*A Spatial Model for Legislative Roll Call Analysis**

Keith T. Poole, *Carnegie-Mellon University*
Howard Rosenthal, *Carnegie-Mellon University*

A general nonlinear logit model is used to analyze political choice data. The model assumes probabilistic voting based on a spatial utility function. The parameters of the utility function and the spatial coordinates of the choices and the choosers can all be estimated on the basis of observed choices. Ordinary Guttman scaling is a degenerate case of this model. Estimation of the model is implemented in the NOMINATE program for one dimensional analysis of two alternative choices with no nonvoting. The robustness and face validity of the program outputs are evaluated on the basis of roll call voting data for the U.S. House and Senate.

Introduction

> *One way to try to account for political choices is to imagine that each chooser occupies a fixed position in a space of one or more dimensions, and to suppose that every choice presented to him is a choice between two or more points in that space . . .*
>
> *One of the most difficult problems of defining dimensions in this way centers about the operational definition of distance. . . Scales of the sort we have used . . . appear to define only an ordering relation rather than an interval scale . . . The definition of distance therefore marks a crucial gap between the model we shall propose and the data we have presented.*
>
> MacRae (1958, pp. 355-356)

This essay bridges MacRae's "crucial gap." Using solely the nominal data of observed political choices, we are able to estimate metric spatial distances. Our methodology estimates spatial coordinates for the choosers and the choices on the basis of observed choices. These methods can be applied to the analysis of voting in popular elections and other forms of political choice behavior when the choices form a finite set of alternatives. In this paper we develop the methodology for the simplest choice situation—a one-dimensional space with only two possible choices. We apply this methodology to voting in the U.S. Senate from 1979 through 1982 and the U.S. House in 1957 and 1958. The choosers are either representatives or senators, the choices are yea and nay on each roll call vote, and the observed choices are the recorded roll call votes.

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A long-standing (e.g., Rice [1924]) research method applied to roll call voting was to create a Euclidean representation of either the choices or the legislators. To create the spatial representations, various methods, such as factor analysis and nonmetric scaling, were applied in an essentially black box, statistical-method-driven fashion to measures of association, such as Yule's Q or ϕ/ϕ_{max}, computed between legislators or roll calls. (For examples, see Weisberg, 1968, and Warwick, 1977.)[1]

Over a decade ago, researchers began to realize that, if choice behavior is consistent with the elementary spatial model (Davis, Hinich, and Ordeshook, 1970), these methods would inaccurately recover the true underlying Euclidean coordinates. In particular, Morrison (1972) showed that measures of association based on the proportion of the total votes on which two legislators disagreed can serve neither as a general measure of angle nor as one of distance.[2] Since the black boxes assume inputs which are either distances or angles, they are unlikely to recover the true Euclidean space.

Independently, Weisberg (1968) presented a discussion similar to Morrison's and also covered roll-call-by-roll-call analysis. In addition, Weisberg addressed how error would affect the black box methods. In an errorless world, a legislator will always vote for the closest alternative, assuming sincere voting. That is, the legislator votes for the alternative with highest utility. But suppose these utilities are subject to error (perhaps from perceptual error or from omitted, idiosyncratic dimensions), so that the legislator no longer always chooses the closest alternative. In that case, citing an abundant psychometric literature, Weisberg shows that the black box methods will generally find a space with more dimensions than "truly" exist.

The problems that Weisberg and Morrison pointed out with the various multidimensional "black box" procedures also occur with Guttman scaling, a procedure even more widely used by political scientists. To see the relationship of Guttman scaling to spatial analysis, first assume a unidimensional space where the yea and nay alternatives are points on the continuum.[3] Assume further that each legislator votes without error for the alternative closest to his or her ideal point; that is, each legislator has a symmetric, single-peaked utility function over the dimension. In this case, the "cutting point" equidistant from the two alternatives for each roll call will divide the legislators into "left" and "right" camps, and one obtains a perfect "Guttman scale" even though the underlying dimension is not a true dominance scale.[4] When this occurs, we can never hope to learn anything about the spatial position of legislation since all pairs of alternatives with the same cutting point generate the same roll call be-

[1] Weisberg (1968) contains a comprehensive review of the literature up to 1968.

[2] Even when legislators always vote for the closest alternative, the proportion of disagreement depends upon both the distance between the two legislators, the angle they form with the (arbitrary) origin of the space, and the distribution of cutting lines of bills.

[3] While MacRae (1958) should be credited with the model that each roll call is *two* points on the continuum, his roll call analysis methods do not recover the points.

havior. We can, at least ordinally, identify the cutting points, but we can never, in this perfect world, learn where the alternatives are. Somewhat paradoxically, we need error to learn about the location of alternatives.

Now if there is error but only one "true" dimension and we insist upon Guttman scaling (or related techniques such as MacRae's [1970] *Q*-cluster analysis) not all the roll calls will form a single scale. In fact, as acknowledged by Clausen (1967, p. 1023) in his discussion of Lingoes' multiple scalogram analysis, we might well find several scales and conclude that there are multiple dimensions or issue areas when in fact only one exists.

When the true space is multidimensional, Guttman scaling will also exaggerate dimensionality for another reason. To see this, consider a two-dimensional space where choice is again without error and legislators vote for the closest alternative. Now yea and nay voters are separated by a cutting line; that is, the perpendicular bisector of the line joining the two alternatives. Draw any line through the space. All roll calls with cutting lines perpendicular to this line will form a perfect Guttman scale. These roll calls will generally not scale with roll calls whose cutting lines are not perpendicular to the line. As we try a variety of lines, we may find many "Guttman scales," although the space is only two-dimensional. When we have both error and multidimensionality, we have two effects that cause ordinary Guttman scaling to exaggerate the true dimensionality.

To summarize the preceding discussion, the multivariate black box methods are not based upon a spatial model of choice while ordinary Guttman scaling is based on a very limited model. Consequently, it is not surprising that traditional analyses often have to segregate the data by political party (MacRae, 1958, 1967), thus obscuring an overall picture of the legislature, or find a relatively large number of dimensions (Clausen, 1973).

While helping us to understand the perils of black boxes, Weisberg (1968) took a "least evil" approach in his dissertation. He sought to find which inputs would cause the fewest problems to the black boxes. In contrast, in his seminal piece, Morrison began the quest for a procedure that would be model-driven. By model-driven, we mean beginning with a model of individual choice behavior, then drawing the implications of the model for how such observed data as roll call votes will be generated, and finally developing methods for recover-

[4] What distiguishes the model we develop from classical Guttman scaling is that we assume a space composed of *proximity* dimensions rather than *dominance* dimensions. Guttman scaling or scalogram analysis was developed in the context of ability and trait testing and was later extended to attitude testing. Depending upon how the end points are defined, an item on the scale dominates all items to the left or right on the scale. Thus if you can work a difficult math problem you should be able to work an easier math problem or if you do not object to one of your children (if you are white) marrying a black then you should not object to sitting next to a black person on a bus, and so on. In terms of utility theory, individuals have monotonically increasing utility functions over a dominance dimension. In contrast, on a proximity dimension, individuals have single-peaked utility functions. The two models are functionally equivalent when there are only two choices, which is why Guttman scaling has been a popular methodology in research on roll call voting.

ing the unknown Euclidean coordinates from the observed data in a manner that is consistent with the underlying choice model. Morrison's approach was based upon very restrictive assumptions, such as error-free choice and a symmetric distribution of cutting lines.

