The empirical application of the spatial theory of voting in multiparty systems with random utility models

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Abstract

Spatial models of voting predominate in the formalization of political decisions and continue to be a growth industry in political science. But strict empirical applications of this theory have been rare. Only recently, conditional discrete choice models have been proposed to fill the gap between formal and empirical models and to predict the individual voting decision in multi-candidate/multiparty contests on the basis of the spatial model. This article highlights several flexible features of these models that are well known in transportation economics and applied marketing science but not yet discussed in the electoral studies community. Empirical illustrations are provided on the basis of (nested) multinomial logit. © 2000 Elsevier Science Ltd. All rights reserved.

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1. The theory of spatial voting: background

Since the seminal work of Harold Hotelling (1929), Duncan Black (1958) and Anthony Downs (1957), spatial models of voting have been the dominant paradigm in mathematical political theory. In such models, policy options are represented by points in a finite-dimensional vector space. Each voter has a utility function on this space, which is commonly assumed to be a decreasing function of the Euclidean distance from the voter’s ideal point. The spatial conceptualization of politics allows us to visualize easily a substantial number of classes of examples. Theoretically, the
main thrust of these models has been to define conditions under which majority rule yields an equilibrium on the political market. The most important result from this body of work is Black’s Median Voter Theorem. Its key assumptions are that policies are defined along a single dimension and voters have single-peaked preferences. The model is in general restricted to two-candidate, simple majority elections.¹

The least satisfying aspect of spatial voting theory is that it generally confines itself to theory. As stated by Mueller: “…the sophistication and elegance of the theoretical models of public choice far exceed the limits placed by the data on the empirical models that can be estimated” (Mueller, 1989, p. 193). So far, only a few examples of rigid empirical applications of spatial voting models exist.

In the following, I shall first critically discuss recent developments in the empirical applications of the spatial voting model in multi-candidate/multiparty² systems. Then I present the theoretical and econometric approach to predict individual decisions as well as market shares of parties. Third, based on a subsample of the 1990 German national election study, I shall illustrate several flexible methodological features of these models that have not been considered in electoral research.

2. Empirical applications of the spatial theory of voting

There are only a few empirical applications of the spatial theory of voting based on policy distances and representing a multi-candidate contest. Due to methodological restrictions, earlier studies (Aldrich, 1975; Enelow and Hinich, 1985; Enelow et al. 1986, 1995) limited their analyses to binary contests. Therefore, an assessment of the overall competitive structure was not possible. The market analogy was lost in the empirical modeling. In more recent papers, however, there is a tendency to use statistical models that are able to handle polychotomous alternative sets (Iversen, 1994; Whitten and Palmer, 1996; Thurner and Eymann, 1997; Alvarez and Nagler, 1998; Dow, 1998; Quinn and Martin, 1998; Schofield et al., 1998; Thurner, 1998; Thurner and Pappi, 1998a,b). However, a number of methodological questions remain to be answered. As a contribution to the ongoing discussion, I would like to discuss several innovative theoretical and methodological features of the (conditional) discrete choice model.

1. Recent empirical applications of the spatial voting model have, first, to be distinguished as to whether they are oriented toward the classical Davis–Hinich–Ordeshook (Davis et al., 1970) variant where political dimensions are essentially constituted by policies, or whether they are oriented toward the neo-dowsonian

¹ During the last years there have been considerable efforts to relax certain of these restrictive assumptions. For example, attempts were made to determine the general implications of multi-candidate competition (cf. Eaton and Lipsey, 1975; Shepsle, 1991; Anderson et al., 1994) under diverse electoral systems (Cox, 1990). Rather sobering have been the results when exploring equilibria in the more general, and more realistic, K-dimensional policy space, cf. Plott (1967) and Schofield (1978); Mueller (1989) provides an overview of this literature.

² In the following I use the terms ‘multiparty’ and ‘multi-candidate’ interchangeably.
approach (Enelow and Hinich, 1984) where latent dimensions are essentially assumed to be ideological. Iversen (1994), Thurner and Eymann (1997), Alvarez and Nagler (1998) follow the classical approach; Dow (1998), Quinn and Martin (1998) and Schofield et al., 1998 follow the neo-downsonian approach and combine it with conditional discrete choice theory. A discussion of both approaches is provided by Quinn and Martin (1998), concluding with a preference for the latent space approach. However, this conclusion will not attract unanimous approval. Despite the admitted ‘common scale’ problem of the classical approach, the classical approach conceived as a reduced-form model (cf. Enelow and Hinich, 1985) seems more promising. The latent space approach suffers at least from the following shortcomings: (a) variants that construct the political space on the basis of a single input indicator (Dow, 1998) run the risk of interpreting a general psychological dimension as decision criterion for a rational choice; (b) constructing a multidimensional space on the basis of one (!) indicator signals more the presence of a heterogeneous sample population than an individually perceived multidimensional space; (c) the common space is constructed on the basis of inter-individual correlations; (d) the extraction of factors always implies a substantial amount of unexplained variance; (e) variants matching mass and elite survey data (Iversen, 1994; Schofield et al., 1998; Quinn and Martin, 1998) on multiple issue preferences may present an advantage for a graphical representation of voters and candidates in the same low-dimensional space. However, for the explicitly stated purpose of explaining individual voter choices (cf. Quinn and Martin, 1998), this approach seems inappropriate as well, since the voters’ decision calculuses—including well-known biases such as projection and persuasion which nevertheless have behavioral consequences—are simply cut.

2. Another question concerns the specification of the voters’ preferences. Iversen (1994), following the formal model of Davis et al. (1970), uses composed policy distances—i.e., distances added up over $k$ dimensions are composed in a single predictor, thereby assuming identical weights for each dimension. Thurner and Eymann (1997) and Alvarez and Nagler (1998) use decomposed distances—i.e., every policy distance is represented by at least one predictor, and the weights of every dimension are estimated. A further distinction has to be made concerning the assumed metric: most ideal-point models assume that preferences are negatively related to the square of the Euclidean distance between the parties’ positions and the individual’s ideal point. However, as Ordeshook (1986) and Laver and Hunt (1992) have shown, distances can also be assumed to be City-block versus Euclidean or linear versus squared loss functions. Furthermore, the classic spatial voting model includes the restrictive assumption that the weights ascribed to the policy dimensions are identical across all voters, which implies that the electorate is homogeneous (cf. Davis et al., 1970, p. 434). Relaxing this restrictive assumption would allow the voters to have individual-specific or at least group-specific saliency weights for an issue. Assump-

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3 Most often this is constituted by the feeling thermometer scores.

4 For a criticism of this procedure in the context of attitude consistency research, see Judd and Krosnick (1989).
tions over spatial preferences should be empirically tested, for “…assumptions about the relationship between dimensions are assumptions about the preferences of individuals in the system, not assumptions about the system itself” (Laver and Hunt, 1992, p. 18).

3. As an explicitly multidimensional model, the classic approach is open for all kinds of decision criteria. Davis et al. (1970) pleaded for integrating non-policy criteria such as partisan attachment into the spatial voting model. Their statement “the relative importance of issues, compared to image and partisan bias remains an open question” (Davis et al., 1970, p. 429) applies also to recent empirical contributions.

