Members’ Participation in Local Union Activities: Measurement, Prediction, and Replication

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Past research on members’ participation in union activities has been characterized by the lack of a psychometrically adequate criterion for participation and an exploratory focus on the correlates of participation. We report two studies addressing these issues. In the first study, we found that participation was best understood as a unidimensional and cumulative construct. Across three large data sets (N = 229, 551, and 413), a Guttman scale model provided a good fit to the data. The scale demonstrated temporal reliability, and correlations with external criteria supported the construct validity of the measure. In the second study, we formulated a model to predict members’ participation on the basis of the theory of reasoned action and the partisan model of political participation. After slight modification, the proposed model fit the data for the study (n = 202) and for replication samples (n = 147). Directions for future research on union participation, conceptual implications for the understanding of union participation, and practical implications for union governance are discussed.

Labor unions have been described as utilitarian and normative organizations (Schein, 1980). As utilitarian organizations, unions provide members with the benefits of collective bargaining (e.g., job security, wages, fringe benefits, and access to grievance procedures; see Barling, Fullagar, & Kelloway, 1992; Freeman & Medoff, 1984; Kochan, 1980). As normative organizations, the existence of labor unions is predicated on members’ wanting to belong to the union and wanting to fulfill their roles in the organization, that is, on moral involvement in the union (Schein, 1980). The utilitarian and normative nature of unions has led to the paradox of members’ participation in union activities (Olson, 1971). According to Olson, “over 90 per cent of union members] will not attend meetings or participate in union affairs; yet over 90 per cent will vote to force themselves to belong to the union and make considerable dues payments to it” (Olson, 1971, p. 86). For over 45 years (for reviews see Barling et al., 1992; Perline & Lorenz, 1970; Spinrad, 1960), behavioral scientists have attempted to resolve this paradox through a search for the correlates of members’ participation and, in particular, why active unionists participate in union activities (Kolchin & Hyclak, 1984). We extend this research by (a) examining the psychometric properties of a measure of members’ participation in local union activities and (b) developing, evaluating, and replicating a structural model to predict such participation.

The widespread lack of members’ involvement in union activities is a prominent source of concern for North American (Huszczo, 1983; Miller, Zeller, & Miller, 1963; Spinrad, 1960) and British (Sherman, 1986) labor leaders. As a visible manifestation of individual attachment to labor unions (Barling et al., 1992), the involvement of the rank and file in union activities is thought to reflect members’ support for union goals (Form & Dansereau, 1957) and members’ satisfaction with the union (Strauss, 1977) and, most important, is thought to provide a fundamental test of union democracy (Gordon & Nurick, 1981; Strauss, 1977).

Though definitions of union democracy abound (Nicholson, 1978), little attention has been paid to the definition of union participation. There is agreement that members’ participation is a behavioral construct, requiring the expenditure of time on union affairs (e.g., Sayles & Strauss, 1952). Perhaps as a result of this simplistic definition, the construct of participation has been defined primarily through operationalization (Kryl, 1990), with little attempt to understand how various forms of participation are related. Meeting attendance, voting in union elections or other union votes, serving on union committees, holding union office, and use of the grievance procedure are the most commonly used indicators of members’ participation (Spinrad, 1960). As Kahn and Tannenbaum (1954) noted, these are all measures of formal participation. In addition to being highly visible behaviors, formal activities are regulated or controlled to some extent by the structure or constitution of the union (Miller & Young, 1955; Shepard, 1949; Steele, 1951). In contrast, frequently used measures of informal participation,
such as reading union literature (Anderson, 1979; Hagburg, 1966; Sayles & Strauss, 1952) and talking to shop stewards or other union members about union or work issues (Blyton, Nicholson, & Ursell, 1981; Huszczko, 1983; Miller & Young, 1955), are less visible and less regulated by union policies.

Some general conclusions may be drawn from this listing of participation measures. First, the types of participation considered in previous research are not independent, and the behaviors chosen to represent the construct of union participation imply differential levels of involvement in the union (Nicholson, 1978). The differential requirements of participation have been largely ignored, with Spinrad (1960) claiming that research findings have not been affected by the particular criterion chosen to reflect members' participation. Similarly, the interdependence of different forms of participation has been ignored in previous research. Second, with some recent exceptions (e.g., Gallagher, McLean-Parks, & Wetzel, 1987; McShane, 1986a), the clear focus of the research has been on participation at the local, rather than the national or regional, level of the union. The first goal in this research was to develop a psychometrically adequate measure of members' participation that reflected known patterns of participation in local union activities.

Research on union participation has proceeded piecemeal with little or no attempt to integrate findings across studies or to examine the relative impact of different variables on members' participation in the union. Accordingly, our second goal was to develop a literature-based model predicting members' participation in union activities. We suggest that members' participation is most appropriately viewed as the outcome of a socialization process leading to the development of individual commitment to the union. It is this suggestion that provided the main impetus for the second study.

A characteristic problem of research on members' participation in union activities is reliance on data from one particular research setting in any given study. Such reliance leaves the generalizability of any research findings open to question and maximizes the possibility of capitalization on chance in the identification of measurement or predictive structures (Cook & Campbell, 1979). To minimize the capitalization on chance and to maximize the possibility of obtaining stable and generalizable results in both studies, we used data drawn from different local unions to assess the measurement and structural models.

Study 1

The goal of the first study was to develop and evaluate a criterion for members' participation in local union activities. A review of the empirical research on members' participation suggests two predominant definitions of members' participation, a unidimensional and a multidimensional definition. Both of these definitions are predicated on the common factor model, in which participation in various union activities is assumed to be the result of one or more common factors. In contrast, we propose a third unidimensional model based on a cumulative definition of unidimensionality.

Much of the research on members' participation has implicitly or explicitly adopted a unidimensional definition of participation. Researchers have operationalized participation with both single-item measures (e.g., Glick, Mirvis, & Harder, 1977; Kolchin & Hyclak, 1984) and composite indexes (e.g., Chacko, 1985; Fullagar & Barling, 1989; Huszczko, 1983). The unidimensional view of participation emerges directly from Spinrad's (1960) comment that research findings have not been affected by the nature of the criterion.

Both McShane (1986a) and Klandermans (1986) challenged the unidimensional view, with McShane citing modest indices of internal consistency (e.g., .50 to .75; Anderson, 1979; Fullagar & Barling, 1989; Nicholson, Ursell, & Lubbock, 1980) and mixed evidence for parallelism (Hunter & Gerbing, 1982) as indications of multidimensionality. Three studies have explicitly tested the dimensionality of members' participation in union activities. Both McShane (1986a) and Gallagher et al. (1987) reported support for a three-factor solution, though an attempt to replicate McShane's (1986a) findings using the same participation measure resulted in a unidimensional solution (Kuruvilla, Gallagher, Fiorito, & Wakabayashi, 1990). The interpretability of these factor analytic studies is clouded by numerous methodological issues, including the application of factor analysis to dichotomously scored behavioral data (Comrey, 1978), the interpretation of "doublets" as meaningful factors (Nunnally, 1978), the instability of derived solutions across subsamples (e.g., Gallagher et al., 1987) and populations (e.g., Kuruvilla et al., 1990) and the choice of highly unrealistic factor models (Barling et al., 1992).

Members' Participation as a Cumulative Construct

The move toward a more multidimensional definition of members' participation in union activities is explicitly predicated on the differential requirements of various forms of participation (Klandermans, 1986; McShane, 1986a). Specifically, it is widely accepted that there are different degrees of union participation (Nicholson, 1978; Sayles & Strauss, 1952; Tannenbaum & Kahn, 1958), and McShane (1986a) criticized the practice of deriving composite indexes of participation for failing to take these degrees of participation into account. However, factor analytic research on members' participation has not recognized the measurement implications of these degrees of involvement.

On the basis of Kahn and Tannenbaum's (1954) observation that there are different degrees of involvement associated with various forms of participation, Nicholson (1978) proposed a continuum of union participation. He suggested that the continuum of participation might extend from taking a direct role in the leadership of the union through holding union office and various levels of involvement (e.g., serving on committees, attending union meetings), to passive participation in the form of reading union newsletters and paying union dues. This continuum is ordered along a dimension Nicholson (1978) referred to as involvement in the union. There is also evidence suggesting that the continuum is ordered along a dimension representing opportunities to participate in the union. For example, holding union office is not a form of activity available to most union members at any given time. Typically, a local union will have
between four and six members who serve on the executive committee and a larger group of individuals who act as shop stewards. A widely held rule of thumb would suggest that there is one steward for every 50 members (Barling et al., 1992). Thus, in a union local comprising 500 members, 15 to 20 members (i.e., 3–4%) could possibly hold a union office. Although all members have the opportunity to attend union meetings, estimates of meeting attendance range from 10–15% (Gordon, Phlipot, Burt, Thompson, & Spiller, 1980) to 20–25% (Roy, 1964) of the membership. Substantially more members turn out for union elections, with Roy (1964) reporting an average of 40–45% of the membership casting a ballot. There are few data on which to base estimates of participation rates for the other activities mentioned by Nicholson (1978), although in the Canadian context, 100% of the membership is required to pay union dues. Union officials estimate that 90% of the members read union literature (Davis, 1980).

