Advance Access publication March 25, 2015

*Political Analysis* (2015) 23:385–399
doi:10.1093/pan/mpv004

# Competing Gridlock Models and Status Quo Policies

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Edited by R. Michael Alvarez

Spatial theories of lawmaking predict that legislative productivity is increasing in the number of status quo policies that lie outside the gridlock interval, but because locations of status quo policies are difficult to measure, previous empirical tests of gridlock theories rely on an auxiliary assumption that the distribution of status quo points is fixed and uniform. This assumption is at odds with the theories being tested, as it ignores the history dependence of lawmaking. We provide an alternative method for testing competing theories by estimating structural models that explicitly account for temporal dependence in a theoretically consistent way. Our analysis suggests that legislative productivity depends both on parties and supermajority pivots, and we find patterns of productivity consistent with a weaker, contingent form of party influence than found in previous work. Parties appear to exert agenda power only on highly salient legislation rather than strongly influencing outcomes through voting pressure and party unity.

## 1 Introduction

Understanding the role that parties play in shaping policy outcomes is central to understanding fundamental issues concerning the distribution of power and its consequences. Cleanly differentiating between competing partisan and pivot theories, however, has proven to be a difficult methodological problem (Krehbiel, Meirowitz, and Woon 2005; Clinton 2007). Under many conditions the theories turn out to be observationally equivalent, and the unobservability of key elements of the theories raises challenging inferential problems. A particularly vexing problem is the fact that the locations of current policies—referred to as "status quo" policies—relative to legislators' preferences are unknown and difficult to estimate (Poole and Rosenthal 1997). Previous empirical tests often assume that the status quo distribution is fixed, which might seem to be innocuous, but turns out to be problematic because it is theoretically inconsistent with the theories being tested, for it denies the temporal dependence of policy-making in which previous outcomes (laws and their effects) shape current decisions.

We propose a novel method for comparing alternative theories using a structural model of legislative productivity. Our approach seeks to place tests of competing gridlock theories on solid theoretical ground with respect to the role of status quo policies. The key feature of our structural model is that it combines an explicit, yet flexible, model of history-dependent status quo policies with spatial models of lawmaking, thereby avoiding problems that undermine the validity of inferences made in previous work. Building on the analytical approach proposed by Krehbiel (1998, 2006a, 2006b), we assume that status quo policies are partly endogenous: They are inherited

*Authors' note:* A previous version was presented at the Positive Political Theory Mini-Conference at the 2011 Annual Meeting of the Southern Political Science Association, New Orleans, LA. Replication material can be found at the Political Analysis Dataverse (Woon and Cook 2015). Supplementary materials for this article are available on the *Political Analysis* web site.

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https://doi.org/10.1093/pan/mpv004 Published online by Cambridge University Press

from policy-making in previous periods but subject to exogenous stochastic shocks. By combining the model of stochastic shocks with various assumptions about the role of parties (positive agenda power, negative agenda power, direct pressure on roll call voting), we generate estimated distributions of status quo policies and outcomes, which lead to a series of statistical models of legislative productivity. Importantly, our method does not rely on the need to measure the spatial location of bills or status quo points directly. Using Monte Carlo methods and Common Space ideal point estimates (Carroll et al. 2010) to generate likelihood distributions for counts of legislative productivity, we then fit the models to data on salient legislative enactments in the 80th–110th Congresses via maximum likelihood, which allows us to estimate parameters and to compare models.

Our analysis yields two substantive findings and suggests that the status quo assumptions underlying empirical tests are indeed consequential. First, hybrid pivots-plus-party agenda control models provide the best fit to the data for highly salient legislation. While we find evidence consistent with party influence, the form of influence we find stands in contrast to previous work on gridlock that finds that the source of party influence is party unity via strong pressure on voting behavior (in which intra-party heterogeneity is irrelevant, as in Chiou and Rothenberg 2003; Richman 2011) or negative agenda control without pivots (in which the minority party plays no role, as in Cox and McCubbins 2005; Stiglitz and Weingast 2010). Second, we find that the best model varies with the salience of legislation, with the non-partisan pivot model providing the best fit for the least salient legislation. Thus, parties and pivots *both* matter. Parties exercise agenda power, but they must do so within the constraints dictated by their members' preferences and institutional voting rules, choosing to exercise their agenda power only on the most salient issues of the day.

The article proceeds as follows. In Section 2, we compare spatial models of lawmaking and briefly review the literature. In Section 3, we explain why the fixed, status quo distribution is both substantively and methodologically problematic: The theories emphasize history dependence, but relaxing the assumption of a fixed status quo distribution causes problems for empirical tests because the relationship between the key variables of interest becomes indeterminate. Section 4 describes our structural model, estimation procedure, and data. Section 5 presents the results, and Section 6 concludes.

## 2 Competing Models of Party Influence and Gridlock

Spatial models of lawmaking yield precise predictions about legislative outcomes and the frequency of policy change. We consider four models that vary in their assumptions about the nature of party influence, comparing a non-partisan pivotal politics baseline against a set of hybrid pivot-plus-party models.[1] The baseline model in our analysis is the *non-partisan pivot* (NP) model in which parties have no influence and outcomes are constrained by supermajority voting rules (Krehbiel 1998; Brady and Volden 2006). The *party unity* (PU) model is a hybrid pivot-plus-party model with the strongest form of party influence in which parties dictate the locations of pivots (Chiou and Rothenberg 2003). The majority party has weaker influence in the *party agenda setter* (AS) model, wielding positive agenda power but not influence over members' voting behavior. The *party gatekeeping* (GK) model involves the weakest form of party influence in our anaysis and is a hybrid version of the NP model with a prior GK stage following the cartel agenda model (Cox and McCubbins 2005).

Figure 1 illustrates the differences between the four models' basic predictions regarding outcomes as a function of the status quo, holding the ideal points of key players constant. The key players are the majority party median $P_{\mathrm{M}}$, the minority party median $P_{\mathrm{N}}$, the veto pivot $V$, the filibuster pivot $F$, and the floor median $M$.[2] The figure shows the obvious differences between the models' predictions about the width of gridlock intervals and also illustrates the severity of the observational equivalence problem. For instance, all four models predict gridlock if the status quo lies between $V$ and $F$, while the AS, GK, and NP models predict the same outcomes for status quo points lying between $2F - M$ and $F$. The figure also highlights the fact

[1] See the Online Appendix for a detailed discussion of these models and analyses of additional models.
[2] The figure corresponds to unified Republican government. Similar figures can be made for other cases.

