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Exploring the relationship between job satisfaction and organizational commitment: An instrumental variable approach

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Abstract

The possible role of job satisfaction (JS) on organizational commitment (OC) has been a very important and hotly debated topic among experts. However, existing studies have yielded mixed results potentially due to utilization of small datasets, different methodological designs, estimation techniques that do not control for potential endogeneity between the variables, or a combination of these issues. Using a large matched employer-employee dataset from Britain (WERS2011), we find that increases in employees' JS positively influence OC. We also show that this relationship holds when an instrumental variable framework (IV ordered probit/IV probit) is adopted to take into account the potential endogeneity of JS. However, throughout the analysis, the IV estimates are smaller in magnitude in comparison to where JS is considered as an exogenous variable. Moreover, utilising a two-stage probit least square (2SPLS) estimator, we support our previous findings i.e. increased JS is likely to lead to enhanced OC, but we also show that greater OC leads to higher levels of JS suggesting that JS and OC are likely to be reciprocally related. Overall, the IV estimates confirm the importance of addressing the endogeneity issue in the analysis of the relationship between JS and OC.

Keywords: Job satisfaction; Organizational commitment; Endogeneity; Instrumental variable framework; WERS, Britain

Introduction

Job satisfaction (JS), which is commonly referred as an emotional state emanating from an individual's evaluation of his or her experiences at work (Locke, 1976), has been widely discussed in both the organizational psychology and labour economics literatures. Most of the existing studies deal with self-reported, subjective measures at the individual level, and assume that reported subjective JS is a satisfactory empirical approximation to individual utility (Frey and Stutzer, 2002a, 2002b).[1] This work has generally shown that JS is closely related to job turnover, absenteeism, supply of effort and propensity to take industrial action, which in turn may influence firm performance and profits. In the organizational literature and especially in the labour economics literature, however, much less attention has been given to the concept of organizational commitment (OC) and its possible link with JS, which is the issue that we focus on in this study.

The literature on OC is not only more limited in scope than the literature on JS, it is also more fragmented and less coherent. To begin with, three broad types of OC have been identified (Allen & Meyer, 1990). The first type is normative commitment, which 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. The second type is continuance commitment, which refers to the perceived costs of leaving an organization, or the risk of losing valued "side bets" (Becker, 1960), such as pension entitlement. The third type is affective commitment, which refers to the desire to

[1] A large number of measures of JS have been developed but there appears to be no consensus on how to measure JS. There is also debate as to whether single item questions are adequate, or whether it is better to conceptualize JS as multi-dimensional and to employ facet measures (van Saane 2003; Wanous, Reichers, & Hudy, 1997; Judge & Kammeyer-Muller, 2012). Van Saane (2003) reviewed and evaluated 29 JS measures published between 1988 and 2001 and found only seven met their reliability and validity criteria. The recently renewed debate about how precisely to define JS (Judge & Kammeyer-Muller, 2012) also implies earlier measures may no longer be considered adequate.

belong to an organization and more specifically the extent to which an individual identifies with a given organization (Mowday, Steers, & Porter, 1979). For example, an 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. Taken together this body of work, which has mostly been developed by organizational psychologists, shows that affective commitment, which is the most studied kind of OC (Allen & Meyer, 1990), has important economic significance since it is related to essential organizational outcomes similar to those reported in the JS literature (e.g. Fabi, Lacoursiere, & Raymond, 2015; Allen, Shore, & Griffeth, 2003; Sagie, 1998). Work by labour economists (e.g. Brown, McNabb & Taylor, 2011; Green, 2008) is also in line with this finding.

However, in the OC literature, the relationship between JS and affective commitment either tends to be ignored (e.g. Brown et al., 2011; De Clercq & Belausteguigoitia Rius, 2007) or remains ambiguous and inconclusive (Huang, You, & Tsai, 2012), suggesting that additional research in this area is needed. Rayton (2006), for example, reviewed the social scientific evidence and noted that four distinct sets of findings have been identified in the literature. The first finding, which is also the most commonly held view in the Human Resource Management (HRM) field, is that JS predicts OC (e.g. Top, Akdere, & Tarcan, 2015; Froese & Xiao, 2012; Malhotra, Budhwar, & Prowse, 2007; Bakan, Suseno, Pinnington, & Money, 2004; Elangovan, 2001; Mathieu, 1991). Highly satisfied employees are more likely to be committed to the organization than those who are less satisfied, thereby reducing employee turnover and withdrawal behaviours, and increasing job performance (Fabi et al., 2015; Brunetto, Teo, Shacklock, & Farr-Wharton, 2012). The second finding is that high levels of OC enhance JS (e.g. Imran, Afrif, Cheema & Azeem, 2014; Indartono & Chen, 2011; Paik, Parboteeah, & Shim, 2007; Lund, 2003; Vandenberg & Lance, 1992; Bateman & Strasser, 1984). The third finding is that OC and JS are reciprocally related (e.g.

Huang & Hsiao, 2007; Mathieu, 1991; Lance, 1991; Farkas & Tetrick, 1989)[2]. And finally, there is no relationship between JS and OC (e.g. de la Torre-Ruiz, Vidal-Salazar & Cordon-Pozo, 2017; Brunetto et al., 2012; Nawab and Bhatti, 2011; Cramer, 1996). The latter was empirically supported by Rayton (2006), but he recognized that his findings were not based on a dataset that was representative of a larger population[3]. In a word, no consensus has been reached with regard to the causality of JS and OC.

Nevertheless, it is important to clarify the association between JS and OC, given that both variables figure in a variety of models of individual work behaviours as explanatory or as predictor variables. Such models include those for turnover and turnover intention (e.g. de la Torre-Rius et al., 2017; Brunetto et al., 2012), job performance (e.g. Kooij, Guest, Clinton, Knight, Jansen, & Dikkers, 2013; Boxall, Ang, & Bartram, 2010), organizational citizenship behaviours (e.g. Paille, Grima, & Dufour, 2015; Ko and Smith-Walter, 2013), in-role behaviours (e.g. Gregory, Albritton, & Osmonbekov, 2010), public service motivation (e.g. Vandenabeele, 2009), and service effort level (e.g. Humborstad & Perry, 2011; Testa, 2001). Without a clear understanding of the nature of the inter-relationship between JS and OC, models of effects in which both appear could be incorrectly represented at micro/employee level. For example, in the field of strategic human resource management (SHRM), JS and OC are identified as the two most widely studied employee motivation variables in understanding the linkage between HRM and organizational performance (Jiang, Lepak, Hu, & Daer, 2012). Hence, the precise causal relationship between JS and OC has important implications for both theory and research concerning the employment relationship, and for HRM/managerial

[2] Mathieu (1991) and Lance (1991) both found that the influence of satisfaction on commitment was higher than the influence of commitment on satisfaction. Farkas and Tetric (1989) and Huang and Hsiao (2007), on the other hand, found the relationship to be broadly symmetrical.

