak trade-off between the extent to which citizens’ policy preferences and performance evaluations are reflected in election outcomes. The contexts around the top left corner of Figure 3 are characterized by strong performance-oriented and weak policy-oriented voting, and include a conspicuously large number of non-European cases: the two Taiwanese elections, the Philippines, Hong Kong, Mexico, English-speaking Canada, Japan, and Chile. On the opposite, bottom-right quadrant of the figure we see an equally remarkable concentration of European multiparty-systems, with Norway, Poland, the French-speaking part of Belgium, Italy, Finland, the Netherlands and Portugal in 2002 as probably the most extreme cases on this end. It is tempting to infer that ideological self-placements (particularly on the left-right scale) are probably much less appropriate proxies of policy preferences in these non-European contexts, but at this point this remains merely a speculation.

Table 2 presents our statistical evidence about how type of democracy may alter the responsiveness of election outcome to policy preferences in the electorate. Here, we regress the bootstrapped estimates of the Impact of Policy Preferences on various sets of our key independent variables. Note again that the level-1 margin of error regarding the true Impact of Policy Preferences in individual elections is factored in the analysis using Rubin’s (1987) rule and the mim package of Carlin et al. (2008), which makes these results technical equivalents of estimated macro-macro effects in a single-step multilevel model.

The findings are largely consistent with our theoretical expectations. The executive-parties dimension of consensus democracy is positively and significantly associated with the responsiveness of election outcomes to policy preferences in the electorate, and this effect retains borderline statistical significance at $p=0.06$ even when ideological polarization between the parties is controlled for (data not shown), and $p$ remains 0.06 when both polarization and country location on the federal-unitary dimension are controlled for (see the rightmost panel in Table 2). The bivariate impact of the executive-parties dimension on polarization is positive (with

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7 As in all aggregate level analyses reported in this paper, here too we report estimates aggregated across the 200 sets of bootstrapped estimates following Rubin’s (1987) rules, i.e. as if we were dealing with a dataset that with 200 multiply imputed values for each observation.
R²=0.09, data not shown) and borderline significant (p=0.06). Polarization has a substantial and significant direct effect on the Impact of Policy Preferences net of other model variables. However, a little over half the total effect of the executive-parties dimension is independent of polarization. This may support our expectation that country locations on the executive-parties dimension also impact the clarity with which party positions are communicated to voters and/or voters’ motivation to respond to party positions on divisive policy issues.

Somewhat unexpectedly, country locations on the federal-unitary dimension show a similarly sizeable effect on the Impact of Policy Preferences (see Table 2) as well as on Polarization (data not shown). Federalist traits are associated with less policy voting and less polarization than unitary ones, though the differences regarding polarization do not reach conventional levels of statistical significance in either the bivariate or multivariate specifications that we tried (data not shown). The negative effect of federalism on Impact of Policy Preferences is also just borderline significance (see Table 2).

5. Discussion
The institutional features of consensus democracy appear to affect voting behaviour, and there is a moderately strong tradeoff between performance- and policy-based voting across democracies. Both theoretical and empirical uncertainty remains though about the influence of country locations along the federal-unitary dimension on the responsiveness of national election outcomes to policy preferences in the electorate. If there is an effect at all, then, according to our analysis, the impact is more likely to be negative than positive. In contrast, features associated with the executive-parties dimension should, on theoretical grounds, make consensus democracies more responsive to electoral preferences than majoritarian democracies. Our empirical analysis suggests that this is indeed the case. The effect is at least partly mediated by consensus democracy facilitating slightly bigger ideological polarization between parties than majoritarian systems. Yet, while the statistical significance of our findings is not very impressive, on both theoretical and empirical grounds we are inclined to think that consensus democracy also influences the degree of policy-based voting in the electorate independently of party polarization.

The 0.05 regression coefficient shown in the leftmost panel of Table 2 suggests though that these normatively desirable effects of consensus democracy are very-very modest. On this basis, we would expect the difference on the executive-parties dimension between Britain (with a score of -2.44) on one extreme and Belgium (with a score of 1.62) on the other to produce just about one percent bigger aggregate change in election results among non-partisan voters in the latter country when a massive two standard deviation change occurs in the left-right policy preferences of the electorate. It is hard to see by what standard this tiny expected difference would be worth attention from institutional designers.

