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# The Causal Relation Between Job Attitudes and Performance: A Meta-Analysis of Panel Studies

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Do job attitudes cause performance, or is it the other way around? To answer this perennial question, the author conducted meta-analytic regression analyses on 16 studies that had repeatedly measured performance and job attitudes (i.e., job satisfaction or organizational commitment). The effect of job attitudes on subsequent performance, with baseline performance controlled, was weak but statistically significant ($\beta = .06$). The effect was slightly stronger for commitment than for satisfaction and depended negatively on time lag. Effects of performance on subsequent job attitudes were elusive ($\beta = .00$ across all studies), which suggests that job attitudes are more likely to influence performance than vice versa.

*Keywords:* job satisfaction, organizational commitment, performance, organizational citizenship behavior, meta-analysis

Few topics in industrial and organizational psychology have received as much attention as has the relation between job attitudes and performance (e.g., Brief & Weiss, 2002; Judge, Thoresen, Bono, & Patton, 2001; Meyer & Allen, 1997; Mowday, Porter, & Steers, 1982; Staw, Sutton, & Pelled, 1994). Numerous meta-analyses (e.g., Cooper-Hakim & Viswesvaran, 2005; Harrison, Newman, & Roth, 2006; Judge et al., 2001; Meyer, Stanley, Herscovitch, & Topolnytsky, 2002; Riketta, 2002) have demonstrated that positive job attitudes, such as commitment and satisfaction, are accompanied by better work outcomes. Although the existence of positive correlations is well established, the causal relationship between job attitudes and performance is still unclear. Do job attitudes increase performance? Is it the other way around? Or are the frequently observed correlations between job attitudes and performance spurious (e.g., due to common causes)? The vast majority of empirical studies on job attitudes and performance are mute on these issues because of their cross-sectional designs. The same is true of the aforementioned meta-analyses. Thus, the longstanding debate about the causal relationship between job attitudes and individual performance (e.g., Harrison et al., 2006; Judge et al., 2001; March & Sutton, 1997; Organ, 1977; Schwab & Cummings, 1970) is far from being resolved.

The goal in this article is to contribute to this debate by providing the most controlled (to date) meta-analytic test of causal links between job attitudes and performance. This article is built around a meta-analysis of panel studies on these two constructs. These studies permit the extent to which job attitudes predict performance to be disentangled from the extent to which performance predicts job attitudes. In this article, meta-analytic regression analysis is applied to the aggregated correlations to estimate the unique effect of job attitudes on performance (with baseline performance controlled) and the unique effect of performance on job attitudes (with baseline job attitudes controlled). Differences between forms of job attitudes (organizational commitment and job satisfaction) and performance (in-role and extra-role), as well as the moderating role of measurement interval, are explored.

This article provides the first meta-analysis that estimates longitudinal effects between job attitudes and performance while controlling for baseline scores (for similar methods that examine team cohesion instead of job attitudes, see the meta-analysis by Mullen & Copper, 1994; for a meta-analysis of zero-order longitudinal correlations between job attitudes and performance, see Harrison et al., 2006). The present meta-analysis therefore extends previous meta-analyses of the job attitude–performance relationship by way of a more rigorous test of causal hypotheses.

## Definitions and Theoretical Models

Throughout this article, the term *job attitude* refers to the evaluation or personal importance of job-related targets (e.g., organization, work group, job as a whole). The two most frequently investigated job attitudes probably are job satisfaction, defined as a cognitive and/or affective evaluation of one's job as more or less positive or negative (Brief & Weiss, 2002), and attitudinal or affective organizational commitment, defined as "the relative strength of an individual's identification with and involvement in a particular organization" (Mowday, Steers, & Porter, 1979, p. 226; see also Allen & Meyer, 1990). The following arguments and empirical analyses refer to these job attitudes only. Other forms of job attitudes include organizational identification (see Riketta, 2005); job involvement (see Brown, 1996); continuance and normative commitment (Allen & Meyer, 1990); and satisfaction and (affective) commitment with reference to targets other than job or organization, such as work group, career, or occupation (see, e.g., Becker, 1992; Cohen, 2003; Meyer, Allen, & Smith, 1993). These

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forms of job attitudes are not considered within this article, due to a lack of published investigations that contain these constructs.[1]

In the literature, performance is usually divided into in-role performance (similar to task performance), defined as fulfillment of tasks that are required by the formal job description, and extra-role performance (similar to organizational citizenship behavior or contextual performance), defined as behavior that is beneficial to the organization and goes beyond formal job requirements (e.g., helping colleagues at work, working extra hours, making suggestions for improvement; Borman & Motowidlo, 1997; Organ, 1988). This meta-analysis considers both forms of performance.

At least four interpretations of positive correlations between job attitudes and performance are possible. Because these viewpoints have been laid down many times (e.g., Brief & Weiss, 2002; Brown & Peterson, 1993; Harrison et al., 2006; Judge et al., 2001; Meyer & Allen, 1997; Mowday et al., 1982; Staw et al., 1994), only a brief summary is given below. This meta-analysis tests all four cases by estimating the unique effects of job attitudes on later performance (with baseline performance controlled) and of performance on later job attitudes (with baseline job attitudes controlled).

Case 1: Job attitudes cause performance. Arguments that support this view usually refer to the functions of attitudes as guidelines and facilitators of behavior (e.g., Eagly & Chaiken, 1993; Fishbein & Ajzen, 1974; see Judge et al., 2001); the energizing and facilitative effects of positive affect (as one component of satisfaction) at the workplace (e.g., Staw et al., 1994); or the motivational effects of the personal importance or identification with the job or organization (e.g., as a component or consequence of commitment; see, e.g., Meyer, Becker, & Vandenberghe, 2004). In this meta-analysis, this view would receive support if job attitudes predicted later performance.

