Abstract

Background: Recent research postulated that organizational identification plays an important role in employees' health and well-being. Building on the Social Identity Approach as a framework, we test the so-called *social cure* hypothesis, according to which group-based processes of social support should reduce employees' psychological distress.

Design and Methods: While there is a considerable amount of cross-sectional evidence concerning the positive role played by organizational identification in this dynamic, there is a lack of full panel studies. This study tries to fill this gap by using data from a sample of technical and administrative staff of a University in Italy at three time points (N = 96). Data were analyzed using Autoregressive Cross-Lagged Panel models.

Results: We found support for the hypothesized longitudinal mediational model. Specifically, strongly identified employees tend to receive more social support, and this in turn reduces psychological distress over time.

Conclusions: This study is the first test of the social cure hypothesis in an organizational context that uses a panel study design. We discuss the theoretical and practical implications for management.

Keywords: Social identity model of stress, organizational identification, psychological distress, social support, panel design, social cure hypothesis

The mediational effect of social support between organizational identification and employees' health: A three-wave study on the social cure model

Since Ashforth and Mael (1989) published their seminal study on the application of the Social Identity Theory in order to understand and manage many important organizational phenomena, hundreds of papers have been published on issues such as voluntary turnover or extra-role behavior (Lee et al., 2015; Riketta, 2005). According to Ashforth and Mael (1989), organizational identification represents "the perception of oneness or belongingness to some human aggregate" (p. 21). Several meta-analyses have highlighted the role played by organizational identification to explain the employees' intention to remain, their affective commitment and involvement, as well as many important organizational outcomes, such as job satisfaction, extra-role behaviors, and more recently employees' health and well-being (e.g., Lee et al., 2015; Ng & Allen, 2018; Steffens et al., 2017). Despite this large amount of empirical research, longitudinal evidence is still missing, and this poses a conceptual as well as empirical problem. Indeed, even if it is often assumed that the results of cross-sectional studies will also hold longitudinally, this is not always the case. Indeed, robust cross-sectional evidence of a direct relation between predictor and criterion variables could disappear in a longitudinal model (Maxwell & Cole, 2007).

In the limitations section of his meta-analysis on the relation between commitment, trust, and identification with several outcomes (e.g., job satisfaction, turnover intentions, citizenship behaviors), author concluded that most studies utilized cross-sectional designs (Ng, 2015). Ng and Allen (2018) drew the same conclusion in their meta-analysis of over 400 independent samples on the association between organizational attachment and better subsequent health. Lee and colleagues' (2015) meta-analysis on the relation between identification and attitudes/behaviors (such as satisfaction and performance) conducted on 114 independent samples, found only one study using panel data and 10 studies using longitudinal data with time separation between identification and outcomes. Finally, a more recent meta-analysis exploring the association between identification and employees' well-being conducted on 58 samples, found only five longitudinal

studies (Steffens et al., 2017). The authors concluded their limitations section by stating: “There is a need for more studies that employ experimental and intervention as well as longitudinal designs to examine the impact of organizational identifications on individuals’ health” (p. 25).

Hence, the present study is an attempt to respond to the call for more longitudinal investigations in this area of research. In a three-wave study, we will test the hypothesis that the effect of organizational identification on employees’ psychological distress is mediated by social support. By definition, “mediation implies change over time” (Maxwell & Cole, 2007, p. 24). Thus, three-wave longitudinal studies are essential to fully probe a mediation model. Furthermore, by using this research design we will be able to also investigate the bidirectional and reverse causality among variables (see Saeri et al., 2017).

Confirming this mediational path longitudinally could have important theoretical and practical implications. For example, the enhancement of an employee's sense of being identified to their organization could drive performance, but at the same time better social connection among employees. In this vein, colleagues’ support may be a potential resource for helping employees to deal with stressful events or conditions, reappraising them in fruitful ways, and preventing them to report unhealthy consequences. This could be especially true for less experienced employees who face many difficulties in their early career, but also for employees in general, when their organizations undergo substantial change.

Social Identity Approach in Organizational Contexts

The Social Identity Approach consists of two related but independent theories: Social Identity Theory and Self-Categorization Theory (Haslam, 2004). It argues that groups are not mere external characteristics of the social environment; rather, they are internalized in a group member’s social identity and as a consequence groups affect how individuals perceive, react and manage social reality itself. Being a member of a specific group partly answers the individual’s question of ‘Who am I?’, contributing to his/her self-definition. Thus, social identities represent that part of an individual’s identity, which derives from his/her social affiliations (Tajfel & Turner, 1979). People

belong to many groups simultaneously and they shift from one identity to another according to social cues, which are able to activate one identity rather than another. However, some identities tend to be more accessible than others because they are more valued or important for that person (Hogg & Terry, 2000; Turner, Hogg, Oakes, Reicher, & Wetherell, 1987). Being a member of a certain group enables the in-group members to understand the otherwise incomprehensible and chaotic world around them, to be aware of the norms and constraints regulating correct action within that world, and to be able to mobilize resources from the group to manage the difficulties and the tasks. This approach predicts that there is a qualitative difference between behaviors that are based on the activation of one's personal identity versus his/her social identity.

In organizational contexts, this implies that when employees perceive themselves as members of an organization, department, or team, their attitudes and behaviors will be strongly influenced by those affiliations (Ashforth et al., 2008). Indeed, strongly identified employees tend to work harder and to put effort into achieving the *organizational* goals and aims, because these become their *personal* goals and aims. Identification acts as an employee's personal driver. It increases employees' motivation and organizational loyalty; and at the same time, it decreases their intention to quit, given that it represents a strong link between employees and their organization. There is abundant empirical evidence showing the positive impact of strong identification on important job-related attitudes and behaviors, such as commitment, job involvement, turnover intention, in- and extra-role behaviors (Lee et al., 2015; Ng & Allen, 2018; Riketta, 2005). Identification plays a role in terms of not only employees' performance but also their (ill)-health as outlined in the social cure model of social identity.

The *Social Cure* model

Being a member of a group also activates the so-called *social cure* function of one's social identity (Jetten et al., 2012; Haslam et al., 2018). The core assumption of this model is that the people's group memberships – and the related social identities – "have important consequences on their health and well-being" (Jetten et al., 2017, p. 790). Belonging to social groups "makes people

stronger and healthier because they [*group memberships*] provide them with, among other things, self-esteem, belonging, meaning, and a sense of purpose, control, and efficacy in life" (Jetten et al., 2017, p. 972). In particular, people could benefit from their memberships, only to the extent they develop strong identification with that social group. For example, Saeri et al. (2017) found that social connectedness – comprising belongingness, the absence of loneliness, and perceived social support – predicted health over time in a cross-lagged design in a large national sample in New Zealand (N = 21,227).

