## Rating Politics? Partisan Discrimination in Credit Ratings in Developed Economies

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**Abstract**: How does government partisanship influence sovereign credit ratings of developed countries? Given the convergence of fiscal and monetary outcomes between left and right governments in the past decades, credit rating agencies (CRAs) should in principle not discriminate according to ideology. However, we hypothesize that CRAs might lower ratings for left governments as a strategy to limit negative policy and market surprises as they strive to keep ratings stable over the medium term. A panel-analysis of Standard and Poor’s, Moody’s and Fitch’s rating actions for 23 OECD countries from 1995 to 2014 shows that left executives and the electoral victory of non-incumbent left executives are associated with significantly higher probabilities of negative rating changes. We find no evidence of similar systematic partisan bias in spreads on government bonds, but spreads do adjust to partisan-biased downgrades. This suggests that CRAs introduce partisan discrimination into sovereign credit markets.

Credit rating agencies (CRAs) have emerged as important gatekeepers to sovereign debt markets over the past three decades. The power of the largest CRAs – Moody’s, Standard and Poor’s (S&P) and Fitch[[3]](#footnote-3) – over the fiscal fortunes of prosperous developed countries was thrown into sharp relief when they initiated a series of downgrades that triggered and perpetuated a succession of sovereign debt crises in Europe between 2008 and 2012. While CRAs’ decisions are mostly motivated by the emergence of news about a country’s fiscal and economic performance, anecdotal evidence (such as the 2011 downgrading of the United States over the debt-ceiling debate or the 2012 collective downgrade of nine Eurozone member states over perceived governance problems within Europe’s common currency area) indicates that CRAs incorporate political considerations into their assessment of highly rated (and politically stable) sovereigns’ debt.

Although sovereign ratings received ample and growing scholarly attention in the past two decades, it remains unclear what sort of political factors influence the sovereign credit rating scores of prosperous developed countries. Most studies have concentrated on identifying key economic and institutional factors that have the greatest effect on rating scores (e.g. Cantor and Packer, 1996; Afonso, 2003; Afonso, Gomes, Rother, 2007). Scholars who examined the influence of political factors on CRA decisions did so strictly in the context of developing economies, whose access to global financial markets is especially sensitive to the pronouncements of rating agencies. Most focused on large-scale systemic factors, such as regime type or political stability, which show greater variation in (and therefore, have particular relevance for) developing countries (e.g. Haque et al, 1998; Block and Vaaler, 2004 and 2006; Vaaler, Schrage and Block, 2006; Archer et al, 2007; Beaulieu et al, 2012; Biglaiser and Staats, 2012).

We explore whether government partisanship has an effect on the sovereign credit ratings of prosperous developed countries. Thereby, we seek to make three contributions. First, we focus on sovereign credit rating activity in an important, but so far scantily explored, group of countries (developed economies) whose vulnerability to changes in credit ratings has been highlighted by the unprecedented frequency of downgrades since 2008. While emerging market economies are more likely to encounter debt crises and defaults due to their underdeveloped financial markets (Reinhart and Rogoff, 2008), recent events suggest that sovereigns with fully developed financial markets are not as immune to sovereign risk as previously thought. Because developed countries characteristically have more stable economies and better access to financial markets than developing economies, domestic politics is likely to play a more significant role in the variation of perceived sovereign creditworthiness. Second, we shift the attention from systemic political factors (which are at the focus of most existing scholarship) to the role of day-to-day politics (elections and government partisanship) in sovereign credit ratings. Two studies suggested that CRAs discriminate against the center-left, particularly after elections in developing economies (Block and Vaaler, 2004; Vaaler, Schrage and Block, 2006). We seek to establish the presence or absence of such partisan discrimination in developed countries, which (by virtue of being stable democracies) provide particularly fertile ground for exploring how everyday politics enters assessments of creditworthiness, independently of concerns about political instability.

Third, by exploring whether and why CRAs treat executives of different partisan color differently in developed countries, we contribute to the debate over how markets influence and constrain politics in advanced economies. Evidence of partisan discrimination would lend support to claims that economic globalization places governments into a “golden straitjacket”, curtailing the scope for democratic choices in return for providing abundant access to financing (Rodrik, 2000). The absence of partisan discrimination, on the other hand, would suggest that developed countries (unlike their developing counterparts) have substantial “room to move” in their political and policy choices even in a globalized financial world (Mosley, 2000). Furthermore, exploring whether and how CRAs (highly visible market actors specializing in judging creditworthiness) discriminate between governments of different ideologies sheds light on the mechanisms through which politics may influence market outcomes.

We argue that although CRAs should in principle not differentiate on the basis of government partisanship once relevant policy differences are controlled for, they might espouse the expectation that left governments pursue less credit-friendly policies as a mechanism to insure against negative policy and market surprises over the medium term, since they cannot frequently adjust their ratings to incorporate new information as it arises. Keeping ratings stable and reliably reflective of all potential risks to credit quality is essential if a CRA wants to ensure that its ratings are used as third-party credit risk measures in financial governance documents, which is crucial for retaining their credit rating market share. CRAs themselves emphasize in their official communications the need to combine reliable assessment of risks with relative stability. A conservative strategy to limit the negative surprises arising from the uncertainty involved in making rating decisions over the medium term is to preemptively incorporate all perceived risks into ratings, which implies assigning lower ratings to governments who have less political incentive to pursue policies that CRAs deem conducive to good credit quality. Bond markets, in contrast, do not operate under similar medium-term “stability” constraints, because they can make adjustments upon receiving new information about governments’ actual policy choices instantaneously.

In our quantitative analysis, we test for the presence of partisan bias in the ratings of each of the big three agencies as well as the market’s pricing of sovereign risk in order to gauge the extent to which different market actors employ partisan discrimination. We find that CRAs discriminate against the left, but markets do so only insofar as they incorporate partisan-biased credit ratings into their decisions. We employ a linear probability model panel analysis of negative ratings decisions by S&P, Moody’s and Fitch for 23 OECD economies from 1995 to 2014[[4]](#footnote-4), and then compare these results to an identical (OLS) panel analysis of spreads on long-term government bonds (a proxy of market attitudes towards sovereign creditworthiness) for the same sample. We find that, *ceteris paribus,* downgrades were more likely from S&P and Moody’s under left-controlled executives (i.e. executives headed by left prime ministers). Fitch does not display similar biases against left prime ministers in their (negative) ratings decisions, but they do exhibit partisan bias in downgrades against executives with strongly left-oriented manifestos and cabinets with high shares of seats occupied by left parties (i.e. single-party left administrations). In addition, the probability of negative rating changes is even higher for S&P and Moody’s upon the election of non-incumbent left governments. While Moody’s rating penalties against left governments is weak to robustness checks, and appears to be driven by the European debt crisis, S&P’s left government and left electoral victory biases precede 2009 and are robust to numerous alternative model specifications. In contrast, spreads on long-term government securities are not directly influenced by the partisanship of ruling executives, or by (non-incumbent) left electoral victories. Yet bond spreads *are* impacted by CRA negative rating decisions, which implies that CRAs’ ratings changes may be a crucial transmitter of partisan effects into sovereign bond markets.

The evidence of systematic negative discrimination by CRAs against left governments in developed countries contradicts the notion that developed countries are unencumbered by the political influence of markets. Furthermore, the absence of similar discrimination in spreads suggests that CRAs may be a source of the partisan impetus, rather than markets at large. In an effort to avoid to frequent rating changes as well as negative policy and market surprises, credit ratings systematically give markets partisan signals, which are not linked to actual differences in fiscal performance under the left and the right. Thereby, they introduce theoretically questionable factors into assessments of creditworthiness. This reinforces already existing concerns about giving credit ratings a central role in prudential regulations.

**Partial arbiters? Creditworthiness, partisanship and business strategies in the credit rating market**

Sovereign credit rating agencies wield considerable influence over countries’ access to credit through their pronouncements about creditworthiness. Although investors are in principle free to ignore CRAs’ opinion (and sometimes do[[5]](#footnote-5)), credit ratings systematically influence the price of government debt (Brooks et al. 2004, Afonso et al 2015; also see our results below). This is partly due to the authority and visibility of the three large CRAs, but also because upgrades and downgrades automatically trigger automatic portfolio adjustments by investors who need to comply with portfolio regulations (Sinclair 2008, De Haan and Amtenbrink 2011).

The practical implications of a potential partisan effect on sovereign credit ratings are clear. If left governments consistently face higher probability of downgrades than right ones *all else being equal*, they face more difficult market conditions and political challenges. First, left governments could experience higher borrowing costs and, consequently, tighter constraints on their fiscal policies. Second, since perceived sovereign creditworthiness can have knock-on effects on the private sector’s access to foreign capital, lower ratings can generate more adverse financial conditions for businesses under left governments independently of the government’s policy performance. Third, given that downgrades can trigger capital flight at times of economic and financial turmoil, countries may be at greater risk of plunging into a debt-crisis under left governments. Finally, beyond these financial disadvantages, partisan discrimination in ratings can also create political costs for left governments if proneness to negative rating actions is interpreted as a sign of inferior ability to manage the economy.

