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The Partisan Consequences of Turnout Revisited*

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Abstract

Why does the leftist party vote increase when turnout increases in some countries and not in others? Why does this happen in some instances in time but not in others? Thus far there exists no academic consensus on the relationship between turnout and electoral results. This paper argues that in order to adequately address these questions we need to focus on three elements: class voting, the mechanisms behind whether the correlation is observed over the short or long-term, and the use of more rigorous model specifications. By looking at the cases of Spain and Portugal, we find a correlation in the short and long-term for Spain but not for Portugal and this is due namely to the prominence of class voting in the former.

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1. Introduction

One of the most controversial propositions in the voting behaviour literature is that the higher the electoral turnout is the higher the left party vote share will be. The argument is parsimonious and persuasive since it is generally accepted that citizens with a higher socio-economic status (SES) are more likely to vote than those with a lower SES. Further it is assumed that citizens will have different ideological leanings based on their SES which is likely to manifest into party support. With this in mind, the argument proposes that low turnout biases election outcomes such that right-wing parties gain at the expense of left-of-centre alternatives (Rubenson et al, 2007: 595). However, the empirical evidence supporting the partisan consequences of turnout is far from conclusive. In a piece using data on national elections in 23 OECD countries, Fisher (2007) only finds statistically significant positive correlations between left share of the vote and turnout in five countries. Thus, after decades of research "there exists little scholarly agreement about either the partisan consequences of high turnout or its effect on incumbents in general" (Hansford and Gomez, 2010: 268).

In this paper we argue that this disagreement has to do with three elements: the assumption that class voting exists with the same strength everywhere[1], the different mechanisms behind the correlation between turnout and electoral results depending on whether the correlation is examined in the short or long-term, and the use of inappropriate statistical models to account for the partisan consequences of turnout. First, class voting differs across countries. All else equal, the correlation between turnout and electoral results should be higher as class voting increases. Furthermore, if class voting is not important at all, then voting for one party or for the other does not make a difference in terms of class, therefore we should not expect a significant correlation. However, class inevitably is assumed to mold individual voting behavior in all societies in the same way. Lipset's (1981 [1960]: 220-224) words that "in virtually every economically developed country the lower income groups vote mainly for parties of the left" captures the conventional wisdom. Consequently, as one's socio-economic status is positively linked with voting, leftist parties should always benefit with high turnout. But why does turnout have partisan consequences in some countries, but not in others, during some periods but not during others?

Second, as explained by Fisher (2007: 598-600), when assessing the relationship between turnout and the left share of vote, there are two possible questions: whether the left share of the vote is higher or not when turnout is relatively high (a long-term relationship) and whether the left share of the vote tends to increase between elections when turnout rises (a short-term relationship).[2] We argue that the answers to these two questions are not necessarily similar. As Pacek and Radcliff (1995) or Martinez and Gill (2005), our argument is that the long-term relationship is only a function of class voting; in the long-run leftist and parties alternate in government and then the impact of short-term factors, mainly the incumbent effect, cancels out or at least tends to zero. On the contrary, given the existence of class voting and therefore a correlation between turnout and the left share of the vote, the short-term relationship depends on who the ruling party is as higher turnout is associated with lower vote share for the incumbent. In sum, it is not possible to a have a compelling conclusion about the correlation in a given election in a particular country without clarifying if it is examined in the short or the long-term, and secondly, if the short-term is selected, without taking into account the anti-incumbent effect. In methodological terms, the main implication is that there are a variety of model specifications that could seem to be reasonable for the correlation between turnout and the left share of the vote. Given that any particular specification rests on assumptions about how the two variables are connected (for instance, if they are differenced we are assuming that the relationship is in the short-term, but not in the long-term), a compelling test demands not having ex ante assumptions.

We test our argument with a comparison between Portugal and Spain, two third-wave democracies with strong differences in class voting; irrelevant in Portugal and significant in Spain. Relying on cross-sectional and time series cross-section analyses in which assumptions about whether the partisan consequences take place in the short or long-term are examined, our findings show that the correlation between turnout and the left share of the vote exists in Spain both in the long and short-terms, but not in Portugal, neither in the short nor in the long-term.

The paper is organized as follows. In the next section we present our argument explaining cross-national differences in the correlation between turnout and the left share of the vote. The following sections describe the data and methods, the results of the empirical analysis using aggregated data, and an individual-level data analysis of the causal mechanisms driving the relationship between turnout and leftist support. The final section concludes and offers some empirical extensions.

2. Arguments

The argument linking voter turnout and electoral results rests on three substantive assumptions. First, there exists some degree of social and/or economic inequality in countries. If there are no inequalities in society then we cannot expect to find a positive correlation between turnout and electoral results because the differential benefit in terms of class for an individual associated with the election of various legislators/governments would be zero. All else equal, the higher the degree of social and/or economic inequalities, the higher the correlation between turnout and electoral results. Second, voters and nonvoters can be identified by their SES. Accordingly, people of a lower SES have a lower propensity to vote than those of a higher socioeconomic status. If the SES of voters and nonvoters is the same, then turnout levels should not affect electoral results. Third, class voting takes place in countries and therefore people of a lower socioeconomic status have a higher propensity to cast ballots for leftist parties than those with a higher SES. If, when citizens do vote, they are not voting according to their SES, then there would not be a turnout-electoral results correlation. [3]

While the first assumption is always met, the other two are more problematic. The positive relationship between SES and turnout has been repeatedly demonstrated (see Blais, 2000), although, as Nevitte et al (2009) or Gallego (2010) show, there are significant differences in the extent to which SES accounts for the variance in non-voting across countries. However, this correlation is influenced by the existence of strategic behaviours encouraged by electoral systems (Cox, 1997 1999) and the number of viable parties in a district/polity. Two individuals with the same SES may vote or abstain depending on whether parties and voters behave strategically or sincerely. As a consequence, the socioeconomic gap between voters and nonvoters would be reduced as well as the correlation between turnout and electoral results. The level of elite mobilization effort is predicted to increase in closer elections. Elite effort boosts turnout because voters respond to the act-contingent incentives, those marshaled by political parties as part of explicit get-out-the-vote efforts (Cox, 1999: 389-90). Given that closeness varies across districts within countries, the cost of voting can be different for two individuals with the same socioeconomic status and therefore their probability of voting. This impact of closeness is exacerbated if there is strategic abstention, with voters not showing up for non-competitive elections.[4]

More importantly, the strength of the relationship between turnout and electoral results should be negatively correlated with the number of viable parties in a district or polity, If there are only two parties, one rightist and the other leftist, voters dissatisfied with his/her "natural" party will have a higher probability of being abstainers (and the other way around) than when there are other (minor) parties that can channel the dissatisfaction. In other words, the correlation between turnout and electoral results should be higher when a voter has two possible actions –loyalty or exit- instead of three –loyalty, voice or exit.

Similarly, social class does not shape voting behaviour by default. As explained by Przeworski and Sprague (1986: 7-9, 11), "class, religion, ethnic, race, or nation do not happen spontaneously as reflections of objective conditions in the psyches of individuals … The organization of politics in terms of class is not inevitable … the salience of class as political behaviour can be attributed to the strategies pursued by political parties, especially parties of the Left". More recently, Anderson and Beramendi (2012) have shown that countries' income distributions and the presence of left party competition provide different incentives for left parties to mobilize lower income voters: the association between income inequality and turnout is muted by the presence of several parties on the left side of the political spectrum. Accordingly, whether citizens vote according to their SES is an empirical issue and clearly it should vary across countries and over time.

