GRADUATE SCHOOL
OF INDUSTRIAL ADMINISTRATION
WILLIAM LARIMER MELLON, FOUNDER

REPRINT NO. 1221

Analysis of Congressional Coalition Patterns: A Unidimensional Spatial Model

By

Keith T. Poole and Howard Rosenthal

1987

Carnegie-Mellon University
PITTSBURGH, PENNSYLVANIA 15213

KEITH T. POOLE
HOWARD ROSENTHAL
Carnegie-Mellon University

Analysis of Congressional Coalition Patterns: A Unidimensional Spatial Model

This article proposes stochastic, spatial model of roll-call voting as a benchmark for evaluating more complex models of congressional behavior. On the basis of several evaluative criteria, this model outperforms Weisberg's two-party and three-party models and Hammond and Fraser's random-voting models. Particularly, the spatial model successfully recovers the occurrences of hurrah, party unity, conservative coalition, and Civil War votes as well as the marginal distribution of yea votes. At the same time, the spatial model is parsimonious, requiring that more parameters be estimated than the two-party model does but fewer than the three-party model.

There are a number of criteria for evaluating statistical models of roll-call voting. These models should do more than just classify individual votes relatively correctly. They should be consistent with observed patterns of coalition voting (Hammond and Fraser, 1983); they should correctly reproduce the marginal distributions of yea and nay voting (Weisberg, 1983); and they should accomplish these tasks with relative parsimony in conceptual formulation and in parameter estimation. Moreover, as Hammond and Fraser (1983) emphasize, a behavioral model of roll-call voting should outperform models based on purely random processes such as votes determined by flips of fair coins.

In this paper, we conduct multiple evaluations of NOMINATE, a one-dimensional spatial model of roll-call voting presented in detail in an earlier publication (Poole and Rosenthal, 1985a). NOMINATE is a parsimonious model in the sense that it needs to have fewer parameters estimated from the data than does the simple three-party benchmark model proposed by Weisberg (1978). Since NOMINATE performs well in these multiple evaluations and in comparison to several other models, it is arguably a strong candidate to serve as a benchmark for evaluating more complex models of roll-call voting.

We begin our presentation, in the next section, by briefly reviewing the NOMINATE procedure. Then we outline various ways of evaluating roll-call models. Finally, we present a detailed analysis of data for the House and Senate and relevant Monte Carlo experiments.

NOMINATE, Unidimensional Spatial Model

The NOMINATE model is conceptually simple. Both legislators and roll calls occupy positions on an unidimensional continuum best thought of as liberal-conservative. Each legislator is represented by a single point on this continuum, corresponding to his/her ideal point in spatial theory, which we denote as x_i ($i = 1, \ldots, p$ where p is the number of legislators). Each roll call is represented by two points, corresponding to the yea and nay outcomes. We denote these two points as z_{yl} and z_{nl} respectively ($l = 1, \ldots, q$ where q is the number of roll calls). The distance of a legislator to a roll-call outcome is $d_{ijl} = |x_i - z_{jl}|$ where j indexes yea and nay. We assume that voting is sincere. Accordingly, legislators with ideal points such that $d_{iyl} < d_{inl}$ should vote yea; those with ideal points at the midpoint ($d_{iyl} = d_{inl}$), an empirically irrelevant situation in our model, might flip a fair coin; and those legislators with ideal points such that $d_{iyl} > d_{inl}$ should vote nay. (We do not allow for abstention because of its relative rarity in congressional voting.) If the above model were true, the observed data would form a perfect Guttman scale (MacRae, 1958).

For a variety of reasons, such as perceptual error engendered by the difficulty of being informed about roll-call alternatives or the presence of multiple dimensions, voting along the dimension is apt to be noisy. Therefore, NOMINATE allows for a form of probabilistic voting consistent with the formulation of Coughlin and Nitzan (1981). In this formulation, voters to the yea side of the midpoint are simply more likely to vote yea. We accomplish this by allowing each legislator to have an interval-level quasi-concave utility function which is composed of a fixed component and a stochastic component; that is, we define the utility of legislator i for alternative j on roll call l to be

$$U_{ijl} = \beta \exp[-d_{ijl}^2/8 + \epsilon_{ijl}$$

where β is a parameter which we estimate, the "8" represents a preset scaling factor, and the ϵ_{ijl} are the error terms which we assume to be independently distributed as the logarithm of the inverse exponential (i.e., the logit distribution; Dhrymes, 1978, pp. 341-2). The parameter β scales the logit error relative to the deterministic portion of the utility function. If β is large, legislators are almost always voting for the closest alternative. As β approaches zero, the voting probabilities approach .5 and the legislators are, in effect, voting by flipping fair coins.

Given the assumption that the ϵ_{ijl} have a logit distribution, the probability that legislator i votes yea/nay on roll call l can be written as

$$P_{ijl} = \frac{e^{u_{ijl}}}{e^{u_{iyl}} + e^{u_{inl}}}$$

where u_{ijl} is the deterministic portion of the utility function. The likelihood function is the product of the P_{ijl} corresponding to the actual choices made by the legislators. Because this model is nonlinear in its parameters, standard linear logit packages cannot be used to maximize the likelihood function. In addition to the nonlinearity, the number of parameters is typically so large (p + 2q + 1) that simultaneous estimation of them is simply impractical. Accordingly, we developed an algorithm that alternates between estimating the roll-call outcomes, the legislator coordinates, and finally the utility function parameter, holding the other two sets of parameters constant while one set is being estimated. Following standard practice, in each phase we estimate parameters which maximize the logarithm of the likelihood of the observed choices of the legislators. The successive estimations of the z_{jl}, the x_i and β define a global iteration. We define convergence as a squared Pearson correlation of .99 or better between all coordinates (both the x_i and z_{jl}) estimated in the current global iteration with those estimated in the previous iteration (see Poole and Rosenthal, 1985a, for details). The NOMINATE acronym denotes *Nomina*l *T*hree Step *E*stimation. Alternating algorithms of this type are common in psychometric applications (e.g., Carroll and Chang 1970; Takane, Young, and deLeeuw, 1977).