The social cure model comprises a complex and large body of theoretical and empirical research. For example, it predicts many conditions under which and mechanisms through which social identities could affect a person's well-being and health. For example, the high status of the group to which people belong, in comparison with the status of other relevant out-groups, can be beneficial for their well-being. On the contrary, when people perceive a threat against their social identities (e.g., social discrimination) this could have negative consequences on their health. The social cure model also predicts a series of expectable behaviors in terms of social creativity, social competition, or personal mobility depending on the perceived permeability of group boundaries (Jetten et al., 2017).

Particularly interesting for organizational contexts and for our hypotheses, organizational identification affects employees' health and well-being both by satisfying important needs, and because a strong sense of identification between colleagues increases team cohesiveness which in turn increases social support among team members and a sense of collective self-efficacy (van Dick & Haslam, 2012).

Being a member of a social group satisfies important human needs, such as the need to belong, the need for uncertainty reduction, and that of achieving collective self-esteem among others (Ashforth et al., 2008). Researchers have shown that an increase in identification with a group or gaining a new group identity leads to satisfaction of human needs, such as control over

one's own environment, belonging needs or meaningful existence needs, and this in turn reduces depression and increases life satisfaction among students (Greenaway et al., 2016; Kyprianides et al., 2019). Coherently, in organizational literature there is a positive relation between identification and employees' health, satisfaction, and well-being as shown in a recent meta-analysis conducted on about 60 studies (Steffens et al., 2017). Moreover, strong identification increases cooperation and collaboration among in-group members, and this is at the basis of collective action. Employees should be more likely to interpret and manage job tasks and workloads collectively - as "regarding all of us" - instead of individually - "regarding only me" (Junker et al., 2019; van Dick & Haslam, 2012). While many efforts have been made to test the relation between identification and health and well-being, less empirical evidence has been obtained in order to understand "the psycho-social mechanisms underlying the social identity-(ill-)health link" (Junker et al., 2019, p. 991).

In particular, many mechanisms relating to identification and employees' well-being have been hypothesized, and these mechanisms are not necessarily in contrast to each other. Indeed, identification was predicted to increase employees' well-being by increasing received support (Haslam et al., 2005). Following Haslam et al.'s (2005) arguments, social support can reduce the adverse effects of stress on employees' health through four functions. Specifically, social support could provide: a) a sense of acceptance, b) connection with others, c) practical assistance, and d) information on how to cope with stressors. For these reasons, strong identification increases the probability of receiving (Levine et al, 2005) and positively interpreting (Frisch et al., 2014) support from colleagues, and strong support represents itself as a coping strategy to face both stressors and strain (Viswesvaran et al., 1999). In fact, employees who receive support from colleagues can mobilize group resources against stressors when needed. Furthermore, receiving encouragement and aid from co-workers acts as an emotion regulation strategy, helping employees to reappraise negative conditions or events in more constructive terms. Thus, we predict that social support from colleagues acts as a mediator in the relation between employees' organizational identification and health (Haslam, 2004; van Dick & Haslam, 2012).

Although there is empirical evidence of the mediating role of social support, we noted a lack of full longitudinal design studies that tested these relationships. In the present study, we will test the hypothesis based on a full cross-lagged design across three waves. In particular, according to this hypothesis strong organizational identification should directly increase colleague support and indirectly reduce psychological distress. We tested this hypothesis using a full longitudinal design, in which each variable was collected on each occasion for three time points.

Method

Organizational Context

Following various normative changes in public administration in Italy, in 2012 the national anti-corruption authority (ANAC, *Agenzia Nazionale Anti-Corruzione*) was asked by the Italian Ministry of Economy and Finance to coordinate the assessment of employees' well-being in Italian public services. A commission established by ANAC developed a survey in order to investigate many important work-related aspects (e.g., well-being, assessment system sharing, and supervisor evaluation). This survey was thought and designed for practical purposes rather than research ones, but a group of Italian researchers has recently provided an initial evaluation in terms of its psychometric characteristics (see Loera et al., 2018). We participated in the coordinating team of an Italian University which implemented the (ANAC) survey among its employees. We also had the chance to include some additional scales to the standard ANAC survey (e.g., general health questionnaire) to the final questionnaire.

The University conducted the first evaluation in 2014. Then, a new evaluation was conducted each following year. Unfortunately, in 2015 for a mere technical problem, the participants' identification code was not used, and thus we were unable to use the data of that year. Thus, our total data comprises Time 1 (2014) data and two further waves in 2016 and 2017, respectively.

Participants

The survey was sent to all the employees (technical and administrative staff) working at a University in Italy (697 employees in the first year). Questionnaires were matched by using anonymous codes that respondents created from personal information. This "personal code" consisted of alphanumeric characters that only the participants could know and was simple enough to be remembered (e.g., "*write the first two letters of your mother's maiden name*"). Participation was voluntary and 412 employees completed the questionnaire at Time 1, 385 at Time 2, and 425 at Time 3. For each occasion (T1, T2, and T3), employees were invited to participate in the survey by means of an email sent by the Delegate for Organizational Wellbeing of the University. The email included a statement in which they were asked to voluntarily participate in the survey (through a link) and it was specified that data would be anonymized. The total longitudinal sample comprised 96 employees who participated in each survey at the three time points (13.77% response rate). Most of the samples consisted in women (68.8%) and age was distributed as follows: 40.6% aged under 40, 44.8% between 40 and 50, and 12.5% were aged above 51 years. Of the final sample, 34.4% had less than 10 years of work experience, 41.7% between 11 and 20 years, and 21.9% more than 20 years.

In order to assess potential differences between employees who completed all surveys (T1, T2, and T3) and those who participated only at Time 1, a series of T-tests were conducted on organizational identification ($t_{(185.595)} = 1.031$, $p = .304$), colleague support ($t_{(375)} = 0.799$, $p = .425$), and psychological distress ($t_{(426)} = -0.032$, $p = .974$). Furthermore, two χ^2 tests were performed to determine whether participants' distribution with respect to age ($\chi^2_{(4)} = 7{,}863$, $p = .097$) and gender ($\chi^2_{(1)} = 0{,}199$, $p = .656$) varied between waves. Such analyses didn't show statistically significant results. Thus, the final longitudinal sample did not deviate from the larger sample at any of the study variables.