Case 2: Performance causes job attitudes. Two common arguments that support this view are (a) that performance often leads to internal and external rewards (e.g., pay, recognition, feeling good at work), which in turn may foster positive job attitudes (e.g., Lawler & Porter, 1967), and (b) that people adjust their attitudes to their behavior, due to strivings for cognitive consistency or as a rationalization for their actions (e.g., as assumed by psychological theories of cognitive dissonance and of self-perception, Festinger, 1957, and Bem, 1972, respectively; see, e.g., Staw, 1975). This view would receive support if performance predicted later job attitudes.

Case 3: Performance and job attitudes cause each other. This case results from the combination of Cases 1 and 2.

Case 4: Performance and job attitudes are causally unrelated. In this case, the positive concurrent correlations between them would be due to research artifacts (e.g., common source bias) or to third variables that influenced both constructs (see, e.g., Brown & Peterson, 1993; Judge et al., 2001). Although it is virtually impossible to rule out Case 4 with correlational data, this meta-analysis tests one possible implication of this case: that job attitudes and performance do not predict each other over time.

This study explores the moderating roles of type of job attitude (satisfaction vs. commitment), performance type (in-role and extra-role performance), and timing (shorter vs. longer intervals between measurement waves). In theory, panel designs require that the time between two measurement waves matches the time that the effects under investigation presumably take to unfold. However, little is known about this process (e.g., how long it takes for satisfaction to influence performance or vice versa). Because of this lack of knowledge and the common constraints of field studies (e.g., the organization may provide access only at certain intervals), the time lags in most studies seem to be set independently of theoretical considerations, as reviews of panel studies have admonished (e.g., Williams & Podsakoff, 1989). Through its comparison of attitude–performance effects between different time lags, this meta-analysis may inform decisions on timing issues in future panel studies.

## Method

### *Study Search and Coding*

Studies had to meet the following criteria to be included in the meta-analysis:

1. Participants were employees in an organization. Thus, studies in other contexts, such as classrooms, sports teams, or artificial environments (e.g., laboratory, scenario), were excluded (e.g., Dorfman & Stephan, 1984; Grieve, Whelan, & Meyers, 2000).

2. The study examined job satisfaction or organizational commitment (attitudinal or affective).

3. The study examined job performance.

4. Job attitudes and performance were measured at each of at least two measurement waves. That is, the study had a panel design.

5. No major changes in the work environment, such as an organizational merger or a change in the task of the participants, occurred between the measurement waves (e.g., Jetten, O'Brien, & Trindall, 2002).

6. Data were analyzed at the individual level rather than at the group level. This criterion was included because most theoretical accounts of the job attitude–performance relation refer to individual processes and because individual-level correlations are not comparable with group-level correlations.

7. The complete matrix of the zero-order correlations for job attitude and performance was available for at least two measurement waves. Thus, the report of the study had to contain the two synchronous correlations, the two cross-lagged correlations, and the two stabilities for job

[1] The literature search for this meta-analysis did refer to studies on all of the mentioned job attitudes. After the search, all job attitudes were to be excluded for which fewer than five independent studies were available. This step would reduce the heterogeneity of the data set for the meta-analysis and would ensure that the meta-analytic results for single job attitudes were generalizable to some extent. Only job satisfaction and organizational commitment met this criterion (and the other inclusion criteria described in the Method section). For ease of presentation, the excluded constructs are not mentioned further.

attitude and performance. Only studies that reported the complete set of correlations were considered in the meta-analysis. This criterion ensured that the attitude–performance and the performance–attitude paths within an analysis were from the same studies and thus ruled out differences between studies as confounds of the observed effects. In other words, studies with missing correlations were excluded listwise.

This study used several strategies for identification of published and unpublished studies that met these criteria. A range of electronic databases was searched, including ABI/Inform (covering published articles and unpublished dissertations, some of them in full text); Business Source Premier (covering published articles); PsycINFO (covering published articles, chapters, and books and unpublished dissertations); and Web of Science (the former Social Sciences Citation Index; covering published articles). The following search terms, decomposed into smaller search terms as necessary, were used: (*satisfaction* or *commitment*) and (*work* or *job* or *organization*) and (*cross-lagged* or *longitudinal*) and (*performance, in-role,* or *extra-role;* or *citizenship* or *effort* or *productivity* or *work motivation*). Moreover, the lists of studies included in previous meta-analyses and qualitative reviews on satisfaction, commitment, and performance were checked, as were the references of several papers on cross-lagged panel analysis (most notably, Clegg, Jackson, & Wall, 1977; Williams & Podsakoff, 1989). Further, colleagues who research actively in the area of job attitudes were asked if they knew of relevant research; a request for unpublished data was sent via the mailing list of the German association of industrial and organizational psychologists (in September 2006); a request for unpublished data was posted on the web page of the Society for Industrial and Organizational Psychology (in September 2006); the abstracts of recent annual conferences of that society (2005–2007) and of the Academy of Management (2000–2007) were searched, and several papers were requested; and authors of published studies that met all but the last inclusion criterion (i.e., that failed to report the complete correlation matrix) were contacted and asked for the missing correlations. The references of each relevant paper retrieved were scanned for additional studies. Study search was completed in June 2007.

All usable studies were coded by Michael Riketta and a doctoral student into two categories according to the nature of performance the studies measured (in-role and extra-role). When a performance measure included items that referred to both performance types, it was coded into the category to which most of its items referred. Intercoder agreement was 100%, and each study could be unambiguously assigned to one of the two categories. Because all other data to be coded did not require subjective judgments, they were coded by Michael Riketta.