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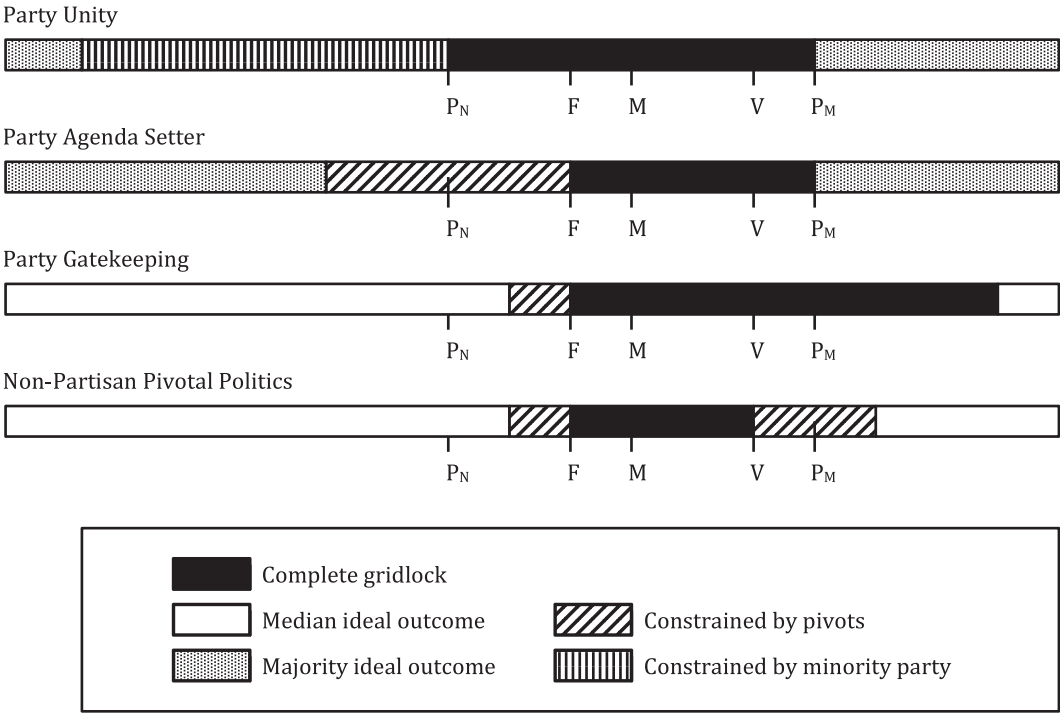



**Fig. 1** Comparison of theoretical models.

**Table 1** Previous empirical tests

| *Article/book* | *Non-partisan pivots* | *Gatekeeping* | *Agenda setting* | *Party unity* | *Status quo* | *Dependent variable* | *Supported Theories* |
|---|---|---|---|---|---|---|---|
| Krehbiel (1998) | X | | | | Fixed | Productivity | Non-partisan pivots |
| Chiou and Rothenberg (2003) | X | | x | X | Fixed | Productivity | Party unity |
| Covington and Bargen (2004) | X | X | | | Fixed | Roll call voting | Gatekeeping |
| Cox and McCubbins (2005) | X | X | | | Limited dependence | Roll rates | Gatekeeping |
| Krehbiel, Meirowitz, and Woon (2005) | X | X | | | None | Cutpoints | Ambiguous |
| Chiou and Rothenberg (2006) | X | | X | X | Fixed | Productivity | Party unity |
| Krehbiel (2006b) | X | X | | | History dependence | Roll rates | Ambiguous |
| Clinton (2007) | X | X | | | None | Cutpoints | None |
| Chiou and Rothenberg (2009) | | | | X | Fixed | Productivity | Party unity |
| Stiglitz and Weingast (2010) | X | X | | | Fixed | Cutpoints | Gatekeeping |
| Richman (2011) | X | | | x | Estimated | NPAT, fiscal policy | Party unity |

that the models make different predictions about what will happen to policies when the status quo lies outside the gridlock interval. For example, the PU and AS models predict that extreme status quo policies lead to outcomes at the majority party median, while the GK and NP models imply that such outcomes will instead lie at the floor median. These differences provide additional leverage that our method uses to discriminate between the models, unlike the gridlock interval test used in previous work.

Table 1 summarizes previous empirical tests of spatial gridlock theories and illustrates that our analysis improves on previous scholarship in several ways. In particular, previous empirical tests

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rely on strong assumptions about the status quo distribution. Some tests assume that status quo policies are normally distributed around the chamber median (Cox and McCubbins 2005; Lawrence, Maltzman, and Smith 2006), while others assume that status quo policies are uniformly distributed independent of the configuration of legislative preferences (Chiou and Rothenberg 2003, 2006, 2009; Covington and Bargen 2004). Most existing tests relying on the fixed status quo assumption seem to find evidence in favor of an extremely strong form of party influence (PU). In these models, there is no role for intra-party heterogeneity, which largely renders pivots irrelevant as well because gridlock occurs only whenever the parties disagree. In contrast to this work, we confront the problem of unobservable status quo policies in a novel, theoretically consistent way, leading us to find evidence for weaker forms of party influence. Our approach is also flexible in that it is straightforward to incorporate alternative assumptions about the source of stochastic variation in our model, and we allow for theoretical model parameters to be estimated from the data rather than imposing any a priori restrictions. As Table 1 shows, our set of models is also more comprehensive than in previous work. Most of this work either compares pure pivot models with pure GK (cartel) models or ignores GK entirely, while ours is the only article to compare all four types of models (NP, AS, PU, and GK).

## 3 Status Quo Policies and Gridlock Predictions

In this section, we review the gridlock interval test and explain why it is invalid under the more general assumption of history dependence. The origin of the test is Krehbiel's insight that pivotal politics theory generates an *equilibrium gridlock interval* (EGI), the set of status quo policies for which attempted policy-making will be unsuccessful, and the implication that "Policy change requires that the status quo must lie outside the gridlock interval" (1998, 47). The following prediction is an immediate corollary.

