# The Hunt for Party Discipline in Congress

NOLAN McCARTY *Princeton University*
KEITH T. POOLE *University of Houston*
HOWARD ROSENTHAL *Princeton University*

*We analyze party discipline in the House of Representatives between 1947 and 1998. The effects of party pressures can be represented in a spatial model by allowing each party to have its own cutting line on roll call votes. Adding a second cutting line makes, at best, a marginal improvement over the standard single-line model. Analysis of legislators who switch parties shows, however, that party discipline is manifest in the location of the legislator's ideal point. In contrast to our approach, we find that the Snyder-Groseclose method of estimating the influence of party discipline is biased toward exaggerating party effects.*

In the past several years there has been renewed scholarly investigation of how political parties and their leaders influence legislative institutions and behavior (see Aldrich 1995; Cox and McCubbins 1993; Rohde 1991; Sinclair 1995). Much of the focus is on the collective action problems that are inherent in the legislative and electoral processes. Cox and McCubbins (1993), who conceptualize political parties as cartels that direct legislative activity to enhance the collective electoral fortunes of their members, provide a typical variant on this theme but by no means the only one. The primary function of such a cartel is to build a collective reputation on which its members can run. They argue, however, that without strong leadership members have individual incentives to engage in legislative activities (such as pork) that diminish the collective reputation.

The focus on collective action has generated much interest in the cohesiveness of parties as floor coalitions.[1] The principal prediction is that a party produces a more cohesive coalition than would be possible if members were to act on their individual preferences. Rohde (1991) uses evidence of greater party cohesion since 1975 to demonstrate an increasing role of party in the postreform House. Aldrich, Berger, and Rohde (1999) choose party voting as the main dependent variable to test the predictions of "conditional party government," and Cox and McCubbins (1993) use member support on leadership votes to test for the role of leaders in creating voting coalitions. Furthermore, some scholars, including Rohde (1991), see a reassertion of party strength behind the growing cohesiveness and polarization of congressional parties since the mid-1970s (for alternative explanations, see King 1998; McCarty, Poole, and Rosenthal 1997; and Poole and Rosenthal 1984).

As Krehbiel (1993, 1998) points out, however, the patterns of behavior uncovered in these empirical studies are consistent with both theories of strong, influential parties and nonpartisan models in which member preferences are sorted along party lines. This dilemma is exacerbated by the difficulty of measuring legislative preferences. Ideally, some exogenous measure of member preferences should be used to test party theories. Voting behavior under the null hypothesis of no party influence then could be compared with actual voting behavior. The problem is that the usual measures of legislative preferences are derived from the voting behavior itself.

We attempt to untie this Gordian knot. We begin by reviewing the evidence from work on congressional roll call votes that hypothesizes sincere spatial voting. We discuss how this evidence suggests the presence of some party discipline. We next examine the party discipline model of Snyder and Groseclose (2000) and argue that their method both seriously biases the estimate of ideal points for ideological moderates and overestimates the extent of party discipline. We provide a compelling theoretical illustration of the bias. A more extensive technical discussion can be found in McCarty, Poole, and Rosenthal (2000, Appendix B).

To assess party discipline (or pressure) properly, we propose an alternative approach. The basic idea is very simple. We start with Krehbiel's (1993, 1998) proposal that the spatial model of purely preference-based voting is the appropriate benchmark for evaluation of models that incorporate party effects. In one dimension, the spatial model asserts that, on each roll call, "yea" and "nay" voters are separated by a cutpoint on the liberal-conservative continuum. If the Republicans apply pressure to their membership, some moderates to the left of the "sincere" cutpoint will vote with the conservative wing. Republicans will have a cutpoint to the left of the sincere cutpoint, and Democrats will have a cutpoint to the right of it.

To illustrate this point, we assume some overlap

Nolan McCarty is Associate Professor of Politics and Public Affairs, Princeton University, Princeton, NJ 08544-1012. Keith T. Poole is Kenneth L. Lay Professor of Political Science, University of Houston, Houston, TX 77204-3472. Howard Rosenthal is Roger Williams Straus Professor of Social Sciences and Professor of Politics, Princeton University, Princeton, NJ 08544-1012.

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[1] Cohesiveness is not the only or even the most active focus of this line of inquiry. Much of it explores the role of party leaders in setting the legislative agenda.

between the parties: Certain Democrats, historically largely southerners, are more conservative than the most liberal Republicans. A single cutpoint model would always predict that either some Republicans vote with the majority of Democrats or vice versa. But if enough pressure were applied to turn a roll call into a straight party vote, the voting pattern would be captured better by a Democratic cutpoint to the right of the most conservative Democrat and a Republican cutpoint to the left of the most liberal Republican. More generally, when one or both parties apply pressure, the voting patterns should look as if there were separate cutpoints for each party, with the Democratic cutpoint being to the right of the Republican cutpoint. Therefore, if pressure is important, we should find a better fit to the data when we estimate two cutpoints rather than one. (We will return to this point in discussing Figure 4.)

To keep the analysis as simple as possible, we use nonparametric optimal classification analysis; legislator ideal points and roll call cutpoints are jointly rank ordered to maximize predictive success on roll call votes (Poole 2000). By classifying the voting of each party independently and then comparing the results to classifying both parties together, we can evaluate the maximum possible improvement in correct classification attributable to party discipline. An advantage of the cutpoint approach is that it does not require any assumptions about which specific roll calls are subject to party pressure. A particular advantage of the nonparametric approach is that it assumes only that the amount of pressure applied to individual members does not change the order of their induced ideal points. It does not require making parametric assumptions about how pressure varies with the ideal point of the individual member, such as equal pressure being applied to all. On the basis of our optimal classification analysis, we conclude that allowing for party discipline affords only a very marginal improvement over the sincere spatial model, particularly in recent Congresses.

Where, then, is party discipline? We argue that the main influence of party discipline is not on the votes on specific roll calls but on the choice of ideal point made by the representative. The smoking gun is provided by the great change in ideological position demonstrated by the few legislators who have switched parties. Wayne Morse and Strom Thurmond are two well-known examples in the postwar Senate. The Democrats who defected to the Republicans after the 1994 election made equally dramatic shifts. Our finding that parties shape ideal points ends our hunt for party discipline in roll call voting.

## INDEPENDENT VOTING ON THE FLOOR: THE EVIDENCE FROM THE SPATIAL MODEL

The standard spatial model provides a benchmark approach to independent floor voting. Poole and Rosenthal (1991, 1997) demonstrate that the model is quite successful in accounting for floor decisions. With two dimensions, one can correctly predict roughly 85% of the individual decisions—even on close roll calls—between 1789 and 1985. McCarty, Poole, and Rosenthal (1997, 7) report additional results for 1947–95.[2] In recent Congresses, a one-dimensional model classifies nearly 90% of the individual decisions.

The spatial estimates strongly suggest that party influence underpins much of this remarkable classification success.[3] First, in Congresses in which voting is largely one dimensional, party-line votes are along the main dimension. The distribution of ideal points is strongly bimodal. The two parties appear as two very distinct "clouds" that barely overlap, particularly in recent years (As an illustration, see McCarty, Poole, and Rosenthal 1997, 11.) The presence of a "channel" between the clouds suggests that party affiliation may discipline the roll call voting behavior of members. The main dimension of political conflict clearly appears to reflect partisan conflict. Parties perhaps also influence their members' votes on specific roll calls.

Second, in Congresses in which voting is two dimensional, there also are two distinct clouds separated by a channel. Party-line votes are no longer on the main dimension but are a blend of the first and second dimensions. (See Poole and Rosenthal 1991, 233, or 1997, 44, for an example.) An interpretation of such plots is that ideal points projected onto roughly a 45° line represent the ideological (liberal-conservative) dimension. The orthogonal projection, roughly at −45°, represents a party loyalty or valence dimension. Most votes occur along the main, 0° dimension. On these votes, the legislator's decision depends both on ideology and on party loyalty.

Although the spatial model shows that the structure of voting coalitions in Congress coincides strongly with party affiliation, it does not prove that party per se has any influence on voting behavior. Party-line voting is, of course, consistent with both strong party models and ideological models in which preferences are sorted by parties. In the sections that follow, we review a recent attempt to separate partisan effects from preferences and then propose a method of our own.

## THE SNYDER-GROSECLOSE MODEL OF PARTY DISCIPLINE

One inherent problem in identifying the effects of party is that we observe only behavior, which is presumably a mix of individual preferences and party influence. This problem is particularly acute with congressional voting data. If party discipline is exercised on floor votes, the ideal points estimated on the assumption of independent spatial voting may be very biased estimates of legislator preferences. If party influences these estimates, it is inappropriate to use it as a control for preferences when testing for a party effect. Snyder and

[2] These authors use the Poole and Rosenthal (1991, 1997) NOMINATE methods. Both NOMINATE and the Heckman and Snyder (1997) method are parametric. The results of the two approaches are very similar, particularly on the first and second dimensions.

[3] The spatial model does, however, strongly outperform a model of straight party voting.