Finally, when analyzing the partisan consequences of turnout, controlling for the anti-incumbent effect is crucial. As outlined by Grofman et al (1999), all else equal, higher turnout will be associated with lower vote share for the incumbent party, independently on whether it is a leftist or a rightist party. There are two mechanisms for this expectation (Hansford and Gomez, 2010: 270-1). First, the conditions that cause voters to reject the incumbent party may also cause more voters to turn out at the polls. Second, since core voters are on average more supportive of the governmental status quo than peripheral voters, the more peripheral voters are involved in an election, the worse the incumbent party's candidate will do.

Taking into account simultaneously the SES model and the anti-incumbent effect, the partisan consequences of turnout in the short-term will be particularly important when both variables push in the same direction, that is, when turnout is high and the incumbent party is rightist. On the contrary, the sign of the correlation between turnout and the left share of the vote will be not clear when turnout is high and the incumbent party is leftist, since the SES model predicts a positive correlation and the anti-incumbent effect a negative one. In sum, as shown in Table 1, and assuming the existence of class voting, there is a clear interaction between the SES model and the anti-incumbent effect when linking voter turnout and electoral results.

Table 1:
The partisan consequences of turnout in the short-term
(assuming the existence of class voting)*

		Turnout (mobilization of peripheral voters)	
		High	Low
Incumbent	Leftist	? (?)	? (?)
	Rightist	+ (–)	– (+)

*In each cell, the first sign is the expected impact for the left share of the vote; in parentheses, the impact for the right share of the vote.

On the basis of these arguments, the correlation between turnout and the left share of the vote in the long and short-terms can be formulated as follows:

- The left share of the vote is higher (lower) on average when turnout is high (low) if and only if social class shapes voting behaviour. If *n* elections are studied, the incumbent effect is canceled or at least tends to zero because there are alternating leftist and rightist governments. That is, *class voting is a necessary and sufficient condition for observing the correlation between electoral results and turnout in the long-term.*

- The left share of the vote does not increase (decrease) between elections when turnout rises (decreases) if social class does not shape voting behaviour, independently of the incumbent effect. That is, *no class voting is a necessary and sufficient condition for not observing the correlation between electoral results and turnout in the short-term.*

- The left share of the vote increases (decreases) between elections when turnout rises (decreases) if social class shapes voting behaviour and the leftist party is the challenger. However, if the leftist party is the incumbent, the sign of the correlation is not clear, since the impact of class voting and the incumbent effect go in opposite directions. That is, *class voting plus a leftist challenger generate a positive correlation between electoral results and turnout in the short-term, while class voting plus a leftist incumbent generate a weaker correlation with an unpredictable sign.*

In Table 2 these different combinations of the long and short-term correlations between the left share of the vote and turnout in three hypothetical countries are displayed. In Country C the correlation exits both in the long and short-term, in Country A only in the long-term, and in Country B neither in the long nor the short-tem. Accordingly, model specification in statistical analyses has to respond to these different patterns. For instance, a model in which the left share of the vote and turnout are differenced partially captures the relationship between the two variables in Country C, but would lead to wrongly conclude that the correlation does not exist in Country A, In sum, four conclusions emerge from here: (i) there is not a straightforward relationship between turnout and electoral results; (ii) the determinants of the correlation in the long and short-term are not identical, (iii) class voting is a necessary and sufficient condition for the correlation a in the long-term, but only necessary in the short-term, and (iv) the selection of a particular model specification is not an option for the researcher, but imposed by data.

Table 2. Three hypothetical cases of correlation between turnout and the left share of the vote

Countries	(%)	Elections			Correlation	
		t	t+1	t+2	Long-term	Short-term
A	Turnout	80	78	80	Yes	No
	Left share of the vote	45	46	45		
B	Turnout	50	52	50	No	No
	Left share of the vote	45	45	45		
C	Turnout	78	80	78	Yes	Yes
	Left share of the vote	45	46	45		

3. Data and measures

In order to better address the puzzle of the partisan consequences of turnout, we look at data at the district level in Lower House elections within two individual countries, Portugal (1975-2009)[5] and Spain (1977-2008).[6] There are four reasons for this research design. First, cross-sectional studies of turnout are subject to limitations, particularly the omission of important factors (Blais, 2006). Second, because it is easier to register in some countries than in others, turnout measures are not strictly comparable (Blais and Aarts, 2006). Third, to the best of our knowledge, data measuring the strength of class voting across a significant number of countries and over time (decades) are not available. Fourth consequently, the selection of Portugal and Spain allows us to focus on the variation in class voting amongst others that do not differ. The two countries are third-wave democracies with similar electoral systems, but with a considerable variation in the impact of social class on party choice.[7]

Unfortunately, the first comprehensive scientific national election survey in Portugal was not conducted until 2002 (Lewis-Beck and Lobo, 2011: 294). Therefore, instead of measuring class voting over time using logistic modeling techniques such as the Kappa or the Lambda indexes, for instance, (Evans, 2000), the different saliency of class voting in Portugal and Spain will be shown according to the existing research. Using different methods and data, the finding that class voting is weak or even absent in Portugal and stronger in Spain is largely consensual in the literature. First, according to Freire (2006: 364-365), the weight of social class (a typology based on occupation and number of employees) in explaining individual left-right placement is three/four times higher in Spain than in Portugal in different moments in time. Not surprisingly, Portugal is at the bottom in the sample of 12 countries and Spain is at the top. Second, Gunther and Montero (2001: 120) show that class explains the 16 percent of the vote in 1983 in Portugal and the 6 percent in 1993, while in Spain the percentages are 19 and 10, respectively. Third, relying on multilevel models of voting behaviour in Southern Europe in the period 1985-1999, Freire and Costa Lobo (2005: 510-11) conclude that "in the Portuguese case, the impact of social class' indicators on the vote is never significant ... in the Spanish case ... cleavage voting is more important than in Portugal: contrary to the latter, in the former case both education and head of household income have a significant impact on the vote ... class cleavage is more important in Spain than in Portugal". In the same vein, according to Knutsen and Scarbrough (1998: 504-505), the coefficients for both the bivariate and the "controlled" effects of social class on party choice –measured by occupation, education, and household income- is more than double in Spain than Portugal. While Portugal shows the weakest correlations in the sample of 13 countries (0.06), Spain is in the middle (0.12, 0.14). Fifth, in a recent piece estimating dynamic, multi-equation models with two-stage, instrumental variable regression procedures, Lewis-Beck and Lobo (2011: 299:301) found that social structures measures (education and income) does not affect voting in the 2005 legislative election. Finally, using qualitative evidence, Fishman (2011) in his comparison of democratic outcomes in Spain and Portugal post dictatorship, also concludes that Spain and Portugal have different underlying societal class dynamics. He finds that the inverse hierarchical class structure that was established in Portugal during the revolution with its consensual and inclusive elements, cooled animosities among classes which still persists today. Whereas, in Spain, Fishman argues, the democratization was state-constructed and less-consensual, and therefore caused the marginalization of certain groups. This theory complements Chhibber and Torcal´s assertion that it was the political parties in Spain (particularly the socialist party) that eventually tapped into these societal class differences and managed to capitalize on them through mobilization and the politicization of social divisions.

