THE MASS POLITICS OF FINANCIAL MARKETS

A Dissertation

by

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ABSTRACT

Research on the role of the democratic process—and particularly that of mass political opinion—in financial markets is scarce. Extant published research that either principally or tangentially addresses this subject matter commonly focuses on elections as sources of political risk in financial markets. In this work, I consider the relationship between political dynamics and financial market outcomes beyond elections. I focus on the connection between mass political opinion and treasury rates as well as the connection between mass political opinion and stock market volatility. In both cases, I propose theories that are conditional on government ideology and institutional factors. In a methodological chapter, I explore an empirical challenge I came across when developing an empirical strategy in the previous chapters: modeling conditional theories with error correction models. Ultimately, four substantive questions are answered in this study: Does mass political opinion influence financial markets beyond elections? Do these effects hold across different categories of investment assets? Under what conditions does mass political opinion influence markets? Why and when do markets respond to mass political opinion? Broadly, the work presented here contributes to an understanding of the intricate relationship between political risk and financial markets.

DEDICATION

To my parents: José Flávio and Maria Bernadette.

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1. INTRODUCTION

1.1 Background

Research on the role of the democratic process—and particularly that of mass political opinion—in financial markets is scarce. Extant published research that either principally or tangentially addresses this subject matter commonly focuses on elections as sources of political risk in financial markets (Roberts, 1990; Herron et al., 1999; Herron, 2000; Leblang and Mukherjee, 2005; Jensen and Schmith, 2005; Fowler, 2006; Füss and Bechtel, 2008; Goodell and Bodey, 2012; Sattler, 2013). However, studies have not identified other phenomena, such as incumbent approval, as sources of political risk.

The claim that political and economic elites respond to (and perhaps *fear*) mass political opinion is not new (Hibbs, 1977; Persson and Tabellini, 2012). Markets likely perceive mass political opinion as crucial for policymakers' political capital and standing. It influences the kinds of economic and regulatory policies the latter are likely to pursue and whether they are likely to succeed in enacting their agenda. Importantly, markets recognize that mass preferences exert pressure on incumbent policymakers to implement policies that might benefit or harm the interests of capital owners (Alesina and Perotti, 1996; Dalio et al., 2017). The underlying rationale is that changes in regulatory codes and enforcement priorities can substantially impact the business climate and, subsequently, returns on investment.

The absence of research discussing non-electoral political phenomena as political risk is likely due to a few different reasons. First, the implications of political risk are most prominent outside of political science (e.g., economics, finance), even though political scientists are likely best suited to address questions about politics. Second, financial market researchers tend to oversimplify and underspecify politics and political risk in order to parse them out of theoretical and statistical models, rather than fully explore the role they play in markets. Third, most works on the role of politics in financial markets overwhelmingly focus on developed countries where non-electoral political risk is assumed to play a diminished role in markets.

1.2 Research Questions and Chapter Outline

Four principal questions are asked in this dissertation: Does mass political opinion influence financial markets beyond elections? Do these effects hold across different categories of investment assets? Under what conditions does mass political opinion influence markets? Why and when do markets respond to mass political opinion?

To answer these questions, I propose theories that integrate literatures from multiple disciplines, namely political science, economics, and finance. I build on foundational theories including work on the political economy of debt and equity markets (Mosley, 2003; Leblang and Mukherjee, 2005; Fowler, 2006; Füss and Bechtel, 2008; Bechtel, 2009; Goodell and Bodey, 2012; Breen and McMenamin, 2013), the economic and financial implications of democracy and electoral competition (Saiegh, 2005; Ballard-Rosa, Mosley and Wellhausen, 2019), mass political opinion and policymaking (Risse-Kappen, 1991; Wlezien, 1995; Soroka and Wlezien, 2005; Jennings and John, 2009; Lax and Phillips, 2012; Wlezien and Soroka, 2016), institutional and ideological constraints on policymaking (Shepsle, 1979; Hibbs, 1989; North, 1991; Powell and Whitten, 1993; Tsebelis, 1995, 2002, 2011), the structural dependence of the state on capital (Przeworski and Wallerstein, 1988), and market efficiency (Fama, 1970, 1995, 2021).

In the chapter that follows, "The Mass Politics of Debt Markets," I argue that mass approval of the chief executive influences investor behavior and, ultimately, open-auction rates on treasury bills. It is noteworthy that the theory I put forth complements, what I label, the *good-payer* story about credible commitment of repayment with a *steady-hand* story about the role of economic policymaking in debt markets. According to the *steady-hand* story, market investors use information about mass political opinion to evaluate policymaking incentives and forecast whether incumbents are likely to pursue policy changes. I also consider the role of government ideology and institutional constraints in moderating the relationship between mass political opinion and treasury rates. Given that higher popular approval is associated with increased political capital, my theory leads me to expect that in some contexts an increase in incumbent approval is associated with market-unfriendly policies.[1] Hence, investors penalize a country's sovereign debt with higher rates at open auction. Conversely, my theory leads me to expect that in other contexts an increase in incumbent approval is associated with market-friendly policies. Hence, investors reward a country's sovereign debt with lower rates at open auction.

In the third chapter, "The Mass Politics of Stock Market Volatility," I focus on the effect of government approval on stock market volatility. I argue that, as with debt markets, the effect of politics and policymaking on stock markets is unlikely to be limited to elections given that uncertainty about a policymaker's choices remains throughout her time in office. I further argue that the effect of mass political opinion on stock market volatility is likely to be stronger in contexts of relatively less developed financial market institutions. I also consider the moderating role of government ideology. Namely, I argue that an increase in approval of a left-wing chief executive—which markets are predisposed to disliking—signifies less desirable economic policies. Investors,

[1]Following Putnam (1993), Booth and Richard (1998), Sørensen and Torfing (2003) Nee and Opper (2010), Thrower (2017), I define political capital as the goodwill, prestige, and the influence a political actor enjoys in her interactions with other political actors.

therefore, depress stock trading and speculative behavior leading to lower stock market volatility. On the other hand, an increase in approval of a right-wing chief executive—which markets are predisposed to preferring—signifies more desirable economic policies. Investors, therefore, raise stock trading and speculative behavior leading to higher stock market volatility.

In the fourth chapter, I address a methodological issue I came across when developing an empirical strategy in the previous chapters: modeling conditional theories with error correction models. Here, I explore this infrequently used empirical technique and draw insights about the implications of employing distinct modeling specifications in these contexts. I make recommendations about interacting the lagged dependent variable with an exogenous variable in error correction models and use a series of Monte Carlo simulations to show the value of following these recommendations when theoretically and statistically warranted.

In the last chapter, I summarize my findings and conclusions, and discuss them in the broader context of the relevant literatures they contribute to. I give particular consideration to their place in the political economy literature on financial markets. Finally, I explore potential avenues of further research.

2. THE MASS POLITICS OF DEBT MARKETS

2.1 Introduction

Scholars of public opinion have found that mass preferences can influence the economic policies policymakers pursue and the economic outcomes they deliver (Wlezien, 1995; Stimson, 2015). Elections and mass partisanship have also been found to influence investor behavior and market outcomes in various financial domains (Leblang and Mukherjee, 2005; Goodell and Bodey, 2012; Sattler, 2013). Yet, we know little about the role of mass preferences in financial markets outside of elections. Mass approval of the chief executive, for example, varies considerably between elections and is likely to influence perceptions of political risk (Brace and Hinckley, 1992; ICRG Researchers, 2013). Although research on political risk in financial markets is almost always evaluated within the scope of elections, uncertainty about a policymaker's policy choices remains after an election has been decided and policymakers have been seated (Fowler, 2006). Can public opinion influence markets outside of elections?

To answer this question I focus on the market for government debt[1], which is well-known for being particularly sensitive to variation in perceptions of political risk (Bekaert et al., 2014). The market for government debt is also especially important to policymakers, given that it impacts the state's access to capital. I argue that debt market investors expect democratic policymakers to respond to the incentives (and disincentives) provided by mass preferences (Risse-Kappen, 1991; Brace and Hinckley, 1992; Wlezien, 1995; Lax and Phillips, 2012; Stimson, 2015). Hence, mass approval of the chief executive influences investor behavior and, ultimately, open-auction rates

[1] I use the expressions "market for government debt" and "market for treasury bills" interchangeably in this paper.

on treasury bills. Importantly, I consider the role of government ideology and institutional arrangement (i.e., effective number of veto players) as moderators of the effect of mass approval on treasury rates.

It is noteworthy that I depart from, what I label, the *good-payer* story, which political scientists commonly use to explain the role of politics in the market for government debt. (Saiegh, 2005; Ballard-Rosa, Mosley and Wellhausen, 2019). The *good-payer* story argues that governments that can credibly commit to repaying their debt experience lower treasury rates compared to governments that cannot credibly commit to repaying their debt. Instead, I propose the *steady-hand* story where debt markets, having already priced in the credibility of repayment, seek to forecast whether an incumbent will be a steady or an unsteady hand guiding the economy and mitigating macro-economic pressures (Mosley, 2003).

In order to maximize leverage in the variation of mass approval, country-year risk, government ideology, institutional arrangement, and economic development, I use a time-series cross-sectional sample of developed and developing countries. I find considerable support for my theoretical argument. The results indicate that public opinion can in fact influence the market for government debt beyond elections. This article contributes to the broader political economy literature in four major ways. First, this is the first political science article that I am aware of that considers the degree to which the masses can influence markets and, thus, capital owners' vulnerability to the politics of democratic accountability. Second, it contributes to an understanding of democratic accountability beyond elections. Third, it examines the extent to which government ideology and institutional arrangements constrain the influence of public opinion on markets. Finally, this article contributes to an understanding of the intricate relationship between politics and investor behavior.

In the sections that follow, I begin with a discussion of the extant literature on mass politics and markets, economic voting, government debt, and policymaking. After that, I propose a theory for how mass approval affects treasury rates and then outline the resulting hypotheses. In the section after that, I describe my empirical strategy before discussing my findings. Finally, I conclude with suggestions for future research.

2.2 Literature

Mass Preferences and Policymaking

The influence of mass preferences on public policy is well documented in the literature. Researchers have noted that mass preferences serve as a "thermostat" of current policy and, thus, a guide to future policy (Wlezien, 1995; Soroka and Wlezien, 2005; Jennings, 2009; Wlezien and Soroka, 2016). In fact, in the US context, some have warned that modern "presidents govern increasingly on the strength of support from (little understood) public-opinion polls" (Brace and Hinckley, 1992). Issue areas spanning a gap as wide as that between affirmative action and foreign policy have been shown to be influenced by public opinion (Risse-Kappen, 1991; Lax and Phillips, 2012). Some tout this as evidence that democracy works as intended (Przeworski, Stokes and Manin, 1999; Stimson, 2015).

Yet the problems with the connection between mass preferences and democratic governance are not hard to come by. For one, translating the preferences of disparate groups into policy is a difficult task that is largely conditional on the formal and informal institutions within which the system is embedded (Shepsle, 1979; North, 1991; Austen-Smith and Banks, 1996; Jacobs and Matthews, 2017). Democratic preference aggregation depends on the will of the majority and gives disproportionate power to "median" or "swing" voters (Black, 1948*b*; Riker, 1982; Feddersen and Pesendorfer, 1996). The democratic process also hinges on the existence of competitive elections which, despite their advantages, incentivize manipulative and utilitarian policymaking behavior (Kayser, 2005; Krishna and Morgan, 2015; Schleiter and Tavits, 2016; Beckman and Schleiter, 2020). Additionally, election timing, incumbency advantage, and distrust among coalition partners can prevent policymaking from effectively reflecting the will of the masses (Martin and Vanberg, 2004; Lipsmeyer and Pierce, 2011; Martin and Vanberg, 2013; de Benedictis-Kessner, 2018).

Despite these problems, financial market investors are aware that in established democracies mass preferences ultimately influence policy outcomes. Investors gauge political uncertainty in the economy, for example, by using information about the probability of particular parties winning an election (Leblang and Mukherjee, 2005; Jensen and Schmith, 2005; Fowler, 2006; Füss and Bechtel, 2008; Bechtel, 2009; Goodell and Bodey, 2012; Sattler, 2013). To the extent that elections are likely to bring changes to the *status quo*, they induce uncertainty about political and economic conditions. New policymakers, after all, can invigorate hope for capital-friendly conditions or dismantle capitalists' optimism altogether.

Investors might also gauge political uncertainty by looking at the incentives policymakers have to pursue changes to public policies once in office. Mass preferences can be a source of risk to capital owners, who stand to experience many of the immediate effects of new policies (Fowler, 2006; Breen and McMenamin, 2013; Brooks, Cunha and Mosley, 2015). While investors in established democracies may not be at risk of outright expropriation, changes in monetary and fiscal policy, regulatory codes, and enforcement priorities can still impact the business climate and, subsequently, returns on investment (Glick and Leduc, 2012). All else equal, policy change is risky and can be a reason for skepticism among investors (Citron and Nickelsburg, 1987; Brewer and Rivoli, 1990; Busse and Hefeker, 2007; Fortunato and Turner, 2018).

Mass preferences likely influence policymaking and the perception of policymaking risk differently depending on the ideology of those in power.[2] The extensive research on economic voting finds that left-wing governments are often associated with labor-friendly causes while right-wing governments are associated with capital-friendly ones (see Hibbs, 1989, for example). This is because left-wing policymakers represent relatively labor-rich voters while right-wing policymakers represent relatively capital-rich voters. One would expect voters to punish or reward elected officials differently based on voters' own ideology and the ideology of those in power (Hibbs, 2005). Left-wing politicians, thus, have an incentive to pursue policies that prioritize employment even if it comes at the cost of higher inflation. Inversely, right-wing politicians prioritize low inflation even when it achieves a higher unemployment rate. Many scholarly works corroborate the expectation that investors are aware of these dynamics and respond accordingly (Alesina, 1987; Leblang and Mukherjee, 2005; Fowler, 2006; Füss and Bechtel, 2008; Bechtel, 2009; Sattler, 2013).

The degree to which institutional arrangements (such as the number of legislative chambers, the existence of sub-federal veto players, which parties control government institutions, and the size of their coalitions) constrain policymakers and contribute to economic stability can also moderate the role of mass preferences in markets. Institutions determine whether effecting policy change is possible. Some scholars suggest that political information has distinct effects on market volatility depending on institutional arrangements and the level of institutional development (Leblang and Bernhard, 2006; Breen and McMenamin, 2013). In some institutional contexts, for example, there is evidence that investors weight news of protests more heavily than opinion polls because the former can be a more reliable measure of policy effectiveness (Hays, Freeman and Nesseth, 2003). Additionally, democratic institutional arrangements have been found to foster economic stability to

[2]When referring to government ideology in this paper, I focus exclusively on its traditional left-right economic dimension.

the extent that they encourage cooperation among adversaries and lower the risk of expropriation (Rodrik, 2000; Ballard-Rosa, Mosley and Wellhausen, 2019). Hence, political shocks are likely to affect markets differently depending on the institutions in place and which groups of elites control these institutions.

Debt Markets

Studying the role of mass preferences in debt markets in particular is especially beneficial. For one, it is hard to overstate the premium governments place on their ability to issue debt. Public debt is a critical tool for governments to boost the economy in bad times or, in good times, to invest in resources that lead to long-term gains (Alesina and Perotti, 1995; Alesina, Perotti et al., 1999). The cost of accessing debt can greatly limit the ability of governments to deliver public services and maintain economic stability. Their goals stand in contrast to those of lenders, who seek to negotiate higher interest rates in order to secure higher returns.[3] As such, countries that pose a higher risk of default or inflation growth, for instance, are charged higher risk premiums (Edwards, 1983; Saiegh, 2005). These premiums are continuously adjusted based on lenders' expectations about the government's ability to manage the economy and service its debt.

Studies on the politics of debt markets mostly look at governments' credible commitment to repaying debt (Citron and Nickelsburg, 1987; Brewer and Rivoli, 1990; Saiegh, 2005; Ballard-Rosa, Mosley and Wellhausen, 2019). This is what I label the *good-payer* story. Under this view, countries where economic policies and institutions provide credible assurance of stability and low risk are the ones most likely to enjoy the lowest interest rates. Established industrialized democracies with a good-payer track record are seen as low risk and, thus, access debt on better

[3] Investors *negotiate* these rates at open auction of government securities (see Fleming, 2007).

terms. In these countries, high demand for public debt leads to low interest rates. In some cases, countries with a record of low inflation and low default risk are charged negative rates—where investors pay to lend money to the government.[4]

2.3 Theory

While the *good-payer* story has an intuitive logic, I argue that it is incomplete. Even in *riskier* developing countries, I expect government debt to hold a *safe-haven* status relative to other investment instruments. This is because higher overall political risk is already accounted for across different families of investment assets. Investments backed by the state, such as treasury bills, are likely to present relatively lower risk than domestic equities, for example. I argue that domestic investors, thus, primarily seek to forecast a government's (mis)handling of the economy—which I label the *steady-hand* story. Once investors form an opinion of a country's overall level of risk, forecasted changes in economic policy—and expectations about a leaders' ability to manage currency and inflationary pressures (Mosley, 2003)—they adjust their preferred interest rates. Importantly, deterioration in expectations about economic policy management leads to demands for higher interest rates. Improvement in expectations about economic policy management leads to demands for lower interest rates.

Investors continually adjust their demand for a collection of investment assets in order to hedge against the risk of policy change. When changes in risk are perceived, they re-calibrate their investment positions.[5] And they seek to do so as far ahead as possible—in line with the goal of

[4]Negative real rates on government securities are an unusual and fairly recent phenomenon. They appeared after the Great Recession largely as a response to a slow economic recovery and expectations of deflation. See Anderson, Liu et al. (2013) for a brief but informative discussion of this topic. See Jackson (2015) and Arteta et al. (2016) for more detailed expositions.

[5]This re-calibration occurs at different intervals depending on how often treasury bills and bonds are commercialized at open auction. In some countries these auctions take place weekly, in others bi-weekly, monthly, and at wider or irregular intervals.

out-investing the crowd. In any political system, they do this by predicting the moves of those in power and forecasting subsequent effects. As in any statistical model, the forecaster seeks to minimize the unexplained—and uncertain—error term. *Ceteris paribus*, uncertainty is always less preferred than certainty. As such, I assume that domestic investors prefer no policy change over any change at all (Brewer and Rivoli, 1990; Fortunato and Turner, 2018). If a sign of change is on the horizon, investors downgrade economic expectations due to greater uncertainty—which is reflected in interest rates.

In order to forecast policy change in an inter-election period, investors consider that incumbents sometimes have an incentive to respond to mass preferences. Executive approval ratings, in particular, serve as a sign of support for a government and can be a reason for those in power to pursue policy changes or to hold off on them. Thus, executive approval is an important determinant of the *steady-hand* story. I argue that other key determinants are the ideology of those in power and the institutional arrangements within which policy changes become feasible or unfeasible. The interaction between executive approval and government ideology and the interaction between executive approval and institutional arrangements have implications for policy change. Thus, they also have implications for investors seeking to forecast policy change. Hence, this is reflected in interest rates on government securities. I present this argument in Figures 2.1 and 2.2, where the solid arrows represent separate causal effects that moderate one another.

In Figure 2.1, I use Arrow (I) to demonstrate the connection between executive approval and interest rates on government debt. Recall that this connection stems from policymakers' interest in pleasing the electorate in order to win reelection and stay in power (Hibbs, 2005). To optimize their chances of reelection, policymakers might implement policies they otherwise would not. In some cases, they might refrain from pursuing policies they would normally prefer. This could play out in

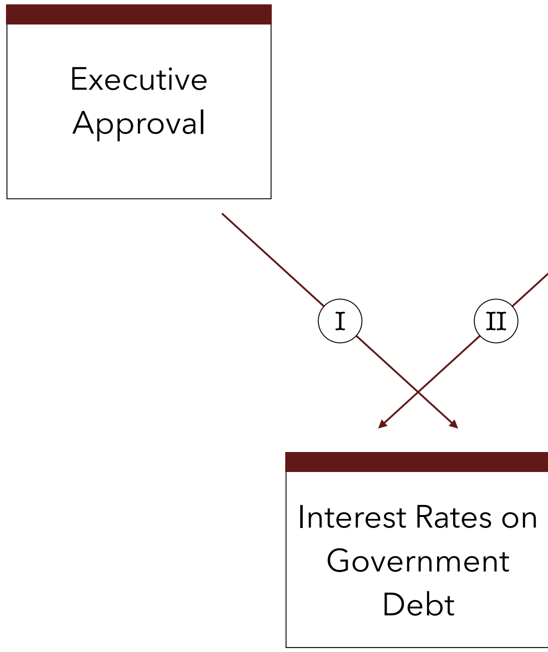


Figure 2.1: Theorized role of executive approval and government ideology as determinants of interest rates on government debt. (Solid arrows represent separate causal effects that moderate one another.)

one of at least two ways. First, policymakers could make concessions to the opposition in order to gain reciprocal concessions elsewhere. This is more likely to occur under divided government or in legislative contexts where a supermajority is necessary (Alesina, Rosenthal et al., 1995). Second, they could pursue (or refrain from pursuing) specific policies in order to increase the probability of winning over the median voter during the next electoral cycle. In any case, policymakers update their expectations, and change their actions, to meet constituent expectations (Downs et al., 1957; Black, 1948*b*; Wlezien, 2004; Stimson, 2015).

I use Arrow (II) to show an additional relationship between government ideology and interest rates on government debt. Previous research has shown that markets punish left-wing governments relative to right-wing governments (Sattler, 2013). Specifically in the context of debt markets, one would expect that investors would perceive right-wing governments as more capital friendly and fiscally disciplined than their left-wing counterparts, warranting relatively better lending terms during the tenure of right-wing governments.

I argue that executive approval and government ideology moderate the effect of one another on interest rates. This is shown by the intersection of Arrows (I) and (II) in Figure 2.1. As approval increases, governments are more likely to implement their preferred policies for various reasons. First, they have more political capital—defined as prestige and political influence (see Putnam, 1993; Booth and Richard, 1998; Sørensen and Torfing, 2003; Nee and Opper, 2010; Thrower, 2017). Second, they are empowered relative to institutional veto players. Third, they are more confident that the public approves of their policies. On the other hand, as approval decreases, the government is weakened relative to institutional veto players and is less likely to pursue its set of preferred policies. Hence, an increase in approval of a left-wing chief executive—which markets are predisposed to disliking—signifies less desirable economic policies. Investors therefore demand relatively higher interest rates. On the other hand, an increase in approval of a right-wing chief executive—which markets are predisposed to preferring—signifies more desirable economic policies. Investors therefore demand relatively lower interest rates.[6]

As in Figure 2.1, I use Arrow (I) in Figure 2.2 to show the theorized connection between executive approval and interest rates on government debt. As before, investors expect chief executives to respond to variation in public opinion. In addition, institutional arrangements determine the ability of policymakers to meet constituent expectations. I assume that investors are aware of this and theorize that they adjust their demand for government debt in response. This connection, por-

[6] I thank an anonymous reviewer at the *Journal of Politics* for helping me reframe and rephrase the arguments in this paragraph.

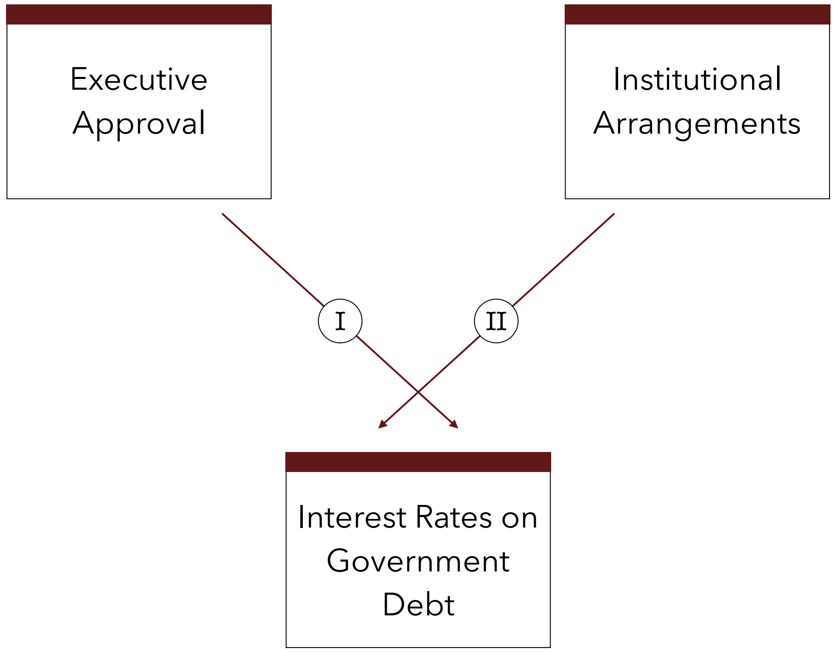


Figure 2.2: Theorized role of executive approval and institutional arrangements as determinants of interest rates on government debt. (Solid arrows represent separate causal effects that moderate one another.)

trayed by Arrow (II) in Figure 2.2, is due to the norms and procedures that pave the way for policy proposals and funnel them into law. Traditionally, this argument revolves around the role of veto players (Tsebelis, 2002). The greater their number, the harder it is to make changes to the *status quo*. Keep in mind that this also implies greater policy stability and credibility (Keefer and Stasavage, 2003). Institutional norms can also influence agenda setting, having major ramifications for policy changes (McKelvey, 1976).

Institutional arrangements may also affect the extent to which popular approval matters in financial markets. This is because incumbents are likely to respond to approval differently depending on the set of veto players constraining them. It is easy to imagine an institutional context that provides considerable flexibility for incumbents to respond to decreases in approval with desperate policy changes. Another institutional context might hinder the ability of those in power to carry out their agenda. I argue that investors adjust their demand for government securities accordingly, thus the moderating relationship suggested by the intersection of Arrows (I) and (II) in Figure 2.2. It is important to keep in mind that an equivalent relationship exists where public opinion also modifies the effect of institutions on demand for government securities. High executive approval might weaken commitment to democratic norms, for example.

