The Endogeneity Bias in the Relationship between Employee Commitment and Job Satisfaction[^]

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Abstract

The effect of job satisfaction on organisational commitment has attracted a great deal of interest in the human resource management and organizational psychology literatures. However, much of the empirical work is based on relatively small sample sizes and neglects the possibility of endogeneity of job satisfaction, potentially leading to biased estimates of the effect of job satisfaction on organizational commitment. Using a large matched employee-employer dataset this paper re-examines the relationship between the two work-related attitudes by employing instrumental variables approach to correct for endogeneity. We show that estimating the determining factors of employee commitment without correcting for simultaneity understate the association.

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1. Introduction

Job satisfaction has been widely discussed in the economic and organizational psychology literature. Most empirical studies deal with self-reported, subjective measures at the individual level, and assume that reported subjective job satisfaction is a satisfactory empirical approximation to individual utility (Frey and Stutzer, 2002). In economic work, however, much less attention has been given to the concept of organizational commitment. This is perhaps surprising given that the dominant view among organizational researchers and human resource professionals is that employees who are satisfied with their jobs are more likely to be committed to the firm than those who are less satisfied, thereby reducing employee turnover[1] (e.g. Gaertner, 1999; Wallace, 1995; Lincoln and Kalleberg, 1990; Mowday et al., 1982). However, the existing empirical work on the relationship between job satisfaction and employee commitment, which is based on relatively small sample sizes, is not satisfactory because it mostly neglects the potential problem of endogeneity of job satisfaction. This suggests that if job satisfaction is not an exogenous variable in the organizational commitment equation, most of the existing empirical studies provide biased estimates of the effect of job satisfaction on organizational commitment.

There are two potential sources of bias that are discussed here. First, the relationship between satisfaction and commitment might reflect some unmeasured third variable - a variable that is hard or impossible to observe. An example is omitted personality traits in an employee commitment equation, where individuals' levels of job satisfaction are likely to be correlated with unobserved personality. If so, the findings of previous research that has investigated the effect of job satisfaction on organizational commitment are based on estimates that suffer from omitted variable bias (for discussion see Rayton, 2006). Second, there is the possible effect of simultaneity (e.g. Farkas and Tetric, 1989; Lance, 1991; Mathieu, 1991). Increased job satisfaction is likely to increase employee commitment, but there is also little doubt that increases in employees' levels of commitment will translate into larger levels of job satisfaction. For instance, employees may alter their levels of job satisfaction so that they match the degree to which they are committed to their employing organization (e.g. Bateman and Strasser, 1984; Vandenberg and Lance, 1992; Lund, 2003).

[1] Furthermore, Brown et al. (2006) found that employee commitment is positively associated with firm financial performance.

In empirical terms, this means that there is a feedback relationship between both variables causing, for example, probit estimates of the probability to be committed to organization to be biased. Both omitted variables and simultaneity may appear in the same equation.

The objective of this paper is to re-examine the employee commitment-job satisfaction relationship by using the 2004 Workplace Employment Relations Survey (WERS) - the latest in the series of cross-section matched employer-employee datasets - and utilizing econometric techniques to correct for endogeneity. Specifically, in the cross-section context of WERS, simultaneity presents severe problem of identification. Unless there are a priori restrictions available, it will not be possible to distinguish empirically an explanation based on two-way causation from an explanation based on one-way causation. The main example of such prior restriction would be variable exclusions. Hence, to correctly identify one equation from the other one needs a valid instrumental variable (IV) that is excluded from either the job satisfaction equation or the employee commitment equation.

The paper is structured as follows. We begin by deriving the direction of the bias. We continue by describing the database that we used, and explain how we constructed our measures. Following on from this, we present our empirical analysis. We conclude by summarizing our results and their implications.

2. Deriving the direction of the bias

2.1 Omitted variable bias

Consider, for example, the variable C_{fi} that indicates the employee *i* level of commitment to organization *f* and S_{fi} that indicates the employee *i* overall job satisfaction in organization *f*. Suppose personality trait, P_{fi}, is unobservable, but influence employee commitment:

$$C_{f_i} = aS_{fi} + bP_{fi} + u_{fi} \tag{1}$$

where u_{fi} is zero-mean white noise disturbance. If P_{fi} is omitted, the estimated $\hat{a}$ equals the true effect plus a potential bias term:

