*U.S. Presidential Elections 1968–80: A Spatial Analysis**

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

A methodology is presented for empirical estimation of spatial models of voting in mass elections. The basic choice model posits utility functions that depend on spatial distance and a random error term. The parameters of the utility functions are estimated by a polytomous logit model, which is applied to spatial maps of voters and candidates for all presidential elections since 1968. These maps indicate that presidential candidates take positions at the periphery of the distribution of voters. Results based on these maps support the hypothesis of sincere, spatial voting in two-candidate elections. In three-candidate elections, the coefficients of the voter utility function adjust to proxy for the lesser viability of the minor party candidate. In contrast to choice, turnout shows only weak spatial effects. The data are inconsistent with the proposition that turnout is low because voters are not offered distinct alternatives. While the utility function used by voters appears stable through time and candidate and voter positions appear stable in the last few months prior to an election, voter and candidate positions exhibit substantial change over longer time periods.

Introduction

Spatial models dominate theoretical analysis of voting in mass elections. However, empirical estimation of spatial models has not kept pace with the proliferation of theoretical papers. We here present a methodology that overcomes several current methodological problems and can be used to examine many substantive issues in political science.

The virtue of spatial theory is its attempt to explain electoral behavior on the basis of a consistent model of self-interest. However, in its initial formulation, beginning with Downs (1957), spatial theory for the most part assumed a limited electoral environment. Voter behavior was characterized by the assumption of fixed spatial preferences; the candidate "closest" to a voter's fixed ideal point received the voter's vote. Candidate behavior was characterized by a competitive "game" in which the only parameters were the rules of voter behavior. Much of the subsequent effort has not gone to enriching this environment but to discovering the existence or, typically, the nonexistence of equilibrium when competition is in multidimensional spaces.

In our view, theory has emphasized the dimensionality problem at the expense of enriching the environment in two substantively relevant directions. The first direction would recognize a likely instability in voter preferences, even during the course of a campaign. The second would recognize that candidate posi-

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tions are determined by support (interest) group preferences as well as those of the mass electorate.

An emphasis on changing voter preferences can be found in the discussion of party realignment by Burnham (1970) and Sundquist (1973). A reasonable spatial interpretation of this literature would portray the parties as adopting separate and distinct positions relative to a unimodal electorate distributed over a low-dimensional space. A realignment occurs when voter preferences change and the mass electorate becomes more polarized than the parties.

Recently, Hinich and Pollard (1981) have developed spatial theory that allows for changing voter preferences over the space that defines electoral competition. In the Hinich-Pollard model, the candidates are viewed as having fixed positions on a small number of ideological dimensions. The basic policy issues that affect voters vary in salience across elections. Voters oftentimes cannot predict candidate performance on these issues so they "project" to the issues from the candidates. The changing saliences of the policy issues results in a changing projected ideal point of the voter on the ideological space. In short, in contrast to previous spatial theory, the candidates remain fixed (at least in the short run) and the voters move.

In addition to voter mobility, the party realignment literature emphasizes relatively fixed and *distinct* party and candidate positions even given a unimodal electorate. The spatial model literature has as yet not produced a convincing explanation for this presumed stability. Our view, to which we return shortly, is that this party stability arises because of the relatively stable position of support groups with extreme spatial positions. Candidates need to appeal to these groups in order to be nominated. Vote totals in the general election in turn depend not only on voter preferences but also on resources furnished by these groups. Resource seeking "pulls" the candidates away from seemingly vote-maximizing positions in the center of the distribution.

Much empirical work needs to be done in order to characterize the electoral environment and to test spatial theory. Presently, there are two main entry points to empirical spatial analysis. One is the unfolding literature. Rabinowitz (1976), Cahoon (1975), Wang, Schonemann, and Rusk (1975), Poole (1981), and Poole and Daniels (1980, 1982) have used distance measures to produce Euclidean (spatial) representations. Thermometer ratings of candidates by a sample of voters are used to produce a Euclidean representation of both voters and candidates. Interest group ratings of congressmen are used to produce a Euclidean representation of both the congressmen and the interest groups. Aldrich and McKelvey (1977) have used perceptual data to produce Euclidean representations of candidates on issue dimensions. The other entry point is the estimation of spatial voting models with aggregate data by Rosenthal and Sen (1973, 1977).

Several features of our analysis differ from these earlier dimensional analyses of presidential voting.

1. We achieve our Euclidean representation of candidates and voters by a straightforward metric least-squares unfolding. After a seminal paper by Weisberg and Rusk (1970), which located candidates but not voters, Rabinowitz (1976)

located candidates with a nonmetric procedure and then, in a second stage, voters. Our procedure locates voters and candidates simultaneously. Use of metric versus nonmetric procedures is largely a trade-off between the generality of the nonmetric procedure versus the greater precision (assuming the actual analysis is robust to the metric assumptions) obtained by making metric assumptions a priori. The robustness of our procedure allows us to include nearly all respondents in the analysis, whereas Rabinowitz and, especially, Cahoon, Hinich, and Ordeshook (1978) treated a larger number as missing data.[1] As our unfolding procedure has had previous development in this journal (Poole, 1981), the remainder of our analysis emphasizes methods for analyzing the actual voting behavior reported by respondents.

2. Once one has obtained a Euclidean representation of preferences, one is naturally interested as to how this representation relates to choice behavior. In both our plots (Figures 1–4) and those of Rabinowitz, cutting lines (bisectors) separating the candidates partition the space in a manner that discriminates well between, say, reported McGovern voters and reported Nixon voters. However, voters close to the cutting lines are not discriminated nearly as well as voters distant from the cutting lines. This observation suggested to us that voting should be modeled as a probabilistic phenomenon. We allow voters to share a common stochastic utility function but to differ in their ideal points. We estimate the parameters of the utility function via a polytomous logit model. The estimates of the parameters can be of substantial interest to theoreticians since many theoretical results depend upon assumptions (e.g., concavity) about the form of the utility function.

3. Models of nonvoting from alienation and indifference are readily specified within the logit framework. This allows us to treat participation and candidate choice as a single decision problem, whereas these decisions are treated sequentially in earlier analyses such as Brody and Page (1973) and Rosenthal and Sen (1977).

Our empirical analysis is based on the CPS U.S. presidential election surveys for 1968, 1972, 1976, and 1980. In the following sections we develop the use of this data base to calculate Euclidean coordinates for candidates and voters, explain how these coordinates are used in a logit model of choice, and estimate a probabilistic model of choice on the basis of these coordinates.

The major substantive conclusions drawn from the work include: (1) on at least one dimension of competition, the presidential candidates are widely separated, with positions at the periphery of the distribution of the voter ideal points; (2) that there is substantial short-term but little long-term stability in the spatial

[1] Aldrich and McKelvey (1977) obtain a set of unidimensional scales, one for each issue, whereas the other procedures, including our own, are multidimensional. The work of Wang, Schonemann, and Rusk (1975) is conceptually based on a notion of probabilistic voting similar to but more restrictive than the one we develop. Their paper focuses mainly on the method for obtaining a Euclidean configuration and does not lead to explicit analysis of probabilistic voting or to the development of substantive implications.

maps; (3) that the less viable third-party candidates, George Wallace and John Anderson, have, as expected, voting probability contours quite different than those for major-party candidates. In particular, Wallace support in 1968 declined sharply as voter ideal points became distant from the Wallace position; (4) although some mild spatial effects on reported participation were found, the primary determinants of turnout would appear to be nonspatial.

Spatial Configurations for Presidential Elections

The Data

The Center for Political Studies at the University of Michigan began asking for "thermometer" evaluations of candidates in its presidential election studies beginning in 1968. Respondents are asked to rate the candidates on a zero to one hundred scale, with one hundred being the most favorable evaluation. Not only the major-party candidates but also roughly ten other prominent politicians are rated. Thus, politicians such as Edward Kennedy, Nelson Rockefeller, and Edmund Muskie have been included.

In 1968, thermometers were included as part of a single post-election survey. In 1972, thermometers appeared in both waves of a two-wave panel study. The panel was first interviewed in September and then again in November. Some respondents were interviewed in 1972, 1974, and 1976, forming a panel. In both presidential years, this panel was part of a larger cross-section. Thermometer data for 1976 were gathered only in a September pre-election wave of a two-wave panel. For 1980, panel interviews in February, June, and October gave us three observations of thermometers.

Polarized Candidates and a Unimodal Electorate: Unfolding Results

Using metric multidimensional unfolding techniques developed by Poole (1978, 1981, 1982), we have computed one-, two-, and three-dimensional Euclidean coordinates for the candidates for each of the seven interviewing dates (one in 1968, two in 1972, one in 1976, three in 1980). For a given set of coordinates, one can compute Euclidean distances between respondents and candidates. For each candidate-respondent pair, one can define prediction error as the difference between the distance and the normalized quantity $(100 - T)/50$, where T is the original thermometer rating. The coordinates are then chosen to minimize the sum of squares of these errors.

