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# A bifactor exploratory structural equation modeling representation of the structure of the basic psychological needs at work scale

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ABSTRACT

This study investigates the structure of employees' ratings of the satisfaction of their basic psychological needs for autonomy, competence, and relatedness at work using the newly developed bifactor exploratory structural equation modeling (ESEM) framework. Using a sample of 366 exercise professionals who completed the new Portuguese version of the Basic Psychological Needs at Work Scale (BPNWS: Brien et al., 2012), the results demonstrated the superiority of a Bifactor-ESEM representation of BPNWS ratings when compared to alternative representations of the data (first-order and bifactor confirmatory factor analyses, and first-order ESEM). The results also supported the composite reliability, measurement invariance across gender, and nomological validity (in relations to measures of psychological wellbeing and distress at work) of BPNWS ratings. Importantly, these results demonstrated the importance of relying on measurement models providing a way to achieve a proper disaggregation of employees' global levels of need satisfaction relative to the satisfaction of their more specific needs for autonomy, competence, and relatedness.



Using a sample of 35,765 workers from 35 European countries, the European Foundation for the Improvement of Living and Working Conditions (EUROFOUND, 2015) observed that 20% of European workers reported poor psychological wellbeing at work. Psychosocial risk factors at work are multiple and varied, including job strain, low decision latitude, low social support, high psychological demands, effort–reward imbalance, and high job insecurity (Stansfeld & Candy, 2006). In this regard, several studies conducted under the Self-Determination Theory (SDT) framework (Deci & Ryan, 2000, 2008) have demonstrated the critical role of the satisfaction of basic psychological needs at work in the relation between these risks factors and employees' psychological wellbeing and distress (Boudrias et al., 2014; Brien et al., 2012; Deci et al., 2001; Desrumaux et al., 2015). Given the important role of the satisfaction of basic psychological needs at work, the current study aims to provide an improved representation of the structure of the Basic Psychological Needs at Work Scale (BPNWS) while relying on a bifactor exploratory structural

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equation modeling analytical framework. As a further test of the validity and generalizability of this improved representation of the BPNWS structure, we also verify the extent to which it generalizes across gender through tests of measurement invariance, and assess the nomological validity of the BPNWS factors in relation to measures of psychological wellbeing and distress.

## 1. Basic psychological needs at work

SDT defines basic psychological needs as the "nutrients that must be procured by a living entity to maintain its growth, integrity and health" (Deci & Ryan, 2000, p. 326). SDT focuses more specifically on the needs for autonomy, competence and relatedness, which are believed to generalize across cultures and life domains (Deci & Ryan, 2000). It is important to keep in mind that, although the current study focuses on SDT as applied to the work domain (Gagné & Deci, 2005), SDT can be considered as a pan-human theory of motivation with well-established applications across life domains including sport (e.g., Gillet, Vallerand, & Paty, 2013), education (e.g., Ryan & Deci, 2009), health (e.g., Ng et al., 2012), and multiple domains perspectives (e.g., Standage, Gillison, Ntoumanis, & Treasure, 2012). According to SDT, the *need for autonomy* is defined as the need to experience choice and volition in one's actions, rather than feeling controlled by external forces. When their needs for autonomy at work are fulfilled, workers experience a sense of personal control over their own work-related behaviors. The *need for competence* is defined as the need to experience effective interactions with one's environment. When their needs for competence are fulfilled, workers feel that that they have the ability to achieve desired work-related outcomes, and experience a sense of mastery and accomplishment at work. The *need for relatedness* is defined as the need to feel closeness, connection, and belongingness with others in one's environment. When their needs for relatedness are fulfilled, workers' feel included and accepted by significant others in their workplaces. Importantly, research has shown that employees' need satisfaction tends to be positively associated with wellbeing outcomes, such as optimism, vigor, self-esteem, positive affect, or life satisfaction, and negatively associated with psychological distress outcomes, such as emotional exhaustion, distress, negative affect, depression, or anxiety (Brien et al., 2012; Deci et al., 2001; Desrumaux et al., 2015; Longo, Gunz, Curtis, & Farsides, 2016; Van den Broeck, Vansteenkiste, De Witte, & Lens, 2008; Van den Broeck, Ferris, Chang, & Rosen, 2016; Vansteenkiste et al., 2007).

## 2. Measurement of the psychological needs at work

Over the past decades, several measures have been proposed to evaluate the satisfaction of basic psychological needs in work-related contexts, including the Basic Need Satisfaction at Work Scale (BNSW-S: Deci et al., 2001) and the Work-Related Need Satisfaction Scale (W-BNS) (Van den Broeck, Vansteenkiste, De Witte, Soenens, & Lens, 2010). Despite the widespread use of these measures, psychometric evidence regarding the generalizability of their psychometric properties to additional samples of employees beyond those used in the initial development studies remains lacking. Similarly, questions have been raised regarding the content of some of the items included in these instruments (e.g., Brien et al., 2012; Van den Broeck et al., 2010). Finally, the inclusion of negatively-worded items also raises concerns in terms of psychometric complexity (Marsh, Scalas, & Nagengast, 2010) and cross-cultural generalizability (Schmitt & Allik, 2005; Watkins & Cheung, 1995).

For these reasons, the present study focuses on yet another measure, the BPNWS (Brien et al., 2012). The BPNWS was developed to address some of the aforementioned limitations, and focuses on the assessment of the satisfaction of employees' basic needs for autonomy, competence, and relatedness at work.[1] Using three samples of employees from Canada and France (total $N = 1122$), Brien et al. (2012) results supported the a priori 3-factor structure of ratings on this instrument, as well as their scale score reliability ($\alpha = 0.84$ to 0.90) and nomological validity in terms of relations between the BPNWS dimensions and measures of intrinsic motivation, wellbeing, distress, optimism, and procedural justice. They also demonstrated the cross-cultural generalizability of their factor structure to two samples of Canadian participants (Study 1 and Study 2), and one sample of French participants (supporting the measurement invariance of the model in Study 2). Given our objective of presenting a Portuguese adaptation of a measure of needs satisfaction, this pre-existing evidence of the cross-cultural generalizability of the psychometric properties of the BPNWS made this instrument particularly well-suited to the present study.

## 3. The bifactor exploratory structural equation modeling (Bifactor-ESEM) framework

One critical limitation of most prior research focusing on the structure of measures of employees' need satisfaction is their reliance on the implicit assumption that their various subscales would be perfectly unidimensional psychometrically – which is a key assumption of confirmatory factor analyses (CFA). As noted by Morin, Arens, and Marsh (2016, p. 117), CFA "fail to account for at least two sources of construct-relevant psychometric multidimensionality, and might thus produce biased parameter estimates as a result of this limitation" (for similar arguments, see Morin, Arens, Tran, & Caci, 2016; Morin, Boudrias, Marsh, Madore, & Desrumaux, 2016; Morin, Boudrias, Marsh, McInerney et al., 2016).

[1] It is important to acknowledge that Longo et al. (2016) recently proposed the Need Satisfaction and Frustration Scale (NSFS) to achieve a more comprehensive assessment of needs satisfaction and frustration. In the work context, the authors found support for an a priori six-factor model differentiating the satisfaction and frustration of employees' needs for autonomy, competence, and relatedness. Unfortunately, this instrument was published after the current data collection was conducted, and thus could not be used in the present study. Furthermore, our goal was to focus on need satisfaction, and so far evidence regarding the distinctive nature of the need satisfaction and frustration constructs is limited to this single recent study (Longo et al., 2016).

The first of these two sources of construct-relevant psychometric multidimensionality is related to the assessment of hierarchically-ordered constructs, so that each item can be assumed to simultaneously contribute to the assessment of one specific construct (e.g., autonomy, competence and relatedness) and one global construct (e.g., global need satisfaction). In measures of need satisfaction, it is usual to rely on hierarchical factor models (e.g., Ntoumanis, 2005; Stebbings, Taylor, Spray, & Ntoumanis, 2012; Van den Broeck et al., 2008). In hierarchical CFA, each item is specified as loading on one a priori first-order factor (e.g., the satisfaction of the basic need for autonomy, competence and relatedness), and these first-order factors are specified as loading on a higher-order factor (e.g., global need satisfaction). However, as recently discussed by Gignac (2016), hierarchical models rely on one very stringent assumption that is very seldom tested in practice. More precisely, these models assume that the relations between the items and the higher-order factor are indirect and mediated by the first-order factors. Similarly, these models assume that the relation between the items and the unique part of the first-order factor is indirect and mediated by the first-order factor. Indirect effect are calculated as the product of two coefficients (i.e., the item loading on the first-order factor x the first-order factor loading on the higher-order factor; the item loading on the first-order factor x the disturbance of the first-order factor), so that for all items associated with a single first-order factor, the second term of this multiplication is a constant for both variance components (i.e., unique and global). Thus, the ratio of global/specific variance is assumed to be the same for all items associated with a single first-order factor.

