Research Note

Job satisfaction and organizational continuance commitment: a two-wave panel study

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Summary

The temporal relationship between job satisfaction and organizational continuance commitment over 13 months was examined in 295 professional employees of a British engineering company using LISREL with latent variables analysis. The cross-lagged path coefficients in the LISREL models were not significantly positive, suggesting that the relationship between these two variables was spurious and due to error variance. Test–retest coefficients for both variables were moderately positive, showing that the relative ranking of individuals on these variables was fairly stable over time.

Introduction

Job satisfaction has usually been defined as the extent to which an employee has a positive affective orientation or attitude towards their job, either in general or towards particular facets of it (Smith, Kendall and Hulin, 1969). It has been differentiated from job involvement which has been seen as the degree to which an individual identifies psychologically with their job (Locke, 1976; Lodahl and Kejner, 1965). Several conceptions of organizational commitment have been distinguished (Allen and Meyer, 1990; Griffin and Bateman, 1986; Mathieu and Zajac, 1990; Reichers, 1985). The concept most frequently studied has been attitudinal or affective commitment. This has been defined as the extent to which an employee identifies and is involved with a particular organization and has been most commonly measured by the Organizational Commitment Questionnaire (Porter, Steers, Mowday and Boulian, 1974). The second most frequently investigated concept has been variously referred to as behavioural (Salancik, 1977), calculative (Griffin and Bateman, 1986; Hrebiniak and Alutto, 1972) or continuous (Meyer and Allen, 1984) commitment. It has usually been conceptualized in terms of the costs of leaving an organization and has been most often assessed with the scale developed by Hrebiniak and Alutto (1972). This scale operationalizes commitment as the level of various inducements needed to leave an organization.

O’Reilly and Chatman (1986) distinguished three kinds of commitment based on the psychological processes of compliance (or exchange), identification (or affiliation), and internalization (or value congruence). A later factor analysis on a larger and more representative sample revealed two rather than three factors, called instrumental (compliance) commitment and
normative (identification and internalization) commitment (Caldwell, Chatman and O'Reilly, 1990). Furthermore, some have suggested that the various foci of commitment such as the particular individuals and groups within the organization need to be specified (Becker, 1992; Reichers, 1985).

There is evidence that job satisfaction, job involvement and attitudinal organizational commitment can be discriminated (Brooke, Russell and Price, 1988; Mathieu and Farr, 1991). Factor analytic studies of Meyer and Allen's (1984) affective and continuance commitment scales suggest that while affective commitment can be distinguished from continuance commitment, the continuance scale itself may consist of two dimensions reflecting few alternatives and high personal sacrifice respectively (Hackett, Bycio and Hausdorf, 1994; McGee and Ford, 1987; Meyer, Allen and Gellatly, 1990).

Interest in the causal nature of the relationship between job satisfaction and organizational commitment stems partly from the presumed role these two variables play in conceptual models for predicting turnover (Farkas and Tetrick, 1989; Williams and Hazer, 1986). Most models of turnover assume that greater job satisfaction leads to greater organizational commitment (Bluedorn, 1982; Marsh and Mannari, 1977; Mobley, 1977; Price and Mueller, 1981). The primary reason for this causal order appears to be that job satisfaction is a more immediate affective response to one's work which is established more quickly after joining an organization whereas organizational commitment is slower to develop since it is based not only on the job but on others aspects of working for the organization such as its goals and values (Porter et al., 1974).

However, Bateman and Strasser (1984), on the grounds of self-perception theory and research, suggested that greater organizational commitment may produce increased job satisfaction since commitment may initiate a rationalization process in which attitudes are made consistent with behaviour. They claimed empirical support for this causal sequence in a five-month longitudinal study of 129 nursing employees which found that the cross-lagged path coefficient between earlier organizational commitment and later job satisfaction was significantly positive while that between prior job satisfaction and subsequent organizational commitment was not significant. This finding suggests that job satisfaction results from organizational commitment.

