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by

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ABSTRACT

Today’s complex sociopolitical context features an increasing determent of fundamental bipartisan principles and negotiation at both the federal and state levels of government. A competitive political environment akin to post-Civil War times, amplified by growing partisan polarization and politicians’ quest for party allegiance and self-reward, pervasively discourages productive compromising efforts to work across the isle. We believe this hinders government’s sole and rather straightforward fiscal duty: to provide stable, healthy, and predictable economic conditions for its constituents. Credit ratings offer a window into the interaction of public policy, political uncertainty, and economic performance, which all lie at the nucleus of the political economy. Present literature provides a variety of studies concerning the adverse economic effects of partisan gridlock on fiscal efforts. Very few, however, go so far as to specifically examine this at the state level, let alone the inaugural political realm for many of today’s congressional members: the state house.

We reference previously successful models similar to this nature and uniquely construct a polychotomous ordinal dependent variable for several multinomial ordered probit models. These models seek to explain the economic side effects of three indicators of political instability: state government competitiveness, polarization in the lower house, and party control in the lower house. We conclude substantial insignificance for two out of the three indicators, primarily due to the structure of our dependent variable. Our analysis also provides evidence that, on average, Republican control of the lower house increases the chances of a higher credit rating and economic stability within a given state.
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“There is an unavoidable fact about legislating in a democratic system. No single person, faction, or interest can get everything it wants. Legislating inevitably means compromising, except in the rare circumstances when consensus is so strong that one dominant view can prevail with ease.”

– Robert Kaiser 2013, pp. 174
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I. INTRODUCTION

Today’s complex sociopolitical context features an increasing determent of fundamental bipartisan principles and negotiation at both the federal and state levels of government. The American system – separation of powers, bicameralism, and all – is unique to any other government infrastructure in the world in that it requires an intentional but extraordinary level of consensus to pass new legislation\(^1\), a process that partisan gridlock only makes less efficient. Given today’s poignant government role in virtually every capacity of the economy, this political phenomenon poses a substantial and growing risk to the economic stability of a government’s constituency. Before exploring its exact ramifications, we first need to look at what gridlock looks like and what causes it.

1. Partisan Gridlock in the U.S. Congress

The United States Congress serves as the arena of American democracy; a place where senators and representatives have come together for centuries to passionately discuss solutions to our nation’s most pressing issues. It is the heart and soul of our federal political infrastructure and the exclusive domain where citizens bind their closest connection to the federal government. The bimodal body’s constant rotation of 535 members from 435 distinct districts and 50 unique states has beckoned uninterrupted ridicule in support of Zelizer (2004)’s modern impression: “an amorphous, messy, and chaotic body […] full of corrupt individuals and crooked interest groups.” In order to prevent what the Founders called a “mischief of faction,” they intuitively designed such a

\(^1\) Krehbiel (1999).
structure to preserve the status quo and bring out the best – and only the best – of societal consensus. Even so, the difficulty of passage has grown to a gross, unnecessary magnitude that has debilitated our national economy and contributed to one of the slowest economic recoveries in our nation’s history.

Only recently has gridlock become a warranted occurrence for those in Congress. The body fluctuated from extreme fragmentation to strong partisan centralization starting the late 1970s. Optimistic approaches to end growing partisanship, like reforms to reduce filibusters, were unsuccessful. In the mid-1980s, “party leaders replaced committee chairs as the center of decision-making in the House and the Senate” (Zelizer, 2004, p. 618), making each party more or less homogenous ideologically. Southern conservative Democrats diminished, GOP moderates became almost extinct, and both parties shifted to their left and right, respectively. By the late 1980s, there was clear evidence that congressional polarization had begun. House Speaker Jim Wright (D-Texas) threatened to have Chairman of Ways and Means Dan Rostenkowski (D-Ill.) removed from his position for publishing a newspaper article that criticized a Democratic tax initiative. A young Republican representative from Georgia, Newt Gingrich, publicly blasted party member and president, George H. W. Bush, after Bush agreed to a tax increase in 1990. This was just the beginning to a now obvious national embarrassment.

According to Zelizer (2004), congressional majorities offer a concrete basis for analyzing gridlock because of data availability and measurement simplicity. Since the 92nd Congress (1971-1973), the average Senate majority has been 9.8 senators – half the average of the period between the 57th United States Congress (1901-1903) and the 92nd

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Congress (19.4 senators). A similar trend can be observed in the House of Representatives, where the average majority since the 92nd Congress (66.9 members) is also smaller than the average majority up until 1971 (79.5 members).

Nearly every federal election cycle since the 1970s has narrowly held out the possibility of an institutional shift in partisan control. We live in a political era that’s reminiscent of the Gilded Age (1876-1896) – a time when alternating party majorities, close elections, and ferocious party conflict (Binder and Lee, 2013) hinged on issues of social conflict and states’ rights following a civil war that nearly ripped apart the nation. Despite the era’s rapid economic growth due to industrialization and an immigrant labor force, this isn’t exactly the political comparison we should be proud of.

2. Partisan Polarization Pathology: Cause and Effect

The causes of partisan polarization are not straightforward, which makes the phenomena both difficult to measure and remedy. The general public seems to believe that the cause lies somewhere between the amount of money in the current political system and gerrymandering4, yet academics are still speculating whether campaign finance corruption is a significant factor. They also claim only ten to fifteen percent of increased polarization can be attributed to the redrawing of congressional districts (Hetherington, 2012). Perhaps it’s the general public itself postulating polarization’s existence. Pew’s (2012) study shows that the average partisan gap amongst citizens has

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nearly doubled over a twenty-five-year period – from ten percentage points in 1987 to 18 percentage points in 2012 – yet many are still critical of their respective party’s work.\textsuperscript{5}

\textbf{Cultural Polarization.} Social psychology’s morality principle tells us that our ethics and fundamental distinction between what is right and wrong binds us into like-minded groups. Like Pew (2012), Bishop (2008) finds that Americans have been residentially segregating themselves by political party over the past thirty years. With heterogeneity and political coexistence quickly becoming a thing of the past, like-minded groups have assumed a disregard of objective reality and increased polarity. These groups become more extreme than any one individual within it. This is an unfortunate since humans are the species on the planet that have the intellectual capacity for large-scale cooperation without a need for kinship (Haidt 2012; Orzag 2012).

Other scholars attribute gridlock to the realignment of Southern politics,\textsuperscript{6} others to increasing economic and social inequality,\textsuperscript{7} or party-enhancing legislative reforms.\textsuperscript{8} Barone (2001) points to growing religious and cultural divides as the polarizer and instigator of a more passionate and heated political discourse. Fiorina (2005) challenges this conventional wisdom by arguing an “inside the beltway” phenomenon is taking place. From this perspective, the mass isn’t polarized at all; it is closely divided, but not deeply divided. Fiorina and others\textsuperscript{9} maintain that political moderation and pragmatism is only at risk at the elite level; yet state legislatures have polarized as well (roughly two-

\footnotesize
\begin{itemize}
\item \textsuperscript{5} Pew’s (2012) data claims that nearly 71 percent of Republicans and 58 percent of Democrats say that have not stood up for their traditional positions.
\item \textsuperscript{6} McCarty, Poole, and Rosenthal (1997; 2003); Perlstein (2002); Stonecash, Brewer and Mariani (2003).
\item \textsuperscript{7} McCarty, Poole, and Rosenthal (1997; 2003); Stonecash, Brewer, and Mariani (2003); Hicks (2003); Rosenthal (n.d.).
\item \textsuperscript{8} Rohde (1991); Synder and Grosclose (1999).
\item \textsuperscript{9} Hetherington (2001) and McCarty, Poole, and Rosenthal (2003).
\end{itemize}
thirds of them\textsuperscript{10}). Such polarization implies a fundamental structural problem at the mass level.