Table 1 displays the unfolding results for the four presidential elections of our study. We used the squared Pearson correlation coefficient between the actual and reproduced thermometer scores as our measure of fit. Thus, the two-dimensional Euclidean representation of 1,399 respondents and 12 candidates in the 1968 election survey explained nearly 60 percent of the variance of the thermometer scores.

To obtain a rough guide as to the substantive content underlying the dimensions recovered by the unfolding, we estimated a set of ordinary least-squares

TABLE 1

Unfolding Results for U.S. Presidential Elections

	1968	1972[a]	1976	1972–76 C	1972–76 I	1980 P1	1980 P2	1980 P3	1980 C	1980 I
One dimension	.402	.428	.329	.382	.318	.312	.316	.275	.295	.283
Two dimensions	.587	.601	.475	.521	.431	.484	.450	.405	.439	.381
Three dimensions	.662	.681	.558	.631	.514	.598	.572	.544	.561	.462
Number of candidates	12	12	15	18	31	13	14	14	14	41
Number of respondents	1,399	2,643	2,158	3,822	1,317	967	817	746	2,530	992
Number of thermometers	16,140[b]	27,349	27,447	34,695	34,695	9,932	9,098	9,540	28,570	28,570

NOTE: All entries are Pearson r-squares of computed Euclidean distances with original thermometer scores.

[a] Pre-election panel results. The post-election panel results were virtually identical to these.

[b] This number is less than the number of candidates times the number of respondents because of missing data.

P1, first wave of panel unfolded separately.

P2, second wave of panel unfolded separately.

P3, third wave of panel unfolded separately.

C, all panel thermometers unfolded simultaneously, candidates held constant.

I, all panel thermometers unfolded simultaneously, individuals held constant.

regressions. The dependent variables were the party identification scales and issue scales in the various CPS studies. In addition to party identification, we examined 2 seven-point scales from 1968, 18 from the pre and post waves of the 1972 panel, 11 from 1976, 5 from the second wave of the 1980 panel, and 12 seven-point scales and 4 four-point scales from the pre and post waves of the 1980 cross-section. The independent variables were the unfolding coordinates. Thus, we attempted to see how well we could explain issue positions solely from knowledge of spatial positions estimated from the thermometer responses.

The results indicate clearly that the first dimension is a basic liberal-conservative dimension. When issues other than party identification correlate at all with spatial position, the first dimension has the larger regression coefficient. (Comparisons using standardized beta coefficients are even more strongly in favor of the first dimension.) Issues with simple correlations in excess of .3 with the first dimension include urban unrest and Vietnam in 1968; Vietnam, jobs, busing, civil rights, foreign aid, urban unrest, student unrest, and liberal/conservative in 1972; jobs, busing, foreign aid, urban unrest, liberal/conservative, and health services in 1976; and defense, government services, government spending, busing, the Equal Rights Amendment, and liberal/conservative in 1980. That is, there is a broad spectrum of issues—some foreign policy, some domestic, some economic, some social—that relate to the first dimension. (Of course, issues that do not differentiate the thermometer stimuli, e.g., feminist issues before 1980, do not relate to either dimension.) When it was included, the general liberal/conservative issue had the highest of these correlations. Therefore, ''liberal/conservative'' appears to be a highly appropriate description for the main dimension.

The second dimension generally has very low correlations with all of the issues and tends to have insignificant regression coefficients. The only case where the second dimension had a greater coefficient than the first was for the two inflation questions in 1972, but here the multiple squared correlations were below .05. All second-dimension simple correlations were below .11 in 1968, .24 in 1972, .17 in 1976, and .07 in 1980.

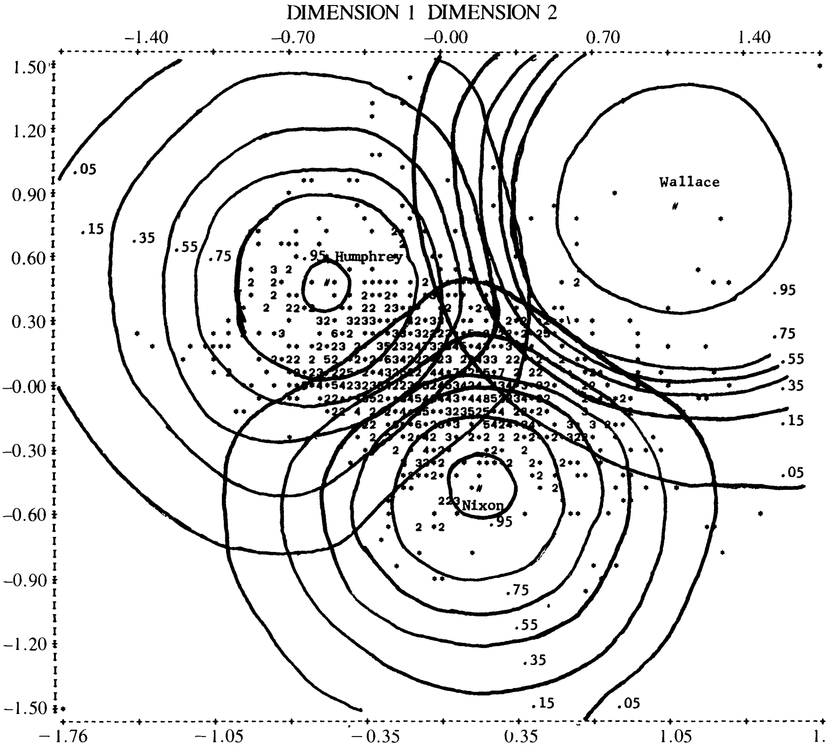
In contrast to the substantive issue scale results, the second dimension does have a regression coefficient on party identification that equals that for the first dimension in 1972 and 1976, exceeds it in 1968, but is essentially zero in 1980. Our speculative interpretation of this finding is that the second dimension captures the traditional identification of southern conservatives with the Democratic party. This aspect of party identification is gradually disappearing as the South realigns itself in terms of the national liberal/conservative polarity. The dimension was heightened by the Wallace candidacy in 1968, and its gradual disappearance was attenuated by a regional attachment to Carter in 1976, but by 1980 Reagan's brand of conservatism and Carter's record in office had served to erase the relevance of the second dimension. The contrast between 1980 and earlier years is consistent with our finding, below, that a one-dimensional voting model is preferred for 1980 while two dimensions are better for earlier years.

Our unfolding results are basically similar to those of Rabinowitz (1978). First, most of the variation of the thermometers is accounted for by three dimensions. Second, the candidates appear near the periphery of the space, although voters are unimodally distributed about the center. (See Figures 1–4. Figures 1 and 3 are the plots obtained for voters, Figures 2 and 4 the plots obtained for nonvoters. Similar plots for 1972 and 1980 are available from the authors on request. Major candidate positions are similar across years. In 1980, Anderson is close to Carter. The contour lines are explained below.)

This second result is at variance with some simple spatial theories which would hold that the candidates should converge to a point in the center of the distribution. One might be inclined to ascribe the extreme placement of the candidates to either lack of gradated discrimination by respondents or to methodological artifact.[2] Two corroborative pieces of evidence, however, buttress the

FIGURE 1

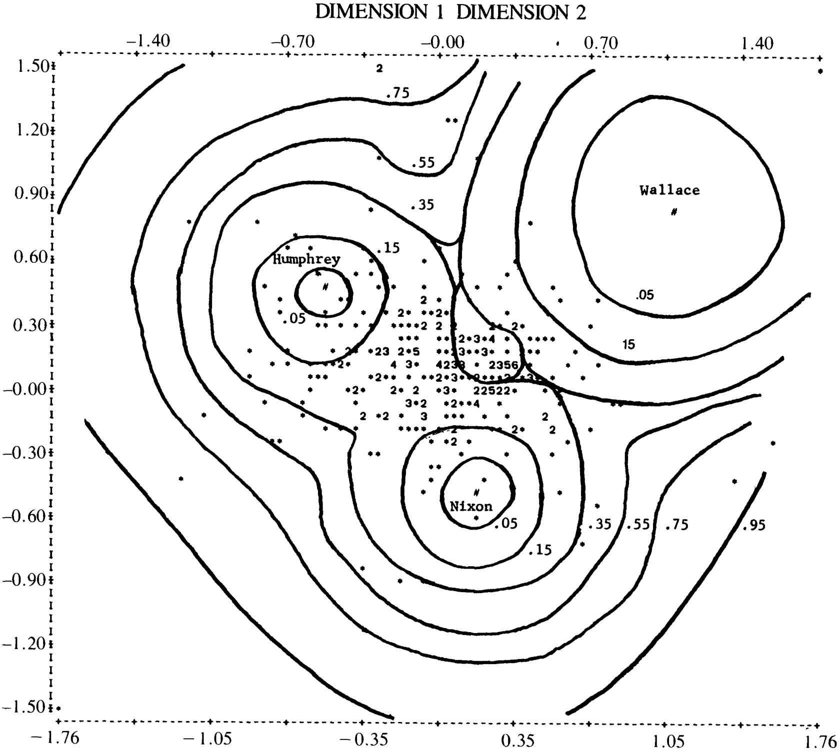
1968 Voters



[2] One possibility for methodological artifact can be ruled out. Suppose the thermometers measured not distance but utility and that reported utilities were some nonlinear function of distance, say distance squared. Suppose further that the candidates had converged to a common position. Using

FIGURE 2

1968 Nonvoters



peripheral placement. First, independent study of candidate placement on issue scales by Brody and Page (1973) and Page (1978) discloses that while candidates converge on some issues they diverge on others. To the extent thermometer scores are issue-related, they will be determined by the issues on which the candidates are in fact differentiated. In the multidimensional spatial representation, we can always define one axis to be the line connecting the two major party candidates. This axis represents an issue dimension on which separation occurs. However, from the specification of the first axis, all other axes must represent dimensions on which convergence occurs. Thus, Figures 1–4 are quite consistent with the issue analyses.

