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ORIGINAL PAPER

Exploratory Structural Equation Modeling Analysis of the Self-Compassion Scale

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Abstract The present research investigated the construct validity and psychometric properties of the Self-Compassion Scale (SCS) with a new and advanced statistical procedure, exploratory structural equation modeling (ESEM), in order to contribute to the ongoing discussion about its dimensionality by employing a bifactor-ESEM framework. A Hungarian online representative sample (N = 505, N_{female} = 265, M_{age} = 44.37) filled out the Hungarian version of the SCS. Confirmatory factor analysis (CFA) and ESEM methods were employed, and first-order and bifactor solutions were examined and compared. The bifactor ESEM model demonstrated the best fit to the data with the joint presence of the general self-compassion factor and the specific factors. Internal consistency was adequate in all cases. Reliability indices—omega and omega hierarchical—showed that not all specific factors had unique contributions over and above the general factor. High levels of gender invariance were also achieved with females having lower general self-compassion and self-judgment latent means, while having higher self-kindness scores. The findings shed new light on the underlying theory behind the SCS and proved the usefulness of the bifactor ESEM framework in the investigation of multidimensional constructs.

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Keywords Bifactor · Confirmatory factor analysis (CFA) · Exploratory structural equation modeling (ESEM) · Invariance · Self-compassion scale (SCS)

Introduction

Over the last decade, self-compassion has increasingly become a topic of investigation in the mindfulness literature. Self-compassion involves the recognition of one's own suffering and the desire to ease it with understanding and kindness (Neff 2003). It has been positively linked with well-being, happiness, and life-satisfaction (e.g., Leary et al. 2007; Neff 2011; Neff and Germer 2013; Zessin et al. 2015) and negatively with different aspects of psychopathology such as depression, anxiety, stress, or eating disorders (e.g., Ferreira et al. 2013; Friis et al. 2016; Krieger et al. 2013; MacBeth and Gumley 2012). In order to measure these dimensions, the Self-Compassion Scale (SCS) was created (Neff 2003). The SCS is a 26-item instrument that measures self-compassion through the three hypothesized dimensions with their negative counterparts: self-kindness versus self-judgment, common humanity versus isolation, and mindfulness versus over-identification.

Regarding its psychometric properties, in the original validation study, Neff (2003) identified a six-factor first-order structure with the six intercorrelated factors and a six-factor second-order (the term "second-order" or "higher-order" can be used interchangeably) model with a general self-compassion factor behind the six components and subsequently chose the latter as a default model. Several replications have been performed in different countries (see Table 1), but the results are not without contradictions. Although the original six-factor second-order solution (as proposed by Neff 2003) has been investigated in multiple countries—for instance in

Table 1 Prior validity and reliability characteristics of the Self-Compassion Scale[a]

