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THE ITALIAN REVISED WORK-FAMILY CULTURE SCALE: A MULTISAMPLE STUDY

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Although the Work-Family Culture Scale (Thompson, Beauvais, & Lyness, 1999) has been widely used in work-family research, its factor structure remains largely understudied and varies across countries. This study aims to examine both its construct validity by testing its measurement invariance through confirmatory factor analysis and its concurrent validity. Study 1 was carried out on a sample of 556 workers, while Study 2 on 202 workers. The original three-factor structure was supported in the 17-item Italian version of the scale. It also showed partial metric and scalar invariance (Study 1) supporting its construct validity. Moreover, the scale showed expected correlations with work-family balance, conflict and enrichment, job and family satisfaction (Study 2). Although the fit indices for the two-factor and the second-order three-factor models were satisfactory, the original three-factor structure of the scale prevailed as the best factor solution. Construct and concurrent validity were supported. Practical implications concerning the use of the scale are discussed.

Key words: Work-family culture; Work-family balance; Work-family conciliation; Factor analysis; Measurement invariance.

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Recent decades have seen significant shifts in the labor force that have blurred the boundaries between work and family domains. The increased proportions of women, dual-earner and single-parent families, as well as employed individuals with child- and elder-care responsibilities, have attracted considerable research interest concerning the question of how organizations can avail themselves of supportive programs and policies to help accommodate the different needs of today's workforce (Duxbury & Gover, 2011). These initiatives, usually referred to as formal organizational work-family support, include a wide array of options and benefits especially designed to alleviate the burden inherent in managing multiple life roles (e.g., flextime, job sharing, support leave policies, child-care referrals, etc.; Thompson, Beauvais, & Lyness, 1999). However, sometimes these policies do not have the effect intended, as they do not bring with them a similar progressive change in the organizational norms and values (Allen, 2001). In such cases, employees may fear the potential negative repercussions associated with the advantage of such initiatives (e.g., chance of facing negative judgements of being signaled as a worker with low commitment or less desire for upward mobility) and, consequently, will be unlikely to take advantage of them so as not to jeopardize their own careers. Therefore, simply offering family-friendly arrangements cannot truly benefit individuals if the organizational culture does not legitimate and encourage their effective uptake (Thompson et al., 1999). In this sense, in many if not most situations, a key role is played by the so-called informal organizational work-family support, defined as "organizational phenomena that are not mandated or proscribed by formal organizational policies or programs" (Behson, 2005, p. 488), and operationalized through, among others, concepts such as work-family culture (Thompson et al., 1999) or family-supportive organizational perceptions (Allen, 2001).

Thompson et al. (1999) were among the first to suggest that a *work-family culture* in an organization can be a crucial element for enabling employees to adequately balance their multiple obligations. They defined a work-family culture as "the shared assumptions, beliefs, and values regarding the extent to which an organization supports and values the integration of employee's work and family lives" (p. 394). Moving beyond this definition, these authors claimed that the perception of a supportive work-family culture would have been significantly related to a number of positive outcomes, such as greater affective commitment, less intention to leave the organization, and less work-to-family conflict. A large body of subsequent empirical evidence (for reviews of outcomes, see Kinnunen, Mauno, Geurts, & Dikkers, 2005; Mesmer-Magnus & Viswesvaran, 2006) supported the predictive power of work-family culture in relation to improved attitudinal employee outcomes, such as job satisfaction, organizational commitment, lower turnover intentions (Allen, 2001; Dikkers, Den Dulk, Geurts, & Peper, 2005; Mauno, Kinnunen, & Piitulainen, 2005), lower work-to-family conflict (Allen, 2001; Dikkers et al., 2005; Mennino, Rubin, & Brayfield, 2005), and higher work-to-family enrichment (Lo Presti & Mauno 2016; Peeters, Wattez, Demerouti, & de Regt, 2009). A few studies also lent credit to the association between work-family culture and employee well-being (e.g., better self-reported health and lower level of physical symptoms; Mauno, Kinnunen, & Piitulainen, 2005; Mauno, Kinnunen, & Pyykkö, 2005; Mauno, Kinnunen, & Ruokolainen, 2006).

Yet, work-family culture is still a relatively novel concept and many unresolved issues, primarily methodological ones, limit our understanding of this phenomenon (Duxbury & Gover, 2011; Mauno, 2010). Moreover, current evidence suggests that some work-family culture aspects may enhance the use of family-friendly solutions and benefits. Dikkers et al. (2007) found that work-family support provided by supervisors and colleagues resulted in higher flextime, part-time and subsidized child-care among employees, while Breaugh and Frye (2008) found that a supportive supervision was positively associated with use of flextime and family leaves. Thus, a valid and reliable measure of work-family culture is essential to both scholars and practitioners in order to evaluate the role of this variable. Work-family culture may have a role in supporting or hindering the effectiveness of organizational efforts in helping individuals strike a balance between work and family through an increased individual effort in work-family conciliation or segmentation, and through the organizational implementation of flexible work arrangements and other work-family supportive programs.

THE WORK-FAMILY CULTURE SCALE

A pre-eminent line of argument relates to the operationalization and measurement of work-family culture as it has been conceptualized and measured both as a one- and multi-dimensional construct (Kinnunen et al., 2005). The pioneering study of Thompson et al. (1999) suggested that there were at least three components of work-family culture. The first component concerned *organizational time demands*, which referred to the rules concerning the "face time" employees are expected to devote to work, prioritizing this latter over family as a demonstration of dedication and commitment to own career. The second component captured the perception of the *negative career consequences* associated with utilizing the available work-family benefits or as a result of devoting excessive time and efforts to one's personal life. Finally, the last component related to *managerial support*, namely the extent to which managers are perceived to be supportive and sensitive to the employees' needs of finding a balance between work and family life.

On this theoretical grounding, Thompson et al. (1999) created a 20-item scale that mirrored such a threefold characterization of work-family culture in the United States: (1) *organizational time demands*, which comprised four items (e.g., "Employees are often expected to take work home at night and/or on weekends"); (2) *career consequences*, which encompassed five items (e.g., "To turn down a promotion or transfer for family-related reasons will seriously hurt one's career progress in this organization"); and (3) *managerial support*, which covered 11 items (e.g., "Higher management in this organization encourages supervisor to be sensitive to employees' family and personal concerns"). Alpha for the total scale was .92, while the reliability of the subscales ranged between .74 and .90.

Since it was first delineated, Thompson et al.'s (1999) questionnaire, called Work-Family Culture Scale (henceforth WFCS), has become the most widely utilized work-family culture measure and has been employed in different fashions by a number of researchers (e.g., Bragger, Rodriguez-Srednicki, Kutcher, Indovino, & Rosner, 2005; de Sivatte & Guadamillas, 2012; Mauno, 2010; Mauno, Kinnunen, & Piitulainen, 2005). Nevertheless, there has been relatively little research about its adaptation and validation in other countries and not even one in the Italian context, as opposed to other well-known constructs such as work-family conflict and enrichment (Colombo & Ghislieri, 2008; Ghislieri, Martini, Gatti, & Colombo, 2011). Furthermore, the few existing studies in the literature showed no consensus about the number of dimensions constituting the construct.