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

practice - whether practitioners and managers should seek first to enhance JS, or OC, or whether the order of the activities does not matter (Mathieu, 1991).

We believe that the root of the reported discrepancies may lie in the potential problem of endogeneity of JS and therefore in the failure of previous studies to adopt appropriate modelling strategies to surmount this problem (see, for example, the early work of Bateman & Strasser, 1984 and Curry, Wakefield, Price, & Mueller, 1986). Hence, the purpose of the paper is threefold. First, we estimate the effects of JS on OC using ordered probit and probit regression estimators assuming that JS is exogenous. Second, we implement an instrumental variables (IV) estimator that addresses the concern about the endogeneity of JS in the estimation equation of OC. This allows us to examine the validity of the estimates that have relied on the assumptions about the exogeneity of JS. Third, we complement our analysis by applying a two-stage probit least squares (2SPLS) estimator introduced by Madala (1983) to estimate OC and JS simultaneously, i.e. increased JS is likely to enhance OC which in turn simultaneously translates into higher levels of JS. Though the endogeneity of JS has been acknowledged and investigated in the labour economics literature (e.g. Bockerman & Ilmakunnas, 2012; Schneider, Hanges, Smith, & Salvaggio, 2003; Judge, Parker, Colbert, Heller, & Ilies, 2001), existing HRM research is dominated by treating JS as an exogenous variable. But if JS is not an exogenous variable in the OC equation, most of the existing evidence may have provided biased inferences about the relationship between JS and OC, which could potentially have contributed to sources of the discrepancies in the JS-OC linkage.

We use a large matched employer-employee dataset collected from the British labour market and the above mentioned micro-econometric analytical techniques to re-examine the relationship between JS and OC. To the best of our knowledge, this is the first study that uses such data to empirically examine the link between OC and JS, as the existing empirical work primarily relied on employee respondents only, and small sample sizes, that makes generalization of the results impossible (e.g. de la Torrez-Ruis et al., 2017; Kontoghiorghes, 2016; Jayasingam & Yong, 2013; Froese & Xiao, 2012; Malhotra et al., 2007; Rayton, 2006; Bakan et al., 2004). One primary value of matched employer-employee datasets is to provide important information on both employer and employee, and thus to aid analysts in the separation of employer and employee effects (Jensen, 2010). The overarching conclusion drawn upon the utilisation of matched employer-employee dataset is that both firms and workers play an important role in explaining observed differences in work attitudes of individual workers. Ignoring the effects of one would be to overstate the effects of the other. To this end, we are able to control for a wide range of employee level and firm level characteristics when utilizing micro econometric techniques to estimate models for OC and JS.

The paper is structured as follows. We begin by discussing the existing literature on the conceptualisation of JS and OC, four hypothetical models of the JS-OC relationship, and the endogeneity issue of JS arising from omitted variables and simultaneity. We continue by describing the database and explaining the construction of OC and JS. Next, we present estimates' results and discuss our findings. This is followed by a discussion of limitations and directions for future research. Finally, we conclude the paper.

Literature review

Research on JS

Systematic research into JS began in the 1930s (Locke, 1976) exploring employees' evaluations of their job across dimensions of satisfaction/contentment/liking with their job, motivated by the idea that job attitudes affect productivity and performance, amongst other outcomes (Judge, Thoresen, Bono, & Patton, 2001; Weiss, 2002). Locke's review and synthesis of JS research provided a widely used definition of JS as "*a pleasurable or positive emotional state resulting from the appraisal of one's job or job experiences*." (Locke, 1976, p.1300. Original author's italics). Locke's definition included both affect and cognition making his definition consistent with those of attitudes more generally (Judge et al., 2001; Brief, 1998) but measures of JS emphasized cognition, neglected affect, and obscured differences amongst evaluations of, beliefs about, and affective experiences of jobs (Weiss & Cropanzano, 1996; Weiss, 2002). More recent efforts to define JS have explicitly drawn on attitude literature, and included both affect and cognition (Brief & Weiss, 2002). JS refers to "*an internal state that is expressed by affectively and/or cognitively evaluating an experienced job with some degree of favor or disfavor*" (Brief, 1998, quoted in Brief & Weiss 2002, p.283. Original author's italics). Judge and his colleagues (Judge & Klinger, 2008; Judge & Kammeyer-Muller, 2012) noted that JS is regarded as a complex social attitude, and defined it as "an evaluative state that expresses contentment with and positive feelings about one's job" (Judge and Kammeyer-Muller, 2012, p.347).

Federici and Skaalvik (2012) regard JS as an emotional response to all of the factors that an individual experiences in the placement of employment. Indeed, the concept and operationalisation of JS involves one's subjective evaluation of a wide range of work-specific factors, such as promotional opportunities, pay and benefits, work relationships, job autonomy and participation in decision making (Wood & Ogbonnaya, 2016; David, Gidwani, Birthare, & Singh, 2015; Rayton, 2006). These work-specific variables are generally categorized as either intrinsic or extrinsic. Intrinsic factors pertain to higher order variables such as desire for recognition, personal accomplishment and advancement (Nawab & Bhatti, 2011), while extrinsic factors refers to external environment elements including compensation, physical work environment, and quality of leadership (Suki & Suki, 2012). Empirical studies, on the other hand, can emphasise the evaluation of one's satisfaction with the entire work experience, or merely focus on satisfaction with a single or narrow aspect of the job. For instance, Bakan et al. (2004) investigate the effects of employees' satisfaction level with contingent pay schemes (i.e. profit sharing and save-as-you-earn schemes) on work attitudes. Top et al. (2015) examine the multiple facets of intrinsic and extrinsic job satisfaction including pay, promotion, supervision, fringe benefits, contingent rewards, operational procedures, co-workers, nature of work and communication. The focus of the research is of significant importance because people may place different degrees of importance on the various facets of the work that contribute to JS (Federici & Skallvik, 2012). For example, an individual may report a high level of dissatisfaction with one particular aspect of the job but is not dissatisfied with the job overall. In the present study, our estimates use employee's satisfaction with nine aspects of job characteristics embracing both extrinsic and intrinsic satisfaction as an instrument for JS.

Research on OC

The concepts "commitment" and "organizational commitment" date back to the 1950s (Becker, 1960; Gouldner, 1960) and refer to consistency in people's behaviour towards other people, institutions or organizations. Synthesising earlier work on commitment, Mowday, Steers and Porter (1982) suggest two alternative perspectives to conceptualize the notion of OC: attitudinal and behavioural. This distinction is relevant here because each provides grounds for contrasting hypotheses concerning the relationship between OC and JS. They suggested that attitudinal OC is a process concerned with how employees "come to identify with the goals and values of their organization" and wish to maintain membership (Mowday et al., 1982, p.24). It was defined formally as "the relative strength of an individual's identification with and involvement in a particular organization" (p.27) characterized in terms of the belief in and acceptance of the organization's values, willingness to exert effort on behalf of the organization, and strong desire to maintain membership (Mowday et al, 1982). Behavioural commitment, on the other hand, focused on how behaviours serve to bind the employee to the organization (Mowday et al. 1982; Salancik, 1977). Mowday et al. (1982) saw the two perspectives as complementary, but as their definition, and associated measurement tool (the OC questionnaire (OCQ) – Mowday et al., 1979) emphasized attitudes, this perspective has prevailed in management literature (e.g. Rode, Huang, & Flynn, 2016; Bakan et al., 2004; Allen et al., 2003).

