

Strategic Partisanship and Left-Wing Policy Efficiency

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Abstract

It is common for political scientists to investigate the degree to which partisanship affects government policies — especially expenditure. This paper advances the view that such partisanship effects are likely to be conditioned by strategic calculations about the probability of future policy changes. Defining the partisanship effect as the difference between the policy that would be implemented by left-wing and right-wing governments, a formal model is developed that relates the size of this effect to two parameters: the electoral bias faced by left-wing parties and the degree to which today's policies are likely to be rolled back by a future government. In some cases when they face a negative electoral bias, left-wing parties spend 'inefficiently' highly, which itself exacerbates their electoral difficulties. The model yields an unanticipated non-linear hypothesis that finds support when tested with welfare expenditure data across 'developed democracies'.

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It is common for political scientists to investigate the degree to which partisanship affects government policies — especially expenditure. This paper advances the view that such partisanship effects are likely to be conditioned by strategic calculations about the probability of future policy changes. Defining the partisanship effect as the difference — more specifically, the ratio — between the policy that would be implemented by left-wing and right-wing governments, a model is developed that relates the size of this effect to two parameters: the electoral bias faced by left-wing parties and the degree to which today’s policies are likely to be rolled back by a future government. The model yields unanticipated hypotheses that find support when tested with a variety of welfare expenditure data across ‘developed democracies’.

While the partisanship effect is defined as the policy difference between parties from opposing blocs, theoretically, my focus is on ‘left-wing’ parties. This is so because I consider that, as compared to ‘right-wing’ parties, theirs is the more interesting position. For the Right, the task is to shrink government expenditure and redistribution to their preferred relatively low levels. For left-wing parties, however, the task is to install policies that redistribute from the relatively rich (right-wing constituency) to the relatively poor (left-wing constituency) *that last beyond today’s government*. Thus, crudely, right-wing parties will simply tend to cut expenditure to their preferred levels, whilst left-wing parties face the more complicated task of deciding how much to spend so as to both get elected and maximise the probability that redistribution will survive the possible predations of future right-wing governments.

The model set out here shows how, as future right-wing cuts become more likely — say, because right-wing government itself becomes more likely — left-wing parties will tend to move towards more inefficient policies as a counter-measure. However, this leads to a vicious cycle in which they are electorally punished for such a move. In this way, a secondary mechanism of electoral disadvantage for left-wing parties can be layered on top of more direct theoretical mechanisms of the sort set out by Iversen and Soskice (2006) and Ticchi and Vindigni (2010): a propensity to lose induces a policy defensiveness that exacerbates the original propensity to lose.

Building on the seminal work by Hibbs (1977), the political science literature is replete with quantitative empirical studies analysing the impact that partisanship has on policy outputs. Despite a voluminous literature, until relatively recently, there has been a tendency to ignore the context in which parties operate. Earlier researchers have implicitly assumed that left-wing parties seek the same policies in any political environment. This assumption is exhibited in the plethora of time-series cross-section (TSCS) studies, pooling across OECD countries, that model welfare expenditure (e.g. Pampel and Williamson, 1988; Hicks and Swank, 1992; Iversen and Cusack, 2000; Huber and Stephens, 2001; Franzese, 2002; Swank, 2002; Allan and Scruggs, 2004).

Of course, some political scientists have emphasised the importance of context in mediating political decisions. For example, Garrett (1998, Chapter 4) investigates how and why left-wing

and right-wing parties react differently to the onset of globalisation. Rueda (2005, 2007) argues that the division between ‘insider’ and ‘outsider’ labour can have important consequences for the kinds of policies pursued by left-wing parties, which, in turn, can have important consequences for inequality. His argument rests on economic preferences of particular sections of society, not party strategy, though.

Largely separate from the partisanship literature, another branch of research has focused on the influence of institutions of a constitutional nature — most relevantly, work on the implications of credible commitment mechanisms for the nature and structure of public sectors. Moe (1990) proposes that ‘political uncertainty’ should be considered an important factor in policy-making, with Horn and Shepsle (1989) making a similar point. Essentially, they argue that protection of policy gains is an inherent part of the political process. Formalising the theory, de Figueiredo (2002) shows when politicians will choose to insulate their policies. In a similar vein, Acemoglu and Robinson (2001) provide a model showing that an absence of credible commitment mechanisms will lead governments to adopt policies that increase the likelihood of them winning future elections — with inefficiencies in public policy as the by-product. However, those earlier papers were largely concerned with US congressional politics. While Moe (1990, 238-248) provides some theoretical analysis for how the impact of ‘political uncertainty’ will vary from presidential to parliamentary systems, there is generally a dearth of empirically-grounded comparative work that applies the theoretical insight. Furthermore, none of these authors pay attention to how their theories will interact with partisanship, and thus they fail to draw conclusions about why parties of similar ideological origin will pursue different policies across countries.

1 The Model

1.1 A Standard Public Finance Model

Consider an arbitrary number of voters with utility functions given by,

$$u(t, y) = (1 - t)y + H(g) , \tag{1}$$

where t denotes the tax rate, y the (pre-tax) income of a voter, and $H(g)$ the utility derived from government expenditure (g) on a public good/service. The government budget constraint is given by

$$t\mu = g , \tag{2}$$

where μ denotes the mean income of the population. Combining (1) and (2) yields the indirect utility function,

$$v(y, g) = \left(1 - \frac{g}{\mu}\right) y + H(g) . \quad (3)$$

To facilitate simple closed-form solutions, we now specify that:

$$H(g) = h \log(g) . \quad (4)$$

Now, the optimal level g as a function of income is simply found from the relevant first-order condition of $v(\cdot)$:

$$g^*(y) = \frac{h\mu}{y} \quad (5)$$

For completeness, this implies that

$$t^*(y) = \frac{h}{y} . \quad (6)$$

Naturally enough, preferred government spending (and, therefore, preferred tax rates) are decreasing in pre-tax income (y).

1.2 A Two-Period Partisan Model with Electoral Uncertainty

Now consider that there are two political parties, indexed by $x \in \{L, R\}$, in electoral competition. The parties represent individuals with particular income levels, such that

$$y^L < \mu < y^R . \quad (7)$$

Following Alesina (1988), assume that, while there is electoral competition, binding pre-election policy commitments are not possible. The electorate knows what income level each party represents and therefore knows what policy to expect them to implement if they are elected. Thus, if we consider a one period game with an election preceding it to determine which party forms the government, the electorate know that they are choosing between parties that will implement the policy platform given by $g^*(y^x)$. This is simply the standard result of left-wing parties spending more than right-wing parties when in power.

