# Estimating Party Influence on Roll Call Voting: Regression Coefficients versus Classification Success

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*Contrary to the claims of McCarty, Poole, and Rosenthal, our method does not estimate the ideal points of moderates significantly less accurately than the ideal points of extremists. This is true for at least two reasons: (1) there is significant randomness in voting; as a consequence, on a lopsided vote moderates often vote with the extremists; and (2) our data set includes some roll calls that require a supermajority for passage; for these we define a 50%–50% roll call as lopsided. We also show that the classification-success method of McCarty, Poole, and Rosenthal severely understates the presence of party influence. Furthermore, we show that a proper interpretation of some of their results reveals a significant amount of party influence in Congress.*

In "The Hunt for Party Discipline in Congress," McCarty, Poole, and Rosenthal (2001) do three things. First, they critique the estimation strategy proposed in Snyder and Groseclose (2000), arguing that it overstates the actual amount of party influence on roll call voting. Second, they propose a different approach for estimating party influence, which compares the classification success achieved using a one-cutline model and a two-cutline model that allows parties to have distinct cutlines on each roll call. Third, they study congressional party switchers and show that they quickly change their roll call voting behavior, moving sharply away from the center of their old party and toward the center of their new party. This is interpreted as evidence that "party discipline . . . is manifest in the location of the legislator's ideal point" (McCarty, Poole, and Rosenthal 2001, 686).

In this response, we do four things. First, we counter the main criticism of our estimation strategy, showing that it does not measure poorly the ideal points of moderates. Second, we discuss an independent measure of ideal points, which is not derived from roll calls, that produces results remarkably similar to those obtained when we estimate ideal points using lopsided roll calls. Third, we show that the basic conclusions in our earlier article are not driven by functional form. Specifically, we obtain similar results whether we use a linear model (as in our earlier article) or Poole and Rosenthal's W-NOMINATE model. Finally, we show that the method proposed by McCarty, Poole, and Rosenthal is severely biased toward not finding any party influence. It will often fail to detect such influence even when it is quite strong. In fact, some simple calculations suggest that there is no contradiction between the findings reported by McCarty, Poole, and Rosenthal and our earlier article.

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## LOPSIDED ROLL CALLS PROVIDE AN ACCURATE MEASURE OF IDEAL POINTS

We will first respond to McCarty, Poole, and Rosenthal's key example, which describes conditions that cause our method to yield poor ideal point estimates for moderate legislators. We will show that these conditions are not empirically relevant. Next, we will present independent evidence that supports our use of lopsided roll calls to estimate ideal points. We then will show that some of the puzzling findings reported by McCarty, Poole, and Rosenthal are actually consistent with the existence of party influence.

### Congressional Roll Call Voting Is Probabilistic

McCarty, Poole, and Rosenthal (2001, Figure 1) present a simple example to show how the lopsided roll call estimation strategy we propose cannot distinguish among moderate legislators.[1] Because their example assumes errorless voting, there cannot be any roll calls on which, say, legislator 3 votes on the left side of the issue but legislator 1 votes on the right side. In practice, however, such cases occur frequently. To demonstrate, we gathered data for the 90th–105th Congresses (1967–98) and constructed a list of moderate House members based on Heckman and Snyder's (1997) linear factor model.[2] Specifically, we rank representatives by their first-dimension scores and call them "moderate" if they fall between the 45th and 55th percentiles.[3]

We then examined all roll calls on which 68–72% of

[1] See Snyder 1992 for an earlier study of a closely related phenomenon.

[2] Using Poole and Rosenthal's W-NOMINATE scores to rank representatives produces similar results.

[3] If McCarty, Poole, and Rosenthal are correct, then ideal point estimates should be based on all roll calls (not just lopsided roll calls), so the ranks shown on the horizontal axis are based on all roll calls. Of course, if McCarty, Poole, and Rosenthal are correct and voting is errorless, then scores from metric scaling techniques—including the linear model and NOMINATE—are inaccurate, since these methods are based on probabilistic voting models. We would prefer to use the ranks produced by McCarty, Poole, and Rosenthal's one-cutline method, but do not have them. In any case, we strongly doubt it would affect the results, given the high correlations typically found among various first-dimension scalings.

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the House voted on the winning side. On the one hand, if the House fits McCarty, Poole, and Rosenthal's example—a one-dimensional spatial model with errorless voting—then a moderate member will never vote on the losing side. On the other hand, if the House fits a world in which all legislators vote completely randomly, then each moderate will vote on the losing side about 30% of the time. Our findings show that the actual world of the House is halfway between these two hypothetical worlds. In the Congresses sampled, moderates voted on the losing side about 16% of the time.[4] Thus, unlike the McCarty, Poole, and Rosenthal example, moderates frequently vote on the losing side of an issue (about half as often as extremists do). Therefore, our strategy of estimating ideal points only from lopsided votes can make distinctions among moderate legislators.

Next, if the McCarty, Poole, and Rosenthal example is empirically accurate, and congressional roll call voting is not probabilistic, then using only lopsided roll calls will produce the same ideal point estimate for all legislators with an ideal point between the 35th and 65th percentiles. Thus, if we order legislators according to their ideal point rank on the horizontal axis and plot the estimated ideal points based on lopsided roll calls on the vertical axis, the resulting graph will appear as in Figure 1A.[5] The actual pattern is quite different, however. For instance, in Figure 1B we plot for the 103d House (1993–94) the rank of representatives' ideal points, on the horizontal axis, and the ideal point estimates based on lopsided roll calls, on the vertical axis.[6] The relationship is nothing like the inverted S-shape that appears in Figure 1A. Most important, the slope of the figure in the moderate range of ideal points is far from horizontal. Thus, it is very different from the prediction of the McCarty, Poole, and Rosenthal example. In fact, the slope is much more similar to the slope that occurs in the two extremist ranges of the figure than it is to being horizontal. This is exactly the result that one should expect if our method estimates moderate and extremist ideal points with approximate equal accuracy. The patterns are similar in other Congresses.[7]

