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# EXTREMISM, INTENSITY, AND PERCEPTION IN CONGRESSIONAL VOTING

by

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# EXTREMISM, INTENSITY, AND PERCEPTION IN CONGRESSIONAL VOTING

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Social psychology, since the days of Thurstone, has emphasized the study of attitude distributions on quantitative dimensions. The discipline has, moreover, emphasized the study of individual differences. In particular, social psychologists--and sociologists and political scientists as well--have emphasized that the behavior of extremists differs sharply from that of moderates.

Economists, in contrast, have been concerned with the development of mathematical models of choice along a dimension based on the concept, in one dimension, of a single-peaked indirect utility function. At the same time, empirical work by economists has neglected individual differences at the psychological level. In economic models, individuals typically differ only in their ideal points (which may differ because of budget constraints or socioeconomic characteristics) but not in the form of the utility function.

In this paper, we use the findings on extremism developed in social psychology to inform a choice model commonly used in econometrices. This is the logit or double exponential model. In fact, as shown by Yellott (1977), the double exponential distribution is the only distribution that satisfies Thurstone case V discriminal processes, equivalent to Luce's choice axiom, for all possible choice experiments. Consequently, the research reported here accomplishes a form of Thurstone scaling that allows for the behavior of extremists to differ from that of moderates.

Thurstone scaling can be viewed as a vehicle for obtaining spatial representations. Students of public choice have had a persistent interest in obtaining spatial representations of voters and the alternatives they face. Work has taken place both in the setting of mass elections (Weisberg and Rusk, 1970; Rabinowitz, 1976; Wang, Schonemann and Rusk, 1975; Cahoon, Hinich, and Ordeshook, 1978; Poole and Rosenthal, 1984; Rosenthal and Sen 1973, 1977) and in the setting of legislative roll call voting (MacRae, 1958, 1967, 1970; Weisberg, 1968; Morrison, 1972; Poole, 1981, 1984; Poole and Daniels, 1984; Poole and Rosenthal, 1983). Much of this work derives from the stimulus provided by Louis Guttman's seminal development of unidimensional scaling of survey items (Guttman, 1941, 1950). As orginally formulated by Guttman, Scalogram analysis (later to be named after its inventor) produced a scale which measured movement away from or towards a single stimulus. The classic example is racial prejudice where the items on the scale measure the "social-distance" of, say, whites from blacks.

In the decade following the publication of the American Soldier, MacRae (1958) developed the connection between Guttman scaling and spatial voting models[1]. Suppose the voters have symmetric, single peaked utility functions over a continuum and each vote is between a pair of alternatives represented as points on the continuum. If there is no error, then the voters will vote for the alternative closest to them. For example, the “Center Moderate" shown in Figure 1 would vote for an alternative at y over one at y'. Accordingly, a series of such roll calls will produce a perfect Guttman scale. The "items" on the scale will be the midpoints of the roll call outcome pairs. [In the example, the midpoint is (y+y')/2.] Numerous researchers have followed MacRae's lead. Unidimensional Guttman scales pervade the substantive literature on roll-call voting (e.g., Miller and Stokes, 1963; Wood, 1968; Clausen, 1973).

Choice with Error

Several important observations in the early work by Guttman and his colleagues have not, however, influenced the methodology of voting analysis. First, Guttman realized that scales might not always be perfect. They might contain error. He referred to scales with error as "quasi-scales.” Rather than explicitly model error, however, most researchers have continued to try to form near perfect scales. In an earlier paper (Poole and Rosenthal, 1983), we developed a method for spatial analysis of roll call data that explicitly allowed for error in the form of a choice model with stochastic utility functions. The level of error was constant across roll calls. That paper strongly supports the contention by Judd and Krosnick (1982) that most issues (e.g., oil windfall profits, the Panama canal treaty) can be mapped into a single liberal-conservative dimension. This dimension correctly classifies over 80 per cent of all roll call votes. The earlier paper provides further details on the estimation procedure, model testing, and substantive implications for political economy.

This paper extends the method to allow for varying error levels across roll calls. Such variation might well occur if only because legislators are better informed about some bills than they are about other bills.

It is important to emphasize, however, that factors other than perception can result in errors. For example, assume responses are partly determined by factors outside the dimension that appear random. As an individual tends to indifference, in terms of position on the dimension between alternatives, these outside elements will dominate his or her choice. This might explain the finding, reported by Clausen (1967), that even on near perfect Guttman scales, "moderates" make more errors than “extremists”. But if the midpoints generated by survey items tend to lie closer to the ideal points of moderates than to those of extremists, we can expect more errors by moderates even if their utility functions and levels of perceptual error are identical. Similarly, if because of the competition induced by majority rule, most of the midpoints in a legislature fall close to the center of the space, we can expect, ceteris paribus, more "errors" by moderates. In addition, when we allow for different error levels for different roll calls we are perhaps allowing as much for different levels in omitted dimensions as in differences in perception.