As is common in meta-analyses, the moderator effects were explored by repeating the analyses for discrete values of the moderator (Hunter & Schmidt, 1990). Thus, the moderator variable *time lag* was categorized into 1–6 months, 7–12 months, and 13+ months. These categories were chosen as a compromise between the two conflicting demands of (a) having a reasonable number of studies in each category (which could be best fulfilled by a small number of categories) and (b) having a category system differentiated enough for detection of nonmonotonic moderator effects (which could be best fulfilled by a large number of categories). Given the small number of available studies, a larger number of time lag categories did not seem meaningful. (Nonetheless, a more differentiated analysis for the total sample is reported in Footnote 3.)

### Features of the Analyzed Studies

The literature search yielded 16 usable studies (see Table 1). The average sample size of these studies was 192, with a range from 35 to 526. Mean time lag between the coded waves of measurement averaged 9.2 months, with a range from 1 month to 18 months. Mean organizational tenure of participants at the beginning of data collection was 4.5 years ($k = 10$). Mean proportion of women was 55% ($k = 11$). The majority of the studies were conducted in English-speaking countries (10 in the United States, 2 in the United Kingdom, 1 in Australia). Two studies were conducted in Germany; for another study, the country was not stated. Eight studies were conducted in service organizations, 3 were conducted in manufacturing organizations, and 4 examined participants from multiple organizations and industries (e.g., alumni of the study authors' institutions). One study was conducted in an organization within an unspecified industry.

Of the studies, 14 examined job satisfaction with a variety of measures. The most frequent measure was the Job Descriptive Index (Smith, Kendall, & Hulin, 1969), which was used in 3 studies. Five studies examined organizational commitment, measured with the Organizational Commitment Questionnaire (Mowday et al., 1982). In-role and extra-role performance were measured by 11 and 5 studies, respectively. Extra-role performance was measured with self-reports in 4 studies and with both self-reports and peer ratings in the remaining study. In-role performance was measured with supervisor ratings in 6 studies, with objective indicators in 3 studies, with both supervisor ratings and objective indicators in 1 study, and with self-reports in another study. No single performance measure was used more than twice.

### Data Aggregation

One requirement of a meta-analysis is independence of the aggregated data points (here, correlations). Thus, a study must not contribute more than one correlation to each aggregated correlation. When a study provided correlations for (a) more than one job attitude or performance form or (b) more than one measure for the same job attitude or performance form, the correlations were averaged, such that the study contributed no more than one set of correlations (two stabilities, two synchronous correlations, two cross-lagged correlations) to each of the following analyses. For example, when a study provided correlations for commitment and satisfaction that had the same outcome, these correlations were averaged for the overall analyses. The single correlations for commitment and satisfaction were used, however, in the separate analyses for commitment and satisfaction.

The issue of independent correlations is also relevant to studies with more than two waves of measurement (here, three studies with three waves). To ensure independence of data points, this analysis used only the data from the first two measurement waves, except in the analyses that compared different time lags. In this latter case, if a study reported correlations for more than one time-lag category (i.e., <7, 7–12, and 13+ months), these corre-

Table 1
*Studies Included in the Meta-Analysis*

| | | | | | | Coded correlations | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Study | *N* | Participants | Country | Lag | Variables | $A_1A_2$ | $P_1P_2$ | $A_1P_1$ | $A_2P_2$ | $A_1P_2$ | $P_1A_2$ |
| Ashforth & Saks (1996) | 222 | University graduates | U.S. | 6 | JS, IP | .64 | .69 | .11 | .21 | .14 | .20 |
| Bateman & Organ (1983) | 77 | Nonacademic university staff | U.S. | 1.5 | JS, EP | .71 | .80 | .41 | .41 | .43 | .39 |
| Bechtold et al. (1981) | 64 | Medical center employees | U.S. | 18 | JS, IP | .53 | .57 | .15 | .21 | .17 | .19 |
| Bond & Bunce (2003) | 412 | Call-center employees | U.K. | 12 | JS, IP | .66 | .21 | .26 | .17 | .05 | .66 |
| Borrill et al. (2003) | 370 | Hospital employees | U.K. | 12 | JS, EP | .67 | .48 | .28 | .25 | .20 | .17 |
| Crampon et al. (1978) | 46 | Management trainees | U.S. | 2 | OC, IP | .72 | .78 | .16 | .16 | .36 | .00 |
| Donaldson et al. (2000) | 157 | Nonprofessionals | U.S. | 6 | OC, EP | .71 | .25 | .20 | .25 | .14 | .71 |
| Griffin (1991)[a] | 526 | Bank tellers | U.S. | 18 | JS, IP | .61 | .53 | .04 | .06 | −.02 | .06 |
| | | | | | OC, IP | .74 | .53 | −.03 | .05 | .05 | .03 |
| Maier & Rosenstiel (2006) | 185 | University graduates | Germany | 14 | JS, EP | .57 | .57 | .11 | .195 | .09 | .17 |
| | | | | | OC, EP | .75 | .57 | .18 | .185 | .24 | .18 |
| Maier & Rosenstiel (2006) | 216 | University graduates | Germany | 14 | JS, EP | .64 | .65 | .16 | .16 | .14 | .14 |
| | | | | | OC, EP | .84 | .65 | .18 | .14 | .12 | .13 |
| Nathan et al. (1991) | 300 | Managers and professionals | Not stated | 3.5 | JS, IP | .56 | .23 | .06 | .17 | .14 | −.02 |
| Sheridan & Slocum (1975) | 59 | Machine operators | U.S. | 11 | JS, IP | .45 | .50 | −.03 | .15 | −.08 | −.06 |
| Sheridan & Slocum (1975) | 35 | Managers | U.S. | 12 | JS, IP | .68 | .49 | .20 | .21 | .21 | .24 |
| Szilagyi (1980) | 128 | Controllers and accountants | U.S. | 3 | JS, IP | .62 | .65 | .09 | .05 | .09 | .09 |
| Tharenou (1993) | 200 | Electrical apprentices | Australia | 12 | JS, IP | .48 | .64 | .19 | .08 | .11 | .08 |
| Wanous (1974) | 80 | Telephone operators | U.S. | 2 | JS, IP | .73 | .44 | .09 | .15 | .18 | .24 |