Measures

Organizational identification. Organizational identification was measured with four items gathered from the dimension of the survey "*sense of belonging*"[1]. Responses were given on a 6-point scale, ranging from 1 (*not at all*) to 6 (*completely*). Sample items were: "*I am proud when my organization achieves a good result*", "*The values and behaviors practiced in my organization are consistent with my personal values*" (α_{T1} = .79; α_{T2} = .90; α_{T3} = .89).

Colleague support. Colleague support was measured with four items gathered from the dimension of the survey "*my colleagues*". Responses were given on a 6-point scale, ranging from 1 (*not at all*) to 6 (*completely*). Sample items were: "*I am valued and treated with respect by my colleagues*", "*In my group, those who have information make it available to everyone*" (α_{T1} = .83; α_{T2} = .79; α_{T3} = .85).

Psychological distress. Psychological distress was assessed with the GHQ-12 (General Health Questionnaire; Goldberg, 1972; Italian version by Piccinelli et al., 1993). This scale consists of 12 items with responses ranging along a 4-point Likert scale, with different labels for items worded in the negative (0 = *not at all*, 3 = *much more than usual*) or in positive terms (0 = *better than usual*; 3 = *much less than usual*). By using the conventional binary scoring, the "0" and "1" answers were coded as 0, meaning "*no symptom or healthy situation*" while answers from "2" to "3" were coded as 1, meaning "*presence of symptom or unhealthy situation*". High scores corresponded to high psychological distress. An example of a negative item is: "*Have you recently*

[1] To further test the validity of this measure, we administered this scale to 102 master students through an online procedure, together with the most commonly used six-item scale by Mael & Ashforth (1992) plus one item of organizational citizenship behaviors ("*I help new colleagues to be guided even if not asked*"). Some example items for this scale are: "*I am proud when my university achieves a good result*", "*The values and behaviors practiced in my university are consistent with my personal values*" (α = .87). Some example items for Mael & Ashforth's scale are: "*When someone criticizes my university, it feels like a personal insult*", "*When I talk about this university, I usually say 'we' rather than 'they'*" (α = .86). The correlation between these two identification scales was substantial and significant (r = .86, p < .001), and both scales correlated to about the same extent with the OCB item (r = .46, and r = .55, p < .001, for the Mael & Ashforth, and for this scale, respectively). Thus, our scale showed an acceptable degree of construct validity. With regards to (lack of) discriminant validity, we demonstrated equivalence of the two scales via CFA approach. In particular, we ran a CFA with two latent factors representing the two scales (i.e., one latent factor was composed by the six items of the Mael and Ashforth's scale and the other latent factor was composed by the four items of the scale used in this study). The 2-factor model returned a correlation between latent factors of .992 (98.4% of shared variance). Moreover, the 1-factor model is equivalent but more parsimonious then 2-factor model (2-factor model: χ^2 = 84.954, df = 34 p < .001; CFI = .926, AIC = 2926.088. 1-factor model: χ^2 = 85.086, df = 35 p < .001; CFI = .927, AIC = 2924.220) and shown a lower AIC. Hence, this analysis supports the two scales are highly overlapped.

lost confidence in yourself?"; an example of a positive item is: "*Have you recently been able to concentrate on what you were doing?*" (α_{T1} = .90; α_{T2} = .87; α_{T3} = .88).

Covariates. We also controlled for age and gender, because previous studies had found effects of both variables on employees' psychological distress (e.g., Ng & Feldman 2010). We did not use organizational tenure since this variable was strongly correlated with age (r_s = .66). Following Becker's (2005) suggestions, we also conducted all analyses with and without controls, which yielded very similar results (see below).

Data Analytic Strategy

In testing our theoretical model, we used a three-wave autoregressive cross-lagged panel model with observed variables. In this mediational model the indirect effect is computed as the product of two cross-lagged paths, namely (a) the path from predictor variable at T1 (i.e., organizational identification) to mediator T2 (i.e., colleague support) and (b) the path from mediator T2 to the criterion variable at T3 (i.e., psychological distress).

Because of the small sample size, we decided to reduce the number of freely estimated parameters by using the observed variables (e.g., Balducci et al., 2020; Xanthopoulou et al., 2009). Nonetheless, before testing our mediational model, we checked the measurement model and the metric invariance across time for each variable. Since we used the conventional mode to compute the GHQ score, the values of our items were binary (0 or 1); hence for this variable we tested the metric invariance by using the weighted least squares mean and variance adjusted (WLSMV) estimator.

We then tested a series of competing models. We started by testing the model in which all autoregressive paths and all cross-sectional covariances among observed variables at the same time, and all cross-lagged paths were estimated, without any constraints (M1). In the subsequent three models, we constrained one pair of paths at a time in order to test whether the reversed models were good models as well. For example, a possible reversed model was the one in which psychological distress was the predictor, colleague support the mediator, and organizational identification the

criterion variable. We controlled for all possible reversed models. In particular, in the second model (M2) we fixed to zero the paths from colleague support T1 to organizational identification T2 and from colleague support T2 to organizational identification T3. In the subsequent model (M3) we fixed to zero the path from psychological distress T1 to organizational identification T2 and from psychological distress T2 to organizational identification T3. In the fourth model (M4) we fixed to zero the path from psychological distress T1 to colleague support T2 and from psychological distress T2 to colleague support T3. This approach starts from the less parsimonious model and then moves toward more parsimonious models only after investigating that fixing paths to be zero did not worsen the model. In this way, no cross-lagged relationship (even those hypothesized to be non-significant) was neglected (i.e., fixed to be zero) without attesting that the zero-constrain did not worsen the whole model.

The Statistical Package for Social Science (SPSS) 25 and M*plus* 8 (Muthén & Muthén, 1998-2017) were used to perform the analyses. The goodness of fit of each model was evaluated using the χ^2 test, the Comparative Fit Index (CFI), the Tucker-Lewis Index (TLI), the Root-Mean-Square Error of Approximation (RMSEA), and the Standardized Root Mean Square Residual (SRMR). For CFI and TLI we considered as acceptable values > .90, for RMSEA and SRMR values < .08. For model comparison purposes, we accepted a nested model if constraints did not significantly worsen the previous model, as evidenced by a non-significant likelihood ratio test ($\Delta\chi^2$) and by a difference in CFI (ΔCFI) lower than .01 (Cheung & Rensvold, 2002). Finally, the critical values for the upper and lower confidence limits for indirect effects were computed using the bias-corrected bootstrap method on 5000 resamplings. We considered an effect being significant, when 95% confidence intervals not including the value of 0 (Preacher & Hayes, 2008).