**FIGURE 1. A Six-Member Legislature**

| | Left | | | | | Right |
|---|---|---|---|---|---|---|
| **Legislator** | 1 | 2 | 3 | 4 | 5 | 6 |
| **Party** | D | D | D | R | R | R |

Groseclose (2000) note this potential for bias. They propose a method for estimating unbiased ideal points and then using them to estimate the effect of party discipline.[4]

The basics of the one-dimensional Snyder-Groseclose model are as follows. On roll call $j$, a legislator $i$, if a Republican has an induced ideal point $x_{ij} = x_i$; on roll call $j$, a legislator $i$, if a Democrat, has an induced ideal point $x_{ij} = x_i + \gamma_j$. In other words, the true ideal points of the Democrats, the $x_i$, are displaced by the amount of party pressure given by $\gamma_j$.[5] It turns out that only the relative amount of party pressure matters in the model, so the ideal points of the Republicans can just be given by their true values. For the difference in pressure to be consistent with discipline, we would expect that pressure must move Democrats in a liberal direction relative to Republicans. Thus, pressure works to increase the separation of the parties. If preferences are viewed on a scale, such that the left end is liberal and the right end conservative, then we would expect $\gamma$ to have a negative sign.

Snyder and Groseclose argue that, because there is little need to apply party discipline on votes not expected to be close, ideological position taking should occur on lopsided votes. These votes, such as those with margins greater than 65 to 35, could be used to estimate the true ideal points. On these votes, the $\gamma_j$ would be zero. The true ideal points could then be used to estimate the $\gamma_j$ on close votes, such as those with margins less than 65 to 35.

In brief, their procedure is as follows. Stage 1: Use votes with margins greater than 65 to 35 to estimate the ideal points, $x_i$. Stage 2: On the remaining votes, for each roll call $j$, estimate the following ordinary least-squares (OLS) model:

$$Y_{ij} = \beta_0 + \beta_1 x_i + \beta_2 D_i, \qquad (1)$$

where $D_i = 1$ if legislator $i$ is a Democrat, 0 if Republican, and $Y_{ij} = 1$ if $i$ votes "yea," 0 if $i$ votes "nay." Equation 1 is equivalent to $Y_{ij} = \beta_0 + \beta_1 x_{ij}$ when

$$\gamma_j = \beta_2/\beta_1. \qquad (2)$$

To see this, note that for Democrats equation 1 implies $Y_{ij} = \beta_0 + \beta_1 x_i + \beta_2 = \beta_0 + \beta_1(x_i + \beta_2/\beta_1)$. As we noted above, the party pressure model predicts a negative estimate for $\gamma$ when preferences are scaled with Democrats on the left (as we assume they are). Therefore, the two estimated $\beta$s should be of opposite sign.

This method is likely to generate the inference that party pressure is substantial even when all voting is preference based. Consider, for example, a six-member legislature with the party affiliations and spatial preferences given in Figure 1. If all voting in this legislature is spatial without error, there are only twelve possible voting configurations, which are given in Figure 2.

Stage 1 of the Snyder-Groseclose method estimates a preference score using only voting patterns 1–10. But voters 3 and 4 cast identical votes in each of these patterns, so any scaling procedure will estimate them as having the same position. Thus, stage 1 provides biased estimates of the preferences of moderates. There is not enough information in the lopsided votes to discriminate "left" moderates from "right" moderates. The preference ordering that maximizes the classification of votes is shown in Figure 3.

In stage 2, the preferences in Figure 3 and party affiliation are used to explain vote patterns 11 and 12. The votes of legislators 1, 2, 5, and 6 are correctly classified on the basis of the preference estimates, but the votes of legislators 3 and 4 cannot be. Yet, since 3 and 4 are members of different parties, adding the party variable to the model increases its explanatory power even though voting is purely preference driven.

Our example extends to larger legislatures. In general, with perfect spatial voting, a first stage based only on lopsided votes will produce identical preference estimates for all members in the interval between the 35th and 65th percentiles. The second stage will produce a spurious party effect so long as party and ideology are correlated within this interval.[6] We now present an alternative procedure that maintains the essential features of the Snyder-Groseclose model of discipline.

## A NONPARAMETRIC MODEL

All specifications of a spatial model of voting have two critical elements: ideal points for the legislators and

[4] See Jenkins 1999 for an application of this method.

[5] Snyder and Groseclose (2000) allow the displacement to be other than a constant, but their empirical work relies on the simple constant displacement model. They also allow for multiple dimensions, but the unidimensional case gives the intuition of their more general model.

[6] Given the assumptions of no voting error and no overlap of preferences between the parties, this example is somewhat special. In McCarty, Poole, and Rosenthal 2000, Appendix B, we present Monte Carlo evidence that shows how this result extends to large legislatures in which, as in the Snyder and Groseclose approach, there is some error in voting and the distribution of preferences of each party overlaps.

**FIGURE 2. Perfect Spatial Voting in a Six-Member Legislature**

| Vote on Roll Call | Legislator: Left 1 | 2 | 3 | 4 | 5 | Right 6 |
|---|---|---|---|---|---|---|
| 1 | Y | Y | Y | Y | Y | Y |
| 2 | N | N | N | N | N | N |
| 3 | Y | N | N | N | N | N |
| 4 | N | Y | Y | Y | Y | Y |
| 5 | Y | Y | N | N | N | N |
| 6 | N | N | Y | Y | Y | Y |
| 7 | Y | Y | Y | Y | N | N |
| 8 | N | N | N | N | Y | Y |
| 9 | Y | Y | Y | Y | Y | N |
| 10 | N | N | N | N | N | Y |
| 11 | Y | Y | Y | N | N | N |
| 12 | N | N | N | Y | Y | Y |

*Note*: Y = "yea" vote, N = "nay" vote.

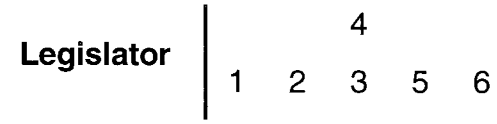
cutpoints (or separating hyperplanes) for the roll calls. The Snyder-Groseclose model, with a discipline parameter to each roll call, is isomorphic with one in which each party has its own cutting line (see Appendix A). That is, moving the ideal points for all Democrats to the left by a magnitude $\gamma_j$ is equivalent to moving the cutpoint for Democrats to the right by the same amount. Party discipline generally involves getting moderates to vote with extremists.[7] Therefore, if there is party discipline, the cutpoint for the Democrats should be to the right of the cutpoint for Republicans.[8]

Consider a one-dimensional spatial configuration. When the cutpoint is constrained to be the same for both parties, this produces the standard spatial model. For example, in Figure 4, with a common cutpoint, there are three classification errors: legislators 3, 11, and 15. When each party can have its own cutpoint, this produces a model that allows for party discipline. Moderate Democrats to the right of some Republicans can vote with the majority of their party. Moderate Republicans to the left of some Democrats can vote with the majority of their party. The best cutpoint for the Republicans in Figure 4 remains the common cutpoint (the solid line). Legislator 15 is the only Republican classification error. But the best cutpoint for the Democrats is to the right of the common cutpoint (the dashed line), which leaves only legislator 3 as a classification error. Rather than estimate either

[7] Poole and Rosenthal (1997, 155–7) document that there are very few "both ends against the middle" votes in which extremists defect.

[8] The predicted order of cutpoints is equivalent to the prediction that $\gamma$ is negative.

**FIGURE 3. Preference Order Based on Lopsided Votes**




*Note*: Since voter 3 and voter 4 voted identically on lopsided votes, any estimation procedure will generate identical ideal points.

the one- or the two-cutpoint model via a metric technique, such as Poole and Rosenthal's (1991) NOMINATE or the Heckman and Snyder (1997) method, one can simply find the joint rank order of legislators and cutpoints that will minimize classification error. Poole (2000) presents an efficient algorithm that very closely approximates the global maximum in correct classification.[9] Note that this method, in contrast to equation 1, does not require a uniform adjustment in the ideal points of all party members. Only moderates need to be disciplined. All that is required is a displacement of the cutpoint.

Our empirical analysis involves not only testing the implications of our methodological critique of Snyder and Groseclose but also testing the implications of their model of party discipline with our two-cutpoint model. We begin by stating three methodological predictions. Each is consistent with the mismeasurement of preferences in the Snyder-Groseclose framework under the hypothesis of purely spatial voting. In only one case is the prediction also consistent with their theoretical model. Therefore, verification of these relationships illustrates the inability of Snyder-Groseclose to distinguish party pressure from mismeasurement of preferences. The three methodological predictions are as follows.