Recently, a two-dimensional approach has gained support by which the first factor related to managerial support as a work-family support facet and the second one comprised both the negative career consequences and organizational time demands (which usually showed a robust intercorrelation) as work-family barriers or constraints (Dikkers et al., 2007; Mauno, 2010; Mauno, Kinnunen, & Feldt, 2012). This line of reasoning is also coherent with Voydanoff's (2005) model, whereby managerial support could be considered as a boundary-spanning resource (i.e., a resource facilitating a positive work-family interface), whereas expected career consequences and organizational time demands are conceived as boundary-spanning demands (i.e., forces hampering a positive work-family interplay). Apart from initial validation in the U.S., these studies have been carried out in some European countries (i.e., Finland and Spain).

Mauno (2010) carried out a series of confirmatory factor analyses (CFA) on a 17-item version of the WFCS (three items from the original managerial support scale had to be removed) among a sample of 409 Finnish healthcare workers. A two-factor structure was supported against the original three-factor one: managerial work-family support (eight items) and work-family barriers (nine items). The model included three estimated error variances (Items 15, 16, and 17) between the observed variables and showed acceptable goodness-of-fit indices (CFI = .934; RMSEA = .064). Later, Mauno et al. (2012) examined the factor structure of the WFCS with a sample of over 2000 Finnish employees from different economic sectors, and found a two-factor structure

(.88 < CFI < .95; .062 < RMSEA < .084) similar to the one found by Mauno (2010): a five-item factor named work-family support and encompassing only managerial support items, and a nine-item factor called work-family barriers. Both studies did not report final Cronbach's alphas.

Other studies, however, support a three-factor solution. De Sivatte and Guadamillas (2012) accounted for three factors on a modified version of the original scale capturing particular Spanish organizational facets, and developed a 20-item questionnaire (from an initial pool of 53 items, of which 14 had been self-developed and several had been adapted from Shinn, Wong, Simko, & Ortiz-Torres, 1989). Career consequences and organizational time demands collapsed in a unique factor, as in the previous mentioned studies, which was labeled career consequences (10 items, $\alpha = .95$), while supervisor support (five items, $\alpha = .95$) and managerial support (five items, $\alpha = .94$) constituted two distinct factors (the former referring to the direct supervisor, while the latter to middle/top managers).

Beléndez Vázquez, Martin Llaguno, and Hernández Ruiz (2013) provided evidence of the original threefold factor structure of the WFCS (CFI = .99; RMSEA = .038). Using a 16-item version of the scale (Poelmans & Chinchilla, 2001), they found a final 11-item version of the questionnaire with three distinct factors: managerial support (four items; $\alpha = .73$), organizational time demands (three items; $\alpha = .69$), and career consequences (four items; $\alpha = .65$).

This brief literature review of psychometric properties of the WFCS was focused on the few studies that, as a primary purpose or not, provided evidence about the factor structure of the scale, along with its validity and, sometimes, reliability. Evidence is mixed, with both the original three-factor and the more recent two-factor solutions being supported. Moreover, these studies often included additional items, removed original items, or collapsed them within different factors on the basis of the exploratory and confirmatory factor analyses carried out. It derives that an agreement among scholars about how to evaluate work-family culture is far from definitive and it seems that scholars interested in evaluating work-family culture are in need of more tailored scales, justified by differences between nations or cultural/economic areas concerning: labor markets, availability of formal organizational work-family benefits, cultural norms and values attached to work-family conciliation (Allen et al., 2014; Masuda et al., 2012).

Because of these inconclusive and mixed results and the absence of an Italian validated scale evaluating informal organizational work-family support, this study had two purposes. First, we aimed at examining the construct validity of the Italian version of the WFCS through a series of CFA and measurement invariance tests (Study 1), contrasting two- and three-factor structures (first- and second-order), as differently found in other studies. In the present study two groups, female and male workers, were involved. The issue of work-family balance and its related variables have always been examined from a gender perspective (Colombo & Ghislieri, 2008). Although the results of the studies have often been conflicting, the gender dimension can play a role in modulating experiences related to reconciliation of work and family. This may be particularly true in Italy, where differences in male and female participation in work-family balancing are well recognized (Donati & Prandini, 2008). Accordingly, we believed that it would be worthwhile to use these two groups in order to test if the scale was suitable for both. Second, we aimed at verifying its concurrent validity, as resulted from Study 1, with other constructs salient to work-family conciliation research (Study 2). In particular, on the basis of the available evidence and in order to check for its concurrent validity, we expected work-family culture to be positively related to job satisfaction (Allen, 2001) and work-to-family enrichment (Lo Presti & Mauno 2016; Peeters et al., 2009). We also expected work-family culture to be negatively related with work-to-family conflict (Allen, 2001; Dikkers et al., 2005; Mennino et al., 2005). Despite the rather scant literature on the subject, we assumed that work-family culture is also positively associated with work-family balance (Carlson, Grzywacz, & Zivnuska, 2009) and family-life satisfaction (Kobau, Sniezek, Zack, Lucas, & Burns, 2010). The former is defined as the extent to which an individual meets negotiated role-related expectations in both work and family domains, and the latter as the degree of individual satisfaction with his/her own family life. We believe it is reasonable to expect that, as the two variables are related to the positive interweaving between life and work domains, they could benefit from a supportive organizational culture and thus show a positive association. In sum, we selected variables that have traditionally been considered as proximal (i.e., work-to-family conflict, enrichment, and balance) and distal (i.e., job satisfaction, family-life satisfaction) outcomes of informal work-family organizational support. Thus, we hypothesized that:

H1: work-family culture will be positively associated with (a) job satisfaction, (b) work-to-family enrichment, (c) work-family balance, and (d) family-life satisfaction.

H2: work-family culture will be negatively associated with work-to-family conflict.

Study 1

Method

Participants and Procedure

Data were collected via self-report paper and pencil questionnaires among the employees of nine organizations. We adopted a maximum variation sampling strategy in order to collect data from several diverse organizational settings and identify common phenomena concerning work-family conciliation.

Frequencies and descriptive statistics will be reported for the global sample (N = 556). Three hundred twenty-three men (58.1%) and 233 women (41.9%) were interviewed. Among those, 253 (45.5%) were employees from the industry sector, 270 (48.5%) were employees from public and private service sector, and 33 (6%) were from the agriculture sector or other.

Seventy-one were single/unmarried (12.8%), 437 married/cohabitating (78.6%), 48 held other marital statuses (i.e., widow, divorced; 8.6%). Then, 150 (27%) had no children, whereas 405 (72.8%) had at least one child (1 missing value, 0.2%). Mean age was 42.12 years (SD = 9.63), while average tenure was 14.72 years (SD = 9.05).