$$\hat{a} = \frac{\operatorname{cov}(S_{fi}, C_{fi})}{\operatorname{var}(S_{fi})} = \frac{1}{\operatorname{var}(S_{fi})}\left\{\operatorname{cov}(S_{fi}, aS_{fi}) + \operatorname{cov}(S_{fi}, bP_{fi}) + \operatorname{cov}(S_{fi}, u_{fi})\right\} = \\ a + b\frac{\operatorname{cov}(S_{fi}, P_{fi})}{\operatorname{var}(S_{fi})} + \frac{\operatorname{cov}(S_{fi}, u_{fi})}{\operatorname{var}(S_{fi})} \tag{2}$$

$$\text{Taking expectations,} \quad E(\hat{a}) = a + b\frac{\text{cov}(S_{fi}, P_{fi})}{\text{var}(S_{fi})} \tag{3}$$

Knowing the sign of b and the sign of the covariance between S_{fi} and P_{fi} tells us the direction of the bias. If $b > 0$ and $\text{cov}(S_{fi}, P_{fi}) > 0$ [2], the bias will be positive. The effect of S_{fi} will be over-estimated. If $b > 0$ and $\text{cov}(S_{fi}, P_{fi}) < 0$, the bias will be negative. The effect of S_{fi} will be under-estimated etc.

2.2 Simultaneity bias

We now consider two-equation structural model:

$$C_{f_i} = aS_{fi} + b'X_{fi} + u_{fi} \tag{4}$$

$$S_{fi} = \gamma C_{fi} + \delta' X_{fi} + v_{fi} \tag{5}$$

where X_{fi} is the vector of exogenous variables (firm, industry and worker characteristics) ; u_{fi} and v_{fi} are zero-mean white noise disturbance. Estimating models (4) and (5) individually will give us bias estimates for the coefficients a and γ, respectively.

To make this argument clearer, let's focus, for example, on estimating the equation (4). The reduced form equation for S_{fi} is:

$$S_{fi} = \pi X_{fi} + e_{fi} \quad \text{where} \quad \pi = \frac{b\gamma + \delta}{1 - a\gamma}\text{;} \; e = \frac{\gamma u_{fi} + v_{fi}}{1 - a\gamma} \; \text{and} \; a\gamma \neq 1 \tag{6}$$

Assuming that X_{fi} and u_{fi} are uncorrelated, we examine whether S_{fi} and u_{fi} are uncorrelated. The reduced form equation (6) suggests that S_{fi} and u_{fi} are correlated if and only if u_{fi} and e_{fi} are correlated. If we assume that u_{fi} and v_{fi} are uncorrelated[3] then e_{fi} and u_{fi} must be correlated whenever $\gamma \neq 0$.

Hence, estimating a single-equation model for C_{fi} will potentially lead to bias estimates. By assuming that $\sigma_{uv} = 0$ the covariance between S_{fi} and u_{fi} is:

$$Cov(S_{fi}\, u_{fi}) = E[S_{fi} - E(S_{fi})][u_{fi} - E(u_{fi})] = E(S_{fi}\, u_{fi}) = \frac{\gamma E(u_{fi}^{2})}{1 - a\gamma} = \frac{\gamma}{1 - a\gamma}\sigma_{u_{fi}}^{2} \tag{7}$$

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[2] If, $\text{cov}(S_{fi}, P_{fi})$ is zero the bias term disappears.

[3] This rules out omitted variables or measurement error in u_{fi} that are correlated with S_{fi}.

If $\gamma > 0$, $a > 0$ and $a\gamma < 1$ the asymptotic bias in the OLS estimate of the coefficient (a) of C_i will be positive[4]. In other words, if $a = 0$ we would, on average, estimate a positive effect of job satisfaction on employee commitment (the estimator of a is attenuated toward zero).

3. Data

We use data from the WERS 2004, which is a nationally representative cross-section survey based on a stratified random sample of UK establishments and a sample of employees at those establishments (see Chaplin et al., 2005 for comprehensive discussion regarding the sample design and selection). The WERS comprises an employee questionnaire (EQ) that consists of a self-completion questionnaire distributed to a random sample of up to 25 employees in workplaces[5] with 5 or more employees located in Britain (*n*=22,451; response rate 60.4%). Also, the WERS comprises a management questionnaire (MQ) that consists of face-to-face interviews with senior managers dealing with industrial, employee or personnel relations at the workplace (*n*=2,295; response rate=64%). Hence, the WERS differs from its predecessors in collecting matched employer-employee information, and therefore this is the first study that uses such a data to empirically examine the link between employee commitment and job satisfaction.