Using the LISREL program for linear structural relationship analysis with latent variables, Williams and Hazer (1986) re-analysed the cross-sectional data of Bluedorn (1982) and Michaels and Spector (1982) to examine the causal nature of the relationship between organizational commitment and job satisfaction in four nested models of turnover. They interpreted their analyses as providing greater support for the causal order of job satisfaction leading to organizational commitment than the reverse sequence. However, they acknowledged that their data did not enable them to test for a reciprocal relationship between the two variables as such a model would have been underidentified and that assessing directionality with cross-sectional data is a 'nearly impossible task' (p. 229), which needs to be investigated with longitudinal data.

Subsequently, Farkas and Tetrick (1989) tested longitudinal analogues of the models proposed by Williams and Hazer (1986) on a four-year longitudinal study of 440 naval recruits. Because two of their variables (job satisfaction and re-enlistment intention) were only measured with two items, these variables had to be assessed with single rather than multiple indicators. However, in contrast to Williams and Hazer (1986), they suggested that their analyses indicated a cyclical or possibly a reciprocal relationship between job satisfaction and organizational commitment. Although they argued that the reciprocal relationship between these two variables could not be tested directly because of underidentification at each time period, it could have been assessed if they had restricted their model to the cross-lagged relationship between these two variables (Greenberg, Kessler and Logan, 1979; Cramer, 1994; Cramer, Henderson and Scott, 1995). Moreover, a re-analysis of their data in which cross-lagged paths between job satisfaction and organizational
commitment were postulated indicated that these paths became non-significant when error terms (many of which were significant) were free to correlate (Anderson and Williams, 1992).

Curry, Wakefield, Price and Mueller (1986) analysed with LISREL the temporal relationship between job satisfaction and organizational commitment over a seven-month period in 508 female nurses. Although multiple indicators were available, these were not used to define the latent variables. Instead, the loading of the indicator representing each latent variable was set equal to the square root of the reliability coefficient. They found no relationship between job satisfaction and organizational commitment over time. Vandenberg and Lance (1992) examined the relationship between job satisfaction and organizational commitment at the same point in time on two occasions five-months apart in 100 management information system professionals. The two exogeneous variables of bonus equity and value congruence were used to identify the reciprocal pathways from job satisfaction and organizational commitment respectively. Once again, multiple indicators were not used to define the variables. Their results suggested that organizational commitment causes job satisfaction. However, as was true for Williams and Hazer (1986), trying to determine the causal relationship between two variables measured at the same point in time would appear not to be possible. Furthermore, as Kenny (1979) has noted, correcting correlations for attenuation means that the resulting parameter estimates cannot be tested for significance in the usual way.

Because of differences in the way the data from the longitudinal studies were analysed, it is difficult to compare their results. The most appropriate method for examining the temporal relationship between job satisfaction and organizational commitment seems to be latent variable structural equation modelling in which the statistical fit of models postulating reciprocal, unidirectional and no cross-lagged relationships between job satisfaction and organizational commitment are compared. None of the previous studies has done this.

However, Curry et al. (1986) found that the cross-lagged path coefficients between these two variables were not significant as did Anderson and Williams (1992) when the error terms were free to covary, implying that there is no temporal relationship between the two variables. In both analyses, organizational commitment was measured with the nine-item Organizational Commitment Questionnaire (Porter et al., 1974), which is seen as being a measure of organizational affective commitment. In the study of Curry et al. (1986) job satisfaction was assessed with six of the 18 items from the Brayfield and Rothe (1951) scale. However, in the study of Anderson and Williams (1992) it was only measured with two items involving general satisfaction with the Navy. It is not known to what extent referring to the Navy in this way measured organizational satisfaction rather than job satisfaction and so it is unclear whether job satisfaction was confounded with organizational satisfaction.