Bifactor models provide a more flexible alternative to hierarchical models. In Bifactor-CFA, all items are directly allowed to load on one global G-factor (e.g., global need satisfaction) and on one specific S-factor (e.g., autonomy, competence and relatedness), and all factors are set to be orthogonal (e.g., Gignac, 2016; Morin, Arens, & Marsh, 2016; Reise, 2012). This specification allows for the total item covariance matrix to be directly separated into one global component (i.e., the G-factor) underlying responses to all items, and a series of specific components (i.e., the S-factors) specific to subsets of items but not explained by the global component. In the top section of Fig. 1, we provide a graphical illustration of first-order, hierarchical, and bifactor CFA.

The second source of construct-relevant psychometric multidimensionality is related to assessment of conceptually-related constructs which, coupled with the fallible nature of typical psychometric ratings, suggest that it is reasonable to expect ratings of a specific target construct to also present significant associations with non-target, yet conceptually-close, constructs. In CFA, cross-loadings between items and non-target factors are fixed to be exactly zero. Statistical research has shown that, whenever cross-loadings between items and non-target factors (even as small as 0.100) are present in the population model, forcing them to be exactly zero leads to biased estimates of factor correlations (for a review of this research, see Asparouhov, Muthén,

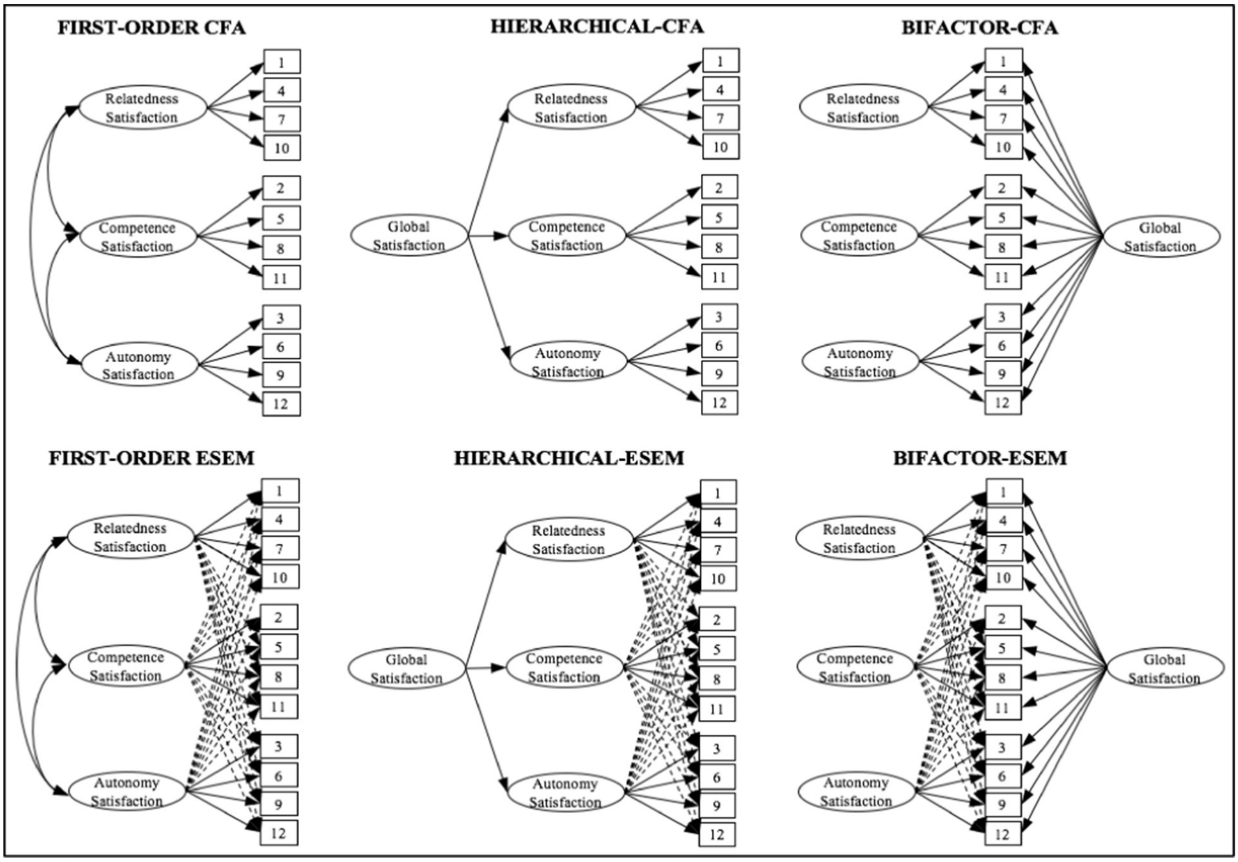


**Fig. 1.** Graphical representation of the alternative models tested. Note. CFA = confirmatory factor analyses; ESEM = Exploratory factor analyses.

& Morin, 2015). However, this same research shows that relying on measurement models allowing for the free estimation of all cross-loadings between items and non-target factors still results in unbiased estimates of factor correlations even when no cross-loadings are present in the population model. This form of construct-relevant multidimensionality thus calls for the reliance on classical Exploratory Factor Analyses (EFA; e.g., Marsh, Morin, Parker, & Kaur, 2014; Morin, Marsh, & Nagengast, 2013), which have recently been integrated with CFA into a single overarching exploratory structural equation modeling (ESEM) framework (Asparouhov & Muthén, 2009; Myers, Chase, Pierce, & Martin, 2011). Target rotation also makes it possible to rely on a "confirmatory" approach to the estimation of ESEM factors through the pre-specification of target loadings in a confirmatory manner while cross-loadings are freely estimated but "targeted" to be as close to zero as possible (Asparouhov & Muthén, 2009).

ESEM has also been integrated with bifactor modeling in an even more comprehensive Bifactor-ESEM framework, which allows for the simultaneous consideration of construct-relevant psychometric multidimensionality related to the presence of hierarchically- and conceptually- related constructs in a single model (Morin, Arens, & Marsh, 2016; Morin, Arens, Tran, & Caci, 2016). Orthogonal bifactor target rotation even makes it possible to rely on a confirmatory approach to the estimation of such models (Reise, 2012). Statistical research further shows the importance of systematically contrasting these alternative models (CFA, Bifactor-CFA, ESEM, Bifactor-ESEM) given that each of them can absorb unmodeled sources of construct relevant multidimensionality (e.g., Asparouhov et al., 2015; Morin, Arens, & Marsh, 2016; Murray & Johnson, 2013). Thus: (a) CFA will absorb unmodeled cross-loadings or G-factors through the inflation of factor correlations; (b) Bifactor-CFA will absorb unmodeled cross-loadings through the inflation of loadings on the G-factor; (c) ESEM will absorb unmodeled G-factors through the inflation of factor correlations and/or cross-loadings. A graphical representation of first-order, hierarchical, and bifactor ESEM models is provided in the bottom section of Fig. 1.

Interestingly, prior research has demonstrated the superiority of a Bifactor-ESEM representation of SDT-based measures of motivation in the work (Howard, Gagné, Morin, & Forest, 2016) and sport (Gunnel & Gaudreau, 2015) contexts, showing that it provided a direct estimate of participants' global levels of self-determination disaggregated from the specific quality of their individual motivational characteristics. Similar results have also been found for SDT-based measures of coaching in the sport area (Appleton, Ntoumanis, Quested, Viladrich, & Duda, 2016; Stenling, Ivarsson, Hassmén, & Lindwall, 2015). Even more relevant to the present investigation, Myers, Martin, Ntoumanis, Celimli, and Bartholomew (2014) demonstrated that the Bifactor-ESEM framework was particularly well-suited to analyses of the underlying structure of the Psychological Need Thwarting Scale among a sample of 643 adolescent athletes (Bartholomew, Ntoumanis, Ryan, & Thøgersen-Ntoumani, 2011), resulting in the estimation of a global measure (G-Factor) of need thwarting properly disaggregated from athletes' individual levels of autonomy, competence, and relatedness in the sport context.