In a two period game, however, we can introduce some dependence between policy choices across the periods that creates more political complexity.

$$g_2(y, g_1) = \text{Max}[d \cdot g_1, g^*(y)] . \quad (8)$$

The idea is that period 2 expenditure (g_2) must be at least as large as $d \cdot g_1$, where d captures

the degree of ‘stickiness’ of government expenditure between periods. A higher d means that it is *possible* to reduce expenditure by less in a subsequent period. Of course, if the optimal expenditure ($g^*(y)$) is greater than $d \cdot g_1$, then that is chosen instead.

Now, it follows that the expected utility for party L in power in the first period is given by

$$\begin{aligned} E(u^L) &= v(y^L, g) \\ &\quad + p^L(e, b) v[y^L, g_2(y^L, g)] \\ &\quad + p^R(e, b) v[y^L, g_2(y^R, g)] , \end{aligned} \tag{9}$$

where $p^L(e, b)$ denotes the probability of party L winning the election between periods 1 and 2.¹ More specifically, we write the probability of left-wing electoral success as

$$p^L(e, b) = 1 - p^R(e, b) = \frac{be}{2} , \tag{10}$$

where e denotes the (to-be-endogenised) relative ‘efficiency’ of the policies chosen by party L and b denotes an exogenous bias in favour of or against party L . Note that the case of no bias is when $b = 1$. In this situation, when party L makes policy that is as efficient as party R , $p^L = 1/2$. As efficiency decreases, party L loses ground (linearly).²

The first-order condition for expenditure is:

$$0 = \frac{2h\mu - 2gy^L + \begin{cases} 2(h\mu - dgy^L) & \text{if } dg^* \geq \frac{h\mu}{y^L} \text{ \& } dg \geq \frac{h\mu}{y^R} \\ (e + 2b - 2)(dgy^L - h\mu) & \text{if } dg^* \geq \frac{h\mu}{y^R} \\ 0 & \text{otherwise} \end{cases}}{g\mu} . \tag{11}$$

Of the three conditions above, the first is substantively irrelevant as party L would never have an incentive to pick a level of g_1 that binds *itself* in period 2. To do so could only be with the motive of forcing party R to spend more than $g^*(y^L)$, but that would not be rational. The third of the conditions corresponds to the apolitical optimum $g^*(y)$ given in (5). Thus, the most interesting is the second condition, when $dg \geq \frac{h\mu}{y^R}$. In this case, party R is constrained by g_1 chosen in the first period, but party L is not. Let us refer to the g that results from this case as the ‘political’ optimum, g^p . Note that the political outcome is more likely as μ/y^R gets smaller — which is to say, the political outcome becomes more likely as the income represented by party R gets larger relative to mean income.

Solving the first-order condition for the case where μ/y^R yields

¹Time discounting has been ignored.

²As p^L is a probability, there is an implied constraint of $0 \leq be \leq 2$.

$$g^p(y^L) = \frac{h\mu}{y^L} \frac{(be - 4)}{(d(be - 2) - 2)}. \quad (12)$$

From this, we can see that the optimal g in this case is the ‘apolitical’ $g^*(y)$, scaled by a function of d , e , and b . Naturally enough, when $d = 1$, this scaling function reduces to 1 as there is no trade-off for party L between the period 1 and period 2 choice: picking g_1 simply imposes that choice on period 2, no matter what.

At this point, e is still an exogenous parameter to the model. To endogenise it, we first define the popular perception of e in the following way:

$$e = \frac{g^*(y^L)}{g} = \frac{h\mu}{gy^L} \quad (13)$$

The idea is simply that the further the chosen g is from the ‘apolitical’ optimum, $g^*(y^L)$, the more inefficient the electorate perceives the policy choice to be.³ Now, in equilibrium, the electorate has rational/accurate expectations about the efficiency with which L will operate policy. Thus, we can write

$$\begin{aligned} e &= \frac{g^*(y^L)}{g^p(y^L)} \\ &= \frac{(d(be - 2) - 2)}{(be - 4)} \end{aligned} \quad (14)$$

Solving for e gives equilibrium efficiency as a function of the exogenous parameters:

$$e^* = \frac{4 + bd - \sqrt{16 + b(bd^2 - 8)}}{2b}. \quad (15)$$

The interesting result is that equilibrium policy efficiency is increasing in the bias (b) in favour of party L :

$$\frac{\partial e^*}{\partial b} = \frac{2(4 - b - \sqrt{16 + b(bd^2 - 8)})}{b^2 \sqrt{16 + b(bd^2 - 8)}} > 0. \quad (16)$$

³N.B. As written, this expression for e implies that choosing levels of g below $g^*(y^L)$ leads to perceptions of greater efficiency. From the setup of the model, this is not a case that is of concern as there are only incentives to over-spend, not under-spend, on g . However, such a formulation is not necessarily unreasonable. One could argue that an electorate seeing party L spending less than it wished in a move towards the median voter would lead to more positive views of the party.

As the bias against party L increases, its policy choice becomes progressively more inefficient. Of course, as $p^L(e, b)$ is a function of e , it is clear that there will be electoral consequences to this inefficiency beyond those stemming directly from a decrease in b . To see this, we substitute (15) into (10), giving:

$$p^L(e^*, b) = \frac{1}{4} \left(4 + bd - \sqrt{16 + b(bd^2 - 8)} \right) . \quad (17)$$

Now we can see directly that

$$\frac{\partial p^L(e^*, b)}{\partial b} = \frac{1}{4} \left(d + \frac{4 - bd^2}{\sqrt{16 + b(bd^2 - 8)}} \right) > 0 . \quad (18)$$

Weak left-wing parties lose both because they are weak and because they can't be trusted when they reach office.

Of course, we can now write the political equilibrium public expenditure $g^p(y^L)$ as a function of only exogenous parameters by substituting (15) into (12):

$$g^p(y^L) = \frac{h\mu \left(4 + bd + \sqrt{16 + b(bd^2 - 8)} \right)}{y^L 4(1 + d)} \quad (19)$$

$$= \frac{h\mu}{y^L} A \quad (20)$$

where A is simply defined as the factor of left-wing expenditure determined by strategic considerations. Given the previous results, it is unsurprising that the expenditure is decreasing in the degree to which party L is electorally favoured, b :

$$\frac{\partial g^p(y^L)}{\partial b} = \frac{h\mu}{y^L} \left(d + \frac{bd^2 - 4}{\sqrt{16 + b(bd^2 - 8)}} \right) < 0 . \quad (21)$$

While the 'political' level of left-wing government expenditure ($g^p(y^L)$) has been characterised, (11) shows that equilibrium policy is not always $g^p(y^L)$, sometimes it is $g^*(y^L)$. Thus, taking the conditions from (11), we can write the expression for equilibrium left-wing government expenditure as

$$g^\lambda(y^L) = \begin{cases} g^p(y^L) & \text{if } dg^p(y^L) \geq \frac{h\mu}{y^R} \\ g^*(y^L) & \text{otherwise} \end{cases} \quad (22)$$

It is clear that there is a discontinuity in $g^\lambda(y^L)$. Intuitively, this stems from the idea that, for low levels of d , over-spending provides no benefit as it does not constrain R in the next period. However, once d reaches a critical point such that a constraint is imposed on R , it is rational to over-spend rather heavily as the level of d still allows R to make *some* reductions in expenditure. From that point on, a rising d reduces the need for some of the over-spending as a higher proportion of current spending will survive into the next period.