*Note.* Lag = time lag between the coded measurement waves in months; $A_1$ and $A_2$ = job attitude at first and second coded wave, respectively; $P_1$ and $P_2$ = performance at first and second coded wave, respectively; JS = job satisfaction; IP = in-role performance; EP = extra-role performance; OC = organizational commitment.
[a] Only Time 2 and Time 3 were coded because of an intervention between Time 1 and Time 2.

lations were included in the analyses for the respective categories. Again, when a study reported more than one set of correlations relevant to the same time-lag category, only the set of correlations for the two earliest measurement waves was considered.

In longitudinal studies, changes in reliability between the measurement waves can bias estimates of cross-lagged effects (Kenny, 1975; Williams & Podsakoff, 1989). To correct for this, the correlations were disattenuated. The study-specific reliability estimates for the relevant measurement waves were used if available. These reliability estimates were internal consistencies in all cases. When reliability information was lacking, imputed estimates made the analysis for these studies more comparable with those for the other studies. Specifically, when reliability information was available for only one measurement wave (as was the case in one study), this value was imputed as the reliability estimate for the second measurement wave. Reliabilities of single-item rating scales were set at .70 (Wanous & Hudy, 2001). In all other cases, missing reliabilities for job satisfaction, organizational commitment, and (in- or extra-role) performance were set at .83, .83, and .85, respectively. These values were the average reliabilities (mostly internal consistencies) from a recent, extraordinarily large meta-analysis on these constructs (Cooper-Hakim & Viswesvaran, 2005; these estimates were based on 949, 311, and 159 studies, respectively).[2]

The next step consisted of averaging the disattenuated correlations across studies, after weighting them with the product of sample size (to correct for sampling error) and the squared disattenuation factor (i.e., the square of the ratio of uncorrected to corrected correlation; Hunter & Schmidt, 1990). The resulting weighted correlation was an estimate of the mean population correlation. Its standard error was computed as the standard deviation of the corrected correlations divided by the square root of the number of studies. Thus, as recommended by Hunter and Schmidt (1990), this meta-analysis used a random effects model. The variance of the population correlations was computed as the dif-

[2] Schmidt and Hunter (1996) and Viswesvaran, Ones, and Schmidt (1996) argued that interrater reliabilities are better estimates of measurement error than are internal consistencies. In a meta-analysis, Viswesvaran et al. estimated the interrater reliability of supervisor ratings of performance at .52 and suggested that this estimate be used for disattenuation in meta-analyses. When their estimate (rather than the internal consistencies from the original studies) was used in this study for supervisory performance ratings, the paths between job attitudes and performance tended to be slightly weaker but showed patterns largely similar to those in the present analysis. In particular, the job attitudes–performance and performance–job attitudes effects were, respectively, $\beta$ = .04 and .00 overall; .02 and .00 for satisfaction; .07 and .02 for commitment; .4 and .01 for in-role performance; .05 and −.02 for extra-role performance; .11 and .03 for a time lag of 1–6 months; .01 and −.08 for 7–12 months; and .01 and .05 for 13+ months. A limitation of this method is that it does not consider changes in reliability within studies, although this is advisable for the analysis of panel data (Kenny, 1975). Moreover, several authors have argued that the use of interrater reliabilities in disattenuation may bias correlations (Murphy & De Shon, 2000; Sackett, Laczo, & Arvey, 2002). Hence, only the analysis that used internal consistencies is reported in the text.

ference between the variance of the corrected correlation coefficients and their average squared standard errors (Hunter & Schmidt, 1990). Heterogeneity of population correlation was tested with Hunter and Schmidt's chi-square test (*Q* test). A significant result would indicate that there was more than one population correlation.

### Meta-Analytic Regression Analysis

For the causal analyses, the matrix of the corrected mean correlations served as input into a meta-analytic regression analysis (Judge & Piccolo, 2004; Viswesvaran & Ones, 1995). The software MPlus 4.2, using maximum likelihood estimation, was used for these computations. To increase the sensitivity of significance tests, the sum of the sample sizes of the relevant studies (rather than, e.g., the average) was used to compute the standard errors for the regression coefficients (see Cheung & Chan, 2005). Specifically, across all studies, performance or job attitude at the second coded measurement wave was regressed on both performance and job attitudes at the first coded measurement wave. The standardized regression coefficients provided by this analysis estimated how well job attitudes and performance predicted each other, with baseline scores of the criterion variable being controlled. These analyses were conducted across all job attitudes and performance forms, irrespective of time lags (called *overall analysis* hereafter), for each combination of satisfaction or commitment with in-role or extra-role performance and with each time-lag category.

## Results

Table 2 shows the mean corrected correlations (as estimates of the mean population correlations; symbolized by $r$ in the following text); their standard errors (as indicators of the precision with which the mean population correlations were estimated); and the estimated standard deviations of the population correlations (as estimates of the extent to which the population correlations vary around the mean population correlation). Data are shown for the overall analysis and for each category of each moderator (job attitude, performance type, time lag, and all possible combinations thereof).

