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RESEARCH REPORT

Revisit the Causal Inference Between Organizational Commitment and Job Satisfaction: A Meta-Analysis Disentangling Its Sources of Inconsistencies

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Existing theories and studies have exclusively focused on the direct temporal ordering of organizational commitment (OC) and job satisfaction (JS). However, their ordering varies highly across empirical studies. It is unclear whether such high variation is caused by statistical artifacts (i.e., spurious variation) or substantive moderators (i.e., true variation). Therefore, to disentangle artificial and substantial sources of this variation, we adopted traditional and full information meta-analytic structural equation modeling based on 71 independent samples ($N = 16{,}698$) with panel designs. After correcting for statistical artifacts, there was moderate heterogeneity in the lagged effects of OC and JS, suggesting that their ordering was not direct or simple but moderated by other variables. Further, the conceptualizations and/or measurements of OC and JS, time lag, and the timing of the investigation (e.g., newcomers vs. nonnewcomers) moderated their ordering. Finally, different specifications of the ordering between OC and JS resulted in differential cross-lagged relationships among JS, OC, and turnover intentions ($k = 31, N = 6{,}876$). Our meta-analysis provides new insights into the conceptualizations and/or measurements of OC and JS, adds some theoretical clarification to the dissenting theories by incorporating the time element, nuanced differences in the OC conceptualizations and the timing of the investigation, and raises concern over theory and research informed by studies overlooking alternative orderings of JS and OC. Looking forward, we suggest several promising future directions for explaining the heterogeneous cross-lagged relationships between OC and JS.

Keywords: organizational commitment, job satisfaction, conceptual clarity, contamination and deficiency, time

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Job satisfaction (JS) and organizational commitment (OC) are often modeled as common antecedents, mediators, or outcomes of different organizational phenomena (e.g., job performance, abusive supervision, workplace deviance, and sleep). However, researchers have diverse specifications of the temporal ordering between JS and OC (e.g., Douglas & Leite, 2017; Li et al., 2018; Podsakoff et al., 2007). This may generate seemingly equivocal findings, cast doubt on the validity of research models, and consequently impede theory building and knowledge accumulation. This issue is further fueled by dissenting theories and conflicting findings on the temporal ordering of OC and JS (Klein & Brinsfield, 2016; Meyer, 2016). Theoretically, there are four competing perspectives: (a) JS causes OC;[1] (b) OC causes JS; (c) JS and OC have reciprocal causation; and (d) JS and OC have no causality. Empirically, a few cross-sectional studies examine their temporal ordering but reveal large inconsistencies, providing support for each of the four dissenting theoretical perspectives (e.g., Huang & Hsiao, 2007; Saridakis et al., 2020; Tarigan & Ariani, 2015). Consistently, panel studies demonstrate high inconsistencies.[2] For instance, some studies support no causality (e.g., Currivan, 1999), some support that OC or JS precedes the other (e.g., Shin & Reyes, 1991; Vandenberg & Lance, 1992), and one supports reciprocation (Abdelmoteleb, 2019). Thus, one critical question arises: What cause these empirical inconsistencies?

There are two possible explanations. First, statistical artifacts (e.g., sampling and measurement errors) could result in seemingly contradictory findings across studies—an illusion of inconsistencies (i.e., spurious variation; Schmidt & Hunter, 2015). Researchers have well recognized the limitations of any single study, such as poor population representativeness due to sampling error which could generate complex, ambiguous findings across studies and called for multiple, diverse samples to address the ordering of OC and JS (e.g., Currivan, 1999; Shin & Reyes, 1991).

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[1] Causality refers to causal inference other than actual causality, and this study focuses on the temporal precedence.

[2] There are 11 panel studies examining this issue (see Supplemental Table S1).

Second, inconsistencies across studies may reflect true variation. That is, the ordering between JS and OC might not be simple or direct, but moderated by other variables. Several variables may moderate the ordering between JS and OC. First, constructs are building blocks of theories (Suddaby, 2010). Divergent theoretical perspectives and empirical evidence for the ordering of JS and OC might be due to the neglection of the nuanced yet critical differences in construct conceptualizations and operationalizations. For instance, affective attachment theory (Lawler, 1992) predicts that JS causes OC because JS forms earlier (i.e., soon after organizational entry) than OC (i.e., long after organizational entry). However, normative commitment can develop *before* organizational entry (Meyer & Allen, 1997), and therefore, JS may not be its cause. Moreover, proper construct operationalization is critical for theory testing and development (Edwards & Bagozzi, 2000). Some OC and JS measures (e.g., Ironson et al., 1989; Porter et al., 1974) may suffer from contamination or deficiency and present threats to construct validity and thus contribute to inconsistent findings. Second, theoretically, time is the medium through which causal processes unfold (e.g., causes take time to exert effects). The dynamics of causal relationships over time and the critical role of time in advancing strong and cumulative theories have been highlighted by the fundamental question of "when events occur, when they change, and how quickly they change" (Mitchell & James, 2001, p. 533). Finally, the timing of the investigation might be relevant to the ordering of JS and OC. For instance, during organizational entry, newcomers "learn the ropes" and acquire an understanding of the job and the organization (Wanous, 1992). That is, organizational entry is the critical time during which JS and OC forms. Thus, tracking newcomers over time allows for testing the ordering between JS and OC from their early formation. This becomes more crucial, when considering that the formation of JS and OC at different time points *during organizational entry* is the theoretical foundation of JS causing OC (Lawler, 1992; Vandenberg & Lance, 1992).

To date, there is no empirical synthesis disentangling artificial and substantive sources of inconsistencies in the ordering between JS and OC. Thus, we use traditional and full information meta-analytic structural equation modeling (MASEM and FIMASEM) to synthesize existing panel studies ($k = 71$, $N = 16{,}698$) to (a) correct for statistical artifacts and quantify the *true* variation in the ordering of JS and OC, (b) explore time lag, conceptualizations and/or measurements of OC and JS, and the timing of the investigation (e.g., newcomer status) that may explain some true variation if there is any, and (c) examine whether different specifications of the ordering between JS and OC influence the conclusions of the cross-lagged relationships among JS, OC, and turnover intentions (TI). Thus, we provide the first comprehensive integration of the fragmented literature, disentangle true and artificial sources of variation, and answer repeated calls for moving beyond the direct ordering between JS and OC to explore its contingencies with multiple, diverse samples (e.g., Vandenberg & Lance, 1992). Doing so will guide theory development and future research in different directions (cf. Aguinis et al., 2011; Schmidt & Hunter, 2015). Specifically, if there were substantial true variation, existing theories exclusively focusing on the *direct* ordering would not be generalizable, and a need for moderator-based theories and research focusing on moderators would be vindicated. In contrast, the lack of true variation would suggest that *one* theoretical perspective should be selected over other competing perspectives (i.e., *theory pitting and pruning*, Leavitt et al., 2010; Vandenberg & Grelle, 2008), and researchers should specify the ordering consistently across studies to remove "unnecessary" confusion in the literature and prevent downstream effects of model misspecification. Collectively, identifying sources of inconsistencies can significantly advance theories and research, as it helps test the generalizability of theoretical propositions; select a superior theory over alternatives; expand, correct, or overturn existing theories in the light of new findings, such as the boundaries of a theory (cf. Leavitt et al., 2010; Vandenberg & Grelle, 2008). Further, by showcasing how different orderings between JS and OC influence their cross-lagged relationships with TI, we demonstrate downstream effects of misspecifying the ordering between JS and OC on theory and research. In short, our meta-analysis provides improved focus on theory and research by looking forward and backward in the literature.

Competing Theoretical Models

There are four competing models regarding the ordering of OC and JS. The first model, based on affective attachment theory, proposes that JS forms sooner than OC and thus causes OC (Lawler, 1992). Specifically, with a micro-orientation, JS varies instantaneously due to changing work experiences and consequently forms soon after organizational entry. In contrast, due to its macro-orientation, OC develops gradually, after one acquires a firm understanding of organizational cultures, values, and the implications of maintaining organizational membership. Thus, OC forms sometime after organizational entry with JS as one determinant. The second model, grounded in cognitive dissonance theory, proposes that OC initiates rationalizing processes through which employees adjust JS levels to be consistent with OC levels in order to reduce cognitive dissonance (e.g., Vandenberg & Lance, 1992). Because the first and second models are both theoretically defensible and receive empirical support, the third model proposes that JS and OC have reciprocation. The final model proposes that their observed relation is due to shared antecedents (e.g., negative affectivity; Thoresen et al., 2003). Further, some scholars argue that OC and JS are indicators of a higher order factor or "A-factor" (e.g., Harrison et al., 2006). Essentially, propositions of shared antecedents and "A-factor" both suggest that a common factor(s) explains the observed relation between JS and OC (no causality; Finkel, 1995).