4. It has been argued by Alvarez and Nagler (1998) that the simple multinomial logit (MNL)⁵ or probit model ‘collides’ with the inherent competitive logic of the spatial model. This is correct as long as it is intended to assess hypothesized movements of parties, since only conditional discrete choice models⁶ are able to model parties’ address change in the political space by considering their changing attributes/positions. However, it is quite possible by means of simple MNL and MNP to produce a state picture of the multiparty competition with a given distribution of voters’ ideal points and fixed party positions. First, MNL and conditional MNL models are algebraically ‘totally equivalent’ (Maddala, 1983, p. 42). Following transportation economics, which has established a useful terminology,⁷ conditional logit and simple MNL are special cases of a model with alternative-specific variables only (Kühnel, 1993). The fact that competitive aspects have been lost in MNL or MNP applications of electoral choice is not due to the inherent model logic but more so to the exclusive presentation of MNL beta coefficients.⁸ Due to the nonlinearity of the model, these coefficients are rather inappropriate for substantial conclusions. As a supplement to King et al.’s (1998) suggestions, so-called effect coefficients (Long, 1987; Kühnel, 1993) as well as partial derivatives and elasticities should be considered for use in electoral studies. Then, static substitutional patterns between parties can be described and explained also within the simple MNL and MNP framework, respectively.

5. Discrete choice models (simple as well as conditional variants) allow for varying individual alternative sets (McFadden, 1984, p. 1416). Therefore, if voters do not recognize all alternatives, it is possible to use their relevant alternative set and to preclude a modeling of competition, where in fact there is no competition at all. In a first application of this feature Thurner and Pappi (1998a) estimate simultaneous

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⁵ In the following I refer to multinomial logit models as MNL, to multinomial probit models as MNP.

⁶ Conditional multinomial logit and probit describe models where, in addition to individual characteristics, alternative-specific attributes can also be specified.

⁷ It differentiates between individual-specific (socio-economic or attitudinal) variables, alternative-specific variables and generic variables (Ben-Akiva and Lerman, 1985; Wrigley, 1985). Attributes of alternatives are captured by ‘generic variables’. However, this specification “imposes restriction of equality of coefficients on a more general model with alternative-specific attributes” (Ben-Akiva and Lerman, 1985, p. 22). Which specification actually applies has to be tested.

⁸ For a criticism of this practice, see King et al. (1998).
reaction functions for a national sample despite specifying varying alternative sets for regional subsamples. Future studies should invent survey instruments in order to directly elicit the voter’s evoked set.

6. Last but not least, Alvarez and Nagler’s (1998) plea for multinomial probit against MNL and nested multinomial logit models (NMNL) seems too apodictic and rash. Despite the recent improvement of MNP estimation techniques, there are also reasonable arguments for the adequacy of NMNL: (a) McFadden’s generalized extreme value model maintains the extreme value distribution. However, it relaxes the assumption of identically and independently distributed unobserved utility components. Therefore, it constitutes an ideal compromise; (b) it is sometimes useful to group together alternatives according to their degree of substitution and to make transparent the underlying decision structure and not to hide it away in the flexible correlation structure of disturbances in the MNP model; (c) there are a number of specification tests to determine whether MNL or NMNL applies and which nesting structure is adequate (McFadden 1981, 1984; Hausman and McFadden, 1984). Their combined application ensures a correct model specification; (d) the extreme value distribution has also some theoretical appeal: “If choices are made according to maximum utility, the extreme value distribution seems a rather natural choice for a discrete choice model, more so than the normal distribution” (Boersch-Supan, 1990b, p. 203); (e) empirical comparisons of MNP versus NMNL (Alvarez and Nagler, 1994; Quinn and Martin, 1998) do not yet provide unambiguous evidence for rejecting one or the other assumption on error distribution. Therefore, the statistical and theoretical philosophy of the NMNL model and its wide application for travel demand studies is worth being transferred to and tested in political science applications.

Before giving several empirical applications in the framework of the classical Davis–Hinich–Ordeshook model, I first briefly outline the formal model.

3. The spatial voting model formulated within multiattributive random utility theory

Starting with the standard spatial model I assume that all voters participate. Let \( Z \) be a subset of an \( N \)-dimensional Euclidean space, with \( Z \subset R^N \) representing the policy space. Each individual voter, denoted by \( i \), is assumed to have a well-defined utility function over this space. Her policy preferences are characterized by a finite point of maximum utility \( x_i \in Z \), called her ideal point or bliss point. Let the party system consist of \( n \geq 2 \) parties and the choice set therefore be \( A=\{a_1, a_2, \ldots, a_j, \ldots, a_J\} \). Each party \( j \), conceived here as a unitary actor, takes policy positions \( z_{jk} \in Z \) in the \( K \)-dimensional policy space \( Z \) with the dimensions being separable. Voters are assumed to have identical choice sets. They base their evaluation of the parties on the policy platforms of these parties. Partial utility \( i \) deriving from

\[ ^9 \text{The more general model would allow the voters to have individual-specific choice sets } A_i. \]
a perceived party’s policy position $z_{ijk}$ is denoted $u_i(z_{ijk})$. All voters vote for the platform closest to their most-preferred position and the utility of a policy location $z \in \mathcal{Z}$ decreases with the distance from $z$ to $x_i$. Hence, the utility associated with a party is the negative of the minimum distance. It is now commonplace in the theoretical literature to assume weighted Euclidean distances for their mathematical properties and for the intuitively appealing visualization of Euclidean spaces, but as Ordeshook (1986, p. 22ff.) and Laver and Hunt (1992, p. 15ff.) have pointed out, this is far from being the only possible assumption. Therefore, I suppose for the total utility function $U_i(z_{ij})$ the general Minkowski-metric, containing the negative of any disutility function that is radially symmetric about an ideal point $x$:

$$U_i(z_{ij}) = -\left(\sum_{k=1}^{K} \beta_{gk}|x_{i_k} - z_{ijk}|^r\right)^{1/r},$$

where $\beta_{gk}$ represents a weighting constant that determines the saliency of each of the $k$th policy dimensions and is assumed to be common to a mutually exclusive segment $g$. $r$ represents the order of the metric; so, for example, if $r=2$ the Euclidean metric is specified, while if $r=1$ the City-block metric is specified. Weights and the appropriateness of metric order should be determined empirically. They indicate how the voters trade-off closeness in one dimension against distance in another when evaluating the different polices on offer. I shall call this a decompositional multiattributive utility function. Another possible variant, chosen by Iversen (1994), is to determine a single, unweighted measure of the distance between ideal points and parties’ policy positions in multidimensional policy space, which would imply the following general specification:

$$U_i(z_{ij}) = -\left(\beta \sum_{k=1}^{K} |x_{i_k} - z_{ijk}|^r\right)^{1/r},$$

where $\beta$ represents a constant that weights the impact of the platform as a whole, assuming equal weights of each dimension. I refer to this as a compositional, unweighted multiattributive utility function. It will be an empirical question which type of utility function will adequately reproduce the voters’ preferences.