It is critical to realize that the activities listed in the previous paragraphs are not independent. Perhaps the best predictor of meeting attendance is whether or not the individual holds a union office (Barling et al., 1992). Presumably, members who display enough interest in the union to come out to a union meeting will also turn out for union votes, and those who vote can be expected to read union newsletters and other literature to become informed on the issues or candidates. Individuals who report involvement at a specific level of the continuum are also likely to be involved in activities ranked lower on the continuum. Thus, the participation activities identified by Nicholson (1978) are ordered along a single, and cumulative, dimension.

Several measurement implications ensue from this unidimensional and cumulative view of participation. First, a cumulative measurement model calls for rejection of the common factor model, in which responses to individual items are assumed to be the result of one or more underlying common factors or dimensions. Instead, items ordered along a single and cumulative dimension follow a deterministic scale model (Nunnally, 1978) first identified by Guttman (1944, 1947). The correlation matrix of Guttman-scalable items is a function of an underlying causal structure in which items are ordered in a causal sequence and manifest a simplex structure (Guttman, 1954); that is, correlations decrease as they move away from the main diagonal. Exploratory factor analytic techniques applied to Guttman-scalable data may result in misleading results (Jöreskog, 1978). Factor analysis of dichotomously scored, Guttman-scalable data will result in the extraction of multiple common factors, with the number of common factors increasing as the number of items on the scale increases (Weisberg, 1984).

We propose that the appropriate model for measuring members’ participation in local union activities is the Guttman scale model. In this study, we evaluated the psychometric properties of a Guttman scale of members’ participation in local union activities. One question with Guttman scales is their variation across samples (Babbie, 1973). Accordingly, we tested the scalability of participation items across three large samples drawn from independent unions. Use of three independent samples allows greater confidence in the generalizability of the results (Cook & Campbell, 1979).

Klandermans (1986) pointed out that participation in union activities is stable across time. Except in times of internal dissent or conflict with management, union members display a stable pattern of participation in union activities. Therefore, we assessed the stability of participation across six months in one sample as evidence for the temporal reliability of the scale.

Assessment of Construct Validity of the Cumulative Union Participation Measure

Assessing the scalability of participation items provided an indication of the measurement structure of the scale. It was equally important to establish the construct validity of the measure. Although the literature on members’ participation in local union activities is fragmented, sufficient data exist to propose criteria that correlate with members’ participation. The measure of participation was expected to correlate positively with three measures of union commitment: willingness to work for union, responsibility to the union, and union loyalty (Fullagar, 1986; Fullagar & Barling, 1989; Gordon et al., 1980; Kelloway, Catano, & Southwell, 1992). Participation was also expected to correlate positively with perceptions of union instrumentality (Chacko, 1985; Fullagar & Barling, 1989; Kelloway et al., 1992), satisfaction with the union (Glick et al., 1977), general union attitudes (McShane, 1986b), and union tenure (Kelloway & Catano, 1988; Spinrad, 1960).

Negative correlations were hypothesized between participation in union activities and both intrinsic and extrinsic job satisfaction (Fullagar & Barling, 1989; Huszczko, 1983). Neither respondents’ age nor their education were expected to correlate with the measure of participation in local union activities (Kelloway & Catano, 1988).

Method

Subjects Subjects for this study were members of one of three large Canadian unions. Data from Samples 1 and 2 have been previously reported by Kelloway et al. (1992), and data from Sample 3 was originally reported by Fullagar and Barling (1991). Both studies focused on union commitment. The first (Kelloway et al., 1992) focused on the dimensionality of a shorter union commitment scale, and the second (Fullagar & Barling, 1991) focused on the development of dual commitment. In neither of these studies was the scalability of participation items examined.

Sample 1 was drawn from a survey of 630 unionized clerical and technical university employees. Usable responses were obtained from 36% of the union members (n = 229). Data made available by the union allowed for the comparison of the respondent group with the union membership in terms of both sex and occupational classification. In neither case were differences between the two groups significant (p > .05). The median age of the sample was 30 years, the average level of formal education was 16 years, and the average length of union membership was 5 years.

Sample 2 was drawn from a large national survey of approximately 5,800 unionized airline employees. Usable responses were obtained from 551 (9.5%). Although no direct comparisons between respondents and nonrespondents were possible, the ratio of female to male respondents (77:23) reflects the composition of the union as a whole. Analysis of the returned survey packages suggested that the preponderance of returns were from Ontario and British Columbia, with no replies from certain provinces of the country that had large member-
ship concentrations, suggesting that some locals refused to participate in the study. Therefore, the response rate of individuals who actually received the questionnaire is probably higher than 9.5%. The median age of the sample was 35 years, the average level of formal education was 16 years, and the average length of union membership was 6 years.

Sample 3 was drawn from a three-phase longitudinal study of members of two unions of university employees in two different universities in Ontario (a faculty and a staff union). The participation measure at Time 1 differed from that used at Time 2 and Time 3. For purposes of this study, data are taken from the second and third phases of data collection. Questionnaires were distributed to 1,920 members at Time 1, with 120 being returned as nondeliverable. At Time 1, 464 (25.5%) of the distributed questionnaires were returned. At Time 2, 426 of the 1,800 questionnaires (22.2%) were returned. At Time 3, questionnaires were distributed only to individuals who had responded at Time 1 or Time 2 (n = 603); 323 (53.6%) respondents completed questionnaires at Time 3. Of the two unions constituting the sample, members of the faculty union were older (M = 47.3 years vs. M = 37.6 years), were better educated (M = 20.8 years vs. M = 15.1 years), and had greater union tenure (M = 8.8 years vs. M = 6.8 years) than members of the staff union. Male respondents made up 72% of the faculty sample and 33% of the staff sample. Although comparison of the respondent group with nonrespondents was not possible, a comparison of the Time 1 respondents with subjects who provided data for more than one time period did not reveal any significant differences between these groups. Because initial inspection of the data did not reveal any differences in patterns of union participation across the two samples, data from both samples were combined for the current study.

Measures. Respondents in Sample 1 completed single-item measures of holding union office, serving on union committees, attendance at union meetings, voting in union elections, talking to union representatives about problems at work (e.g., grievances), and reading the union newsletter. The measures of holding union office, serving on union committees, and reading the union newsletter were dichotomously scored (1 = yes, 0 = no). All other measures had a 5-point frequency response scale (ranging from 1 = never to 5 = always). Respondents in Sample 1 also completed a 30-item union commitment scale (Gordon et al., 1980). Subsequent analysis suggested that the construct validity of this scale was threatened by the inclusion of negatively worded items, which introduced construct-irrelevant covariance (Kelloway et al., 1992). Therefore, a shorter 19-item version of the scale measuring the three dimensions of union commitment—union loyalty (5 items, α = .85), responsibility to the union (7 items, α = .79) and willingness to work for the union (3 items, α = .84)—was used. Respondents also completed Glick et al.'s (1977) 6-item union satisfaction measure (α = .79) and demographic items (age, education, and union tenure).

Members of Sample 2 completed the same union participation items as did members of Sample 1, as well as a 20-item union commitment scale (Friedman & Harvey, 1986). Although this scale was originally designed to reflect Friedman and Harvey's proposed two-factor definition of union commitment, subsequent analysis suggested that negatively worded items should be removed (Kelloway et al., 1992). The remaining 13 items reflect three dimensions of union commitment—union loyalty (6 items, α = .92), responsibility to the union (4 items, α = .76), and willingness to work for the union (3 items, α = .85). Because of length limitations imposed by the union, respondents in Sample 2 completed a two-item version of Glick et al.'s (1977) union satisfaction scale (the two items correlated .57) and provided information on age, education, and union tenure.

Members of Sample 3 completed a checklist of dichotomously scored (1 = yes, 0 = no) participation items. These included holding union office, serving on a union committee, attending union meetings, voting in union elections, participating in other union votes, and reading the collective agreement. Respondents also completed a 28-item union commitment scale (Ladd, Gordon, Beauvais, & Morgan, 1982). For this study, we again used only three commitment dimensions—union loyalty (10 items, α = .93), responsibility to the union (4 items, α = .71), and willingness to work for the union (4 items, α = .88). Respondents also completed Warr, Cook, and Wall's (1979) 15-item measure of job satisfaction, which assesses extrinsic (8 items, α = .69) and intrinsic (7 items, α = .78) job satisfaction. Brett's (1980) 8-item union attitudes scale (α = .86) and Chacko's (1985) 7-item measure of perceived union instrumentality (α = .79) were completed in addition to questions on age, tenure, and education.

Data analysis. The primary goal of this first study was to examine the hypothesis that participation in union activities is a unidimensional and cumulative construct that can be most appropriately represented by a Guttman (1944, 1947) scale. One method of evaluating this hypothesis is through scalogram analysis (Guttman, 1947), which results in a coefficient of reproducibility ranging from 0 to 1. Values exceeding .90 indicate that the items are consistent with the Guttman scale model (Guttman, 1947). Because coefficients of reproducibility can emerge through capitalization on chance, Green (1956) proposed an index of scalability, which accounts for artificial effects emerging from the judicious selection of items. Reproducibility coefficients greater than .90 and scalability coefficients greater than .60 (Green, 1956) are typically accepted as the minimum standards for scalability (Mc이ver & Carmines, 1981).