Gouldner (1960) had found it possible to empirically distinguish commitment to specific values of an organization from commitment to the organization itself. Drawing on the idea that organizations can usefully be conceptualized as comprising multiple constituencies, Reicher (1985) proposed a multiple commitment perspective that employees can be committed to different foci of an organization, such as top management, co-workers, customers, occupation and so on. Support has been found for the idea, thus calling into question the idea of a unidimensional ‘OC’ perspective. Further reformulation of the concept emerged later (Allen & Meyer, 1990; Meyer & Allen, 1991).

Meyer and Allen (1991, 1997) suggested that conceptualizations of attitudinal commitment contained three core components: affective, continuance and normative commitment. *Affective commitment* denotes the desire to belong to an organization, and reflects the extent to which an individual identifies with the values and goals of a given organization (Meyer & Allen, 1991). It concerns "employees’ emotional attachment to, identification with, and involvement in the organization" (Meyer & Allen, 1991, p.67). Highly affectively committed employees remain members of the organization simply because they want to (Meyer & Allen, 1997). *Continuance commitment* refers to the perceived costs of leaving an organization, or the risk of losing valued ‘side bets’ (Becker, 1960). Employees with high levels of continuance commitment retain membership because they need to stay with the organization for the time being until they find a better or more suitable employment opportunity somewhere else (Meyer & Allen, 1997). *Normative commitment* is regarded as a desire to remain part of an organization due to feelings of moral obligation based on personal values and beliefs (Manion, 2004; Meyer & Herscovitch, 2001). Employees with high levels of normative commitment stay in the organization because they believe they ought to (Meyer & Allen, 1997).

Researchers have continued to reformulate the concept of OC. For example, Judge and Kammeyer-Muller (2012, p.343) defined it as "an individual's psychological bond with the organization, as represented by an affective attachment to the organization, a feeling of loyalty towards it, and an intention to remain as part of it". Klein, Molloy, & Brinsfield (2012, p.137. Original authors' italics) conceptualized OC as '*a volitional psychological bond reflecting dedication to and responsibility for a particular target*'. However, the common attribute of these emerging definitions remains consistent with the general view of OC being concerned with affect and behaviour directed by an employee towards their employing organization. The attitude perspective in particular has guided most empirical research on OC, primarily relying upon the work of Mowday et al. (1979; 1982) (e.g. Rode et al., 2016; Huang & Hsiao, 2007; Bakan et al., 2004; Allen et al., 2003; Elangovan, 2001) and Meyer and Allen (1991; 1997) (e.g. Ogbonnaya, Daniels, Connolly, & van Veldhoven, 2017; Kooij et al., 2013; Si & Li, 2012; Markovits, Davis, Fay, & van Dick, 2010; Malhotra et al., 2007).

The relationship between JS and OC

Model 1: JS is antecedent to OC.

Research advocating that JS is a predictor of OC is based on an exchange of resources between the organization and its members (Martin & Bennett, 1996). Specifically, a prospective member brings needs and goals to an organization and agrees to supply her or his knowledge, skills and abilities (KSAs) in exchange for organizational resources capable of satisfying his/her needs and goals (Angle & Perry, 1983). This argument is consistent with the tenets of social exchange theory (SET) (Blau, 1964) and the norm of reciprocity (Coyle-Shapiro & Conway, 2005; Gouldner, 1960). The norm of reciprocity postulates that a rewarding activity, gift or favour received by one party is expected to be returned in kind to the other party. In an organizational setting, JS reflects an individual's affective response to specific work-related facets, and is determined only by a subset of personal and organizational factors, e.g. job characteristics (Huang & Hsiao, 2007); whereas OC represents one's affective reaction to the whole organization (Martin & Bennett, 1996). As resources, manifested in perceived equitable and favourable treatment of the individual, provided by the organization satisfy individual needs, the resulting satiated state appears to align with a focal organization. In other words, employees who are satisfied with their jobs are strongly prone to remain in the organization, leading to a positive effect on OC (Malik, Nawab, Naeem, & Danish, 2010).

This model has received considerable empirical support. For instance, Top and Gider (2013) find a positive and substantive relationship between overall JS and OC among nurses and medical secretaries in Turkish hospitals, and state that JS explains 36% of total variance of the OC scores. Findings of white-collar-workers employed by foreign-invested companies in China (Froese & Xiao, 2012) show that various dimensions of JS including job autonomy, appraisal and pay satisfaction, influence OC, with job autonomy satisfaction being a stronger predictor of OC than pay and appraisal satisfaction. Aydogdu and Asikgil (2011) identify a strong positive correlation between JS and OC with data from employees in both manufacturing and service industry, and show that JS is a predictor variable of OC. The results of Elangovan's (2001) study across part-time students indicate that there are strong casual links between satisfaction and commitment (i.e. lower satisfaction leads to lower commitment). In addition to these, the idea that JS is a function of OC has been widely evidenced in a large body of HRM research (e.g. Top et al., 2015; Tarigan & Ariani, 2015; Chan & Qiu, 2011; Liao, Hu, & Chuang, 2009; Bakan et al., 2004), lending further empirical support to this model.

Model 2: OC is antecedent to JS.

A reverse causal ordering in which OC is causally antecedent to JS has also been proposed. The rationale of this model is based on cognitive dissonance theory (Festinger, 1957), in which 'a cognitive outlook such as commitment is rationalized by subsequent attitudes of job satisfaction' (Bateman & Strasser, 1984, p.97). It is believed that individuals make sense of the situation by developing a level of JS consistent with the level of OC to reduce cognitive dissonance (Huang & Hsiao, 2007). In this sense, people are committed to an organization ultimately because they join the organization, and this act, along with other conditions, subsequently shapes their attitude toward the work (Vandenberg & Lance, 1992). This reasoning is similar to the idea that 'individuals may develop commitment during their initial entry to the organization and subsequently interpret job experience (e.g. satisfaction) in light of their level of commitment' (Mathieu, 1991, p.609). Bateman and Strasser's (1984) longitudinal study of 786 nurses demonstrated that OC emerges before JS, so did Vandenberg and Lance's (1992) empirical findings on 455 employees of a multinational software R&D

company. The perspective of the causal precedence of OC to JS has also been documented in a number of more recent studies (e.g. Imran et al., 2014; Indartono and Chen, 2011; Paik et al., 2007). For instance, based on the data collected from teachers, Imran et al. (2014) examine the OC-JS relationship and confirm that an increase in OC leads to an increase in JS. Indartono and Chen (2011) find that OC influences JS by investigating the influence of perception of organizational politics on employee work attitudes.