**Prediction 1**
*Legislative productivity is a strictly increasing function of the frequency with which status quo policies lie outside the gridlock interval.*

Although Prediction 1 is a direct consequence of the spatial pivotal politics model, it cannot be tested directly unless status quo policies can be identified relative to the EGI. Instead, Krehbiel tests the hypothesis that changes in the *width* of the (non-partisan pivotal politics) gridlock interval are inversely related to legislative enactments. Chiou and Rothenberg (2003) note that the EGI concept and gridlock interval test also apply to alternative models of lawmaking and provide a basis for assessing competing models. However, because the gridlock interval test ignores the distribution of status quo policies, it is not quite a test of Prediction 1—the relationship between status quo policies, EGI, and gridlock—but rather a test of whether there is a significant, positive relationship only between gridlock and the width of the EGI. We state the basis for the gridlock interval test as Prediction 2.

**Prediction 2**
*Gridlock is a strictly increasing function of the width of the gridlock interval.*

Importantly, Chiou and Rothenberg also point out that the hypothesized relationship between the width of the gridlock interval and legislative productivity (what we label Prediction 2) depends on a *critical identifying assumption*. They explicitly assume that "status quo policies in each period are identically and uniformly distributed, which makes comparing the widths of different gridlock intervals at the same time or different points in time possible," and note also that "Krehbiel did not make this assumption explicitly, but his hypothesis requires it" (2003, 511). More precisely, previous work relied on the assumption of a fixed, uniform distribution of status quo policies because it is sufficient for Prediction 2 and thus guarantees the validity of the gridlock interval test.

To see why the assumption implies a negative, monotonic relationship between gridlock interval width and legislative enactments, let the gridlock interval be $\text{EGI}_t = [L_t, R_t]$ in period $t$ and let status quo policies be uniformly distributed over the interval $[q_0, q_1]$, which is constant in every

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period. Define the predicted level of gridlock to be the proportion of status quo policies that fall within the interval, provided that $q_0 < L_t$ and $R_t < q_1$:

$$G_t = \frac{R_t - L_t}{q_1 - q_0}. \quad (1)$$

Since the denominator is a constant, it clearly follows that $G_t$ is increasing in the width of the EGI, $R_t - L_t$. Previous research relied extensively on the assumption of this fixed, uniform distribution of status quo policies to guarantee the validity of Prediction 2 as the basis for comparing a variety of generalized pivot-plus-party models (Chiou and Rothenberg 2003, 2006, 2009; Covington and Bargen 2004; Stiglitz and Weingast 2010).

### 3.1 *The Problem of History Independence*

The assumption of a fixed distribution of status quo policies may seem like a reasonable solution to the identification problem, but there are two issues. First, it represents a strong substantive assumption of history independence that is inconsistent with the spirit and intent of spatial theories of lawmaking. Second, when plausible alternative assumptions are considered, the relationship between the width of gridlock intervals and legislative productivity varies depending on changes in the distribution of status quo policies. That is, Prediction 2 does not hold generally and the validity of the gridlock interval test cannot be assured when the distribution of status quo policies is not assumed to be static.

An exogenously fixed distribution of status quo policies means that policy-making at time $t$ has *no bearing* on the extent of gridlock at time $t+1$. In contrast, spatial lawmaking theorists emphasize the *history dependence* of status quo policies (even if status quo policies are, strictly speaking, exogenous to the *formal* representations of the theories). For example, in applying the pivotal politics theory to interpret the broad historical contours of policy change from the late 1970s through the 1990s, Krehbiel writes that "Carter equilibria $x_1$ become Reagan-Bush status quo points $q_2$" and the "funneling effect of liberal policies toward the [Reagan-Bush] median creates Reagan-Bush outcomes $x_2$ which serve as status quo points $q_3$ for Clinton" (1998, 43). Brady and Volden are also explicit about the intertemporal dependence of status quo policies, emphasizing that "policy gridlock depends on both the *size* and the *shifting* of the gridlock region" (2006, 26, emphasis original). Similarly, party theorists such as Cox and McCubbins explicitly recognize that "status quo policies reflect bills enacted in the previous legislative period, as well as exogenous shocks" (2005, 174).

To illustrate how history dependence and independence have drastically different implications for policy change and gridlock, we consider two examples. First, suppose the series of gridlock intervals for three periods of lawmaking is as depicted in Fig. 2.[3] If status quo polices are history independent so that in every period they have the same distribution over the interval shown in the upper-most (unshaded) horizontal bar, then the rate of gridlock will be the *same* in all three periods because the widths of the intervals are the same.[4]

Now suppose instead that status quo policies are history dependent so that all out-of-gridlock status quo policies are brought into the interior of the gridlock interval at time $t$. Thus, status quo policies at time $t+1$ must lie in the time $t$ gridlock interval. In this case, there is complete gridlock at $t=2$ because all of the status quo policies at $t=2$ (which lie within the gridlock interval from $t=1$) also fall within the gridlock interval at $t=2$. But at $t=3$, the gridlock interval shifts to the right, which "releases" a set of status quo policies (those that lie within the $t=2$ interval but outside the $t=3$ interval), which now meet the condition for policy change. Thus, in contrast with history independence and Prediction 2, which imply a constant rate of gridlock, gridlock *decreases* from

[3]Whether the gridlock model is partisan or non-partisan is immaterial. Our argument depends only on the existence of a gridlock interval.

[4]In other words, because Prediction 2 states that gridlock is an increasing function of EGI width, no change in the argument (EGI width) implies no change in the value of the function (gridlock).

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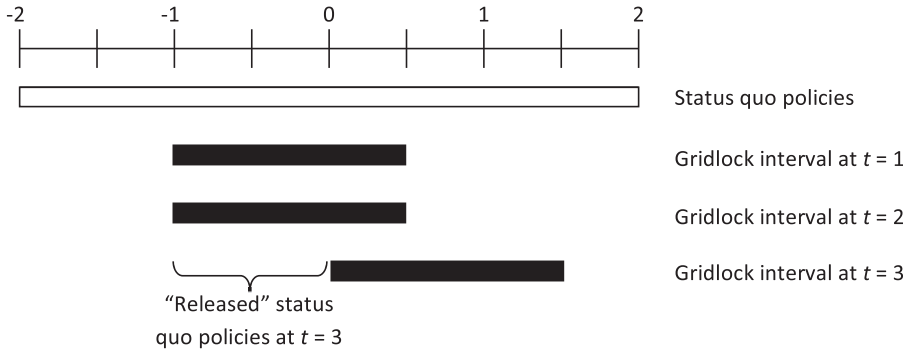



**Fig. 2** Shifting gridlock intervals.

$t=2$ to $t=3$ even though there is *no change* in its width. This is because the gridlock interval test based on Prediction 2 ignores *shifts* of the location in the gridlock interval.