Alternative Approaches to Evaluating Roll-Call Models

In evaluating NOMINATE or other models of roll-call analysis, one can use several criteria.

Ability to Classify or Fit Individual Votes

For each legislator voting on each roll call, the model predicts a vote. One can compare actual votes to predicted votes and use the percentage correctly predicted as a measure of fit. For a successful model, this percentage obviously should substantially exceed 50. In addition, in a legislature where there are many yea votes with large majorities, the percentage correctly predicted should substantially exceed the average percentage yea.

In probabilistic limited dependent variable models like standard logit or probit or NOMINATE, one may predict yea if the model's probability that a legislator votes yea on a particular roll call exceeds 0.5. Obviously, converting the probabilities into binary choices results in a loss of information. As a result, a frequently used alternative is the average probability assigned to the observed choices.

An alternative to using average probability is simply to divide the total log-likelihood by sample size to get an average log-likelihood. This quantity can then be exponentiated to yield the geometric mean probability.

Comparing the geometric mean to the average probability is analogous to comparing squared error to absolute deviation error. The geometric mean heavily weights poor observations. Thus, if two observed choices have assigned probabilities of .1 and .9, the average probability is 0.5 while the geometric mean is only 0.3.

Amemiya (1981) discusses other methods for evaluating limited dependent variable models on a case-by-case basis. In Poole and Rosenthal (1985a), we evaluated the model exclusively with the percentage correctly classified and the geometric mean. For the House in 1957-1958 and the Senate from 1979 through 1982, we found classification percentages in the range of 79% to 83% and geometric mean probabilities ranging from 0.63 to 0.69.

We now turn to other criteria which will be the emphasis of this paper.

Accounting for Coalition Behavior

Hammond and Fraser (1983) recently suggested evaluating substantive models of coalition formation in terms of their ability to outperform purely random baseline models of coalition formation. In a two-party model, one can consider "party unity" votes, in which a majority of Democrats oppose a majority of Republicans, and "hurrah" votes, in which majorities of both parties are the same side of the issue. A three-party model gives rise to two additional patterns. In "conservative coalition" votes, southern Democrats and Republicans vote on one side, in opposition to a majority of northern Democrats. In "Civil War" votes, a majority of southern Democrats is opposed by a majority of Republicans and a majority of northern Democrats.

For each type of roll call, Hammond and Fraser suggest two questions. First, does the model reproduce the marginal percentages of each type of roll calls? Second, when these roll calls occur, does the model correctly "predict" which side wins?

Of course, very little can be learned about actual coalition behavior or logrolling from such an analysis. If, for example, logrolling takes place among spatially adjacent legislators, so that a few moderate conservatives who might have voted nay instead vote yea on a given roll call, the observations will simply appear as if the midpoint has shifted toward the conservative end of the spectrum. As a result, one cannot differentiate this type of logrolling from pure spatial voting. Similarly, the relative frequency of conservative coalition votes and party unity votes does not tell us whether coalitions actually form or whether the votes are a pure spatial response to a set of midpoints appearing on the agenda.

In contrast, the case of Civil War votes is particularly instructive. Civil War votes represent both ends opposing the middle, in spatial terms. Such voting is clearly inconsistent with the model posited in NOMINATE. As a result, Hammond and Fraser have proposed a useful check. One could correctly classify most votes but fail with Civil War votes, a result which suggests that either an additional dimension was very important or that the behavioral assumptions underlying NOMINATE (roll-call voting viewed as independent dichotomous choice with no logrolling) were seriously inadequate.

Accounting for the Marginal Distribution of Yea Votes

Hammond and Fraser (1983) demonstrated that the distribution of actual congressional roll calls over the various coalition patterns and the frequency of wins for the various coalitions could be accounted for to a large degree if one simply assumed that each member of Congress flipped a fair coin on every vote. Weisberg (1983) pointed out, however, that fair coin flips would predict 50% yea votes on average whereas the actual percentage yea in Congress exceeded 60% in certain years. When biased coin-tossing models were evaluated by Weisberg, they failed to reproduce the coalition patterns. It is thus important to evaluate a model in terms of its ability to capture not only the patterns of coalition voting but the yea marginals as well. Hammond and Fraser's results matched some of the coalition patterns, but their marginals differed substantially from the real data. Weisberg did better on the marginals, but his results in some cases did not account as well for the coalition patterns. As we show below, NOMINATE is an improvement on both counts.

Checking for Method Artifact

Roll-call analysis as performed by NOMINATE is in fact a form of scaling for nominal data. Because nonlinear scaling algorithms tend to produce misleading results, the results must be checked. Two related questions can be asked. First, can the scaled distributions of legislators result from purely random-choice models of the type proposed by Hammond and Fraser (1983)? Second, if Monte Carlo choice data is generated from an environment obeying the assumptions of the model, will the true positions of the legislators and roll calls be recovered?