It is well understood that institutions influence policymaking by incentivizing or deterring political accountability. Since voters have imperfect information on policies and the policymaking process, those in power have a unique opportunity to strategically filter information (Przeworski, Stokes and Manin, 1999; Jones and Baumgartner, 2005; Ashworth, 2012). They can use it with varying degrees of biases and selectivity to set an agenda that furthers their goals. Institutional arrangements can thus sharpen or mask policy choices and actions. This builds on the concept of *clarity of responsibility*—first introduced in Powell and Whitten (1993) and further explored in Hibbs (2005). When voters cannot attribute policy changes to specific policymakers, these policymakers are better able to deceive voters by using potentially incompetent means to deliver electorally advantageous results—like running unnecessary deficits or abusing discretionary power. Less clear contexts, thus, provide a platform for unpredictable behavior—the unsteady hand. This implies greater policy risk. Otherwise, policymakers are beholden to *clearer* scrutiny by their constituents and must seek to act more pragmatically. In Figure 2.2, Arrow (II) reflects both the more traditional *veto-player* role of institutions and this more nuanced *clarity* effect.

Following these expectations, I put forth the following complementary hypotheses. Together, they capture the nuances of public opinion and the risks public opinion instills in debt markets in the short run, between election cycles.

H_{1a}: Under left-wing governments, an increase in approval leads to expectations of relatively less desirable economic policies. Investors therefore demand relatively higher interest rates on government debt.

H_{1b}: Under right-wing governments, an increase in approval leads to expectations of relatively more desirable economic policies. Investors therefore demand relatively lower interest rates.

H_{2a}: Under few institutional constraints, an increase in approval leads to expectations of more policy changes. Investors therefore demand relatively higher interest rates on government debt.

H_{2b}: Under many institutional constraints, an increase in approval leads to a null effect on expectations about policy changes and, thus, a null effect on interest rates.

2.4 Empirical Model

Data

Time-series cross-sectional data would be ideal for testing the hypotheses I put forth. Leveraging variation across time *and* space is rarely an option when dealing with public opinion data. In part, this is because of inconsistent or relatively low-frequency coverage in some countries—particularly developing countries. But this is also because of the challenges involved with comparing available data over time and across countries. Improvement in the quality and consistency of time-series cross-sectional data now available to political scientists makes the analysis of these data possible.

To test my hypotheses, I estimate a time-series cross-sectional model with data from nine developed and developing countries: Australia, Brazil, Canada, Mexico, Philippines, New Zealand, UK, USA, and Uruguay. While these are the only countries for which appropriate data are available, this sample provides a good opportunity to evaluate my theoretical argument because of the variation in approval and institutional development across these countries. The time period covered varies by country and always falls between 1975 and 2016. The regression for Brazil, for example, sees coverage between 1988 and 2016, while that for the United States includes 1975 through 2016.

The dependent variable, short-term treasury bill interest rates, was obtained from the International Monetary Fund's (IMF) International Financial Statistics dataset. These are quarterly data representing the expected returns on short-term treasury securities.

I use Carlin et al.'s (2019) quarterly approval data, for either percentage "government" or "executive" approval. This dataset assembles various series of publicly available variables reporting popular opinion concerning chief executives (1). To measure ideology, I use the World Bank's Database of Political Institutions ranging across left, right, and center. I also use Henisz's (2000) political and institutional constraints variable to measure political context. This is a time-series measure of the "feasibility of policy change" and ranges from 0 (no constraints) to 1 (fully constrained). For the time period included in each analysis, all countries have an average Henisz score between 0.6 and 0.9.

I control for the central bank rate and the consumer price index to capture inflation. I also include an elections dummy in the model.[7,8] It is set to 0 in quarters when no election happens and 1 when an election takes place. The central bank rate and consumer price index were both obtained from the IMF's International Financial Statistics dataset.

[7]The elections dummy variable accounts for presidential elections in presidential systems and legislative elections in parliamentary systems.

[8]I experimented with a continuous variable measuring the number of quarters left until the next elections. Results were statistically and substantively similar to what I obtained while including an elections dummy instead.

Modeling Strategy

I use lagged dependent variable models to test my hypotheses. Lagged dependent variable models include temporal lags of the dependent variable, estimated using ordinary least squares (OLS) regression. These dynamic models also allow for the estimation of persistence in the dependent variable. As with most financial data, I expect movement in treasury rates to be robustly persistent over time. I am, however, agnostic about the different short- and long-run implications for the relationships theorized in the previous section.

I specify the separate moderating effects of government ideology and institutional constraints on the effect of executive approval on treasury rates by including, separately, two multiplicative interactions in the model. These specifications follow Brambor, Clark and Golder's (2006) recommendations. My model specifications are as follows:

$$\begin{aligned}\text{Interest}_{i,t} = \alpha_0 + \alpha_1 \text{Interest}_{i,t-1} + \\ + \alpha_2 \text{Approval}_{i,t} \times \text{Ideology}_{i,t} + \\ + \alpha_3 \text{Approval}_{i,t} + \alpha_4 \text{Ideology}_{i,t} + \\ + \sum_{j=5}^{P} \alpha_p \text{Controls}_{p,i,t} + \varepsilon_{i,t}\end{aligned} \tag{2.1}$$

$$\begin{aligned}\text{Interest}_{i,t} = \beta_0 + \beta_1 \text{Interest}_{i,t-1} + \\ + \beta_2 \text{Approval}_{i,t} \times \text{Institutions}_{i,t} + \\ + \beta_3 \text{Approval}_{i,t} + \beta_4 \text{Institutions}_{i,t} + \\ + \sum_{j=5}^{P} \beta_p \text{Controls}_{p,i,t} + \mu_{i,t}\end{aligned} \tag{2.2}$$

Where,

- $\text{Interest}_{i,t}$ is a continuous variable, interest rates on treasury bills in percentages, for country i at time t
- $\text{Approval}_{i,t}$ is a continuous variable, incumbent approval in percentages, for country i at time t
- $\text{Ideology}_{i,t}$ is an ordinal variable, government ideology where 0 is left, 1 is center, and 2 is right, for country i at time t
- $\text{Institutions}_{i,t}$ is a continuous variable, institutional feasibility of policy change where greater values represent more constraints, for country i at time t
- α_0 and β_0 are vertical intercepts in their respective equations,
- $\alpha_1, \alpha_2, \alpha_3, ..., \alpha_{P-1}, \alpha_P$ and $\beta_1, \beta_2, \beta_3, ..., \beta_{P-1}, \beta_P$ are parameter estimates for the effect of right-hand-side variables, and
- $\varepsilon_{i,t}$ and $\mu_{i,t}$ are white-noise error terms that vary across each country, i, and time, t, in their respective equations.

2.5 Results

In Figure 2.3, I present the average marginal effects of executive approval on treasury rates across left, center, and right governments. These results are partially consistent with my theoretical expectations. A one-percent increase in government approval under left governments leads to a small but positive and statistically significant effect on treasury rates at the 95% confidence level. However, this effect is not distinguishable from that produced by an increase in approval under center or right governments. This is partially consistent with hypothesis H_{1a}, which states that under left-wing governments, an increase in approval leads to an increase in interest rates. This increase in treasury rates occurs because an increase in approval of left-wing governments leads to expectations of relatively less desirable economic policies. Hence, investors demand relatively higher interest rates on government debt.

Under right governments, an increase in approval has no statistically significant effect on treasury rates. This finding is not consistent with hypothesis H_{1b}, which states that under right-wing governments, an increase in approval leads to relatively lower interest rates. Perhaps, because investors are predisposed to preferring right-wing governments and their policies, they are less likely to respond meaningfully to changes in approval and potential policy change when the right is in power.

I did not have specific expectations about the effect of approval under center governments at the outset. Yet, I find that under center governments, approval has a positive and statistically distinguishable effect on treasury rates at the 90% confidence level. Interestingly, the parameter estimate recovered is smaller than that recovered under left governments and greater than that recovered under right governments. However, this effect is not statistically different from that under left governments or right governments.

In Figure 2.4, I show the average marginal effects of executive approval on treasury rates across levels of institutional constraints. Lower values along the horizontal axis represent fewer constraints and higher values along the horizontal axis represent more constraints (Henisz, 2000). Figure 2.4 is consistent with hypotheses H_{2a} and H_{2b}. Under few constraints, an increase in approval is associated with a positive and statistically significant effect on treasury rates. As is the case under left governments, markets are predisposed to punishing governments in political con-

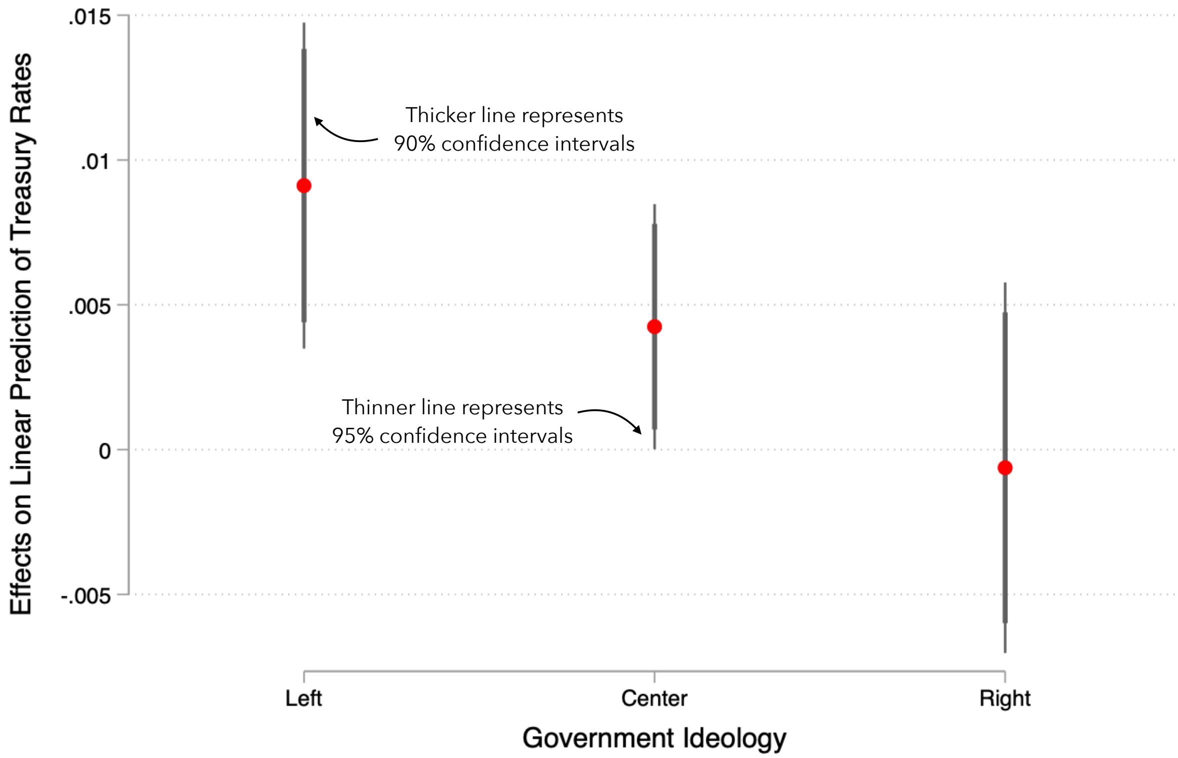


Figure 2.3: Average marginal effects of executive approval on treasury rates across different government ideological orientations.

texts where institutional constraints are few. Naturally, investors forecast higher uncertainty with regards to policy change and demand higher interest rates to compensate for higher risk.

As shown in Figure 2.4, this effect becomes smaller as institutional constraints increase. Under many institutional constraints, the effect disappears altogether as in the case for right governments in Figure 2.3. Under many institutional constraints, markets are likely less concerned about approval and its ability to signal demand for policy change because the institutional context itself is able to constrain demand for policy change.

In both models (see regression Table A.1 in Section A.1 of Appendix A, the lagged dependent variable is positive and statistically different from zero at the 99.9% confidence level. This estimate

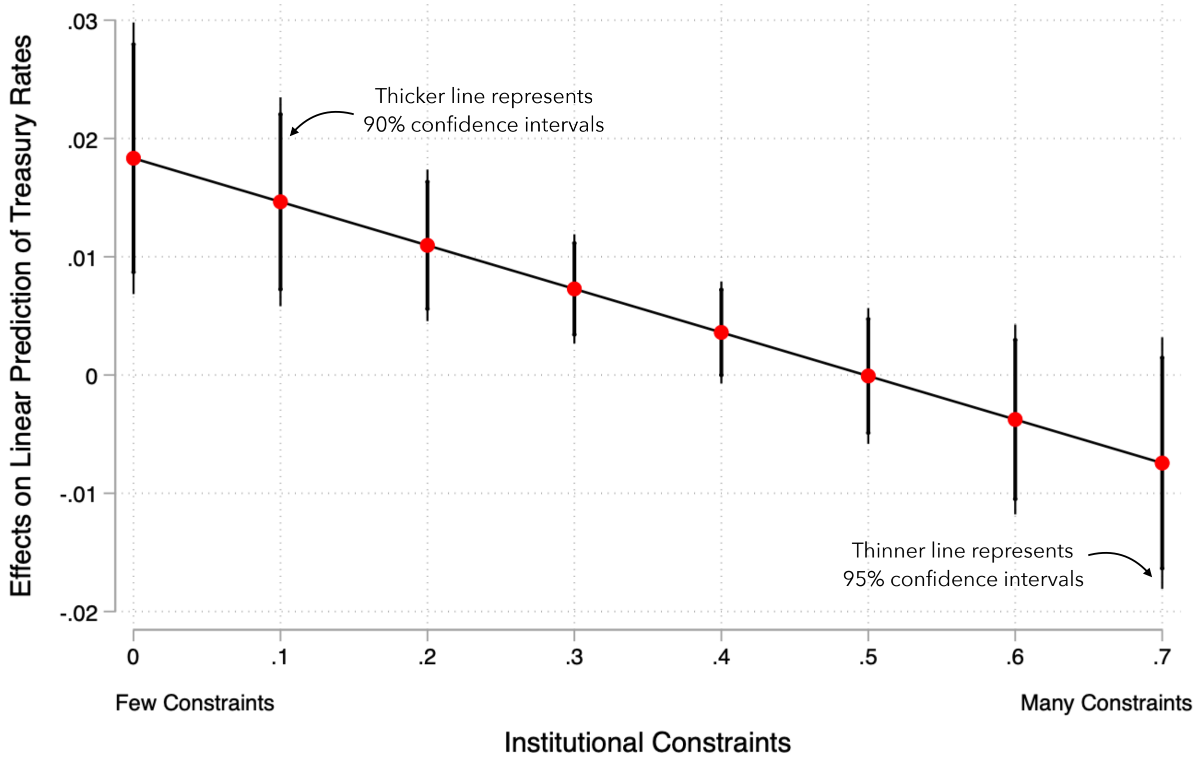


Figure 2.4: Average marginal effects of executive approval on treasury rates across levels of institutional constraints.

suggests fairly strong persistence in treasury rates between quarters, which is frequently observed in financial data more broadly.[9]

2.5.1 Model-Based Monte Carlo Simulations

I assess the dynamic effects of approval on treasury rates by producing a series of Monte Carlo simulations using the regression results reported in Table A.1. To do this, I (1) estimate the model using an ordinary least squares strategy, (2) simulate the parameter estimates 1000 times drawing from a sampling distribution with mean equal to the recovered parameters and variance equal to

[9]Following the recommendations of Keele, Stevenson and Elwert (2020), I do not provide an interpretation of the effects on control variables.

the variance-covariance matrix of estimates, (3) set baseline values of continuous control variables to their means and dichotomous control variables to zero, (4) set values of "shock variables" of interest before and after "shock times," and (5) simulate predicted values of treasury rates over ten time periods.[10] I introduce two back-to-back shocks since I find it more plausible that approval generally moves gradually, rather than abruptly, in any given quarter.[11] I repeat this procedure in six different scenarios for the model specified in Equation 2.1 as well as six different scenarios for the model specified in Equation 2.2.

In Figure 2.5, I show simulations based on the model specified in Equation 2.1, which includes a multiplicative interaction between approval and government ideology. In panels 2.5a and 2.5b, I show the simulated effects of shocks to approval when left governments are in power. In panel 2.5a, I induce a half-standard-deviation ($+6.2\%$) shock to approval at time $t = 3$ and an additional half-standard-deviation shock to approval at time $t = 5$. The initial shock does not lead to a statistically distinguishable effect on treasury rates—immediately or in the next time period. However, an additional shock at time $t = 5$ leads to an increase of approximately one tenth of a percentage point in treasury rates, relative to baseline. This effect is statistically significant at the 95% confidence level. In panel 2.5b, I repeat the procedure in panel 2.5a except with one-standard-deviation ($+12.4\%$) shocks at time $t = 3$ and $t = 5$. Compared to baseline, the first shock leads to a 0.1% increase in treasury rates in the next time period, $t = 4$. The additional shock at $t = 5$ leads to a further 0.1% immediate increase in treasury rates. Simulated treasury rates reach equilibrium around time $t = 8$, roughly 0.3 percentage points higher than at baseline.

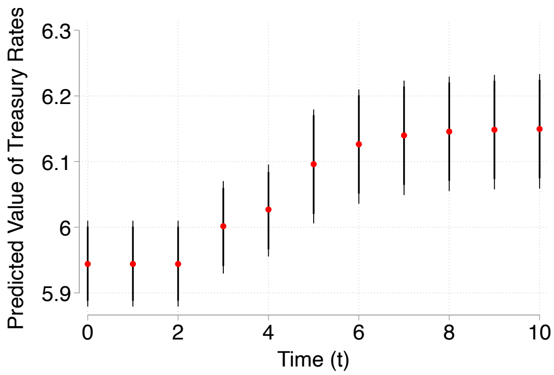
In panels 2.5c and 2.5d, I repeat the procedures followed in the simulations shown in 2.5a

[10] I use Tomz, Whittenberg and King's (2003) `clarify` package to perform parts 1-3.

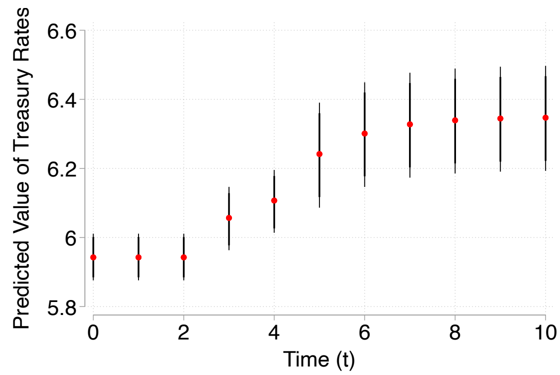
[11] Importantly, the size or repeated introduction of shocks does not influence the value of simulated parameter estimates or standard errors.

and 2.5b except for setting government ideology to “center.” The two half-standard-deviation shocks to approval shown in 2.5c do not lead to a statistically distinguishable effect on treasury rates at the 95% or, even, at the 90% confidence levels. Similarly, the one-standard-deviation shocks at time $t = 3$ and $t = 5$, shown in 2.5d, do not lead to statistically significant effects immediately. At equilibrium (around $t = 8$), a 0.1-percentage-point increase in treasury rates can be observed, though this is only statistically significant at the 90% confidence level.

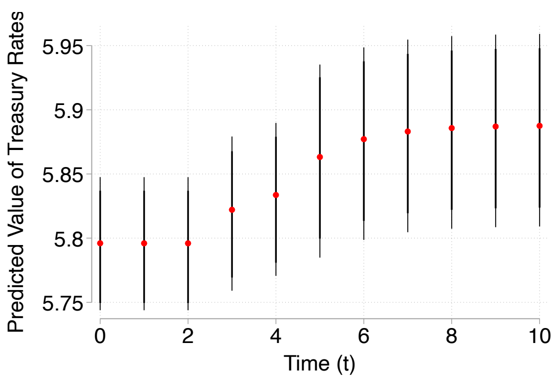
In panels 2.5e and 2.5f I show simulated effects of shocks to approval when government ideology is set to “right.” The two half-standard-deviation shocks shown at $t = 3$ and $t = 5$ in 2.5e have no statistically significant effect on treasury rates at any point after the shock. Similarly, the one-standard-deviation shocks shown at $t = 3$ and $t = 5$ in 2.5f have no statistically significant effect on treasury rates.



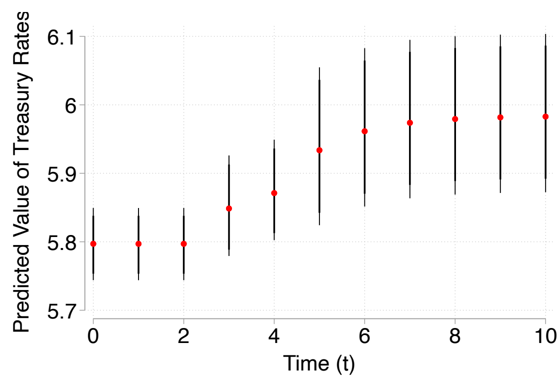
(a) +1/2 SD at t=3, +1/2 SD at t=5
Left



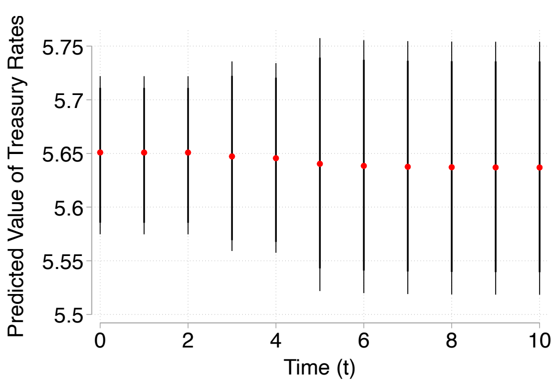
(b) +1 SD at t=3, +1 SD at t=5
Left



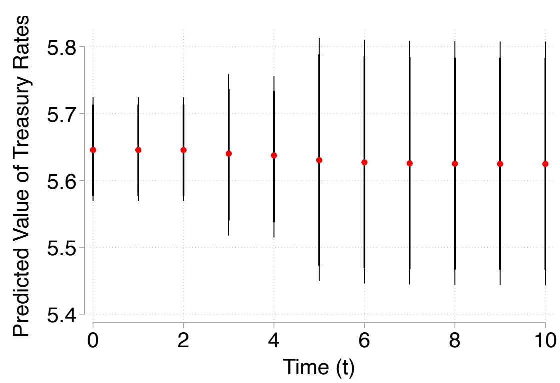
(c) +1/2 SD at t=3, +1/2 SD at t=5
Center



(d) +1 SD at t=3, +1 SD at t=5
Center



(e) +1/2 SD at t=3, +1/2 SD at t=5
Right



(f) +1 SD at t=3, +1 SD at t=5
Right

Figure 2.5: Simulated effect of an increase in incumbent approval on treasury rates under left, center, and right governments

2.6 Conclusion

In this study, I argue that mass approval of the chief executive influences investor behavior and, ultimately, open-auction rates on treasury bills. Market investors use information about mass political opinion to evaluate policymaking incentives and forecast whether incumbents are likely to pursue policy changes. Importantly, the theory I put forth complements the typical *good-payer* story about credible commitment of repayment with a *steady-hand* story about the role of economic policy management in debt markets.

To test the implications of my theory, I consider the role of mass political opinion as a source of policymaking risk within and well beyond electoral cycles. I argue that incumbent popular approval affects the incentives incumbents have to pursue policy changes which, in turn, affect treasury rates. These effects depend on (1) the ideology of the government in power and (2) institutional constraints on the incumbent's ability to effect policy changes.

Using a time-series cross-sectional sample, I find overall support for my theory. Additional model-based Monte Carlo simulations also produce results that are consistent with my theoretical expectations and the inferences drawn from regression findings. Ultimately, this study contributes to the literature by (1) considering the relationship between mass political opinion and public debt, (2) evaluating these effects in times away from elections, (3) examining the role of ideology and institutions in the aforementioned relationship, and (4) connecting policymaking and investor behavior.

While this work is a valuable addition to the political economy literature, it says little about the connection between mass preferences and other investment assets (e.g., equity markets). It also fails to explore other sources of policymaking risk (e.g., populist tendencies, democratic recession) or other measures of mass preferences (e.g., social media presence, news media coverage). Future work should extend this research along those lines.

3. THE MASS POLITICS OF STOCK MARKET VOLATILITY

3.1 Introduction

Empirical evidence suggests that stock markets respond negatively to news of left-wing candidates winning elections (e.g. Leblang and Mukherjee, 2005; Bechtel, 2009; Sattler, 2013). In theory, investors believe that right-wing policymakers are more likely than left-wing policymakers to pursue policies that will optimize macroeconomic outcomes and maximize stock market returns. On net, markets effectively "punish" left-wing election victories by decreasing trading and speculation, and "reward" right-wing election victories by increasing trading and speculation (Sattler, 2013).

But the effect of politics and policymaking on stock markets—particularly stock market volatility—is unlikely to be limited to elections. Elected officials may modify parts of their agenda over time and adjust their commitment to specific policy issues depending on popular demand. In addition, optimal policy choices will change in light of socio- and marcoeconomic conditions. As such, uncertainty about a policymaker's policy choices remains throughout her time in office (Fowler, 2006).

Do stock markets remain more likely to "punish" left-wing governments, relative to right-wing governments, outside of election cycles? To answer this question, I focus on the effect of mass political opinion on stock market volatility. I argue that mass political opinion embodies the core political capital necessary for democratic policymaking. Popular approval emboldens an incumbent government to pursue its policy agenda. By this logic, stock markets will "punish" relatively more popular left-wing governments — by engaging in lower trading volume and inducing lower volatility — and will "reward" relatively more popular right-wing governments — by engaging in greater trading volume and inducing greater volatility. These relationships will differ depending on the level of development of a country's financial institutions.

With a sample of emerging market economies that vary in their level of financial institutional development, I test the impact of executive approval on stock market volatility and find support for my theoretical arguments. Results suggest that stock market investors "punish" more popular left-wing governments within and well beyond election cycles and do so differently in light of a country's financial institutional development. These findings contribute to the extant literature in three primary ways. First, I present a theory for the effect of politics and policymaking on financial market volatility outside of election cycles. Second, I develop the first article on the effect of mass political opinion on stock market volatility. Third, I test a theory on the politics of stock markets — often exclusively tested using data from developed markets — with data from emerging markets.