## 4. The present study

### 4.1. A Bifactor-ESEM representation of the BPNWS

A key objective of this study is to develop, and validate, a Portuguese adaptation of the BPNWS (Brien et al., 2012). However, an even broader objective is to extend prior SDT research (Myers et al., 2014), to investigate the underlying structure of employees' ratings of the satisfaction of their basic psychological needs for autonomy, competence, and relatedness at work. More specifically, the present study evaluates the extent to which the Bifactor-ESEM framework is able to provide a satisfactory representation of BPNWS ratings provided by a sample of exercise professionals. In line with mounting statistical evidence showing the superiority of an ESEM, relative to CFA, representation of the underlying structure of similar multidimensional constructs (Asparouhov et al., 2015; Morin, Arens, & Marsh, 2016; Morin, Boudrias, Marsh, McInerney et al., 2016) we expected the ESEM solution to provide a superior representation of BPNWS responses than CFA, as illustrated by an improved level of fit the to the data and lower factor correlations. Furthermore, based on Myers et al. (2014), we expected the Bifactor-ESEM solution to provide an even more optimal representation of BPNWS ratings than the alternative models (CFA, Bifactor-CFA, and ESEM solutions), as illustrated by an improved level of fit the to the data, reduced cross-loadings, and a well-defined G-factor.[2]

### 4.2. Generalizability of the BPNWS structure across gender

A second objective of this study is to assess the extent to which the structure of the BPNWS ratings remains invariant across samples of males and females respondents and the presence of latent mean differences as a function of gender. Tests of measurement invariance as a function of participants' gender are conducted to systematically test the generalizability of the results across meaningfully distinct groups of participants (e.g., Marsh, Parker, & Morin, 2016; Millsap, 2011). It is well known that any investigation aiming to provide an improved psychometric or statistical representation of latent constructs should systematically investigate the extent to which the results generalize to new samples of participants. Traditionally, this has been done either by the arbitrary creation of randomized subgroups of participants or by the costly recruitment of new samples of participants. However, these approaches are inherently flawed, for different reasons. On the one hand, the randomized division of participants into

[2] In light of statistical recommendations arguing against hierarchical models (e.g., Gignac, 2016; Morin, Arens & Marsh, 2016; Morin, Arens, Tran, & Caci, 2016; Morin, Boudrias, Marsh, Madore, & Desrumaux, 2016; Morin, Boudrias, Marsh, McInerney et al., 2016; Reise, 2012), we do not consider hierarchical CFA or hierarchical ESEM in this study. Furthermore, in CFA or ESEM models including only three first-order factors (such as those considered here), first-order and hierarchical models are identical and impossible to distinguish in terms of model fit (replacement of three factor correlations by three higher-order factor loadings).

distinct subgroups is, by definition, random. As such, failure to replicate the results calls more into question the process of randomization (if both samples are purely random, then there is no reason for the results not to replicate) than the generalizability of the results themselves. On the other hand, new groups of participants often end up differing from the original sample in more than one way (e.g., demographics, work status), some of which typically remain undocumented (e.g., testing procedure and context). A far more potent test of the extent to which the results generalize is provided by systematic tests of measurement invariance conducted across predetermined, and meaningful, subgroups of participants (Marsh et al., 2016; Millsap, 2011; Morin, Boudrias, Marsh, McInerney, et al., 2016; Morin, Meyer, Creusier, & Biétry, 2015). In the present study, given the sample size and composition, the most realistic way to assess the extent to which the results generalized to distinct sample of participants was to form these samples based on gender. This is also well-aligned with recent calls made in the organizational literature to more systematically test the extent to which results can be expect to generalize as a function of participants' age, gender, profession, culture, or other forms of diversity (e.g., Ayman & Korabik, 2010; Lukaszewski & Stone, 2012).

Still, gender was not only selected for convenience purposes. A recent meta-analysis of basic psychological needs at work (Van den Broeck et al., 2016) based on 42 studies showed women to present significantly higher levels of satisfaction of their need for relatedness relative to men, which is aligned with the fact that women generally tend to value social relationships more than men (Cross & Madson, 1997; Helgeson, 1994; Hyde, 2014). In contrast, no gender differences were found on the satisfaction of the needs for autonomy and competence. However, to be able to conclude with confidence that observed group-based mean-level differences are meaningful, it is first critical to establish that the measurement properties of the instrument remains unchanged across these groups through tests of measurement invariance (Marsh et al., 2016; Millsap, 2011). Interestingly, the few previous studies in which the measurement invariance of need satisfaction measures was tested as a function of gender supported the complete invariance of these measures (Vlachopoulos, 2008; Wilson, Rogers, Rodgers, & Wild, 2006). As such, evidence of measurement invariance across gender would provide additional support to the construct validity of the BPNWS.

### *4.3. Nomological validity of the BPNWS*

A last objective of the present study is to assess the nomological validity of the BPNWS, through associations with measures of psychological wellbeing and distress. SDT argues that “needs specify innate psychological nutriments that are essential for ongoing psychological growth, integrity, and wellbeing … and specify the necessary conditions for psychological health or wellbeing and their satisfaction is thus hypothesized to be associated with the most effective functioning” (Deci & Ryan, 2000, p. 229). Thus, it seems logical to think that workers (a) who can take on responsibilities, make decisions, and use their judgement for executing their tasks freely (autonomy), (b) who have the ability to do their work well, to feel competent, and are able to solve problems at work (competence), and (c) who feel understood, heard, and part of a well-integrated social network (relatedness) will tend to experience greater feelings of professional efficacy (e.g., feelings of being energized and exhilarated at work), and be less prone to experience high levels of burnout at work (e.g., being emotionally drained and exhausted, and experiencing feelings of depersonalization, or disconnection from others). Our decision to rely on both facets of psychological wellbeing and distress stems from research evidence showing that both poles are necessary to obtain a complete representation of employees' psychological health (Morin, Boudrias, Marsh, Madore et al., 2016; also see World Health Organization, 2014). Prior research leads us to expect that scores on the BPNWS would be negatively related to emotional exhaustion and depersonalization, and positively related to professional efficacy (Brien et al., 2012; Deci et al., 2001; Desrumaux et al., 2015; Longo et al., 2016; Van den Broeck et al., 2008, 2010, 2016; Vansteenkiste et al., 2007).

## 5. Method

### *5.1. Participants and procedures*

The sample includes 366 (172 females, 193 males, 1 did not specify gender) exercise professionals currently working in health and fitness settings. Participants' age varied between 18 and 58 years ($M = 34.16$), and they had between 1 and 35 years of work experience ($M = 7.70$). The sample comprise exercise professionals mostly working in branches of national or international chains of health clubs located in Portugal in the municipalities of Lisboa, Porto and Coimbra. Most professionals are working full-time in the context of a fixed-term employment contract and are responsible for the prescription and control of exercise plans.

Participants were recruited online through two contact lists obtained from associations of exercise professionals. Participants were informed about the study and assured that their participation was anonymous and that their responses would remain confidential. All participants signed a consent form prior to the online data collection. Online assessments were preferable to in-person assessments due to the sensitive nature of the questions, particularly prone to social desirability. Individuals completing computerized evaluations tend to be more willing to provide sensitive, personal information compared to regular in-person assessments (Schulenberg & Yutrzenka, 2004). Approval for this study was obtained from the University's research ethics council.

## 5.2. Measures

### 5.2.1. Satisfaction of basic psychological needs

The Basic Psychological Needs at Work Scale (BPNWS: Brien et al., 2012) was used to assess the professionals' need satisfaction in the work context. The BPNWS comprises 12 items (see Appendix A), representing the three factors of autonomy need satisfaction (items 3, 6, 9, and 12), competence need satisfaction (items 1, 4, 7, and 10), and relatedness need satisfaction (items 2, 5, 8, and 11). Responses are provided on a 6-point Likert- type scale ranging from 1 ("Strongly Disagree") to 6 ("Strongly Agree"). As noted above, Brien et al.' (2012) results provided support for the scale score reliability ($\alpha = 0.84$ to 0.90), factor validity, generalizability, and nomological validity (with measures of intrinsic motivation, wellbeing, distress, optimism, and procedural justice) of the BPNWS. Since then, additional studies conducted on the BPNWS have generally tended to support these initial results (Desrumaux et al., 2015; Kong, 2015; Shuck, Zigarmi, & Owen, 2015; Waddimba et al., 2016).

A standardized back-translation procedure (Hambleton & Kanjee, 1995) by a panel of experts was used to adapt the original BPNWS to Portuguese. The original items were first translated into Portuguese separately by two bilingual experts. Thereafter, translation discrepancies between the two versions were discussed to develop an initial Portuguese version of the BPNWS. An additional bilingual translator not involved in the first steps then back-translated this initial Portuguese version to the original language. The back-translated version was then compared with the original version and any inconsistencies were highlighted. These inconsistencies were removed in a further translation and the back-translation comparison process was repeated until the versions were identical.