In contrast to all of these earlier approaches, we here develop a method that derives from the basic spatial model of choice, allows for error, and makes no assumptions regarding the distribution of either legislator ideal points or the Euclidean coordinates of alternatives. Like the earlier analyses, we assume that the observations are independent across individuals and over time and that, on each roll call, sincere (in the usual sense of nonstrategic) voting prevails. Based on a model of probabilistic voting akin to Coughlin (1982) and Hinich (1977), our procedures permit *simultaneous* recovery of the Euclidean coordinates of both individuals and choices and the parameters of a utility function for the individuals. (In contrast, most conventional approaches do not place both the choices and the choosers in a common space.) The simultaneous recovery of coordinates for both legislators and choices is what distinguishes our work from previous research. One-dimensional coordinates for legislators can be gotten quite simply by using ADA scores. Even better estimates can be gotten by using metric unfolding on the ratings of a set of interest groups (Poole, 1981, 1984; Poole and Daniels, forthcoming [1985]). ADA scores and the coordinates recovered from a set of interest groups' ratings all have correlations above .9 with the coordinates we recover in our analysis of the 1979-82 Senates below. It is good that such widely different methodologies yield basically the same coordinates for legislators. However, our procedure not only produces these legislator coordinates, it also produces coordinates for the policy outcomes of the roll call votes. In psychometric parlance, we have developed an *unfolding* methodology for nominal level data.

A Unidimensional Spatial Model of Roll Call Voting

Along the lines of the spatial model of electoral competition, we assume that each legislator has a most-preferred position or ideal point in the unidimensional space. We further assume that each roll call is a choice between two points on the dimension – one point represents the outcome corresponding to a yea vote and the other point represents the outcome corresponding to a nay vote. The number of legislators is denoted by p, and the position of the ith legislator ($i = 1,\ldots, p$) is denoted by x_i. The number of roll calls is denoted by t, and the positions of the yea and nay outcomes are denoted by z_{yl} and z_{nl} ($l = 1, \ldots, t$) where "y" stands for yea and "n" nay.[5] The distance of the ith legislator to one of the outcomes of the lth roll call is therefore

$$d_{ijl} = \left| x_i - z_{jl} \right|, \quad j = y,n \tag{1}$$

[5] Several individuals have suggested that we consider an alternative model where each roll call is represented by a single point rather than a pair of points. Legislators close to the point vote "Yea" and legislators far from it vote "Nay." Such a model might occasionally apply to some congressional roll calls, such as those pertaining to final passage. One prediction of a "single-point" model is that we would observe, on some issues, the most liberal and most conservative legislators voting together. This in fact happens rarely in Congress. Empirically, a one-point model is undoubtedly easily outperformed by a two-point model.

Each legislator is assumed 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

$$U_{ijl} = \beta \exp\left[\frac{-\omega^2 d_{ijl}^2}{2}\right] + \varepsilon_{ijl} \tag{2}$$

where β and ω are parameters which we estimate, d_{ijl} is as given in (1), and the ε_{ijl} are the error terms which, for purposes of tractability, are assumed to capture both spatial and nonspatial aspects.[6] When there is no error, equation (2) is simply a normal distribution multiplied by a constant. We assume that the error terms are independently distributed as the logarithm of the inverse exponential (i.e., the logit or Weibull distribution; cf. Dhrymes, 1978, pp. 341–42). Our assumption of independence in this context means that the error a legislator makes on any particular roll call is (1) independent of the errors he/she makes on other roll calls; (2) independent of the errors other legislators make. The error distribution closely resembles the normal, and its use is without major consequence for the type of empirical work discussed here. Its great advantage over a normal distribution model of error is that the Weibull distribution function allows us to solve analytically for the probability a legislator votes yea/nay.

If $U_{iyl} > U_{inl}$ then legislator i votes yea on roll call l; conversely, if $U_{iyl} < U_{inl}$ the legislator votes nay.[7] Given the assumption that the ε_{ijl} have a Weibull distribution, the probability that legislator i votes yea/nay on roll call l is

$$P_{ijl} = \frac{\exp\left[\beta \exp\left[\frac{-\omega^2 d_{ijl}^2}{2}\right]\right]}{\varphi_{il}}$$

where (3)

$$\varphi_{il} = \exp\left[\beta \exp\left[\frac{-\omega^2 d_{iyl}^2}{2}\right]\right] + \exp\left[\beta \exp\left[\frac{-\omega^2 d_{inl}^2}{2}\right]\right]$$

Thus, the probability that a legislator votes yea/nay depends not only on how close the legislator is to the yea outcome but also how far apart the yea and nay outcomes are.

[6] Technically, spatial error should appear in d_{ijl} in the exponent term of (2). For example, in the case of perceptual error, an individual might use $z_{jl} + \alpha$, where α is the perceptual error, instead of z_{jl} to compute d_{ijl}. We avoided this complex specification in order to make the problem tractable. We do not think this is a serious problem, however. In our Monte Carlo work we found that the recovery of the x_i and the z_{jl} to be reasonably robust to a misspecification of the form of the utility function.

[7] Because the ε_j have a continuous distribution, equal utilities can be ignored.

The likelihood of the observed choices of the legislators is therefore

$$L = \prod_{i=1}^{P} \prod_{l=1}^{t} \prod_{j=1}^{2} P_{ijl}^{c_{ijl}} \tag{4}$$

where $c_{ijl} = 1$ if legislator i voted for outcome j on roll call l and 0 otherwise. Following standard practice, we obtain estimates of the parameters which maximize the logarithm of the likelihood function, namely

$$\ln L = \sum_{i=1}^{p} \sum_{l=1}^{t} \sum_{j=1}^{2} c_{ijl} \exp\left[\frac{-\omega^2 d_{ijl}^2}{2}\right] - \sum_{i=1}^{p} \sum_{l=1}^{t} \ln \varphi_{il} \tag{5}$$

To estimate β, ω, and the x_i and z_{jl}, we have developed the NOMINATE program, a constrained nonlinear maximum likelihood procedure (based in part on the method of Berndt et al. [1974]). Details of the procedure and Monte Carlo testing results can be found in Poole and Rosenthal (1983).[8] The next section provides a brief overview of the major theoretical issues we had to cope with in developing NOMINATE.

Theoretical Problems of Estimation

Perfect Roll Calls

In order to recover a Euclidean configuration of legislators and roll call outcome pairs from purely nominal (yea, nay) data, a number of theoretical problems must be dealt with in the estimation procedure. We briefly discussed the first of these problems, perfect roll calls, in the Introduction.

Assume the legislator positions are known. Say that on a given roll call every legislator to the left of a certain point on the dimension voted yea and every legislator to the right of this point voted nay. That is, suppose we observe

YYYYY•••YYYNNNN•••NNN

Then the midpoint of the yea and nay outcomes is clearly between the rightmost Y and the leftmost N. However, any pair of outcomes equidistant from the midpoint could have produced the observed pattern if there is no error.

If we observe a set of perfect or near perfect roll call responses and attempt to estimate outcome locations for fixed legislator locations and a fixed, stochastic utility function, we will estimate a midpoint corresponding to a Guttman scale cutting line. But where will we place the outcome coordinates? Clearly we will not place them close to the midpoint since all legislators would then be predicted to vote yea or nay with probability .5. Similarly, we will not place one outcome far to the left of the leftmost legislator and the other outcome far to the right of the rightmost legislator. Given the functional form of our utility function, all legislators would be close to indifferent between these two distant alternatives and would vote yea/nay with probabilities near .5. The likelihood function will be maximized by placing the yea and nay alternatives at an

[8] This paper is available from the authors on request.

intermediate solution that assigns a high yea probability to actual yea voters and similarly a high nay probability to nay voters.[9] But these placements are purely artifactual. In fact, for a given midpoint, recovered outcome locations for a perfect roll call will always be identical and do not depend on the true outcome locations.

It should be emphasized, however, that the problems that beset recovery of the outcome locations do not affect the ordinal recovery of the *midpoints*. In fact, NOMINATE recovers the correct ordering of not only the midpoints but also the legislators when there is no error. In other words, when the data contain no error, the algorithm will produce a perfect Guttman scale.[10]

When error is present, however, metric information can be recovered. This is so because the presence of error in effect constrains the placements of the coordinates of the legislators and roll call outcomes. From the basic model (2), legislators nearly midway between two outcomes are relatively likely to make errors in voting whereas legislators near one outcome but distant from the other outcome are unlikely to make errors in voting. The *pattern* of voting across a large set of roll calls thus tends to pin down the locations of the legislators precisely. Conversely, the *pattern* of voting across a large set of legislators tends to pin down the locations of the outcomes—and especially the midpoint—precisely.