Following an approach originally proposed by McFadden (1974) for modeling discrete choice behavior, voters are assumed to vote probabilistically as a function of their policy preferences and perceived distances. One of the central features of this so-called conditional logit model, also labeled ‘random utility model’, is that it considers the effects of alternatives’ characteristics as determinants of choice probabilities. The model will therefore be also called multiattributive random utility model. Random utility models capture the candidates’ uncertainty about voters. The

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10 For a critical discussion of this assumption of ‘sincere voting’ in multiparty systems with coalition governments, see Shepsle (1991, p. 63ff.).
11 See the corresponding proposition by Laver and Hunt (1992, p. 15ff.).
12 For a general introduction into multiattributive decision theory, see Keeney and Raiffa (1992).
derivation of qualitative choice models from utility theory is based on a precise distinction between the decision-maker’s behavior and the researcher’s analysis. Following McFadden (1974) and Manski (1977), the assumption that utility is a random function “…does not reflect a lack of information in the decision-maker but reflects a lack of information regarding the characteristics of alternatives and/or decision-makers on the part of the observer” (Manski, 1977, p. 229).13

Let us therefore assume that the parties are not certain how voters will react to the parties’ locations in the policy space. However, each party has the same subjective expectations about the random behavior of voters. These expectations are represented by probabilistic voting functions. The probability \( P_{ij} \) that voter \( i \) chooses alternative \( j \) of the set of alternatives \( A \) with \( A = \{(P_0, P_1, P_2, \ldots, P_J) | P_0 + P_1 + P_2 + \ldots + P_J = 1, \text{each } P_j \in [0, 1]\} \) depends on the individually observed characteristics \( z_{ij} \) of the alternative \( j \) compared with the characteristics of each of the other alternatives. Utilities are assumed to consist of a deterministic component \( V_{ij} \) and a random component \( \varepsilon_{ij} \):

\[
U_{ij} = U_{ij}(V_{ij}, \varepsilon_{ij}) \quad \text{with} \quad V_{ij} = V_{ij}(z_{ij}).
\] (3)

The probability that alternative \( j \) is chosen is the probability that the utility of alternative \( j \) is higher than that of any other alternative. The random utility model specifies this probability as a parametric function of the general form:14

\[
P_{ij} = f(z_{ij}, z_{ih} \text{ for all } j, h \in A \text{ and } j \neq h, \beta),
\] (4)

where \( f \) is the function that relates the observed data to the choice probabilities. This function is specified up to some vector of parameters, \( \beta \), representing the relative importance of the characteristics. Assuming that the random components \( \varepsilon_{ij} \) are identically and independently distributed with a Gumbel distribution,15 McFadden (1974) derived the conditional logit model with the following choice probabilities:

\[
P_{ij} = \frac{\exp(V_{ij})}{\sum_{h=1}^{J} \exp(V_{ih})} \quad \text{for all } j, h \in A.
\] (5)

Luce (1959) originally derived this model by starting from the axiom of Independence of Irrelevant Alternatives. The McFadden model is a special case of the Luce model insofar as the representative component \( V_{ij} \) is a linear function of the attributes \( z_{ij} \).

The assumption of the irrelevance of independent alternatives (IIA) is, however, the primary drawback of these models when choice situations with more than two alternatives are considered. It requires that the odds of a particular choice are unaffected by the presence or absence of additional alternatives. Suppose for example the following hypothetical choice situation, where German leftist parties are being per-

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13 For models assuming candidate uncertainty about voters’ choices, see the overviews provided by Calvert (1986) and Coughlin (1992).
14 Cf. Train (1986, p. 8).
15 With \( F(\varepsilon_1, \ldots, \varepsilon_J) = \exp[-\exp(-\varepsilon_1), \ldots, \exp(-\varepsilon_J)] \).
received as more similar as compared with rightist parties. The tree structure in Fig. 1 reflects that nested similarity structure of the alternatives in the choice set. As a consequence, different patterns of intergroup versus intragroup substitutions can be observed. In this case, parties’ choice probabilities are dependent of the set in which they are contained. Therefore, the assumption that any two random utility terms are identically and independently distributed may be too restrictive. McFadden (1978, 1981) proposed a more general random utility model that is able to accommodate different degrees of cross-alternative substitution by partitioning the choice set into nests where IIA holds within nests but not across nests. This so-called generalized extreme value model, or a special form of it, the nested multinomial logit model, is a generalization of the multinomial logit model. The NMNL maintains the extreme value distribution whereas the assumption of identically and independently distributed unobserved components is no longer necessary. The choice probability for an elemental alternative \( j \) in the leftist subset \( l \) of our example can then be expressed as the product of the probability that the alternative within the subset \( l \) is chosen and the probability that alternative \( j \) is chosen given that subset \( l \) is chosen.

As we assume that the utility function is representative for the whole population/sample, each voter’s choice probability equals the aggregate share for each choice alternative (Ben-Akiva and Lerman, 1985; Cooper and Nakanishi, 1988). This assumption is unlikely to be empirically tenable as most populations are structured. As a consequence, more complex modeling will have to take account of popu-

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17 The general extreme value distribution of the random utility terms is given by: 
\[ F(e_1, ..., e_J) = \exp\{-G(\exp(-e_1), ..., \exp(-e_J))\}, \] where \( G \) is a non-negative function, homogeneous of degree \( \mu \). It is assumed that the scale parameter \( \mu \) of the Gumbel distribution is constant within subsets, yet may vary across choice subsets. This so-called dissimilarity parameter reflects the degree of similarity between choice subsets and has to be in the unit interval in order to be compatible with the assumption of random utility maximization (McFadden, 1981; Boersch-Supan, 1990a). For \( \mu = 1 \) the NMNL collapses to simple MNL. For further formal and practical details of NMNL, see Ben-Akiva and Lerman (1985).

18 Formally: 
\[ P_{ij} = P_{ij} \cdot P_i, \] where \( P_{ij} \) is the conditional probability of choosing party \( j \) given that an alternative in the subset \( l \) is chosen and \( P_i \) is the marginal probability of choosing a party in \( l \).
lation heterogeneity by appropriate segmentation. For this aim the market can be partitioned into a finite number of homogeneous, mutually exclusive voter segments differing in both party preference and sensitivity to policy changes (cf. Ben-Akiva and Lerman, 1985, p. 134).

The multiattributive random utility model is therefore a suitable theoretical and econometric tool to model market reaction functions for markets with product differentiation. It provides a powerful tool for deriving demand functions whose arguments were variables other than prices, such as advertising and product qualities, but also policy positions, candidate images and government intervention, as in the case of political competition, which is a competition on non-price variables since “there is no direct analogue in politics to prices” (cf. Shepsle, 1991, p. 43).

4. Operationalization and measurement

The following empirical analyses are based on the German part of the international ‘Comparative National Election Project’ (CNEP) containing the national study of the first all-German general election on December 2, 1990. For simplification, the analyses will be restricted to the first wave of the West German study with a target sample size of N=1400. Data were collected from a representative random sample carried out as face-to-face interviews.

The data set contains four bipolar issue scales for German unification, immigration policy, abortion and nuclear energy. These scales have labeled end-points suggesting a bipolar policy continuum. The respondent’s placement of herself and of each of the parties on each of the policy scales allows for the computation of the respective distances presented in the formal model.

In order to take account of population heterogeneity, mutually exclusive voter segments differing in both party preference and sensitivity to policy changes will be created. Numerous bases for segmentation can be advanced, each with its own set of advantages and disadvantages for particular types of campaign issue. Differing longstanding party preferences will be captured by the concept of party loyalty as measured by a version of the classic party identification question which explicitly accentuated its long-term aspect.