PREDICTION 1. *Estimate the rank order of ideal points by one-dimensional optimal classification first using all roll call votes and then using only lopsided votes. The correlation between the two rank orders will be greater for extremists (the first and last one-third of the all votes distribution) than for moderates (the middle one-third). This prediction is consistent with the Snyder-Groseclose assertion that party pressure primarily affects moderates, but it also follows from our claim that, if there is preference-based voting, ideal points of moderates will be inaccurately recovered if only lopsided votes are used to estimate ideal points.*

PREDICTION 2. *Similarly, when the rank order is estimated first on all roll call votes and second on only close votes, the correlation between the two rankings will be greater for moderates than for extremists. The motivation for this prediction is similar to that of the first. If there is preference-based voting, the ideal points of extremists will be inaccurately recovered if only close votes are used to estimate ideal points. This prediction is inconsistent with the Snyder and Groseclose model, which implicitly assumes that extremists will have preference estimates on pressured votes similar to those on unpressured votes.*

PREDICTION 3. *The correlation between the two rank orders for moderates will be higher if there is preference-based voting, lower if there is party discipline. The reason is that, if there is discipline only on close votes, as claimed by Snyder and Groseclose, the all-votes estimates will mix preference-based lopsided votes and disciplined close votes. The estimates of close votes will have more distortion of the true ideal points.*

We now turn to testing hypotheses from the party discipline model. In all cases, the null hypothesis of preference-based voting predicts no difference.

HYPOTHESIS 1. *Classification should be substantially higher with a two-point model than with a one-point model. Note that classification cannot be lower with the two-point model.*

HYPOTHESIS 2. *The improvements in classification should be greater on close votes. Since the Snyder-Groseclose model predicts that rational parties will whip close votes, the incremental predictive power of the two-cutpoint model should be higher on those votes.*

HYPOTHESIS 3. *The rank order of the legislators should disclose more separation of the parties in the one- than in the two-cutpoint model because the former ignores party discipline. Moving Democrats to the left and Republicans to the right should pick up some of the effects of party pressure. In contrast, in the two-cutpoint model, each legislator's ideal point can take on its true rank order position, because the cutpoints can pick up the effects of party discipline.*

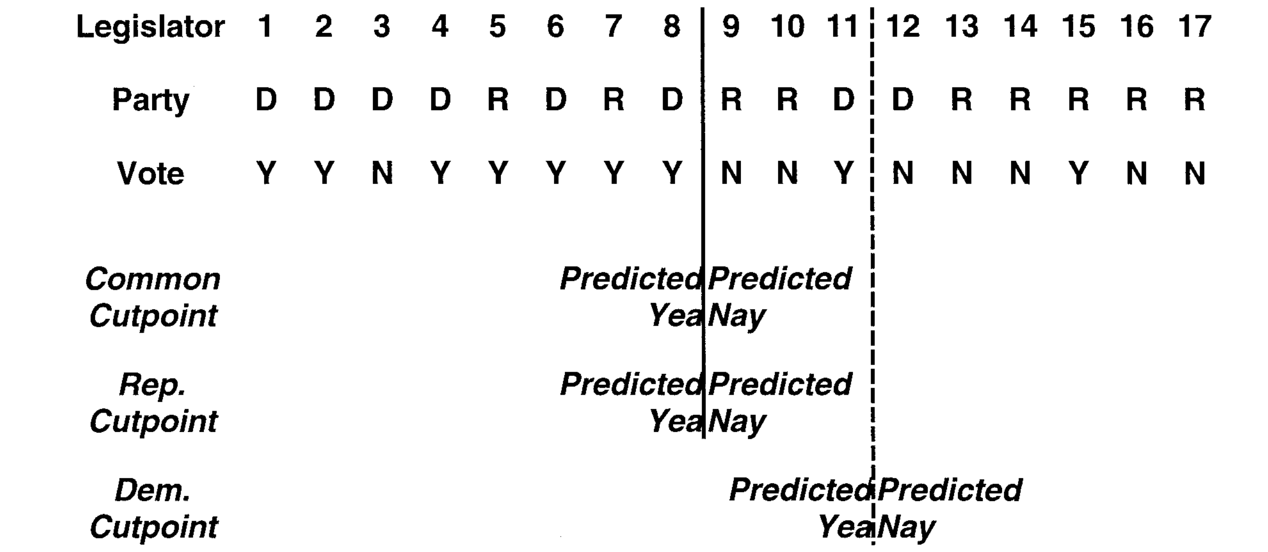
HYPOTHESIS 4. *The separation of the cutpoints should be greater on close votes. The identifying assumption of the Snyder-Groseclose model is that party pressures are more likely on close votes. Therefore, under their assumptions, the distance between the Democratic and the Republican cutpoint should be greatest on those votes.*

HYPOTHESIS 5. *The estimated cutpoint for the Democrats should be to the right of the estimated cutpoint for the Republicans.*

It should be noted that some instances of party pressure may be masked. Consider a legislature with no party overlap—all Democrats are to the left of all Republicans. Suppose that, were there no pressure, a Republican bill would be rejected by a majority composed of all Democrats and some moderate Republi-

[9] Although the underlying assumptions are very different, in one dimension this method is essentially equivalent to classical Guttman scaling.

**FIGURE 4. Cutpoint Models**




can defectors. If the Republicans then apply pressure to the defectors, the resulting party-line vote will appear to be consistent with preference-based voting. When ideal points are estimated correctly, the true explanatory power of party may be masked. Indeed, when there is no overlap in the distribution of party ideal points and there is errorless spatial voting, it is impossible to identify party pressure effects. This masking is inherent to spatial analysis. It would confound the Snyder-Groseclose model as well as our optimal classification method. Albeit important, the question we can ask is limited to whether allowing for party discipline can improve on the classification of a purely preference-based model.[10]

With our optimal classification method, it is possible to calculate an upper bound for the amount by which party pressure can increase vote classification. This bound depends on the overlap between the two parties. This upper bound represents the classification on a strict party vote of a two-cutpoint model (perfect classification) minus the classification of a strict party vote using a single cutpoint.[11] When there is no overlap between the parties, a single cutpoint correctly classifies a strict party vote, so there can be no classification gain for the two-cutpoint model. The greater the party overlap, the worse a one-cutpoint model does in explaining a strict party vote. Thus, the maximum classification gain increases in the overlap.

If we use the configurations of preferences that emerge from optimal one-cutpoint classification to measure overlap, the maximum classification gain from cutpoints consistent with party pressure (i.e., $D > R$) ranges from 0 in the 80th House (where there is zero overlap) to 16% in the 92d House. The average upper bound over all the Congresses we analyze is 5%. It is important to remember that these upper bounds are simply for roll calls consistent with party pressure (i.e., Democratic cutpoint to the right of the Republican cutpoint). Perfect classification is the upper bound if we allow other cutpoint configurations (e.g., the Republican cutpoint on the right). Furthermore, as we discuss below, optimal classification with a single cutpoint will underestimate party overlap, which would lead to the underestimation of these upper bounds.

## TESTS USING THE NONPARAMETRIC MODEL

### Classification with One Cutpoint

We begin with the three predictions concerning the correlation of ideal points, which we tested with actual data. We first performed optimal one-cutpoint classification using all roll call House votes from the 80th through the 105th Congress. If the basic spatial model is correct, this procedure should produce a rank order of legislator ideal points that is very close to the true order. Next, we did optimal classification using only lopsided votes, or margins greater than 65 to 35. Finally, we did optimal classification using only close votes, or margins of 65 to 35 or less.

We then computed Spearman correlations between the lopsided and the all-votes rank orders for leftwingers, the one-third of legislators farthest to the Left in the all-votes classification; for moderates, the middle one-third; and for rightwingers, the one-third farthest to the Right. We expect these correlations to be high for the extremes but low for the moderates because the lopsided votes provide little information about the ideal points of moderates (prediction 1). Conversely, in

[10] This focus is consistent with a key point of Krehbiel (1998). He argues that the main empirical question should not be whether parties influence legislative behavior but whether partisan models represent a substantial improvement over those that assume autonomous legislators.

[11] Pressure beyond that necessary to generate a strict party vote cannot further increase classification.

**TABLE 1. Average Spearman Rank-Order Correlations of Legislator Ideal Points from Optimal Classification Analyses**

| | Estimates for Lopsided Votes versus All Votes | | | Estimates for Close Votes versus All Votes | | |
|---|---|---|---|---|---|---|
| Houses | Left Third | Middle Third | Right Third | Left Third | Middle Third | Right Third |
| 80th–90th (1947–68) | .86 (.07) | .44 (.30) | .91 (.05) | .58 (.13) | .97 (.03) | .51 (.17) |
| 91st–105th (1969–98) | .94 (.01) | .77 (.07) | .94 (.07) | .54 (.14) | .97 (.02) | .60 (.17) |
| 80th–105th (1947–98) | .90 (.07) | .63 (.26) | .93 (.06) | .56 (.14) | .97 (.03) | .56 (.17) |

*Note*: In each House, each 1/3 represents an $N$ of at least 145. Actual $N$'s are typically slightly larger because some districts have two distinct representatives serving in a given House as a result of deaths, replacements, etc. The averages are then computed as unweighted averages across the indicated set of Houses. Standard deviations are in parentheses.

the case of close votes and all votes, we expect the correlations to be high for moderates but low for the extremes.[12]

The hypothesized patterns occur, as shown in Table 1. Indeed, for the lopsided-all comparison, in every postwar House but one, the correlation for moderates is lower than for leftwingers and rightwingers.[13] Table 1 indicates that the middle correlation is particularly low in the years preceding passage of the major civil rights bills of the 1960s. In this period, an important second dimension (Poole and Rosenthal 1997) confounds the recovery of moderate positions on the first dimension. When the second dimension vanishes, even the middle correlations are reasonably high because the "errors" in voting provide some information about moderates. That is, for example, a relatively liberal moderate is still less likely to vote with the rightwingers than is a relatively conservative moderate, even on lopsided votes. Nonetheless, in accord with prediction 2, correlations for moderates are lower than for extremists.[14]

As predicted, these results reverse for the close-all comparison. The moderates always produce a correlation above 0.9. The leftwinger and rightwinger correlations are always below 0.9, usually much below, and in one case the correlation is negative.