The theoretical relevance of potential partisan discrimination in sovereign credit ratings lies in shedding light on the markets-politics nexus and of the distinctive role of CRAs in shaping that nexus. Since governments tapped into international credit markets to gain access to more abundant and cheaper funding, scholars have debated to what extent their political and policy choices would be constrained by markets. Some argued that increased trade and capital mobility has limited the autonomy of governments in macroeconomic policy, and highlighted the convergence in macroeconomic indicators, especially deficits and inflation, as a sign of these constraints (Garrett and Lange, 1991; Boix, 2000). Others went further to argue that markets place governments in a “golden straitjacket”: they severely limit the role of democratic politics in policy choice in general (Rodrik, 2000). Mosley qualified this hypothesis and claimed that markets treat developed and developing countries differently. She argued that “financial markets’ influences on developed democracies are somewhat strong, but somewhat narrow”: while they pay attention to macroeconomic indicators, markets are less concerned about supply-side policies such as welfare provisions or the size of the public sector (Mosley 2000, p766 and 2004). Prosperous developed countries are given greater “room to move”, because their ability to repay their debt is not fundamentally questioned (Mosley 2005). However, concerns about the power of financial markets over democratically elected (but underfunded) governments strongly resurfaced in the wake of the global financial and economic crisis (Streeck 2014).

Market constraints manifest themselves as partisan discrimination, if governments of different partisan color pursue systematically different policies. The rational partisanship theory posits that as representatives of labor, left parties are more likely to pursue expansionary policies to keep unemployment low, at the expense of higher inflation and deficits (Hibbs, 1977). By the same token, they are assumed to have stronger commitment to generous welfare policies than to fiscal stability. The expectation of higher deficits from the left received some empirical support in the past (De Haan and Strum, 1994; Boix, 2000), but has lost empirical validation in recent decades. Newer studies report little systemic relationship between left incumbents and deficits or the willingness to implement fiscal cuts to consolidate public finances (Mierau et al., 2007, Hübscher 2016). Other political factors (i.e. power-sharing in multiparty or minority governments) have been found to be more important in explaining macroeconomic and government spending outcomes than partisanship (Bawn and Rosenbluth, 2006; Breen and McMenamin, 2013). Indeed, our own difference-in-means tests demonstrate that left executives in our OECD sample presided over similar levels of average debt and deficits levels as right executives since 1995. While there may have been a relationship between partisanship and policy choices in the past, this relationship has faded over time in developed countries (Cusack, 1997).

If there is no systematic connection between partisanship and fiscal and monetary outcomes, markets should in principle not discriminate between left and right; but evidence suggests they do. Markets have been shown to react to the prospect of a left government in office with drops in stock prices (Bechtel, 2009; Sattler 2013), increased inflationary expectations (Fowler, 2006), and greater exchange rate volatility (Bernhard and Leblang, 2002) even in developed countries. However, sovereign bond yields (indicators of sovereigns’ creditworthiness) do not seem to be systematically affected by government partisanship over time according to recent political economy literature. The impact of ideology on yields is modulated by the extent to which institutional constraints allow governments of different partisan color to pursue significantly differing policies (Breen and McMenamin, 2013). The conundrum between immediate market reactions to changes in government partisanship and the lack of systematic effect on bond yields in the long run can be resolved if we consider that markets flexibly adjust prices as information about actual policy choices becomes available, diminishing the initial effects of partisan expectations over the longer term.[[6]](#footnote-6)

In contrast to markets, CRAs potentially have motivation to systematically employ partisan expectations in their rating decisions due to a combination of three factors: an express preference for social and economic policies (above and beyond maintaining fiscal and monetary stability) that right-wing governments tend to embrace; the inability to adjust ratings upon new information as flexibly as markets do; and potentially the urge to insure against downside policy and market risks when dealing with the uncertainty entailed in making rating decisions over the medium run. We elaborate on all three of these factors in turn.

Although none of the CRAs discuss the effect of government ideology on sovereign creditworthiness, S&P and Fitch expresses preferences for economic and social policies commonly associated with the right in their sovereign credit rating methodologies.[[7]](#footnote-7) All iterations of the big three’s sovereign rating methodologies[[8]](#footnote-8) steer clear of commenting on the effect of government partisanship on sovereign creditworthiness, not even mentioning the words “ideology”, “party”, “left” or “right”.[[9]](#footnote-9) At the same time, all CRAs’ methodologies express a preference for conservative macroeconomic policies, such as low and stable inflation, and low deficits and public debt. Moreover, earlier editions of S&P and Fitch’s methodology (Fitch, 2002; S&P, 2006 and 2008) also clearly advocated trade, labor market and financial liberalization; praised small governments and low taxation; cautioned against large state-owned enterprises and entitlements; and urged welfare reforms to increase fiscal sustainability amidst population aging (Moody’s methodologies and later iterations of S&P’s and Fitch’s methodologies do not explicitly comment on these economic and social policy issues). Such preferences make it plausible that CRAs prefer political forces that embrace such policies, although in principle they should only make a distinction between right and left insofar as the relevant policy variables actually differ significantly under the tenure of the two sides.

CRAs might have incentives to use ideological labels as indicators of future policy choices – instead of waiting to see the actual policy choices – because they cannot flexibly adjust ratings every time there is a new policy development. CRAs need to keep their ratings stable (and reliably reflective of all possible downside risks), because rating stability and reliability is key to the use of their ratings as third-party credit risk measures in financial governance documents like portfolio mandates, contractual stipulations or even public regulations. Widespread use of a CRA’s ratings in such documents is crucial to retaining market share. CRAs offer their services to borrowers who want to advantageously place their bonds on the market. While this creates short-term incentives for CRAs to inflate ratings to attract bond issuers (Becker and Milbourn 2010), CRAs can only retain market share in the long-term by maintaining investors’ willingness to use their ratings to orient investment decisions, because even the most favorable rating is useless for a bond issuer unless it induces investors to hold the bond. The dominant channel through which a given CRA’s ratings motivate large-scale buying and selling decisions is their use in financial governance rules. Most mutual funds, pension funds, insurance companies, private endowments, and foundations use credit ratings to set minimum credit standards (IMF 2010, p92), and the portfolios of such institutional investors automatically have to be adjusted when ratings change. (Ratings are unlikely to achieve the same effect purely through providing risk analysis to investors, as most investors rely on in-house research to orient their decisions besides ratings). Making a CRA’s ratings suitable for use in financial governance documents requires not only that they reliably reflect all reasonable risks to credit quality, but also that they cannot change too frequently, because the transaction costs associated with automatic portfolio adjustments would make the given agency’s ratings prohibitively costly to use (Cantor and Mann 2007).

A conservative approach to dealing with the trade-off between rating stability and reliability is to incorporate all perceived possible future threats to credit quality into ratings ahead of time. Of the three rating agencies, S&P explicitly embraces such an approach. In its Credit Stability Criteria of 2016 (S&P 2016), S&P explains that it assigns lower ratings (to any type of issuer) than would arise from the analysis of current credit quality, if the issuer’s creditworthiness could slightly/significantly deteriorate in one/three years when encountering hypothetical scenarios representing “moderate stress” on credit quality. The document does not spell out the range of hypothetical scenarios (for any class of issuers), so it remains unspecified what sources of stress are taken into consideration when testing the susceptibility of sovereign credit quality. Presumably, however, they include significant adverse changes in the variables and policies specified in S&P’s sovereign rating methodology. This would not only imply policy choices disapproved by CRAs, but also potential adverse knock-on effects on economic and financial variables as a result of market responses to such policy choices.

The other two CRAs’ approach to the stability-reliability trade-off is more difficult to gauge. In a Special Comment (Cantor and Mann 2007), Moody’s justifies its commitment to ensuring rating stability, provides estimates of the losses of accuracy that result from greater stability and concludes that the trade-off seems to suit investors’ needs. However, Moody’s does not disclose whether this trade-off is achieved by factoring contingent future risks into ratings (similarly to S&P’s practice) or by foregoing adjustments upon receiving new information as long as it does not drastically affect credit quality. We have not been able to locate similar official communication about rating stability and reliability from Fitch. Therefore, while we expect S&P to systematically discriminate between left and right governments in its rating decisions, it is not clear whether the other two would translate their similar policy preferences into preemptive rating changes linked to partisanship.