Model 3: OC and JS are reciprocally related.

The third model posits that JS and OC are reciprocally associated. In this case, the theoretical arguments of the above two hypothetical model drawing upon the social exchange (Blau, 1964) and cognitive dissonance perspectives (Festinger, 1957) can both be utilised to justify the reciprocal relationship. Earlier work (e.g. Farkas & Tetrick, 1989; Williams & Hazer, 1986) suggest that JS and OC are either cyclically or reciprocally related. Lance (1991) and Mathieu (1991) found support for a reciprocal linkage between JS and OC, with JS affecting OC more strongly than the reverse. Using data collected from 3,037 Taiwanese employees, Huang and Hsiao (2007) compared the four models of the JS-OC relationship and found that the reciprocal relation model fitted the data best.

Model 4: OC and JS are independent.

Finally, some scholars have found no empirical support for any of the previously proposed causal relationships between JS and OC. Results of these empirical studies support neither the assertion that OC has a direct causal effect on JS nor that which holds JS to be a direct predictor of OC (e.g. de la Torrez-Ruiz et al., 2017; Brunetto et al., 2012; De Gieter, Hofmans, & Pepermans, 2011; Aghdasi, Kiamanesh, & Ebrahim, 2011; Currivan, 1999;

Cramer, 1996; Curry et al., 1986). Such findings may be attributed to two alternative explanations. First, JS and OC are independent constructs explained by the same antecedent variables. For instance, findings of De Gieter et al.'s (2011) study show that JS and OC are independent predictors of turnover among nurses, and the individual characteristics (i.e. personality and demographic antecedent variables) suggest that JS and OC are endogenous constructs. Similarly, Nawab and Bhatti (2001) find no interaction between JS and OC among university facility staff, and JS and OC are independent variables dependent on similar explanatory variables such as compensation. Second, the causal relationship between JS and OC is subject to the influence of a mediating or moderating variable. For example, de la Torrez-Ruiz et al. (2017) examine consequences of three aspects of employees' benefit satisfaction (i.e. benefit level, benefit determination, and benefit administration) on organizational commitment among Spanish workers and find an indirect-only mediation effect (Zhao, Lynch, & Chen, 2010): the effect of benefit satisfaction on OC is fully mediated by perceived organizational support. Another group of researchers (e.g. Akomolafe & Olatomide, 2013; Aghdasi, et al., 2011) find that the relationship between JS and OC is subject to the influence of the antecedent variable of emotional intelligence, which moderates the linkage between JS and OC among employees.

Methodological issues: endogeneity bias

With mixed or inconclusive findings being reported to support all four hypothetical models, the current paper considers the need for further analysis of the relationship between JS and OC. In part, this discrepancy may be caused by small sample sizes and methodological problems in the form of variations in study designs, measures of JS and OC and/or estimation techniques.[4] In the present study, we attempt to shed more light on the JS-OC relationship using a large scale cross-sectional matched employer-employee dataset and well established measures of JS and OC (e.g. Federici & Skaalvik, 2012; Rayton, 2006; Mowday et al., 1982; Green 2008), and applying IV techniques (IV ordered probit, IV probit and 2SPLS) to account for potential endogeneity. In order to be valid, the instruments are required to be strongly associated with JS but exogenous to OC; to this respect, we test for exogeneity and validity of instruments using standard over-identification test methods.

Regarding the latter, we argue that two potentially prevalent sources of bias contribute to this endogeneity issue: omitted variables, and simultaneity (for a detailed derivation of the endogeneity of JS, see Appendix A). First, the relationship between satisfaction and commitment might reflect some third variable. An example is omitted personality traits in an OC equation, where individuals' levels of JS are likely to be correlated with unobserved personality. If so, the findings of previous HRM research that has investigated the effect of JS on OC are drawn upon estimates that suffer from omitted variable bias (for discussion, see Rayton, 2006). Second, there is the possible effect of simultaneity. Increased JS is likely to increase employee commitment, but there is also little doubt that increases in employees' levels of commitment will simultaneously convert into higher levels of JS (see model 3; Huang & Hsaio, 2007; Lund, 2003; Vandenberg & Lance, 1992; Bateman & Strasser, 1984). In combination with the prior discussion with respect to the relationship between JS and OC, we are thus left with no clear conclusion as to the relationship between JS and OC.

[4] The empirical work is extensive and covers an extensive spectrum of methods (see, for example, De La Torrez-Ruiz et al., 2017; Top et al., 2015; Buonocore & Russo, 2013; Top and Gider, 2013; Forese & Xiao, 2012; Brunetto et al., 2012; Markovits et al., 2010; Rayton, 2006; Bakan et al., 2004; Huang & Hsiao, 2007; Elangovan, 2001; Wong, Chun, & Law, 1995; Mathieu, 1991; Curry et al., 1986; Williams & Hazer, 1986; Bateman & Strasser, 1984). It can be argued that the empirical strategy adopted here can deal with statistical issues concerning for example, endogeneity and the measurement levels of the examined variables (see for example, Bollen, 2001).

Methodology

Data

We use data from the WERS 2011, which is a nationally representative dataset integrating cross-section and panel samples of workplaces providing a sample of employees in them. The survey population includes all workplaces in Britain with five or more employees operating in Sections C to S of the *Standard Industrial Classification (SIC2007)*, which accounts for 35% of all workplaces and 90% of all employees in Britain. These workplaces were drawn from the official business register – Interdepartmental Business Register (IDBR) maintained by the Office for National Statistics. The WERS 2011 includes four components, including *Management Questionnaire (MQ)*, *Employee Questionnaire (EQ)*, *Worker Representative Survey* and *Financial Performance Questionnaire*.

Our empirical analysis exploits data drawn from the MQ and EQ. Specifically, interviews were conducted with the most senior managers who are responsible for employment relations, human resources or personnel. The *MQ* yields 2,680 workplaces with a response rate of 46%. Next, a self-completion questionnaire was randomly distributed to a maximum of 25 employees at the participating workplace after permission had been sought from the manager. In total, 21,981 employee questionnaires were returned, yielding a response rate of 54%. However, after eliminating observations with incomplete data, the sample used in the econometric analysis discussed below was reduced to 17,616 employees working in 1,820 workplaces.