To this point, only left-wing policy has been characterised. However, as noted above, the theoretical quantity of interest is the partisanship effect. There are, however, at least two obvious ways to operationalise this concept and they each have different implications for subsequent empirical work. Working with a partisanship effect defined simply as $g^\lambda(y^L) - g^*(y^R)$ suffers from the problem that exogenously higher equilibrium values of g are likely to lead to larger absolute differences, but not necessarily larger relative differences. The result is somewhat empirically intractable as it becomes necessary to control for the hard-to-observe level of h in estimated equations.⁴ The alternative is to define the partisanship effect as a ratio. Specifically, we can write:

$$\Pi \equiv \frac{g^\lambda(y^L)}{g^*(y^R)} \quad (23)$$

The predictions of the model can helpfully be visualised by plotting the predicted partisanship effect (Π) in the parameter space created by the electoral bias parameter (b) and the spending ‘stickiness’ parameter (d). Figure 1 does this and makes clear how the effect of the two parameters differs.

As can be seen, there is a clear non-linear relationship between the size of the partisanship effect and d . On the whole, Π is largest for intermediate values of d . For low values of d , spending must be raised so far into the ‘inefficient’ realm in order for subsequent right-wing governments to be constrained to spend more than they otherwise would that the electoral costs to left-wing governments from this over-spending outweigh the spending gains. Meanwhile, for high values of d , very little over-spending is required to force the hand of subsequent right-wing governments. In the limit, when $d = 1$, no over-spending is required at all as today’s choice must simply be maintained in the next period.

⁴Specifically, $\Pi \equiv g^\lambda(y^L) - g^*(y^R) = \mu \left(\frac{h_{t+1}A}{y^L} - \frac{h_t}{y^R} \right)$, at which point it is difficult to disentangle changes in h from the core partisanship effect, as well as imposing a requirement to measure μ .

For the electoral bias parameter (b), the pattern is different. For low values of d , there is no relationship between Π and b . For high values of d , there is some effect from b , but it is muted by the mechanism described above. Conditional on being in the range where the d parameter renders strategic partisanship potentially desirable for left-wing governments, Π is decreasing in b . Indeed, when the bias is maximally in favour of the left-wing party — $b = 2$ — $g^p(y^L) = g^*(y^L)$ and Π is at its minimum, regardless of d . If party L knows it is going to win the next election, then there is no need to spend to a level that would constrain the choices of party R after that election.

It is these predictions about the relationships between Π , b and, d that I take to data.

2 Empirics

The challenges in empirically testing the model's predictions are far from trivial. First, the non-linearity of the predicted size of Π in b - d parameter space presents difficulties. Second, while the general shape of the partisanship effect is clear, the theory is not precise enough to provide predictions about where exactly we should expect the effect to arise as d increases.⁵ Finally, these problems are compounded by the lack of accurate measures for b and d with which to empirically isolate the relevant political effects. Despite these difficulties, my approach is to estimate a series of models of changes in various types of government expenditure that should plausibly allow the b - d parameter space to be explored.

An important issue that must be resolved before estimating any models is what, exactly, is the unit of analysis. The empirical partisanship literature cited above has tended use country-years — presumably because this fits with the format of much spending data and so maximises the number of observations available to the analyst. However, this poses a problem for a test of the hypotheses here. The theory is built upon time periods for which a change of government is a possibility. While such changes can, in theory, occur in parliamentary democracies at any point, they are most common after elections. Furthermore, the use of country-years as the unit of observation potentially biases against support for the theory. Consider a government in power for four years and assume that the theory is correct. In this case, it seems reasonable to expect that the government will make the theoretically justified spending adjustment in one of the early years. After that, spending is at the politically optimum level so no more changes are required, but with country-years as the unit of observations, those later years will appear to be observations when predictions of spending increases were not followed. For this reason, I take country-governments as the unit of observation. This limits difficulties of this sort, although it does not completely remove them if governments survive across elections.

While the theoretical formulation of the partisanship effect given in (23) is clear, the classic

⁵While figure 1 shows it at around $d = 0.5$, this is entirely contingent on somewhat arbitrarily chosen values of other parameters.

empirical challenge is that we only observe one of $g^*(y_t^R)$ and $g^\lambda(y_t^L)$. Thus, it is necessary to estimate the partisanship effect from observed differences in the government expenditure. To this end, we can write

$$\begin{aligned}\tilde{\Pi} &= \frac{g_t}{g_{t-1}} \\ &= \frac{h_t}{h_{t-1}} \frac{y_{t-1}^G}{y_t^G} I(G_t; A)\end{aligned}\tag{24}$$

where $G \in \{L, R\}$ denotes the party in government, y_t^G denotes the income level represented by the party in government, and $I(G_t; A)$ is a function that returns 1 if $G_t = R$ and A otherwise. The $I(\cdot)$ function captures how the strategic partisanship effect is only relevant when a left-wing party is in power. Note that this formulation assumes that a left-wing government will, in a single period, not be able to raise expenditure all the way to its preferred level. If it could, then we would only expect the first of two successive left-wing governments to exhibit a strategic partisanship effect. However, raising taxes is often politically difficult, and even raising expenditure may require time for policy development and subsequent bureaucratic changes. To give just one example, the Labour government of 1997–2010 adopted large spending increases that were implemented in both its second and third terms. Beyond these policy-related reasons, the data on lengths of government lend further weight to the view that the empirically relevant case is of governments that tend not to have time to achieve their optimal spending levels. Figure 2 presents data on the length of governments in the sample that I use below. It shows that around half of the governments last for three years or less and, indeed, a non-negligible proportion last for two years or less.