These results show that congressional roll call voting is best described by a probabilistic rather than an errorless voting model. They also show that when voting is probabilistic, moderates do not all vote alike, even when the cutline is far from the center of distribution. In fact, the law of large numbers implies that if voting is probabilistic, and if there are "enough" roll calls, then all legislators' ideal points can be accurately estimated, regardless of the distribution of cutlines. How many is enough? In our earlier article we ran a number of simulations and found that 200 or even 100 lopsided roll calls is usually sufficient to obtain accurate estimates of both moderates and extremists (Snyder and Groseclose 2000, Appendix A).[8]

In fact, in previous work, Poole and Rosenthal argue forcefully that congressional roll call voting looks probabilistic. They are also highly critical of empirical applications of deterministic voting models. For example: "Perfect (errorless) spatial voting is not the appropriate benchmark with which to compare scaling results. . . . In a nutshell, while roll call voting is nicely captured by a spatial model, there is significant nonspatial, random-looking behavior" (Rosenthal 1992, 31–2). Similarly: "The null model of errorless voting in seven or fewer dimensions is not even remotely supported by the data" (Poole and Rosenthal 1991, 956). More recently: "Even with two dimensions, however, about 15 percent of the individual votes fail to fit a simple structure. . . . A probabilistic model accounts for this structure" (Poole and Rosenthal 1997, 20). We agree.

Finally, the McCarty, Poole, and Rosenthal example ignores another factor that allows our method to distinguish among moderates on lopsided roll calls. This is that some roll calls in our sample, such as a suspension of the rules of the House and veto overrides, require a supermajority. On these we define a lopsided roll call as one where the winning side received less than 51.7% or more than 81.7% of the vote. Thus, for these roll calls a 50%–50% division qualifies as lopsided. Consequently, in most Congresses there are actually several lopsided roll calls that have cutlines within the group of moderate ideal points.

## An Independent Preference Measure: The NPAT Survey

Ansolabehere, Snyder, and Stewart (n.d.) use survey data from the National Political Awareness Test (NPAT) to measure the ideal points of House members and test for the independent influence of party on roll call voting. They study the 103d–105th Congresses and find a large amount of party influence. More important, they find that the results obtained using NPAT-based estimates are very similar to the results obtained using lopsided roll calls.[9]

[4] The percentages range from 8.4% in the 100th Congress to 22.6% in the 95th Congress. On average, there were 58 roll calls with 68–72% of the House on the majority side.

[5] This assumes the number of roll calls with cutlines to the left of the 35th percentile is approximately equal to the number with cutlines to the right of the 65th percentile.

[6] Again, the ranks are based on the Heckman and Snyder (1997) linear factor scores computed using all roll calls.

[7] Another implication of the McCarty, Poole, and Rosenthal example is that our method should (incorrectly) find similar amounts of party influence across all first-dimension issues. In fact, in our earlier article we found that party influence is frequent on items such as budget bills, tax bills, and rules on bills, but almost nonexistent on issues such as abortion, school prayer, and gun control. (Examination of classification success and cutline angles shows that these are all basically first-dimension issues.) This heterogeneity strongly suggests that forces other than preferences influence representatives' roll call voting on at least some issues, and party is probably one of these forces.

[8] As Snyder (1992) points out, one problem with interest group ratings is that they are usually based on only 10–25 roll calls. With so few, the distribution of scores may be quite sensitive to the distribution of cutlines.

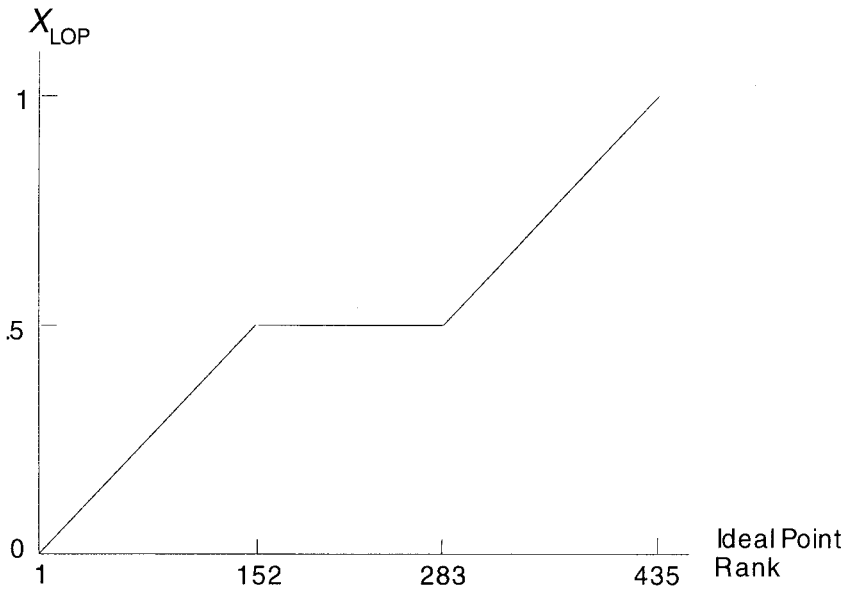
[9] The NPAT survey is administered by Project Vote Smart, which can be found at http://www.vote-smart.org. See Erikson and Wright (1997) and Ansolabehere, Snyder, and Stewart (n.d.) for details about the survey.



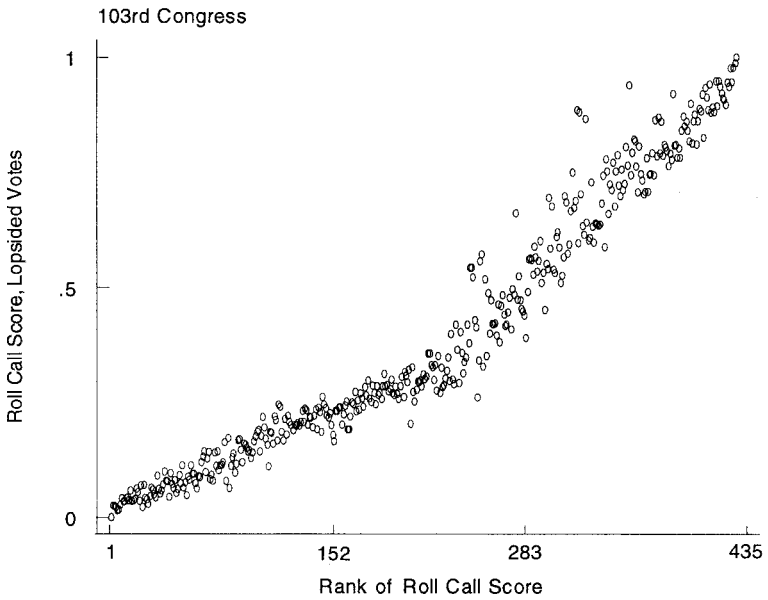
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**FIGURE 1. McCarty, Poole, and Rosenthal's Example**