Questionnaires were distributed by trained survey researchers along with letters of invitation inside sealable envelopes. Participants had the right to not participate and return blank questionnaires, sealed inside the envelopes.

Measures

Work-Family Culture Scale (Thompson et al., 1999). The original scale was translated from English to Italian and from Italian to English following the procedure by Brislin (1970). Responses were collected through a 5-point Likert scale ranging from 1 = *completely false* to 5 = *completely true*. Items from the original career consequences and organizational time demands subscales had been reversed so that higher scores indicated lower career consequences and organizational time demands.

Data Analysis

Confirmatory factor analysis (CFA) through structural equation modeling using maximum likelihood (ML) method was conducted in order to test the dimensionality of the scale on the overall sample. CFA represents a kind of structural equation modeling that deals specifically with measurement models. The aim of this analysis is to examine the relationship between observed measures and latent variables, or factors. Through a CFA, a researcher can specify the number of factors and the pattern of indicator-factor loadings in advance as well as other parameters such as those bearing on the independence or covariance of the factors and indicator unique variance. The prespecified factor solution can be evaluated in terms of how well it reproduced the sample covariance matrix of the measured variables. CFA is a robust analytic tool for construct validation, since its results can provide compelling evidence of the convergent and discriminant validity, which are adjusted for measurement error theory (Brown & Moore, 2012). According to Byrne (2001), the use of ML method in estimating CFA assumes that the following conditions have been met: a) the sample is very large (asymptotic); b) the distribution of the observed variables is multivariate normal; c) the hypothesized model is valid (West, Finch, & Curran, 1995); and d) the scale of the observed variables is continuous. In the current study, conditions a, c, and d were met given that: sample size is large enough to test the model ($N = 556$); the hypothesized model has already been validated in literature (Thompson et al., 1999) and the scale of the observed variables is continuous. Condition b, regarding the normality distributions of the items, will be presented at the beginning of the results section.

Multigroup confirmative factor analysis (MCFA) through structural equation modeling using maximum likelihood (ML) method was conducted in order to test configural and measurement invariance. MCFA consists of simultaneous CFAs in two or more groups, offering a strong analytic framework for evaluating the equivalence of measurement models across distinct groups (Brown & Moore, 2012). Both CFA and MCFA were carried out through the structural equation modeling method using AMOS 20. Descriptive analysis and reliability analysis were conducted using SPSS 20.

The indices of the model fit considered for testing the dimensionality and the configural invariance of the Italian version of the Work-Family Culture Scale were: the comparative fit index (CFI), the root mean square error of approximation (RMSEA), the normed chi-square (χ^2/df), and the Akaike information criterion (AIC). CFI assesses the extent to which the tested model is superior to an alternative model in reproducing the observed covariance matrix (Bentler, 1990; McDonald & Marsh, 1990). The CFI index varies from 0 to 1 and a cut-off criterion of CFI > .90 is needed in order to ensure that mis-specified models are not accepted (Hooper, Coughlan, & Mullen, 2008). The RMSEA introduces a correction for lack of parsimony since, all other things being equal, more complex models are penalized. A cut-off value close to .06 (Hu & Bentler, 1999) or a stringent upper limit of .08 (Steiger, 2007) seems to be the general consensus among researchers. The normed chi-square, or the chi-square to degrees of freedom ratios (χ^2/df), is a further version of the traditional chi-square. The advantage of the normed chi-square is that it might be less sensitive to the sample size. Schumacker and Lomax (2004) suggested that a normed chi-square lower than 5 is an index of a good fit. The AIC is a comparative measure of fit. According to Burnham and Anderson (2004), lower values indicate a better fit and so the model with the lowest AIC is the best fitting model.

After the examination of the factor structure of the three-factor original model, the fit of this model was contrasted to the fit of a two-factor model in order to test which model represents better the Italian version of the scale. Furthermore, given that the composite measure of Work-Family Culture Scale is very often used in research in this field (e.g., Thompson et al., 1999) and that a proper test of the second-order model, which statistically would support the use of the composite measure, is lacking, the original three-factor model was contrasted with its second-order model.

Then, measurement invariance analysis was conducted on the three-factor model in order to further test construct validity. Measurement invariance test consists of several steps including the examination of configural, metric, and scalar invariance. Configural invariance is achieved when the model holds on the different groups included in the analysis (Byrne, 2004). Then, in order to test full metric invariance, the fit of a constrained model including all the factor loadings fixed is compared to the fit of a free-to-vary model. Following Chen (2007) and Cheung and Rensvold (2002), CFI and RMSEA were used to test measurement invariance. The cut-off points for rejecting measurement invariance are established as an increase of RMSEA by .015 and a decrease of CFI by .01. If the fit difference between the models falls into the threshold for rejecting the full metric invariance, partial metric invariance could still be explored leaving at least two factor loadings fixed in a construct, or in a factor when a construct is composed by several factors. Once that at least partial metric invariance has been established, it is important to explore whether the scores from different groups have the same origin, thus whether the intercept across the group is the same (scalar invariance). As to metric invariance, if full scalar invariance cannot be established, partial scalar invariance can be explored by relaxing constraints on the intercepts one by one (Byrne, Shavelson, & Muthén, 1989; Millsap & Meredith, 2007).

Results

First, we examined descriptive statistics and normality of the variable distributions through the calculation of mean values, standard deviations, skew and kurtosis. Item mean values ranged from 2.58 to 3.40 (mean = 2.95), whereas standard deviations ranged from 1.23 to 1.61 (mean = 1.34). Since the sample was quite large, the skew and kurtosis analysis were conducted through the interpretation of their absolute values (Kline, 2005). All in all, results reported skewness and kurtosis values within the normality range, with the absolute skewness values ranging from 0.14 to 3.72 (mean = 1.31) and a kurtosis with the absolute kurtosis values ranging from 3.97 to 6.64 (mean = 5.40). According to Kline, a skewness > 3 and a kurtosis > 7 are indices of some nonnormality. Since in this study just a negligible skewness effect was found in three out of twenty variables, ML could be carried out without any potential statistical bias.