On the other hand, elections in Spain and Portugal are held by D´Hondt formula in one-tier electoral systems with closed party lists. The 52 districts in Spain (2008 election) range from 1 to 35 seats, while in Portugal (resident in 2009 election) the 20 districts range from 2 to 47 seats. Mean district magnitude is 6.7 Spain and 11.3 in Portugal.[8] While there is a sizable difference in the mean district magnitudes of Spain and Portugal, we do not foresee this being a problem. If the effective number of parties at the district level were linearly correlated with district magnitude, we may expect turnout in districts to increase with the number of seats and the number of parties competing, however Selb and Grofman (2011) have recently found that this is not the case and this relationship is non-linear. They find that district magnitude plays a role in shaping the relationship between turnout and the effective number of parties when a district magnitude is equal to one or when a district magnitude is greater than one. But it is not expected that turnout will increase as district magnitude increases. Additionally, there is a (virtually irrelevant) 3% threshold at the district level in Spain, but not in

Portugal. In sum, although party systems in Portugal and Spain are not identical at the national level nor at the district level (more national leftist parties in Portugal than in Spain and strong sub-national parties in Spain and not in Portugal), they share the crucial characteristic demanded by our argument: the existence of one main party on the left and the right and at least one challenger on the left channelling the dissatisfaction with the Socialist Party instead of staying at home. Similarly, as they have quite similar electoral systems, the impact of closeness is constant.

Nevertheless, there are some obvious differences in the institutional arrangements of Portugal and Spain. While Spain is parliamentary and decentralized, Portugal is semi-presidential and unitary. Whether a country is more decentralized does not seem to have an impact on turnout in national elections (Blais and Carty 1990; Black 1991), so we do not see this as impacting our results. Research is thin on presidential systems and the potential impact on turnout. As Blais et al (2011: 301) summarize, "no one work has carefully tested whether turnout declines in legislative elections when there is a powerful president". For instance, Anduiza (1999:162 ff) considers the institutional relevance of national parliaments as inversely related to the existence of an elected president, regional parliaments with political autonomy, and direct democracy institutions. However of these variables only the last one remains a significant predictor of turnout in national elections in Western Europe. Things are even less clear when the question is in what way –if any– turnout may be affected by a semi-presidential system.

In sum, given that we are not accounting for differences in turnout between the two countries, but differences in the correlation between turnout and the left share of the vote, with the exception of the party system, institutional variables do not play any role as they are constant over time. For instance, the impact of having a parliamentary or a semi-presidential system should be the same in the founding election as in the following elections.

According to the previous discussion, and given the different saliency of class voting in both countries, our expectation is that the correlation between turnout and the left share of the vote should be weak in Portugal both in the long and the short-term. In Spain it should be much stronger in the long-term and, when the Socialist Party is the challenger, also in the short-term.

4. Estimation methods and results

Testing the argument that left parties benefit from high turnout requires that we properly identify whether the causal effect of turnout rates on the left share of the vote takes place in the short or long-term or both. To address this issue in our estimations, cross-sectional and time series-cross section models are estimated.

Cross-sectional analyses

When analyzing the long-term relationship between turnout and the left share of the vote, we run the following model:

$$Left_i = \alpha + \beta Turnout_i + \varepsilon_i \qquad [1]$$

Here *Left* is the share of the vote of the main leftist party in national elections, the Socialist Party, PS in Portugal and PSOE in Spain, respectively;[9] *Turnout* is the percentage of registered voters who cast votes in national elections, and ε is a residual error term. Districts are indexed by *i* = 1,… J. As the two variables are district-level averages for all the elections, this model captures whether the left share of the vote is higher on average when turnout is high. The number of observations is 20 in Portugal and 52 in Spain.

On the other hand, when analyzing the short-term relationship, we run three models:

$$\Delta Left_{it} = \propto + \beta \Delta Turnout_{it} + \varepsilon_{it} \qquad [2]$$

This is the same as Model 1 but with the dependent and independent variables differenced (by subtracting the value at the previous election). Districts are indexed by *i* = 1, … J and elections are indexed by *t* = 1,… As the two variables are differenced, this model captures whether the left share of the vote increases between elections when turnout rises. The incumbent effect is not controlled here. The number of observations is 220 (20 districts × 11 elections) in Portugal and 468 (52 districts × 9 elections) in Spain.[10]

$$\Delta Left_{it} = \propto + \beta \Delta Turnout_{it} + \gamma Governing_{it} + \varepsilon_{it} \qquad [3]$$

Here the variable *Governing* (1 if the Socialist Party is the governing party; 0 otherwise) is added to the previous specification. That is, the incumbent effect is included in the model. The number of observations is the same as in model [2].

$$\Delta Left_{it} = \propto + \beta\Delta Turnout_{it} + \gamma Governing_{it} + Interaction_{it} + \varepsilon_{it} \quad [4]$$

Here an interactive term (between $\Delta Turnout + Governing)$ is included. We are testing then to what extent the correlation between turnout and the left share of the vote changes depending on whether the socialist party is the incumbent or the challenger. The number of observations is the same as in models [2] and [3].

Descriptive statistics are shown in Table 3. The most relevant thing is that the within-variation in *Turnout* in Portugal is double than in Spain, but the between-variation is lower in the former. In other words changes across elections in *Turnout* are higher but more homogeneous across electoral districts in Portugal than in Spain.

Table 3: Descriptive statistics
Spain (52 districts ×10 elections = 520 observations); Portugal (20 districts ×12 elections = 240 observations)

Variable	Mean	Std. deviation (overall)	Std. deviation (within)	Std. deviation (between)	Minimum (overall)	Maximum (overall)
Spain						
Left	38.31	9.75	6.77	7.07	14.50	63.67
Turnout	73.68	7.05	4.78	5.23	42.20	87.60
Governing	0.60	0.49	0.49	0.00	0	1
Portugal						
Left	33.55	9.76	8.76	4.39	13.20	56.00
Turnout	71.33	11.34	10.73	3.76	43.90	95.27
Governing	0.55	0.50	0.50	0.00	0	1

Results

The least squares method is highly unsatisfactory due to the presence of outliers which can be supposed in the analysis of the level of nationalization in the sample of countries. The residuals plotted against the fitted values exhibited some outliers. In such a case, the robust regression is an acceptable and useful tool because it provides a good fit to the bulk of the data and exposes the outliers quite clearly.[11]

The estimation results of model [1] presented in Table 4 strongly support our argument. As predicted, in Spain the left share of the vote is significantly higher when turnout is high: one point increase in turnout increases the vote share of the Socialist Party by 0.51. The variable is statistically significant at the 0.01 level. However, in Portugal the relationship, although positive, is not statistically significant. Consequently, the partisan consequences of turnout are not particularly relevant in Portugal.

