The Interest Premium for Left Government: Regression Discontinuity Estimates

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Abstract

This paper employs a regression-discontinuity design (RDD) to ascertain the effects of left government on the interest-rate premium that markets build into government-bond prices. One advantage of this approach is that RDD does not require, as have some previously employed strategies, strong assumptions about how market actors form political expectations, about the quality and dissemination of political information, or about functional forms or explanatory-variable selection. We expand from previous RDD studies in exploring effect heterogeneity, namely whether particular political-economic conditions produce larger or smaller interest costs of left government. Our findings suggest no or very small and insignificant partisan-government effects except under specific circumstances: sharp governing alternatives (low fragmentation, high polarization), in certain eras (around the 1950s-1970s), and for a short-term (about one year). Under these conditions of stark differences between alternative left/right governments and relatively great domestic policy autonomy, however, there is a statistically discernible and substantively notable government-bond yield increase after left parties enter government following close elections.

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1 Introduction

How do financial markets react to the partisan composition of government? What cost, if any, do citizens pay in higher government bond interest rates for electing left-leaning and social democratic parties?

Because social-democratic (left) parties tend to support policies that involve redistribution of wealth, many scholars expect financial markets will react strongly when such parties are elected to government. The \textit{locus classicus} on the topic is Lindblom (1977) \textit{Politics and Markets}, which emphasizes the “privileged position of capital”, i.e., capital’s ability to withhold investment as a credible threat against governments thinking to act against investors’ interests. Proponents of rational partisan theory also argue that financial markets impose a high price on citizens who choose center-left representatives (Alesina, Roubini and Cohen, 1997; Herron, 2000; Fowler, 2006; Bechtel, 2009; Sattler, 2013; Barta and Johnston, 2018).

Others theories, however, predict little or no financial market response to government partisanship. Some scholars contend that traders only worry about macroeconomic performance and not policy or partisanship \textit{per se} (Garrett, 1998; Mosley, 2000). Since—according to a ‘Varieties of Capitalism’ version of this view for example—social-democratic government is at worst not systematically related to macroeconomic performance, and in particular to inflation, the interest-rate cost of social democracy may be very low, zero, or even negative. Still others maintain that the globalization of financial markets places a ‘Golden Straight-jacket’ on governments, shrinking the policy space between the right and left (Rodrik, 2000), or that democratic competition induces Hotelling-Downsian convergence in the macroeconomic policies of left and right governments (Clark, 2003). In either case, financial markets will be indifferent to the partisan composition of government.\footnote{More recently, Hübscher (2016) finds no partisan differences with respect to fiscal policies. According to her empirical analysis, left and right governments are equally likely to adopt fiscal consolidation reforms.}

Previous attempts to resolve this question empirically have faced three challenges. First,
omitted variable bias: many potentially observable variables can influence both bond-market performance and the electoral success of parties. For example, oil-price shocks might affect both interest rates and the effectiveness of economic policies in a party’s policy toolkit, and thereby its probability of winning elections. Second, reverse causality: bond prices affect right- and left-party core constituencies differently and therefore affect the probability of left or right government election wins. For example, a hypothetical exogenous bond-price appreciation (interest-rate decrease) is macroeconomically stimulatory, and such a stimulus may alter the relative appeal of leftist or rightist platforms and so the probability of a left party election victory.

A third challenge arises because financial markets are composed of forward-looking political economic actors. Even in the presence of large treatment effects, naïve pre-post comparisons might fail to find partisan-government effects if shifts in government partisanship are anticipated by traders. To address this, structural-estimation approaches must rest on specific, restrictive assumptions about how traders form political expectations (Alesina, Roubini and Cohen, 1997; Pástor and Veronesi, 2013), about the quality and dissemination of political information (Herron, 2000), or various specification decisions related to functional forms and variable selection (Garrett, 1998; Franzese, 2002; Clark, 2003; Mosley, 2003; Bernhard and Leblang, 2006). As a result, these approaches leave conclusions about ‘the price of social democracy’ more contestable. Studies of single elections can forego these restrictive assumptions if there are idiosyncratic fluctuations in market expectations as measured by political prediction markets (Snowberg, Wolfers and Zitzewitz, 2007), but data availability limits this approach mostly to recent US presidential elections.

The regression discontinuity design (RDD) offers a method to redress these challenges. RDD identifies the effect of treatment at a threshold point by exploiting a discontinuous break in the probability of treatment at that threshold (Hahn, Todd and Van Der Klaauw, 2001; Imbens and Lemieux, 2008; Calonico, Cattaneo and Titiunik, 2014). In our context, parliamentary elections provide such a discontinuity: a party becomes sharply more likely
to enter government when it crosses the plurality threshold of parliamentary seats, i.e. the party crosses from holding second-most to most seats in parliament. The discontinuity occurs because being the largest party confers several distinct advantages. For one, the largest party may hold an absolute majority of seats, in which case it can, and virtually always does, form a single-party government. Even short of majority, the largest party is typically nominated first as the *formateur*, granting it great first-proposal power in forming government (Baron and Ferejohn, 1989). The largest party is also most likely to be at dimensional medians and necessary to coalitional majorities, which contribute further great bargaining power to enter any government that forms (Laver and Shepsle, 1996). In any case, the validity of this discontinuity can be tested directly, and in Section 4 we demonstrate that plurality status does indeed yield a discontinuous break in the probability of left parties entering government, but clearly so only in countries with low party-system fragmentation (both as noticed also in Powell Jr (2000), see esp. Figures 6.3 & 6.4).