Parsimony in Parameters and Concepts

Models of congressional voting vary in their conceptual simplicity. At one extreme are the simple coin-tossing models of Hammond and Fraser. Also simple are the benchmark two-party and three-party models of Weisberg. In these models, one merely predicts that each legislator votes with the majority of his or her party. One-dimensional spatial voting, as in NOMINATE, is also straightforward conceptually. Other conceptually direct work assumes that, on each roll call, a legislator's utility for a yea vote is approximated by a linear function of independent variables describing characteristics of his/her constituency (See Kalt and Zupan, 1984; Peltzman, 1984).

Multidimensional spatial voting models, as in Clausen (1973) or Warwick (1977) are obviously more complex than NOMINATE. We argue elsewhere (Poole and Rosenthal, 1985a; see also Morrison, 1972) that multidimensionality is at least partially the consequence of failing to treat error explicitly, as in Clausen's work, or of method artifact, as in factor analyses performed by Warwick (1977). In any event, we do not consider multidimensional models in this paper.

Far more complex models are found in various simulation models of roll-call voting, such as Cherryholmes and Shapiro (1969) and Matthews and Stimson (1975).

Parsimony in estimated parameters is related (but not identical) to conceptual simplicity. Coin tossing is conceptually simple and involves no estimation. The two- and three-party models estimate 2q and 3q parameters respectively, because each roll call must be examined in order to determine how the majority of each "party" voted. NOMINATE has one parameter for the utility function, p for the legislators, and 2q for the roll calls for a total of p + 2q + 1. The constituency variables models typically employ many independent variables. Thus, in studying the Senate, Peltzman (1984) has a total of 18q parameters. Since the number of legislators is, particularly for the Senate, generally fewer than the number of roll calls, NOMINATE falls in between the two-party and three-party models in the number of its parameters, while the constituency variables models estimate far more than either NOMINATE or the party models.[1]

All of the models just mentioned have many fewer parameters than the simulation models. Matthews and Stimson (1975) require far more than pq parameters. In other words, they require more parameters than the actual number of individual votes.

Considering the number of parameters, we can reject the simulation models outright. One should always be estimating fewer, preferably substantially fewer, parameters than the number of observations. The constituency-variables models, given the substantial number of parameters estimated, would, in order to be useful, have to substantially outperform both NOMINATE and Weisberg's benchmarks. Evidence that they do not add substantially to the predictions of NOMINATE is presented in Poole and Rosenthal (1985b).

Prediction

It is important to point out that none of the models considered is useful for prediction simply because (except for the coin tosses) they all involve parameters that are roll-call specific. Given a sampling of legislators' intentions on a key roll call, both the party benchmarks and NOMINATE could be used to forecast the final vote. The constituency-variables models would be far less reliable because of the instability of multiple regression estimates when the coefficients of many independent variables are estimated from small samples.

Internal Validity

The simulation models would appear to capture most realistically the cue taking and coalition formation that takes place in actual legislative settings. As we indicated above, they are in fact too realistic to be useful models. The constituency-variables models also point to a plausibly critical influence on roll-call behavior. However, the highly aggregated variables they use are generally implausible measures of constituency preferences (Fiorina, 1974). NOMINATE captures just one aspect of roll-call behavior, namely, liberal-conservative ideology. To the extent that the "parties" are just averages of positions on the liberal-conservative continuum, the two-party and three-party benchmarks would approximate aggregated versions of the same unidimensional model. Coin tossing is totally implausible.

In summary, coin tossing is an obviously inadequate description of purposive behavior and the simulation models are too rich in parameters to be of much use. However, both the Weisberg models and NOMINATE can be recommended as comparative benchmarks, since they both involve only a relatively small set of parameters. In earlier work (Poole and Rosenthal, 1985a), we demonstrated that NOMINATE outperformed the two- and three-party models in predicting votes.

Specifically, for the data sets analyzed in the paper, the spatial model makes fewer errors than the two-party model in classifying the votes of individual legislators on 73% of the roll calls and fewer errors than the three-party model on 67% of the roll calls.[2] Consequently, we now evaluate how well NOMINATE fits the marginals and coalition patterns. We also mention several tests that show the results of NOMINATE are not method artifacts.

The Liberal-Conservative Continuum in Congress: Estimated Legislator Positions and Possible Method Artifact

The distribution of legislator positions estimated by NOMINATE for the House in 1957-1958 and the Senate in 1979, 1980, and 1981 are shown in Figure 1. (In all years, the space has been normalized so that the most liberal legislator is at -1 and the most conservative is at $+1$.)

Could these distributions be artifactual in the sense that they could have arisen from coin flipping? Figure 1 also shows the distribution estimated by NOMINATE from a Monte Carlo study in which 100 "senators" tossed coins on 297 roll calls (i.e., 29,700 random numbers were generated). Unlike our estimates from actual roll-call data, the distribution is strongly unimodal and tightly grouped in the center. Each bar graph in the figure represents a division of the $[-1, +1]$ interval into ten equally spaced intervals. The outer four bars of the coin-toss experiment always have less of the distribution than do the corresponding four bars for the houses of Congress. Similarly, far more of the distribution is concentrated in the center two bars on the coin-toss experiment data than in actual data. The result for the fair coin toss is accented by the result for a similar run when a weighted coin is flipped with yea probability .624, the proportion of yea votes reported by Weisberg (1983) for the 1977-1978 Senate. The distribution is very tightly packed. Indeed, as the probability approaches 1.0, the distribution can be expected to become even more concentrated.

If we convert the log-likelihood computed by NOMINATE to a geometric mean probability, we obtain values of .507 for the .50 coin flip and .522 for the .624 coin flip.[3] For the actual roll-call vote data, our geometric means range from .63 to .69. Both the distribution of coordinates and measures of fit, therefore, indicate that the NOMINATE results are very unlikely to arise by chance.