Table 4:

The partisan consequences of turnout in the long-term for the Socialist Party in Portugal and Spain

	Models	
Variables	Portugal	Spain
Turnout	0.45	0.51***
	(0.28)	(0.19)
Constant	1.66	-0.30
	(20.24)	(12.42)
F	2.50	7.55***
N	20	52

Notes: Robust regression. Standard errors are in parentheses. ***p<0.01.

The short-term relationship between the left share of the vote and turnout (models [2], [3], and [4]) is displayed in Table 5. Again, as expected, there is considerable support for the partisan consequences of turnout in Spain, but not in Portugal. In Spain, all of the model specifications indicate that left share of the vote is significantly correlated with turnout. According to the first model, the difference in turnout in a given district is statistically significant at the 0.01 level and has the expected positive sign: one point increase in turnout increases the vote share for the Socialist Party by 0.53. The coefficient for the difference of turnout and its statistical significance do not change appreciably when controlling for whether the Socialist Party enters an election as the governing or an opposition party. As shown in model 2, when the PSOE enters an election as the governing party, its results are six points worse than when it enters an election from opposition. The variable is statistically significant at the 0.01 level. Finally, model 3 shows the interaction between if the Socialist Party enters the election governing or not and the difference in participation. The interaction is statistically significant, as well as its constitutive elements. Figure 1 shows how the marginal effect of the difference in turnout changes depending whether the Socialist Party enters an election governing or not.

Table 5:

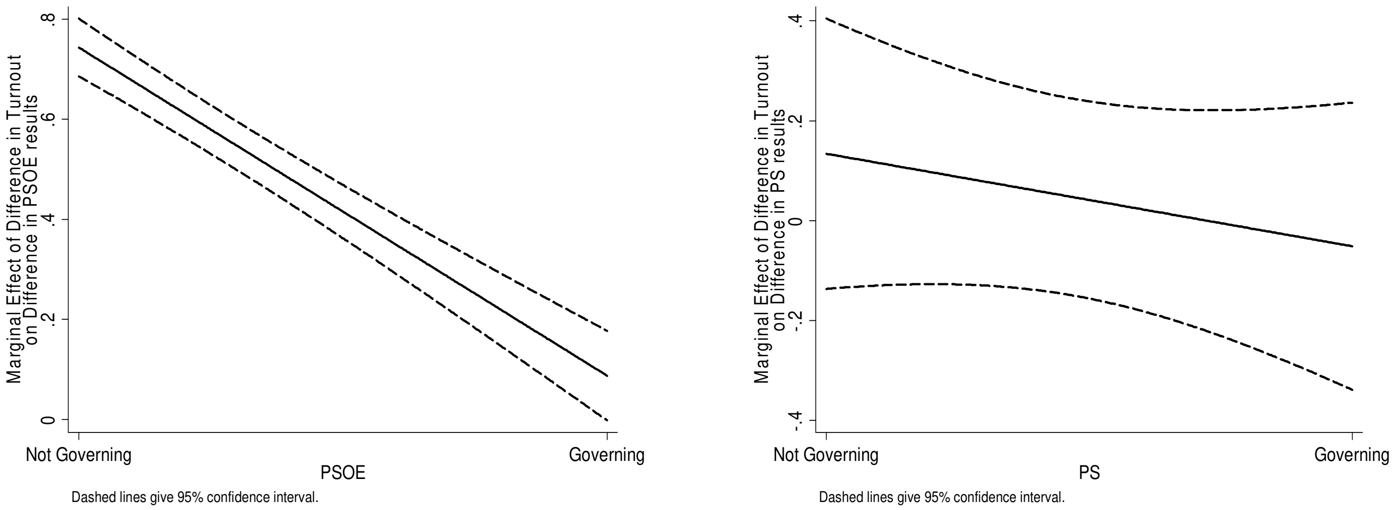
The partisan consequences of turnout in the short-term for the Socialist Party in Portugal and Spain

	Portugal			Spain		
Variables	1	2	3	1	2	3
Δ Turnout	-0.14	0.06	0.13	0.53***	0.54***	0.74***
	(0.14)	(0.10)	(0.14)	(0.04)	(0,03)	(0.03)
Governing		-15.28***	-15.70***		-6.00***	-6.12***
		(0.88)	(1.00)		(0.45)	(0.38)
Interaction			-0.19			-0.66***
			(0.20)			(0.05)
Constant	-0.25	8.02***	8.26***	1.49***	4.96***	4.82***
	(0.80)	(0.69)	(0.75)	(0.27)	(0.33)	(0.28)
F	1.01	152.33***	101.49***	214.99***	260.72***	298.89**
N	220	200	200	468	468	468

Notes: Robust regression. Standard errors are in parentheses. ***p<0.01.

In Portugal, as expected, the results of the Socialist Party do not depend on the level of turnout. The difference in turnout is not statistically significant in any of the three models and it only has the expected positive sign when controlling for whether the Socialist Party enters an election as the governing or an opposition party. Finally, the interaction between turnout and being the governing party is not statistically significant. As shown in Figure 1, this means that turnout does not play any role when explaining the vote share for the PS.

Figure 1: The marginal impact of the change in turnout on the change in the left share of the vote in Spain (left) and Portugal (right)



Time series-cross section analyses

When using averages for every district in the whole period it could happen that results were spurious if turnout and the left share of the vote are dominated by long-term social trends. And even if the results are not spurious, cross-section estimates do not inform us about the existence and the velocity of adjustments in *Left* to changes in exogenous variables. Similarly, when using differences in turnout and the left share of the vote some parametric assumptions are imposed without being tested. First, the coefficient of the lag of the dependent variable is equal to 1. As explained by Fisher (2007), this is a problematic assumption, since parties that did well in one election are more likely to go back down at the next election rather than continue to rise. Second, the coefficients of turnout and the lag of turnout are equal in absolute values, but with opposite signs. If these two assumptions are not true, estimates would be biased. Accordingly, we have explored the robustness of our findings replacing cross-sectional estimates with time series-cross section (TSCS) analyses.

The point of departure to analyze the dynamic relationship between turnout and the left share of the vote is the following general specification in which no assumptions are imposed:

$$Left_{it} = \alpha_i + \rho Left_{it-1} + \beta Turnout_{it} + \gamma Turnout_{it-1} + \varepsilon_{it} \qquad [5]$$

Where districts are indexed by i = 1, … J and elections are indexed by t = 1,… α_i are the individual fixed effects. Since the districts are a complete set and not a random sample from a wider population, a fixed effects model is more appropriate than a random effects model. As is well-known, when the lagged endogenous variable is on the right side of the equation, the initial impact or the short-term impact of the change in the regressor x is given by its coefficient, while the steady-state or long-term impact depends on the value of ρ. In [5], the short-term effect of changes in *Turnout* on *Left* is β and the long-term impact is $\frac{\beta}{1-\rho}$.

Departing from [5] three models can be derived:

1. If $\rho = 1$ and $\gamma = -\beta$, then we obtain the "differenced model":

$$\Delta Left_{it} = \alpha_i + \beta \Delta Turnout_{it} + \varepsilon_{1t} \quad [6]$$

In this model both the dependent and independent variables are differenced by subtracting the value at the previous election. As the two variables are differenced, this model captures whether the left share of the vote increases between elections when turnout rises. This is our previous model [2].