Thus, the first step of the factor structural analysis consisted of running a CFA based on the three factors by Thompson et al. (1999). The results of the three-factor model showed a poor fit (χ^2/df = 5.234; CFI = .815; RMSEA = .087). Moreover, although all the lambdas were statistically significant, three of them were particularly low indicating a form of their weakness. At this stage of the analysis, when results of CFA indicate some critical points, Byrne (2001) argues that a researcher can decide whether or not to respecify and re-estimate the model by conducting a post hoc analysis. Accordingly, it is important to note that the analyses are now framed within an exploratory, rather than a confirmatory, approach. In other words, once a hypothesized CFA model, for example, has been rejected, this spells the end of the confirmatory factor-analytic approach, in its truest sense. Although CFA procedures continue to be used in any respecification and re-estimation of the model, these analyses are exploratory in the sense that they focus on the detection of termed "specification searches" (see MacCallum, 1986). According to this rationale and in order to achieve a better fit of the model, the three most problematic items (factor loading < .35) were removed from the following analysis. However, before removing these items a deeper item analysis was carried out in order to better understand the reason for their weakness. In particular, two items were ambiguous because in Italian they could relate both to managerial support and organizational time demands. The two items were: "In this organization it is very hard to leave during the workday to take care of personal or family matters," and "This organization encourages employees to set limits on where work stops and home life begins".

Moreover, also in the results of the factor analysis conducted on the original version of the scale by Thompson et al. (1999) there was a trace of such ambiguity, given that these two items loaded better on managerial support, but showed also a weaker but acceptable loading on organizational time demands. The third item ("Employees are often expected to take work home at night and/or on weekend") represents a typical "white collar" situation that is difficult to generalize to all the organizations and job roles, such as, for example, hospitals or factories. After deleting the ambiguous and weak items according to the values of their factor loadings, a 17-item version of the scale emerged as the best factor solution (χ^2/df = 4.789; CFI = .873; RMSEA = .083). However, although all the factor loadings ranged from .37 to .77, few modification indices still indicated the need to freely covariate three pairs of errors. Because of the high value of these modification indices (> .30) and the possible overlapping of meanings between the related items, we decided to set the three pairs of errors free to covariate. Thus, the three-factor model (Model 1), including the 17-items, showed a very good fit (χ^2/df = 3.023; CFI = .934; RMSEA = .06; AIC = 421.601). However, since recent literature reports also a two-factor version of the scale that has gained empirical support (e.g., Mauno, 2010), in the current study the fit of the three-factor model was contrasted with the fit of the two-factor model (Model 2). According to the literature, the two factors were work-family support and work-family barriers; the latter factor emerged from collapsing organizational time demand and career consequences. Despite the fit of the two-factor model being only slightly lower than that of the three-factor model, the results of the fit indices showed a very good fit (χ^2/df = 3.085; CFI = .930; RMSEA = .061; AIC = 431.895). Thus, it seemed that both the three-factor and the two-factor models could represent well the Italian version of the scale. However, since the three-factor model reported the lowest AIC, we should assume that the three-factor structure shows better fit, though negligibly, than the two-factor model. In order to provide further information for disambiguating the use of the three-factor or the two-factor models of the Italian version of the WFC scale, we followed the procedure proposed by Miyake, Friedman, Emerson, Witzki, and Howerter (2000), and also implemented by Bobbio and Manganelli (2009). Thus, we compared a model (Model 3) where the correlation between the two highly correlated factors was 1 and the correlations between the two highly correlated factors and the other third factor were constrained. Results (showed in Table 2) indicated that, although the fit indices of Model 3 were adequate, they were worse than those of the original three-factor model (χ^2/df = 3.485; CFI = .918; RMSEA = .067; AIC = 476.718). Thus, a three-factor solution was preferable.

Finally, since in the literature the result of the composite measure of the WFC scale is often provided (e.g., Thompson et al., 1999), we also tested a second-order model, assuming that the latent construct "work-family culture" may account for the three subdimensions. Although it would be very important in order to empirically support the use of the composite measure of the scale, as far as we know there is no trace of the test of the second-order model in the literature. Results of the second-order three-factor model (Model 4) reported satisfactory fit (χ^2/df = 3.027; CFI = .934; RMSEA = .060; AIC = 422.076). In particular, although all the fit indices considered were very similar to those of the original three-factor model, the original three-factor model still showed the lowest AIC.

In sum, we could conclude that both the two-factor and the second-order three-factor models reported satisfactory fit indices, very similar to those of the original three-factor model. In particular the difference between the original three-factor model and the second-order model is paltry, allowing the use of the composite measure of work-family culture for research and practical purposes in Italy. Nevertheless, according to the AIC values, the best factor solution for the Italian version of the scale seemed to be the original three-factor model. Cronbach's alpha coefficients for the three factors were: managerial support = .89; organizational time demands = .64; career consequences = .71; overall scale = .78. Intercorrelation coefficients among the three factors were all statistically significant (p < .01): managerial support/career consequences = .29; managerial support/organizational time demands = .31; career consequences/organizational time demands = .92. This latter intercorrelation showed to be robust and deserves further inquiry, although it is consistent with previous findings (Mauno, 2010; Mauno et al., 2012); in those cases, this high intercorrelation was considered a weakness because of the factors measuring common, rather than separate, subconstructs. Table 1 depicts the definitive item list and their respective factor loadings.

Then, this factor model was used to perform the MCFA and test its measurement invariance. According to the measurement invariance literature, the first step consisted of examining the configural invariance. Configural invariance is achieved when a model holds on all the groups in the study (Byrne, 2004). Two different groups of workers were used: 233 women and 323 men. The results of the MCFA reported an excellent fit (χ^2/df = 2.336; CFI = .916; RMSEA = .049) and, thus, the next step of the measurement invariance path could be conducted. In order to examine the full metric invariance, the free-to-vary model and the constrained model were compared. The constrained model comprised all the factor loadings fixed (CFI = .883; RMSEA = .055). The comparison of the CFI and the RMSEA values of the two models showed that full metric invariance of the model could be established only for the RMSEA (ΔRMSEA = .006), whereas it could not be established for CFI (ΔCFI = .033) (Chen, 2007; Cheung & Rensvold, 2002). However, metric scalar invariance could still be achieved by releasing some of the fixed factor loadings in the constrained model. According to Byrne (2004), in order to test partial metric invariance at least two factor loadings in each construct should be held fixed. The results of the difference between the CFI (ΔCFI = .013) and RMSEA (ΔRMSEA = .003) of the compared models confirmed a partial metric invariance (CFI = .903; RMSEA = .052).

Then, the next step was to test the scalar invariance. Accordingly, the model was compared to a constrained model where all the intercepts were fixed (full scalar invariance). The results (CFI = .838; RMSEA = .063) showed that the full scalar invariance was achieved only for the difference between the RMSEA values (ΔRMSEA = .014), while it could not be established for the CFI values (ΔCFI = .078). However, partial scalar invariance could still be achieved by releasing some of the fixed intercepts in the constrained model. The results of the difference between the

TABLE 1
CFA definitive items' list and their respective factor loadings

	Factor 1 Work-family managerial support	Factor 2 Organizational time demands	Factor 3 Negative career consequences
Item 1	.673		
Item 2	.744		
Item 3	.607		
Item 5	.760		
Item 8	.727		
Item 16	.772		
Item 17	.517		
Item 18	.715		
Item 20	.652		
Item 10		.727	
Item 11		.712	
Item 6		.396	
Item 12			.598
Item 13			.497
Item 9			.372
Item 7			.572
Item 19			.467

CFI (ΔCFI = .018) and RMSEA (ΔRMSEA = .003) reported a ΔCFI slightly above the threshold of invariance acceptance and an excellent ΔRMSEA. Thus, the results of the compared models confirmed a partial scalar invariance (CFI = .898; RMSEA = .052). Table 2 presents the results for the fit indices of the models tested on the overall sample, on each of the two sub-samples, on the multigroup, and for the invariance measurement test. Therefore, the construct validity of the Italian 17-item version of the Work-Family Culture Scale was supported.