Potential Moderators Underlying the Inconsistencies

Conceptualizations and Measurements of OC

There is no consensus on the OC conceptualization (e.g., Klein et al., 2012; Meyer & Herscovitch, 2001). Porter et al. (1974; Mowday et al., 1979) defined OC as one's identification with and involvement in the organization and developed a unidimensional Organizational Commitment Questionnaire (OCQ). However, there are criticisms over its contamination with JS and TI (Bozeman & Perrewé, 2001). Because the "intended" OC is misrepresented, the theoretical grounds for the ordering of JS and OC may not hold for OCQ. For instance, the "JS" captured by OCQ could make OC appear to form sooner, and thus the ordering between JS and OC may not be consistent with the prediction of affective attachment theory. However, the "TI" captured by OCQ could make OC appear

to be the outcome of JS. Taken together, the overlap between OCQ and JS/TI (i.e., construct-irrelevant variance) should influence the proportion of "valid" variance in JS and OC to be explained and thus distort the cross-lagged relationship between JS and OC.

Meyer and Herscovitch (2001) defined OC as a force binding an employee to a course of action relevant to the organization, including affective, normative, and continuance commitment (AC, NC, and CC). AC refers to one's emotional attachment to and identification with the organization; CC is one's awareness of the associated costs when leaving the organization; NC reflects one's felt obligation to the organization (Meyer & Allen, 1991). Despite being the dominant view of OC (Meyer & Morin, 2016), this model is criticized by the AC–NC convergence and the irrelevance of CC (e.g., Bergman, 2006; Klein et al., 2012; Solinger et al., 2008). However, recent person-centered research supports the AC–NC distinction and the relevance of CC (for a review, see Meyer & Morin, 2016). For instance, commitment profiles with distinctive levels of AC and NC are detected and have nuanced relationships with other variables (e.g., Kabins et al., 2016; Kam et al., 2016; Meyer & Morin, 2016).

Klein et al. (2012) defined commitment as a psychological bond reflecting one's volitional dedication to and responsibility for a target. Klein et al. (2012) argued that commitment has the same meaning across different targets such as organization, supervisor, team (i.e., a target-free approach). Subsequently, Klein et al. (2014) developed a four-item *u*nidimensional, *t*arget-free commitment measure (KUT). Since its development, KUT has received considerable attention and been translated into different languages (Klein, 2022).

Nuanced differences in OC conceptualizations might influence the ordering between JS and OC. First, NC could develop before organizational entry through familial and cultural socialization (Cohen, 2007; Meyer & Allen, 1997). Thus, contradicting to the prediction of affective attachment theory, JS forming later than NC, may not cause NC. In contrast, AC, OCQ, and KUT reflecting emotional attachment to, involvement in, or dedication to and responsibility for the organization require firm understanding of the organization (e.g., its cultures, values), and CC reflects an accumulation of investments—all could only be built up over time (Klein et al., 2012; Meyer & Herscovitch, 2001; Porter et al., 1974). Thus, JS may act as their determinant. Second, according to cognitive dissonance theory, freedom of choice is a necessary condition for dissonance arousal to occur (Festinger, 1957). Individuals without free choices will not experience dissonance arousal and thus no pressure to reduce cognitive dissonance. CC, a need to be committed to the organization, can be construed as an absence of free choice. Thus, CC may not elicit pressure for initiating rationalizing processes to adjust JS levels. In contrast, because AC, OCQ, KUT, and NC all reflect one's freedom of choice (e.g., *voluntary* attachment to, or dedication to the organization due to value congruence, or felt obligation due to personal exchange orientation and willingness to repay), they can elicit pressure to reduce cognitive dissonance and adjust JS to be commensurate with OC levels.

Conceptualizations and Measurements of JS

Scholars have conceptualized JS as an emotional state, a cognitive evaluation, or both. For example, Locke (1969, p. 317) defined JS as a "pleasurable *emotional state* resulting from the appraisal of one's job as achieving or facilitating one's job values," while Weiss (2002, p. 175) defined JS as "a positive (or negative) *evaluative judgment* one makes about one's job or job situation." Yet, Judge et al. (2017, p. 357) defined JS as "the assessment of the favorability of a job" and "containing both *cognitive and affective* elements." However, the affective element is missing from most JS measures (cf. Judge et al., 2020). Most importantly, it is theoretically and methodologically difficult to separate cognition and affect measures (cf. Judge et al., 2017, 2020). Thus, we are not able to address whether JS conceptualizations influence the ordering of JS and OC due to misalignments between construct conceptualizations and operationalizations.

JS can be measured by composite and global measures. Composite measures use the sum of ratings with multiple facets (e.g., Minnesota Satisfaction Questionnaire, Weiss et al., 1967). Global measures use globally worded items (e.g., Michigan Organizational Assessment Questionnaire, Cammann et al., 1983). Research suggests that composite JS measures cannot capture all aspects of a job and thus miss important JS variance (e.g., Dalal, 2013; Ironson et al., 1989). That is, composite measures may fail to capture the full domain of JS or misrepresent JS (i.e., measure deficiency). Statistically, deficient measures miss some valid variance in JS that may explain OC at a later time point and/or be explained by OC at an earlier time point (i.e., *attenuated* cross-lagged relationships). That is, the underrepresentation of JS may lead to a disconnection between theory and theory testing, and distort the ordering between JS and OC.

Time Lag

Time is a critical element for determining causal inferences (Mitchell & James, 2001). Specifically, when time lags are too short, the effect may not have sufficient time to occur; when time lags are too long, the effect may vanish due to individual adaptation or intervening variables. Further, a relationship may change over time. For instance, cause (effect) may become effect (cause) over time (e.g., Taris & Kompier, 2014). Thus, existing "competing" models may not be conflicting with one another because each may capture a snapshot of the *evolving* relationship between JS and OC over time. That is, each model may be valid during a specific time period such that the effect(s) of JS and/or OC may be observed as nonexistent, weaker, or stronger over the course of time, contributing to inconsistencies across studies.

Newcomer Status

Affective attachment theory suggests that JS forms sooner than OC after organizational entry (Lawler, 1992; Vandenberg & Lance, 1992). Therefore, tracking newcomers from organizational entry may provide "a theoretically meaningful starting point" for testing the ordering of JS and OC (cf. Vandenberg & Lance, 1992). Perhaps the effect of JS on OC occurs soon after organizational entry. Yet, long after organizational entry when newcomers become "insiders," OC forms, stabilizes, and may exert a reversed effect on JS via rationalization processes. That is, the proposition of affective attachment theory may be valid only for newcomers, whereas the proposition of cognitive dissonance theory may be valid for established employees. Thus, comparing newcomers with established employees may help address the inconsistencies in the ordering between JS and OC across studies (Vandenberg & Lance, 1992).

OC, JS, and TI

We test whether different orderings between OC and JS change the cross-lagged relationships among OC, JS, and TI. This will provide straightforward evidence that inaccurately specifying the ordering has implications on the validity of theories and models with both JS and OC as causes, mediators, or outcomes of various phenomena (e.g., performance, deviance, sleep).

Method

Transparency and Openness

We describe our search strategy, inclusion criteria, and analytical approaches. We adhered to the *Journal of Applied Psychology* methodological checklist. The summary of included studies, meta-analytic codes, coding procedures, average reliabilities, and additional analyses are available at (https://osf.io/5dg8b/?view_only=35d31d570424498ca2f312cafce12153). Meta-analyses were conducted with Microsoft Excel based on Schmidt and Hunter (2015). We ran traditional MASEMs with Mplus 8.6 (Muthén & Muthén, 2017) and FIMASEMs based on the codes adapted from Cheung (2018) and Yu et al. (2016) with R 4.2.0 for MacOS. The meta-analytic scope, design, and analyses were not preregistered. Additional materials and syntaxes are available upon request from the first author.

Literature Search

We conducted an extensive literature search for relevant articles. First, we conducted a search in PsycINFO, Abstracted Business Information/Inform, The Institute for Scientific Information Web of Science, ProQuest Dissertation Abstracts, and Google Scholar using broad keywords: combining commitment AND satisfaction with cross-lagged, or wave, or longitudinal, or panel, or experimental, or intervention for studies available before April 2022. Second, we examined the reference sections of relevant meta-analyses (i.e., Judge et al., 2001; Kabins et al., 2016; Meyer et al., 2002; Riketta, 2008). Third, we examined all available conference programs for the Society for Industrial and Organizational Psychology (2004–2021) and Academy of Management (1998–2021). Fourth, the authors of studies that met inclusion criteria but failed to report the complete correlation matrix were contacted for missing information. Finally, the references of each relevant article were examined for additional studies.

