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POLITICAL INVOLVEMENT AND ATTITUDE STRUCTURE IN THE GENERAL PUBLIC*

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Differences in attitude structure as a function of political involvement are examined using a structural equation procedure with latent variables. In addition, criticisms of an earlier Judd and Milburn (1980) paper are examined. Findings suggest that the less involved are less likely to express political attitudes and that when they do their expressions contain more residual variance or error. At the same time, the attitudes of both the involved and uninvolved show evidence of a single underlying ideology that is remarkably stable in all groups. In arriving at these conclusions, a number of points are made concerning the procedures to be followed in comparing latent variable models between groups.

During the past twenty years sociologists and political scientists have studied the ideological organization of political attitudes in the general public. More specifically, they have examined differences in the degree of ideological organization between groups that differ in levels of political activity, interest in politics, and political sophistication (Achen, 1975; Converse, 1964; Converse and Markus, 1979; Judd and Milburn, 1980; Nie, Verba, and Petrocik, 1976). Philip Converse (1964) originally proposed that the organization of political attitudes differs between these groups in three ways: the degree to which attitudes reflect an underlying ideological predisposition, the degree to which attitudes are consistent with each other, and the degree to which attitudes are stable. The original research supporting the notion that attitude organization differs between groups reported correlations between attitudes (to examine consistency and stability) and re-

sponses to open-ended questions (to examine the presence of ideological thought).

A number of methodological problems in such procedures have recently been pointed out (Achen, 1975; Judd and Milburn, 1980). First, it is unclear whether responses to open-ended questions can be used to determine whether closed-ended attitude questions are ideologically based. A better procedure would be to look for an underlying ideological factor or construct that explains covariance in responses to the closed-ended questions. Second, comparisons of correlations between samples are affected by between-sample differences in the variances of the variables involved (Barton and Parsons, 1977). A better procedure would involve the comparison of unstandardized regression or structural coefficients (Blalock, 1967). Finally, two sorts of stabilities are confounded in correlations between the same attitude at two time points: ideological stability and the stability of residual or issue-specific variation. Judd and Milburn (1980) discussed these three problems at length and employed the technique of linear structural modeling in an attempt to overcome them. We examined the organization of five attitudes at three points in time in two samples that differed in educational attainment.

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A number of criticisms have been raised about the Judd and Milburn (1980) approach (Converse, 1980a; Converse, 1980b). First, it has been suggested that our requirement of complete data on all fifteen attitude items (5 issues by 3 times) for inclusion in the sample had a large and differential impact upon the two samples. It is reasonable to assume that more respondents were deleted from the low-education group than from the high-education group due to missing data. Assuming that individuals with missing data have less organized attitude systems, it seems reasonable that the low-education group would have exhibited less stability and consistency had individuals with missing data somehow been included.

Second, it has been suggested that the amount of residual variance in the indicators of the Judd and Milburn (1980) model is an important indicator of attitude structure. The residual variance in an indicator reflects the degree to which variance in the measured attitude is due to factors other than the time-specific ideology and the issue-specific construct. In unstandardized models, a larger residual variance does not necessarily imply less impact of the ideological factor or of the issue-specific construct, but it does show that there is a greater amount of error in the observed attitude. The estimates of residual variance presented by Judd and Milburn (1980) are greater on the average for the low-education group than for the high-education group. We did not test the significance of this difference, however.

Finally, it has been suggested that Judd and Milburn's (1980) use of education to define the samples does not adequately address the hypothesis. Converse (1964:229) suggested originally that the samples should be defined according to "political interest." He later stated that measures of political involvement empirically discriminate between political "elites" and the general public better than does education (Converse, 1979, 1980a).

One purpose of this paper is to assess the validity of these criticisms of the Judd and Milburn (1980) paper. We explore the missing data problem, test the magnitude of residuals, and examine stability and consistency in groups that differ on politi-

cal involvement as well as on education. Like Judd and Milburn (1980) we estimate the unstandardized coefficients of a latent variable structural model to overcome the problems in the original correlational procedures. We use the same data set as Judd and Milburn (1980): the 1972-1974-1976 election panel sample collected by the Center for Political Studies at the University of Michigan.