The correlation matrix of items that form a Guttman scale manifests simplex structure (Guttman, 1954), with the interitem correlations decreasing as they move away from the main diagonal. Structural modeling techniques provide explicit tests of the structure of a correlation matrix (Joreskog, 1970, 1978; Joreskog & Sörbom, 1984). This is a confirmatory approach in that an explicit causal structure giving rise to the simplex pattern is posited and then evaluated as to whether the data fit the proposed model. Given the self-report nature of the participation measures used in the current study, an added benefit was the possibility of estimating a quasisimplex model that allows for measurement error in the variables (Joreskog, 1970, 1978).

The models estimated for each sample shared a similar structure but varied according to the nature of the variables in the data set. For example, in Sample 1 holding union office was hypothesized to predict serving on union committees. Committee service predicted attending union meetings, which in turn predicted talking to union representatives about problems at work. This variable predicted voting in union elections, which predicted participation in other votes. Finally, participation in other votes predicted the reading of union literature. To allow the model to be identified, we constrained the unique variances of holding union office and serving on union committees to be equal, as were the unique variances of reading union literature and participation in union votes (Joreskog, 1970, 1978). Similar quasisimplex models were estimated within each sample using maximum likelihood estimation as implemented in LISREL VI (Joreskog & Sörbom, 1984).

Results

Scalability of participation items. Correlations for all participation items from the three samples are presented in Tables 1 and 2. Within each sample, the correlation matrix conformed to the hypothesized simplex structure, with the correlations decreasing as they move away from the main diagonal.

For Sample 1 the proposed simplex model provided an excellent fit to the data, χ²(10, N = 229) = 8.35, p > .05 (goodness-of-fit index [GFI] = .99, adjusted goodness-of-fit index [AGFI] = .97, and root mean square residual [rmsr] = .03). The Guttman scale of participation manifested a coefficient of reproducibility of .92 and a coefficient of scalability of .67.
Table 1
Correlations Between Participation Measures: Sample 1 (n = 229) and Sample 2 (n = 541)

<table>
<thead>
<tr>
<th>Measure</th>
<th>1</th>
<th>2</th>
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<th>4</th>
<th>5</th>
<th>6</th>
<th>7</th>
</tr>
</thead>
<tbody>
<tr>
<td>Hold union office</td>
<td>.64</td>
<td>.59</td>
<td>.35</td>
<td>.32</td>
<td>.24</td>
<td>.02</td>
<td></td>
</tr>
<tr>
<td>Serve on committees</td>
<td>.56</td>
<td>.63</td>
<td>.35</td>
<td>.35</td>
<td>.22</td>
<td>.05</td>
<td></td>
</tr>
<tr>
<td>Attend union meetings</td>
<td>.26</td>
<td>.30</td>
<td>.28</td>
<td>.39</td>
<td>.24</td>
<td>.08</td>
<td></td>
</tr>
<tr>
<td>Talk to union leaders</td>
<td>.10</td>
<td>.12</td>
<td>.27</td>
<td>.36</td>
<td>.23</td>
<td>.04</td>
<td></td>
</tr>
<tr>
<td>Vote in elections</td>
<td>.11</td>
<td>.19</td>
<td>.21</td>
<td>.10</td>
<td></td>
<td>.53</td>
<td>.06</td>
</tr>
<tr>
<td>Vote in other votes</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>.14</td>
</tr>
<tr>
<td>Read union literature</td>
<td>.09</td>
<td>.10</td>
<td>.03</td>
<td>.06</td>
<td>.14</td>
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</table>

*Note. Correlations for Sample 1 are presented above the diagonal, and correlations for Sample 2 are presented below the diagonal. Whether union members voted on issues other than elections was not assessed in Sample 2.*

For Sample 2 the simplex model also provided an excellent fit to the data, \( \chi^2(6, N = 542) = 4.60, p > .05 \) (GFI = .99, AGFI = .98, rmrs = .03). The coefficient of reproducibility was .91, and the coefficient of scalability was .60.

For Sample 3, we analyzed the Time 2 and Time 3 data separately. At Time 2, the model provided an excellent fit to the data, \( \chi^2(6, N = 413) = 8.10, p > .05 \) (GFI = .99, AGFI = .98, rmrs = .03). The Guttman scale manifested a coefficient of reproducibility of .91 and a coefficient of scalability of .62. A similar pattern of results emerged at Time 3. Again the model provided an excellent fit to the data, \( \chi^2(6, N = 313) = 2.34, p > .05 \) (GFI = 1.00; AGFI = .99; rmrs = .02); the scale obtained a coefficient of reproducibility of .91 and a coefficient of scalability of .63.

Test–retest reliability: For Sample 3, union participation was stable across 6 months. The correlation between participation at Time 2 and participation at Time 3 was .70, \( p < .01 \).

Construct validity: Adherence to the Guttman scale model provided initial evidence for the construct validity of the scale; the data support the existence of the hypothesized structure. Additional information about construct validity emerged from the correlations between the measure and external criteria.

Table 3 presents the correlations between the participation-measure criterion variables for Samples 1 and 2. As hypothesized, participation in union activities was significantly \( p < .01 \) correlated with measures of union loyalty, willingness to work for the union, responsibility to the union, and union satisfaction. The participation measure also was correlated significantly \( p < .05 \) with union tenure but was not significantly related to either age or education.

Table 4 presents the correlations between the relevant criteria and the participation measure for Sample 3 at both time periods. At both time periods, participation in union activities correlated significantly and positively \( p < .01 \) with the three dimensions of union commitment: union loyalty, willingness to work for the union, and responsibility to the union. Members' participation also was significantly related to union attitudes and instrumentality perceptions at Time 2. Contrary to our hypotheses, members' participation was not correlated with instrumentality at Time 3. Union tenure correlated \( p < .05 \) with participation in union activities, as did respondents' age. As hypothesized, level of education was not significantly correlated with participation in union activities. Both extrinsic and intrinsic job satisfaction correlated significantly with members' participation in local union activities.

Discussion

The results of this first study strongly suggest that items measuring members' participation in union activities are ordered along a single and cumulative dimension. Across three large and independent samples, a Guttman scale provided an accurate depiction of the data; in each sample the coefficient of reproducibility exceeded .90 and the coefficient of scalability exceeded .60. Taken together, these findings suggest that the Guttman scale provided a good fit to the data and that the results cannot be attributed to capitalization on chance within a single research setting. Replication of the Guttman scale across different unions and different measures of participation suggests that these findings may be taken as robust.

We interpret these findings as supporting the unidimensional and cumulative nature of members' participation in local union activities. We recognize, however, that alternate hypotheses are also plausible. As Nunnally (1978) pointed out, Guttman scalability is a nonsufficient condition for an inference of unidimensionality. In particular, by selecting items that vary greatly in difficulty and administering them to a diverse population, the simplex structure characteristic of a Guttman scale can be reproduced. Thus, our interpretation of unidimensiona-
Table 3
Criterion Correlations for Sample 1 (n = 229) and Sample 2 (n = 541)

<table>
<thead>
<tr>
<th>Criterion</th>
<th>1</th>
<th>2</th>
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<th>7</th>
<th>8</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. Participation in local union activities</td>
<td>.49</td>
<td>.55</td>
<td>.38</td>
<td>.28</td>
<td>.17</td>
<td>.12</td>
<td>.08</td>
<td></td>
</tr>
<tr>
<td>2. Union loyalty</td>
<td>.45</td>
<td>.54</td>
<td>.62</td>
<td>.72</td>
<td>.01</td>
<td>.18</td>
<td>.16</td>
<td></td>
</tr>
<tr>
<td>3. Willingness to work for the union</td>
<td>.63</td>
<td>.54</td>
<td>.48</td>
<td>.26</td>
<td>.01</td>
<td>.11</td>
<td>.12</td>
<td></td>
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<tr>
<td>4. Responsibility to the union</td>
<td>.36</td>
<td>.49</td>
<td>.46</td>
<td>.42</td>
<td>.01</td>
<td>.17</td>
<td>.06</td>
<td></td>
</tr>
<tr>
<td>5. Union satisfaction</td>
<td>.24</td>
<td>.62</td>
<td>.36</td>
<td>.29</td>
<td>.02</td>
<td>.11</td>
<td>.01</td>
<td></td>
</tr>
<tr>
<td>6. Union tenure</td>
<td>.10</td>
<td>.05</td>
<td>.10</td>
<td>.07</td>
<td>.08</td>
<td>.09</td>
<td>.06</td>
<td></td>
</tr>
<tr>
<td>7. Age</td>
<td>-.05</td>
<td>.07</td>
<td>.06</td>
<td>.03</td>
<td>.06</td>
<td>-.61</td>
<td>-.01</td>
<td></td>
</tr>
<tr>
<td>8. Education</td>
<td>.09</td>
<td>-.03</td>
<td>.12</td>
<td>.06</td>
<td>.01</td>
<td>-.15</td>
<td>.15</td>
<td></td>
</tr>
</tbody>
</table>

Note. Correlations based on Sample 1 are presented above the diagonal, and correlations based on Sample 2 are presented below the diagonal. For Sample 1, r > .13, p < .05; r > .18, p < .01. For Sample 2, r > .10, p < .05; r > .15, p < .01.