Another example illustrates the severity of the theoretical indeterminacy. Suppose that there are three periods and that $\text{EGI}_1 = [-1, 1]$, $\text{EGI}_2 = [-1, 0.5]$, and $\text{EGI}_3 = [-1, 0.25]$, as depicted in Fig. 3. Note that because each successive gridlock interval is strictly nested in, and therefore smaller than, the preceding gridlock interval, the uniform status quo assumption implies that gridlock is *strictly decreasing* from $t=1$ to $t=3$.

Alternatively, consider the case of complete history dependence. For purposes of illustration, suppose that status quo policies at time $t+1$ are distributed uniformly over $[L_t, R_t]$, which is the previous period's EGI.[5] Under this assumption, the predicted level of gridlock is not proportional to the size of the EGI $[L_t, R_t]$ but instead depends on how much of the previous EGI at time $t$ falls in the new EGI at time $t+1$. Given that the gridlock interval at $t=1$ is $[-1, 1]$, the status quo policies at $t=2$ are uniformly distributed over $[-1, 1]$. Since the gridlock interval at $t=2$ is $[-1, 0.5]$, the proportion of gridlocked policies is 0.75. At time $t=3$, status quos are distributed over $[-1, 0.5]$ and the gridlock interval is $[-1, 0.25]$, so the proportion of gridlock is 0.83. Thus, given these intervals and the alternative assumption that the distribution of status quos depends completely on history, gridlock is *increasing* even though the width of the intervals is decreasing—precisely the opposite of Prediction 2.

These examples illustrate that history dependence can invalidate the gridlock interval test under substantively plausible conditions.[6] The gridlock interval test relies on the maintained assumption of a monotonic relationship between EGI width and gridlock, which is guaranteed by a fixed, uniform distribution of status quo policies. But this is a strong assumption in which policy-making is completely independent of the past. If the distribution of status quo points instead varies over time depending on the collective choices made in previous Congresses, as both pivotal politics and partisan theorists emphasize, then the relationship between EGI width and gridlock is indeterminate, as it can be increasing or decreasing. Previous regression analyses of gridlock or legislative productivity using the width of the gridlock interval as the independent variable (e.g., Krehbiel 1998; Chiou and Rothenberg 2003) relied on an inappropriate method for comparing competing models.

[5]We use this assumption for ease of exposition. A slightly different assumption is that status quo policies at $t+1$ depend on equilibrium lawmaking outcomes from time $t$ as in Krehbiel's approach, which is a special case of our structural model, but the gridlock predictions are more complicated to describe.

[6]Although we focus primarily on history independence, the assumption of a uniform distribution is also important for the gridlock interval test. To see this, suppose that the distribution of status quo policies follows a standard normal distribution in every period of lawmaking. Contrast the gridlock intervals $I = [-1.64, 1.64]$ and $I' = [0, 4]$: 90% of status quo policies lie in the first interval while less than 50% of status quo policies lie in the second, so gridlock is decreasing even as the width of the interval increases.

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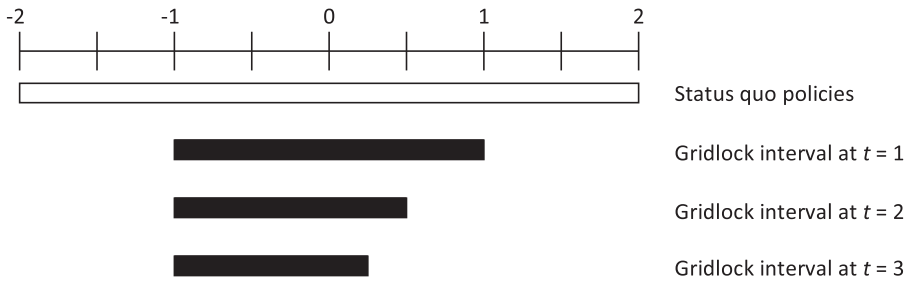



**Fig. 3** Shrinking gridlock intervals.

## 4 A Structural Model Linking Theory and Data

Without a clear hypothesis about the direction of the relationship between gridlock intervals and legislative productivity or direct measures of status quo policies, how can the theories be tested? We solve this problem by taking a structural approach that is theoretically consistent with Prediction 1. Instead of assuming that status quo policies come from a fixed distribution, we assume that the distribution of status quo policies varies and is itself stochastic. More important, the distribution depends on history: It is partially the result of prior lawmaking and partially the result of stochastic shocks. Policy change in this model therefore results from both changes in the gridlock interval (size and shift) as well as exogenous changes in policy (e.g., from policy implementation, advances in technology, or the resolution of policy uncertainty).[7]

To construct a model of the data-generating process, we begin with a theoretical model of legislative productivity that embeds the generalized pivot-plus-party models within a stochastic, dynamic framework. This model generates a probability distribution over status quo policies, and by averaging over simulated distributions of status quo policies, we then generate a likelihood function for the dependent variable (levels of legislative productivity). The resulting structural model is one in which the theoretical and statistical models are much more closely integrated (Morton 1999; Signorino 1999, 2003) than in a reduced form regression approach (e.g., Carrubba, Yuen, and Zorn 2007).[8]

Our structural model is a more precise implementation of the "quasi-dynamic" application of pivotal politics proposed and employed by Krehbiel (1998, 2006a, 2006b).[9] Although the model is not fully game theoretic, it incorporates the equilibrium results of spatial policy-making models (including, but not limited to, pivot models) as part of the stochastic model. In order to generate a likelihood function, the model assumes that there are many issues and that status quo policies are subject to random shocks. Using standard principles of maximum likelihood, we can fit the parameters of each generalized pivot model and then compare the different theoretical models.

### 4.1 *The Model*

Suppose there are $N$ issues and $T$ periods. For any period $t$, let the EGI be $[L_t, R_t]$, and the proposer be $P_t$, where $L_t \leq P_t \leq R_t$. In the NP model, $L_t$ and $R_t$ are the relevant filibuster and veto pivots

[7]The stochastic shocks in our model differ from the uniform status quo assumption in another important way. Generating Prediction 2 requires viewing the uniform status quo distribution as deterministic (i.e., there are an infinite or nearly infinite number of policy issues) or as stochastic but only in large samples (i.e., if the status quo is a random variable, Prediction 2 holds in the limit). In contrast, our method is more flexible, as it does not require an asymptotic or large sample argument, and generates likelihood distributions for any number of policy issues.