Authors	Nation	Sample	Analysis	Characteristics	Self-kindness	Self-judgment	Common humanity	Isolation	Mindfulness	Over-identification	Σ	CFI	TLI	RMSEA	Final model
Arimitsu (2014)	Japan	N = 366	CFA	N of items	5	5	4	4	4	4	26	0.86	0.83	0.066	6-factor first-order
		M_{age} = 19.6		Average loadings	0.69	0.62	0.68	0.62	0.64	0.64					
				Cronbach's alpha	0.82	0.76	0.77	0.72	0.73	0.73					
Azizi et al. (2013)	Iran	N = 265	CFA	N of items	5	5	4	4	4	4	26	0.90	–	0.08	6-factor first-order
		M_{age} = 22.1		Average loadings	0.67	0.86	0.83	0.83	0.81	0.91					
				Cronbach's alpha	0.79	0.79	0.93	0.90	0.88	0.88					
Benda and Reichová (2016)[b]	Czech Republic	N = 5368	CFA	N of items	–	–	–	–	–	–	20	0.929	0.913	0.059	6-factor first-order
		M_{age} = -		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	0.86	0.67	0.75	0.78	0.65	0.69					
Bento et al. (2016)	Portugal (pregnant women)	N = 417	CFA	N of items	–	–	–	–	–	–	–	0.94	0.93	0.00	6-factor first-order
		M_{age} = 33		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	Ranged from 0.77 to 0.87										
Castilho et al. (2015)[c]	Portugal (non-clinical)	N = 1128	CFA	N of items	5	5	4	4	4	4	26	0.92	0.91	0.07	6-factor second-order
		M_{age} = 24.5		Average loadings	Ranged from .56 to .86										
				Cronbach's alpha	Ranged from 0.70 to 0.88										
	Portugal (clinical)	N = 316	CFA	N of items	5	5	4	4	4	4	26	0.84	0.82	0.08	6-factor second-order
		M_{age} = 28.69		Average loadings	Ranged from 0.51 to 0.81										
				Cronbach's alpha	Ranged from 0.70 to 0.88										
Chen et al. (2011)	China	N = 660	EFA & CFA	N of items	–	–	–	–	–	–	–	–	–	–	6-factor first-order
		M_{age} = —		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	–	–	–	–	–	–					
Costa et al. (2015)[d]	Portugal (clinical)	N = 361	CFA	N of items	13			13			26	0.880	0.868	0.070	2-factor first-order
		M_{age} = 25.19		Average loadings	0.72			0.68							
				Cronbach's alpha	0.91			0.89							
Cunha et al. (2016)[e]	Portugal (adolescent)	N = 3165	CFA	N of items	5	5	4	4	4	4	26	0.92	0.91	0.05	6-factor second-order
		M_{age} = 15.49		Average loadings	0.64	0.63	0.57	0.68	0.61	0.66					
				Cronbach's alpha	0.77	0.77	0.72	0.79	0.70	0.75					
de Souza and Hutz (2016)[f]	Brazil	N = 432	CFA	N of items	5	5	4	4	4	4	26	0.937	0.928	0.071	6-factor first-order
		M_{age} = 32.5		Average loadings	0.72	0.68	0.60	0.76	0.72	0.69					
				Cronbach's alpha	0.81	0.77	0.66	0.79	0.77	0.76					
Deniz et al. (2008)[g]	Turkey	N = 341	EFA	N of items	24						24	–	–	–	1-factor
		M_{age} = 19.81		Average loadings	0.55										
				Cronbach's alpha	0.89										
Dundas et al. (2016)	Norway	N = 277	CFA	N of items	5	5	4	4	4	4	26	0.82	—	0.079	6-factor second-order
		M_{age} = 22.9		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	0.79	0.80	0.80	0.82	0.86	0.71					
Garcia-Campayo et al. (2014)	Spain	N = 268	CFA	N of items	5	5	4	4	4	4	26	0.95	—	0.06	6-factor first-order
		M_{age} = 20.54		Average loadings	0.70	0.69	0.68	0.68	0.68	0.68					
				Cronbach's alpha	0.79	0.76	0.72	0.77	0.73	0.76					
Hupfeld and Ruffieux (2011)	Germany	N = 561	ESEM	N of items	5	5	4	4	4	4	26	0.98	—	0.031	6-factor first-order
		M_{age} = 26.04		Average loadings	Ranged from 0.37 to 0.94										
				Cronbach's alpha	0.83	0.74	0.75	0.80	0.66	0.70					
Kotsou and Leys (2016)[h]	France	N = 1554	CFA	N of items	5	5	4	4	4	4	26	0.92	0.91	0.047	6-factor bifactor-order
		M_{age} = 42.92		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	0.88	0.85	0.74	0.79	0.81	0.77					
Lee and Lee (2010)	Korea	N = 1554	CFA	N of items	–	–	–	–	–	–	—	—	—	—	6-factor first-order
		M_{age} = —		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	–	–	–	–	–	–					
López et al. (2015)[i]	The Netherlands	N = 1643	EFA	N of items	12			12			24	–	–	–	2-factor first-order
		M_{age} = 54.9		Average loadings	0.59			0.65							
				Cronbach's alpha	0.86			0.90							
Mantzios et al. (2015)	Greece	N = 556	EFA	N of items	5	5	4	4	4	4	26	–	–	–	6-factor first-order
		M_{age} = 24.43		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	0.70	0.77	0.72	0.71	0.72	0.76					
Neff (2003)[j]	USA	N = 391	CFA	N of items	5	5	4	4	4	4	26	0.90	0.92	–	6-factor second-order
		M_{age} = 20.91		Average loadings	0.74	0.72	0.72	0.68	0.71	0.70					
				Cronbach's alpha	0.78	0.77	0.80	0.79	0.75	0.81					

Table 1 (continued)