Inclusion Criteria and Features of the Studies Included

Studies must meet three criteria to be included: (a) participants must be employees, including employed students (e.g., working MBAs); (b) we only included individual-level relationships (i.e., no group-level relationships); (c) the zero-order correlation matrix of JS and OC is available for at least two waves. This yielded 71 independent samples ($N = 16{,}698$) with two repeated measurements of JS and OC, and 31 independent samples ($N = 6{,}876$) with two repeated measurements of JS, OC, and TI (Supplemental Table S2). Six studies tested the temporal ordering of JS and OC, but none of them provided a definitive test with experimental designs.

Data Analyses

Because CC, NC, and KUT are conceptually different from AC and OCQ, the overall analysis and moderation analyses included only AC and OCQ to rule out confounds among OC conceptualizations.[3] Time lags were divided into three subgroups (i.e., ≤1 month: $M = 0.76$, $SD = 0.41$; >1 month and ≤6 months: $M = 4.43$, $SD = 1.43$; >6 months: $M = 15.30$, $SD = 12.54$) by making a compromise between two conflicting goals (cf. Riketta, 2008): (a) a reasonable number of samples in each subgroup and (b) different time lags for detecting nonmonotonic effects. JS measures and newcomer status were divided into two subgroups, respectively (i.e., composite vs. global JS, newcomers vs. nonnewcomers).

We meta-analyzed correlations using Schmidt and Hunter's (2015) random-effects, individual-correction method by correcting for unreliability in JS, OC, and TI. Then, following scholars' recommendation (Cheung, 2018; Oh, 2020), we ran traditional MASEMs to obtain cross-lagged path and standard error estimates using meta-analytic correlation (i.e., ρ) matrices. Specifically, we tested three alternative models: Model 1 with freely estimated cross-lagged paths between JS and OC, Model 2 with these path estimates constrained to be equal, and Model 3 with JS and OC as indicators of a higher order or common factor. This allowed us to test whether the cross-lagged paths between JS and OC were equal or not and whether JS and OC should be treated as indicators of a higher order or common factor. Models were compared (Hu & Bentler, 1999) using chi-squared difference tests along with comparative fit index (CFI > .95), root-mean-square error of approximation (RMSEA < .06), and standardized root-mean-square residual (SRMR < .08). Finally, we utilized FIMASEM with 10,000 bootstrap samples (Yu et al., 2016) to estimate the heterogeneity of effect sizes (i.e., 80% credibility intervals [CVs]) using matrices of ρ and SD_ρ *while accounting for random-effects covariance* (Cheung et al., 2019). A width of 80% CVs < 0.18 suggests small heterogeneity, between 0.18 and 0.54 indicates moderate heterogeneity, and >0.54 indicates large heterogeneity (Bosco et al., 2015).

Results

Table 1 displays the mean corrected correlations between JS and OC. Overall and across different moderation levels (including the models involving TI), the models with JS and OC loaded on a higher order or common factor had the worst fit, suggesting that JS and OC were distinct constructs (Supplemental Tables S5 and S7). Overall (Table 2), there was a reciprocation between JS and OC such that the lagged effect of OC on JS was not significantly different from that of JS on OC (βs = 0.06, $SE = .004$, $p < .05$). Most importantly, the 80% CV widths of these lagged effects were both 0.32 (moderate heterogeneity), suggesting that there were moderators (Table 2). Figures 1 and 2 show path estimates of the final models for the overall analysis and across different subgroups. Across different subgroups (Table 2), the 80% CV widths ranged from .18 to .52 (small to moderate), suggesting that there may be other moderators beyond the ones examined in the current meta-analysis.

Conceptualizations and Measurements of OC

The lagged effect of AC on JS (β = 0.10, $SE = 0.01$, $p < .05$) was stronger than that of JS on AC (β = 0.04, $SE = 0.01$, $p < .05$). There was a lagged effect of NC on JS (β = 0.14, $SE = 0.02$, $p < .05$) but no effect of JS on NC (β = −0.001, $SE = 0.02$, $p = .97$). No

[3] We thank an anonymous reviewer for this suggestion.

Table 1

Meta-Analytic Results for the Autocorrelations, Synchronous, and Cross-Lagged Correlations of Job Satisfaction and Organizational Commitment

Correlations	N	k	$\bar{r}$	SD_r	ρ	SD_ρ	CV_{10}	CV_{90}	CI_L	CI_U	% var.
Overall (AC and OCQ only)											
JS1–JS2	15,084	62	0.55	0.20	0.66	0.21	0.38	0.93	0.64	0.67	5.89
OC1–OC2	15,084	62	0.60	0.20	0.71	0.21	0.43	0.98	0.69	0.72	4.83
JS1–OC1	15,084	62	0.55	0.16	0.66	0.17	0.45	0.88	0.65	0.68	9.42
JS2–OC2	15,084	62	0.57	0.17	0.68	0.16	0.47	0.89	0.66	0.69	8.56
JS1–OC2	15,084	62	0.42	0.17	0.50	0.18	0.28	0.73	0.49	0.52	11.28
OC1–JS2	15,084	62	0.39	0.19	0.47	0.20	0.22	0.72	0.45	0.49	9.70
OCQ											
JS1–JS2	5,595	25	0.55	0.16	0.66	0.18	0.43	0.89	0.64	0.68	8.97
OC1–OC2	5,595	25	0.58	0.16	0.66	0.18	0.43	0.89	0.64	0.68	7.26
JS1–OC1	5,595	25	0.51	0.14	0.61	0.16	0.40	0.82	0.59	0.63	11.72
JS2–OC2	5,595	25	0.54	0.13	0.62	0.15	0.42	0.81	0.59	0.64	11.55
JS1–OC2	5,595	25	0.39	0.13	0.45	0.14	0.28	0.63	0.43	0.48	18.76
OC1–JS2	5,595	25	0.35	0.14	0.40	0.15	0.21	0.60	0.37	0.43	16.78
Affective commitment											
JS1–JS2	9,888	38	0.56	0.23	0.67	0.24	0.36	0.97	0.65	0.68	4.33
OC1–OC2	9,888	38	0.62	0.21	0.74	0.23	0.45	1.03	0.73	0.76	3.92
JS1–OC1	9,888	38	0.58	0.17	0.70	0.16	0.50	0.90	0.68	0.71	9.00
JS2–OC2	9,888	38	0.60	0.18	0.72	0.16	0.51	0.92	0.70	0.73	8.12
JS1–OC2	9,888	38	0.45	0.20	0.54	0.19	0.29	0.79	0.52	0.56	8.60
OC1–JS2	9,888	38	0.43	0.20	0.52	0.21	0.26	0.78	0.50	0.54	7.85
Normative commitment											
JS1–JS2	1,815	11	0.70	0.14	0.76	0.12	0.60	0.92	0.74	0.79	11.16
OC1–OC2	1,815	11	0.69	0.08	0.86	0.08	0.76	0.96	0.83	0.89	31.17
JS1–OC1	1,815	11	0.54	0.20	0.64	0.20	0.38	0.90	0.60	0.68	9.59
JS2–OC2	1,815	11	0.58	0.17	0.68	0.17	0.47	0.89	0.64	0.71	11.59
JS1–OC2	1,815	11	0.47	0.22	0.55	0.22	0.26	0.83	0.51	0.59	9.26
OC1–JS2	1,815	11	0.48	0.16	0.57	0.15	0.38	0.77	0.53	0.61	18.07
Continuance commitment											
JS1–JS2	1,761	12	0.73	0.13	0.82	0.10	0.69	0.94	0.79	0.84	16.38
OC1–OC2	1,761	12	0.62	0.12	0.77	0.13	0.60	0.95	0.74	0.81	18.49
JS1–OC1	1,761	12	−0.07	0.29	−0.10	0.32	−0.51	0.31	−0.15	−0.04	8.56
JS2–OC2	1,761	12	−0.04	0.28	−0.06	0.31	−0.46	0.34	−0.12	−0.01	8.99
JS1–OC2	1,761	12	−0.05	0.24	−0.08	0.26	−0.41	0.26	−0.13	−0.02	12.41
OC1–JS2	1,761	12	−0.06	0.26	−0.08	0.29	−0.46	0.29	−0.14	−0.03	10.11
KUT (Klein et al., 2012)											
JS1–JS2	1,014	4	0.75	0.08	0.82	0.06	0.74	0.90	0.79	0.85	19.49
OC1–OC2	1,014	4	0.69	0.06	0.75	0.06	0.67	0.83	0.71	0.78	23.29
JS1–OC1	1,014	4	0.65	0.11	0.72	0.12	0.56	0.87	0.68	0.75	9.32
JS2-OC2	1,014	4	0.69	0.06	0.75	0.05	0.69	0.81	0.72	0.79	37.60
JS1–OC2	1,014	4	0.58	0.11	0.64	0.09	0.51	0.76	0.59	0.68	19.03
OC1–JS2	1,014	4	0.61	0.10	0.66	0.10	0.53	0.79	0.62	0.70	15.90
Global measures of JS (AC and OCQ only)											
JS1–JS2	8,711	42	0.56	0.23	0.65	0.24	0.34	0.96	0.64	0.67	4.91
OC1–OC2	8,711	42	0.61	0.23	0.70	0.25	0.38	1.01	0.68	0.71	3.97
JS1–OC1	8,711	42	0.60	0.16	0.70	0.17	0.48	0.93	0.69	0.72	7.98
JS2–OC2	8,711	42	0.65	0.15	0.74	0.15	0.55	0.93	0.72	0.75	8.79
JS1–OC2	8,711	42	0.46	0.19	0.53	0.20	0.27	0.79	0.51	0.55	9.05
OC1–JS2	8,711	42	0.44	0.20	0.51	0.21	0.25	0.78	0.50	0.53	8.88
Composite measures of JS (AC and OCQ only)											
JS1–JS2	3,812	18	0.61	0.15	0.72	0.16	0.52	0.93	0.70	0.75	9.26
OC1–OC2	3,812	18	0.66	0.12	0.76	0.15	0.57	0.95	0.74	0.78	8.41
JS1–OC1	3,812	18	0.51	0.14	0.61	0.14	0.43	0.79	0.58	0.64	15.87
JS2–OC2	3,812	18	0.53	0.12	0.61	0.13	0.44	0.78	0.58	0.63	15.91
JS1–OC2	3,812	18	0.41	0.12	0.49	0.12	0.33	0.64	0.46	0.52	23.85
OC1–JS2	3,812	18	0.37	0.16	0.43	0.17	0.20	0.65	0.39	0.46	13.62