Liability is based on both the empirical results of our analyses and prior conceptual development that posits a unidimensional ordering of participation measures (Nicholson, 1978). Although our data are consistent with this interpretation, alternative models may provide an equivalent or better fit to the data.

The findings of this first study also support the reliability and construct validity of the Guttman scale of members' participation in union activities. First, consistent with Klandermans' (1986) analysis, participation emerged as a highly stable construct over time. For Sample 3, the correlation of participation scores across 6 months was .70. Second, with the exception of some demographic variables, the Guttman scale of participation correlated with external criteria in a predictable fashion. The scale was positively correlated with measures of union loyalty, willingness to work for the union, responsibility to the union, satisfaction with the union, union instrumentality, union attitudes, and union tenure. Consistent with our hypotheses, extrinsic and intrinsic job satisfaction were negatively correlated with members' participation in union activities. Finally, as predicted, education was not significantly associated with participation in union activities. Respondents' age was not associated with participation in two of the three samples, although it was significantly correlated with participation for members of Sample 3.

In summary, the results of the current study offer support for the suggestion that members' participation in union activities can be appropriately conceptualized and measured as a unidimensional and cumulative scale.

Study 2

The primary goal of the second study was to develop and evaluate a predictive model of members' participation in local union activities. There are at least three theoretical perspectives from which potential predictors of members' participation in union activities may be derived. First, research has consistently identified union commitment as a correlate of (e.g., Fullagar,
1986; Gordon et al., 1980; Kelloway et al., 1992), and a precursor to (Fullagar & Barling, 1989; Kryl, 1990), members' participation. Second, the theory of reasoned action (Fishbein & Ajzen, 1975) has been applied to the prediction of union commitment (Fullagar, McCoy, & Shull, 1992) and union voting intentions (Fullagar & Howland, 1990; Montgomery, 1989) and can be plausibly extended to the prediction of members' participation in union activities. Third, consistent with the partisan model of political participation (Fishbein & Ajzen, 1981), research on both union commitment (e.g., Fullagar & Barling, 1989; Fullagar et al., 1992; Gordon et al., 1980) and union participation (e.g., Anderson, 1979; Blyton et al., 1981; Kelloway & Catano, 1988; Spinrad, 1960) has suggested the importance of socialization influences, such as those that occur in the first year in the union (Fullagar & Barling, 1989; Fullagar et al., 1992; Gordon et al., 1980), and perceptions of union leadership styles (Fullagar et al., 1992; Kahn & Tannenbaum, 1954).

In Study 2, we incorporated specific suggestions from each of these three theoretical perspectives in a predictive model of union participation. Specifically, we hypothesized that, although each of these theoretical propositions provides an adequate prediction of members' participation in the union, none of them provides a sufficient description. Rather, each of these views contributes to the prediction of members' participation in union activities. The proposed model, incorporating relations identified from each theoretical position, is presented in Figure 1. Specific rationales for the paths that make up the model are presented in the following paragraphs.

Union Commitment, Union Participation, and Job Satisfaction

The relationships between union commitment, union participation, and job dissatisfaction have been the focus of empirical enquiry for some time. Traditionally, job dissatisfaction is posited as a precursor to union commitment, which in turn is used to predict members' participation in union activities (e.g., Fullagar & Barling, 1987, 1989). In contrast, the model proposed here incorporates elements of a conflict model (Kryl, 1990) of members' participation in union activities. Kryl contrasted six competing models in her longitudinal study of members' participation in union activities. Although her results are not unambiguous, strongest support emerged for the conflict model that posited members' participation as an outcome of union commitment and a predictor of job dissatisfaction. Specifically, Kryl suggested that the conflict inherent in the practice of industrial relations (Bluen & Barling, 1988) implies that active unionists (i.e., those who are likely to experience this conflict most directly) will experience increased job dissatisfaction as a result of their involvement. The conflict model is also consistent with the voice effect of unionization, whereby union members experience more dissatisfaction with their jobs than do comparable nonunionized workers (e.g., Freeman & Medoff, 1984). Freeman and Medoff made a distinction between true and voiced dissatisfaction and suggested that union members express more dissatisfaction because of the politicization of the workforce in union activities. Logically, the primary vehicle for this politicization is through participation in union activities. As union members attend union meetings, read union literature, and become more active in union leadership, they also become more aware of the costs of their jobs.

We extended Kryl's (1990) formulation of the conflict model by incorporating the causal structure of union commitment as identified by Tetrick (1989) and by examining potential moderators of the posited relationships. Tetrick proposed a specific causal ordering of commitment dimensions. Specifically, union loyalty predicts responsibility to the union, and willing-

![Figure 1. Proposed overidentified model of union participation.](image-url)
ness to work for the union is predicted by both responsibility to the union and union loyalty (Tetrick, 1989). Thacker, Fields, and Barclay (1990) also suggested that responsibility to the union and willingness to work for the union are consequences of union loyalty. Thus, in our model union loyalty was proposed as an antecedent of the willingness to work for the union, and responsibility to the union was posited as a consequence of union loyalty and an antecedent of willingness to work for the union. Within this framework, only willingness to work for the union was posited as exerting a direct effect on members' participation in local union activities.

However, we also predicted the existence of some barriers to participation (Chaissou & Andiappan, 1989; Wertheimer & Nelson, 1975), which we expected to moderate this relationship. Specifically, role constraints (i.e., interrole conflict and shiftnwork) were expected to inhibit the translation of attitudes (e.g., union loyalty) or behavioral intentions (e.g., willingness to work for the union) into actual participation in the union. Because much union participation must take place after regular working hours, family responsibilities can restrict the extent to which some individuals are able to participate in the union. Interrole conflict reflects the feeling of being caught between two competing role demands (Barling, 1990). Being involved in regular shift work might rule out members' being unable to attend union meetings or participate in other union activities because of conflicting role demands. Our proposed model suggested that both employment-related commitments (working shifts) and conflict with other role commitments (interrole conflict) would moderate the relationship between union commitment and members' participation.

**Predicting Union Commitment—Two Complementary Specifications**

Previous research provides little information on the predictors of union commitment. In the current study, we extended the model by identifying predictors of union commitment and members' participation derived from the theory of reasoned action (Fishbein & Ajzen, 1975) and the partisan model of political participation (Fishbein & Ajzen, 1981).

A theory of reasoned action. Fishbein and Ajzen's (1975) theory of reasoned action has been applied to many contexts involving the prediction of behavior by attitudinal variables. Briefly, their theory posits that the best predictor of behavior is the intention to perform the behavior. Behavioral intentions, in turn, are predicted by both subjective norms and attitudes. Finally, beliefs about the behavior are posited as exerting a direct effect on attitudes and indirect effects on behavioral intentions and behavior. In our study, willingness to work for the union was posited as reflecting behavioral intentions to participate in the union, and union loyalty was posited as reflecting attitudes toward the union.

Fishbein and Ajzen's (1975) formulation has been applied previously to the study of union-relevant criteria. Montgomery (1989) applied Fishbein and Ajzen's theory to union voting intent in union certification elections and found general support for the model. Intentions to vote for union representation were predicted by union attitudes, beliefs, and normative pressures and were in turn predictive of actual voting behavior (Montgomery, 1989). Fullagar et al. (1992) incorporated elements of the theory of reasoned action in their study of the development of union commitment. On the basis of Fullagar et al.'s (1992) findings, we proposed one addition to the theory of reasoned action. Fishbein and Ajzen posited subjective norms as a predictor of behavioral intentions but not of attitudes. Fullagar et al. (1992), however, reported that subjective norms exerted direct effects on union attitudes and union loyalty. Therefore, we posited a direct relationship between subjective norms and union loyalty.

We hypothesized that the belief that participation in the union is instrumental to the attainment of individual needs would predict union loyalty directly. This is consistent with Fishbein and Ajzen's (1975) model in that beliefs are posited as exerting a direct effect on attitudes and an indirect effect on behavioral intentions. Perline and Lorenz (1970) identified the satisfaction of individual needs as the payoff for participation in union activities. Similarly, Nicholson et al. (1980) identified the perceived instrumentality of union participation as a strong predictor for union members' engagement in union activities.

In addition, on the basis of findings regarding the economic instrumentality of the union (Fullagar & Barling, 1989), we suggested that the effect of such instrumentality beliefs on members' participation would be mediated by union attitudes and behavioral intentions to participate in the union.

Thus, in the proposed model members' participation was hypothesized to be an outcome of behavioral intentions (willingness to work for the union). In turn, willingness to work for the union was posited to be an outcome of subjective norms (Fishbein & Ajzen, 1975), union loyalty (Fishbein & Ajzen, 1975; Tetrick, 1989), and responsibility to the union (Tetrick, 1989).

Both the perceived instrumentality of participation in union activities (Fishbein & Ajzen, 1975) and subjective norms (Fullagar et al., 1992) were posited as predictors of union loyalty. Finally, both intrinsic and extrinsic job dissatisfaction were posited as outcomes of participation in union activities (Kryl, 1990).