By studying the reactions of financial markets following closely contested (in terms of seats between the largest two parties) parliamentary elections, we avoid the need to assume or specify and estimate how traders process information and form expectations. As the formal model in Section 2 demonstrates, it suffices for identification to assume merely that close elections imply greater *ex ante* uncertainty than elections where one party wins a plurality of seats by a large margin. In addition, RDD does not require restrictive pre-specification of functional forms and controls to eliminate omitted-variable-bias and reverse-causality concerns. Instead, the central identifying assumption (for the treatment effect at the threshold) is that the probability of treatment is the only variable that changes discontinuously at the threshold. If all other covariates change smoothly at the threshold, then a discontinuous change in the outcome cannot be attributable to those factors. This too can be evaluated empirically for observed factors, though it must be assumed for unobserved.

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2Neither do we need to assume away other sources of uncertainty about who will govern or what their policies will be, rather only that close elections entail greater uncertainty than lopsided ones. After all, these other sources of uncertainty are in general *additional* to the uncertainty related to the electoral margin.
The remainder of the paper proceeds as follows. Section 2 presents a formal-theoretical model which demonstrates how RDD recovers an estimate of the bond-price premium of left government (at the plurality threshold) without strong assumptions about how traders form expectations. The conditional-expectation function implied by this model differs from that typically seen in RDD studies. Section 3 describes our dataset of close parliamentary elections and government-bond yields. Section 4 explores the testable RDD assumptions and estimates the local average treatment effect (LATE) of left-party entry to government on bond yields in high- and low-fragmentation party systems. In further extensions, we also explore potentially heterogeneous LATEs by historical era and by party ideology. The estimated LATE is larger in the 1950s-1970s, before highly liberalized capital markets, and where differences between the main left and right parties’ platforms is greater.\(^3\)

\section{Theoretical Model}

A strength of RDD for estimating the effect of left government on interest premia is that one needn’t specify an empirical model of how traders form expectations. Instead, RDD relies only on a simpler, highly plausible assumption: that close elections imply greater \textit{ex ante} uncertainty, on average, than elections decided by large margins,\(^4\) as the following formal model demonstrates.

In our model, the market values a government’s bonds at price \(P_L\) when the left party, and \(P_R\) when the right party, controls parliament. The quantity that we would like to estimate empirically is \(P_L - P_R\), the bond market response to left-party control, but we cannot observe

\(^3\)Other bases for heterogeneous treatments and effects across our sample countries and time periods are surely also present: e.g., variations in central-executive control of government, in government control of bond-price relevant policy, \&/or in policy maneuverability or efficacy (Franzese Jr, 2002). Practical limitations set by the number of close elections that occur limit how many such variations one can effectively consider. We think partisan polarization and exchange-rate regimes and capital mobility variations, which latter we proxy by time periods, are among the more important ones. In any case, our estimated Local \textit{Average} Treatment Effects average appropriately across any such unconsidered variations.

\(^4\)See note 2 for elaboration. Also see the Supplementary Materials for a test of this assumption, building on a measure of electoral risk from (Kayser and Lindstädt, 2015).
this quantity directly, as we cannot simultaneously observe $P_L$ and $P_R$. So we must estimate an average difference across multiple elections instead. Let us define $V$ as the size of the left-party seat-plurality, measured as a percentage of the top-two parties’ total seats. When $V > 0$ the left party is the largest in parliament, yielding left-party government. When $V < 0$ the right party is largest, yielding right-party government.\footnote{The existence of a sharp discontinuity at the plurality threshold can be evaluated empirically; in Section 4, we show that such sharp discontinuity clearly manifests in some country-years but not in others. Specifically: in two-party or low-fragmentation systems, the plurality party always or almost always forms government, but in more-fragmented multiparty systems the share of largest parties entering government increases much less sharply, i.e. not clearly discontinuously, at that plurality threshold.}

Prior to the election, markets are uncertain about the future value of $V$ but receive information from numerous polls. The Central Limit Theorem suggests the expected seat shares across these polls will be distributed (approximately) normally around the true value $V$. After observing the polls, this distribution becomes the market’s prior belief, which we denote $f_V = N(V, \sigma^2)$. Therefore, the market’s prior expected probability of right government is equal to the integral (cumulative distribution) of $f_V$ evaluated at 0, denoted $F_V$.