2. If $\rho = 0$ and $\gamma = 0$, we have the "model in levels":

$$Left_{it} = \alpha_i + \beta Turnout_{it} + \varepsilon_{1t} \quad [7]$$

Here the adjustment of *Left* to changes in *Turnout* is instantaneous, that is, short-term and long-term multipliers are the same.

3. Finally, if $0 < \rho < 1$ and $\gamma = -\beta$ then we obtain the "semi-differenced model":

$$Left_{it} = \alpha_i + \rho Left_{it-1} + \beta \Delta Turnout_{it} + \varepsilon_{1t} \quad [8]$$

In this model short-term and long-term multipliers differ, while the level of *turnout* is irrelevant. Only changes in this variable has an impact on *Left* .

Hence, the models [6], [7], and [8] are specifications nested in the general model [5]. Given that there are no *ex ante* reasons to select one of them, some preliminary tests are necessary before imposing constraints on the parameters ρ and γ. Finally, the variable *Governing* (1 if the Socialist Party is the governing party; 0 otherwise) is added to the previous specifications as a control variable to capture the "incumbent effect".

The first step is to test for unit root processes in both *Left* and *Turnout* in order to determine: (i) if we are dealing with integrated or stationary series, (ii) if the order of integration is the same, (iii) if they are cointegrated or not, and (iv) if differencing *Left* is appropriate or not. Two unit root tests for panel data have been run. The Levin-Lin-Chu (2002) test or LLC assumes that each individual unit in the panel shares the same AR(1) coefficient, while the Im, Pesharan and Shin (2003) test or IPS allows for different AR(1) coefficients in each series. Both tests allow for individual effects, time effects and possibly a time trend and assume that all series are non-stationary under the null hypothesis. The null hypothesis in both cases is that series are integrated of order 1 or I(1). In table 6, individual and time effects are included, but not time trends or lags of the dependent variable. The p values and the t-star statistic and the W[t-bar] when suing the LLC test and the IPS test, respectively, are shown.

Table 6: Unit root tests: Series are I(1) under the null hypothesis in all cases

Variable	LLC t-star and p-value	IPS W[t-bar] and p-value	Observations (t*N)
Spain			
Left	-8.58 (0.0000)	-4.65 (0.000)	9*52=468
Turnout	-9.33 (0.0000)	-5.81 (0.000)	9*52=468
Portugal			
Left	-6.68 (0.0000)	-3.78 (0.000)	11*20=220
Turnout	-7.09 (0.0000)	-4.51 (0.000)	11*20=220

Clearly, the null hypothesis has to be rejected meaning both variables are stationary in both countries. Hence, the problem of spurious regressions and the potential lack of cointegration are not a concern. Additionally, using the lagged endogenous variable in levels on the right-side of the equation (as in specifications [5] and [8]) is more appropriate than differencing it (as in specification [6]).

Results for Spain

The results of the estimates of models [5] and [8] are displayed in Table 7. Individual fixed effects are highly significant, according to an F-test on the null hypothesis of irrelevance: $F(51, 413) = 3.23$ p-value<0.0000). Following Greene (1997: 598), we have calculated a modified Wald statistic for groupwise heteroskedasticity in the residuals. According to the results, the null hypothesis of homoskedasticity can be rejected (p-value < 0.0001). Moreover, we have computed the Breusch-Pagan statistic for cross-sectional independence in the residuals of a fixed effect regression model (Greene, 1997: 601). The null hypothesis can be rejected (p-value<0.0001).

However, serial correlation of residuals does not seem problematic. When computing the modification of the Breusch-Godfrey test proposed by Greene (1997: 517), the existence of a common AR(1) process in residuals may be discarded.[12] Contrary to what Hansford and Gomez (2010) argue, endogeneity of variable *Turnout* may be also rejected according to the Hausman test, while multicollinearity is not a serious concern according to estimates of multiple correlations among regressors. For each regressor the coefficient of determination of the auxiliary regression on the rest of right-hand variables was calculated. All of them were below 0.59.

Finally, three more problems have been addressed: (i) possible biases in the coefficients due to the estimation of first-order autoregressive models with fixed effects (Nickell, 1981); (ii) panel heteroskedasticity; and (iii) contemporaneous cross-correlation. To deal with problems (ii) and (iii) Panel-Corrected Standard Errors (PCSE) can be used instead of Ordinary Least Squares (OLS) standard errors, following the methodology proposed by Beck and Katz (1995). In Table 7 t-statistics computed using PCSE are shown in brackets. While PCSE are substantially higher than standard errors, all independent variables are significant at the 0.05 level or better. Moreover, according to the Wald tests, we may assume the hypothesis $\gamma = -\beta$ and the hypothesis $0<\rho<1$. In other words, instead of using *Turnout* and $Turnout_{t-1}$ in levels, first differences can be used. On the contrary, differencing *Left* is not supported by the data.

Second, the coefficients do not change appreciably depending on whether biases are corrected following the proposal by Kiviet (1995) in columns 2 and 3 of Table 7.[13] Insofar as T=9, a bias of order T^{-1} is not as problematic as the most often cases of T=3 or T=6 when working with microdata (Beck and Katz, 2011). Using estimates of coefficients ρ and β in column (3), one point increase in turnout increases by 0.473 points the vote share for the Socialist Party in the short-term (the same election) and by.473 / (1-.465) = 0.88 points in the long-term. The lag of the dependent variables and turnout, independently on how it is defined, are statistically significant at the 0.05 level or better in all columns.

Additionally, to explore the robustness of our findings we have calculated the system GMM estimator in column 4.[14] We compute the two-step estimator and the covariance matrix robust to any pattern of heteroskedasticity and autocorrelation within panel. Unfortunately, we cannot correct for contemporaneous correlation across panels. The only endogenous variable is *Left*, which is instrumented with its lagged values. We also include as additional instruments the first and second lags of *ΔTurnout*, and a time trend. The results for both the Arellano-Bond test for AR(2) in first differences, and the Hansen test of overidentification restrictions discard problems in both senses. The results do not change appreciably: the short-term effect is weaker (0.28), while the long-term is stronger (0.66).

Finally, in columns 5 and 6 the variable *Governing* is added to the previous models. In column 5 the LSDV estimator with PCSE is used, while in column 6 it is replaced with the Kiviet's bias correction. The variable *Governing* is statistically significant at the 0.01 level and increases the R^2 from 0.752 to 0.809. Using the corrected coefficients in column 6, when *Governing* = 1 (i.e., when the Socialist Party is the governing party), the left share of the vote increases by 5.8 points. Not surprisingly, the inclusion of *Governing* reduces the magnitude of the effect of both the lagged endogenous and *ΔTurnout*. The short-term effect is now 0.25 and the long-term effect, 0.36.

In sum, as in our cross-sectional analysis in Table 4 and 5, our results strongly support the partisan consequences of turnout in Spain; the left share of the vote is significantly correlated with turnout. According to the Wald test, the best specification to deal with this correlation in Spain is [8]. The difference in turnout in a given district is statistically significant at the 0.01 level and has the expected positive sign. When the variable *Governing* is included, and the potential bias in the coefficients is corrected, one point increase in *ΔTurnout* increases the left share of the vote by 0.255 in the short term and by 0.356 in the long-term.