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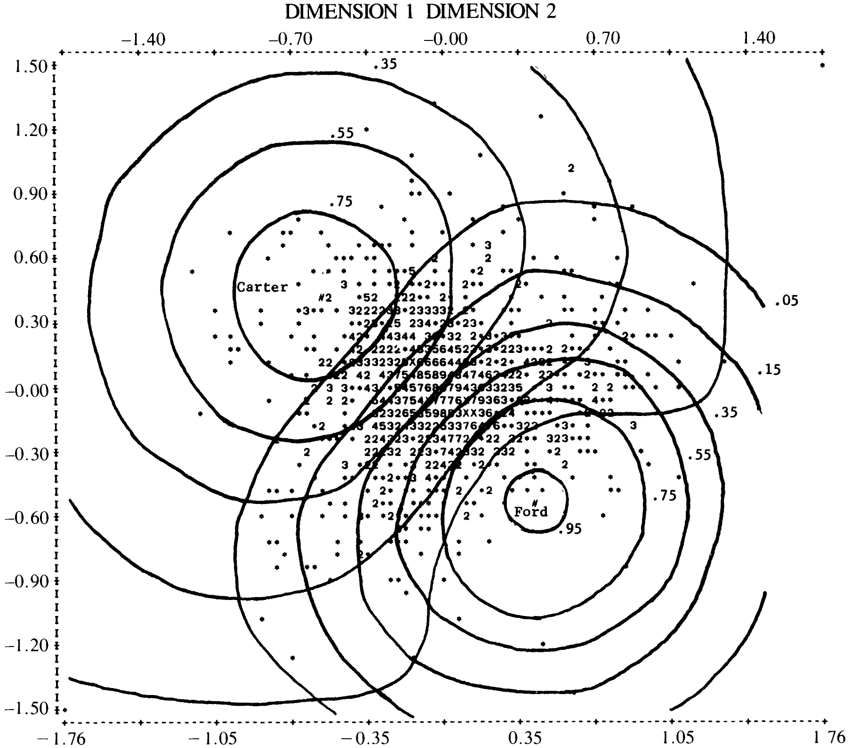
these incorrect measures of distance, would our methodology produce widely separated candidates? Fortunately, the answer is no. While the recovered distances of voters to candidates would be deformed, the candidates themselves would be placed at a common position (or close together with ''noisy'' reported distances). In fact, attempts at psychophysical measurement by Andrus and Lodge (1982) indicate linearity for thermometers over the range of alternatives represented by presidential candidates. Andrus and Lodge obtain nonlinearities with thermometers only when they include strongly negative stimuli such as the Ayatollah Khomeini.

FIGURE 3

1976 Voters



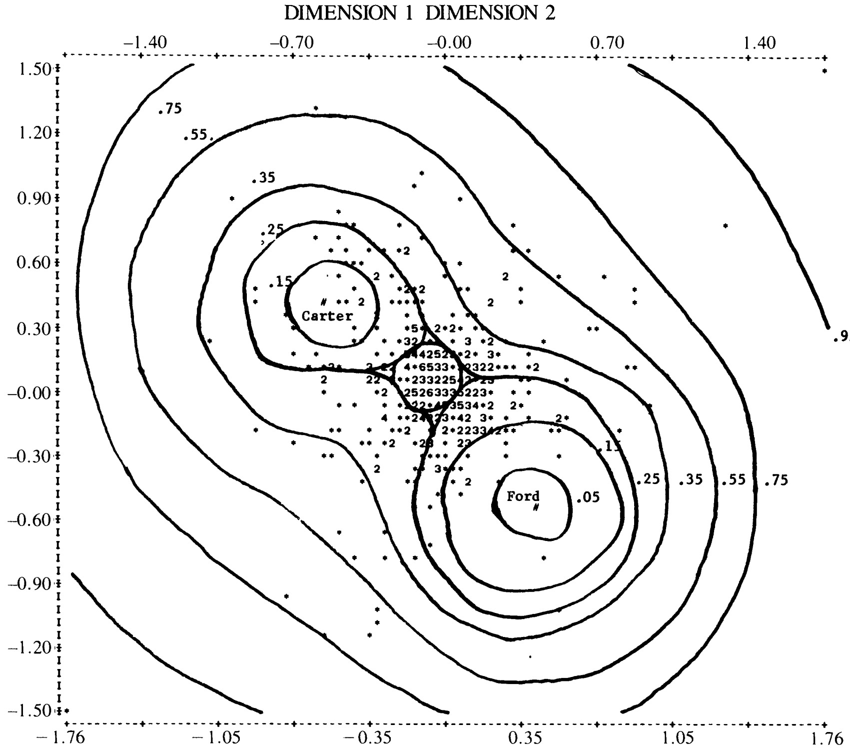
Our second set of evidence on polarization comes from studies of Congress. Fiorina (1974), replicating an earlier study by Strain, found dramatic changes in the roll call positions of a constituency when its representative was replaced by a representative from the opposite party. Similarly, using the same metric unfolding technique used here, but this time on interest group ratings of Congress, Poole (1981) and Poole and Daniels (1982) found a polarized, bimodal Congress.[3] Using their one-dimensional unfolding values, we have confirmed the Strain-Fiorina results for the Senate (see Poole and Rosenthal, 1983; see also Bullock and Brady, 1983). Senators from the same state but from different parties are far apart on the dimension, but senators from the same party and the same state tend to be spatial clones.

Even though congressional data and survey data on issue placement of presidential candidates both point to spatial polarization, we remain struck by the extent to which the candidates are at the periphery of the space. Indeed Burnham

[3] In distinction to the thermometer data, the interest group ratings contain a full range of numerical values and no piling up of implicit "don't knows" on a score of 50. Hence, the polarized positions of the candidates are surely not simply a consequence of these quirks of survey thermometer data.

FIGURE 4

1976 Nonvoters



(1970) and Sundquist (1973) write of distinct, but not extreme, positions. Why the candidates would be pulled so far apart is indicated in the congressional unfolding results. Not only is Congress polarized, but the interest groups are even more polarized than Congress; more than half of the roughly 60 interest groups analyzed by Poole and Daniels (1982) have more extreme positions than almost all members of Congress. Groups like the National Association for the Advancement of Colored People, the Committee on Political Education, American Conservative Union, American Taxpayer's Union, and Americans for Democratic Action tend to be very extreme. To the extent that presidential candidates respond to these groups, they will be drawn to polar positions. To summarize, we equate the polarization of presidential candidates to the polarization of support groups.[4]

[4] Our interpretation thus differs somewhat from that of Rabinowitz (1978), who emphasizes more of an attitudinal disposition to extreme placement in the mass public. Wittman (1983) emphasizes policy preferences of candidates themselves and cites other empirical evidence of divergent candidate positions.

The Choice Model

The general form of a stochastic utility function consistent with spatial theory is:

$$U(\text{individual } i \text{ for candidate } j) = f(d_{ij}) + e_{ij}$$

where d_{ij} is the distance from individual i ($i = 1, \cdots, p$) to candidate j ($j = 1, \cdots, q$), $f(d_{ij})$ is a decreasing function, and e_{ij} is an appropriately defined disturbance term. The specific utility function that we estimate below is

$$U_{ij} = \beta_{1j} + \beta_{2j}d_{ij}^{2\gamma} + e_{ij} \tag{1.1}$$

where U_{ij} represents the utility of individual i for candidate j. The coefficients β_{1j}, β_{2j}, and γ are to be estimated, and the e_{ij} terms are independently distributed as the logarithm of the inverse exponential (cf. Dhrymes, 1978, pp. 341–42). Without loss of generality, we use not voting to define the origin of the utility function:

$$U(\text{individual } i \text{ not voting}) = U_{i,q+1} = 0 + e_{i,q+1}. \tag{1.2}$$

The model of voting is not a standard linear logit model, but one in which restrictions on the β's are imposed by spatial theory. The most important restriction is that there is implicity a zero β on the distance to any other choice in the equation for a given choice, (1.1). In effect, we are assuming that an individual's utility for a candidate depends only on his/her distance to that candidate alone. The utility does not depend upon the distances to the other candidates. An additional hypothesis is that $\beta_{2j} < 0$. In two-candidate races, where only sincere voting should occur, the hypothesis that $\beta_{11} = \beta_{12}$ and $\beta_{21} = \beta_{22}$ can be tested as an additional restriction.