## Intensity and Choice

A second observation by Guttman and his colleagues was that people who cared strongly about an issue tended to have extreme scale values. This was referred to as the U-shaped curve of intensity (Suchman, 1950)[2]. That people with relatively extreme positions differ systematically from moderates is well-documented. In another classic study, Lerner (1957) found that Iranian extremists of both the left and right persuasions had shared sociodemographic and attitudinal characteristics. Similarly, Ladd (1982) has shown that strong conservatives and strong liberals in the United States are both disproportionately college-educated.

That the choice behavior of extremists may differ systematically from that of moderates is perhaps most forcefully developed in social psychology in the social judgment theory of Sherif and Hovland (1961)[3]. To structure the theory in terms of our model, we rely on the development of Keisler et al. (1969). There are at least two aspects of the theory that are relevant for choice models. First, the theory explicitly posits that humans order stimuli on dimensions even when there is no natural or explicit ordering [Keisler et al., 1969, p.241].Second, behavior is related to "involvements" [Keisler et al., 1969, p.243]. That is, to translate, more involved people may have different utility functions. Moreover, operationally, Sherif et al. (1965) define involvement as membership in a group with a position on an issue. The issue in their case was in fact a political one, prohibition. Poole (1981) has documented that interest groups take very distinctive positions on political issues. They are found disproportionately at the extreme ends of the political spectrum.

Sherif et al. (1965) postulate that individuals will partition the dimension into the three latitudes of acceptance, rejection, and non-commitment. If a model of continuous utility seems to us to be conceptually preferable to this trichotomization, the partitioning is much in the spirit of our model. In fact, in an institutional setting other than the U.S. Congress, one might think of the latitude of acceptance as referring to choices for which a "yea” vote was most likely; similarly, "nay" would be most likely for those choices in the rejection latitude; and "Abstention" would correspond to non-comittment. Sherif et al. claim that involvement increases the latitude of rejection. To translate to our model, utility will fall to a low level more rapidly as a function of spatial distance for "involved" individuals. (Compare the two extremists to the moderate in Figure 1.) A series of empirical studies tHovland and Sherif (1952); Hovland et al. (1957); and Sherif et al. (1965)] all developed the finding that "those with extreme positions use broader categories for rejection than for acceptance and that their category for rejection is wider than the rejection category of more moderate subjects" (Keisler et al., 1969, p.252). However, the experimental data appears as compatible with a continuous, stochastic utility model as it does with three discontinuous "latitudes." For example, in Hovland and Sherif's well-known experiments on the prohibition issue, percentages favorably evaluating communications vary smoothly as a function of spatial position [see Figure 6.3, p.253 in Keisler et al. (1969)].

The claim by Hovland and Sherif that the perceptions of choices (stimuli) vary as a function of spatial position has been disputed by Hinckley (1963), Zavalloni and Cook (1965), and Upshaw (1965). But in arguing that “judges do not differ in the astuteness of their discriminative abilities, but rather that they only differ in their judgmental language”(Keisler et al., 1969, p.265), Upshaw does not contradict the proposition that utility functions (read "judgmental language") vary with spatial position. Our model in fact embodies Upshaw's conclusions since we make the stochastic disturbance (including perceptual error) dependent on specific roll calls (stimuli) but independent of spatial position of the legislator whereas the parameters of the individual utility function vary with spatial position.

Of course, the controversy over whether perceptions depend upon spatial position is not definitively resolved. But Keisler et al. make the interesting argument that the variation in experimental findings may in fact result from the influence of the eventual implications of choice on the reporting of perceptions. That is, reporting may be affected by "the extent to which judges believe that social consequences will ensue as a function of how one judges things. Thus the involved Biro her may indeed be able to discriminate among types of liberals and radicals but may still tend to place them in one category because of his concern about socialism"[5] (Keisler, et al., 1969, p.276).

This phenomenon pops up often in national politics. During the Republican national convention in 1980, Jesse Helms and his allies described George Bush as liberal. In fact, Bush's voting record as a member of the House (1967-1970) was to the right of Gerald Ford's [Poole and Daniels, 1984]).  In

American political discourse, what is meant by liberal and conservative is a function of how liberal and conservative you are. If Jesse Helms sees George Bush and Edward Kennedy as liberals, Edward Kennedy undoubtedly sees both Bush and Helms as conservatives. But it would be naive to assume that Helms, for example, is unable to differentiate the liberal-conservative positions of Bush and Kennedy. They just appear at such low positions on his utility scale that the differences do not affect his evaluation. Both Bush and Kennedy fall into Helms' “latitude of rejection." We do not need to argue perceptual differences to explain differences in strategic political discourse.