Given these beliefs, markets will price government bonds at their expected value, an $F_V$-weighted average of $P_L$ and $P_R$:

$$P_{\text{before}} = (1 - F_V)P_L + F_V P_R$$

The election reveals the true value of $V$, and markets price bonds to $P_L$ or $P_R$ accordingly:

$$P_{\text{after}} = \begin{cases} P_L & \text{if } V > 0 \\ P_R & \text{if } V \leq 0 \end{cases}$$

The expected difference between ex ante bond prices before the election and ex post bond prices after the election will be given by the following function:\footnote{For expositional clarity, the model isolates electoral uncertainty exclusively; as previously noted, the qualitative results depend only upon closer elections entailing greater uncertainty than lopsided ones.}
\[ \Delta P = \begin{cases} 
F_V(P_L - P_R) & \text{if } V > 0 \\
(1 - F_V)(P_R - P_L) & \text{if } V \leq 0 
\end{cases} \] (1)

Figure 1 illustrates this conditional expectation function. Note that, due to the presence of forward-looking traders, the shape of this function differs from what is typically seen in regression discontinuity studies. When \( V \) is far from zero, the expected post-election bond price movement is zero, because traders are more certain of the election outcome, and price bonds accordingly. As \( V \) approaches zero, the function diverges sharply. Because \( F_0 = \frac{1}{2} \), the right side approaches \( \frac{P_L - P_R}{2} \) in the limit as \( V \) approaches 0, and the left side approaches \( \frac{P_R - P_L}{2} \). Subtracting these two limits yields \( P_L - P_R \), our quantity of interest and the focal estimand of a regression discontinuity design.

Figure 1: The expected shape of the conditional expectation function. Taking the difference between the two limits as they approach \( V = 0 \) recovers the quantity of interest, \( P_L - P_R \).
3 Data

Our data on parliamentary election results are from the ParlGov database compiled by Döring and Manow (2018). ParlGov contains data on parties, elections, and cabinets from all EU and OECD parliamentary democracies from 1948 to 2015. We gather from ParlGov data on each party’s number of seats, the party composition of each post-election cabinet, the family classification of each party, and a 0 to 10 measure of each party’s left-right ideology.\(^7\)

We adjust these ParlGov data to incorporate information on pre-electoral coalitions (PECs), as opposed to individual parties. PECs are sets of parties that pledge—critically: publicly, credibly, and exceedingly rarely broken—to form one government together if elected (Golder, 2006). Absent this information, some elections might appear close that were not (e.g., Germany’s 1961 election, where SPD had a plurality, but the CDU/CSU coalition easily out-sized them), and some elections might appear lopsided that were quite close (e.g., Germany’s 1976 election, where a pre-electoral coalition of SPD and FDP narrowly overtook the plurality Christian Democrats). For data on PECs, we rely on the dataset of Golder (2006) through 1999, which we extend from primary data through 2015.\(^8\)

For each election, we identify the largest (by seats) left party or PEC and the largest other party or PEC. We then compute the seat gap, \(V\), as a percentage of these top-two parties’ total parliamentary seats. The final dataset contains 576 elections, for which we were able to obtain applicable interest-rate data, our dependent variable, in 335 cases. Specifically, we use the long-term (10-year) interest rate on government bonds, reported monthly from the OECD (2018) and the IMF’s International Financial Statistics database (IMF, 2018).\(^9\)

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\(^7\) We define ‘Left’ parties as those labeled “Social Democracy” or “Communist/Socialist” in ParlGov. ParlGov’s ideology score is drawn from multiple studies estimating party ideology on a left-right scale, including Castles and Mair (1984), Huber and Inglehart (1995), and Benoit and Laver (2006). We assign pre-electoral coalitions the seat-weighted mean ideology score of their parties.

\(^8\) Unlike Golder (2006), our PEC data are not exhaustive. Because our analysis requires vote shares only of the largest left party and largest other party, we take particular care to identify PECs that affect these two values. These include, most prominently, the Liberal/National coalition in Australia, the CDU/CSU alliance in Germany, and the Red-Green / Centre-Right pre-electoral alliances in Sweden. We can ignore most of the more-numerous smaller-party PECs formed to overcome electoral vote-share thresholds.

\(^9\) We combine these sources to maximize country-year coverage, using IMF data where (rarely) they dis-
bond market response to the election is computed by taking the difference between this rate at \( t_0 \) (election month) and \( t_m \) (\( m \) months following the election).

4 Empirical Analysis

4.1 Evaluating the Mechanism

Throughout our empirical analysis, we follow current state-of-the-art practices in regression discontinuity design, as suggested by Calonico, Cattaneo and Titiunik (2014). This procedure (hereafter CCT) estimates two low-order local polynomial regressions (often linear) on each side of the threshold, using a triangular kernel to place greater weight on observations close to the threshold and dropping data outside of a bandwidth selected to minimize mean squared error of the RD estimator.\(^{10}\) The estimated treatment effect is the difference between the limits of these two regressions as they approach the threshold. Because this approach aims to be nonparametric (i.e., not assume a data-generating process), the CCT estimate subtracts a bias-correction term for any misspecification error in the estimation procedure and constructs robust confidence intervals centered around this bias-corrected estimate.