[8]Signorino (2003, 318) defines a "theoretical" model as one "that results from theory construction," while he defines a "statistical" model as one "that guarantees positive probability over all outcomes—for example, there is some random component in the model that induces a probability distribution over the outcomes." Thus, although Signorino's own models typically involve game theoretic equilibria, his definition of "theoretical" is quite broad and precludes neither partial equilibrium approaches like ours nor other non-equilibrium models.

[9]Our formalization is closest to that in Krehbiel (2006b), but differs in that he examines roll rates and uses a different empirical method. Krehbiel "bins," or discretizes, the policy space in his simulations, then evaluates model fit using deviations from predictions rather than maximum likelihood.

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while $P_t$ is the floor median.[10] In the PU model, the endpoints of the gridlock interval are defined by the majority and minority party medians while $P_t$ is the majority party median. In the AS model, $P_t$ and one of the EGI endpoints are both defined as the majority party median while the other EGI endpoint is the relevant filibuster or veto pivot. Finally, in the GK model, $L_t$, $R_t$, and $P_t$ are the same as in the NP model but there is an additional player $G_t$, which is the House majority party median.

For any issue $i \in \{1, \ldots, N\}$ and period $t \in \{1, \ldots, T\}$, denote the status quo by $q_t^i$ and the policy outcome by $x_t^i$. In the initial period, $t=1$, status quo policies are independently and identically distributed (according to distributions described below). Within every period $t$, policy-making is independent across issues and the outcome follows from standard subgame perfect equilibrium analysis of the model. In the NP, PU, and AS models, the outcome function is

$$x_t^i(q_t^i) = \begin{cases} P_t & \text{if} \quad q_t^i \le 2L_t - P_t \\ 2L_t - q_t^i & \text{if} \quad 2L_t - P_t < q_t^i \le L_t \\ q_t^i & \text{if} \quad L_t < q_t^i \le R_t \\ 2R_t - q_t^i & \text{if} \quad R_t < q_t^i \le 2R_t - P_t \\ P_t & \text{if} \quad 2R_t - P_t < q_t^i \end{cases}. \tag{2}$$

In the GK model, the outcome function depends on the location of $G_t$ relative to the NP gridlock interval. If $L_t \le G_t \le R_t$, then the outcome is the same as in equation (2) because the gatekeeper always weakly prefers the outcome of the lawmaking game to the status quo. However, if $G_t$ lies outside the interval $[L_t, R_t]$, then there is an additional region of gridlocked status quo points. If $G_t < L_t$, the outcome function is

$$x_t^i(q_t^i) = \begin{cases} P_t & \text{if} \quad q_t^i \le 2G_t - P_t \\ q_t^i & \text{if} \quad 2G_t - P_t < q_t^i \le R_t \\ 2R_t - q_t^i & \text{if} \quad R_t < q_t^i \le 2R_t - P_t \\ P_t & \text{if} \quad 2R_t - P_t < q_t^i \end{cases}, \tag{3}$$

while if $G_t > R_t$, it is

$$x_t^i(q_t^i) = \begin{cases} P_t & \text{if} \quad q_t^i \le 2L_t - P_t \\ 2L_t - q_t^i & \text{if} \quad 2L_t - P_t < q_t^i \le L_t \\ q_t^i & \text{if} \quad L_t < q_t^i \le 2G_t - P_t \\ P_t & \text{if} \quad 2G_t - P_t < q_t^i \end{cases}. \tag{4}$$

Once the policy outcome $x_t^i(q_t^i)$ is determined in period $t$, the status quo for issue $i$ in period $t+1$ is a function of the previous policy and a random shock. Substantively, if Congress enacts a new law, the location of the outcome $x_t^i$ should be thought of as a statute or law while the random shock should be thought of as changes that occur outside the legislative process that affect the location of policy relative to legislators' preferences. Such changes can occur for many reasons: technical or scientific breakthroughs, economic changes, judicial rulings, bureaucratic implementation, shifts in public attitudes, or because experience showed that the policy did not work quite as intended.

[10] We follow Chiou and Rothenberg's (2003) Propositions 1–4 for determining the locations of the pivots $L_t$ and $R_t$ in the NP, PU, and AS models, and we adopt their convention in assuming that the proposer $P_t$ is the Senate floor median in the NP model and that it is the Senate majority party median in the PU and AS models.

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We consider two alternative assumptions about the distribution of shocks and the way in which shocks combine with previous policy. For *normal additive* policy shocks, the status quo is an additive function of the policy outcome and a normally distributed shock,

$$q^i_{t+1} = x^i_t(q^i_t) + \varepsilon^i_t\,, \tag{5}$$

where $\varepsilon^i_t$ is independently and identically distributed following a normal distribution with mean 0 and standard deviation $\sigma$. We assume that in the initial period (without prior lawmaking) $q^i_1$ is also distributed normally with mean 0 and standard deviation $\sigma$. The normal distribution is reasonable because it is natural to assume that most exogenous shocks are relatively small while allowing large shocks to occur with positive probability. We emphasize that our point is not that there cannot be a source of stochastic variation or that the functional form of this variation should matter, but rather that there is an important, theoretically relevant distinction between a uniform distribution of *status quo policies* ($q_{t+1}$) and a uniform distribution of *random shocks* ($\epsilon_t$).

For *uniform weighted average* shocks, the location of the status quo is a weighted average of the outcome and a uniformly distributed shock,

$$q^i_{t+1} = \alpha x^i_t(q^i_t) + (1-\alpha)\upsilon^i_t\,, \tag{6}$$

where $\upsilon^i_t$ is independently and identically distributed uniformly over the interval $[-\delta, \delta]$ and $\alpha \in [0, 1]$ is the weight put on previous policy outcomes relative to the strength of the policy shock. In the initial period, $q^i_1$ is also distributed uniformly over $[-\delta, \delta]$. Although the additive assumption in equation (5) seems most reasonable to us, the weighted average in equation (6) has a nice property in that it encompasses both a model of complete status quo dependence (i.e., $q^i_t = x^i_{t-1}$ when $\alpha = 1$) and the fixed, uniform status quo assumption (i.e., $q^i_t \sim U[-\delta, \delta]$ when $\alpha = 0$). Our estimate of $\alpha$ therefore serves to measure how close the data are to one of the two extreme models.[11]

Finally, in order to apply the model to legislative productivity (our dependent variable), we must aggregate across all issues $i \in \{1, \ldots, N\}$. Let $b^i_t$ be an indicator variable denoting whether or not a bill passes. Formally, $b^i_t = 1$ if $x^i_t = q^i_t$ and 0 otherwise, and let the total number of bills that pass at time $t$ be $B_t = \sum_{i=1}^{N} b^i_t$. Since the model is stochastic and $\varepsilon^i_t$, $\upsilon^i_t$, $q^i_t$, and $x^i_t$ are all random variables, it follows that $b^i_t$ and $B_t$ are random variables as well. Thus, the model implies a probability distribution for $B_t$ for each period $t$ that depends on the relevant ideal points $L_t$, $R_t$, $P_t$, and $G_t$, as well as the ideal points in periods prior to $t$, the number of issues $N$, and the stochastic shock parameters $\sigma$ (for the normal additive version) or $\alpha$ and $\delta$ (for the uniform weighted average version). We do not impose any a priori assumptions on $N$, $\sigma$, $\alpha$, or $\delta$, instead estimating these parameters from the data.