Perceptual reporting is affected by information as well as by strategic considerations. As noted initially by Berelson et al. (1954) and recently by Granberg et al. (1981), voters are likely to assimilate the issue position of a politician they favor to their own position. "Misperception” of this type appears less likely as information improves. For example, Judd and Johnson (1981) contrasted women inolved in the feminist movement with uninvolved women and found no polarization effects in perception where information was likely to be high and substantial effects where information was likely to be low. Since members of Congress can be expected to be exceptionally well-informed, it would be reasonable to assume undeformed perceptions in our observations.

To summarize our understanding of the literature on extremism, extremists do differ substantially from moderates in important ways, as indicated by the evidence on intensity. Whether there are differences in perception is ambiguous. Consequently, our modeling decision has been to omit distinctions in perception and to use only the utility function to model distinctions between extremist and moderate legislators. Specifically, we hypothesize that extremists will have more sharply peaked utility functions than moderates. This has immediate implications for choice. As an illustration, consider a roll call with both alternatives to one side of a legislator and in the concave region of the utility function. Imagine such a roll call with the two alternatives at given distances from an extremist legislator, such as the points x and x' in Figure 1. Imagine another roll call with its alternatives at the same distances from a moderate legislator, such as the points y and y'. The difference in the two utilities is obviously greatest for the extremist legislator. The extremist legislator will thus be less likely to make a voting error on the first roll call than will the moderate on the second. In fact, for the illustration, the extremist will choose x over x' with probability .86 while the moderate chooses y over y' only with probability .72[6]. As a result, their behavior will look like the behavior of two individuals with the same utility function but with different levels of perceptual error. The example illustrates the point that it is very difficult to identify the effects of perception separately from those of “intensity” in the utility function. We have loaded the distinctions between moderate and extremist legislators into the utility function largely as a matter of economy in model-building.

The work presented here is based on a one-dimensional model. We have a multidimensional extension in progress. However, in terms of our data base, the U.S. Congress, a one dimensional model accounts for most behavior. Moreover, allowing for variation in utility functions and error levels can be accomplished by adding only four parameters to the basic one dimensional model. In contrast, adding a dimension would add hundreds of parameters. We think, therefore, that we have taken a parsimonious next step.

We next present the methodology, followed by the results. We find that allowing for variable error levels and utility functions makes a highly statistically significant improvement to the likelihood function. However, increases in explantory power are quite modest. The real gain comes in increased substantive validity of the estimated coordinates for roll calls and legislators. In other words, in estimating coordinates, there are identification or "colinearity" problems which we detail below. The extensions to our model help to pick a "reasonable" set of estimates.

Method

Consistent with spatial theory we assume that each legislator has a most preferred position or ideal point on the continuum. We also assume that each roll call is a choice between two point 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,...,p) is denoted by xi. The number of roll calls is denoted by t and the positions of the yea and nay outcomes are denoted by $Z_{y\ell}$ and $Z_{n\ell}$. ($\ell$ =1,...,t) where “y” stands for yea and “n" nay. The distance of the ith legislator to one of the outcomes on the $\ell$th roll call therefore is

$$d_{ij\ell} = | x_I - z_{j\ell} | , \ j = y,n \tag{1}$$

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 $\ell$ to be:

$$U_{ij\ell} = \beta_\ell \exp\left[\frac{-w_i^2 d_{ij\ell}^2}{2}\right] + \epsilon_{ij\ell} \tag{2}$$

where $\beta_\ell$ and $w_i$ are estimated, $d_{ij\ell}$ is as given in (1), and the $\in_{ij\ell}$ are the error terms which we assume to be independently distributed as the logarithm of the inverse exponential (i.e., the logit or Weibull distribution; cf. Dhrymes, 1978, pp.341-342). Consistent with our assumption that extremists will have more sharply peaked utility functions than moderates, we set

$$w_i = w_o + w_1 x_i + w_2 x_i^2 \qquad (3)$$

that is, the utility function is a function of the spatial position of the legislator. The parameter $w_i$ is our measure of the "intensity" of legislator i's preferences. $W_i$ depends solely on the legislator's spatial location, $x_i$, and three estimated parameters that are common to all legislators. We include the linear term, $w_1$, in (3) to test our hypothesis that extremists are more intense than moderates. If we are correct, $w_1$ should be near zero and unsignificant and $w_2$ should be positive and significant.