STUDY 2

Method

Participants and Procedure

Data were collected via self-report paper and electronic questionnaires among 202 employees of seven small and medium enterprises. Thus, we adopted a maximum variation sampling strategy in order to collect data from several diverse organizational settings and identify

TABLE 2
Fit indices for the factor solutions for the Italian version of the WFC Scale

Groups	χ^2/df	CFI	RMSEA	χ^2/df	AIC
M1. Overall three-factors	357.895 (116)	.934	.060	3.023	421.601
M2. Overall two-factors	341.601 (113)	.930	.061	3.085	431.895
M3. Overall three-factors (correlation constrained)	400.718 (115)	.918	.067	3.485	476.718
M4. Second-order three-factors	342.076 (113)	.934	.060	3.027	422.076
Women three-factors		.938	.060	1.829	
Men three-factors		.899	.085	2.181	
Configural invariance		.916	.049	2.336	
Full metric invariance		.883	.055		
Partial metric invariance[1]		.902	.052		
Full scalar invariance		.838	.063		
Partial scalar invariance[2]		.898	.052		

Note. WFC = Work-Family Culture; CFI = comparative fit index. RMSEA = root mean square error of approximation. AIC = Akaike information criterion. [1] The parameters constrained invariant for the partial metric invariance model are the following items' factor loadings presented in Table 1: 1, 2, 10, 11, 12, 13. [2] The parameters constrained invariant for the partial scalar invariance model are the intercepts of the following items presented in Table 1: 1, 20, 10, 6, 12, 9.

common phenomena concerning work-family conciliation. In particular, 101 men (50%) and 101 women (50%) were interviewed. Among those, 38 were single/unmarried (18.8%), 153 married/cohabitating (75.8%), 11 held other marital statuses (5.4%). The number of children within the household ranged between 0 and 5 (M = 1.85, SD = 0.97). Mean age was 41.79 years (SD = 10.8), while average tenure was 12.8 years (SD = 9.43). Finally, 178 respondents (88.5%) had permanent employment contracts, while 23 had temporary ones (11.4%; 1 missing value, .01%).

Paper and pencil questionnaires were distributed by trained survey researchers along with letters of invitation inside sealable envelopes. The same questionnaire was also administered using an online form (LimeSurvey version 1.92+). The questionnaire link was sent to participants with a detailed introduction and strict instructions. In both cases, participants had the right to not participate and no compensation was provided for the participation in research.

Measures

Work-Family Culture Scale (Thompson et al., 1999). We used a single composite score for the WFCS because: 1) scholars (e.g., de Sivatte & Guadamillas, 2012; Mauno, 2010) provided different work-family culture scales (in terms of number of factors, respective items' meaning and number) so the comparison between separate work-family subscales could return biased results due to significant differences between studies; using a composite score as a basis for present and future comparisons would return less biased results; 2) for the same reason, future comparisons with similar constructs and measures such as family-supportive organizational perceptions (Allen, 2001), work-family climate (O'Neill et al., 2009), and so forth, would be easier; 3) from an empirical point of view, the WFCS second-order model emerged in Study 1 showed almost interchangeable goodness-of-fit indices with the three-factor model. Work-family culture's Cronbach's alpha was .85. Scores ranged between 17 and 85.

Work-Family Balance Scale (Carlson et al., 2009). It refers to the extent to which an individual meets negotiated role-related expectations in both the work and family domains. It comprised six items (e.g., "I am able to accomplish the expectations that my supervisors and my family have for me") with a 5-point Likert scale from 1 = *completely disagree* to 5 = *completely agree*. Cronbach's alpha for this scale was .93. Scores ranged between 6 and 30.

Work-to-Family Conflict Scale (Netemeyer, Boles, & McMurrian, 1996; Italian version by Colombo & Ghislieri, 2008). It concerns how often the participation in the family role is made more difficult by virtue of participation in the work role and included five items (e.g., "Due to work-related duties, I have to make changes to my plans for family activities") with a frequency scale from 1 = *never* to 6 = *always*. Cronbach's alpha for this scale was .88. Scores ranged between 5 and 30.

Work-to-Family Enrichment Scale (Carlson, Kacmar, Wayne, & Grzywacz, 2006; Italian short version by Ghislieri et al., 2011). It assesses the extent to which experiences in the work domain improve the quality of life in the family domain. It comprised three items (e.g., "At work I feel positive emotions and this helps me to be a better family member") with a 5-point Likert scale from 1 = *completely disagree* to 5 = *completely agree*. Cronbach's alpha for this scale was .89. Scores ranged between 3 and 15.

Job Satisfaction Scale (Pejtersen, Kristensen, Borg, & Bjorner, 2010). It consisted of a general question ("Thinking about your job, please indicate your satisfaction with regard to the following aspects") related to five different job features (e.g., "physical working conditions") with a 5-point scale from 1 = *not satisfied at all* to 5 = *completely satisfied*. Cronbach's alpha for this scale was .89. Scores ranged between 5 and 25.

Satisfaction with family life scale (Kobau et al., 2010). It refers to the extent to which the respondent is satisfied with his/her own family life. It comprised five items (e.g., "In most ways my family life is close to my ideal") with a 7-point Likert scale from 1 = *completely disagree* to 7 = *completely agree*. Cronbach's alpha for this scale was .94. Scores ranged between 5 and 35.

Results

Zero-order correlations (Pearson's r) were used to examine the associations between the variables. Table 3 depicts zero-order correlations between our Study's 2 variables.

Concerning work-family culture, it positively correlated with work-family balance (r = .29, $p < .001$), work-to-family enrichment (r = .38, $p < .001$), job satisfaction (r = .32, $p < .001$), and family satisfaction (r = .19, $p < .01$). It negatively correlated with work-to-family conflict (r = –.14, $p < .05$). These results, that will be discussed later, provided support for the concurrent validity of the Italian version of the Work-Family Culture Scale.