Measuring JS

In the *EQ*, employees were asked to evaluate their JS using a five point scale, where (5) represents the maximum (i.e. '*strongly agree*') and (1) the minimum (i.e. '*strongly disagree*'), on nine aspects of their job, including

(i) *The sense of achievement* (mean: 3.846);

(ii) *The scope for using your own initiative* (mean: 3.899);

(iii) *The amount of influence you have over your job* (mean: 3.636);

(iv) *The training you receive* (mean: 3.414);

(v) *The opportunity to develop skills* (mean: 3.408)

(vi) *The amount of pay you receive* (mean: 3.047);

(vii) *The job security* (mean: 3.488);

(viii) *The work itself* (mean: 3.864); and

(ix) *Involvement in decision-making* (mean: 3.289)

Since employees were not asked to evaluate their overall JS we adopt a hybrid combination of the nine survey questions by generating an additive scale based upon Cronbach's alpha (α) ranging from 1 to 5, where the scale of reliability is 0.880 implying a good level of reliability and the mean 3.543, and treat this variable (S_{fi}) as continuous[5]. The rationale behind this measure of JS is provided by Rose (2007), and the same instrument of JS has been used in a number of WERS-based studies, including Wood and Ogbonnaya (2016), Lai, Saridakis & Johnstone (2017) and Bryson, Cappellari & Lucifora (2010). Figure B1 in Appendix B shows the Kernel density estimated distribution of the employee overall JS.

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

Measuring OC

One question in the *EQ* provides information about an individual's identification with their organization, and appears consistent with Mowday et al.'s (1982) definition of OC, emphasizing the belief in and acceptance of the organization's values. Specifically, employees were asked to indicate the degree of agreement with the following statement: '*I share many of the values of my organization*'. A similar operationalization of OC has also been used in previous studies, such as Brown et al. (2011), Green (2008) and Rayton (2006). This question calls for a qualitative response ranging from (1) *"strongly disagree"* to (5) *"strongly agree"*, from where a five point index was constructed as follows:

$$C_{fi} = \begin{cases} 5 = \text{Strongly agree} & (16.167\%) \\ 4 = \text{Agree} & (49.720\%) \\ 3 = \text{Neutral} & (26.796\%) \\ 2 = \text{Disagree} & (5.983\%) \\ 1 = \text{Strongly disagree} & (1.334\%) \end{cases} \tag{1}$$

Results

Ordered probit and RE ordered probit

Our analysis begins by estimating an empirical model of OC in which JS is assumed to be an exogenous variable, controlling a wide range of organizational and employee level variables in the OC equation. Table B1 in Appendix B summarises these data. More specifically, employee demographics such as age, gender, ethnicity, marital status, education, job tenure, work contract, supervisory duties, trade union membership and weekly wage are controlled (also see Tarigan & Ariani, 2015; Top et al., 2013; Malhotra et al., 2007; Bakan et al., 2004). In addition to this, the matched dataset also allows us to control for firm-level characteristics, including firm size (e.g. Brown et al., 2011; Storey, Saridakis, Sen-Gupta, Edwards, &

Blackburn, 2010), types of sector (e.g. Markovits et al., 2010) and establishments with recognised trade unions (e.g. Brown et al., 2011).

We then conduct a regular ordered probit analysis to explore the determinants of the employee commitment index:

$$C^*_{fi} = aS_{fi} + b'X_{fi} + u_{fi} \quad (2)$$

where C^*_{fi} represents 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:

$$C_{fi} = 1 \text{ if } C^*_{fi} \leq \mu_1 \quad (3)$$

$$C_{fi} = 2 \text{ if } \mu_1 < C^*_{fi} \leq \mu_2 \quad (4)$$

$$C_{fi} = 3 \text{ if } \mu_2 < C^*_{fi} \leq \mu_3 \quad (5)$$

$$C_{fi} = 4 \text{ if } \mu_3 < C^*_{fi} \leq \mu_4 \quad (6)$$

$$C_{fi} = 5 \text{ if } \mu_4 < C^*_{fi} \quad (7)$$

where a, b and μ are the parameters to be estimated[6].

We also employ a random effects (RE) ordered probit estimator to correct for intra-firm correlation among employees nested within the same workplace, given that multiple employee respondents are drawn in some workplaces.[7] The ordered probit and RE estimation coefficient results for the OC model are presented in columns 1 and 2 of Panel A in Table 1, respectively. Both coefficients are found to be positive and statistically significant, and the magnitude of the coefficients is very close (α=0.879 vs α_{RE}=0.889; p<0.01). Also, the value

[6] $0 < \mu_1 < \mu_2 < \mu_3 < \mu_4$.

[7] 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^2_η; ϑ_f is the firm specific unobservable effect capturing differences in satisfaction across firms with mean 0 and variance σ^2_ϑ, and it is assumed to be independent of S_{fi} and X_{fi}.

of ρ, is found to be statistically significant (p<0.01) but relatively small (0.076) implying little unobservable intra-firm correlation in the determinants of commitments. Overall, these results suggest that there is a positive association between high levels of JS and OC.

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 (Greene, 2003). The marginal effects (MEs) relating to JS are presented in Panel B of Table 1 where it can be seen that JS has a negative influence on being in the relatively low OC categories and a positive influence on being in the relatively high OC categories. Moreover, the effects are found to be highly statistically significant. It is evident, for example, that JS, evaluated at the mean, increases the probability that OC is at higher commitment category "strongly agree" by approximately 17 percentage points.

[Table 1 about here]

For brevity, Table 1 only presents the results relating to the job satisfaction variable. The results relating to the other control variables accord with the existing literature. For example, we find that the coefficients of female (e.g. Forkuoh, Affum-Osei, Osei, & Addo Yaw, 2014), married individual (e.g. Salami, 2008) and older employees (e.g. Salami, 2008; Dodd-McCue & Wright, 1996) to be positive and statistically significant. Similarly, leadership responsibility (e.g. Valentine, 2001) and higher wages (also see Al-Kahtani, 2012; Steers, 1977) increase the probability of OC. We also find that OC decreases with increasing tenure (e.g. Nifadkar and Dongre, 2014) and, surprisingly, permanent employees to be associated with reporting lower levels of OC (e.g. Foote, 2004). Moreover, we find that private sector employees experience greater affective commitment than their counterparts in the public sector (e.g. Zeffane, 1994). Also, employees in smaller organizations are more likely to report higher levels of OC than larger firms, also showing consistency with earlier studies (e.g. Storey et al., 2010; Forth, Bewley, & Bryson, 2006).

IV ordered probit

Due to the likely overlap in unobserved characteristics that determine both OC and JS and simultaneity (i.e. endogeneity of JS), there is potential bias in a. [8] 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 the C^*_{fi} equation. To overcome these problems we replicate the above analysis based on an instrumental variable framework (see Roodman, 2011). Thus we estimate the following joint model:

$$C^*_{fi} = \tilde{a} S_{fi} + \pi' X_{fi} + v_{1,fi} \quad (8)$$

$$S_{fi} = \lambda' X_{fi} + \omega' \Theta_{fi} + v_{2,fi} \quad (9)$$

Given that the dependent variable in Eq. (8) is an ordered outcome and the dependent variable in Eq. (9) is continuous, the model is estimated using a conditional (recursive) mixed process estimator (CMP).[9]

The set of instruments[10] included in Θ_{fi} are: (1) flexible working arrangements related to working time and day schedule. It is expected that there is a direct association between flexible working arrangements and job satisfaction (e.g. Wheatley, 2017; Possenriede &

[8] This potential for unobserved heterogeneity will result in the error term, u_{fi} in model (2), being correlated with S_{fi}. The correlation between u_{fi} and S_{fi} may also result in biased estimates of the other coefficients.