METHOD

Samples

It is now well established that there exist multiple dimensions of political involvement (Verba and Nie, 1972). Because we wish to compare groups that differ in involvement, it was necessary to identify the dimensions of involvement for the present sample. Hence, eleven involvement variables from the 1976 wave of the panel survey were factor analyzed, using a principle factoring procedure with iterated communalities. The eleven variables included measures of interest in the election campaign, self-reported efforts to persuade others, whether the respondent had signed petitions, whether he or she had written to newspapers or governmental officials, reported interest in following national affairs, the interviewer's judgment of the respondent's political knowledge, level of education, and a number of other potential indicators of political involvement. Three major factors accounted for the majority of variance in the unrotated factor solution, and hence three factors were rotated to enhance interpretation, using an oblique (Oblimin) rotation procedure.

The three involvement factors that emerged were readily interpretable. Loading highly on the first were campaign interest, national affairs interest, and whether the respondent reported trying to persuade others' electoral choices. This factor represents general interest and campaign activity, corresponding well with Verba and Nie's (1972) campaign activity factor. Loading highly on the second factor were the variables that assessed whether the respondent had written letters to political officials and/or newspapers. This factor seemed quite close to Verba

and Nie's (1972) contacting dimension. Loading on the third and final factor were respondent's education and the interviewer's judgment of respondent's political information. Thus, this factor seemed to be a general education or political information dimension.

In this paper we examine differences between groups that are high and low on each of these dimensions of involvement. To define the groups, we split the sample on three different variables, one loading highly on each of the three involvement factors. The three variables are expressed interest in the presidential campaign, whether or not the respondent reported writing to a public official, and education. To better distinguish extreme groups, we omitted those who fell in the middle of the three distributions. To further isolate extreme individuals on the first two dimensions, we selected respondents who were relatively stable in their level of interest and action during the four years of the panel study.

If we pay no attention to whether the attitude questions were answered and split the full panel sample of 1320, 308 respondents indicated in both 1972 and 1976 that they were "very much" interested in the presidential campaign. Of the remaining 1012 respondents, 597 reported that they were very interested in neither of the two years. The first group is defined as the high-interest group and the second as the low-interest group. We found that 264 respondents indicated in both 1972 and 1976 that they had written a public official on some political issue and 659 respondents did not indicate that they had written a public official either year. We defined the first sample as the high-write group and the second as the low-write group. We found that 257 respondents had attended four years of college or more by 1976 and 841 respondents had not attended college at all by 1976. The first group was defined as the high-education group, and the second was defined as the low-education group.

Measures

Judd and Milburn (1980) studied responses to five attitude questions at each

of three time points. In order to conduct the present analysis with the largest possible samples, we specified a simpler model than the one used by Judd and Milburn (1980). The linear structural model can be estimated with as few as three different attitudes measured at two time points. Hence, in the present analysis, we examined attitudes in only 1972 and 1976, choosing the three attitude issues that resulted in the largest samples with complete data.

The attitude item on which there was by far the most missing data was the liberal/conservative self-rating scale used by Judd and Milburn (1980). Hence, it was omitted in the present analysis. As we would expect, many more individuals in the low interest, write, and education groups had missing data on this item than in the high groups. The large amount of missing data on this item, especially in the low groups, supports Converse's (1964) notion that the general public, particularly the less politically involved members of that general public, have a great deal of trouble expressing an ideological position.

In addition to deleting the liberal/conservative item, we did not use the measure of respondents' attitude about the rights of those accused of committing crimes, both because it had a substantial amount of missing data and because Judd and Milburn (1980) found it to be less consistent with the underlying ideology than were the remaining three indicators.

The three attitude measures that we did use are the following:

(1) Some people think achieving racial integration of schools is so important that it justifies busing children to schools out of their own neighborhoods. Others think letting children go to their neighborhood schools is so important that they oppose busing. Where would you place yourself on this scale (1 = Bus to achieve integration; 7 = Keep children in neighborhood schools)?