Midpoints are more precisely estimated than the outcome pairs. An extreme illustration of this point was previously presented in the case of perfect roll calls. There, midpoints can be identified (at least up to a monotone transformation) while the outcomes cannot be. More generally, different patterns of voting can be associated with the same midpoint but produce different pairs of outcome estimates. For example, suppose that the legislators are uniformly

[9] This problem is one of the reasons why we do not utilize a quadratic utility function as in Poole and Rosenthal (1984). It can be shown that the outcome locations for perfect roll calls cannot be identified with a quadratic utility function. In addition, we think that a quasi-concave utility function is more realistic behaviorally. Finally, as a practical matter, the estimated β's and ω's in our empirical analyses result in utility functions which are for the most part concave over the length of the recovered dimension.

[10] Although we pointed out earlier (n. 4) that the motivation underlying Guttman scaling is quite distinct from our choice model, the fact that NOMINATE produces a Guttman scale in the case of errorless voting shows that, in a technical sense, Guttman scaling is a special case of NOMINATE. Since Guttman scaling is also known to be a special case of latent structure analysis (Lazarsfeld, 1950), it is interesting to outline the relationship of NOMINATE to latent structure analysis. Like latent structure analysis, NOMINATE assumes an underlying continuum (liberal-conservative). Also like latent structure analysis, we assume that, conditional on position on the latent continuum (x_i), the probability of a "Yea" vote on a particular roll call is statistically independent of the probability of a "Yea" vote on any other roll call. The probability of a "Yea" vote on a particular roll call as a function of x_i, that is, the trace line in latent structure analysis, is given by equation (3). This is distinct from latent structure analysis where the trace lines are polynomial functions of x_i. Our assumed trace lines and error distributions have led to an effective strategy for computing the latent continuum in the form of x_i values. In contrast, computation

distributed across the dimension and we observe the following two roll calls:

YYY • • • YYYYYNYNY | NYNYNNNNN • • • NNN
YYY • • • YYNNYNYNY | NYNYNYYNN • • • NNN

The same midpoint will be estimated for both roll calls, but the outcome locations will be recovered closer to the midpoint in the second roll call than in the first roll call. This is because in the second roll call the voting error is more dispersed around the midpoint so the two outcomes must be closer together to force the probabilities nearer to .5 over a broader range than in the first roll call where the voting error is concentrated near the midpoint.

Random Roll Calls and Extreme Placements

The analysis above cannot be continued indefinitely. At some level, there is so much error that it is meaningless to think of it as dispersed around a midpoint with the outcome locations being recovered ever closer together to account for the dispersal. For example, assume that on a given roll call the yea and nay alternatives were identical. Then, in our model, legislators would be effectively flipping coins to make vote decisions. Moreover, any common pair of outcome locations would lead to this behavior. When the observed responses appear as randomly distributed along the dimension, our estimation method will find it difficult to identify outcome locations. It will either put the alternatives close to each other at a variety of locations, including locations outside the range of legislators, or, if unconstrained, make the alternatives very distant from one another.

Even if our model is correct – roll call voting is neither perfect nor purely random – ill-behaved roll calls can arise as a matter of chance, since the utility function has a stochastic as well as a deterministic component. When they do arise, attempts at strict maximum likelihood estimation of these ill-behaved roll calls can result in coordinate estimates that are far from the limits of the space defined by the legislators. Political theory, however, suggests that one alternative should always lie within the space of legislators and that the midpoint should also fall within this space. NOMINATE imposes these constraints. Coordinate estimates for those roll calls with constraints imposed should, however, be viewed as unreliable.

Perfect Legislators

One can conceptualize a legislator who is similar to a perfect roll call. This individual always votes liberal on roll calls with midpoints to his/her right and

of a continuous polynomial latent structure has generally been attempted only for linear polynomials (Lazarsfeld, 1961). Linearity would imply that the probability of a "Yea" vote is monotonically increasing or decreasing in spatial position, clearly an unreasonable assumption in a legislative setting. Most latent structure analysis has in fact dealt not with a latent continuum but with the more restrictive situation where a discrete number of latent classes are assumed (Lazarsfeld and Henry, 1968).

conservative on those roll calls with midpoints to his/her left. That is, we would observe:

CCCCCCCCCCLLLLL • • • LLLL

This legislator would be located between the rightmost C and the leftmost L and is easily identified. Thus a perfect legislator is much like the midpoint of a perfect roll call. However, if a legislator always votes liberal or always votes conservative, then the legislator is like a unanimous roll call and the legislator's position cannot be identified. For a perfect liberal, all we know is that this legislator is to the left of all the midpoints. As a consequence of this identification problem, we will obtain relatively imprecise estimates of the locations of legislators at the periphery of the space. In particular, one can obtain a large gap between the positions of the leftmost and second leftmost legislators. A similar situation can hold at the right end of the dimension. As noted above, however, this is not a problem for the interior legislators.

Let us summarize the discussion of this section. When the error level is very high, approaching random voting, we will find it difficult to estimate the legislator and roll call coordinates. When the error level is very low, approaching perfect spatial voting, we will find it difficult to estimate the location of the alternatives, but we can readily order the legislators and midpoints. We will find it more difficult to estimate the location of legislators (and midpoints) in extreme positions than to estimate coordinates in the center of the space.

Use of NOMINATE in Practice

Having discussed the theory of nominal unfolding of choice data, we turn to a presentation of the implementation of our methodology in the NOMINATE program. Subsequent sections then compare the results to those developed by alternative methodologies, discuss the substantive validity of the results, and summarize extensive testing of our procedure.

Prior to beginning the estimation, pairs and announced positions are recoded as yea and nay votes. Absences, which are generally few in number and usually without political significance, are treated as missing data. Thus the input to NOMINATE is a matrix of roll call votes where the rows are legislators and the columns are roll calls. Our largest single run of NOMINATE has been for the House of Representatives in the 85th Congress where we estimated the parameters of the utility function, coordinates for 440 representatives, and 344 coordinates corresponding to 172 roll calls. The data set contains 68,284 individual voting choices (440 × 172 − missing data). As it is impractical to estimate nearly 800 parameters simultaneously, we first estimated the utility function parameters, then the legislators, and then the roll call parameters. The NOMINATE acronym thus denotes *Nominal Three-step Estimation*. These successive estimations define a global iteration. After initial generation of starting values, global iterations continue until the parameters from one iteration correlate at a user-defined level with those from the previous iteration. We use the

.99 level. After each global iteration, the legislator space is normalized to be two units in length, with the most liberal legislator at −1 and the most conservative at +1. Alternating algorithms of this type are common in psychometric applications (e.g., Carroll and Chang, 1970; Takane, Young, and de Leeuw, 1977).

Use of the alternating, three-step procedure introduces important economies. First, in the utility function phase, only two parameters are estimated, β and ω, while the roll call coordinates, (z_{jl}) and the legislator coordinates (x_i) are held fixed. The whole matrix of roll calls is used to estimate the utility function parameters. We found that, while some values of ω are clearly bad, β and ω were highly collinear in the neighborhood of global convergence. Consequently, we have included an option to preset either parameter. All our results in the next section are based on $\omega = 1/2$. Second, when the legislator coordinates are estimated and the other parameters are held constant (β, ω, and the z_{jl}), each legislator's estimates are independent of all others, so we can estimate one legislator at a time. In fact, we found that, conditional on all other parameters being held constant, the likelihood function is empirically globally convex in the legislator coordinate, which leads to very rapid convergence.[11] For each legislator, every roll call for which he/she had a recorded position was used in the estimation of x_i for that legislator. As a general rule, the estimate of x_i is most determined by roll calls with widely separated coordinates. To see why this is so, consider the extreme case where $z_{yl} = z_{nl}$. Since all probabilities must be .5, the likelihood will not vary with x_i. In practice, there were only a few roll calls with near identical outcome coordinates. Third, similar to the legislator situation, we can do the roll calls one at a time. For each roll call, every legislator with a recorded position was included in the estimation of z_{yl} and z_{nl}.