Taking logs yields a model that is closer to being empirically viable:

$$\begin{aligned}\log \tilde{\Pi} = \Delta \log g_{i,t} &= \log \left(\frac{h_t}{h_{t-1}} \right) + \log \left(\frac{y_{t-1}^G}{y_t^G} I(G_t; A) \right) \\ &= \Delta \log h_t + \log \left(\frac{y_{t-1}^G}{y_t^G} I(G_t; A) \right)\end{aligned}\tag{25}$$

The remaining issue is how to deal with the final term of (25). Data for y^G — the level of income represented by a government — is not commonly available. Indeed, in a world that has coalition governments, it would be necessary to have measures for the individual parties composing each government, and then to formulate some combination of these to construct y^G . In the absence of such data, I follow a standard approach in the partisanship literature of taking a binary coding of parties into ‘left’ and ‘right’ (or ‘non-left’), and then using a variable that measures the share

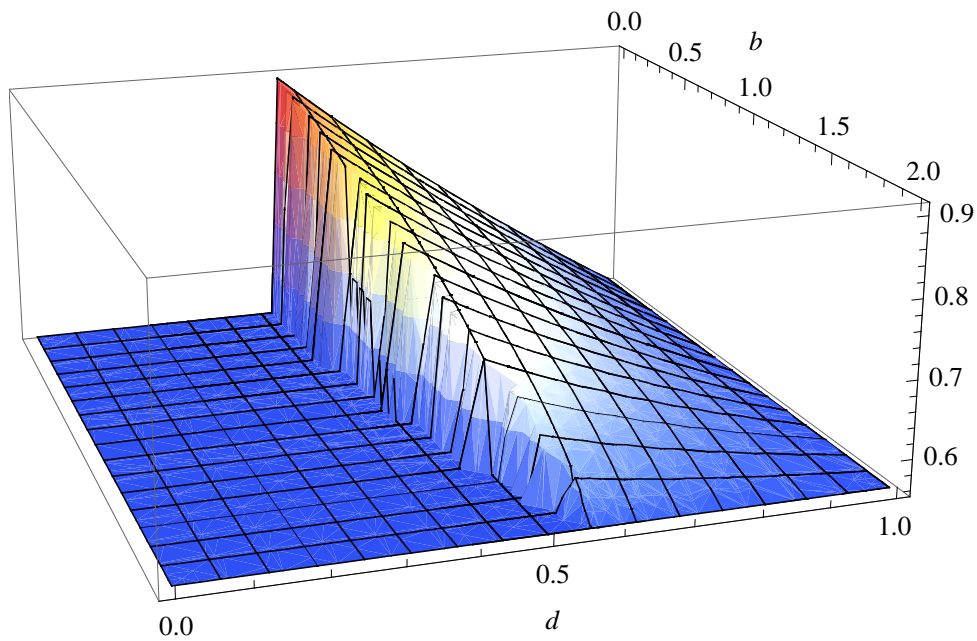


Figure 1: A plot of the predicted partisanship effect (Π) against the degree to which the left-wing party is not electorally disadvantaged (b) and the 'stickiness' of existing spending (d). The plot is drawn with: $h = 1$ $\mu = .5$ $y^L = .4$ $y^R = .7$.

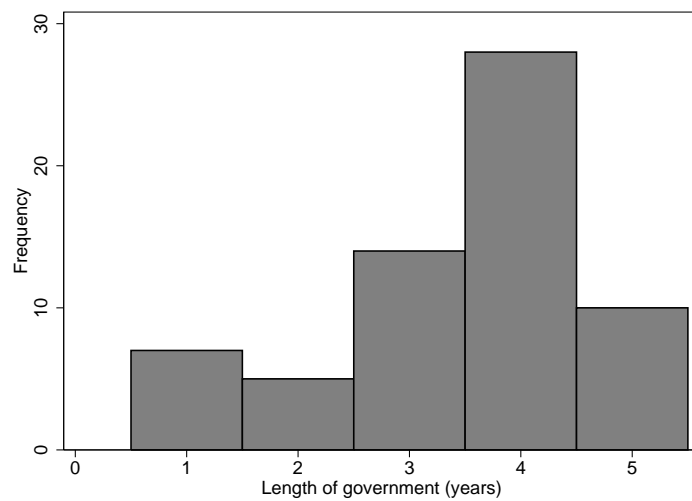


Figure 2: Histogram of the length of governments (in years).

of cabinet seats held by left-wing parties as the measure of their governmental strength. Denote this variable $Left_t$. We can now write

$$\log \left(\frac{y_{t-1}^G}{y_t^G} I(G_t; A) \right) \approx \beta_L Left_t + \beta_{LL} Left_t \cdot \log A \quad (26)$$

where $\beta_L Left_t$ captures the direct effect of changes in the income level represented by successive governments and $\beta_{LL} Left_t \log A$ is used to approximate $I(G_t; A)$ so that partial changes in the complexion of government can be accommodated,⁶ rather than the simple binary shift that was assumed theoretically.⁷ Substituting (26) into (25) implies the following general model to estimate:

$$\Delta \log g_{i,t} = \beta_0 + \Delta \log h_t + \beta_L Left_t + \beta_{LL} Left_t \cdot \log A \quad (27)$$

Note how the partisanship effect is decomposed into a component relating to changes in h and a component that stems directly from the logic of strategic spending embedded in the theory. This is convenient as it allows us to control for those parts of the partisanship effect that result from changes in h , for which there are plausible proxies.

Before proceeding, I shall briefly discuss the dependent variable for the models estimated below. For partisanship to be of relevance, there clearly needs to be disagreement between parties on preferred expenditure. As the theoretical model outlined above has income inequality at the core of its left-right dimension of politics, this leads naturally to a spending measure capturing redistributive effort as the dependent variable. Thus, I employ social expenditure on cash transfers as a percentage of GDP ($SocExp$) as the dependent variable. The data comes from the OECD $SocExp$ database, via Armingeon et al. (2007).

With these preliminary issues dealt with, it is only necessary now to consider how to model the effect of the $\log A$ interaction on the partisanship effect. As this varies depending on whether the focus is on variation in d or in b , I consider the specifics of this modeling choice in the successive sections below.

2.1 Variation in spending ‘stickiness’

In this section, I study how the partisanship effect varies with the d parameter. The aim is to find a measure and a functional form that corresponds to the theoretically predicted interaction effect from $\log A$. From figure 1, it is clear that a primary challenge is how to capture the

⁶E.g. from $Left = 60\%$ to $Left = 80\%$.

⁷See appendix A for more details on the derivation of the approximation.

non-linear relationship between Π and d . I discuss this below, but it is first necessary to find a reasonable proxy measure for d .