A further potential problem with the study by Curry et al. (1986) is that the interval between the two waves may not have been sufficiently long for enough change in the two variables of job satisfaction and organizational commitment to have occurred. The relatively high test–retest path coefficients for job satisfaction (0.81) and organizational commitment (0.84) implies that insufficient change may have taken place. Previous longitudinal research indicates that determining the minimum interval necessary for studying the temporal relationship between job satisfaction and organizational commitment may be problematical. Test–retest correlations have varied for job satisfaction from a low of 0.35 over 6–8 months (Farkas and Tetrick, 1989) to a high of 0.68 over five months (Bateman and Strasser, 1984) and for organizational commitment from a low of 0.37 over 6–8 months (Farkas and Tetrick, 1989) to a high of 0.74 over five months (Vandenberg and Lance, 1992). All these studies used either the nine or 15-item Organizational Commitment Questionnaire (Porter et al., 1974) which is viewed more as a measure of affective commitment. Meyer et al. (1990) reported, for their two consecutive five-month intervals,
test–retest path coefficients of between 0.75 and 0.86 for affective continuance and between 0.83 and 0.92 for continuance commitment.

The explanation that job satisfaction influences organizational commitment implies that job satisfaction is less stable than organizational commitment (Porter et al., 1974). One index of stability is the test–retest correlation of a measure. However, the presumption that organizational commitment takes more time to develop than job satisfaction means that measures of organizational commitment taken shortly after employees have joined an organization are likely to be less well established than those taken later on when a person has had more experience on which to base their commitment to the organization. Consequently, during the initial period of belonging to an organization it is not clear whether the test–retest correlation for organizational commitment would be expected to be more positive than that for job satisfaction as implied by Curry et al. (1986). However, once a person’s organizational commitment has had an opportunity to develop, then the test–retest correlation for organizational commitment should be more positive than that for job satisfaction.

The expectation that the test–retest correlation for organizational commitment would be less positive in the earlier compared to the later stage of belonging to an organization was supported in the study by Farkas and Tetrick (1989) who measured job satisfaction and organizational commitment two months, 8–10 months and 20–21 months after enlistment. However, in the second interval when an individual’s organizational commitment would be expected to be better established than in the first interval, the test–retest correlation for organizational commitment was no more positive than that for job satisfaction, implying that job satisfaction was no less stable than organizational commitment. As mentioned previously, since job satisfaction was measured with only two items which referred to general satisfaction with the Navy, it is not clear to what extent job satisfaction was confounded with organizational satisfaction.

The main aim of this paper is to explore the temporal nature of the relationship between job satisfaction and organizational continuance commitment (as opposed to organizational affective commitment) in a two-wave panel study over a longer period than the study of Curry et al. (1986) using linear structural equation analysis with latent variables. Unlike previous studies which have investigated this relationship over time with LISREL (Curry et al., 1986; Farkas and Tetrick, 1989), latent variables based on three indicators each were used to measure these constructs. The statistical fit of models having reciprocal, unidirectional and no cross-lagged relationships between job satisfaction and organizational continuance commitment were compared. An interval of 13 months was chosen to increase the likelihood that sufficient change in these two variables would have occurred to be better able to explore their temporal relationship. Continuance commitment was chosen because it appears to show higher temporal stability than affective commitment (Meyer et al., 1990), which itself seems to show similar stability to job satisfaction (Bateman and Strasser, 1984; Farkas and Tetrick, 1989; Vandenbergh and Lance, 1992). Consequently, it was predicted that if continuance commitment is indeed more stable than job satisfaction when error variance is controlled, then prior continuance commitment is more likely to affect subsequent job satisfaction than prior job satisfaction is to affect subsequent continuance commitment.