In the sections that follow, I begin with a discussion of the literature on the political economy of stock markets. I then propose a theory for how mass political opinion affects stock market volatility. After that, I present my research design and modeling strategy. I test my theoretical expectations with data from four emerging markets. Finally, I conclude with suggestions for future research.

3.2 Background

Stock market investors derive expectations about a firm's performance, price them accordingly and, in turn, trade "shares" of ownership at public stock exchanges (e.g., NASDAQ). Such expectations are generally derived in reference to a *status quo* performance and a respective valuation of the firm. If market signals indicate that a firm will perform better than its *status quo* performance, then investors bid a higher price for a share of that firm. If instead market signals indicate that a firm will perform worse than its *status quo* performance, then investors bid a lower price for a share of that firm. In some cases where limited information about a firm is available (e.g., penny stocks), little or no bidding leads to unchanged prices for a share of that firm. According to the efficient markets hypothesis, stock prices incorporate all publicly available information relevant to firms (Fama, 1970, 1995, 2021).

Information about markets is often abundant and diverse. For one, government regulations around the world generally require firms listed on public stock exchanges to disclose performance information. Standardized briefings are periodically issued to report a firm's financial health, market expectations, risks, goals, and strategies, among other things. Withholding information from mandatory reports or prevaricating on these reports can be punished with hefty fines as well as civil and criminal lawsuits (see Dechow, Sloan and Sweeney, 1996; Karpoff, Lee and Martin, 2014). In addition, mainstream media sources dedicate substantial resources to report on industry health, predict moves by executives at large publicly traded firms, and even speculate on individual stocks (Engle and Ng, 1993; Jones, Lamont and Lumsdaine, 1998). Media reports on the impact of elections and government policy on stocks are also widespread.

Research shows that, around elections, financial markets evaluate the expected benefits and costs of politics and policymaking to the price of investment assets (Fowler, 2006; Füss and Bechtel, 2008; Bechtel, 2009; Goodell and Bodey, 2012; Ballard-Rosa, Mosley and Wellhausen, 2019). At a macro level, investors are interested in predicting how the actions of policymakers — both at the executive and legislative branches — will impact the health of the economy. At a micro level, investors seek to gauge the extent to which public policies will boost or hinder the performance of specific industries and firms. On both ends, the goal is to anticipate fiscal, monetary, and regulatory policy changes likely to affect investments.

Empirical analyses of the impact of politics on stock markets have largely focused on election-related phenomena — mostly in the United States and other developed economies (e.g. Roberts, 1990; Herron et al., 1999; Herron, 2000; Leblang and Mukherjee, 2005; Goodell and Bodey, 2012). The predominant argument is that markets usually prefer particular candidates or political parties to win elections. Such candidates are more likely to frame their campaigns as "pro-capital" and to defend supply-side economic policies, such as low taxes (Akard, 1992; Prillaman and Meier, 2014). Scholars often suggest that uncertainty about who will win an election drives volatility in stock markets (Jensen and Schmith, 2005; Füss and Bechtel, 2008).

Leblang and Mukherjee (2005), in particular, provide a comprehensive discussion, as well as formal and empirical models of the relationship between electoral expectations and stock market volatility. Their claims are founded on the assumption that trading volume is directly related to stock volatility. In unpacking a causal argument about the role of partisanship and electoral expectations in financial markets, they identify three important implications of their study. (1) Expectations of left-wing victories, vis-à-vis right-wing victories, lead to lower stock prices but higher market stability (lower volatility). (2) Markets (in the US and the UK) are generally sensitive to political phenomena. (3) Right-wing governments and candidates have an advantage over left-wing governments in "priming the economy" ahead of elections (800). A broader claim in this study is that all governments have an incentive to exert influence on markets considering that markets themselves influence macroeconomic outcomes.

Market expectations about government policy are rooted in the political economy literature where it is well established that the ideology of those in power is a useful signal to investors[1]

[1]When discussing government ideology, I focus exclusively on its traditional left-right economic dimension.

(Hibbs, 1989). On one hand, left-wing policymakers largely represent labor-rich constituencies and defend labor-friendly causes, such as increases to statutory minimum wages and expansion of workers' benefits. On the other hand, right-wing policymakers largely represent capital-rich constituencies and defend capital-friendly causes, such as lower corporate taxes and limited industry regulation. The economic voting literature suggests that, in the presence of competitive elections, these constituencies punish or reward policymakers for their economic policies and the macroeconomic outcomes they deliver (Hibbs, 2005). Thus, rational policymakers have an incentive to pursue policies that prioritize the causes of their respective constituencies. It is understood that investors believe that right-wing policymakers are more likely than left-wing policymakers to pursue policies that will maximize stock market returns (Herron, 2000; Bechtel, 2009; Leblang and Mukherjee, 2005; Sattler, 2013).

That said, political-economic dynamics constrain the extent to which a policymaker translates her ideology into actual policy. The Median Voter Theorem indicates that policymakers ultimately have to win over median voters to win an election; the Meltzer-Richard model suggests that socioeconomic conditions may shift the median voter's position over time (Black, 1948*a*; Downs et al., 1957; Meltzer and Richard, 1981; Alesina and Rodrik, 1994; Congleton, 2004). Recessions and inflation, for instance, will influence popular demand for redistributive policies and may drive a "swing to the left" or "swing to the right" among the electorate (Lewis-Beck and Stegmaier, 2000; Duch and Stevenson, 2010; Lindvall, 2014, 2017).

Of course, enacting policy is also a function of the kind and quality of political institutions in place. Unfriendly veto players in control of competing institutions, including the judiciary, can (and often will) dilute the government's policies or block them altogether (Henisz, 2000, 2004; Tsebelis, 1995, 2002, 2011). The clarity of responsibility for economic conditions and policies of certain institutional players may also affect the electoral incentives policymakers have to pursue an agenda (Powell and Whitten, 1993). Financial markets are aware of these dynamics and respond accordingly (Breen and McMenamin, 2013; Ballard-Rosa, Mosley and Wellhausen, 2019). Importantly, investors may themselves possess a veto player role in the kinds of economic policies policymakers pursue, given the former's ability to threaten the latter with capital flight (Przeworski and Wallerstein, 1988).

Political economists have argued that markets are prone to "punishing" left-wing governments in relation to right-wing governments (Leblang and Mukherjee, 2005; Bechtel, 2009; Sattler, 2013). Because stock market investors believe that left-wing policies are less favorable for firms than right-wing policies, they are more likely to expect weaker firm performance and undervalue stocks when the left is in power. On net, the literature suggests that average stock trading (volume) is more likely to be lower under left governments than under right governments.

It stands to reason that investors consider political risks in light of the market's ability to weather these risks. Larger, more diverse, and better connected markets provide a wider set of asset alternatives and more hedging opportunities (Merton, 1990). Specifically, financial institutional development—defined in relation to financial depth (value of financial sector), access to financial markets (relative number of participants), efficiency (overall sector profitability), and value stability—influences investor expectations and behavior (World Bank, 2019). Financial institutional development also speaks to "market quality" which, among other things, comprises market liquidity and market heterogeneity (Holmström and Tirole, 1993; Chordia, Roll and Subrahmanyam, 2001; Baker and Stein, 2004; Brunnermeier and Pedersen, 2009). Almost certainly, financial institutional development continuously impacts investment decisions and the extent to which these decisions incorporate information about politics.

Yet, scholars have almost exclusively focused on elections or election-related phenomena (e.g., pre-election polling), overlooking the role of politics in financial markets beyond elections (Fowler, 2006). There is also ample evidence that emerging market economies are relatively more sensitive to politics and policy changes (Diamonte, Liew and Stevens, 1996; Bilson, Brailsford and Hooper, 2002). Still, no work in comparative politics develops and tests a theory about stock markets that incorporates the continuous role of mass political opinion, incumbent ideology, and financial institutional development.

3.3 Theory

Policymakers ultimately need a degree of political capital to pursue and enact policies (Putnam, 1993; Booth and Richard, 1998; Sørensen and Torfing, 2003; Nee and Opper, 2010; Thrower, 2017). In this context, I define political capital as the goodwill, prestige, and influence a political actor enjoys in her interactions with other political actors. Such political capital stems from (1) support from political elites, including those in control of competing government institutions, (2) support from economic elites, and (3) support from the masses.

These sources of political capital are not exogenous to one another (Brody, 1991; Zaller et al., 1992; Converse, 2006). One one hand, there is abundant evidence that both political and economic elites seek to influence mass preferences in order to induce issue convergence and elite-mass congruence (Herman and McChesney, 1997; Keller, 2017). In no small part, elites are interested in influencing mass preferences because the masses determine the outcome of elections and, thus, public policies.[2] On the other hand, mass preferences have been found to influence elite behavior and government policy (Risse-Kappen, 1991; Wlezien, 1995; Soroka and Wlezien, 2005; Jennings,

[2]Less relevant to this discussion: economic elites also attempt to influence mass *consumer* preferences in order to maximize revenue for specific firms and, thus, maximize financial returns.

2009; Lax and Phillips, 2012; Wlezien and Soroka, 2016). Importantly, the dynamic interaction between elite-mass preferences is a continuous process at play within and well beyond election cycles.

With this in mind, I argue that mass political opinion embodies *core* political capital — the kind that also underpins political and economic elite support — necessary for democratic policymaking. Executive approval, in particular, empowers the incumbent government to pursue its policy agenda. This occurs because popular governments are more likely to also enjoy greater popular support for their policy agenda. Hence, (1) they are more likely to be rewarded by voters at the next election cycle for pursuing their policy agenda, while (2) those opposing such policies are more likely to be punished by voters at the next election cycle. By this logic, popular governments are less likely to face policy opposition from elected officials in control of competing government institutions. The policies of popular governments, if already in place, are more likely to remain in place; if not in place, their policy proposals are more likely to become law.

Stock market investors form expectations about the odds that an incumbent government's policy platform will remain or become law. They adjust their demand for individual stocks in light of their expectations about the impact of the incumbent government's policies on financial returns. Investors seek to understand whether government policy will influence the health of the economy as well as that of specific economic sectors and firms.

The extent to which investors form expectations about executive approval and adjust demand for stocks will differ in light of financial institutional development. As discussed in the last section, this is ultimately about the quality of financial market institutions and varies even within separate groups of emerging or developed market economies. I argue that investors are more sensitive to executive approval in relatively less developed financial markets. First, in these contexts, implementing public policies to the detriment of financial markets impacts a narrower segment of the electorate and is less likely to produce wide-ranging electoral repercussions. Second, less developed financial market institutions are less capable of weathering the effect of policymaking on the market. Third, less developed financial market institutions reflect lower capital intensity and provide fewer investment alternatives; this exacerbates investor fears of the repercussions of political risk in markets.

In theory, market investors "punish" less-favorable governments by engaging in less trading (i.e., lower trading volume) and diminished speculative behavior[3]. This deflates capital markets and depresses economic activity. Lower trading volume may reflect a less heterogeneous pool of investors mostly comprised of those with the highest risk appetites. Given scholars' findings of a strong positive correlation between trading volume and volatility in stock markets (Gallant, Rossi and Tauchen, 1992; Kothari and Shanken, 1992; Leblang and Mukherjee, 2005), lower volatility ensues. On the other hand, market investors "reward" more-favorable governments by engaging in more trading (i.e., higher trading volume) and increased speculative behavior. This inflates capital markets and catalyzes economic activity. Higher trading volume may reflect a more heterogeneous pool of investors, decreasing the investor pool's average risk appetite. Hence, higher volatility ensues.

It should be noted that the positive correlation between trading volume and stock market volatility, corroborated with empirical evidence in the financial literature, suggests that buoyant markets are more volatile than moribund markets. Specifically, buoyant markets—those with a positive outlook for average returns—are more attractive to larger pools of investors who in turn are more

[3] I define "speculative behavior" as high risk-high reward behavior. Speculative markets likely have greater availability of capital and more potential for *favorable* macroeconomic outcomes.

willing to take greater risks. These markets thus experience higher volatility. Conversely, moribund markets—those with a negative outlook for average returns—are less attractive to larger pools of investors who in turn are willing to take fewer risks. These markets thus experience lower volatility. Although potentially counterintuitive, this implies that markets "punish" less-favorable governments (which provide a less positive outlook for average returns, given their policies) with lower volatility and "reward" more-favorable governments (which provide a more positive outlook for average returns, given their policies) with higher volatility.

In emerging markets in particular, expectations about political phenomena may also impact trading volume and stock volatility without necessarily affecting average stock prices. This is particularly the case in countries with relatively lower financial institutional development. In these countries, stock market investment is less ubiquitous and almost entirely in the hands of corporate investors. These investors are more likely to use tiered investing structures (i.e., selling or buying a large number of shares gradually) to avoid influencing market prices to their own detriment or to avoid greater tax liabilities (Gurley-Calvez et al., 2009).

Following the political economy literature, I argue that investors prefer right-wing policies over left-wing policies, given the right's representation of capital interests compared to the left's representation of labor interests. Specifically, investors perceive right-wing policies as more capital friendly and fiscally disciplined than left-wing policies. As such, investors expect right governments to pursue policies that maximize financial market returns relative to left governments. Hence, they "punish" left-wing governments and "reward" right-wing governments.

I now turn to Figure 3.1 to summarize the theoretical argument developed above. I use Arrow (I) to demonstrate the connection between executive approval and stock market volatility. This connection stems from the role of incumbent popularity as a source of political capital which, in

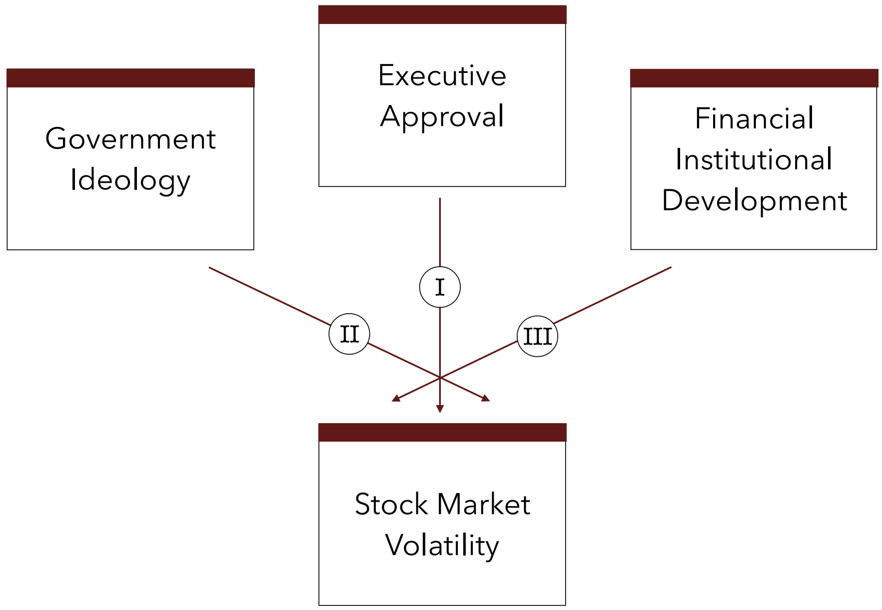


Figure 3.1: Theorized role of executive approval, government ideology, and financial institutional development as determinants of stock market volatility. (Solid arrows represent separate causal effects that moderate one another.)

turn, affects the incumbent's incentives to pursue her policy agenda. As a result, expectations about the incumbent's probability of pursuing a specific policy agenda affects stock market investors' speculative behavior and, ultimately, stock market volatility.

I use Arrow (II) to represent the effect of government ideology on stock market volatility. This effect occurs because stock market investors have different expectations about left-wing and right-wing policies. Specifically, investors expect right-wing policymakers, relative to left-wing policymakers, to pursue capital-friendly policies that will maximize stock market returns. Thus, government ideology affects stock market investors' willingness to engage in speculative behavior and, hence, stock market volatility. Importantly, the intersection of Arrows (I) and (II) in Figure 3.1 represents the moderating effect of executive approval and government ideology on the effect of one another on stock market volatility.

I use Arrow (III) to demonstrate that the effects of executive approval and government ideology on stock market volatility are further moderated by financial institutional development. Financial institutional development influences investors' willingness to engage in speculative trade given financial institutions' ability to weather the effects of politics and political risk on financial markets.

Hypotheses

Based on the framework outlined above, I put forth the following hypotheses about the effect of executive approval on stock market volatility:

H_1: Under left-wing governments, relative to center and right-wing governments, an increase in executive approval leads to relatively lower stock volatility.

H_2: Under right-wing governments, relative to center and left-wing governments, an increase in executive approval leads to relatively higher stock volatility.

H_3: Regardless of government ideology, an increase in executive approval has a weaker effect on stock volatility under relatively more developed financial market institutions than under relatively less developed financial market institutions.

3.4 Research Design and Modeling Strategy

Stock exchanges list stocks from potentially thousands of firms that vary in share price, price variance, trading volume, trading history, price-to-equity ratio, liquidity, profitability, revenue, and market capitalization, among other dimensions. Given the idiosyncrasies of individual firms and the different mix of firms in different stock exchanges, it is difficult to compare across countries' stock exchanges as a whole. Instead, scholars regularly use a stock exchange's benchmark index—usually a summary measure that is representative of listed stocks—and therefore reflect the overall health and performance of the stock market.

While summary measures of stock markets are available for a wide range of stock exchanges around the world, these indices are hardly comparable. They amalgamate "leading" stocks listed on a stock exchange based on various criteria and consider factors such as a constituent firm's market capitalization, trading history, and liquidity. Different indices will also use these and other criteria to determine the proportional contribution of a constituent firm's stock price to the index as a whole. Furthermore, indices are often calculated based on a country's local currency and are, in turn, affected by currency exchange rates and purchasing power. While daily stock market data are widely reported by the mainstream news media — including those in emerging markets — the same cannot be said of other political and macroeconomic data relevant for this analysis.

To test the implications of the theory developed in the last section, one would ideally restrict empirical analyses to a set of democracies with relatively similar government structures (e.g., parliamentary/presidential, unicameral/bicameral legislature), electoral systems (e.g., proportional/plurality), and market dynamics (e.g., state/private ownership of key economic sectors, market classification[4], currency-exchange regime). Importantly, the sample of countries included in the analysis should comprise competitive electoral democracies (as defined in Alvarez et al., 1996) with recorded alternation of power between ideologically distinct political parties and sufficient variation in incumbent popularity. In addition, these countries should possess a major stock exchange with daily variation in stock trading across broad economic sectors. While many countries meet these conditions, limitations in the comparability of public opinion data restrict my sample to a set of Latin American countries. Given limited comparability across stock exchanges, I limit my analysis to four country-cases: Argentina, Brazil, Chile, and Mexico.

Although a broader cross-country analysis with comparable stock market indices would pro-

[4]By market classification I refer to "developed," "emerging," and "frontier" markets.

vide greater statistical power, these four countries provide a unique opportunity to test the hypotheses outlined in the last section. As I show in Table 3.1, between the 1990s and 2016, all four countries were full presidential systems and at least "partially free," as classified by Freedom House. In such settings, investors interpret the role of executive approval on stocks in relatively comparable ways. While Chile and Mexico[5] have been OECD members since 2010 and 1994, respectively, both are classified as emerging markets by MSCI, Inc. — a leading American finance company providing stock market indices and classifications — along with Argentina and Brazil. Importantly, the World Bank's (2019) "Global Financial Development Database" classifies Argentina, Brazil, and Mexico as upper-middle-income economies — with relatively less developed financial institutions than Chile, a high-income economy.

Table 3.1: Sample of countries

	Full Presidential	Democracy Status (Freedom House)	OECD Member	Emerging Market (MSCI)	World Bank Income
Country					
Argentina	Yes	Always Free	No	Yes	Upper-Middle
Brazil	Yes	Always Free	No	Yes	Upper-Middle
Chile	Yes	Always Free	Yes	Yes	High
Mexico	Yes	Sometimes Free	Yes	Yes	Upper-Middle

Note: OECD membership and emerging market classifications are in reference to the full period or any subset of the period between 1991-2016.

In light of the data limitations discussed above, I estimate a separate model for each of the four country-cases and then compare results with those of Chile, given different levels of financial institutional development. According to Franzese and Kam (2009, 103-111), this estimates a

[5]It should be noted that while Mexico is an OECD member, since its accession to the bloc it has performed at or near the bottom in a variety of socioeconomic indicators, such as per capita income and the Gini coefficient, compared to other OECD members.

"separate-sample" interaction (as opposed to a "pooled-sample" multiplicative interaction). Note that I still include a conventional multiplicative interaction between executive approval and government ideology, thus avoiding a mathematically complex three-way interaction or the assumptions traditionally imposed on pooled time-series models.

Scholars interested in modeling volatility over time often employ some variation of an autoregressive conditional heteroskedasticity (ARCH) model (e.g. Engle, 1982; Maestas and Preuhs, 2000; Leblang and Mukherjee, 2005; Jensen and Schmith, 2005; Leblang and Bernhard, 2006; Füss and Bechtel, 2008; Benton and Philips, 2020). This strategy allows researchers to relax the Markov assumption of constant variance in the conditional mean equation by specifying an additional equation to model the conditional variance of a series. This is convenient because researchers can still specify a time-series model of choice, such as an autoregressive distributed lag (ARDL) model, to estimate the conditional mean. The choice to specify an additional equation for the conditional variance amends the assumption of constant variance in the error term in the original equation.

In this analysis, the conditional mean equation is specified as an ARDL as follows:[6]

$$\Delta \mathrm{y}_t = \alpha + \phi \Delta y_{t-1} + \boldsymbol{\beta} \mathbf{K_{t-1}} + \varepsilon_t \tag{3.1}$$

where Δy_t is the time-differenced stock market index, α is the y-intercept, ϕ is the parameter effect on the lagged dependent variable, $\boldsymbol{\beta}$ is a vector containing parameter effects, $\mathbf{K_{t-1}}$ is a vector containing the explanatory variables, and ε_t is an error term. Importantly, the error term has non-constant variance.

[6]I ran a series of unit root tests and calculated time differences when appropriate. Importantly, the dependent variables across all four country cases contain unit roots.

To account for and model heteroskedasticity (and, hence, volatility), I also estimate a conditional variance equation of the type ARCH(1) as follows:

$$\sigma_t^2 = \omega\varepsilon_{t-1}^2 + \exp(\gamma x_{t-1} + \eta z_{t-1} + \lambda x_{t-1}z_{t-1} + \boldsymbol{\rho}\mathbf{m_{t-1}}) \tag{3.2}$$

where the variance σ_t^2, is a function of the lagged error term ε_{t-1}^2, as well as lagged approval x_{t-1}, lagged ideology z_{t-1}, the interaction between approval and ideology $x_{t-1}z_{t-1}$, and a vector of control variables $\mathbf{m_{t-1}}$. In addition, ω, γ, η, λ, and $\boldsymbol{\rho}$ represent parameter effects of these variables on the variance. Both the conditional mean and the conditional variance equations are estimated using maximum likelihood estimation.

3.5 Data

Dependent Variable

Argentina

Argentina's flagship stock market index is the S&P MERVAL, comprised of stocks in the S&P/BYMA Argentina General Index listed at the Buenos Aires Stock Exchange. It includes stocks traded in at least 95% of sessions for firms with market capitalization superior to 2.5 billion Argentine pesos (approximately 25 million US dollars as of December 2021). The index is calculated daily (excepting non-trading days such as weekends and holidays), though only the last observation per month is included in the analysis due to limitations in the frequency of the independent variables. Between 1991 and 2016, the index varies between a low of 202.45 points in November of 2001 and a high of 13724.07 points in April of 2016, the last observation included in the analysis. The index is recalibrated and reweighted twice a year, in February and August.

Brazil

The Bovespa Index is the primary indicator of the São Paulo Stock Exchange — Latin America's largest. As of November 2021, the index includes 88 individual stocks traded in Brazilian reais and is representative of various economic sectors. The index represents a total market capitalization of almost 700 billion US dollars, comprised of stocks with the highest tradability ratios traded in at least 95% of sessions. The index is calculated daily (excepting non-trading days such as weekends and holidays), though only the last observation per month is included in the analysis due to limitations in the frequency of the independent variables. Between 1998 and 2016, the index varies between a low of 0.0001 point in May 1998 and a high of 72592.5 points in May 2008. The index is recalibrated and reweighted every four months, in January, May, and September.

Chile

The S&P/CLXA Index is a benchmark indicator for the leading stocks listed on the Santiago Stock Exchange. All high-liquidity stocks listed on the exchange, excepting some pension funds, are eligible. To be included in the index, stocks must be traded in excess of 300 thousand US dollars annually. Its 61 constituent stocks are representative of various economic sectors, though more than half of its capitalization corresponds to the materials and financial sectors. As was the case for the S&P MERVAL and the Bovespa, this index is calculated daily (excepting non-trading days such as weekends and holidays), though only the last observation per month is included in the analysis due to limitations in the frequency of the independent variables. Between 1993 and 2016, the index varies between a low of 3200.2 points in October 1993 and a high of 22979.22 points in December 2010. The index is recalibrated and reweighted annually in March.

Mexico

The S&P/BMV IPC Index is a benchmark indicator for the leading stocks listed on the Bolsa

Mexicana de Valores. All high-liquidity stocks listed on the Exchange, excepting real estate investment trusts and mortgage trusts, are eligible. In addition, constituent stocks must have been traded in at least 95% of trading sessions over a period of six months with a median daily trading value of 1.4 million US dollars as of December 2021. Its 35 constituent stocks are representative of various economic sectors, though more than half of its capitalization corresponds to the consumer goods and materials sectors. As was the case for the indices described above, this index is calculated daily (excepting non-trading days such as weekends and holidays), though only the last observation per month is included in the analysis due to limitations in the frequency of the independent variables. Between 1989 and 2016, the index varies between a low of 261.7 points in April 1989 and a high of 45881.08 in March 2016. The index is recalibrated and reweighted in March, September, and December.