Conclusions

Lewin's (1951) famous maxim "there is nothing so practical as a good theory" recalls the impossibility of conducting research without reference to a theory, which also has to be a solid

TABLE 3
Descriptive statistics and zero-order correlations

	M (*SD*)	1	2	3	4	5
1) Work-family culture	53.11 (11.93)					
2) Work-family balance	22.72 (5.37)	.29***				
3) Work-to-family conflict	13.67 (5.40)	–.14*	–.17*			
4) Work-to-family enrichment	10.07 (3.13)	.38***	.52***	–.15*		
5) Job satisfaction	17.11 (4.47)	.32***	.52***	–.16*	.69***	
6) Family satisfaction	24.77 (7.61)	.19**	.54***	–.14*	.32***	.47***

Note. *** $p < .001$. ** $p < .01$. * $p < .05$.

one. Nonetheless, to test and advance theory, researchers also need a “good method” (Greenwald, 2012), in the sense of psychometrically sound measures to guide useful applications of their hypotheses.

When it comes to work-family culture, for many years, despite the burgeoning literature about how work and family may intersect and the role recognized to organizational culture in influencing this relationship in a very positive way, researchers were limited in their ability to test hypotheses about this concept and its potential antecedents, consequences, moderators, and mediators by the very lack of well-established measures of work-family culture (Andreassi & Thompson, 2008). Undoubtedly, since it was published, Thompson et al.’s (1999) WFCS has stood out in its field as the measure that best modeled the concept of work-family culture. However, the review of existing empirical studies that profited from its use suggests that there is still no agreement on the dimensionality of the work-family culture construct as touted by Thompson and colleagues.

Thus, the aim of this paper was to adapt the original 20-item factor structure of Thompson et al.’s (1999) questionnaire in Italy[1], examining first its construct validity (Study 1). In order to pursue this objective and to cope with the inconsistent measurement issues around the WFCS, we firstly carried out a CFA to test competing models of its latent structure. Despite both the three-factor, the two-factor, and second-order three-factor structures showing a very similar good fit, the three-factor solution prevailed, providing confirming evidence of the WFCS’ original structure in the Italian context. Work-family managerial support focused on the perceived easiness and supportiveness of balancing work and family within the organization, managerial empathy toward employees’ conciliation needs, the importance attached to the role of supervisors in supporting employees’ work-family conciliation, and so forth. Organizational time demands concerned the organizational-induced need to prioritize work over family, for example working overtime and putting job ahead of own family and personal life. Finally, negative career consequences pertained to organizational and colleagues’ ostracism as well as the anticipation of damaged career progress as a consequence of participating in work-family programs, use of flextime, and so forth. The three subscales’ internal reliabilities, as measured by Cronbach’s alpha values, were satisfactory and results from MFCA showed that the measure demonstrated a promising partial scalar invariance. The results led to a shortened version of the original questionnaire, with the same factor structure as the original questionnaire by Thompson et al. (1999). It must be noted that, consistently with the literature (Allen et al., 2014; Masuda et al., 2012), national cultural peculiarities concerning work organization, labor market, welfare policies, and so forth, may have played a significant role in shaping individual experiences and thus returning this reduced three-factor structure of the work-family culture construct with respect to Italian respondents. Such peculiarities could explain the broad differences concerning work-family culture's operationalization with respect to specific nations: Finland (Mauno, 2010), Spain (de Sivatte & Guadamillas, 2012), USA (Thompson et al., 1999).

Then, in order to examine its concurrent validity (Study 2), we evaluated if the WFCS was truly capable of measuring something unique from other relevant concepts in the work-family literature also concerning organizational supportiveness toward employees' family-related needs; thus, the scale as emerged from Study 1 was correlated with work-to-family enrichment (H1b), work-family balance (H1c), and work-to-family conflict (H2), as well as job (H1a) and family satisfaction (H1d). Consistently with reasonable theoretical expectations, higher work-family culture was associated with higher work-family balance and enrichment (Lo Presti & Mauno, 2016; Peeters et al., 2009), as well as higher job and family satisfaction (Dikkers et al., 2005), while it showed a negative, albeit lower, association with work-to-family conflict (Allen, 2001), thus providing full support to our hypotheses. In sum, work-family culture has stronger associations with variables referring to a positive proximal work-family interface (i.e., work-family balance, work-to-family enrichment) and then to more distal outcomes (i.e., job satisfaction; to a lesser extent family satisfaction). Moreover, it showed a weak significant association with a proximal negative outcome like work-to-family conflict, lower than Thompson et al.'s (1999) evidence ($r = -.55$).

Taken together, these results further underlined the importance of work-family culture in examining work-family conciliation and provided ample support for using Thompson et al.'s (1999) WFCS, in this present revised version, as a valid measure in the work-family research. Moreover they stress the importance of those organizational phenomena that are not mandated or proscribed by formal organizational policies or programs but rather are an important part of the organizational culture, and that can make a difference in the success of formal organizational work-family support interventions and tools. The ubiquitous message from this study and the rather scant literature on this matter is that, for work-family supportive policies to be effective, organizations should not only produce surface changes but also get at the heart of shared values and assumptions, without which being a real family-friendly organization will be much more difficult (Duxbury & Gover, 2011).

This study comes with some limitations that should be mentioned. In particular, as clarified by Beléndez Vázquez et al. (2013), a limitation concerns issues beyond the specific objectives of WFCS' validation studies as it regards the theoretical grounding on which this measure was first outlined. Indeed, the items in this measure neglected the increased proportion of labor force comprising single people and/or those who do not have children, but who also need to strike an appropriate balance between work and life responsibilities, therefore failing to properly examine and represent the different experiences of the actual workforce. Therefore, to use a more inclusive measure, researchers might consider the possibility of adding specific items related to other non-work needs, also taking into account cultural and national differences about work-family conciliation issues. Moreover, both being a cross-sectional study and the absence of objective indicators, such as the individuals' recourse to organizational family-friendly benefits and programs, limited the predictive validity of our results. What emerges as a suggestion for future research and a recommendation for potential organizational interventions, is that the Work-

Family Culture Scale should be used when assessing the success of formal work-family support initiatives within organizations, in order to verify if and how this recourse is sustained by the organizational culture or is prompted by other incidental factors. Nowadays, there is a large consensus among scholars about the importance played by work-family culture in promoting positive outcomes of employees' work-family conciliation (Mennino et al., 2005; Peeters et al., 2009) which, in turn, may have positive effects for the organization in terms of commitment, satisfaction, and so forth (Allen, 2001; Dikkers et al., 2005). It derives that the availability of a valid and reliable measure of such construct represents the conditio sine qua non for implementing and evaluating effective and useful family-friendly organizational programs.

Finally, as emerged from Study 1, the organizational time demands reliability score was lower than expected (.64). So, further effort is needed in order to collect more evidence and verify if such a result was due to a specific sampling error or to more general work-family Italian peculiarities.

NOTE

1. The Italian version of the items is available upon request from the first author.

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REFERENCES

Andreassi, J. K., & Thompson, C. (2008). Work-family culture: Current research and future directions. In K. Korabik, D. Lero, & D. Whitehead (Eds.), *Handbook of work-family integration: Research, theory, and best practices* (pp. 331-351). New York, NY: Elsevier Academic Press.