### 4.2 *Estimation*

We compute the distribution of $B_t$ using Monte Carlo simulation methods.[12] While it is possible in principle to derive the distribution of $B_t$ analytically, the results would be quite complicated—there certainly would not be a single parametric equation that holds for any configuration of preferences. In order to set the ideal points $L_t$, $R_t$, $P_t$, and $G_t$, we use Common Space scores to compute the values corresponding to each pivot-plus-party model. The time periods $t \in \{1, \ldots, T\}$ are the Congresses corresponding to those for which we have legislative productivity data. In addition, we use one Congress immediately preceding the earliest one in the data as $t = 0$ as a starting period. This ensures that the distribution of status quos for the first period of data is not entirely random but is dependent on previous lawmaking to some extent. For example, if the actual legislative productivity data correspond to the 80th ($t = 1$) through 106th ($t = 27$) Congresses, then the initial values (for $t = 0$) correspond to the 79th Congress.

[11] In addition, in the absence of new law (when there is gridlock such that $x_t(q_t) = q_t$), the normal additive assumption is equivalent to a random walk while the uniform weighted average assumption is similar to an AR(1) process.

[12] Replication material including data and MATLAB code (Woon and Cook 2015) can be found at the Political Analysis Dataverse.

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The number of issues $N$ and the status quo parameters $\sigma$ and $\delta$ are "nuisance" parameters to be estimated.[13] For each shock assumption and set of parameters, $(N, \sigma)$ or $(N, \delta, \alpha)$, we run $K = 10{,}000$ iterations of the simulation and then use the results to compute the probability mass function (PMF) for each $B_t$. Let $m_t(n)$ denote the number of iterations for which $B_t = n$. The PMF is computed as

$$p_t(B_t = n) = \frac{m_t(n) + 1}{K + N} \tag{7}$$

for each $n \in \{1, \ldots, N\}$, where the $t$ subscript for $p$ indicates that each $p_t$ is a distinct distribution. The calculation of the likelihood is a version of Laplace's Law of Succession. It differs from a simple proportion to deal with a computational issue known as the "zero frequency problem" (Witten and Bell 1991) in which a rare event may not be observed in a finite number of trials even though it is known to occur with positive probability.[14]

Once we compute the likelihood function for a given set of parameters, we can then compute the likelihood for any data set on legislative productivity. Given a set of count data $y_1, \ldots, y_T$, the log likelihood is

$$\ln L(y_1, \ldots, y_T) = \sum_{t=1}^{T} \ln p_t(B_t = y_t)\,. \tag{8}$$

In our analysis, we use several measures of legislative productivity summarized in Table 2. We use counts of legislative enactments from Binder (2003), which span the 80th through 106th Congresses (1947–2000) and for which there are five different levels of salience. An enactment is counted as salient and included in the data if there is at least one *New York Times* editorial mentioning the issue, and higher levels of salience correspond to a greater number of editorial mentions. We also use a count of enactments based on the updated version of Mayhew's (1991) Sweep One series, which covers the 80th–110th Congresses (1947–2008). A legislative enactment is included in the series if it is discussed in end-of-year assessments of legislative accomplishments. Note that the average and maximum number of enactments are relatively small, especially for the higher salience measures, which suggests that it is important to take into account variation due to uncertainty in small samples, which our method does.

We maximize the likelihood function using a simple grid search for each pivot-plus-party model and for each assumption about the stochastic shocks. For parameters, $N$ ranges from 25 to 105 in increments of 5, $\sigma$ ranges from 0.025 to 0.55 in increments of 0.025, $\delta$ ranges from 0.35 to 1.25 in increments of 0.025, and $\alpha$ ranges from 0 to 1 in increments of 0.05.[15]

## 5 Results

As a starting point, Table 3 shows the fit of each version of our structural model to Binder's level 4 productivity data, which is the level of salience used in Chiou and Rothenberg's (2003) analysis. Several findings are noteworthy. First, the single best-performing gridlock model is the party AS

[13] An "issue" represents a specific value of a status quo policy at any given time, including both explicit agenda items (e.g., policy areas mentioned in the *New York Times* or the kinds of year-end legislative summaries examined by Mayhew) as well as "latent" or "new" issues not yet noticed by the news media. Treating $N$ as a free parameter is important. Imposing a priori values of $N$ possibly biases our estimates of the other parameters in the same way that imposing a value for the intercept in a bivariate regression will bias the slope estimate. We treat it as a free parameter to be estimated so that our results do not depend on an arbitrarily imposed value.

[14] In practice, we frequently encounter the zero frequency problem. If we were to use the simple proportion $m_t(n)/K$, we often end up with likelihoods of 0 because one or more of the frequencies $m_t(n)$ corresponding to some observed value $n$ is 0.

[15] To make the length of time needed to run the simulations manageable, we conducted the grid search in a two-step process. First, we used a "coarse" grid search with larger increments. We then performed a "finer" grid search with the smaller increments in neighborhoods around the maxima from the "coarse" search.