(*table continues*)

Table 1 (*continued*)

Correlations	N	k	$\bar{r}$	SD_r	ρ	SD_ρ	CV_{10}	CV_{90}	CI_L	CI_U	% var.
Time lag: ≤1 month (AC and OCQ only)											
JS1–JS2	2,264	10	0.75	0.16	0.82	0.17	0.60	1.04	0.80	0.84	3.49
OC1–OC2	2,264	10	0.73	0.18	0.82	0.21	0.55	1.08	0.79	0.84	2.81
JS1–OC1	2,264	10	0.70	0.14	0.78	0.15	0.60	0.97	0.76	0.81	6.39
JS2–OC2	2,264	10	0.74	0.11	0.82	0.10	0.69	0.95	0.80	0.84	10.13
JS1–OC2	2,264	10	0.62	0.14	0.70	0.14	0.51	0.88	0.67	0.72	9.02
OC1–JS2	2,264	10	0.61	0.16	0.68	0.17	0.46	0.91	0.65	0.71	6.78
Time lag: >1 month and ≤6 months (AC and OCQ only)											
JS1–JS2	6,399	32	0.57	0.19	0.67	0.21	0.41	0.94	0.65	0.69	6.87
OC1–OC2	6,399	32	0.65	0.15	0.75	0.17	0.53	0.96	0.73	0.76	7.40
JS1–OC1	6,399	32	0.54	0.16	0.64	0.18	0.41	0.87	0.62	0.66	10.24
JS2–OC2	6,399	32	0.58	0.17	0.67	0.19	0.42	0.91	0.65	0.68	7.72
JS1–OC2	6,399	32	0.43	0.16	0.50	0.16	0.29	0.71	0.48	0.53	14.97
OC1–JS2	6,399	32	0.41	0.16	0.47	0.18	0.24	0.70	0.45	0.49	13.19
Time lag: >6 months (AC and OCQ only)											
JS1–JS2	6,421	20	0.45	0.17	0.56	0.19	0.32	0.80	0.54	0.59	7.87
OC1–OC2	6,421	20	0.51	0.20	0.61	0.23	0.33	0.90	0.59	0.64	4.71
JS1–OC1	6,421	20	0.51	0.14	0.64	0.14	0.46	0.81	0.62	0.66	12.27
JS2–OC2	6,421	20	0.52	0.14	0.63	0.12	0.48	0.78	0.61	0.65	14.59
JS1–OC2	6,421	20	0.34	0.13	0.42	0.13	0.25	0.59	0.39	0.44	17.24
OC1–JS2	6,421	20	0.31	0.14	0.38	0.15	0.19	0.57	0.35	0.41	14.50
Newcomers (AC and OCQ only)											
JS1–JS2	1,382	9	0.46	0.13	0.61	0.11	0.47	0.75	0.56	0.66	34.88
OC1–OC2	1,382	9	0.56	0.16	0.66	0.19	0.41	0.90	0.61	0.70	10.48
JS1–OC1	1,382	9	0.56	0.10	0.72	0.12	0.56	0.87	0.67	0.76	26.05
JS2–OC2	1,382	9	0.64	0.11	0.77	0.11	0.63	0.91	0.73	0.81	21.29
JS1–OC2	1,382	9	0.37	0.07	0.47	0.00	0.47	0.47	0.41	0.53	100
OC1–JS2	1,382	9	0.35	0.07	0.43	0.04	0.38	0.48	0.37	0.49	82.49
Non-newcomers (AC and OCQ only)											
JS1–JS2	13,702	53	0.56	0.21	0.66	0.22	0.38	0.94	0.65	0.67	5.08
OC1–OC2	13,702	53	0.61	0.20	0.71	0.22	0.43	0.99	0.70	0.72	4.43
JS1–OC1	13,702	53	0.55	0.17	0.66	0.17	0.44	0.87	0.65	0.67	8.58
JS2–OC2	13,702	53	0.57	0.17	0.67	0.17	0.46	0.88	0.65	0.68	8.06
JS1–OC2	13,702	53	0.42	0.18	0.50	0.18	0.27	0.74	0.49	0.52	9.68
OC1–JS2	13,702	53	0.40	0.19	0.47	0.20	0.21	0.73	0.46	0.49	8.41

Note. JS = job satisfaction; OC = organizational commitment; AC = Meyer and Allen's affective commitment; OCQ = Organizational Commitment Questionnaire; KUT = Klein et al.'s (2012) target-free unidimensional measure of organizational commitment; 1 = Time 1; 2 = Time 2; $\bar{r}$ = mean sample size-weighted r value; SD_r = sample size-weighted observed standard deviation of r value; ρ = r corrected for unreliability; SD_ρ = standard deviation of ρ; CV_{10} and CV_{90} = 10% and 90% credibility values, respectively; CI_L and CI_U = lower and upper bounds, respectively, of the 95% confidence interval around the corrected mean correlation; % var. = percentage of variance attributable to sampling error.

cross-lagged relationship existed between CC and JS (βs = 0.00, SE = 0.01, p = .98). There was a lagged effect of JS on OCQ (β = 0.08, SE = 0.01, p < .05) but no lagged effect of OCQ on JS (β = −0.004, SE = 0.01, p = .74). A symmetric, reciprocal relationship existed between KUT and JS (βs = 0.17, SE = 0.02, p < .05).

Measurements of JS

There was a lagged effect of composite JS on OC (β = 0.04, SE = 0.01, p < .05) but no effect of OC on composite JS (β = −0.02, SE = 0.01, p = .30), whereas there was a reciprocal relationship between global JS and OC (βs = 0.09, SE = 0.01, p < .05).

Time Lag

There was a reciprocal relationship between JS and OC (βs = 0.13, SE = 0.01, p < .05) for time lags ≤1 month. However, the lagged effects of JS and OC decreased for time lags longer than 1 month (βs = 0.05, SE = 0.01, p < .05) and time lags longer than 6 months (βs = 0.04, SE = 0.01, p < .05).