The restrictions imposed by spatial theory help to identify the model. In a standard logit model, the coefficients on any independent variable are identified only up to addition of a constant. In our model, this is true only of the intercepts, which have been normalized by defining the utility of nonvoting to be zero. Had the utility of nonvoting been declared some other constant, this constant would be added to the utility of all other choices. In contrast to the intercepts, the distance coefficients are fully identified by the restrictions that there are zero β's on the distances to other choices. Of course, if the units used to measure distances were changed, the distance β's would change inversely. If all distances were doubled, then β_2 would be halved (if $\gamma = 1/2$).

When there are more than two candidates, strategic voting must be considered. The literature does not provide any theoretical guide as to how to specify the relationship between utility and choice when strategic voting is possible. Rather than forego any opportunity to make empirical observations on the substantively interesting three-candidate races, we treat (1.1) as a pseudo-utility function that proxies for the effects of strategic voting. In the context of U.S. presidential elections, we would still expect equal β's for the major candidates but distinct β's for the less viable third-party candidates. This expectation is supported by the empirical results presented below. In addition, since we "predict" the 1968 election with Wallace as well as the later two-candidate races, our use of (1.1) appears acceptable, even with strategic voting.

We now turn to the motivation for the nonvoting portion of the model, (1.2). Without the stochastic disturbance, if he or she votes, the individual will choose that candidate for which $\beta_{1j} + \beta_{2j}d_{ij}^{2\gamma}$ is greatest in magnitude. Assume the β coefficients do not depend on the candidate, i.e., if $\beta_{1j} = \beta_1$ and $\beta_{2j} = \beta_2$ for all j. In this case, an individual will, in the absence of stochastic disturbance, vote for the closest candidate provided the utility of the closest candidate exceeds the utility of not voting. Precisely, for some candidate j, $\beta_1 + \beta_2 d_{ij}^{2\gamma} > 0$. Since nonvoting is essentially dictated by the distance of the closest candidate, system (1) is a version of the Hinich-Ordeshook (1969) alienation model.

In two-candidate races, nonvoting from indifference can be modeled as:

$$U(\text{individual } i \text{ not voting}) = U_{i,q+1} = \delta_1 + \delta_2|U_{i1} - U_{i2}| + e_{i,q+1}$$

which, subtracting the nonstochastic portion from (1.1) and (1.2) to renormalize, yields:

$$U_{ij} = \beta'_{1j} + \beta_{2j}d_{ij}^{2\gamma} - \delta_2|U_{i1} - U_{i2}| + e_{ij} \qquad (j = 1, 2).$$

When $\beta_{11} = \beta_{12} = \beta_1$ and $\beta_{21} = \beta_{22} = \beta_2$, this becomes

$$U_{ij} = \beta'_1 + \beta_2 d_{ij}^{2\gamma} - \delta'_2|d_{i1}^{2\gamma} - d_{i2}^{2\gamma}| + e_{ij} \tag{1.1'}$$

where $\beta'_1 = \beta_1 - \delta_1$ and $\delta'_2 = \delta_2\beta_2$. We also have, after renormalization,

$$U_{i,q+1} = 0 + e_{i,q+1}. \tag{1.2'}$$

Thus, the indifference model differs from the alienation model only by adding an additional "regressor," the absolute value of the difference in $d_{ij}^{2\gamma}$.

The coefficient γ is an exponent parameter on the utility function. Because estimating γ would result in a nonconvex likelihood function, we restrict estimation to certain prespecified γ. The following empirical results are based on $\gamma = ½$, or utility linear in distance. As long as the utility function was concave ($\gamma > ½$), results were not overly sensitive to the specification of γ.[5] Details of estimation appear in the appendix.

A Probabilistic Choice Model of Presidential Elections

From Preferences to Choices

Having argued that the preference data represented by the thermometers portray the American electorate as a mass largely interior to two polarized major-

[5] We also experimented with $\gamma = 1$ and with $\exp(-d)$ and $\exp(-d^2)$ instead of $d^{2\gamma}$ in (1.1). Results were robust to these changes in specification, with the geometric mean probabilities not varying by more than .01. However, experimentation with $\gamma = ¼$ showed an appreciable deterioration in fit for some runs. On balance, the various estimates indicate that the model is robust to specification as long as the utility function remains concave or quasi-concave. Utility functions that are strictly convex near the ideal point, such as $\gamma = ¼$, should be avoided in empirical work as they are in theoretical work.

party candidates, we now investigate how these preferences relate, via a spatial utility model, to electoral choice between candidates and nonvoting. Specifically, we estimate the previously introduced models using the Euclidean distance produced by the unfolding. (We thus treat the computed distances as if they were errorless measurements.)

Since both the choice and the underlying thermometer ratings are measured by survey, one might think we would obtain overly favorable results since the thermometers, rather than being pure measures of affect or, in spatial terms, distance, are contaminated by the choice decision and strategic behavior. Such a possibility is mitigated by two factors. First, except for the 1968 data, the thermometers were collected considerably in advance of the elections, while the choice variable was measured in a post-election interview. Second, the distances generated by unfolding are based on the respondent's evaluation of a larger set of 12 or more stimuli; it is unlikely that the evaluations of noncandidates are affected by strategic considerations.

Moreover, the results of the logit estimation strongly indicate that thermometers are not just proxying for choice. Whereas the coefficients on distance in two-candidate races are virtually identical for the two candidates, coefficients differ between major and minor candidates in three-candidate races, indicating that considerations of viability intervene between the affect-distance measure (thermometers and unfolded distances) and the choice measures. We now turn to these results.

Basic Results of the Logit Estimation

We begin the presentation of results by indicating the estimates of the stochastic utility functions developed above. These estimates are shown in Table 2. With the exception of John Anderson in 1980, all the estimates are 8 to 23 times their estimated standard errors. The precision of the estimates simply reflects our very large sample sizes. As a consequence of the sample sizes, little of substantive interest can be learned from classical significance tests.

As an appropriate alternative to significance tests, we focus on the *geometric mean probability*. This is developed from the log likelihood, which is the sum of the logarithms of the probabilities given by the logit model for the actual reported voting decisions. To obtain the geometric mean, we divide the log likelihood by sample size to get an average log probability. Then, to return to a probability, we take the exponential of this average. The result is the geometric mean. It should be noted that the geometric mean, somewhat in the spirit of squared error, penalizes serious prediction errors. Thus, if two individuals voted for Reagan and the model said one would do so with probability .9 and the other with probability .1, the geometric mean of these two probabilities is only .30 and not .50. When we compare models, we will rely mainly on comparisons of the geometric means. If the geometric mean probability of one model is .44 and another .43, say, we will not conclude that there are important differences in explanatory power even if a statistical (likelihood ratio) test is highly significant.

TABLE 2

Coefficients of the Utility Function

Year		Intercept (β_1 or $\beta_1 + \delta_1$)	Distance (β_2)	Indifference ($\delta_2\beta_2$)	*N*
1968	Humphrey, Nixon	3.416	–5.260	—	1,348
		(0.161)	(0.270)		
1968	Wallace	7.515	–7.842	—	
		(0.851)	(0.818)		
1972	McGovern, Nixon	2.603	–3.629	0.321	2,115
		(0.137)	(0.158)	(0.161)	
1976	Carter, Ford	2.944	–3.538	–0.347	1,747
		(0.171)	(0.181)	(0.194)	
1976	Carter, Ford	2.7368	–3.432	—	
		(0.122)	(0.167)		
1980	Carter, Reagan	3.907	–4.4702	—	681
		(0.273)	(0.335)		
1980	Anderson	–0.190	–1.541	—	
		(0.681)	(0.770)		

NOTE: All results are for linear distance, two-dimensional unfolding, and equal coefficients on major candidates. In 1968, 1972, and 1976 the results are based on all respondents with a full set of thermometer ratings for the actual candidates. The 1972 pre-election unfolding was used. The 1980 results are based on the second-wave unfolding. All respondents with a full set of distances for all three waves were included. Standard errors appear in parentheses.