(2) Some people feel that the government in Washington should see to it that every person has a job and a good standard of living. Others think that the government should just let each person get ahead on his own. Where would you place yourself on this scale (1 = Government see to job and good standard of living; 7 = Government let each person get ahead on his own)?

(3) Some people feel that the government in Washington should make every possible effort to improve the social and economic position of blacks and other minority groups. Others feel that the government should not make any special effort to help minorities because they should help themselves. Where would you place yourself on this scale (1 = Government should help minority groups; 7 = Minority groups should help themselves)?

Table 1 presents the sample sizes for each of the six involvement groups. In the first column of Table 1, the sizes of the six groups are presented when we do not require any attitude responses. In the second column, we present the six sample sizes that result if we adopt Judd and Milburn's (1980) strategy of requiring complete data on five items at three time points. Below the sample sizes in this column are the percentages of the full samples in the six involvement groups with complete data. The third column presents the sample sizes when complete data are required on the three items used in this analysis at only two time points. Also presented in this column are the percentages of the full samples in the six groups having complete data on the six attitude questions. The models estimated in this paper used the samples in this third column.

It should be clear from this table that our samples represent a substantial improvement over those used by Judd and Milburn (1980). Nevertheless, we still must delete a substantial number of cases, particularly in the low-involvement groups, when complete data are required on only six attitude items. The dispropor-

tionate amount of missing data in the low-involvement groups supports Converse's (1980b) notion that the politically uninvolved are less likely to express political attitudes.

The Model

As we have said, the model used in this paper is a simplified version of Judd and Milburn's (1980) structural model. It is one of a set of models for panel data that has been suggested by Wheaton, Muthén, Alwin, and Summers (1977).

The model is defined by two structural equations:

$$X_{ij} = \alpha_{ij}T_j + \beta_{ij}I_i + U_{ij} \quad (1)$$

$$T_j = \gamma T_1 + \zeta \quad (2)$$

In the first equation, called the measurement model, X_{ij} is a specific attitude of the three ($i = 1$ to 3) measured in either 1972 or 1976 ($j = 1, 2$). T_j is the single time-specific latent variable that is assumed to explain all the covariation between the X_{ij} at any one point in time, and I_i is an item-specific construct which explains covariance between the same attitudes across time. For conceptual purposes, the T_j can be thought of as an underlying ideology. The α_{ij} is a loading coefficient of X_{ij} on T_j , and β_{ij} is a loading coefficient of X_{ij} on I_i . U_{ij} is residual variation in X_{ij} uncorrelated with T_j and I_i . These residuals (U_{ij}) are assumed to be uncorrelated with one another.

The second equation defines the structural effects of the latent ideology in

Table 1. Sample Sizes of Involvement Groups with and without Complete Data

	Total Sample	Judd & Milburn (1980) Complete Data (5 questions × 3 waves)	Present Model Complete Data (3 questions × 2 waves)
High Interest	308	163 (52.9%)	234 (76.0%)
Low Interest	597	178 (29.8%)	315 (52.8%)
High Write	264	158 (59.8%)	209 (79.2%)
Low Write	659	180 (27.3%)	355 (53.9%)
High Education	257	160 (62.3%)	205 (79.8%)
Low Education	841	226 (26.9%)	439 (52.2%)