A: How Ideal Points and Ranks Should Appear if the Example Is Accurate




B: How Ideal Points and Ranks Actually Appear

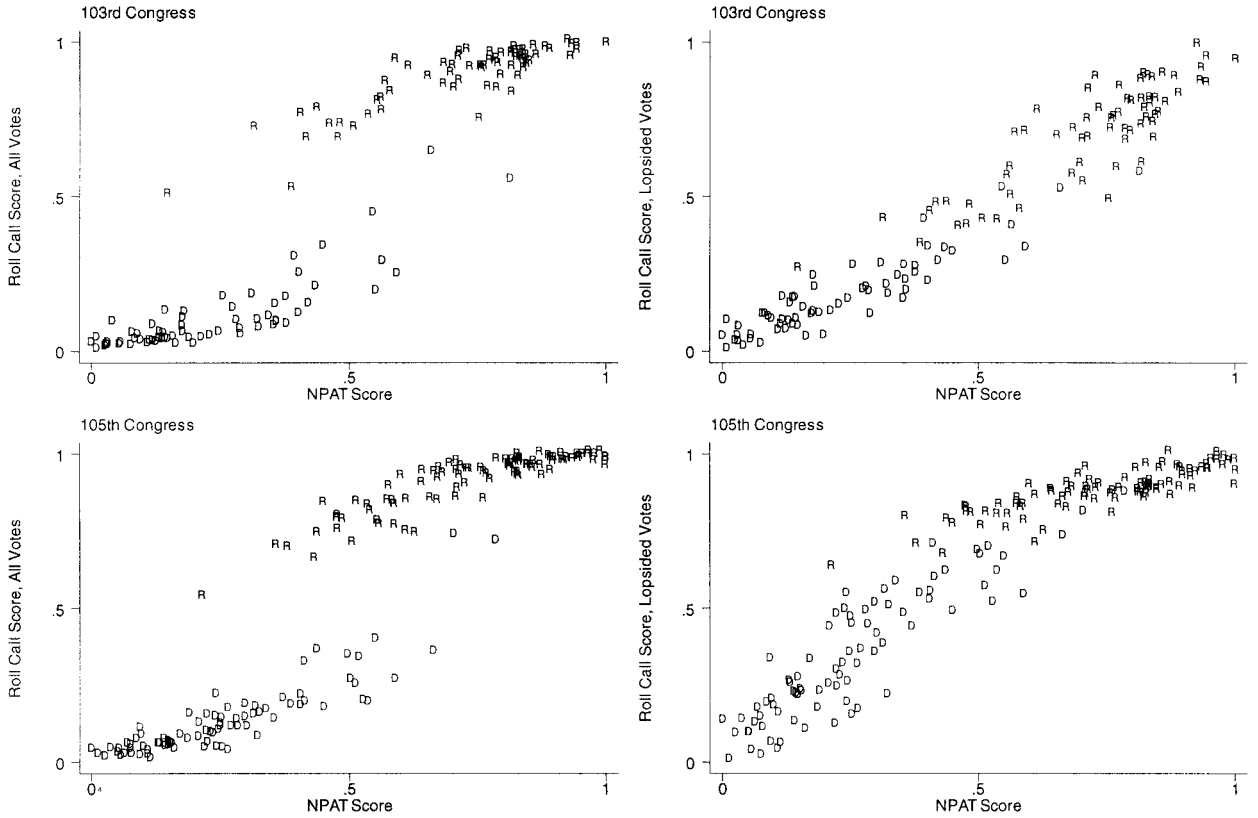



Here, we present some other comparisons between the ideal point estimates based on NPAT and lopsided roll calls. As in our earlier article, we use Heckman and Snyder's (1997) linear factor model to estimate representatives' ideal points. We applied this model to five different sets of data: the 1996 and 1998 NPAT surveys and the roll call voting records of the 103d, 104th, and 105th Congress. We also conducted a parallel analysis using Poole and Rosenthal's W-NOMINATE model in place of the linear model.

We focus on the first dimension. Let $X_{ALL}$ be the ideal point estimates based on all roll calls, let $X_{LOP}$ be the estimates based on lopsided roll calls and let $X_{NPAT}$ be the estimates based on the NPAT responses. We

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**FIGURE 2. Ideal Point Estimates using NPAT, All Roll Calls, and Lopsided Roll Calls**




match the ideal points derived from the 1996 NPAT with the roll calls for the 103d and 104th Congress, and those derived from the 1998 NPAT with the roll calls for the 105th Congress.[10]

Figure 2 shows plots of $X_{ALL}$ versus $X_{NPAT}$ and $X_{LOP}$ versus $X_{NPAT}$ for the 103d and 105th Congress (the plots for the 104th Congress are similar). The relationship between $X_{NPAT}$ and $X_{ALL}$ is strongly positive, but it exhibits a large "jump," which is clearly related to party affiliation. In contrast, the relationship between $X_{LOP}$ and $X_{NPAT}$ is not only strong and positive but also quite smooth—nearly linear, in fact. Clearly, if $X_{NPAT}$ represents an accurate estimate of members' first-dimension ideal points, then $X_{LOP}$ does as well.

Table 1 presents correlations among the variables $X_{ALL}$, $X_{LOP}$, and $X_{NPAT}$. The correlations are high, which is not surprising in view of Figure 2. More important, in each Congress the correlation between $X_{NPAT}$ and $X_{LOP}$ is noticeably higher than the correlation between $X_{NPAT}$ and $X_{ALL}$. This is true even for "moderate" legislators.[11] It is also true whether the estimates of $X_{ALL}$, $X_{LOP}$, and $X_{NPAT}$ are obtained using the linear factor model or W-NOMINATE (compare the top and bottom halves of the table).