### 5.2.2. Psychological wellbeing and distress

Professional Efficacy, Emotional Exhaustion, and Depersonalization were measured using the Portuguese version (Melo, Gomes, & Cruz, 1999) of the Maslach Burnout Inventory (MBI: Maslach, Jackson, & Leiter, 1996). This scale includes of 22 items assessing respondents' professional efficacy (8 items; $\alpha = 0.765$; e.g., I deal very effectively with the problems of my recipients), emotional exhaustion (9 items; $\alpha = 0.827$; e.g., I feel emotionally drained from my work), and depersonalization (5 items; $\alpha = 0.604$; e.g., I feel I treat some recipients as if they were impersonal objects). Responses are provided on a 6-point scale ranging from 1 ("Strongly Disagree") to 6 ("Strongly Agree"). Research has extensively documented the MBI's psychometric qualities, supporting the scale score reliability (generally $0.70 \leq \alpha \leq 0.80$), factor validity, nomological validity in relation to a wide variety of measures, and cross-cultural generalizability of responses to this instrument (for meta-analytic reviews, see Aguayo, Vargas, de la Fuente, & Lozano, 2011; Poghosyan, Aiken, & Sloane, 2009; Wheeler, Vassar, Worley, & Barnes, 2011; Worley, Vassar, Wheeler, & Barnes, 2008). Similar results have been obtained by Melo et al. (1999) regarding the scale score reliability ($0.70 \leq \alpha \leq 0.80$), factor validity, and nomological validity (in relation to measures of stress, job satisfaction, and physical health) for Portuguese version used in the present study.

## 5.3. Data analysis

### 5.3.1. Model estimation

All models were estimated using the robust maximum likelihood (MLR) estimator available in Mplus 7.3 (Muthén & Muthén, 1998–2016), which provides standard errors and fit indices that are robust to non-normality and to the Likert nature of items including five or more response categories (Finney & DiStefano, 2013). The small amount of missing data present at the item level (0% to 5.46% per item, $M = 3.78\%$) was handled with Full Information Maximum Likelihood (FIML; Enders, 2010; Graham, 2009). In a first stage, CFA, Bifactor-CFA, ESEM, and Bifactor-ESEM representations of participants' responses to the BPNWS were estimated. In first-order CFA, items were allowed to define their a priori factor, no cross-loading was allowed, and all three factors were allowed to correlate. In Bifactor-CFA, all items were allowed to define one general factor (G-factor) and one out of three a priori specific factors (S-factors), no cross-loading was allowed, and all factors were specified as orthogonal. In first-order ESEM, oblique target rotation was used to allow for the free estimation of the main loadings of the items on the three a priori factors, while all cross-loadings were estimated but targeted to be as close to zero as possible. Finally, the Bifactor-ESEM solution was estimated using orthogonal target rotation, resulting in the free estimation of items loadings on the G-factor and on one out of three S-Factors, while all cross-loadings were freely estimated but targeted to be as close to zero as possible.

For all models, we report standardized factor loadings ($\lambda$: reflecting the strength of association between ratings on each specific item and the underlying factors) and uniquenesses ($\delta$: reflecting the proportion of variance that is unique to the rating of each specific items and which incorporate item-specific random measurement error). In addition to these parameter estimates, we also report model-based omega coefficients of composite reliability, calculated as (McDonald, 1970): $\omega = (\Sigma|\lambda_i|)^2 / ([\Sigma|\lambda_i|]^2 + \Sigma\delta_{ii})$ where $\lambda_i$ are the factor loadings and $\delta_{ii}$ the error variances. Compared to classical estimates of scale score reliability (e.g. Cronbach's $\alpha$), $\omega$ has the advantage of taking into account the strength of association between the items and the latent factors ($\lambda_i$), as well as item-specific measurement errors ($\delta_{ii}$) (e.g., Dunn, Baguley, & Brunsden, 2014; Reise, Bonifay, & Haviland, 2012; Sijtsma, 2009).

### 5.3.2. Measurement invariance

Using the most appropriate model from the previous analyses, tests of measurement of invariance across gender were conducted in the following sequence (Millsap, 2011; Marsh et al., 2009; Morin, Arens, & Marsh, 2016): (1) configural invariance; (2) weak invariance (invariance of the factor loadings/cross-loadings); (3) strong measurement (invariance of the factor loadings/cross-loadings, and intercepts); (4) strict invariance (invariance of the factor loadings/cross-loadings, intercepts, and uniquenesses); (5) latent variance-covariance invariance (invariance of the factor loadings/cross-loadings, intercepts, uniquenesses, and latent variances-covariances); (6) latent means invariance (invariance of the factor loadings/cross-loadings, intercepts, uniquenesses, latent variances-covariances, and latent means). In practical terms, measurement invariance assesses the presence of different types of measurement biases in the context of group comparisons. Weak invariance assesses whether the latent constructs are defined in the same manner by their items across groups, and represents a prerequisite to all form of group comparisons and to all subsequent tests of measurement invariance. Strong invariance assesses whether mean differences in observed scores can be fully explained by mean differences at the construct level and represents a prerequisite to the comparison of latent means across groups. Strict invariance assesses whether item-specific measurement errors remain comparable across groups and represents a prerequisite to all forms of group comparisons relying on scale scores (rather than latent variables). In contrast, the last two steps do not test for the presence of measurement biases, but rather for the presence of meaningful, and unbiased (as these steps come after tests of weak, strong, and strict invariance), group differences occurring at the level of the latent variances, covariances, and means.

#### 5.3.3. Nomological validity

For tests of nomological validity, additional latent CFA factors representing professional efficacy, emotional exhaustion, and depersonalization were included to the final model. This model allowed us to obtain estimates of correlations disattenuated for measurement errors between the BPNWS factors and the convergent measures (e.g., Bollen, 1989). Given the low level of reliability of the depersonalization scale ($\alpha = 0.604$), it was particularly important to rely on a statistical approach providing a direct control for unreliability in the estimation of the relations among the various constructs considered in the present study.

#### 5.3.4. Goodness-of-fit assessment

Because the chi-square ($\chi^2$) test of exact fit tends to be oversensitive to sample size and minor model misspecifications, we relied on the following common goodness-of-fit indices: the comparative fit index (CFI), the Tucker-Lewis Index (TLI), the root mean square error of approximation (RMSEA). According to typical interpretation guidelines (e.g., Hu & Bentler, 1999; Marsh, Hau, & Grayson, 2005; Marsh, Hau, & Wen, 2004), values > 00.90 and 0.95 for the CFI and TLI are respectively considered to indicate adequate and excellent fit to the data, whereas values smaller than 0.08 or 0.06 for the RMSEA respectively support acceptable and excellent model fit. Nested models in measurement invariance tests were compared via consideration of changes (Δ) in goodness-of-fit indices, with decreases CFI and TLI of at least 0.010 or increases in RMSEA of at least 0.015 indicates a lack of invariance across groups (Chen, 2007; Cheung & Rensvold, 2002). It is important to keep in mind that goodness-of-fit indices corrected for parsimony (TLI, RMSEA) can improve with the addition of model constraints. Although $\chi^2$ and CFI should be monotonic with complexity, they can still improve with added constraints when the MLR scaling correction factors differ across models. These improvements should be considered to be random.

## 6. Results

### 6.1. Factor structure and reliability

The descriptive statistics and item correlations for participants' response to the BPNWS are reported in Table 1. The goodness-of-fit statistics of the alternative measurement models estimated based on these responses are reported in Table 2. The first-order CFA failed to achieve an acceptable level of fit to the data based on all goodness-of-fit indices (CFI and TLI ≤ 0.900; RMSEA ≥ 0.080), while the Bifactor-CFA achieved an adequate degree of fit to the data (CFI and TLI ≥ 0.900; RMSEA ≤ 0.080). The ESEM solution achieved a comparable level of fit to the data than the Bifactor-CFA solution according to the TLI (ΔTLI =