An alternative interpretation of the results from our coin-toss experiments recognizes that, in (1), β is a parameter that scales the noise level relative to the nonstochastic portion of utility. One scenario that could underlie 0.5 coin tosses would be for β to be very, very small relative to the nonstochastic portion. As β approaches 0, the voting prob-

FIGURE 1
Liberal-Conservative Distributions

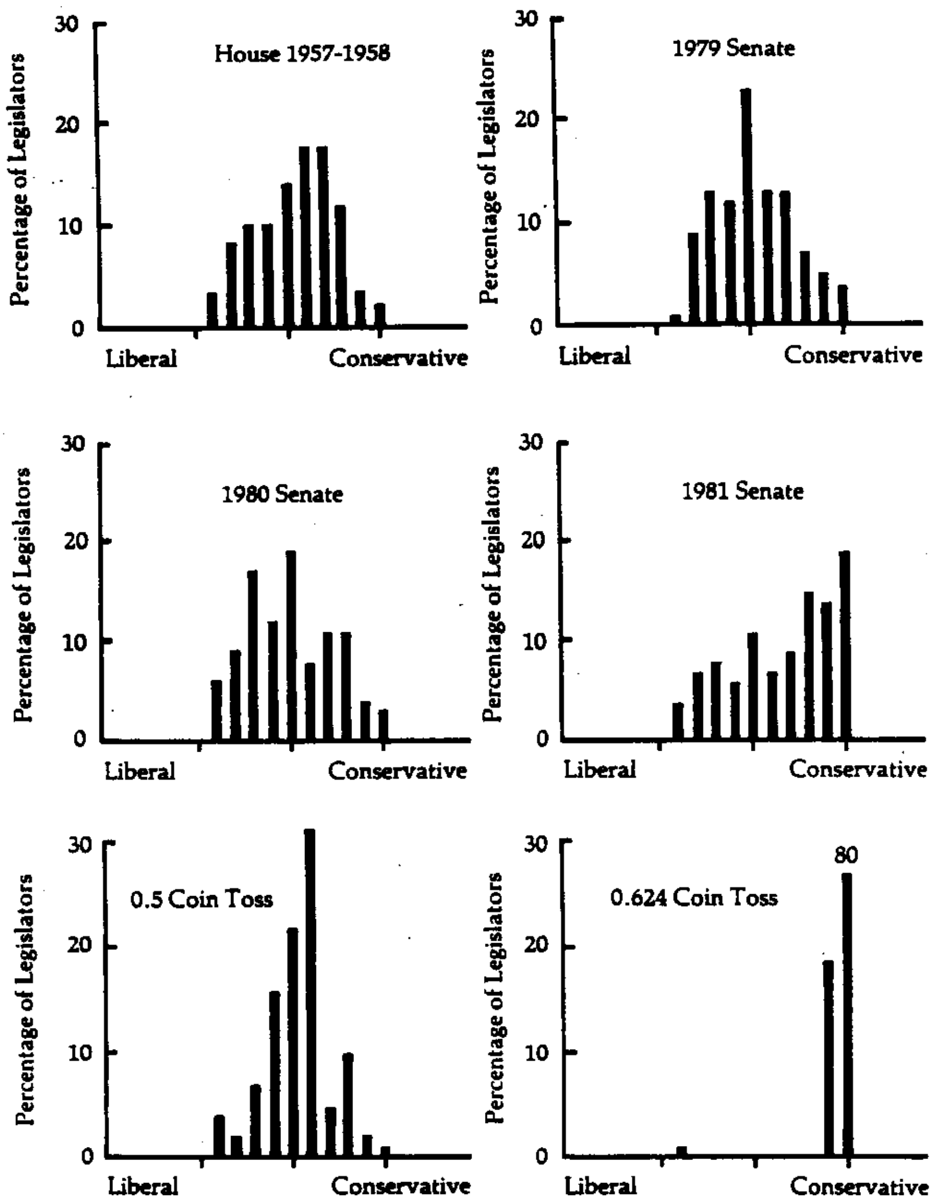

TABLE 1
Monte Carlo Results for Senators and Utility Function

Run	β	Data Configuration	R^2[a]	Standard Error of Recovery[b]	Recovered β
A	15.00	Midpoints of 97 roll calls	.990	.047	16.47
B	22.50	generated at midpoints of	.990	.047	27.46
C	7.50	adjacent senators. Three liberal	.980	.068	8.72
D	18.25	coordinates per midpoint. Total	.990	.048	20.53
E	11.75	of 291 roll calls.	.988	.053	12.45
F	15.0	Liberal coordinates and midpoints uncorrelated.	.987	.055	12.45
G	15.0	Midpoints throughout but	.988	.054	17.96
H	15.0	concentrated in center.	.986	.057	19.87
I	15.0	Liberal coordinates generated by random process.	.987	.055	19.87
50 Senators, β = 15.0, run A structure			.991	.047	16.88

[a] R^2 is the squared correlation between true and recovered coordinates.

[b] The standard error of the regression of true β on recovered β. Since a space is defined only up to a linear transformation, it is appropriate to pick such a transformation before assessing recovery error.

abilities will approach 0.5 for all legislators on all roll calls. For congressional roll-call data, estimates of β are on the order of 15. The contrast between results of a β of 15 and a β of 0 is amply illustrated in Figure 1.

In other Monte Carlo experiments, we generated random logit errors and then generated utilities using (1). We used the 98 interior senator coordinates recovered by NOMINATE from the 1979 Senate data and tried various distributions of roll-call coordinates over the space. We also ran a Monte Carlo experiment with only 50 senators.