As always, the possibility remains that these results are merely reflecting the impossibility of identifying, across such a heterogeneous sample of democracies, either the extent of policy-based voting with the help of the left-right dimension, or clear effects of institutional design. However, in further analyses we failed to find significantly stronger effects of the executive-parties dimension on the Impact of Policy Preferences either in European than non-European countries, or in old than in new democracies. Nevertheless, we suspect that the relative irrelevance of the left-right semantics outside of Europe does influence our findings: even with Israel and the Anglo-Saxon world included, the impact of the executive-parties dimension on our measure of policy-based voting is essentially zero, and, in fact, negative. Yet,
while the *Impact of Policy Preferences* appears notably higher within Europe, it does so almost irrespectively of institutional setup, with even the bivariate impact of the executive-parties dimension dropping to 0.03 and staying there when we only look at Europe’s older democracies.  

All in all, the theories explored in this chapter receive some support from the empirical data, but add little to explaining why the world’s democracies show such a striking variation in policy responsiveness as Figure 3 suggested above.

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8 The associated p-value is 0.11 in the analysis of all 26 European elections in our analysis, and 0.13 when only the 13 elections in older European democracies are considered.
Figure 1: The effect of changes in policy demand and performance evaluations on the vote share of UK parties in 2005. 200 bootstrap estimates of the change in each party’s fractional share of all votes.
Figure 2: The effect of changes in policy demand and performance evaluations on the vote share of Swiss parties in 2003. 200 bootstrap estimates of the change in each party's fractional share of all votes.

The impact of a two sample standard deviation shift to the right in all voters' left-right position.
Figure 3: The party-level effects of changing policy demand and performance evaluations aggregated with the Pedersen-index.
200 bootstrap estimates for each election.
Table 1: The estimated Impact of Policy Preferences and Impact of Performance Evaluations across 42 political contexts