[9] CMP is a limited-information maximum likelihood (LIML) estimator where the first stage parameters are structural and the second stage parameters are reduced form. The error terms $v_{1,fi}$ and $v_{2,fi}$ are assumed to be jointly normally distributed. For further discussion see Roodman (2011).

[10] For the validity of the instruments see the discussion in the next section where the model is re-examined within an IV probit framework.

Plantenga, 2011), but not to how employees feel about the fundamental goals and values of the organization (see Lum, Kervin, Clark, Reid, & Sirola, 1998; Bateman & Strasser 1984). (2) We also include length of working hours. The length of working hours should reduce JS satisfaction (see Bakker, Demerouti and Verbeke, 2004; Demerouti, Bakker, Nachreiner, & Schaufeli, 2001), but is exogenous to OC. On the other hand, committed employees may not resent longer working hours if it is for the sake of the company-wide objectives, because they tend to gradually depersonalize and de-emphasize their self-interest in place of organizational interest and values (Mael & Ashforth, 1992). (3) Finally as additional instruments we include various HR responses as a result of the recent recession, such as cost cutting related to work recognition, increased workload, job rotation, pay freeze or cut, reduced non-wage benefits, and reduced contracted hours (Lai, Saridakis, Blackburn, & Johnstone, 2016). Such HR practices and measures are expected to significantly affect job satisfaction (Osterloh, Frey, & Frost, 2002; Ryan & Deci, 2000). On the other hand, they may exert a very limited effect on affective commitment because employees may view such responses as necessary for maintaining and achieving long-term organizational objectives and aims.

The last column of Table 1 presents the coefficient of JS, where JS is treated as an endogenous variable (column 3). As found previously, higher levels of JS are associated with higher levels of OC. The estimated JS coefficient, however, is found to be smaller in magnitude in comparison to one that is estimated using an ordered probit and treating JS as an exogenous variable ($\tilde{\alpha} = 0.755$ vs $\alpha = 0.879$). Also, the ME of endogenous JS on the probability of reporting the "strongly agree" in OC question is found to be 0.150, which is about 14 percent smaller that the ME of exogenous JS (0.174).

IV Probit and 2SPLS

In this section, OC 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, and use probit regression in order to examine the potential relationships between OC and JS. Thus Eq. (2) can be written:

$$c^{*}_{fi} = \breve{\alpha} S_{fi} + \breve{b}' X_{fi} + \upsilon_{fi} \qquad (10)$$

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

$$c_{fi} = \begin{cases} 1, & \text{if } c^{*}_{fi} > 0, \\ 0, & \text{otherwise} \end{cases} \qquad (11)$$

As noted earlier, to overcome the endogeneity problem, we use an instrumental variable IV probit model (see Amemiya, 1978; Rivers & Vuong, 1988). Here we formally investigate the null hypothesis of exogenous JS using the Smith–Blundell test. The Smith–Blundell test indicates that JS is endogenous (Chi squared (1) =5.359, p= 0.021). We further examine the validity of the same set of instruments discussed in the previous section using the Amemiya-Lee-Newey test of over-identifying restrictions. The tests of over-identifying restrictions indicates that the null hypothesis cannot be rejected (Chi squared (8) = 6.176, p-value= 0.628) and thus, exclusion of the additional instruments from the primary equation is valid. Finally, the instruments are found to be individually and jointly statistical significant with the F-statistic (F-statistic=163.770, p<0.01) to be in excess of the minimum threshold recommended by Stock, Wright, & Yogo (2002), and thus rule out weak instruments concerns.

The results from the IV probit are shown in Table 2 (column 3), along with a single-equation probit (column 1) and RE probit (column 2) models. The results from these models all suggest that a higher value of JS is associated with an increased probability of reporting high levels of OC. The IV coefficient on JS in column 3 is 0.696 (p<0.01) with ME of 0.243 and similarly to previous findings is smaller than the corresponding probit estimate ($\breve{\alpha}$= 0.819 and ME=0.286; p<0.01). Although small, we also find that there is a positive correlation, ρ , between the error terms of the instrument equation and the OC equation. This confirms the role of unobserved variables influencing both JS and OC.

[Table 2 about here]

The above IV probit, however, does not estimate OC and JS simultaneously, but instrument JS in the probit model. As the next step, we supplement our analysis by following a method similar to that described in Madala (1983**),** which allows simultaneous estimation of both variables. Hence, we estimate a 2SPLS model. 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:

$$c_{fi}^{*}=\xi' X_{fi}+\mu_{fi} \tag{12}$$

$$S_{fi}=\zeta' X_{fi}+\tau_{fi} \tag{13}$$

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:

$$c_{fi}^{*}=\bar{\alpha}\hat{S}_{fi}+\overleftarrow{b}X_{1,fi}+u_{fi} \tag{14}$$

$$S_{fi}=\gamma\hat{C}_{fi}+\delta' X_{2,fi}+v_{fi} \tag{15}$$

Again we estimate equation (14) via probit and equation (15) via OLS[11]. The 2SPLS method gives us an unbiased and efficient estimator of each parameter in the equations[12]. Column 4 in Table 2 shows the results, which are generally in line with the IV probit model presented earlier. Specifically the coefficient of JS is found to be positive, statistically significant and similar in magnitude to the estimate from the IV probit ($\overleftarrow{\alpha}$ =0.628 vs $\breve{\alpha}$ = 0.696). Furthermore, the coefficient on OC in the JS equation is positive and statistically significant (γ =0.469, p<0.01), suggesting that there is some evidence that JS increases with employees being committed to the organization.

Discussion

The relationship between JS and OC has been a hotly debated topic in organizational psychology research. Generally, four alternative relationships have been proposed: 1) JS predicts OC; 2) OC predicts JS; 3) JS and OC are reciprocally related; and 4) JS and OC are independent. However, findings of available research in organizational psychology and HRM literature have produced mixed and conflicting results as all four hypothetical models have received either strong or modest support. These variations may be caused by the utilization of small datasets, different methodological designs and inappropriate estimation modelling. On the other hand, advances in econometric methods and understanding along with the accessibility of appropriate large datasets provides invaluable opportunities to mitigate or overcome the methodological limitations that have been largely overlooked in the empirical studies, such as endogeneity of JS.