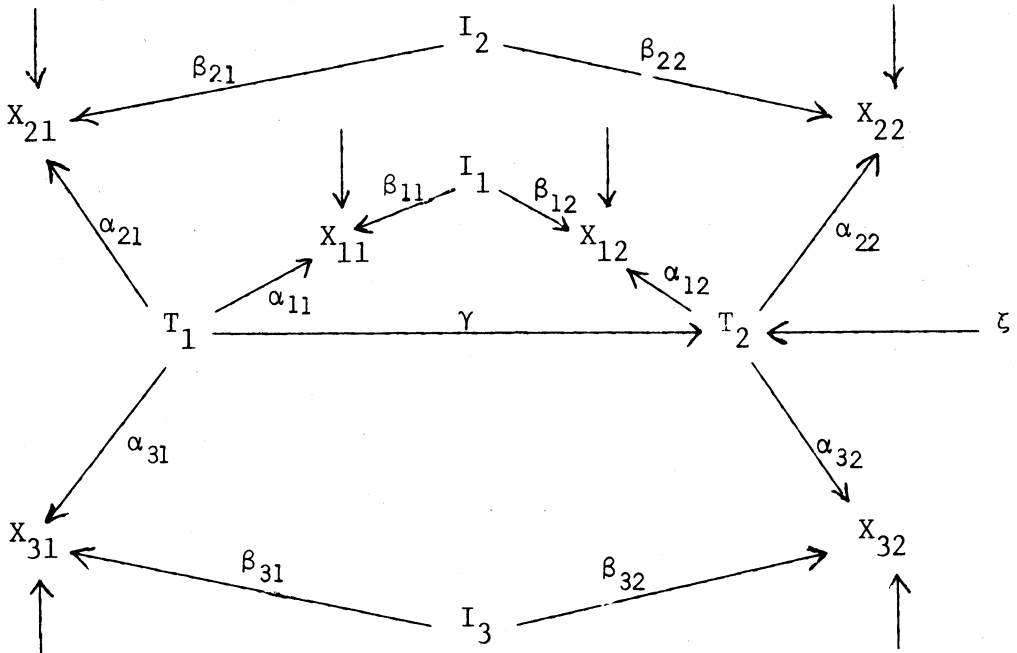


Figure 1. Diagram of the Measurement and Structural Models

1972 (T_1) on the latent ideology in 1976 (T_2). The stability coefficient is γ , and ζ is disturbance variance in T_2 , uncorrelated with T_1 . The measurement and structural models are diagrammed in Figure 1.

Estimation

The unknown parameters of the model were estimated using the LISREL procedure (Version IV; Jöreskog and Sörbom, 1978). This procedure generates maximum likelihood parameter estimates under the assumption of multivariate normality. The following unknown parameters of the model were estimated:

- the α_{ij} loading coefficients;
- the stability coefficient, γ ;
- The U_{ij} disturbance variances;
- the ζ disturbance variance; and
- the variances of the latent variables ($\sigma_{T_1}^2, \sigma_{T_2}^2, \sigma_{I_1}^2, \sigma_{I_2}^2, \sigma_{I_3}^2$).

In order to estimate these parameters, the model must be identified; that is, there must be sufficient information in the variance-covariance matrix. In the present model, two kinds of restrictions or constraints must be put on the model in order for it to be identified. First, to estimate the variances of the latent con-

structs, one structural coefficient from each to an X_{ij} must be constrained at unity (Kenny, 1979). Second, since the same attitude is only measured twice, to estimate the variances of the item-specific factors (T_i), both structural coefficients from any I_i must be constrained at unity (i.e., $\beta_{i1} = \beta_{i2} = 1$).¹

Assuming that the parameters of the model are identified, LISREL estimates them from the sample variance-covariance matrices. In the present case, the parameters of the model are overidentified and, hence, a goodness-of-fit test can be conducted to determine whether the hypothesized model is consistent with the data. For the present model the resulting χ^2 has 5 degrees of freedom.

In addition, the parameters of the model can be estimated simultaneously in multiple groups, and certain parameters can be constrained to be equal between groups.

¹ Instead of estimating the variances if I_i , they could have been fixed at some value and the β_{ij} coefficients could have been estimated, assuming that $\beta_{i1} = \beta_{i2}$. Such a model, however, is equivalent to the one we estimated, both structurally and substantively. A single structural parameter can be estimated for each I_i and it makes no difference whether it is the variance of I_i or β_{ij} .