The close-all correlations for moderates are strikingly high, predicted by preference-based voting but not by voting subject to party discipline (prediction 3). If the party discipline effect were important, we would expect lower rank order correlations, particularly for the House before 1980, when there was still considerable overlap in the ideal point distributions of the two parties.

[12] We focus on the rank-order correlation results since they are most consistent with our optimal classification approach. When we conducted each of these experiments using standard correlations, we found little substantive difference.

[13] A simple sign test for the observation of 25 successes in 26 trials has a $p$-value $< 10^{-6}$.

[14] Since hypothesis 1 may be consistent with either party or preference voting, we generated Monte Carlo data that imposed preference voting without party discipline. These results, listed in columns i and j of Table A1 in McCarty, Poole, and Rosenthal 2000, show that under pure preference voting the correlation between the true and estimated preferences is lower when only lopsided votes are used.

## Classification with Two Cutpoints

We now assess the ability of a party discipline model to improve on a preference-based model. Our criterion is percentage of votes correctly classified.

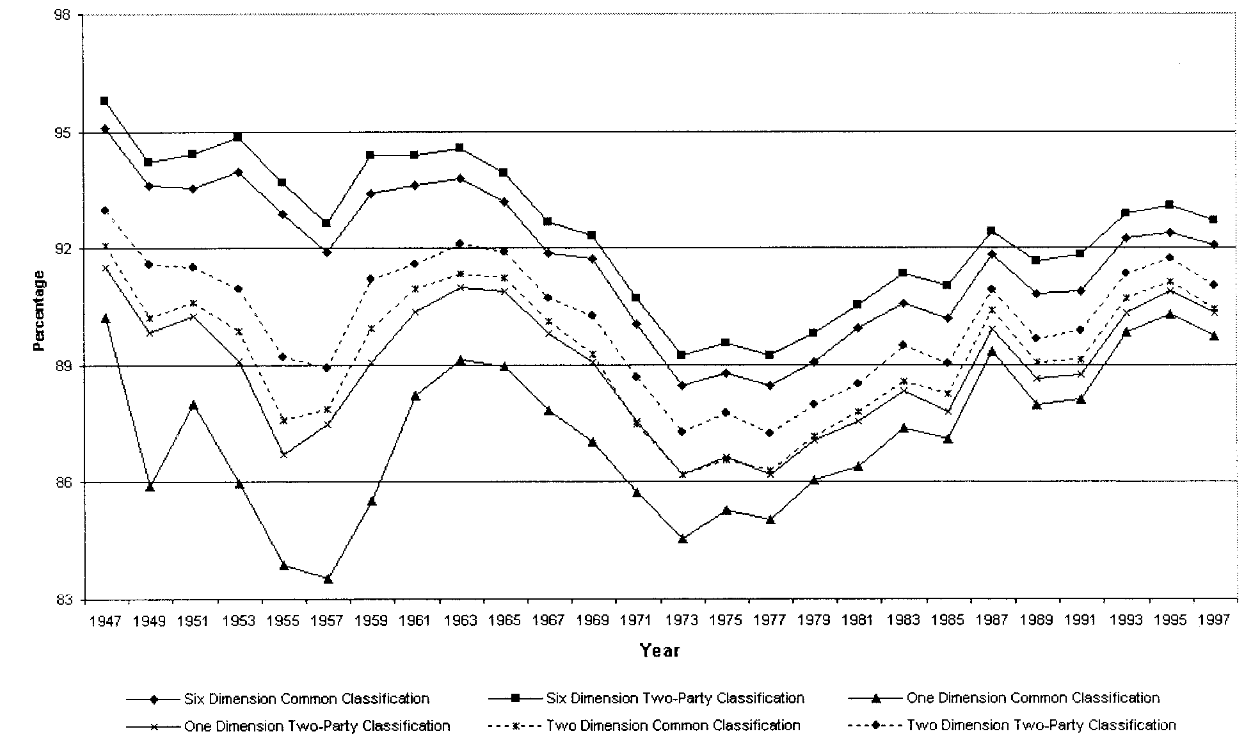
To find the highest classification possible for a party discipline model, there is a simple solution: Classify each party separately. This allows the cutpoint on each roll call to adjust to pressures internal to the party. Because the cutpoints can adjust, the true intraparty rank order of the ideal points can be found. The classification from this model then can be compared to that of a single-cutpoint model.

The outcome of this exercise appears in Figure 5, which shows results for a one-, two-, and six-dimensional model. The latter parallels the high dimensionality used by Snyder and Groseclose in their empirical work. For multiple dimensions, the cutpoint is replaced with a separating hyperplane.

In the case of one dimension, it is apparent that a two-party model adds little, particularly in recent Congresses. The improvement in the earlier Houses is at the level that results when a two-dimensional model with one cutpoint is used. The two cutpoints allow for southern Democrats to vote with northern Democrats on some issues as well as a coalition of conservative Republicans and southern Democrats in opposition to liberal Republicans and northern Democrats. Since the Democrats were the majority party in the conservative coalition era, these votes demonstrate a breakdown in party discipline that is exactly opposite the basic assumption of the Snyder-Groseclose model.

In the case of two or six dimensions, when two (as against one) separating hyperplanes are allowed, there is even less improvement than in the one-dimensional case. The improvement is almost always less than 1% for all postwar Houses. The reason is that, with one dimension, "party" picks up some effects that can be accounted for just as well as by a multidimensional preference-based voting model. The strength of the results in Figure 5 is further emphasized by two observations. First, some of the increase in fit is simply noise fitting due to the extra degrees of freedom. Second, classifying each party separately allows for "both ends against the middle" voting, when liberal

**FIGURE 5. Correct Classification in One, Two, and Six Dimensions**




*Note*: "Common" refers to optimal classification when all representatives are scaled together, and "two-party" refers to optimal classification of each party separately. The total classifications for the two cases are equal. The percentage correct for the "two-party" must exceed the percentage correct for the "common."

Democrats and conservative Republicans vote together. This last problem and other considerations led us to adopt a slightly different approach.

The remainder of the analysis in this section follows a two-stage procedure. First, using optimal classification, we estimate a one-dimensional spatial model that has a single cutpoint, common to both parties. Second, holding the rank order positions of the legislators constant at the positions produced by step 1, we then estimate separate cutpoints for the two parties. The two cutpoints must be placed to maintain polarity. That is, unlike the separate scalings reported in Figure 5, we did not consider improving classification by allowing moderates to be opposed by extremists at both ends of the spectrum. Bob Barr and Maxine Waters cannot vote together against Connie Morella. This constraint is fully consistent with the Snyder-Groseclose approach, which calls for an order-preserving shift in a party's ideal point distribution but not for a flip-flop.

The motivation for this two-step approach is that it is not possible to estimate jointly a single order for the legislators and two cutpoints for each roll call. The reason is that the rank order of the legislators within each party is pinned down only by the cutpoints for that party. Therefore, it is impossible to rank order either the legislators of a party or the cutpoints for that party with those for the other party. In contrast, once we fix the rank order of the legislators, we can estimate separate cutpoints and test theoretical predictions about them. We cannot directly test hypothesis 3, however, that preferences will show less party overlap in a one-point model than in a two-point model. That hypothesis could be tested only indirectly, by our test of prediction 3.

To justify holding the legislators constant, we computed within-party Spearman correlations between the rankings of the single-cutpoint model and the rankings when optimal classification is applied to the party separately. Recall that this separate classification is consistent with a party pressure model—there is a true underlying order of ideal points, but cutpoints are adjusted to reflect party pressure. As Table 2 shows, these correlations are remarkably high. For both parties, they have exceeded 0.95 since the mid-1960s. (Before then, some correlations were lower due to an important second dimension, as noted earlier.) The single-cutpoint ratings, particularly for the past 30 years, are likely to provide accurate rankings of the "true" ideal points within each party.

Note that Table 2 informs us that the relative order of legislators within each party is insensitive to whether we just assume pure preference-based voting or explicitly account for party pressure. The results do not rule out party pressure; rather, consistent with equation 1, party pressure is unlikely to change the relative order of induced ideal points. The results also do not rule out polarization due to party pressures, and the lack of overlap observed in the 1990s may be the outcome of party pressure. We will return to this point.