<table>
<thead>
<tr>
<th>Country</th>
<th>Acronym</th>
<th>Impact of Policy Preferences 95% confidence interval of the bootstrapped mean</th>
<th>Impact of Performance Evaluations</th>
</tr>
</thead>
<tbody>
<tr>
<td>ALBANIA (2005)</td>
<td>AL</td>
<td>0.34 to 0.42</td>
<td>0.14 to 0.23</td>
</tr>
<tr>
<td>AUSTRALIA (2004)</td>
<td>AU</td>
<td>0.19 to 0.28</td>
<td>0.42 to 0.48</td>
</tr>
<tr>
<td>BELGIUM, DUCH SPEAKERS (2003)</td>
<td>BE-F</td>
<td>0.20 to 0.27</td>
<td>0.29 to 0.34</td>
</tr>
<tr>
<td>BELGIUM, FRENCH SPEAKERS (2003)</td>
<td>BE-W</td>
<td>0.34 to 0.42</td>
<td>0.08 to 0.15</td>
</tr>
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<td>BRAZIL (2002)</td>
<td>BR</td>
<td>0.10 to 0.16</td>
<td>0.17 to 0.22</td>
</tr>
<tr>
<td>BULGARIA (2001)</td>
<td>BG</td>
<td>0.26 to 0.33</td>
<td>0.12 to 0.17</td>
</tr>
<tr>
<td>CANADA W/O QUEBEC (2004)</td>
<td>CA-E</td>
<td>0.02 to 0.14</td>
<td>0.20 to 0.31</td>
</tr>
<tr>
<td>QUEBEC, CANADA (2004)</td>
<td>CA-Q</td>
<td>0.27 to 0.35</td>
<td>0.21 to 0.28</td>
</tr>
<tr>
<td>CHILE (2005)</td>
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<td>0.12 to 0.21</td>
<td>0.42 to 0.47</td>
</tr>
<tr>
<td>TAIWAN (2001)</td>
<td>TW01</td>
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<td>CR</td>
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<td>0.24 to 0.37</td>
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<tr>
<td>DENMARK (2001)</td>
<td>DK</td>
<td>0.40 to 0.44</td>
<td>0.23 to 0.28</td>
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<tr>
<td>FINLAND (2003)</td>
<td>FI</td>
<td>0.37 to 0.43</td>
<td>0.14 to 0.22</td>
</tr>
<tr>
<td>FRANCE (2002)</td>
<td>FR</td>
<td>0.25 to 0.38</td>
<td>0.15 to 0.26</td>
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<tr>
<td>WEST GERMANY (2002)</td>
<td>DE-W</td>
<td>0.16 to 0.24</td>
<td>0.43 to 0.48</td>
</tr>
<tr>
<td>EAST GERMANY (2002)</td>
<td>DE-E</td>
<td>0.23 to 0.29</td>
<td>0.37 to 0.42</td>
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<tr>
<td>HONG KONG (2004)</td>
<td>HK</td>
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<td>0.33 to 0.38</td>
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<td>0.23 to 0.28</td>
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<td>ISRAEL (2003)</td>
<td>IL</td>
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<td>ITALY (2006)</td>
<td>IT</td>
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<td>JAPAN (2004)</td>
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<td>KOREA (2004)</td>
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<td>0.25 to 0.32</td>
<td>0.21 to 0.28</td>
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<tr>
<td>MEXICO (2003)</td>
<td>MX</td>
<td>0.03 to 0.11</td>
<td>0.21 to 0.27</td>
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<tr>
<td>NETHERLANDS (2002)</td>
<td>NL</td>
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<td>NEW ZEALAND (2002)</td>
<td>NZ</td>
<td>0.28 to 0.34</td>
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<td>NORWAY (2001)</td>
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<td>0.48 to 0.52</td>
<td>0.09 to 0.14</td>
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<tr>
<td>PERU (2006)</td>
<td>PE</td>
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<td>0.05 to 0.11</td>
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<td>0.07 to 0.15</td>
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<td>PORTUGAL (2002)</td>
<td>PT02</td>
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<td>0.23 to 0.31</td>
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<td>PORTUGAL (2005)</td>
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<td>0.26 to 0.32</td>
<td>0.21 to 0.26</td>
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<td>ROMANIA (2004)</td>
<td>RO</td>
<td>0.08 to 0.27</td>
<td>0.16 to 0.29</td>
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<tr>
<td>RUSSIA (2004)</td>
<td>RU</td>
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<td>0.14 to 0.19</td>
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<td>SLOVENIA (2004)</td>
<td>SI</td>
<td>0.35 to 0.42</td>
<td>0.25 to 0.32</td>
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<td>SPAIN (2004)</td>
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<td>0.19 to 0.26</td>
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<td>SWEDEN (2002)</td>
<td>SE</td>
<td>0.42 to 0.47</td>
<td>0.29 to 0.36</td>
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<td>Country/Year</td>
<td>Code</td>
<td>Lower 95%</td>
<td>Upper 95%</td>
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<td>SWITZERLAND (2003)</td>
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<td>GREAT BRITAIN (2005)</td>
<td>GB</td>
<td>0.25</td>
<td>0.33</td>
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<tr>
<td>UNITED STATES (2004)</td>
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<td>0.19</td>
<td>0.27</td>
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</table>

*: Confidence interval of the estimate is obtained through the fifth lowest and the fifth highest estimate across 200 resamplings from the individual-level survey data.
Table 2: Multivariate regressions of the Impact of Policy Preferences on political system characteristics

<table>
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<tr>
<th></th>
<th>b</th>
<th>(s.e.)</th>
<th>beta</th>
<th>b</th>
<th>(s.e.)</th>
<th>beta</th>
<th>b</th>
<th>(s.e.)</th>
<th>beta</th>
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</thead>
<tbody>
<tr>
<td>Executive/parties</td>
<td>0.05***</td>
<td>(0.02)</td>
<td>0.42</td>
<td>0.05**</td>
<td>(0.02)</td>
<td>0.37</td>
<td>0.03*</td>
<td>(0.02)</td>
<td>0.26</td>
</tr>
<tr>
<td>Federal/unitary</td>
<td>-0.04**</td>
<td>(0.02)</td>
<td>-0.35</td>
<td>-0.03*</td>
<td>(0.02)</td>
<td>-0.23</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Polarization</td>
<td></td>
<td></td>
<td></td>
<td>0.08***</td>
<td>(0.03)</td>
<td>0.44</td>
<td></td>
<td></td>
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</tr>
<tr>
<td>Constant</td>
<td>0.30***</td>
<td>(0.02)</td>
<td>0.30***</td>
<td>(0.02)</td>
<td>0.14**</td>
<td>(0.05)</td>
<td></td>
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Adjusted R^2           | .18   | .30    | .45  |

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