Independent Variables

The main independent variable is Carlin et al.'s (2019) monthly "Executive Approval Project" data. This dataset incorporates publicly available public opinion data concerning the chief executive's job. While annual and quarterly data are available for a broader sample of countries, monthly data are exclusively available for a select subset of Latin American countries. Values represent the percentage of the population approving of the president's job.

To measure government ideology, I use the World Bank's "Database of Political Institutions"' indicator of whether the incumbent government is left, center, or right. In the analysis, this variable takes a value of 0 under left governments, 1 under center governments, and 2 under right governments.

I control for inflation using the consumer price index produced by each country's main statistical agency and reported by Trading Economics. I also include an elections dummy to account for months when a national-level presidential or congressional election takes place. In this analysis, this variable takes a value of 0 in months when no election takes place and 1 in months when an election takes place.

3.6 Results

In this section, I first report individual results for my country-cases with relatively less developed financial institutions: Argentina, Brazil, and Mexico. These results speak to hypotheses H_1 (under left-wing governments, relative to center and right-wing governments, an increase in executive approval leads to relatively lower stock volatility) and H_2 (under right-wing governments, relative to center and left-wing governments, an increase in executive approval leads to relatively higher stock volatility). I then turn to my relatively more developed financial market: Chile. Finally, I summarize results contrasting those of the first three country-cases with those of Chile, evaluating hypothesis H_3 (regardless of government ideology, an increase in executive approval has a weaker effect on stock volatility under relatively more developed financial market institutions than under relatively less developed financial market institutions.)

Argentina

In Table 3.2, I report regression results for Argentina. In this Table, rows are organized under two primary sections: one with parameter estimates recovered from the conditional mean equation and one with parameter estimates recovered from the conditional variance equation. Two models are reported. Model 1 is of the type ARCH(1), with a conditional mean equation specified according to Equation 3.1 and a conditional variance equation specified according to Equation 3.2. Model 2 is similar to model Model 1 except that its conditional variance equation omits the ARCH(1),

ε^2_{t-1}, from the right-hand side.

In Model 1, none of the parameter estimates in the conditional mean equation are statistically significant at conventional 95% confidence levels. In the conditional variance equation, however, I recover statically significant parameters on lagged ideology, the ARCH(1) term, and the constant. The statistically significant parameter estimate on the ARCH(1) term suggests that the variance is a function of the previous term's stochastic shocks. Its magnitude, however, is exceptionally high (greater than 1). Ideally, one would explore this further by adding additional ARCH terms (e.g., ARCH(2), ARCH(3)) to the conditional variance equation. However, when doing so the model fails to converge via maximum likelihood estimation.

Table 3.2: Δ S&P MERVAL (Buenos Aires) regression results.

	(Model 1)		(Model 2)	
Mean Equation				
Δ Equities$_{t-1}$	-0.00160	(0.0130)	0.0102	(0.0431)
Approval $_{t-1}$	0.828	(0.555)	0.529	(1.111)
CPI$_{t-1}$	0.915	(0.587)	0.873	(1.636)
Ideology$_{t-1}$	-10.75	(10.33)	-12.46	(17.30)
Election$_{t-1}$	-8.317	(17.49)	-4.895	(38.47)
Constant	-1.294	(40.53)	19.80	(64.46)
Variance Equation				
Approval $_{t-1}$	-0.00395	(0.0280)	-0.00443	(0.0101)
CPI$_{t-1}$	0.0216	(0.0115)	0.0607***	(0.00595)
Ideology$_{t-1}$	-1.419*	(0.604)	-2.189***	(0.230)
Election$_{t-1}$	-1.633	(1.226)	-1.047*	(0.433)
App.$_{t-1}$ $\times$ Ideol.$_{t-1}$	0.00740	(0.0130)	0.0227***	(0.00427)
Constant	11.26***	(1.528)	12.47***	(0.602)
ARCH(1)	1.129***	(0.183)		
Log lik.	-1774.9		-1842.6	
Observations	280		280	

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

As in Model 1, in Model 2 none of the parameter estimates in the conditional mean equation are statistically significant at conventional 95% confidence levels. However, most parameter estimates in the conditional variance equation are statistically significant. To interpret theoretically relevant results using these parameters (and following the recommendations of Clark, Gilligan and Golder 2006, Franzese and Kam 2009, and Hainmueller, Mummolo and Xu 2019), I now turn to Figure 3.2.

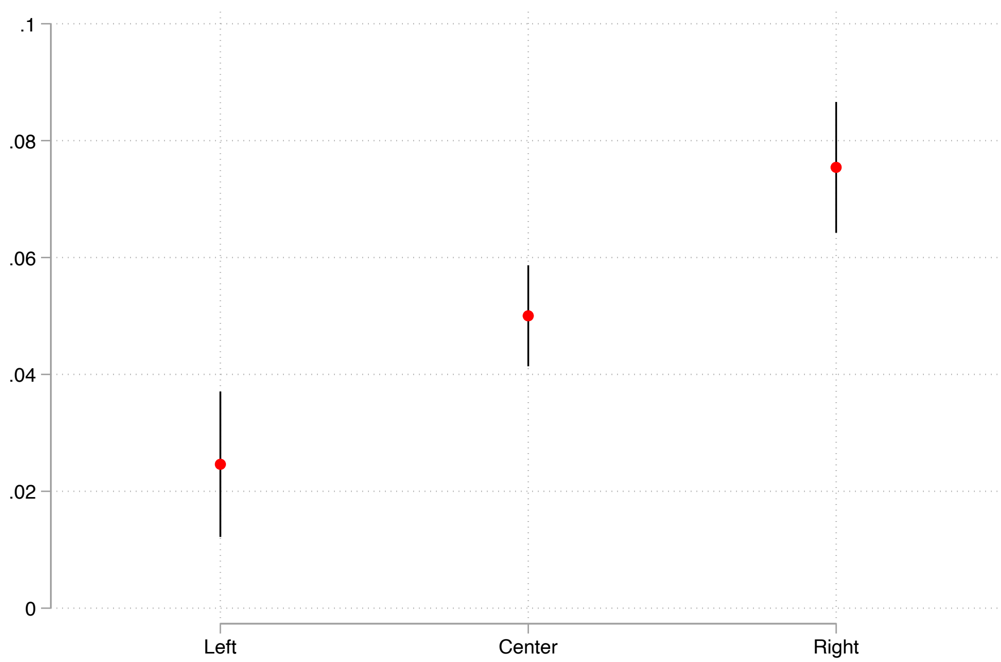


Figure 3.2: Estimated marginal effects of a one-percentage increase in executive approval on the variance of the S&P MERVAL Index across distinct government ideologies (95% CI).

In Figure 3.2, I show the estimated marginal effects of a one-percentage increase in executive approval on the variance of Argentina's S&P MERVAL Index across distinct government ideologies (left, center, and right governments are observed in Argentina in 1991-2016). These results support hypotheses H_1 and H_2, and suggest that stock market investors are more likely to "punish" left-wing governments, relative to center or right-wing governments, in Argentina. Similarly, stock market investors are more likely to “reward” right-wing governments, relative to center or left-wing governments, in Argentina. As evidenced by the positive and statistically significant effect across distinct government ideologies, higher executive approval is always associated with higher stock volatility in Argentina. However, an increase in approval under center governments raises stock volatility by a greater and statistically distinguishable magnitude than an increase in approval under left governments. In addition, an increase in approval under right governments raises stock volatility by a greater and statistically distinguishable magnitude than an increase in approval under center governments. This is evidence that markets engage in higher trading volumes and induce higher stock volatility in light of an increase in popularity of right-wing governments than in light of an increase in popularity of center or left-wing governments. Similarly, markets engage in lower trading volumes and induce lower stock volatility in light of an increase in popularity of left-wing governments than in light of an increase in popularity of center or right-wing governments.

Brazil

In Table 3.3, I report results from a single Model 1 that includes both a conditional mean equation and a conditional variance equation for Brazil. In the conditional mean equation, only the lagged dependent variable is statistically significant. Unfortunately, the conditional variance equation in Model 1 does not include ARCH terms because such specifications failed to converge via maximum likelihood estimation.

In the conditional variance equation in Model 1, all parameter estimates are statistically significant at conventional 95% confidence levels. To interpret theoretically relevant results using these parameters, I now turn to Figure 3.3.

Table 3.3: Δ iBovespa (São Paulo) regression results.

	(Model 1)	
Mean Equation		
Δ Equities$_{t-1}$	0.249***	(0.0485)
Approval $_{t-1}$	0.00318	(0.561)
CPI$_{t-1}$	-0.00000177	(0.00548)
Ideology$_{t-1}$	-62.42	(80.70)
Election$_{t-1}$	133.0	(763.7)
Constant	187.2	(243.4)
Variance Equation		
Approval $_{t-1}$	-0.228***	(0.0142)
CPI$_{t-1}$	0.00745***	(0.000283)
Ideology$_{t-1}$	-18.60***	(0.546)
Election$_{t-1}$	-1.232	(0.712)
App$_{t-1}$ $\times$ Ideol$_{t-1}$	0.249***	(0.0133)
Constant	33.16***	(0.609)
Log lik.	-2558.5	
Observations	299	

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

In Figure 3.3, I show the marginal effects of a one-percentage increase in executive approval on the variance of Brazil's Bovespa Index across distinct government ideologies (only left-wing and right-wing governments are observed in Brazil in 1988-2016). These results support hypotheses H_1 and H_2, and suggest that stock market investors are more likely to "punish" left-wing governments than right-wing governments in Brazil. As evidenced by the positive and statistically significant effect across distinct government ideologies, higher executive approval is always associated with higher stock volatility in Brazil. However, an increase in approval under right-wing governments raises stock volatility by a greater and statistically distinguishable magnitude than an increase in approval under left-wing governments. This is evidence that markets engage in higher trading volumes and induce higher stock volatility in light of an increase in popularity of right-wing governments than in light of an increase in popularity of left-wing governments.

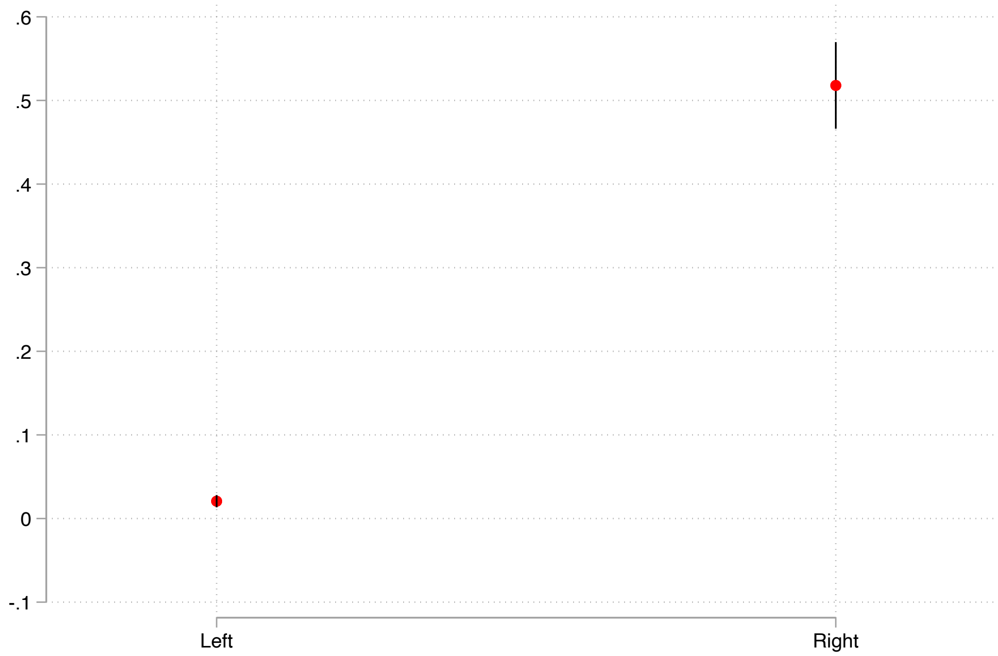


Figure 3.3: Estimated marginal effects of a one-percentage increase in executive approval on the variance of the iBovespa Index across distinct government ideologies (95% CI).

Mexico

In Table 3.4, I report results from a single Model 1 that includes both a conditional mean equation and a conditional variance equation for Mexico. In the conditional mean equation, none of the parameter estimates is statistically significant at the 95% confidence level. Unfortunately as was the case with Brazilian data, the conditional variance equation in Model 1 does not include ARCH terms because such specifications failed to converge via maximum likelihood estimation.

In the conditional mean equation in Model 1, all parameter estimates are statistically significant at conventional 95% confidence levels. To interpret theoretically relevant results using these parameters, I now turn to Figure 3.4.

In Figure 3.4, I show the marginal effects of a one-percentage increase in executive approval

Table 3.4: Δ S&P/BMV IPC (Mexico City) regression results.

	(Model 1)	
eq		
Δ Equities$_{t\text{-}1}$	-0.0213	(0.0462)
Approval $_{t\text{-}1}$	-0.192	(3.032)
CPI$_{t\text{-}1}$	0.784	(2.287)
Ideology$_{t\text{-}1}$	231.6*	(111.2)
Election$_{t\text{-}1}$	-74.15	(185.3)
Constant	-410.3	(386.3)
HET		
Approval $_{t\text{-}1}$	-0.535***	(0.0707)
CPI$_{t\text{-}1}$	-0.115***	(0.00987)
Ideology$_{t\text{-}1}$	-11.37***	(1.966)
Election$_{t\text{-}1}$	-0.942	(0.688)
App.$_{t\text{-}1}$ $\times$ Ideol.$_{t\text{-}1}$	0.211***	(0.0324)
Constant	42.89***	(4.356)
Log lik.	-2541.6	
Observations	325	

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

on the variance of Mexico's S&P/BMV IPC Index across distinct government ideologies (only center and right-wing governments are observed in 1989-2016). Overall, these results support hypothesis H_2, and suggest that stock market investors are more likely to "reward" right-wing governments relative to center governments in Mexico. Furthermore, it appears that in Mexico an increase in executive approval is associated with a decrease in stock market volatility under center governments, but associated with an increase in stock market volatility under right-wing governments. While these results stand out from those of Argentina and Brazil, it is still the case here that an increase in approval under right-wing governments affects stock volatility by a greater and statistically distinguishable magnitude than an increase in approval under center governments. This is evidence that markets engage in higher trading volumes and induce higher stock volatility in light of an increase in popularity of right-wing governments than in light of an increase in

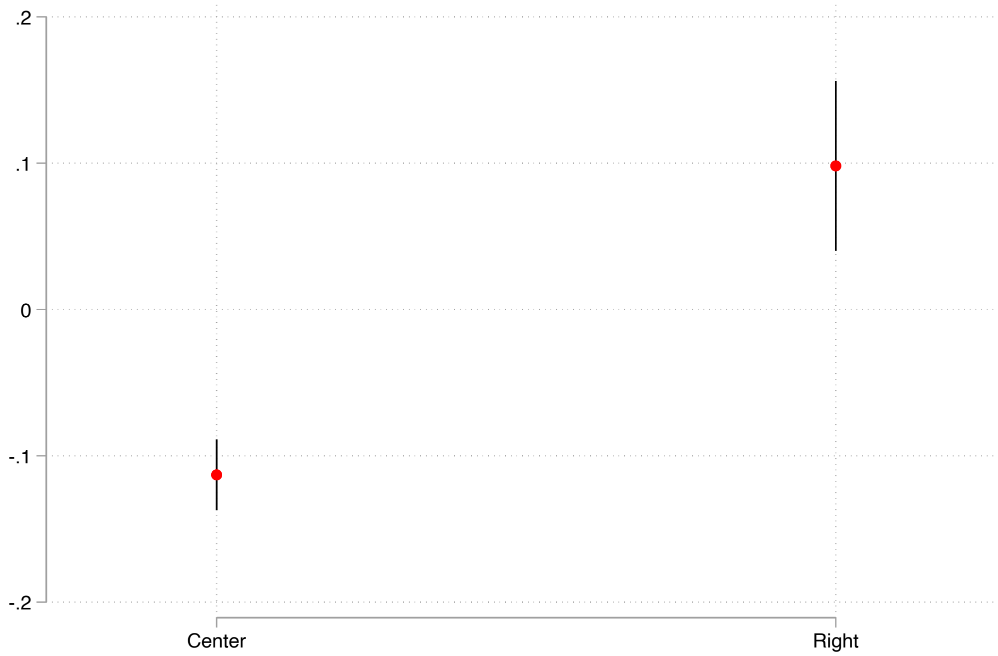


Figure 3.4: Estimated marginal effects of a one-percentage increase in executive approval on the variance of the S&P/BMV IPC Index across distinct government ideologies (95% CI).

popularity of center governments.

Chile

In Table 3.5, I report results from Chile, the one country-case with relatively more developed financial institutions. Two models are reported. Model 1 is of the type ARCH(1). Model 2 is similar to Model 1 except that its conditional variance equation omits the ARCH(1), ε^2_{t-1}, from the right-hand side.

In the conditional mean question in Model 1, only parameter estimates for lagged approval and lagged inflation are statistically significant at the 95% confidence level, suggesting that both of these variables affect changes in stock prices. However, in the conditional variance equation, only the constant is statistically significant.

Table 3.5: Δ IGPA (Santiago) regression results.

	(Model 1)		(Model 2)	
Mean Equation				
Δ Equities$_{t-1}$	0.132	(0.0705)	0.135**	(0.0519)
Approval $_{t-1}$	4.028*	(2.005)	5.922*	(2.301)
CPI$_{t-1}$	-16.99*	(6.664)	-20.72**	(6.982)
Ideology$_{t-1}$	-51.25	(133.6)	-11.92	(130.3)
Election$_{t-1}$	57.99	(247.5)	36.31	(237.5)
Constant	-5.368	(200.1)	-111.4	(196.3)
Variance Equation				
Approval $_{t-1}$	-0.00776	(0.0893)	0.00453	(0.0608)
CPI$_{t-1}$	0.0295	(0.0425)	-0.0529	(0.0283)
Ideology$_{t-1}$	1.989	(2.916)	1.849	(1.926)
Election$_{t-1}$	0.618	(1.594)	0.346	(1.268)
App.$_{t-1}$ $\times$ Ideol.$_{t-1}$	-0.0219	(0.0876)	-0.0204	(0.0593)
Constant	10.73***	(3.062)	11.07***	(2.068)
ARCH				
L.arch	0.391***	(0.0893)		
Log lik.	-2014.7		-2027.2	
Observations	271		271	

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

In the conditional mean question in Model 2, parameter estimates for the lagged dependent variable, lagged approval, and lagged inflation are statistically significant. But as with Model 1, in the conditional variance equation, only the constant is statistically significant.

In Figure 3.5, I show the marginal effects of a one-percentage increase in executive approval on the variance of Chile's S&P/IGPA Index across distinct government ideologies (only left-wing and center governments are observed in Chile in 1994-2016). These results are not consistent with hypotheses H_1 and H_2. As shown, the effect of a one-percentage increase in executive approval on the variance of the S&P/IGPA Index is indistinguishable from zero and indistinguishable across government ideologies at the 95% confidence level.

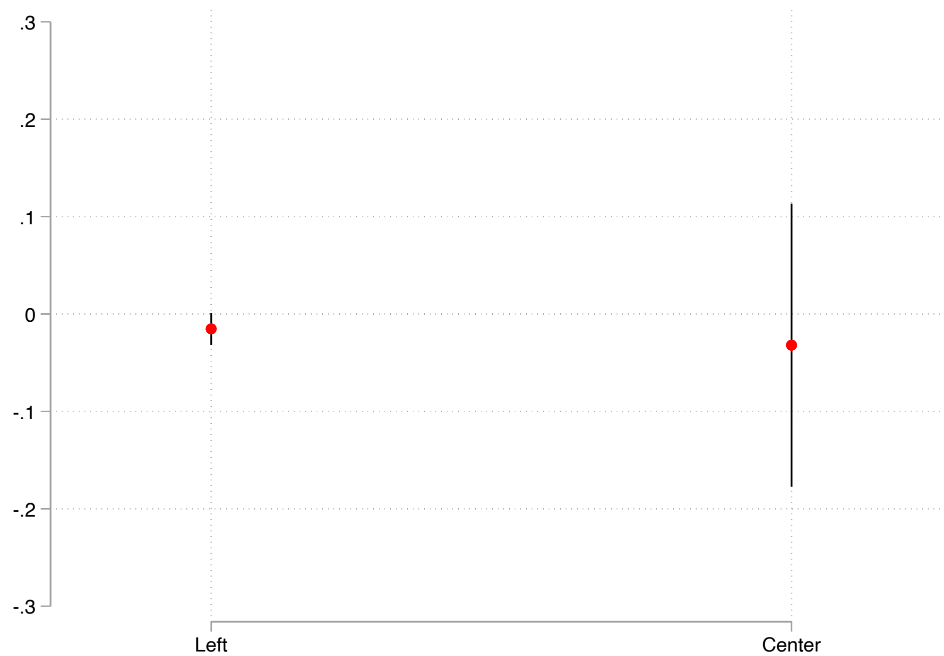


Figure 3.5: Estimated marginal effects of a one-percentage increase in executive approval on the variance of the S&P/IGPA Index across distinct government ideologies (95% CI).

Synthesis of Results

Overall, results suggest that stock market investors "punish" left-wing governments, relative to center or right-wing governments. Specifically, increases in the approval of right-wing governments lead to higher stock volatility relative to increases in the approval of center or left-wing governments. Similarly, increases in the approval of left-wing governments lead to lower stock volatility relative to increases in the approval of center or right-wing governments. Hence, I find support for hypotheses H_1 and H_2 in Argentina, Brazil, and Mexico.

The caveat is that these relationships are conditional on a country's level of financial institutional development. Specifically, given Chile's "advantage" as a relatively less risky financial market — with relatively higher capital intensity (Caselli and Feyrer, 2007) — and relatively less speculation compared to the other three country cases, my results suggest support for hypothesis H_3.

3.7 Conclusion

In this article, I propose a theory that mass political opinion affects stock market volatility. Specifically, I argue that more popular governments are more likely to implement their preferred policies. This occurs because they have more political capital and are, thus, emboldened in relation to other government institutions. Hence, pursuing and implementing a particular policy agenda is more *affordable* for more popular governments. Stock market investors are aware of these dynamics and respond accordingly.

Because markets expect right-wing governments, relative to center or left-wing governments, to pursue more capital-friendly policies that are likely to maximize investment returns, they engage in more trading and speculative behavior under right-wing governments. Higher volatility ensues. Conversely, they engage in less trading and speculative behavior under left-wing governments, relative to center and right-wing governments. Lower volatility ensues. I further argue that these effects are more likely to hold in contexts of relatively less developed financial institutions compared to contexts of more developed financial institutions.

Overall, I find support for the theory I propose with the broad implication that stock markets "punish" left-wing governments by engaging in less stock trading activity (thus depressing capital availability and potentially depressing macroeconomic performance under left-wing governments). Conversely, stock markets "reward" right-wing governments by engaging in more stock trading activity (thus inflating capital availability and potentially catalyzing macroeconomic performance under right-wing governments). These findings are consistent with previous research suggesting that markets punish left-wing election victories (e.g. Fowler, 2006; Sattler, 2013) and contributes to an understanding of the political economy of stock market volatility beyond elections.

This piece can be extended to consider a larger and more diverse cross-country sample of countries—specifically, a sample that includes both developed and emerging markets. With more expansive data, researchers may also consider alternative measures of volatility and potentially new ways to capture stock market health and performance. In addition, future research should consider additional asset types (e.g., currency-exchange markets, commodity markets), and other measures of mass politics (e.g., social media, corporate news media).

4. CONDITIONAL RELATIONSHIPS IN ERROR CORRECTION MODELS

4.1 Introduction

Conditional relationships are prominent in political science (Franzese and Kam, 2009). They reflect complex phenomena where the effect of an explanatory variable on a dependent variable varies across values of additional moderating variables. Over time, scholars have developed useful tools and recommendations to (1) model these types of theories empirically, (2) adjudicate between alternative modeling strategies, and (3) intuitively present and interpret results (Brambor, Clark and Golder, 2006; Berry, Golder and Milton, 2012; Hainmueller, Mummolo and Xu, 2019). Applying these tools and recommendations to time-series analysis is generally straightforward. One exception to this occurs in the context of error correction models (ECMs).

ECMs are useful when researchers are interested in estimating the speed at which a non-stationary dependent variable returns to equilibrium in response to shocks to an explanatory variable (Engle and Granger, 1987). Recently, these models have grown in popularity in the discipline and have been the topic of substantial debate among political methodologists (De Boef and Keele, 2008; Keele, Linn and Webb, 2016; Grant and Lebo, 2016; Webb, Linn and Lebo, 2020). Interactions within ECMs have also grown in popularity (e.g. Morgan and Kelly, 2013; Enns et al., 2014; Soroka, Stecula and Wlezien, 2015; Herwartz and Theilen, 2017; Ezrow, Hellwig and Fenzl, 2020). But it can be challenging to represent conditional theories in ECMs; it is not always clear which monomials within an ECM best represent the theoretical relationships the researcher is interested in. Unfortunately, researchers are left without a straightforward overview of the implications of alternative specifications of multiplicative interactions within these complex and relatively inflexible models.

In this paper, I present an exposition of four alternative specification strategies one might contemplate when testing conditional theories in the context of ECMs. First, I consider interactions between a moderating variable and a time-differenced explanatory variable. Second, I consider interactions between a moderating variable and a time-lagged explanatory variable. Third, I consider interactions between a moderating variable and the lagged dependent variable. Finally, I consider interactions between a moderating variable and the equilibrium component of the model—the lagged components in the RHS.

In the sections that follow, I begin with a discussion of the extant literature on multiplicative interactions and ECMs. I then consider four plausible rival specification strategies one might consider when testing conditional relationships in the context of ECMs. After that, I present a series of Monte Carlo experiments to evaluate the bias and efficiency trade-offs across these rival specification strategies. I illustrate my recommendations with an empirical application that examines whether media content influences the public's economic sentiment in the United States. Finally, I discuss my findings and conclude with suggestions for future research.

