Article

The Dimensionality of Congressional Voting Reconsidered

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Jason M. Roberts[1], Steven S. Smith[2], and Stephen R. Haptonstahl[3]

Abstract

This article reports findings for a decomposition of the roll-call voting record of the U.S. Congress to determine the effect of the level of aggregation on the observed dimensionality of the policy space. In doing so, we identify some but certainly not all of the ways in which the aggregation of the voting record affects the observed dimensionality of the policy space. For the 1955 to 2008 period (84th-110th Congresses), we apply optimal classification (OC) to votes aggregated to the level of the individual bill and policy area to measure dimensionality. We examine the marginal proportional reduction in error (MPRE) across dimensions. Our results demonstrate that complexity in voting patterns of individual bill episodes is the norm, that aggregating to higher levels reduces the observed dimensionality, and that the liberal–conservative dimension appears more dominant in more highly aggregated analyses. These results call into question many of the conclusions from the theoretical and empirical literature on the U.S. Congress that uses a unidimensional model.

Keywords

Congress, roll call voting, dimensions

[1]University of North Carolina at Chapel Hill, USA
[2]Washington University in St. Louis, MO, USA
[3]National Public Radio, Reston, VA, USA

Corresponding Author:
Jason M. Roberts, University of North Carolina at Chapel Hill, 361 Hamilton Hall, CB 3265, Chapel Hill, NC 27599, USA.
Email: jroberts@unc.edu

Introduction

The dimensionality of congressional voting behavior is critical to the development and application of spatial voting theories to congressional policy making. Yet, the dimensionality of voting in the U.S. Congress remains a contested empirical issue. Numerous studies in the literature on the U.S. Congress assume, either implicitly or explicitly, that one or at the most two dimensions can characterize the underlying policy space. However, as we demonstrate in this article, the clear patterns of unidimensionality that are often observed in voting aggregated to whole Congresses are not duplicated when the unit of analysis is the individual bill or subsets of bills. In fact, we find considerable evidence that multidimensionality is the norm for most major bills and policy areas across both the House and Senate. This is an important point, as most theories and applications of spatial voting models to Congress deal with particular bill episodes or policy areas, not aggregations of votes over an entire Congress. Our findings suggest that much of this research incorrectly assumes low dimensionality, rendering many of the empirical findings potentially suspect.

In this article, we report findings for a decomposition of the roll-call voting record. In doing so, we identify some but certainly not all of the ways in which the aggregation of the voting record affects the observed dimensionality of the policy space. For the 1955 to 2008 period (84th-110th Congresses), we apply optimal classification (OC) algorithms to votes aggregated to the level of the individual bill policy area to measure dimensionality. We examine the marginal proportional reduction in error (MPRE) across dimensions. Our results demonstrate that complexity in voting patterns of individual bill episodes is the norm, that aggregating to higher levels reduces the observed dimensionality, and that the liberal–conservative dimension appears more dominant in more highly aggregated analyses.

Dimensionality and Theories of Legislating

In studies of legislative behavior, it is well understood that there are fundamental differences between unidimensional and multidimensional spaces. Unidimensional policy spaces are well-behaved—pivotal legislators such as the median can be readily identified. Multidimensional policy spaces, however, generally do not have identifiable pivots. Consequently, outcomes are predicted with great precision in unidimensional space, whereas in most cases multidimensional spaces do not offer precise predictions and can lead to the possibility of cycling or chaos. For this reason, most empirical applications of spatial theory assume unidimensionality (Cameron, 2000; Krehbiel, 1998).

This was not always the case. Through the 1950s and 1960s, political scientists posited multidimensionality as the basic expected pattern in a pluralistic polity. Perhaps most prominently, Truman (1951), MacRae (1958), Miller and Stokes (1963), Lowi (1964), and Clausen (1973) argued that the forces influencing legislators varied from issue to issue as did the voting alignments that they produced. Party, constituency, and ideology were emphasized as the primary forces in the equation, but considerable attention was also given to presidential influence. Variation in the character of lobbying groups was given attention, although systematic study of group influence was very limited (see Bauer, de Sola Pool, & Dexter, 1972). The relevance of these large forces in American politics was thought to vary systematically across large domains of public policy.

In more recent studies, the theoretical convenience of assuming a unidimensional space is often given empirical justification by reference to the findings of Poole and Rosenthal (1997, 2007). Poole and Rosenthal observe that, in analyses in which the DW-NOMINATE method is applied to whole Congresses, it is difficult to improve on the predictive value of a single liberal–conservative dimension in most Congresses (2007, p. 34). The aggregate proportional reduction in error (APRE), taken as the difference between errors based on the DW-NOMINATE estimates and those based on the marginal vote distribution, averages about 0.5 for the first dimension with only a slight improvement, about 0.06 on average, by adding a second dimension. Poole and Rosenthal (2007) draw the following inferences:

- A reasonable fit is obtained from a one-dimensional (1-d) model in which each legislator's position is constant throughout his or her career. (33)
- Introducing more parameters in a dynamic model—through extra dimensions or higher order polynomials (to capture changing ideal points)—does not appreciably add to our understanding of the political process. (35)
- A 1-d model typically provides a good fit to the data, with a second dimension needed in periods when race issues are distinct from economic ones. (59)

Poole and Rosenthal reinforce these inferences by observing that the correlations among first-dimension W-NOMINATE scores for four large policy categories (each encompassing hundreds of votes) are very high. The concluding chapter of the 2007 edition of the Poole–Rosenthal book captures their views well in its title, "The Unidimensional Congress."