Although there are only two parameters per roll call, the roll call phase is the most difficult part of the estimation. Even the conditional likelihood is not globally convex, and the problems of (near) perfect and (near) random roll calls are encountered. NOMINATE contains heuristics to deal with these problems.

To use NOMINATE, one traditional and critical decision must be made. One must fix a cutoff level, in terms of minority voting, that determines whether a given roll call is included. This involves an important tradeoff. If the cutoff level is a high one, so that, for example, roll calls with 10 percent or fewer of the legislators voting in the minority are excluded, this tends to create many perfect legislators. The high cutoff levels don't allow for enough differentiation between the most conservative legislators and between the most liberal legislators. So high cutoff levels worsen the legislator estimates.

On the other hand, very low cutoff levels, say 1 percent and below, lead to poor roll call estimates. This is because roll calls with low minorities tend to be noisy. Noisiness implies a high level of error, which implies a low value of β. To see this point, note that a fully equivalent representation of the model (2) is:

[11] This is true within the constrained space $[-1, +1]$.

$$U^*_{ijl} = \exp\left[\frac{-\omega^2 d^2_{ijl}}{2}\right] + \frac{\varepsilon_{ijl}}{\beta} \tag{2'}$$

so that the magnitude of the error is now ε_{ijl}/β. For errors to increase in magnitude, β must decrease.[12] Thus, when the cutoff level is lowered, the lower value of β forces an adjustment in the roll call coordinates for those less noisy roll calls above the previous cutoff level. The outcome coordinates tend to drift off the ends of the dimension, forcing greater use of the heuristic constraints.

Use of the heuristic constraints in fact appears more frequently with actual data than would be suggested by our Monte Carlo studies that generate data as if our one-dimensional, stochastic model were "truth." This could reflect the need for a multidimensional model.[13] What to us is a more parsimonious model would allow for variation in error across roll calls—some roll calls are better perceived than others—and for variation in legislator utility functions—extremists might have more sharply peaked functions than moderates. We can in fact incorporate variation across both roll calls and legislators by adding just two parameters to our total of several hundred. Such a change eliminates most of the ill-behaved estimates. To simplify presentation, however, we defer development of both a unidimensional model with variation and multidimensional models to later papers. For the model used here, we have found a cutoff level of 2.5 percent, used throughout the next section, to be a good tradeoff between the quality of legislator coordinates and the quality of roll call coordinates. The simple, unidimensional, common utility function model successfully accounts for the data, as we now proceed to demonstrate.

Applications to Congressional Roll Call Voting

In this section, we conduct a variety of analyses to argue the substantive validity of our approach. First, we compare our analysis to earlier analyses of the U.S. House for 1957-58. The comparison returns us to two related points made earlier: (1) that more than one Guttman scale can be recovered from a single dimension when voting error occurs; (2) that previous techniques tend to find too many dimensions. Second, through an analysis of the classification of individual votes, we indicate why it is important to locate midpoints and individual legislator coordinates in the analysis of roll call voting. Third, since the major innovation of our method is the estimation of outcome coordinates and not just the legislators and midpoints, we briefly discuss the substance of roll calls with similar midpoints but different liberal outcome estimates. Fourth, we develop the geometric mean probability as an alternative to simply counting prediction errors in assessing the results. We use the geometric mean to in-

[12] The quantity β also controls the maximum choice probability. This probability is simply $e^{\beta}/(e^{\beta} + 1)$.

[13] On the other hand, Poole and Daniels (forthcoming [1985]) report that only 3 percent more of the votes are correctly classified when a two-dimensional interest group scaling is compared to a one-dimensional scaling.

terpret our results for legislators and roll calls. Fifth, we examine the relationship between the position of the median legislator and mean roll call coordinates in the light of an elementary spatial model of legislative behavior. Finally, we examine the intertemporal stability of our unfolding. This analysis suggests that omitted dimensions, if any, are not stable temporally.

The House of Representatives in the 85th Congress

To compare NOMINATE to standard methods of roll call analysis, we conducted an analysis of voting in the U.S. House of Representatives during the 85th Congress (1957-58). Roll calls from this Congress were analyzed by Guttman scaling in a well-known paper by Miller and Stokes (1963) and later were subjected to a careful application of a variety of methods by Weisberg (1968).

Our analysis was based on essentially all the relevant data. Eliminating only roll calls with less than 2.5 percent of the House on the minority side, we analyzed 172 of the 193 recorded roll calls in this Congress. Our one-dimensional model correctly classified 78.9 percent of the 68,284 individual votes. While there are undoubtedly multidimensional issues, roll call voting behavior can be largely accounted for by this one-dimensional liberal-conservative pattern, whereas, as argued in the Introduction, traditional techniques have exaggerated dimensionality.

Miller and Stokes used three scales based on 21 roll calls: *social welfare, foreign policy,* and *civil rights.* As shown in Table 1, the midpoints calculated by NOMINATE reproduce exactly the item order in the social welfare scale. The same is true for the foreign policy scale, although one item has an unreliable constrained estimate. Thus, both foreign policy and social welfare can be thought of as liberal-conservative issues although, by the usual criteria of coefficients of reproducibility, the foreign policy and social welfare scales could not be combined into one large Guttman scale. The two sets of items fit into the liberal/conservative dimension, however, if we allow for errors in individual voting.

In contrast, the civil rights scale does represent a separate dimension. The recovered midpoints tend to extreme, unreliable values, and the rank order of the scale is not recovered. So although civil rights must be treated separately, we can regard social welfare and foreign policy as part of the same basic dimension. This result is partially echoed by Weisberg's (1968, p. 208) factor analysis of these 21 roll calls. Although he found three factors and although the civil rights items load distinctly on the first factor, there is less clear-cut separation of the foreign policy and social welfare items on the second two factors.

We suspect many other "issue scales" can be mapped onto the basic liberal-conservative dimension. For example, Weisberg formed an 11-item mutual security scale using 3 of the Miller-Stokes foreign policy items and 8 other roll calls. As shown in Table 2, midpoints along our dimension perfectly reproduce the ordering of this scale.

TABLE 1

The Miller-Stokes Scales

Bill	Year	*CQ* Number	Content	Liberal-Conservative Midpoint
			Social Welfare	
HR 9955	1958	3	Passage of $5 billion debt limit.	0.35
HR 675	1958	73	Open rule for National Defense Education Act.	0.34
S 4035	1958	80	Passage of Housing Act under rules suspension.	0.23
HR 13247	1958	74	Recommit Defense Education Bill.	0.20
HR 682	1958	78	Open rule for depressed areas aid.	0.13
S 3683	1958	79	Recommit depressed areas resolution.	0.07
HR 1	1957	56	Kill School Construction Bill.	0.06
HR 6287	1957	17	Cut Labor Department appropriation.	0.04
HR 6287	1957	19	Cut unemployment funds for federal employees.	−0.08
HR 6287	1957	20	Cut Mexican farm labor program funds.	−0.48
			Foreign Policy	
HR 8922	1957	70	Congress not required to approve International Atomic Energy Agency transfers of fissionable materials.	0.61
HR 13192	1958	56	Passage of Mutual Security Act.	0.29
S 2130	1957	79	Adoption of conference report on mutual security.	0.17
HR 6871	1957	30	Cut funds for international organizations.	0.12
HR 9302	1957	81	Recommit Mutual Security Act to restore cut.	−1.00
			Civil Rights	
HR 6127	1957	96	Adopt jury trial provision.	0.96
HR 6127	1957	95	End debate on amendment involving jury trial.	0.79
HR 6589	1958	22	Civil Rights Commission appropriation.	1.00
HR 259	1957	40	Open rule for debate of Civil Rights Bill.	0.42
HR 6127	1957	42	Passage of Civil Rights Act.	1.00
HR 6127	1957	41	Recommit to modify jury trial provision.	−0.02

NOTE: Description of the bills taken from Weisberg (1968).