**Method**

**Participants**

Questionnaires were distributed by the company on two occasions to all professional employees of a British engineering company manufacturing aerospace and industrial power systems. These
were returned by post direct to the author. Because of the need for anonymity and in order to try to match questionnaires across the two waves, participants were invited to give the first two letters of their first forename, the first two letters of their mother's first forename and the day in the month of their mother's birthday. At wave 1, the questionnaire was given to 1522 employees, of whom 1074 (71 per cent) completed and returned it and of whom 865 (57 per cent) provided one or more of the questionnaire matching details. At wave 2, the questionnaire was distributed to 1500 employees, of whom 1068 (71 per cent) filled and returned it and of whom 879 (59 per cent) gave one or more of the questionnaire matching details. Across the two waves, questionnaires from 295 individuals could be matched. Of these, 239 were male and 55 were female. Gender was missing for one person. On the first occasion the mean age of these 295 employees was 25.24 years (S.D. = 1.80) and the mean duration of their employment with the firm was 4.59 years (S.D. = 2.21). The mean interval between the completion dates of the two questionnaires was 378.36 days (S.D. = 13.04). At the time of the study, about 10 per cent of the workforce were laid off in a period of general national economic recession (1991–92).

Measures

This study was part of a larger one commissioned by the company to monitor sources of dissatisfaction thought to lead to turnover. A requirement of the study was that questions would be specifically developed for this survey in conjunction with personnel staff and that the use of existing scales was not permitted. The questionnaire used had been previously constructed for the company to assess among other factors job satisfaction and organizational continuance commitment (Cramer, 1993). Items measuring these two variables were answered on a five-point scale, ranging from 'strongly agree' to 'strongly disagree'. The six items which correlated most highly and consistently on the two orthogonal factors representing job satisfaction and organizational continuance commitment on the two occasions were selected to measure these two variables. Responses on these items were used as raw scores and were simply summed where appropriate. Cattell's (1966) scree test was used to determine the number of factors extracted by principal axis factoring. Orthogonal and oblique rotation of these factors gave essentially similar results. The highest an item from one scale loaded on the other was less than ±0.30.

For job satisfaction, the items (with their factor loadings on the oblique factors given in brackets first for wave 1 and second for wave 2) were 'I find my work interesting' (0.72, 0.78), 'My job gives me much pleasure' (0.67, 0.84), 'I do not feel very involved with my job' (−0.63, −0.75), 'I am generally satisfied with my job' (0.69, 0.76), 'I generally like working here' (0.53, 0.75) and 'Most of the tasks I carry out are lower than I should be doing' (−0.47, −0.53). The items for organizational continuance commitment were 'Changing companies is important for my further development' (0.79, 0.76), 'I intend to work for another organization within the next five years' (0.75, 0.69), 'Working for the same firm makes one too complacent' (0.57, 0.64), 'It is unrealistic to expect someone to work for a long time in the same organization' (0.57, 0.63), 'I could have a better career outside the company' (0.53, 0.53) and 'I could not satisfy my career aspirations any better anywhere else' (−0.45, −0.46). The correlation between the oblique factors representing job satisfaction and organizational continuance commitment was −0.16 at wave 1 and −0.24 at wave 2. Item responses rather than factor loadings were used to create scales and indicators. Items were coded so that higher scores represented greater job satisfaction and organizational continuance commitment. The alpha reliabilities for the two scales on the two occasions are shown in Table 1. They are high, ranging from a low of 0.79 for organizational
continuance commitment at time 1 to a high of 0.87 for job satisfaction at time 2, and indicate that all four variables were reliably measured.

In their study of the discriminant validity of measures of organizational commitment, job involvement and job satisfaction, Mathieu and Farr (1991) found that job involvement but not job satisfaction was significantly positively correlated with age, position tenure and organizational tenure. In the present study, job satisfaction was also not significantly correlated with these three variables, suggesting that this measure may be more appropriately described as an index of job satisfaction than of job involvement. Mathieu and Zajac (1990) in their meta-analysis of correlates of attitudinal and calculative commitment reported that the mean weighted correlation corrected for attenuation between position tenure and attitudinal commitment ($\hat{r}_t = 0.152$) was significantly more positive than that between position tenure and calculative commitment ($\hat{r}_t = 0.025$). In the present study, organizational continuance commitment was not significantly correlated with position tenure, implying that this scale may be more a measure of calculative or continuance commitment than affective commitment. Furthermore, the synchronous correlation between job satisfaction and organizational continuance commitment at wave 1 ($r = 0.35$) and wave 2 ($r = 0.37$) was closer to the corrected mean weighted correlation between work satisfaction and calculative commitment ($\hat{r}_t = 0.332$) than between work satisfaction and attitudinal commitment ($\hat{r}_t = 0.629$) reported by Mathieu and Zajac (1990).