The Poole–Rosenthal findings are foundational for an important body of theory and associated empirical tests, yet we think in many ways they have been misapplied. The scores derived from NOMINATE are by definition summaries of behavior over an entire 2-year Congress, yet most of our theories and applications apply to individual bills. The books and papers that use their scores in this manner are far too numerous to cite here, so we focus on two prominent examples. Krehbiel (1998) uses first-dimension DW-NOMINATE scores to test a 1-d model that identifies super-majority pivots in the contexts of veto overrides and Senate cloture votes. In all cases, Krehbiel considers how the identity of pivotal actors will affect the outcome of an individual bill. Similarly, Cameron (2000) limits his analysis of presidential leverage gained through vetoes to a cutting-point analysis based on first-dimension NOMINATE scores. Veto-bargaining also, by definition, occurs in the context of an individual bill, not an entire Congress of votes. The theory in both studies is unidimensional but as we show below, both analyze bill episodes that are in many cases multidimensional. Nevertheless, both studies use first-dimension NOMINATE scores to test some aspects of their unidimensional theory. Some of the predictions in both studies would likely have been very different if a multidimensional space was assumed.

We are not alone in questioning the assumption of unidimensionality. Heckman and Snyder (1997) devise a simpler linear estimation procedure that yields similar estimates to the non-linear model of Poole and Rosenthal when the estimates are constrained to two dimensions. A six-dimension principal component model, Heckman and Snyder find, provides a substantially better fit and the 1-d model. On average for the 80th to 100th Congresses, the proportional reduction in error increased by more 50% by moving from a one- to a six-dimension model. By finding that distinctive subsets of votes load heavily on the various dimensions, Heckman and Snyder infer that the additional factors represent systematic differences in legislators' preferences across issues.

While the Heckman–Snyder results cast doubt on the aggregate-level claim of unidimensionality, they do not address the dimensionality of the policy space relevant to spatial theories of legislative politics—the individual bill episode. Appearing to realize this, Heckman and Snyder continue by exploring the voting record on civil rights legislation. Their findings are in line with most investigators who focus on legislative episodes associated with individual bills rarely find a unidimensional account adequate. There are too many accounts of congressional politics to mention in this regard, but a handful of studies warrant brief discussion.

Jones (1961) examined the expressed preferences and behavior of House Agriculture Committee members on a farm subsidy bill. He observed that legislators approved of federal support for agricultural interests in their own districts but otherwise followed their parties. We can readily infer that multiple dimensions were present—the dimensions representing the multiple agricultural commodities that were the subject of the bill and the one dimension, and a liberal–conservative dimension that produced a division between the parties.

The Heckman and Snyder (1997) account of civil rights voting shows that it is useful to consider both the liberal–conservative dimension (the first dimension) and the civil rights dimension (the second dimension). The weakness of this analysis is that it is based on all roll-call votes, not just those related to the civil rights bills, so we are left with substantial uncertainty about the relevance of the factors applied. Subsequent studies are more persuasive.

Hurwitz, Moiles, and Rohde (2001) argue that three theories of legislative organization—distributive, informational, and partisan—can be readily accommodated if we realize that legislators often perceive multiple dimensions in the bills they construct and vote on. Examining the agricultural appropriations and farm bills of 1996, this team found the bills to be complex and the voting alignments to vary across parts of the bills in a way that defined distinctive dimensions. Some votes on key amendments generated the "first dimension" liberal–conservative or partisan alignment, whereas others yielded more parochial or distributive divisions. Moreover, committee members showed distinctive support for farm interests. Plainly, the bills, which incorporated several distinguishable issues, generated multiple dimensions in the observed voting behavior of legislators.

Smith (2007) examined voting for all bills during the 2001 to 2005 period that were subject to a "key vote," identified by Congressional Quarterly as a vote on an issue of national significance. For 23 of 98 bills, Smith found that the House key vote represented a dimension of voting different from the final passage vote. To be sure, 75 key and final passage votes appeared to reflect the same underlying dimension, but nearly a quarter of the most important votes cast in the House, spanning a wide range of subject matter, appeared to reflect dimensions other than the one represented in the vote on final passage.[1]

Similarly, Crespin and Rohde (2010) break apart the roll-call record into subsets consisting of votes on the 13 Appropriations bills that must pass in some form during each 2-year Congress. They then estimate the dimensionality of the policy space using DW-NOMINATE and find multiple dimensions. They also find that the 1-d model rarely explains half of the variance in roll-call voting when the record is decomposed in this manner. They conclude that while the unidimensional model does a good job of explaining votes over an entire Congress, it misses much of the complexity in voting on specific subsets of bills.

In more recent work, Aldrich, Montgomery, and Sparks (2014) demonstrate that high levels of party polarization bias estimates of dimensionality downward. Their simulations reveal that scaling votes *within* party caucuses returns more than one dimension. They conclude that the evidence implying a low dimensional policy space for the U.S. Congress may simply be an artifact of high levels of party bloc voting.

There are good reasons to expect that multidimensionality is present in most bill episodes. Many bills today run to thousands of pages, with hundreds of provisions that affect millions of Americans. Routine appropriations legislation, for example, spends billions of dollars on numerous programs and individual projects that likely cut across multiple dimensions. This is likely exacerbated when multiple bills are combined into omnibus legislation. The Affordable Care Act had multitudes of provisions that cut across the liberal–conservative dimension including abortion, health care for immigrants, and what level of government should administer the program. If the studies of individual legislative episodes are a guide, we should be careful about limiting our theory to unidimensional spatial models and about empirical claims that complex models add little to our understanding of the political process. Our working hypotheses are simple:

Proposition 1: The observation of a powerful first dimension in highly aggregated analyses of roll-call voting is explained by the high frequency of bills for which the liberal–conservative dimension of conflict emerges.
Proposition 2: The observation of multiple dimensions in many, if not most, individual bill episodes is explained by the high frequency of bills for which infrequently occurring, non-liberal/conservative dimensions of conflict emerge.