The conclusion is simple: The ideal point estimates obtained using lopsided roll calls are very similar to those obtained using the NPAT survey. In fact, the estimates derived from the NPAT survey are more similar to those based on lopsided roll calls than they

**TABLE 1. Comparison of Ideal Point Estimates, Linear Model and W-NOMINATE**

| | Corr. between $X_{NPAT}$ and $X_{ALL}$ | Corr. between $X_{NPAT}$ and $X_{LOP}$ |
|---|---|---|
| Linear Model, All Reps. | | |
| 103d Congress ($N$ = 132) | .92 | .96 |
| 104th Congress ($N$ = 170) | .92 | .94 |
| 105th Congress ($N$ = 176) | .92 | .94 |
| Linear Model, Moderates | | |
| 103d Congress ($N$ = 65) | .85 | .90 |
| 104th Congress ($N$ = 84) | .82 | .85 |
| 105th Congress ($N$ = 87) | .84 | .86 |
| W-NOMINATE, All Reps. | | |
| 103d Congress ($N$ = 132) | .91 | .95 |
| 104th Congress ($N$ = 170) | .90 | .96 |
| 105th Congress ($N$ = 176) | .90 | .95 |
| W-NOMINATE, Moderates | | |
| 103d Congress ($N$ = 65) | .80 | .92 |
| 104th Congress ($N$ = 82) | .81 | .90 |
| 105th Congress ($N$ = 86) | .80 | .87 |

[10] The ideal points estimated using NPAT are quite stable over time. For the 135 representatives who completed both the 1996 and 1998 NPAT, the correlation between the two scores is more than .97.

[11] Moderates are those whose $X_{NPAT}$ score lies in the middle 50% obtained by eliminating the upper and lower quartiles.



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**TABLE 2. Comparison of Party Influence Estimates, Linear Model and W-NOMINATE**

| | Dependent Variable = $X_{ALL}$ | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | 103d Congress | | | 104th Congress | | | 105th Congress | | |
| Linear Model | | | | | | | | | |
| $X_{NPAT}$ | .41 (.02) | — | — | .38 (.02) | — | — | .42 (.02) | — | — |
| $X_{LOP}$ | — | .51 (.02) | .47 (.01) | — | .45 (.02) | .46 (.01) | — | .47 (.02) | .44 (.01) |
| Republican | 1.24 (.04) | 1.03 (.04) | 1.13 (.03) | 1.33 (.03) | 1.16 (.04) | 1.16 (.02) | 1.23 (.04) | 1.10 (.05) | 1.17 (.03) |
| Constant | .62 (.03) | .51 (.02) | .67 (.01) | .54 (.02) | .47 (.02) | .54 (.01) | .57 (.02) | .51 (.02) | .56 (.01) |
| $N$ | 132 | 132 | 427 | 170 | 170 | 426 | 176 | 176 | 425 |
| $R^2$ | .98 | .98 | .98 | .98 | .99 | .99 | .98 | .98 | .98 |
| W-NOMINATE | | | | | | | | | |
| $X_{NPAT}$ | .48 (.04) | — | — | .41 (.03) | — | — | .46 (.04) | — | — |
| $X_{LOP}$ | — | .63 (.04) | .66 (.02) | — | .55 (.04) | .54 (.02) | — | .67 (.04) | .66 (.03) |
| Republican | 1.06 (.08) | .75 (.08) | .70 (.05) | 1.22 (.07) | .91 (.08) | .93 (.05) | 1.10 (.07) | .70 (.07) | .68 (.05) |
| Constant | .52 (.05) | .37 (.04) | .41 (.03) | .50 (.03) | .37 (.04) | .43 (.02) | .52 (.04) | .31 (.04) | .32 (.03) |
| $N$ | 132 | 132 | 427 | 170 | 170 | 426 | 176 | 176 | 425 |
| $R^2$ | .92 | .95 | .94 | .93 | .94 | .94 | .92 | .95 | .94 |

are to those based on close votes. Moreover, this is true for moderate representatives as well as extremists.[12]

These results are strongly contrary to what we would expect if there were no party influence on any roll calls. In a world with no party influence, we would expect $X_{ALL}$ to be a better estimate of members' true ideal points than $X_{LOP}$, if only because it is based on a much larger sample. So, assuming that the NPAT responses yield a good measure of ideal points, the results in Table 2 and Figure 2 strongly suggest the presence of some party influence (or some other influence that is correlated with party). Something is "distorting" members' choices on roll calls that are not lopsided.

Finally, it is important to note that the distribution of implied "cutlines" on the NPAT questions is relatively uniform. Thus, it is not true that there is a high correlation between $X_{NPAT}$ and $X_{LOP}$ simply because NPAT has few questions with cutlines near the ideal points of moderates. NPAT includes many items on which respondents split nearly evenly in their answers. For example, in 1996 nearly one-third of the items (88 out of 268) were "close" questions, to which 35–65% of the representatives answered "yes."[13]

## MANY FINDINGS OF McCARTY, POOLE, AND ROSENTHAL ARE CONSISTENT WITH PARTY INFLUENCE

McCarty, Poole, and Rosenthal claim that their Figure 8 implies very little party influence on close votes. For instance, they note that since the 91st House (1969–70) the average difference between Republican and Democratic cutlines on close roll calls has been about .02, or a distance of only about 8 or 9 out of 435 ideal points. But their results for lopsided roll calls tell a very different story: The average distance between the Democratic and Republican cutlines is −.15. This translates into a distance of about 65 ideal points of House members. Most interesting is the sign of this measure, which indicates negative party influence. That is, their findings for lopsided roll calls are consistent with, for instance, a Democratic leader who pressures the rank-and-file Democrats to vote for the conservative alternative.[14] Also interesting is the fact that the party influence that McCarty, Poole, and Rosenthal find on lopsided roll calls is many times the effect they find on close roll calls.