Validation and Stability of the Basic Choice Model

We have presented Table 2 for the case where the coefficients for major party candidates are constrained to equality. In fact, the geometric mean probabilities corresponding to the estimates in Table 2 are improved by less than .01 when the constraint is relaxed.[6] In Figures 1 and 3 we have plotted the probability contours derived from the coefficients when the constraint is not imposed. Nonvoting contours appear in Figures 2 and 4. It can be seen that the contours are still reasonably symmetric about the major candidates. For all practical purposes, we can take the major party coefficients to be equal, thus supporting our earlier theoretical position that the absence of strategic voting in two-candidate races should lead to equal coefficients.

The equality of coefficients, expected theoretically, is not, we point out, expected on the basis of empirical experience. In his detailed review of logit models, Amemiya (1981, p. 1491) indicates that where similar theoretical restrictions exist in economic models the fit is usually substantially improved by relax-

[6] Moreover, there is a considerable gain in the precision with which the utility function is estimated. When the coefficients are constrained, the estimated standard errors of the coefficients are roughly one-third less than the standard errors of unconstrained coefficients.

ing the constraint. Finding equal coefficients on the major candidates thus provides a fairly strong indication that our unfolding recovered a reasonable spatial map for presidential elections.[7]

It is also of interest that the coefficients for the pair of two-candidate elections, 1972 and 1976, are very similar despite the facts that no candidate repeated, Watergate intervened, and one election was a landslide, the other a dead heat in popular vote. This suggests that there is some stability not only in how the CPS samples treat the thermometer question but also in how thermometers relate to voting behavior. (We later discuss a quite different question, whether individual ideal points and candidate positions are stable.)

Comparison of the Spatial Model with Alternative Models

Having found that the coefficients conform to theoretical expectations, at least in the two-candidate case, we now assess the explanatory power of the model. As benchmark lower bounds for the geometric means one would expect to obtain, we consider *equiprobability* and *marginal* probabilities. Under equiprobability, with two candidates and abstention, for example, the probability of each choice (and the geometric mean) would be .333. In contrast, using the marginals, one assigns all individuals choice probabilities equal to the marginal proportions in the sample. In addition to these benchmarks, we consider a serious competitor. Rather than computing Euclidean distances via unfolding, we insert

[7] In recent theoretical and empirical work, Hinich and colleagues have attempted to incorporate valence and ascriptive issues, issues over which the candidates have no control of their position, into spatial models. Enelow and Hinich (1982) have presented a model in which voter utility depended on one issue dimension and one non-issue dimension. On the issue dimension, voters were placed into two groups, all voters in the same group having identical ideal points. The non-issue part of the model followed a random process. When the variance of this process is zero, the winning candidate has a position equal to the ideal point of one group. When the variance is large, candidates converge to a weighted mean of the two group positions. Because of the random process, however, we believe the Enelow-Hinich model is empirically indistinguishable from our stochastic utility model of pure spatial voting. If we had no abstention and just two groups of voters in our model, $\beta_1 = 0$, one group located at -1, another group at least equal to the first group's size located at $+1$, and $\gamma = 1$, one can readily show that there exists a symmetric equilibrium position $x \in [0,1)$ and $dx/d\beta_2 \leq 0$, reproducing the Enelow-Hinich result exactly. As to the empirical paper by Cahoon, Hinich, and Ordeshook (1978), their unfolding method treats valence issues by simply allowing for an intercept in the metric. Whereas we have the relationship that $d^2 = (x - z)'(x - z)$, where x and z are the Euclidean coordinates of the voters and candidates, they allow for $d^2 = a + (x - z)'(x - z)$, where a is a valence intercept. Clearly, the extra degress of freedom will improve the fit. We have omitted them for parsimony. However, even if included, we would still argue that utility should depend solely on distance as calculated from the unfolding program and not on other characteristics.

Our logit results indeed support our not including a valence dimension in the original unfolding. If there was an important bias in favor of one candidate because of valence considerations, the bias should have been reflected in the utility estimates. Moreover, the estimates provide some support for using metric as against nonmetric unfolding since Rabinowitz (1978) used nonequidistant cutting lines for his voting predictions. The cutting lines corresponding to a logit model are equidistant if and only if the coefficients are equal.

the raw thermometers for each candidate directly into the d terms in the expressions introduced above.

Comparisons between the various models are shown in Table 3. Both the spatial and raw thermometer estimates substantially outperform the marginals and equiprobability. The more interesting comparisons are between the spatial and thermometer models.

Analyzing similar thermometer data collected in France in 1967, Pierce (1981) argued that direct preference measures based on raw thermometers would give better predictions of choice than a spatial model based on left-right perceptions. Rosenthal (1981) demonstrated that Pierce's methodology was not appropriate for evaluating either type of model. Moreover, Rosenthal argued that a Euclidean representation of the thermometer data might lead to a spatial model that predicts as well as raw thermometers.

For American data, our results inform this debate. The spatial model outperforms raw thermometers in four of the seven interviewing situations, and it does better in the two cases where the geometric mean probabilities differ by more than .01 (1972, pre-election, and 1980, first wave). Furthermore, the preferred spatial models are also of low dimensionality. Except in 1976, two-dimensional spatial models outperform three-dimensional models. For 1980, we even found an interesting case where a one-dimensional model gave the best fit. Because both Carter and Reagan were at more extreme positions than nearly all voters and because Anderson's spatial position was extremely close to Carter's, one-dimensional distances were highly intercorrelated. Singularity conditions pre-

TABLE 3

Geometric Mean Probabilities for Best Fitting Spatial and Thermometer Models

	1968 Post	1972 Pre	1972 Post	1976 Pre	1980[a] 1	1980 2	1980 3
Spatial	.433	.476	.484	.423	.413	.396	.380
Thermometer	.436	.434	.477	.418	.391	.401	.386
Marginal	.278	.347	.348	.337	.349	.309	.297
Equiprobability	.250	.333	.333	.333	.333	.250	.250
Number of candidates	3	2	2	2	2	3	3

NOTE: All results are for models with coefficients unconstrained. All 1968 and 1972 spatial results are for two-dimensional squared distances. The 1972 thermometer results are for linear distance. The 1976 results are for linear distance and for a three-dimensional spatial configuration. All 1980 thermometer results and the spatial results for the first two panels are for linear distance. The spatial result for the third panel is for the exponential of distance and for linear distance on the major candidate and squared distance for Anderson (two models tied for best likelihood). All 1980 spatial results are based on two-dimensional configurations.

[a] John Anderson was not included in the thermometers on the first wave of the 1980 panel.

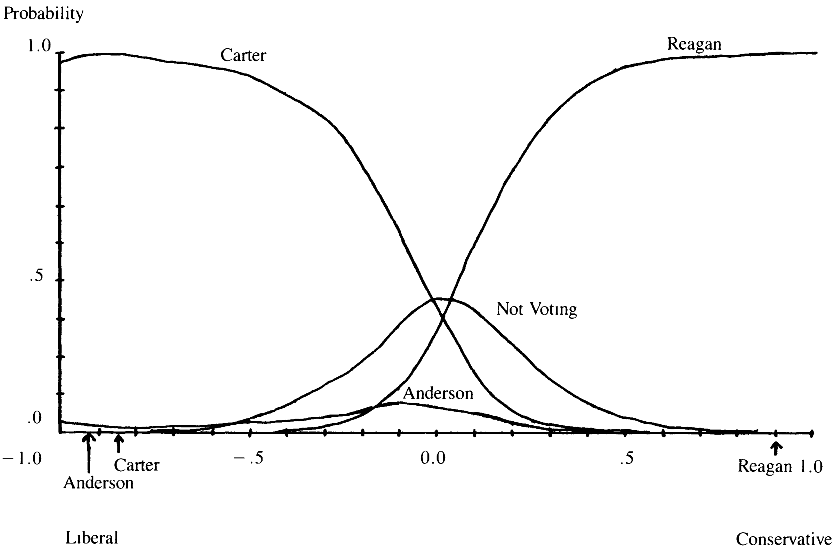
vented convergence in all runs except one. In that case, on the second wave, a one-dimensional model gave the highest geometric mean. The estimated choice probabilities along the dimension are graphed in Figure 5.

We conjecture that spatial models are as successful as raw thermometers in explaining electoral choice because the thermometers are noisy measurements. In positioning a respondent relative to a given candidate on the basis of the respondent's *set* of thermometer rankings and not just the ranking for the candidate, one smooths the error. The basic, commonly shared ideological space is, we further conjecture, of very small dimension. Thus, when one unfolds the thermometer data into higher dimensional spaces, one is basically reintroducing the noise in the original measurements.[8] These conjectures are consistent with the observation that the three-dimensional model typically does not predict as well as a two-dimensional model. The fact that the spatial models do not do even better than thermometers probably reflects the ability of the thermometers to measure, as Pierce (1981) and Fiorina (1981, p. 154) suggest, nonspatial aspects of preference as well as noise.