In short, we agree with Aldrich et al. (2014) that the inference of unidimensionality is largely an artifact of the estimation process aggregated to the 2-year Congress. Although we agree that a unidimensional model can characterize general patterns of behavior, most individual bills have multiple dimensions of conflict, which can and do affect legislative outcomes.[2] In this article, we decompose the voting record to show that, despite the presence of a powerful first dimension that drives aggregate measures of roll-call voting, most individual bill episodes are multidimensional in nature.

Levels of Aggregation and Dimensionality

We have conducted a dimensional analysis for the 27 Congresses of the 1955 to 2008 period to explore the dimensionality of the policy space exhibited in voting associated with three levels of aggregation: (a) 2-year Congresses; (b) broad policy domains within Congresses (the Clausen domains, Clausen [1973)] and Clausen and Van Horn [1977]); and (c) individual bills. We use OC (Poole, 2000, 2005) to explore the underlying dimensionality of congressional voting.[3]

Estimating Dimensionality With MPRE

Consider, for example, a Senate vote that is 60 "aye" and 40 "nay." A "zero"-dimensional model that predicts all senators vote with the majority will have 40 classification errors, exactly the size of the minority (Hammond & Fraser, 1983; Weisberg, 1978). Suppose a 1-d OC analysis correctly classifies 75 senators' votes, and a two-dimensional (2-d) OC analysis correctly classifies 85 votes.

The APRE is defined as

$$\text{APRE}=\frac{\sum_{j=1}^{q}[\text{minority vote-classification errors}]_j}{\sum_{j=1}^{q}[\text{minority vote}]_j}.$$

There are two reasons for using APRE instead of just measuring the fraction of votes classified correctly. First, we do not want to give "credit" to the 1-d model for explaining the 60 votes of the majority because those can be classified correctly just by noting that the measure passed, so we only count the improvement in the number of correctly classified votes, "minority vote - classification errors." Second, the most improvement the 1-d model can show is to explain all of the 40 minority votes misclassified in the simpler model; dividing by the number of minority votes puts APRE on a (0,1) scale.

We proceed by comparing the fraction classified correctly and APRE in our example. The fraction correctly classified by the 0-d model is .600 and the fraction correctly classified by the 1-d model is .750, a difference of .150. The APRE for the 1-d model is .375, which is the fraction of votes that were misclassified under the 0-d model but that are correctly classified under the 1-d model. Is this a good measure? That depends on the inference one wishes to draw. APRE has face and construct validity when used to describe *how much better the 1-d model is than the 0-d model* because it chooses quantities for both the numerator and denominator relating to the inference in question (the difference between the two models.)

Moving on to the 2-d model, Poole and Rosenthal (1997, 2007, Chapter 3) measure the contribution of adding a second dimension by calculating the APRE for both models and subtracting. APRE_1 is .375 and APRE_2 is .625, a difference of .250. The problem with using this difference, in our view, is that it is a measure of how much *more* the 2-d model improves on the 0-d model than the 1-d model does; it is not a measure of how much the 2-d model improves on the 1-d model. For the same reason, the 1-d model is not a .250 improvement on the 0-d model; it simply reflects the change in the fraction correctly classified.

To measure the improvement of 2-d over 1-d, we want to use the quantity

$$\frac{25-15}{25}=\frac{2}{5}=.400,$$

which is the fraction of the classifications the 2-d model could make correctly above and beyond those classified by the 1-d model. Generalizing APRE, we define the MPRE from Model A to Model B as

$$\mathrm{MPRE_{AB}}=\frac{\sum_{j=1}^{q}[\text{errors by A-errors by B}]_j}{\sum_{j=1}^{q}[\text{errors by A}]_\mathrm{j}}.$$

In terms of APREs, this can be calculated using

$$\mathrm{MPRE}_{AB}=\frac{(1-\mathrm{APRE}_A)-(1-\mathrm{APRE}_B)}{1-\mathrm{APRE}_A}.$$

A minimum MPRE must be set above which a dimension is said to be significant. We use the same standard that Netflix[4] did and assume that a 10% MPRE is significant. Specifically, we say the dimensionality of set of votes is the smallest number of dimensions for which the MPRE is greater than 10% for each dimensions.[5]

In some cases, we report that five dimensions are manifest. There could be more dimensions manifest in the data, but we only fit up to five-dimensional model and therefore are only able to detect up to five dimensions.

When a model fits perfectly, that is when a model of k dimensions correctly predicts all of the votes, the MPRE for the k-dimensional model is 100% and we would say that k dimensions are manifest in the data.[6] This is intentional; the k-dimensional model is a complete improvement on the k-1-d model, even if the APRE increases only slightly.

Estimating Dimensionality With Eigenvalues

While eigenvalues are computationally convenient quantities to examine when assessing the dimensionality of roll-call data, they do not provide nor rely on measures of fit of OC or NOMINATE. The eigenvalues reported by OC and NOMINATE are the eigenvalues of the double-centered squared distance matrix.

Other than using a different default threshold, there is no difference between the eigenvalues reported by OC and NOMINATE. The eigenvalues have nothing directly to do with the models assumed by OC or NOMINATE; rather, they depend on the assumptions of principal component analysis.

Harding (2008) shows that eigenvalue analysis of finite samples bias toward the identification of unidimensionality. This suggests that an eigenvalue-based analysis of dimensionality will *underestimate* the number of dimensions.

Why use eigenvalues instead of some measure of dimensionality based directly on OC or NOMINATE? Our experience suggests that it is computationally cheaper, while yielding the same substantive answers,[7] at least when the number of votes is large. However, it is not clear that this is the correct approach when the number of votes is smaller, such as when we focus on the votes for a single bill. We therefore focus on results based on MPRE, but note here that the results are the same and even more pronounced when based on eigenvalue analysis.