As I have already noted, candidates to serve as accurate measures of d are not immediately obvious. What variable can capture the degree to which a particular type of spending can be cut in the future? Given this difficulty, I take a slightly different tack in this section. Rather than attempting to measure d for a particular type of spending, I seek to proxy d with a variable that captures the difficulty that a future government will face in adopting *any* reform. Specifically, following the veto player logic associated with Tsebelis (2002), I suggest that the extent to which opposition parties are divided (or fractionalised) offers a plausible measure of this concept. A more divided opposition bloc will tend to contain more veto players if/when they form a government than an opposition bloc united under a single party leadership. Thus, I employ a variable taken from the World Bank’s *Database of Political Institutions* (Kefer and Stasavage, 2002) that provides a measure of ‘opposition fractionalization’ (*OppFrac*), defined as the probability of two randomly chosen opposition legislators being from different parties.

To see that using *OppFrac* as a proxy for future government fractionalization might be appropriate, consider table 1, which provides the correlations between *OppFrac* and an equivalently defined measure for governing bloc fractionalization, *GovFrac*. Given the common (electoral) institutional underpinnings of party fractionalization within each country, it is unsurprising that *OppFrac_{i,t}* is positively correlated with *GovFrac_{i,t}*. More relevantly, this correlation rises from 0.37 to 0.53 when considering government fractionalization in the *next* period. Opposition fractionalization today is correlated with government fractionalization tomorrow. The correlation is not perfect both because elections do not always result in opposition parties forming a new government and because, even when they do, not all such parties enter government.

| | <i>GovFrac_{i,t+1}</i> | <i>GovFrac_{i,t}</i> | <i>OppFrac_{i,t}</i> |
|--------------------------------|--------------------------------|------------------------------|------------------------------|
| <i>GovFrac_{i,t+1}</i> | 1 | 0.64 | 0.53 |
| <i>GovFrac_{i,t}</i> | | 1 | 0.37 |
| <i>OppFrac_{i,t}</i> | | | 1 |

Table 1: Correlation matrix of government and opposition party fractionalization.

With this proxy measure in place, the question arises of how to model the non-linear relationship exhibited in figure 1 between d and Π . As d rises, Π initially remains unchanged, then takes a discontinuous/step rise to its maximum value, and finally reduces continuously to a minimum value again. The step change is difficult to model as the theory lacks precision on what value of d this should occur at — not to mention the uncertain mapping between that critical value of d and the equivalent for its proxy measure, *OppFrac*. For this reason, I model the relationship with a quadratic functional form. This will tend to smooth over the step change, but should fit the latter part of d variation reasonably well. Obviously, this quadratic relationship

is conditional on government partisanship, so it is interacted with a partisanship variable.

As a final note on the use of *OppFrac* as a proxy for d , it may be thought that *OppFrac* will also partially proxy for electoral bias, b . A more divided opposition may imply that it will be less effective at winning elections. While this may be true, I suggest that it does not pose a problem for my empirical strategy. Correlation with b as well as d simply leads the use of *OppFrac* to imply a tracing out of the diagonal in the b - d space depicted in figure 1,⁸ but this yields essentially the same functional form to approximate for Π . Furthermore, the functional form is theoretically driven by variation in d , so there is no reason to think that any empirical results are driven by a different mechanism from the model.

In sum, I estimate an empirical model of the following form:

$$\begin{aligned} \delta SocExp_{i,t} = & \beta_i + \beta_D \log PublicDebt_{i,t-1} + \beta_U \delta Unempl_{i,t} + \beta_D \delta Deind_{i,t-1} \\ & + \beta_L Left_{i,t} + \beta_O OppFrac_{i,t} + \beta_{O2} OppFrac_{i,t}^2 \\ & + \beta_{LO} (Left_{i,t} \cdot OppFrac_{i,t}) + \beta_{LO2} (Left_{i,t} \cdot OppFrac_{i,t}^2) \\ & + \epsilon_{i,t} \end{aligned} \tag{28}$$

where δ denotes the first difference of the log of a variable. In this specification, h is proxied with public debt (*PublicDebt*), unemployment (*Unempl*), and deindustrialization (*Deind*) (Iversen and Cusack, 2000). To resolve estimation problems stemming from any remaining serial correlation, the models are estimated using the Prais-Winsten procedure. As my focus is on the effects of political agency, I remove federal countries from the sample, leaving only unitary states where there is a clear line of responsibility from national-level partisanship to spending decisions. This gives a sample of 11 countries⁹ for the period 1981–2005.

Table 2 presents the results from estimating this model using *Left_t* as the partisanship variable. The first, *SX1*, presents a base model with the relevant interaction effects. The second, *SX2*, additionally includes a direct effect of government fractionalization (*GovFrac_t*) and its interaction with *Left_t*.¹⁰

The raw results are very encouraging as the models appear well specified. As expected, public debt is a drag on social expenditure, while unemployment and deindustrialisation lead to higher spending in this area. For the results relating directly to the claims in this paper, the evidence is, again, supportive. The slightly complex interaction effects are difficult to interpret directly, so figures 3 and 4 depict how the estimated partisanship effect varies with opposition

⁸In the extreme, from $b = 0, d = 0$ to $b = 2, d = 1$.

⁹Austria, Denmark, Finland, France, Greece, Ireland, Italy, Netherlands, Norway, Portugal, Sweden, and the UK. XXX[Norway is excluded as it appears to be a large outlier on the dependent variable.]

¹⁰This is implied by the fact that the justification for using *OppFrac* is that it correlates with future fractionalized governments that will find it more difficult to adopt reforms. Logically, then *current* government fractionalization is important.

| | (SX1) | | | (SX2) | | |
|------------------------------------|--------|------|------|---------|------|------|
| | b | se | p | b | se | p |
| $\delta PublicDebt_{i,t-1}$ | -0.171 | 0.05 | 0.00 | -0.172 | 0.05 | 0.00 |
| $\delta Unemp_{i,t}$ | 0.163 | 0.03 | 0.00 | 0.161 | 0.03 | 0.00 |
| $\delta Deind_{i,t-1}$ | 0.465 | 0.34 | 0.17 | 0.481 | 0.33 | 0.14 |
| $Left_{i,t}$ | -0.343 | 0.17 | 0.05 | -0.344 | 0.17 | 0.04 |
| $OppFrac_{i,t}$ | -1.958 | 0.61 | 0.00 | -1.988 | 0.62 | 0.00 |
| $OppFrac_{i,t}^2$ | 2.114 | 0.65 | 0.00 | 2.109 | 0.65 | 0.00 |
| $Left_{i,t} \cdot OppFrac_{i,t}$ | 2.488 | 0.83 | 0.00 | 2.509 | 0.80 | 0.00 |
| $Left_{i,t} \cdot OppFrac_{i,t}^2$ | -2.941 | 0.85 | 0.00 | -2.914 | 0.79 | 0.00 |
| $GovFrac_{i,t}$ | | | | 0.0465 | 0.10 | 0.65 |
| $Left_{i,t} \cdot GovFrac_{i,t}$ | | | | -0.0551 | 0.15 | 0.71 |
| N | 64 | | | 64 | | |
| Countries | 12 | | | 12 | | |
| R^2 | 0.663 | | | 0.661 | | |

Notes: Prais-Winsten estimation with panel-corrected standard errors.
Country fixed effects not reported.