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[11]To help identify the simultaneous system of equations in vector X_2 we include the previously discussed instruments along with a set of control variables. In contrast the vector X_1 includes religious denomination (Farrukh, Wei Ying, & Abdallah Ahmed, 2016; Guiso, Sapienza, & Zingales, 2003), type of contract and various control variables. Treating OC as continuous, the Hansen *J* statistic is found to be 3.559, which is insignificant at the 5% level. Also, standard F-tests indicate the joint significance of these variables in the OC model (F(2, 13422)=25.790; p<0.01).

[12] For further discussion of methods to adjust the standard errors see Keshk (2003).

The present study sought to contribute to overcoming this deficiency, and to provide possibly more refined evidence of the JS-OC relationship. We utilised a large cross-sectional matched employer-employee dataset (WERS2011), and employed micro-econometric techniques (i.e. IV ordered probit/ IV probit and 2SPLS estimator) to control for potential endogeneity arising from omitted variables and simultaneity biases. Our findings drawn upon the probit/ordered probit and respective random effect estimators assuming JS is exogenous show that individuals with higher levels of JS are more likely to report higher levels of OC. This finding corroborates prior research that JS is a critical work lever and should be given priority in managerial practices deigned to foster OC (e.g. Fabi et al., 2015; Gibbs & Ashill, 2013; Boxall & Macky, 2007). HRM practices signal that organizational resources invested in a subset of personal and organizational factors of an individual's job satisfy one's needs and increase JS, and ultimately OC (Martin & Bennett, 1996). Using an IV estimator to control for endogeneity bias, our findings also lead to a positive and significant impact of JS on OC, but the IV estimates are smaller than those without instrumenting. In addition to this, empirical evidence from the 2SPLS shows not only that increased JS is likely to lead to enhanced OC but also that greater OC simultaneously contributes to higher levels of JS. This evidence seems to lend further support to previous studies that JS and OC are reciprocally related (e.g. Huang & Hsiao, 2007; Allen et al., 2003).

The present study makes two important contributions to the understanding of the JS-OC relationship in theory and practice. First, most HRM research involving JS and OC has been dominated by the perspective that JS is the precursor to OC in the estimation model, and many studies do find a positive and strong effect of JS on OC (e.g. Kontoghiorghes, 2016; Top et al., 2015; Jayasingam & Yong, 2013; Top & Gider, 2013; Chan & Qiu, 2011; Liao et al., 2009; Malhotra et al., 2007; Bakan et al., 2004). However, this stream of research has theoretically and empirically ignored the endogeneity of JS, potentially resulting in an incorrect classification of the relationship between JS and OC. Our findings show that the quantitative magnitude of the estimates without correcting endogeneity of JS may have generated biased inference about the relationship between JS and OC, and thus misleading implications for human behaviour outcomes (Huang & Hsiao, 2007). Hence, this finding raises important methodological implications for future work that aims to explain the linkage between JS and OC.

Second, our analysis has important implications for the formulation of HRM strategy, policy and practices, because the issue concerning which work attitudinal variable should be focused on for organizational interventions in the form of people management practices seems to be rendered moot (Mathieu, 1991). This is due to the reciprocity of JS and OC, i.e. changing either variable will also affect the other. More specifically, HRM practices influence a set of job characteristics (Ogbonnaya & Vallizade, 2016) that mirror both intrinsic and extrinsic dimensions of JS (e.g. work itself, job autonomy, training and development, promotion and contingency pay), improving employee satisfaction and in turn leading to improved OC (Fabi et al., 2015). Alternatively, these HRM practices also can be used to align with the key components of employee commitment such as shared values and emotional bond between an individual and his/her employing organization, because they relay positive signals about the extent to which employees are integral to organizational success and growth (Bowen & Ostroff, 2004). Improved OC is thus subsequently transferred into higher levels of jobs satisfaction.

Limitations and future research directions

Though we believe the current study makes important contributions to the JS-OC literature, we recognize that it has limitations that suggest avenues for future research. First, the

WERS2011 only provides information on OC that resembles Mowday et al.'s (1982) conceptualisation of commitment, which appears theoretically most aligned with affective commitment proposed by Meyer and Allen (1991; 1997). It would be interesting for future research to consider the three dimensional conceptualisation of OC, particularly normative commitment and continuance commitment in conjunction with affective OC, in order to develop a complete picture of the JS-OC relationship. Relatedly, the lack of consistency in defining and constructing JS and OC in the literature and empirical studies may potentially hinder the comparison between empirical studies. To this end, more attention should be directed to advance the understanding the concepts of OC/JS in terms of providing a more coherent and widely accepted conceptualisation and measurement of OC and JS.

Second, the analysis is based on cross-sectional designs and datasets, which may limit conclusions regarding the direction of the causality between JS and OC. Hence, the results from the current investigation should be interpreted with caution. We encourage future research which examines the causal ordering between JS and OC and other relevant variables (e.g. HRM policies and practices) using longitudinal research designs (Wright, Gardner, Moyniham, & Allen, 2005), panel data and recent advances in panel econometric analysis (Wooldridge, 2002) to provide evidence of more robust casualty relationships between JS and OC.

Third, potential mediators and moderators may exist in the causal relationship between JS and OC. De la Torrez-Ruiz et al. (2017) find that the pathway from an individual's benefit satisfaction to OC is fully mediated by perceived organizational support. Researchers have also suggested that the JS-OC relationship is conditional upon the influence of some moderator variables, such as level of knowledge work (e.g. Jayasingam & Yong, 2013) and emotional intelligence (e.g. Akomolafe & Olatomide, 2013; Aghdasi et al., 2011). Examining these moderating and mediating models of the JS-OC relationship seems to be a potentially fruitful avenue for future research.

Finally, our analysis relies on British employer and employee data. Evidence has suggested work-related attitudes including JS and OC may vary among societal and national cultures (e.g. Kirkman & Shapiro, 2007; Cheng & Stockdale, 2003). Future research that sheds light on the JS-OC relationship in different cultural contexts would therefore improve the understanding of the nature of the JS-OC relationship.

Conclusion

This study employs micro-econometric techniques (i.e. probit/ordered probit estimator, random effects estimator, IV ordered probit/IV probit and 2SPLS) and a large matched employer-employee dataset to re-examine the relationship between JS and OC by correcting for potential endogeneity of JS arising from omitted variables and simultaneity in the OC equation. Findings from ordered probit and probit model specifications as well as RE probit/ordered probit estimators assuming employee satisfaction is exogenous show that JS has a positive and significant effect on OC. However, the magnitude of this relationship becomes smaller when an IV estimator is used to correct for endogeneity. Moreover, utilising the 2SPLS estimator that allows simultaneous estimation of both attitudinal variables, we find that JS and OC are potentially reciprocally related.