Although Reichers (1985) has suggested that it may be necessary to distinguish desire to stay with the organization from organizational commitment, measures of organizational commitment, particularly those assessing continuance commitment, have traditionally included items referring to staying with or leaving the organization. For example, three of the 15 items of the Organizational Commitment Questionnaire (Mowday, Steers and Porter, 1979) contain this reference (‘4. I would accept almost any type of job assignment in order to keep working for this organization’; ‘9. It would take very little change in my present circumstances to cause me to leave this organization’; and ‘11. There’s not too much to be gained by sticking with this organization indefinitely’). All eight items of Meyer and Allen’s (1984) Continuance Commitment Scale (McGee and Ford, 1987) include this reference (e.g. ‘1. Right now, staying with my organization is a matter of necessity as much as desire’ and ‘5. It would be very hard for me to leave my organization right now, even if I wanted to’) while all four items of Hrebiniak and Alutto’s (1972) Index of Organizational Commitment are answered in terms of a general question incorporating this content (‘Would you leave your present organization under any of the following conditions?’). The finding that items containing this reference did not emerge as a separate factor in this study suggests that it may be difficult differentiating desire to leave from organizational continuance commitment.

**Model development and evaluation**

Linear structural relationship analysis (LISREL) with latent variables was used to determine the statistical fit of four competing models (Jöreskog and Sörbom, 1989). For each scale the responses of the six items were summed in pairs to form three observed indicators to represent that latent variable. The loading of the first pair was fixed as unity. The two items with the highest intercorrelation formed the first pair, followed by two items with the next highest intercorrelation. All four models postulated that the variable at time 2 was a function of the same variable at time 1. The error term of the same indicator at the two waves was free to covary. For instance, the first indicator of job satisfaction at time 1 was free to correlate with the first indicator of job satisfaction at time 2. This was done because Anderson and Williams (1992) have
shown that models which specify errors to be uncorrelated result in bias of the parameter estimates representing reciprocal effects. The first model was the most comprehensive and held that both cross-lagged paths were functional. In other words, it assumed that the two variables were reciprocally related in that time 2 job satisfaction was influenced by time 1 organizational continuance commitment and time 2 organizational continuance commitment was affected by time 1 job satisfaction.

The remaining three models were subsets of the first. The second and third models held that only one or other of the cross-lagged paths was operative while the fourth model posited no cross-lagged paths. The second and third models assumed a unidirectional relationship between the two variables. The second model held that time 2 job satisfaction was a function of time 1 organizational continuance commitment whereas the third model postulated that time 2 organizational continuance commitment was determined by time 1 job satisfaction.

The statistical fit of a model which is a subset of another one can be compared with that model. Consequently, the first model can be compared with the others, while the second and third model can be compared with the fourth. A model with a significantly lower $\chi^2$ as shown by a $\chi^2$ difference test provides a better fit to the data. Where no significant difference exists between any two of the models, the model with the fewer significant path coefficients offers the most parsimonious solution.