3.1 Organizational commitment

Three are three broad types of organizational commitment.[6] First, normative commitment refers to a desire to remain part of an organization due to feelings of moral obligation (Wiener, 1982). For example, an individual who has begun an important project may feel a sense of obligation to finish it. Continuance commitment refers to the perceived costs of leaving an organization, or the risk of losing valued "side bets" (Becker, 1960) such as pension entitlement. Affective commitment is the desire to belong to an organization, and more specifically the extent to which an individual identifies with a given organization (Mowday et al., 1982). For example, an

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[4] The asymptotic bias in the estimate of the coefficient (γ) of S_{fi} will be also positive.

[5] Our analysis is concerned with private-owned UK workplaces only.

[6] Each refers to "a psychological state that binds the individual to the organization" (Allen and Meyer, 1990: 14), but each is empirically distinguishable. While all three can be distinguished empirically, affective and continuance commitment do appear to be somewhat related, although the precise nature of the relationship between them is a matter of some debate (see Bergman, 2006 for a review).

employee who works for a charity that supports disadvantaged children may be committed to an organization in part because s/he identifies with the group it seeks to support. Affective commitment is the most studied kind of organizational commitment (Allen and Meyer, 1990), and the one that we focus on in this study.

We have identified the following survey question included in the EQ that provides information about an individual's identification with their organization and hence. It can therefore be used as a proxy of (affective) employee commitment. Specifically, employees were asked to indicate the degree of agreement with the following statement: *I share many of the values of my organization*. This question is calling for a qualitative response ranging from (1) *"strongly disagree"* to (5) *"strongly agree"*, and from where a five point index was constructed as follows:

$$C_{fi} = \begin{cases} 5 = \text{Strongly agree} & (11.31\%) \\ 4 = \text{Agree} & (41.83\%) \\ 3 = \text{Neutral} & (34.25\%) \\ 2 = \text{Disagree} & (9.73\%) \\ 1 = \text{Strongly disagree} & (2.88\%) \end{cases} \tag{8}$$

3.2 Job satisfaction

While the concept of job satisfaction is not uncontested (e.g. Hodson, 1991), there is general agreement in the literature that it refers to an emotional state emanating from an individual's evaluation of his or her experiences at work (Locke, 1976). In the EQ, employees were asked to evaluate their job satisfaction using a five point scale, where (5) represents the maximum and (1) the minimum, on seven aspects of their job:

(i) *The sense of achievement* (mean: 3.744);

(ii) *The scope for using your own initiative* (mean: 3.802);

(iii) *The amount of influence you have over your job* (mean 3.560);

(iv) *The training you receive* (mean: 3.309);

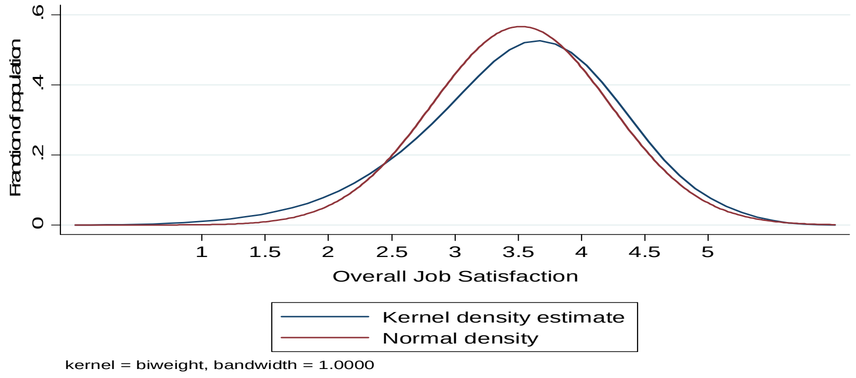
(v) *The amount of pay you receive* (mean: 2.885);

(vi) *The job security* (mean: 3.558); and

(vii) *The work itself* (mean: 3.763).

Since employees were not asked to evaluate their overall job satisfaction we have computed an average value of satisfaction for each employee (S_{fi}) and treat this variable as a continuous[7]. Figure 1 shows the Kernel density estimated distribution of the employee overall job satisfaction.