Results for the "senator" coordinates from these experiments (see Poole and Rosenthal, 1983, for greater details) are shown in Table 1. It can be seen that coordinates are accurately recovered. While the senators are distributed over an interval of length 2.0, the standard error of the recovery is only on the order of 0.05. In addition, the 50 senators are recovered as accurately as the 98, since the effective sample size in both cases is essentially the number of roll calls.

At a given effective sample size, recovery of the midpoints is about as accurate as recovery of the legislators. This result simply verifies that legislators and midpoints play symmetric roles in the NOMINATE model. Thus with 50 senators, midpoints will be recovered less accurately than with 100. In general, NOMINATE results appear to satisfy the basic statistical fold theorem that standard errors of recovery decrease linearally with the square root of the effective sample size. As a consequence, we obtain very reliable midpoint estimates with the larger sample size of 435 representatives in the House.

Although the midpoints are reliably recovered, the yea and nay coordinates must be used with some caution. This is so because the recovery of the outcome coordinates is sensitive to the level of noise in the roll call. Suppose voting on a particular roll call were perfect; that is, suppose we observe YYY . . . YYYNNN . . . NNN. The midpoint location is immediately pinned down to the interval between the coordinates corresponding to the rightmost Y and the leftmost N. In contrast, any pair of outcomes equidistant from the midpoint could have produced the voting pattern if there were no error. In effect, in order to recover outcome coordinates, we need error. This problem becomes less serious as the size of the legislature increases because, at a fixed level of error, there will be more "mistakes" in the voting and it becomes easier to identify the outcome locations. Consequently, the estimation of the roll-call outcomes for the House will be more reliable than those estimated for the Senate.

Up to this point, we have shown that NOMINATE will not artifactually produce good fits to a purely random world and that the procedure accurately recovers unidimensional, stochastic voting. Would we also get good fits if legislative behavior was more complex than that posited by the NOMINATE model?

Consider, first, a two-outcome, two-dimensional spatial model with sincere voting. Suppose the two dimensions are equally salient and the dimension recovered by NOMINATE represented the first axis. If a roll call has yea and nay outcome points that are perpendicular to the first axis—that is, the yea and nay outcome points have the same coordinate on the first dimension—then voting, when viewed from the perspective of the first dimension, will appear to be random. Consequently, the estimation of the outcome points for such a roll call will be very poor. (Note that, in this case, one can maximize classification accuracy by predicting that all legislators will vote yea if the actual percentage yea exceeds 50.) For roll calls with outcomes that are on a line that is at an angle, but not orthogonal to, the dimension estimated by NOMINATE, we would obtain a relatively noisy fit—the larger the angle, the noisier the fit. Consequently, although the overall fit of the model will not be terribly high, it will not be terribly low either and it will be considerably better than that for a random set of votes.

Consider, however, what would happen in an n-dimensional, two-outcome spatial model with sincere voting and all dimensions of equal salience. Now, almost all the roll calls will be nearly orthogonal to any one dimension. As a result, nearly all voting will look like coin flips when the roll-call outcome points are projected onto a single dimension. Given that our assumptions are corect, we can use the NOMINATE results to reject the hypothesis of very high dimensionality in congressional voting. At the same time, we cannot reject the hypothesis that NOMINATE successfully approximates only a two- or three-dimensional pattern of congressional voting.

In addition to multidimensionality, another complexity missing from NOMINATE is logrolling. We earlier noted that logrolls among spatially adjacent legislators cannot be differentiated from sincere voting. Other forms of strategic behavior would reduce the fit of the model. Some strategic behavior would be manifest as Civil War votes. Our results (see further details below) suggest, given that our assumptions are correct, that such strategic behavior does not occur with a frequency sufficient to upset the fit of a unidimensional model.

To summarize this discussion of method artifact, our Monte Carlo experiments have shown that random coin flips cannot generate the actual data results of NOMINATE and that NOMINATE will accurately recover spatial locations when the data generated has the considerable amount of random noise that is implied by our probabilistic model of spatial voting. Theoretical considerations further permit us to reject a highly multidimensional world as a source of the NOMINATE results. We now move on to consider whether NOMINATE recovers the patterns of coalition voting in Congress.

Explaining the Distribution of Coalition Patterns

Method

To conduct the evaluations suggested by Hammond and Fraser (1983) and Weisberg (1983), we analyzed the House in the 85th Congress (1957-1958) and annual data for the Senate for 1977, 1979, 1980, 1981, and 1982. The number of years is limited because estimating the model involves expensive, iterative, nonlinear techniques. Processing any one of the data sets requires about two hours of cpu time on a VAX 11/780. In our analysis, we restrict ourselves to votes where the actual minority (including pairs and announced votes) exceeded 2.5%. We exclude unanimous and near-unanimous votes from NOMINATE for technical reasons (see Poole and Rosenthal, 1985a). However, we could readily include them for present purposes by assigning the midpoints to the left of the most liberal or to the right of the most conservative legislator, as appropriate, and predicting unanimity. All such near unanimous votes are fact hurrah votes, so we would trivially classify these correctly. Such a procedure would only improve the fits shown in Tables 2 and 3.

To assess the fit to the actual percentages of hurrah, party unity,

TABLE 2
The Marginal Percentage Distribution of Yes Votes[a]
(in percentages)

Data Set	Average of Actual Marginals	Average of Predicted Marginals
House 1957-1958	59.0	62.0
Senate 1977	57.0	58.7
Senate 1979	56.9	58.4
Senate 1980	59.8	61.2
Senate 1981	54.0	55.6
Senate 1982	57.7	60.0

[a]Pairs and announced votes are excluded; unanimous and nearly unanimous roll calls (those with less than 2.5% minority) are also excluded.