4.2 A Brief Overview of Multiplicative Interactions and ECMs

Political scientists are often interested in phenomena where the effect of an independent variable on a dependent variable is conditional on values of a third variable. The complex and multivariate nature of politics makes it fertile ground for work that considers these types of relationships (Brambor, Clark and Golder, 2006; Franzese and Kam, 2009). Many seminal works in the discipline gained prominence with theories that accounted for conditional relationships scholars had previously ignored (e.g. Powell and Whitten, 1993; Burkhart and Lewis-Beck, 1994; O'Toole Jr and Meier, 1999; Fearon and Laitin, 2003; Taber and Lodge, 2006).

Over the years, the complexity of modeling conditional relationships has fueled vigorous debates on the appropriate strategies to do so (Wright Jr, 1976; Friedrich, 1982; Aiken, West and Reno, 1991). Many of these debates were settled by the canonical recommendations of Brambor, Clark and Golder (2006) and Franzese and Kam (2009), which provide general guidelines for the use of multiplicative interactions in linear models testing conditional relationships of the type:

$$y = \theta_0 + \theta_1 x + \theta_2 z + \theta_3 (x \cdot z) + \varepsilon \tag{4.1}$$

where $\varepsilon \sim N(0, \sigma^2)$. Equation 4.1 implies that the effect of x on y is composed of θ_1 and $\theta_3 z$, which is conditional on values of z. Thus, $\partial y / \partial x = \theta_1 + \theta_3 z$. It is important to keep in mind that z has a symmetric effect on y conditional on x, such that $\partial y / \partial z = \theta_2 + \theta_3 x$ (Berry, Golder and Milton, 2012). Brambor, Clark and Golder (2006) emphasize the need to model the individual constituent terms of the interaction ($\theta_1 x$ and $\theta_2 z$ in Equation 4.1), and to always interpret the conditional marginal effects of interacted variables (as opposed to their individual parameter estimates).

More recently, Hainmueller, Mummolo and Xu (2019) expand these recommendations by stressing the need to diagnose the presence of (1) non-linear interaction effects, and (2) common support to warrant the estimation of marginal effects across particular values of the variables included in multiplicative interactions. They also expand on the complexities of modeling interactions composed of at least one continuous variable, contending that the inclusion of continuous variables in interactions makes them especially vulnerable to misspecification bias and model dependency.

The recommendations above are commonly understood to apply to time-series modeling as well. It is ordinary practice in the time-series literature (and the time-series-cross-sectional literature; see Williams and Whitten 2012; Jung et al. 2020) to include the constituent terms of any interactions modeled, to consider the range of values providing common support, and to report marginal effects of explanatory variables of interest (e.g. Copelovitch, 2010; Whitten and Williams, 2011; Lowande, 2019; Compton and Lipsmeyer, 2019; Abott and Magazinnik, 2020; Johnson and Strother, 2020). Researchers have considerable latitude to interact these lags and differences based on theory and statistical fit. Any time lags and time differences of the same variable are handled as individual variables. Recent work has sought to establish diagnostic and specification guidelines specific to time-series models that include interaction terms (Warner, Working Paper).

Although including interactions in some of the most popular time-series models is relatively straightforward, it becomes more complicated when dealing with error correction models (ECMs). ECMs are the preferred modeling choice when cointegration is present (Engle and Granger, 1987).[1] They allow researchers to (1) control for persistence in left-hand-side (LHS) and right-hand-side (RHS) variables, (2) account for the shared stochastic trend between them (Davidson et al., 1978; Stock and Watson, 1988; Durr, 1992), and (3) estimate the error correction rate in nonstationary series[2]. The trade-off is that ECMs are relatively inflexible and econometrically complex. In the simplest case where both dependent and independent variables are each integrated of order 1, I(1), and cointegrated, a correctly specified general ECM requires all continuous variables to be lagged and differenced once:

$$\Delta y_t = \alpha_0 + \alpha_1 y_{t-1} + \alpha_2 \Delta x_t + \alpha_3 x_{t-1} + \varepsilon_t \tag{4.2}$$

[1]However, recent scholarship suggests that the presence of cointegration may not be a prerequisite for running error correction models (Keele, Linn and Webb, 2016; Webb, Linn and Lebo, 2020).

[2]The terms “error correction rate” and “rate of adjustment to equilibrium” are used interchangeably in the literature and in this piece.

where y_t is the dependent variable, x_t is an explanatory variable, α's are parameter estimates, and $\varepsilon_t \sim N(0, \sigma^2)$. Equation 4.2 also assumes that only the current- and last-period x affect y. We would denote this as an ARDL($p = 1$,$q = 1$,$r = 1$) if referring to its autoregressive-distributed-lag equivalent, where p, q, and r are the number of lags on y, the number of lags on x, and the number of exogenous regressors, respectively[3].

As is the case in other time-series models, non-continuous explanatory variables may also be included in ECMs (Hardy, 1993; Enders, 2008; Pickup, 2014). Three types of non-continuous variables are especially common: time trends, seasonal dummies, and intervention dummies.[4] Time trends are indices (e.g., 1, 2, 3, ..., T) that account for linear or quadratic time patterns in the data, which will violate covariance-stationarity assumptions. Seasonal dummies account for periodicity in the data, which will also violate covariance-stationarity assumptions. Intervention dummies account for mean-level changes in the data, often representing specific "events," which can violate mean-stationarity assumptions.

In the context of ECMs (and related models, such as vector error correction models), scholars habitually include time trends and seasonal dummies in levels (and, thus, omit any time differences; e.g. Durr, 1992; Kulendran and King, 1997). Either lagged levels or current levels are used. Yet intervention dummies are often included in both lagged and differenced forms (e.g., Faricy, 2011; Rickard, 2012; Philips, Rutherford and Whitten, 2015; Whiteley et al., 2016; Philips, Rutherford and Whitten, 2016). Unfortunately, limited guidance exists on the appropriate use of time-lagged and time-differenced interventions; few works explore the implications of lagging and differencing

[3]I demonstrate the equivalence between a general ECM (GECM) and the ARDL in Section B.1. of Appendix B

[4]Scholars oftentimes also include fixed-effects dummies to account for heterogeneity in intercepts. In the context of ECMs, fixed-effects dummies are included in current (as opposed to lagged) levels (e.g. Kang and Powell Jr, 2010). It is worth noting, however, that recent work warns of potential misspecification bias when fixed effects are included in dynamic models (Plümper and Troeger, 2019).

distinct types of interventions.

Regardless of the types of explanatory variables included, Engle and Granger (1987) show that Equation 4.2 contains an equilibrium component (composed of its time-lagged variables) and a disequilibrium component (composed of its time-differenced variables). Shocks to the equilibrium component of the model affect its disequilibrium component. The speed at which disequilibrium components return to equilibrium in response to these shocks is measured by the error correction rate.

ECMs are useful in a variety of substantive domains. Some notable applications in political science include: public opinion and party support (e.g. Durr, 1993; Clarke and Stewart, 1994; Jennings and John, 2009; Ura and Ellis, 2012; Ramirez, 2013), foreign policy and conflict (e.g. Moore and Lanoue, 2003), spending and redistribution (e.g. Iversen and Soskice, 2006; Rickard, 2012; Morgan and Kelly, 2013; Doyle, 2015), income inequality (e.g. Kelly, 2005; Philips, Souza and Whitten, Forthcoming), mass incarceration (e.g. Enns, 2014), and the political economy of financial markets (e.g. Leblang and Mukherjee, 2005), among others.

Given the widespread applications of ECMs in the discipline, political methodologists have developed work to show the contexts in which these models are appropriate and their equivalence to other time-series models. De Boef and Keele (2008) show the equivalence between ECMs and autoregressive distributed lag (ARDL) models; they argue that ECMs can be used with or without cointegration. Disagreeing, Grant and Lebo (2016) defend the usage of ECMs exclusively in the presence of cointegration, emphasizing the need to guarantee equation balance. Following Keele, Linn and Webb's (2016) findings, it is now understood that balanced ECMs can be used even if cointegration is not present. Additional work on stationarity and unit root testing has also been done (Freeman, 2016; Webb, Linn and Lebo, 2020). Some authors have further discussed the importance of equation balance (Enns and Wlezien, 2017) and suggested more comprehensive cointegration testing strategies (Philips, 2018).[5]

In spite of the great progress on the methodological issues discussed above, there has been insufficient discussion of the alternative strategies and implications of modeling multiplicative interactions in ECMs. This is in spite of the growing number of political science papers that already employ interactions within ECMs (e.g., Clarke, Ho and Stewart, 2000; Nooruddin and Simmons, 2006; Kono, 2008; Philips, Rutherford and Whitten, 2015). This is what I expand on in the next section.

4.3 Modeling Interactions Within ECMs

Theoretically interesting conditional relationships can lead scholars to include multiplicative interactions in error correction models (ECMs). As noted above, ECMs are complex and relatively inflexible, which complicates the inclusion of interaction terms. In particular, it can be challenging to identify which monomials within an ECM best represent the theoretical relationships the researcher believes are conditional on a third variable. This warrants an exposition of the alternative specification strategies one might find appropriate when testing conditional theories with ECMs. In this introductory treatment of this topic, I focus on two-way interactions since interactions composed of more than two variables are much less common in political science.

Before I present a discussion of alternative specification strategies for ECMs with interactions, it is worthwhile to consider three potential *types* of interactions one is likely to encounter: (1) interactions between continuous variables (or non-continuous variables treated as continuous[6]), (2) interactions between dummy variables, and (3) interactions between continuous and dummy

[5] Scholars have also discussed fractional integration methods and their relevance to ECMs (DeBoef and Granato, 1997; Esarey, 2016; Helgason, 2016; Keele, Linn and Webb, 2016).

[6] For example, ordinal explanatory variables are non-continuous but are commonly treated as continuous.

variables. Interactions between continuous variables are the most theoretically and statistically complex. They are likely to introduce non-linearity, lack common support for marginal effects calculations, and suffer from misspecification bias and multicollinearity (Jaccard, Wan and Turrisi, 1990; Hainmueller, Mummolo and Xu, 2019). Interactions between dummy variables are the least complex and oftentimes easier to interpret. Hainmueller, Mummolo and Xu (2019) recommend that scholars model these as fully saturated models (and not as multiplicative interactions), where a dummy variable is included for each of the potential combinations of values of the moderator and of the moderated variable. Interactions between continuous and dummy variables are still somewhat theoretically and statistically complex but are less likely to be unusually model dependent and less likely to suffer from model misspecification. In this piece, I focus on this third type of interactions.

I begin with Equation 4.3—a general error correction model where the dependent variable, y_t, and the main explanatory variable of interest, x_t, are each continuous, I(1), and cointegrated. I also include z_t, defined as a multi-period intervention dummy, in its contemporaneous form in the right-hand side (RHS) of Equation 4.3. Since z_t is a multi-period intervention and, thus, cannot be associated with a permanent intercept shift, I opt to omit a time-differenced term in the RHS of Equation 4.3.

$$\Delta y_t = \alpha_0 + \alpha_1 y_{t-1} + \alpha_2 \Delta x_t + \alpha_3 x_{t-1} + \alpha_4 z_t + \varepsilon_t \tag{4.3}$$

In each of the alternative scenarios I describe in Subsections 4.3.1–4.3.4 below, z_t is also part of the interaction terms. These can be thought of as statistical relationships moderated by a condition z set to “on” when z_t equals 1 and “off” when z_t equals 0. The insights derived from the discussion below can be extended to contexts where the moderator is time lagged (e.g., z_{t-1}) or time differenced (e.g., Δz_t).

4.3.1 Interactions between z_t and Δx_t

$$\Delta y_t = \alpha_0 + \alpha_1 y_{t-1} + \alpha_2 \Delta x_t + \alpha_3 x_{t-1} + \alpha_4 z_t + \alpha_5 \Delta x_t \times z_t + \varepsilon_t \tag{4.4}$$

In Equation 4.4, I show an ECM identical to that shown in Equation 4.3 except that it also includes an interaction term between z_t and Δx_t. Such an interaction is an appropriate representation of a theory predicting that the effect of changes to the current values of a RHS variable (Δx_t) is conditional on values of a moderating variable (z_t). Wrongly omitting an interaction between z_t and Δx_t will induce bias in short-run parameter estimates but will not affect any long-run parameter estimates or the error correction rate.

The marginal effect of Δx_t on Δy_t in Equation 4.4 is given by $\partial \Delta y_t / \partial \Delta x_t = \alpha_2 + \alpha_5 z_t$. The long-run effect of x_{t-1} on the outcome variable remains unaffected by z_t in this scenario.

4.3.2 Interactions between z_t and x_{t-1}

$$\Delta y_t = \alpha_0 + \alpha_1 y_{t-1} + \alpha_2 \Delta x_t + \alpha_3 x_{t-1} + \alpha_4 z_t + \alpha_6 x_{t-1} \times z_t + \varepsilon_t \tag{4.5}$$

In Equation 4.5, I show an ECM identical to that shown in Equation 4.3 except that it also includes an interaction term between z_t and x_{t-1}. Such an interaction is appropriate when the researcher expects the effect of previous values of a RHS variable (x_{t-1}) to be conditional on values of a moderating variable (z_t). Wrongly omitting an interaction between z_t and x_{t-1} will induce bias in long-run parameter estimates but will not affect any short-run parameter estimates or the error correction rate.

In Equation 4.5, the marginal effect of x_{t-1} on Δy_t is given by $\partial \Delta y_t / \partial x_{t-1} = \alpha_3 + \alpha_6 z_t$. The short-run effect of Δx_t on the outcome variable remains unaffected by z_t in this scenario.

4.3.3 Interactions between z_t and y_{t-1}

$$\Delta y_t = \alpha_0 + \alpha_1 y_{t-1} + \alpha_2 \Delta x_t + \alpha_3 x_{t-1} + \alpha_4 z_t + \alpha_7 y_{t-1} \times z_t + \varepsilon_t \tag{4.6}$$

In some cases, researchers might consider an interaction between z_t and the lagged dependent variable, y_{t-1}. This is appropriate if the researcher expects y_t's time persistence to be conditional on values of a moderating variable z_t. For instance, an increase in the number of institutional veto players (the moderating variable) following an election—likely implying persistence in macroeconomic policy—can increase the effect of the previous year's budget on the current year's budget.

As I show in Equation 4.6, accounting for these types of interactions will influence the error correction rate (and, by definition, the estimate of the effect of y_{t-1} on y_t). Following De Boef and Keele (2008), the error correction rate is given by $\partial \Delta y_t / \partial y_{t-1} = \alpha_1 + \alpha_7 z_t$. The estimate of the effect of y_{t-1} on y_t can be obtained by calculating $1 - \partial \Delta y_t / \partial y_{t-1}$.

It is important to keep in mind that allowing the error correction rate to take on different values conditional on z_t will also affect the calculation of any long-run parameters on RHS variables. For example, in Equation 4.6 the long-run effect of x_{t-1} on the outcome variable will be given by:

$$\frac{\partial \Delta y_t / \partial x_{t-1}}{\Delta y_t / \partial y_{t-1}} = \frac{\alpha_3}{\alpha_1 + \alpha_7 z_t}.$$

4.3.4 Interactions between z_t and all equilibrium-component variables

Researchers might find it appropriate to consider cases where the cointegrated relationship itself differs across values of another variable. Notice the distinction between proposing that a dummy, z_t, moderates the effect of x_{t-1} on Δy_t versus proposing that it moderates the *speed* at which Δy_t responds to shocks in x_{t-1}. Ignoring this possibility leads to problems that should be of theoretical and methodological interest to the researcher. Theoretically, there is a depth of information about the error correction rate that is left out of the model. By ignoring this interaction, the researcher is constraining the long-run dynamics to fixed values, whereas they might vary in theoretically interesting ways. Methodologically, we run the risk of inducing omitted variable bias—particularly with respect to the long-run estimates.

Selectively interacting theoretically relevant RHS variables is unlikely to solve the problem. This is because cointegration and equation balance should be treated in the context of the set of variables included in the model—not between individual RHS variables and the dependent variable (Enns and Wlezien, 2017).

To show the strategy I propose, I follow Arnade, Kuchler and Calvin's (2011) derivation, based on Engle and Granger's (1987) two-step method. The first step is writing an equilibrium equation, in levels:

$$y_t = \omega + \beta_1 x_t + \mu_t \tag{4.7}$$

Where y_t and x_t contain a unit root, and $\mu_t \sim N(0, \sigma^2)$. Equation 4.7 can, of course, be rewritten one-step back and solved in terms of the disturbance, μ_{t-1}.

$$y_{t-1} = \omega + \beta_1 x_{t-1} + \mu_{t-1}$$

$$\mu_{t-1} = y_{t-1} - \omega - \beta_1 x_{t-1} \tag{4.8}$$

Because the variables in Equation 4.7 contain a unit root, the standard errors recovered from an estimation of Equation 4.7 cannot be used for inference (Pickup, 2014). Thus, in the second step we rewrite Equation 4.7 in differences.

$$\Delta y_t = \gamma + \beta_2 \Delta x_t + \varepsilon_t \tag{4.9}$$

This is the disequilibrium equation, where $\varepsilon_t \sim N(0, \sigma^2)$. Engle and Granger's (1987) method requires that we add the error term from the equilibrium equation, $s_{t-1} \equiv \hat{\mu}_{t-1}$, to the disequilibrium equation:

$$\begin{aligned} \Delta y_t &= \gamma + \eta s_{t-1} + \beta_2 \Delta x_t + \varepsilon_t \\ &= \gamma + \eta(y_{t-1} - \omega - \beta_1 x_{t-1}) + \beta_2 \Delta x_t + \varepsilon_t \end{aligned} \tag{4.10}$$

Note that long-term effects, including the error correction rate, come from shocks to the equilibrium component.

If we believe that cointegration itself is moderated by a dummy variable, z_t, then it follows that z_t interacts with the equilibrium component of the model as a whole—not single RHS variables of interest. Hence, we should model the interaction as follows:

$$\Delta y_t = \gamma + \eta(y_{t-1} - \omega - \beta_1 x_{t-1}) \times (1 + \psi z_t) + \beta_2 \Delta x_t + \varepsilon_t \tag{4.11}$$

I multiply the equilibrium component of the model by $(1+\psi z_t)$, instead of ψz_t, in order to include all constituent terms, as suggested in Brambor, Clark and Golder (2006).

For simplification, I set:

$$\gamma + \eta\gamma = \alpha_0$$

$$\eta = \alpha_1$$

$$\beta_2 = \alpha_2$$

$$-\eta\beta_1 = \alpha_3$$

$$-\eta\omega\psi = \alpha_4$$

$$-\eta\beta_1\psi = \alpha_6$$

$$\eta\psi = \alpha_7$$

Equation 4.11 becomes:

$$\Delta y_t = \alpha_0 + \alpha_1 y_{t-1} + \alpha_2 \Delta x_t + \alpha_3 x_{t-1} + \alpha_4 z_t + \alpha_6 x_{t-1} \times z_t + \alpha_7 y_{t-1} \times z_t + \varepsilon_t \tag{4.12}$$

Equation 4.12 demonstrates that the dummy variable, z_t, moderates all lagged variables in the model, including the lagged dependent variable. An added benefit of this strategy is that additional time-lagged RHS variables will also be moderated by z_t. This implies that any long-term effects of exogenous variables will be different from those recovered from a model where an existing interaction is omitted. This occurs for two reasons. First, the error correction rate is itself different because it is conditional on values of z_t ($\partial \Delta y_t / \partial y_{t-1} = \alpha_1 + \alpha_7 z_t$). Second, the effect of a lagged exogenous variable, such as x_t, is also conditional on values of z_t ($\frac{\partial \Delta y_t / \partial x_{t-1}}{\Delta y_t / \partial y_{t-1}} = \frac{\alpha_3 + \alpha_6 z_t}{\alpha_1 + \alpha_7 z_t}$). Recall from De Boef and Keele (2008) that the parameter on lagged exogenous variables in an ECM contains both the short- and long-run effects added in a single value. From Equation 4.12, the short-run parameters on the differenced RHS variables will not be affected by the interaction with z_t.

4.4 Monte Carlo Experiments

Researchers may contemplate the four specification strategies discussed above when using an ECM to model conditional theories. Considering that we are never completely sure of the underlying DGP and these four strategies are plausible in various contexts, I now turn to evaluating the bias and efficiency implications of each. To do this, I present four sets of Monte Carlo experiments. In each set, I simulate data from one (and only one) of the DGPs discussed in Subsections 4.3.1-4.3.4. I then estimate four models, each assuming one of these DGPs. I then evaluate the bias in (1) the short-run parameter on the continuous explanatory variable of interest, (2) the long-run parameter on the continuous explanatory variable of interest, and (3) the error correction rate. I also evaluate the efficiency trade-offs from estimating a model other than that which reflects the correct DGP.

In each set, I employ 1000 Monte Carlo simulations with 40 observations each[7]. The continuous explanatory variable x_t is I(1) and z_t is drawn from a latent continuous variable that is itself

[7]I produced additional simulations with 100 and 200 observations. Results from these additional simulations are reported in Section B.2 of Appendix B. Overall, results are similar.

drawn from a normal distribution. Variables x_t and z_t are correlated by design, since x_t and z_t's latent continuous variable are drawn using a correlation matrix where off-diagonal elements equal 0.5. The error term, ε_t, is normally distributed, has constant variance, and is i.i.d. By construction, each observation is autocorrelated. Therefore, only weak exogeneity can be assumed. If correctly specified, weakly exogenous models produce asymptotically unbiased parameter estimates (Enders, 2008; Pickup, 2014).

DGP 1: $\mathbf{z_t \times \Delta x_t}$

The first DGP and estimation strategy reflect a general ECM equivalent to an ARDL($p = 1$, $q = 1$, $r = 1$) where z_t, a dummy variable, moderates the effect of Δx_t on Δy_t. As discussed in Subsection 4.3.1, this DGP implies that the contemporaneous effect of Δx_t on Δy_t is conditional on the values of a third variable, z_t. This interaction bears no consequence to the long-run effect of x_{t_1} on Δy_t. In the simulations I present below, DGP 1 can be mathematically represented as follows:

$$\Delta y_t = 0.75 y_{t-1} + 4\Delta x_t + 8x_{t-1} + 2z_t + 6z_t \times \Delta x_t + \varepsilon_t \tag{4.13}$$

DGP 2: $\mathbf{z_t \times x_{t-1}}$

The second DGP and estimation strategy reflect the case where z_t, a dummy variable, moderates the effect of x_{t-1} on Δy_t. As discussed in Subsection 4.3.2, this DGP implies that the long-run effect of x_{t-1} on Δy_t is conditional on values of a third variable, z_t. This interaction bears no consequence to the short-run effect of Δx_t on Δy_t. In the simulations I present below, DGP 2 can be mathematically represented as follows:

$$\Delta y_t = 0.75 y_{t-1} + 4\Delta x_t + 8x_{t-1} + 2z_t + 6z_t \times x_{t-1} + \varepsilon_t \tag{4.14}$$

DGP 3: $\mathbf{z_t \times y_{t-1}}$

The third DGP and estimation strategy reflect the case where z_t, a dummy variable, moderates the effect of the lagged dependent variable, y_{t-1}, on the outcome, Δy_t. As discussed in Subsection 4.3.3, this DGP implies that persistence in the dependent variable is itself conditional on values of a third variable, z_t. This means that both the error correction rate and any long-run effects of RHS variables are also conditional on values of z_t. This interaction bears no direct consequence to the short-run effect of Δx_t on Δy_t but its incorrect omission is likely to lead to serially correlated errors and, potentially, biased and inefficient parameter estimates due to autocorrelation. In the simulations I present below, DGP 3 can be mathematically represented as follows:

$$\Delta y_t = 0.5 y_{t-1} + 4\Delta x_t + 8x_{t-1} + 2z_t + 0.25 z_t \times y_{t-1} + \varepsilon_t \tag{4.15}$$

DGP 4: Equilibrium Component Interaction

The fourth DGP and estimation strategy reflect the case where z_t, a dummy variable, moderates the effect of the equilibrium component of the model. This DGP is appropriate in contexts where the researcher expects the long-term dynamics—and cointegration itself—to be conditional on values of a third variable, z_t. The incorrect omission of interactions between z_t and the equilibrium component of the model is likely to induce bias in the error correction rate and any estimates of long-run relationships. In addition, RHS estimates are likely to be inefficiently estimated. In the simulations I present below, DGP 4 can be mathematically represented as follows:

$$\Delta y_t = 0.5 y_{t-1} + 4\Delta x_t + 8x_{t-1} + 2z_t + 6z_t \times x_{t-1} + 0.25 z_t \times y_{t-1} + \varepsilon_t \tag{4.16}$$

4.4.1 Short-Run Effects of the Explanatory Variable of Interest

I begin by examining kernel density plots of the short-run effects of Δx_t, the continuous explanatory variable of interest, on the outcome, Δy_t. In Figure 4.1, I show results from 16 estimated models. Each column represents one of the DGPs discussed above and each row represents one of the respective estimation strategies. In the main diagonal of sub-panels I depict the correct estimation strategies across the respective DGPs. The correctly specified models lead to relatively unbiased estimates that perform similarly to or better than alternative estimations, though they do not perform as well as one would expect of correctly specified models.

True parameter values are depicted by a vertical dashed line. Under DGP 1, which includes an interaction between z_t and Δx_t, the short-run marginal effect of Δx_t can be calculated from Equation 4.13 as follows: $\partial \Delta y_t / \partial \Delta x_t = 4 + 6z_t$. In Figure 4.1, I set $z_t = 1$, when applicable, such that the true parameter value equals 10. Under DGPs 2, 3, and 4, the short-run effect of Δx_t on Δy_t can be readily obtained from Equations 4.14, 4.15, and 4.16: $\partial \Delta y_t / \partial \Delta x_t = 4$.

Except under DGP 1, the first estimation strategy produces estimates that slightly underestimate the short-run effect of Δx_t on Δy_t. As shown in Figure 4.1, these estimates are substantially less tightly distributed around the true parameter estimate than estimates obtained from other estimation strategies. The second estimation strategy underestimates the short-run effect of Δx_t on Δy_t under DGP 1 but produces unbiased estimates under DGPs 2, 3, and 4. But with this estimation strategy, parameter estimates are less tightly distributed around true values under DGPs 3 and 4. Similarly, the third estimation strategy underestimates the short-run effect of Δx_t on Δy_t under DGP 1 but produces unbiased estimates under DGPs 2, 3, and 4. With this estimation strategy, parameter estimates are less tightly distributed around true values under DGPs 2 and 4.