Figure 1. Kernel Density Estimated Distribution of the Employee Overall Job Satisfaction



3.3 Other variables

The EQ collects detailed information about characteristics of workers such as gender, age, ethnicity, marital status, educational qualification, gross weekly pay, number of working hours, job tenure, job contract, managerial duties, use of computer, union membership, job flexibility and a proxy of job suitability. Additionally, MQ allows us to extract useful information about industry and workplace characteristics including organizational size, establishment age, a proxy of management formality, percentage of employees working part-time and percentage of non-white employees. The covariates used in this study and their construction are summarized in Table 1. Given the focus of the paper, however, we focus our discussion on the empirical association between C_{fi} and S_{fi} [8].

[7] Recent work by Ferrer-i-Carbonell and Frijters (2004) found that assuming ordinality or cardinality of happiness scores makes little difference.

[8] Overall the signs and significance of all control variable coefficients match those reported in the existing literature. Also, the existing literature deals with issues related to endogeneity of the wage variable (e.g. Brown et al., 2006) and they are not therefore discussed here.

Table 1: Description of Covariates

Variables	Mean (SD)	Value
Male	0.52	0-1
Age (Less than 21)		0-1
Age (22-49)	0.67	0-1
Age (50 and over)	0.25	0-1
White	0.94	0-1
Single	0.25	0-1
No qualification	0.18	0-1
Ln(Gross weekly pay)	5.64(0.75)	3.22-6.82
Ln(Number of working hours)	3.54(0.52)	0-4.57
Tenure (More than 10 years)		
Less than 1 year	0.18	0-1
1-2 years	0.13	0-1
2-5 years	0.27	0-1
5-10 years	0.19	0-1
Job Contract (Permanent)		
Temporary	0.04	0-1
Fixed	0.03	0-1
Managerial duties	0.34	0-1
No need to use computer	0.28	0-1
Recent Member of Trade Union	0.24	0-1
Job flexibility[2]	0.259(0.592)	0-3
Job suitability[3]	3.706(0.879)	1-5
Organization Size (250+)		
Less than 50	0.17	0-1
50-249	0.14	0-1
Ln(Establishment age)	3.21(1.06)	0-6.80
Management formality[1]	9.58(2.02)	0-12
% of Employees working part time	22.31(26.22)	0-100
% of non-white Employees	6.84(13.63)	0-100
Sector (Other business and community services)		
Manufacturing	0.20	0-1
Electricity, gas and water	0.02	0-1
Construction	0.06	0-1
Wholesale and retail	0.14	0-1
Hotels and restaurants	0.04	0-1
Transport and communication	0.07	0-1
Financial services	0.09	0-1
Education	0.04	0-1
Health	0.10	0-1

[1]We use twelve binary variables that indicate a formal structure or process (Person mainly concerned with HR issues; Existence of a formal strategic plan; Investors in People; Presence of tests at induction as part of recruitment; Any communication channels; Any meeting between management and employee; Presence of a dispute procedure; Presence of an equal opportunity policy; Presence of a grievance policy; Presence of a performance appraisal programme; Formal target; and Any non-payment benefits). We then sum over the twelve variables to get an overall management formality score that ranges between 0 (not formally structured) to 12 (very formally structured): see Storey et al. (2009).

[2]We sum over three binary variables that indicate flexibility in workplace in terms of working time, parental leave and workplace nursery (0 indicates no job flexibility and 3 highly job flexibility).

[3]Workers were asked to how well the skills they personally have match the skills they need to do the job (1:much lower-5:much higher).

4. Empirical analysis

The analysis begins by estimating an empirical model of employee commitment in which job satisfaction is assumed to be an exogenous variable. We conduct an ordered probit analysis[9] to explore the determinants of the employee commitment index:

$$C_{fi}^{*} = aS_{fi} + b'X_{fi} + u_{fi} \tag{9}$$

where C_{fi}^{*} represent the latent variable denoting the unobserved propensity of worker *i* in firm *f* to be committed to firm *f*. Although, C_{fi}^{*} is unobserved, we observe C_{fi} such that:

$$C_{fi} = 1 \text{ if } C_{fi}^{*} \leq \mu_1 \tag{10}$$

$$C_{fi} = 2 \text{ if } \mu_1 < C_{fi}^{*} \leq \mu_2 \tag{11}$$

$$C_{fi} = 3 \text{ if } \mu_2 < C_{fi}^{*} \leq \mu_3 \tag{12}$$

$$C_{fi} = 4 \text{ if } \mu_3 < C_{fi}^{*} \leq \mu_4 \tag{13}$$

$$C_{fi} = 5 \text{ if } \mu_4 < C_{i}^{*} \tag{14}$$

where the a, b and μ are the parameters to be estimated[10].