In sum, while there is no robust empirical evidence that currently links left governments to inferior fiscal outcomes, we expect that (some) CRAs might still systematically discriminate against left governments in an effort to manage the uncertainty of having to keep ratings stable over the medium term, given their preference for a range of economic and social policies usually associated with right governments. In order to better explore the validity of this reasoning, we test a series of hypotheses that examine the effect of (new and incumbent) left electoral victories and the left’s tenure in power on rating decisions. We differentiate between change in government and tenure in government in order to better understand the effect of political uncertainty around elections and separate it from a more generalized mistrust towards left governments. Our analysis incorporates the possibility that left governments that have been in power for a while benefit from having an established policy record, whereas newly incoming left governments are more subject to negative rating movements.

At the same time, we also test two auxiliary hypotheses about investor reactions to partisanship and about the relationship between investors' decisions and ratings in an effort to better locate any potential source of partisan discrimination within sovereign credit markets. First, we investigate whether partisanship affects the pricing of sovereign debt to test the results of previous studies that fail to show systematic differences in perceived creditworthiness that are driven by government partisanship, *once controlling for other factors*. Second, we probe our assumption that ratings have an impact on prices and are, therefore, materially consequential.

*Hypothesis 1: Ceteris paribus, negative rating actions[[10]](#footnote-10) (downgrades and worsened outlooks) are more likely*

*(a.) after the election of a new left executive;*

*(b) upon the re-election of an incumbent left executive;*

*(c) during the term an incumbent left executive.*

*Hypothesis 2 (Auxiliary hypothesis): Ceteris paribus, spreads are higher*

*(a.) after the election of a new left executive,*

*(b) after the re-election of an incumbent left executive and*

*(c) during tenure of a left executive.*

*Hypothesis 3 (Auxiliary hypothesis): Spreads are higher in the wake of negative rating actions.*

**Partisan discrimination in sovereign ratings: Empirical evidence from 23 OECD economies**

We test the hypotheses above employing a series of panel analyses to determine the extent to which government partisanship affects ratings, and to infer how CRAs’ and markets’ views of government partisanship are related. Therefore, we have two dependent variables. Our primary dependent variable is the likelihood of negative rating actions. Our secondary dependent variable is spreads. Spreads are defined as the interest rate on the long-term treasury bill of the given country minus the interest rate for the 10-year US Treasury bond.[[11]](#footnote-11) Rating actions are CRA communications about a sovereign’s creditworthiness. They can be downgrades, confirmations or upgrades, but also revisions in the outlook assigned to the rating. Outlooks can be positive, negative and stable, signaling the likely changes in the country’s rating score over the course of the next two years. Outlooks are not necessarily followed by a rating change and do not trigger automatic portfolio changes like downgrades or upgrades. However, they transmit important information to markets about the changing opinions of the CRA about a country. In fact, the IMF (2010; pg 105) finds that CRAs influence market prices to a greater extent through outlooks and credit watches than through actual rating changes.

We focus on rating actions as our dependent variable rather than actual rating scores for two reasons. First, the ratings of developed sovereigns are characterized by severe clustering at the top end of the rating scale. AAA ratings make up roughly 55%, 65% and 55% of country-years in our panels for S&P, Moody’s and Fitch, respectively. (Moreover, only 1.5%, 2.1% and 2.2% of country-years for our S&P, Moody’s and Fitch panels, respectively, are speculative grade). This produces extreme “ceiling effects”. Extreme ceiling effects lead to biased parameters, particularly for binary independent variables, which our political variables are coded as (see Wang et al, 2007). In this respect, our developed country sample generates methodological challenges that comparable studies working with broader samples of sovereign ratings do not face.[[12]](#footnote-12) Second, focusing on rating actions allows us to take into consideration those signals of changing CRA opinion (specifically worsened outlooks) that do not necessarily result in actual downgrades, but are nevertheless consequential for prices (as the above-referenced IMF study suggests). Despite these methodological considerations, we also test our hypotheses about partisanship and elections using rating scores instead of ratings changes as the dependent variable, and report our results in Appendix A.[[13]](#footnote-13)

Our analysis examines the partisan determinants of rating actions for each CRA individually rather than collectively, in order to better understand each agency’s strategy towards government partisanship. The source of information on the date and nature of rating actions is S&P’s Sovereign Rating and Country T&C Assessment Histories (September 4, 2013) and various rating reports, Moody’s Rating List[[14]](#footnote-14) and Fitch’s Sovereign Ratings History[[15]](#footnote-15). Between 1995 and 2014, Fitch, Moody’s and S&P initiated 39, 39 and 52 downgrades and 26, 25 and 50 worsened outlooks, respectively (see Figure 1).

**Figure 1: Rating Events by Credit Rating Agency (23 OECD countries, 1995-2014)**



We use two codings for our “rating action” dependent variable. One measures whether a country experienced a downgrade in year t (1 if yes, 0 if otherwise), while the other measures whether a country experienced a negative rating event more broadly (a downgrade OR a worsened outlook decision) in year t (1 if yes, 0 if otherwise). Table 1 provides descriptive statistics of these dependent variables for our annual and quarterly data, as well as our independent variables outlined below.

**Table 1: Descriptive Statistics**

|  |  |  |
| --- | --- | --- |
|  | 1995-2014 (Annual Data) | 1995-2014 (Quarterly Data) |
|   | Mean | Std Dev | Mean | Std Dev |
| **Dependent variables** |   |   |   |   |
| Downgrade, S&P’s (1=yes) | 0.078 | 0.269 | 0.026 | 0.159 |
| Downgrade, Moody’s (1=yes) | 0.059 | 0.235 | 0.020 | 0.140 |
| Downgrade, Fitch (1=yes) | 0.060 | 0.238 | 0.020 | 0.141 |
| Downgrade or worsened outlook, S&P’s (1=yes) | 0.135 | 0.342 | 0.048 | 0.213 |
| Downgrade or worsened outlook, Moody’s (1=yes) | 0.085 | 0.279 | 0.034 | 0.180 |
| Downgrade or worsened outlook, Fitch (1=yes) | 0.098 | 0.298 | 0.034 | 0.180 |
| Spread on government bonds | 0.374 | 2.067 | 0.174 | 1.977 |
| **Independent Variables** |   |   |   |   |
| Left controlled cabinet (1=yes) | 0.399 | 0.490 | 0.391 | 0.484 |
| Cabinet manifesto score | -0.335 | 16.489 | NA |
| Government debt (% of GDP) | 67.327 | 39.005 | NA |
| Δ in debt | 1.128 | 5.871 | NA |
| Net government lending (% of GDP) | -1.646 | 4.929 | NA |
| Δ in government lending | 0.136 | 2.603 | NA |
| Trade balance (% of GDP) | 3.241 | 7.910 | NA |
| Unemployment | 7.120 | 3.821 | NA |
| Inflation | 2.073 | 1.545 | 2.074 | 1.642 |
| Election year (1=year) | 0.272 | 0.445 | 0.068 | 0.252 |
| Left incumbent electoral victory (1=yes) | 0.054 | 0.227 | 0.014 | 0.116 |
| Left non-incumbent electoral victory (1=yes) | 0.046 | 0.209 | 0.011 | 0.106 |
| Minority government (1=yes) | 0.233 | 0.423 | 0.234 | 0.421 |
| Coalition majority government (1=yes) | 0.563 | 0.497 | 0.564 | 0.493 |
| Single-party majority government (1=yes) | 0.193 | 0.395 | 0.191 | 0.392 |
| Government expenditure (% of GDP) | 44.758 | 7.072 | NA |

Data sources: Standard and Poor’s Sovereign Rating and Country T&C Assessment Histories, EU’s AMECO Database, OECD, Armingeon et al (2014).

Our data cover 23 OECD countries[[16]](#footnote-16) from 1995 to 2014. Although the earliest sovereign credit ratings were issued in the mid-1970s, we conduct our analysis from the mid-1990s onwards for two reasons. First, one of the “big three”, Fitch, entered the ratings market relatively late, issuing its first ratings in 1994. Moody’s and S&P also extended their sovereign rating portfolio gradually. Although they covered most of the developed world by the late 1980s, some outliers (e.g. Luxemburg) were only rated as late at 1994. As a result, 1994 is the first year for which all sovereigns are rated by the “big three” and, therefore, 1995 is the first year for which ratings changes can be observed. Our second, and more significant reason for starting our investigations in the mid-1990s is that it allows us to observe the “mature phase” of the sovereign ratings market. From the 1970s to the mid-1990s, the nature of the sovereign rating market transformed considerably as the weight of institutional investors increased, more than doubling both as a share of the economy and a share of claims held (Davis and Steil, 2004: pp 5-7). This shift coincided with increased economic volatility in developed countries, exemplified by the 1992 EMS crisis, arguably increasing the appeal of more conservative rating strategies. Starting our investigations in the mid-1990s allows us not only to work with relatively constant set of countries, but also to observe CRAs’ behavior under uniform financial market characteristics and economic conditions.