These findings are puzzling. In two important re-

[12] It is possible that many respondents tailored their NPAT responses in order to make themselves appear more moderate than they really are. We find it remarkable, however, that the first-dimension scores recovered from the NPAT responses are so similar to those recovered from the lopsided roll calls. It seems that a more likely explanation for the close correspondence is that both are fairly good measures of representatives' actual (revealed) ideal points.

[13] This ignores extremely lopsided questions to which more than 95% answered "yes" or "no."

[14] Alternatively, as McCarty, Poole, and Rosenthal note, this result could occur if Democratic leaders pressure the Republican rank-and-file to vote in the liberal direction. Yet, for the effect to cause a negative sign, the Democratic leaders must pressure Republicans more strongly than they pressure Democrats. Furthermore, no matter what the reason for the negative sign, it indicates that the rank-and-file is pressured to vote against the direction its party prefers.

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spects they contradict the way it seems a rational party leader would act. First, they suggest that party leaders concentrate their resources not on close roll calls but on lopsided roll calls. Second, they suggest that on lopsided roll calls leaders pressure legislators to vote against their party, not with it.

The most likely explanation for these puzzling findings is not that party leaders are acting irrationally but that McCarty, Poole, and Rosenthal use an inappropriate baseline in their analysis. They note that a two-cutline method cannot identify the overlap in ideal points of the two parties. Consequently, to measure this overlap, McCarty, Poole, and Rosenthal use a one-cutline model in the first stage of their method. This forms the baseline used in the second stage to measure party influence. They implicitly assume that the baseline reflects no party influence, but they use all roll call votes to form this baseline. If parties exert some influence on some of these roll calls, then the baseline reflects ideal points that are shifted from their true values by an amount that corresponds to the average degree of party influence. Thus, if party influence exists, then the baseline is tainted with that influence and cannot be used to measure it. Using this baseline to assess party influence on a subset of roll calls will understate the true degree of influence when it exists, and there will be a spurious finding of negative influence when it does not exist.

We believe this is exactly what happens in McCarty, Poole, and Rosenthal's Figure 8. Since the true amount of party influence on lopsided roll calls is zero, or at least below the average (baseline) amount, the method spuriously shows negative influence. Next, although the method shows positive influence on close roll calls, it shows not the true degree but the difference between the true degree and the average (baseline) degree.

It is likely that McCarty, Poole, and Rosenthal's Figure 9 does reflect a better baseline and, therefore, approximately captures the true extent of party influence in Congress. In that figure McCarty, Poole, and Rosenthal have shifted ideal points so that on lopsided roll calls the two parties have identical cutlines. This effectively sets the baseline as the ideal points that one would estimate from a sample of lopsided roll calls. Figure 9, like our results, indicates a large amount of party influence in Congress. For the average Congress it indicates that party influence shifts ideal points by about 50 ranks. If no overlap exists between the parties, then this is about half the effect of pulling the median member of the House to the ideal point of the median of the majority caucus. If there is significant overlap, and the parties are distributed fairly uniformly within the overlap region, then 50 ranks translates into about one-quarter the effect of pulling the median of the House to the median of the majority party.

McCarty, Poole, and Rosenthal (2001) note that in the most recent Congresses "the order of change is of 100 ranks, or about half the Democratic membership" (p. 684). They note that this places "the most moderate Republicans in the middle of the Democratic Party" (p. 684), which "is seriously lacking in face validity" (p. 684). Four points should be noted in response. First, this is the maximum effect they find among modern Congresses, and it occurs only recently, in Congresses noted for their partisanship. Second, a shift of 100 ranks does not necessarily imply that the majority party is pulling its members 100 ranks. It could arise because, say, the majority party pulls its members 60 ranks in one direction, and the minority party pulls its members 40 ranks in the other direction. Third, a shift of 100 ranks, even if only one party is exerting pressure, is not overly large. For instance, if there is enough overlap, it is about two-thirds the effect of a "strong" party, that is, a party that can make its most moderate member vote like the median of the party.[15] Fourth, the effect is smaller than McCarty, Poole, and Rosenthal's own estimate of the effect of a switch in party. Specifically, when a switch occurs, the switcher's ideal point shifts on average 28% of the total ranks, which in the House translates to 122 legislators.

## THE LINEAR FUNCTIONAL FORM DOES NOT DRIVE OUR RESULTS: A COMPARISON WITH W-NOMINATE

We now show that the basic conclusions in our earlier article are not simply the result of the specific functional form used in that analysis. We do this by comparing the linear model with Poole and Rosenthal's W-NOMINATE model. We also complete the analysis of the previous section by showing that the estimates of party influence are the same whether we use the lopsided roll calls or the NPAT responses to estimate ideal points.

Consider estimates of the following equations:

$$X_{ALL} = b_0 + b_1 X_{LOP} + c\ Republican + \epsilon; \text{ and}$$

$$X_{ALL} = b_0 + b_1 X_{NPAT} + c\ Republican + \epsilon.$$

Table 2 presented the estimates of these equations for the 103d–105th Congress. In the top half of the table, $X_{ALL}$, $X_{LOP}$, and $X_{NPAT}$ are all estimated using the linear factor model. In the bottom half, W-NOMINATE is used. To facilitate comparisons, we standardize $X_{ALL}$, $X_{LOP}$, and $X_{NPAT}$ so that each has a mean of zero and a standard deviation of one in each Congress. Thus, the estimates of $b_1$ represent the expected change in $X_{ALL}$, measured in standard deviations, produced by a change of one standard deviation in $X_{NPAT}$ or $X_{LOP}$. We do not standardize the *Republican* dummy variable, so the estimates of $c$ represent the expected change in $X_{ALL}$, measured in standard deviations, of switching parties.

[15] To see this, consider the following example. Suppose there are 218 Republicans and 217 Democrats. Suppose that the 167 most liberal members are Democrats. For the rest, Republicans and Democrats alternate in their ranks. That is, the 168th most liberal member is a Republican, the 169th most liberal is a Democrat, and so on. This means that the 267th most liberal member is a Democrat, and all remaining members are Republicans. The median of the Republicans is the 326th most liberal member. To achieve "strong" party influence, the Republican leaders must make the 168th member vote like the 326th member, a shift of 158 ranks. A shift of 100 ranks is a little less than two-thirds this effect.