Table 2: Models estimating the partisanship effect on total social expenditure (*SocExp*) across a range of ‘developed democracies’.

fractionalization for the two models. As can be seen, both specifications yield the predicted non-linear relationship between Π and d . In general, then, the theory finds support in the data on the d dimension of variation

2.2 Variation in electoral bias

In this section, I study how the partisanship effect varies with the b parameter. As figure 1 shows, the predicted effect of b on Π varies with d . For some values of d , b is predicted to have no effect. At other values its effect is predicted to be negative and large. Finally, at other values, its effect is predicted to be negative and small. Mathematically, the appropriate test of the theory including the predictions relating to b would be to estimate a model that had all of the *Left–OppFrac* interactions, the *Left–b* interaction, and the cross interactions between *Left*, *OppFrac*, and b . This yields 11 parameters to estimate from a sample of only around 65 observations, with several of the interaction effects being three- or four-way multiplications. As I do not believe the data can bear this weight, I concentrate instead on estimating the effect of b averaged over all values of d . As a result, the expectation is for the negative interaction effect between *Left* and proxies for b to be somewhat weaker than would otherwise be the case.

Similar to the previous section, a primary aim is to find a measure that corresponds with the theoretical parameter, b . Given the difficulty of this, the approach I take is to estimate several models with different plausible proxies for b . Perhaps the most obvious proxy is a measure

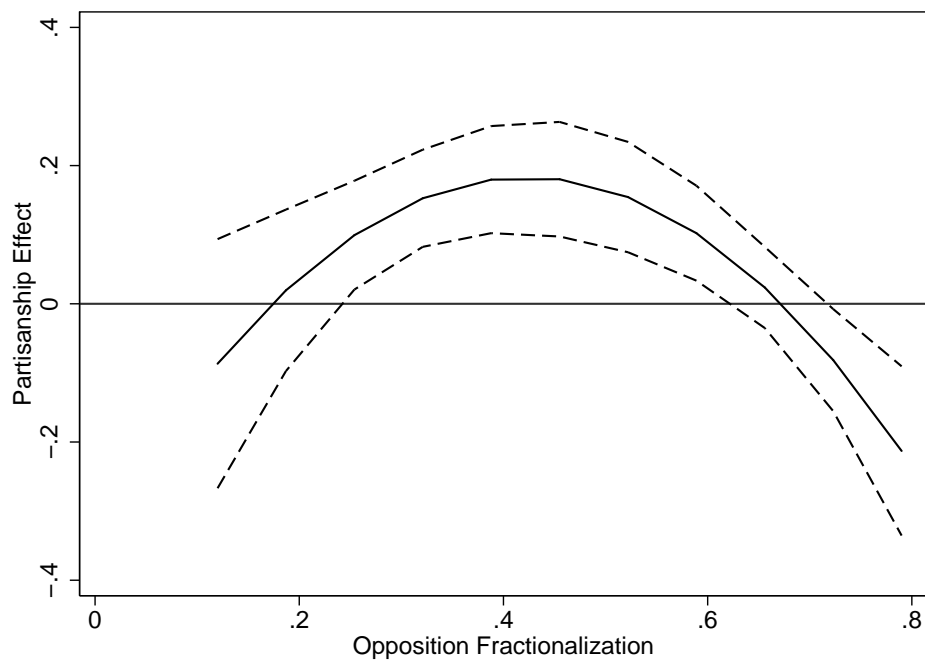


Figure 3: The partisanship effect estimated by model *SX1*, for varying levels of opposition fractionalization (*OppFrac*).

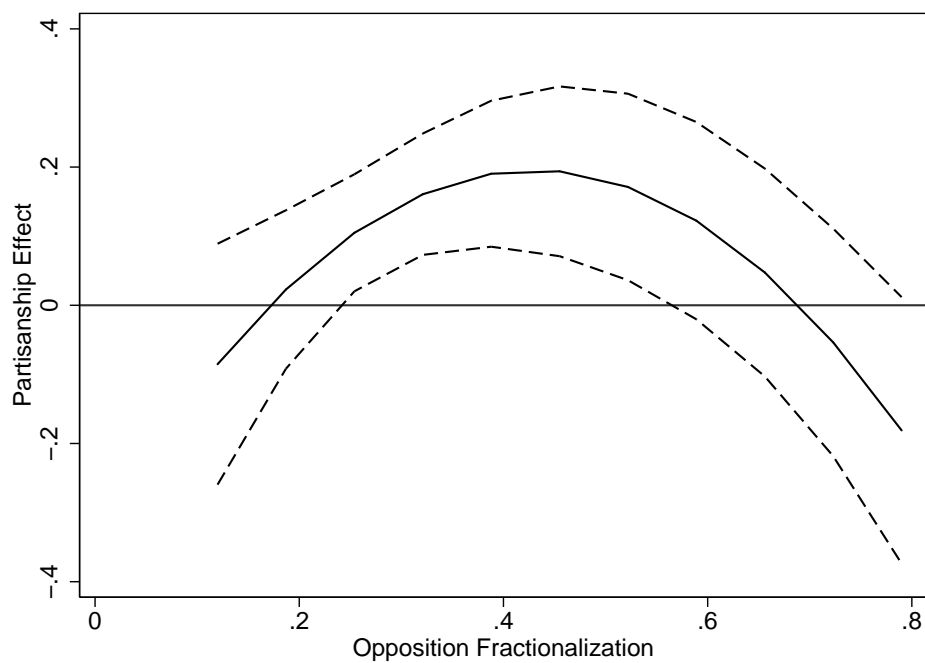


Figure 4: The partisanship effect estimated by model *SX2*, for varying levels of opposition fractionalization (*OppFrac*).

derived from the electoral strength of right-wing parties. When left-wing parties are in power, they may reasonably expect a more right-wing government in the next period when more votes were cast for right-wing parties in the current period. To capture this, I calculate the share of votes won by right-wing parties in the election at the start of each governing period, and use this as a proxy for b . In order to match the scaling of b from low levels being bias against left-wing parties and high values being bias in favour, I actually use $VoteShare^{-R} = 100 - VoteShare^R$.