For differing sensitivity to policy changes, I propose a specific political segmentation into highly involved issue publics measuring the individual saliency of one of

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19 Cf. Cooper and Nakanishi (1988) and Hanssens et al. (1989).
20 For a discussion of the most commonly used approach, the address or characteristics approach, to modeling demand for differentiated products and the close affinities to multiattributive random utility models, see Anderson et al. (1992).
21 The German part of the project has been coordinated by Professors Kaase, Klingemann and Pappi, and financed by the Deutsche Forschungsgemeinschaft. Data are available at the Zentralarchiv für Empirische Sozialforschung, University of Cologne.
22 This is due to differently structured party systems in East and West Germany. For a detailed comparative analysis of voting decisions in East and West Germany, cf. Thurner (1998).
the four position issues considered. Voters have been classified into a dummy variable according to their first answer to the open-ended question on the most important problem facing the country. The naming of one of the four issues makes it possible to test whether the assumption of a representative utility function is appropriate and to show whether voters highly involved in one political objective make statistically significant differing evaluations of attributes.

The voting choice is conceptualized by the respondent’s prospective report of her vote for one of the parties, resulting in a multicategorial-dependent variable. The German electoral rules allow the voter to cast two votes, and the CNEP study, therefore, differentiated between the first vote and the second vote. As the share of seats in the federal parliament, the Bundestag, depends on the second vote where the voter marks his preference for a party list, the second vote will be predicted in the empirical analysis. Owing to the very limited number of minor party responses, the following analysis is carried out for four alternatives: SPD, CDU/CSU, FDP and the GREENs.

5. Empirical applications

5.1. The ideal-points-only model

What choice probability-change results if we wander along the voters’ ideal points on a policy dimension and assume fixed party positions? Which substitutional patterns can be observed? To address this question I first present average partial derivatives of a simple MNL (Table 1). Simultaneous Wald tests of each MNL parameter across choice groups confirm that all policy dimensions have effects which are significantly different from zero, with immigration and nuclear energy showing the highest impact on the change of choice probabilities. Policy preferences in the case of achieving unification show the slightest effect. Of particular interest are the patterns of cross-substitution which can be detected: each one-unit change on the immigration scale toward a more restrictive legislation increases the choice probability of the CDU/CSU by 7.4% on average, when pari passu the choice probabilities of the SPD diminish by 2.7% and the choice probabilities of the GREENs by 4.2%.

23 The following analyses have been conducted with the software package LIMDEP 7.0 and its full information maximum likelihood (FIML) estimator for NMNL; for a discussion of sequential estimation techniques, see McFadden (1981, 1984).

24 Partial derivatives:

$$\frac{\partial P_j}{\partial X} = \beta_j P_j - P_j \sum_{h=1}^{J} \beta_h P_h = P_j \left( \beta_j - \sum_{h=1}^{J} \beta_h P_h \right).$$

Normally, direct and cross-market share elasticities are considered when regarding demand systems. Market share elasticities can be informally defined as the ratio of the relative change in a market share corresponding to a relative change in a marketing-mix variable. I prefer to present partial derivatives instead of elasticities, because it is not useful to measure the cross-substitutional effects of a 1% change in the case of the seven-point issue scale.
Table 1
Policy preferences and vote in the multinomial logit model: average partial derivatives (West Germany) (N=865)

<table>
<thead>
<tr>
<th>Policy</th>
<th>SPD</th>
<th>CDU/CSU</th>
<th>FDP</th>
<th>GREENs</th>
</tr>
</thead>
<tbody>
<tr>
<td>Unificationa</td>
<td>−0.5</td>
<td>2.8</td>
<td>−0.6</td>
<td>−1.7</td>
</tr>
<tr>
<td>Immigrationa</td>
<td>−2.7</td>
<td>7.4</td>
<td>−0.6</td>
<td>−4.2</td>
</tr>
<tr>
<td>Abortiona</td>
<td>3.4</td>
<td>−3.2</td>
<td>−0.3</td>
<td>0.1</td>
</tr>
<tr>
<td>Nuclear energya</td>
<td>2.8</td>
<td>−5.4</td>
<td>−0.4</td>
<td>2.9</td>
</tr>
</tbody>
</table>

*a Simultaneous Wald tests on MNL coefficients across choice groups, significant on the 5% level.

Only the CDU/CSU profits from supporters of strong state intervention for the economic reconstruction of East Germany.

Preferences for a more liberal abortion law diminish the chances of the CDU/CSU. But only the SPD makes capital from these losses, whereas in the case of the GREENs the internal fragmentation into value conservatives on the one hand and radical adherents of women’s rights completely neutralizes any effect. Changes on the nuclear energy scale also lead to notable consequences. As is to be expected, the GREENs make the most of preferences for a complete closure of nuclear plants, albeit competing in this segment with the policy offers of the SPD. Their gains are exclusively compensated by losses of the CDU/CSU.

This presentation of direct and cross partial derivatives demonstrates that we can gain important insights into substitutional patterns of the policy competition of the parties considered.

5.2. Individual policy preferences as alternative-specific variables

In this section the individual ideal points will be allowed to have differential impacts upon the odds of choosing one alternative rather than another. For this aim, the multiattributive random utility model will be applied and the coefficient vector will be made alternative-specific.25 J−1 alternative-specific variables will be specified by using the Free Democrats as the reference party (Table 2).

Estimates of the alternative-specific constants indicate that there are strong, highly significant non-policy biases in this model toward each party. Several alternative-specific variables fail to achieve statistical significance at the 5% level: SPD–Unification, GREENs–Unification, SPD–Immigration, CDU/CSU–Abortion, CDU/CSU–Nuclear Energy. Looking at the unification issue, it becomes obvious that this criterion influenced in a significant way only the choice of the Christian Democrats, whereas the SPD and the GREENs cannot capitalize on this issue. Voters of these parties frequently did not appreciate the policy positions of their parties in this dimension or they have been indifferent in this dimension. But as before, even the effect for the CDU/CSU fares poorly compared with the other issues, especially with the

Table 2
Policy preferences as alternative-specific variables and vote decision (West Germany) (N=865)