The partisan model. In reviewing various models of voting in American elections, Fishbein and Ajzen (1981) identified a model that focuses on psychological predictors as determinants of voting behavior, namely, the role of partisanship in predicting voting behavior. Three general partisan motivators have been identified: personal attachment to a political party (party identification), individual orientation to campaign issues (issue partisanship), and individual attitudes toward political candidates (candidate partisanship).

The identification of partisan influences on voting intentions is similar to Spinrad's (1960) comments regarding the importance of the socialization process in fostering active participation in union activities. Spinrad suggested that any factor that increased the identification of the individual with his or her work group would also increase his or her participation in union activities. Presumably, socialization practices foster greater attachment to the union and, in turn, active participation in union activities. From the literature on members' participation in the union and commitment to the union, several such socializing influences can be identified that parallel Fishbein and Ajzen's (1981) partisan model of political participation. In
the current study, three socialization processes were identified and posited to affect the development of union loyalty.

First, the Marxist work belief is very close to the raison d’être of labor unions (Fullagar & Barling, 1987) and, in this sense, corresponds to Fishbein and Ajzen’s (1981) concept of issue partisanship. Studies of preemployed individuals (Barling, Kel- loway, & Agar, 1991; Barling, Kelloway, & Bremermann, 1991) have suggested that the Marxist work belief develops prior to entry into the workforce and predicts union attitudes. Early research into union participation suggested that Marxist be-

liiefs were characteristic of union activists (Buchholz, 1978, 1979; Spinrad, 1960) and were associated with union loyalty for Black respondents in South Africa (Fullagar & Barling, 1989) and with pronoun attitudes in Canada and the United States (Barling, Laliberté, Fullagar, & Kelloway, 1993). Perhaps more important from a socialization perspective, Marxist work be-

liiefs have predicted union attitudes and willingness to join a union in individuals not yet employed on a full-time basis (Bar-

ling, Kelloway, & Agar, 1991; Barling, Kelloway, & Bremer-

mann, 1991). In our proposed model, Marxist work beliefs were posited as exerting a direct effect on an individual’s loyalty to the union.

Paralleling Fishbein and Ajzen’s (1981) description of candidate partisanship, we focused on members’ evaluations of their shop stewards’ leadership styles. Shop stewards have been consistently identified as playing a vital role in building and maintaining unions (Barling et al., 1992; McCarthy, 1967). As the representative of the union on the shop floor, the shop steward has been referred to as “the rank-and-file leader in the shop” (Peck, 1963, p. 15) and the “transmission belt between union officers and the membership” (Nash, 1984, p. 9). Even earlier, Kahn and Tannenbaum (1954) suggested that members’ participation in the union is directly associated with stewards’ interpersonal skills. Nicholson et al. (1981) concluded that steward performance is the link that translates individual motivation into actual participation. Arguably then, perceptions of shop stewards’ leadership styles may serve as a proxy for candidate partisanship. Supporting this argument, Gallagher and Clark (1989) emphasized members’ perceptions of the availability and the skill of shop stewards as important in fostering positive union attitudes and active involvement in the union.

We posited that members’ perceptions of the transformational leadership qualities of their shop stewards would have a direct effect on union loyalty (Fullagar et al., 1992). Moreover, following Kahn and Tannenbaum (1954), we posited a direct link between shop stewards’ leadership qualities and participation in union activities. Though stewards facilitate satisfaction with the union through effective performance of their duties (Barling et al., 1992), they also encourage membership support and involvement through their role as socialization agents (Fullagar et al., 1992; Gallagher & Clark, 1989). Bass (1985) defined transformational leadership as going beyond the reward-based transactional leadership style to include the ability to inspire a sense of pride in the organization and a sense of mission, the ability to provide followers with intellectual stimulation, and the encouragement of individuals.

The socialization practices of unions are explicitly designed to foster identification with the union, and therefore they correspond to Fishbein and Ajzen’s (1981) concept of party identification. Socialization experiences during the first year of union membership correlate consistently with union loyalty (Fullagar & Barling, 1989), other dimensions of union commit-

ment (Gordon et al., 1980), and general union attitudes (Fulla-

gar et al., 1992). Social integration in the union has been consis-

tently related to greater participation in union activities (e.g., Anderson, 1979; Dean, 1954; Kelloway & Catano, 1988; Kol-

chin & Hyclak, 1984). The effect of these early socialization experiences was posited as acting directly on the formation of union loyalty and indirectly on participation in the union.

Thus, the primary goal of this second study was to evaluate the proposed model presented in Figure 1. To do so, we assessed the specific predictions and overall fit of the model to the data in more than one sample. Specific unions can vary widely in the degree to which members’ involvement is encouraged or desired; thus, failure to replicate results across samples could result from differences in union structure or constitution rather than from any instability of the proposed relationships. Accord-

ingly, we initially developed and validated the model with data from a sample of union members and replicated the results with data from two other locals of the same national union. This strategy provided for the assessment of the model in diverse populations while controlling for differences in union struc-


Method

Subjects. Participants in the study were members of three locals of a large union of government employees in Ontario. Members of Local 1 were clerical and maintenance employees. Members of Locals 2 and 3 were guards and rehabilitation staff employed in a correctional facility. A total of 1,112 surveys were distributed, with 53 being returned as undeliverable. Three hundred and seventy-seven completed surveys were returned for a usable response rate of 35.6%. Descriptive information for respondents in each of the three locals as well as the sample as a whole is presented in Table 5. Members in Local 1 were somewhat older, F(2, 356) = 11.54, p < .0001, and had less union tenure, F(2, 365) = 3.23, p < .05, and less education, F(2, 366) = 5.85, p < .01, than members of either Local 2 or Local 3. These differences were anticipated as the employer of Local 1’s members has a policy of hiring retired military service people.

Measures. All respondents completed a checklist of seven participation items, one of which was deleted from the analysis. During the data collection process, the union called for a strike vote. As the ques-

Table 5

<table>
<thead>
<tr>
<th>Variable</th>
<th>Local 1</th>
<th>Local 2</th>
<th>Local 3</th>
<th>Full sample</th>
</tr>
</thead>
<tbody>
<tr>
<td>Age (in years)</td>
<td>48.50</td>
<td>42.73</td>
<td>43.70</td>
<td>46.27</td>
</tr>
<tr>
<td>Education (in years)</td>
<td>11.85</td>
<td>12.59</td>
<td>12.71</td>
<td>12.19</td>
</tr>
<tr>
<td>Union tenure (in years)</td>
<td>10.78</td>
<td>11.53</td>
<td>12.93</td>
<td>11.42</td>
</tr>
<tr>
<td>Male:female ratio</td>
<td>1.36:1</td>
<td>3.14:1</td>
<td>3.68:1</td>
<td>3.94:1</td>
</tr>
<tr>
<td>No. of surveys returned</td>
<td>220</td>
<td>63</td>
<td>94</td>
<td>377</td>
</tr>
<tr>
<td>Response rate (%)</td>
<td>38.0</td>
<td>30.0</td>
<td>37.0</td>
<td>35.5</td>
</tr>
</tbody>
</table>
tionnaire returns took place over a 4-week period, some members may have returned the questionnaire prior to the vote, which may have resulted inambiguous responses to the item "participated in other union votes." Analysis of scale responses supported this possibility. Despite the consistent ordering of responses for the two voting items in Study 1, the proposed Guttman scale model did not fit the data for the current sample. Analysis of item responses indicated that the largest source of error in the scale was attributable to the item referring to participating in other union votes. Accordingly, this item was excluded from further analysis. The remaining six scale items fit the simplex model, \( x^2(6, N = 377) = 6.51, p > .05 \) (GFI = .99; AGFI = .98; rmr = .03). Accordingly, a Guttman scale of participation was constructed ranging from 0 (no participation) to 6 (holding union office). The coefficient of reproducibility for the scale was .90, and the coefficient of scalability was .61.

Respondents completed a 13-item checklist of first-year socialization experiences (\( z = .82 \)) based on that used by Fullagar and Barling (1989). Respondents indicated whether they had experienced each act-
vity (e.g., a 5-point scale (yes = 2, not sure = 1, and no = 0).

The perceived instrumentality of participation in union activities was measured with seven items, each scored along a 5-point scale ranging from strongly disagree (1) to strongly agree (4). Items were based on Nicholson et al.'s (1981) Need for Involvement scale. Coefficient alpha for this scale was .82.

On the basis of results reported by Kelloway et al. (1992), respondents completed a 15-item measure of union commitment. The 15 items measured union loyalty (7 items; \( \alpha = .86 \)), responsibility to the union (4 items; \( \alpha = .80 \)) and willingness to work for the union (4 items; \( \alpha = .86 \)).

In operationalizing subjective norms, we followed Fishbein and Aj-
zen's (1975) definition, which was previously used by Fullagar et al. (1992). Respondents were asked to rate the extent to which three refer-
ence groups (family, co-workers, and people important to the respond-
ent) supported their active participation in the union. Normative compliance was assessed with three items asking the extent to which individuals wanted to comply with the wishes of the three reference groups. An overall subjective norms scale was calculated by multiplying each normative item by the normative compliance item for each group and summing across the three products. The scale demonstrated adequate internal consistency (\( \alpha = .82 \)).