The two-cutpoint model creates only minor gains in

**TABLE 2. Correlations of Legislator Ideal Points from One- and Two-Cutpoint Models**

| | Average Within-Party Rank Order Correlations | |
|---|---|---|
| Houses | Democrats | Republicans |
| 80th–90th (1947–68) | .94 (.06) | .93 (.03) |
| 91st–105th (1969–98) | .99 (.01) | .98 (.01) |
| 80th–105th (1947–98) | .97 (.05) | .96 (.03) |

*Note*: The averages are unweighted averages across the indicated set of Houses. Standard deviations are in parentheses.

classification of all roll call votes, which is evident in Figure 6. As the second dimension has diminished in importance, these gains have declined to less than 0.5% in the last eight Congresses. In other words, adding a second cutpoint typically allows correct classification of only an additional 2 of the 435 representatives (assuming full turnout). Note that (1) the classification must get better with a second cutpoint, (2) the second cutpoint can just fit noise in the data (see Poole and Rosenthal 1997, 156), and (3) much of the improvement in classification occurs from using two cutpoints that have the Democratic cutpoint counterhypothesis to (i.e., left of) the Republican cutpoint, as will be seen in Table 3. An improvement of less than 1 percent is very minor.[15] Hypothesis 1 is not supported.

Figure 7 shows the gains for the two-cutpoint model for close and lopsided roll calls. It contains a little good news for party pressure advocates. The classification gain is greater on close votes than on all roll calls, but only since the mid-1960s. The evidence for the earlier Congresses reinforces our contention that the larger improvements in classification for those years, shown in Figure 5, are the work of a second dimension. If party discipline were at work, then the gain should not occur on lopsided roll calls, and in later Congresses there is systematically a greater gain on close than on lopsided votes. Some of the gain on close votes, however, must result from nondiscipline factors—such as noise fitting—that affect lopsided as well as close votes. The increase in the gain on close versus lopsided votes is roughly 1 percent, which suggests that party pressures are changing only about four votes per roll call on the close votes. At best, hypothesis 2 is weakly supported.

Another hypothesis derived from the party pressure model is also weakly supported. To test hypothesis 4, we computed for each House the average of the difference between the rank of the Democratic cutpoint and the rank of the Republican cutpoint and then divided by the number of legislators in the House. This procedure normalized the difference in the rank orders to a $-1$ to $+1$ scale, so that the Houses could be more easily compared. We used the difference rather than the magnitude of the difference between the ranks because the party pressure model predicts that the Democratic cutpoint will be greater than the Republican cutpoint ($D > R$).

We classified all roll calls into three types. For the first type, in line with hypothesis 5, the Democratic cutpoint is greater than the Republican cutpoint ($D > R$). Note that whenever there is some overlap in the ideal point ranks of the two parties, straight party votes are counted as $D > R$. The second type is clearly counterhypothesis roll calls, that is, $R > D$. Finally, for many roll calls (see Table 3) the relative location of the two party cutpoints is ambiguous, and we term this third type "undecided." Note that cutpoints interior to the legislators of a party can be identified for only a subset of roll calls.[16] A portion of our analysis will be restricted to such roll calls.[17]

When the ideal point distributions of the two parties have no overlap, as happened in the 80th House (1947–48), we cannot identify any roll calls as $D > R$, so the average difference must be less than zero. In contrast, when there is substantial party overlap, as in the 1970s, the party pressure model predicts that the average difference for close votes will be greater than zero and be greater than the average difference for lopsided roll calls. The average difference for lopsided roll calls should be near zero. The results, computed for all roll calls with interior cutpoints in both parties, appear in Figure 8.

The average difference for the close votes is indeed above zero for 19 of the 26 Houses. Since the 91st House, however, the average difference has been very close to zero, hovering around .02, or an average difference of about 8 or 9 ranks. In only three Houses, all in the two-dimensional 1950s and 1960s, does the figure exceed 0.1, or 10% of the House membership. To benchmark this difference, note that the normalized difference or overlap between the third rightmost Democrat and the third leftmost Republican averages 46% of the House membership for the 26 Houses we analyze; it exceeds 32% in all but the 80th, 84th, and 100th to 105th Houses. Moreover, note that this average difference is highly biased in favor of the party pressure model in that it does not include undecided roll calls, for example, votes on which the Republicans were unanimous but the Democratic cutpoint was to the left of the leftmost Republican. On such roll calls, most likely party discipline broke down among the Democrats, so that $D < R$.

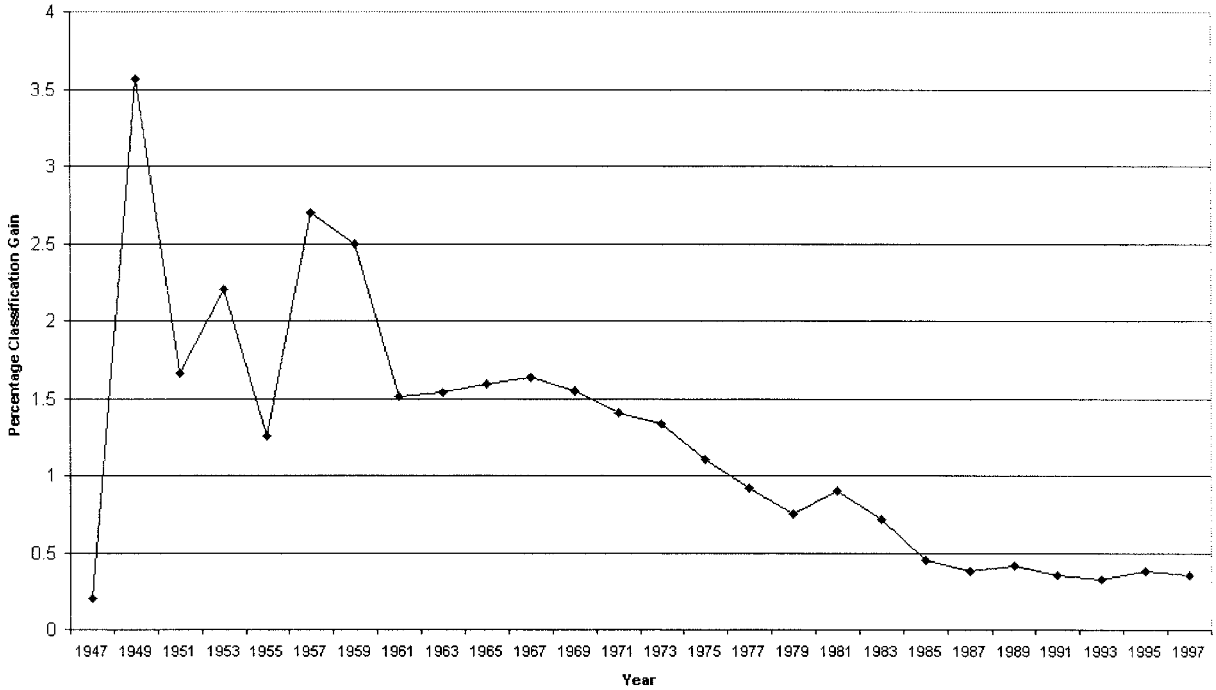
The average difference for the lopsided roll calls is

[15] An improvement of 0.5% may well be statistically significant. In column (k) of table A1 in McCarty, Poole, and Rosenthal 2000, which shows simulations for preference-based voting, classification gains are shown for various one-dimensional specifications. In the first three (low overlap) rows, similar to actual overlap in the past eight Congresses, the gains range from 0.10% to 0.26%, all considerably less than 0.5%. Of course, the gains from "fitting" an extra hyperplane in a multidimensional model would be expected to be even higher. In any event, an improvement of 0.5% may lack substantive import.

[16] A cutpoint is interior whenever at least one legislator is to the left of it and one legislator is to the right. Otherwise, the cutpoint is exterior.

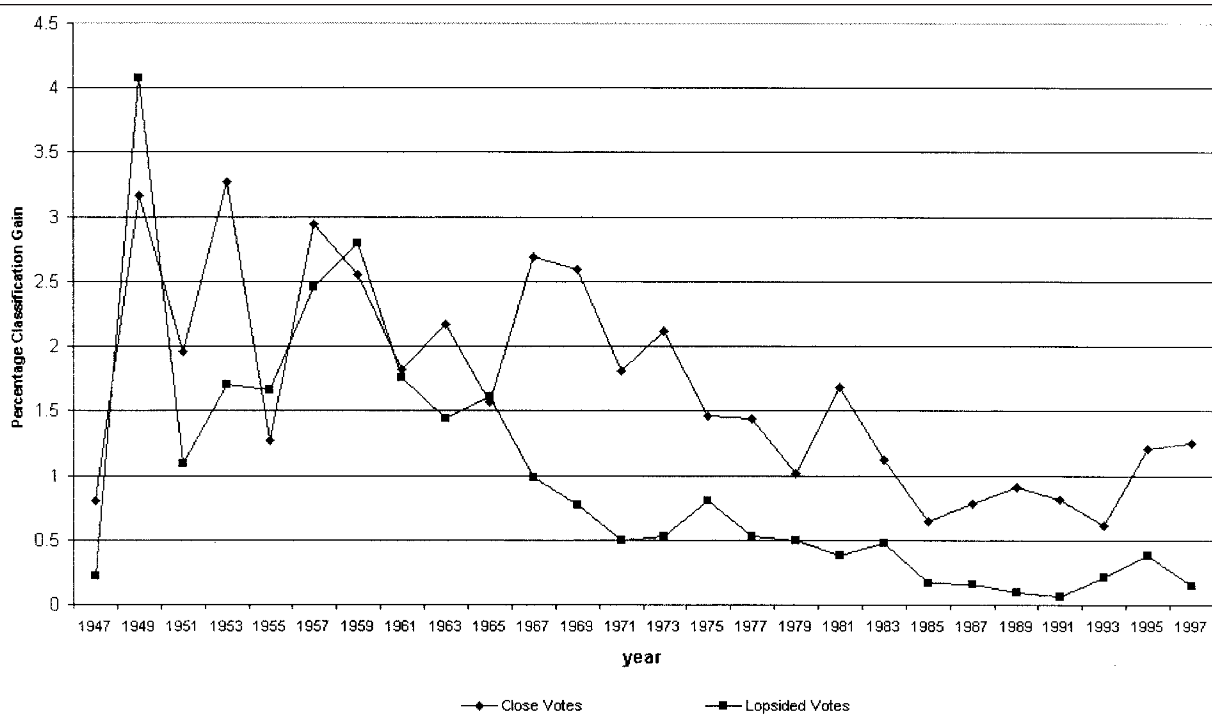
[17] In Appendix B, we outline our procedure for determining roll call cutpoints, classifying roll calls into the three categories, and computing the differences in ranks.