TABLE 2

Mutual Security Votes

Bill	Year	*CQ* Number	Content	Liberal-Conservative Midpoint
HR 9302	1957	82	Passage of Mutual Security Appropriations of 1958.	0.29
HR 13192	1958	56	Passage of Mutual Security Appropriations of 1959.	0.29
HR 12181	1958	31	Passage of Mutual Security Act of 1958.	0.28
HR 12181	1958	51	Recommit Mutual Security Act of 1958 to conference.	0.27
S 2130	1957	54	Passage of Mutual Security Act of 1957.	0.26
HR 9302	1957	100	Conference report on 1958 appropriations.	0.21
S 2130	1957	79	Conference report on 1957 Mutual Security Act.	0.17
S 2130	1957	53	Recommit 1957 Mutual Security Act to delete Development Loan Fund.	0.16
HR 13192	1958	55	Recommit 1959 Appropriations to increase defense funds.	−0.53
HR 9302	1957	81	Recommit and cut 1958 appropriations.	−1.00

NOTE: Description of the bills taken from Weisberg (1968).

After forming the mutual security scale, Weisberg conducted a factor analysis of 26 foreign policy items, including the mutual security roll calls. He found four factors. We have found that all these items generally fit well on the liberal-conservative dimension except for three votes on the Eisenhower Mid-East doctrine that load highest on Weisberg's third factor. Except for the Mid-East issue, foreign policy votes can be interpreted within the liberal-conservative dimension since NOMINATE represents a model that allows for the "errors" captured in three factors by Weisberg.

Weisberg also performed a factor analysis on the 140 roll calls with a minority of at least 15 percent. He found 5 factors. He found 10 factors in an alternative analysis of only 97 of these roll calls. Finally, performing a cluster analysis, Weisberg found 14 scales, although these scales could account for only 54 of the 140 roll calls. Judging from our analysis of the civil rights, social welfare, foreign policy, and mutual security Guttman scale items, we believe that all of these results overemphasize the dimensionality of congressional voting.

Classification of Voting Outcomes

Our analysis of midpoint locations for the 1957-58 House has suggested that a one-dimensional spatial model that allows for error may reduce much of the multidimensional complexity that existed in previous statistical roll call analyses. We now provide some indication of the importance in this model of two key elements: roll call midpoints and legislator coordinates.

A method frequently used to assess probabilistic, binary choice models is to predict the choice assigned the higher probability. [14] For our model, this is equivalent to predicting that, if the legislator is to the left of the midpoint, the legislator's vote is liberal, while the vote is conservative if the legislator is to the right of the midpoint. The results of applying this maximum probability approach can be seen in the "midpoint" column of Table 3. We typically correctly classify over 80 percent of the individual votes.

TABLE 3

Percentage of Individual Votes Correctly Classified

Year	Model						Totals	
	Majority	Party	(*N* Dem.)	Liberal-Conservative	(*N* Lib.)	Midpoint	*N*	Votes
				House				
1957-58	65.6%	67.3%	(238)	70.2%	(260)	78.9%	440	68,284
				Senate				
1979	68.2%	66.5%	(58)	70.1%	(66)	80.3%	100	40,554
1980	68.7%	66.5%	(59)	69.4%	(66)	80.6%	101	41,234
1981	67.3%	71.4%	(46)	72.2%	(43)	83.2%	100	37,175
1982	68.9%	66.4%	(45)	68.2%	(52)	81.7%	100	39,572

NOTE: Figures in this table refer to *all* roll calls with at least a 2.5 percent minority. Total *N* can exceed size of House or Senate as a result of replacements.

We can eliminate use of the estimated midpoint from the prediction in the following manner. For each roll call, we can identify the liberal outcome and the conservative outcome. For each legislator, we can then compute, over all roll calls, whether the legislator more frequently votes on the liberal side or on the conservative side. Ignoring the midpoints, we predict that liberals always vote liberal and conservatives always vote conservative. The "Liberal-Conservative" column of Table 3 shows that we now correctly classify only about 70 percent of the individual votes. So identifying the midpoint of the roll call

[14] Ties can be dealt with, say, by random assignment.

in addition to the location of the legislator reduces classification errors by about 10 percent.

In a similar fashion, we can eliminate the use of the individual legislator location information by making identical predictions for all legislators of the same political party and predicting that Democrats always vote liberal and Republicans always vote conservative. The results of this exercise, shown in the "Party" column, show further deterioration in classification ability.[15]

Locating the Yea and Nay Coordinates for Individual Roll Calls

While NOMINATE can be seen to provide accurate locations for the midpoints and legislators, to some degree these tasks have been accomplished by older methodologies. Where NOMINATE provides a distinctly new methodology is in its location of the yea and nay or "liberal" and "conservative" coordinates for each roll call. Use of these coordinates can potentially prove useful in the analysis of policy outcomes or of congressional agenda strategy. Poole and Smith (1983), for example, have compared the coordinates of amendments to the coordinates of the senators introducing the amendments. In this paper, we briefly suggest that these coordinates have face validity.

A relevant illustration is found in Table 4. There we list six 1981 Senate roll calls whose midpoints are all in the center of the space but whose liberal (and conservative) coordinates vary over the entire space. Highly polarized coordinates occur for visible, broadly ideological issues of the social welfare or economic liberal-conservative variety such as the two votes on taxes, a central theme in the 1981 Congress. Coordinates occur close together for geographic distribution issues that are not broadly liberal-conservative. Voting on these issues appears nearly random, as in the peanut case, and the coordinates are not differentiated, leaving all senators with voting probabilities close to .5. Not surprisingly, an issue that combines both rich-poor redistribution and geographic distribution, such as subsidies for heating to low-income families, has coordinates that have an intermediate degree of separation. It can be seen that as the percentage of classification error falls, the coordinates generally become more separated. This separation raises the estimated probability of voting liberal for liberals and conservative for conservatives. With near perfect roll calls, raising these probabilities increases the likelihood.

[15] We have also compared our classification with those of the two-party and three-party baseline models developed by Weisberg (1978). The differences here are minor. We typically only classify about 1 percent more of the votes correctly than does the three-party model and 3 percent more than the two-party model. Much of the reason for the small improvement lies in the fact that the two-party and three-party models are estimated with the direct objective of minimizing classification errors. One predicts each senator will vote with the majority of his/her party. As shown in Poole and Daniels (forthcoming [1985]), estimating a one-dimensional spatial model of senator locations and midpoint locations with the objective of minimizing classification errors has about 3 percent fewer errors than NOMINATE. The subsequent text provides further discussion of why, as a maximum likelihood approach, NOMINATE does not seek to minimize classification errors.

The relationship of classification error to locations, however, is not perfect. Observe that the vote to table the Hart amendment has less separated coordinates than the Metzenbaum amendment even though there are slightly fewer classification errors for the Hart amendment. This is because, as discussed in section 3, the pattern of votes informs us about the coordinates. The errors on the Metzenbaum amendment were concentrated in the middle of the space, indicating that the roll call coordinates should be widely separated. The errors on the Hart amendment, while few in number, were more widely dispersed, leading to more centrally located coordinates.