Scholars have paid attention to gridlock’s rather intuitive effects on sociopolitical elements like the level of trust in government and public policy outcomes; however, virtually none have focused specifically on its effect on economic stability in U.S. states.

3. Partisan Gridlock in the States: An Empirical Analysis

One of Congress’ most recent debacles was the debt ceiling standoff in 2011 that rattled markets and resulted in a S&P (Standard and Poor’s) downgrade of the government’s AAA credit rating\textsuperscript{11,12}. This political impasse proved that gridlock unequivocally poses potential risk to securing fiscal policy and thus economic stability. This ascertains the primary motivation of our empirical work and hypothesis: does gridlock – characterized by competitiveness, dividedness, and party control – pose a similar risk for the U.S. states?

Both Poterba (1994) and Alt and Lowry (1994) show that increased propensity for gridlock in state legislatures can pose significant consequences for policy outcomes. This is especially apparent when responding to economic shocks, because of the need for additional (a) political bargaining, and (b) heightened disagreement over appropriate responses.

Pallay’s (2014) recent study shows the division of state governments in real numbers. From 1992-2013, the average state government was divided over half the time

\textsuperscript{10} Orzag (2012).
\textsuperscript{11} Lawder (2013).
\textsuperscript{12} This alone broadly justifies the economic significance between the relationship of partisan polarization and economic condition. We will attempt to provide an argument for statistical significance through empirical analysis.
(57.2 percent), while Republican (24.9 percent) or Democratic trifecta\(^{13}\) (22.4 percent) characterized the remaining composition. Over the same period, the average state legislature was divided nearly one-quarter of the time (23.3 percent), while Democrats (41.1 percent) and Republicans (35.6 percent) controlled the rest\(^{14}\). Democrats controlled state houses for 53.5 percent of the time, while Republicans held the lower chambers for 44.8 percent of the time.

This represents just a portion of the literature pertaining to the study of gridlock at the U.S. state level and we hope to add to it by studying this intriguing relationship. In order to empirically consider gridlock using congressional bodies, our analysis requires more than one observation (the U.S. Congress), and thus, we turn to state governments. This comes at a cost, however. Since each state government is economically unique, its controlling political parties hold unequal ideologies (i.e. Massachusetts Democrats vs. Mississippi Democrats). Its governments also consider drastically different policies that seek to remedy drastically different political and economic issues\(^{15}\).

\(^{13}\) A trifecta refers to a particular time when one party controls the governorship and both chambers of the state’s legislature (Pallay 2014).

\(^{14}\) The most competitive state legislatures have been Delaware, Nevada, and New York, while the most noncompetitive legislatures have been Hawaii (D), Idaho (R), Maryland (D), Massachusetts (D), Rhode Island (D), Utah (R), West Virginia (D), and Wyoming (R) (Pallay 2014).

\(^{15}\) Shor (2014).
II. LITERATURE REVIEW AND SCOPE

1. Measurements of Economic Stability

Economic stability in the U.S. states is difficult to measure. The existing literature, its various units of measurement, and proposed strategies fortify this observation.

Pallay (2014) explores the correlation between partisan control and a state’s economic performance, proxied by a “state quality-of-life” (SQLI) metric. This index includes the following ranked indicators:

- state credit ratings;
- governance ranking;
- employment/unemployment;
- high school graduation rates;
- poverty rates;
- personal income per capita;
- real GDP per capita;
- state government spending to GDP;
- state and local tax burden;
- and unfunded pension liabilities due per capita.

He uses an ordinary least squares (OLS) bivariate regression to find correlations between state economic performance rankings and partisan control. He then introduces multiple

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16 Pallay (2014) emphasizes repeatedly that the trends in the data are intended to describe correlation not causation, however they do provide some insight.
17 Pallay (2014) uses varying weights for theses and doesn’t seem to address exactly how these weights originated or how they affect his analysis. The reasoning behind these different weights is heavily influential in his results.
18 Badenhausen (2013); Cohn (2013); Donlon (2013); Sauter et al. (2013).
19 BLS (2013); BLS and DOL (2014).
20 UHF (2013).
21 Census and BLS (2012).
22 BEA (2013).
26 Eucalitto (2014).
fixed-effects probit regressions\textsuperscript{27}, using a dependent variable of state quality-of-life ranking for each year, and both contemporaneous and lagged independent ranking of overall partisanship.

Arnett’s (2014) method draws on an approach laid out by Groves, Godsey, and Shulman (1981). Using fiscal year 2012, he creates “cash, budget, long-run, and service-level solvency indices to measure the dimensions of [a state’s] fiscal condition.” The cash solvency component reflects the government’s liquidity, overall cash-management, and ability to pay interest payments on time. Budget solvency concerns the government’s ability to avoid deficits by meeting revenue with spending obligations. Long-run solvency represents the government’s capacity to address all of its costs, including long-term obligations like pensions. Service-level solvency is determined by a number of short- and long-run factors, including the size of the state’s revenue base. Her model seeks to describe a state government’s ability to meet financial and service obligations, much like credit ratings, but instead a “transparent and nuanced measure” (Arnett, 2014).

\textbf{Credit Ratings}. Credit ratings offer an independently opinionated, yet credible, measurement of a state’s economic stability and outlook in a given year. More specifically, credit ratings of general obligation bonds (GBO) are based on the ability and willingness of an issuer, such as a state government, “to meet its financial obligations in full and on time\textsuperscript{28}” (S&P, 2011, p. 3). Three rating agencies dominate the credit rating

\textsuperscript{27} The ranking form of Pallay (2014)’s SQLI index severely restricts the analytical power of a series regression using panel data.

\textsuperscript{28} For reasonable protection against criticism, credit agencies like S&P (2011) claim that market participants who are using ratings for investment and business decisions can’t fully rely on them since the ratings don’t explicitly guarantee repayment or no default. We adopt this notion as well, in that because our analysis is dependent on the plethora of different opinions analyzing such credit risks, we can’t ultimately guarantee a completely unbiased dependent variable.
market and provide the majority of GBO ratings: Standard and Poor’s, Fitch, and Moody’s. Since all are evenly well-established rating options, there are several approaches that seek to converge the three.

Depken and LaFountain (2006) and Schelker (2009; 2012) simulate economic performance and stability with the construction of a single credit rating measure based on ratings from the three principal rating agencies (S&P, Moody’s, and Fitch). After assigning a quantitative categorical measure to each rating, the numerical rating for state $i$ in year $t$ by rating agency $j$ is $R_{itj} \in \{1, ..., N_j\}$, where 1 refers to the highest GOB rating and $N_j$ the lowest rating by agency $j$. Next, scores $R_{itj}$ are normalized by dividing them by the number of possible rating, $N_j$, for each rating agency $j$ ($R'_{itj} = R_{itj}/N_j$).

The scores are then averaged to obtain a normalized overall rating $R'_{it}$ for each state $i$ in year $t$. We follow a similar approach in constructing our dependent variable in an attempt to incorporate the varying degrees of analysis conducted by each agency.

Significant issues with data arise when considering the difference in analyses by each of the three rating agencies. Ammar et al. (2001) observe that S&P typically relies on economic factors such as employment, income growth, income inequality, and poverty, while Moody’s and Fitch consider factors like fiscal policies and institutional considerations. Additionally, the analytical weight delegated to each factor relies on the analysts’ judgment. These factors change overtime and some factors are only considered in specific states (e.g. tourism in Maine).

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29 Depken and LaFountain (2006) change the sign on these measures to negative in order to provide a more intuitive interpretation of the regression results. We find it easier to leave the categorical measures as is and account for this accordingly in our interpretation of the model’s estimates.
31 Uyar and Escarraz (1995); Ammar et al. (2001).
32 Ammar et al. (2001).
Most agencies carry out a number of precautions\(^{33}\) to avoid conflicts of interest, since the issuer typically pays the agency to rate their issuances. One clear bias arises when considering the lack of a diversity within the credit rating agency industry, where S&P, Moody’s, and Fitch account for somewhere between 90 to 96 percent of all outstanding financial obligation ratings\(^{34}\). Although the three agencies disagree at times, they are generally very close to each other in consensus; thus we believe our approach to aggregating ratings like Depken and LaFountain (2006) won’t present any significant risk of bias.