Allen, T. D. (2001). Family-supportive work environments: The role of organizational perceptions. *Journal of Vocational Behavior*, *58*, 414-435. doi:10.1006/jvbe.2000.1774

Allen, T., Lapierre, L., Spector, P., Poelmans, S. A. T., O' Driscoll, M., Sanchez, J. I., . . . Woo, J. (2014). The link between national paid leave policy and work-family conflict among married working parents. *Applied Psychology: An International Review*, *63*, 5-28. doi:10.1111/apps.12004

Behson, S. J. (2005). The relative contribution of formal and informal organizational work-family support. *Journal of Vocational Behavior*, *66*, 487-500. doi:10.1016/j.jvb.2004.02.004

Beléndez Vázquez, M., Martín Llaguno, M., & Hernández Ruiz, A. (2013). The factor structure of the Spanish version of the Work-Family Culture Scale in a sample of workers from the advertising sector. *Psicothema*, *25*, 232-237. doi:10.7334/psicothema2012.176

Bentler, P. M. (1990). Comparative fit indexes in structural models. *Psychological Bulletin*, *107*, 238-246.

Bragger, J. D., Rodriguez-Srednicki, O., Kutcher, E. J., Indovino, L., & Rosner, E. (2005). Work-family conflict, work-family culture, and organizational citizenship behavior among teachers. *Journal of Business and Psychology*, *20*, 303-324. doi:10.1007/s10869-005-8266-0

Breaugh, J. A., & Frye, N. K. (2008). Work-family conflict: The importance of family-friendly employment practices and family-supportive supervisors. *Journal of Business and Psychology*, *22*, 345-353. doi:10.1007/s10869-008-9081-1

Brislin, R. W. (1970). Back-translation for cross-cultural research. *Journal of Cross-Cultural Psychology*, *1*(3), 185-216.

Byrne, B. M. (2001). *Structural equation modeling with AMOS: Basic concepts, applications, and programming*. Mahwah, NJ: Laurence Erlbaum Associates.
Byrne, B. M. (2004). Testing for multigroup invariance using AMOS graphics: A road less travelled. *Structural Equation Modeling*, *11*, 272-300. doi:10.1207/s15328007sem1102_8
Byrne, B. M., Shavelson, R. J., & Muthén, B. O. (1989). Testing for equivalence of factor covariance and mean structures: The issue of partial measurement invariance. *Psychological Bulletin*, *105*, 456-466. doi:10.1037/0033-2909.105.3.456
Brown, T. A., & Moore, M. T. (2012). Confirmatory factor analysis. In R. H. Hoyle (Ed.), *Handbook of structural equation modeling* (pp. 361-379). New York, NY: Guilford.
Bobbio, A., & Manganelli, A. M. (2009). Leadership Self-Efficacy Scale. A new multidimensional instrument. *TPM – Testing, Psychometric, Methodology in Applied Psychology*, *16*, 3-24. doi:10.4473/TPM.16.1.1
Burnham, K. P., & Anderson, D. R. (2004). Multimodel inference. Understanding AIC and BIC in model selection. *Sociological Methods Research*, *33*, 261-304. doi:10.1177/0049124104268644
Carlson, D. S., Kacmar, K. M., Wayne, J. H., & Grzywacz, J. G. (2006). Measuring the positive side of the work-family interface: Development and validation of a work-family enrichment scale. *Journal of Vocational Behavior*, *68*, 131-164. doi:10.1016/j.jvb.2005.02.002
Carlson, D. S., Grzywacz, J. G., & Zivnuska, S. (2009). Is work-family balance more than conflict and enrichment? *Human Relations*, *62*, 1459-1486. doi:10.1177/0018726709336500
Chen, F. F. (2007). Sensitivity of goodness of fit indexes to lack of measurement invariance. *Structural Equation Modeling*, *14*, 464-504. doi:10.1080/10705510701301834
Cheung, G. W., & Rensvold, R. B. (2002). Evaluating goodness-of-fit indexes for testing measurement invariance. *Structural Equation Modeling*, *9*, 233-255. doi:10.1207/S15328007SEM0902_5
Colombo, L., & Ghislieri, C. (2008). The work to family conflict: Theories and measures. *TPM – Testing, Psychometrics, Methodology in Applied Psychology*, *15*, 35-55. doi:10.4473/TPM.15.1.3
de Sivatte, I., & Guadamillas, F. (2012). The measurement of work-family culture in Spain and some applications to other economies. *The International Journal of Human Resource Management*, *23*, 1930-1949. doi:10.1080/09585192.2011.602021
Dikkers, J. S. E., Den Dulk, L., Geurts, S. A. E., & Peper, B. (2005). Work-nonwork culture, utilization of work-nonwork arrangements, and employee-related outcomes in two Dutch organizations. In S. Poelmans (Ed.), *Work and family – An international research perspective* (pp.147-173). Mahwah, NJ: Lawrence Erlbaum Associates.
Dikkers, J., Geurts, S., den Dulk, L., Peper, B., Taris, T., & Kompier, M. (2007). Dimensions of work-home culture and their relations with the use of work-home arrangements and work-home interaction. *Work and Stress*, *21*, 155-172. doi:10.1080/02678370701442190
Donati, P., & Prandini, R. (2008). *La cura della famiglia e il mondo del lavoro* [Family care and the world of work]. Milano, Italy: FrancoAngeli.
Duxbury, L., & Gover, L. (2011). Exploring the link between organizational culture and work-to-family conflict. In N. M. Ashkanasy, C. P. M. Wilderom, & M. F. Peterson (Eds.), *The handbook of organizational culture and climate* (pp. 271-290). London, UK: SAGE Publications.
Ghislieri, C., Martini, M., Gatti, P., & Colombo, L. (2011). The “bright side” of the work-family interface: A brief work-family enrichment scale in a sample of health professionals. *TPM – Testing, Psychometrics, Methodology in Applied Psychology*, *18*, 211-230. doi:10.4473/TPM.18.4.2
Greenwald, A. G. (2012). There is nothing so theoretical as a good method. *Perspectives on Psychological Science*, *7*, 99-108. doi:10.1177/1745691611434210
Hooper, D., Coughlan, J., & Mullen, M. R. (2008). Structural equation modelling: Guidelines for determining model fit. *Electronic Journal of Business Research Methods*, *6*(1), 53-60.
Hu, L., & Bentler, P. M. (1999). Cutoff criteria for fit indexes in covariance structure analysis: Conventional criteria versus new alternatives. *Structural Equational Modeling: A Multidisciplinary Journal*, *6*, 1-55. doi:10.1080/10705519909540118
Kinnunen, U., Mauno, S., Geurts, S., & Dikkers, J. (2005). Work-family culture in organizations: Theoretical and empirical approaches. In S. Poelmans (Ed.), *Work and family: An international research perspective* (pp. 87-121). Mahwah, NJ: Lawrence Erlbaum Associates.
Kline, R. B. (2005). *Principle and practice of structural equation modelling* (2nd ed.). New York, NY: Guilford.
Kobau, R., Sniezek, J., Zack, M. M., Lucas, R. E., & Burns, A. (2010). Well-being assessment: An evaluation of well-being scales for public health and population estimates among US adults. *Applied Psychology: Health and Well-Being*, *2*, 272-297. doi:10.1111/j.1758-0854.2010.01035.x

Lewin, K. (1951). *Field theory in social science: Selected theoretical papers*. New York, NY: Harper&Row.