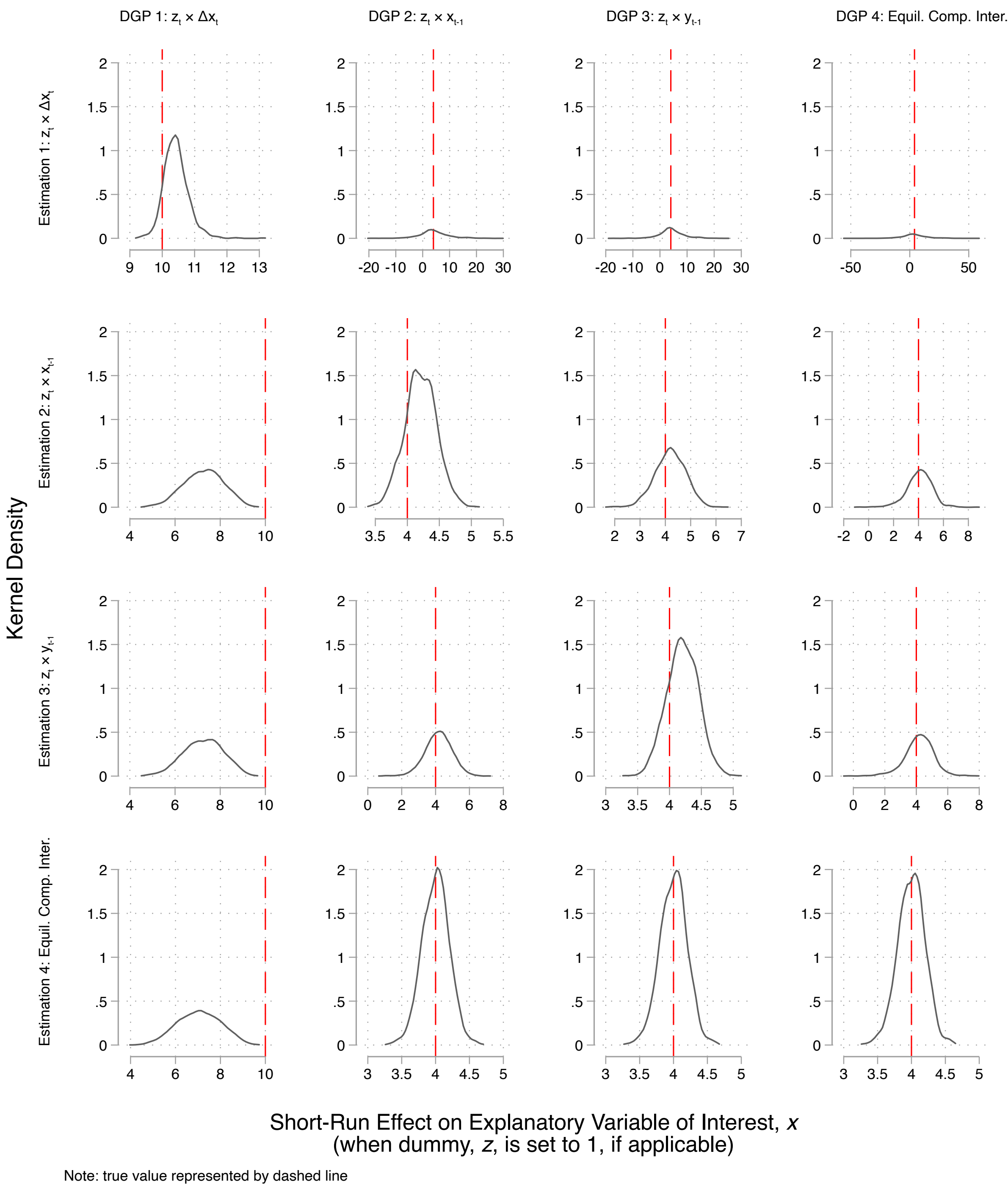


Figure 4.1: Kernel Density plots of the short-run effects of the explanatory variable of interest, Δx_t, from 40-observation Monte Carlo simulation estimation results. (Columns represent different data generating processes and rows represent the respective estimation strategies.)

The fourth estimation strategy also underestimates the short-run effect of Δx_t on Δy_t under DGP 1 but performs well under DGPs 2, 3, and 4. Parameter estimates under these DGPs are unbiased and tightly distributed around true values.

4.4.2 Long-Run Effects of the Explanatory Variable of Interest

The ECM, like any dynamic model, has the advantage of producing parameter estimates that can be easily used in the calculation of long-run effects. In Figure 4.2, I show the kernel density plots of the estimated long-run effect of x_{t-1}, a continuous explanatory variable of interest included in the model. As in Figure 4.1, the sub-panels in Figure 4.2 are organized such that each column represents a different DGP and each row represents an alternative estimation strategy. The correctly specified strategy is shown along the main diagonal.

In DGP 1, the long-run effect of x_{t-1} on Δy_t can be calculated by dividing the parameter on x_{t-1} by the error correction rate (the parameter on y_{t-1}). From Equation 4.13, this is given by: $\partial \Delta y_t / \partial x_{t-1} = 8/0.75 = 10.7$. In DGP 2, the long-run effect of x_{t-1} on Δy_t can be calculated by dividing the marginal effect of x_{t-1} in Equation 4.14 ($\partial \Delta y_t / \partial x_{t-1} = 8 + 6z_t$) by the error correction rate ($\partial \Delta y_t / \partial y_{t-1} = 0.75$). When $z_t = 1$, the true long-run effect of x_{t-1} on Δy_t simplifies to $14/0.75 = 18.7$. In DGP 3, the long-run effect of x_{t-1} on Δy_t can be calculated by dividing the parameter on x_{t-1} ($\partial \Delta y_t / \partial x_{t-1} = 8$) by the marginal effect of y_{t-1} in Equation 4.15 ($\partial \Delta y_t / \partial y_{t-1} = 0.25 + 0.5z_t$). When $z_t = 1$, the true long-run effect of x_{t-1} on Δy_t simplifies to $8/0.75 = 10.7$. Finally, in DGP 4 the long-run effect of x_{t-1} on Δy_t can be calculated by dividing the marginal effect of x_{t-1} ($\partial \Delta y_t / \partial x_{t-1} = 8 + 6z_t$) by the marginal of effect of y_{t-1} ($\partial \Delta y_t / \partial y_{t-1} = 0.25 + 0.5z_t$) in Equation 4.16. When $z_t = 1$, the true long-run effect of x_{t-1} on Δy_t simplifies to $14/0.75 = 18.7$.

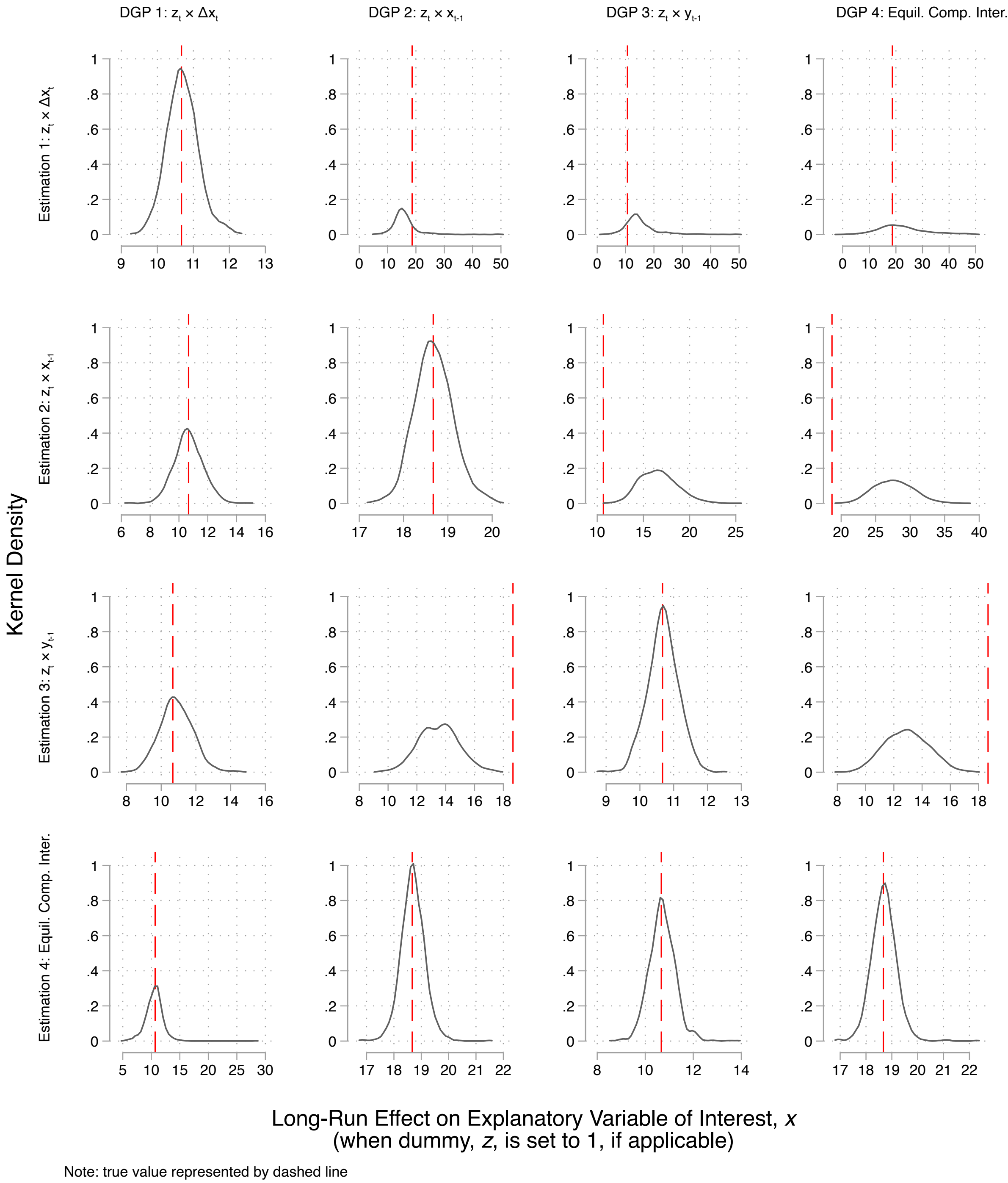


Figure 4.2: Kernel Density plots of the long-run effects of the explanatory variable of interest, x_{t-1}, from 40-observation Monte Carlo simulation estimation results. (Columns represent different data generating processes and rows represent the respective estimation strategies.)

As expected, the results along the main diagonal of sub-panels in Figure 4.2 show that the correct estimation strategy produces results that are comparable to or less biased than those from alternative strategies across their respective DGPs. The first estimation strategy, in particular, often produces biased estimates of the long-run effects of x_{t-1} on Δy_t. Both the second and third estimation strategies also often produce biased estimates when wrongly employed, though less so under DGP 1. The fourth estimation strategy performs well across the board, though less so under DGP 1. It is noteworthy, however, that the bias incurred from employing the fourth estimation strategy under DGP 1 is substantially attenuated when compared to the bias incurred when using the first estimation strategy under DGPs 2, 3, and 4.

These results suggest that it may be less pernicious to wrongly employ the fourth estimation strategy than to wrongly employ any of the other three estimation strategies. More specifically, the fourth estimation strategy tends to produce unbiased estimates of the long-run effect of x_{t-1} on Δy_t under DGPs 2, 3, and 4 though some bias is incurred when using this estimation strategy under DGP 1.

4.4.3 Error Correction Rate

The error correction rate is essential when calculating long-run effects in ECMs (Engle and Granger, 1987). In fact, correctly estimating this rate is likely the central motivating factor behind a researcher's use of the ECM. When a theoretical framework suggests the presence of a multiplicative interaction in an ECM, it is paramount that the researcher consider the possibility that the error correction rate and, more broadly, the equilibrium component of the model are conditional on values of a moderating variable.

In Figure 4.3, I show kernel density plots of the error correction rate recovered from the alter-

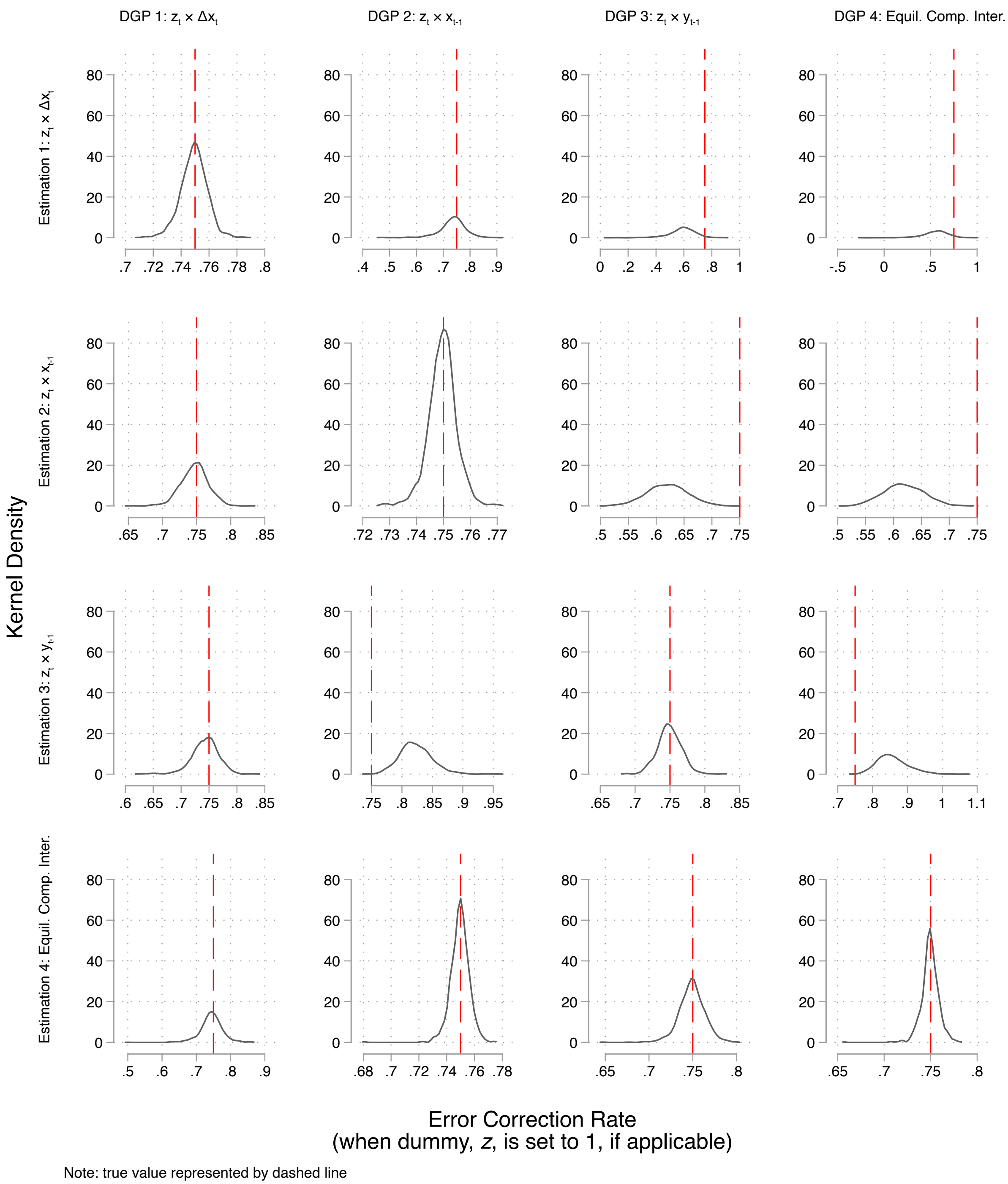


Figure 4.3: Kernel Density plots of the error correction rate from 40-observation Monte Carlo simulation estimation results. (Columns represent different data generating processes and rows represent the respective estimation strategies.)

native estimation strategies across different DGPs. Under DGP 1 and DGP 2, which do not include an interaction between the lagged dependent variable and z_t, the error correction rate can be readily obtained from the parameter on the lagged dependent variable. The true error correction rate in these DGPs is $\alpha_1 = 0.75$, from Equations 4.13 and 4.14. In the third and fourth DGPs, which include an interaction between the lagged dependent variable and z_t, the true error correction rate is given by: $\partial \Delta y_t / \partial y_{t-1} = 0.5 + 0.25 z_t$, as shown in Equations 4.15 and 4.16.

Along the main diagonal of the matrix of sub-panels shown in Figure 4.3 are the kernel density plots of the error correction rate from correctly specified models. In comparison to alternative estimations, the correctly specified models lead to error correction rates that are centered around the true value, indicated by a vertical dashed line. In every case, the correct specification leads to results that are either comparable to or less biased than alternative specifications. It is noteworthy, however, that all estimation strategies lead to an unbiased estimate of the error correction rate under DGP 1 where z_t moderates Δx_t. In addition, the first and second estimation strategies tend to underestimate the error correction rate in DGPs 2-4 and DGPs 3-4, respectively. The third estimation strategy tends to overestimate the error correction rate in DGP 2 and DGP 4. Interestingly, while the first three estimation strategies do not perform well under DGP 4, the fourth estimation strategy produces unbiased estimates under all four DGPs.

These results suggest that the researcher is less likely to incur substantial bias in the estimation of the error correction rate from incorrectly including an interaction between a dummy and the equilibrium component of the model than from omitting this interaction when one is truly present. In the absence of strong *a priori* expectations, the researcher should give strong consideration to the interaction with the equilibrium component of the model in an ECM.

4.4.4 Root Mean Square Error

I now turn to assessing the efficiency of each of these alternative estimation strategies with data from each of the DGPs discussed. The root mean square error (RMSE) is useful when evaluating the overall efficiency of a model (though it also incorporates bias). It measures how far, on average, the data points are from the regression line. Lower RMSE values reflect better fit than higher RMSE values. In practice, a regression with no error (perfect fit) produces an RMSE value equal to zero.

I show the kernel density plots of the RMSEs recovered from each of the 16 estimations shown in the sub-panels included in Figure 4.4. As was the case in Figures 4.1, 4.2, and 4.3, the main diagonal of the matrix of sub-panels in Figure 4.4 shows the correct estimation strategies for the respective DGPs. As should be expected, the results along the main diagonal show that those estimations were comparable to or more efficient than alternative estimations. All estimation strategies also produced comparably efficient estimates under DGP 1.

The second and third estimation strategies perform relatively well throughout. However, when used incorrectly (when not reflecting the true DGP) they tend to be relatively less efficient. The fourth estimation strategy performs remarkably well on data from all four DGPs.

The results shown in Figure 4.4 suggest that the inclusion of an interaction between a dummy and the equilibrium component of the model is unlikely to lead to more uncertainty in the results. Quite the opposite—the omission of such an interaction when one is truly present leads to higher uncertainty in the estimates.

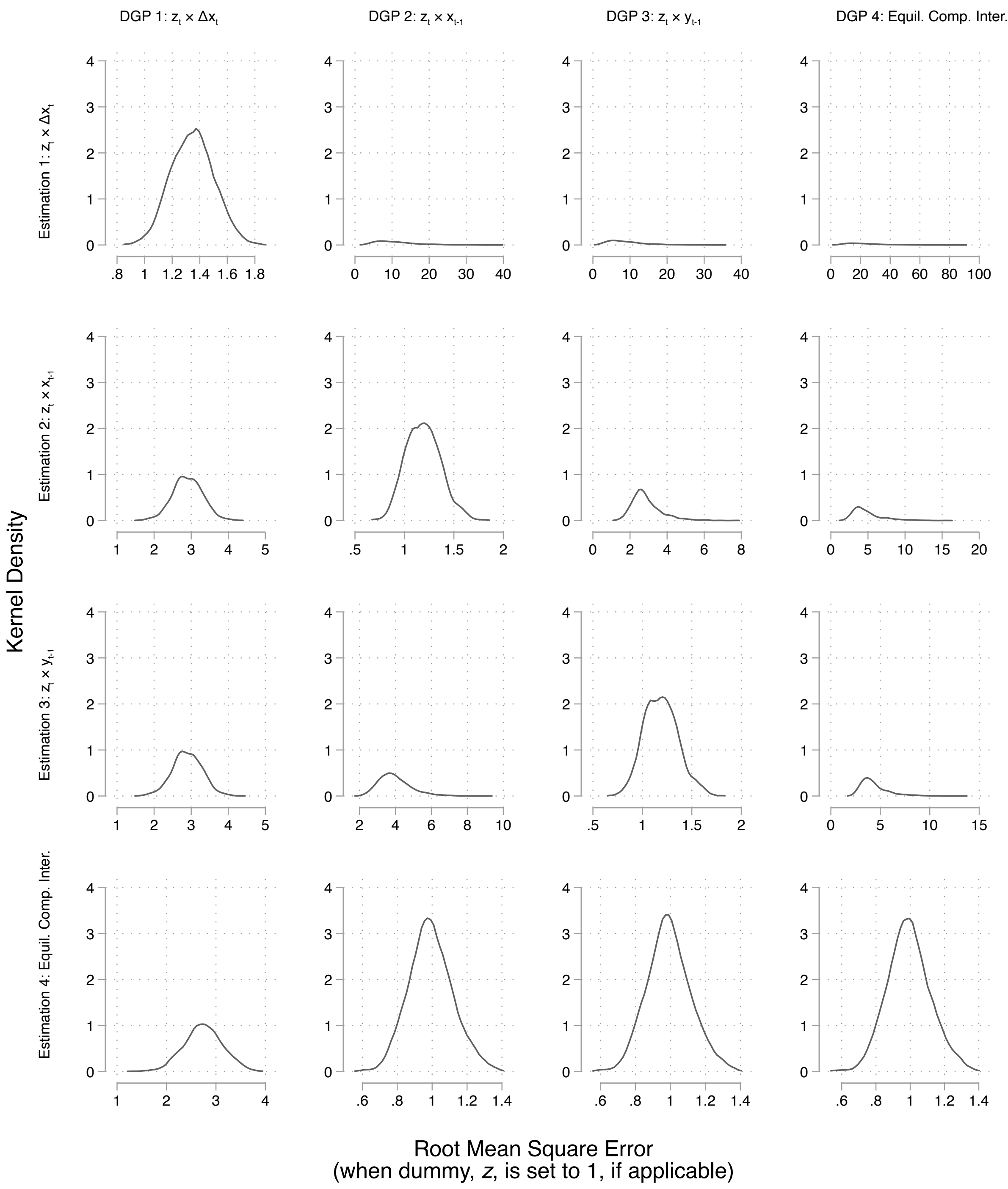


Figure 4.4: Kernel Density plots of the root mean square error (RMSE) from 40-observation Monte Carlo simulation estimation results. (Columns represent different data generating processes and rows represent the respective estimation strategies.)

4.5 Empirical Application

Media Coverage and Retrospective Economic Evaluations in the US

To illustrate the value of modeling multiplicative interactions within ECMs, I replicate Soroka, Stecula and Wlezien (2015), who examine whether media content influences the public's economic sentiment in the United States. I focus on the authors' (retrospective) "Responsiveness of Economic Evaluations to Media" baseline model, where they employ a general ECM using data from 1980 to 2011.[8] The dependent variable utilized is consumer sentiment from Thomson Reuters/University of Michigan Surveys of Consumers. The RHS variables include volume of economic news coverage (media count) and tone of media coverage. Both of these are from published editions of the *New York Times* and *Washington Post* in the Lexis-Nexis database. In addition, the authors control for economic conditions with a composite index of economic indicators from the Conference Board.

I reproduce Soroka, Stecula and Wlezien's (2015) findings and report them as Model 1 in Table 4.1. These findings indicate a high degree of persistence in the dependent variable and relatively slow error-correction dynamics (Error Correction Rate $= -0.0821$). Although the recovered parameter estimates on ΔMedia Count and Media Count$_{t-1}$ are negative, the results are not statistically distinguishable from zero at the 95% confidence level. The recovered parameters on ΔMedia Tone and Media Tone$_{t-1}$ are both positive and statistically distinguishable at the 99% confidence level. Only the short-run parameter on the composite economic indicator index (ΔEconomic Indicator) is statistically distinguishable from zero.

I extend Soroka, Stecula and Wlezien's (2015) model by including a multi-period interven-

[8] This model is shown in the first column of Table 7 in Soroka, Stecula and Wlezien (2015).

DV: Δ Retrospective Evaluations

	Model 1	Model 2	Model 3
DV_{t-1}	-0.0821***	-0.0967***	-0.0859***
	(0.0170)	(0.0178)	(0.0180)
Δ Media Count	-0.0422	-0.0367	-0.0326
	(0.0251)	(0.0250)	(0.0247)
Media Count_{t-1}	-0.0374	-0.0318	-0.0308
	(0.0216)	(0.0216)	(0.0213)
Δ Media Tone	8.977***	8.533***	8.144***
	(1.670)	(1.668)	(1.655)
Media Tone_{t-1}	9.670***	8.958***	8.875***
	(1.890)	(1.898)	(1.877)
Δ Economic Indicator	3.836***	3.122***	3.334***
	(0.843)	(0.885)	(0.878)
$\text{Economic Indicator}_{t-1}$	-0.0886	-0.125	-0.0998
	(0.169)	(0.169)	(0.167)
Recession Dummy		-4.366*	-14.15***
		(1.749)	(3.656)
Recession $\times$ DV_{t-1}			-0.169**
			(0.0555)
Constant	0.0949	0.340	0.244
	(1.735)	(1.725)	(1.707)
N	380	380	380
Adj. R^2	0.216	0.227	0.244

Standard errors in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

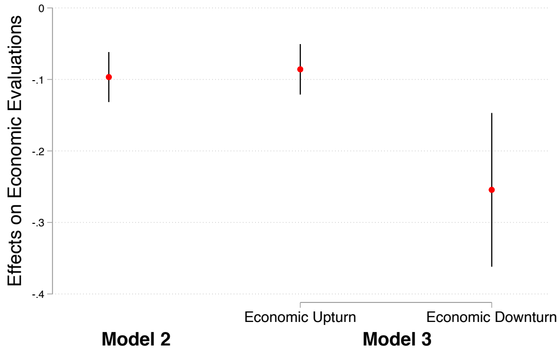
Table 4.1: Regression results from Model 1 (original baseline model reported in Soroka, Stecula and Wlezien (2015), Model 2 (which includes a recession dummy but no interaction), and Model 3 (which includes an interaction between a lagged-dependent variable and a recession dummy).

tion variable that takes the value 1 when the economy is in a recession and 0 otherwise. Following Lewis-Beck and Stegmaier (2000); Duch and Stevenson (2010); Lindvall (2014, 2017), it is reasonable to expect that economic recessions both influence news media coverage as well as the public's economic sentiment. Importantly, recessions can moderate the rate of adjustment to equilibrium in the public's economic sentiment.