Five other indices of fit were also given, the first three of which are calculated by LISREL and described by Jöreskog and Sörbom (1989). The goodness of fit index represents the proportion of variance explained. The closer this value is to 1.00, the better the fit. The adjusted goodness of fit index adjusts the goodness of fit index for the degrees of freedom used to estimate the free parameters of the model. The root mean squared residual is a measure of the average residuals between the observed covariance matrix and the estimated matrix. The smaller this index is, the better the fit. To calculate the other two indices of fit, a structural null model needs to be estimated (Bentler and Bonett, 1980) although some feel this unduly restrictive model should only be used in exploratory studies where neither previous theory nor research is available as a guide (Sobel and Bohrnstedt, 1985). The normed fit index is $\left(\frac{\chi^2_{null} - \chi^2_{model}}{\chi^2_{null}}\right)$ (Bentler and Bonett, 1980) whereas the parsimonious fit index is the normed fit index multiplied by $\frac{df_{model}}{df_{null}}$ (James, Mulaik and Brett, 1982). The four models were estimated with the maximum likelihood method.

### Results

All analyses involved the full sample of 295 participants. The means (averaged across items), standard deviations, alpha reliabilities and intercorrelations for job satisfaction and organizational continuance commitment on the two occasions are presented in Table 1. Two-tailed related $t$-tests indicated that both job satisfaction ($t = 2.38, p < 0.05$) and organizational continuance commitment ($t = 4.40, p < 0.001$) were significantly lower at time 2 than at time 1.

All four variables were significantly and positively intercorrelated, ranging from a low of 0.26 between job satisfaction at time 1 and organizational continuance commitment at time 2 to a high of 0.70 between organizational continuance commitment at time 1 and time 2.

The $\chi^2$ test of each model and its degrees of freedom are shown in Table 2 together with the other goodness of fit indices and the standardized beta coefficients. The $\chi^2$ difference test revealed no significant differences between any of the nested models, suggesting that no model was superior to the other three: M4–M1 ($\chi^2 = 1.32$, n.s.); M4–M2 ($\chi^2 = 1.14$, n.s.); M4–M3
Table 1. Means, standard deviations, alpha reliabilities and intercorrelations of job satisfaction and organizational continuance commitment on two occasions (n = 295)

<table>
<thead>
<tr>
<th>Variables</th>
<th>JS1</th>
<th>JS2</th>
<th>OCC1</th>
<th>OCC2</th>
</tr>
</thead>
<tbody>
<tr>
<td>M</td>
<td>3.58</td>
<td>3.49</td>
<td>2.89</td>
<td>2.75</td>
</tr>
<tr>
<td>S.D.</td>
<td>0.68</td>
<td>0.77</td>
<td>0.65</td>
<td>0.71</td>
</tr>
<tr>
<td>JS1</td>
<td></td>
<td></td>
<td>0.66*</td>
<td>0.35*</td>
</tr>
<tr>
<td>JS2</td>
<td></td>
<td></td>
<td></td>
<td>0.26*</td>
</tr>
<tr>
<td>OCC1</td>
<td></td>
<td></td>
<td>0.36*</td>
<td></td>
</tr>
<tr>
<td>OCC2</td>
<td></td>
<td></td>
<td>0.70*</td>
<td></td>
</tr>
<tr>
<td>Alpha Reliability</td>
<td>(0.85)</td>
<td>(0.87)</td>
<td>(0.79)</td>
<td>(0.80)</td>
</tr>
</tbody>
</table>

JS1 and JS2 = job satisfaction at times 1 and 2 respectively; OCC1 and OCC2 = organizational continuance commitment at times 1 and 2 respectively; alpha reliabilities are in brackets along the diagonal.

*p < 0.001 (one-tailed).

Table 2. Goodness of fit indices and standardized beta coefficients for the four competing models (n = 295)