A short form (15 items) of Bass's (1985) Multifactor Leadership Ques-
tionnaire was developed for the current study and measured percep-
tions of the leadership qualities of shop stewards. Although Bass pro-
posed a three-dimensional interpretation of the scale (Intellectual Stimulation, Charisma, and Individual Consideration), the high inter-
correlations (\( r > .80 \)) among these dimensions in the current study indicated the advisability of interpreting the scale as a single dimen-
sion. Coefficient alpha for the 15-item scale was .97.

Marxist work beliefs were measured with Buchholz's (1978, 1979) 10-item scale (\( \alpha = .75 \)). The validity of the scale has been established (e.g., Barling, Lalive, Fullagar, & Kelloway, 1993; Barling, Kelloway, & Bremermann, 1991).

Interrole conflict was assessed with Kopelman, Greenhaus, and Connolly's (1983) eight-item scale. A recent multitrait-multimethod analysis supports the construct validity of the scale (Barling & Mc-
Ewen, 1988), which demonstrated adequate internal consistency in the current study (\( \alpha = .89 \)).

Warr et al.'s (1979) measures of extrinsic (eight items; \( \alpha = .78 \)) and intrinsic (seven items; \( \alpha = .90 \)) job satisfaction were also completed. Extensive psychometric information is available for this scale (e.g., Cook, Hepworth, Wall, & Clegg, 1981).

Procedure. All members of the three locals whose names appeared on the union's membership list were sent copies of the survey package. The package included an introductory letter from the researcher, a letter of support signed by the presidents of the three locals (indicating the support of the local, regional, and national levels of the union), the questionnaires, and a prepaid return envelope. Approximately one week after the distribution of the surveys, reminder cards were sent to all members of the three locals. Union members were further encour-
gaged to participate in the survey by local union officials, who reiter-
ated the union's support for the survey at union meetings and individu-
ally encouraged members to participate.

Data analysis. Prior to evaluating the structural models presented earlier, we divided the sample into two subsamples: a study sample and a replication sample. The study sample comprised all respondents from Local 1 who provided complete responses to all of the study variables (\( n = 202 \)). The replication sample was formed by combining the responses of respondents from Locals 2 (\( n = 58 \)) and 3 (\( n = 89 \)). Although the division was initially based on the similarity of occupa-
tions for Samples 2 and 3, tests for the equality of covariance matrices (Jöreskog & Sörbom, 1984) between samples supported the validity of this procedure. There were no significant differences between the co-
variance matrices derived from Local 2 and Local 3 respondents, \( x^2(66, N = 147) = 81.15, p > .05 \), indicating that data from these two samples could be combined without introducing sample-specific bias into the analysis. However, there were significant differences between the covariance matrices derived from the study and the replication samples, \( x^2(66, N = 347) = 126.02, p < .001 \), indicating the advisability of analyzing these samples separately and enhancing any inferences about generalizability if the model was supported in both samples.

We first used multiple regression to evaluate the proposed modera-
tor effects. Although it is possible to examine moderator or interaction effects using LISREL (Kenny & Judd, 1984), the procedure is awkward and complicated by the inability to define a variable as the product of two other variables in the LISREL environment (Hayduk, 1987).

We also compared the magnitude of the estimated parameters across samples. To do so, we followed the procedure described by Jöreskog and Sörbom (1984) and elaborated on by Hayduk (1987) for testing equivalence hypotheses. First, using LISREL's capacity for mul-
tisample analyses, we estimated the proposed model simultaneously in both samples on the basis of the covariance matrices. Next we re-
estimated the model, constraining all the directional estimates for the replication sample to be equal to the corresponding parameters esti-
mated for the study sample. As the two models (freely estimated vs. equality constraints) are nested, a chi-square difference test (Long, 1983) was used to test the hypothesis that all parameter estimates were invariant across the two samples. Given a significant omnibus test, a series of tests of specific parameters was used to identify those that differed between the two samples (Hayduk, 1987).

Results

Descriptive statistics and intercorrelations for all study vari-
ables in both samples are presented in Table 6.

Moderators. Prior to the evaluation of model fit, the spe-
cific moderated relationships posited in the model were evalu-
at the multiple regression analyses. None of the proposed moderator effects were statistically significant (\( p > .05 \)). Nei-
ther interrole conflict nor shiftwork nor the product of these variables with willingness to work for the union made a signif-
ificant contribution to the prediction of union participation. Nei-
ther shiftwork nor interrole conflict nor the product of either of these two variables with union loyalty made a significant con-
tribution to the prediction of willingness to work for the union.

Therefore, all moderated relationships were deleted from the
model, and only the direct and mediated relations proposed earlier were included in subsequent analyses.

**Prediction.** For the study sample, the proposed model provided an acceptable, but not outstanding fit to the data, \( \chi^2(31, N = 202) = 69.41, p < .001 (\chi^2/df = 2.24, \text{GFI} = .95, \text{AGFI} = .89) \), normed fit index (NFI) = .89, Tucker–Lewis index (TLI) = .91). Inspection of the model parameters (see Figure 2) suggested that 11 of the 13 hypothesized relations were significant in the expected direction. Participation in union activities was predicted by willingness to work for the union (B = .34, Z = 5.29, \( p < .01 \)) and perceptions of shop stewards’ transformational leadership styles (B = .32, Z = 4.90, \( p < .01 \)). In turn, participation in union activities predicted both intrinsic (B = −.25, Z = −3.67, \( p < .01 \)) and extrinsic (B = −.29, Z = −4.17, \( p < .01 \)) job satisfaction. As hypothesized, willingness to work for the union was predicted by union loyalty (B = .47, Z = 7.21, \( p < .01 \)), responsibility to the union (B = .13, Z = 2.19, \( p < .01 \)), and perceived subjective norms (B = −.26, Z = 4.68, \( p < .01 \)). Responsibility to the union was also predicted by union loyalty (B = .56, Z = 9.64, \( p < .01 \)). In turn, union loyalty was predicted by subjective norms (B = .20, Z = 3.43, \( p < .01 \)), socialization experiences in the first year of union membership (B = −1.9, Z = −3.05, \( p < .01 \)), and perceptions of the instrumentality of union activism (B = .42, Z = 6.94, \( p < .01 \)). Contrary to our predictions, union loyalty was predicted by neither Marxist work beliefs (B = −.09, Z = 1.56, \( p > .05 \)) nor perceptions of shop stewards’ transformational leadership styles (B = .09, Z = 1.39, \( p > .05 \)).

**Model modification.** Although these results indicate the acceptability of the proposed model, it did not provide a good absolute fit to the data, which suggested that there were some misspecification in the model. Accordingly, we improved the fit of the model by theory trimming (Pedhazur, 1982) in an exploratory–confirmatory approach. First, using data from the study sample, we calculated the just-identified model (in which all possible direct effects are calculated) and then trimmed all paths that did not meet at least two of Pedhazur’s (1982) criteria: statistical significance (\( p < .05 \)) and meaningfulness and magnitude (\( B > .05 \)). Because theory trimming is an optimizing technique involving considerable capitalization on chance (Cliff, 1983), we then evaluated the stability of the resulting parameters by estimating the resultant model with data from the replication sample.

Theory trimming resulted in the identification of a better fitting model that retained all paths described earlier for the study sample. Although no paths were deleted from the original model, four paths were added (i.e., Marxist work beliefs predicted intrinsic job satisfaction, extrinsic job satisfaction, and responsibility to the union, and perceptions of shop stewards’ transformational leadership characteristics predicted responsibility to the union). Standardized parameter estimates for the revised model are presented in Figure 3. Again, participation in union activities predicted both intrinsic (B = −.22, Z = −3.25, \( p < .01 \)) and extrinsic (B = −.25, Z = −3.74, \( p < .01 \)) job satisfaction. Marxist work beliefs also predicted intrinsic (B = −.25, Z = −3.69, \( p < .01 \)) and extrinsic (B = −.27, Z = −4.01, \( p < .01 \)) job satisfaction. Participation in union activities was predicted by willingness to work for the union (B = .34, Z = 5.26, \( p < .01 \))

### Table 6

<table>
<thead>
<tr>
<th>Variable</th>
<th>Study sample M (SD)</th>
<th>Replication sample M (SD)</th>
</tr>
</thead>
<tbody>
<tr>
<td>1. Participation</td>
<td>3.68 (1.24)</td>
<td>3.68 (1.24)</td>
</tr>
<tr>
<td>2. Willingness to work for the union</td>
<td>11.94 (4.08)</td>
<td>11.94 (4.08)</td>
</tr>
<tr>
<td>3. Responsibility to the union</td>
<td>15.09 (5.74)</td>
<td>15.09 (5.74)</td>
</tr>
<tr>
<td>4. Union loyalty</td>
<td>24.97 (8.82)</td>
<td>24.97 (8.82)</td>
</tr>
<tr>
<td>5. Extrac. job satisfaction</td>
<td>37.41 (5.83)</td>
<td>37.41 (5.83)</td>
</tr>
<tr>
<td>6. Subjective norms</td>
<td>21.54 (11.66)</td>
<td>21.54 (11.66)</td>
</tr>
<tr>
<td>7. Leadership</td>
<td>43.08 (6.26)</td>
<td>43.08 (6.26)</td>
</tr>
<tr>
<td>8. Replication</td>
<td>25.20 (5.59)</td>
<td>25.20 (5.59)</td>
</tr>
<tr>
<td>9. Perceived instrumentality</td>
<td>20.71 (5.69)</td>
<td>20.71 (5.69)</td>
</tr>
</tbody>
</table>