The mean cross-sectional correlations between job satisfaction and organizational commitment with performance were weakly positive and statistically significant ($r$s between .10 and .21, $p$s $<$ .05). These findings are consistent with those of previous meta-analyses (e.g., Cooper-Hakim & Visweswaran, 2005; Judge et al., 2001; Meyer et al., 2002; Riketta, 2002). The stabilities of job attitudes and performance were remarkably high ($r$s $>$ .52 across all time lags, i.e., for an average time lag of 9 months).

Table 3 shows the results of the meta-analytic regression analyses conducted on the correlations from Table 2. The upper panel of Table 3 shows general analyses, which averaged across at least one of the moderator variables. Overall, job attitudes were weak predictors of performance ($\beta = .06$, $p < .001$). This effect tended to be stronger for commitment than for satisfaction ($\beta = .08$ vs. .03, $p$s $<$ .05) and stronger for shorter than for longer time lags ($\beta = .12$, $p < .001$, for 1–6 months; $\beta = .02$, *ns*, for 7–12 months; $\beta = .03$, *ns*, for more than 12 months). It did not differ between in-role and extra-role performance ($\beta$s $= .05$, $p$s $< .05$).

Effects of performance on job attitudes in the general analyses were more elusive. The only significant effect was a negative effect for moderate time lag ($\beta = -.08$, $p < .001$, 7–12 months). Because no studies of organizational commitment were available for this time-lag category, this effect was entirely due to satisfaction. All other effects of performance on job attitudes in the general analyses were nonsignificant ($\beta$s $< .04$).[3]

Table 3, lower panel, shows more specific analyses, which examined all possible combinations of job attitude, performance type, and time lag. Because most of these analyses were conducted on three samples or fewer, they have to be interpreted with great caution. The most remarkable finding may be that the tendency for higher effects of job attitudes on performance for shorter time lags was replicated for all four job attitude–performance type combinations.

## Discussion

With its use of meta-analytic regression analysis and its exclusive focus on studies with repeated measurements, this meta-analysis accomplished a more rigorous test of causal relations between job attitudes and performance than did previous meta-analyses on these relations. The results provide some support for the common assumption that job attitudes influence performance. Across job attitudes and performance forms, the effect was weak but significant ($\beta = .06$). The effect was significant for both satisfaction and commitment and for both in- and extra-role performance. The effect tended to be stronger for shorter time lags between measurement waves and for commitment rather than for satisfaction. Almost no statistically significant evidence for the reverse causal direction emerged, with the effect size in the overall analysis being $\beta = .00$. This finding suggests that job attitudes are more likely to influence performance than vice versa.

### Limitations

Before the implications of the findings for research and practice are discussed, several limitations of this meta-analysis should be noted. First of all, the number of studies was small, especially for commitment and extra-role performance (only five studies each) and for the moderator analysis for time lag (fewer than eight studies in each category).

Moreover, almost all studies on extra-role performance measured this construct with self-reports only. Common source bias and socially desirable self-presentation may have distorted the

[3] Additional analyses explored the effects of time lag in more detail by repeating the computations (across job attitudes and performance types) for every available time lag (rounded to months). The job attitudes–performance and performance–job attitudes effects were, respectively, $\beta = .18$ and .06 for a time lag of 2 months ($k = 3$, $n = 203$); .03 and .03 for 3 months ($k = 1$, $n = 128$); .20 and $-.12$ for 4 months ($k = 1$, $n = 300$); .06 and .08 for 6 months ($k = 2$, $n = 379$); $-.03$ and .01 for 9 months ($k = 1$, $n = 64$); $-.08$ and $-.05$ for 11 months ($k = 1$, $n = 59$); .02 and $-.08$ for 12 months ($k = 4$, $n = 1017$); .01 and .03 for 14 months ($k = 2$, $n = 401$); and .02 and .06 for 18 months ($k = 2$, $n = 590$). Weighted linear regression analyses on these data (with the inverse of $n$ as weights) revealed that time lag related strongly and negatively to the job attitudes–performance effects ($\beta = -.69$, $p = .04$) but not to the performance–job attitudes effects ($\beta = .13$, $p = .73$).