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The findings are straightforward. First, the estimates of party influence are similar whether we use $X_{LOP}$ or $X_{NPAT}$ to control for preferences. If anything, party influence appears even more important when $X_{NPAT}$ is used. This is true for both the linear model and the W-NOMINATE model. Second, the estimated effect of party influence is substantively large and highly statistically significant regardless of whether we use the linear or W-NOMINATE functional forms. The W-NOMINATE model produces smaller estimates of party influence, but the qualitative conclusions are the same. Finally, the third, sixth, and ninth columns in Table 2 show that the estimates for the restricted sample of representatives with NPAT scores are similar to the estimates for the full sample of all representatives with scores based on roll calls.[16]

## A FOCUS ON CLASSIFICATION SUCCESS UNDERSTATES PARTY INFLUENCE

We noted above a difficulty with the baseline used by McCarty, Poole, and Rosenthal to measure party influence. Even if this problem is ignored, however, their method is flawed because it will frequently miss party influence, even when it is present. A simple example reveals the problem. The example shows that even if the majority party exerts a large amount of influence—enough to move policy from the chamber median to the party median, sometimes called a "strong" party influence—adding a second cutline typically will lead to only a tiny gain in classification success. In fact, in typical cases the gain will be similar to or even less than the 1–2% gain reported by McCarty, Poole, and Rosenthal. In many cases, adding a second cutline will not improve classification success at all, despite the presence of substantial party influence.

Consider voting over one "canonical" bill, which pits the majority party median against the floor median. We will calculate the classification success using a one-cutline model. Classification success using a two-cutline model (one for each party) will be 100%.[17]

There are two parties, $L$ and $R$. Each has a symmetric, triangular distribution of ideal points, and for simplicity we assume the triangles are similar.[18] Without further loss of generality, fix the support of the $R$ distribution at [0, 1]. Let the support of the $L$ distribution be $[l_0, l_1]$, with $l_0 < 0$. Let $N_R$ be the measure of the set of legislators in $R$, let $N_L$ be the measure of the set of legislators in $L$, and let $\psi = N_L/N_R < 1$ (so $R$ is the majority party). Normalize by setting $N_R = 1$. Thus, the density of the $R$ distribution, $f_R$, is given by $f_R(z) = 4z$ for $z \leq 1/2$ and by $f_R(z) = 4(1- z)$ for $z > 1/2$. Denote the cumulative distribution functions of the ideal point distributions by $F_L$ and $F_R$; so, $F_L(l_0) = 0$, $F_L(l_1) = \psi$, $F_R(0) = 0$, and $F_R(1) = 1$. Denote the "overlap" between the parties by $\phi$. For convenience we assume that the median of party $L$ is to the left of all members of party $R$ and that the median of party $R$ is to the right of all members of party $L$. That is, we assume $(l_0 + l_1)/2 \leq 0$ and $l_1 \leq 1/2$. When $l_1 \leq 0$, $\phi = 0$. When $l_1 > 0$, $\phi$ is equal to the fraction of legislators whose ideal points lie in $[0, l_1]$, that is, $\phi = [F_L(l_1) - F_L(0) + F_R(l_1) - F_R(0)]/(1 + \psi)$.[19] Denote the chamber median by $\mu$. The status quo is at $\mu$. The bill is at the median of the $R$ distribution, 1/2; so the cutline for legislators who are free from party influence is $c = (2\mu + 1)/4$.

One other point is crucial, which we denote $c_R$. This is the point at which, if all members of $R$ with ideal points in $[c_R, c]$ are persuaded to vote for the legislation, the bill will exactly pass. So $[1 - F_R(c_R)] + [\psi - F_L(c)] = F_R(c_R) + F_L(c) = (1 + \psi)/2$. Note that $F_L(c)$ might be equal to $\psi$, in which case only members of $R$ vote for the bill, and $c_R = F_R^{-1}((1 - \psi)/2)$.

We assume that $R$ only exerts influence on $R$-affiliated legislators with ideal points in $[c_R, c]$ and persuades all of them to vote for the bill. It passes, and policy is moved from the chamber median to the $R$ caucus median, a large shift. Denote the fraction of $R$ affiliates who "switch" their vote as a result of pressure by $s_R = F_R(c) - F_R(c_R)$.

Let $c_1$ be the optimal cutline using a one-cutline model. That is, taking the distributions of ideal points as known, $c_1$ is chosen to minimize the fraction of votes that are misclassified. Let $M$ be the fraction of misclassified votes using a one-cutline model. Clearly, there is no misclassification using a two-cutline model: Setting the $L$ cutline at $c$ and the $R$ cutline at $c_R$ produces a perfect classification. To calculate $c_1$ and $M$, we must consider five cases. A few straightforward calculations yield the following:

1. $\mu \geq l_1$. Then $c_1 = c_R$, and $M = 0$.
2. $\mu < l_1, l_1 < c$, and $l_1/2 \leq c_R$. Then $c_1 = c_R$, and $M = 2(l_1 - c_R)^2/(1 + \psi)$.
3. $\mu < l_1, l_1 < c$, and $l_1/2 > c_R$. Then $c_1 = l_1/2$, and $M = (l_1^2 - 2c_R^2)^2/(1 + \psi)$.
4. $\mu < l_1, l_1 > c$, and $l_1/2 \leq c_R$. Then $c_1 = c_R$, and $M = 2[(l_1 - c_R)^2 - (l_1 - c)^2]/(1 + \psi)$.

[16] The full sample does not contain exactly 435 observations because of intrasession turnover.