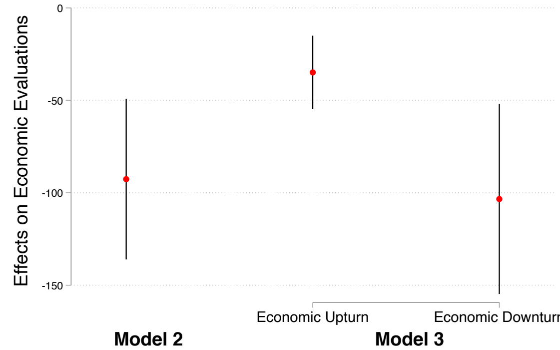
In Model 2 (shown in Table 4.1), the recession dummy is included additively to show that results are substantively and statistically consistent with those of Model 1. In Model 3, I interact the recession dummy with the lagged dependent variable assuming a DGP akin to DGP 3 in the Monte Carlo experiments explored in Section 4.4. The inclusion of an interaction between a recession dummy and the lagged dependent variable assumes that the persistence of retrospective economic evaluations itself varies depending on whether times are good (i.e., economic upturn) or bad (i.e., economic downturn).

In Model 3, reported in Table 4.1, the parameter estimates on the recession dummy and on the interaction term are negative and statistically significant suggesting that, indeed, persistence in retrospective economic evaluations itself varies based on the presence or absence of a recession. It is important to keep in mind that the presence of heterogeneous effects on the lagged dependent variable also means that (1) the error correction rate is conditional on values of the moderating variable and (2) long-run effects of RHS variables may also be conditional on values of the moderating variable.

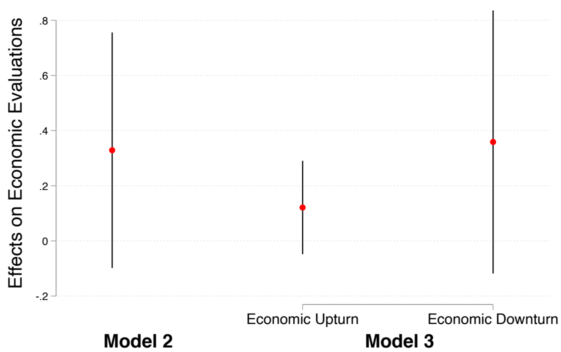
I explore the implications of estimating homogeneous effects (in Model 2) and heterogeneous effects (in Model 3) with marginal effects graphs in Figure 4.5. In panel 4.5a, I show the error correction rates recovered from Models 2 and 3. (I omit results from Model 1 in Figure 4.5 because they are substantively and statistically similar to those from Model 2.) In Model 2, I recover an



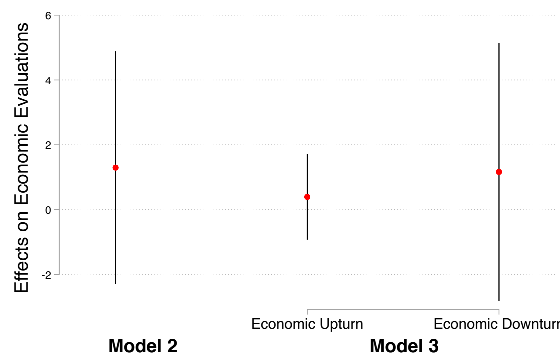
(a) Error Correction Rate



(b) LR Effect of Media Tone



(c) LR Effect of Media Count



(d) LR Effect of (Composite) Economic Indicator

Figure 4.5: Estimated error correction rate and long-run effects from Model 2 (no interaction) and Model 3 (which includes an interaction between the lagged-dependent variable and a recession dummy).

error correction rate of -0.1. Since in this model the lagged dependent variable is not part of a multiplicative interaction, the error correction rate is homogeneous across values of the recession dummy. In Model 3, on the other hand, the error correction rate is conditional on values of the recession dummy. Specifically, the error correction rate is greater during economic upturns (-0.09) than during economic downturns (-0.25). This implies that error correction takes places more speedily during economic downturns than during economic upturns. These results are statistically significant at the 99% confidence level.

In panel 4.5b, I show the long-run effects of media tone on retrospective economic evaluations. The results here suggest that omitting the interaction can be problematic for inference of long-run effects. In the case of media tone, estimating Model 2 over Model 3 would lead to the conclusion that these effects are homogeneous (-92.6) whereas Model 3 shows that media tone has a more negative effect during economic downturns (-103.3) than during economic upturns (-34.8). These findings are statistically distinguishable from zero at the 99% confidence level. In panels 4.5c and 4.5d, I show the long-run effects of media count and the composite indicator of economic conditions. In both models, these variables are not statistically distinguishable from zero at the 90% confidence level.

4.6 Discussion

With the discussion and findings above, I demonstrate the importance of appropriately modeling conditional theories in general ECMs in order to draw accurate inferences. Importantly, misspecifying conditional relationships can be especially costly to inferences of the error correction rate and long-run parameter estimates on RHS variables. Given the centrality of the error correction rate (for calculating long-run relationships), this implies that neglecting the theoretically appropriate alternatives in which to specify these relationships will lead to misusing these models and misinterpreting results.

If scholars believe (1) that an error correction model is theoretically and statistically warranted, and (2) that a conditional relationship exists between RHS variables, then they must give careful consideration to the particular ways in which the moderator interacts with moderated variables. Here, I simplify my recommendations to contexts in which we assume that the moderator is a contemporaneous dummy variable. In this paper, I consider contexts in which the moderator interacts (1) with the first difference of a RHS variable (Δx_t), (2) with a time lag of a RHS variable (x_{t-1}), (3) with a time lag of the dependent variable (y_{t-1}), and (4) with the equilibrium component of the model (i.e., all time-lagged terms in the RHS).

In the presence of strong *a priori* theoretical expectations in favor of one of these contexts, it certainly makes sense to model that particular context. However, scholars may find it beneficial to model the entire equilibrium component of the model. Results from the Monte Carlo experiments above suggest that modeling the equilibrium component interaction—the most comprehensive of the alternatives considered—recovers (mostly) relatively unbiased and efficient estimates throughout even when it does not reflect the correct DGP. When it does produce some bias in estimates (for instance, when the correct DGP includes a conditional relationship between a moderator and the first difference of a RHS variable), it attenuates the magnitude of estimates. Importantly, the magnitude of bias produced by incorrectly specifying an interaction with the equilibrium component of the model is smaller than the bias produced by incorrectly using alternative specifications.

This piece's central contribution to the time-series literature is in calling attention to heterogeneous error correction rates. Whereas most applications of ECMs assume that a shock to an explanatory variable of interest affects the dependent variable and returns to equilibrium at a constant rate, I show that this rate may actually be conditional on values of a moderator. Assuming that theories undergirding the use of ECMs are often about long-run relationships, it stands to reason that scholars should consider the possibility that the error correction rate is not homogeneous.

My findings suggest that researchers should keep the following considerations in mind when developing conditional theories in contexts where ECMs are theoretically and statistically warranted. First, researchers should consider the dynamic nature of the conditional theory—is it about a short-term relationship or a long-term relationship? Second, researchers should consider the extent to which the conditional relationship applies to a single RHS variable or multiple RHS variables. Third, when they suspect that the conditional theory applies to multiple RHS variables, they should ask themselves whether the theorized relationship is specifically about the effect of a predictor on the outcome or about the speed at which the outcome returns to equilibrium following shocks to the predictor.

4.7 Conclusion

Given the complex and multivariate nature of political science, conditional relationships are prominent in the field. Over time, helpful guidelines have been developed to appropriately model these types of relationships—though not necessarily for time-series analysis. While applying these guidelines to time series is relatively straightforward, one exception occurs in the context of ECMs. Here, I present a study of four potential specification strategies of multiplicative interactions between a moderating variable and RHS terms in a general ECM. Namely,

1. An interaction between a dummy moderator (z_t) and a time-differenced RHS variable (Δx_t)
2. An interaction between a dummy moderator (z_t) and a time-lagged RHS variable (x_{t-1})
3. An interaction between a dummy moderator and the lagged dependent variable (y_{t-1})
4. An interaction between a dummy moderator (z_t) and the equilibrium component of the model (all time-lagged RHS terms: y_{t-1}, x_{t-1})

Findings from a series of Monte Carlo experiments and an empirical application with data from the United States suggest that misspecifying the appropriate interaction can lead to substantial bias and inefficiency in the long-run parameter estimates. Importantly, these findings call attention to the presence of heterogeneous error correction rates—a context often neglected by researchers modeling conditional relationships in ECMs.

In discussing these findings, I provide some considerations scholars should keep in mind when modeling conditional relationships in ECMs. These considerations can be summarized with the following questions.

1. Is the conditional theory about a short-term relationship or a long-term relationship?

2. Does the conditional theory apply to a single RHS variable or multiple RHS variables?

3. Is the conditional theory about the individual effect of predictors on the outcome or about the speed at which the outcome returns to equilibrium following shocks to predictors?

While a contribution in its own right, this paper is but a first step in discussing conditional relationships in time-series analysis. In seeking to fill this void in the literature, I focus on a particular set of models: error correction models. I also simplified the analysis to four likely data generating processes one might encounter and assume the use of a dummy moderator within a two-way interaction. Further research should consider (1) the use of continuous moderators, (2) multiplicative interactions between more than two variables, (3) constrained or expanded ECMs (with fewer or more terms, including additional time lags when theoretically warranted), (4) comparative contexts in which cointegration is and is not present, (4) contexts in which endoegeneity is introduced. Furthermore, diagnostic guidelines should be provided to warrant the use of multiplicative interactions in ECMs as well as in other time-series models.

5. CONCLUSIONS AND FUTURE RESEARCH

"*Buy on the rumor, sell on the news.*"

The relationship between politics and financial markets is frequently about rumors. Markets continuously and, oftentimes, frantically seek to incorporate all available information when trading and pricing assets (Fama, 1970, 1995, 2021). In democracies, this process does not stop when elections are over and winners have been declared. In a way, studying the mass politics of financial markets is like studying the role of political rumors in financial markets.

The broader claim presented in this dissertation is that incumbent popularity is a source of political capital and, therefore, affects the set of policy choices policymakers are willing and able to pursue. Specifically, I argue that incumbent popularity—conditional on factors such as government ideology and institutional constraints—affects expectations about the extent to which governments are more likely to pursue policy changes and whether they are more likely to pursue market-friendly or market-unfriendly policies. Whether higher incumbent popularity *actually* produces market-friendly or market-unfriendly policies is only tangential to this discussion. It is a valid question, of course. But it is a question that likely belongs in a different, future project. What is relevant to this discussion is the *rumor* that incumbent policymakers will pursue certain types of policies. Mass political opinion informs markets—albeit imperfectly—on the rumors about policymakers' likely policy choices.

Research has shown that individual investors do not necessarily understand the ins and outs of why certain pieces of information matter for investment returns (Ballard-Rosa, Mosley and Wellhausen, 2019). Investors might categorize different groups of countries (e.g., democracy/autocracy, emerging/developed, high/low income) in order to screen information and make systematic investment decisions. Importantly, information—even the *rumorous* kind—that is priced in may have a permanent effect on investments that is not fully unwound once new information comes out and decidedly disproves old information. This implies that high approval of a left-wing incumbent, for example, may still frighten markets (i.e., a residual effect) long after that incumbent's market-unfriendly policy proposal loses a vote in the legislature and fails to become law.

Indeed, in Chapters 2 and 3 I find evidence in support of my theory that mass political opinion is associated with sovereign debt market interest rates and stock market volatility. In the former chapter, I argue that the effect of incumbent approval on treasury rates is conditional on government ideology and political-institutional constraints. It is important to keep in mind that when discussing the role of government ideology, I focus exclusively on its traditional left-right economic dimension. Under left-wing governments, higher approval is associated with an increased probability of market-unfriendly policies and, therefore, investors penalize sovereign debt markets with relatively higher interest rates. Under right-wing governments, higher approval is associated with an increased probability of market-friendly policies though, interestingly, investors do not seem to reward sovereign debt markets with lower interest rates. Perhaps, because investors are predisposed to preferring right-wing governments and their policies, they are less likely to respond meaningfully to changes in approval and potential policy change when the right is in power. In this Chapter, I also find that higher approval is more likely to lead to higher interest rates under contexts of few institutional constraints than in contexts of many institutional constraints.

In Chapter 3, I argue that effect of incumbent approval on stock market volatility is conditional on government ideology and financial market institutional development. As in Chapter 2, I focus exclusively on the economic dimension of government ideology. I find evidence that a rise in approval of left-wing governments, associated with a higher probability of market-unfriendly policies, leads markets to engage in less trading and speculative behavior. As a result, stock market volatility declines. On the other hand, a rise in right-wing incumbent approval, associated with a higher probability of market-friendly policies, leads markets to engage in more trading and speculative behavior. As a result, stock market volatility rises. Importantly, these results hold in countries with relatively less developed financial market institutions but not in countries with relatively more developed financial market institutions.

It is common for political scientists—political economists even—to question the value of studying the role of mass political opinion in financial markets. After all, the masses are not, in any coordinated manner, adjusting their views of the incumbent in order to manipulate asset prices.[1] Much more likely, they evaluate incumbents based on a series of micro- and macro-socioeconomic factors that happen to influence policymaking—and rumors about policymaking—which in turn influence financial markets. In addition, it is widely assumed that one's subject of research is her dependent variable. Hence, some argue, research employing financial market indicators as dependent variables cannot truly belong in political science.

At least three counterarguments can be provided here. First, the claim that mass political opinion affects financial markets is inherently about the implications of democratic politics far beyond election times and legislative chambers. It speaks to the ability of democracy to influence the economy continuously and dynamically. Second, while unintentional and uncoordinated, the effect of mass political opinion on markets is useful to students of politics. In fact, even if an unintended consequence of democracy, the ins and outs of this relationship are critical for policy and institutional design. Third, financial market success is important to elected officials. Governments place

[1]Although this would certainly be a phenomenal conspiracy theory to expound on in the future.

a premium on their ability to issue debt, access capital for public projects, maintain economic stability, and privatize or nationalize economic sectors.

The substantive work presented in this dissertation also contributes to various literatures within the study of political economy. First, it provides evidence that democratic policymaking affects financial markets and has potential ramifications for the economy more broadly. This contributes to the literature on the political economy of democracy (e.g. Przeworski and Wallerstein, 1988; Saiegh, 2005; Ballard-Rosa, Mosley and Wellhausen, 2019). Second, my findings speak to the role of political institutions and political ideology in moderating market behavior. This contributes to the literature on the political economy of political institutions (e.g. North and Thomas, 1973; Shepsle, 1979; Hibbs, 1989; North, 1991; Powell and Whitten, 1993; Tsebelis, 1995, 2002, 2011). Third, this discussion adds to the understanding of political risk in financial markets within and well beyond elections. This contributes to the literature on the political economy of financial markets (e.g. Mosley, 2003; Leblang and Mukherjee, 2005; Fowler, 2006; Füss and Bechtel, 2008; Bechtel, 2009; Goodell and Bodey, 2012; Breen and McMenamin, 2013). Fourth, in my empirical analyses I extend the study of political risk in financial markets to the context of emerging markets. This contributes to the broader study of comparative political economy.

The work developed in the fourth chapter—discussing conditional relationships in ECMs—speaks to the methods literatures on multiplicative interactions (e.g. Brambor, Clark and Golder, 2006; Franzese and Kam, 2009; Hainmueller, Mummolo and Xu, 2019) and ECMs (e.g. Engle and Granger, 1987; De Boef and Keele, 2008; Keele, Linn and Webb, 2016; Grant and Lebo, 2016; Philips, 2018; Webb, Linn and Lebo, 2020). It provides the most comprehensive discussion that I am aware of on modeling conditional theories with ECMs and considers the ramifications of employing rival specification strategies. The recommendations I provide are useful for time-series analyses in a variety of domains within and outside of political science.

I close with the following observations. The substantive theories developed in this dissertation are founded on two primary assumptions. (1) Rises in incumbent popularity are associated with greater political capital. (2) Rises in incumbent popularity are associated with greater policy uncertainty. Together, these assumptions imply that rises in incumbent popularity increase policymakers' willingness to pursue policy changes and exacerbate the potential political-economic consequences of enacting those changes. Additional work is required to relax these assumptions and model the subsequent implications. Empirically, one might consider the extent to which ideology and institutional constraints moderate the effect of mass political opinion on financial markets *simultaneously*. Unfortunately, limited data (and statistical power) make this difficult in the current dissertation. As new time-series data become available, scholars should extend this research along those lines. In addition, much work remains to be developed considering the role of political risk (and mass political opinion) across various financial market assets (e.g., commodities, real estate, currency markets) as well as considering alternative dimensions of mass politics (e.g., digital media, populist rhetoric, alternative media sources).

The methodological work developed in Chapter 4 can be extended to consider conditional relationships in time-series more broadly. In addition, the study of error correction models in time-series cross-sectional samples—as well as the use of non-continuous variables in time-series and time-series cross-sectional samples—would be a major contribution to the literature.

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APPENDIX A

SUPPLEMENTARY INFORMATION FOR CHAPTER 2

A.1 Regression Results from Models 1 and 2

DV: Treasury Rates	Model 1	Model 2
Treasury Rates$_{t-1}$	.445***	.446***
	(23.83)	(23.88)
Approval$_t$	.0140**	.0183**
	(3.03)	(3.13)
Ideology$_t$	.138	-.0814**
	(1.34)	(-3.03)
Institutions$_t$	.605**	2.277**
	(2.59)	(3.12)
Central Bank Rate$_t$	.559***	.560***
	(28.09)	(28.14)
Election Dummy	-.0124	-.0150
	(-.12)	(-.15)
Approval$_t\times$ Ideology$_t$	-.00487*	
	(-2.23)	
Approval$_t\times$ Institutions$_t$		-.0368*
		(-2.47)
Intercept	-.884***	-1.074***
	(-3.60)	(-3.66)
Adj. R^2	.98	.98
N	690	690

t statistics in parentheses

* $p < .05$, ** $p < .01$, *** $p < .001$

Table A.1: Chapter 2 regression results from Model 1 (reflecting Equation 2.1) and Model 2 (reflecting Equation 2.2).

APPENDIX B

SUPPLEMENTARY INFORMATION FOR CHAPTER 4

B.1 Equivalence between ARDL(1,1,1) and ECM

Let's assume an ARDL(1,1,1) as follows, where $\varepsilon_t \sim N(0, \sigma^2)$:

$$y_t = \phi_0 + \phi_1 y_{t-1} + \phi_2 x_t + \phi_3 x_{t-1} + \phi_4 z_t + \phi_5 z_{t-1} +$$

$$+ \phi_6 x_t z_t + \phi_7 x_{t-1} z_t + \phi_8 x_t z_{t-1} + \phi_9 x_{t-1} z_{t-1} + \varepsilon_t$$

If we subtract y_{t-1} from both sides of the equation, we're left with the following.

$$y_t - y_{t-1} = \phi_0 + \phi_1 y_{t-1} - y_{t-1} + \phi_2 x_t + \phi_3 x_{t-1} + \phi_4 z_t + \phi_5 z_{t-1} +$$

$$+ \phi_6 x_t z_t + \phi_7 x_{t-1} z_t + \phi_8 x_t z_{t-1} + \phi_9 x_{t-1} z_{t-1} + \varepsilon_t$$

$$\Delta y_t = \phi_0 + (\phi_1 - 1) y_{t-1} + \phi_2 x_t + \phi_3 x_{t-1} + \phi_4 z_t + \phi_5 z_{t-1} +$$

$$+ \phi_6 x_t z_t + \phi_7 x_{t-1} z_t + \phi_8 x_t z_{t-1} + \phi_9 x_{t-1} z_{t-1} + \varepsilon_t$$

We can then add and subtract the following monomials from the right-hand side of the equation: $\phi_2 x_{t-1}$, $\phi_4 z_{t-1}$, $\phi_6 x_{t-1} z_t$, $\phi_8 x_{t-1} z_{t-1}$.

$$\Delta y_t = \phi_0 + (\phi_1 - 1) y_{t-1} + \phi_2 x_t + \phi_3 x_{t-1} + \phi_2 x_{t-1} - \phi_2 x_{t-1} +$$

$$+ \phi_4 z_t + \phi_5 z_{t-1} + \phi_4 z_{t-1} - \phi_4 z_{t-1} +$$

$$+ \phi_6 x_t z_t + \phi_7 x_{t-1} z_t + \phi_6 x_{t-1} z_t - \phi_6 x_{t-1} z_t +$$

$$+ \phi_8 x_t z_{t-1} + \phi_9 x_{t-1} z_{t-1} + \phi_8 x_{t-1} z_{t-1} - \phi_8 x_{t-1} z_{t-1} + \varepsilon_t$$

We then collect similar terms.

$$\begin{aligned}
\Delta y_t = \phi_0 + (\phi_1 - 1)y_{t-1} &+ \phi_2(x_t - x_{t-1}) + (\phi_3 + \phi_2)x_{t-1} + \\
&+ \phi_4(z_t - z_{t-1}) + (\phi_5 + \phi_4)z_{t-1} + \\
&+ \phi_6(x_t z_t - x_{t-1} z_t) + (\phi_7 + \phi_6)x_{t-1} z_t + \\
&+ \phi_8(x_t z_{t-1} - x_{t-1} z_{t-1}) + (\phi_9 + \phi_8)x_{t-1} z_{t-1} + \varepsilon_t \\
= \phi_0 + (\phi_1 - 1)y_{t-1} &+ \phi_2 \Delta x_t + (\phi_3 + \phi_2)x_{t-1} + \\
&+ \phi_4 \Delta z_t + (\phi_5 + \phi_4)z_{t-1} + \\
&+ \phi_6 \Delta x_t z_t + (\phi_7 + \phi_6)x_{t-1} z_t + \\
&+ \phi_8 \Delta x_t z_{t-1} + (\phi_9 + \phi_8)x_{t-1} z_{t-1} + \varepsilon_t
\end{aligned}$$

Now I add and subtract the following two monomials: $\phi_6 \Delta x_t z_{t-1}$ and $(\phi_7 + \phi_6) x_{t-1} z_{t-1}$.

$$
\begin{aligned}
\Delta y_t = \phi_0 + (\phi_1 - 1) y_{t-1} &+ \phi_2 \Delta x_t + (\phi_3 + \phi_2) x_{t-1} + \\
&+ \phi_4 \Delta z_t + (\phi_5 + \phi_4) z_{t-1} + \\
&+ \phi_6 \Delta x_t z_t + (\phi_7 + \phi_6) x_{t-1} z_t + \\
&+ \phi_8 \Delta x_t z_{t-1} + (\phi_9 + \phi_8) x_{t-1} z_{t-1} + \\
&+ \phi_6 \Delta x_t z_{t-1} - \phi_6 \Delta x_t z_{t-1} + (\phi_7 + \phi_6) x_{t-1} z_{t-1} - (\phi_7 + \phi_6) x_{t-1} z_{t-1} + \varepsilon_t \\
= \phi_0 + (\phi_1 - 1) y_{t-1} &+ \phi_2 \Delta x_t + (\phi_3 + \phi_2) x_{t-1} + \\
&+ \phi_4 \Delta z_t + (\phi_5 + \phi_4) z_{t-1} + \\
&+ \phi_6 \Delta x_t (z_t - z_{t-1}) + (\phi_7 + \phi_6)(x_{t-1} z_t - x_{t-1} z_{t-1}) + \\
&+ (\phi_8 + \phi_6) \Delta x_t z_{t-1} + (\phi_9 + \phi_8 + \phi_7 + \phi_6) x_{t-1} z_{t-1} + \varepsilon_t \\
= \phi_0 + (\phi_1 - 1) y_{t-1} &+ \phi_2 \Delta x_t + (\phi_3 + \phi_2) x_{t-1} + \\
&+ \phi_4 \Delta z_t + (\phi_5 + \phi_4) z_{t-1} + \\
&+ \phi_6 \Delta x_t \Delta z_t + (\phi_7 + \phi_6) x_{t-1} \Delta z_t + \\
&+ (\phi_8 + \phi_6) \Delta x_t z_{t-1} + (\phi_9 + \phi_8 + \phi_7 + \phi_6) x_{t-1} z_{t-1} + \varepsilon_t
\end{aligned}
$$

We can simplify the equation by setting:

$$\phi_1^* = \phi_1 - 1$$

$$\phi_2^* = \phi_2$$

$$\phi_3^* = \phi_3 + \phi_3 + \phi_2$$

$$\phi_4^* = \phi_4$$

$$\phi_5^* = \phi_5 + \phi_4$$

$$\phi_6^* = \phi_6$$

$$\phi_7^* = \phi_7 + \phi_6$$

$$\phi_8^* = \phi_8 + \phi_6$$

$$\phi_9^* = \phi_9 + \phi_8 + \phi_7 + \phi_6$$

Hence,

$$\Delta y_t = \phi_0 + \phi_1^* y_{t-1} + \phi_2^* \Delta x_t + \phi_3^* x_{t-1} + \phi_4^* \Delta z_t + \phi_5^* z_{t-1} +$$
$$+ \phi_6^* \Delta x_t \Delta z_t + \phi_7^* x_{t-1} \Delta z_t + \phi_8^* \Delta x_t z_{t-1} + \phi_9^* x_{t-1} z_{t-1} + \varepsilon_t \qquad \blacksquare$$

B.2 Additional Monte Carlo Simulation Results

Short-Run Effects of the Explanatory Variable of Interest, 100 observations

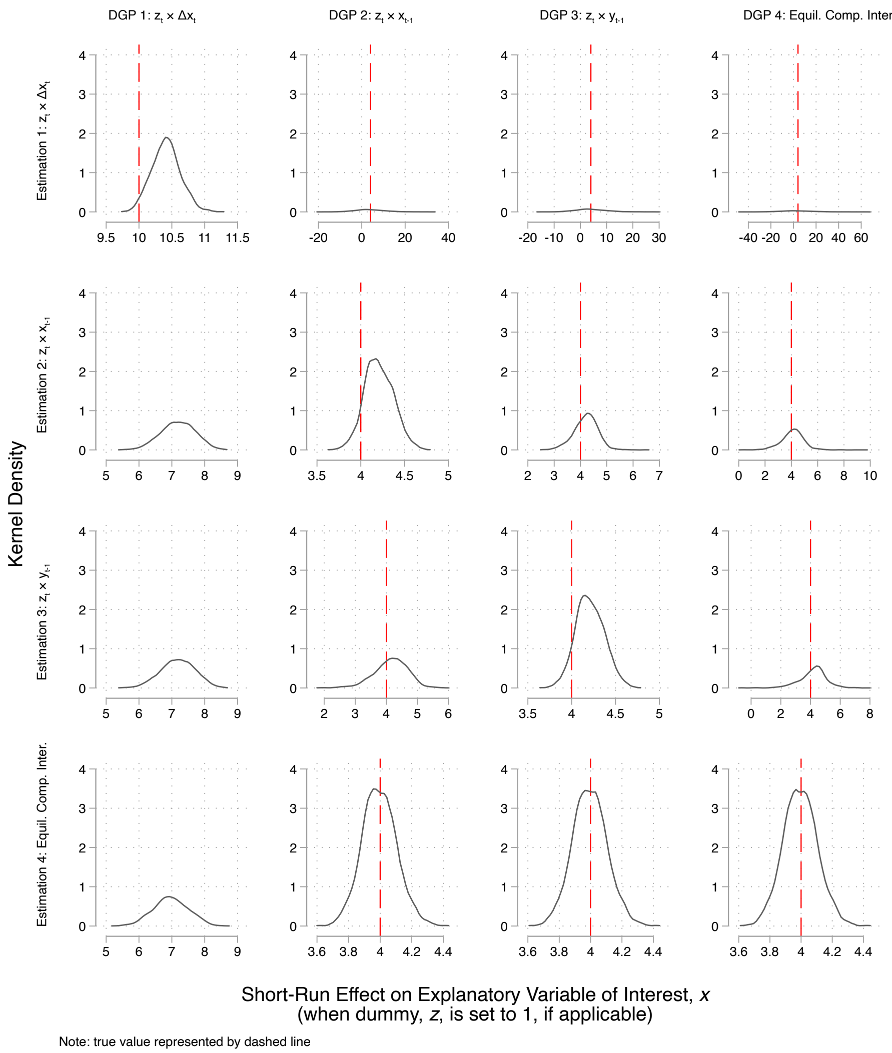

Figure B.1: Kernel Density plots of the short-run effects of the explanatory variable of interest, Δx_t, from 100-observation Monte Carlo simulation estimation results. (Columns represent different data generating processes and rows represent the respective estimation strategies.)