Authors	Nation	Sample	Analysis	Characteristics	Self-kindness	Self-judgment	Common humanity	Isolation	Mindfulness	Over-identification	Σ	CFI	TLI	RMSEA	Final model
Neff et al. (2008)	Taiwan	*N* = 164	CFA	*N* of items	–	–	–	–	–	–	24	0.87	–	0.06	6-factor first-order
		M_{age} = 20.5		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	0.68	0.68	0.69	0.74	0.65	0.70					
Neff et al. (2008)	Thailand	*N* = 223	CFA	*N* of items	–	–	–	–	–	–	25	0.88	–	0.06	6-factor first-order
		M_{age} = 19.8		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	0.72	0.70	0.60	0.74	0.66	0.75					
Petrocchi et al. (2014)[k]	Italy	*N* = 424	CFA	*N* of items	4	5	3	4	4	4	24	0.90	0.97	0.08	6-factor first-order
		M_{age} = 36.53		Average loadings	0.79	0.77	0.75	0.83	0.69	0.76					
				Cronbach's alpha	0.83	0.85	0.71	0.84	0.73	0.82					
Williams et al. (2014)[l]	UK (adult)	*N* = 821	CFA	*N* of items	5	5	4	4	4	4	26	0.895	0.877	0.071	6-factor first-order
		M_{age} = –		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	0.86	0.84	0.81	0.81	0.76	0.79					
	UK (meditator)	*N* = 211	CFA	*N* of items	5	5	4	4	4	4	26	0.804	0.772	0.100	6-factor first-order
		M_{age} = –		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	0.84	0.82	0.79	0.79	0.73	0.80					
	UK (clinical)	*N* = 390	CFA	*N* of items	5	5	4	4	4	4	26	0.852	0.828	0.077	6-factor first-order
		M_{age} = –		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	0.81	0.78	0.79	0.76	0.74	0.71					
Zeng et al. (2016)[m]	China (Buddhist)	*N* = 179	CFA	*N* of items	–	–	–	–	–	–	13 + 13	0.94	0.93	0.072	3-factor and 1-factor
		M_{age} = 35.5		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	–	–	–	–	–	–					
	China (non-Buddhist)	*N* = 232	CFA	*N* of items	–	–	–	–	–	–	13 + 13	0.96	0.94	0.074	3-factor and 1-factor
		M_{age} = 3.1		Average loadings	–	–	–	–	–	–					
				Cronbach's alpha	–	–	–	–	–	–					

M_{age} mean age; *N* number of participants; *CFA* confirmatory factor analysis; *EFA* exploratory factor analysis, *ESEM* exploratory structural equation modeling, *N of items* number of items, *Σ* total number of items in the final version of the Self-Compassion Scale, *CFI* comparative fit index, *TLI* Tucker–Lewis Index, *RMSEA* root-mean-square error of approximation

[a] Literature search was performed on July 20, 2016

[b] Benda and Reichová (2016) also identified a single higher-order factor model

[c] Castilho et al. (2015) found support for a second-order model in both the non-clinical and the clinical samples as well

[d] Although Costa et al. (2015) examined alternative structures (i.e., six-factor first-order and higher-order models), they concluded that the two-factor solution—a positive and a negative factor—is the most appropriate solution

[e] Cunha et al. (2016) investigated a first-order solution as well, which showed similar fit to the one reported in the table

[f] Apart from the first-order model, de Souza and Hutz (2016) examined the second-order and bifactor solutions as well with both demonstrating worse fit than the first-order solution

[g] Deniz et al. (2008) performed CFA on the original factor structure, resulting in unsatisfactory fit. They then performed EFA, indicating five factors on the basis of eigenvalues. However, based on the scree plot, they concluded that the scale is unidimensional

[h] Kotsou and Leys (2016) also examined a first-order and a second-order solution and while the first-order one showed better and the second-order one showed worse fit than the bifactor one, they conclude that a general self-compassion factor can be hypothesized based on the bifactor solution

[i] According to López et al. (2015), the original second-order structure could not be estimated due to identification issues. As a consequence, EFA was performed, resulting in a positive and a negative factor

[j] Although the results supported the first-order model, Neff (2003) chose the higher-order solution as the final one

[k] Petrocchi et al. (2014) reported that the second-order model had inadequate fit, thus they retained the first-order solution

[l] Upon comparing first-order and second-order solutions in different samples, Williams et al. (2014) concluded that the higher-order model might not be the best representation of this construct

[m] After having problems of misspecification with the original structure, Zeng et al. (2016) separated the positive and negative factors and identified two models: a three-factor first-order model for the positive factors and a one-factor model combining the negative factors in both of their samples. In the table above, although the fit indices were averaged based on the samples, all four showed at least acceptable goodness of fit

Portugal (Castilho et al. 2015; Cunha et al. 2016), Norway (Dundas et al. 2016), Germany (Hupfeld and Ruffieux 2011), France (Kotsou and Leys 2016), the Czech Republic (Benda and Reichová 2016), Brazil (de Souza and Hutz 2016), the Netherlands (López et al. 2015), Italy (Petrocchi et al. 2014), and the UK (Williams et al. 2014)—most of these results suggest that the higher-order model might not be the most appropriate representation of the concept of self-compassion, questioning whether a single superordinated construct can adequately explain the six facets. Alternative solutions have also been suggested with the most often replicated one being the six-factor first-order model: Japan (Arimitsu 2014), Iran (Azizi et al. 2013), Spain (Garcia-Campayo et al. 2014), Korea (Lee and Lee 2010), Greece (Mantzios et al. 2015), Taiwan (Neff et al. 2008), Thailand (Neff et al. 2008), Portugal (Bento et al. 2016), and China (Chen et al. 2011). Additionally, other solutions have also been proposed, such as a one-factor model with a general self-compassion factor (Deniz et al. 2008), a two-factor solution with one positive and one negative factor (Costa et al. 2015), and a model where the three positive intercorrelated factors are present and a general negative factor representing all negative aspects (Zeng et al. 2016).