<table>
<thead>
<tr>
<th>Model</th>
<th>1</th>
<th>2</th>
<th>3</th>
<th>4</th>
<th>Null</th>
</tr>
</thead>
<tbody>
<tr>
<td>JS ← OCC</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>OCC → JS</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>JS → OCC</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>JS–OCC</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>df</td>
<td>42</td>
<td>43</td>
<td>43</td>
<td>44</td>
<td>66</td>
</tr>
<tr>
<td>$\chi^2$</td>
<td>125.33*</td>
<td>125.51*</td>
<td>126.40*</td>
<td>126.65*</td>
<td>1817.28*</td>
</tr>
<tr>
<td>$\chi^2/df$</td>
<td>2.984</td>
<td>2.919</td>
<td>2.940</td>
<td>2.878</td>
<td>27.535</td>
</tr>
<tr>
<td>GFI</td>
<td>0.939</td>
<td>0.939</td>
<td>0.938</td>
<td>0.938</td>
<td>0.375</td>
</tr>
<tr>
<td>AGFI</td>
<td>0.886</td>
<td>0.889</td>
<td>0.887</td>
<td>0.890</td>
<td>0.261</td>
</tr>
<tr>
<td>RMSQ</td>
<td>0.063</td>
<td>0.063</td>
<td>0.064</td>
<td>0.065</td>
<td>0.358</td>
</tr>
<tr>
<td>NFI</td>
<td>0.931</td>
<td>0.931</td>
<td>0.930</td>
<td>0.930</td>
<td></td>
</tr>
<tr>
<td>PFI</td>
<td>0.592</td>
<td>0.607</td>
<td>0.606</td>
<td>0.620</td>
<td></td>
</tr>
</tbody>
</table>

Standardized beta coefficients

| JS1–JS2        | 0.71*| 0.73*| 0.71*| 0.72*|      |
| OCC1–OCC2      | 0.76*| 0.76*| 0.80*| 0.80*|      |
| OCC1–JS2       | 0.07 | 0.07 |      |      |      |
| JS1–OCC2       | 0.03 |      |      |      |      |
| GFI = goodness of fit index; AGFI = adjusted goodness of fit index; RMSQ = root mean square residual; NFI = normed fit index \([\chi^2_{null} - \chi^2_{model}] / \chi^2_{null}\); PFI = parsimonious fit index \((NFI \times df_{model}/df_{null})\); JS1 and JS2 = job satisfaction at times 1 and 2 respectively; OCC1 OCC2 = organizational continuance commitment at times 1 and 2 respectively.

Discussion

The results of this longitudinal study found that the correlation between job satisfaction and organizational continuance commitment was significantly positive when both variables were...
measured at the same point in time and when they were assessed about 13 months apart, suggesting that the two variables may be causally related. The LISREL analyses, however, revealed that neither cross-lagged path coefficient was significantly positive, indicating that there was no temporal relationship between these two variables. Measuring affective rather than continuance commitment, Curry et al. (1986) and Anderson and Williams (1992) also concluded that there was no temporal relationship between job satisfaction and organizational commitment.

These results appear to differ from those of Farkas and Tetrick (1989), who investigated with LISREL the relationship between job satisfaction, organizational commitment and several other variables over three waves. They argued that their analyses implied that the relationship between job satisfaction and organizational commitment may be cyclical or even reciprocal. A direct comparison of the analyses of the two studies is not possible because the same generic models were not examined. However, three differences in particular between the analyses of the two studies should be noted.

First and foremost, the models tested by Farkas and Tetrick (1989) do not appear to have allowed the error terms of the latent variables to intercorrelate as pointed out by Anderson and Williams (1992). As a consequence of this omission, the association between job satisfaction and organizational commitment reported in their study may have been due to error common to these two variables. In accord with other investigations (Jöreskog and Sörbom, 1989), permitting correlations between error terms in the LISREL analyses in the present study produced models with a substantially better fit to the data. Anderson and Williams (1992) found that when error terms were allowed to covary over time the cross-lagged path coefficients were not significant.

Second, even though these models appeared to provide a fit as good if not better than other models considered, Farkas and Tetrick (1989) rejected models in which either earlier job satisfaction was free to covary with subsequent organizational commitment or earlier organizational commitment was free to covary with job satisfaction because the beta coefficients were negative whereas the zero-order correlations had been positive. In other words, the reciprocal relationship between job satisfaction and organizational commitment was assumed to be instantaneous. In contrast to their results, it should be noted that both the cross-lagged beta coefficients in the present study were positive, although neither were significantly so. And third, as they acknowledged, their models did not enable both of the synchronous paths between these two variables to be estimated at the same time. Although the models presented in this paper did not permit the synchronous beta coefficients between job satisfaction and organizational continuance commitment to covary, it did allow both of the cross-lagged paths from time 1 to time 2 to be estimated in the same model, thereby enabling them to be compared in the same analysis.