Long-Run Effects of the Explanatory Variable of Interest, 100 observations

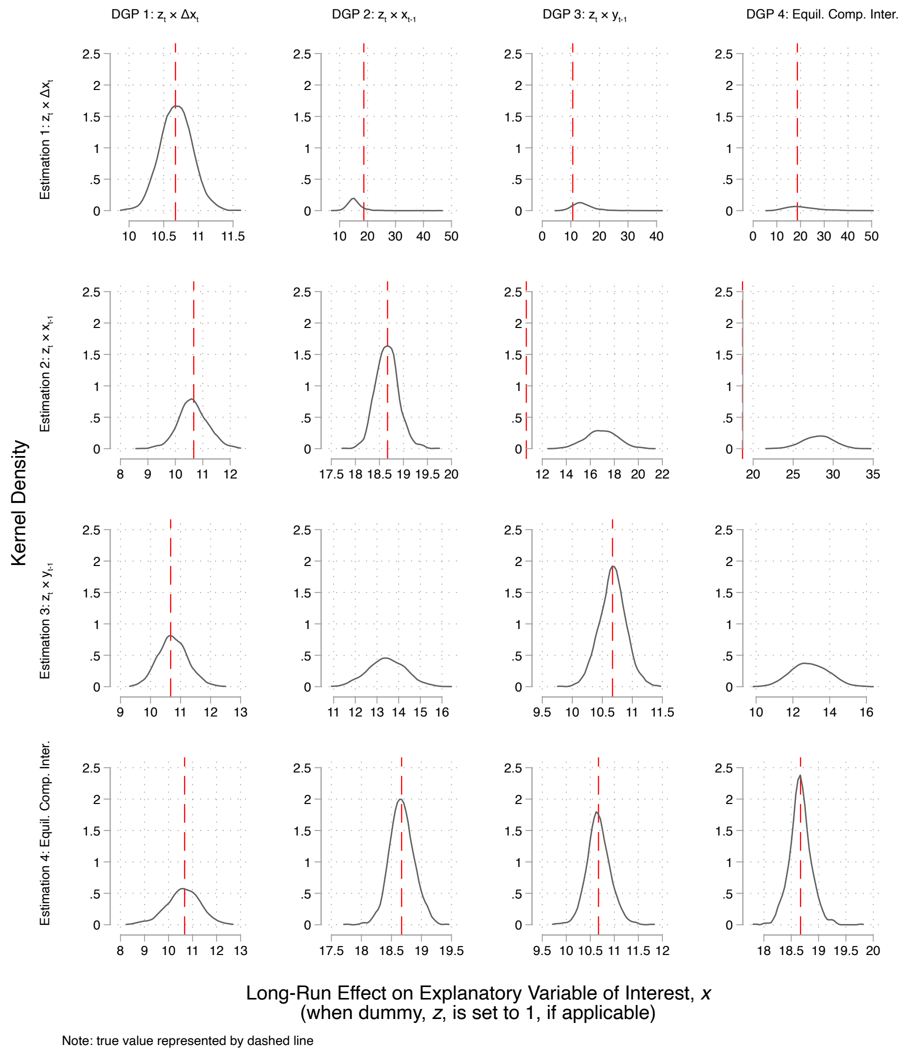

Figure B.2: Kernel Density plots of the long-run effects of the explanatory variable of interest, x_{t-1}, from 100-observation Monte Carlo simulation estimation results. (Columns represent different data generating processes and rows represent the respective estimation strategies.)

Error Correction Rate, 100 observations

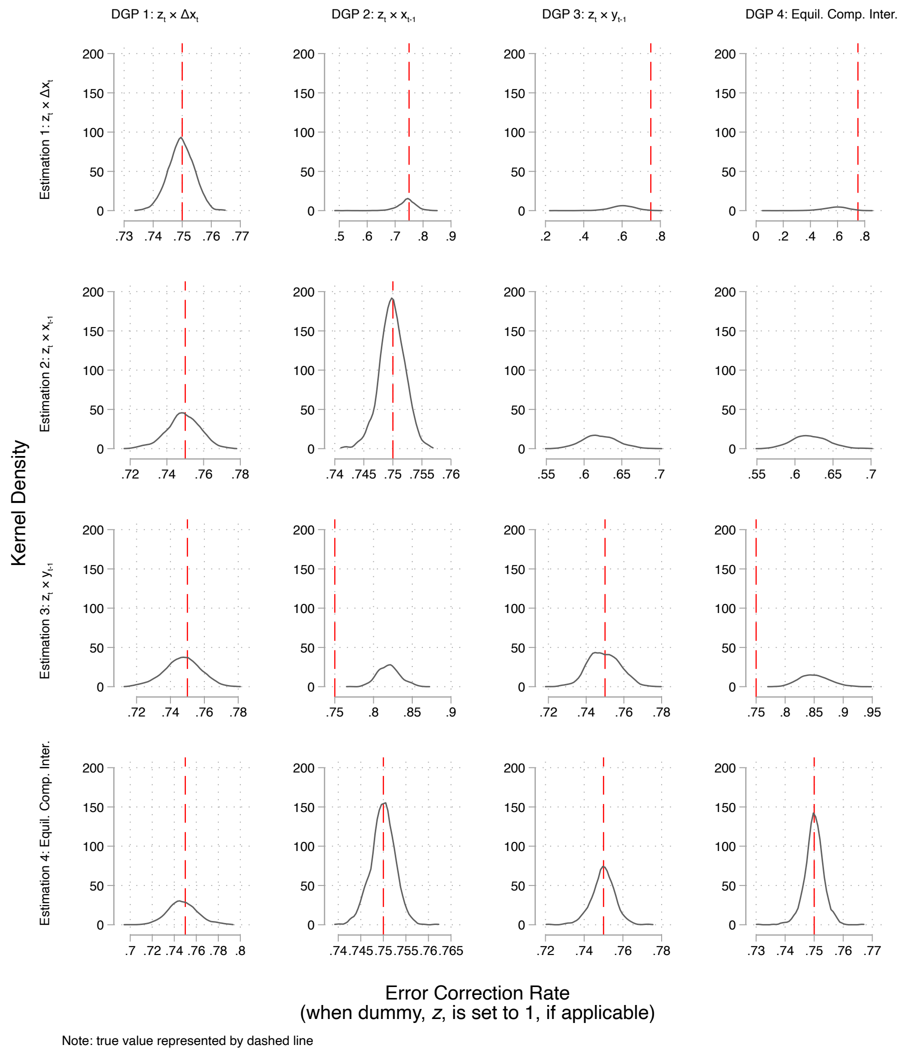

Figure B.3: Kernel Density plots of the error correction rate from 100-observation Monte Carlo simulation estimation results. (Columns represent different data generating processes and rows represent the respective estimation strategies.)

Root Mean Square Error, 100 observations

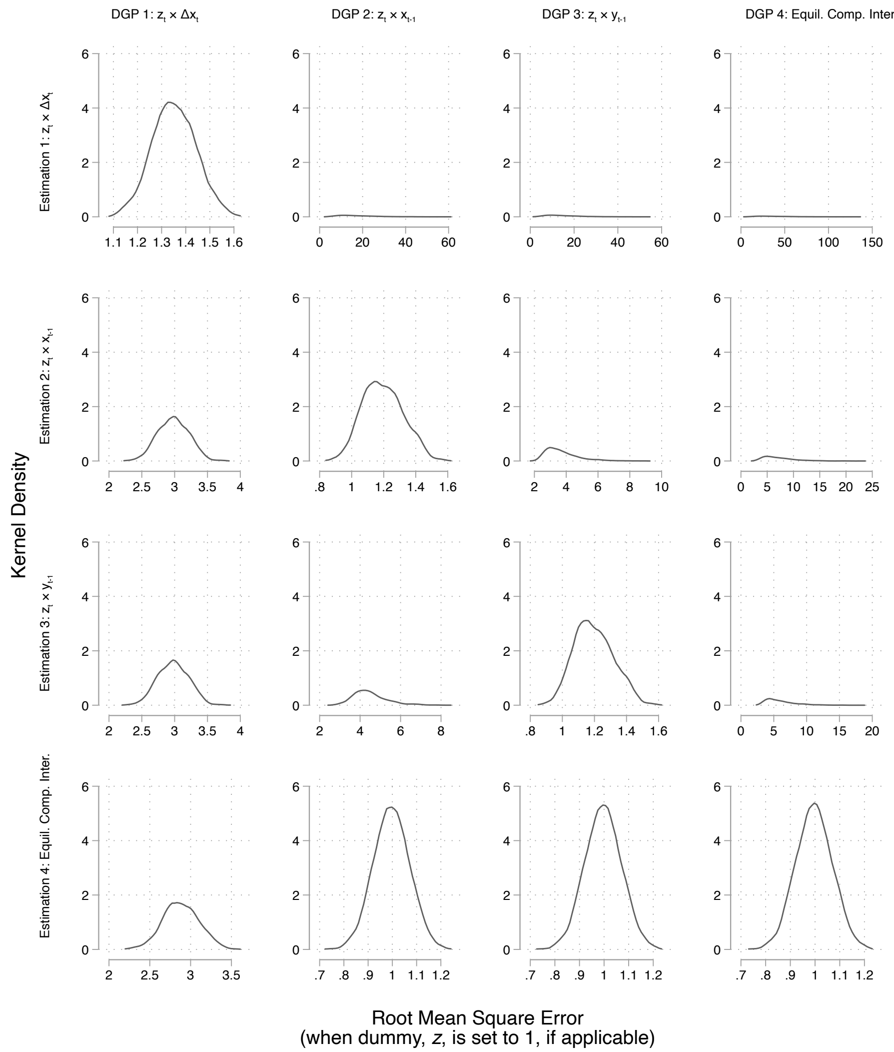

Figure B.4: Kernel Density plots of the root mean square error (RMSE) from 100-observation Monte Carlo simulation estimation results. (Columns represent different data generating processes and rows represent the respective estimation strategies.)

Short-Run Effects of the Explanatory Variable of Interest, 200 observations

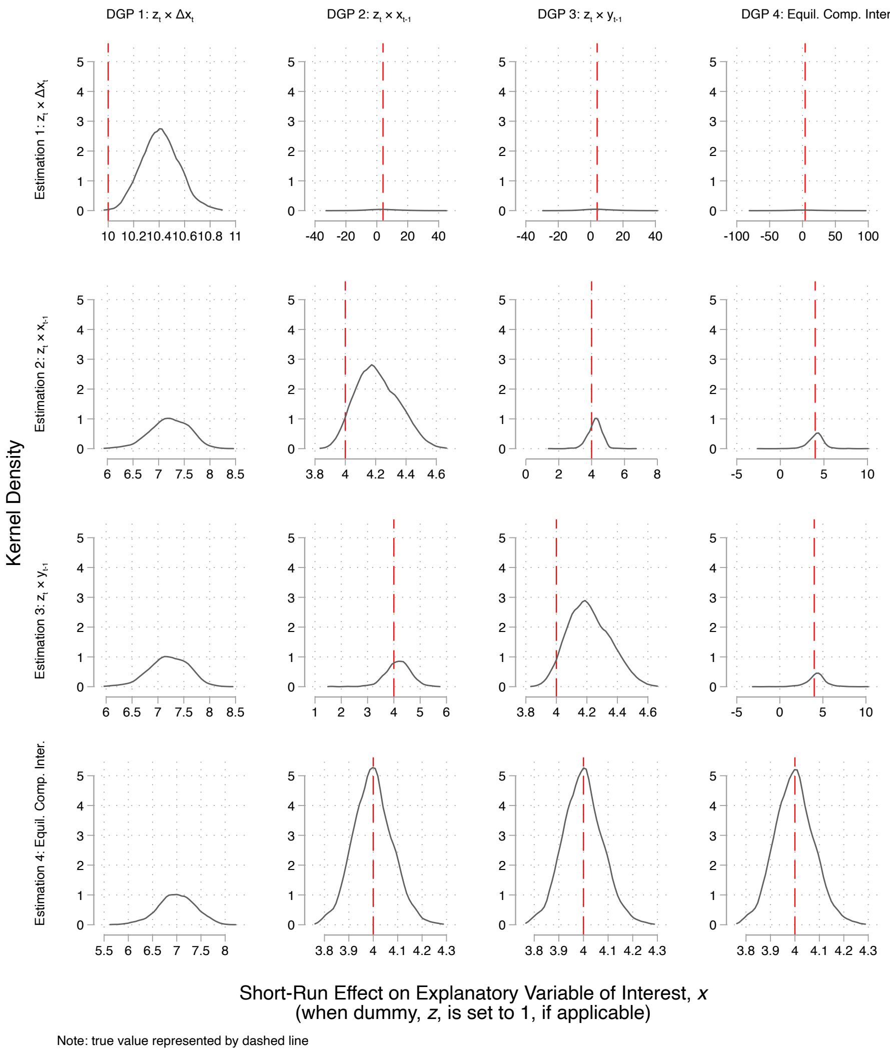

Figure B.5: Kernel Density plots of the short-run effects of the explanatory variable of interest, Δx_t, from 200-observation Monte Carlo simulation estimation results. (Columns represent different data generating processes and rows represent the respective estimation strategies.)

Long-Run Effects of the Explanatory Variable of Interest, 200 observations

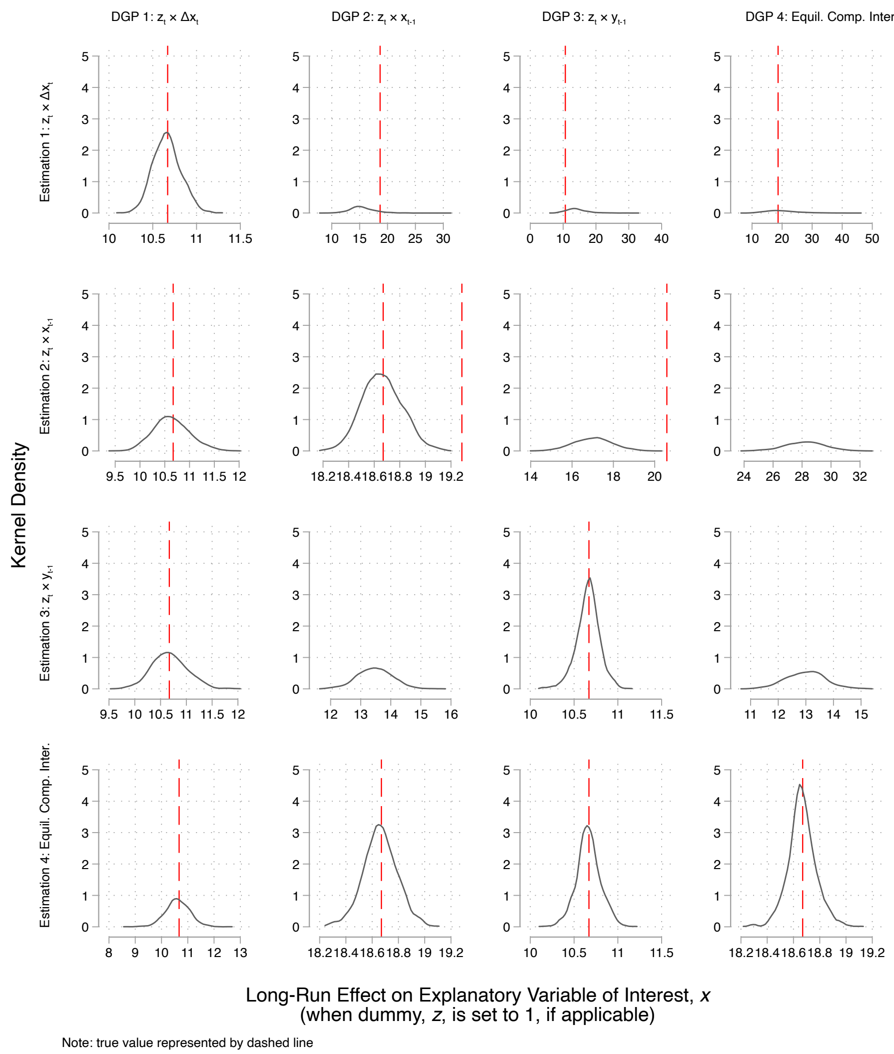

Figure B.6: Kernel Density plots of the long-run effects of the explanatory variable of interest, x_{t-1}, from 200-observation Monte Carlo simulation estimation results. (Columns represent different data generating processes and rows represent the respective estimation strategies.)

Error Correction Rate, 200 observations

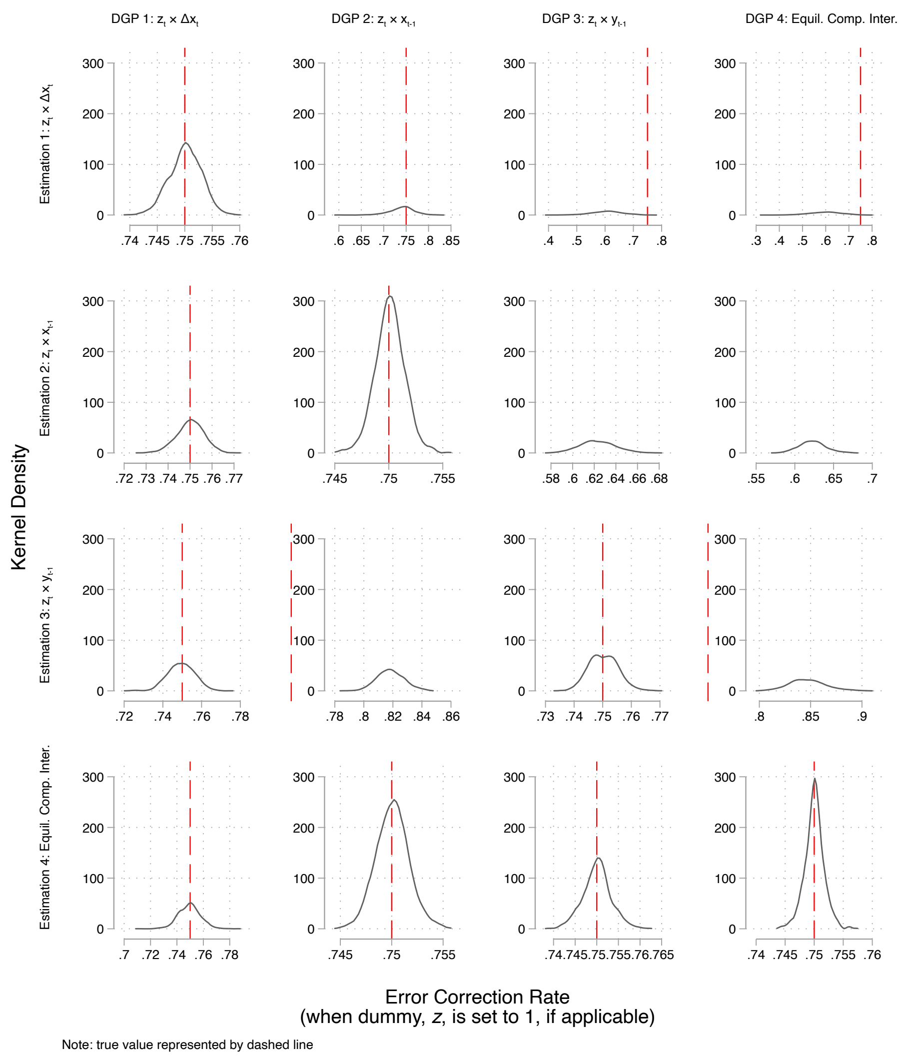

Figure B.7: Kernel Density plots of the error correction rate from 200-observation Monte Carlo simulation estimation results. (Columns represent different data generating processes and rows represent the respective estimation strategies.)

Root Mean Square Error, 200 observations

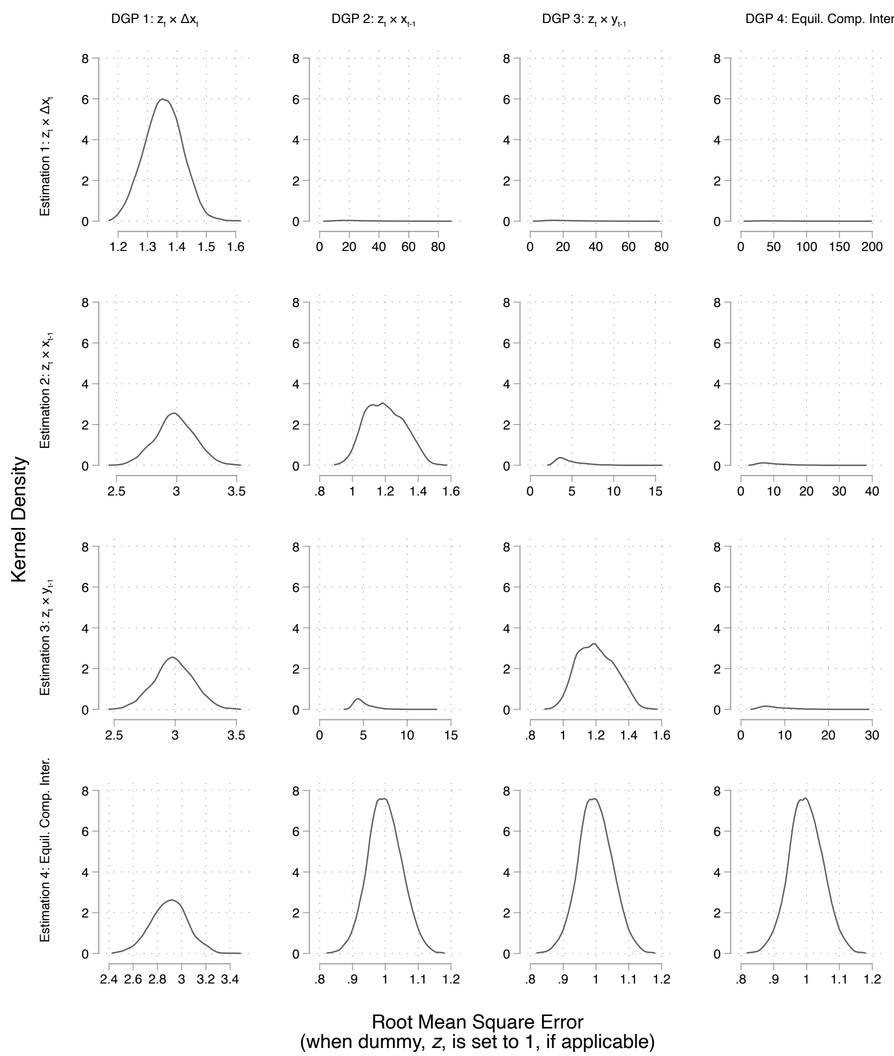

Figure B.8: Kernel Density plots of the root mean square error (RMSE) from 200-observation Monte Carlo simulation estimation results. (Columns represent different data generating processes and rows represent the respective estimation strategies.)