Table 2. Maximum Likelihood Parameter Estimates; No Between-Group Constraints

	High Interest	Low Interest	High Write	Low Write	High Education	Low Education
α_{11}	.943	.537	1.011	.600	1.150	.543
α_{21}	1.007	.580	1.008	.775	.954	.823
α_{31}	1.000 ^a	1.000 ^a	1.000 ^a	1.000 ^a	1.000 ^a	1.000 ^a
α_{12}	.852	.472	.841	.584	1.038	.534
α_{22}	.861	.766	.981	.844	1.005	.937
α_{32}	1.000 ^a	1.000 ^a	1.000 ^a	1.000 ^a	1.000 ^a	1.000 ^a
γ	1.018	.855	.936	.990	.881	.889
ζ	.121	.768	.232	.445	.309	.416
$\sigma_{T_1}^2$	1.623	1.887	1.440	1.601	1.443	1.776
$\sigma_{T_2}^2$	1.801	2.148	1.492	2.014	1.429	1.820
$\sigma_{U_{11}}^2$.664	.624	.606	.659	.840	.427
$\sigma_{U_{12}}^2$.423	.477	.110	.406	.071	.484
$\sigma_{U_{13}}^2$.189	.036	.494	-.100 ^b	.255	.108
$\sigma_{U_{21}}^2$	1.165	.941	1.003	1.088	1.227	.977
$\sigma_{U_{21}}^2$	1.709	2.247	1.589	2.549	1.405	2.525
$\sigma_{U_{31}}^2$	1.648	1.381	1.092	2.062	.882	1.985
$\sigma_{U_{12}}^2$.856	1.162	.915	1.122	.748	1.245
$\sigma_{U_{22}}^2$	1.559	1.346	1.372	1.606	1.008	1.503
$\sigma_{U_{32}}^2$	1.443	1.190	1.109	1.746	.913	1.657
$\chi^2(10)$	6.699		11.875		14.187	
p	.754		.294		.165	

^a These coefficients were constrained at unity.

^b This negative item-specific variance is nonsignificant. It indicates a nonsignificant negative relationship between X_{31} and X_{32} when ideology is controlled.

Such simultaneous estimates also yield goodness-of-fit tests that can be used to determine whether a model with between-group equality constraints is less consistent with the sample data than a model without such constraints.

RESULTS

As discussed in the previous section, the pattern of missing data is consistent with Converse's (1964) hypothesis that the less involved have difficulty expressing political attitudes. While many respondents have missing data on the attitude items, especially in the low-involvement groups, we can still examine whether attitude structures differ by political involvement among the majority in each group from whom we do have complete data.

When discussing estimation, we said that one structural coefficient from each of the latent ideologies (T_j) must be constrained at unity to derive their variances. In essence this defines the metric of the latent construct to be the same as the variable whose loading has been constrained. Judd and Milburn (1980) had a theoretical rationale for their choice of metric for the

latent ideology. We constrained the loading of the liberal/conservative item, since the ideology was assumed to represent an underlying liberal/conservative ideological predisposition. In the present case, we have no theoretical rationale for the choice of metric of the latent ideology. Hence, we randomly chose to fix the loadings of the government aid to minorities items.

Table 2 presents the maximum likelihood parameter estimates from the six involvement groups. Also included in this Table are the χ^2 goodness-of-fit statistics for the simultaneous fit of the models in the high- and low-involvement groups with no between-group constraints. Regardless of how we define involvement, the hypothesized model is consistent with the data in the high- and low-involvement groups.²

The first step in examining whether the

² These goodness-of-fit statistics test whether the departures of the sample data from the variance/covariance matrices predicted by the model are significant. If they are not, and thus the χ^2 is not significant, the model is said to be consistent with the data (Kenny, 1979).

attitude structures differ between high- and low-involvement groups is to determine whether the measurement models are equivalent in the high and low groups. This involves constraining the α_{ij} coefficients to be invariant across involvement level and then examining whether the resulting models fit significantly less well than the unconstrained models. The difference between the χ^2 s for the unconstrained and constrained models is itself a χ^2 and can be used to determine whether the constrained model fits less well. For the high- and low-interest groups, the difference $\chi^2(4) = 14.577$, $p < .01$; for the high- and low-write groups, the difference $\chi^2(4) = 7.415$; $p = .11$; and for the education groups, the difference $\chi^2(4) = 24.271$; $p < .001$. Hence, for interest and education, the α_{ij} coefficients are significantly different in the high- and low-involvement groups. For the write groups, the difference approaches significance. In the case of all three dimensions of involvement, then, the conclusion seems to be that the α_{ij} coefficients, the loading coefficients of the specific attitudes on the latent ideology, differ by level of involvement.