Newcomer Status

There was no cross-lagged relationship between JS and OC for newcomers (βs = −0.02, SE = 0.004, p = .27), whereas there was a reciprocal relationship between JS and OC for nonnewcomers (βs = 0.06, SE = 0.004, p < .05).

Turnover Intentions

Table 3 shows the mean corrected correlations among OC, JS, and TI. As shown in Table 4, when the cross-lagged paths between JS and OC were freely estimated, JS and OC had the same lagged effect on TI (βs = −0.06, SE = 0.01, p < .05). When the cross-lagged paths between JS and OC were constrained to be equal, their lagged effects on TI were different ($\beta_{JS \rightarrow TI}$ = −0.05, SE = 0.01, p < .05; $\beta_{OC \rightarrow TI}$ = −0.08, SE = 0.01, p < .05).[4] However,

[4] Our findings differed from Tett and Meyer's (1993) cross-sectional meta-analysis (see Supplemental Materials).

Table 2

Estimates of the Cross-Lagged Relationships Between Organizational Commitment and Job Satisfaction for the Final Models Based on the Traditional MASEM and FIMASEM Results

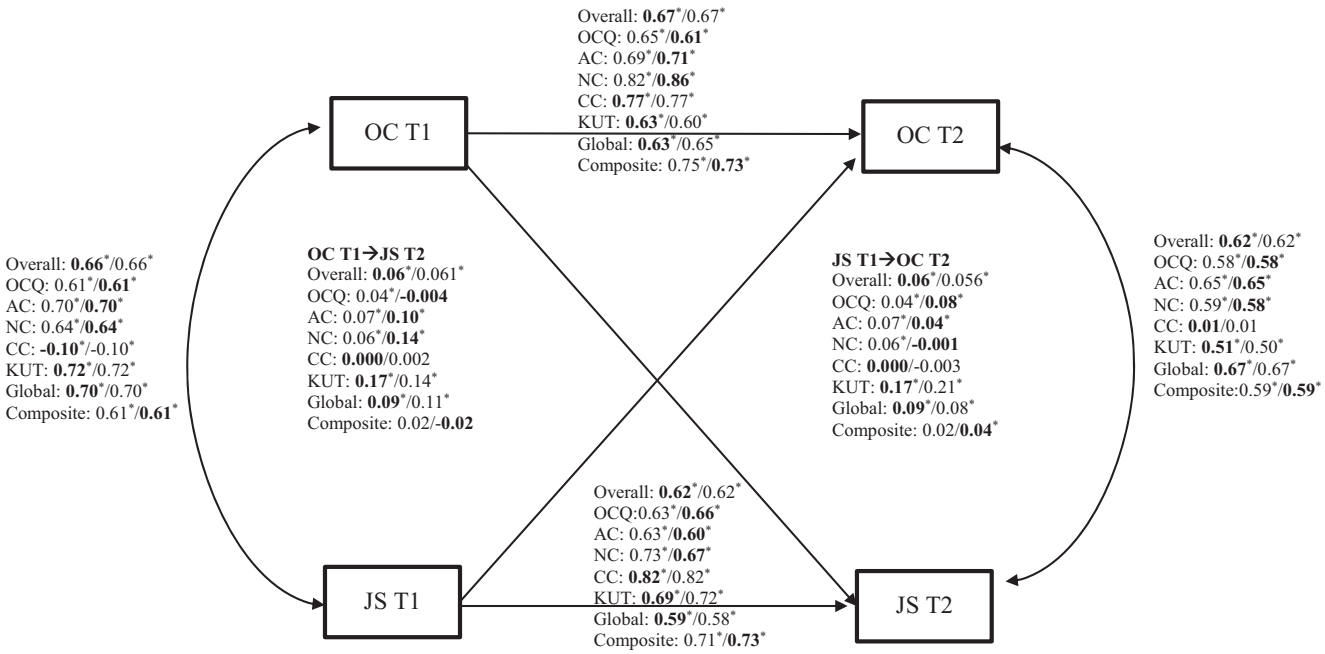
Cross-lagged paths	β	*SE*	SD_β	80% CV_β	80% CV_β width
			Overall (AC and OCQ only)		
JS1 → OC2	0.06*	0.004	0.14	−0.10, 0.22	0.32
OC1 → JS2	0.06*	0.004	0.14	−0.10, 0.22	0.32
			OCQ		
JS1 → OC2	0.08*	0.01	0.25	−0.13, 0.32	0.45
OC1 → JS2	−0.004	0.01	0.25	−0.23, 0.23	0.46
			Affective commitment		
JS1 → OC2	0.04*	0.01	0.24	−0.19, 0.33	0.52
OC1 → JS2	0.10*	0.01	0.16	−0.06, 0.26	0.32
			Normative commitment		
JS1 → OC2	−0.001	0.02	0.23	−0.16, 0.19	0.35
OC1 → JS2	0.14*	0.02	0.12	0.06, 0.24	0.18
			Continuance commitment		
JS1 → OC2	0.000	0.01	0.11	−0.10, 0.11	0.21
OC1 → JS2	0.000	0.01	0.11	−0.10, 0.11	0.21
			KUT		
JS1 → OC2	0.17*	0.02	0.12	0.06, 0.26	0.20
OC1 → JS2	0.17*	0.02	0.12	0.06, 0.26	0.20
			Global measures of JS (AC and OCQ only)		
JS1 → OC2	0.09*	0.01	0.16	−0.09, 0.27	0.36
OC1 → JS2	0.09*	0.01	0.16	−0.09, 0.27	0.36
			Composite measures of JS (AC and OCQ only)		
JS1 → OC2	0.04*	0.01	0.19	−0.14, 0.22	0.36
OC1 → JS2	−0.02	0.01	0.19	−0.21, 0.18	0.39
			Time lag: ≤1 month (AC and OCQ only)		
JS1 → OC2	0.13*	0.01	0.16	−0.01, 0.36	0.37
OC1 → JS2	0.13*	0.01	0.16	−0.01, 0.36	0.37
			Time lag: >1 month and ≤6 months (AC and OCQ only)		
JS1 → OC2	0.05*	0.01	0.15	−0.09, 0.20	0.29
OC1 → JS2	0.05*	0.01	0.15	−0.09, 0.20	0.29
			Time lag: >6 months (AC and OCQ only)		
JS1 → OC2	0.04*	0.01	0.13	−0.09, 0.19	0.28
OC1 → JS2	0.04*	0.01	0.13	−0.09, 0.19	0.28
			Nonnewcomers (AC and OCQ only)		
JS1 → OC2	0.06*	0.004	0.15	−0.11, 0.22	0.33
OC1 → JS2	0.06*	0.004	0.15	−0.11, 0.22	0.33
			Newcomers (AC and OCQ only)		
JS1 → OC2	−0.02	0.01	0.14	−0.11, 0.22	0.33
OC1 → JS2	−0.02	0.01	0.14	−0.11, 0.22	0.33

Note. Standardized path coefficients (i.e., β) and the standard error of the path coefficients (i.e., *SE*) are from the traditional MASEM results, while the distribution of the path coefficients (i.e., SD_β, 80% CV_β, and 80% CV_β width) is from Yu et al.'s (2016) FIMASEM results—bootstrapping simulation technique with 10,000 samples while accounting for random-effects covariance among the variables. JS = job satisfaction; OC = organizational commitment; AC = Meyer and Allen's affective commitment; OCQ = Organizational Commitment Questionnaire; 1 = Time 1; 2 = Time 2; SD_β = standard deviation of the distribution of β coefficients; 80% CV = 80% credibility interval; MASEM and FIMASEM = full information meta-analytic structural equation modeling.
* $p < .05$.

because these two models had significant lagged effects of both JS and OC, they represented the same ordering (either symmetric or asymmetric reciprocation). That is, comparing these two did not allow for testing the implications of varying orderings. Thus, we tested three additional models, in which the lagged effect of JS or/and OC was fixed to be zero to represent the three alternatives in the literature. Different orderings between JS and OC resulted in *differential* lagged effects of JS and OC on TI (i.e., OC → JS: $\beta_{JS \to TI} = -0.004$, $SE = 0.01$, $p = .73$; $\beta_{OC \to TI} = -0.09$, $SE = 0.01$, $p < .05$; JS → OC: $\beta_{JS \to TI} = -0.08$, $SE = 0.01$, $p < .05$;

Figure 1

Cross-Lagged Relationships Between Job Satisfaction and Organizational Commitment for the Overall Analysis and Different Measurements of Organizational Commitment and Job Satisfaction



Note. JS = job satisfaction; OC = organizational commitment; T1 = Time 1; T2 = Time 2; AC = affective commitment; NC = normative commitment; CC = continuance commitment; OCQ = Organizational Commitment Questionnaire; KUT = Klein et al.'s (2012) target-free unidimensional measure of organizational commitment; Global = global measures of job satisfaction; Composite = composite measures of job satisfaction. Standardized path coefficients reported before the/were for the model with the cross-lagged paths between JS and OC constrained to be equal, whereas path coefficient after the/were for the model with the freely estimated cross-lagged paths between JS and OC. Coefficients in bold are the ones for the final models.