To allow for different error levels for different roll calls we set

$$\beta_\ell = \beta_o + \beta_1^\lambda{}_\ell \qquad (4)$$

where $\lambda_\ell$ is a measure of the level of error in the roll call. To arrive at $\lambda_\ell$ we first count the number of legislators voting "incorrectly." For example, if a legislator voted yea on roll call $\ell$ but is closer to $z_{n\ell}$ than $z_{y\ell}$, then the vote is "incorrect." We then subtract the number of incorrect votes from the minority vote on the roll call and then divide the difference by the minority vote. This produces a number between O and 1. For example, suppose we have a 65-35 vote in the senate and suppose that $z_y\lambda$ and $z_{n\ell}$ are estimated such that 90 senators are classified correctly and 10 incorrectly. Then $\lambda_\ell$

would be (35-10)/35=.714. Letting $MV_\ell$ stand for the minority vote on roll call $\ell$ and $VE_2$ stand for the number of classification errors on roll call $\ell$ resulting from the estimated locations of the $x_i$'s and $z_{j\ell}$'s, we can express this symbolically as

$$\lambda_\ell = \frac{MV_\ell - VE_\ell}{MV_\ell} \tag{5}$$

The reason we use $\lambda_\ell$ as a measure of error rather than, say, the log likelihood of the roll call or the raw number of classficiation errors, is that $\lambda_\ell$ is not affected by the vote margin of the roll call. For example, suppose we have two roll call--65-35 and 95-5 --and the estimated locations of the $x_i$'s and $z_{j\ell}$'s result in 5 misclassified senators for each roll call. Clearly, making only 5 errors on a 65-35 vote is much more impressive than making 5 errors on a 95-5 vote. This is captured by the $\lambda_\ell$ statistic which, for this example, is .857 and .000 respectively.

If $U_{iy\ell} > U_{in\ell}$ then legislator i votes yea on roll call $\ell$; conversely, if $U_{iy\ell} < U_{in\ell}$ the legislator votes nay. Given the assumption that the $\in_{ij\ell}$ have a Weibull distribution, the probability that legislator i votes yea/nay on roll call $\ell$ is

$$P_{ij\ell} = \frac{\exp[\beta_\ell \exp(\frac{w_i^2 d_{ij\ell}^2}{2})]}{\Phi_{i\ell}}$$

where (6)

$$\Phi_{i\ell} = \exp[\beta_\ell \exp(\frac{w_i^2 d_{iy\ell}^2}{2})] + \exp[\beta_\ell (\frac{-w_i^2 d_{in\ell}^2}{2})]$$

To estimate $\beta_\ell$, $w_i$, and the $x_i$ and $z_{j\ell}$, we have developed the NOMINATE program, a constrained non-linear maximum likelihood procedure. We estimate 5 parameters for the utility function ($w_0$,$w_1$,$w_2$, $\beta_0$, $\beta_1$), p parameters for the legislators (the $x_i$), and 2t parameters for the roll calls (the $z_{y\ell}$ and $z_{n\ell}$) for a total of p+2t+5 parameters. For example, our largest single run of NOMINATE has been for the House of Representatives in the 85th Congress where we estimated the 5 parameters for the utility function, coordinates for 440 representatives, and 344 coordinates for 172 roll calls for a total of 789 parameters. The total number of roll calls taken in the 85th House was 192. We used all those roll calls with at least 2.5% in the minority (i.e., 424-11 or better)--this left us with a total of 172. The total number of individual voting choices was 68,284 (435x172 - missing data). Pairs and announced positions were counted as votes.

As it is impractical to estimate nearly 800 parameters simultaneously, we first estimate the roll call coordinates, then the legislator coordinates, and finally the utility parameters. The NOMINATE acronym thus denotes Nominal Three-step Estimation. These successive estimations define a global iteration. In practice we found that the $\beta$ and w parameters were highly colinear in the neighborhood of global convergence. Consequently, we fix $w_o$, $w_1$, and $w_2$ and perform global iterations estimating the $z_j$, the $x_i$, and $\beta_o$ and $\beta_1$ until convergence and then perform another set of global iterations re-estimating the $z_{j\ell}$, the $x_i$, and estimating $w_o$,$w_1$, and $w_2$ until convergence. This process defines a solar iteration. Solar iterations are continued until overall convergence is achieved . We define convergence as a squared Pearson correlation of .99 or better between all coordinates (both the $x_i$ and $z_{j\ell}$) estimated in the current iteration with those estimated in the previous iteration. This criteria is used to stop both the global iterations as well as the solar iterations. After each global iteration, the legislator space is normalized to be two units in length, with the left-most legislator at -1 and the right-most legislator at +1. Alternating algorithms of this type are common in psychometric applications [e.g., Carroll and Chang (1970); Takane, Young, and DeLeeuw (1977)].