Power Consideration

In order to test the adequacy of the statistical power of our sample for investigating the hypothesized mediational paths (organizational identification T1 → colleague support T2; colleague support T2 → psychological distress T3), we ran a simulation analysis. In particular, we specified

an autoregressive cross-lagged panel model with three constructs across 3 time-points. Latent variables were specified to be composed by one observed indicator, latent variances were all fixed to be 1, and residual variance for each observed indicator was fixed to be 0; in this way unstandardized and standardized estimates (for both autoregressive and cross-lagged paths) have equal values, hence facilitating interpretation. We specified all autoregressive paths to be 0.50 (|.50| corresponds to a large effect size according to Cohen, 1992), whereas hypothesized cross-lagged paths for "organizational identification T1 → colleague support T2" and "colleague support T2 → psychological distress T3" to be 0.25 and -0.25, respectively (|.25| corresponds to low/medium effect size, according to Cohen, 1992). This analysis was run with pwrSEM (Wang & Rhemtulla, in press). A spreadsheet for replicating this analysis including the full results obtained is reported in supplementary material (Appendix A). Results showed that with 1000 simulations, N = 96, and α = .05, power for organizational identification T1 → colleague support T2 was .86, whereas power for colleague support T2 → psychological distress T3 was .85, thus over the recommended threshold of .80 (Boomsma, 2013; Muthén & Muthén, 2002; Wolf et al., 2013).

Results

Table 1 reports the means, standard deviations, and correlations among the variables studied. Regarding concurrent correlations, at the cross-sectional level all correlations were in the expected direction with organizational identification positively related to colleague support, and with both organizational identification and colleague support negatively related to psychological distress. All but one correlations were significant, ranging from r = -.18, p = .08 (organizational identification T1 with psychological distress T1) to r = .57, p < .001 (organizational identification T3 with colleague support T3). Regarding the longitudinal correlations, each variable showed a certain level of construct stability with re-test correlations ranging from r = .50 (colleague support T1 with colleague support T2) to r = .82 (organizational identification T2 with organizational identification T3), p_s < .001. Finally, regarding the cross-lagged correlations, from T1 to T2 they ranged from r = -.10, p = .322 (psychological distress T1 with organizational identification T2) to r = .39, p < .001

(organizational identification T1 with colleague support T2), and from T2 to T3 they ranged from r = -.16, p = .119 (organizational identification T2 with psychological distress T3) to r = .50, p < .001 (colleague support T2 with organizational identification T3).

Regarding the measurement models, we computed two separate models, one for organizational identification and one for colleague support, with four items each loading on a single latent construct. We found evidence of configural invariance across the three time points, with acceptable fit for both organizational identification (χ^2 = 59.484, df = 39, p = .019; CFI = .975; TLI = .958; RMSEA = .074; SRMR = .061), and colleague support (χ^2 = 68.607, df = 39, p = .002; CFI = .952; TLI = .918; RMSEA = .089; SRMR = .067). Each factor loading was statistically significant at p < .001, and the size of the factor loadings across time ranged from .57 to .95 for organizational identification and from .52 to .92 for colleague support. Furthermore, by constraining the factor loadings for each item to be equal across three time points, we found also evidence for metric invariance. In particular, the metric invariance model for organizational identification showed an adequate fit to the data (χ^2 = 62.145, df = 45, p = .045; CFI = .979; TLI = .970; RMSEA = .063; SRMR = .066), and it did not worsen in comparison with the configural invariance model fit: $\Delta\chi^2$ = 2.661, df = 6, p = .850; ΔCFI = -.004. Moreover, for colleague support the metric invariance model showed an adequate fit to the data (χ^2 = 77.670, df = 45, p = .002; CFI = .947; TLI = .922; RMSEA = .087; SRMR = .080), and it did not worsen in comparison with configural invariance model fit ($\Delta\chi^2$ = 9.063, df = 6, p = .170; ΔCFI = .005).

Regarding GHQ, the fit of the configural model was good (WLSMV-based χ^2 = 745.651, df = 591, p < .001; CFI = .954; TLI = .951; RMSEA = .052). Factor loadings were all significant at p < .001 and ranged from .62 to .96. By constraining the factor loadings for each item to be equal across three time points, we found also evidence for metric invariance. In particular, the metric invariance model for GHQ showed an adequate fit to the data (WLSMV-based χ^2 = 762.074, df = 613, p < .001; CFI = .956; TLI = .955; RMSEA = .050), and it did not worsen in comparison with the configural invariance model fit: ΔWLSMV-based χ^2 = 32.058, df = 22, p = .076; ΔCFI = -.002.

The results of the hypothesis test is depicted in Table 2. As can be seen, the model without any constraint (M1) showed a good fit to the data (χ^2 = 12.794, *df* = 9, *p* = .172; CFI = .99; TLI = .97; RMSEA = .066; SRMR = .022). In Model 2 (M2) we constrained the paths from colleague support to organizational identification in both time lags (T1-T2 and T2-T3) to zero; both such paths were not statistically significant in M1 (β = .09, *p* = .343, β = .00, *p* = .996, for T1-T2 and T2-T3 path, respectively). The fit of the model was adequate and the constraints did not significantly worsen the model fit ($\Delta\chi^2$ = 1.025, *df* =2, *p* = .599). In Model 3 (M3) we constrained the paths from psychological distress to organizational identification in both time lags (T1-T2 and T2-T3) to zero; in M2 the first path (T1-T2) was not statistically significant (β = .00, *p* = .881), while the second (T2-T3) was small but significant (β = -.05, *p* = .026). Again, the model fit did not significantly worsen ($\Delta\chi^2$ = 4.865, *df* = 2, *p* = .088).

In the fourth model (M4) we constrained the path from psychological distress to colleague support in both time lags (T1-T2 and T2-T3) to zero; in M3 the first path (T1-T2) was small but statistically significant (β = -.07, *p* = .006), while the second (T2-T3) was small and not significant (β = -.05, *p* = .070). However, in this case the model fit significantly worsened ($\Delta\chi^2$ = 10.410, *df* =2, *p* = .005).