The ordered probit estimates of the employee commitment model are given in Table 2. The magnitude of the ordered probit coefficient does not have a simple interpretation since the sign of the coefficient only uniquely determines the change in probability at the top and bottom categories of the dependent variable, and it may not determine the effect for the intermediate outcomes (see Greene, 2003). For this reason, we also present the marginal effects. The ordered probit coefficient, which is found to be statistically significant and great in magnitude, carries a positive sign suggesting a positive effect of job satisfaction on employee commitment. Turning to the marginal effects presented in Table 2, we estimate that the effects of job

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[9] We also adopt a random effects estimator to allow for the fact that the employee level data are drawn from a number of workplaces. In this case u_{fi} is decomposed into independent components as follows: $u_{fi} = \vartheta_f + \eta_{fi}$ where η_{fi} is a random error term with mean 0 and variance σ_{η}^{2}; ϑ_f is the firm specific unobservable effect capturing differences in satisfaction across firms with mean 0 and variance σ_{ϑ}^{2}, and it is assumed to be independent of S_{fi} and X_{fi}. Using the random effects estimator the magnitude of the coefficients were altered slightly. Finally the value of *ρ*, was found to be statistically significant but relatively small (0.06) implying little unobservable intra-firm correlation in the determinants of commitments. .

[10] $0 < \mu_1 < \mu_2 < \mu_3 < \mu_4$.

satisfaction on the probability that an employee is in the higher commitment category, i.e. *"strongly agree"*, is at 41.3% (evaluated at the sample mean, 3.5).

However, due to the likely overlap in unobserved characteristics that determine both employee commitment and job satisfaction and simultaneity there is potential bias in a.[11] Although the source of bias caused by omitted variables is different from that of simultaneity the result is the same, that is S_{fi} is correlated with u_{fi} in C^*_{fi} equation. To overcome these problems we use an instrumental variable (IV) probit model (see Amemiya, 1978; Rivers and Vuong, 1988)[12]. For simplicity, in model (9) commitment is treated as a binary variable (c_{fi}) which takes the value of one if the individual either *agrees* or *strongly agrees* with the commitment question. Thus (9) can be written:

$$c^*_{fi} = aS_{fi} + b'X_{fi} + u_{fi} \tag{15}$$

where the latent variable c^*_{fi} drives the observed outcome of being committed to the organization, c_{fi}, through the measurement equation:

$$c_{fi} = \begin{cases} 1, & \text{if } c^*_{fi} > 0, \\ 0, & \text{otherwise} \end{cases} \tag{16}$$

The instrumented equation is:

$$S_{fi} = \zeta' X_{fi} + \omega' \Theta_{fi} + v_{fi} \tag{17}$$

where ζ and ω are the vectors of coefficients, Θ_{fi} is the vector of identifiers and v_{fi} the error term. A proxy of flexible patterns working arrangements[13] is included in Θ_{fi} , which we assume that it is justifiable on a priory grounds. For instance, we argue that if co-workers are randomly assigned then one might end up working with someone who she/he does not like resulting in reducing *S* but not *c*. The assumption that employees' relationships with their co-workers affect job satisfaction is consistent

[11] This potential for unobserved heterogeneity will result in the error term, u_{fi} in model (9), being correlated with S_{fi}. The correlation between u_{fi} and S_{fi} may also result in biased estimates of the other coefficients.

[12] Rayton (2006) used a bivariate probit estimation technique, which allows for interaction between the error processes of the employee commitment and job satisfaction equations. However, the second dependent variables did not appear on the right-hand side of the first equation (recursive bivariate model, see Greene, 2003).

[13] Specifically, employees were asked to indicate whether changing working patterns (including shifts) would be available to them if they personally needed (1-0).

with the dominant view in the literature; it is unsurprising the nature of social interactions among employees has a major bearing on a given organizational member's experience of a given job. On the other hand, we suggest that while dysfunctional relationships with other organizational members have a bearing on the employment experience, they do not affect how members feel about the fundamental goals and values of the organization (see Lum et al. 1998; Bateman and Strasser 1984). Θ_{fi} also includes gender, marital status, job contract, trade union membership, number of working hours and a proxy of job suitability. Each of these variables was found to have an insignificant effect on employee commitment but a significant effect on job satisfaction. Likelihood ratio tests were used to confirm the validity of their inclusion[14].