We conduct our panel analysis for this time period using both annual data (results in Tables 2 and 3) and quarterly data (results in Tables 4). We chose to complement our annual panel with one using quarterly data, in order to better capture the effect of within-year events, like elections and changes in government composition.[[17]](#footnote-17) Our results based on annual and quarterly data are largely consonant. In our discussion below, however, we prioritize results from the annual data for two reasons. First, quarterly economic data is less available than annual data. Crucially, fiscal indicators such as gross debt and net lending, which are arguably the single most important determinants of sovereign creditworthiness, are only available on an annual basis in all major macroeconomic databases (OECD, Eurostat, the IMF, etc). Excluding fiscal data from the analysis poses a significant omitted variable bias problem (especially if fiscal balances are impacted by partisan decisions). While the main economic indicators that we use in our annual panel analysis are available on a quarterly basis, some (notably unemployment) demonstrate a higher frequency of missing observations for quarterly data than annual data. Second, our annual data demonstrates more variation both in downgrades and negative ratings decisions, as well as electoral events than quarterly data (see Table 1).

*Model Specification and Estimator*

We employ a fixed-effects, ordinary least squares estimator for both rating actions and spreads, in order to keep the results on the two different dependent variables comparable. For rating actions, an OLS model is the equivalent to a linear probability model (LPM) as the variable is binary. While logistic regression overcomes several problems with using ordinary least squares to estimate binary outcomes, we opt for a LPM in our credit rating decision analysis for two reasons. First, when using a logistic estimator, we suffered quasi-separation problems (results either did not converge or some of our independent variables were dropped) in some of our models because several country and time dummies perfectly predicted zero outcomes (see Carter and Signorino, 2010). Second, related to separation problems, employing a fixed effects model in logistic regression would result in the dropping of panels where there are no downgrades (affecting 10 of the 23 countries in our OECD sample) or negative rating events (affecting 2 of the 23 countries within our sample). Consequently, if we employed fixed effects for both the spreads dependent variable using OLS, and the credit rating decisions using logistic regression, we would effectively be comparing two different samples.

We model credit rating decisions as follows:

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 is the binary manifestation of the two rating actions that we describe above in country i at time t. is a binary variable that indicates whether a left party controls the executive in country i at time t: it equals 1 if the prime minister is from a left party, 0 if otherwise. We also use the proportion of cabinet seats occupied by left parties and the manifesto scores of incumbent governments as alternative measurements of partisanship (we present these results in Appendix C).[[18]](#footnote-18) Executive partisanship data stems from Armingeon et al (2014), while manifesto scores are taken from the Wissenschaftszentrum Berlin für Sozialforschung’s Manifesto Project Database. For election years, Armingeon and his co-authors weight their partisanship data by the number of days an executive is in office. If a rating change happens during an election year (or election quarter), however, we code executive partisanship solely as the executive in power when the rating change happens.

 is a vector of electoral controls and their interaction with the left executive dummy which enables us to test our two electoral hypotheses; whether CRAs are more likely to impose negative ratings when a *new* left government comes into office (hypothesis 1a) and when a left incumbent is re-elected (hypothesis 1b). These controls include: an election year/quarter dummy (1 if the year/quarter is an election year, 0 if not); an interaction term between an election year/quarter and whether a left incumbent won control of the executive (1 if yes, 0 if no), and; an interaction term between an election year/quarter and whether a left non-incumbent won control of the executive (1 if yes, 0 if no). Non-election years/quarters serve as the baseline category of all of these dummies. The left incumbent and non-incumbent interaction terms highlight the *additional* (or *conditional*, if the electoral dummy is non-significant) negative rating penalty associated with an election year/quarter. Hence, the beta coefficient on the election dummy is indicative of a *right* executive electoral victory; the betas on the left incumbent/non-incumbent interaction terms indicate the *added* electoral effect for left government victories. Election data also stems from Armingeon et al (2014).

 is a vector of economic and fiscal controls that CRAs claim impacts their ratings within their methodologies for assessing countries’ solvency. These include the debt to GDP ratio, net government lending as a percentage of GDP (positive/negative values indicate an annual fiscal surplus/deficit), the trade balance (positive/negative values indicate an annual trade surplus/deficit) as a percentage of GDP, unemployment and inflation. We also include year-on-year differences debt and net lending, as CRAs are concerned not only about debt and deficit levels, but also sudden negative changes in them. To test Mosley’s “room to move” hypothesis, we control for the size of the government (government spending as a percentage of GDP).[[19]](#footnote-19) Because we were unable to obtain fiscal and government spending indicators or trade balance data from the OECD on a quarterly basis, these variables are absent from our quarterly analysis.

Multicollinearity problems within our economic controls prompted us to exclude real GDP growth as a control from our *annual* panel dataset (even though all agencies cite it as an important variable), as it shares significant correlations with our fiscal variables. However, we incorporate real GDP growth into our quarterly panel dataset, because fiscal controls were not available (moreover, we use real GDP growth in the place of unemployment in the quarterly panel, because quarterly real GDP data is complete for all our countries, while quarterly unemployment succumbs to missing data problems[[20]](#footnote-20)). We stress that including GDP growth in our models for the annual data does not alter the results of our political variables.[[21]](#footnote-21) In order to avoid multicollinearity problems between the government debt and net public lending variables, we included these controls in separate regressions (Tables 2 and 3 incorporate public debt, while Appendix B incorporates net government lending instead of debt for the same models). Debt[[22]](#footnote-22), unemployment, and trade balance data were taken from the EU’s Annual Macroeconomic (AMECO) database, while inflation, net government lending[[23]](#footnote-23) and government expenditure[[24]](#footnote-24) data were taken from the OECD.

 is a vector of government type dummies (minority, coalition majority or single-party majority, the latter serves as the baseline category). Government type data also stems from Armingeon et al (2014). Additionally, we incorporated a number of political institution controls that may influence debt ratings: capital account liberalization, a country’s central bank independence (CBI) index, and the World Bank political stability and absence of violence index (higher values indicate greater political stability). These controls could only be included in our annual results, as they are not available on a quarterly basis. We also included a control variable for whether a country is an EU member in time t, given previous scholarly findings that membership regularizes market expectations about a country’s future policy choices (Gray, 2009 and 2013), as well as whether a country was an EMU member in time t, as the creation of the Euro eliminated currency risk for countries with the common currency. In order to preserve space we do not present these results below, but they are provided in Appendix D, where we offer a fuller interpretation of their impacts on negative ratings and spreads. We discuss how the inclusion of these variables affects our results below.

 is a vector of (n-1) time dummies (in order to control for omitted time shocks that may impact ratings), as well as four path dependency dummies for CRAs: whether a CRA bestowed a downgrade upon country i in the last 3 years (1=yes, 0 if no), a worsened outlook in the last 3 years (1=yes), an improved outlook in the last 3 years (1=yes), and an upgrade in the last 3 years (1=yes). is a vector of country-specific fixed effects, which account for omitted variables that are constant over time but differ across countries (we test the sensitivity of our results to random effects in Table 3). Finally, Wooldridge statistics indicated that first-order serial correlation was present in the majority of our models. Consequently, we incorporated a panel specific AR(1) disturbance, in addition to panel corrected standard errors to correct for heteroskedasticity within panels (see Beck and Katz, 1995; Table 3 also provides results if a panel specific AR(1) disturbance is excluded, as well as if a common AR(1) disturbance is used).

In testing our auxiliary hypotheses about spreads, we use the following model specification:

 =

 is the spread of country i’s nominal interest rate on long term government bonds vis-à-vis the US’s nominal interest rate on 10 year T-bills at time t (higher values indicate a country’s long term government bond is more risky relative to the US’s). Though we use the spread *level* for our regressions below, we stress that our results are similar if we use *changes* in spreads as the dependent variable (partisanship does not impact spread changes, but negative ratings do). Interest rate data for long-term government bonds stems from the OECD.

In order to test the effects of rating changes on spreads (hypothesis 3), we include credit rating agencies’ decisions as an independent variable in modeling spreads. is vector of rating changes bestowed on country i in year t from at least one of the big three (our results are identical if we examine ratings changes of each CRA separately). This vector includes four rating decisions (all coded as binary if one or more of the big three bestow them in a given year): downgrades, worsened outlooks, improved outlooks and upgrades. All other partisan, electoral, economic, fiscal and institutional explanatory variables in our spreads models are the same as in the model for rating actions, as are the time and country fixed-effect dummies and error structures (however, the CRA path dependency dummies in vector are not included in the spreads models, because these actions are incorporated into vector ).