**FIGURE 6. Classification Gain of Two Cutpoints versus One Cutpoint for All Roll Call Votes**



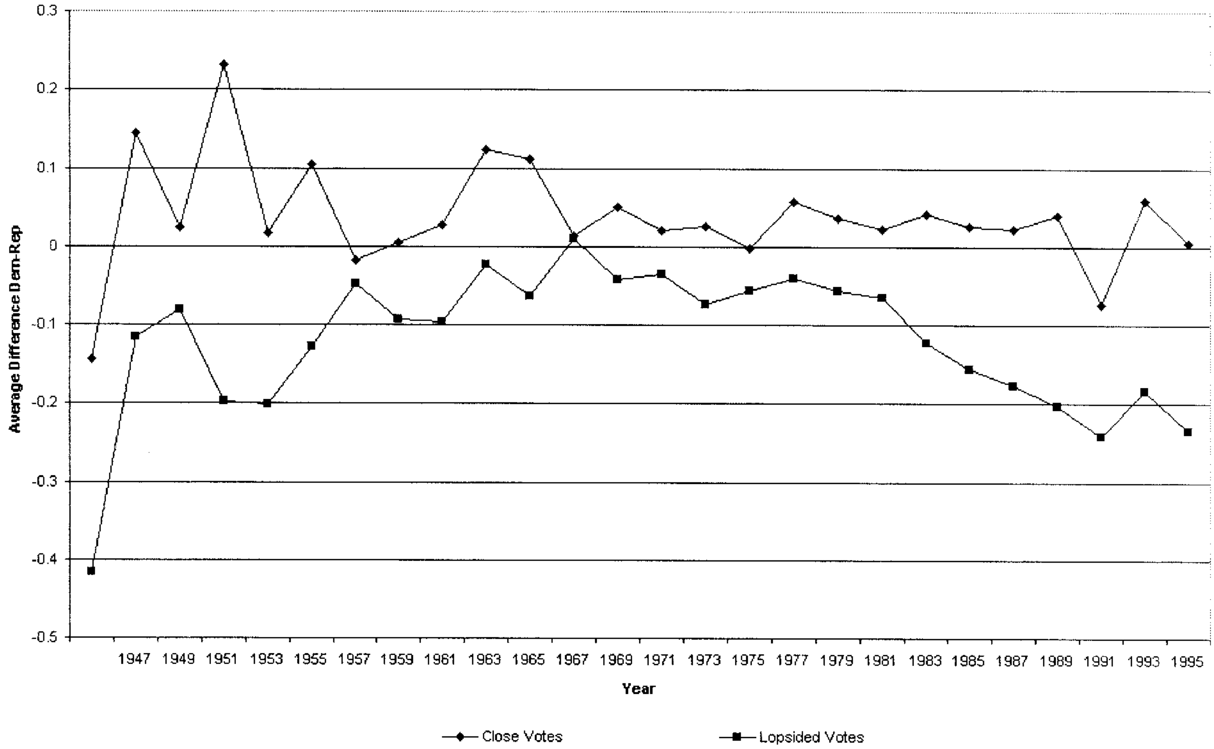

*Note*: The classification gains are for a one-dimensional voting model. All representatives were scaled together, as in the "common" scaling of Figure 4. With the ideal point orders from the "common" scaling held fixed, a separate cutpoint was then estimated for each party. Comparison to Figure 4 shows that the classification gains are similar to those in the "two-party" scalings where each party has an independent rank order of ideal points as well as a separate cutpoint.

**FIGURE 7. Classification Gain of Two Cutpoints versus One Cutpoint for Close and Lopsided Votes**




*Note*: The gains from the common scaling (see note to Figure 6) have been broken down into those for close votes and those for lopsided votes.

**FIGURE 8. Normalized Differences in Ranks between Party Cutpoints for Close and Lopsided Votes**




*Note*: For each House, the rank orders were normalized to run from 0 to 1. For example, if 438 legislators served in a House, the rank order was normalized to 0/437 . . . 437/437. The "average difference" is the average of the differences between the normalized rank of the Democrat cutpoint and the normalized rank of the Republican cutpoint.

negative for all 26 Houses. The negative sign probably reflects "both ends against the middle" voting. If the six most liberal Democrats and the six most conservative Republicans cast protest votes on final passage and these are the only negative votes, with fixed polarity, one of the party cutpoints will be near an end of the dimension, and the other party cutpoint will be near the middle. Therefore, the difference in ranks will be negative and large in magnitude. The negative differences can reflect a few conservative Republicans and a few liberal Democrats voting against a lopsided majority.

Hypothesis 5, which predicts that the Democratic cutpoint will be to the right of the Republican cutpoint, is not supported, as shown in Table 3. The pattern, except for the no or low overlap Congresses (80th and 103d–105th), is quite stable, so we present results in tabular form. Recall that low overlap means there will be very few or no roll calls with $D > R$. Even when there is overlap, the pattern runs counter to the Snyder-Groseclose model; $R > D$ roll calls outnumber the straight party $D > R$ roll calls by more than three to two.

**TABLE 3. Order of Cutpoints on Close Roll Calls**

| House | $D > R$ | Undecided | $R > D$ |
|---|---|---|---|
| 81st–102d | 41.7% | 2.7 | 55.6 |
| 80th, 103d–105th | 18.0% | 40.2 | 41.7 |
| All (80th–105th) | 38.1% | 8.5 | 53.4 |

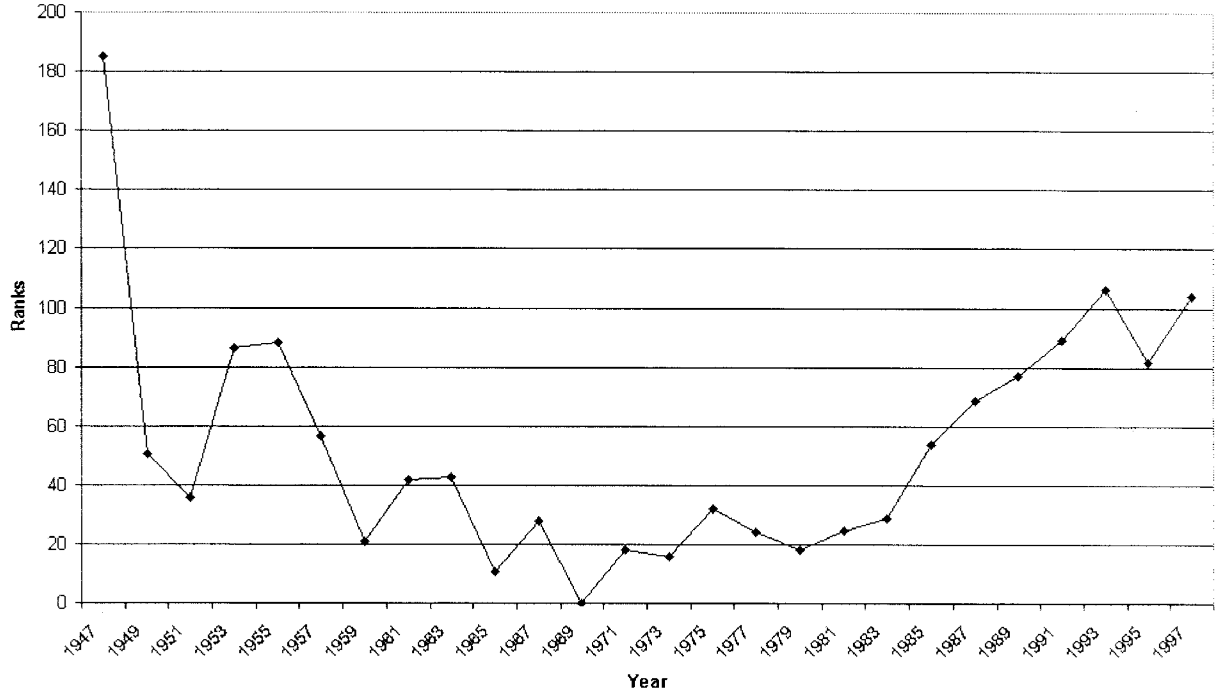
*Note*: Entries are the percentage of close roll votes that exhibit the indicated pattern, e.g., 53.4% of all close roll call votes in the 80th through 105th Congress had an $R > D$ pattern.