The first condition for validity of an RD design is that the discontinuity exists: in our application, e.g., that a left-party/PEC crossing to a seat plurality yields a discontinuous jump in the probability that left-party/PEC enters government. This assumption is easily verified directly. Figure 2 shows the relationship between the largest left-party/PEC plurality margin (\( V \)) and its inclusion in the cabinet. The figure reveals clearly a sharper discontinuity in some countries than in others. Countries with few parties in parliament exhibit a very sharp discontinuity (left panel), whereas in countries with more-fragmented party systems, and so more potential coalitions, a left plurality is less predictive of left government. In

\(^{9}\) agree. (Only Iceland exhibits any appreciable discrepancies.) All results are robust to reasonable alternative choices in these regards.

\(^{10}\) This procedure should be flexible enough to properly capture the theoretically expected form of the discontinuity seen in Figure 1.
what follows, we will define party-system fragmentation based on the Effective Number of Parliamentary Parties (ENPP). ¹¹ When fragmentation is low (ENPP < 3.5), there is a very sharp discontinuity at the plurality threshold, but where high (ENPP > 3.5), no statistically significant discontinuity manifests (see Table 1). ¹² Because there is no discontinuity in high-fragmentation party-systems, we can use this set of countries as a sort of placebo group. A narrow plurality for left parties should only cause market reactions if it provides new information about government formation. In high-fragmentation countries, it does not, so we should not expect to observe discontinuous bond-price movements in those countries. Note that the low-fragmentation countries are not exclusively majoritarian single-party governments. They also include some countries with proportional representation, like Spain and Portugal, and countries with frequent coalition governments, like Germany. (Figure 7 in the Supplementary Materials give descriptive statistics on party fragmentation over time.)

Table 1: First-stage regression discontinuity estimates (bias-corrected) with 95% confidence intervals (robust standard errors) in brackets.

<table>
<thead>
<tr>
<th>Fragmentation:</th>
<th>All</th>
<th>Low</th>
<th>High</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
</tr>
<tr>
<td>Discontinuity Estimate</td>
<td>0.233</td>
<td>0.575</td>
<td>0.037</td>
</tr>
<tr>
<td></td>
<td>[-0.035, 0.50]</td>
<td>[0.35, 0.80]</td>
<td>[-0.35, 0.42]</td>
</tr>
<tr>
<td>Observations</td>
<td>551</td>
<td>309</td>
<td>242</td>
</tr>
<tr>
<td>Observations Within Bandwidth</td>
<td>256</td>
<td>158</td>
<td>127</td>
</tr>
<tr>
<td>Bandwidth (h)</td>
<td>0.175</td>
<td>0.218</td>
<td>0.181</td>
</tr>
</tbody>
</table>

¹¹ENPP = \( \frac{1}{\sum p_i^2} \), where \( p_i \) is the seat share of party \( i \).

¹²Henceforth, we use this threshold (ENPP = 3.5) as the cutoff between Low and High party fragmentation. Appendix A, demonstrates that the results are robust to varying this choice. Appendix A also lists the country-years above and below this cutoff.
Figure 2: Left party plurality margin plotted against binary indicator of left party entering cabinet; curves are LOESS fits. Where party-system fragmentation is low ($ENPP < 3.5$), there is a sharp discontinuity in the probability of left parties entering government as they achieve a plurality of seats in parliament. Where party-system fragmentation is high ($ENPP > 3.5$), there is no such discontinuity.

4.2 Balance Tests

A crucial identifying assumption for the RD design is that the treatment must be the only variable that changes discontinuously at the threshold. If any other covariates do so as well, then one cannot unequivocally attribute a discontinuity in the outcome to the treatment alone. We subject a number of pre-treatment covariates to tests of this continuity condition. For each covariate, we estimate a local-linear RD (triangular kernel), testing whether the difference in expected value on either side of the threshold differs significantly from zero. The covariates we test include GDP per capita, Polity score, population, central government debt per capita, government expenditures, tax revenue per capita, annual inflation, and OECD average bond yields. That last test serves to ensure that our main results are not being driven by global movements in the bond market, but by country-specific bond-price changes.

Figure 3 uncovers no significant discontinuities for any of these covariates, except govern-
Figure 3: For nearly every pre-treatment covariate, there is no significant discontinuity at the threshold. Note that these tests are conducted for elections with low party fragmentation, as in the primary analysis, but the finding holds when looking at the entire sample as well.

ment expenditures; after elections yielding slight left-party pluralities, central government expenditures as a percent of GDP are slightly and marginally significantly lower. Although this seems unlikely to be the cause of a discontinuity in bond-yield increases, a robustness check considered in Appendix A deploys a variant of the RD estimator that conditions on covariates, as proposed by Calonico et al. (2018). These conditional-RD results are very similar to those of the primary analysis presented here.