TABLE 4

Six 1981 Roll Calls with Midpoints Near Zero

ICPSR Number	Content	Error Percent	Geometric Mean Probability	Coordinate Liberal	Coordinate Conservative
232	Riegle Amendment to Tax Bill	9	.78	–.94	+.99
113	Metzenbaum Amendment on Commodity Tax Straddles	13	.76	–.82	+.78
595	Motion to Table Hart Amendment Halting Libyan Oil Imports	10	.71	–.65	+.67
84	Biden Amendment on Low-Income Fuel Assistance	15	.71	–.60	+.65
275	Dole Motion to Table Zorinsky Amendment on Commodity Programs	27	.61	–.38	+.34
269	Mattingly Amendment on Peanut Supports	47	.51	+.03	+.19

The results in Table 4 indicate a general pattern. The liberal-conservative dimension generally does poorly on those pork barrel, regional, and special interest issues that will always lie outside of any low-dimensional spatial model. These include tobacco subsidies, solar power in California, the Tombigbee waterway, pay for members of Congress and the federal Civil Service, Amtrak service, D.C. airports, Mt. St. Helens relief, etc. Such roll calls either have coordinates that are quite close to each other or have constrained estimates. In contrast, votes on the key policy issues of each session generally show strong separation of voting along the dimension. In addition to the budget cuts and tax bill under Reagan, we find that the windfall profits tax and the Taiwan issue gave rise to widely separated coordinates in 1979 as did the Federal Trade Commission and other votes in 1980 that eroded the welfare spending and regulation of the previous decade. Such social control issues as abortion and school prayer typically occupy more intermediate positions.

There are also a few, striking, essentially unscalable votes with constrained estimates. These occur when members of the majority party are cross-pressured between ideology and support for the President. Examples are MX in 1979, the draft in 1980, and sugar subsidies in 1981. As can be seen from the success with which NOMINATE classified individual votes, such roll calls are rare. On the whole, our procedure appears to give a sensible Euclidean representation of the roll calls.

Evaluation of the Results by a Probabilistic Measure

Until now, we have used classification errors as a vehicle for evaluating NOMINATE. While readily interpretable, classification errors are not a fully satisfactory means of evaluating a probabilistic model like ours. NOMINATE in fact seeks not to minimize ex post prediction errors but to estimate the parameters of a structural model. The reason that the midpoints chosen by NOMINATE are not the error-minimizing ones is that NOMINATE essentially weights errors in terms of distance. To see this point, consider the estimated senator coordinates shown in Table 6. Assume Heinz votes liberal on a roll call and Kennedy votes conservative. To maximize the likelihood, we may move a midpoint from slightly to the right of Heinz to slightly to the left. This move creates a classification error, but it may raise the probability of Kennedy's conservative vote far more than it lowers the probability of the liberal vote by Heinz. Put somewhat differently, rather than stating that a legislator will definitely vote yea on a certain roll call, our model only states that the legislator will vote yea with a certain probability. The sum of the logarithms of all these probabilities is the log likelihood, $\ln L$, which is expressed as equation (5). The log-likelihood statistic itself, while useful for certain hypothesis tests, is not useful as a descriptive statistic. Its value is a function of the number of legislators and roll calls so that two analyses are not comparable unless p and t are the same. The average log likelihood is better but not easily interpretable. Instead, we use the *geometric mean probability* which is calculated by taking the exponential of the average log likelihood; that is

$$\bar{P} = \exp(\ln L/A)$$

where A is the total number of choices made by all legislators on all roll calls.[16] It should be noted that $\bar{P}$ is a "conservative" statistic and is always less than the mean probability of the actual choices. It "penalizes" actual choices with

[16] We can also compute a geometric mean for an individual legislator by dividing the legislator's contribution to the likelihood by his total recorded votes.

low probabilities.[17] Thus, if one vote occurred with probability .9 and another with probability .1, the geometric mean would be $\exp\{[\ln(.9) + \ln(.1)]/2\} = .3$, not .5.

Using $\bar{P}$ generally gives results quite similar to the use of classification error, as can be seen in Table 4. For a given midpoint, the geometric mean tends to increase with separation of coordinates. In fact, as can be seen in Table 5, the geometric means are higher for roll calls that are "unscalable" and have constrained coordinates than they are for roll calls whose coordinates both fall inside the dimension (first versus eighth columns). The constrained roll calls are of two types. One type has a well-defined midpoint, but the outcome coordinates are constrained. In this case, we have a near perfect roll call problem and cannot identify the outcome coordinates, but the geometric mean is quite high. The second type has a midpoint constrained to one end. These represent random roll call problems, but many of these represent lopsided votes of the 97-3 variety. Here randomness implies that we model each legislator as flipping an *un*fair coin. With lopsided votes, the geometric mean will still be high.

Both types of constrained roll calls arise less frequently in the House estimation than in the four estimations for the Senate, as seen in Table 5. From the viewpoint of our model, this result is explicable from a simple sample size argument. With 435 voters against 100, a given stochastic realization is far less likely to generate a pattern of votes that looks nearly perfect or nearly random. Although the larger sample size diminishes the constrained roll call problem in the House, the general pattern is true in both Houses: high geometric means are associated with extreme placement of at least one of the two outcome coordinates.

The relationship of geometric means to spatial position is at least as evident for legislators as it is for roll calls. In fact, the plot of the geometric means of legislators versus the spatial positions of the legislators discloses a tight V shape. Legislators in the middle have low geometric means, near .5; legislators at the ends have geometric means near .8. Similarly, we make far fewer classification errors with the extreme legislators. This result is consistent with simple ideas of competition within the legislature. Most midpoints will fall near the center of the legislature. Legislators close to these midpoints will be less predictable. On the other hand, when compared to the various benchmark models, our model makes the most difference for these legislators. Whereas Kennedy and Helms are almost as predictable as the tides, we make substantial improvements at the center.

Spatial Behavior in the Aggregate

In a unidimensional legislature with probabilistic voting, majority leadership should plan votes such that midpoints lie somewhat away from the median voter. By moving a slight distance away from the median voter, the probability of passage can be increased substantially. Thus, when the Democrats

[17] For further approaches to summarizing the results of logit estimation, see Amemiya (1981).

TABLE 5

Summary of Estimates

Year	Geometric Mean Probability[a]	Median Legislator: Name	Median Legislator: Coordinate	Mean Coordinate of Legislators[b]	Mean Midpoint	Mean Liberal Coordinate[b]	N^c	Geometric Mean Probability[d]
				House				
1957-58	.652	Corbett	+.07	+.02	+.15	–.34	172/132	.642
				Senate				
1979	.666	Bentsen	–.05	–.03	+.03	–.48	448/346	.654
1980	.664	Hollings	–.06	–.10	–.01	–.56	480/320	.638
1981	.692	Heinz	+.33	+.21	+.05	–.47	397/249	.657
1982	.673	Proxmire	+.26	+.22	+.14	–.42	421/244	.637

[a] This geometric mean was calculated using all roll calls (first number in *N* column).
[b] Unconstrained roll calls only.
[c] The first number is the total number of roll calls. The second number is the number estimated without constraints.
[d] This geometric mean was calculated using only those roll calls which were not constrained (second number in *N* column).

control a house of Congress, the mean midpoint should be to the right of the median legislator; when the Republicans control, it should be to the left. As Table 5 shows, the empirical results correspond with this spatial model.

If we define a liberal as a legislator with a majority of liberal votes, then when the Democrats are in control, as in the House and the first two years for the Senate, midpoints are chosen sufficiently far to the right that the number of liberals actually exceeds the number of Democrats. Two influences appear to be relevant to this pattern. On the one hand, legislation is tailored to be acceptable to moderate Republicans; on the other, these legislators "go along to get along." The same type of pattern prevailed when the Republicans took control of the Senate in 1981. Now the number of conservatives (a legislator with a majority of conservative votes) exceeded the number of Republicans. But the pattern failed to hold in 1982. Although the *mean* midpoint fit the expected pattern, as shown in Table 5, the entire distribution of midpoints was such that liberals slightly outnumbered conservatives. We attribute this to the presence of a large number of roll calls not oriented to the passage of legislation but to allowing social conservatives to be counted.

Intertemporal Stability of Scaled Positions

In the discussion above, we examined how well the data corresponded to some very simple spatial ideas of how the leadership would place typical votes that come before a legislature. Spatial theory also says that legislators will adjust to perceived changes in opinion. Having conducted a separate estimation for each of four years of Senate voting, we can't yet study dynamics of this type in an *absolute* sense. But we can study it in a relative sense. We can ask whether a senator is closer to the conservative *end* of the dimension (whatever that end may represent) in 1981 than he or she was in 1979. And we can see how stable is the placement of the senators on the main dimension, whatever the meaning of the dimension in various years.