Thus, Model 3 is the preferred model (see Figure 1); M*plus* syntax for replicating Model 3 is available in supplementary material (Appendix B). As can be seen in Figure 1, organizational identification at T1 significantly predicted colleague support at T2 (β = .21, *p* = .024), and in turn, colleague support at T2 significantly predicted psychological distress at T3 (β = -.21, *p* = .016). The sign and the significance of these paths supported our mediational hypothesis. To corroborate our results, findings from a bootstrap analysis with 5000 resamples attested to the significance of the indirect effect (unstandardized indirect effect = -0.158; 95% CI: -0.459, -0.015). The final model, without covariates, explained a substantial amount of variance in psychological distress T3, i.e. 51.8% (R^2 = .518, *z* = 7.157, *p* < .001). None of the theoretically alternative mediational models was

supported, considering that – for each alternative model – either path *a* (from predictor T1 to mediator variable T2) or path *b* (from mediator T2 to criterion variable T3) was not significant.

We tested our final model (M3) also with gender and age as covariates, but the results remained substantially unchanged and confirmed our expectations for both model fit (χ^2 = 18.60, *df* = 13, *p* = .136; CFI = .99; TLI = .95; RMSEA = .068; SRMR = .035) and indirect effect (-.183; 95% CI: -0.477, -0.035). Specifically, age was uncorrelated to any of the other variables across times, while gender (1 = *female*, 2 = *male*) showed a significant relation only with organizational identification T1 (β = .26, *p* = .008) and with psychological distress T1 (β = -.25, *p* = .008).

Discussion

A large amount of evidence originating from different literature highlights the role of social factors for (good) human health. For example, incorporating the results from about 150 studies, authors showed that social integration and social support are two of the most important predictors of lower mortality, as evidenced by odd-ratios indicating a large effect size and larger than other important and more established risk factors such as smoking, physical activity, or obesity (Holt-Lunstad et al., 2010). The authors concluded that "social relationship factors [*have*] to be added to" the list of more traditional risk factors such as nutrition, smoking or exercising (p. 14). Work and organizational psychology literature has extensively studied psychosocial factors with respect to their potential impact on both employees' performance and health – however, appropriate research designs to establish the temporal order of the relations have rarely been used (see for an exception: Crane et al, 2018).

Following the Social Identity Approach, we tested a full cross-lagged mediation model in which colleague support mediated the relationship between organizational identification and employees' health. The results fully supported our expectations and are in line with existing cross-sectional evidence. A strong sense of organizational identification indirectly reduced psychological distress over time, increasing the support perceived from colleagues. In this way, we provided evidence for a psychosocial mechanism postulated by several authors in previous theoretical and

empirical papers (i.e., Jetten et al., 2017; Junker et al., 2019; van Dick & Haslam, 2012), but we did so by using a more appropriate and relatively rare design. In fact, to the best of our knowledge, this is the first time that a full longitudinal design with three waves across four years has been used to test the aforementioned hypothesis in organizational literature. Organizational identification provides employees with a "psychological basis for receiving and benefiting from the support of other in-group members" (Haslam et al., 2005, p. 365). This is because employees will tend to recognize, accept, and correctly interpret more easily the received support when it comes from in-group members (Frisch et al., 2014; Levine et al., 2005). Colleague support in turn, provides the employee with additional resources to cope with stress, in terms of emotional, practical, and instrumental means.

Our findings highlight the importance of considering, in both evaluation and activities aiming at the prevention of work-related stress, the impact of group processes on employees' health and well-being. Sometimes, organizations seek to maximize the competition among individuals or groups in order to achieve the best performance – for instance by individual awards ("employee of the month") or bonus systems based on individual performance only. However, this orientation may have a negative role in terms of stress and employees' well-being, reducing the possibility to develop team cohesiveness and organizational identity. An employee's sense of organizational identification is an important driver of his/her efforts and extra-role behaviors, and at the same time it increases cooperation and reciprocal support among colleagues. The perception of a supportive work environment increases employees' confidence that they can successfully deal with stressors at work. Furthermore, colleague support can act as an emotion regulation strategy, helping employees cope with negative events, reappraising them in more positive and constructive terms. This could be particularly useful for younger employees in their socialization phase or for colleagues having trouble in their lives.

Some evidence of the reverse direction from psychological distress to colleague support was also apparent in our final model. In particular, there was a significant negative path from

psychological distress T1 to colleague support T2. This reciprocal relationship can mean that employees with poor psychological health tend to perceive less support over time from their colleagues. People suffering from depression or with cognitive difficulties (i.e., difficulties in concentrating, poor ability to make decisions, sleeping difficulties, and so on) may be more isolated, and it is likely that they are less able to seek support or to interpret it correctly when received, or finally, they could withdraw from social situations in general. In another context, authors found similar results, with a reciprocal relationship from distress to social connectedness being significant over time, even if lower in magnitude than the hypothesized path – from social connectedness to distress (Saeri et al., 2017). Our findings can also mean, however, that there is less actual support because distressed employees lack the capacity to recognize the support provision offered by their colleagues or, when they do perceive that support is needed, lack the capacity to actually request help. The direction of the link between support and membership on the one hand and employees' health on the other should be clarified in future experimental studies, even if it may be reasonable to assume a reciprocal influence. However, compared with alternative mediational models, only the one we predicted showed a significant indirect effect. This means that, at least in our sample, we did not find evidence either for the model in which social support mediates the relation between psychological distress to organizational identification or for the model in which organizational identification mediates the relation between psychological distress to social support. All in all, with these results we are confident in saying that a lack of connectedness related to poor health conditions, rather than representing an outcome of employees' distress, should be better considered as a health risk factor (Saeri et al., 2017). This would imply that the creation of workplace environments with high level of social identification and supportive climate should be contemplated as employees' health and well-being prevention and promotion activity.

This study has also some limitations, which may suggest new directions for future research. First, we chose the time lag of our research for convenience reasons and not for theoretical ones. A time lag of one year could be a reasonable amount to ascertain our hypothesized relations, but more

theoretical efforts should be devoted to understanding the actual role played by time in order to identify the best time lag. Moreover, Zapf et al., (1996) suggested using the same time interval between all waves in a full longitudinal design. Unfortunately, we were unable to do this, therefore our results may be at least partially biased by this non-equivalent time lag across the waves. Thus, in future research we suggest the use of other and equivalent time lags for comparison purposes. Another limitation concerns the self-reported nature of our data, which increased the likelihood of common method variance effects. However, given that our design was longitudinal, this should have reduced the risk (Doty & Glick, 1998). Nevertheless, future studies could use also objective measures: For example, a checklist to observe and quantify actual support exchanged among colleagues (Panari et al., 2012). Third, our final longitudinal sample was relatively small. Even if this situation is not so rare in longitudinal design (Ford & Tetrick, 2011; Simbula et al., 2011), future studies should test this hypothesis with larger samples. In particular, the low sample size may raise concerns in regards to adequate statistical power. Although we ran a power analyses that returned satisfactory levels of power for our hypothesized mediational paths, we cannot exclude that the low sample size may have had an impact on the failure to find an effect for those paths hypothesized to be non-significant (e.g., the path from colleague support to organizational identification). Hence, despite our efforts to investigate power analyses through a theory-driven approach, we called for future studies that use bigger sample in order to better investigate the tenability of our findings. Finally, we used psychological distress as a criterion variable that represents a measure not specifically related to work. Future studies might use more work-related measures of distress, such as burnout or psychophysical complaints.