Fehr (2012) provides clear insight into the relationship between a state’s economic environment and its credit rating by studying S&P ratings from 2001-2012. His work concludes that rating upgrades (or downgrades) can be attributed to positive (or negative) changes in a particular state’s economy. For example, Massachusetts improved from AA- in 2001 to AA+ in 2012 by diversifying its manufacturing sector. Louisiana improved from A in 2001 to AA in 2012 because of an infusion of federal dollars following the Hurricane Katrina relief effort. Conversely, Michigan’s shrinking auto industry brought its AAA rating in 2001 down to AA- in 2012. California’s continued struggle to balance its budget (following strict tax limits in 1978) saw its credit rating fall from A+ in 2001 to A- in 2012.

Credit ratings capture a host of economic indicators of stability and risk\(^{35}\).

Scholars have used ratings to try and extrapolate these indictors, which include:

\[
\text{annual government finance figures}^{36,37};
\]

\(^{33}\) For S&P (2011), these safeguards include physical separation between the analysts and personnel who negotiate business terms (similar to newspapers’ editorial and advertising functions) and an extensive committee review and assessment procedure.

\(^{34}\) White (2013).

\(^{35}\) Ingram, Brooks, and Copeland (1983).
political market indicators\textsuperscript{38};
fiscal transparency\textsuperscript{39};
political turnover\textsuperscript{40};
corruption\textsuperscript{41};
new debt interest costs\textsuperscript{42};
and levels of debt\textsuperscript{43}.

Debt plays an essential role in establishing credit ratings. This is because state governments need access to debt in order to hedge fluctuations of tax revenues over the course of one or more fiscal years. Such flexibility allows governments to meet short-term and long-term expenditure obligations. A state government’s cost of borrowing credit has a significant impact on a state’s fiscal health or condition. Perceived risk of default accompanies a government’s ability to repay the debt’s principal and interest\textsuperscript{44}, and when the risk of default is high, credit ratings are worse. We make sure to include a measure debt in our analysis to account for this established relationship.

2. Measurements of Political Stability

\textbf{Competitiveness and Divided Government.} A central research topic within the study of state politics has been the degree to which state political systems exhibit

\textsuperscript{36} For examples, see the following: Aronson and Marsden (1980); Parry (1983); Sharp (1986); Bahl and Duncombe (1993); Clingermayer and Wood (1995); and Lowry and Alt (2001).
\textsuperscript{37} Indicators include such things as debt to revenue, expenditures to population, intergovernmental aid as a proportion of total revenue, and the existence of budget deficits (Krueger and Walker 2008).
\textsuperscript{38} Perry and Robertson (1998).
\textsuperscript{39} Arbatli and Escolano (2012).
\textsuperscript{40} Krueger and Walker (2008).
\textsuperscript{41} Depken and LaFountain (2006).
\textsuperscript{42} Johnson and Kriz (2005).
\textsuperscript{43} Liu and Thankor (1984); Clingermayer and Wood (1995); Uyar and Escarraz (1995); Kiewiet and Szakaly (1996); Ellis and Schansber (1999).
\textsuperscript{44} Stallman, Deller, Amiel, and Maher (2012).
competitiveness, from the consequences of varying levels, to how to measure it\textsuperscript{45}. Taken together, the literature presents the notion that political institutions with more frequent changes in party control (i.e. divided or more competitive) pose an adverse effect to budget and legislative outcomes (although some scholars conclude otherwise\textsuperscript{46}).

Holbrook and Van Dunk (1993) offer the literature’s first applicable measurements by focusing on electoral competitiveness through average margins of victories in legislative elections. A district level measure (K. E. Klarner, personal communication, March 19, 2014) called the Ranney index\textsuperscript{47} offers a similar approach, but adds a slight complexity by using four separate components:

a. the proportion of legislative seats Democrats held in a state’s senate;
b. the proportion of legislative seats Democrats held in a state’s lower house;
c. the proportion of votes attained by the Democratic candidate for governor in the last gubernatorial election;
d. the average number years Democrats have controlled all three components (senate, house, and governorship) over the time period to which the indicator is being constructed for.

The Ranney index has been utilized as an explanatory measure of political competitiveness for analyses attempting to explain various state political and economic phenomena. These topics include:

- welfare policy\textsuperscript{48};
- budget amounts\textsuperscript{49};
- organizational strength of state parties\textsuperscript{50};

\textsuperscript{45} Examples include: Alt and Lowry (1994); Anderson, Lassen and Nielsen (2010); Binder (1999); Bohn and Inman (1994; 1996); Clarke 1998; Clingermayer and Wood (1995); Conley (2002); Cummins (2012); Klarner, Phillips, and Muckler (2010); Kousser (2010); McCubbins (1991); Mayhew (2005); Poterba (1994; 1996); and Poterba and Rueben (1999).
\textsuperscript{46} Gilligan and Matsusaka (1995; 2001).
\textsuperscript{47} Ranney and Kendall (1954); Ranney (1965)
\textsuperscript{48} Carmines (1974); Dawson and Robinson (1963).
\textsuperscript{49} Sharkansky (1968); Cnudde and McCrone (1969).
lobbyist regulation, diversity, density, and mobilization\textsuperscript{51};
policy liberalism\textsuperscript{52};
child support enforcement\textsuperscript{53};
bureaucracy\textsuperscript{54};
and partisan composition, competitiveness, and gridlock\textsuperscript{55}

Not the strongest of measurements, the index poses a host of potential problems if used as an overall measurement of competitiveness. This includes an oversimplification of four distinctly unique components. Klarner’s (2009) provides a felicitous example of this. Suppose that Democrats control 67 percent the senate; they also control 51 percent of the lower house; the last gubernatorial election saw the Democratic candidate attain 70 percent of the vote; and Democrats have held power in all three institutions, continuously, over the past six years. This situation would return a Ranney index score of 0.72, the same scores as a different situation in which Democrats control both legislative chambers by 59 percent. The later is much different than the former and poses a significantly biased measure of a state’s political competitiveness.

The index’s fourth component poses a similar inconsistency. Again, Klarner (2009) provides a relevant example. Republicans hold the state senate, Democrats hold the state house, and there is a Democrat governor over a four-year period. This may seem like a less competitive circumstance than a situation where party control of both the senate and the house switches in the middle of the period and the party of the governor is still Democratic; but the Ranney index score for both circumstances is identical. In

\textsuperscript{50} Cotter et al. (1984).
\textsuperscript{51} Opheim (1991); Lowery and Gray (1998); Lowery, Gray, Anderson, and Newark (1998); Berkman (2001).
\textsuperscript{52} Jackson (1992).
\textsuperscript{53} Keiser and Meier (1996).
\textsuperscript{54} Huber, Shipman, Pfäehler (2001).
\textsuperscript{55} Bowling and Ferguson (2001); Klarner (2009); Stallman, Deller, Amiel and Maher (2012).
addition, the index’s four components measure separate, unsystematic proportions (i.e. one measures party control, while another measures electoral margins).

**Partisan Ideology and Polarization.** As with political competitiveness and divided government, researchers have also studied the effects of another political phenomena: the distances between legislators’ ideologies. Numerous studies have argued that this “polarization” plays a significant role in creating budget gridlock and political instability.\(^{56}\)

Poole and Rosenthal (1984)\(^ {57}\) offer a multidimensional scaling measure called NOMINATE\(^ {58}\), that when dynamically manipulated and weighted (DW-NOMINATE), ranks ideological homogeneity based on the distance between parties. The measure analyzes individual voting behavior and preferences to infer degrees of polarization. Since data isn’t readily available for state legislators over a long period of time, the DW-NOMINATE method can only be used when analyzing polarization at the federal level (Congress\(^ {59}\)) and not state governments.