We emphasize, though, that regardless of nonspatial aspects, thermometer data are consistent with the basic assumptions of spatial theory. For 1980, for example, we do as well by replacing three thermometer ratings for the candidates by a single coordinate, based on the *theoretical restrictions of the Euclidean metric and a common perception of candidate locations for all individuals*. Sim-

FIGURE 5

One Dimension, 1980



[8] See Weisberg (1968) for a discussion of literature demonstrating that when error is present a variety of multivariate techniques will find excessive dimensionality.

ilarly, in 1968, we replace three thermometers with two coordinates. True, in 1972 and 1976, we require as many coordinates as thermometers (plus four candidate coordinates), but, unlike the thermometers, the permissible "distances" are highly constrained by the theoretical restrictions. In brief, by matching the thermometer probabilities in a relatively parsimonious fashion, we have shown that the relationship of thermometers to choice is predominantly accounted for by a spatial model.

Prediction Analysis of Individual Cases

While we have analyzed our spatial model as a probabilistic choice model, we have not yet addressed whether the model succeeds in predicting individual observations. Nor have we indicated the relative success of the model in predicting major-party candidates, third-party candidates, and nonvoting.

One way of seeing how well the "actuarial" predictions (see Hildebrand, Laing, and Rosenthal, 1977, pp. 112–14) do in predicting individual outcomes is the maximum probability approach. For each respondent, one finds the choice with the greatest predicted probability; one predicts this choice will be the actual choice. The cross-classifications of predicted versus actual choice are shown in Tables 4–7.

Inspection of the tables reveals that the predictions are quite accurate when one deletes the row and column corresponding to nonvoting and, in 1980, when one also deletes the row and column for Anderson. In other words, the predictions

TABLE 4

1968, Linear Distances, Two Dimensions, All Respondents

Contingency Table

	Predicted Choice				
Actual Choice	Nonvoting	Humphrey	Nixon	Wallace	*N*
Nonvoting	44	130	140	23	337
Humphrey	28	387	33	3	451
Nixon	41	20	392	10	463
Wallace	31	0	17	59	107
					1,358

∇_{ij} Analysis

	Predicted Choice			
Actual Choice	Nonvoting	Humphrey	Nixon	Wallace
Nonvoting	—	.02	.03	.02
Humphrey	.41	—	.83	.90
Nixon	.16	.89	—	.69
Wallace	−1.73	1.00	.63	—

TABLE 5

1972, Linear Distance, Two Dimensions, Thermometer Respondents

Contingency Table

Actual Choice	Predicted Choice: Nonvoting	McGovern	Nixon	N
Nonvoting	129	149	290	568
McGovern	74	411	67	552
Nixon	81	30	884	995
				2,115

∇_{ij} Analysis

Actual Choice	Predicted Choice: Nonvoting	McGovern	Nixon
Nonvoting	—	.06	.13
McGovern	0.0	—	.79
Nixon	0.39	.89	—

TABLE 6

1976, Linear Distance, Three Dimensions, All Respondents

Contingency Table

Actual Choice	Predicted Choice: Nonvoting	Carter	Ford	N
Nonvoting	46	246	173	465
Carter	69	495	96	660
Ford	65	62	514	641
				1,766

∇_{ij} Analysis

Actual Choice	Predicted Choice: Nonvoting	Carter	Ford
Nonvoting	—	−.16	.16
Carter	−.03	—	.67
Ford	.01	.79	—

of the model are highly accurate when they concern choice among candidates, major or minor, whose positions have been well established.

The separation between the Humphrey and Wallace electorates is particularly sharp. No one predicted to vote for Humphrey claimed to have voted for Wallace,

TABLE 7

1980, Second Panel, Linear Distance, One Dimension, All Respondents, Reagan Coefficients = Carter Coefficients

Contingency Table

	Predicted Choice				
Actual Choice	Nonvoting	Carter	Reagan	Anderson	N
Nonvoting	28	63	95	0	186
Carter	35	143	36	0	214
Reagan	24	18	260	0	302
Anderson	1	22	13	0	36
					738

∇_{ij} Analysis

	Predicted Choice			
Actual Choice	Nonvoting	Carter	Reagan	Anderson
Nonvoting	—	−.02	.07	—[a]
Carter	−.37	—	.69	—[a]
Reagan	.33	.82	—	—[a]
Anderson	.77	−.83	.34	—[a]

[a] ∇ not defined because of zero marginal.

and only three voters predicted to vote for Wallace reported themselves as having voted for Humphrey. There is somewhat more intermingling for the Nixon and Wallace electorates, suggesting that Nixon and Wallace competed primarily with one another. Without the Wallace candidacy, Nixon may well have won a victory in 1968 comparable to his 1972 landslide.[9]

Third Party Candidates

Our ability to explain the Wallace vote as well as the vote for the major candidates indicates that any effects from strategic voting are proxied for adequately by allowing separate coefficients for minor-party candidates. For Wallace, these coefficients result in his probability contours (see Figure 1) forming a *mesa*. Those individuals with ideal points close to the Wallace position were very likely to vote for him, but the probability drops abruptly once a distance threshold is passed. Compared to Humphrey and Nixon, there is a very rapid drop from the .75 to the .35 contour for Wallace.

The rapidity with which Wallace was deserted has the obvious interpretation that he was a nonviable candidate. That he had the level of support he did obtain

[9] Kiewiet (1979) develops similar findings from the analysis of ordinal rankings based on raw thermometers.

may have resulted from his supporters tending to have more sharply peaked utility functions (corresponding to the -7.8 coefficient on distance in Table 2 for Wallace compared to the -5.3 for Humphrey and Nixon). Flash parties may draw extremists, and the more rigid ideology of extremists (Lerner, 1957) may result in sharply peaked utility functions. While we can't disentangle the various effects, further support for the viability hypothesis comes not only from Wallace's drop in the polls prior to the elections but also from Kiewiet's (1979) analysis of voters who defected from Wallace even though they assigned him a higher thermometer than Humphrey or Nixon. Kiewiet found that the nondefectors were disproportionately concentrated in states like Alabama where Wallace had a strong chance of carrying the state. If individuals with ideal points close to the Wallace position are also disproportionately concentrated in those states, our results would be consistent with Kiewiet's.

In contrast to Wallace, we account for none of the Anderson votes since, by the maximum probability criterion, no one is predicted to vote for Anderson (Table 7). This fact is represented in Figure 5, which contains the curve of predicted probabilities for Anderson.[10]

Anderson is also characterized, again in distinction to Wallace, by a flat response to distance. In fact, his distance coefficient (-1.5) is by far the smallest in magnitude of any of our estimates (see Table 2). To some extent, the contrast with Wallace may reflect Anderson's supporters not being as extremist as Wallace's were—Brie-and-Chablis instead of redneck. But we believe a more relevant explanation of the results lies in Anderson's ephemeral appearance on the national political stage. Anderson was not salient even to the CPS pollsters until the second panel wave in July. At that time, only one-third of the respondents could assign him a thermometer rating. His unfolding fits, .10 in two dimensions and .15 in three, are very poor, especially in comparison to the fits averaged over all 14 candidates (Table 1). Moreover, unlike Wallace's differentiated position, Anderson's position is so close to Carter's (even in two- and three-dimensional fits), that it is not possible to discriminate his electorate. As should be expected, the spatial model works well only when candidates establish sharp public images.

Nonvoting

In distinction to the model's performance for established candidates, we had only limited success in seeking to explain nonvoting with a spatial model. We will only comment briefly on nonvoting from indifference. As can be discerned from Table 2, adding indifference terms did not always result in coefficients with the expected positive sign. Moreover, improvements in the geometric mean probability were miniscule. Like Brody and Page (1973) and Rosenthal and Sen (1973), we have encountered difficulty in separating indifference from alienation effects. In our case, the reason for this result is quite clear. Since the candidates

[10] The peaking of the Anderson probability to the right of his position occurs because his estimated distance coefficient has a smaller magnitude than that of Carter (see Table 2).

are at the periphery of the distribution of voters, the voters who are most distant (alienated) from any candidate also tend to be those who have equal utilities (indifferent) for the candidates. Consequently, we will continue solely with a discussion of the alienation model, bearing in mind that alienation and indifference are virtual synonyms in the present context.

Our observation about our relative lack of success in predicting nonvoting can be made more precise via a ∇_{ij} analysis (Hildebrand, Laing, and Rosenthal, 1977, p. 94). Note that in the upper sections of Tables 4–7 we are predicting that the cases will lie on the major diagonal. The other cells are error cells. The proportionate reduction in error for each error cell can be measured as:

$$\nabla_{ij} = 1 - \frac{f_{ij}}{f_i \cdot f_{\cdot j}} = 1 - \frac{\text{Observed Error}}{\text{Expected Error}}$$

where f_{ij} is observed frequency in the cell and $f_i\cdot$ and $f_{\cdot j}$ are the corresponding row and column frequencies. When no errors occur, ∇ is 1. When an independence model holds for the sample, $\nabla = 0$. Negative ∇ can result when the predictions are worse than the independence benchmark.