**Table 1**
Descriptive statistics and correlations for the items.

| Item | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 | 9 | 10 | 11 | 12 |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| 1. Relatedness 1 | – | | | | | | | | | | | |
| 2. Competence 1 | 0.33 | – | | | | | | | | | | |
| 3. Autonomy 1 | 0.38 | 0.37 | – | | | | | | | | | |
| 4. Relatedness 2 | 0.64 | 0.35 | 0.48 | – | | | | | | | | |
| 5. Competence 2 | 0.34 | 0.72 | 0.40 | 0.36 | – | | | | | | | |
| 6. Autonomy 2 | 0.28 | 0.37 | 0.41 | 0.36 | 0.38 | – | | | | | | |
| 7. Relatedness 3 | 0.53 | 0.28 | 0.31 | 0.56 | 0.29 | 0.38 | – | | | | | |
| 8. Competence 3 | 0.34 | 0.42 | 0.37 | 0.37 | 0.41 | 0.50 | 0.43 | – | | | | |
| 9. Autonomy 3 | 0.25 | 0.44 | 0.44 | 0.35 | 0.43 | 0.41 | 0.34 | 0.69 | – | | | |
| 10. Relatedness 4 | 0.45 | 0.26 | 0.23 | 0.48 | 0.22 | 0.24 | 0.55 | 0.34 | 0.32 | – | | |
| 11. Competence 4 | 0.30 | 0.55 | 0.39 | 0.37 | 0.51 | 0.36 | 0.34 | 0.54 | 0.52 | 0.37 | – | |
| 12. Autonomy 4 | 0.30 | 0.32 | 0.44 | 0.44 | 0.32 | 0.50 | 0.37 | 0.43 | 0.48 | 0.35 | 0.44 | – |
| Mean | 4.50 | 5.48 | 5.17 | 4.66 | 5.42 | 4.77 | 4.44 | 5.17 | 5.36 | 4.82 | 5.14 | 4.77 |
| Standard deviation | 1.01 | 0.68 | 0.93 | 1.08 | 0.74 | 1.05 | 1.12 | 0.84 | 0.82 | 1.11 | 0.90 | 1.09 |

Note: All correlations are significant at the $p < 0.01$ level.

**Table 2**
Goodness-of-fit statistics for the estimated models.

| Model | $\chi^2$ | *df* | CFI | TLI | RMSEA [90% CI] |
|---|---|---|---|---|---|
| M1. First-Order CFA | 302.63* | 51 | 0.868 | 0.829 | 0.116 [0.104, 0.129] |
| M2. Bifactor CFA | 125.22* | 42 | 0.938 | 0.903 | 0.074 [0.059, 0.089] |
| M3. ESEM | 100.29* | 33 | 0.950 | 0.900 | 0.075 [0.058, 0.092] |
| M4. Bifactor ESEM | 59.29* | 24 | 0.974 | 0.928 | 0.063 [0.043, 0.084] |

| Model | $\chi^2$ | *df* | CFI | TLI | RMSEA [90% CI] | CM | $\Delta\chi^2$ | $\Delta df$ | ΔCFI | ΔTLI | ΔRMSEA |
|---|---|---|---|---|---|---|---|---|---|---|---|
| Measurement invariance | | | | | | | | | | | |
| M5. Configural invariance | 94.67* | 48 | 0.967 | 0.908 | 0.073 [0.051, 0.095] | – | – | – | – | – | – |
| M6. Weak invariance | 138.18* | 80 | 0.958 | 0.931 | 0.063 [0.045, 0.081] | M5 | 46.77 | 32 | −0.009 | +0.023 | −0.010 |
| M7. Strong invariance | 149.41* | 88 | 0.956 | 0.934 | 0.062 [0.044, 0.079] | M6 | 10.81 | 8 | −0.002 | +0.003 | −0.001 |
| M8. Strict invariance | 171.72* | 100 | 0.949 | 0.932 | 0.063 [0.046, 0.078] | M7 | 21.78 | 12 | −0.007 | −0.002 | +0.001 |
| M9. Var-Cov invariance | 176.94* | 110 | 0.952 | 0.943 | 0.058 [0.042, 0.073] | M8 | 10.40 | 10 | +0.003 | +0.011 | −0.005 |
| M10. Latent means invariance | 189.54* | 114 | 0.946 | 0.937 | 0.060 [0.045, 0.075] | M9 | 12.76 | 4 | −0.006 | −0.006 | +0.002 |

Note: CFA = confirmatory factor analyses; ESEM = exploratory factor analyses; $\chi^2$ = scaled chi-square test of exact fit; *df* = degrees of freedom; CFI = -comparative fit index; TLI = Tucker-Lewis index; RMSEA = root mean square error of approximation; 90% CI = 90% confidence interval of the RMSEA; Var-Cov = variance-covariance; CM = comparison model; Δ = change in fit information relative to the CM.
* $p < 0.01$.

−0.003) and RMSEA (ΔRMSEA = +0.001), but achieved excellent fit to the data according to the CFI (CFI ≥ 0.950; ΔCFI = +0.012). Like the ESEM solution, the bifactor ESEM solution achieved an excellent level of fit to the data according to the CFI (≥0.950), an adequate level of fit to the data according to the TLI (≥0.900) and RMSEA (≤0.080), and a substantially higher level of fit to the data than the ESEM solution (ΔCFI = +0.024; ΔTLI +0.028; ΔRMSEA = −0.012). This statistical information supports the superiority of the Bifactor-ESEM solution, and the need for the model to account for both sources of construct-relevant psychometric multidimensionality. However, each alternative model is able absorb unmodeled sources construct-relevant psychometric multidimensionality, thus hiding sources of misfit behind apparently similarly fitting models (e.g., Asparouhov et al., 2015; Morin, Arens & Marsh, 2016; Murray & Johnson, 2013). For this reason, Morin and colleagues (Morin, Arens, & Marsh, 2016; Morin, Arens, Tran, & Caci, 2016; Morin, Boudrias, Marsh, McInerney et al., 2016) note that this comparison of goodness-of-fit indices is not a sufficient basis in the selection of the model providing the optimal representation of the data, and needs to be complemented by a comparison of parameter estimates. These parameter estimates associated with these models are reported in Table 3 for the first-order CFA and ESEM solutions, and Table 4 for the Bifactor-CFA and Bifactor-ESEM solutions.

We first turn our attention to the comparison of the first-order CFA and ESEM solutions. In CFA, all factors were well defined by high and significant factors loadings (λ = 0.64–0.81; *M* = 0.71; $p < 0.01$), resulting in satisfactory model-based coefficients of composite reliability (ω = 0.76 to 0.82, *M* = 0.80) for all factors. However, ESEM revealed a few unexpected results, consistent with the idea that ESEM may sometimes prove to be particularly helpful in revealing sources of misfit in psychometric measures that would otherwise remain hidden in CFA (Morin & Maïano, 2011; Morin et al., 2013). First, the autonomy factor appeared to be

**Table 3**
Standardized factor loadings (λ) and uniquenesses (δ) for the first-order CFA and ESEM solutions.

| Indicator | CFA | | ESEM | | | |
|---|---|---|---|---|---|---|
| | λ | δ | RS λ | CS λ | AS λ | δ |
| Relatedness | | | | | | |
| 1 | **0.74** | 0.45 | **0.79** | *0.05* | −0.14 | 0.41 |
| 4 | **0.81** | 0.35 | **0.82** | *0.04* | *−0.04* | 0.33 |
| 7 | **0.73** | 0.46 | **0.70** | *−0.07* | 0.13 | 0.47 |
| 10 | **0.64** | 0.59 | **0.59** | *−0.06* | *0.13* | 0.61 |
| ω | 0.82 | | 0.82 | | | |
| Competence | | | | | | |
| 2 | **0.76** | 0.42 | *−0.06* | **0.93** | *−0.08* | 0.26 |
| 5 | **0.75** | 0.44 | *0.03* | **0.90** | *−0.09* | 0.29 |
| 8 | **0.68** | 0.54 | *0.07* | **0.19** | 0.68 | 0.32 |
| 11 | **0.74** | 0.46 | *0.07* | **0.47** | 0.30 | 0.50 |
| ω | 0.82 | | | 0.82 | | |
| Autonomy | | | | | | |
| 3 | **0.64** | 0.59 | 0.30 | 0.24 | ***0.18*** | 0.65 |
| 6 | **0.64** | 0.59 | 0.19 | 0.21 | **0.34** | 0.64 |
| 9 | **0.71** | 0.49 | *−0.03* | 0.25 | **0.69** | 0.32 |
| 12 | **0.67** | 0.55 | 0.29 | *0.10* | **0.37** | 0.60 |
| ω | 0.76 | | | | 0.53 | |

Note: CFA = confirmatory factor analyses; ESEM = Exploratory factor analyses; ω = omega coefficient of composite reliability; RS = relatedness satisfaction; CS = competence satisfaction; AS = autonomy satisfaction; bold = target factor loadings. Non-significant parameters ($p \geq 0.05$) are marked in italics.