**Results**

Table 2 presents our results for the annual panel data for the big three’s rating decisions and for bond spreads when debt is used as the primary fiscal control. Appendix B provides our results when annual public lending is used as the primary fiscal control. The dependent variable in models I-III in Table 2 and Appendix B are downgrades for S&P, Moody’s and Fitch, while the dependent variable in models IV-VI are negative rating actions (downgrades or worsened outlooks) for S&P, Moody’s and Fitch. The dependent variable in Model VII in Table 2 and Appendix B is government bond spreads. Appendix C provides our results when the proportion of cabinet seats occupied by left parties (Models I-IV) and manifesto scores (Models V-VIII) are used to measure partisanship. Tables D.1, D.2, and D.3 in Appendix D provide our results for the rating decisions of S&P, Moody’s and Fitch respectively, when alternative political institutional controls are included (capital account openness, central bank independence, political stability, EU and EMU membership) while Table D.4 provides these results for bond spreads. Results for credit rating decisions can be interpreted in terms of changes in probabilities; i.e. if the beta coefficient is 0.03, this translates to an increased probability of a downgrade (or worsened outlook) by 3 percentage points.

**Table 2: Political and economic determinants of negative credit ratings and bond spreads (annual data)**

|  |  |  |  |
| --- | --- | --- | --- |
|  | DV is Downgrade | DV is Downgrade and Worsened Outlook | DV is Spreads |
|  | S&P | Moody's | Fitch | S&P | Moody's | Fitch | on Gov't Bonds |
|  | I | II | III | IV | V | VI | VII |
| **Partisan Controls** |   |   |   |   |   |   |   |
| Left-Controlled Executive (1=yes) | 0.071\*\*\* | 0.034\* | 0.037 | 0.003 | 0.019 | 0.006 | -0.148 |
|  | (0.009) | (0.077) | (0.200) | (0.927) | (0.353) | (0.853) | (0.218) |
| Election year (1=yes) | 0.002 | 0.01 | 0.017 | -0.034 | 0.029 | 0.051\* | 0.127\* |
|   | (0.949) | (0.538) | (0.427) | (0.277) | (0.146) | (0.079) | (0.088) |
| Election win for incumbent left  | 0.001 | -0.028 | -0.029 | 0.011 | -0.025 | -0.049 | -0.015 |
| Executive (1=yes)  | (0.977) | (0.353) | (0.432) | (0.824) | (0.409) | (0.313) | (0.926) |
| Election win for non-incumbent left  | 0.113\*\* | 0.084\*\* | 0.046 | 0.106\* | 0.058 | 0.004 | -0.213 |
| Executive (1=yes)  | (0.027) | (0.016) | (0.350) | (0.096) | (0.164) | (0.948) | (0.248) |
| **Economic and Fiscal Controls** |   |   |   |   |   |   |   |
| Public Debt (% of GDP) | -0.001 | 0.001 | -0.002\*\* | 0.001 | 0.001 | 0.001 | 0.012\*\* |
|   | (0.635) | (0.513) | (0.012) | (0.439) | (0.381) | (0.246) | (0.021) |
| Trade Balance (% of GDP) | -0.018\*\*\* | -0.002 | 0.003 | -0.009\* | 0.003 | -0.009 | -0.045\* |
|   | (0.000) | (0.559) | (0.589) | (0.069) | (0.438) | (0.242) | (0.084) |
| Unemployment | 0.025\* | 0.010 | 0.001 | 0.010 | 0.013\*\* | 0.000 | 0.338\*\*\* |
|   | (0.077) | (0.242) | (0.973) | (0.417) | (0.039) | (0.993) | (0.000) |
| Δ Public Debt (% of GDP) | 0.010\*\*\* | 0.008\*\*\* | 0.004 | 0.008\*\* | 0.007\*\*\* | 0.006\* | -0.027\* |
|   | (0.003) | (0.001) | (0.220) | (0.024) | (0.005) | (0.085) | (0.095) |
| Inflation  | 0.007 | 0.031\*\*\* | -0.003 | 0.006 | 0.045\*\*\* | 0.009 | 0.238\*\*\* |
|   | (0.663) | (0.007) | (0.844) | (0.705) | (0.000) | (0.670) | (0.000) |
| Government size (% of GDP)  | 0.013\*\* | 0.019\*\*\* | 0.025\*\*\* | 0.015\*\* | 0.016\*\*\* | 0.031\*\*\* | 0.009 |
|  | (0.038) | (0.000) | (0.000) | (0.011) | (0.001) | (0.000) | (0.730) |
| **Government Type** |   |   |   |   |   |   |   |
| Minority Government | -0.105\*\* | 0.027 | 0.124\*\* | -0.158\*\*\* | 0.016 | 0.026 | 0.516\*\* |
|   | (0.030) | (0.552) | (0.029) | (0.004) | (0.761) | (0.740) | (0.021) |
| Coalition Majority | 0.003 | 0.085\*\* | 0.108\*\*\* | -0.112\*\* | 0.05 | 0.011 | 0.606\*\* |
|   | (0.944) | (0.017) | (0.009) | (0.043) | (0.228) | (0.830) | (0.019) |
| **Credit rating controls (for spreads only)** |   |   |   |   |   |   |   |
| Downgrade by big three (1=yes)  |  |  |  |  |  |  | 0.494\*\* |
|  |  |  |  |  |  |  | (0.019) |
| Worsened outlook by big three (1=yes)  |  |  |  |  |  |  | -0.076 |
|  |  |  |  |  |  |  | (0.619) |
| Improved outlook by big three (1=yes)  |  |  |  |  |  |  | -0.277\* |
|  |  |  |  |  |  |  | (0.060) |
| Upgrade by big three (1=yes)  |  |  |  |  |  |  | 0.056 |
|  |  |  |  |  |  |  | (0.629) |
| N | 424 | 424 | 417 | 424 | 424 | 417 | 399 |
| R-squared | 0.405 | 0.483 | 0.409 | 0.509 | 0.525 | 0.442 | 0.648 |
| Chi-squared (p-value) | 0 | 0 | 0 | 0 | 0 | 0 | 0 |

Estimators used were linear probability models (for ratings changes), and OLS (for spreads), for 23 OECD countries from 1995-2014. All models include a panel-specific AR1 disturbance. N-1 country, time dummies, path dependency dummies (for credit rating actions regressions) and constant term included but not shown. P-values are in parenthesis (panel corrected standard errors used). \*, \*\*, and \*\*\* indicate significance on a 90%, 95% and 99% confidence level.

*Government partisanship:*

The results for our left executive dummy in Models I-III in Table 2 and Appendix A indicate that two of the big three CRAs (S&P and Moody’s) were more likely to downgrade left-controlled executives. Left governments were, 7.1% and 3.4% more likely to witness downgrades from S&P’s and Moody’s respectively than their right counterparts. The magnitude of these downgrade penalties is slightly larger for S&P in Appendix B when net government lending is used as a control, but becomes insignificant for Moody’s. Executive partisanship’s impact on downgrades also holds when using alternative measures for left partisanship, examined in Appendix C. *For all three* CRAs, the higher proportion of cabinet seats occupied by left parties, the more likely a downgrade (cabinets that are solely occupied by left parties exhibit downgrade likelihoods of 10% from all three CRAs). For Moody’s and Fitch, higher manifesto scores (indicative of more *right-wing* executives) are associated with lower probabilities of downgrades. A one standard deviation increase in an incumbent’s manifesto score (16.489) is associated with a 4.9% and 3.3% lower likelihood of a downgrade from Moody’s and Fitch, respectively.

The effect of government partisanship is very robust to different model specifications for S&P. Table 3 provides the beta coefficients, and their associated significance levels, for the left-controlled executive dummy (and the non-incumbent left election victory dummy, which we discuss below) for the following alternative specifications of our models in Table 2: 1.) if random effects are used; 2.) if Greece is excluded (downgrade and worsened outlooks were relatively numerous under its left executives from 2009 onwards); 3.) if the US is excluded (we do this because our spreads regressions omit the US, given that its bond yield is used as the baseline for spreads); 4.) if our economic variables are lagged; 5.) if our time period spans 1999-2008 (excluding the European debt crisis); 6.) if the panel-specific AR(1) disturbance is excluded, and; 7.) if a common AR(1) disturbance is used. S&P’s (significant) left-executive downgrade penalty is robust to all re-specifications in Table 3. Moreover, S&P’s results for downgrade premiums associated with left executives is robust when central bank independence, capital account liberalization, political stability, EU and EMU membership are controlled for (see Appendix D.1) and when we used quarterly data (see Table 4 below).

The effect of government partisanship is not as robust for Moody’s. In Table 3, partisanship holds its significance and sign for Moody’s if the US is excluded, and if a PSAR(1) or common AR(1) error term is used. However, Moody’s left executive downgrade penalties lack significance for the other specifications in Table 3 and if other political controls are included in our models (see Models I-V in Appendix Table D.2). Intriguingly, left executives are *less* likely to witness downgrades from Moody’s between 1995-2008, indicating that this CRA’s left executive bias in downgrade decisions is moved solely by the inclusion of the years of the debt crisis within our models. We provide an explanation for this finding in our conclusion below.