Congresses

We find considerable variance in the dimensionality of voting over the 27 Congresses in our sample.[8] Figure 1 shows that the dimensionality manifest in voting is not fixed over time, even when measured at a highly aggregated level. The number of manifest dimensions is decreasing over time in the House ($p < .001$) and very likely in the Senate ($p < .1$). With fewer dimensions manifest, the first (liberal–conservative) dimension may be capturing more of the voting behavior. The figure suggests that we must exercise care in assuming that the first dimension is uniformly strong, even over the period since the 1950s.

Figure 2 shows the MPREs by chamber for each of the first five dimensions in three Congresses, 1973-1974, 1993-1994, and 2007-2008. In each Congress, the two chambers are remarkably similar. In both Congresses, the first dimension is much stronger than the second dimension, although the figure shows a stronger first dimension and weaker second and third dimensions in the more recent Congress. The figure is consistent with Poole–Rosenthal NOMINATE findings, which show some variation in the strength of a second dimension but with a dominant first dimension.

The average number of manifest dimensions across Congresses in the House is 2.2, whereas the average in the Senate is 3.1, a difference that is

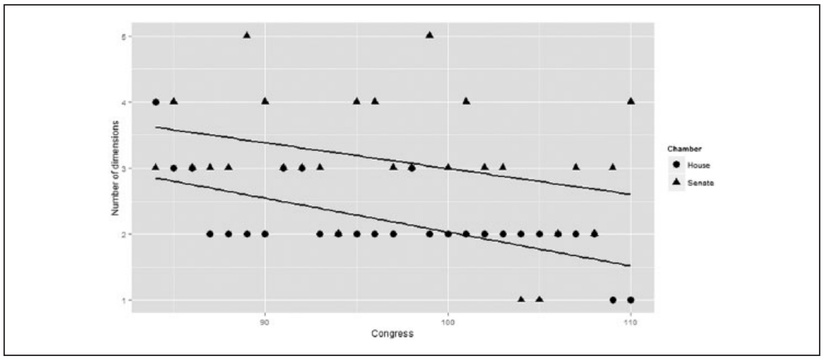

Figure 1. Dimensions manifest in each Congress-Chamber.
Note. Dimensions contributing at least 10% MPRE. Slope for House negative with $p < .001$; slope for Senate likely negative with $p < .1$. MPRE = marginal proportional reduction in error.

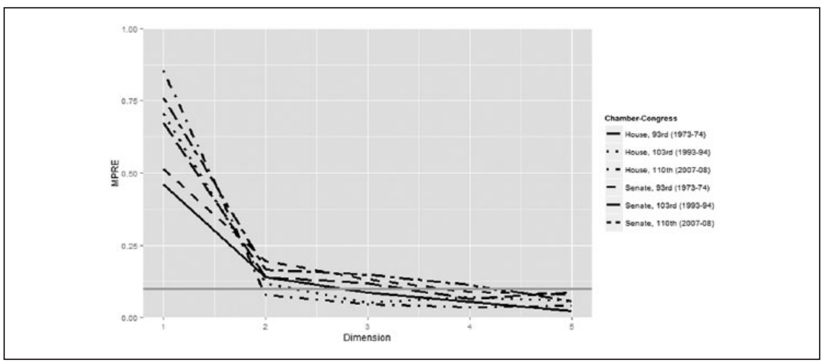

Figure 2. MPRE by dimension by chamber for selected Congresses.
Note. Dotted line indicates 10% standard for significant dimensions. MPRE = marginal proportional reduction in error.

likely the result of the fact that the Senate does not have a general germaneness rule and that House party leaders have far more control over the agenda than do their Senate counterparts.

Clausen Domains

The policy domains, conceptualized by Clausen (1973) and coded by Poole and Rosenthal, exhibit more widely varying dimensionality. Figure 3 breaks out the manifest dimensions by Congress-Chamber for each of Clausen's six

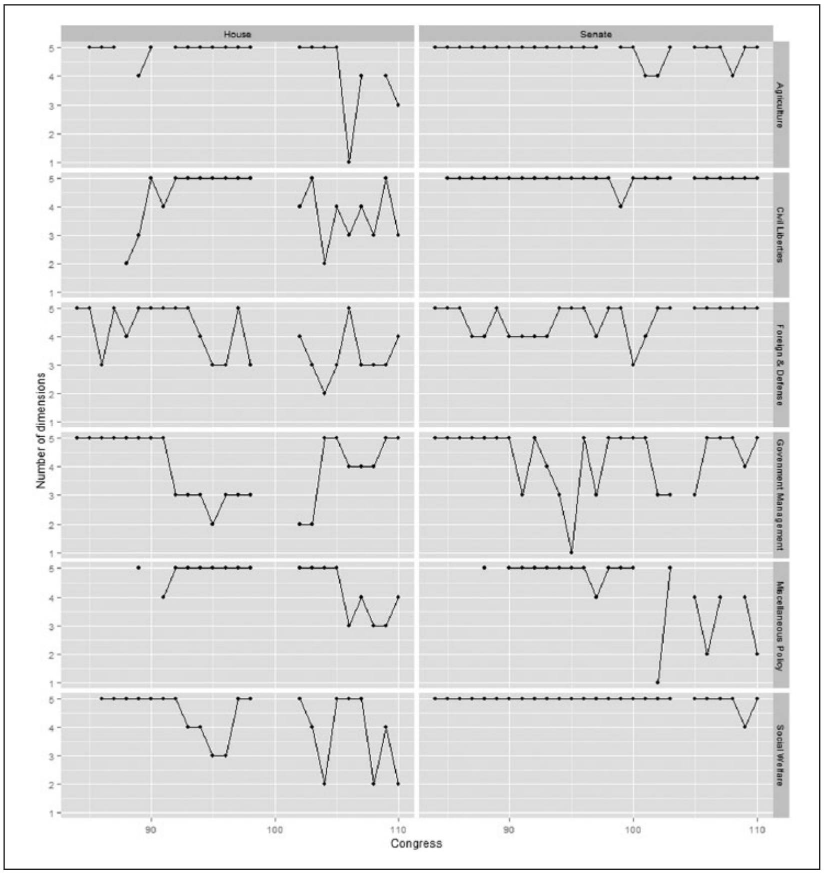

Figure 3. Dimensions manifest in each Congress-Chamber for each policy domain.
Note. Dimensions contributing at least 10% MPRE. MPRE = marginal proportional reduction in error.