* $p < .05$.

$\beta_{OC \rightarrow TI} = -0.03$, $SE = 0.01$, $p < .05$; no causality: $\beta_{JS \rightarrow TI} = -0.02$, $SE = 0.01$, $p = .19$; $\beta_{OC \rightarrow TI} = -0.05$, $SE = 0.01$, $p < .05$), and *differential* lagged effects of TI on JS and OC (e.g., OC → JS: $\beta_{TI \rightarrow JS} = -0.03$, $SE = 0.01$, $p < .05$; JS → OC: $\beta_{TI \rightarrow JS} = -0.02$, $SE = 0.01$, $p = .07$).

Discussion

After correcting for statistical artifacts, there is moderate heterogeneity in the lagged effects of JS and OC, suggesting the existence of moderators. Indeed, our meta-analysis indicates that conceptualizations and/or measurements of OC and JS, time lag, and the timing of the investigation (e.g., newcomer status) moderate the temporal ordering of JS and OC.

Theoretical Implications

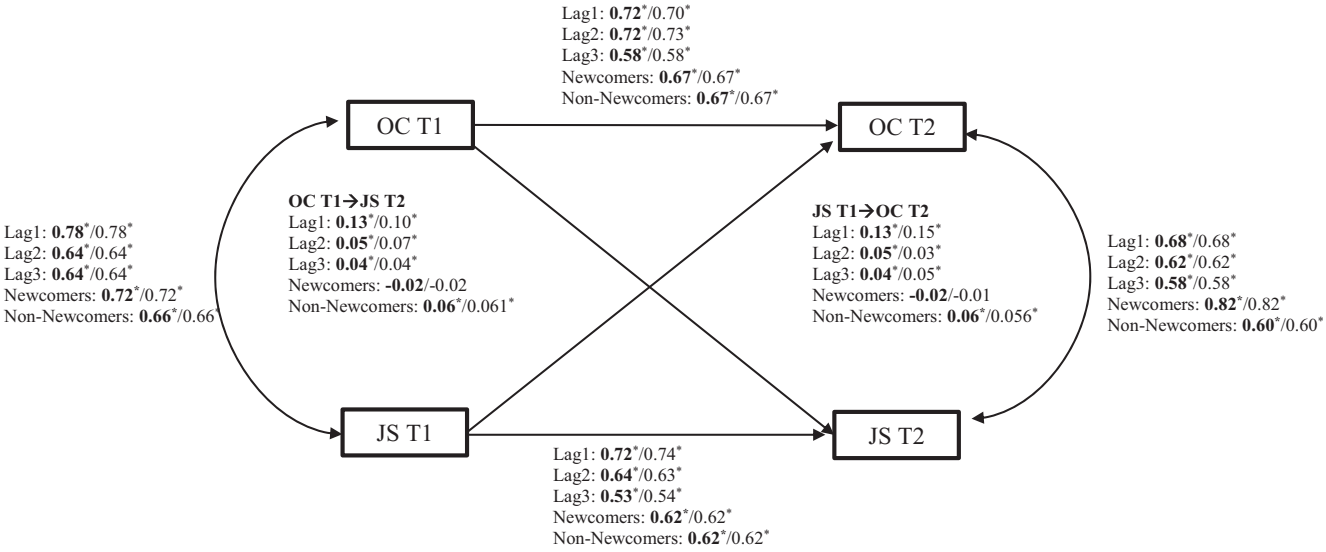
Measurement issues (i.e., deficiency and contamination) influence the ordering of JS and OC, suggesting that poor construct operationalization introduces inconsistent findings, creates a disconnection between theory and theory testing, and hinders theory development and knowledge accumulation (Edwards & Bagozzi, 2000). First, there is a lagged effect of OC on global JS but not composite JS, and the lagged effect of global JS on OC is larger than that of composite JS. One explanation for these discrepancies is that the cross-lagged relationship between composite JS and OC is deflated. Research suggests that composite measures may miss relevant facets (e.g., Highhouse & Becker, 1993; Ironson et al., 1989). The "deflated" cross-lagged relation between composite JS and OC is consistent with the proposition that composite measures may not capture the full domain of JS and miss unique variance in JS that could be accounted by OC at an earlier time point and/or explain the variance in OC at a later time point (e.g., Highhouse & Becker, 1993). Indeed, effect size deflation could be evidence for measure deficiency. However, one alternative explanation is that the cross-lagged relation between global JS and OC is inflated.[5] Global JS measures may prompt employees to include their thoughts and feelings about the organization (e.g., satisfaction with the organization) when evaluating JS, resulting in an "inflated" relationship with OC. Closely examining all JS measures in our meta-analysis reveals that job description form (JDF; Neuberger & Allerbeck, 1978) includes an item *explicitly* assessing satisfaction with the organization itself. In contrast, global measures do not have items *explicitly* assessing satisfaction with the organization and thus may or may not prompt employees to include satisfaction with the organization when evaluating their JS. That is, including satisfaction with the organization may be a possibility for global measures but a certainty for JDF. Thus, the inflation might be larger for JDF than global measures. However, the supplementary analyses[6] for JDF ($k = 3$)

[5] We thank an anonymous reviewer for raising this interesting point.

[6] See the additional analyses on JDF in the Supplemental Materials on pp. 21–22.

Figure 2

Cross-Lagged Relationships Between Job Satisfaction and Organizational Commitment for the Time Lag and the Timing of the Investigation



Note. JS = job satisfaction; OC = organizational commitment; T1 = Time 1; T2 = Time 2; Lag1: ≤1 month; Lag2: >1 month and ≤6 months; Lag3: >6 months. Standardized path coefficients reported before the/were for the model with the cross-lagged paths between JS and OC constrained to be equal, whereas path coefficient after the/were for the model with the freely estimated cross-lagged paths between JS and OC. Coefficients in bold are the ones for the final models.

* $p < .05$.

indicate that there is a lagged effect of OC on JDF but no lagged effect of JDF on OC. Thus, the (null) lagged effect of JDF on OC seems to be "deflated" compared to the (significant) lagged effect of global JS on OC. This suggests that even including satisfaction with the organization, composite JS measures might still suffer from measurement deficiency by missing important facets. Nevertheless, the finding of JDF should be interpreted with caution due to its small *k*.

Second, OCQ mispresents OC by including irrelevant variance of TI ("noise") which is demonstrated to have no lagged effect on JS (see Figure 2). Further, OCQ's contamination with JS increases the concurrent correlations between JS and OC across time, which

Table 3

Meta-Analytic Results for the Autocorrelations, Synchronous and Cross-Lagged Intercorrelations Among Job Satisfaction, Organizational Commitment, and Turnover Intentions

Correlations	*N*	*k*	$\bar{r}$	SD_r	ρ	SD_ρ	CV_{10}	CV_{90}	CI_L	CI_U	% var.
JS1–JS2	6,876	31	0.59	0.21	0.69	0.21	0.42	0.96	0.67	0.71	5.61
OC1–OC2	6,876	31	0.62	0.23	0.71	0.26	0.38	1.04	0.69	0.73	3.30
JS1–OC1	6,876	31	0.56	0.19	0.66	0.20	0.40	0.92	0.64	0.68	6.47
JS2–OC2	6,876	31	0.60	0.17	0.68	0.18	0.45	0.92	0.67	0.70	6.79
JS1–OC2	6,876	31	0.45	0.19	0.53	0.20	0.27	0.78	0.51	0.55	8.81
OC1–JS2	6,876	31	0.43	0.21	0.49	0.22	0.21	0.78	0.47	0.52	7.51
TI1–TI2	6,876	31	0.56	0.21	0.64	0.23	0.34	0.93	0.62	0.66	4.94
TI1–JS1	6,876	31	−0.50	0.24	−0.59	0.26	−0.91	−0.26	−0.61	−0.56	5.01
TI1–OC1	6,876	31	−0.45	0.22	−0.52	0.24	−0.83	−0.21	−0.54	−0.50	6.04
TI2–JS2	6,876	31	−0.49	0.30	−0.56	0.34	−0.99	−0.12	−0.58	−0.53	2.92
TI2–OC2	6,876	31	−0.47	0.25	−0.53	0.28	−0.89	−0.17	−0.55	−0.51	4.26
TI1–JS2	6,876	31	−0.36	0.23	−0.42	0.25	−0.73	−0.10	−0.44	−0.39	6.94
TI1–OC2	6,876	31	−0.34	0.23	−0.38	0.26	−0.71	−0.06	−0.41	−0.36	6.65
JS1–TI2	6,876	31	−0.37	0.21	−0.44	0.23	−0.74	−0.15	−0.46	−0.42	7.88
OC1–TI2	6,876	31	−0.35	0.19	−0.40	0.21	−0.67	−0.12	−0.42	−0.37	9.09