4.3 Effects of Left-Party Entry to Government on Bond Markets

Table 2 reports, and Figure 4 illustrates, the estimated Local Average Treatment Effect (LATE) of left-party entry to government following closely contested winning of parliamentary plurality, in samples of high party-system fragmentation, in low-fragmentation systems,
and in all systems. In High-Fragmentation ($ENPP > 3.5$) countries, our placebo group, we see no statistically significant discontinuity in bond yields at the threshold. In contrast, in Low Fragmentation ($ENPP < 3.5$) countries, there is a roughly half-a-percentage-point increase in bond yields one month after a left party narrow plurality win. Precisely as our mechanism predicts, and striking in magnitude: left-party government, at least in low-fragmentation contexts where governments tend to be relatively efficacious single- or few-party majorities, are estimated to “cost” roughly half a percentage point higher interest-premia on government debt. Finally, with the benefit of comparison of results from low-fragmentation and high-fragmentation samples, the entire-sample LATE estimates can be seen as influenced by these heterogeneous-treatment effects to suggest only a marginally insignificant interest-rate increase less-than one-quarter as large as that in low-fragmentation contexts (and the whole-sample estimate is about 85% noisier proportionately to effect-size).

Table 2: 1-month bond yield regression discontinuity estimates (bias-corrected) with 95% confidence intervals (robust standard errors) in brackets.

<table>
<thead>
<tr>
<th>Fragmentation:</th>
<th>All</th>
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<th>High</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
</tr>
<tr>
<td>Local Average Treatment Effect</td>
<td>0.145</td>
<td>0.592</td>
<td>-0.048</td>
</tr>
<tr>
<td></td>
<td>[-0.12, 0.41]</td>
<td>[-0.01, 1.19]</td>
<td>[-0.32, 0.23]</td>
</tr>
<tr>
<td>Observations</td>
<td>316</td>
<td>179</td>
<td>137</td>
</tr>
<tr>
<td>Observations within Bandwidth</td>
<td>135</td>
<td>70</td>
<td>63</td>
</tr>
<tr>
<td>Bandwidth ($h$)</td>
<td>0.137</td>
<td>0.141</td>
<td>0.125</td>
</tr>
</tbody>
</table>

4.4 Further Exploration of Heterogeneous Treatment Effects

The strength of the bond-market response to left government likely depends on the counterfactual, namely, in this context, on how far that left government is from the alternative that would have formed had the left not gained plurality. We would expect bond-market movements to be strongest when the party that lost the plurality was strongly conservative.
relative to a strongly left party that won it. Whereas, if both parties vying for the plurality were broadly parties of the center-left, then the election outcome would have revealed little information to financial markets about likely bond-price relevant policies of the new government: bond prices would hardly move. To explore this hypothesis, we test how the RD estimate varies with the absolute difference between the plurality-contending parties’ ParlGov ideology score.\(^{13}\) Figure 5 plots these estimates (low-fragmentation elections only), showing further evidence that the treatment effect arises from the information-revelation that occurs when a left party narrowly gains plurality status: the LATEs are clearly larger when we restrict the sample to cases where the absolute difference in ideology score is large (e.g., greater than 3).\(^{14}\)

Figure 6 illustrates how the estimated LATE varies over time periods; each plotted point reports an estimate and confidence interval from a 30-year window of data starting from the year indicated on the X-axis (low-fragmentation elections only). This subdivision of

\(^{13}\)In so doing, we sacrifice some of the nonparametric character of our design to gain further insight (and efficiency and precision) by using these measures of ideology.

\(^{14}\)A caveat, however: the sample sizes for those estimates are relatively small (n = 69).
Figure 5: When we restrict our estimates to low-fragmentation elections where the two largest parties are ideologically distant, the estimated treatment effect grows.

the data further reduces statistical power, and should be interpreted cautiously; however, interestingly, the interest-rate effect of left government is largest and most discernible around the Bretton Woods era of fixed exchange-rates and low capital-mobility (from early 1950s through mid-to-late 1970s). This would be consistent with Rodrik’s “golden straightjacket” hypothesis—i.e., with considerable globalization-and-capital-mobility constraint on domestic governments’ policymaking autonomy—if the reason we are not finding interest-rate effects of left government in the later period is because left-government policies do not (or are not expected to) differ much from other governments. After all, furthermore, Mundell-Fleming suggests that monetary and fiscal policy should have been at least moderately maneuverable and effective in that 1950s-1970s era of limited capital-mobility and (so) imperfectly fixed exchange-rates; and both monetary and fiscal policy should have grown less maneuverable and/or effective as capital grew highly mobile and the remaining dependent central banks
increasingly independent (although exchange-rate-fixing efforts lessened this fiscal straightjacketing some for a time) (Clark, 2003; Franzese, 2003).