policy domains. The overall trend that the Senate exhibits more dimensions that the House is clear here just as in Figure 1, but difference across policy domains are readily apparent. In the Senate, Agriculture, Civil Liberties, and Social Welfare are fairly uniform in manifesting a high number of dimensions whereas Government Management tends to have fewer dimensions. In the House, Agriculture and Civil Liberties tend to show a higher number of dimensions during the 93rd to 98th Congresses, which coincide with decreased numbers of dimensions in Government Management. This mixed

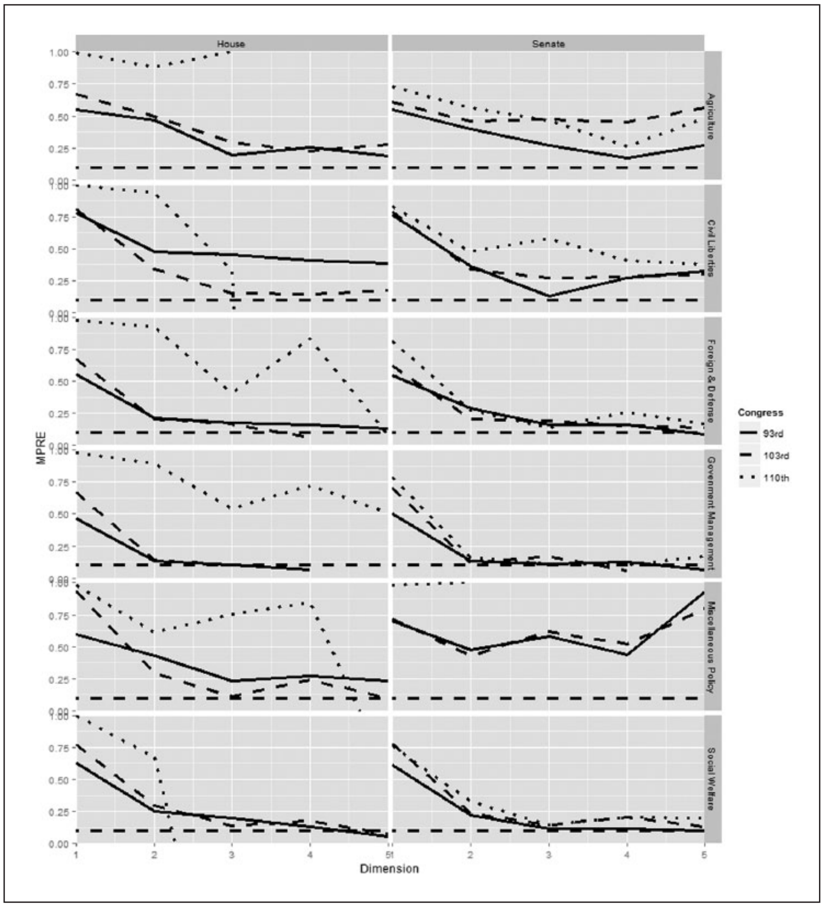

Figure 4. MPRE by dimension by chamber for selected Congresses by policy domain, 93rd Congress (1973-1974), 103rd Congress (1993-1994), and 100th Congress (2007-2008).
Note. Dotted line indicates 10% standard for significant dimensions. MPRE = marginal proportional reduction in error.

pattern comports with the Clausen argument that congressional behavior is not necessarily consistent across domains. It also hints that dimensional analyses aggregated to the 2-year Congress may mask important variation.

Due to space constraints, we forgo a presentation of the MPRE plot by dimension for six policy domains over 27 Congresses. Instead, in Figure 4 we show the MPRE plot by dimension for the same three Congresses considered

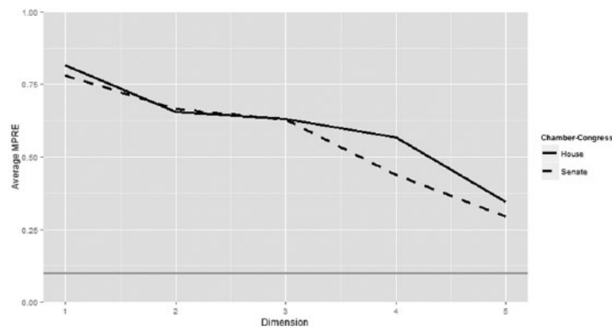

Figure 5. Average MPRE by chamber and dimension across Congresses for bills with five or more votes.
Note. Dotted line indicates 10% standard for significant dimensions. MPRE = marginal proportional reduction in error.

in Figure 2. In the House, the dimensions seem to have consistently decreasing MPRE across dimensions for the 93rd and 103rd Congresses, but there is wide variation across policy domains for the 110th Congress. In the Senate, four of the six policy domains show consistently decreasing MPRE across dimensions but Agriculture consistently shows an uptick in the fifth dimension. All of this is consistent with having a strong first dimension—the MPRE for the first dimension is consistently strong—but also suggests that in some cases like Agriculture in the Senate the other dimensions can have an impact similar to the first dimension.