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Table 1. Ordered probit, RE ordered probit and IV ordered probit estimates

Panel A	Column 1		Column 2		Column 3	
	Ordered probit		Random effects ordered probit		IV ordered probit	
	Coef.	Std. Err.	Coef.	Std. Err.	Coef.	Std. Err.
Overall job satisfaction	0.879***	0.013	0.889***	0.014	0.755***	0.046
Controls	Yes		Yes		Yes	
Cut point 1	1.757	0.121	1.724	0.138	1.322	0.194
Cut point 2	2.692	0.120	2.689	0.136	2.254	0.195
Cut point 3	3.958	0.121	4.000	0.137	3.486	0.198
Cut point 4	5.667	0.123	5.780	0.140	5.187	0.203
Log-likelihood	-18444.488		-18318.123		-27841.620	
LR/Wald Chi squared (35)	6049.250		5347.720		1591.360	
Pseudo R squared	0.141					
ρ			0.076	0.007	0.085	0.030
Number of observations	17616		17616		13467	
Panel B						
	ME	Std. Err.			ME	Std. Err.
Category 1	-0.009***	0.001			-0.008***	0.001
Category 2	-0.062***	0.002			-0.057***	0.004
Category 3	-0.235***	0.005			-0.198***	0.013
Category 4	0.131***	0.004			0.114***	0.008
Category 5	0.174***	0.003			0.150***	0.009

***p<0.01.

Table 2. Probit, RE probit, IV probit and 2SPLS estimates

Panel A	Column 1		Column 2		Column 3		Column 4			
	Probit		Random effects probit		IV probit		2SPLS			
							Probit		OLS	
	Coef.	Std. Err.	Coef.	Std. Err.	Coef.	Std. Err.	Coef.	Std. Err.	Coef.	Std. Err.
Overall job satisfaction	0.819***	0.019	0.837***	0.020	0.696***	0.057	0.628***	0.048		
Organizational commitment									0.469***	0.094
Controls	Yes		Yes		Yes		Yes		Yes	
Log-likelihood	-7064.170		-7012.718		-20720.983		-8060.055			
LR/Wald Chi squared (35/36)	3026.600		2249.900		986.580		1034.830			
F (43, 13456)									64.920	
(Pseudo) R squared	0.176						0.060		0.172	
ρ			0.094	0.011	0.089	0.038				
Number of observations	13467		13467		13467		13467		13467	
Panel B										
	ME	Std.Err.			ME	Std.Err.	ME	Std.Err.		
	0.286***	0.007			0.243***	0.020	0.225***	0.017		

***p<0.01.

Appendix A: 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 influences 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:

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

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

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 $\text{cov}(S_{fi}, P_{fi}) > 0$ [13], 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.

2.2 Simultaneity bias

We now consider two-equation structural model:

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

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

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

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 us focus, for example, on estimating the equation (4). The reduced form equation for S_{fi} is:

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

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[14] 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:

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

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[15]. 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).

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

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

Appendix B: Figures and Tables

Figure B1: Kernel density estimate of JS

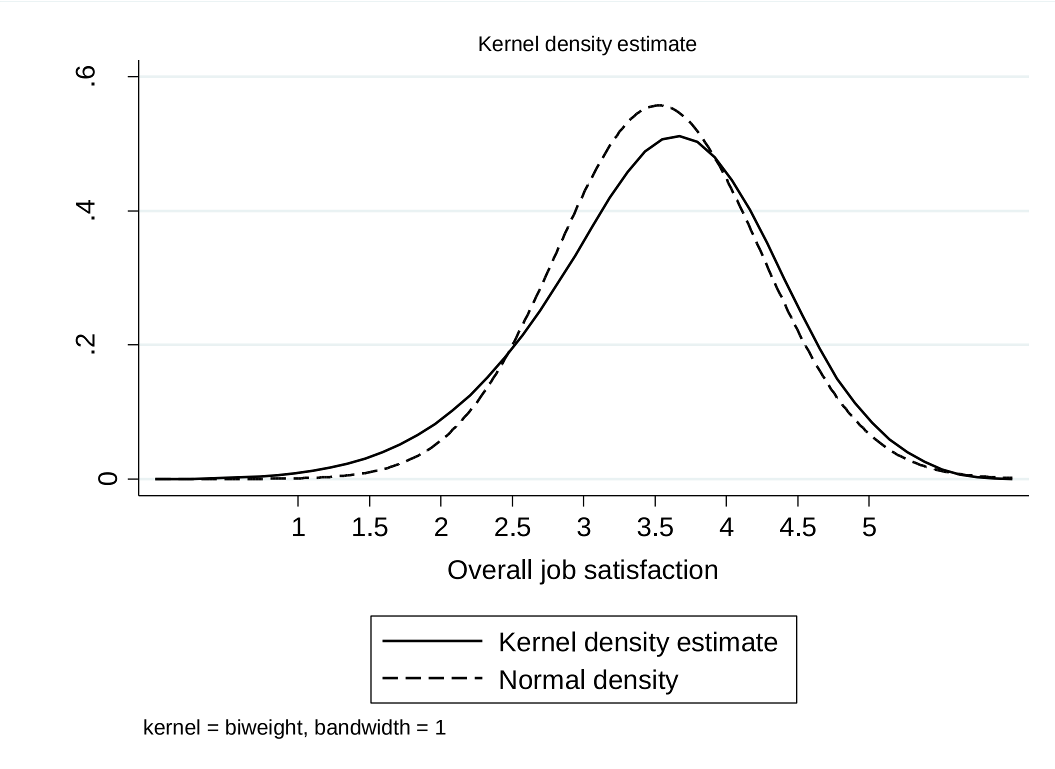

Table B1: Summary statistics of the variables used in the empirical analysis

Study Variables	Mean	S.D.	%
Overall job satisfaction score	3.543	0.717	
Affective organization commitment (I share many of the values of my organisation)			
Strongly disagree			1.334
Disagree			5.983
Neutral			26.796
Agree			49.720
Strongly agree			16.167
Employee characteristics			
Age			
16-21yrs			4.536
22-29yrs			17.054
30-39yrs			23.042
40-49yrs			26.859
50-59yrs			21.535
60-65+yrs			6.974
Female			50.598
British			85.603
Married			69.056
Academic qualification			95.065
Job tenure			
less than 1yr			13.128
1 to less than 2yrs			10.766
2 to less than 5yrs			24.375
5 to less than 10yrs			24.140
10yrs or more			27.591
Permanent			93.159
Supervisory responsibility			34.758
ln(midpoint weekly wage)*	5.913	0.754	
Member of a trade union or staff association			28.630
Organization characteristics			
Private sector			76.263
Firm size			
Small firms (n<50)			17.244
Medium-size firms (49<n<250)			11.744
Large firms (n>=250)			71.012
Recognized trade union or staff association			55.751
ln(1+firm age in years)*	3.232	1.010	

The mean of the 9 items formed the overall job satisfaction measure. All estimates computed using sample weights and based on sample size of 17614 observations. We also consider industry classifications (SIC 2007), but they have been excluded from the table for simplicity.

*ln() denotes natural log.