**Table 4**
Standardized factor loadings (λ) and uniquenesses (δ) for the Bifactor CFA and ESEM solutions.

| Indicator | B-CFA | | | B-ESEM | | | | |
|---|---|---|---|---|---|---|---|---|
| | G λ | S λ | δ | G-λ | S-RS λ | S-CS λ | S-AS λ | δ |
| Relatedness | | | | | | | | |
| 1 | **0.65** | **0.42** | 0.40 | **0.44** | **0.60** | 0.10 | *0.09* | 0.43 |
| 4 | **0.60** | **0.53** | 0.37 | **0.53** | **0.59** | *0.06* | 0.26 | 0.31 |
| 7 | **0.52** | **0.51** | 0.47 | **0.54** | **0.52** | −0.07 | *−0.08* | 0.42 |
| 10 | **0.46** | **0.44** | 0.60 | **0.48** | **0.47** | −0.05 | *−0.15* | 0.53 |
| ω | | 0.66 | | | 0.74 | | | |
| Competence | | | | | | | | |
| 2 | **0.70** | **0.57** | 0.19 | **0.59** | *0.01* | **0.67** | *−0.02* | 0.20 |
| 5 | **0.58** | **0.56** | 0.35 | **0.57** | *0.02* | **0.58** | *0.07* | 0.34 |
| 8 | ***−0.04*** | **0.80** | 0.37 | **0.83** | *−0.07* | ***−0.10*** | *−0.10* | 0.29 |
| 11 | **0.25** | **0.66** | 0.50 | **0.67** | *0.02* | **0.22** | *−0.01* | 0.50 |
| ω | | 0.83 | | | | 0.65 | | |
| Autonomy | | | | | | | | |
| 3 | **0.17** | **0.57** | 0.65 | **0.53** | 0.12 | 0.09 | **0.50** | 0.45 |
| 6 | **0.45** | **0.62** | 0.42 | **0.59** | *0.04* | *0.02* | ***0.19*** | 0.61 |
| 9 | ***−0.19*** | **0.81** | 0.30 | **0.80** | −0.15 | *−0.05* | ***0.03*** | 0.33 |
| 12 | **0.25** | **0.63** | 0.54 | **0.60** | *0.11* | *−0.05* | **0.25** | 0.56 |
| ω | 0.82 | 0.78 | | 0.91 | | | 0.33 | |

Note: B-CFA = bifactor confirmatory factor analyses; B-ESEM = bifactor exploratory factor analyses; G = global factor estimated as part of a bifactor model; S - = specific factor estimated as part of a bifactor model; ω = omega coefficient of composite reliability; RS = relatedness satisfaction; CS = competence satisfaction; AS = autonomy satisfaction; bold = target factor loadings. Non-significant parameters ($p \geq 0.05$) are marked in italics.

globally less well defined ($\lambda = 0.18$ to 0.69; $M = 0.39$) than the other factors, with most items presenting a variety of cross-loadings on the other factors ($|\lambda| = 0.03$ to 0.30; $M = 20$). This result suggests that these items may be more relevant to the assessment of global need satisfaction than of a specific and distinct factor tapping into the satisfaction of employees' need for autonomy, reinforcing the need to pursue a bifactor representation of the data. Second, item 8 (I am able to solve problems at work) had a stronger cross-loading on the autonomy factor ($\lambda = 0.68$; $p < 0.01$) than on the a priori competence factor ($\lambda = 0.19$; $p < 0.05$). This ESEM result suggests that problem solving ability might potentially be more relevant to employees' ability to satisfy their needs for autonomy at work than their needs for competence, at least in the present sample of Portuguese exercise professionals. Still, it is important to acknowledge that this strong cross-loading could also reflect the presence of an unmodeled G-factor. Whereas the first interpretation suggests the need to further explore the reasons underlying this surprising result (e.g., content validity, culture/language, type of employment, etc.) and the possibility to investigate alternative measurement models, the second hypothesis suggests that the bets-fitting Bifactor-ESEM solution should first be examined. With these exceptions, the relatedness and competence factors were well defined by satisfactory factor loadings from the remaining items (0.47 to 0.93; $M = 0.74$; $p < 0.01$) and reasonably low cross-loadings ($|\lambda| = 0.03$ to 0.30, $M = 0.09$).

Consistent with these results, the ESEM-based coefficients of composite reliability proved to be much higher for the relatedness ($\omega = 0.82$) and competence ($\omega = 0.82$) factors than for the autonomy factor ($\omega = 0.52$). Finally, the observation of much reduced factors correlations in ESEM ($r_{autonomy\text{-}competence} = 0.48$; $r_{autonomy\text{-}relatedness} = 0.46$; $r_{competence\text{-}relatedness} = 0.55$; $M = 0.50$; $p < 0.01$), relative to CFA ($r_{autonomy\text{-}competence} = 0.85$; $r_{autonomy\text{-}relatedness} = 0.69$; $r_{competence\text{-}relatedness} = 0.61$; $M = 0.72$; $p < 0.01$) lends additional support to the importance of relying on an model allowing for the consideration of construct-relevant psychometric multidimensionality related to the assessment of conceptually-related constructs.

The next step of this comparison involves the best fitting Bifactor-ESEM solution. Supporting the superiority of this solution, the results revealed a highly reliable G-Factor ($\omega = 0.91$) defined by strong and positive loadings from all items ($\lambda = 0.44$ to 0.83, $M = 0.60$; $p < 0.01$). Further attesting to the superiority of this solution, cross-loadings remained small ($|\lambda| = 0.01$ to 0.26, M = 0.06), and smaller than in ESEM ($|\lambda| = 0.03$ to 0.68, M = 0.15). Consistent with our prior interpretation of the ESEM solution, item 11 appeared to provide a clearer representation of the need satisfaction G-Factor ($\lambda_8 = 0.83$; $\lambda_{11} = 0.67$) than of the a priori competence S-Factor ($\lambda_8 = -0.10$; $\lambda_{11} = 0.22$). Similarly, all autonomy items tended to provide a better representation of the need satisfaction G-Factor ($\lambda = 0.53$ to 0.80, $M = 0.63$) than of the autonomy S-Factor ($\lambda = 0.03$ to 0.50, $M = 0.24$), and to retain only a very limited level of specificity once the variance explained by the G-factor is taken into account. Finally, item 8 also appeared to provide a better representation of the need satisfaction G-Factor ($\lambda_8 = 0.83$) than of the competence S-Factor ($\lambda_8 = -0.10$), suggesting that the previously elevated cross-loading of this item observed on the autonomy factor in the context of the ESEM solution was in fact related to the presence of an unmodeled G-factor. Overall, these results support the superiority of the Bifactor-ESEM solution, which is retained for further analyses.

Consistent with these observations, the model-based coefficients of composite reliability proved to be much lower for the autonomy S-Factor ($\omega = 0.325$), then for the relatedness ($\omega = 0.740$) and competence ($\omega = 0.651$) S-factors, which appeared to tap into meaningful specificity once the variance explained by the G-factor is taken into account. Indeed, the relatedness and competence S-Factors were defined by satisfactory factor loadings from most remaining items (0.47 to 0.67; M = 0.57; $p < 0.01$), supporting their consideration in the context of analyses of relations among latent constructs.

### 6.2. Measurement invariance

Starting from the retained Bifactor-ESEM solution, we proceeded to tests of measurement invariance of the BPNWS ratings across samples of males and females participants. The results from these tests are reported in the bottom section of Table 3. The model of configural invariance provided an adequate fit to the data (CFI = 0.967, TLI = 0.908, RMSEA = 0.073). Invariance constraints were then progressively added to the factor loadings (weak invariance), intercepts (strong invariance), uniquenesses (strict invariance), latent variance-covariance matrix, and latent means. None of these steps resulted in a decrease in model fit exceeding the recommended guidelines (ΔCFI and ΔTLI ≥ 0.01, ΔRMSEA ≥ 0.015, and non-significant $\Delta\chi^2$), supporting the complete equivalence of BPNWS ratings across genders. Although the changes in fit indices associated with the invariance of the latent means remained under the recommended guidelines, these changes remained high for both the CFI (−0.006) and TLI (−0.006). Thus, based on our theoretical expectation that gender differences in mean levels should be present on the BPNWS factors, as well as on statistical evidence suggesting that typical guidelines may not be stringent enough for tests of latent mean differences (Fan & Sivo, 2009), we decided to interpret latent mean differences obtained from the model of latent variance-covariance invariance. In multi-group models, latent means are constrained to be zero in a referent group for identification purposes, and freely estimated in the other groups (e.g., Morin et al., 2013). These freely estimated latent means provide a direct estimation of the magnitude of the difference between the target group and the referent group, expressed in SD units, and are accompanied by tests of statistical significance. In the current study, when females' latent means were constrained to be zero for identification purposes, males' latent means proved to be significantly lower (−0.33 SD; $p \leq 0.01$) on the need satisfaction G-factor, but not significantly different for any of the S-factors ($p \geq 0.05$).