Table 3 lends much less support than Figure 8 to the party pressure model because of the fact that many of the Houses scores for a handful of southern Democrats fell in the midst of the Republican scores, and those of a handful of liberal Republicans fell in the range of Democratic scores. Therefore, on strict party or near strict party votes $D > R$, and the difference in ranks was quite large. The difference was smaller in magnitude on counterhypothesis $R > D$ votes, but such votes are typically a majority of the roll calls.[18]

Some of the counterhypothesis $R > D$ votes almost certainly indicate a true breakdown in party discipline. A breakdown can occur, for example, when the majority is subject to a few defections by its own moderates but offers bills or makes promises that buy the support of moderates in the opposite party. The seduction of minority party moderates is a scenario that seems to fit the two Gingrich Houses, in which, in the single-cutpoint analysis, the modal cutpoint fell interior to the Democratic Party (see McCarty, Poole, and Rosenthal

[18] In addition, the ordinal comparisons involve some roll calls with exterior cutpoints. See Appendix B.

**FIGURE 9. Ranks Shifted to Satisfy Average $D = R$ Condition on Lopsided Roll Calls**




1997, 12). (The two Gingrich Houses are the last two points in every plot.)

These results about cutpoints are, however, subject to the warning that the single-cutpoint estimation of ideal point ranks may show too much separation of the parties. We therefore calculated how far the ideal points of Republicans would have to shift leftward to reach the average difference for lopsided roll calls of zero. In that case, a new version of Figure 8 would have a flat line through zero for lopsided votes. That is, the shift would force the average pattern for lopsided votes to match the theoretical level in the Snyder-Groseclose model.[19]

The results of this exercise are shown in Figure 9. When the lopsided vote difference is just slightly negative, as in the late 1960s, very few ranks need to be shifted. In these cases the close vote difference is near zero, and $R > D$ roll calls outnumber $D > R$, so the Snyder-Groseclose model is not supported. When the lopsided vote difference is sharply negative (see Figure 8), in the late 1940s and in the 1990s, many ranks have to be shifted to force the lopsided votes to show a zero average. In the most recent Congresses in our time series, the order of change is of 100 ranks, or about half the Democratic membership. A shift of 100 ranks, which would place the "true" ideal points of the most moderate Republicans in the middle of the Democratic Party, is seriously lacking in face validity. The amount of overlap in the ideal point distribution is simply too great to make a party pressure model credible. In fact, the amount of shifting needed matches the decrease in party polarization in the postwar period and its increase since the late 1960s (McCarty, Poole and Rosenthal 1997) as measured by NOMINATE scores. The increasing separation of the parties is, in our view, much more likely to reflect fundamental political changes, such as a large increase in southern Republican representatives, than an increase in party discipline within Congress.

Because our initial ideal point distribution has greater face validity than the shifted distribution, we use the initial distribution to ask whether party discipline makes a difference in outcomes. We assess this in two ways. (1) We assume the true cutpoint is the minority cutpoint. Pressured voters are those majority party voters with ideal points between the minority and majority cutpoints. The benchmark is that all pressure is exerted by the majority party. Would the outcome have changed if their votes were reversed? (2) We assume that the true cutpoint is the average of the two party cutpoints, which reflects equal pressure exerted by both parties. Pressured voters are those with ideal points between their party cutpoint and the average. Would the outcome have changed if their votes were reversed?

The results vary substantially from one Congress to the next, which in part is a function of the separation of party ideal points. We find that, averaged across Con-

[19] We thank Tim Groseclose for suggesting the adjustment. The algorithm we developed to implement the suggestion is as follows. If the average difference in ranks for lopsided roll calls is nonnegative, then no shift is required. Otherwise, shift every Republican leftward by a number of ranks equal to the average difference in ranks. This procedure implicitly assumes that the ranks are interval measurements. By shifting the Republicans leftward, we are compressing the space. In the original estimates, the unnormalized space will extend from 1 to $N$, where $N$ is the number of scaled legislators in the House. In the shifted estimates, the space will run from 1 to $N - A$, where $A$ is the number of ranks shifted.

**TABLE 4. Rank Order Positions of Legislators in Their Old and New Party**

| Party | | | | | | Normalized Rank | |
|---|---|---|---|---|---|---|---|
| Old | New | State | Name | First Congress in New Party | Last Congress in Old Party | Old | New |
| Senate | | | | | | | |
| R | D | OR | Morse, W. | 82 | 83 | .127 | .004 |
| D | R | SC | Thurmond, S. | 88 | 89 | .988 | .824 |
| D | R | AL | Shelby, R. | 103 | 104 | .439 | .856 |
| D | R | CO | Campbell, B. | 104 | 104 | .355 | .603 |
| House | | | | | | | |
| D | R | SC | Watson, A. | 89 | 90 | .840 | .860 |
| R | D | NY | Reid, O. | 92 | 93 | .085 | .082 |
| D | R | OK | Jarman, J. | 93 | 94 | .517 | .796 |
| R | D | NY | Peyser, P. | 94 | 96 | .269 | .172 |
| D | R | PA | Atkinson, E. | 96 | 97 | .473 | .539 |
| D | R | AZ | Stump, B. | 97 | 98 | .898 | .991 |
| D | R | TX | Gramm, P. | 97 | 98 | .918 | .969 |
| D | R | FL | Ireland, A. | 98 | 99 | .495 | .884 |
| D | R | FL | Grant, B. | 100 | 101 | .442 | .641 |
| D | R | AR | Robinson, T. | 100 | 101 | .486 | .659 |
| D | R | LA | Hayes, J. | 104 | 104 | .480 | .890 |
| D | R | TX | Laughlin, G. | 104 | 104 | .455 | .881 |
| D | R | MS | Parker, M. | 104 | 104 | .500 | .876 |
| D | R | GA | Deal, N. | 104 | 104 | .474 | .925 |
| D | R | LA | Tauzin, B. | 104 | 104 | .506 | .872 |

*Note*: The orderings were normalized to 0–1 by dividing the raw ranks by 472 for the Senate and 2,326 for the House.

gresses, discipline makes a difference, under the first assumption, on 16.97% of close roll calls (std. dev. 9.17) and, for the second assumption, on 11.07% (std. dev. 8.03%). These numbers are substantial, but they are well below the proportion of significant *t*-statistics reported by Snyder and Groseclose. Moreover, they are almost certainly overestimates. One qualification is that the first assumption is extreme, since only the majority party exerts pressure. Another is that some of the pressured voters might not have changed their votes even if pressure were removed. This is because under the null hypothesis of a single cutpoint, errors in voting will result in some legislators who are on the "yea" side of the cutpoint who vote "nay," and vice versa. Similarly, under the alternative hypothesis of two cutpoints, there will be two types of legislators between the cutpoints—those who are pressured and those who vote with their party for idiosyncratic reasons. The first assumption mistakenly counts both types of legislators as pressured.

This section, in summary, has established that allowing for party discipline does not make an important contribution to classification; and those improvements in classification that do occur are, more frequently than not, the result of using cutpoints that are inconsistent with the party pressure model.

## IDEAL POINT CHANGES IN PARTY SWITCHERS

There is little evidence that many ideal points are displaced on individual votes, but there is very substantial evidence that party affiliation has a strong influence on ideal points. To see this, we again used the procedure of Poole (2000) to obtain rank orders of the ideal points in separate estimations for the House and Senate. This time we pooled all roll calls from 1947 to 1998.[20] Each member was constrained to have a constant ideological position in his or her career, except that party switchers were allowed to have two positions, one before and one after the switch. There were 472 senators and 2,326 representatives (counting the party switchers as two individuals). The orderings were normalized to 0–1 by dividing the raw ranks by 472 for the Senate and 2,326 for the House.

When legislators switch from Republican to Democrat, they should have a lower rank. The reverse should hold for Democrat to Republican switchers. There were 19 legislators who both switched parties and remained in the same chamber in the period of our analysis. They are listed in Table 4. In 18 of 19 cases the rank changed as expected. The exception is Strom Thurmond, whose slightly more moderate position as a Republican reflects his more moderate views on race relations in the past 20 years. A simple sign test is overwhelmingly significant. Induced ideal points respond to party affiliation.[21]

We have shown that party switchers generally move

[20] In an earlier version, we used the McCarty, Poole, and Rosenthal (1997) DW-NOMINATE procedure to obtain metric estimates of the magnitude of changes induced by party switching. The metric assumptions in the NOMINATE procedure lead to sensible results; for example, there is less distance between the median and the 9th decile in the Gingrich Houses than between the 1st decile and the median. We conducted the metric analysis in two dimensions. Switchers from *R* to *D* were expected to become more negative on the first dimension and more positive on the second, and vice versa for *D* to *R* switchers. All movement on both dimensions was in accord with the hypothesis. For more details, the earlier version can be accessed at http://voteview.uh.edu/d011000merged.pdf.