Despite these limitations, we think that our paper contributes to the current literature by answering the call for full longitudinal design in order to provide strong empirical evidence of theoretically-based hypotheses. Our findings suggest that organizations should direct their efforts to improving employees' health and well-being not only individually, for example by promoting personal training for specific employees or through work design initiatives, but also by increasing

the identification with the organization. In particular, organizations could boost team cohesiveness by providing team-directed rewards and incentives and assigning team goals. Moreover, they could encourage initiatives designed to increase their employees' feelings of being "at home" in the workplace and part of a family, through ceremonies and other common events (Steffens et al., 2017). Again, organizations could choose and promote positive communication and leadership style which are able to increase levels of organizational identification among employees (van Dick et al., 2018). Such interventions have recently been demonstrated by a meta-analysis (Steffens et al., in press). They showed, in a summary of 27 intervention studies, that identity-increasing measures had a significant effect on a range of health and clinical outcomes in various groups of vulnerable people, such as depressive patients. We believe that our results highlight that also non-clinical populations such as employees in work settings would benefit from leaders paying more attention to group identities and team cohesiveness.

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Table 1

Means, Standard Deviations and Correlations among study variables

	M	*SD*	1	2	3	4	5	6	7	8	9
1. Organizational Identification T1	4.18	0.99	1								
2. Colleague Support T1	4.02	1.09	.46**	1							
3. Psychological Distress T1	4.63	3.52	-.18	-.28**	1						
4. Organizational Identification T2	4.35	1.07	.64**	.36**	-.10	1					
5. Colleague Support T2	4.09	1.07	.39**	.50**	-.33**	.56**	1				
6. Psychological Distress T2	4.59	3.38	-.21*	-.22*	.55**	-.30**	-.45**	1			
7. Organizational Identification T3	4.23	1.22	.59**	.31**	-.18	.82**	.50**	-.36**	1		
8. Colleague Support T3	3.96	1.24	.27**	.38**	-.31**	.41**	.64**	-.47**	.57**	1	
9. Psychological Distress T3	4.85	3.61	-.04	-.15	.51**	-.16	-.42**	.71**	-.30**	-.54**	1

Note. $^{*}p < .05$, $^{**}p < .01$, $^{***}p < .001$

Table 2

Model fits of the predicted and alternative model tests

	χ^2	*df*	*p*	CFI	TLI	RMSEA	SRMR	AIC	MC	$\Delta\chi^2$	Δdf	*p*
M1: No constraint	12.794	9	.172	0.991	0.967	0.066	0.022	2897.154				
M2: CS → OI fixed to be zero	13.819	11	.243	0.993	0.980	0.052	0.027	2894.179	M2 Vs M1	1.025	2	0.599
M3: PD → OI fixed to be zero	**18.684**	**13**	**.133**	**0.986**	**0.966**	**0.067**	**0.041**	**2895.044**	**M3 Vs M2**	**4.865**	**2**	**0.088**
M4 PD → CS fixed to be zero	29.094	15	.016	0.966	0.926	0.099	0.073	2901.454	M4 Vs M3	10.410	2	0.005

Note. CS = Colleague Support; OI = Organizational Identification, PD = Psychological Distress. CFI = Comparative Fit Index; TLI = Tucker-Lewis Index; RMSEA = Root Mean Square Error of Approximation; SRMR = Standardized Root Mean Square Residual; MC = Model Comparison. M1 = no constraint were imposed; M2 = the paths from colleague support T1 to organizational identification T2 and from colleague support T2 to organizational identification T3 were fixed to be zero; M3 = the paths from psychological distress T1 to organizational identification T2 and from psychological distress T2 to organizational identification T3 were fixed to be zero; M4 = the paths from psychological distress T1 to colleague support T2 and from psychological distress T2 to colleague support T3 were fixed to be zero.

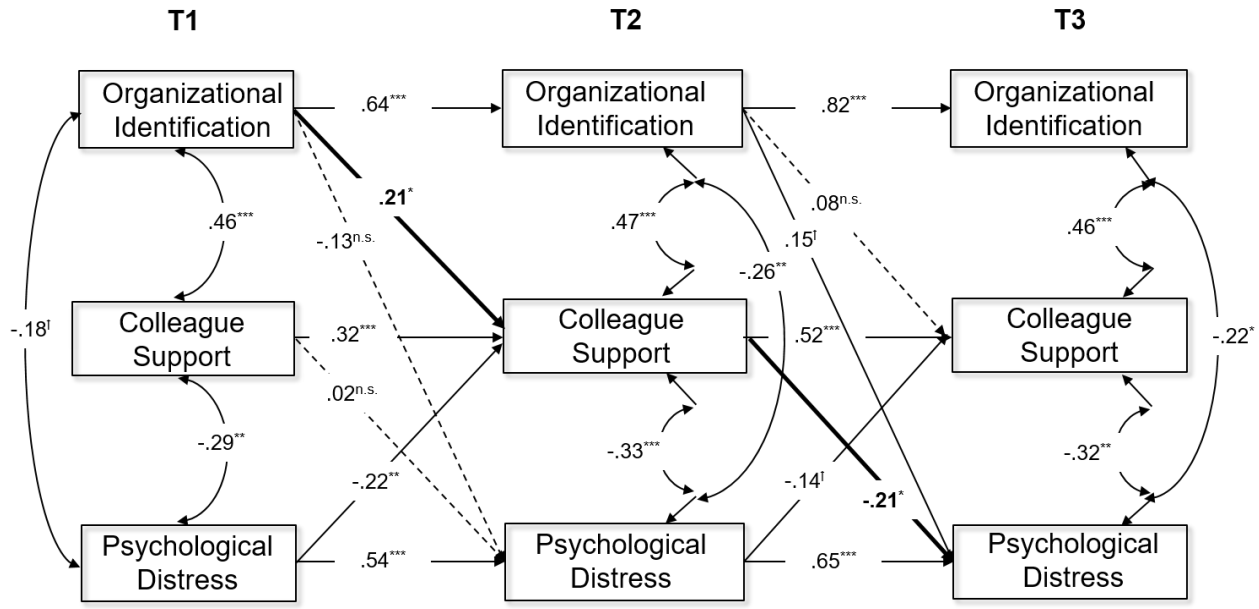


Figure 1. Best fitting model. Mediation paths were highlighted in bold. Non-significant paths are represented by dotted lines. $^{n.s.}$ not statistically significant, $^{\dagger}p$ <.10, $^{*}p$ <.05, $^{**}p$ <.01, $^{***}p$ <.001.