Shor and McCarthy (2011; 2013) offer an alternative approach to measuring polarization at the state level. Their research seeks to explain how trends in state legislatures are similar in many respects to Congress, but vary in terms of ideological spread and pace of polarization. Their method starts by formalizing McCarty, Poole, and

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\(^{56}\) Binder (1999); Clarke (1998); Cummins (2012); Kousser (2010); Masket (2007).

\(^{57}\) Poole and Rosenthal (1984; 1997); McCarty, Poole, and Rosenthal (1997).

\(^{58}\) An acronym for “Nominal Three-Step Estimation.”

\(^{59}\) For an example of such an analysis, see Theriault (2008); Layman, Carsey, and Horowitz (2006); and Thomson (2010).
Rosenthal’s (2009) two distinctions of polarization (intradistrict divergence and sorting). They calculate the difference in mean ideological points as

\[ E(x|R) - E(x|D) = \int\left[ E(x|R, z) \frac{p(z)}{\bar{p}} - E(x|D, z) \frac{1 - p(z)}{1 - \bar{p}} \right] f(z) dz, \]  

(1)

where \( x \) is an ideological point; \( R \) and \( D \) indicate what party the representative is from; \( z \) is a vector of district characteristics, which is distributed according to the density function \( f \); \( p(z) \) represents the probability that a district with characteristics \( z \) is Republican; and \( \bar{p} \) represents the average probability of electing a Republican. The average ideological difference between a Republican and Democrat who represent a district with characteristics \( z \) is \( E(x|R, z) - E(x|D, z) \). This captures intradistrict divergence, while \( p(z) \) captures the sorting effect. When there is a no sorting effect, \( p(z) = \bar{p} \) for all \( z \), so that

\[ E(x|R) - E(x|D) = \int [E(x|R, z) - E(x|D, z)] f(z) dz. \]  

(2)

The right side of Equation (2) represents the average intradistrict divergence component (AIDD) of polarization between the parties. When making cross-state comparisons, Shor and McCarty (2011) use a ratio of AIDD to total polarization, which “captures the amount of polarization that can be attributed to [ideological] divergence.”

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60 McCarty, Poole, and Rosenthal (2009) define intradistrict divergence as the difference between how Republican and Democratic legislators would represent the same district (Shor and McCarty 2011).

61 McCarty, Poole, and Rosenthal (2009) define the remainder of the polarization measure (or sorting effect) as the propensity for Republicans to represent conservative districts and Democrats to represent Democratic ones (Shor and McCarty 2011).

62 Ideal points are estimated by the Project Vote Smart National Political Awareness Test (NPAT) (a survey that measures a candidate’s self-assessed ideology), and supplemented by roll call voting data (Shor and McCarty 2011).

63 Characteristics include presidential vote, median family income, poverty rate, percentage of African-Americans and Hispanics, percentage of college graduates, and percentage of renters.

64 We utilize this measure in Equations (8) and (11) using Shor and McCarty (2013) data.
Their results reveal that intradistrict divergence accounts for a much larger proportion of legislative polarization at the state level, than it does in Congress. They also discover that California is by far the most polarized state legislature\textsuperscript{65}, while Rhode Island\textsuperscript{66} and Louisiana\textsuperscript{67} are the least polarized. Overall, states are distributed fairly evenly around congressional polarization\textsuperscript{68}. Roughly half are more polarized and some are less polarized. On average, Republicans have been getting extreme more rapidly than Democrats.

\textsuperscript{65} California is a unique case in which it is so polarized, but controlled predominately by one party. Democrats enjoy so much control that polarization’s typical effects of gridlock, stalemate, and instability are obsolete – the party has no obstacle to actual lawmaking in the legislature (Shor 2014).
\textsuperscript{66} In Rhode Island, most Democrats and Republicans are considered liberal (Shor and McCarty 2011).
\textsuperscript{67} Conversely, in Louisiana, most Democrats and Republicans are considered conservative (Shor and McCarty 2011).
\textsuperscript{68} This confirms that state legislatures are an adequate proxy for Congress in this analysis.
III. METHODOLOGY AND APPROACH

The influence of political instability and gridlock is notably absent in the study of bond ratings, except in the few cases of work such as Krueger and Walker (2008), Cheung (2008), Loviscek and Crowley (1990). Our main intention is to add as much commentary to the relationship as possible in order to propagate more evaluation and analysis into this most imperative topic. The following section provides insight into our overall process of finding data, constructing variables, and formulating models that best support our empirical hypothesis.

1. Data

**Dependent Variable (RATING).** We construct our baseline model around a dependent variable, RATING, which represents a panel data of long-term\(^69\) state general obligation bond ratings from 1992 to 2010\(^70,71\). We collected data primarily from Prunty and Sugden (2014), Moody’s (2014)\(^72\), and Fitch (2013)\(^73\), recording the standing rating\(^74\) for every applicable state in each year of the period. Inapplicable\(^75,76\) states include

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\(^69\) Outstanding long-term debt accounted for the majority (99.3 percent) of total outstanding debt incurred by states in 2007 (Stallman, Deller, Amiel, Maher 2012), which gives us reason to believe that long-term bond ratings provide a better measurement than short-term bond ratings.

\(^70\) Additional data was made available but sample periods were constrained by the model’s explanatory variables and their respective data ranges. We believed it was important to compile as much data for as many years as possible in order to overcome any time-related biases.

\(^71\) This study observes a wider range than any seen in present literature.

\(^72\) Some ratings were redefined due to the introduction of modifiers and some were recalibrated to global scale from municipal finance scale (Moody’s 2014).

\(^73\) Physical U.S. Census reports from each year of the period were used to confirm all ratings and to fill in any missing values.

\(^74\) Not simply rating changes or reevaluations.

\(^75\) States were deemed inapplicable if they had no or insufficient GOB data.

\(^76\) The selected contiguously inapplicable states pose a selection bias on the Northeast and South ratings considering four states each from the West and Midwest regions were omitted.
Arizona, Colorado, Iowa, Idaho, Indiana, Kansas, Nebraska\(^7\), and Wyoming, plus the exclusion of noncontiguous states Alaska and Hawaii\(^7\). This brought our total population to a total of 40 states and 608 observations\(^7\). Table 1 explains our process of grouping panel data together in ordered categories. For a more an explanation of this process, see the Model Specification section.

Table 1: Dependent Variable Data Classifications and Comparisons

<table>
<thead>
<tr>
<th>Rating Agency</th>
<th>Quantitative Value</th>
<th>S&amp;P Definition(^8)</th>
<th>Ordered Category</th>
<th>Aggregated Observations(^8)</th>
</tr>
</thead>
<tbody>
<tr>
<td>S&amp;P</td>
<td>Moody’s</td>
<td>Fitch</td>
<td></td>
<td></td>
</tr>
<tr>
<td>AAA</td>
<td>Aaa</td>
<td>AAA</td>
<td>Extremely strong capacity to meet financial commitments</td>
<td>Prime (1)</td>
</tr>
<tr>
<td>AA+</td>
<td>Aa1</td>
<td>AA+</td>
<td>Very strong capacity to meet financial commitments</td>
<td>High Grade (2)</td>
</tr>
<tr>
<td>AA</td>
<td>Aa2</td>
<td>AA</td>
<td>Mid-High Grade (3)</td>
<td>206</td>
</tr>
<tr>
<td>AA-</td>
<td>Aa3</td>
<td>AA-</td>
<td>High to Medium-Grade (4)</td>
<td>181</td>
</tr>
<tr>
<td>A+</td>
<td>A1</td>
<td>A+</td>
<td>Strong capacity to meet financial commitments, but somewhat susceptible to adverse economic conditions</td>
<td></td>
</tr>
<tr>
<td>A</td>
<td>A2</td>
<td>A</td>
<td></td>
<td></td>
</tr>
<tr>
<td>A-</td>
<td>A3</td>
<td>A-</td>
<td></td>
<td></td>
</tr>
<tr>
<td>BBB+</td>
<td>Baa1</td>
<td>BBB+</td>
<td>Adequate capacity to meet financial commitments but more subject to adverse economic conditions</td>
<td></td>
</tr>
<tr>
<td>BBB</td>
<td>Baa2</td>
<td>BBB</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