Lo Presti, A., & Mauno, S. (2016). Are support and control beneficial stress buffers in the presence of work-family barriers? Findings from Italy. *International Journal of Stress Management*, *23*, 44-64. doi:10.1037/a0038440

MacCallum, R. C. (1986). Specification searches in covariance structure modeling. *Psychological Bulletin*, *100*(1), 107-120.

Masuda, A., Poelmans, S., Allen, T., Spector, P. E., Lapierre, L. M., Cooper, C. L., . . . Moreno-Velasquez, I. (2012). Flexible working arrangements availability and their relationship with work-to-family conflict, job satisfaction, and turnover intentions: A comparison of three country clusters. *Applied Psychology: An International Review*, *61*, 1-29. doi:10.1111/j.1464-0597.2011.00453.x

Mauno, S. (2010). Effects of work-family culture on employee well-being: Exploring moderator effects in a longitudinal sample. *European Journal of Work and Organizational Psychology*, *19*, 675-695. doi:10.1080/13594320903106588

Mauno, S., Kinnunen, U., & Feldt, T. (2012). Work-family culture and job satisfaction: Does gender and parenting status alter the relationship? *Community, Work and Family*, *15*, 101-129. doi:10.1080/1366 8803.2011.598733

Mauno, S., Kinnunen, U., & Piitulainen, S. (2005). Work-family culture in four organizations in Finland: Examining antecedents and outcomes. *Community, Work and Family*, *8*, 115-140. doi:10.1080/1366 8800500049563

Mauno, S., Kinnunen, U., & Pyykkö, M. (2005). Does work-family conflict mediate the relationship between work-family culture and self-reported distress? Evidence from five Finnish organizations. *Journal of Occupational and Organizational Psychology*, *78*, 509-530. doi:10.1348/096317905X37082

Mauno, S., Kinnunen, U., & Ruokolainen, M. (2006). Exploring work-and organization-based resources as moderators between work-family conflict, well-being, and job attitudes. *Work & Stress*, *20*, 210-233. doi:10.1080/02678370600999969

McDonald, R. P., & Marsh, H. W. (1990). Choosing a multivariate model: Noncentrality and goodness of fit. *Psychological Bulletin*, *107*, 247-255. doi:10.1037/0033-2909.107.2.247

Mennino, S. F., Rubin, B. A., & Brayfield, A. (2005). Home-to-job and job-to-home spillover: The impact of company policies and workplace culture. *The Sociological Quarterly*, *46*, 107-135. doi:10.1111/j.1533-8525.2005.00006.x

Mesmer-Magnus, J. R., & Viswesvaran, C. (2006). How family-friendly work environments affect work/family conflict: A meta-analytic examination. *Journal of Labor Research*, *27*, 555-574. doi:10.1007/s12122-006-1020-1

Millsap, R. E., & Meredith, W. (2007). Factorial invariance: Historical perspectives and new problems. In R. Cudeck & R. C. MacCallum (Eds.), *Factor analysis at 100: Historical developments and future directions* (pp. 131-152). Mahwah, NJ: Lawrence Erlbaum Associates.

Miyake, A., Friedman, N. P., Emerson, M. J., Witzki, A. H., & Howerter, A. (2000). The unity and diversity of executive functions and their contributions to complex "frontal lobe" tasks: A latent variable analysis. *Cognitive Psychology*, *41*, 49-100. doi:10.1006/cogp.1999.0734

Netemeyer, R. G., Boles, J. S., & McMurrian, R. (1996). Development and validation of work-family conflict and family-work conflict scales. *Journal of Applied Psychology*, *81*, 400-410. doi:10.1037/002 1-9010.81.4.400

O'Neill, J. W., Harrison, M. M., Cleveland, J. N., Almeida, D. M., Stawski, R. S., & Crouter, A. C. (2009). Work-family climate, organizational commitment, and turnover: Multilevel contagion effects of leaders. *Journal of Vocational Behavior*, *74*, 18-29. doi:10.1016/j.jvb.2008.10.004

Peeters, M., Wattez, C., Demerouti, E., & de Regt, W. (2009). Work-family culture, work-family interference and well-being at work. Is it possible to distinguish between a positive and a negative process? *Career Development International*, *14*, 700-713. doi:10.1108/13620430911005726

Pejtersen, J. H., Kristensen, T. S., Borg, V., & Bjorner, J. B. (2010). The second version of the Copenhagen psychosocial questionnaire. *Scandinavian Journal of Public Health*, *38*, 8-24. doi:10.1177/1403 494809349858

Poelmans, S., & Chinchilla, M. N. (2001). *The adoption of family-friendly HRM policies. Competing for scarce resources in the labor market* (IESE Research Paper No. 438). Retrieved from the University of Navarra, IESE website http://www.iese.edu/research/pdfs/DI-0438-E.pdf

Schumacker, R. E., & Lomax, R. G. (2004). *A beginner's guide to structural equation modeling*. New Jersey, NJ: Lawrence Erlbaum Associates.

Shinn, M., Wong, N. W., Simko, P. A., & Ortiz-Torres, B. (1989). Promoting the well-being of working parents: Coping, social support, and flexible job schedules. *American Journal of Community Psychology*, *17*, 31-55. doi:10.1007/BF00931201

Steiger, J. H. (2007). Understanding the limitations of global fit assessment in structural equation modeling. *Personality and Individual Differences*, *42*, 893-898. doi:10.1016/j.paid.2006.09.017

Thompson, C. A., Beauvais, L. L., & Lyness, K. S. (1999). When work-family benefits are not enough: The influence of work-family culture on benefit utilization, organizational attachment, and work-family conflict. *Journal of Vocational Behavior*, *54*, 392-415. doi:10.1006/jvbe.1998.1681

Voydanoff, P. (2005). Consequences of boundary-spanning demands and resources for work-to-family conflict and perceived stress. *Journal of Occupational Health Psychology*, *10*, 491-503. doi:10.1037/1076-8998.10.4.491

West, S. G., Finch, J. F., & Curran, P. J. (1995). Structural equation models with nonnormal variables: Problems and remedies. In R. H. Hoyle (Ed.), *Structural equation modelling: Concepts, issues, and applications* (pp. 56-75). Thousand Oaks, CA: Sage Publications.