The results of this study also clearly show that the relative ranking of individuals with respect to job satisfaction and organizational continuance commitment was fairly stable over the 13-month period examined. The test–retest beta coefficients in this study are similar to those found by Curry et al. (1986) but are considerably higher than those reported by Farkas and Tetrick (1989) which varied quite considerably as a function of the model being tested. The test–retest correlations in the present study are also higher than those reported by Farkas and Tetrick (1989) but similar to those found by Bateman and Strasser (1984) and Vandenberg and Lance (1992) over the shorter period of five months. Consequently, it is possible that the interval of 13 months used in the present study was not long enough for sufficient change in job satisfaction and organizational continuance commitment to have occurred for the temporal nature of their relationship to be investigated.

One reason for the lower stability found in the Farkas and Tetrick (1989) study may have been due to the fact that their sample consisted of a cohort of younger individuals who had enlisted at the same time and who were followed up shortly after enrolling whereas participants in the present sample had not joined the organization at the same time and were older. As a
consequence, individuals in the present study had had more experience of the work and the organization on which to judge their satisfaction and commitment. A limitation of the present study is the relatively small number of individuals in the sample who provided the necessary information for matching their questionnaires from the two waves and who may not have been representative of the larger sample. Another potential weakness may have been the conceptualization of continuance commitment more in terms of the intention to leave the organization than the costs of leaving it. The construct validity of this measure needs to be further investigated.

Mean job satisfaction and organizational continuance commitment was significantly lower on the second than on the first occasion. Mean life satisfaction over this period, however, showed a significant increase (Cramer, 1995), implying that the change in job satisfaction and organizational continuance commitment may have been work-related and may have been associated with company changes leading to lay-offs in the workforce. Farkas and Tetrick (1989) also found job satisfaction and organizational commitment declined significantly over the three waves of their study. Bateman and Strasser (1984), on the other hand, implied that there was no significant change in these two measures over a five-month period which was confirmed when the present author checked their figures. Vandenberg and Lance (1992) also reported no significant change in these measures over a five-month period although the author's own re-analysis of their values indicated a significant increase in both organizational commitment ($t = 13.25$, two-tailed $p < 0.001$) and job satisfaction ($t = 4.74$, two tailed $p < 0.001$). The mean values of these variables appear similar over the seven-month period of the study by Curry et al. (1986) but this impression could not be confirmed because the test–retest correlations were not supplied. Although these studies vary in numerous ways and are too few to make meaningfully comparisons, there does not appear to be any relationship between the pattern of change in the two variables and the nature of the cross-lagged relationships obtained between them.

The explanation that job satisfaction affects organizational commitment implies that job satisfaction is less stable than organizational commitment (Porter et al., 1974). Stability is typically measured in terms of the test–retest correlation. In the present study a person's organizational continuance commitment was likely to have been established since employees had been with the firm, on average, for over four years. Nonetheless, the test–retest correlation for job satisfaction was not significantly less positive than that for organizational continuance commitment. In other words, there is no evidence to suggest that job satisfaction may be less stable than organizational continuance commitment in the present study. Consequently, this explanation for job satisfaction leading to organizational continuance commitment may have to be re-evaluated.

The fact that this study together with that of Curry et al. (1986) and Anderson and Williams (1992) found no temporal relationship between job satisfaction and organizational commitment has the important practical implication that simply increasing job satisfaction is unlikely to result in greater organizational commitment. Furthermore, it indicates that the observed cross-lagged correlation between job satisfaction and organizational commitment is spurious and is the result of common factors. What these factors are remains to be determined.

References