\(^7\) Nebraska was not included because of its unicameral legislature that includes no lower house distinction.  
\(^8\) This is the general practice for studies analyzing fiscal conditions in U.S. states (Alt et al. 2002) and it does not affect the results associated with political instability terms used in this study. Alaska and Hawaii are seen as outliers in many respects because these states depend disproportionately on federal transfers (Schelker 2012).  
\(^8\) Equations (8) and (11) also exclude California, in addition to the baseline exclusions, due to its outlying polarization referenced in Shor (2014).  
\(^8\) These observation counts represent the total number of calculated median credit ratings in each category over period 1992-2010, excluding those ratings from Arizona, Colorado, Iowa, Idaho, Indiana, Kansas, Nebraska, and Wyoming.
**Explanatory Variables of Interest.** Our three focal explanatory variables attempt to capture: a government’s competitiveness, its ideological spread, and its partisan preference. We focus our analysis on lower house portion of state legislatures for two out of the three variables for three reasons. First, election cycles are fairly consistent\textsuperscript{82} and this makes for a more varying measure (dissimilar to a legislature’s upper house). Second, the lower house mirrors our more familiar federal counterpart, Congress, as the heart and soul of state democracy. Lastly, the literature shows that U.S. state lower houses are rarely used when analyzing the relationship between political gridlock and economic stability.

1. *Political competitiveness (COMP).* We assume a variation of the Ranney index measure – a folded Ranney index\textsuperscript{83} – to measure the frequency at which states change partisan control. This folded feature uses the same Ranney index discussed in the Literature Review section, but calculates the distance of the Ranney score away from “0.5”\textsuperscript{84} C.E. Klarner claims is the best way to measure a state’s competitiveness (personal communication, March 19, 2014).

We hypothesize that political competition fuels partisan gridlock by raising the stakes on every policy dispute. This leads to an increase in the risk of economic instability, on average, in the state where high political competitiveness exists. When institutional majorities hang on small political advantages (e.g. avoiding government shutdown), party members and leaders are more likely to reject opportunities that grant

\textsuperscript{82} Term lengths are typically two years except for four states that have a term length of four years (Alabama, Louisiana, Maryland, Mississippi).

\textsuperscript{83} Klarner Politics data (Klarner, n.d.).

\textsuperscript{84} C.E. Klarner claims this is the best way to measure a state’s competitiveness (personal communication, March 19, 2014).
political legitimacy to the opposite party, even when the economic stakes are insurmountable and the standing option is a viable one. Instead, a competitive political environment initiates debilitating political dysfunction and instability and amplifies the partisan differences voters perceive between their party and the opposition (Binder and Lee 2013).

2. Partisan polarization (POLAR). We construct our measure of polarization in the state’s lower house using data from Shor and McCarty (2013). Polarization is represented by how far the center (median) of each party\textsuperscript{85} differs from one another\textsuperscript{86} \textsuperscript{87}. Since a larger divergence equates to increased polarization, we speculate that a larger POLAR will lead to a worse credit rating, and thus more economic instability.

3. Partisan control (PART). We implement Ranney data from Klarner (n.d.) to construct our partisan control variable. This variable represents the proportion of Democrats in control of the state’s lower house legislature in a given year. A score of “1” is defined as complete Democratic control and score of “0” is defined as complete Republican control.

On average, Democrats are more permissive, or tolerant, of government expenditures than Republicans\textsuperscript{88}. We argue that this leads to an increased risk of default on government financial obligations. Assuming this generalization of each party’s fiscal

\textsuperscript{85} Democrats and Republicans.
\textsuperscript{86} This is a common measure of legislative polarization (Shor 2014).
\textsuperscript{87} See Section II.2 for more information on exactly how this measurement is calculated.
\textsuperscript{88} We assume this action outweighs default mitigation by Democrats through enacting policies that raise tax revenue.
platforms, we hypothesize that an increase in the proportion of Democrats will lead to a decrease in the state’s credit rating.

4. Regional indicators (Dummies). The following codes were assigned to each state according to the four census regions (Klarner, n.d.):

\[
\begin{align*}
SOUTH &= 1; \\
WEST &= 2; \\
MIDWEST &= 3; \\
NORTHEAST &= 4^{89}. \\
\end{align*}
\]

Structural Explanatory Variables. Unlike Pallay (2014) or Arnett (2014), we only include a couple of fiscal control variables to avoid “overfitting” the model and ultimately invalidating each estimated effect due to multicollinearity\(^{90}\) (Schelker 2009; 2012). The literature suggests cases where researchers have employed several variables to account for various state economic figures. This is a rather intuitive approach considering the economic data is readily available and these figures certainly play a role in the credit rating process. However, state economic conditions are also correlated with our political measures of interest, and therefore, we decided to leave many of them out\(^{91}\). In order to control for differences in rating mechanisms of the three principal rating agencies and state-specific covariates, all regressions include two additional explanatory variables.

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\(^{89}\) The Northeast region indicator variable is omitted to define the reference (or base) group and avoid violation of perfectly multicollinear variables (MR5) (a.k.a. the dummy variable trap).

\(^{90}\) Schelker (2009; 2012).

\(^{91}\) As economic conditions worsen, political competitiveness and partisan polarization increase, and vice versa (Cummins, 2012).
1. Per capita income (PCINC). We follow Krueger and Walker (2008) in excluding a structural variable that captures changes in employment. Instead, we include a measure of personal income, which we believe serves as a better measure of long-term state economic health. Employment primarily captures short-term economic fluctuations, which doesn’t match our long-term dependent variable.

A lower credit rating also equates to a higher borrowing cost to taxpayers. This requires a higher interest-rate-premium to account for the increased level of default risk, which decreases disposable income, and thus nominal per capita income. In the long-run, a fall in personal income will impact consumption levels within a state. Decreased consumption will hurt the state’s economy and credit rating. Using a real measure of income did not change our empirical findings.

We again utilize the state economic data provided by Klarner (n.d.) and propose nominal per capita income (in $1s) as the first of our two structural variables. A term to account for population was not included, however the per capita measure allows us to account for population differences across states, and income carries a nonlinear relationship with credit rating, we make the model a partial linear-log function by taking PCINC’s natural logarithm. We suggest that a one percent change in income will increase the state’s probability of achieving a higher credit rating.

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92 This decision is going against the grain of many competent studies in the literature where unemployment rate is one of the strongest variables (Krueger and Walker 2008).
93 This is partially due to the fact that NPCI partially accounts for per capita discrepancies across states. Schelker (2012) claims that estimating population effects at such an aggregated level may not be representative for all classes of jurisdictions. He says that the “optimal size of jurisdictions depends on the context and analyzed dimension”
2. *Debt to gross state product* (*DEBTGSP*). Klarner (n.d.) provides us with our second structural variable: a debt to gross state product ratio. Since states take on debt in both good and bad markets\(^{94}\), this structural variable will not take away any influence of other explanatory variables. It accounts for the differences in each state’s leveraging habits, while also controlling for states that produce a lot more than others due to other, unspecified variables.

Table 2 summarizes statistical characteristics and a description of each variable included in our model.