Table 2: Ordered probit estimates for employee commitment

Employee commitment model	Ordered probit estimates					
		Marginal Effects				
	Coefficient	C_{fi} =1: Strongly disagree	C_{fi} =2: Disagree	C_{fi} =3: Neutral	C_{fi} =4: Agree	C_{fi} =5: Strongly agree
α (std.err.)[1]	0.905 (0.027)	-0.022 (0.002)	-0.122 (0.005)	-0.214 (0.009)	0.241 (0.010)	0.118 (0.005)
Controls	Yes					
μ_1	2.209 (0.293)	-	-	-	-	-
μ_2	3.213 (0.294)	-	-	-	-	-
μ_3	4.453 (0.297)	-	-	-	-	-
μ_4	6.062 (0.301)	-	-	-	-	-
Prob. at means	-	0.009	0.079	0.367	0.478	0.067
Log-likelihood	-5,789.764					
Wald Chi2 (34)	1,777.49					
Pseudo R2	0.165					
Observations	5,148					

[1]Robust standard errors are reported in brackets.

[14] This is a satisfactory way of proceeding assuming that the model is genuinely overidentified - that is there is an exclusion restriction that seems likely to be true from theoretical perspective. We use a likelihood ratio test to compare the values of the log-likelihood functions for the constrained and unconstrained models. The p-value of the resulting likelihood ratio test statistic was found to be LR chi2(7)=7.37 (prob.=0.391). This means that these exogenous variables have been appropriately omitted from the equation under examination.

The parameters of the probit (15) and the instrument equation can be jointly estimated by maximum likelihood. We write the joint density $f(c_{fi}, S_{fi} | X_{fi}, \Theta_{fi})$ as $f(c_{fi} | S_{fi}, X_{fi}, \Theta_{fi})\, f(S_{fi} | X_{fi}, \Theta_{fi})$ and thus, the log likelihood worker *i* in firm *f* is:

$$\ln L_{fi} = c_{fi} \ln \Phi(m_{fi}) + (1 - c_{if}) \ln[1 - \Phi(m_{fi})] + \ln \varphi\left(\frac{S_{fi} - \zeta' X_{fi} - \omega' \Theta_{fi}}{\sigma}\right) - \ln \sigma \qquad (18)$$

$$\text{where } m_{fi} = \frac{aS_{fi} + b' X_{fi} + \rho(S_{fi} - \zeta' X_{fi} - \omega' \Theta_{fi}) / \sigma}{(1 - \rho^2)^{1/2}}, \qquad (19)$$

$\Phi(\cdot)$ and $\varphi(\cdot)$ are the standard normal distribution and density functions, respectively. σ is the standard deviation of v_{fi} and ρ is the correlation coefficient between u_{fi} and v_{fi}. The results are shown in Table 3. Testing for the exogeneity of job satisfaction using IV Probit does reject the null hypothesis at the 10% significance level, suggesting that the endogenous regressor should not be treated as exogenous. There is a small and significant negative correlation (at the 10% significance level), rho, between the error terms of the instrument equation and the employee commitment equation. Given these results, we should expect the estimates for job satisfaction not to be of comparable magnitude across instrumented and non-instrumented specifications. In fact, comparing the coefficients for IV Probit (column B) and Probit (column A) yields larger estimates.

Table 3:

Model:	A) Probit estimates	B) IV probit[1]	C) Two-stage probit least squares[2]	
Dependent variable:	Commitment	Commitment	Satisfaction	Commitment
Coefficient: α	0.836	1.053	-	0.911
(Std. err.)	(0.032)	(0.129)		(0.121)
Coefficient: γ	-	-	0.249	-
(Std. err.)			(0.130)	
Controls	Yes	Yes	Yes	Yes
Log likelihood	-2,878.02	-7,155.98	-	-3,232.62
Wald Chi2(28)	1,083.01	2,245.46	-	641.52
F(32, 4488)	-	-	28.02	-
Rho	-	-0.16	-	-
Chi2 (Wald test of exogeneity)		Prob.=0.093	-	-
Observations	5,233	4,521	4,521	

[1]Instruments: gender, marital status, job contract, trade union membership, number of working hours and proxies of job suitability and flexible patterns working arrangements. Bootstrap standard errors are reported in brackets.