Examining the error cells for the top row and leftmost column for each table for the four years indicates that the ∇ values involving nonvoting are much worse than those involving both major candidates. In fact, several negative values appear. Note that in 1968 predicted nonvoters, in fact, voted much more heavily for Wallace than expected. In 1976 and 1980, the predicted nonvoters voted more heavily for Carter than expected. Only in 1972, do we find a pattern of positive ∇'s for nonvoting, but even here they are far below the levels obtained for the candidate predictions.

The mixture of positive and negative ∇'s for nonvoting obscures the modest success the spatial model does have in predicting nonvoting. Rather than considering individual error cells, we can consider the set of error cells for the first column in the cross-classifications. These are the errors for the prediction "Respondents whose highest probability is nonvoting will not vote." This ∇ is, in fact, just a weighted average of the ∇_{ij} values for error cells in the column, the weights being the precision of each cell (Hildebrand, Laing, and Rosenthal, 1977, p. 95). The ∇'s for this prediction for the four elections are, successively, +.08, +.25, −.01, and +.09. In turn, we can compute the same ∇ statistic for all four elections pooled. Again, this ∇ is a weighted average of the ∇'s for the four separate elections. This global ∇ is .128. As its estimated standard error is only .002, there is no doubt that on the whole the spatial alienation model is relevant to nonvoting, but the proportionate reduction in error is low, and in the case of 1976 even slightly negative.

The maximum probability approach, however, has a limitation with respect to events that are relatively infrequent. Since their probability is rarely high, they are underpredicted. In examining Tables 4–7, it can be seen that there are always far fewer "predicted" nonvoters than actual nonvoters. To see why this occurs, consider Figure 5. There it can be seen that not voting has the maximum probability only for voters whose ideal points are located in a very narrow range at the center of the space. All other voters, by the maximum approach, are predicted to vote for Carter or Reagan.

As an alternative to maximum probability analysis, we cross-tabulated voter participation with the probability of nonvoting estimated by the logit model. The results are shown in Table 8. The results for the first three elections are consistent with the ordering of the overall ∇ values—nonrelationship in 1976, a weak positive association in 1968, and a stronger positive association in 1972. Voters are relatively frequent when nonvoting probabilities are below .15, and nonvoters are relatively frequent when the probabilities are above .30.

A striking contrast with the maximum probability results occurs for 1980, which now shows a relatively strong linkage between the nonvoting probability and actual turnout. The contrast found for 1980 but not for the other years relates to the fact that a one-dimensional model was used only in 1980. Because in one dimension respondents cannot be located in "corners" of the space remote from the candidates, nonvoting probabilities never get very large. No nonvoting probability exceeded .35 for 1980, while values as high as .90 occurred for 1968 and 1972 and as high as .95 for 1976. (See the probability contours in Figures 2 and 4.) Most of the 1980 middle-of-the-road voters with nonvoting probabilities in the .30 to .35 range would have been predicted to vote by the maximum probability approach.

To summarize our results on turnout, examining the probabilities of nonvoting directly has strengthened our findings based on the maximum probability approach. Nonetheless, spatial alienation makes at best a modest impact on turnout. Perhaps our inability to discover stronger effects resides in the notorious overreporting of participation by survey respondents, as discussed in Wolfinger and Rosenstone (1980). Yet the unequivocal relationships of education and income to turnout reported by these authors suggest that the primary determinant of turnout resides in long-term sociological factors. Indeed, a similar conclusion was reached by Rosenthal and Sen (1973) using aggregate data—hence, free from the overreporting problem—for French legislative elections. Although they found that absentions responded to the competitiveness of each constituency's contest, spatial factors were not reflected in abstentions but only in spoiled ballots, a form of "non" voting not present in the United States. Our findings run strongly counter to the belief that turnout in the United States is low because the candidates do not represent sharply differentiated alternatives (e.g., socialism). First of all, we argue from our unfolding that the candidates are sharply differentiated with respect to the American electorate (if not a European one). Second, even though the candidates are diffentiated, the people most distant from them are not overly likely to abstain.

The Dynamics of Voter Choice

Until recent work by Hinich and Pollard (1981), spatial theory had been developed on the strong assumption that voters had fixed spatial positions. Can-

TABLE 8

Voters and Nonvoters by Predicted Probability of Not Voting

Probability of Not Voting	1968		1972, Pre		1976		1980, P 2	
	Voters	Nonvoters	Voters	Nonvoters	Voters	Nonvoters	Voters	Nonvoters
0–.15	22	17	10	4	11	9	21	10
.15–.30	53	46	46	34	74	79	42	27
.30–1	25	38	45	62	16	12	37	62
Total (%)[a]	100	101	101	100	101	100	100	99
N	1,021	337	1,577	599	1,301	465	552	210

Nø2: Results based on constrained coefficients, linear distance, two dimensions (except 1980, one dimension), alienation model. All entries are percentages.
[a] Entries may not sum to 100 due to rounding.

didates then competed against the backdrop of fixed voter positions. An alternative, similarly strong, view is that candidates have little spatial mobility as a result of commitments to interest groups, the nominating process, and public positions taken prior to candidacy. Any changes in voting behavior, then, would be due to changes in individual spatial positions.

Our work can begin to examine these two contrasting views of the electoral process. Since (1) the orientation of the unfolding axes is entirely arbitrary and (2) the origin of the space is defined simultaneously by the positions of voters and candidates, we can't proceed by comparing the separate unfoldings previously analyzed. Panel data do permit us to study the dynamics of spatial positions. Ideally, one would like to be able to estimate the movement of both candidates and voters simultaneously. However, when one allows both candidate positions and voter positions to vary over waves of a panel, one is simply unfolding each panel separately. The results remain noncomparable. To make over-time comparisons, something must be held constant.

The two obvious solutions for this problem are either to estimate a constant set of candidate coordinates for the panel and allow individual coordinates to vary across waves or to estimate a constant set of individual coordinates and allow candidate coordinates to vary. We used two sets of panel data. One was a panel the CPS interviewed in both 1972 and 1976. The other was the panel interviewed on three occasions prior to the 1980 election. We carried out both the candidates constant (C) and individuals constant (I) unfoldings for these two situations. For the 1972–76 panel, however, even when other candidates' positions were held constant, Richard Nixon's was allowed to vary since we believe that Watergate made a constant position for Nixon unrealistic.[11]

From an unfolding viewpoint, there is little or no deterioration in fit in assuming that *candidates have constant spatial positions*. The 1972–76 C unfolding gives a fit intermediate between the fits for the two years unfolded separately. (See Table 1 for the numerical results.) The 1980 C fit is also intermediate, being close to the average of the P2 and P3 fits. It compares somewhat less well to the P1 fits, but it should be remembered that P1 did not include the noisy thermometer for Anderson. Consequently, assuming that candidates have constant spatial positions—even across periods as long as four years—does not affect our ability to recover thermometer scores.

When individuals are held constant, there is much more deterioration in fit. With the exception of the one-dimensional unfolding of the third 1980 panel, the individual constant unfolding always fits worse than the separate unfoldings. This is to be expected, however, since we are estimating far fewer parameters in an I unfolding than in a C unfolding (see Table 1). The individual "movements" may be without great significance for predicting choice behavior. Moreover, our basic

[11] Note that for some stimuli (candidates) we have only one set of distances. For example, no thermometer was elicited for Carter or Ford in 1972. This is no problem for the analysis as long as we have some stimuli (Humphrey, Rockefeller) who were included on each wave.

concern is to explain choice and not just to represent thermometers spatially. Consequently, we turn to examining how well the C and I unfoldings perform in the logit model.

In analyzing the logit results for the panel analysis, we begin with the short-run dynamics of the 1980 campaign. Although important events may well have influenced voting behavior between July and October, such events are not reflected in our estimates. Constraining candidates to be constant, constraining individuals to be constant, or allowing both to move (separate unfolding) produces no important difference in results between the second and third waves. In fact, the geometric mean probabilities are, for some comparisons, higher for the second wave than for the third. (See Table 9. See also Table 2, which reports results for all respondents in a wave and not just those present in all three waves.)