Bill Episodes

In Figure 5, the mean MPRE of each dimension through the fifth dimension is plotted for bills that received at least five recorded votes.[9] House and Senate MPRE values are about the same, well above the 10% threshold across all five estimated dimensions. Even tripling the threshold would not bring the average MPRE below the threshold across dimensions. The proportion of five-vote bills manifesting at least three dimensions is 95% in the House and 91% in the Senate. This basic result confirms the observation in case studies of legislative battles that it is common that legislators' expressed preferences are more complex than is fairly characterized by a single dimension.

Figure 6 shows the distribution of dimensionality for bill receiving at least five votes by chamber. Almost all of these bills show multidimensionality. Only 2 of 297 bills in the House and 4 of 830 bills in the Senate manifest a

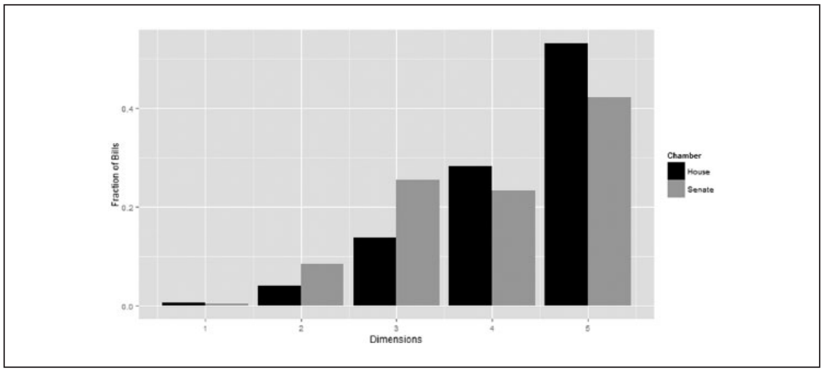

Figure 6. Frequency of dimensions manifest in bills with five or more votes, by chamber.
Note. Dimensions contributing at least 10% MPRE. MPRE = marginal proportional reduction in error.

single dimension. At the level of individual bills then we conclude that complexity is the norm and unidimensionality is rare.

These bill-level findings lead us to point out that assuming unidimensionality at the bill level can be problematic. One example that Krehbiel (1998) uses to illustrate his pivotal politics theory is the 1993 Economic Stimulus package proposed by President Clinton. In the Senate, this bill received 21 recorded votes. The MPRE for the first dimension is quite large (94%) suggesting that the left–right dimension was an important part of the conflict, but the MPRE for the second dimension is 76%, suggesting a second dimension at play. Similarly, in Cameron's (2000) analysis of veto bargaining episodes, we find that 15 bills that are classified as veto chains reach our threshold of at least 5 votes and exhibit multidimensionality. These bills are described in Table 1. Seeing that bills as complex as agriculture subsidies, civil rights legislation, and campaign finance regulations are multidimensional should come as no surprise to even casual observers of politics. There is little doubt that most, if not all, of these bills have a strong liberal–conservative dimension of conflict, but this alone is not the sole basis of conflict or compromise. Given that a number of the analyzed bills are multidimensional, Cameron's use of a 1-d cut point analysis is perhaps not the correct technique for these bills. Changes in bills that Cameron regards as "concessions" may simply be movement of the bills along other dimensions. Movements along other dimensions are still concessions, but those issues may or may not be as contentious as the liberal–conservative dimension.

Table 1. Multidimensional Veto Chains From Cameron (2000).

Congress	Chamber	Bill description	Number of dimensions	Number of votes
84	Senate	1956 Agriculture Act	4	7
85	Senate	Postal Rate Increase	5	9
85	Senate	1958 Agriculture Act	2	7
86	Senate	Airport Construction	2	6
87	Senate	Postal Rate Increase	3	8
90	Senate	Federal Pay Raise	4	13
91	Senate	Labor Appropriations	5	9
92	Senate	Emergency Employment Act	3	5
92	Senate	Federal Election Campaign Act (FECA)	3	14
94	Senate	Nuclear Insurance	3	5
94	Senate	Oil Price rollback	3	19
95	House	Cancel B1 Bomber	5	8
100	Senate	Omnibus Trade Act	3	46
100	Senate	Plant Closings	2	21
102	Senate	1991 Civil Rights Act	5	8

Identity of the First Dimension

A final but essential consideration is a comparison of the substantive identity of the first dimension generated at the three levels of aggregation. We are confident that the first-dimension OC results for the entire Congress define a liberal–conservative dimension. For each chamber over time, the Spearman's r between the first-dimension OC scores and the first-dimension DW-NOMINATE scores, calculated across all roll-call votes, exceeds 0.98. The issue is how those first-dimension OC scores are correlated with the first-dimension scores for the policy domains and bills. We summarize those correlations in Figure 7, which reports the mean absolute correlation[10] over time in each policy domain, and in Figure 8, which reports the mean absolute correlation over time for bills with more than five votes.

The first-dimension scores derived from votes associated with individual bills exhibit a much lower mean correlation with liberal–conservative scores than the scores derived from the more highly aggregated policy domains. In fact, the bill correlations are very modest whereas the policy domain correlations vary from uniformly high (government management of the economy) to fluctuating and often low (agriculture). Plainly, as we increase the level of aggregation from bills to policy domains to Congresses, we reduce the inferred dimensionality of the policy space and enhance the strength of the liberal–conservative dimension.