[2]Corrected standard errors are reported in brackets (see Keshk, 2003).

Based on our bias coefficient analysis due to simultaneity presented earlier, we would have expected an upward bias. The observed downward bias might be caused by omitted variables (possibly personal characteristics), which are negatively correlated with job satisfaction, i.e. this negative correlation explains the increase in the magnitude of the coefficient estimates for commitment in the IV probit model compared with this for the univariate probit analysis. The above IV probit, however, does not estimate employee commitment and job satisfaction simultaneously, but instrument job satisfaction in the probit model. As a next step, we follow a method similar to that described in Madala (1983**),** which allows simultaneous estimation of both variables. Specifically, in the first stage the two models are fitted using all of the exogenous variables to eliminate the likely correlation between the endogenous explanatory variables and the stochastic disturbance terms in each equation, which violates the assumptions of the classical OLS and probit methods:

$$c_{fi}^{*} = \xi' X_{fi} + \mu_{fi} \tag{20}$$

$$S_{fi} = \zeta' X_{fi} + \tau_{fi} \tag{21}$$

Equation (20) is estimated via probit and equation (21) is estimated via OLS. From these reduced-form estimates, the predicted values from each model are obtained. In the second stage, the endogenous variables are replaced by their respective fitted values:

$$c_{fi}^{*} = a\hat{S}_{fi} + b' X_{1,fi} + u_{fi} \tag{22}$$

$$S_{fi} = \gamma \hat{C}_{fi} + \delta' X_{2,fi} + v_{fi} \tag{23}$$

Again we estimate equation (22) via probit and equation (23) via OLS[15]. The two-stage probit least squares method gives us an unbiased and efficient estimator of each

[15]The explanatory variables in vector X_1 do not include controls for gender, marital status, job contract, trade union membership, number of working hours and proxies of job suitability and flexible patterns working arrangements. Since job tenure was found to be only a strong predictor of employee commitment, this variable does not appear in (X_2) in equation (23). The latter is consistent with Porter et al.'s (1974) model which assumes that (1) organizational commitment develops over a longer period than job satisfaction, and (2) organizational commitment is a more stable construct than job satisfaction. Because it is less volatile and more enduring, commitment is more likely to drive any decision to remain with an organization than is satisfaction. As Lum et al. (1998) note, commitment can be considered "more global" than job satisfaction, which is concerned the task environment of a particular employee. From this perspective, it is an employee's connection to the goals and values of the organization that determine job tenure, not the "transitory events" which affect job satisfaction and which "should not cause an employee to re-evaluate his or her attachment to the overall organization" (p.317).

parameter in the equations. For further discussion and method to adjust the standard errors see Keshk (2003). Column (C) in Table 3 presents the results. The results are generally in line with the IV probit model presented earlier. Interestingly, the coefficient on commitment in the job satisfaction equation is statistically significant, which means that there is evidence that job satisfaction increases with employees being committed.

5. Conclusion

The relationship between job satisfaction and employee commitment has received a great deal of attention from organizational researchers, but the extant research has produced conflicting results. In our study we sought to clarify the relationship between the two work-related attitudes using micro-econometric techniques that allow us to control for possible endogeneity arising from omitted variables and simultaneity. In doing so we used the Workplace Employment Relations Survey 2004, a nationally representative cross-section survey based on a stratified random sample of UK establishments and a sample of employees at those establishments. We show both theoretically and numerically that ignoring the endogeneity can lead to biased inferences about the relationship between job satisfaction and employee commitment. Specifically, in the binary response model with one continuous endogenous regressor, the effect of job satisfaction on employee commitment is higher than the estimate predicted by the univariate probit-model. The observed downward bias is not expected when taking account for simultaneity. We suggest that this could result from omitted variables (possibly personal characteristics) that are negatively correlated with employee commitment.

While we believe our study makes an important contribution to the literature, we recognize that it has limitations which suggest avenues for future research. First, we encourage future analyses which examine the relationship between these variables using panel data and recent advances in panel econometric analysis (see Wooldridge, 2002). Second, it would be interesting to consider normative commitment and continuance commitment, both separately and in conjunction with affective commitment, in order to develop a more complete picture of the satisfaction-commitment relationship. Finally, our study relies on UK data only. However, there is evidence that work-related attitudes may vary between cultures (see Cheng and Stockdale, 2003). Further research which sheds light on the job satisfaction-employee commitment relationship in different cultures would therefore make an important contribution to the literature.

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