### 6.3. Nomological validity

Starting again from the retained Bifactor-ESEM solution, we proceeded to test of nomological validity of the BPNWS ratings through the addition of additional CFA factors representing professional efficacy, emotional exhaustion, and depersonalization to the model. For illustration purposes, we report similar relations estimated from the CFA, Bifactor-CFA, and ESEM solutions. However, it is important to reinforce that our results clearly supported the superiority of the Bifactor-ESEM solution, meaning that any observed difference across solutions is likely to reflect the presence of biases in the CFA, Bifactor-CFA, and ESEM solutions (e.g., Morin, Arens, Tran, & Caci, 2016).

The latent correlations between these additional latent factors and the BPNWS G- and S-Factors are reported in Table 5. When the results are contrasted across alternative solutions, it is very interesting to note that both the CFA and ESEM solutions revealed that all three need satisfaction factors were similarly related to higher levels of professional efficacy and to lower levels of emotional exhaustion and depersonalization. Although these results showed slightly stronger relations involving the competence factor, and slightly weaker relations involving the autonomy factor (particularly in ESEM where this factor was more weakly defined), they still suggested that all three need satisfaction factors presented a generally undifferentiated pattern of associations with all convergent measures.

The results from the Bifactor-CFA and Bifactor-ESEM solutions were highly similar to one another. These results showed that the apparently undifferentiated pattern of associations was in fact related to the presence of an unmodeled global need satisfaction factor presenting clear associations with all three convergent measures. Consistent with the importance of taking into account employees' levels of global need satisfaction in such predictive models, the G-factor presented a significant positive relation with

**Table 5**
Nomological validity.

| | Professional efficacy | Emotional exhaustion | Depersonalization |
|---|---|---|---|
| CFA solution | | | |
| Relatedness satisfaction | 0.44** | −0.35** | −0.25** |
| Competence satisfaction | 0.57** | −0.27** | −0.36** |
| Autonomy satisfaction | 0.45** | −0.32** | −0.32** |
| Bifactor-CFA solution | | | |
| Global need satisfaction | 0.47** | −0.24** | −0.31** |
| Relatedness satisfaction | 0.16 | −0.21** | −0.05 |
| Competence satisfaction | 0.25** | −0.09 | −0.19* |
| Autonomy satisfaction | 0.01 | −0.24 | 0.01 |
| ESEM solution | | | |
| Relatedness satisfaction | 0.43** | −0.36** | −0.25** |
| Competence satisfaction | 0.53** | −0.26** | −0.36** |
| Autonomy satisfaction | 0.33** | −0.17* | −0.22* |
| Bifactor-ESEM solution | | | |
| Global need satisfaction | 0.47** | −0.26** | −0.31** |
| Relatedness satisfaction | 0.15 | −0.22** | −0.03 |
| Competence satisfaction | 0.26** | −0.06 | −0.19* |
| Autonomy satisfaction | 0.02 | −0.17 | −0.07 |

* $p < 0.05$.
** $p < 0.01$.

professional efficacy, and significant negative relations with emotional exhaustion and depersonalization. However, these results also showed a well-differentiated pattern of associations between the S-factors and the convergent measures once the variance explained by the G-factor was taken into account. More precisely, these results showed that the competence S-Factor presented a significant negative association with depersonalization and a significant positive relation with professional efficacy, whereas the relatedness S-factor presented a significant negative relation with emotional exhaustion.

Finally, these results also revealed non-significant relations between the autonomy S-Factor and the convergent variables, which is also consistent with the fact that this S-Factor retained a much lower level of specificity than the other S-factors, and was associated with a very low level of reliability ($\omega = 0.325$). In practical terms, it makes sense to observe that a factor including only a negligible amount of residual specificity, once the variance in items' ratings explained by the G-factor is taken into account, would present no significant associations with the convergent measures. It is important to note that this result does not suggest that the satisfaction of employees' needs for autonomy does not share meaningful relations with the convergent measures. Rather, it suggests that employees' ratings of the satisfaction of their need for autonomy provide a direct estimate of their global level of need satisfaction at work, and retain no meaningful specificity once this more global level of need satisfaction is taken into account.

It is important to reinforce that this similarity of results observed between the Bifactor-CFA and Bifactor-ESEM solutions does not suggest that both solutions are equivalent. Indeed, Table 4 clearly reveals that the G-factor and relatedness S-factor are more clearly defined in the Bifactor-ESEM solution, while the competence and autonomy S-factor are more clearly defined in the Bifactor-CFA solution. Similarly, Table 2 shows that the Bifactor-CFA solution provided a drastically reduced level of fit to the data when compared to the Bifactor-ESEM solution, supporting the need to account for cross-loadings in the model (e.g., Morin, Arens, & Marsh, 2016; Morin, Arens, Tran, & Caci, 2016). Simply, these results demonstrate that the nomological validity of both solutions is similar in relation to the convergent measures considered here.

## 7. Discussion

Following from Myers et al. (2014), the present study aimed to investigate the structure of employees' ratings of the satisfaction of their basic psychological needs for autonomy, competence, and relatedness at work using the newly developed Bifactor-ESEM framework. This overarching framework is specifically designed to assess, and explicitly disaggregate, two sources of construct-relevant psychometric multidimensionality in employees' ratings of their need satisfaction at work and related to: (1) the assessment of conceptually-related constructs; and (2) the hierarchical nature of multidimensional need satisfaction measures which can logically be used to assess employees' global levels of need satisfaction as well as their more specific level of autonomy, relatedness, and competence satisfaction. Whereas the identification of the first source of construct-relevant psychometric multidimensionality can be achieved by the comparison of first-order CFA and ESEM solutions, the second source can be identified by comparing first-order and bifactor solutions.

The first of these comparisons supported the superiority of the ESEM representation of BPNWS ratings, which resulted in a higher level of fit to the data and less correlated (i.e., more differentiated) factors relative to CFA. As noted above, ESEM has been shown to provide more precise estimates of factor correlations whenever cross-loadings are present in the population model, yet to remain unbiased in the absence of cross-loadings (Asparouhov et al., 2015). As such, the observation of reduced ESEM, relative to CFA, factor correlations represents a strong source of support to the superiority of the ESEM solution (e.g., Morin, Arens, & Marsh, 2016; Morin, Boudrias, Marsh, McInerney et al., 2016). Starting from the retained ESEM solution, the second comparison then supported the superiority of the Bifactor-ESEM representation of BPNWS ratings, both in providing a higher level of fit to the data, but also in revealing a well-defined G-Factor accompanied by substantially reduced cross-loadings.

Beyond these statistical considerations, it is important to keep in mind that when Morin and colleagues (Morin, Arens, & Marsh, 2016; Morin, Arens, Tran, & Caci, 2016) proposed the concept of construct-relevant psychometric multidimensionality, they explicitly insisted on the fact that these two components (i.e., cross-loadings and G-Factor) did not simply refer to some form of random noise to be controlled for in the model, but were substantively meaningful in their own right. Morin, Arens, Tran, and Caci (2016, p. 2–3) note that "In CTT [Classical Test Theory], ratings are assumed to reflect a combination of true score variance and random measurement error (estimated in reliability analyses). By definition, "random" measurement error is unrelated to other constructs, leading to its absorption within the indicators' uniquenesses in CFA. CTT further distinguishes among construct-relevant and construct-irrelevant forms of true score variance, a distinction covered in discussions of validity. This distinction makes it obvious that indicators are expected to include at least some degree of true score association with other constructs."

Although it is easy to make sense of the G-factor as reflecting the global extent to which employees feel that their basic needs are met at work, cross-loadings are less straightforward to interpret. Or are they? As noted by Asparouhov et al. (2015), because the meaning of latent constructs is related to the way they relate to other constructs, the fact that allowing for cross-loadings to be freely estimated results in a more exact depiction of the true underlying factor correlations suggests that cross-loadings help to achieve a more exact estimation of the factors. In fact, the observation of cross-loadings suggests that some item ratings simultaneously tap into the satisfaction of more than one basic need, albeit at different levels. This is perfectly consistent with the idea that, to some extent, autonomy helps to ensure that employees' can express their competencies in their workplaces, or have the latitude required to build positive relationships. Similarly, having strong and positive relationships, or high levels of competencies, may in turn favor the emergence of greater levels of autonomy. Finally, competent employees' may mentor their peers, encouraging the development of positive relationships, just like strong relationships increase one's accessibility to positive mentoring.