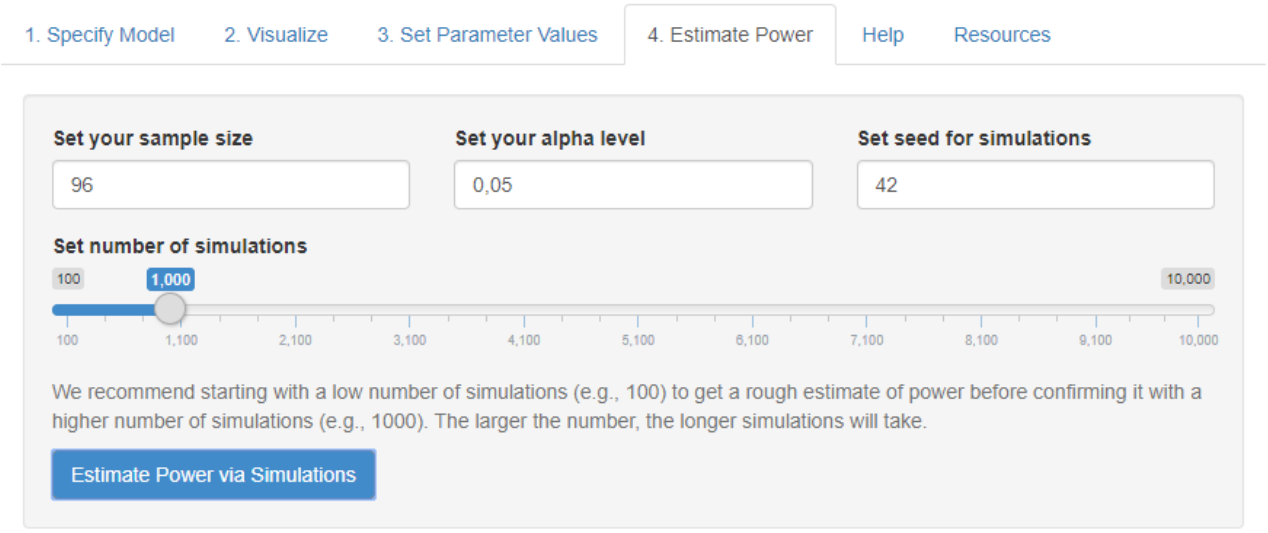
Appendix A

As following, we provided information for replicating our power analysis on pwrSEM (Wang & Rhemtulla, in press).
pwrSEM is freely available at https://yilinandrewang.shinyapps.io/pwrSEM/

```
#Hypothesized Population Model
#Measurement models
SUPP14 =~ 1*SUPP14_obs
IDORG14 =~ 1*IDORG14_obs
GHQ_14_T=~ 1*GHQ_14_T_obs
SUPP16=~ 1*SUPP16_obs
IDORG16=~ 1*IDORG16_obs
GHQ_16_T=~ 1*GHQ_16_T_obs
SUPP17=~ 1*SUPP17_obs
IDORG17=~ 1*IDORG17_obs
GHQ_17_T=~ 1*GHQ_17_T_obs

#Covariances
SUPP14 ~~ IDORG14 + GHQ_14_T
IDORG14 ~~ GHQ_14_T

SUPP16 ~~ IDORG16 + GHQ_16_T
IDORG16 ~~ GHQ_16_T

SUPP17 ~~ IDORG17 + GHQ_17_T
IDORG17 ~~ GHQ_17_T

#Autoregressive paths
IDORG16 ~ IDORG14
IDORG17 ~ IDORG16
SUPP16 ~ SUPP14
SUPP17 ~ SUPP16
GHQ_16_T ~ GHQ_14_T
GHQ_17_T ~ GHQ_16_T

#Cross-lagged paths
#IDORG16 ~ SUPP14 + GHQ_14_T
SUPP16  ~ a*IDORG14 #+ GHQ_14_T
#GHQ_16_T ~ SUPP14 + IDORG14

#IDORG17 ~ SUPP16 +  GHQ_16_T
#SUPP17  ~ IDORG16 + GHQ_16_T
GHQ_17_T ~ b*SUPP16# + IDORG16

#Indirect effect
ab := a*b
```

Row	Parameter	Label	Description	Value	Type	Effect	Free
1	SUPP14 =~ SUPP14_obs		SUPP14 is measured by SUPP14_obs	1	factor loading		
2	IDORG14 =~ IDORG14_obs		IDORG14 is measured by IDORG14_obs	1	factor loading		
3	GHQ_14_T =~ GHQ_14_T_obs		GHQ_14_T is measured by GHQ_14_T_obs	1	factor loading		
4	SUPP16 =~ SUPP16_obs		SUPP16 is measured by SUPP16_obs	1	factor loading		
5	IDORG16 =~ IDORG16_obs		IDORG16 is measured by IDORG16_obs	1	factor loading		
6	GHQ_16_T =~ GHQ_16_T_obs		GHQ_16_T is measured by GHQ_16_T_obs	1	factor loading		
7	SUPP17 =~ SUPP17_obs		SUPP17 is measured by SUPP17_obs	1	factor loading		
8	IDORG17 =~ IDORG17_obs		IDORG17 is measured by IDORG17_obs	1	factor loading		
9	GHQ_17_T =~ GHQ_17_T_obs		GHQ_17_T is measured by GHQ_17_T_obs	1	factor loading		
10	SUPP14 ~~ IDORG14		Variance of SUPP14 covaries with variance of IDORG14	0.457	covariance		1
11	SUPP14 ~~ GHQ_14_T		Variance of SUPP14 covaries with variance of GHQ_14_T	-0.286	covariance		2
12	IDORG14 ~~ GHQ_14_T		Variance of IDORG14 covaries with variance of GHQ_14_T	-0.178	covariance		3
13	SUPP16 ~~ IDORG16		Residual of SUPP16 covaries with residual of IDORG16	0.469	residual covariance		4
14	SUPP16 ~~ GHQ_16_T		Residual of SUPP16 covaries with residual of GHQ_16_T	-0.325	residual covariance		5
15	IDORG16 ~~ GHQ_16_T		Residual of IDORG16 covaries with residual of GHQ_16_T	-0.259	residual covariance		6
16	SUPP17 ~~ IDORG17		Residual of SUPP17 covaries with residual of IDORG17	0.464	residual covariance		7
17	SUPP17 ~~ GHQ_17_T		Residual of SUPP17 covaries with residual of GHQ_17_T	-0.315	residual covariance		8
18	IDORG17 ~~ GHQ_17_T		Residual of IDORG17 covaries with residual of GHQ_17_T	-0.215	residual covariance		9
19	IDORG16 ~ IDORG14		IDORG16 is regressed on IDORG14	0.5	regression coefficient		10
20	IDORG17 ~ IDORG16		IDORG17 is regressed on IDORG16	0.5	regression coefficient		11
21	SUPP16 ~ SUPP14		SUPP16 is regressed on SUPP14	0.5	regression coefficient		12
22	SUPP17 ~ SUPP16		SUPP17 is regressed on SUPP16	0.5	regression coefficient		13
23	GHQ_16_T ~ GHQ_14_T		GHQ_16_T is regressed on GHQ_14_T	0.5	regression coefficient		14