It is tempting to compare the α_{ij} coefficients of Table 2 between involvement levels and conclude that specific attitudes are less ideological in the low than in the high groups, since the α_{ij} coefficients are without exception smaller in the low-involvement groups. Such comparisons, however, must be made relative to the attitude whose loading has been constrained at unity. In other words, the unequal α_{ij} coefficients in Table 2 permit the conclusion that, relative to the aid-to-minorities issue, the busing and guaranteed-job issues are less ideological in the low-involvement than in the high-involvement groups. If another X_{ij} had been used to define the metric of the latent ideology, very different α_{ij} coefficients result. For instance, if the loading coefficients for the busing attitude are constrained at unity, the remaining α_{ij} coefficients are all larger in the low-involvement groups than in the high groups. In essence, then, the unequal loading coefficients permit us only to examine differences between involvement levels in

the relative impact of ideology across issues. For the low groups, the busing issue is consistently the least ideological, while the aid-to-minorities issue is consistently the most. For the high groups, this ordering of issues is reversed: busing is the most ideological and aid to minorities the least. In the absence of a theoretical rationale for defining the metric of the latent ideology, comparisons of the overall impact of ideology across issues cannot be made between groups.

Examination of the coefficients of Table 2 suggests that the latent ideological factor is remarkably stable over the four-year interval in all six groups. All of the stability coefficients are greater than .85. These coefficients can be interpreted just as unstandardized regression coefficients are interpreted: for every one unit difference in the 1972 ideology we can expect .85 or greater units difference in 1976.

Is the latent ideology less stable in the low-involvement groups than in the high groups? Inspection of the coefficients of Table 2 suggests they are not. In the write groups, the stability coefficient is larger in the low group than in the high group. In the education groups, the stability coefficients are essentially equal. Only in the interest groups does there seem to be less stability among the less involved. However, such comparisons are also contingent upon the issue used to define the metric of the latent ideology so long as equal measurement models cannot be assumed between groups. With equal measurement models, the results of comparisons of stability coefficients are invariant across different metrics of the latent ideological factor. However, when the measurement models differ between groups, as we know they do in the present case, stability comparisons are contingent upon the metric used for the latent construct.

In essence, then, we can adopt one of three strategies in comparing stability coefficients between latent constructs in unstandardized models: (1) If we have a theoretical rationale for defining the metric of the latent factor, we can test for differences in stability under that metric (this is the strategy used by Judd and Milburn, 1980); (2) If the measurement model

is invariant across groups, we can compare stability coefficients between groups no matter which indicator is used to define the metric of the latent factors; (3) In the absence of a theoretical rationale for defining the metric, and when measurement models are known to differ between groups, comparisons of stability coefficients must be conducted defining the factor metrics all possible ways, i.e., using each indicator to fix the factor metric. If stability comparisons between groups yield essentially the same result regardless of the factor metric, we can have some confidence in the conclusions.

In the present situation, the third strategy must be adopted since we have shown unequal measurement models across involvement levels and since no theoretical rationale exists for the choice of the factor metric. Hence, for each of the three involvement dimensions, three different tests of equal stability in the high and low groups were computed, defining the metric of the factors using each of the three attitude items. The χ^2 tests of these nine comparisons are presented in Table 3. Each of the nine difference χ^2 s in this table has a single degree of freedom. None of them approach significance. Hence, there is no evidence of unequal stability in high- and low-involvement groups, regardless of the dimension used to define involvement and regardless of the attitude item used to determine the metric of the latent ideology.