TABLE 3
Two-Party Comparisons
(in percentages)

	All Votes That are Party Unity Votes		Party Unity Votes Won by Democrats		
Data Set	Actual	NOMINATE	Actual	NOMINATE	Total Roll Calls
Pairs and Announced Votes Included					
House 1957-1958	55.8	59.6	66.7	59.0	172
Senate 1977	54.6	64.6	63.3	60.2	518
Senate 1979	53.0	52.1	79.8	82.4	448
Senate 1980	52.2	52.4	71.3	72.2	480
Senate 1981	59.3	61.6	18.0	21.1	377
Senate 1982	49.6	51.1	28.2	29.0	421
Pairs and Announced Votes Excluded					
House 1957-1958	55.5	57.6	68.6	57.8	172
Senate 1977	53.9	64.8	62.0	60.1	518
Senate 1979	53.0	52.1	79.4	82.4	448
Senate 1980	51.9	52.2	71.1	72.9	480
Senate 1981	59.3	61.3	17.2	19.9	397
Senate 1982	49.6	51.1	26.8	28.4	421

conservative coalition, and Civil War votes, we had to develop predicted percentages from NOMINATE. This is easy. On a given roll call, if a legislator is to the left of the midpoint, he or she is recorded as having voted on the liberal side. Otherwise, his or her vote is conservative. So, for example, if both over 50% of all Democrats and over 50% of all Republicans are to the left of the midpoint estimated by NOMINATE, the roll call is classified as a hurrah vote for the two-party case.

We carried out two types of analyses. First, as is our standard procedure with NOMINATE, we recoded announced votes and pairs as if they were actual votes. Second, emulating Hammond and Fraser (1983), we treated pairs and announced votes as missing data. The NOMINATE computations for pairs and announced votes excluded in Tables 2 through 5 are based on predictions for only those legislators who actually voted on the roll calls considered.

Results

As a preliminary, we point out that NOMINATE closely matches the marginal distribution of yea votes (on roll calls with more than 2.5% minorities). This can be seen in Table 2. There is a slight upward bias because, when a roll call fits the unidimensional model poorly, NOMINATE leads to predictions that nearly everyone will vote with the majority. Since yea votes are empirically more frequent than nays, these predictions force the observed upward bias.

In general, we fit the two-party distributions quite well, as can be seen in Table 3. The distribution of party unity votes (and, by implication, hurrah votes) is closely approximated by NOMINATE. The largest difference is 10% (for the 1977 Senate), with the next largest difference being only 4%. Similarly, the percentage of party unity votes won by Democrats is closely tracked. The only large deviation is for the House in the 85th Congress, where we are off by 7.7%. But even this deviation is substantially less than that of the coin-toss model, which appears to be off by about 16% (Hammond and Fraser, 1983, Figure 2b, p. 647). After unidimensional spatial voting has been taken into account, any theory of conscious coalition formation is not left with much variance to explain, even less than what is left after what would have occurred randomly (with fair coins) is taken into account.

Unlike the two-party results, the three-party results shown in Table 4 contain some systematic deviations between NOMINATE and the actual data. Most importantly, NOMINATE never predicts a Civil War vote, in which both ends (liberals and conservatives) voting against the middle (moderates) on one dimension. However, in the recent Senate data, Civil War votes are sufficiently rare that a one-dimensional model

TABLE 4
Three-Party Occurrence Comparisons

Data Set	Hurrah		Party Unity		Civil War		Conservative Coalition	
	Actual	NOMINATE	Actual	NOMINATE	Actual	NOMINATE	Actual	NOMINATE
Pairs and Announced Votes Included[a]								
House 1957-1958	28.8	37.8	36.6	46.5	14.8	0.0	19.8	15.7
Senate 1977	43.1	42.6	23.6	26.9	5.6	0.0	38.0	40.2
Senate 1979	38.7	43.3	37.2	39.1	5.7	0.0	18.4	17.6
Senate 1980	40.7	45.0	34.2	35.5	4.7	0.0	20.4	19.5
Senate 1981	31.6	34.9	41.8	45.5	4.2	0.0	27.7	25.5
Senate 1982	41.0	42.9	32.9	32.3	4.9	0.0	21.3	24.8
Pairs and Announced Votes Excluded								
House 1957-1958	28.8	39.0	36.0	45.3	15.4	0.0	19.8	15.7
Senate 1977	43.9	42.4	22.8	26.9	5.5	0.0	37.5	40.3
Senate 1979	38.2	43.2	37.3	39.3	5.9	0.0	18.6	17.5
Senate 1980	40.8	45.2	34.0	35.4	4.7	0.0	20.5	19.4
Senate 1981	31.8	34.4	41.8	45.5	4.0	0.0	27.7	25.5
Senate 1982	41.0	43.2	33.0	32.3	4.9	0.0	21.1	24.5

[a]For total roll calls, see Table 1.