5 Conclusion

A discontinuous increase in the probability of a party’s entry to government at the parliamentary seat plurality threshold offers quasi-experimental identification of the causal effect, in that plurality-threshold vicinity, of that party on outcomes we think it may influence in government. We focused on the interest-rate-premium cost of left-party government that financial markets add to government-bond yields.

Applying a regression discontinuity design, we estimate that left government indeed car-
ries a sizable interest-rate cost, over 0.5 percentage points, but only in the short term—for about a year or a little more, peaking around 10 months—under specific conditions—relatively polarized left- and right-party governing alternatives in low-fragmentation party-systems—and limited time periods—namely, the Bretton Woods era from the early 1950s through the mid-to-late 1970s, when capital mobility was limited and exchange rates were imperfectly fixed (among our sample countries, \textit{de facto}, to the (extra-sample) U.S. dollar). Under these prevailing conditions, autonomous domestic government monetary and fiscal policy maneuverability and efficacy was relatively high. As these (and other: e.g., high and rising central-bank independence) conditions changed, domestic-government policy autonomy, maneuverability, and efficacy will have faded, which could explain the reduced magnitude and certainty of left interest-rate cost estimated in later periods.

Some important potential limitations of our approach merit mention. One issue surrounds our reliance on an \textit{ex post} measure of \textit{ex ante} uncertainty. We used an \textit{ex post} measure of how uncertain was the election outcome—how close to 50-50 was the proposition of left-party being largest—whereas the actually relevant concept for market reaction is how \textit{surprisingly} left was the government. To construct such an \textit{ex ante} measure of government-partisanship surprise, we would need pre-election polls or forecasts and party partisanship measures across our entire sample of country-years, and some manners of translating those to expectations, and a mapping from party to government partisanship; and each of these components would require further structural specifications. We believe our approach goes as far as possible, insofar as we wish to retain non-parametric causal-inference robustness over structural-specification efficiency. Moreover, we are confident our \textit{ex post} measure of electoral uncertainty is at least unbiased with respect to \textit{ex ante} electoral uncertainty. Some forecasted-close elections are landslides and some forecasted landslides are close, but both are rare and, we strongly expect, orthogonally random.

Another limitation is the small number of close elections. We count 135 close elections (ones with top two parties’ seat-differentials within the CCT optimal bandwidth) in parlia-
mentary democracies 1948-2015 for which we have data on bond yields. Even this relatively small sample-size, however, proved adequate given the large market reactions at the threshold in low fragmentation, stark alternative contexts. Null results, conversely, should be viewed with these same small-sample and low-power considerations in mind: emphasis on the fact that ‘absence of evidence is not evidence of absence’.

Finally, we must pool countries with widely disparate institutions and draw cases from an extended historical period. Although such broad pooling is common practice in comparative political economy, it does entail complications (Beck and Katz, 1995). Pooling majoritarian and proportional parliamentary democracies, e.g., may mask consequential other differences beyond that the former tend to produce single-party left and right governments whereas the latter tend to yield less-stark alternation of coalition governments. We explored this latter directly across our broad pool of democracies, saw that it generated heterogeneous effects, and we highlighted the important implications arising from this heterogeneity. Likewise, our time-period varying estimates were suggestive of important international-economic contextual heterogeneity (we believe prominently: variations in exchange-rate regimes and capital mobility). Other bases for heterogeneous treatment effects may also occur across this wide diversity of parliamentary democracies and eras. Unaccounted treatment-effect heterogeneity, akin to specification error, induces inefficiency at best (if the heterogeneity is unrelated to treatment) and bias at worst (if related).

To explore the accuracy of our interpretations in these regards, future analyses could explore bond-relevant economic-policy differences near the discontinuity. Are there differences in policy that could explain the variation in the estimated interest-premium effects corresponding to differences in domestic and international political-economic institutions and structure that we suggest are driving these heterogeneous effects we discovered? Inter alia, these analyses will help distinguish whether we are observing Downsian convergence in policies due to democratic competition and so in financial-market outcomes; convergence in policies and outcomes from globalization-induced policy competition; convergence in outcomes
but not in policy, indicating lack of market concern about those policies or suggesting some non-policy-related market reaction to left government; and/or some other political-economic institutional-contextual conditioning of partisan-government effects on policy and/or outcomes that could be driving these varying interest-premium costs of left government.

References


OECD. 2018. “Long-term interest rates (indicator).”.


Appendix A Supplementary Materials

A.1 Descriptive Statistics

Figure 7 plots party fragmentation – measured by ENPP – for each country-year in our dataset. As the figure makes clear, the Low Fragmentation countries are not exclusively majoritarian single-party governments. They include countries with proportional representation, like Spain and Portugal, and countries with frequent coalition governments, like Germany. Figure 8 plots the ideology of the largest left party and largest non-left party in
parliament, as measured by ParlGov’s left-right score. Countries with low party fragmentation and large distances between the largest left and non-left parties (the sharpest test for our theory) include the United Kingdom, Australia, Canada, Hungary, Japan, Spain, Portugal, France, Germany, Luxembourg, and pre-2000 Austria and Sweden.