While using $VoteShare^{-R}$ is intuitive, it is potentially confounded by a number of factors — such as whether high right-wing vote share today really implies *higher* vote share tomorrow. A high share may simply correspond to a high-water mark from which the Right can be expected to drop next time. My second proxy takes a more systemic view of relative left- and right-wing electoral bias. Recent theory has suggested that greater electoral proportionality is advantageous to left-wing parties (Iversen and Soskice, 2006; Ticchi and Vindigni, 2010). Specifically, Iversen and Soskice (2006) argue that majoritarian electoral systems have an inherent bias against left-wing parties as the middle class fear giving unconstrained power to a left-wing party that may turn out to be dominated by those sections of the Left that wish to ‘soak’ both the rich and the middle classes. By contrast, under proportional electoral systems, the middle class will tend to have their own party representation and thus not face such a risk as their party can always withdraw support from a centre-left coalition government if necessary. From this theory, we should expect that as electoral proportionality (PR) rises, then so does left-wing electoral strength.¹¹ Even if the theoretical link between proportionality and left-wing strength is questioned, Iversen and Soskice (2006) show that there is an empirical association that should be sufficient for my purposes.

In summary, denoting each of the proxies for b as B , I estimate models of the following form, where the prediction is for negative values of β_{LB} :

$$\begin{aligned} \delta SocExp_{i,t} = & \beta_i + \beta_D \log PublicDebt_{i,t-1} + \beta_U \delta Unempl_{i,t} + \beta_D \delta Deind_{i,t-1} \\ & + \beta_L Left_{i,t} + \beta_B B_{i,t} + \beta_{LB} (Left_{i,t} \cdot B_{i,t}) \\ & + \epsilon_{i,t} \end{aligned} \tag{29}$$

Table 3 presents the results from estimating the model with each proxy, in turn. Once again, the model performs robustly as each of the effects from the control variables accords with expectations. Interpretation of the partisanship interaction effects is more nuanced than normal given that the effect is expected to be rather weak. Nonetheless, as predicted, each of the interaction effects is correctly (negatively) signed, indicating that the partisanship effect declines as the proxies left-wing electoral strength get larger. Figures 5(a) and 5(b) depict these estimated partisanship effects and show that, the evidence is indeed supportive of a weak

¹¹Electoral proportionality is measured as 100 minus the Gallagher (1991) disproportionality index.

negative interaction.

| | (SX4) | | | (SX5) | | |
|---|----------|------|------|-----------|------|------|
| | b | se | p | b | se | p |
| $\delta PublicDebt_{i,t-1}$ | -0.140 | 0.06 | 0.03 | -0.146 | 0.07 | 0.02 |
| $\delta Unemp_{i,t}$ | 0.170 | 0.04 | 0.00 | 0.178 | 0.04 | 0.00 |
| $\delta Deind_{i,t-1}$ | 0.398 | 0.42 | 0.34 | 0.572 | 0.41 | 0.16 |
| $Left_{i,t}$ | 0.435 | 0.17 | 0.01 | 0.408 | 0.28 | 0.14 |
| $VoteShare^{-R}$ | 0.000706 | 0.00 | 0.56 | | | |
| $Left_{i,t} \cdot VoteShare_{i,t}^{-R}$ | -0.00555 | 0.00 | 0.03 | | | |
| $PR_{i,t}$ | | | | -0.000334 | 0.00 | 0.78 |
| $Left_{i,t} \cdot PR_{i,t}$ | | | | -0.00379 | 0.00 | 0.22 |
| N | 66 | | | 66 | | |
| Countries | 12 | | | 12 | | |
| R^2 | 0.555 | | | 0.535 | | |

Notes: Prais-Winsten estimation with panel-corrected standard errors. Country fixed effects not reported.

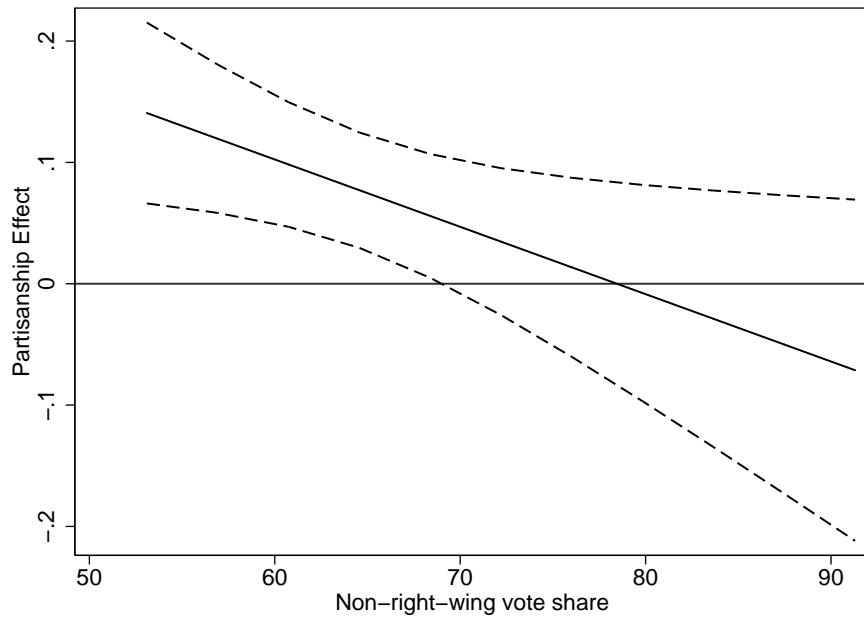
Table 3: Models estimating the partisanship effect on total social expenditure (*SocExp*) across a range of ‘developed democracies’.