24	GHQ_17_T ~ GHQ_16_T		GHQ_17_T is regressed on GHQ_16_T	0.5	regression coefficient		15
25	SUPP16 ~ IDORG14	a	SUPP16 is regressed on IDORG14	0.25	regression coefficient	true	16
26	GHQ_17_T ~ SUPP16	b	GHQ_17_T is regressed on SUPP16	-0.25	regression coefficient	true	17
27	SUPP14_obs ~~ SUPP14_obs		Residual variance of SUPP14_obs		residual variance		
28	IDORG14_obs ~~ IDORG14_obs		Residual variance of IDORG14_obs		residual variance		
29	GHQ_14_T_obs ~~ GHQ_14_T_obs		Residual variance of GHQ_14_T_obs		residual variance		
30	SUPP16_obs ~~ SUPP16_obs		Residual variance of SUPP16_obs		residual variance		
31	IDORG16_obs ~~ IDORG16_obs		Residual variance of IDORG16_obs		residual variance		
32	GHQ_16_T_obs ~~ GHQ_16_T_obs		Residual variance of GHQ_16_T_obs		residual variance		
33	SUPP17_obs ~~ SUPP17_obs		Residual variance of SUPP17_obs		residual variance		
34	IDORG17_obs ~~ IDORG17_obs		Residual variance of IDORG17_obs		residual variance		
35	GHQ_17_T_obs ~~ GHQ_17_T_obs		Residual variance of GHQ_17_T_obs		residual variance		
36	SUPP14 ~~ SUPP14		Total variance of SUPP14	1	total variance		18
37	IDORG14 ~~ IDORG14		Total variance of IDORG14	1	total variance		19
38	GHQ_14_T ~~ GHQ_14_T		Total variance of GHQ_14_T	1	total variance		20
39	SUPP16 ~~ SUPP16		Residual variance of SUPP16	0.69	residual variance		21
40	IDORG16 ~~ IDORG16		Residual variance of IDORG16	0.75	residual variance		22
41	GHQ_16_T ~~ GHQ_16_T		Residual variance of GHQ_16_T	0.75	residual variance		23
42	SUPP17 ~~ SUPP17		Residual variance of SUPP17	0.75	residual variance		24
43	IDORG17 ~~ IDORG17		Residual variance of IDORG17	0.75	residual variance		25
44	GHQ_17_T ~~ GHQ_17_T		Residual variance of GHQ_17_T	0.69	residual variance		26
45	ab := a*b	ab	Labelled parameter	-0.063	labelled parameter		



Parameter	Value	Median	Power	Power (All Cases)
SUPP16 ~ IDORG14	0.25	0.25	0.86	0.86
GHQ_17_T ~ SUPP16	-0.25	-0.25	0.85	0.85

Convergence rate is 1. Value is the population parameter value as set in Step 3. Median is the median of simulated estimates of a parameter. Power is estimated from all simulations with converged models. Power (All Cases) is estimated from all simulations, including those with non-converged models (which had no parameter estimates and were counted as failure to reject the null).

Appendix B

As following, we provided M*plus* code for replicating our best fitting model (Model 3). In order to compute indirect effect confidence intervals with bootstrap estimation, it is necessary to remove the "!" in the following two codes: `!bootstrap=5000;` and `!cinterval(bcbootstrap);`

```
title: Model 3

data: file is Dati.dat;

variable: names are
SUPP17 IDORG17 SUPP16 IDORG16 SUPP14 IDORG14
GHQ_14_T GHQ_16_T GHQ_17_T;


MISSING = ALL(-99);

USEVARIABLES ARE
SUPP17 IDORG17 SUPP16 IDORG16 SUPP14 IDORG14
GHQ_14_T GHQ_16_T GHQ_17_T;

analysis: estimator = ml;
!bootstrap=5000;

MODEL:
!covariaces
```

```
SUPP14 with IDORG14 GHQ_14_T;
IDORG14 with GHQ_14_T;

SUPP16 with IDORG16 GHQ_16_T;
IDORG16 with GHQ_16_T;

SUPP17 with IDORG17 GHQ_17_T;
IDORG17 with GHQ_17_T;

!autoregressive
IDORG16 on IDORG14;
IDORG17 on IDORG16;

SUPP16 on SUPP14;
SUPP17 on SUPP16;

GHQ_16_T on GHQ_14_T;
GHQ_17_T on GHQ_16_T;

!cross-lagged
IDORG16 on SUPP14@0
           IDORG14
           GHQ_14_T@0;
SUPP16  on SUPP14
           IDORG14 (a)
           GHQ_14_T;
GHQ_16_T on SUPP14
            IDORG14
            GHQ_14_T;

IDORG17 on SUPP16@0
           IDORG16
           GHQ_16_T@0;
SUPP17  on SUPP16
           IDORG16
           GHQ_16_T;
GHQ_17_T on SUPP16 (b)
            IDORG16
            GHQ_16_T;

OUTPUT:
STANDARDIZED SAMPSTAT TECH1 TECH4 MODINDICES(3.84); !cinterval(bcbootstrap);

model constraint:
new(indirect);
indirect=a*b;
```