<table>
<thead>
<tr>
<th>Variable</th>
<th>( \beta )</th>
<th>T-ratio</th>
<th>P-value</th>
<th>Effect</th>
</tr>
</thead>
<tbody>
<tr>
<td>SPD–Unification</td>
<td>0.123</td>
<td>1.631</td>
<td>0.103</td>
<td>1.131</td>
</tr>
<tr>
<td>CDU/CSU–Unification</td>
<td>0.150</td>
<td>2.017</td>
<td>0.044</td>
<td>1.162</td>
</tr>
<tr>
<td>GREENs–Unification</td>
<td>0.074</td>
<td>0.753</td>
<td>0.451</td>
<td>1.077</td>
</tr>
<tr>
<td>SPD–Immigration</td>
<td>0.023</td>
<td>0.334</td>
<td>0.738</td>
<td>1.023</td>
</tr>
<tr>
<td>CDU/CSU–Immigration</td>
<td>0.272</td>
<td>3.900</td>
<td>0.000</td>
<td>1.313</td>
</tr>
<tr>
<td>GREENs–Immigration</td>
<td>-0.254</td>
<td>-3.023</td>
<td>0.003</td>
<td>0.776</td>
</tr>
<tr>
<td>SPD–Abortion</td>
<td>0.198</td>
<td>2.960</td>
<td>0.003</td>
<td>1.219</td>
</tr>
<tr>
<td>CDU/CSU–Abortion</td>
<td>-0.064</td>
<td>-1.005</td>
<td>0.315</td>
<td>0.938</td>
</tr>
<tr>
<td>GREENs–Abortion</td>
<td>0.302</td>
<td>3.041</td>
<td>0.002</td>
<td>1.353</td>
</tr>
<tr>
<td>SPD–Nuclear energy</td>
<td>0.253</td>
<td>2.945</td>
<td>0.003</td>
<td>1.288</td>
</tr>
<tr>
<td>CDU/CSU–Nuclear energy</td>
<td>-0.137</td>
<td>-1.632</td>
<td>0.103</td>
<td>0.872</td>
</tr>
<tr>
<td>GREENs–Nuclear energy</td>
<td>0.798</td>
<td>6.081</td>
<td>0.000</td>
<td>2.221</td>
</tr>
<tr>
<td>SPD–ASC</td>
<td>3.689</td>
<td>4.069</td>
<td>0.000</td>
<td>–</td>
</tr>
<tr>
<td>CDU/CSU–ASC</td>
<td>5.580</td>
<td>5.971</td>
<td>0.000</td>
<td>–</td>
</tr>
<tr>
<td>FDP–ASC</td>
<td>5.168</td>
<td>4.868</td>
<td>0.000</td>
<td>–</td>
</tr>
</tbody>
</table>

\( 2(\ln L(N) - \ln L(0)) = 304.252, \text{ DF}=12, P-value=0.000, \text{ Pseudo } R^2=14.9\% \)

Percentage correctly predicted prediction table: 43.70
Classification table: 55.70

\( ^a \) ASC: Alternative-specific constant.

The table provides a summary of the policy preferences as alternative-specific variables and vote decision for West Germany, with a sample size of 865. The coefficients, T-ratios, and P-values indicate the influence of various policy issues on voter choice, with the CDU/CSU and the GREENs showing significant effects on immigration and nuclear energy issues, respectively. The table also includes a classification of voters into 43.70% correct predictions and 55.70% correctly classified voters.

The insights from the table highlight the importance of specific policy issues in influencing voter decisions, with the nuclear energy issue showing a significant effect on the GREENs’ choice. The immigration issue, however, did not significantly affect the SPD chances, reflecting the internal fragmentation of the Social Democrats. Only the GREENs and the CDU/CSU were able to reach homogeneous policy preferences among their respective voters on immigration.

Specifying variables in this way, the coefficients represent the difference in the utility of respective alternatives compared with the reference alternative conditional on individually varying ideal points. In interpreting the coefficients, it should be noted that changes of respective variables influence the deterministic utility component of several alternatives. In order to calculate the increase or decrease of the relative chances of one party in comparison with another party, we have, therefore, to take the difference between respective effect coefficients (Long, 1987). For example, a change of ideal point on the nuclear energy dimension by one unit in the direction of closing down the plants, increases the relative chances of the GREENs constantly by the multiplicative factor of \( \exp(0.789) \), at the same time diminishing the chances of the CDU/CSU.

26 Consequently, the party changed its position after the election and cooperated with the new government in tightening the asylum laws in 1993.

27 Relative to the changes of the FDP.
the relative chances of the CDU/CSU by a factor of $\exp(-0.137)$. Together, the chances of voting for the GREENs instead of the CDU/CSU would increase by the factor $\exp(0.935)$; that is, a constant multiplicative factor of 2.6.

Considering the model fit, represented by McFaddens\(^{28}\) Pseudo $R^2=14.9\%$ and two summary measures of prediction success,\(^ {29}\) with the prediction success table [percentage correctly predicted (PCP) =43.7\%] and the classification table (PCP =55.7\%), the result is not satisfactory.

Hausman–McFadden tests\(^ {30}\) yield that the IIA assumption is not violated when removing CDU/CSU and FDP, respectively, from the full choice set. In the case of the SPD and the GREENs the test is inconclusive. Therefore, supplementary tests are required. Likelihood ratio tests of NMNL versus MNL as well as $t$-tests of dissimilarity parameters indicate, however, that a tree structure is not justified.\(^ {31}\) On the basis of this result and assuming for this purpose a homogeneous population, estimated market shares can be calculated and presented graphically. As an example, the shares in the question of nuclear energy are shown in Fig. 2.

Fig. 2 illustrates estimated choice probabilities/market shares conditional on preferences in the nuclear energy dimension, multivariately controlled by the other policy dimensions. Once again the vicinity of GREENs and SPD on ecological questions is evident. These two parties are competing for the segment of voters supporting the closure of nuclear plants. The potential of the CDU is mainly located in the segment of adherents of the further development of nuclear energy.

---

\(^{28}\) \(\rho^2=1-\frac{\ln L(N)}{\ln L(0)}\), also termed as likelihood ratio index, where $\ln L(N)$ is the log likelihood of the unconstrained model and $\ln L(0)$ is the log likelihood of the model defined by the null hypothesis, in most practical applications the constants-only model. McFadden (1979, p. 307) has suggested that $\rho^2$ values between 0.2 and 0.4 could be considered to represent a very good fit.

\(^{29}\) Both summary measures are the result of cross-classifying the actual choice with the predicted choice. Whereas in the prediction success table the predicted choices are represented by the estimated choice probabilities, in the classification table a classification rule has to be defined, which, in our case, accords to the assumption that the category with the highest predicted probability is selected, cf. McFadden (1979, p. 307).

\(^{30}\) The form of the test statistic proposed by Hausman and McFadden (1984) is:

\[ \chi^2(c)-\frac{(\hat{\beta}^r-\hat{\beta})'(\hat{\text{Cov}}(\hat{\beta}^r)-\hat{\text{Cov}}(\hat{\beta}))^{-1}(\hat{\beta}^r-\hat{\beta})}{}, \]

where $\hat{\beta}^r$ is an estimator based on full choice set, $\hat{\beta}$ is an estimator based on restricted choice set and $\hat{\text{Cov}}(\hat{\beta})$ is the estimated covariance matrix of the estimator.

\(^{31}\) Hausman–McFadden test: removing CDU/CSU yields $\chi^2=1.764$, degrees of freedom (DF) =10, $P$-value =1.0; SPD $\chi^2=2.922$, DF=10, $P$-value =0.983; GREENS $\chi^2=0.935$, $P$-value =0.34; FDP $\chi^2=1.764$, DF=10, $P$-value =1.0. Yielding negative test statistics is not uncommon for the Hausman–McFadden test, cf. Small and Hsiao (1985). Tables and test results will be delivered on request.
5.3. Policy distances as party attributes

In the following section the vote will be explicitly modeled as a function of the distances between candidates and the voter, regardless of where the voter is located.

The structure of the multiattributive random utility model makes it possible to treat policy-specific distances to each of the parties as attributes of these parties and to specify them as a generic variable. As indicated in the theoretical model, there are multiple different specification possibilities. When estimating more than one specification it is useful to compare goodness-of-fit measures. The testing of non-nested hypotheses of discrete choice models (cf. Ben-Akiva and Lerman, 1985, p. 171ff.) requires an adjusted likelihood ratio index $\bar{\rho}^2$ which takes account of differing degrees of freedom in the compared models by using the Akaike information criterion (Ben-Akiva and Lerman, 1985, p. 167).