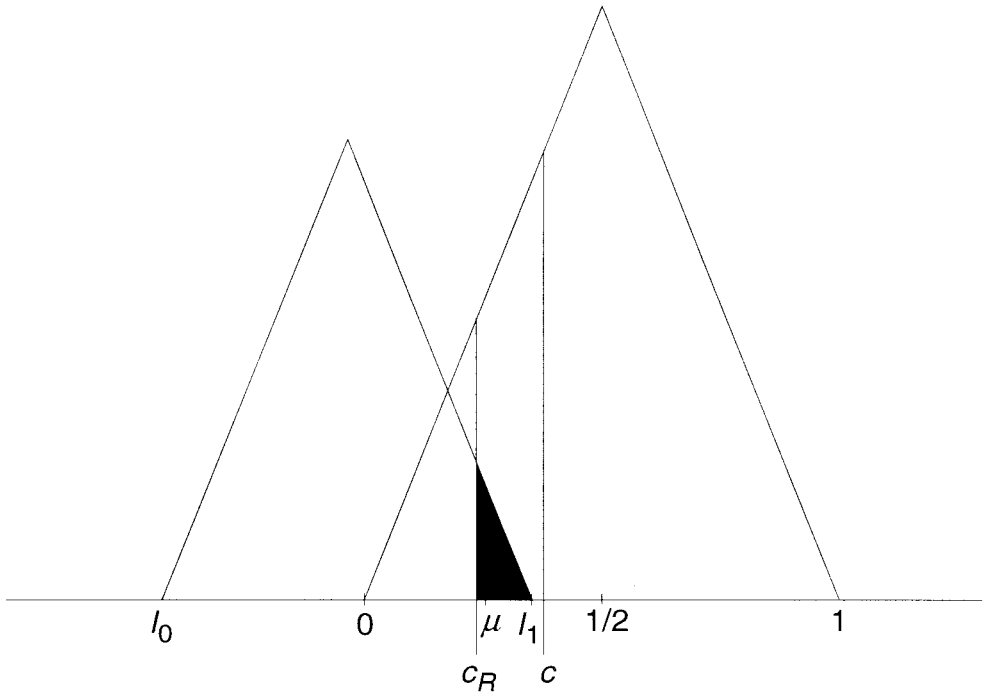
[17] In this example, we assume that voting is deterministic, which we do for analytic convenience only (as do McCarty, Poole, and Rosenthal in their example on page 675). Similar results hold when voting is probabilistic. We cannot find simple closed-form solutions under probabilistic voting, even for simple examples, so we must run simulations. These show that the basic conclusions of this section hold—sometimes even more strongly—when voting is probabilistic. There are two countervailing forces. First, if parties must allocate their "influence" before seeing the member-specific idiosyncratic shocks, then they will want to pad their majorities, which increases the amount of measurable party influence. This will not be true, of course, if parties can observe the member-specific idiosyncratic shocks before allocating their influence. Second, voting "errors" will be especially frequent among legislators close to the cutline, hence among many of the influenced legislators, so adding a second cutline will not improve classification success as much as it does under deterministic voting.

[18] We employ a triangular distribution, because it is an analytically tractable way to capture the distribution of ideal points typically produced by empirical (metric) models of roll call voting, such as Poole and Rosenthal's NOMINATE, D-NOMINATE, and W-NOMINATE, and various linear factor models.

[19] Given our parameterization, $\psi = (l_1 - l_0)^2$ and $\phi = 4l_1^2/(1 + \psi)$.

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**FIGURE 3. Increase in Classification Success When Party Influence Exists**




5. $\mu < l_1, l_1 > c$, and $l_1/2 > c_R$. Then $c_1 = l_1/2$, and $M = [l_1^2 - 2c_R^2 - 2(l_1 - c)^2]/(1 + \psi)$.

Figure 3 gives an example of case 2.

We can treat two parameters as "primitives": the relative size of the minority party $\psi$ and the degree of overlap $\phi$. We can calculate all others, including $M$ and $s_R$, from these. Rather than present a full characterization, we provide a table of calculations to illustrate what happens in a variety of circumstances. In Table 3, the columns labeled $N_R$ and $N_L$ show the number of legislators implied by the value of $\psi$ for a legislature with 435 members.

The conclusion is simple: $M$ is never very large, even when there is substantial overlap between the parties (see the columns labeled $\phi$ and $M$ in Table 3). In many cases, $M = 0$. In all cases, the amount of influence is large: 15–20% of the $R$ legislators switch their vote as a result of party influence, and the policy outcome is moved from $\mu$ to .5 (see the columns labeled $s_R$ and $\mu$).

What does Table 3 tell us about the results reported by McCarty, Poole, and Rosenthal? As they note, even on a straight party vote there is an upper bound on the amount to which a two-cutline model can increase classification over a one-cutline model. They then report (McCarty, Poole, and Rosenthal 2001, 678): "The average upper bound over all the Congresses we analyze is 5%." Given the parameterization in our example, a two-cutline model reduces misclassifications on a straight party vote by $l_1^2/(1 + \psi) = \phi/4$. Setting this equal to .05 yields $\phi = .20$. For the even higher value of $\phi = .25$ (highlighted in Table 3), we see that $M$ ranges from .000 to .023. So, even if $R$ influences its members on all close votes, there is only a gain of 0% to 2.3% in classification success under the two-cutline model. In our earlier article we estimated statistically significant party influence on about 45% of votes in the 82d–105th Congress (1951–98). If $R$ influences its members on only 45% of the close roll calls, and if we use the 2.3% figure and average across all close roll calls, the gain in classification success under the two-cutline model is just 1%.[20] This is even less than what McCarty, Poole, and Rosenthal find. For example, in their Figure 7 for the same set of Congresses, on close votes the average gain in classification success from the two-cutline model is about 1.5%.

## CONCLUSION

McCarty, Poole, and Rosenthal claim that our strategy—the use of lopsided roll calls to estimate ideal points—may not accurately estimate the ideal points of moderate representatives, which would bias our estimates of party influence. We would agree if that were the case, but we believe the use of lopsided roll calls does not lead to poor estimates of moderates' ideal points.

McCarty, Poole, and Rosenthal's argument rests heavily on a simple example with errorless voting. We have shown that this assumption is strongly contradicted by the evidence, which reveals that congressional voting contains a fairly large amount of "error." Therefore, it is much more accurate to view roll call voting in terms of a low-dimensional probabi-

[20] Even this is probably an overestimate. During the 82d–105th Congress the average value of $\psi$ was .71, which corresponds to the bottom of the highlighted rows in Table 3, for which $M = .000$.