Note. JS = job satisfaction; OC = organizational commitment (*only AC and OCQ*); TI = turnover intentions; 1 = Time 1; 2 = Time 2; $\bar{r}$ = mean sample size-weighted *r* value; SD_r = sample size-weighted observed standard deviation of *r* value; ρ = *r* corrected for unreliability; SD_ρ = standard deviation of ρ; CV_{10} and CV_{90} = 10% and 90% credibility values, respectively; CI_L and CI_U = lower and upper bounds, respectively, of the 95% confidence interval around the corrected mean correlation; % var. = percentage of variance attributable to sampling error; AC = Meyer and Allen's affective commitment; OCQ = Organizational Commitment Questionnaire.

Table 4

Estimates of the Cross-Lagged Relationships Among Organizational Commitment, Job Satisfaction, and Turnover Intentions Based on the Traditional MASEM and FIMASEM Results

Cross-lagged paths	β	*SE*	SD_β	80% CV_β	80% CV_β width
The model with free estimates of the cross-lagged paths between JS and OC (asymmetric reciprocation)					
JS1 → OC2	0.12*	0.01	0.25	−0.09, 0.47	0.56
OC1 → JS2	0.06*	0.01	0.27	−0.13, 0.50	0.63
JS1 → TI2	−0.06*	0.01	0.26	−0.30, 0.31	0.61
OC1 → TI2	−0.06*	0.01	0.32	−0.68, 0.13	0.81
TI1 → JS2	−0.008	0.01	0.35	−0.49, 0.35	0.84
TI1 → OC2	0.03*	0.01	0.31	−0.47, 0.25	0.72
The model with equal estimates of the cross-lagged paths between JS and OC (symmetric reciprocation)					
JS1 → OC2	0.09*	0.01	0.20	−0.08, 0.40	0.48
OC1 → JS2	0.09*	0.01	0.20	−0.08, 0.40	0.48
JS1 → TI2	−0.05*	0.01	0.29	−0.33, 0.34	0.67
OC1 → TI2	−0.08*	0.01	0.35	−0.73, 0.13	0.86
TI1 → JS2	−0.008	0.01	0.32	−0.46, 0.32	0.78
TI1 → OC2	0.02	0.01	0.29	−0.43, 0.23	0.66
The model with the lagged effect of JS on OC fixed to be zero (OC → JS)					
OC1 → JS2	0.10*	0.01	0.28	−0.13, 0.53	0.66
JS1 → TI2	−0.004	0.01	0.32	−0.27. 0.45	0.72
OC1 → TI2	−0.09*	0.01	0.34	−0.73, 0.10	0.83
TI1 → JS2	−0.03*	0.01	0.34	−0.48, 0.33	0.81
TI1 → OC2	−0.02	0.01	0.24	−0.44, 0.14	0.58
The model with the lagged effect of OC on JS fixed to be zero (JS → OC)					
JS1 → OC2	0.14*	0.01	0.27	−0.10, 0.50	0.60
JS1 → TI2	−0.08*	0.01	0.27	−0.32, 0.30	0.62
OC1 → TI2	−0.03*	0.01	0.38	−0.67, 0.24	0.91
TI1 → JS2	−0.02	0.01	0.28	−0.48, 0.21	0.69
TI1 → OC2	0.02	0.01	0.30	−0.50, 0.20	0.70
The model with both the lagged effect of OC on JS and the lagged effect of JS on OC fixed to be zero (no causality)					
JS1 → TI2	−0.02	0.01	0.32	−0.28, 0.42	0.70
OC1 → TI2	−0.05*	0.01	0.39	−0.72, 0.21	0.93
TI1 → JS2	−0.06*	0.01	0.27	−0.55, 0.13	0.68
TI1 → OC2	−0.04*	0.01	0.25	−0.48, 0.10	0.58

Note. Standardized path coefficients (i.e., β) and the standard error of the path coefficients (i.e., *SE*) are from the traditional MASEM results, while the distribution of the path coefficients (i.e., SD_β, 80% CV_β, and 80% CV_β width) is from Yu et al.'s (2016) FIMASEM results—bootstrapping simulation technique with 10,000 samples while accounting for random-effects covariance among the variables. JS = job satisfaction; OC = organizational commitment; TI = turnover intentions; 1 = Time 1; 2 = Time 2. SD_β = standard deviation of the distribution of β coefficients; 80% CV = 80% credibility interval; MASEM and FIMASEM = full information meta-analytic structural equation modeling.

* $p < .05$.

remove a larger proportion of variance in JS at Time 2 to be explained by OCQ at Time 1, resulting in a "smaller" lagged effect of OCQ on JS compared to AC as a parallel without contamination (cf. Meyer & Herscovitch, 2001). Despite criticisms over it, OCQ is still commonly used (Newman & Yearick, 2019), perhaps because previous meta-analyses based on cross-sectional data demonstrate that OCQ often has "stronger" correlations with other variables than other OC measures (i.e., *inflated* correlations, e.g., Griffeth et al., 2000; Riketta, 2002). However, our meta-analysis based on panel data challenges previous meta-analyses, showing that OCQ has no lagged effect on JS (i.e., no inflation), strikingly different from AC and KUT. The adoption of panel designs with cross-lagged analyses which control for autoregressive and concurrent effects might correct OCQ's inflated relationships with other variables due to measure contamination.

Our study also provides insights into OC conceptualizations. First, there is a lagged effect of JS on AC but not NC, supporting that positive work experiences (e.g., JS) contribute to the development of AC but not NC (cf. Meyer & Allen, 1991). This is in stark contrast to previous meta-analysis based on cross-sectional data supporting that AC and NC have similar correlation with JS (Meyer et al., 2002). Although AC and NC are expected to develop through distinct routes, shared antecedents can lead to their convergence (Meyer & Allen, 1991). Taking a developmental perspective may provide new insights into their distinction by separating the influence of shared antecedents and common-method variance (Bergman, 2006). Taking a static approach, cross-sectional research cannot control previous levels of the variables and rule out the influence of constant individual differences or shared antecedents (Lang et al., 2011; Nohe et al., 2015), such as negative affectivity and the Big Five, resulting in similar correlations with other variables (e.g., JS) between AC and NC (Meyer et al., 2002). Second, there is no cross-lagged relation between JS and CC, confirming that CC is different from AC and NC. Both AC and NC have a lagged effect on positive work experiences (i.e., JS), but not CC (and no lagged effect of CC on JS). However, person-centered research suggests that the

nature of CC depends on the context created by AC and/or NC (e.g., Kabins et al., 2016; Kam et al., 2016; Meyer & Morin, 2016). Specifically, with the presence of high AC, CC reflects a potential loss of valued opportunities and resources, which may influence the development of JS. Yet, with the absence of high AC, CC reflects a threat of economic costs or a loss of side bets, which may not affect the growth of JS. Thus, the cross-lagged relationship between CC and JS might be more nuanced. Finally, KUT is conceptually different from AC. Specifically, the lagged effect of AC on JS is larger than that of JS on AC (i.e., an asymmetric reciprocation), whereas the lagged effect of KUT on JS is similar to that of JS on KUT (i.e., a symmetric reciprocation).

Collectively, our findings suggest that OC conceptualizations may place boundaries on existing theories that explain the ordering between JS and OC. First, affective attachment theory predicts that forming sooner than OC, JS causes OC. Yet, JS has a lagged effect on OCQ, AC, and KUT but not NC, perhaps because NC could develop before organizational entry (Cohen, 2007; Meyer & Allen, 1997). Thus, affective attachment theory as the theoretical foundation for JS causing OC may not hold for NC which could form before organizational entry. Second, cognitive dissonance theory predicts that OC initiates a rationalization process to adjust JS in order to reduce cognitive dissonance, only when individuals have free choices (Festinger, 1957). Because CC reflects one's need to be committed to the organization due to perceived costs of leaving and no employment alternatives, such absence of free choice may not elicit dissonance arousal, resulting in no lagged effect of CC on JS. In contrast, AC, KUT, and NC reflecting volitional commitment to the organization due to free choices (e.g., value congruency) will elicit pressures to adjust JS to reduce cognitive dissonance. In short, nuanced differences in OC conceptualizations should be considered when theorizing the ordering between JS and OC.