Table 3. Chi Squares Testing Equal Stability

	Item which Fixes Metric		
	Busing Issue	Guaranteed Job Issue	Aid to Minorities Issue
Interest Groups	.281	.253	.004
Write Groups	1.097	1.652	1.037
Education Groups	1.290	.770	.107

NOTE: All χ^2 s have a single degree of freedom. All are nonsignificant.

The residual variances to the indicators ($\sigma^2_{U_{ij}}$) do not depend upon the metric used to define the factors. Hence, comparisons of the amount of residual error in the indicators yield the same result regardless of the factor metric. In Table 2, differences

in the residuals are consistent in direction for both the write and education dimensions. As Converse (1980a) suggested, there seems to be more residual error in the low-involvement groups than in the high ones. This difference is not, however, apparent in the interest groups. Tests of equality of residual variances between high- and low-involvement groups reveal that: (1) in the write groups, residual variances are consistently and significantly larger in the low-involvement group than in the high, $\chi^2(6) = 21.14$, $p < .005$; (2) in the education groups, the residual variances differ significantly, $\chi^2(6) = 41.34$, $p < .001$. All but one of them are larger in the low group than in the high group. (3) In the interest groups, residual variances do not differ significantly by involvement level; $\chi^2(6) = 7.45$; $p > .25$.³

DISCUSSION

Judd and Milburn (1980) have developed an approach for examining differences in attitude structures that has distinct advantages over the original techniques used by Converse (1964) and others. Nevertheless, a number of criticisms have been made of this approach. In this paper, we have tried to evaluate those criticisms while keeping the advantages of Judd and Milburn's (1980) approach. Our results suggest that the issue of interest, differences in attitude structure, is an exceedingly complex one.

The following conclusions summarize

³ One further apparent difference in the models between involvement groups deserves comment. The disturbances to the 1976 latent ideology, ζ , reported in Table 2, are higher in the low than in the high involvement groups. The magnitude of these disturbances, however, depends upon the metrics of the latent constructs. We tested for differences in ζ between groups under all three metrics. For the interest groups, ζ is higher in the low than in the high group under all three metrics, significantly so in two of the three cases. For the write groups, ζ is nonsignificantly higher in the low than in the high group under all three metrics. For the education groups, ζ is nonsignificantly higher in the low than in the high group under two of the three metrics. Under the third metric the high group has a nonsignificantly higher ζ . The general trend in these differences suggests that while ideology is equally stable between involvement groups, there is a tendency towards greater time-specific variation in the latent ideologies of the low involvement groups.

the results that have been presented in this paper:

(a) The pattern of missing data on the attitude questions supports the notion that those who are relatively uninvolved politically tend to have difficulty expressing political positions. This is particularly true when the attitude question asks the respondent to locate him- or herself along a fairly abstract ideological continuum.

(b) Among those who have complete data on the three issues used in the model we have examined, attitude structures differ between more and less involved groups in some ways but not in others. Different issues are ideologically derived to differing degrees in the high- and low-involvement groups. Some issues seem relatively more ideological for those with high involvement, while other issues seem to be relatively more ideological for those who are low on involvement. Thus, consistent with Judd and Milburn's (1980) findings, the role of ideology seems to depend jointly on the nature of the issue and the respondent's level of involvement. No matter how we define the latent ideological construct, there seems to be no evidence in our data for greater ideological stability among those who are more politically involved. At the same time, however, the attitude measures in the less involved groups seem to contain more error of measurement than in the high-involvement groups, so long as we use either education or political contacting to define political involvement.

(c) Generally, the results we have reported, concerning both the pattern of missing data and the attitude structures that emerge among those having complete data, do not depend upon the various ways of defining political involvement. The only exception to this conclusion concerns the fact that the attitude variables in the low-interest group do not have larger residuals than in the high-interest group, although differences in residuals are apparent when political involvement is defined in the other two ways.