<table>
<thead>
<tr>
<th>Variable</th>
<th>Population Min.–Max.</th>
<th>Population Mean (Std. dev.)</th>
<th>Description</th>
</tr>
</thead>
<tbody>
<tr>
<td>State credit rating (Long-term GOB) RATING</td>
<td>–</td>
<td>–</td>
<td>Ordinal polychotomous variable with four categories of aggregated credit ratings from S&amp;P, Fitch, and Moody’s.</td>
</tr>
<tr>
<td>Per Capita Income PCINCOME</td>
<td>$14,651 – $56,959</td>
<td>29,675 (7,889)</td>
<td>Nominal per capita income</td>
</tr>
<tr>
<td>Debt to Gross State Product DEBTGSP</td>
<td>1.59% – 22.76%</td>
<td>7.43% (3.9182)</td>
<td>Percentage of total debt outstanding to gross state product</td>
</tr>
<tr>
<td>State Government Competitiveness COMP</td>
<td>0.65 – 0.99</td>
<td>0.88 (0.0873)</td>
<td>Folded Ranney state competitiveness score (4 Years)</td>
</tr>
<tr>
<td>Lower House Legislature Polarization POLAR</td>
<td>0.45 – 3.21</td>
<td>1.26 (0.4557)</td>
<td>Shor-McCarty median ideological spread in the lower house</td>
</tr>
<tr>
<td>Lower House Legislature Party Control PART</td>
<td>0.23 – 0.92</td>
<td>0.55 (0.1494)</td>
<td>Percentage of Democrats in the lower house</td>
</tr>
</tbody>
</table>

\(^{94}\) Krueger and Walker (2008)
2. Model Specification

Our baseline approach follows a simplified nonlinear version of Cheung’s censored regression (2008) study of provincial credit ratings. We assume an ordered probit model with an ordinal polychotomous dependent variable, $y_{it}$, to estimate the max likelihood (MLE) of state credit ratings resulting from a latent measure of risk ($y^*_{it}$), or what we call, economic stability. The unobserved economic stability is a linear function of explanatory variables $x$, with parameter vector $\beta$, and an error term $\varepsilon$,

$$y^*_{it} = \beta'x + \varepsilon.$$  

(3)

The latent dependent variable, $y^*_{it}$, is characterized by an observed, qualitative and ordinal $y_{it}$ using a multinomial probit model. Ordinary least squares (OLS) regression estimation would misinterpret the differences between each rating because there is a nonlinear, unobserved movement from one qualitative rank to another. Instead, we use an ordered multinomial probit (OMP), which offers an estimation technique that assumes an order. It also assumes that the size of the difference between each ordered rank is not known and does not incorporate this into its configuration (as the OLS method would).

The OMP model observes our dependent variable, $RATING (y_{it})$, a quantitative panel of credit ratings for state $i$ in year $t$. We derive the variable from an aggregated qualitative pool of Fitch$^{95}$, Moody’s$^{96}$, and S&P GOB ratings, which range from AAA to BBB over the years 1992-2010$^{97}$. In order to construct a single-rating measure while

$^{95}$ Fitch Ratings is roughly one-third the size of both S&P and Moody’s in terms of ratings outstanding (White 2013).

$^{96}$ Moody’s was the originator of publicly available ratings, an innovation of its founder, John Moody, in 1909 (White 2013). However, S&P seems to be regarded as the market’s most credible rating agency; hence our preference of its ratings in our dependent variable aggregation process.

$^{97}$ This implies that the range of $y_{it}$ is limited to those ratings assigned to particular states during the years in the sample. Therefore, the probability distribution of a rating outside the range of observed ratings (e.g. BBB-) cannot be estimated (Cheung 1996).
accounting for missing data, we construct panel data using an approach\textsuperscript{98} similar to Depken and LaFountain (2006) and Schelker (2008; 2012). We take the median rating for each state $i$ in each year $t$ (when all three ratings are available), or the S&P rating (when only two ratings are available\textsuperscript{99}). Then, we assign this rating to the corresponding state $i$ in year $t$ in quantitative form with one and nine being the highest and lowest ratings, respectively (i.e. AAA, Aaa = 1; AA+, Aa1 = 2 … BBB, Baa2 = 9). Next, we group these ratings into four custom categories of roughly equal proportion\textsuperscript{100,101} to ease estimation inference. These groups are: (1) Prime (AAA, Aaa); (2) High Grade (AA+, Aa1); (3) Mid-High Grade (AA, Aa2); and (4) Low-High to Medium Grade (AA-, Aa3 to BBB, Baa2). Thus,

\begin{equation}
\begin{align*}
y_{it} &= (1) \text{Prime } \text{if } y_{it}^* \leq u_1, \\
&= (2) \text{High Grade } \text{if } u_1 < y_{it}^* \leq u_2, \\
&= (3) \text{Mid} − \text{High Grade } \text{if } u_2 < y_{it}^* \leq u_3, \\
&= (4) \text{Low} − \text{High to Medium Grade } \text{if } u_3 \leq y_{it}^*,
\end{align*}
\end{equation}

where $u$’s are unknown cut points that define the boundaries between credit rating groups. These parameters are estimated through maximum likelihood in conjunction with vector $\beta$, given that error $\epsilon$ is normally distributed and that its mean and variance is normalized to zero and one.

\textsuperscript{98} This approach, although intuitive and pragmatic, ignores the full information content of the ratings, as well as the differences between the rankings of the three different agencies, potentially inducing biases. A more accurate but complex approach is mentioned in Section V.

\textsuperscript{99} Not all GOBs are always rated by all three rating agencies and, therefore, some observation measurements relied on only one or two agencies. This was uncommon for the majority of all states and years, but it could not be avoided without magnifying the complexity of the model (Schelker 2012).

\textsuperscript{100} See Figures 1 and 2 for observation distributions.

\textsuperscript{101} Prime (195 observations); high grade (233); mid-high grade (365); and low-high to medium grade (210). See Table 1 for visualization.
Where $\phi$ is the cumulative distribution function of a normal distribution and given the probabilities,

\begin{align*}
Prob(y = 1) &= \phi(u_1 - \beta' x), \\
Prob(y = 2) &= \phi(u_2 - \beta' x) - \phi(u_1 - \beta' x), \\
Prob(y = 3) &= \phi(u_3 - \beta' x) - \phi(u_2 - \beta' x), \\
Prob(y = 4) &= 1 - \phi(u_3 - \beta' x),
\end{align*}

a maximum likelihood equation (MLE) can be formed to estimate the unknown $u$ and $\beta$ parameters:

\begin{align*}
L(y/x) &= \sum_{k=1}^{n} \left\{ Y_{1k} \ast \log \phi(u_1 - x_k' \beta) \\
&\quad + \sum_{i=2}^{4} Y_{1k} \ast \log[\phi(u_i - x_k' \beta) - \phi(u_{i-1} - x_k' \beta)] + Y_{4k} \\
&\quad \ast \log[1 - \phi(u_3 - x_k' \beta)] \right\}
\end{align*}

The indicator variable $Y_{tk}$ takes on the value of one if the realization of the $k^{th}$ observation $y_k$ is the $i^{th}$ rating, and zero otherwise. The log-likelihood function stated above is maximized using estimated $u$’s (cutoff points) and $\beta$’s (coefficients).\textsuperscript{102}

\textsuperscript{102} We agree with Cheung (1996) that this model can also be estimated, just as well, with an ordinary ordered logit.
The model describes probabilities of outcomes and therefore doesn’t directly describe the relationship between a $y_{lt}$ and covariates $x_j$. Since there are no obvious regression relationships at work between the observed random variable and the covariates, the effect ($\beta_j$) of a change in an explanatory variable ($x_j$) on the estimated probabilities of the highest and lowest ordered categories is certain. However, the impact on the estimated probabilities of intermediate classifications isn’t clear.