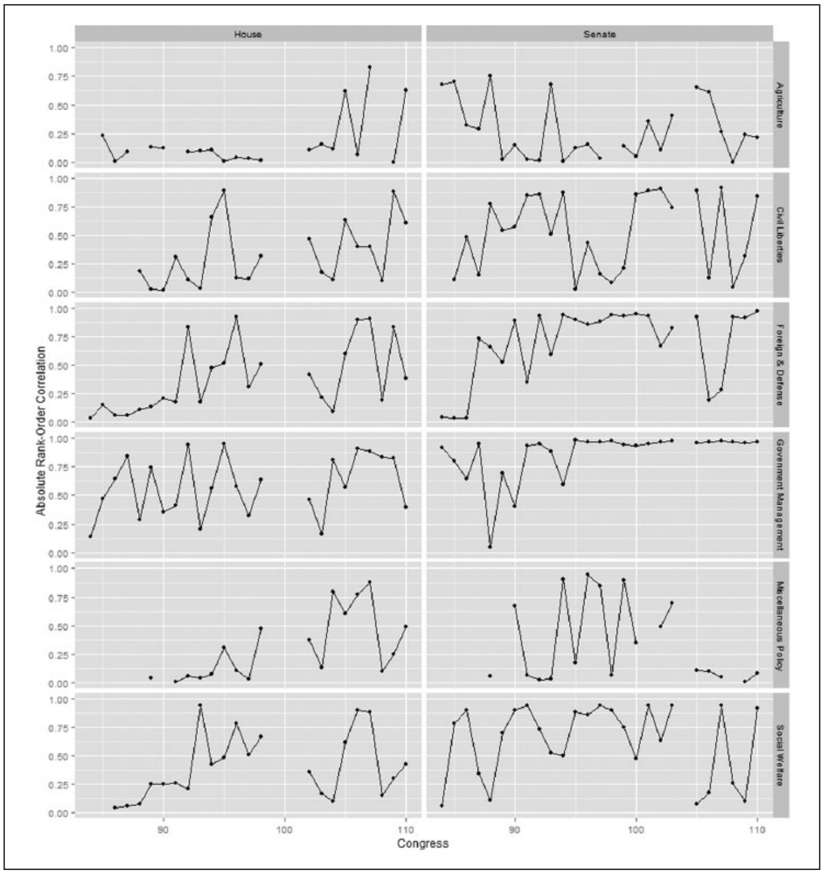

Figure 7. Correlation of overall chamber one-dimensional OC scores with one-dimensional OC scores for policy domains, by Congress-Chamber.
Note. OC = optimal classification.

Dimensionality in the House and Senate

Our main inference—that dimensionality varies with level of aggregation—appears to apply to both chambers with equal force. Yet, there are very good reasons to think that the two chambers would differ systematically in the dimensionality of the observed policy spaces in floor voting.

The rules and practices of the House and Senate differ in several significant ways that might lead to us expect systematic differences in the dimensionality of floor voting. First, until 1971, very few amendments to bills

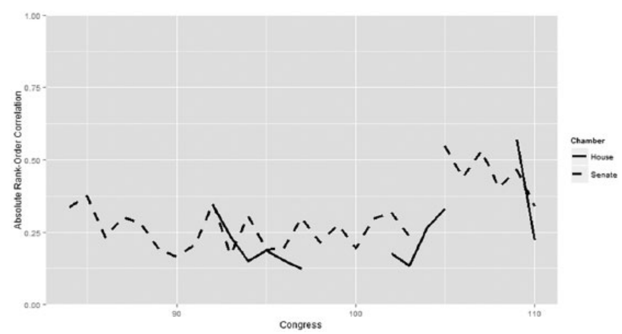

Figure 8. Mean correlation across bills of overall chamber one-dimensional OC scores with one-dimensional OC scores for bills with five or more votes, by Congress-Chamber.

Note. Missing values indicate that there were no bills with five or more votes that were not lopsided, so OC scores could not be reliably estimated. A vote here is considered lopsided if less than 2.5% of legislators vote with the minority. OC = optimal classification.

received recorded votes in the House. In 1971, for the first time, the House allowed recorded votes in the Committee of the Whole where the only votes on most amendments occurred (Roberts & Smith, 2003). The result is that far fewer House bills had as many as five roll-call votes and House bills were subject to far fewer votes, at least on average. With fewer amendments subject to votes, the observed dimensionality in voting associated with House bills would be expected to be lower than that for the companion action in the Senate.

Second, as a matter of standing rules, non-germane amendments are allowed in the Senate but not in the House. This creates the possibility that a wider variety of issues and divisions among legislators will emerge during consideration of a bill in the Senate. The possibility of a wider variety of issues and divisions in the Senate may yield more observed dimensions in voting behavior.

Third, special rules in the House may further limit amendments and the observed dimensionality of the related voting behavior. However, waivers of standing rules also are common in special rules, including waivers that allow some amendments to receive votes that otherwise would not be in order. On balance, we are quite certain that the revolution in the use of special rules in the 1980s (Bach & Smith, 1988; Roberts & Smith, 2007; Smith, 1989) reduced the number and variety of floor amendments and should have reduced the observed dimensionality of floor voting.

To further explore House–Senate differences, we have identified the House and Senate bills associated with the list of important measures on the Clinton–Lapinski list (Clinton & Lapinski, 2006). Of the 500 top measures on the Clinton–Lapinski list, 124 fall in the period studied here and have identifiable companion bills in both chambers, but only 32 pairs of those 124 have a minimum of five votes associated with both bills. We generated OC results for each bill in 32 matched pairs of bills with a minimum of five votes and for all votes associated with the 124 bills. These more limited comparisons give us a basis for hoping that the House and Senate analyses are based on roughly the same range of policies.

Contrary to expectation, there are no significant differences between the House and Senate on these bills. This is true for both the 32 matched pairs with a minimum of five votes per bill and for all votes associated with the 124 pairs. Thus, the tentative inference is that the larger number of Senate votes yields more observable dimensions on these bill episodes.