Studies of Congress have stressed that, perhaps in disagreement with spatial theory, most members of Congress do not make important changes in their voting patterns (e.g. Bullock, 1981; Kuklinski, 1979; Fiorina, 1974; Clausen, 1973). The coordinate estimates for senators for all four years are presented in Table 6, where a generally stable pattern can be observed. Consider the 80 senators who served in all four years. The squared correlation of their positions in 1982 with their positions in 1981 is .95; with 1980, .90; and with 1979, .88. The 1981 positions have a squared correlation of .85 with 1980 and .83 with 1979. Finally 1980 and 1979 have a squared correlation of .95.

In studying the stability of roll call behavior, we note than senators can vary not just in their spatial location on the dimension but in how well this location in fact explains their voting behavior. Thus, another important indication of stability comes from the analysis of geometric mean probabilities. We noted above that there was a V-shaped relationship between geometric means and coordinates. Deviations from this relationship indicate senators who are more or less predictable than is normal for their position. If senators are syste-

TABLE 6

Liberal-Conservative Positions of U.S. Senators

	1979	1980	1981	1982		1979	1980	1981	1982
Kennedy, E	−1.000	−0.939	−0.942	−1.000	DeConcini, D	0.007	−0.167	−0.055	0.127
Tsongas, P	−0.761	−0.864	−0.680	−0.684	Stone, R	0.011	−0.129		
Bradley, W	−0.680	−0.861	−0.513	−0.497	Morgan, R	0.016	0.033		
Williams, H	−0.670	−1.000	−0.616	−0.272	Johnston, J	0.021	−0.071	0.154	0.212
McGovern, G	−0.638	−0.581			Long, R	0.021	0.008	0.275	0.252
Metzenbaum, H	−0.610	−0.616	−0.756	−0.738	Stennis, J	0.022	−0.008	0.265	0.284
Sarbanes, P	−0.609	−0.891	−0.789	−0.866	Durenberger, D	0.035	0.004	0.457	0.166
Levin, C	−0.604	−0.702	−1.000	−0.696	Proxmire, W	0.052	−0.074	0.230	0.265
Riegle, D	−0.595	−0.692	−0.727	−0.779	Danforth, J	0.069	0.157	0.558	0.448
Culver, J	−0.581	−0.887			Heflin, H	0.075	−0.041	0.062	0.182
Nelson, G	−0.530	−0.711			Cohen, W	0.102	0.160	0.441	0.304
Ribicoff, A	−0.527	−0.648			Bellmon, H	0.131	0.237		
Moynihan, P	−0.494	−0.673	−0.533	−0.467	Pressler, L	0.132	0.242	0.343	0.482
Dodd, C, Jr.			−0.866	−0.494	Specter, A			0.321	0.147
Cranston, A	−0.487	−0.756	−0.686	−0.689	Zorinsky, E	0.186	0.198	0.187	0.275
Pell, C	−0.481	−0.624	−0.567	−0.408	Baker, H	0.196	0.346	0.711	0.685
Bayh, B	−0.472	−0.429			Boschwitz, R	0.204	0.233	0.472	0.424
Mitchell, G		−0.528	−0.385	−0.426	Boren, D	0.207	0.099	0.102	0.191
Stevenson, A	−0.423	−0.612			Kassebaum, N	0.224	0.158	0.490	0.467
Muskie, E	−0.411	−0.247			Schweiker, R	0.227	0.247		
Leahy, P	−0.402	−0.538	−0.624	−0.543	Stevens, T	0.230	0.314	0.649	0.651
Jackson, H	−0.397	−0.534	−0.252	−0.273	Young, M	0.249	0.298		
Matsunaga	−0.389	−0.659	−0.481	−0.511	Dole, R	0.300	0.238	0.755	0.691
Inouye, D	−0.387	−0.453	−0.457	−0.452	Domenici, P	0.326	0.334	0.702	0.650
Biden, J	−0.384	−0.465	−0.493	−0.245	Cochran, T	0.328	0.343	0.654	0.649
Javits, J	−0.373	−0.414			Roth, W	0.332	0.465	0.481	0.502

TABLE 6 continued

Baucus, M	−0.356	−0.462	−0.331	−0.242	Hayakawa, S	0.353	0.490	0.869	0.760
Eagleton, T	−0.337	−0.359	−0.702	−0.597	Schmitt, H	0.359	0.430	0.707	0.538
Glenn, J	−0.320	−0.339	−0.228	−0.244	Andrews, M			0.504	0.412
Hart, G	−0.307	−0.360	−0.578	−0.469	Byrd, H	0.418	0.390	0.533	0.667
Durkin, J	−0.298	−0.445			Hawkins, P			0.567	0.455
Magnuson, W	−0.273	−0.490			Warner, J	0.463	0.405	0.727	0.770
Bumpers, D	−0.218	−0.293	−0.554	−0.426	Lugar, R	0.464	0.446	0.775	0.582
Mathias, C	−0.214	−0.482	0.205	0.029	Damato, A			0.611	0.469
Byrd, R	−0.197	−0.329	−0.293	−0.123	Gorton, S			0.615	0.476
Burdick, Q	−0.196	−0.419	−0.174	−0.139	Tower, J	0.491	0.625	0.858	0.800
Chiles, L	−0.185	−0.119	−0.077	0.108	Simpson, A	0.493	0.484	0.760	0.615
Huddleston, W	−0.178	−0.351	−0.197	−0.028	Wallop, M	0.508	0.504	0.801	0.723
Sasser, J	−0.154	−0.248	−0.133	−0.065	Rudman, W			0.563	0.520
Melcher, J	−0.145	−0.272	−0.076	−0.045	Goldwater, B	0.579	0.607	0.844	1.000
Gravel, M	−0.128	−0.468			Kasten, R			0.678	0.596
Church, F	−0.127	−0.391			Jepsen, R	0.599	0.536	0.797	0.591
Weicker, L	−0.121	−0.349	0.194	0.017	Thurmond, S	0.603	0.542	0.817	0.825
Stafford, R	−0.119	−0.146	0.473	0.247	Brady, N				0.614
Dixon, A			−0.077	−0.117	Grassley, C			0.735	0.626
Randolph, J	−0.099	−0.408	−0.300	−0.137	Abdnor, J			0.659	0.641
Cannon, H	−0.097	−0.104	−0.086	−0.070	Murkowski, F			0.655	0.679
Stewart, D	−0.093	−0.224			Mattingly, M			0.832	0.721
Chafee, J	−0.081	−0.175	0.382	0.124	Garn, J	0.760	0.708	0.865	0.784
Percy, C	−0.075	−0.001	0.583	0.391	Laxalt, P	0.777	0.731	0.856	0.776
Pryor, D	−0.066	−0.155	−0.190	−0.045	Quayle, D			0.810	0.784
Hatfield, M	−0.062	−0.047	0.436	0.366	McClure, J	0.784	0.779	0.918	0.894
Bentsen, L	−0.054	−0.138	0.113	0.150	Armstrong, W	0.809	0.819	0.821	0.809
Exon, J	−0.033	0.068	−0.037	0.074	Denton, J			0.838	0.821
Hollings, E	−0.028	−0.063	−0.005	−0.067	Humphrey, G	0.840	0.871	0.778	0.803
Packwood, R	−0.026	−0.136	0.537	0.316	Hatch, O	0.852	0.561	0.846	0.678
Nunn, S	−0.022	0.040	0.071	0.178	Nickles, D			0.921	0.875
Ford, W	−0.009	−0.239	−0.244	−0.108	Symms, S			1.000	0.924
Heinz, J	−0.004	0.034	0.329	0.136	East, J			0.939	0.976
Talmadge, H	−0.001	0.001			Helms, J	1.000	1.000	0.877	0.950

matically deviant, geometric means in previous years should explain variations in current geometric means, even after controlling for current spatial position.