I present both additive linear and additive quadratic utility loss functions in the case of the decompositional multiattributive random utility model. The linear model fares better in terms of model fit (Table 3). Comparing compositional City-block distance with compositional Euclidean distances which are commonly assumed in the theoretical literature surprisingly shows that the former more appropriately reproduces the process of decision-making. Contrasting now decompositional and compositional models shows that the former reproduce the data slightly better. Despite

\[
\bar{\rho}^2 = 1 - \frac{\ln L(N) - K}{\ln L(0)},
\]

with $K$ denoting the number of parameters to be estimated in the restricted model.
Table 3
Comparison of model fit of different specifications of the utility function (West Germany) (N=775)

<table>
<thead>
<tr>
<th>Specification</th>
<th>2(ln L(N)− ln L(0))</th>
<th>DF</th>
<th>P-value</th>
<th>$\rho^2$ (%)</th>
<th>Prediction (%)</th>
<th>Classification (%)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Decompositional multiattributive distances</td>
<td>538.364</td>
<td>4</td>
<td>0.000</td>
<td>28.58</td>
<td>53.70</td>
<td>66.80</td>
</tr>
<tr>
<td>Decompositional squared multiattributive distances</td>
<td>473.672</td>
<td>4</td>
<td>0.000</td>
<td>25.09</td>
<td>51.30</td>
<td>64.40</td>
</tr>
<tr>
<td>Compositional City-block distance$^a$</td>
<td>515.915</td>
<td>1</td>
<td>0.000</td>
<td>27.69</td>
<td>52.80</td>
<td>65.50</td>
</tr>
<tr>
<td>Compositional Euclidian distance$^a$</td>
<td>473.712</td>
<td>1</td>
<td>0.000</td>
<td>25.42</td>
<td>51.30</td>
<td>65.20</td>
</tr>
</tbody>
</table>

$^a$ Unweighted.
losing three degrees of freedom with the decompositional models, the adjusted likelihood ratio index suggests considering separate policy dimensions and rejecting the assumption of equal weights.

Therefore, the decompositional linear model will be used in the following analyses in order to determine the weights of each of the four dimensions. This linear-compensatory functional form supposes a constant marginal disutility represented by the coefficient $\beta$. The relative impacts of the four decision criteria are presented in Table 4. Each of the criteria has an effect significantly different from zero on the decision of the voters. As expected, the signs are all negative: the larger the perceived distance, the smaller the chance of getting the vote. Perceived policy distances have differentiated effects on the chances of different parties of getting the vote in the case of the immigration issue. For example, an additional unit of perceived distance in the immigration issue has the strongest negative impact on the chances of the liberal party FDP. In magnitude, the effect of unification ranks between the estimates for nuclear energy and abortion. Perceived distances in nuclear energy and unification have an impact more than double that of abortion. Expressed in terms of marginal rate of substitution, this would mean that if the distance perceived to a party location on the unification issue is increased by two units, the distance in abortion would have to decrease by two units in order to remain indifferent. Weights of the generic coefficients influence the relative chances of each of the parties in the same way. For example, if the perceived distance to the CDU/CSU in the unification issue

<table>
<thead>
<tr>
<th>Variable</th>
<th>$\beta$</th>
<th>T-ratio</th>
<th>P-value</th>
<th>Effect</th>
</tr>
</thead>
<tbody>
<tr>
<td>Unification</td>
<td>$-0.363$</td>
<td>$-6.895$</td>
<td>$0.000$</td>
<td>$0.696$</td>
</tr>
<tr>
<td>SPD–Immigration</td>
<td>$-0.273$</td>
<td>$-4.624$</td>
<td>$0.000$</td>
<td>$0.761$</td>
</tr>
<tr>
<td>CDU/CSU–Immigration</td>
<td>$-0.326$</td>
<td>$-5.016$</td>
<td>$0.000$</td>
<td>$0.722$</td>
</tr>
<tr>
<td>FDP–Immigration</td>
<td>$-0.464$</td>
<td>$-4.323$</td>
<td>$0.000$</td>
<td>$0.629$</td>
</tr>
<tr>
<td>GREENs–Immigration</td>
<td>$-0.372$</td>
<td>$-4.800$</td>
<td>$0.000$</td>
<td>$0.689$</td>
</tr>
<tr>
<td>Abortion</td>
<td>$-0.177$</td>
<td>$-5.123$</td>
<td>$0.000$</td>
<td>$0.838$</td>
</tr>
<tr>
<td>Nuclear Energy</td>
<td>$-0.439$</td>
<td>$-9.927$</td>
<td>$0.000$</td>
<td>$0.645$</td>
</tr>
<tr>
<td>SPD–ASC</td>
<td>$0.815$</td>
<td>$3.973$</td>
<td>$0.000$</td>
<td>$-$</td>
</tr>
<tr>
<td>CDU/CSU–ASC</td>
<td>$1.411$</td>
<td>$5.842$</td>
<td>$0.000$</td>
<td>$-$</td>
</tr>
<tr>
<td>FDP–ASC</td>
<td>$-0.041$</td>
<td>$-0.150$</td>
<td>$0.880$</td>
<td>$-$</td>
</tr>
</tbody>
</table>

$2(\ln L(N)−\ln L(0))=541.655, \text{ DF}=7, \text{ P-value}=0.000, \text{ Pseudo } R^2=29.2\%$

Percentage correctly predicted: 53.80
Classification table: 66.80

The restriction of equality of coefficients imposed by the generic model on the more general model with alternative-specific attributes has been tested with likelihood ratio tests (cf. Ben-Akiva and Lerman, 1985, p. 168). The results of the likelihood ratio test statistic for the null hypothesis of generic attributes are: unification—$\chi^2=2.29, \text{ DF}=3$; immigration—$\chi^2=7.24, \text{ DF}=3$; abortion—$\chi^2=0.98, \text{ DF}=3$; nuclear energy—$\chi^2=6.03, \text{ DF}=3$. The immigration variable has been split for illustrative reasons; for all other variables the null hypothesis cannot be rejected when using a 10% significance level.
increases by one unit, then the odds of this party are decreasing with relation to each of the other parties by $100[\exp(-0.36)-1]$; i.e., by 30%. This holds for all parties.

Compared with the ideal-points-only model the fit of the distance model fares much better, which points to the fact, noted also by Enelow and Hinich (1985, p. 268), that variables exogenous to the specified model make their influence felt through the variables of the estimated model. These effects should not impair the determination of the relative weights of, and the trade-offs between, policy dimensions as long as we can plausibly assume that these biases turn out to be equal in all dimensions and average out over alternatives (Enelow and Hinich, 1984, p. 171). As in the case of the ideal-point-only model, computing the Hausman–McFadden tests for the distances model is not conclusive, but again additional tests reject the hypothesis of tree-structured choice.\footnote{Removing CDU/CSU yields a chi-square value $c<0$; SPD—$c<0$; FDP—$c=2.769$, $DF=8$, $P$-value=0.948; GREENS—$c=6.071$, $DF=8$, $P$-value=0.639. Tables and test results of likelihood ratio tests of NMNL versus MNL as well as $t$-tests of dissimilarity parameters will be delivered on request.}

Using the results of Table 4, the message of the multiatttributive model can now for exemplary reasons also be graphically represented. Analogous to marketing studies where a brand’s market share is a linear function in marketing-mix variables, in the following analyses the decision for a party will be determined dependent on their evaluation of attributes of the alternatives. Assuming a homogeneous population, the estimated choice probabilities can once again be considered as estimated market shares conditional on perceived distances. In order to determine the conditional aggregate shares of one party, we have to hold constant perceived c.p. distances to all other parties. The c.p. conditions will be defined as the respective average perceived distance to the other parties. This enables us to visualize distance-based policy reaction functions for the immigration issue (Fig. 3).