A.2 Varying the ENPP Threshold

Figure 9 illustrates how the RD estimate and 95% confidence interval changes when we vary the ENPP threshold. The result holds for cutoffs less than 3.5, though when it is less than 3 the effective sample size is much smaller, yielding larger robust standard errors.

A.3 Regression Discontinuity with Covariates

Calonico et al. (2018) propose a procedure to adjust for covariates in a regression discontinuity framework. Estimating using this procedure – conditioning on GDP per capita, log population, expenditures per capita, tax revenue as a percentage of GDP, and inflation – yields the results in Table 3.

Table 3: Regression discontinuity with pre-treatment covariates. Dependent variable = 1-month bond yield regression discontinuity estimates (bias-corrected) with 95% confidence intervals (robust standard errors) in brackets.

<table>
<thead>
<tr>
<th></th>
<th>All</th>
<th>Low</th>
<th>High</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
</tr>
<tr>
<td>Local Average Treatment Effect</td>
<td>0.146</td>
<td>1.01</td>
<td>-0.146</td>
</tr>
<tr>
<td></td>
<td>[-0.14, 0.43]</td>
<td>[0.46, 1.56]</td>
<td>[-0.45, 0.16]</td>
</tr>
<tr>
<td>Observations</td>
<td>236</td>
<td>118</td>
<td>118</td>
</tr>
<tr>
<td>Bandwidth Estimate (h)</td>
<td>0.138</td>
<td>0.135</td>
<td>0.129</td>
</tr>
</tbody>
</table>
A.4 Sensitivity to Bandwidth

As Figure 10 illustrates, our result is somewhat sensitive to choice of bandwidth. Using a smaller bandwidth than the CCT optimum yields significantly fewer observations near the cutoff, increasing standard errors. And the local-linear estimate shrinks with larger bandwidths (as one would expect given the theoretical conditional expectation function presented in Section 2).

A.5 Sensitivity to Polynomial Order and Kernel Function

The \texttt{rdrobust} package defaults to modeling the conditional expectation function with a local-linear and triangular kernel weights. The core result is robust to varying these assumptions, though the 95% confidence interval is wider (and includes zero) when estimated using a second-order polynomial.

Table 4: Regression discontinuity estimates, varying polynomial order and kernel function. Dependent variable = 1-month bond yield regression discontinuity estimates (bias-corrected) with 95% confidence intervals (robust standard errors) in brackets.

<table>
<thead>
<tr>
<th>Fragmentation:</th>
<th>All</th>
<th>Low</th>
<th>High</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
</tr>
<tr>
<td>Linear, Triangular Kernel</td>
<td>0.145</td>
<td>0.592</td>
<td>−0.048</td>
</tr>
<tr>
<td></td>
<td>[-0.12, 0.43]</td>
<td>[-0.01, 1.19]</td>
<td>[-0.32, 0.23]</td>
</tr>
<tr>
<td>Linear, Uniform Kernel</td>
<td>0.152</td>
<td>0.511</td>
<td>−0.077</td>
</tr>
<tr>
<td></td>
<td>[-0.20, 0.39]</td>
<td>[0.10, 0.92]</td>
<td>[-0.39, 0.23]</td>
</tr>
<tr>
<td>Quadratic, Triangular Kernel</td>
<td>0.099</td>
<td>0.646</td>
<td>−0.109</td>
</tr>
<tr>
<td></td>
<td>[-0.20, 0.39]</td>
<td>[−0.11, 1.40]</td>
<td>[-0.43, 0.21]</td>
</tr>
<tr>
<td>Quadratic, Uniform Kernel</td>
<td>0.106</td>
<td>0.598</td>
<td>0.016</td>
</tr>
<tr>
<td></td>
<td>[-0.19, 0.41]</td>
<td>[−0.14, 1.33]</td>
<td>[-0.31, 0.34]</td>
</tr>
<tr>
<td>Observations</td>
<td>316</td>
<td>179</td>
<td>137</td>
</tr>
</tbody>
</table>
A.6 Dynamic RD Estimates

In addition to estimating the short-run effects on bond yields, we can compute a dynamic RD estimate – actually a series of static estimates – by varying the time window with which the dependent variable is measured (Cellini, Ferreira and Rothstein, 2010). Using this method, we can observe whether a left party plurality yields longer-term changes to bond yields. Furthermore, as a placebo test, we can also examine whether there is an estimated effect on bond yields prior to the election. Figure 11 illustrates the results from this moving-window RD analysis; the narrow plurality win of a left party appears to have an effect on bond yields that persist for at least a year, peaking at around 10 months, when party fragmentation is low, and, again, no perceptible effect at any time-window in high-fragmentation contexts.

A.7 Are close elections more ex ante uncertain?