Table 2
*Aggregated Zero-Order Correlations*

| Analysis | $k$ | $n$ | $A_1A_2$ | $P_1P_2$ | $A_1P_1$ | $A_2P_2$ | $A_1P_2$ | $P_1A_2$ |
|---|---|---|---|---|---|---|---|---|
| Overall | 16 | 3,077 | .77 | .67 | .17 | .21 | .17 | .13 |
| | | | (.02, .06) | (.04, .15) | (.03, .09) | (.03, .06) | (.03, .07) | (.03, .05) |
| | | | 37.66$^{**}$ | 114.21$^{***}$ | 33.12$^{**}$ | 23.33 | 26.90 | 22.25 |
| JS | 14 | 2,874 | .75 | .63 | .17 | .21 | .14 | .13 |
| | | | (.02, .06) | (.04, .14) | (.03, .08) | (.03, .06) | (.03, .08) | (.03, .06) |
| | | | 31.59$^{**}$ | 97.24$^{***}$ | 25.98$^{*}$ | 22.24 | 27.07$^{*}$ | 20.56 |
| OC | 5 | 1,130 | .84 | .70 | .10 | .13 | .15 | .10 |
| | | | (.03, .05) | (.06, .13) | (.06, .12) | (.04, .01) | (.05, .09) | (.04, .01) |
| | | | 15.47$^{*}$ | 38.32$^{***}$ | 16.27$^{*}$ | 5.07 | 11.87$^{*}$ | 5.14 |
| In-role | 11 | 2,072 | .75 | .65 | .12 | .18 | .13 | .09 |
| | | | (.03, .07) | (.04, .14) | (.03, .05) | (.03, .05) | (.03, .05) | (.03, .01) |
| | | | 26.24$^{**}$ | 68.03$^{***}$ | 14.59 | 15.34 | 14.7 | 11.17 |
| Extra-role | 5 | 1,005 | .80 | .71 | .29 | .27 | .25 | .21 |
| | | | (.02, .03) | (.07, .16) | (.03, .00) | (.03, .00) | (.04, .02) | (.03, .00) |
| | | | 7.21 | 44.38$^{***}$ | 4.42 | 3.78 | 5.29 | 4.03 |
| Lag, 1–6 | 7 | 1,010 | .79 | .70 | .18 | .23 | .24 | .17 |
| | | | (.04, .09) | (.08, .21) | (.04, .06) | (.04, .00) | (.04, .05) | (.05, .09) |
| | | | 19.59$^{**}$ | 65.52$^{***}$ | 9.06 | 6.71 | 8.65 | 11.98 |
| Lag, 7–12 | 6 | 1,140 | .74 | .67 | .26 | .26 | .19 | .12 |
| | | | (.03, .03) | (.05, .11) | (.03, .00) | (.03, .00) | (.03, .00) | (.03, .00) |
| | | | 7.68 | 29.64 | 5.3 | 5.13 | 4.67 | 5.27 |
| Lag, 13+ | 4 | 991 | .79 | .66 | .07[a] | .13 | .09 | .11 |
| | | | (.03, .04) | (.06, .12) | (.05, .06) | (.04, .02) | (.04, .04) | (.04, .02) |
| | | | 8.36$^{*}$ | 24.94$^{***}$ | 7.15 | 4.42 | 5.17 | 4.22 |
| JS, extra-role | 4 | 848 | .80 | .68 | .28 | .28 | .23 | .22 |
| | | | (.02, .00) | (.08, .16) | (.05, .04) | (.00, .03) | (.05, .05) | (.00, .04) |
| | | | 1.70 | 35.01$^{***}$ | 5.11 | 3.09 | 5.72 | 3.70 |
| Lag, 1–6 | 1 | 77 | .86 | .86 | .47 | .46 | .49 | .45 |
| Lag, 7–12 | 1 | 370 | .79 | .53 | .32 | .28 | .23 | .19 |
| Lag, 13+ | 2 | 401 | .79 | .86 | .19 | .23 | .16 | .20 |
| | | | (.03, .00) | (.03, .00) | (.02, .00) | (.02, .00) | (.03, .00) | (.02, .00) |
| | | | 0.99 | 0.99 | 0.25 | 0.22 | 0.28 | 0.20 |
| JS, in-role | 10 | 2,026 | .73 | .64 | .13 | .18 | .11 | .10 |
| | | | (0.03, 0.07) | (0.05, 0.13) | (0.03, 0.03) | (0.03, 0.06) | (0.03, 0.06) | (0.03, 0.02) |
| | | | 23.68$^{*}$ | 61.06$^{***}$ | 11.09 | 14.71 | 15.66 | 10.46 |
| Lag, 1–6 | 4 | 730 | .79 | .62 | .11 | .20 | .17 | .14 |
| | | | (0.06, 0.1) | (0.11, 0.21) | (0.01, 0.00) | (0.04, 0.00) | (0.02, 0.00) | (0.06, 0.07) |
| | | | 15.74$^{**}$ | 37.22$^{***}$ | 0.10 | 2.71 | 0.90 | 6.10 |
| Lag, 7–12 | 5 | 770 | .71 | .67 | .21 | .23 | .15 | .08 |
| | | | (.00, .03) | (.04, .06) | (.00, .03) | (.00, .04) | (.00, .04) | (.00, .04) |
| | | | 4.49 | 13.86$^{*}$ | 3.46 | 3.94 | 3.77 | 3.23 |
| Lag, 13+ | 2 | 590 | .69 | .58 | .06 | .09 | .01 | .09 |
| | | | (.00, .02) | (.00, .01) | (.00, .03) | (.00, .04) | (.00, .05) | (.00, .04) |
| | | | 0.43 | 0.07 | 0.72 | 1.27 | 1.91 | 0.97 |
| OC, extra-role | 3 | 558 | .84 | .86 | .25 | .21 | .24 | .18 |
| | | | (.05, .08) | (.02, .00) | (.02, .00) | (.02, .00) | (.04, .00) | (.02, .00) |
| | | | 15.37$^{***}$ | 1.14 | 0.53 | 0.45 | 2.19 | 0.35 |
| Lag, 1–6 | 1 | 157 | .70 | .87 | .30 | .24 | .31 | .17 |
| Lag, 13+ | 2 | 401 | .88 | .86 | .22 | .20 | .22 | .19 |
| | | | (.03, .04) | (.00, .03) | (.00, .04) | (.00, .02) | (.00, .05) | (.00, .02) |
| | | | 5.65$^{*}$ | 0.99 | 0.01 | 0.28 | 1.51 | 0.33 |
| OC, in-role | 2 | 572 | .85 | .60 | −.02[a] | .07 | .08[a] | .03 |
| | | | (.00, .00) | (.06, .07) | (.04, .00) | (.02, .00) | (.07, .07) | (.00, .00) |
| | | | 0.07 | 7.27$^{*}$ | 1.51 | 0.54 | 4.23$^{*}$ | 0.01 |
| Lag, 1–6 | 1 | 46 | .87 | .92 | .19[a] | .19[a] | .43 | .01[a] |
| Lag, 13+ | 1 | 526 | .85 | .58 | −.03[a] | .06[a] | .06[a] | .03[a] |