In other words, we know that if coefficient $\beta_j$ is negative, an increase in the conditional mean $\beta'x$ (by way of an increase in $x_j$) definitely decreases a state’s probability of being in the lowest rating group (AA- to BBB) and definitely increases a state’s probability of being in the highest rating group (AAA), but the impact on the estimated probabilities of classifying in between (AA+ or AA) can be in either direction.
The sum of all changes will remain zero, since the new probabilities must still add up to one, and the single-crossing condition applies\(^\text{103}\).

Figure 2: Ordered Category Observation Counts (After Grouping)

Using the variables specified above, we formulate, regress, and estimate the following six equations. Each equation attempts to distinguish between our three measurements of political stability (partisan competitiveness, polarization, and control) with and without regional indicator specification.

\[ y_{it}^* = \beta_0 + \beta_1 \text{COMP}_{it} + \beta_2 \ln(\text{PCINC}_{it}) + \beta_3 \text{DEBTGSP}_{it} + \epsilon_{it} \]  \hspace{1cm} (7)

\[ y_{it}^* = \beta_0 + \beta_1 \text{POLAR}_{it} + \beta_2 \ln(\text{PCINC}_{it}) + \beta_3 \text{DEBTGSP}_{it} + \epsilon_{it} \]  \hspace{1cm} (8)

\[ y_{it}^* = \beta_0 + \beta_1 \text{PART}_{it} + \beta_2 \ln(\text{PCINC}_{it}) + \beta_3 \text{DEBTGSP}_{it} + \epsilon_{it} \]  \hspace{1cm} (9)

\(^{103}\) For a variable with a positive coefficient, effects at \(\text{Prob}(1)\) will begin as negative, then change to a set of positive values, thus changing signs once. See Green and Hensher (2009) for more examination of the single-crossing property in ordered probit models.
\[ y_{it}^* = \beta_0 + \beta_1 \text{COMP}_{it} + \beta_2 \ln(PCINC_{it}) + \beta_3 \text{DEBTGSP}_{it} + \delta_1 \text{SOUTH} + \delta_2 \text{MIDWEST} + \delta_3 \text{WEST} + \epsilon_{it} \] (10)

\[ y_{it}^* = \beta_0 + \beta_1 \text{Polar}_{it} + \beta_2 \ln(PCINC_{it}) + \beta_3 \text{DEBTGSP}_{it} + \delta_1 \text{SOUTH} + \delta_2 \text{MIDWEST} + \delta_3 \text{WEST} + \epsilon_{it} \] (11)

\[ y_{it}^* = \beta_0 + \beta_1 \text{PART}_{it} + \beta_2 \ln(PCINC_{it}) + \beta_3 \text{DEBTGSP}_{it} + \delta_1 \text{SOUTH} + \delta_2 \text{MIDWEST} + \delta_3 \text{WEST} + \epsilon_{it} \] (12)

where \( i \) is an index of applicable states; \( t \) is a time index in years, ranging from 1992 to 2010; \( \text{RATING}_{it} \) is the most recent standing, aggregated credit rating of state \( i \) in year \( t \); \( \text{COMP}_{it} \) is the folded Ranney index score of state government \( i \) in year \( t \), ranging from 1992 to 2010; \( \text{Polar}_{it} \) is the Shor-McCarty AIDD to total polarization ratio measurement value for lower house state legislatures in state \( i \) in year \( t \), ranging from 1998 to 2008; \( \text{PART}_{it} \) is the proportion of Democratic party control in lower house state legislatures in state \( i \) in year \( t \), ranging from 1992 to 2010; \( \ln(PCINC_{it}) \) is the natural log of nominal per capita income for state \( i \) in year \( t \), ranging from 1992 to 2010; \( \text{DEBTGSP}_{it} \) is the debt to gross state product for state \( i \) in year \( t \), ranging from 1992 to 2010; \( \text{SOUTH}, \text{MIDWEST}, \text{WEST} \) are indicator (dummy) variables that are equal to 1 if the indicated state \( i \) is located in the respective region, and equal to 0 if otherwise; and \( \epsilon_{it} \) is assumed to be independently, identically, and normally distributed error terms.
IV. INFERENCE AND RESULTS

1. Inference

Compared to an ordinary regression setting such as OLS, the ordered probit model and maximum likelihood estimation (MLE) inference poses a more complicated interpretation of the coefficients. Where there is no natural conditional mean function, $E[y|x]$, to analyze, direct interpretation of the coefficients is fundamentally ambiguous$^{104}$. The marginal effect of a change in the cumulative estimated probability of moving toward the top or bottom group, due to a unit change in the relevant explanatory variable ($x_j$), isn’t estimated by a $\beta$ value. Instead, it’s estimated by the calculated partial derivative of the probability expression, $Prob(y = i)$, with respect to $x_j$ (a normal density function, $\Phi$, of parameter $\beta_j$). For example, the probability of a state achieving the High Grade rating is

$$
Prob(y = 2) = \Phi(u_2 - \beta'x) - \Phi(u_1 - \beta'x)
$$

The marginal effect ($\delta_j$) of changes in $x_j$ on the probability of having a High Grade (2) rating is

$$
\delta_j(x_j) = \frac{\partial Prob(y = 2)}{\partial x_j} = (\Phi(u_1 - \beta'x) - \Phi(u_2 - \beta'x)) \beta_j
$$

where $\Phi$ is the density function of a standard normal distribution$^{105}$.

Obtaining the marginal effects of a dummy variable in an OMP model entails using a difference of probabilities, rather than just differentiating as if it were a

---

$^{104}$Greene and Hensher (2009).
$^{105}$Cheung (1996).
continuous variable. The latter would result in a finite change computation – a discrete approximation to the derivative. Suppose $D$ is a dummy variable in the model representing the region where state $i$ lies. With all other variables held at their values of interest, the appropriate computation measures the effect of a change in $D$ from 0 to 1,

$$\Delta_i(D) = [\Phi(u_j - \beta'x_i + \gamma) - \Phi(u_{j-1} - \beta'x_i + \gamma)] - [\Phi(u_j - \beta'x_i) - \Phi(u_{j-1} - \beta'x_i)].$$ \hspace{1cm} (15)

2. Results

We use several different regression models to explore tendencies and potential relationships in our analysis. Using panel data, we are able to incorporate the fixed effects associated with each state and control for more lurking or unaccounted variables. State economic stability is influenced by a number of factors that are independent of partisan polarization, such as geography and the weather, so it’s just as likely that the relationship between economic and political stability is exactly the opposite. This observation implies that political stability may be determined by economic stability (credit ratings), rather than vice versa.

Table 3: Estimation Results

| Variable        | Estimate | Standard Error | t-value | Pr(>|t|) | Sample N (states) [period] |
|-----------------|----------|----------------|---------|----------|----------------------------|
| **Equation (7)** |          |                |         |          |                            |
| COMP            | 0.55     | 0.5308         | 1.044   | 0.2964   | 608                        |
| log(PCINCOME)   | -0.97    | 0.1773         | -5.463  | 0.0000   |                            |
| DEBTGSP         | -0.12    | 0.0127         | 9.466   | 0.0000   |                            |
| Cut Point 1-2   | -9.43    | 1.1824         | -5.204  | 0.0000   | (32)                       |
| Cut Point 2-3   | -0.89    | 1.8105         | -0.492  | 0.0000   |                            |
| Cut Point 3-4   | -7.96    | 1.8071         | -4.403  | 0.0000   | [1992-2010]                |
| **Equation (8)** |          |                |         |          |                            |
| POLAR           | -0.13    | 0.1493         | -0.858  | 0.3909   | 388                        |
| log(PCINCOME)   | -0.10    | 0.2876         | -3.405  | 0.0007   |                            |

106 Greene and Hensher (2009).
107 Pallay (2014). For a relevant example, see Bowling and Ferguson (2001).
Our findings\textsuperscript{108} are disappointing considering the evidence provided by some of the literature. Equations (7), (8), (10), and (11) provide consistent evidence that both our

\textsuperscript{108} See Table 4.
COMP and POLAR variables don’t hold statistically significant relationships with state credit ratings given this particular model specification. This infers that our parameters of COMP and POLAR are also most likely assuming the wrong relationship with our dependent variable RATING.