Note: Data for the study sample are presented below the diagonal, and data for replication sample are presented above the diagonal.
and perceptions of shop stewards' transformational leadership characteristics (B = .32, Z = 4.86, p < .01). Willingness to work for the union was predicted by responsibility to the union (B = .13, Z = 2.18, p < .01), union loyalty (B = .47, Z = 7.35, p < .01), and subjective norms (B = .26, Z = 4.65, p < .01). Responsibility to the union was predicted by union loyalty (B = .44, Z = 6.97, p < .01), Marxist work beliefs (B = .21, Z = 3.64, p < .01), and perceptions of shop stewards' transformational leadership.
styles (B = .15, Z = 2.37, p < .01). Finally, as in the original model, union loyalty was predicted by subjective norms (B = .20, Z = 3.43, p < .01), socialization experiences in the first year of union membership (B = .19, Z = 3.05, p < .01), and perceptions of the instrumentality of union activism (B = .42, Z = 6.94, p < .01). Consistent with Pedhazur's (1982) criteria for theory trimming, nonsignificant but meaningful paths were retained between union loyalty and Marxist work beliefs (B = .09, Z = 1.56, p > .05) and between union loyalty and perceptions of shop stewards' transformational leadership styles (B = .09, Z = 1.39, p > .05).

The revised model provided an excellent fit to the data, χ²(27, N = 202) = 29.99, p > .05 (χ²/df = 1.11, GFI = .98, AGFI = .94, NFI = .95, TLI = .99). Because the revised model involved the addition of paths with no deletions from the original model, the original model is nested within the revised model. The chi-square test for the difference between two nested models (Long, 1983) indicated that the revised model provided a significantly better fit to the data than did the original model, χ²(4, N = 202) = 39.42, p < .001.

Replication. To estimate the external validity of the paths in the revised model, we evaluated it again, using the data from the replication sample (n = 147). For the replication sample, the revised model again provided an excellent fit to the data, χ²(27, N = 147) = 39.15, p > .05 (χ²/df = 1.45, GFI = .96, AGFI = .91, NFI = .94, TLI = .97). Moreover, the revised model provided a significantly better fit for the replication sample than did the original model for this sample, χ²(4, N = 147) = 39.96, p < .01.

The standardized parameter estimates for this model are presented in Figure 4. Of the 17 directional relationships represented in the revised model, 15 emerged as significant for the replication sample. As for the study sample, both intrinsic and extrinsic job satisfaction were predicted by participation in union activities (for intrinsic satisfaction, B = -.32, Z = -4.11, p < .01; for extrinsic satisfaction, B = -.45, Z = -6.45, p < .01) and Marxist work beliefs (for intrinsic satisfaction, B = -.18, Z = -2.29, p < .05; for extrinsic satisfaction, B = -.30, Z = -4.25, p < .01). Participation in union activities was predicted by both willingness to work for the union (B = .28, Z = 3.11, p < .01) and perceptions of shop stewards' transformational leadership characteristics (B = .24, Z = 2.63, p < .01). Willingness to work for the union was predicted by union loyalty (B = .55, Z = 7.66, p < .01) and responsibility to the union (B = .27, Z = 4.29, p < .01) but not by subjective norms (B = .08, Z = 1.39, p > .05). Responsibility to the union was predicted by union loyalty (B = .39, Z = 4.14, p < .01) and Marxist work beliefs (B = .32, Z = 4.57, p < .01) but not by perceptions of shop stewards' transformational leadership styles (B = .12, Z = 1.34, p > .05). Finally, union loyalty was predicted by subjective norms (B = .24, Z = 4.28, p < .01), Marxist work beliefs (B = .15, Z = 2.81, p < .01), perceptions of shop stewards' transformational leadership characteristics (B = .43, Z = 6.94, p < .01), perceptions of the instrumentality of union activism (B = .20, Z = 3.16, p < .01), and socialization experiences in the first year of union membership (B = .10, Z = 1.79, p < .05).

Cross-sample comparisons. The results of the omnibus chi-square test indicated a significant difference between the two samples, χ²(17, N = 349) = 52.04, p < .01. Five of the 17 directional parameters were significantly different between the two samples.

Figure 4. Standardized parameter estimates for the revised model (replication sample, n = 147).
The path from participation in union activities to intrinsic job satisfaction was larger for the replication sample ($\beta = -2.81$) than for the study sample ($\beta = -1.22$), $\chi^2(1, N = 349) = 4.17, p < .05$, as was the path from participation to extrinsic job satisfaction (for the study sample, $\beta = -1.30$; for the replication sample, $\beta = -3.97$), $\chi^2(1, N = 349) = 8.97, p < .01$. The prediction of willingness to work for the union by subjective norms was stronger for the study sample ($\beta = .09$) than for the replication sample ($\beta = .03$), $\chi^2(1, N = 349) = 4.81, p < .05$. Conversely, the prediction of union loyalty by perceptions of shop stewards' transformational leadership styles was stronger for the replication sample ($\beta = .03$) than for the study sample ($\beta = .02$), $\chi^2(1, N = 349) = 14.97, p < .01$. Finally, the prediction of union loyalty by perceptions of union instrumentality was stronger for the study sample ($\beta = .42$) than the replication sample ($\beta = .21$), $\chi^2(1, N = 349) = 5.76, p < .05$.

**Discussion**

Our primary goal in Study 2 was to develop and evaluate a structural model predicting members' participation in local union activities. The results of this study offer considerable support for the proposed model. As originally proposed, the model based on the theory of reasoned action (Fishein & Ajzen, 1975) and the partisan model of political participation (Fishein & Ajzen, 1981) offered an acceptable fit to the data. After model modification, a better fitting model was obtained and was substantially replicated in an independent sample. Although some effects in the model varied across the two samples, the overall model provided an excellent fit to the data in both samples. Replication of the results across two independent samples suggests that the results of the current study are robust.

As originally hypothesized, participation in union activities predicted both extrinsic and intrinsic job dissatisfaction, with these effects being stronger for the replication sample than for the study sample. Intrinsic and extrinsic job dissatisfaction were also predicted by Marxist work beliefs in both samples. Participation in union activities was predicted by both willingness to work for the union and perceptions of shop stewards' transformational leadership styles in both samples. Willingness to work for the union was predicted by union loyalty, responsibility to the union, and subjective norms, with the prediction of willingness to work for the union by subjective norms being stronger in the study sample than in the replication sample. Responsibility to the union was predicted by union loyalty, Marxist work beliefs, and perceptions of shop stewards' transformational leadership styles. Finally, union loyalty was predicted by Marxist work beliefs, subjective norms, first-year socialization experiences, perceptions of shop stewards' transformational leadership characteristics, and perceptions of the instrumentality of union activism.

Interestingly, the prediction of union loyalty by shop stewards' transformational leadership styles was stronger for the replication sample, whereas the prediction of union loyalty by instrumentality perceptions was a stronger predictor for the study sample. This finding may emerge from the generally low evaluation of shop stewards expressed by the study sample ($M = 30.40$) compared with the replication sample ($M = 43.08$). Indeed, a substantial number of respondents in the study sample indicated that they had very little contact with their union representatives. Overall, these results suggest that although the same process predicted participation across two samples, the strength of individual relations varied across the samples and may be related to sample-specific characteristics.

The current research both replicates and extends previous research on union members' participation in union activities. First, as in longitudinal research (e.g., Fullagar & Barling, 1989; Kryl, 1990), the results of the current study support the direct and indirect prediction of members' participation by union commitment. The current study extends this finding by incorporating elements of Tetrick's (1989) causal structure of union commitment and Kryl's (1990) identification of a conflict model relating union commitment, participation, and job dissatisfaction. Consistent with the suggestions of these researchers, the dimensions of union commitment were related in a predictable fashion and participation in union activities was predictive of both extrinsic and intrinsic job dissatisfaction.

Second, consistent with the research on union commitment, both first-year socialization experiences (Fullagar & Barling, 1989; Gordon et al., 1980) and perceptions of the instrumentality of union activism predicted union loyalty. This is consistent with previous findings regarding instrumentality perceptions (e.g., Fullagar & Barling, 1989), even though we relied on a different formulation of instrumentality in the current study (e.g., satisfaction of individual needs as opposed to economic instrumentality). Similarly, both Marxist work beliefs (Fullagar & Barling, 1989) and subjective norms (Fullagar et al., 1992) emerged as predictors of union loyalty.