*Note.* Sample-size weighted and disattenuated zero-order correlations. The standard errors of these mean correlations and the standard deviations of the population correlations are given in parentheses. The chi-square values (from the test of heterogeneity of the individual correlations, $df = k - 1$) are given below them. All correlations shown are significant at $p < .05$, except where indicated. $A_1$ and $A_2$ = job attitude at first and second coded wave, respectively; $P_1$ and $P_2$ = performance at first and second coded wave, respectively; JS = job satisfaction; OC = organizational commitment; Lag = time lag between the coded measurement waves in months.
[a] Nonsignificant correlation ($p > .05$).
$^{*}p < .05$. $^{**}p < .01$. $^{***}p < .001$.

Table 3
*Regression Analyses: Subsequent Attitude or Performance Regressed on Preceding Attitude and Performance*

| | | | Criterion: Subsequent attitude | | | Criterion: Subsequent performance | | |
|---|---|---|---|---|---|---|---|---|
| Analysis | *k* | *n* | Attitude | Performance | $R^2$ | Attitude | Performance | $R^2$ |
| General analyses | | | | | | | | |
| Overall | 16 | 3,077 | .77 (65.98)*** | .00 (0.08) | .59 | .06 (4.27)*** | .64 (48.75)*** | .45 |
| JS | 14 | 2,874 | .75 (59.87)*** | .00 (0.21) | .56 | .03 (2.11)* | .65 (44.86)*** | .42 |
| OC | 5 | 1,130 | .84 (51.70)*** | .02 (1.00) | .71 | .08 (3.81)*** | .69 (32.16)*** | .50 |
| In-role | 11 | 2,072 | .75 (51.24)*** | .00 (0.00) | .56 | .05 (3.15)** | .64 (38.37)*** | .43 |
| Extra-role | 5 | 1,005 | .81 (40.84)*** | −.02 (1.22) | .64 | .05 (2.08)* | .70 (30.05)*** | .51 |
| Lag, 1–6 | 7 | 1,010 | .79 (40.06)*** | .03 (1.47) | .63 | .12 (5.23)*** | .68 (30.11)*** | .50 |
| Lag, 7–12 | 6 | 1,140 | .76 (37.08)*** | −.08 (3.79)*** | .55 | .02 (0.74) | .67 (29.24)*** | .45 |
| Lag, 13+ | 4 | 991 | .75 (35.15)*** | .00 (0.12) | .56 | .03 (1.36) | .62 (24.96)*** | .40 |
| Specific analyses | | | | | | | | |
| JS, extra-role | 4 | 848 | .80 (37.33)*** | .00 (0.20) | .64 | .04 (1.64) | .67 (25.51)*** | .46 |
| Lag, 1–6 | 1 | 77 | .83 (12.70)*** | .06 (0.90) | .74 | .11 (1.70) | .81 (12.50)*** | .75 |
| Lag, 7–12 | 1 | 370 | .81 (22.13)*** | −.07 (1.91) | .63 | .07 (1.32) | .51 (9.98)*** | .29 |
| Lag, 13+ | 2 | 401 | .78 (25.10)*** | .05 (1.67) | .63 | .00 (0.14) | .86 (33.16)*** | .74 |
| JS, in-role | 10 | 2,026 | .73 (47.71)*** | .00 (0.33) | .53 | .03 (1.76) | .61 (34.15)*** | .37 |
| Lag, 1–6 | 4 | 730 | .78 (34.47)*** | .05 (2.36)* | .63 | .10 (3.56)*** | .61 (21.01)*** | .40 |
| Lag, 7–12 | 5 | 770 | .73 (28.08)*** | −.07 (2.69)** | .51 | .00 (0.12) | .74 (29.81)*** | .55 |
| Lag, 13+ | 2 | 590 | .69 (23.07)*** | .05 (1.64) | .48 | −.03 (0.74) | .58 (17.32)*** | .34 |
| OC, extra-role | 3 | 558 | .85 (35.81)*** | .03 (1.35) | .71 | .03 (1.20) | .86 (38.29)*** | .74 |
| Lag, 1–6 | 1 | 157 | .71 (11.96)*** | .04 (0.74) | .49 | .05 (1.31) | .85 (20.81)*** | .76 |
| Lag, 13+ | 2 | 401 | .88 (36.23)*** | .00 (0.16) | .77 | .03 (1.24) | .85 (32.71)*** | .74 |
| OC, in-role | 2 | 572 | .85 (38.78)*** | .05 (2.14)* | .73 | .09 (2.77)** | .60 (18.11)*** | .37 |
| Lag, 1–6 | 1 | 46 | .90 (12.84)*** | .16 (2.30)* | .78 | .27 (6.01)*** | .87 (19.74)*** | .91 |
| Lag, 13+ | 1 | 526 | .85 (37.27)*** | .06 (2.43)* | .73 | .08 (2.19)* | .58 (16.46)*** | .34 |

*Note.* Standardized path coefficients, with *t* scores in parentheses. JS = job satisfaction; OC = organizational commitment; Lag = time lag between the coded measurement waves in months.
* $p < .05$. ** $p < .01$. *** $p < .001$.

scores on these measures and may have inflated their correlations with job attitudes in particular, thereby introducing bias in the regression analyses. This possibility, together with the small number of studies available, renders the present findings on extra-role performance particularly tentative. Future research on the causal link between job attitudes and extra-role performance should use observer ratings or objective indicators as alternative or additional measures.