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**TABLE 3. Calculations for Cases 1–5**

| $\psi$ | $N_R$ | $N_L$ | $\phi$ | $l_1$ | Case | $\mu$ | $c$ | $c_R$ | $c_1$ | $M$ | $s_R$ |
|---|---|---|---|---|---|---|---|---|---|---|---|
| .85 | 235 | 200 | .05 | .15 | 1 | .19 | .35 | .19 | .19 | .000 | .17 |
| .85 | 235 | 200 | .10 | .22 | 2 | .19 | .35 | .19 | .19 | .001 | .17 |
| .85 | 235 | 200 | .15 | .26 | 2 | .20 | .35 | .19 | .19 | .005 | .17 |
| .85 | 235 | 200 | .20 | .30 | 2 | .21 | .36 | .19 | .19 | .013 | .18 |
| **.85** | **235** | **200** | **.25** | **.34** | **2** | **.22** | **.36** | **.19** | **.19** | **.023** | **.19** |
| .85 | 235 | 200 | .30 | .37 | 4 | .24 | .37 | .19 | .19 | .035 | .20 |
| .85 | 235 | 200 | .35 | .40 | 5 | .25 | .37 | .20 | .20 | .045 | .20 |
| .78 | 245 | 190 | .05 | .15 | 1 | .24 | .37 | .24 | .24 | .000 | .16 |
| .78 | 245 | 190 | .10 | .21 | 1 | .24 | .37 | .24 | .24 | .000 | .16 |
| .78 | 245 | 190 | .15 | .26 | 2 | .24 | .37 | .24 | .24 | .001 | .16 |
| .78 | 245 | 190 | .20 | .30 | 2 | .24 | .37 | .24 | .24 | .004 | .16 |
| **.78** | **245** | **190** | **.25** | **.33** | **2** | **.25** | **.38** | **.24** | **.24** | **.010** | **.17** |
| .78 | 245 | 190 | .30 | .36 | 2 | .26 | .38 | .24 | .24 | .018 | .18 |
| .78 | 245 | 190 | .35 | .39 | 4 | .27 | .38 | .24 | .24 | .028 | .18 |
| .71 | 255 | 180 | .05 | .15 | 1 | .27 | .39 | .27 | .27 | .000 | .15 |
| .71 | 255 | 180 | .10 | .21 | 1 | .27 | .39 | .27 | .27 | .000 | .15 |
| .71 | 255 | 180 | .15 | .25 | 1 | .27 | .39 | .27 | .27 | .000 | .15 |
| .71 | 255 | 180 | .20 | .29 | 2 | .27 | .39 | .27 | .27 | .001 | .15 |
| **.71** | **255** | **180** | **.25** | **.33** | **2** | **.28** | **.39** | **.27** | **.27** | **.004** | **.15** |
| .71 | 255 | 180 | .30 | .36 | 2 | .28 | .39 | .27 | .27 | .009 | .16 |
| .71 | 255 | 180 | .35 | .39 | 2 | .29 | .39 | .27 | .27 | .016 | .16 |

$\psi$ = (size of minority party)/(size of majority party).
$\phi$ = overlap = fraction of legislators with ideal points in $[0, l_1]$.
$l_1$ = rightmost ideal point in party $L$; 0 = leftmost ideal point in party $R$.
$\mu$ = overall median in legislature.
$c$ = cutline on bill under sincere voting.
$c_R$ = cutline inside $R$ under party influence ($c$ = cutline inside $L$).
$c_1$ = cutline choice to minimize classification errors in one-cutline model.
$M$ = fraction of votes misclassified under one-cutline model.
$s_R$ = fraction of $R$ legislators who change their vote due to party influence.

listic voting model than a deterministic model. Poole and Rosenthal themselves have strenuously argued in support of this view and have mustered a large amount of evidence for it in their important work on NOMINATE. In a probabilistic voting framework, we can estimate the ideal points of moderates as well as extremists, even using a set of roll calls with cutlines that rarely fall between moderates. The Monte Carlo simulations in our previous article demonstrate this. We demonstrate it in another way here, by showing that an independent source of data for measuring ideal points—the NPAT survey—produces estimates that are very close to those produced by our lopsided roll call strategy.

We also show that McCarty, Poole, and Rosenthal's own results, when properly interpreted, actually support our earlier conclusions. In particular, their two-cutline method reveals that party influence is significant on average close votes. Indeed, it is large enough to move the induced ideal point of a legislator across about 11% of the House. We also demonstrate that McCarty, Poole, and Rosenthal's "classification improvement" criterion is inappropriate for assessing party influence. Even when party influence is large—large enough to be consistent with one definition of "strong party government"—the improvement in classification success often will be as small or even smaller than the 1.5% that McCarty, Poole, and Rosenthal find.

Many questions remain open. Perhaps the most important concerns the sources of party influence. McCarty, Poole, and Rosenthal (2001, 674) 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." This may be true, but it is meaningless in the context of their article. After all, their ideal point estimates are defined in terms of roll call votes! Because these are the only data used to estimate ideal points, it is a contradiction to argue that party influence is "manifest in the location of the legislator's ideal point" (p. 686) but that party influence has no effect on the legislator's roll call votes.

More usefully, McCarty, Poole, and Rosenthal (2001, 686) suggest that forces external to Congress, such as "external pressures of campaign donors and primary races," are more important than internal forces. A rough calculation provides a crude idea about the relative magnitude of internal and external forces. McCarty, Poole, and Rosenthal note that when representatives switch parties, on average their ideal point moves across 28% of the ideal points of the legislature. If one accepts our interpretation of their results, then party influence due to forces within the House causes a legislator's ideal point to move across about 11% of the ideal points of the House. This implies that when a legislator switches parties, about 39% (11/28) of the total movement is caused by party influence within the legislature. The remaining 61% is evidently due to external factors, such as new primary voters and campaign donors. Interestingly, this is roughly consistent with Levitt's (1996) study of senatorial voting. Levitt finds that senators place approximately equal weight on

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the positions desired by their primary constituents and the positions desired by their caucus or party leaders. Yet, as McCarty, Poole, and Rosenthal note, most studies find that changing external circumstances, such as redistricting, have little effect on members' voting scores, which suggests little responsiveness to overall district preferences. Taken together, these findings imply that primary electorates and campaign donors have a greater effect than overall constituencies on representatives' roll call voting. Most democratic theorists would probably find this highly disturbing.

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