Our results also revealed that some BPNWS items provided a better reflection of employees' global need satisfaction than of their specific need for relatedness, competence, or autonomy. Yet, these observations also made sense, for instance in suggesting that problem solving skills (item 8) or global feelings of success at work (item 11), might provide a stronger measure of global need satisfaction for this sample of exercise professionals for whom daily work involves relationships with others and a high level of autonomy. In this regard, it is perhaps not surprising for our results to suggest that autonomy ratings play such a central role in the definition of the global need satisfaction factor. Among exercise professionals, work is by nature autonomous, so that a lack of autonomy is likely to create problems at all levels in limiting employees' ability to build positive relationships and to express their competencies at work. Clearly, the present results need to be replicated using more diversified sample of employees in order to test this interpretation.

A key strength of the Bifactor-ESEM framework is that it allows for the examination of the outcomes simultaneously associated with global and specific constructs simultaneously (Myers et al., 2014). In accordance with prior research (Boudrias et al., 2014; Deci et al., 2001; Van den Broeck et al., 2008), our results provided a strong support to the Bifactor-ESEM solution in showing that employees' global levels of need satisfaction at work presented the strongest relations with all three measures of wellbeing and distress considered in this study. Our results also supported SDT assertion regarding the importance of separately considering the satisfaction of employees' basic needs for competence and relatedness (Deci & Ryan, 2000, 2008; Gagné & Deci, 2005) in showing that these S-Factors uniquely predicted lower levels of emotional exhaustion (relatedness) and depersonalization (competence), and higher levels of professional efficacy (competence), over and above the prediction already afforded by employees' global levels of need satisfaction. These results are also consistent with those from prior research (e.g., Brien et al., 2012; Deci et al., 2001; Desrumaux et al., 2015). Unfortunately, analyses of the nomological validity of the factors extracted here did not support the idea that there was sufficient residual specificity associated with the autonomy S-factor to result in significant relations with the wellbeing and distress constructs considered in this study. As noted above, this suggests the need to replicate the present study among samples of employees' for whom autonomy is not such a normative characteristic of the job.

A secondary purpose of this study was to develop a Portuguese adaption of the BPNWS (Brien et al., 2012) and to evaluate its psychometric properties among a sample of exercise professionals. The current study supported the factor validity of scores on this instrument, as well as their composite reliability and nomological validity in relations to measures of psychological wellbeing and distress. As noted above, it is to be expected for S-Factors estimated in the context of bifactor models to be associated with lower levels of composite reliability, due in part to the fact that their items tend to provide a stronger reflection of the G-Factor than or the S-Factor. In this study, this was the case of the autonomy factor. This observation suggests that, pending replication, autonomy items should be mainly used to assess employees' global level of need satisfaction, and not necessarily to provide a separate assessment of autonomy satisfaction.

Finally, a key condition of a successful psychometric validation study is related to the ability to demonstrate that the psychometric properties of scores on an instrument generalize to meaningfully distinct subgroups of participants (Millsap, 2011). In the present study, this generalizability was assessed through systematic tests of measurement invariance of the BPNWS ratings across samples of male and female employees. In line with prior research (Vlachopoulos, 2008; Wilson et al., 2006), our results demonstrated the complete measurement invariance of ratings provided by male and female employees to the Portuguese version of the BPNWS, supporting the applicability of this instrument to males and females, as well as the generalizability of our results to meaningfully distinct samples of employees. However, in contrast with recent meta-analytic evidence (Van den Broeck et al., 2016) suggesting that females' may present higher levels of relatedness need satisfaction that males, we found no gender difference in latent mean levels on any of the S-factors. Rather, our results suggest that, once latent means differences are assessed from a properly invariant measurement model and adequately represented via Bifactor-ESEM, females employees rather tend to present a higher global level of need satisfaction than their males counterparts. In other words, gender differences appear to be located at the level of global needs satisfaction, rather than at the level of the S-factors.

### *7.1. Limitations and directions for future research*

There are some limitations to the present study that should be kept in mind when considering the results. First, although our goal was to provide an improved representation of the BPNWS that could, in theory, be generalized to any employee from any culture, the current study relied on a convenience sample of Portuguese male and females exercise professionals, which restricts the generalizability of our results. Thus, the next step would be to test the extent to which the present results generalize to more diverse linguistic, cultural, and professional contexts. Second, despite the fact that the statistical approach used here to assess the relation between the BPNWS factors and the convergent measures provided a natural control for unreliability, it remains important to keep in mind the fact that the reliability of the depersonalization scale was minimal ($\alpha = 0.60$). As such, future research would do well to more systematically assess the reasons for this high level of unreliability, and to consider the adoption of latent variable models, such as those used in the present study, to provide a way to obtain unbiased estimates of relations among constructs despite the presence of measurement errors. This last conclusion regarding the importance of relying on fully latent models is further reinforced by the fact that the composite reliability of the relatedness ($\omega = 0.74$) and competence ($\omega = 0.65$) S-Factors was also minimal. Third, the current study highlighted the measurement scale as invariant across gender. Future research could also evaluate if the BPNWS works equally well in relation to additional variables, like tenure, workload, familial obligations, etc. Last, but not least, the current study is cross-sectional in nature, precluding any test of the directionality of the associations between the need satisfaction factors and the convergent measures. Even more importantly, there is a clear need to assess the extent to which the proposed "improved" Bifactor-ESEM representation of employees' ratings of the satisfaction of their basic needs at work generalize over time, and the extent to which scores on the various G- and S-Factor will fluctuate over time within specific employees.

## 8. Conclusion

This study aimed to investigate the structure of employees' ratings of the satisfaction of their basic psychological needs for autonomy, competence, and relatedness at work using the newly developed Bifactor-ESEM framework. Our results supported the need to rely on this comprehensive approach to psychometric measurement, which provides a way to estimate need satisfaction factors while simultaneously taking into account the fact that need satisfaction ratings include two sources of construct relevant psychometric multidimensionality related to the assessment of conceptually-related and hierarchically-ordered constructs. Similar results have been previously found in SDT-related research focusing on the structure of measures of motivation (Gunnel & Gaudreau, 2015; Howard et al., 2016), need thwarting (Myers et al., 2014), and coaching (Appleton et al., 2016; Stenling et al., 2015), attesting to the flexibility of the approach. At a more practical level, the present study also demonstrated the role of need satisfaction as a significant predictor of wellbeing and distress, while highlighting the complementary role of global, and specific, need satisfaction. Interestingly, the Bifactor-ESEM approach considered here provides a way to conduct this assessment while relying on a model not tainted by multicollinearity and providing a natural disaggregation of the effects attributable to global need satisfaction relative to the specific needs for autonomy, relatedness and competence. Finally, the present study proposed a valid and reliable Portuguese measure of need satisfaction at work, thus contributing to the feasibility of cross-cultural studies conducted across the European continent. Although this instrument was tested with a sample of exercise professionals, the nature of the items allows future researchers to analyze the need satisfaction in other work-contexts.

## Appendix A. Portuguese and English versions of the Basic Psychological Needs at Work Scale

| Portuguese version | English version |
|---|---|
| 1. Quando estou com as pessoas com quem trabalho, sinto-me compreendido(a). | 1. When I'm with the people from my work environment, I feel understood |
| 2. Sinto que tenho capacidade para desempenhar bem o meu trabalho. | 2. I have the ability to do my work well |
| 3. O meu trabalho permite-me tomar decisões. | 3. My work allows me to make decisions |
| 4. Quando estou com as pessoas com quem trabalho, sinto-me ouvido(a). | 4. When I'm with the people from my work environment, I feel heard |
| 5. Sinto-me competente no trabalho. | 5. I feel competent at work |
| 6. Posso basear-me no meu próprio julgamento para resolver problemas no meu trabalho. | 6. I can use my judgement when solving work-related problems |
| 7. Quando estou com as pessoas com quem trabalho, sinto que posso confiar nelas. | 7. When I'm with the people from my work environment, I feel as though I can trust them |
| 8. Sou capaz de resolver os problemas que surgem no trabalho. | 8. I am able to solve problems at work |
| 9. Sinto que posso assumir responsabilidades no meu trabalho. | 9. I can take on responsibilities at my job |
| 10. Quando estou com as pessoas com quem trabalho, sinto que sou um amigo(a) para elas. | 10. When I'm with the people from my work environment, I feel I am a friend to them |
| 11. Sinto-me bem-sucedido no meu trabalho. | 11. I succeed in my work |
| 12. No meu trabalho, sinto-me livre para realizar as minhas tarefas à minha maneira. | 12. At my work, I feel free to execute my tasks in my own way |

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