One limit to the generalizability of the data is that the majority of the studies were from English-speaking countries, primarily the United States. Thus, the generalizability to other countries is unclear. Moreover, due to the small number of studies, it was not possible to include studies that measured performance at the firm, rather than individual, level. It is still possible that job attitudes and performance show stronger, or different, relations at aggregated than at individual levels (see Koys, 2001, and Schneider, Hanges, Smith, & Salvaggio, 2003, for examples of firm-level analyses).

Finally, because the present analysis is based on correlational rather than experimental data, it allows for only tentative causal conclusions and cannot rule out some alternative causal explanations (e.g., that third variables inflated the cross-lagged paths; see, e.g., Cherrington, Reitz, & Scott, 1971; Brown & Peterson, 1993). Although the present analysis accomplished a more rigorous test for causality than did previous meta-analyses in this domain, it still suffers from the usual weakness of correlational designs. Experiments are required to provide compelling evidence of causal relations.

### Implications for Research

Notwithstanding the limitations mentioned above, this is the first meta-analysis to predict later performance by job attitudes over and above initial performance and vice versa. The findings, obtained using a comprehensive database, support the common view that job attitudes increase performance. They provide weaker support for the alternative view that performance shapes job attitudes. Effects of this latter type remain elusive.

At first glance, this conclusion seems to be at odds with Judge et al.'s (2001) assertion that there is evidence for effects in both directions, which led them to propose a model of reciprocal satisfaction–performance relationships (p. 390). Several points are worth mentioning in this regard. For one, the present findings represent average trends across studies and do not rule out reciprocal or mere performance–attitude effects in specific contexts. Further, many studies that claim to test causal relations between job attitudes and performance are cross-sectional and thus are able to provide only weak evidence of causality. In fact, Judge et al.'s conclusion is based on a qualitative review of 4 longitudinal and 12 cross-sectional studies (pp. 377–379). Finally, even longitudinal studies, especially those published in the 1970s and 1980s, often use inappropriate methods to test for causality (see Williams & Podsakoff, 1989). Rather than taking the analyses and interpretations in the primary studies at face value, the present research reanalyzed their correlation matrices, thus overcoming some limitations of the studies.

Though statistically significant in several cases, the effects of job attitudes on performance were generally weak in the present analyses (e.g., $\beta = .06$ overall, .03 for satisfaction, .08 for commitment). One reason may be that these effects are short lived. As might be expected (see Hulin, Henry, & Noon, 1990), the effects were stronger for shorter time lags between measurement waves, though only slightly so (e.g., $\beta = .11$ for time lags of 6 months or less). It is noteworthy that the time intervals in the analyzed studies were quite long, with a range of 1.5 months to 18 months. Research still has to explore whether stronger effects emerge for shorter intervals (e.g., a few days). In general, to the extent that time lags are longer than the actual duration of the effects of interest, these effects are reflected in the cross-sectional, not the cross-lagged, relationships (Clegg et al., 1977). More theoretical and empirical work on the temporal characteristics of job attitude–performance effects is necessary to help researchers choose optimal time intervals.

Furthermore, cross-lagged effects between job attitudes and performance may be stronger under certain circumstances (see also Judge et al., 2001). Possible moderators of these effects include autonomy at work (Kalleberg & Marsden, 1995; Riketta, 2002); the stimulating nature of jobs (Ivancevich, 1979); organizational tenure (Cohen, 1993; Wright & Bonett, 2002); and the degree to which job attitude and performance are measured with regard to the same, rather than different, targets (Fishbein & Ajzen, 1974; Lee, Carswell, & Allen, 2000; Riketta & Van Dick, 2005; Vandenberghe, Bentein, & Stinglhamber, 2004). The studies included in the present meta-analysis did not provide enough information to test the effects of these moderators. Additional panel studies including these moderators would be valuable.

A theoretical challenge is to explain the counterintuitive negative effect of performance on job satisfaction for moderate time lags ($\beta = -.08$, time lag = 7–12 months). One explanation is that the effect is due to people who perform strongly but who do not perceive themselves to be adequately rewarded for their performance. In this case, high performers may be less satisfied than are low performers, especially when enough time has passed to stifle their hope for performance-adequate rewards. This would explain why the effect is evident only for longer time lags. Thus, reward systems and justice perceptions may be additional moderators, especially for job attitude–performance effects. Because the effect is weak and was based on only six studies, replication attempts are advisable before further interpretation is attempted.

From a practical perspective, the longitudinal effects of job attitudes on performance might be weaker than many practitioners have hoped. Thus, researchers should be cautious with practical recommendations such as, "Increase job satisfaction or commitment to increase productivity." Especially when effect sizes are small, it is important to communicate them so that practitioners understand their practical significance (McCartney & Rosenthal, 2000). A useful tool for this purpose is the binomial effect-size display (Rosenthal & Rubin, 1982). Translated into this display, the present finding of a job attitudes–performance effect of $\beta = .06$ means that, among the half of employees with higher job attitudes, 53% also belong to the half with higher performance; whereas, among the half of employees with lower job attitudes, 47% belong to the half with higher performance. It is up to practitioners to decide, on the basis of this information, whether their organizations would substantially benefit from higher job attitudes.

### *Conclusions*

This meta-analysis provides some support for effects of job attitudes on performance and little support for the reverse effects. In light of the weak effects observed in the present analyses, future research should focus on moderators of the relations between job attitudes and performance. The present findings suggest timing of measurement as one moderator, with effects being more likely to emerge for shorter (vs. longer) time spans. Moreover, research should increasingly use panel designs to broaden the database for follow-up meta-analyses on the causal link between job attitudes and performance.

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