One advantage of our research design is that it does not require strong assumptions about how bond traders form expectations about election outcomes. It only requires that those expectations do not vary discontinuously at the plurality threshold (see Section 2). This assumption is difficult to test. Evidence from prediction markets (Snowberg, Wolfers and Zitzewitz, 2007) are only available for a small subset of elections in our sample. A somewhat larger dataset of ex ante election uncertainty comes from Kayser and Lindstädt (2015), who combine two measures – (1) the observed historical volatility of party vote shares, and (2) the seats-to-votes elasticity – to estimate the probability that largest parliamentary party will lose their plurality in the following election.

This measure is a poor fit for our main analysis for a few reasons. First, it covers fewer than half the elections for which we have bond yield data (150 out of 316). Second, their electoral risk measure is computed as of the previous parliamentary election. Because it does not incorporate any new information or polls that would be available to bond traders, it is a noisy measure of the market’s expectations immediately prior to the election. Finally, it
does not measure electoral risk for pre-electoral coalitions, unlike our analysis.

These caveats aside, it is useful as a robustness test to observe whether this *ex ante* measure of electoral uncertainty varies discontinuously at the plurality threshold. Using the Kayser and Lindstädt (2015) measure, we estimate for each election the probability that a left party will lose its plurality of seats (or, equivalently, that a right party will retain a plurality of seats). Figure 12 plots this probability against the actual left party plurality margin for elections with low party fragmentation (ENPP < 3.5).

As expected, the *ex ante* loss probability declines as the *ex post* seat share rises, with more uncertainty for elections near the plurality cutoff than those farther away. However, there appears to be a small discontinuity in *ex ante* loss probability at the cutoff, though there are so few observations near the cutoff that the estimate is imprecise. This is troubling for the identification assumption of our primary analysis. Perhaps with more data, this apparent discontinuity would disappear, but we are unable to eliminate it as a potential confounder. To ameliorate this concern, we repeat the analysis from Table 3, adding the electoral loss probability as a covariate in the RD analysis. The results, reported in Table 5, are similar to those reported above.

Table 5: Regression discontinuity with pre-treatment covariates, including the estimated loss probability for left parties, estimated by Kayser and Lindstädt (2015). Dependent variable = 1-month bond yield regression discontinuity estimates (bias-corrected) with 95% confidence intervals (robust standard errors) in brackets.

<table>
<thead>
<tr>
<th>Fragmentation:</th>
<th>All</th>
<th>Low</th>
<th>High</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>(1)</td>
<td>(2)</td>
<td>(3)</td>
</tr>
<tr>
<td>Local Average Treatment Effect</td>
<td>0.203</td>
<td>0.665</td>
<td>0.023</td>
</tr>
<tr>
<td></td>
<td>[-0.033, 0.524]</td>
<td>[0.089, 1.669]</td>
<td>[-0.185, 0.321]</td>
</tr>
<tr>
<td>Observations</td>
<td>150</td>
<td>78</td>
<td>72</td>
</tr>
<tr>
<td>Bandwidth Estimate (<em>h</em>)</td>
<td>0.112</td>
<td>0.137</td>
<td>0.173</td>
</tr>
</tbody>
</table>
A.8 Results From Daily Bond Yield Data

Ideally, we would measure our outcome variable using daily bond price data, to observe how traders react immediately following an unexpected left party victory. Unfortunately, these data are not available for every election in our dataset, particularly those during the Bretton Woods period. We cannot match daily bond price data to any election prior to 1980, and can only do so for 12 elections between 1980-1990 and 29 elections between 1990-2000. As we note in Section 4, the period prior to 1980 is of particular interest, and we estimate the largest effect of left party victories on bond prices during that era (see Figure 6). Replicating the regression discontinuity analysis using the available daily data yields Figure 13. Our estimated treatment effect during this period, for both low and high fragmentation elections, is an imprecisely estimated null, consistent with the results from the analysis of monthly data.
Figure 7: Party fragmentation in parliament by country-year. The dashed horizontal line is our cutoff for classifying low vs. high fragmentation elections.
Figure 8: The ParlGov left-right scores of the largest left party and the largest other party in parliament. For pre-electoral coalitions, this value is the seat-share-weighted average score of the coalition members.
Figure 9: Estimated treatment effects and 95% confidence intervals, varying the ENPP threshold. The estimate is largest when party fragmentation is low (ENPP < 3.5)
Figure 10: Sensitivity to choice of bandwidth. Bias-corrected estimates and 95% confidence intervals (robust standard errors). The CCT optimal bandwidth is plotted in red.
Figure 11: Dynamic RD estimates. Top panel is for low fragmentation elections ($ENPP < 3.5$), bottom panel is for high fragmentation elections ($ENPP > 3.5$).
Figure 12: Estimated *ex ante* probability of a left party loss as of the previous election (Kayser and Lindstädt, 2015), plotted against the plurality margin of the largest social democratic party; curves are LOESS fits.
Figure 13: Interest rate change the day before and after election, plotted against the plurality margin of the largest social democratic party; curves are LOESS fits.