**Table 3: Robustness checks for partisanship beta coefficients**

|  |  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- | --- |
|  | **Original Model** | **Random Effects** | **No Greece** | **No US** | **Economic variables lagged** | **1995-2008** | **No PSAR1** | **Common AR1** |
| **STANDARD AND POOR'S** |   |   |   |   |   |   |   |   |
| *DV=Downgrade* |   |   |   |   |   |   |   |   |
| Left gov (1=yes) | 0.071\*\*\* | 0.046\* | 0.072\*\*\* | 0.064\*\* | 0.079\*\* | 0.041\* | 0.084\*\*\* | 0.085\*\*\* |
| New left victory | 0.113\*\* | 0.106\*\* | 0.116\*\* | 0.126\*\* | 0.155\*\*\* | 0.144\*\*\* | 0.112\* | 0.100\* |
| *DV=Downgrade and worsened outlook* |   |   |   |   |   |   |   |   |
| Left gov (1=yes) | 0.003 | -0.021 | -0.004 | 0.004 | 0.021 | -0.007 | 0.013 | 0.012 |
| New left victory | 0.106\* | 0.115\* | 0.111\* | 0.115\* | 0.125\* | 0.153\*\*\* | 0.139\* | 0.122\* |
| **MOODY'S** |   |   |   |   |   |   |   |   |
| *DV=Downgrade*  |   |   |   |   |   |   |   |   |
| Left gov (1=yes) | 0.034\* | 0.023 | 0.027 | 0.037\* | 0.026 | -0.032\*\*\* | 0.033\* | 0.038\* |
| New left victory | 0.084\*\* | 0.045 | 0.065\* | 0.087\*\* | 0.085\*\* | 0.015 | 0.057 | 0.049 |
| *DV=Downgrade and worsened outlook* |   |   |   |   |   |   |   |   |
| Left gov (1=yes) | 0.019 | 0.006 | 0.015 | 0.026 | 0.013 | -0.035\*\*\* | 0.031 | 0.031 |
| New left victory | 0.058 | 0.035 | 0.029 | 0.070\* | 0.069 | -0.018 | 0.001 | -0.0005 |
| **FITCH** |   |   |   |   |   |   |   |   |
| *DV=Downgrade*  |   |   |   |   |   |   |   |   |
| Left gov (1=yes) | 0.037 | 0.023 | 0.036 | 0.044 | 0.045 | -0.011 | 0.039 | 0.033 |
| New left victory | 0.046 | 0.036 | 0.045 | 0.043 | 0.043 | 0.075\*\*\* | 0.027 | 0.016 |
| *DV=Downgrade and worsened outlook* |   |   |   |   |   |   |   |   |
| Left gov (1=yes) | 0.006 | 0.002 | 0.027 | -0.009 | 0.012 | -0.012 | 0.032 | 0.023 |
| New left victory | 0.004 | -0.002 | 0.019 | 0.024 | 0.012 | 0.081 | 0.018 | 0.005 |
| **SPREADS** |   |   |   |   |   |   |   |   |
| Left gov (1=yes) | -0.148 | -0.144 | -0.087 | NA | -0.228\* | 0.044 | 0.032 | -0.146 |
| New left victory | -0.213 | -0.204 | -0.019 |   | -0.42\* | 0.237\*\* | -0.283 | -0.171 |

Note: Model used for the downgrade dependent variables are Models I-III in Table 2. Model used for the downgrade and worsened outlook dependent variables are Models IV-VI in Table 2. Only the beta coefficient is shown. \*, \*\* and \*\*\* denote significance at a 90%, 95% and 99% confidence level.

In contrast to ratings, spreads on government bonds are not impacted by the partisanship of the executive (the left executive dummy is non-significant for spreads in Table 2, Appendix B, C and all model specifications in Appendix Table D.4). We only find partisanship effects under two specifications (when net government lending is used as a fiscal control, see Appendix B, or when all economic variables are lagged, see fifth column in Table 3). However, in these cases, left executives are associated with *lower* bond spreads than right executives.

*Elections and partisanship*

Elections, on their own, do not affect the likelihood of downgrades or negative ratings relative to non-election years, as indicated by the insignificant election year dummy for all CRAs. Election *results,* on the other hand, had significant effects. The re-election of *incumbent* left governments have no effect on the likelihood of negative rating action by any of the big three, as indicated by the insignificance of the left incumbent election win dummy. However, S&P and Moody’s *do* mind left *non-incumbent* electoral victories in their downgrade and negative rating decision, while Fitch only demonstrates a significantly greater likelihood of bestowing downgrades after non-incumbent left electoral victories in the *pre-crisis 2000s* (see results in Table 3 for the 1999-2008 panel). We explain this finding in our conclusion. In other words, the election of right executives (incorporated into the election dummy) and left incumbents exhibit no influence on adverse rating changes. CRAs only appear critical of the election of new left executives.

S&P is *11.3% more likely* to issue a downgrade, and *10.6% more likely* to initiate negative action (downgrade or worsened outlook) if a new left government wins an election*:* this penalty is *in addition to* the 7.1% downgrade penalty applied to left executives in general (see Model I and IV, Table 2). These S&P ratings penalties for non-incumbent left electoral victories are highly robust. With the exception of Model IV in Appendix D.1, the ratings penalty (both for downgrades on their own, and negative ratings more broadly) that S&P associates with non-incumbent left electoral victories remains significant for all alternative model specifications in Table 3, as well as all model specifications in Appendix B, C, and D. The ratings penalty that Moody’s bestows upon non-incumbent left electoral victories is not as extreme or robust as S&P’s reactions. Moody’s is *8.4% more likely* to issue a downgrade upon the election of non-incumbent left executives but these electoral victories do not significantly impact negative ratings more broadly (see models II and V in Table 2.). Moody’s downgrade electoral penalty remains significant if Greece and the US are excluded from the sample or if lagged economic variables are used as controls (see Table 3), if net lending rather than public debt is used as the primary fiscal control (see Appendix B), using alternative measures of left partisanship (Appendix C), and if capital account openness, central bank independence and political stability[[25]](#footnote-25) are included in our models. Outside of these specifications, however, the downgrade penalty Moody’s bestows on non-incumbent left electoral victories is non-significant.

S&P’s and Moody’s downgrade and negative ratings penalties are also present when our regressions are run with quarterly data (results presented in Table 4). As mentioned above, we were limited in terms of the types of controls we could include in our quarterly models, as all of our fiscal and government expenditure data was unavailable on a quarterly basis (similarly, we also omitted unemployment as a control in our quarterly models, due to widespread missing observations for some of our countries, but we included real GDP growth, for which every country had complete data). We structure the models for the quarterly data in exactly the same fashion as our annual data, with one exception – we provide a series of four quarterly lags for our election variables, because elections may not have immediate results on ratings changes (due to the fact that some countries may have transitional or coalition-formation periods between an election and when its victor formally enters office, and/or that CRAs may take some time to observe a newly elected left executive once it has entered office). Lagged non-incumbent left party electoral victories had a significant impact on downgrades for Moody’s (Models II in Table 4; significance was just under the 90% confidence level for S&P, p-value=0.101) and negative ratings more broadly (Models IV-VI) for both S&P and Moody’s. More specifically, electoral victories of new left governments produce a higher probability of a downgrade (8.6% for Moody’s), and a higher probability of a negative ratings change (11.4% more likely for S&P and 9.8% for Moody’s) *two quarters* after they occur (we highlight this result in bold in Table 4 below). In other words, ratings penalties associated with new left government victories do not emerge immediately (as indicated by the insignificant present value of this variable), but rather emerge in the six months afterwards. Likewise, Fitch demonstrates an immediate worsened outlook ratings penalty for a non-incumbent left executive electoral victory in the quarterly data; Fitch is 8.8% more likely to bestow a negative ratings change in the same quarter that new left executives win an election (see Model VI, Table 4).