3 Conclusions

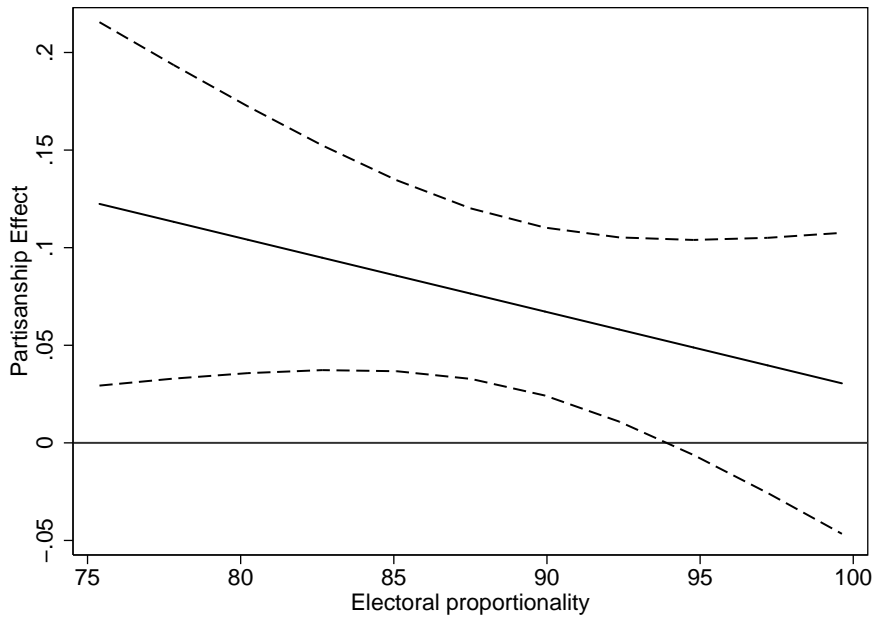
This paper has developed the standard model of partisanship effects on public policy by outlining an implication of the view that parties may make policy choices based on their expectations about future streams of policy pay-offs. With this extension to the standard model, a new set of considerations arise surrounding the likelihood of a future government having different preferences and the likelihood that a such a government would be able to make reforms to reflect those preferences. From these intuitions, a simple theoretical model is built which shows how an electoral bias against (left-wing) parties may be doubly harmful in that it induces them to adopt policies that exacerbate their electoral difficulties. Such a course of action is entirely rational for a policy-seeking party.

The theoretical is developed in such a way that it leads to a set of hypotheses and rather a natural path to empirical models with which to test them. The empirical results are rather supportive of the theoretical model. Interestingly, taken at face value, the results regarding variation in the d parameter from section 2.1 actually suggest that *all* of the partisanship effect can actually be ascribed to the strategic motivations derived from the model developed above. That is, for low and for high levels of the d proxy (*OppFrac*), the estimated partisanship effect is essentially zero. It is only for intermediate levels of d that a partisanship effect emerges. Were there to be a non-strategic component to the partisanship effect, we would expect the estimated curves in figures 3 and 4 to be shifted upwards.

While caution should certainly be exercised when interpreting these findings — not least



(a) The partisanship effect estimated by model *SX4*, for varying levels of non-right-wing vote share ($VoteShare^{-R}$).



(b) The partisanship effect estimated by model *SX5*, for varying levels of electoral proportionality (PR).

because of the measurement challenges for some of the explanatory variables — the apparent importance of the strategic partisanship component is intriguing. It suggests an explanation for previous studies that have found no or muted partisanship effects may be explained as the result of averaging over situations in which left-wing parties have very different levels of incentive to spend differentially from their right-wing opponents.

A Deriving the Partisanship Specification

The question is what to do the final term in (25). We can separate the problem into two parts by writing

$$\log\left(\frac{y_{t-1}^G}{y_t^G}I(G_t; A)\right) = \log\left(\frac{y_{t-1}^G}{y_t^G}\right) + \log(I(G_t; A)) \quad (30)$$

Taking the first of the terms and using the assumption that governing periods are short-lived enough that the partisan optimal expenditure will not be reached, we can write

$$\log\left(\frac{y_{t-1}^G}{y_t^G}\right) \approx \beta_L Left_t \quad (31)$$

The idea here is that, as optimal expenditure is not reached, it is not the difference in the partisanship variable (*Left*) that matters, but level. If the assumption were wrong and optimal expenditure did tend to be reached, we could write

$$\begin{aligned} \log\left(\frac{y_{t-1}^G}{y_t^G}\right) &\approx \log\left(\frac{y^R - \alpha_L Left_{t-1}}{y^R - \alpha_L Left_t}\right) \\ &= \log(y^R - \alpha_L Left_{t-1}) - \log(y^R - \alpha_L Left_t) \end{aligned} \quad (32)$$

where $\alpha_L Left$ captures the (assumed constant) difference between y^R and y^L as the left share of cabinet seats goes from 0% to 100%. This still could not be estimated as we lack data for y^R and the log form is awkward. To avoid these difficulties, we could write

$$\begin{aligned} \log(y^R - \alpha_L Left_{t-1}) - \log(y^R - \alpha_L Left_t) &\approx \beta_0 + \phi(y^R - \alpha_L Left_{t-1}) - \phi(y^R - \alpha_L Left_t) \\ &= \beta_0 + \beta_L \Delta Left_t \text{ where } \beta_L \equiv \phi\alpha_L \end{aligned} \quad (33)$$

This would be a reasonable approximation as long as the difference between y^R and y^L is not *too* large, and yields a far more manageable empirical specification.

Beyond the reasoning offered in the main text, we can assess which of the specifications appears to be more empirically valid by estimating simple baseline partisanship models without interaction effects for each of the two candidate partisanship variables. This should uncover the average partisanship effect across all cases. If one variable performs better than the other, then it can be taken as evidence in its favour for the rest of the models, as well. Table 4 shows the results from this exercise, and reveals that the $Left_t$ model performs better than the $\Delta Left_t$ model.

| | (SXA) | | | (SXB) | | |
|-----------------------------|--------|------|------|--------|------|------|
| | b | se | p | b | se | p |
| $\delta PublicDebt_{i,t-1}$ | -0.154 | 0.06 | 0.02 | -0.151 | 0.07 | 0.03 |
| $\delta Unemp_{i,t}$ | 0.179 | 0.04 | 0.00 | 0.175 | 0.04 | 0.00 |
| $\delta Deind_{i,t-1}$ | 0.632 | 0.41 | 0.12 | 0.720 | 0.42 | 0.08 |
| $Left_{i,t}$ | 0.0562 | 0.03 | 0.03 | | | |
| $\Delta Left_{i,t}$ | | | | 0.0247 | 0.02 | 0.32 |
| N | 66 | | | 66 | | |
| Countries | 12 | | | 12 | | |
| R^2 | 0.529 | | | 0.505 | | |

Notes: Prais-Winsten estimation with panel-corrected standard errors. Country fixed effects not reported.

Table 4: Models estimating the partisanship effect on total social expenditure ($SocExp$) across a range of ‘developed democracies’.

The final issue is how to operationalise $\log I(G_t; A)$. Given that we are assuming (with empirical reason) that governments do not achieve their optimal expenditure in a term, we need a form that equates to 0 when $G_t = R$ and 1 when $G_t = L$. Scaling the intermediate cases of coalition with $Left_t$, we write

$$\log I(G_t; A) \approx \beta_{LL} Left_t \cdot \log A \quad (34)$$

For completeness, if we were assuming that governments do achieve their optimal expenditure in a given term, we would need a form that equates to 0 when $G_t = G_{t-1}$ and $\log A$, otherwise. $\Delta Left_t$ proxies nicely for $G_t = G_{t-1}$, so we could write

$$\log I(G_t; A) \approx \beta_{LL} \Delta Left_t \cdot \log A \quad (35)$$

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