TABLE 5
Three-Party Outcome Comparisons
(in percentages)

Data Set	Party Unity Votes Won by Democrats		Conservative Coalition Votes Won by Conservative Coalition	
	Actual	NOMINATE	Actual	NOMINATE
Pairs and Announced Votes Included				
House 1957-1958	84.1	75.0	67.6	100.0
Senate 1977	90.9	98.9	67.5	70.2
Senate 1979	92.7	99.7	61.4	79.9
Senate 1980	85.8	98.5	65.3	86.8
Senate 1981	23.3	29.2	96.4	99.5
Senate 1982	37.3	44.9	94.5	97.7
Pairs and Announced Votes Excluded				
House 1957-1958	83.9	73.1	64.7	100.0
Senate 1977	90.1	98.9	68.7	70.0
Senate 1979	92.7	100.0	61.5	81.6
Senate 1980	85.1	98.3	67.5	87.2
Senate 1981	22.3	27.4	95.5	99.5
Senate 1982	35.0	44.1	94.4	99.0

is in fact a good first approximation to actual voting. In all cases, even for the House in the 85th Congress, NOMINATE outperforms the simple coin-toss model, which predicts 25% of the votes will fall into each of the four categories in Table 4. The sum, over the four categories, of the absolute deviations between predicted and actual percentage is always less for NOMINATE.

We also appear to do much better than the weighted coin-toss model, particularly for the critical category of the conservative coalition. Under the weighted coin-toss model, the conservative coalition should arise less than 8% of the time for any of the data sets described in Table 4.[4] This rate is far below both the actual rates and the rates predicted by NOMINATE.

We track vote outcomes as well as we track vote occurrences. The vote outcome results are shown in Table 5. NOMINATE appears to overestimate the success of the conservative coalition for some years. This discrepancy can be explained. Although NOMINATE matches the percentage distribution of occurrences in Table 4, the votes classified as conservative coalition votes by NOMINATE are not always actual conservative coalition votes and vice versa. In addition, votes for which we had

TABLE 6
Conservative Coalition Winner Predictions:
Roll Calls Classified by NOMINATE as Conservative Coalition Votes[a]
(in percentages)

Data Set	NOMINATE Predicts Conservative Coalition Victory	Conservative Coalition on Winning Side	Conservative Coalition Defeats Northern Democrats	Total Roll Calls
House 1957-1958	100.0	100.0	85.2	27
Senate 1977	70.7	71.7	68.3	205
Senate 1979	87.7	87.7	82.2	74
Senate 1980	90.0	91.1	88.9	90
Senate 1981	100.0	100.0	92.4	92
Senate 1982	100.0	100.0	91.9	99

[a]Pairs and announced votes are excluded.

TABLE 7
Conservative Coalition Winner Predictions:
Roll Calls That Were Actual Conservative Coalition Votes[a]
(in percentages)

Data Set	NOMINATE Predicts Conservative Coalition Victory	Conservative Coalition Actually Wins	Total Roll Calls
House 1957-1958	82.4	64.7	34
Senate 1977	67.0	69.1	188
Senate 1979	64.6	63.3	79
Senate 1980	68.9	71.1	90
Senate 1981	96.0	97.0	101
Senate 1982	96.6	95.4	87

[a]Pairs and announced votes are excluded.

to break ties in order to classify the coalition pattern are obviously prone to prediction errors. Consequently, we eliminate ties from the ensuing analysis.

On comparable roll calls, NOMINATE closely reproduces the actual frequency of conservative coalition votes. Consider first Table 6. It can be seen that, for NOMINATE's predicted conservative coalition votes, the conservative coalition is on the winning side almost exactly as often as NOMINATE predicted. Moreover, the conservative coalition defeats the northern Democrats less frequently than it wins. This result reflects the fact that a few votes classified by NOMINATE as conservative coalition were actually hurrah votes. On a hurrah vote both sides, the conservative coalition and northern Democrats, "win." Second, consider Table 7. There it can be seen that for actual conservative coalition votes, NOMINATE again accurately captures the percentage of victories.

Conclusion

Weisberg showed that a fair coin failed to reproduce the marginals and that a coin biased to the marginals failed to reproduce the vote occurrences. We have shown, in addition, that coin-flip models of either sort will not generate a distribution of liberal-conservative positions that is even remotely like the estimated distribution for Congress. In contrast, it is possible to reproduce these distributions with Monte Carlo simulations based on stochastic, spatial voting.

What is most striking is that a unidimensional spatial model with sincere voting can largely account for when coalitions occur and who wins when they occur. As mentioned earlier, our model cannot distinguish pure spatial voting from conscious but spatially adjacent logrolling. As such, our parameter estimates may reflect logrolling. The important point, though, is that we do not have to model logrolling to account for the coalition patterns.

The distribution of these patterns, however, depends on the distribution of roll-call midpoints. In other words, it depends on the agenda before Congress. Coalitions may well be important in setting the agenda. Analysis of agenda setting is probably the critical step in progressing to a predictive model from NOMINATE.

In summary, NOMINATE represents a relatively simple model that does an excellent job of recovering the marginals and the occurrences of the various types of votes. Earlier research (Poole and Rosenthal, 1985a) showed that, over all roll calls, it does much better than the standard two-party and three-party benchmarks. Consequently, we would advocate that the one-dimensional liberal-conservative model is now the appropriate benchmark for roll-call analysis. Since the spatial model outperforms both coin flipping and the standard two-party and three-party models, it should furnish a new standard of comparison for more complex models.

If the two-outcome spatial model is a standard of comparison, it is fair to ask what this new standard will mean for our understanding of decision making in Congress. Our short answer is constraint—the positions that legislators take on a wide variety of issues are systematically related (Converse, 1964). Given a legislator's position on the Nicaraguan contras and food stamps, the legislator's position on a test ban treaty with the Soviet Union can be forecast with a reasonable degree of accuracy. If voting in Congress were sincere and occurred over a multidimensional issue space, the presence of constraint means that a substantial majority of the issues lie on a low dimensional hyperplane through the space. What we are doing with NOMINATE is finding the best one-dimensional hyperplane through this issue space.