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**Table 2** Summary statistics for legislative productivity measures

| *Measure* | *Obs.* | *Mean* | *St. Dev.* | *Min.* | *Max.* |
|---|---|---|---|---|---|
| Binder 1 | 27 | 50.11 | 13.26 | 31 | 82 |
| Binder 2 | 27 | 28.93 | 9.85 | 11 | 52 |
| Binder 3 | 27 | 19.52 | 7.28 | 4 | 34 |
| Binder 4 | 27 | 14.93 | 5.58 | 4 | 24 |
| Binder 5 | 27 | 11.56 | 4.37 | 3 | 21 |
| Mayhew | 31 | 10.13 | 3.62 | 4 | 19 |

**Table 3** Maximum-likelihood results, Binder measure (Salience 4), 80th–106th Congresses

| *Shock* | *Model* | *LL* | *Issues* | $\sigma$ *or* $\delta$ | $\alpha$ |
|---|---|---|---|---|---|
| Normal additive | Party Agenda Setter | −92.25 | 40 | 0.200 | — |
| Normal additive | Gatekeeping | −96.32 | 30 | 0.475 | — |
| Normal additive | Non-partisan Pivot | −101.03 | 30 | 0.275 | — |
| Normal additive | Party Unity | −108.70 | 45 | 0.200 | — |
| Uniform weighted average | Party Agenda Setter | −90.44 | 35 | 0.525 | 0.15 |
| Uniform weighted average | Non-partisan Pivot | −94.61 | 35 | 0.375 | 0.05 |
| Uniform weighted average | Gatekeeping | −95.28 | 30 | 0.900 | 0.20 |
| Uniform weighted average | Party Unity | −107.71 | 60 | 0.925 | 0.60 |

model, in which the majority party has a moderate degree of party influence.[16] This holds for both shock assumptions. Second, the model with the strongest level of party influence (PU) performs the worst, again regardless of the shock assumption.[17] The non-partisan baseline (NP) and the weakest version of party influence (GK) fall in between. Under the normal additive assumption, GK clearly outranks NP, while under the uniform weighted average assumption, NP just barely outperforms GK. These results contrast with the results of Chiou and Rothenberg's (2003) gridlock interval tests, which rank PU the highest.

Because all of the models in Table 3 are estimated using the same data, it is appropriate to compare shock assumptions by comparing the likelihoods while holding the theoretical assumptions about party strength constant. For three of the four theoretical models, the uniform weighted assumption produces a better fit. Also, the values of $\alpha$ are low for the AS, GK, and NP models (0.05–0.20) and high for the PU model (0.60). The relatively low values of $\alpha$ for the AS model suggest that most status quo policies are more random than inherited, but the fact that they are non-zero nevertheless also suggests that inherited status quos do play some role in generating the data.[18]

In Table 4, we present the maximum-likelihood results using counts from Mayhew's Sweep One data as the productivity measure. Mayhew's measure tends to count fewer pieces of legislation as highly significant than Binder does, and his data span a longer time period that includes the four Congresses during George W. Bush's administration. Table 4 shows that the ranking of the AS,

[16]Because we compare such a large number of models, we rank them according to the sum of log likelihoods, which is equivalent to using an information criterion such as the AIC or BIC given that our models have the same number of parameters.

[17]These results also hold if we assume that the shock distribution is logistic additive instead of normal additive. The results for the logistic distribution are substantially similar to those for the normal distribution, so we omit them from the presentation.

[18]This result might be due to time constraints and the possibility that the AS chooses some subset of policies outside the gridlock interval to place on the agenda rather than placing all such policies on the agenda. Such a model would be an interesting extension for future work but introduces additional parameters and requires additional assumptions about the agenda-setting process.

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**Table 4** Maximum-likelihood results, Mayhew productivity, 80th–110th Congresses

| *Shock* | *Model* | *LL* | *Issues* | $\sigma$ *or* $\delta$ | $\alpha$ |
|---|---|---|---|---|---|
| Normal additive | Gatekeeping | −103.35 | 25 | 0.375 | — |
| Normal additive | Party Agenda Setter | −105.19 | 25 | 0.250 | — |
| Normal additive | Non-partisan Pivot | −116.16 | 25 | 0.225 | — |
| Normal additive | Party Unity | −121.63 | 25 | 0.275 | — |
| Uniform weighted average | Gatekeeping | −108.63 | 25 | 1.050 | 0.45 |
| Uniform weighted average | Party Agenda Setter | −115.52 | 25 | 1.175 | 0.65 |
| Uniform weighted average | Non-partisan Pivot | −124.46 | 25 | 1.025 | 0.65 |
| Uniform weighted average | Party Unity | −133.16 | 30 | 1.250 | 0.65 |

NP, and PU models are identical to Table 3: the AS model clearly outperforms the basline NP, which outperforms the (worst-fitting) PU model. However, in contrast to the results using the Binder level 4 measure, we find that for the most salient legislation, the GK model—the pivot-cartel hybrid—consistently does the best. These results suggest that parties might exercise a mix of positive and negative agenda control, though our method does not allow us to make distinctions between individual bills.

There are also a few other differences between the results for the Binder and Mayhew measures. For the Mayhew measure, the normal additive shocks outperform their uniform weighted average counterparts for each model. For the GK model, the normal additive shock does better by five points in the log likelihood scale, and for the AS model, it does better by over ten points. In terms of the $\alpha$ parameter, our estimates put much greater weight on $\alpha$ for the Mayhew measure ($\alpha = 0.65$ for AS) than for the Binder measure ($\alpha = 0.15$). This provides evidence against the fixed uniform status quo assumption, but we view it tentatively since there is no clear "winner" across both data series in terms of the random shock assumptions.

We also fit the models to several additional measures of gridlock from Binder (2003) in which the level of salience for inclusion in the count vary. While this analysis serves partly as a robustness check to see whether the ranking of models and the parameter estimates hold across the different measures, it also allows us to examine whether party influence might vary with the salience of legislation. To the extent that party resources are scarce and party leaders wield them primarily to enhance the party's electoral reputation (as emphasized by, e.g., Aldrich 1995; Cox and McCubbins 2005), we would expect the party to conserve its influence only for the most salient legislation. We caution that while our analysis may shed some light on this possibility, it does not serve as a rigorous test of the hypothesis.

The results for the best-fitting model for each data series are summarized in Table 5.[19] Interestingly, it does appear that party influence varies with the salience of legislation in a sensible way. At the lowest level of salience (Binder 1), the model without party influence does the best for both shock assumptions. At the next two levels, the results are not entirely consistent across shock assumptions. The NP model fits the best for Binder 2 and Binder 3 under the uniform weighted average assumption, which outperforms the PU model for Binder 2 and the AS model for Binder 3 under the normal additive assumption. For both Binder 4 and Binder 5, the AS model fits the best, with the results again consistent across shock assumptions, and the only data for which the GK model fits best are Mayhew's. Also, with one exception (Binder 2), the PU model does worse than the NP, AS, and GK models, but the preference-only NP model is a *very close* second for Binder 2, with a log likelihood of –123.91. Overall, these results are consistent with the interpretation that if parties' resources are limited, they use their influence only on high-profile and significant legislation.