Results are quite different for the first wave. The only comparison we can make holds the candidates constant. The other comparisons are ruled out by the absence of Anderson thermometers on the first wave, making it impossible to obtain February coordinates for Anderson. When we do hold the candidates constant and use the individuals' February coordinates in estimating the voting behavior reported in November, the geometric mean probability is roughly .04 lower than when July or October coordinates are used. This result, being based on a C

TABLE 9

Geometric Mean Probabilities from 1980 Panel Analysis

Basis of Estimates	Dimensionality of Unfolding: One	Two	Three
Candidates constant with individual's February positions	—[a]	.325	.331
Separate unfolding of July data	—[a]	.370	.365
Individuals' constant with candidates' July positions	.386	.372	.369
Candidates constant with individuals' July positions	—[a]	.362	.362
Separate unfolding of October data	—[a]	.365	.363
Individuals constant with candidates' October positions	—[a]	.368	.368
Candidates constant with individuals' October positions	.375	.372	.373

NOTE: $N = 681$, linear distance, equal β's for Carter and Reagan; includes respondents who were interviewed in all three waves with no missing distances and who either voted for Carter, Reagan, or Anderson, or did not vote; February results include only candidates constant because Anderson thermometers were not elicited.

[a] Singularity encountered; no estimates available.

unfolding, is certainly consistent with the view that individuals change their position during a campaign.[12] On the other hand, we cannot rule out movement by the candidates, since if candidates move but not individuals, this movement will affect individual coordinates in a C unfolding. We especially do not wish to rule out candidate movement since we conjecture that a major change in a candidate's perceived position occurs when the candidate is actually nominated or perceived to be certain of nomination. Upon nomination, the candidate would tend to move closer to the spatial position that voters generally associate with the candidate's political party. In any event, the positions of voters and candidates, as measured by thermometers, appear to be relatively stable over the last few months of a campaign but not over longer periods.

The absence of any long-term stability is even more apparent when the 1972–76 results are examined, as shown in Table 10. It is again the case that there is no deterioration in the geometric mean probabilities derived earlier from separate unfoldings for each year. As long as we use coordinates corresponding to a given year to estimate that year's choices, we are not harmed by constraining either candidates or individuals to constant positions in time.

Yet one of these two sets of actors must have spatial mobility. Consider estimating the 1976 choices on the basis of 1972 individual coordinates from the C unfolding. In this case, the geometric mean probability declines very substantially, falling by .06 to .07 to a level hardly different from either the marginal or equiprobability. Clearly, there was very substantial change over this four-year period. Unfortunately, as neither Carter nor Ford thermometers had been obtained for 1972, we could not conduct the corresponding forward prediction using 1972 candidate coordinates with constant individual coordinates. In any event, while we cannot pinpoint whether candidates or voters account for most of the changed positions, long-term spatial movement is strongly indicated.

The data provide some indication of the nature of the movement. The geometric mean probabilities deteriorate less for the one-dimensional model than for the higher dimensional models when we substitute 1972 coordinates for 1976 coordinates. The net fall in the geometric mean probability is .072 in two dimensions and .066 in three dimensions but only .045 in one dimension. (Given our sample size, differences of this amount in geometric mean probabilities are significant in terms of the corresponding log likelihoods.) We would thus tentatively argue that spatial dynamics are not so much changes along the basic liberal-conservative dimension but changes on issues that are ephemeral (southern favorite son treatment of Carter and Wallace). In any event, long-run instability cannot

[12] Other evidence that supports the view that politicians are more stable than voters arises in the research of Poole and Daniels (1982) and Rosenthal and Sen (1977). The former report that members of Congress have stable positions over periods as long as 20 years, even when they shift from the House to the Senate. The latter were able to predict voting behavior in French legislative elections assuming fixed spatial positions for parties and stable voter decision rules but changing voter ideal points.

TABLE 10

Geometric Mean Probabilities from the 1972–76 Panel Analysis

	Dimensionality of Unfolding		
	One	Two	Three
1972 Choices (McGovern, Nixon, Nonvoting), N = 1,241			
Separate unfolding	.456	.474	.469
Individuals constant, candidates' 1972 positions	.455	.460	.466
Candidates constant, individuals' 1972 positions	.455	.465	.475
1976 Choices (Carter, Ford, Nonvoting), N = 1,126			
Separate unfolding	.424	.430	.436
Individuals constant, candidates' 1976 positions	.413	.420	.426
Candidates constant, individuals' 1976 positions	.423	.426	.425
Candidates constant, individuals 1972 positions	.378	.354	.359

NOTE: Linear distance, equal β's. For each year, analysis includes respondents making one of the listed choices and with distances to both candidates.

be ignored. Thermometers collected more than six months before an election are virtually unrelated to choice.

Conclusion

If we take a snapshot of the American electorate at any point in time, the behavior of voters appears to correspond reasonably well to the assumptions of spatial theory. Voters choose among candidates, probabilistically, in terms of distance. The underlying spatial utilities appear, from the two-candidate contests, to depend, as spatial theory would claim, only upon distance and not upon the candidate. When a visible but unviable third-party candidate, such as George Wallace, is also present, the coefficients of the utility function adjust to proxy for the viability effect. Nonvoting from alienation, as originally formally modeled by Hinich and Ordeshook (1969), has a modest effect on turnout.

The dynamics of these spatial portraits have proven harder to pin down. Models where either individual positions or candidate positions (but not both!) are held constant in time do as well as models where all positions are free to vary. At least in 1980 the last few months of the presidential campaign appears to exhibit no dynamics at the level of the spatial measurements used here. But as one retreats further from the election in time, one's ability to explain choice diminishes. Individual coordinates from the February wave of the 1980 panel were much less successful in the logit model of choice than were those from July and October, while 1972 individual coordinates barely improved on an equiprobability model for 1976.

Having sought to learn something about the relative magnitude of candidate mobility versus voter mobility, our foray into dynamics has only succeeded in showing that there is important spatial movement from one or both of these two possible sources. This effort, to our knowledge, was the initial empirical spatial model analysis of panel data. We hope it will eventually stimulate an answer to our original question. We also hope others will build upon our introduction of a formal model of choice and our use of the full range of CPS U.S. thermometer data to estimate this model.

Of course, the usual objections to the use of survey data are pertinent—panel effect, bandwagon effect in the post-election survey, chronic underreporting of nonvoting. We have deliberately avoided these issues in order to present a basic methodology for the spatial analysis of choice behavior. We have also avoided these problems because they have no complete solution. One is generally unable to turn partial obfuscation into total truth. We, therefore, see the next challenge as one of examining aggregate election statistics, continuing the Rosenthal and Sen (1973, 1977) research with the aggregate version of the logit model.[13] Encouraged by the ability of the spatial model to explain reported choice, we need to learn more of its ability to explain actual choice.

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APPENDIX

Estimation

The assumption that the e_{ij} terms are independently distributed as the logarithm of the inverse exponential allows the probability of individual i voting for candidate j to be written for the nonvoting from alienation case as:

$$P_{ij} = \frac{\exp(\beta_{1j} + \beta_{2j} d_{ij}^{\,2\gamma})}{1 + \phi} \qquad (j = 1, \cdots, q) \tag{2.1}$$

and the probability of individual i not voting as:

$$P_{i,q+1} = \frac{1}{1 + \phi} \tag{2.2}$$

where

$$\phi = \sum_{j=1}^{q} \exp(\beta_{1j} + \beta_{2j} d_{ij}^{\,2\gamma}).$$

[13] Our earlier working paper (Poole and Rosenthal, 1982) presents the aggregation methods and shows more extensive tables of results for logit estimation of the U.S. survey data.

The likelihood of the individual choices, therefore, is:

$$L = \prod_{i=1}^{p} \prod_{j=1}^{q+1} P_{ij}{}^{C_{ij}}$$

where $C_{ij} = 1$ if individual i chose j, 0 otherwise, so that the log likelihood is

$$\ln(L) = \sum_{i=1}^{p} \sum_{j=1}^{q} C_{ij}\,(\beta_{1j} + \beta_{2j} d_{ij}{}^{2\gamma}) - \sum_{i=1}^{p} \ln(1 + \phi). \tag{3}$$

Taking the partial derivatives yields

$$\partial \ln(L)/\partial\beta_{1j} = \sum_{i=1}^{p} C_{ij} - \sum_{i=1}^{p} \frac{\exp(\beta_{1j} + \beta_{2j} d_{ij}{}^{2\gamma})}{1 + \phi} \tag{4.1}$$

and

$$\partial \ln(L)/\partial\beta_{2j} = \sum_{i=1}^{p} C_{ij} d_{ij}{}^{2\gamma} - \sum_{i=1}^{p} \frac{d_{ij}{}^{2\gamma}\exp(\beta_{1j} + \beta_{2j} + \beta_{2j} d_{ij}{}^{2\gamma})}{1 + \phi}. \tag{4.2}$$

Standard nonlinear maximum likelihood procedures may be applied to (3) and (4) to estimate the β's. Because the function $\ln(L)$ (3) is convex, a global maximum is readily obtained numerically (Dhyrmes, 1978, pp. 347–52). The restrictions on the β's imposed by spatial theory do not affect the convexity of $\ln(L)$. Estimation for the indifference model is a straightforward extension of the above procedure.

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