Table 7: The partisan consequences of turnout in Spain

	Models					
Variables	1	2	3	4	5	6
$Left_{t-1}$	0.372 (10.53)*** [2.54] **	0.455 (7.15)***	0.465 (9.10)***	0.580 (12.50)***	0.234 [2.17]**	0.284 (9.02)***
Turnout	0.587 (11.21)*** [3.27] ***	0.569 (11.78)***				
$Turnout_{t-1}$	-0.354 (7.13)*** [2.03]**	-0.389 (5.62)***				
ΔTurnout			0,473 (19.16)***	0.279 (7.51)***	0.248 [2.11]**	0.255 (6.57)***
Governing					6.035 [3.06]***	5.850 (9.15)***
Wald test. H_0: $\gamma = - \beta$ (p-value)	0.97 [0.325]	2.67 (0.102)				
Wald test. H_0: $\rho=1$ (p-value)	18.35 [0.000]	51.16 (0.000)				
Arellano-Bond test for AR(2) in first differences (p-value)				0.743		
Hansen test of overidentification restrictions (p-value)				0.290		
Observations (T*n=N)	9*52=468	9*52=468	9*52=468	7*52=364	9*52=468	9*52=468
R^2	0.752				0.809	
Method	LSDV	KIVIET	KIVIET	System GMM	LSDV	KIVIET

Notes: t-statistics computed using PCSE in brackets in columns 1, 5 and 6; t-statistics computed using ordinary standard errors in parenthesis in column 1; bootstrapped errors in parenthesis in columns 2, 3, and 6; robust z-statistics in parenthesis in column 4 using the Windmeijer's finite-sample correction for the two-step covariance matrix.
***p<0.01, **p<0.05, *p<0.1.

Results for Portugal

When studying the partisan consequences of turnout in Portugal, there are two crucial differences in comparison with what we have seen in Spain. First, *Turnout* and $Turnout_{t-1}$ are highly correlated (r = 0.91). In order to avoid problems of multicollinearity, the hypothesis $\gamma = - \beta$ has not been tested. Accordingly, in Table 8 the constraint $\gamma = 0$ is imposed in columns 1, 2, and 4 and the constraint $\gamma = - \beta$ is imposed in column 3 to show how the results change when *Turnout* is included in levels or first-differences.

Second, individual fixed effects are not as relevant as in Spain. In column 1 we cannot reject the hypothesis of irrelevance at the 0.05 level and they are even less significant in column 3, when *Turnout* is replaced with *ΔTurnout*.[15]

As in the estimates for Spain, the null hypothesis of cross-sectional independence of residuals is clearly rejected (p-value<0.0001) and serial correlation of residuals is even weaker. On the contrary, heteroskedasticity is not a problem now. The hypothesis of homokedasticity cannot be rejected (p-value=0.22), while residual autocorrelation is even lower than before.

Individual fixed effects are only included in column 1. Columns 2 to 4 are estimated by OLS including t–statistics computed with PCSE instead of OLS standard errors. When comparing the results for Spain with those for Portugal, it seems clear that it is better to start in the latter with the general model specification [5]. The results are very different from those corresponding to Spain. The lagged endogenous variable is statistically significant at usual levels, but not *Turnout* when PCSE are used. Moreover, the latter does hot have expected positive sign. Similarly, the variable *ΔTurnout* is not statistically significant both using PCSE or OLS standard errors, although it has the expected positive sign. These findings strongly support our argument about the crucial role of class voting when determining the partisan consequences of turnout. In Portugal, in congruence with the weak role of class voting, the left share of the vote is not correlated with turnout either in the long and short-terms. On the contrary, *Governing* is statistically significant at the 0.01 level. In sum, there is not evidence of a robust correlation between the left share of the vote and turnout in Portugal.

Table 8: The partisan consequences of turnout in Portugal

	Models			
Variables	1	2	3	[4]
Intercept		33.553	14.870	27.488
		[2.18]**	[2.13]**	[2.40]***
$Left_{t-1}$	0.318	0.493	0.557	0.384
	[1.37]	[2.76]***	[2.90]***	[2.85]***
Turnout	-0.377	-0.227		-0.175
	[1.58]	[1.25]		[1.27]
$Turnout_{t-1}$ [imposed]	[0]	[0]	[$-\beta$]	[0]
Δ Turnout			0.007	
			[0.02]	
Governing				9.83
				[3.25]***
Wald test. H_0: $\rho=1$	8.62			
(p-value)	[0.003]			
Observations (T*n=N)	11*20=220	11*20=220	11*20=220	11*20=220
R^2	0.466	0.384	0.324	0.614
Method	LSDV	OLS	OLS	OLS

Notes: t-statistics computed using PCSE in brackets.
***p<0.01, **p<0.05, *p<0.1.

5. Causal mechanisms and individual data analyses

According to our aggregated data analyses, while we found a positive correlation between turnout and the left party's share of the vote in Spain, this correlation does not exist in Portugal. However, the individual causal mechanisms accounting for this correlation have been hypothesized, but not shown. Based on individual data, in this section we show, first, that there exist social and economic inequalities in both countries, but they are translated into different political preferences (the left-right dimension) in Spain and not in Portugal as a result of the higher importance of class voting in the former. Second, when abstainers are mobilized, they tend to vote according to their socioeconomic status and then are more likely to support leftist parties since they tend to be more ideologically left.

In order to maximize the comparability of the data, we have selected two similar elections in each country. In the first one, the 2000 election in Spain and the 2002 election in Portugal, a rightist party won the election (the Popular Party and the Social Democratic Party, respectively), while in the second, the 2004 election in Spain and the 2005 election in Portugal, the rightist party was defeated by the Socialist Party. Additionally, in both countries turnout was higher in the second election than in the first one (from the 68.7 percent to the 75.7 in Spain and from the 61.5 percent to the 64.3 in Portugal).

Relying on the first and second round of the European Social Survey, Tables 9 and 10 show the placement in the left-right scale and the household's income of Spanish and Portuguese voters and abstainers. Both Tables show that abstainers earn less income than voters in both elections and both countries. That is, voters and non-voters can be identified by their economic status everywhere. The difference is statistically significant at the 0.01 level. However, in Portugal the economic status is not statistically correlated with the placement on the left-right scale, although abstainers are more leftist than voters. Interestingly, when turnout rises (in the 2005 election), the ideological gap between voters and abstainers survives, although decreases. In other words, as our argument suggests and the aggregated analyses have shown, the correlation between turnout and political preferences is weak in Portugal. On the contrary, in Spain there is an ideological gap between abstainers and voters when turnout is low (in the 2000 election), but it disappears when abstainers are mobilized (in the 2004 election). This evidence strongly supports our arguments and previous results.

Table 9: Voters and abstainers in the 2000 and 2004 elections in Spain

	2000 election			2004 election		
	Voters	Abstainers	Difference	Voters	Abstainers	Difference
Placement on left right-scale[a]	4.52 (1060)	4.04 (242)	0.48***	4.38 (1130)	4.57 (207)	-0.19
Feeling about household's income nowadays[b]	1.92 (1217)	2.05 (346)	-0.13***	1.82 (1253)	1.96 (276)	-0.14***

[a] (0, left - 10, right).
[b] (1, living comfortably on present income, 2, coping on present income, 3, finding it difficult on present income, 4, finding it very difficult on present income).
***p<0.01. In brackets, the number of individuals.
Source: European Social Survey, First and Second Round.