There is one obvious deviant senator, William Proxmire, whose reputation for unpredictability in Washington is mirrored by our findings. Proxmire's geometric mean is consistently around .4; no other senator drops below .5. Consequently, we eliminate Proxmire from the ensuing analysis.

Within administrations, there appears to be systematic error by individual senators. To estimate the V-shaped relationship between geometric mean and spatial position, we first ran a quadratic regression of a given year's geometric means on the senator coordinates for that year; then, to see if there was indeed systematic error, we asked whether adding a previous year's geometric mean would improve the fit of the regression. Adding the 1979 geometric mean to a quadratic regression of the 1980 geometric mean on 1980 spatial position improves R^2 from .66 to .75. The coefficient on the 1979 geometric mean is 5.7 times its estimated standard error. Similarly, between 1982 and 1981, the R^2 moves from .49 to .74 and the coefficient is 9.2 times its standard error. In contrast, this systematic error carries much more weakly across administrations. Adding the 1980 geometric mean to the 1981 equation increases R^2 only from .63 to .67. The coefficient is now only 2.7 times its standard error. Further adding the 1979 geometric mean leaves R^2 virtually unchanged. One obvious interpretation of this result is that the systematic error results from multidimensional considerations. However, the fact that this error carries weakly from the 96th Congress to the 97th suggests that any omitted dimensions are not very stable temporally. Rather than being ideological in nature, the omitted dimensions may reflect coalitional considerations, such as loyalty to the leadership of the Senate or to the White House.

Technical Evaluation of the Model

In addition to assessing the substantive validity of our model, it is important to provide a technical assessment of the performance of NOMINATE. There are at least five reasons for caution in the use of NOMINATE: (1) our need to impose constraints as a result of the perfect roll call and random roll call problems; (2) the nonconvexity of the likelihood function; (3) the fact that the expansion of the parameter space as we add roll calls or legislators implies that we cannot rely on the standard proof of consistency of maximum likelihood estimators (Chamberlain, 1980); (4) the technically incorrect computation of standard errors that results from our alternating procedure; (5) misspecification of the model.

We now summarize results concerning these issues. A more detailed analysis is available in Poole and Rosenthal (1983).

Robustness of the Method to Modifications

We developed NOMINATE by extensive testing using the 1979 Senate data on a DEC-2060. We found that our results were quite robust to a set of changes in both the methods for generating starting values and the global iteration algorithm. We also found that results were robust to inclusion or deletion of the one clear outlier in the Senate data, Senator Proxmire. Results were also reasonably robust to inclusion or deletion of our two most nearly perfect senators, Kennedy and Helms. The most sensitive aspect of NOMINATE, as explained in section 4, is in the choice of cutoff level for low minority roll calls. Even here, estimates for legislators and midpoints are quite robust. The choice of the cutoff level does not appreciably affect the recovery of the legislator coordinates and the outcome midpoints. Runs with different cutoff levels show high R^2 values between their sets of legislator and midpoint estimates, with nearly identical linear transformations affecting the two sets of values. Hence, comparisons of legislators to midpoints are quite stable. What changes are the locations of roll call coordinates relative to the legislators.

After the final version of the program was prepared, it was converted to run on a VAX-11/780. Results were replicated. Without further experimentation, the program was applied to the House data set and the other three years of Senate data. In all these cases, the 2.5 percent minority level appeared to give the most sensible results.

Monte Carlo Tests

We also engaged in extensive Monte Carlo tests of the final version of the program. We used one set of 57,036 random numbers to generate data sets for different values of β. In most runs, we used 98 legislators and 291 roll calls; in one run, we used only 50 legislators. We also used three additional sets of 57,036 random numbers to study the effects of varying the distribution of roll call coordinates.

The results were quite encouraging. There is some upward bias in the estimate of β, but the recovered values retain the order of the true values across runs. Estimates of legislator locations and roll call midpoints are highly accurate and essentially unbiased. They are robust to misspecification of the model, at least in a run where voting behavior was generated by a linear utility model rather than by (2).[18]

We recover liberal coordinates less accurately than the midpoints, as expected. In addition, there is some bias toward recovering the liberal coordinates too far to the left (and the conservative coordinates too far to the right). However, the recovery of both the liberal coordinates and the midpoints improved substantially when the number of legislators was increased from 50 to 98. In a legislature as large as the House, the quality of recovery should be excellent as long as the specification is not seriously in error.

Estimation of Standard Errors

In addition to producing point estimates for the coefficients, NOMINATE produces estimates of the standard errors for these coefficients. As these are

[18] We have not investigated such other obvious forms of misspecification as applying a unidimensional model to a multidimensional world, nonindependent errors, etc.

computed separately from the information matrix for each stage of the final global iteration and not from the full information matrix of all parameters, the estimated standard errors might be seriously misleading. Intuition, however, is that the standard errors would be reasonably estimated, since the various stages are only weakly linked. Cross partial derivatives between pairs of legislator coordinates and pairs of coordinates from different roll calls vanish. The cross partials between a legislator coordinate and a roll call coordinate include only a single term corresponding to the legislator's vote on the roll call. They will therefore be small in magnitude.

To test this intuition, we first compared root-mean-square errors of estimates from the Monte Carlo runs to average standard errors estimated by NOMINATE. The two quantities were reasonably similar. We also estimated the 1980 Senate roll call coordinates using the β value and senator coordinate values from 1979 as fixed parameters. In this case, we are computing roll call standard errors using a correct information matrix (under the assumption that the previous estimates are "true" values). There was little difference between these standard errors and those computed when all parameters are estimated. On the whole, the standard errors produced by the program for legislators and unconstrained roll calls are reliable. An overview of the precision of our estimates is provided in Table 7, which shows the average standard errors for our five sets of estimates. A rough guideline for interpreting the numbers in Table 7 is given by the fact that the dimension is normalized to be 2 units in length. Thus a standard error of 0.2 is 10 percent of the length of the dimension. In addition, as the standard deviation, for 1979, of the 100 estimated senator coordinates is 0.41, a standard error of 0.04 for an individual senator represents a very precise estimate. It should be noted that liberal coordinates are always less estimated than midpoints. As would be expected from standard statistical theory, these standard errors are roughly halved as the number of legislators is quadrupled in moving from the Senate estimates to the House estimates. The legislators in the House are less precisely estimated than those in the Senate since we had fewer roll calls for the House.

Conclusion

We have argued, in the preceding sections, that NOMINATE is successful at estimating a unidimensional model of probabilistic roll call voting. Several extensions and refinements are obvious and implementable. These include a multidimensional model and one that allows for variation in utility functions across legislators and in error levels across roll calls. Other extensions are more challenging, including ones that would model correlation in errors across legislators from the same state or cohort and ones that would model temporal variation in the spatial positions of legislators. Still more formidable would be models of agenda control, logrolling, and other forms of strategic behavior.

Rather than conclude with an endorsement of this future research agenda, we would emphasize the usefulness of the present effort. It has successfully accounted for a large share of all the roll call voting data in each of five differ-

TABLE 7

Average Standard Errors Estimated by Nominate

Year	Legislators	Midpoints	Liberal Coordinates	N[a]
		House		
1957-58	0.065	0.053	0.105	172/132
		Senate		
1979	0.037	0.115	0.208	448/346
1980	0.045	0.144	0.258	480/320
1981	0.050	0.135	0.217	397/249
1982	0.046	0.140	0.244	421/244

[a] First number is total roll calls. Second is roll calls estimated without constraints. (Figures in the table for roll calls refer only to estimates without contraints.)

ent congressional voting data sets. It has shown that many of the multiple dimensions claimed in previous research can be interpreted in terms of a single liberal-conservative dimension that allows for voting with error. Clearly, as a first approximation, our spatial model provides a useful description of the congressional roll call voting process. NOMINATE and later evolutions of the program can provide a useful methodology for analyzing the abundant history of roll call votes.

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