To facilitate comparison with previous work that compares "pure" non-partisan with "pure" partisan lawmaking models (see Table 1), we also computed likelihoods for a non-hybridized party

[19]Full results for each salience level, like those given in Tables 3 and 4, can be found in the Supporting Information.

https://doi.org/10.1093/pan/mpv004 Published online by Cambridge University Press

**Table 5** Summary of results, varying levels of salience

| *Data source* | *Shock* | *Best-fitting model* | *LL* | *Issues* | $\sigma$ *or* $\delta$ | $\alpha$ |
|---|---|---|---|---|---|---|
| Binder 1 | Normal additive | Non-partisan Pivot | −156.02 | 90 | 0.325 | — |
| Binder 2 | Normal additive | Party Unity | −123.44 | 80 | 0.200 | — |
| Binder 3 | Normal additive | Party Agenda Setter | −112.06 | 45 | 0.225 | — |
| Binder 4 | Normal additive | Party Agenda Setter | −92.25 | 40 | 0.200 | — |
| Binder 5 | Normal additive | Party Agenda Setter | −81.98 | 25 | 0.250 | — |
| Mayhew | Normal additive | Gatekeeping | −103.35 | 25 | 0.375 | — |
| Binder 1 | Uniform weighted average | Non-partisan Pivot | −155.46 | 85 | 0.725 | 0.25 |
| Binder 2 | Uniform weighted average | Non-partisan Pivot | −119.17 | 55 | 0.45 | 0.10 |
| Binder 3 | Uniform weighted average | Non-partisan Pivot | −106.30 | 40 | 0.475 | 0.20 |
| Binder 4 | Uniform weighted average | Party Agenda Setter | −90.44 | 35 | 0.525 | 0.15 |
| Binder 5 | Uniform weighted average | Party Agenda Setter | −80.87 | 25 | 1.050 | 0.55 |
| Mayhew | Uniform weighted average | Gatekeeping | −108.35 | 25 | 1.050 | 0.45 |

GK model. In this model, which is Cox and McCubbins's (2005) House cartel agenda model, we assume that if the majority party chooses to allow an issue on the agenda, the outcome will be the floor median's ideal point (i.e., under an open rule).[20] Table 6 presents results that compare the NP and GK results with those from the pure cartel model. When restricted to these three possibilities, the NP only model does the best for the three lowest levels of salience (Binder 1, 2, and 3 for both shock assumptions, and also for Binder 4 and 5 under the uniform weighted average assumption). The hybrid pivots-cartel model does the best for the highest levels of salience (Mayhew for both shock assumptions and Binders 4 and 5 under the normal additive assumption). Importantly, we note that the pure cartel GK model performs the worst for all of the data. The results in Table 6 reinforce the importance of supermajoritarian pivots and our finding that party influence depends on the salience of legislation, further supporting our view that party influence plays a role above and beyond, but does not supplant, basic institutional constraints.

## 6 Conclusion

The locations of status quo policies are a key feature of spatial lawmaking theories, but their unobservability makes theory testing difficult. Previous research has solved the problem in the form of an ancillary assumption—a fixed status quo distribution—that is inconsistent with the substance and implications of the theories themselves. We solve the problem by constructing and estimating a structural model, thereby providing a better foundation regarding the treatment of status quo policies for theory testing. The key feature of our approach is to allow the distribution of status quo policies to evolve from prior lawmaking. That is, our structural model explicitly incorporates the kind of history dependence featured in the theoretical analysis of such models and lacking in prior empirical work.

The substantive consistency of our structural model with theoretical gridlock models increases the confidence in the inferences that we can draw from our analysis, and our methodological innovation leads to substantive results that, while broadly consistent with the overall pivots-plus-party story in previous work (Chiou and Rothenberg 2003; Richman 2011), yield distinct conclusions about the nature of party influence. Models of pivotal politics are sometimes presented as incompatible with partisan politics (Krehbiel 1998; Cox and McCubbins 2005), but our evidence suggests that both kinds of models capture important aspects of the lawmaking process. We also show that parties wield influence through agenda control rather than by inducing their members to behave cohesively, as previous research suggests, and we find that the majority party is influential on highly salient legislation and plays little or no role in shaping outcomes on less significant issues. Thus, it appears that parties are constrained and selective in exerting their influence.

[20]See the Supporting Information for results for alternative GK models.

https://doi.org/10.1093/pan/mpv004 Published online by Cambridge University Press

**Table 6** Pivots, cartels, and hybrids

| *Data source* | *Shock* | *Hybrid pivots–cartel* | *Pivots only* | *Cartel only* |
|---|---|---|---|---|
| Binder 1 | Normal additive | −164.26 | −156.02[a] | −181.82 |
| Binder 2 | Normal additive | −135.91 | −123.91[a] | −172.38 |
| Binder 3 | Normal additive | −120.34 | −113.65[a] | −141.15 |
| Binder 4 | Normal additive | −96.32[a] | −101.03 | −123.00 |
| Binder 5 | Normal additive | −86.47[a] | −89.80 | −119.19 |
| Mayhew | Normal additive | −103.35[a] | −116.16 | −134.53 |
| Binder 1 | Uniform weighted average | −164.56 | −155.46[a] | −156.18 |
| Binder 2 | Uniform weighted average | −134.08 | −119.17[a] | −145.73 |
| Binder 3 | Uniform weighted average | −117.91 | −106.63[a] | −131.62 |
| Binder 4 | Uniform weighted average | −95.28 | −94.61[a] | −110.01 |
| Binder 5 | Uniform weighted average | −87.92 | −83.02[a] | −104.56 |
| Mayhew | Uniform weighted average | −108.63[a] | −124.46 | −119.18 |

[a]Best fit.

These findings suggest two directions for continued research. One possibility is to extend our structural model by incorporating various constraints on scheduling or agenda power. For instance, an extended model might explicitly limit the number of bills that can be placed on the agenda, the number of bills that the majority party can bring up under a closed rule, or the number of bills that the legislative median will allow the majority party median to block. Such extensions would add additional parameters and require an explicit model of agenda choice. Another related direction would be to use our structural model to conduct analyses of specific policy areas using more fine-grained data and measures of preferences. This work might find that party influence varies not only across time (in terms of preference shifts caused by elections), but also across issues and policy areas (in terms of relevance to parties and their constituents). Pursuing these lines of research will require advances in theory as well as new sources of data.

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