A particular problem in implementing NOMINATE, detailed in Poole and Rosenthal (1984), involves how to deal with roll calls that are unscalable in two ways. First, if voting is too perfect or error free along the dimension, the roll call coordinates, if unconstrained, will lie outside the space of legislators. This would arise, for example, if the legislators, aligned left to right voted YYY...YYYNNN...NNN. Second, if voting is too random, the midpoint can be placed outside the dimension. NOMINATE heuristically constrains the midpoints and at least one coordinate to lie in [-1,+1]. Empirically the estimates of $\beta_\ell$ and the "spread" $z_n - z_y$ are highly colinear. When $\beta_1 = w_1 = w_2 = 0$, the estimated spreads are quite large and the constraints are invoked relatively frequently. Fortunately, as we shall now see, the variable $\beta$, variable w model greatly lessens the need to rely on these constraints.

## Results

### Reducing the Number of Unscalable Roll Calls

For all five of the data sets we examined, there is little doubt that allowing for variable perception across roll calls and variable utility across senators substantially improves the estimation of a more simple unidimensional model. The major element of improvement is in the reduction of the number of roll calls that are "unscalable" in the simple model. By "unscalable", we mean a roll call whose estimates, if unconstrained, would be unacceptable on theoretical grounds.

As mentioned above, there are two types of unscalable roll calls. First, there are roll calls whose midpoints would be estimated outside the space spanned by the legislator coordinates. Political theory would suggest that all midpoints should be interior. Second, there are roll calls which would have both outcome coordinates estimated outside the space. These roll calls have acceptable estimates of the midpoints, but political theory would suggest that at least one alternative on the agenda should be interior to the space. Of the two types, we regard the second as a more serious problem to a scaling methodology. An external midpoint might simply indicate a roll call dividing legislators on some omitted dimensions, but external outcome coordinates are puzzling when a roll call appears scalable in terms of midpoint location.

As seen in Table 1, the number of unscalable roll calls is cut by roughly 50 per cent when we relax the constraint of common error rates and common utility functions. This is an important gain in the interpretability of outcome coordinates. Moreover, almost all the remaining unscalable roll calls are of the first type. That is, whenever we are able to locate a midpoint interior to the space, we now almost always locate the outcome coordinates interior to the space. Indeed, for the House in 1957-58, there are no roll calls with both outcome coordinates exterior to the space. The House also has, it can be observed, relatively fewer unscalable roll calls than the Senate data. This is because the House coordinates are essentially estimated from 435 observations as against only 100 for the Senate. Consequently, sampling fluctuations produce fewer odd-looking roll calls. The House data are our best indication of how much of all roll call choice can be characterized by a unidimensional model.

Improvement in Fit and Prediction

The improvement in fit and prediction, whether measured in terms of the increase in the geometric mean predicted probability of the observed voting outcomes as shown in Table 1 or in terms of correctly classified votes, is minute. (The geometric mean equals exp(L/N) where L is the log-likelihood and N is the total number of individual votes used in the analysis. It converts an average log-likelihood to a probability.) The probabilities improve only in the third decimal place. However, given our enormous sample sizes, there is little doubt this change is statistically significant. In normal maximum likelihood problems, twice the difference in the log-likelihoods has a chi-square distribution. Since we use heuristic constraints and since the number of parameters we estimate grows as the number of observations grows, it is not technically appropriate to perform the standard chi-square test. For comparative purposes, however, we note that we estimate four additional parameters over those estimated in the simple spatial model. In turn, the 0.01 level is reached with 4 degrees of freedom with a chi-square of 13.3. As shown in Table 2, our "chi-square" values range from 110 to 594.

Importance of the Variable Utility Model

The impact of the variable utility model is first shown in Figure 1. The figure, based on the estimates of the parameter w for the 1957-58 House data, shows that moderates have much flatter utility functions than extreme liberals or extreme conservatives. The estimated parabolas for w [equation (3)] are graphed in Figure 2, where the familiar U-shaped curve of intensity reappears. It can be seen that the estimates for all five data sets are quite similar. In fact, we ourselves were pleasantly surprised in the similarity of estimates for two different legislative bodies separated by over two decades.

One would naturally suspect that these similar results were some consequence of methodological artifact. We, therefore, conducted a Monte Carlo simulation. The values used for senators and roll call coordinates were similar to those recovered for the U.S. Senate. The "true" values of $w_0$, $w_1$, and $w_2$ were 0.5, 0, and 0, respectively. Recovered values were 0.49, .01, and .02. (The Monte Carlo experiment should be quite representative, being based on 60,000 random numbers.) Our recovered values for the Congressional data range from 0.45 to 0.49 for $w_0$, -.01 to .05 for $w_1$, and .14 to .22 for $w_2$. The important quadratic parameter is much larger than that recovered in Monte Carlo experimentation when the data is generated with a zero value. Thus, the U-shaped curve appears authentic.