Conclusion

One of the major developments in the U.S. Congress over the past few decades has been the emergence of an unprecedented level of partisan polarization in the House and Senate. The levels of intra-party voting cohesion and inter-party voting conflict have increased for all four party caucuses that make up the Congress. This has produced an aggregate voting record that shows minimal differences between members of the same party and increasing differences between the two major parties, a process that makes the voting record appear more and more unidimensional. These patterns have been well documented by Keith Poole, Howard Rosenthal, and their colleagues.

We take a different tack in this article by reporting findings for a decomposition of the roll-call voting record of the U.S. Congress to determine the effect of the level of aggregation on the observed dimensionality of the policy space. In doing so, we identify some but certainly not all of the ways in which the aggregation of the voting record affects the observed dimensionality of the policy space. For the 1955 to 2008 period (84th-110th Congresses), we apply OC algorithms to votes aggregated to the level of the individual bill policy area to measure dimensionality. Whether examining eigenvalue or the MPRE, our results demonstrate that complexity in voting patterns of individual bill episodes is the norm, that aggregating to higher levels reduces the observed dimensionality, and that the liberal–conservative dimension appears more dominant in more highly aggregated analyses. To be clear, we are not disputing that first-dimension NOMINATE provide accurate measures of aggregate roll-call behavior, we are simply suggesting that aggregate measures are likely not the correct measure for studying individual bill episodes.

The differences between bill-level and aggregate-level inferences about dimensionality appear to be a product of the high frequency with which one-time issues are considered on the floor. Liberal–conservative, or left–right, divisions are exhibited on votes associated with many, if not most, bills, but other divisions among legislators often are visible on amendment or motion to recommit votes and are unique to one or just a few bills. In aggregate analyses, the liberal–conservative dimension dominates and the other divisions look like minor errors that do not define statistically important dimensions. For legislators seeking to pass a bill, these other divisions, which can be generated by a handful of legislators, may be critical to passing a bill.

Our results suggest that many of the conclusions from the theoretical and empirical literature on the U.S. Congress that uses a unidimensional model may not be based on sound empirical footing. The convenience of treating the congressional voting record as unidimensional is most justified when the analytical purpose is consistent with a high level of aggregation, as when we seek to compare parties across a century of Congresses. It is less justified when we are developing and testing theory about building coalitions, leadership strategies, special rules, or outcomes for individual bills.

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Notes

1. In addition to the studies mentioned in the text, other studies address dimensionality. Potoski and Talbert (2000) and Talbert and Potoski (2002) follow Poole and Rosenthal in arguing that floor voting is characterized by low- or unidimensionality, but that legislation and bill sponsorship shower higher dimensionality. Our hunch, reflected in the argument of this article, is that the high level of aggregation of roll-call votes in these studies masks even greater dimensionality at the level of bill episodes. Jenkins (1999, 2000), Wright and Schaffner (2002) and

Wright and Winburn (2003) find that the presence of strong parties and parity in the two parties in state legislatures and the Confederate Congress reduce the dimensionality of the observed policy space. These studies, too, were conducted on highly aggregated sets of roll-call votes.

2. See Jeong, Lowry, Miller, and Sened (2014) for an excellent analysis of the multidimensional policy space present in votes on the 1978 Energy Act.
3. Optimal classification (OC) is a non-parametric technique for estimating the ideal points of legislators and the underlying dimensions present in voting. Poole (2000, 2005) demonstrates that optimal classification is robust with a small number of votes. For the analysis reported, we eliminated all votes with less than 0.5% percent of legislators on the losing side of the vote. OC, unlike NOMINATE methods, can be used when votes scale perfectly, an important feature when scaling just a few votes. First-dimension OC scores correlate with DW-NOMINATE scores at 0.98 or higher in each of the Congresses in our series.
4. In 2009, the Netflix Prize of US$1 million was awarded to the first team able to beat Netflix's existing movie ratings prediction algorithm by 10%.
5. Of course, marginal proportional reduction in error (MPRE) reaches 100% when we increase the number dimensions to 1 for which all votes are predicted correctly. In the situation when aggregate proportional reduction in error (APRE) increases from 99% to 100%, this small change in APRE could hardly be associated with an additional meaningful dimension. This could in theory lead to concluding that there are multiple meaningful dimensions when only one is substantively meaningful, but that is not the case here. For chamber-Congresses analyses the mean APRE for 1-d models is 61% and all reported APREs are below 90%. For our Clausen groups analyses, the mean APRE for 1-d models is 72% and 84% are below 90%. For individual bill episodes, the mean APRE for 1-d models is 79% and 82% are below 90%. Thus, we think it is clear that findings of multidimensionality are based on substantively significant dimensions and not mere statistical artifacts.
6. We would say there are k dimensions manifest as long as all previous dimensions had MPREs greater than 10%. Unlike APRE, MPRE can increase and decrease and can exceed 10% after declining below 10%. We count the number of dimensions to be the greatest number of dimensions before the *first* time MPRE dips below 10%.
7. See Poole and Rosenthal (2007, p. 144) and Poole, Sowell, and Spear (1992).
8. First-dimension OC scores correlate with DW-NOMINATE scores at 0.98 or higher in each of the Congresses in our series.
9. We examine bills that receive at least five non-lopsided votes. For the analysis reported, we eliminated all votes with less than 0.5% of legislators on the losing side of the vote.
10. We calculate the correlation in each case (policy domain or bill), take the absolute value, and then take the average.

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Author Biographies

Jason M. Roberts is an associate professor in the Department of Political Science at the University of North Carolina at Chapel Hill. His research focuses on the American Congress, congressional elections, and institutional change.

Steven S. Smith is Kate M. Gregg Distinguished Professor of Social Sciences and Director of the Weidenbaum Center at Washington University in St. Louis. His research focuses on the institutional development of the U.S. Congress and the attitudes and behavior of the American public.

Stephen R. Haptonstahl is a data scientist at National Public Radio. His research focuses on mathematical models of politics.