These results support Converse's (1964) notions that there are differences between the more and less involved in the political attitudes they express. The less involved

are more reluctant to express attitudes, and, when they do, those expressions are less reliable, in the sense that they reflect more measurement error. At the same time, however, many of Judd and Milburn's (1980) conclusions are also supported by our results. Among those who do express political attitudes, the presence of an underlying political ideology and the stability of such an ideology do not seem to depend upon the level of political involvement.

In arriving at these conclusions, we have made a number of methodological points concerning procedures for comparing latent structures. These methodological points are implicit in the statistical literature developing the factor-comparison techniques we have used (Jöreskog and Sörbom, 1979). Nevertheless, they are worth reiterating here.

First, comparisons between groups should generally be made with unstandardized coefficients, since such coefficients are relatively unaffected by differences in variability between groups. Second, the results of comparisons of structural relations among latent constructs depend upon the measurement model used to define those constructs. If measurement models are invariant across groups, then comparisons yield identical results regardless of the metric used to derive the variances of the latent constructs. If measurement models are not invariant across groups, then a theoretical rationale must exist for defining the metrics of the constructs. If such a rationale does not exist, then comparisons of structural coefficients between groups should proceed only if different metrics yield consistent results—that is, only if the results of the comparisons are consistent regardless of the metric used. Finally, comparisons of factor loading coefficients between groups must also be done relative to the metric of the factor. All of these points suggest that the researcher must be quite cautious in comparing covariance structures between groups.

While caution is appropriate, such comparisons can nevertheless yield results that are theoretically interesting. In the present case, these comparisons suggest a complex picture concerning the re-

relationship between political involvement and the organization of political attitudes. The less involved seem less able to articulate political attitudes. In addition, when they do, there is more variation in them that is predictable neither from an underlying ideology nor from the same attitudes at other time points. At the same time, however, the ideological base that is reflected in the various attitudes seems to be as reliable and stable in the low-involvement groups as in the high-involvement groups.

REFERENCES

- Achen, Christopher H.
1975 "Mass political attitudes and survey response." *American Political Science Review* 69:1218-31.
- Barton, Allen H. and R. Wayne Parsons
1977 "Measuring belief system structure." *Public Opinion Quarterly* 41:159-80.
- Blalock, H. M.
1967 "Path coefficients versus regression coefficients." *American Journal of Sociology* 72:675-6.
- Converse, Philip E.
1964 "The nature of belief systems in mass publics." Pp. 75-169 in David E. Apter (ed.), *Ideology and Discontent*. New York: Free Press.
1979 "Comment." *Journal of Politics* 41:1182-4.
- 1980a "Comment: rejoinder to Judd and Milburn." *American Sociological Review* 45:644-6.
1980b "Recent thoughts on attitude constraint and stability." American Psychological Association annual meeting, Montreal.
- Converse, Philip E. and Gregory B. Markus
1979 "Plus ça change . . . : the new CPS election study panel." *American Political Science Review* 73:32-49.
- Jöreskog, Karl G. and Dag Sörbom
1978 LISREL: Analysis of Linear Structural Relationships by the Method of Maximum Likelihood. Chicago: National Education Resources, Inc.
1979 *Advances in Factor Analysis and Structural Equation Models*. Cambridge, Ma.: Abt Books.
- Judd, Charles M. and Michael A. Milburn
1980 "The structure of attitude systems in the general public: comparisons of a structural equation model." *American Sociological Review* 45:627-43.
- Kenny, David A.
1979 *Correlation and Causality*. New York: Wiley.
- Nie, Norman H., Sidney Verba, and John R. Petrocik
1976 *The Changing American Voter*. Cambridge: Harvard University Press.
- Verba, S. and N. H. Nie
1972 *Participation in America*. New York: Harper and Row.
- Wheaton, Blair, Bengt Muthen, Duane F. Alwin, and Gene F. Summers
1977 "Assessing reliability and stability in panel models." Pp. 84-136 in D. R. Heise (ed.), *Sociological Methodology 1977*. San Francisco: Jossey-Bass.