Another check on the U-shaped curve is to compute twice the difference between the final log-likelihood after the w parameters have been estimated and the log-likelihood at the point in the solar iteration where just beta was variable. These values range from 90 to 418. In turn, the .01 level for chi-square with 2 d.f. is -9.21.

The estimated U-shaped curve is basically symmetric about the center of the dimension. That is, the linear term is quite small relative to the quadratic term. As such, our data give support to the Rokeach (1956) model of two "rigid" extremes over the Adorno et al. (1950) model of solely a "rigid" extreme right. Tetlock's (1984) value pluralism model blending these two models is weakly supported by the presence of positive linear coefficients for all four years in our Senate estimates but is not supported by the very slightly negative estimate for the House.

Finally, comparing the log-likelihood differences shown in Table 2 indicates that it is, with the exception of the Senate in 1982, variable utility functions rather than variable perceptions that most improve the fit of the model. The comparisons are in fact weighted in favor of variable perceptions since variable perceptions are the first part of the solar iteration. Thus, they can fit the data as well as possible, leaving only the residual fit to the variable utility function model. (For reasons of cost, we have not estimated a model with constant perceptions and variable utility.)

### Is There a "Latitude of Rejection"?

The utility functions graphed in Figure 1 show no evident latitude of rejection, a region which all alternatives have low utility. For the space spanned by the roll call alternatives, we never reach the flat part of the Gaussian curve. In fact, our utility functions are strictly concave up until their inflection points. The inflection point is reached when d=1/w. Even for an extremist with w=0.7, the inflection point is not reached until d=1.43 or about 3/4 of the entire range of our space.

The fact that our utility functions are sharply sloped at points relatively far from ideal points may reflect the fact that the American political spectrum is relatively "short". It may also reflect that the structure of political agendas precludes voting on alternatives in "outer space." Even given these possibilities it is impressive that American politicians discriminate relatively well among the alternatives they face.

## Discussion

The U-shaped intensity model, as discovered empirically in the study of attitudes by Suchman (1950) and advocated theoretically by Rokeach (1956), has added an important element to the analysis of choices as found in Congressional roll call voting. This has implications for the development of political and economic theories that rely on the assumption of single-peaked utility functions. Clearly, theory that assumes that utility functions that differ only in their ideal points is inappropriate. Either the theory should be based on a completely general form for the utility function or should incorporate the U-shaped intensity relationship so pervasive in social psychological research.

Social psychology might further inform the development of the formal theory of public choice if we could accomodate the many suggestions as to how perceptions or perceptual error might depend on the interaction between the true position of the individual and the true position of the choice alternative. From the viewpoint of scaling methodology, such models are far more complex than what we have attacked until the present but perhaps well worth further investigation.

For social psychology, this work presents some methodlogical advantages. It shows how both “attitudes”, in this case liberal-conservative ideology, might be studied without the need for questionnaires, content analysis, or other forms of relatively costly investigation. Similarly, intensity can be measured without recourse to self-reporting as in Suchman (1950). Instead, we can recover both a Thurstone scale and intensity measures from observed choices.

# Footnotes

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1. The American Soldier became the popular name for a four volume set of studies. Guttman (1950) appeared in the fourth volume.

2. Schuman and Presser (1981), in providing a brief review of the literature on intensity, credit Allport and Hartman (1925) with the initial discovery of the linkage between intensity and extremism.

3. Of course, social judgment theory is mainly concerned with attitude change. As this paper is not concerned with attitude change, we are mainly concerned here with the implications of the theory for differential choice behavior of extremists vs. moderates.

4. From the perspective of economics, one might argue that primitive utility functions are identical in form and that involvement results in differing indirect utility functions on the liberal-conservative dimension.

5. This type of behavior would appear to be closely related to accentuation produced by a peripheral dimension (Judd and Harackiewicz, 1980).

6. This assumes $\beta=20$, a representative empirical value. See the Method section for formal development of probability calculations.