Our $t$-statistics in Equations (9) and (12) agree\textsuperscript{109} with Pallay’s (2014) findings, which found that a better quality-of-life is analogous with Republican partisan control. Our model finds that as the proportion of Democrats decreases, the probability of a state reaching the highest credit rating increases. Pallay’s findings also suggest that other, untested variables are important drivers of this relationship and that the explanatory power of these relationships is very limited given little statistical significance. We agree and take our findings with skepticism. Similar to other empirical findings of this nature\textsuperscript{110}, our empirical results do not allow a causal interpretation and inference of the influence of political gridlock, which is undermined by simultaneity and omitted variable biasness.

\textsuperscript{109} We speculate that this statistically significant finding is merely a correlational effect and most likely not a true causal relationship given the complexities of biasness surrounding our complete model.

\textsuperscript{110} Poterba (1994); Bohn and Inman (1996); Poterba and Rueben (2001); Johnson and Kriz (2005); Depken and LaFountain (2006); Schelker (2012); Pallay (2014).
V. SUGGESTIONS FOR FUTURE RESEARCH

1. Dependent Variable

Despite the contributions made in this study, the effects of political gridlock on state credit ratings, especially in the states, merits considerable further investigation and research. There is evidence throughout the literature that the relationship observed in our empirical analysis is conclusive and economically significant (Krueger and Walker 2008). This implies that we took the wrong approach in one or many respects. More specifically, a reworking of our dependent variable is imperative in order to obtain statically significant results in our focal independent variables. There are several advanced methods that can make our dependent variable stronger.

Krueger and Walker (2008) utilize a unique measure of state bond ratings that incorporates Bayesian Marko Chain Monte Carlo (MCMC) simulation to measure its latent construct. A simulation procedure is the optimal approach when dealing with a dependent variable like state credit ratings and would have been implemented in our model if the proper resources were available. Krueger and Walker use censored, or “truncated” data, in a tobit model to describe a continuous, yet unobservable, measure of economic stability ($\Theta_{it}^*$). The credit rating agency doesn’t explicitly provide $\Theta_{it}^*$ through its ratings, but rather, a latent examination ($y_{itb}^*$) of an aggregated order of indicators, where $b = \{Fitch, Moody’s, S&P\}$ and one rating is closest to the state’s true economic stability.
To measure $\Theta_{it}$, we conceptualize the dependent variable $y_{itb}$ from an aggregated $b$ agencies for state $i$ in year $t$ is a function of parameters $\Lambda_b$ and underlying economic stability $\Theta_{it}^*$,

$$ y_{itb} = F(\Lambda_b, \Theta_{it}^*) $$

(16)

Each indicator $b$ has its own intercept, $\alpha_b$, and a particular slope, $\beta_b$, relating the changes in economic stability, $\Theta_{it}^*$, to the probabilities of a specified ordered categories. Defining $F$ as a cumulative standard normal distribution ($\phi$), and $\Lambda_b = \{\alpha_b, \beta_b\}$, we can assume a conventional ordered probit model,

$$ y_{itb}^* = \alpha_b + \beta_b \Theta_{it}^* + \varepsilon_{itb} $$

(17)

with the requirement that,

$$ Pr(y_{itb} = k_b) = \phi(\alpha_b, k_b - \beta_b \Theta_{it}^*) - \phi(\alpha_b, k_{b-1} - \beta_b \Theta_{it}^*). $$

(18)

We also assume $\alpha_b, k_b = \infty$ and $\alpha_{b,0} = -\infty$ to ensure proper probabilities for the discrete values $k_b$ of $y_{itb}$. Unlike a standard probit model, we are interested in the unobserved $\Theta_{it}^*$, not simply the latent $y_{itb}$.

Schelker (2009; 2012) take a different, yet similar approach. They apply Gauss-Hermite quadrature estimation procedures proposed by Frechette (2001) to random effects ordered probit models. Their study claims this is the preferred specification technique, as it allows for individual heterogeneity unlike the fixed effects that we implement.

Schelker (2009) also controls for time effects and additional covariates with variables that have proven to be influential in previous research at the U.S. state level\textsuperscript{111}. Likelihood ratio tests show that the regression model including the time effects fit the

\textsuperscript{111} These variables account for things such as: population density, strict balance budget laws, and initiative rights.
data significantly better than the basic model without the time effects. This fortifies our method of finding structural variables. It proves that taking the time to search for truly statically and economically significant structural variables – while still avoiding overfitting – can make the model much stronger.

If we were to form a similar model again, we would also implement a dependent lag. Benson and Mark’s (2007) study observes that rating agencies are slow to respond to changes in local economic and fiscal conditions. A lag would coordinate data in a position where causal effects would be more visible.

2. Independent Variables

Shor and McCarty (2011) use an advanced algorithm to strengthen their independent variable, which measures polarization. This algorithm works to eliminate districts that have a very high or low propensity to elect a Republican within their AIDD estimates for state lower houses. This trims their sample and allows them to eliminate concerns about including districts that are unlikely to elect a Democrat or Republican.

The Ranney index is the existing standard measure for political competitiveness, but it’s certainly not the best. Most studies that utilize its index are attempting to measure the forecasted probability of a change in party control within state government. This is much different than what the Ranney index actually measures: the makeup of previous partisan coalitions. C. E. Klarner should be coming out with a measure that fixes this problem by the August of 2014 (personal communication, March 19, 2014).

In order to fully utilize the diverse population presented by varying state legislatures, we also suggest considering the trends of third party movements and
independents to capture both polarization and competitiveness within a state government and its lower house. Research by Pew (2012) suggests that the growing partisan divide isn’t simply the result of more Americans classifying themselves as independents over the last twenty-five years. Independents themselves have migrated toward one party or the other. Such analysis of this recent ideological migration could provide more constructive evidence for the adverse effects of partisan gridlock.
VI. CONCLUSION

We have argued that uncertainty of future fiscal policies due to increased political instability, dysfunction, gridlock, competitiveness, and polarization, impose an adverse effect on the corresponding economy by increasing the likelihood of economy instability. We also demonstrated important steps toward finding statistically and economically substantive evidence in support of the view that political instability – through gridlock – has an adverse affect on a state’s economic stability and function.

In summary, we find strong evidence that as more scholarly work is achieved in this area, a more conclusive model will arise. As the dialogue grows, we are confident that is unambiguous evidence such a relationship is a reality and cause for attention. Even if partisan gridlock eases as the country fully recovers from its recent recession, there’s still reason to believe such a debilitating political dysfunction could repeat itself in the future.

Washington faces a corrosive hyper-partisanship problem that must be bridged in order to face the nation’s greatest challenges (Snowe, 2013). Unfortunately, more than half of the Millennial generation\textsuperscript{112} believes that the government is dysfunctional, wasteful, and inefficient. Congress must seek a way to redeem this trust and fortify a constructive, functioning system so that it can face the boundless issues that lie ahead. Putting aside partisan differences and working toward partisan peace is certainly one step in the right direction.

\textsuperscript{112} Up from 31 percent in 2003.


Polarization and Congress. Osher lifelong learning class. Lecture conducted from Vanderbilt University, Nashville, TN.


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Logan Nee was born in Portland, Maine on August 24, 1992. He was raised in Cornish, Maine and graduated from Sacopee Valley High School in 2010. Logan is a financial economics and political science double major and a member of Phi Kappa Phi, Omicron Delta Epsilon, and Pi Sigma Alpha. He competed in varsity track and field for the Black Bears with his identical twin brother, Liam.

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