[21] These results are consistent with Nokken (2000), who also finds significant changes in congressional behavior following a party switch.

in the theoretically expected direction. Do they move very much? The average rank movement was 0.28, or a jump over more than one-fourth of all legislators serving in the period. To benchmark this movement, we reran the analysis for the House allowing two positions not only for the party switchers but also for some legislators who never changed party. Specifically, we picked in the legislator file every 500th legislator among moderates (i.e., those with ideal points between $-0.3$ and $+0.3$) who served in at least two Houses.[22] There were 15 such representatives, matching the number of actual switchers in the House. For each group of 15 we computed the average partisan switch. That is, for Democrats the switch was just the change in the coordinate, as Democratic switchers are expected to increase their rank. For Republicans, we used the negative of the change. Actual switchers moved substantially, a change in normalized rank of 0.281. On average, nonswitchers barely budged, moving only 0.026 in rank. The (one-tail) $t$-statistic for the differences in the means indicates a high level of statistical significance ($p < .0001$).

This evidence is consistent with a party effect, but a couple of caveats are in order. First, because it is silent on the mechanism that generates this effect, the source may not be internal to the legislature. Switchers, after all, have to adapt to a new set of primary constituents and contributors as well as legislative leaders. Second, party switchers are obviously not a random sample of all legislators. In the 104th House, southern Democrats switched to the Republican Party for a reason: They wanted to reflect the increasingly conservative temperament of their district.[23] Therefore, selection bias precludes us from suggesting that the shift in ideal points is an unbiased estimate of party pressure. But even if the selection bias were severe, it is telling that changing party labels was deemed necessary to reflect changing district sentiment.[24] Third, the estimates based on party switchers are almost certainly an upwardly biased measure of the average amount of party discipline. Among those who do not switch there is probably more congruence between their personal/constituency position and the party's desires.[25] In particular, representatives close to the party median are likely to vote "correctly" without any discipline.

[22] The representatives selected were Boland (D-MA), Johnson (D-CA), J. Melcher (D-MT), Button (R-NY), Fallon (D-MD), Traficant (D-OH), Matthews (D-FL), Morella (R-MD), Fountain (D-NC), Taft (R-OH), Lloyd (R-UT), Kasten (R-WI), Haley (D-FL), T. Corcoran (R-IL), and Zion (R-IN).

[23] Yet, studies show that constituency changes do not have much effect on the ideal points of legislators. See Poole and Romer 1993 for House redistricting and Doberman 1997 for House members who moved to the Senate.

[24] Levitt (1996) provides some indication of the relative effects of party versus constituency factors in determining the ideal points of switchers. For the Senate, Levitt models each senator's ideal point (as proxied by ADA rating) as a weighted average of personal ideology, overall state characteristics, support group characteristics, and the "national party line." Although all four of these factors may change for switchers, the main changes are likely to be in the new national party line and in the new support group that is relevant to campaign funding and primaries. Levitt's results put about equal weight on these two factors. Therefore, about half the change in the ideal point would reflect forces internal to Congress.

[25] We thank Larry Bartels for this observation.

## CONCLUSION

In the past decade, theorizing about the influence of parties and leaders on legislative behavior has outstripped progress in solving difficult methodological and measurement problems necessary to test these theories. We have addressed the problems associated with distinguishing party effects from a null hypothesis of individual preference-driven behavior. We began by demonstrating the unattractiveness of regression-based procedures, such as that of Snyder and Groseclose. We find that these methods of estimating the effects of party discipline on individual roll call votes are biased toward exaggerating the party influence. To remedy these statistical problems, we incorporated the theoretical insight of Snyder and Groseclose into the spatial model of voting, which we estimated nonparametrically. We found that empirically a party discipline approach makes, at best, a marginal improvement over the standard spatial model.

We do not conclude, however, that party is irrelevant. Voting behavior changes fairly dramatically when members change parties. Party discipline, we conclude, is manifest in the location of the legislator's ideal point in the standard spatial model. It is not a strategic variable manipulated by party whips from one roll call to another but a part of a legislator's overall environment that forms her induced preferences. The "discipline" that leads a legislator to choose a spatial location may result as much from external pressures of campaign donors and primary voters as from the internal pressures of the congressional party.

Thus, the evidence we present does not suggest that a resurgence of party or party-induced institutional change is responsible for the greater cohesiveness of parties and polarized politics in Congress. Distinct cutting lines (or separating hyperplanes) for Democrats and Republicans do not add substantially to the classification success of the spatial model in the period after World War II. Indeed, the incremental classification success of the second cutting line has fallen throughout this period, both during the period of declining polarization (1947 to about 1975) and during its more recent resurgence.[26]

## APPENDIX A: SHIFTS IN IDEAL POINTS

Let $z_{yj}$ and $z_{nj}$ be the "yea" and "nay" outcomes of roll call $j$. In both the Heckman-Snyder and NOMINATE methods for estimating the spatial model, the nonrandom portion of the utility a legislator $i$ has for roll call outcome $z_j \in \{z_{yj}, z_{nj}\}$, can be expressed as:

$$U_{ijz} = f(d_{ijz}^2), \qquad \textbf{(A-1)}$$

where $f$ is a negative monotonic function, and $d_{ijz}$ denotes the Euclidean distance from $x_i$, $i$'s ideal point, to $z_j$.

[26] On polarization in Congress, see Fiorina 1999; King 1998; Lowry and Shipan 2000; McCarty, Poole, and Rosenthal 1997; and Poole and Rosenthal 1984, 1997, 1999.

Now let the "party-pressured" ideological coordinates for Democrats equal $x_i + \gamma_j$. We obtain:

$$d_{ijz}^2 = (x_i + \gamma_j - z_j)^2. \quad \textbf{(A-2)}$$

But this expression for distance is identical to the expression we would have if the ideal point were unchanged but the "yea" roll call outcome were changed to $z_{yj} - \gamma_j$. The distance to $z_{nj}$ would also be unaffected if it were also changed to $z_{nj} - \gamma_j$. Shifting both roll call outcomes by $\gamma_j$ also shifts the midpoint $(z_{yj} + z_{nj})/2$ by $\gamma_j$. So, for example, a leftward shift in the ideal points for all Democrats is equivalent to a rightward shift in the outcome locations and midpoint for Democrats. The argument extends readily to multidimensional shifts. Since for every ideal point shift there is an equivalent outcome shift, neither Heckman-Snyder nor NOMINATE can discriminate between a model in which a party alters ideal points on a roll call and one in which each party has its own midpoint or separating hyperplane on each roll call.

Now consider the more general situation in which the amount of pressure is not equal for all members but the pressured ideal points maintain the same order as the original members, and the magnitude of the pressure, for Democrats, is increasing in spatial position. Moderates are pressured more than liberals. Since the pressure is not uniform, the shift in ideal points can no longer be captured by a simple shift in outcome locations. Nonetheless, in the map from the pressured ideal points back to the original ideal points, there will continue to be a point at which a party member is indifferent between voting "yea" and "nay." Let this point be the pressured midpoint for the party on the roll call. Optimal classification should be reasonably robust in identifying the pressured midpoint as long as the form of pressure does not depart too strongly from uniform pressure.

## APPENDIX B: PROCEDURE FOR COMPUTING THE DIFFERENCE IN CUTPOINT RANKS

We used the following procedure to determine roll call cutpoints, classify roll calls into the three categories, and compute the differences in ranks.

1. Optimally classify all legislators using a single cutpoint. Rank order the legislators from 1 to $N$, starting at the left.
2. Estimate the two-cutpoint model for roll calls using the rank order of legislators from step 1. (Note that the estimation must "maintain polarity": Classification is optimal subject to making the same prediction for Democrats and Republicans to the left of their party's cutpoint.)
3. Every interior Democratic cutpoint must be between two Democratic legislators. Let their ranks be $i$ and $j$. The rank of the roll call cutpoint is then given as $c_D = (i + j)/2$. When the cutpoint is to the right of the rightmost Democrat, denote the cutpoint by $c_D = d_R$ = rank of rightmost Democrat. When the cutpoint is to the left of the leftmost Democrat, denote the cutpoint by 1. The Republicans are treated similarly; when the cutpoint is to the left of the leftmost Republican, denote the cutpoint by $c_R = r_L$ = rank of leftmost Republican, and to the right of the rightmost Republican, denote the cutpoint by $N$.
4. Score the roll call as follows.
   a. If $c_D = 1$ and $c_R > r_L$ or if $1 < c_D < d_R$ and $c_R > r_L$ and $c_D < c_R$, score the roll call $D < R$.
   b. If $1 < c_D$ and $c_R < N$ and $c_D > c_R$, score the roll call $D > R$.
   c. Otherwise, the roll call is "undecided."
5. For roll calls with interior cutpoints in both parties, the difference in ranks is $c_D - c_R$. Roll calls with one or more party cutpoints exterior are excluded from the difference in ranks computations (Figure 8). (Thus, more roll calls are included in the ordinal comparisons under step 4.)

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