7. Because the model is continuous, we can ignore the case where $U_{iy\ell}=U_{in\ell}$.

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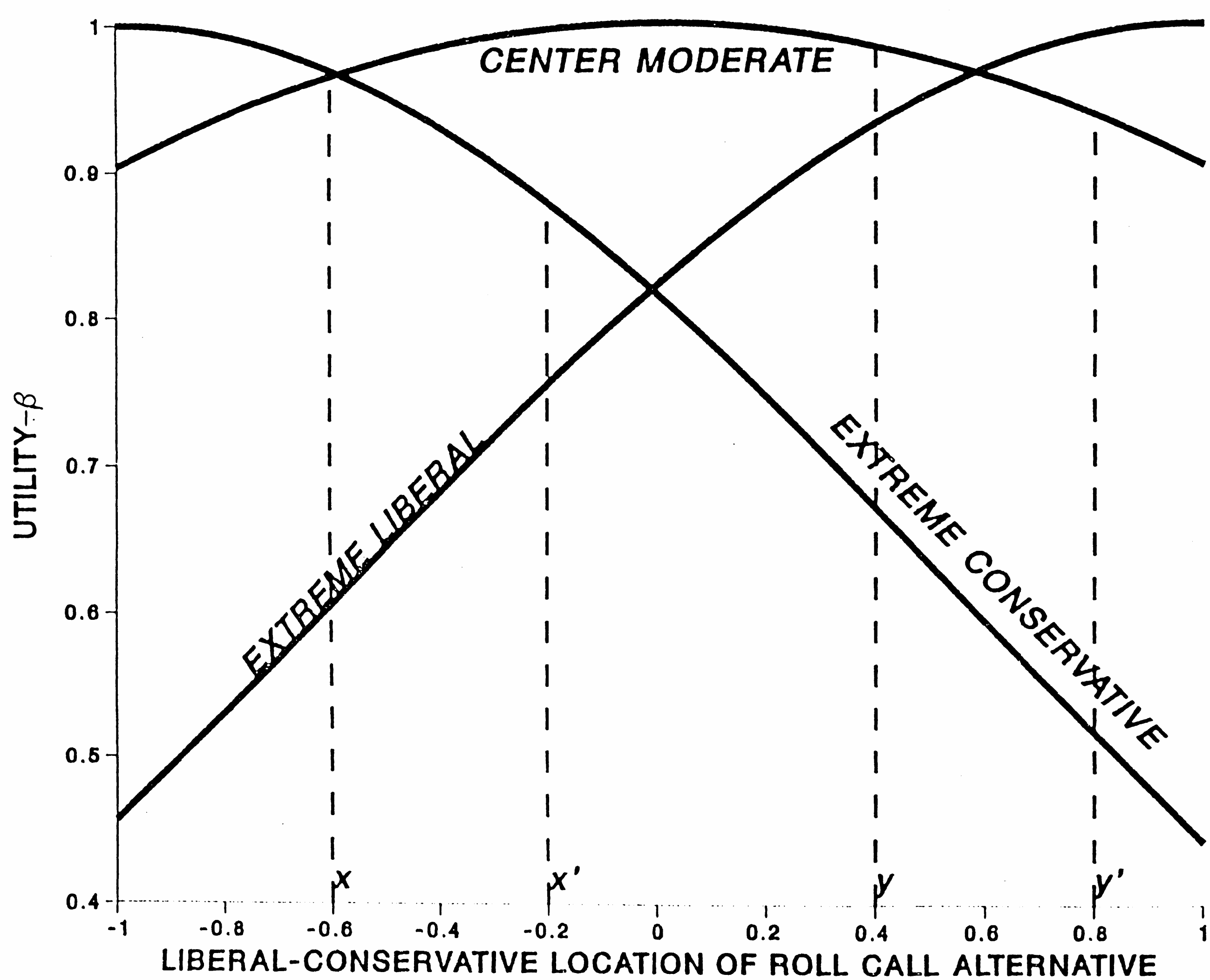
TABLE 1
Estimation Results