**Table 4: Political and economic determinants of negative credit ratings and bond spreads (quarterly data)**

|  |  |  |  |
| --- | --- | --- | --- |
|  | DV is Downgrade | DV is Downgrade and Worsened Outlook | DV is Spreads |
|  | S&P | Moody's | Fitch | S&P | Moody's | Fitch | on Gov't Bonds |
|  | I | II | III | IV | V | VI | VII |
| **Partisan Controls** |   |   |   |   |   |   |   |
| **Left-Controlled Executive (1=yes)** | **0.027\*\*** | 0.003 | 0.004 | -0.004 | 0.002 | -0.009 | -0.137\* |
|  | **(0.031)** | (0.777) | (0.776) | (0.830) | (0.854) | (0.597) | (0.075) |
| Election quarter (1=yes) | 0.000 | 0.021 | 0.014 | -0.018 | -0.007 | 0.007 | 0.151\*\* |
|   | (0.995) | (0.153) | (0.379) | (0.414) | (0.700) | (0.705) | (0.012) |
| Election quarter (t-1) | 0.018 | 0.030\* | -0.004 | 0.018 | -0.002 | 0.003 | 0.189\*\* |
|   | (0.360) | (0.071) | (0.820) | (0.457) | (0.931) | (0.904) | (0.011) |
| Election quarter (t-2) | 0.02 | 0.015 | 0.000 | 0.029 | -0.017 | 0.008 | 0.134\* |
|   | (0.309) | (0.377) | (0.986) | (0.225) | (0.369) | (0.723) | (0.084) |
| Election quarter (t-3) | 0.019 | 0.001 | -0.016 | 0.019 | -0.012 | -0.006 | 0.032 |
|   | (0.313) | (0.970) | (0.377) | (0.421) | (0.542) | (0.790) | (0.662) |
| Election quarter (t-4) | 0.009 | 0.001 | -0.025 | -0.007 | -0.008 | -0.011 | 0.015 |
|   | (0.584) | (0.958) | (0.118) | (0.756) | (0.663) | (0.586) | (0.807) |
| Left incumbent election win  | -0.016 | -0.017 | -0.02 | 0.023 | -0.003 | 0.013 | -0.132 |
|  (1=yes) | (0.589) | (0.534) | (0.530) | (0.552) | (0.934) | (0.744) | (0.270) |
| Left incumbent election win  | -0.035 | -0.029 | -0.002 | 0.018 | 0.029 | -0.013 | -0.133 |
|  (t-1) | (0.284) | (0.324) | (0.959) | (0.678) | (0.396) | (0.758) | (0.346) |
| Left incumbent election win  | -0.028 | -0.024 | 0.022 | -0.01 | -0.003 | 0.007 | -0.153 |
|  (t-2) | (0.399) | (0.420) | (0.540) | (0.810) | (0.935) | (0.877) | (0.330) |
| Left incumbent election win | -0.001 | -0.001 | 0.000 | 0.015 | 0.026 | -0.004 | -0.086 |
|  (t-3) | (0.976) | (0.979) | (0.991) | (0.727) | (0.448) | (0.927) | (0.559) |
| Left incumbent election win  | 0.017 | 0.025 | 0.006 | 0.038 | 0.019 | 0.021 | -0.116 |
|  (t-4) | (0.570) | (0.333) | (0.860) | (0.337) | (0.576) | (0.583) | (0.382) |
| **Left non-incumbent election win**  | 0.013 | -0.017 | 0.053 | 0.017 | 0.007 | **0.088\*** | -0.187 |
|  **(1=yes)** | (0.732) | (0.598) | (0.195) | (0.743) | (0.870) | **(0.081)** | (0.159) |
| Left non-incumbent election win  | 0.004 | -0.044 | -0.015 | 0.018 | -0.037 | 0.017 | -0.157 |
|   (t-1) | (0.923) | (0.208) | (0.742) | (0.735) | (0.406) | (0.755) | (0.334) |
| **Left non-incumbent election win**  | **0.071** | **0.086\*\*** | **0.057** | **0.114\*\*** | **0.098\*\*** | **0.033** | **-0.195** |
| **(t-2)** | **(0.101)** | **(0.018)** | **(0.219)** | **(0.041)** | **(0.033)** | **(0.556)** | **(0.253)** |
| Left non-incumbent election win  | -0.052 | -0.028 | 0.023 | -0.038 | -0.044 | 0.016 | 0.048 |
|  (t-3) | (0.226) | (0.438) | (0.625) | (0.490) | (0.340) | (0.776) | (0.765) |
| Left non-incumbent election win  | -0.063 | -0.049 | -0.018 | -0.015 | -0.007 | 0.046 | 0.023 |
|  (t-4) | (0.111) | (0.144) | (0.667) | (0.769) | (0.864) | (0.371) | (0.861) |
| **Economic Controls** |   |   |   |   |   |   |   |
| Inflation  | -0.002 | 0.009\*\* | -0.001 | 0.007 | 0.006 | 0.000 | 0.052\*\* |
|   | (0.772) | (0.039) | (0.855) | (0.321) | (0.258) | (0.966) | (0.038) |
| Real GDP Growth | -0.011\*\*\* | -0.010\*\*\* | -0.008\*\*\* | -0.008\*\* | -0.013\*\*\* | -0.009\*\*\* | -0.037\*\*\* |
|   | (0.000) | (0.000) | (0.003) | (0.028) | (0.000) | (0.008) | (0.003) |
| **Government Type** |   |   |   |   |   |   |   |
| Minority Government | -0.025 | -0.006 | 0.028 | -0.045 | -0.009 | 0.003 | -0.254\*\* |
|   | (0.294) | (0.780) | (0.377) | (0.124) | (0.735) | (0.938) | (0.035) |
| Coalition Majority | -0.01 | -0.012 | 0.008 | -0.009 | -0.002 | -0.004 | -0.029 |
|   | (0.689) | (0.533) | (0.707) | (0.741) | (0.917) | (0.881) | (0.805) |
| **Credit Rating Actions (for spreads)** |   |   |   |   |   |   |   |
| Downgrade by big three |   |   |   |   |   |   | 0.203\*\* |
|   |   |   |   |   |   |   | (0.042) |
| Worsened Outlook by big three |   |   |   |   |   |   | 0.152\*\* |
|   |   |   |   |   |   |   | (0.033) |
| Improved Outlook by big three |   |   |   |   |   |   | 0.006 |
|   |   |   |   |   |   |   | (0.932) |
| Upgrade by big three |   |   |   |   |   |   | 0.005 |
|   |   |   |   |   |   |   | (0.962) |
| N | 1732 | 1732 | 1700 | 1732 | 1732 | 1700 | 1650 |
| R-squared | 0.201 | 0.27 | 0.161 | 0.197 | 0.229 | 0.168 | 0.401 |
| Chi-squared (p-value) | 0 | 0 | 0 | 0 | 0 | 0 | 0 |

Estimators used were linear probability models (for ratings changes), and OLS (for spreads), for quarterly data for 23 OECD countries from 1995-2014. All models include a panel-specific AR1 disturbance. N-1 country, time dummies, path dependency dummies (for credit rating action regressions only) and constant term included but not shown. P-values are in parenthesis (panel corrected standard errors used). \*, \*\*, and \*\*\* indicate significance on a 90%, 95% and 99% confidence level. t-1, t-2, t-3, and t-4 indicate a one, two, three and four quarterly lag, respectively.

Just as in the case of government partisanship, we find that elections affect bond spreads differently from ratings. Elections do matter for bond spreads, but *not their outcomes*. Results in Table 2, Appendix B, Appendix C, Appendix D and Table 4 demonstrate that election years are associated with increased bond spreads (Table 4, Model VI indicates that these effects are spread over three quarters). However, the electoral victories of left incumbent governments or new left governments *do not* have a significant impact on bond yields (when they do, as in Tables 3 and Appendix D.4, left electoral victories are associated with smaller, rather than larger, spreads). The lack of robust effects for electoral outcomes suggests that markets display greater partisan neutrality than CRAs during elections: while markets appear to price elections into bond yield spreads for our countries, they do not price left electoral victories (either incumbents or non-incumbents) into them.

In regards to other economic and institutional controls, downgrades and worsened outlooks were more likely by the big three during periods of high sudden increases in public debt (Δ debt), high levels of unemployment (for S&P and Moody’s), higher trade deficits (for S&P), higher sudden increases in government deficits (for Fitch), higher inflation (for Moody’s) and higher levels of government expenditure as a percentage of GDP (for all of the big three). Compared to single-party majority governments, minority governments were less likely to encounter downgrades or negative ratings from S&P, but were more likely to encounter these decisions from Fitch, while coalition majority governments were more likely to encounter downgrades from Moody’s and Fitch (but were less likely to encounter negative ratings from S&P). Results for bond yields align with some of these results. Bond spreads are significantly larger for higher levels public debt, unemployment, inflation and trade deficits. Spreads also increase under minority and coalition majority governments (compared to single-party majority administrations). Importantly, spreads increase when downgrades are initiated by the big three. Moreover, the positive impact of downgrades on spreads remains significant for all of our specifications in Tables 2, 3, 4, and Appendix B, C and D.4.

In sum, while CRAs and the markets are aligned in their responses to variation in economic, fiscal and institutional variables, their approach to politics deviates strongly. S&P and Moody’s are more likely to initiate negative rating actions against left governments and even more so when a new left executive comes into office. Markets, on the other hand, are moved by the uncertainty created by elections, but do not *directly* respond to variation in government partisanship. This latter result confirms Breen and McMenamin’s (2013) findings that partisanship does not systematically impact bond yields over time. More importantly, because spreads respond to downgrades, the tendency of CRAs to downgrade left governments more often than right ones indirectly feeds partisanship into market outcomes.