Table 10: Voters and abstainers in the 2002 and 2005 elections in Portugal

	2002 election			2005 election		
	Voters	Abstainers	Difference	Voters	Abstainers	Difference
Placement on left right-scale[a]	5.18 (887)	4.91 (270)	0.27	5.07 (970)	4.87 (275)	0.20
Feeling about household's income nowadays[b]	2.35 (1029)	2.53 (383)	-0.18***	2.36 (1302)	2.47 (512)	-0.11***

[a] (0, left - 10, right).
[b] (1, living comfortably on present income, 2, coping on present income, 3, finding it difficult on present income, 4, finding it very difficult on present income).
***p<0.01. In brackets, the number of individuals.
Source: European Social Survey, First and Second Round.

Given that ideological placements are not the same as party preferences, in Table 11 we show whether low turnout biases election outcomes such that right-wing parties gain at expense of left-of-centre parties in Spain. In this empirical analysis, we have used a post-electoral 2004 survey undertaken by *Demoscopia:* the European Social Survey does not contain voting records from the last two national elections[16]. The evidence is conclusive. The 40 percent of abstainers in the 2000 election voted for the Socialist Party and the 20 percent for the Popular Party in the previous election.

Similarly, the 60 percent of abstainers in the 2000 election voted for the Socialist Party in 2004 and only 20 percent for the Popular Party. In sum, the left share of the vote tends to increase (decrease) between elections when turnout rises (drops).

Table 11: Mobilization and demobilization in the 2000 and 2004 elections in Spain

1996 election	Abstainers in the 2000 election	2004 election	Abstainers in the 2000 election
Socialist Party voters	40 (175)	Socialist Party voters	60 (91)
Popular Party voters	27 (120)	Popular Party voters	20 (30)
Other parties voters	33 (146)	Other parties voters	20 (31)

First, the column percentages. In brackets, the number of individuals
Source: Centro de Investigaciones Sociológicas (2382-2384 study) for the 2000 election and Demoscopia for the 2004 election.

6. Conclusions

In this paper we have tested the partisan consequences of turnout for Portugal and Spain. We have argued and further demonstrated the need for the inclusion of three elements in future studies, from a theoretical and a methodological perspective. As seen in our results, the degree of class voting in a country matters. The expression of the class struggle in the democratic arena is more salient in Spain than in Portugal and this is why we find a strong correlation in Spain and not in Portugal.

We have further demonstrated that consideration of the relevant mechanisms at play in the short and in the long-term for the partisan consequences of turnout are necessary for better explaining fluctuations in the effect. The incumbent effect – whether a party is governing or not - is crucial for explaining the reduction of the magnitude of the effect of turnout on the electoral results for the left party, if they are governing, in the short-term.

Additionally, we have demonstrated how better model specifications can adequately test the assumptions of the model and solve issues related to spuriousness and multicollinearity and address some expressed concerns about endogeneity in the theory of the partisan consequences of turnout. Individual data analyses have shown the causal mechanisms behind the aggregated correlations.

In future research it might prove fruitful to include all leftist parties when analyzing turnout and electoral results in the long-term. Aggregated results for leftist parties may better capture the full logic of theory. Accompanying survey research may also be useful to uncovering additional mechanisms at play in the short and long-term.

7. References

Anderson, Christopher J. and Pablo Beramendi (2012). "Left parties, poor voters, and electoral participation in advanced industrial societies". *Comparative Political Studies* (forthcoming).

Anduiza, Eva (1999). *¿Individuos o sistemas? Las razones de la abstención en Europa occidental*, Madrid: Centro de Investigaciones Sociológicas.

Beck, Nathaniel and Jonathan N. Katz (1995). "What to Do (and Not to Do) with Time-Series Cross-Section Data". *American Political Science Review* 89: 634-47.

Beck, Nathaniel and Jonathan N. Katz (2011). "Modeling Dynamics in Time-Series-Cross-Section Political Economy Data". *Annual Review of Political Science* 14: 331-352.

Black Jerome H. (1991). "Reforming the context of the voting process in Canada: lessons from other democracies". In Herman I Bakvis (ed.). *Voter Turnout in Canada.* Toronto: Dundurn.

Blais André (2000). To Vote *Or Not To Vote? The Merits and Limits of Rational Choice*. Pittsburgh: University Pittsburgh Press.

Blais, André (2006). "What affects voter turnout?" *Annual Review of Political Science* 9:111-125.

Blais, André and Kees Aarts (2006). "Electoral systems and turnout". *Acta Politica* 17:180-196.

Blais. André, Eva Anduiza and Aina Gallego (2011) "Decentralization and turnout". *Enrironment and Planning C: Government and Policy* 29: 297-320.

Blais André and R. Kenneth Carty (1990). “Does proportional representation foster voter turnout?” *European Journal of Political Research* 18:167–81.

Bruno, Giovanni S.F. (2005). “Approximating the bias of the LSDV estimator for dynamic unbalanced panel data models”. *Economics Letters* 87: 361-366.

Calvo, Ernesto and Timothy Hellwig (2011). “Centripetal and Centrifugal Incentives under Different Electoral Systems”. *American Journal of Political Science* 55: 27-41.

Chhibber, Pradeep and Mariano Torcal (1997). "Electoral Strategies, Social Cleavages, and Party Systems in a New Democracy: Spain." *Comparative Political Studies* 30: 27-54.

Cox, Gary W. (1997). *Making Votes Count: Strategic Coordination in the World´s Electoral Systems*. New York: Cambridge University Press.

Cox, Gary W. (1999). "Electoral Rules and the Calculus of Mobilization". *Legislative Studies Quarterly* XXIV, 3: 387-419.

DeNardo, James (1980). “Turnout and the Vote: The Joke’s on the Democrats”. *American Political Science Review* 74: 406-420.

DeNardo, James (1986). “Does Heavy Turnout Help Democrats in Presidential Elections? Turnout and the Vote: The Joke’s on the Democrats”. *American Political Science Review* 80: 1298-1304.

Evans, Geoffrey (2000). “The Continued Significance of Class Voting”. *Annual Review of Political Science* 3: 401-417.

Fisher, Stephen (2007). “(Change in) turnout and (change in) the left share of the vote”. *Electoral Studies* 26: 598-611.

Fishman, Robert (2011). “Democratic Practice After the Revolution: The Case of Portugal and Beyond”. *Politics & Society* 39: 233-267.

Freire, André (2006). “Bringing Social Identities Back In: The Social Anchors of Left-Right in Western Europe”. *International Political Science Review* 27: 359-378.

Freire, André and Marina Costa Lobo (2005). “Economics, ideology and vote: Southern Europe, 1985-2000”. *European Journal of Political Research* 44: 493-518.

Gallego, Aina (2010). “Understanding unequal turnout: Education and voting in comparative perspective”. *Electoral Studies* 29: 239-248.

Greene, William H. (1997). *Econometric Analysis.* New Jersey, Prentice Hall.