Why a high degree of constraint exists, as echoed by the NOMINATE results, is a question beyond this paper's agenda. We note, however, that constraint serves two important purposes in information processing. First, constraint facilitates cue taking by legislators themselves. On issues where legislators are poorly informed, they are able to take cues from adjacent legislators on the liberal-conservative continuum. Second, when constituents are poorly informed, they can learn whether a legislator voted "correctly" by seeing if the legislator voted like spatially adjacent colleagues.

The implication of constraint for roll-call voting is that a legislator's general liberal-conservative orientation—his/her position on the dimension–determines his/her positions on specific issues. Reality, of course, is much more complex than a spatial model and we do not claim that our model is reality. The claim we are making here is that it is very successful in modeling reality. Beyond that we will not go.

Keith T. Poole is Associate Professor of Political Economy and Howard Rosenthal is Professor of Political Science and Industrial Administration, Carnegie-Mellon University, Pittsburgh, Pennsylvania 15213.

NOTES

This work was supported by NSF grant SES-8310390. We thank Thomas Hammond for suggesting the coin flip experiments reported in the second section and for very extensive comments on a draft of the paper.

1. We show elsewhere (Poole and Rosenthal, 1985c) that the year to year configurations of senators and representatives are very stable. Consequently, we could use a previous year's configuration and "predict" voting for the current year, thereby cutting the number of estimated parameters to 2q—the same as the two-party model. However, as a reviewer pointed out to us, this comparison is somewhat misleading. In the current year the only information the two-party model requires is the party and region of the legislator, while the configuration from NOMINATE is produced from all the "information" contained in the previous year's roll calls. The number of estimated parameters would be the same for the current year, but the information needed to implement the estimation is different.

2. The estimates of NOMINATE are maximum likelihood estimates of a structural model. In contrast, the party models seek to minimize classification errors. The classifica-

tions of NOMINATE could in fact be improved somewhat were we simply to estimate legislator coordinates and midpoints that minimized classification errors.

3. It is not surprising that the geometric mean is less than the coin flip probability. As we noted above, the geometric mean penalizes large errors.

4. We also performed the computations outlined by Weisberg (1983, pp. 665-667). Our predicted percentages are somewhat higher than his since our yes vote marginals are less than his, a result of our excluding votes with less than 2.5% in the minority.

REFERENCES

Amemiya, Takeshi. 1981. Qualitative Response Models: A Survey. *Journal of Economic Literature,* 19:1483-1536.

Carroll, J. Douglas, and Jih-Jie Chang. 1970. Analysis of Individual Differences in Multidimensional Scaling via an N-way Generalization of "Eckart-Young" Decomposition, *Psychometrika,* 35:283-319.

Cherryholmes, Cleo H. and Michael J. Shapiro. 1969. *Representatives and Roll-Calls.* Indianapolis: Bobbs-Merrill.

Clausen, Aage R. 1973. *How Congressmen Decide: A Policy Focus.* New York: St. Martins Press.

Converse, Philip E. 1964. The Nature of Belief Systems in Mass Publics. In David E. Apter, ed., *Ideology and Discontent.* New York: Free Press.

Coughlin, Peter, and S. Nitzan. 1981. Directional and Local Electoral Equilibria with Probabilistic Voting. *Journal of Economic Theory,* 24:226-239.

Dhrymes, Phoebus J. 1978. *Introductory Econometrics.* New York: Springer-Verlag.

Fiorina, Morris. 1974. *Representatives, Roll Calls, and Constituencies.* Lexington, MA: Heath.

Hammond, Thomas H., and Jane M. Fraser. 1983. Baselines for Evaluating Explanations of Coalition Behavior in Congress. *Journal of Politics,* 45:635-656.

Kalt, Joseph P., and Mark Zupan. 1984. Amending the Economic Theory of Regulation: Ideology and Capture in Legislative Politics. *American Economic Review,* 74:301-322.

MacRae, Duncan Jr. 1958. *Dimensions of Congressional Voting.* Berkeley: University of California Press.

Matthews, Donald R., and James A. Stimson. 1975. *Yeas and Nays.* New York: Wiley.

Morrison, Richard J. 1972. A Statistical Model for Legislative Roll Call Analysis. *Journal of Mathematical Sociology,* 2:235-247.

Peltzman, Sam. 1984. Constituent Interest and Congressional Voting. *Journal of Law and Economics,* 27:181-200.

Poole, Keith T., and Howard Rosenthal. 1983. A Spatial Model for Legislative Roll Call Analysis. Social Science Discussion Paper, Pasadena: California Institute of Technology.

______. 1985a. A Spatial Model for Legislative Roll Call Analysis. *American Journal of Political Science,* 29:357-384.

______. 1985b. The Political Economy of Roll Call Voting in the 'Multi-Party' Congress of the United States. *European Journal of Political Economy,* 1:45-58.

______. 1985c. The Unidimensional Congress. Working Paper #44-84-85, Graduate School of Industrial Administration, Carnegie-Mellon University.

Takane, Yoshio, Forrest W. Young, and Jan deLeeuw. 1977. Nonmetric Individual Differences Multidimensional Scaling: An Alternating Least Squares Method With Optimal Scaling Features. *Psychometrika,* 42:7-67.

Warwick, Paul. 1977. *The French Popular Front*. Chicago: University of Chicago Press.

Weisberg, Herbert F. 1978. Evaluating Theories of Congressional Roll-Call Voting. *American Journal of Political Science*, 22:551-577.

______ . 1983. Alternative Baseline Models and Their Implications for Understanding Coalition Behavior in Congress. *Journal of Politics*, 45:657-671.