Perceiving a congruence with the SPD in the case of the immigration issue and an average distance to the other parties results in a share of 60% of this party; i.e., the SPD is voted for with a choice probability of 0.6. Contrary to deterministic models, perceived issue congruence and non-indifference to any other party does not lead to an unambiguous decision, but to a probabilistic one. As we have assumed specific c.p. conditions and controlled for other policy attributes, it is obvious that an issue congruence with a party does not lead to a choice probability of one. This effect is amplified in the case of the two parties with minor market shares, where interest congruence leads only to very small choice probabilities. If decision criteria, important for the evaluation of these parties, are not specified in the model, such as strategic calculus, the estimated probabilities are reduced. But the small choice probabilities for the minor parties are also due to the unrealistic assumption about c.p. conditions: perceiving an interest congruence with the GREENs in general implies also a perceived closeness of the SPD and a larger distance on the part of CDU/CSU and FDP. This effect of a genuine individual-level perception of the party system can be illustrated if the reaction function is calculated only for one party by varying c.p. conditions for the other parties, as in Fig. 4.
Fig. 3. Estimated choice probabilities/market shares conditional on distances in immigration policy (West Germany).

(N = 775)

Fig. 4 shows only the conditional shares of the SPD by varying the distance perceived to the CDU/CSU and holding the distances to the remaining parties constant. The lowest curve presents the shares of the SPD with varying distances to this party and a constant distance to the CDU/CSU of zero. In the case of perceiving an interest

(N = 775)

Fig. 4. SPD share conditional on distances in immigration policy and varying CDU distances (West Germany).
with both the SPD and the CDU/CSU in the immigration question, this leads to a share of the SPD of 45%. If we allow the voter to perceive a larger and larger distance to the CDU/CSU — as indicated in the other curves — this results in a monotonically increasing share of the SPD with an estimated maximum of 80% when the location of the SPD is perceived as identical with the bliss point, and at the same time the CDU/CSU is perceived as being located at the other pole of the policy space. This increase of choice probabilities is remarkable as it illustrates very clearly the effect of the individual perception of the party space.

It will be shown in the next section whether the controlling of population heterogeneity (cf. Ben-Akiva and Lerman, 1985, p. 110) improves the model to a substantial degree.

5.4. Controlling for population heterogeneity: party loyalty and highly involved issue publics

By assuming a homogeneous population in the preceding analyses, individual differences have been leveled out and the average responsiveness of the sample has been determined. But, in general, different voters react in a different manner under identical conditions. Unmeasured voter-specific factors may influence voters’ choice behavior. Even with the specification of demographic variables, voters may differ in their responses to marketing instruments of parties. Failure to control for such heterogeneity is likely to yield biased and inconsistent estimates and, more importantly, biased and inconsistent parameter estimates of choice probabilities. It is therefore important to capture systematic taste variations in the utility functions. Understanding and identifying market structures is, therefore, a precondition for the formulation of effective party strategies such as policy positioning, targeting and campaigning. Most of the theoretical literature assumes that the political market has a single structure. Transferring the insights of marketing science suggests, however, that these approaches could lead to incorrect results in a market characterized by heterogeneous structure. In general, asymmetries in markets and competition are reflected in differential cross-effects among parties. Heterogeneous populations differ in both preferences and responsiveness to marketing efforts of parties. Therefore, it is necessary to incorporate multiple sources of heterogeneity.

So, for example, some parties are able to create strong voter loyalties leading to an imparity of competitive interdependencies. This conception of structured markets in politics has also been put forward by Shepsle: “Parties and candidates stake out locations well in advance of any specific election…their respective locational ‘types’ constitute reputations which are relatively durable and not readily altered in the short run of a specific campaign. Positions on the specific issues salient during a given campaign, however, allow for more flexibility… That is, some bases of voter evaluation are fixed and durable (Party ‘types’ are analogous to location-specific capital) while others may be varied by parties (specific issue positions are analogous to product prices)” (Shepsle, 1991, p. 42 and 43). This conceptualization leads to a segmentation of the market into a segment of ‘switchers’ segment, highly responsive to changes in the short-term campaigning variables and another segment of ‘loyal’ vot-
ers, relatively unresponsive to short-term policy programs. Empirical students of voting extensively use the concept of party identification to predict and explain voting behavior. In the context of the Michigan approach, this concept means the longstanding sociopsychological attachment to a party. In rational choice approaches, however, this concept captures the effect of an individual’s past voting behavior on the actual and future votes and the carry-over effects of past campaigning and party reputation (Fiorina, 1981). Fiorina’s ‘running tally’ model conceptualizes the development of party identification as the ongoing result of comparative evaluation of the party platforms and performance of party politicians contributing to a cumulative evaluation of the parties by voters.\footnote{Marketing researchers generally introduce a measure of brand loyalty deduced from scanner panel data, cf. Guadagni and Little (1983). Since the data set does not provide long-term individual vote histories, I content myself with achieving a cross-sectional preference segmentation by the concept of party identification.}

Another way of segmentation proposed in this article is to allow sensitivities for the offered policies in the 1990 national election campaign to vary across segments. Recent literature on voting and elections stresses the growing heterogeneity of the public’s issue interests and a fragmentation into a variety of distinct issue publics. I propose, therefore, a segmentation by specific issues of immediate or personal importance. Different measures of issue saliency (RePass, 1971; Rabinowitz et al., 1982; Niemi and Bartels, 1985) have been proposed with no clear result.\footnote{For an overview, see Niemi and Bartels (1985).} Following RePass (1971) and Rabinowitz et al. (1982) who have found “that any issue singled out as personally most important plays a substantially greater role for those who so view it than it does for others” (Rabinowitz et al., 1982, p. 57), I shall utilize open-ended responses with regard to selective concerns about the policies we have chosen as having been the relevant campaign issues. Specifying policy distances and these so-called highly involved issue publics (HIIP) as additional interaction terms enables us to capture the selective emphasis of voters and thereby relaxing the assumption of common weights across all voters.

The unique feature of our approach is that we are able to determine how parties compete within each structured market. Two distinct groups of explanatory variables are used in the following model. The first group consists of the policies offered in the campaign. The second group consists of voter-specific variables, namely party loyalty and classification of the voters into segments of highly involved issue publics indicating an individual saliency of one of the four policies considered. Table 5 shows the outstanding effect of the loyal segment. This variable has been specified as a generically specified binary variable in order to take into account the comparative character of the ‘running tally’. Therefore, for the first time, we are able to specify the party loyalty variable simultaneously for each of \( n \geq 2 \) parties/candidates. A test on generic variables shows that the decomposition into alternative-specific variables is not justified. Consequently, the effect of being classified in the segment of party loyals is equal for all four parties.