**Discussion and conclusion**

Our results point to an intriguing dynamic in the market-politics nexus. CRAs have systematically discriminated against left governments in the past two decades. S&P, the agency with the most influential and largest number of sovereign ratings (IMF, 2010, pp 87 and 115), has been consistently more likely to downgrade left executives than their right-wing counterparts, and the probability of negative rating actions increased even further when non-incumbent left executives first enter office. Moody’s and Fitch have also penalized left governments, but not as consistently as S&P. While the likelihood of downgrades was correlated with the ideological composition of the executive (and the partisanship of the prime minister) for Moody’s, the partisan stance of the prime minister lacks importance for Fitch. Both Moody’s and Fitch applied downgrade penalties against newly elected non-incumbent left governments, but Fitch only did so in the years preceding the crisis (see Table 3). We found that partisan credit rating cues were incorporated into the pricing of sovereign debt, as markets systematically acted upon downgrades, even though bond prices otherwise did not betray partisan bias. This suggests that CRAs instigate systematic partisan discrimination in sovereign credit markets.

Since CRAs do not acknowledge or justify the influence of partisan considerations on rating decisions in their communications, we hypothesize that discrimination against left governments is a strategy that conservative CRAs can use to minimize downside policy and market risk while keeping ratings stable, in an effort to make them attractive as third-party risk indicators. The pattern of variation in partisan discrimination across CRAs and across time lends some support to this hypothesis. The finding that S&P discriminates most systematically against the left is consistent with S&P’s track record as the most cautious in its risk assessment of the big three – leading the other two 74 percent of the time in downgrading developed economies between 2005 and 2010 (IMF, 2010, p115) – as well as with its professed strategy of preemptively incorporating all potential future risks into ratings. Moody’s less stringent partisan discrimination before the crisis is consistent with a less conservative attitude, reflected in its role as a follower to the other two CRAs in downgrades.[[26]](#footnote-26) The appearance of the left electoral penalty in Moody’s downgrades to new left governments *only* with the inclusion of the years of the crisis likely betrays the adoption of a more conservative approach as a reaction to rating failures during the crisis. In a similar vein, but with opposite effect, Fitch’s abandonment of discrimination against electoral victories of new left governments after 2008 can also be attributed to its effort to regain credibility after the crisis. In 2011, Fitch transformed its sovereign rating methodology, putting it onto a predominantly quantitative footing and making it fully transparent by publishing the models that form the basis of its rating reports[[27]](#footnote-27), which left less room for the incorporation of intangible factors like partisanship.

Our findings about partisan discrimination in credit ratings have several important implications for the role of politics in sovereign debt markets. First, the presence of systematically more adverse market conditions for the left suggest that the “golden straightjacket” exists: markets selectively penalize certain developments in domestic politics. Second, and more surprisingly, the “golden straightjacket” effect does not seem to originate from investors, but from credit rating agencies that provide partisan-discriminated signals to investors, which are then translated into changes in spreads on government debt. Third, the use of credit ratings as third-party risk indicators by institutional investors and public regulators proves to be problematic from more than one perspective. The anomaly of placing unappointed, unelected, unsupervised private actors at the heart of prudential frameworks has already been pointed out (e.g. Sinclair 2008). However, our findings suggest that CRAs efforts to maintain this position motivates them to produce systematically distorted assessments of creditworthiness that generate adverse conditions for governments on the left of the ideological spectrum.

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2. Associate Professor, Department of Political Science and School of Public Policy, Oregon State University. Alison.Johnston@oregonstate.edu [↑](#footnote-ref-2)
3. While there are about 150 credit rating agencies worldwide, Standard & Poor’s, Moody’s and Fitch capture 95 percent of the market. (De Haan and Amtenbrink 2011). [↑](#footnote-ref-3)
4. We justify this time period below. [↑](#footnote-ref-4)
5. #  “Moody’s Gets No Respect as Bonds Shun 56% of Country Ratings” (Bloomberg, December 17, 2012) <https://www.bloomberg.com/news/articles/2012-12-16/moody-s-getting-no-respect-as-bonds-shun-56-of-country-ratings>

 [↑](#footnote-ref-5)
6. We thank one of our anonymous reviewers for calling our attention to this point. [↑](#footnote-ref-6)
7. Sovereign rating methodologies are published to explain the cognitive processes underlying the assessment of sovereign creditworthiness and describe the nature and the source of information that forms the basis of that assessment. They are regularly updated, continuously improved and vetted by experts and investors. [↑](#footnote-ref-7)
8. Fitch, 2002, 2011, 2012 and 2014; Moody’s, 2008 and 2013; Standard and Poor’s, 2006, 2008, 2011 and 2013 [↑](#footnote-ref-8)
9. Although the methodologies emphasize the importance of political factors, they focus on structural, lasting features of politics, such as political institutions and social polarization. In terms of short-term political risks, methodologies only mention the possibility of severe forms of civil conflict associated with critically weakened state control. [↑](#footnote-ref-9)
10. For a number of methodological reasons, our primary analysis focuses on *rating actions* rather than absolute *rating levels*, but we check our results using rating levels as well. [↑](#footnote-ref-10)
11. We use US Treasury securities as our spread benchmark, because US treasury bills are deemed by markets as one of the most secure investments from default and serve as a safe-haven during flights to quality (Hall, 2012: pg 14). This implies that the US is excluded from our spreads analysis. However, our results for rating actions remain unchanged when we exclude the US from the sample, as we highlight in Table 3. [↑](#footnote-ref-11)
12. Afonso et al’s (2007) study, for example, works with a 130 country sample, where between only 10% and 20% ratings scores are AAA for each of the big three, while over a quarter of sovereigns had speculative-grade ratings (pg 53). [↑](#footnote-ref-12)
13. Our results are inconclusive for the full sample when using standard OLS, but when our sample is restricted to sub-AAA country-years to address ceiling effects, we find similar ratings penalties upon the election of non-incumbent left governments for S&P and Moody’s as when we used rating actions as our dependent variable. Moreover, when we employ a tobit model to address ceiling effects, we find ratings penalties for left incumbents from Fitch. [↑](#footnote-ref-13)
14. Accessed June 18, 2015 at [www.moodys.com](https://www.moodys.com/researchandratings/viewall/sovereign-supranational/sovereign/005005001/4294966288%204294963861/4294966848/0/0/-/0/-/-/-/-1/-/-/-/en/emea/pdf/-/rra) [↑](#footnote-ref-14)
15. Accessed June 17, 2015 at <https://www.fitchratings.com/web.../ratings/sovereign_ratings_history.xls> [↑](#footnote-ref-15)
16. Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Iceland, Ireland, Italy, Japan, Luxembourg, the Netherlands, New Zealand, Norway, Portugal, Spain, Switzerland, Sweden, the UK and the US. [↑](#footnote-ref-16)
17. While we examined our panel with monthly data between these years, the use of this smaller time unit saturated our panels with zeros both for rating actions and election variables. [↑](#footnote-ref-17)
18. Our results for S&P and Moody’s do not change if we use the proportion of cabinet seats occupied by left parties instead of a dummy indicating whether a left party holds the executive, while for Fitch, left cabinet seats becomes significantly associated with downgrades. Our results for left executive partisanship are similar in sign and significance for Moody’s and further become significant for Fitch when we use manifesto scores instead of a left executive dummy (negative manifesto score values indicate an incumbent is more left wing, positive values indicate an incumbent is more rightwing). [↑](#footnote-ref-18)
19. We also conducted regressions controlling for social benefit expenditure (as a percentage of GDP – data stemmed from the OECD). We did not include this variable and government expenditure in the same model given its strong correlation with government expenditure. Our partisan results are roughly similar to those when government expenditure is used as a control. We present these results in an online appendix. [↑](#footnote-ref-19)
20. Iceland lacks quarterly unemployment data until 2003, while Switzerland lacks (complete) quarterly unemployment data until 2010. [↑](#footnote-ref-20)
21. We present these results in an online appendix. [↑](#footnote-ref-21)
22. Canada’s (more complete) debt time series data was taken from the OECD. [↑](#footnote-ref-22)
23. France’s (more complete) net government lending data was taken from AMECO. [↑](#footnote-ref-23)
24. Germany’s and Japan’s more complete government spending data was taken from AMECO. [↑](#footnote-ref-24)
25. Moody’s exhibits a significant positive likelihood of negative ratings (downgrades and worsened outlooks) after the election victory of new left executives if political stability is controlled for (see Model VIII in Appendix C.3) [↑](#footnote-ref-25)
26. Between 2005 and 2010, Moody’s trailed one of the other CRAs in its downgrades to developed sovereigns 96 percent of the time (IMF, 2010: p115). [↑](#footnote-ref-26)
27. “Fitch publishes sovereign rating model” (October 17, 2011) <https://www.fitchratings.ru/rws/press-release.html?report_id=730901> [↑](#footnote-ref-27)