Grofman, Bernard and Peter Selb (2011). “Turnout and the (effective) number of parties at the national and district levels: A puzzle-solving approach”. *Party Politics* 17: 93-117.

Gunther, Richard and José Ramón Montero (2001). “The Anchors of Partisanship: A Comparative Analysis of Voting Behaviour in Four Southern European Democracies”. In P. Nikiforos Diamandouros and Richard Gunther (eds.). *Parties, Politics, and Democracy in the New Southern Europe*. Baltimore: Johns Hopkins University Press.

Hansford, Thomas G. and Brad T. Gomez (2010). “Estimating the Electoral Effects of Voter Turnout”. *American Political Science Review* 104: 268-288.

Im Kyung So, M. Hashem Pesaran and Yongcheol Shin (2003). “Testing for Unit Roots in Heterogeneous Panels”. *Journal of Econometrics* 115: 53-74.

Kiviet, Jan F. (1995). “On Bias, Inconsistency, and Efficiency of Various Estimators in Dynamic Panel data Models”. *Journal of Econometrics* 68: 53-78.

Knutsen, Oddbjorn and Elinor Scarbrough (1998). “Cleavage Politics”. In Jan W. van Deth and Elinor Scarbrough (eds.). *The impact of values.* Oxford: Oxford University Press.

Levin, Andrew, Chien-Fu Lin and Chia-Shang James Chu (2002). “Unit Root Tests in Panel Data: Asymptotic and Finite Sample Properties”. *Journal of Econometrics* 108: 1-24.

Lewis-Beck, Michael and Marina Costa Lobo (2011). “Anchoring the Portuguese Voter: Panel Dynamics in a Newer Electorate”. *Political Research Quarterly* 64: 293-308.

Lipset, Seymour M. (1981) [1960]. *Political Man: The Social Bases of Politics*. London: Heinemann.

Martinez, Michael D. and Jeff Gill (2005). “The Effects of Turnout on Partisan Outcomes in U.S. Presidential Elections 1960-2000”. *Journal of Politics* 67: 1248-1274.

Nevitte, Neil, André Blais, Elisabeth Gidengil, and Richard Nadeau (2009). “Socio-economic status and non-voting”. In H.D. Klingemann (ed.). *The Comparative Study of Electoral Systems.* Oxford: Oxford University Press.

Nickell, Stephen (1981). “Biases in Dynamic Models with Fixed Effects”, *Econometrica*, 49: 1417-1426.

Pacek, Alexander and Benjamin Radcliff (1995).”Turnout and the Vote for Left-of-Centre Parties: A Cross-National Analysis”. *British Journal of Political Science* 25: 137-143.

Przeworski, Adam and John Sprague (1986). *Paper Stones. A History of Electoral Socialism*. Chicago, University of Chicago Press.

Rubenson, Daniel, André Blais, Patrick Fournier, Elisabeth Gidengil and Neil Nevitte (2007). "Does Low Turnout Matter? Evidence from the Canadian Federal Election". *Electoral Studies* 26: 589-597.

Soto, Marcelo (2010). "System GMM Estimation with a Small Sample", Instituto de Análisis Económico, Barcelona, (http://soto.iae-csic.org/Text%20Monte%20Carlo.pdf)

FOOTNOTES

[1]Class voting refers to the tendency of voters in a particular class to vote for a specific party, rather than an alternative option, compared with voters in another class or classes (Evans, 2000: 402).

[2] Fisher (2007) points out a third question about the left share of the vote if everyone voted. But given that this is a hypothetical question, while the other two are about average behaviour, the former will be not considered here.

[3] However, this correlation is mitigated by the more volatile behaviour of less frequent voters. As DeNardo (1980, 1986) argues, peripheral voters are more likely to defect from whatever partisan leaning they may possess than core voters.

[4]Additionally, as Calvo and Hellwig (2011: 39) show, strategic behaviors influence party platforms, generating simultaneously centripetal and centrifugal tendencies, based on the party's particular characteristics. While non-proportional rules (and districts) *crowd out* smaller parties of more extreme policy positions, as voters defect from parties that expect fewer seats than their vote share, there is a centripetal effect of plurality-like electoral rules *on favourably biased parties*, e.g., parties positively biased in seats by the electoral rules. If the policies to be adopted will vary substantially depending on who is elected, so does an individual's benefit of voting even when controlling for his/her SES. When Hansford and Gomez (2010) argue that previous work on the electoral consequences of variation in voter turnout does not account for endogeneity between turnout and electoral choice, they are mainly referring to (some) of these mechanisms based on strategic behaviours.

[5] The 1980 election has not been included. The coalition between the Socialist Party (PS) with two minor leftist parties, UEDS and ASDI, in this election with the remaining elections of the period makes the comparison not possible.

[6] Electoral results can be found at www.cne.pt (Portugal), and www.elecciones.mir.es (Spain).

[7] See Blais, Anduiza, and Gallego (2001) and Grofman and Selb (2011) for similar research designs.

[8]An electoral reform in 1991 in Portugal reduced the number of MPs from 250 to 230 thus alter marginally altering district magnitudes. This change in district magnitudes does not change our results appreciably.

[9] Contrary to Fisher (2007), for instance, we do not include the results of more leftist parties, such as the Communist, in the dependent variable. Although the correlation between turnout and electoral results should also work for minor leftist parties, when the Socialist Party is the ruling party the incumbent effect would go against the Socialist Party and in favour of minor leftist parties. Aggregating their results would negatively bias the impact of the incumbent effect.

[10] As the variables are differenced, the first election (1975 in Portugal and 1977 in Spain) is not included.

[11] The results do not change appreciably depending on whether the estimates are OLS or robust. Nor do they if the outliers are simply omitted from the analysis.

[12] When regressing the OLS residuals on the lagged endogenous, the exogenous variables and the lagged residuals, a non significant coefficient for the latter was obtained (p-value=0.40). Robust standard errors to both cross-section heteroskedasticity and contemporaneous correlation were also used in this auxiliary regression.

[13] In order to compute the bias corrected Least Squares Dummy Variables (LSDV) estimators for the standard autoregressive panel data model, we rely on the bias approximations in Bruno (2005). Instead of calculating standard or robust errors, this procedure calculates a bootstrap variance-covariance matrix. While, results did not significantly change with other estimate options, we choose a level of accuracy of $O(T^{-1})$ and the Arellano-Bond consistent estimator to initialize the bias correction.

[14] According to Monte Carlo simulations with small samples performed by Soto (2010), the system GMM estimator has a lower bias and higher efficiency than difference GMM and level GMM.

[15] The corresponding F-statistic in column (1) is $F(19, 197)=1.61$ with p-value=.0576. In column (4), $F(19, 198)=0.56$ with p-value=0.9308. When correcting the estimates in column (1) for the potential bias according to the Kiviet's proposal, the coefficient on was $Left_{t-1}$ 0.425 and the coefficient on *Turnout* was -0.343.

[16] This survey was directed by Richard Gunther and J. R. Montero, and conducted in April-May 2004, covering a representative sample of 2.929 adult Spaniards. The survey

was financed by a consortium of researchers from Ohio State University, the Autonomous University of Madrid, the Autonomous University of Barcelona, the Universitat Pompeu Fabra, the University of Santiago de Compostela, and the Institute of Social Studies of Andalusia.