As indicated by the interaction effects, only in the case of the immigration issue...
can a varying valuation of policy distances in this dimension be observed. But the significance of the varying sensitivities depends on the parties considered. Naming the immigration issue as the most important political problem considerably increases the effect of perceived distances in this segment on the relative chances of the CDU/CSU and the SPD only, whereas in the case of the two other parties this has no significant negative impact on the vote of these parties. At the same time, simple alternative-specific effects of immigration remain significant only in the case of the minor parties. Results of different specification lead to the conclusion to maintain a simple MNL model.37

37 Hausman–McFadden test: removing CDU/CSU yields—\(c=61.359\), DF=15, \(P\)-value=0.000; SPD—\(c=10.303\), DF=15, \(P\)-value=0.801; FDP—\(c=16.110\), DF=15, \(P\)-value=0.375; GREENS—\(c=21.812\), DF=15, \(P\)-value=0.113. In the case of the CDU/CSU the test indicates that the IIA assumption is violated. A hierarchical likelihood ratio test of a respective NMNL versus a MNL version yields a significant model improvement (\(\chi^2=11.29\), DF=1, \(P\)-value=0.001), but the dissimilarity parameter is not in the unit interval, as tested with \(t\)-tests. Since parameter patterns (rank order, weight) are very similar in both models, I prefer to present the MNL results. Tables of the NMNL model will be delivered on request.

Table 5
Policy distances, party loyalty, individual saliency effects and the vote (West Germany) (\(N=775\))

<table>
<thead>
<tr>
<th>Variable</th>
<th>(\beta)</th>
<th>(T)-ratio</th>
<th>(P)-value</th>
<th>Effect</th>
</tr>
</thead>
<tbody>
<tr>
<td>Party loyalty</td>
<td>2.130</td>
<td>16.089</td>
<td>0.000</td>
<td>8.415</td>
</tr>
<tr>
<td>Unification</td>
<td>−0.179</td>
<td>−2.189</td>
<td>0.029</td>
<td>0.836</td>
</tr>
<tr>
<td>CDU/CSU–Immigration</td>
<td>−0.131</td>
<td>−1.590</td>
<td>0.112</td>
<td>0.877</td>
</tr>
<tr>
<td>SPD–Immigration</td>
<td>−0.140</td>
<td>−1.920</td>
<td>0.055</td>
<td>0.869</td>
</tr>
<tr>
<td>FDP–Immigration</td>
<td>−0.311</td>
<td>−2.676</td>
<td>0.007</td>
<td>0.733</td>
</tr>
<tr>
<td>GREENs–Immigration</td>
<td>−0.357</td>
<td>−3.981</td>
<td>0.000</td>
<td>0.700</td>
</tr>
<tr>
<td>Abortion</td>
<td>−0.038</td>
<td>−0.900</td>
<td>0.368</td>
<td>0.963</td>
</tr>
<tr>
<td>Nuclear energy</td>
<td>−0.299</td>
<td>−5.589</td>
<td>0.000</td>
<td>0.742</td>
</tr>
<tr>
<td>Unification×HIIP</td>
<td>0.074</td>
<td>0.627</td>
<td>0.530</td>
<td>1.077</td>
</tr>
<tr>
<td>CDU/CSU–Immigration×HIIP</td>
<td>−0.574</td>
<td>−2.176</td>
<td>0.030</td>
<td>0.563</td>
</tr>
<tr>
<td>SPD–Immigration×HIIP</td>
<td>−0.499</td>
<td>−2.165</td>
<td>0.030</td>
<td>0.607</td>
</tr>
<tr>
<td>FDP–Immigration×HIIP</td>
<td>−0.364</td>
<td>−1.028</td>
<td>0.304</td>
<td>0.695</td>
</tr>
<tr>
<td>GREENs–Immigration×HIIP</td>
<td>−0.174</td>
<td>−0.651</td>
<td>0.515</td>
<td>0.840</td>
</tr>
<tr>
<td>Abortion×HIIP</td>
<td>−0.012</td>
<td>−0.011</td>
<td>0.992</td>
<td>0.988</td>
</tr>
<tr>
<td>Nuclear energy×HIIP</td>
<td>−0.104</td>
<td>−0.623</td>
<td>0.533</td>
<td>0.901</td>
</tr>
<tr>
<td>SPD–ASC</td>
<td>−0.093</td>
<td>−0.366</td>
<td>0.714</td>
<td>–</td>
</tr>
<tr>
<td>CDU/CSU–ASC</td>
<td>0.242</td>
<td>0.822</td>
<td>0.411</td>
<td>–</td>
</tr>
<tr>
<td>FDP–ASC</td>
<td>−0.199</td>
<td>−0.647</td>
<td>0.518</td>
<td>–</td>
</tr>
</tbody>
</table>

\(2(\ln L(N)−\ln L(0))=882.658,\) DF=15, \(P\)-value=0.000, Pseudo \(R^2=47.6\%\).

<table>
<thead>
<tr>
<th>Prediction: 54.10</th>
<th>Classification: 66.70</th>
</tr>
</thead>
</table>
In order to visualize the effects of the preference segmentation, Fig. 5 demonstrates the estimated market shares of the CDU/CSU conditional on perceived distance in the immigration issue and party loyalty. Having a longstanding party loyalty toward the CDU dampens the effect of a perceived disagreement with one’s party on policies of the day. In the segment of party loyals even being in complete disagreement with the CDU/CSU on immigration policy leads to 60% of votes for that party, whereas in the case of the non-loyal switchers this percentage goes down to about 10%. By segmenting the sample it was therefore possible to discover strong preference biases on the one hand and differentiated effects of topical policy instruments on the other.

6. Summary

Political scientists have proceeded to subject the spatial representation of the structure of elections to theoretical analyses of increasing complexity and sophistication in recent years. This paper discusses recent developments of empirically testing this theory. Borrowing from the multiattributive random utility model in its general extreme value variant, I presented the formal model. Then several useful features for electoral studies are illustrated by using survey data of the 1990 German national election campaign. Extensive tests support the conclusion that a simple MNL specification is sufficient for our analysis; i.e., the assumption of a nested structure is not necessary given the data. A comparison of an ideal-points-only model with a dis-

(N = 775)

Fig. 5. CDU/CSU shares conditional on perceived distance in the immigration issue and party loyalty (West Germany).

38 For a study where a tree structure proved to be necessary, see Thurner and Pappi (1998b).
tance model shows that the latter has more explanatory power. This is due to ideo-
syncratic perceptions of the parties. Comparing different specifications of the voter’s  
utility function, it can be proved empirically that quadratic loss functions fare no  
better than linear distances. Specifying decomposed distances improves the model  
fit only slightly. However, this model type provides important information for the  
assessment of the weight of each of the policy dimensions. Furthermore, issue effects  
may vary depending on which party is considered, as has been shown for the immi- 
gration issue. Strong segmentational effects are detected. For the first time, policy  
effects have been controlled for by partisan biases in a multiparty system. As to be  
expected, party loyalty turned out to be the strongest predictor for party choice. But  
policy effects in general and even saliency effects for voters highly involved in the  
immigration issue have also been proved. These tests should be reconsidered in other  
study contexts in order to assess the validity of the spatial theory of voting.

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