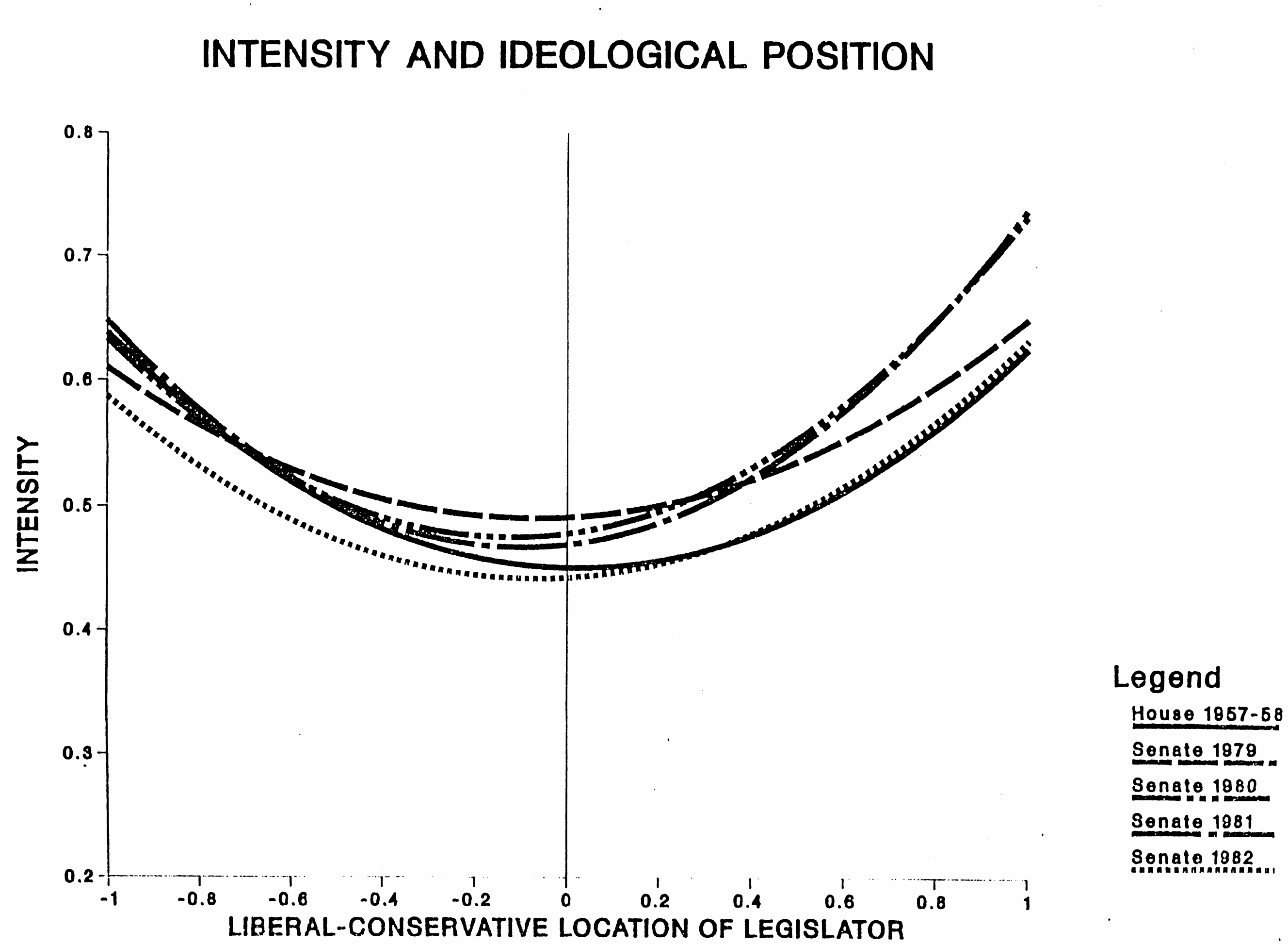
| Data Set | Geometric Mean Probability | | Unscalable Roll Calls | | | Total Roll Calls | Total Individual Votes |
|---|---|---|---|---|---|---|---|
| | Common $\beta$, w | Variable $\beta$, w | Common $\beta$, w | Variable Coordinate $\ell$ $\beta$, w | Outcome Coordinates Outside Space | | |
| House, 57-58 | .6524 | .6536 | 40 | 19 | 0 | 172 | 68,284 |
| Senate, 1979 | .6658 | .6667 | 102 | 43 | 2 | 448 | 40,986 |
| Senate, 1980 | .6639 | .6660 | 160 | 87 | 4 | 486 | 41,951 |
| Senate, 1981 | .6923 | .6981 | 148 | 79 | 15 | 397 | 37,550 |
| Senate, 1982 | .6733 | .6765 | 157 | 103 | 6 | 421 | 40,125 |

TABLE 2
Improvement in Log-Likelihood

| Data Set | 2x Improvement Over Simple Spatial Model | |
|---|---|---|
| | Variable w Phase Only | Total |
| House, 1957-58 | 188 | 262 |
| Senate, 1979 | 90 | 110 |
| Senate, 1980 | 158 | 270 |
| Senate, 1981 | 418 | 594 |
| Senate,1982 | 110 | 376 |

## ESTIMATED SPATIAL UTILITY FUNCTIONS




INTENSITY AND IDEOLOGICAL POSITION
INTENSITY
0.8
0.7
0.6
0.5
0.4
0.3
0.2
-1
-0.8
-0.6
-0.4
-0.2
0
0.2
0.4
0.6
0.8
1
LIBERAL-CONSERVATIVE LOCATION OF LEGISLATOR
Legend
House 1957-58
Senate 1979
Senate 1980
Senate 1981
Senate 1982