Time also places boundaries on the ordering between JS and OC—the lagged effects of both JS and OC decrease over time. Adaptation theory suggests that individuals will adapt to positive and negative experiences such that these experiences only temporally influence individuals due to hedonic adaptation (Helson, 1948, 1964). Thus, employees will get habituated to the effects of JS and OC such that these effects decrease and eventually disappear over time. That is, at the beginning, there may be lagged effects of JS and/or OC; as time goes by, the lagged effects of JS and OC may fade and eventually disappear due to habituation (i.e., no causality). Hence, adaptation theory partially explains the heterogeneous orderings of JS and OC by adding the time element, which is critical for causal processes but has been completely ignored by existing theories (Mitchell & James, 2001).

Contradicting to the predictions of affective attachment theory and cognitive dissonance theory, there is a reciprocation between JS and OC for nonnewcomers, but no cross-lagged relation between JS and OC for newcomers. Perhaps different from established employees, both JS and OC are not stabilized for newcomers who just enter the organization and still need time to have a good understanding of the job and the organization. Indeed, adapting to a new organization and a new job is very complicated because newcomers may have limited knowledge about the organization and unrealistic expectations about the job (Van Vianen & De Pater, 2012). Thus, newcomers' OC and JS are "volatile" due to tumultuous demands such as insecurity, new relationships, and reality shocks, leaving limited room for mutual influence between JS and OC (i.e., the future JS/OC is less well predicted by the past OC/JS). Consistently, our supplementary analyses indicate that there is no cross-lagged relationship between JS and OC for employees experiencing radical organizational change but a reciprocal relationship for those experiencing no organizational change, perhaps this is also because JS and OC are "volatile" due to tumultuous demands during radical organizational change.[7] Collectively, the timing of the investigation is highly relevant to the ordering between JS and OC.

Finally, our results suggest that the specifications of the ordering between JS and OC can change the conclusions regarding whether JS and OC precede TI and whether TI precedes JS and OC. For example, for the model in which OC leads to JS, there is a lagged effect of OC, but not JS on TI, and a lagged effect of TI on JS, but not OC. However, for the model in which JS leads to OC, there are lagged effects of both JS and OC on TI, but no lagged effects of TI on JS or OC. It is worth noting that all alternative specifications of the ordering between JS and OC have good model fit (e.g., CFIs $> .95$). Thus, without considering alterative orderings, theories and practices informed by studies in which JS and OC are modeled as common antecedents, mediators, or outcomes of organizational phenomena (e.g., performance, mistreatment) may be questionable.

Limitations and Future Directions

We acknowledge several limitations. First, we were not able to directly test the ordering between JS and OC with shared antecedents. However, our analyses indicate that a common factor could not explain their relation. Further, cross-lagged analyses control for previous levels of the variables and rule out the influence of constant background variables, such as negative affectivity (cf. Lang et al., 2011; Nohe et al., 2015). However, "real" causation can be established only with experimental designs. Clearly, this is virtually never the case when a relationship is studied in work contexts. Considering that it is often unethical and/or infeasible to manipulate employees' attitudes and behaviors, our meta-analytic estimates are of considerable value. At the very least, our meta-analysis based on panel data improves our understanding of *the temporal precedence* of JS and OC—a necessary condition for causal inference (James et al., 1982). Yet, our meta-analysis based on two-wave data is less informative. For instance, two-wave data does not allow for testing nonlinear change over time. Second, due to the small ks, we could not test confounds and interactions among the moderators (e.g., higher order interactions). For instance, there might be an interaction between time lag and the timing of the investigation (e.g., newcomers). The potential interactions among the moderators raise some concern about the interpretability of the moderation effects discovered in this meta-analysis. Relatedly, the analysis for KUT based on a small number of samples ($k = 4$) may suffer from second-order sampling error. Finally, although our findings have important implications for theory and research, they provide limited practical implications. In particular, the overall lagged effects of JS and OC are relatively small, resulting in limited practical impacts.

[7] We also investigated organizational change as a moderator (see Supplemental Materials on pp. 11–14).

However, the heterogeneity discovered by our meta-analysis in and of itself is an important finding, providing new, fruitful avenues for future research. First, future research can explore other moderators, potential confounds, and interactions among the moderators, such as the interaction between time lag and the timing of the investigation (e.g., organizational entry). Newcomers harbor unrealistic expectations and go through various reality shocks and thus have to frequently reevaluate their JS and OC during the socialization process (Louis, 1980). Thus, unrealistic expectations and socialization experiences become shared determinants of newcomers' JS and OC that drive dramatic changes in their JS and OC, leaving little room for mutual influence between JS and OC. As a result, it takes time for newcomers' JS and OC to stabilize and exert mutual influence on each other, perhaps after the effects of unrealistic expectations and socialization experiences die down. Indeed, research suggests that it takes several months after organizational entry for newcomers' JS and OC to stabilize (e.g., Boswell et al., 2009; Jokisaari & Nurmi, 2009). Therefore, while the cross-lagged effects between JS and OC occur *immediately* for established employees, for newcomers, they may occur several months after organizational entry. Further, because JS (i.e., new job) and OC (i.e., new organization) are relatively "new" to newcomers, the habituation may take longer for newcomers to occur compared to established employees. Thus, the lagged effects of JS and OC may last longer for newcomers than established employees.

Second, future research should take a fine-grained approach to investigate the role of facets of JS. Different composite JS measures include different facets. Thus, how disparities in the facets assessed by these measures influence the temporal relation between JS and OC warrants a systematic evaluation. The bandwidth–fidelity debate (Hogan & Roberts, 1996; Ones & Viswesvaran, 1996) and the principle of correspondence (Ajzen & Fishbein, 1977) suggest that broad constructs more strongly relate to broad constructs, while narrow constructs more strongly relate to narrow constructs. Because facet satisfaction (e.g., pay satisfaction) represents a narrow conceptualization of JS, while OC is a relatively broad conceptualization, the cross-lagged relationship between facet satisfaction and OC might be smaller than that between (overall) JS and OC. Finally, based on the target-similarity model (Lavelle et al., 2007), facet satisfaction such as promotion and pay satisfaction more attributable to the organization might have stronger lagged relationships with OC than facet satisfaction less attributable to the organization.

Third, person-centered approaches may provide new insights into the complex cross-lagged relationship between JS and OC. For instance, growth mixture modeling (Meyer & Morin, 2016; Morin et al., 2018) can identify subgroups of individuals characterized by distinct latent change patterns of JS and OC—whether a single or multiple change patterns adequately characterize the entire population, and explore factors explaining such heterogeneity. Another person-centered approach, latent profile analyses can identify distinct commitment profiles characterized by varying levels of AC, NC, and CC (Meyer & Morin, 2016) and thus helps unpack the nuanced cross-lagged relationship between CC and JS. For instance, if individuals with a profile of high AC and high CC (e.g., high commitment profile) at Time 1 have higher JS at Time 2 compared to those with a profile of high AC and low CC (e.g., AC-dominant profile) at Time 1, it would suggest that CC, reflecting a potential loss of valued opportunities and resources, could influence positive work experiences such as JS. In contrast, if individuals with a profile of low AC and high CC (e.g., CC-dominant profile) at Time 1 experience similar JS at Time 2 to those with a profile of low AC and low CC (e.g., low-commitment profile) at Time 1, it would suggest that CC, reflecting a threat of economic costs or "a feeling of being trapped," may not affect the growth of JS.

Finally, the heterogeneous orderings between JS and OC across different OC conceptualizations may result from distinctive psychological processes. For instance, AC and KUT might influence JS through both cognitive (e.g., cognitive dissonance) and affective (e.g., positive emotion) processes, whereas NC may influence JS only via cognitive processes (e.g., moral rationalization). Future research can explore the mechanisms underlying the heterogeneous cross-lagged relationships between JS and OC to improve our understanding of why JS has a lagged effect on OC and vice versa, as well as how different OC conceptualizations play a role.

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References marked with an asterisk indicate studies included in the meta-analysis.

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