Third, the results of the current research refine several earlier predictions. Whereas Fullagar and Barling (1989) found that Marxist work beliefs predicted union loyalty, simultaneous consideration of other dimensions of union commitment in the current study resulted in the identification of Marxist work beliefs as a predictor of responsibility to the union. Similarly, Fullagar et al. (1992) identified subjective norms as a predictor of union loyalty. Consistent with Fishbein and Ajzen's (1975) formulation, subjective norms also predicted willingness to work for the union in the current study.

Fourth, Fullagar et al. (1992) identified shop stewards' transformational leadership styles as a predictor of union loyalty. The results of the current study argue for a more pervasive effect of union leadership. In addition to a direct effect on union loyalty, respondents' perceptions of shop stewards' transformational leadership styles directly predicted both responsibility to the union and participation in union activities.

In summary, the results of the current study offer strong support for the revised model. Replication of the model in a second independent sample suggests that the relationships posited in the model are robust. Of the 17 relationships in the revised model, 12 were invariant across samples. Combined with the excellent fit demonstrated by the model in both samples, these results support the robustness of the posited effects.

The results of our second study suggest several conceptual and methodological avenues for future research. Perhaps the most striking of these is the pervasive effect exerted by percep-
tions of shop stewards' leadership styles. Members' perceptions of shop stewards' leadership styles predicted members' loyalty, sense of responsibility, and actual participation in union activities. As noted elsewhere (Barling et al., 1992), research on union leadership has been sadly lacking in North America. The findings of the current study point to the need to redress this omission to further the understanding of union participation as well as to enhance union democracy.

With respect to members' participation in union activities, the findings of the current study suggest that interventions at the shop steward level (e.g., leadership training) would have a direct effect on increasing members' participation in union activities. Moreover, the magnitude of the standardized parameter estimates for the leadership-participation relationship (B = .32 and .24 for the study and replication samples, respectively) suggests that more effective leadership would have a meaningful effect on participation in union activities. Although this suggestion requires evaluation in an experimental or quasieperimental design, it points to a specific intervention that may be implemented to enhance both union commitment and union participation.

The results of the current study also point to the need for unions to create a climate that fosters loyalty to and participation in the union. Several theorists have commented that the traditional styles of union governance need to be changed (Sherman, 1986) to retain the support of current members and to attract new members. The current study identifies several likely areas of intervention in support of this endeavor. In addition to intervention at the shop steward level, the results of Study 2 suggest the need for unions to attend to the manner in which they bring new members into the union (i.e., their socialization practices). Moreover, unions are well advised to make members aware of the need for, and benefits of, individual involvement in the union. Further research involving experimental or quasieperimental designs seems to be a likely avenue for identifying the most effective means of achieving these objectives.

At a more general level, our results have implications for the study of attitude-behavior relationships. In particular, the theory of reasoned action and the partisan model of political participation have been presented as competing models (e.g., Fishbein & Ajzen, 1981). The results of our study suggest that the processes involved may in fact be complementary. Given the significant predictions obtained from both models in our study, it is likely that neither model provides a complete description of attitude-behavior relationships. Rather, the models are complementary in that the propositions derived from one model add to the predictive power obtained from the other model.

Finally, despite the support offered for our proposed model in the current study, the possibility that other models provide a better, or equivalent, fit to the data cannot be excluded. Future research should explore alternative formulations and different predictors in an attempt to identify the most efficient predictors of members' participation in union activities.

General Discussion

We have presented two studies that allow for greater understanding of members' participation in local union activities. The results of our first study suggest that members' participation in local union activities can appropriately be viewed as a unidimensional and cumulative construct. These results reflect the representative nature of union democracy (Sherman, 1986), in which members elect representatives who are charged with the day-to-day responsibilities of the union. Attempts to examine or predict members' involvement in the union must recognize the differential nature of various forms of participation as well as the democratic structure of union governance. We suggest that our development of a Guttman scale to measure members' participation in union activities is an appropriate way to account for the nature of union participation. The scale we propose is based on, and mirrors, Nicholson's (1978) description of levels of participation in local union activities.

The results of our second study are consistent with the suggestion that union socialization plays a dominant role in encouraging or inhibiting members' participation in union activities. Participation was predicted most proximally by union commitment, which in turn was predicted by perceptions of union socialization, subjective norms regarding participation, beliefs about the instrumental nature of participation, Marxist work beliefs, and perceptions of shop stewards' transformational leadership characteristics. Taken as a whole, these results suggest several avenues through which unions might increase the commitment and participation of their members; shop steward training and the careful implementation of socialization practices emerge as two of the most likely avenues of intervention. In this sense, our results point the way for more methodologically rigorous research on union commitment. Specifically, the experimental or quasieperimental evaluation of organizational interventions emerges as the next major endeavor in the study of union governance.

One question that the current studies do not address is the degree to which our findings generalize to other industrial relations jurisdictions, such as the United States. There are vast differences in unionization between Canada and the United States: approximately 40% of the Canadian nonagricultural workforce is unionized as compared with approximately 19% in the United States (Barling et al., 1992). Moreover, Canadian industrial relations is characterized by the widespread acceptance of closed-shop agreements in which employees in a unionized workplace must become union members. Despite these differences, previous research has found only trivial differences in union commitment between the two countries (Thacker, Tetrick, Fields, & Rempel, 1991). Moreover, in a study of the willingness to vote for a union (Barling, Laliberte, Fulkagar, & Kelloway, 1993), no differences in the predictors of willingness to vote for a union emerged between comparable Canadian and U.S. samples. Therefore, the available research provides some confidence in the generalizability of our results. The extent to which this confidence is justified remains an empirical question.

Both studies presented here are based on self-report data, leading to two potential limitations on our conclusions. First, the observed relationships may be a function of method bias rather than true relationships between theoretical constructs. The plausibility of this hypothesis is mitigated by the observation that, for the most part, the observed relationships
were predicted a priori in our study. Relationships that were not predicted did not emerge, as would be the case if monomethod bias was artifactually inflating the observed correlations.

The second potential limit to our findings also emerges from a reliance on self-report data. The difficulty in collecting data in a union environment has led to the use of self-report measures of behavioral constructs such as members' participation in union activities. Yet it is possible that such self-reports actually reflect the behaviors in question. In one study that directly addressed this question, McShane (1986a) reported that the correlation between self-reported and actual meeting attendance was .89, sufficient to justify the use of self-report data as a proxy for actual behavior. Indeed, McShane noted that the correlation might have been even higher had accurate union records been kept of participation. Taken in conjunction with the validity evidence we present in Study 1 and the temporal reliability of our scale, we suggest that the use of self-report data to reflect members' participation poses a possible but not a plausible threat to our results.

A third limitation of our results is the use of willingness to work for the union as a measure of behavioral intentions to participate in the union. The problem arises because two items on the measure ("If asked I would run for elected office in the union" and "If asked I would serve on a committee for the union") may be redundant with the participation measure, which includes items assessing holding union office and serving on union committees. Thus, the observed correlation between willingness to work for the union and union participation may reflect item redundancy rather than the relationship between behavioral intent and behavior. The plausibility of this suggestion is limited by two observations. First, previous research has not identified excessively strong correlations between willingness to work for the union and single-item measures of participation (19 < r < .59; Kelloway et al., 1992). Second, when we partialed out the variance attributable to other commitment dimensions, the magnitude of the correlations markedly decreased (13 < r < .31; Kelloway et al., 1992), which would not have been the case had the correlations been entirely attributable to item redundancy. Therefore we suggest that willingness to work for the union is most appropriately viewed as a measure of behavioral intentions that is not redundant with participation per se.

Finally, a fourth potential limitation to both studies reported here is the low response rate to membership surveys. The response rates we obtained are not inconsistent with previous research on union commitment (16%, Fullagar & Barling, 1989: 40%, Gordon et al., 1980). Moreover, in Study 2, union officials anticipated a much lower response rate (i.e., 10–15%) because of their experience with previous internal surveys. Nonetheless, the low response rates in our studies may have resulted in biased parameter estimates, particularly as completion of a union sponsored survey may plausibly be interpreted as a form of participation in the union.

Paradoxically, our development of a Guttman scale assessing members' participation in the union provides a means of identifying issues of sample selection bias in union research. Inspection of the distributions of participation scale scores in conjunction with known patterns of participation in the union (Barling et al., 1992) suggests that our samples overrepresent active unionists and underrepresent inactive members. Two possible limitations emerge from this observation. First, the underrepresentation of inactive union members limits the generalization of our findings to groups other than active participants in the union. Second, the parameter estimates we present may be upwardly biased because of sample selection bias. It remains for future research to validate our findings in more representative samples.

Finally, we have argued elsewhere that the study of union phenomena has much to offer organizational psychology and that organizational psychology, in turn, has much to offer labor unions (Barling et al., 1992). In recent years research on union commitment has provided a theoretical background and the empirical means for organizational psychologists to examine issues of union governance. Having recognized labor unions as a legitimate focus of enquiry, we suggest that it is now time for organizational psychologists to move beyond the notion of labor unions as research subjects to a view of labor unions as research partners. Through joint collaboration, labor unions and organizational psychologists can change their relationship from one of mutual neglect to one of mutual benefit. We believe that this transition is possible, desirable, and long overdue.

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