Based on these contradictions, the dimensionality of the SCS has been called into question. In order to rectify this, Neff (2016a) agreed that a second-order solution might not be the most appropriate representation of an overall self-compassion factor. Additionally, she suggested a bifactor model approach that might be a more realistic representation of the factor structure of the SCS. Supporting evidence was reported in Neff (2016a): they examined the six-factor first-order, the second-order, and the bifactor models as well on five different samples. Overall, they identified the six-factor first-order model as the best-fitting solution, while the other two were either suboptimal or borderline acceptable. Muris et al. (2016) suggested that the six subscales should be analyzed separately without computing a general self-compassion dimension as the negative components go against the notion of measuring self-compassion as a protective mechanism. In response, Neff (2016b) claimed that the bifactor solution is an adequate representation of the underlying theory and it should be thoroughly tested. It seems to be an adequate supposition, because bifactor models can grasp a global construct (in the present case, self-compassion) as well as the smaller, more specific constructs (in the present case, the six dimensions) and all can possibly account for the commonality of the items (Brunner et al. 2012; Chen et al. 2012; Reise 2012). Indeed, several authors concluded that bifactor models might be reasonable representations of complex constructs (e.g., Chen et al. 2012; Reise 2012) and initial supporting evidence has been reported (Kotsou and Leys 2016) in relation to the SCS apart from the one reported by Neff (2016a, b). In sum, in this ongoing scientific debate about the dimensionality of the SCS, a comparison between alternative models—both first-order and bifactor ones—on a comprehensive sample can help in the understanding.

Previous validation and adaptation studies mainly employed confirmatory factor analysis (CFA) and, to a smaller extent, exploratory factor analysis (EFA). In the case of CFA, items can only load on their respective latent factors, whereas cross-loadings are forced to be zero (Marsh et al. 2009). Recently, it has been argued that CFA might be over-restrictive for multidimensional constructs (Marsh et al. 2009; Marsh et al. 2011). One possible side effect of this over-restrictive approach—apart from the unsatisfactory model fit—is that it could result in inflated factor correlations as those are the only ways for items to be expressed which in turn undermines the discriminant validity of the instrument (e.g., Asparouhov and Muthén 2009; Marsh et al. 2009; Marsh et al. 2010; Marsh et al. 2014). Indeed, interfactor correlations were high where CFAs were employed. Typically, the conceptually similar factors had high correlations with each other (e.g., self-kindness and common humanity), but not with their opposing factors (e.g., self-kindness and self-judgment). These relationships could be artificially inflated by this statistical method. On the other hand, in general, EFA demonstrates more exact, less biased parameter estimates as it allows cross-loadings (Asparouhov and Muthén 2009; Marsh et al. 2013). These arguments support the notion that EFA represents a more realistic model for multidimensional constructs. However, CFA has been considered as a superior method due to the statistical advantages associated with it, such as the inclusion of measurement error-corrected latent variables or testing of invariance across different groups (e.g., Asparouhov and Muthén 2009; Marsh et al. 2011; Marsh et al. 2014).

By incorporating aspects from both EFA and CFA, exploratory structural equation modeling (ESEM) has been recommended as a viable alternative (Marsh et al. 2014; Morin et al. 2013). It simultaneously allows a more appropriate examination of factor structure by allowing cross-loadings (EFA aspect) as well as the use of advanced statistical methods (CFA aspects). ESEM has been suggested to result in substantially better fit and less correlated factors than the corresponding CFA solutions (e.g., Arens and Morin 2016; Chiorri et al. 2016; Guay et al. 2015; Joshanloo and Lamers 2016; Marsh et al. 2011; Morin and Maïano 2011). Another advantage of ESEM is that it could easily and directly be compared to previous CFA results as well. Moreover, based on previous research on self-compassion, item cross-loadings are expected between the different factors as supported by the underlying theory; thus, the application of ESEM seems to be appropriate in this context. So far, only the study of Hupfeld and Ruffieux (2011) used this method to examine the SCS. Although their six-factor first-order model was good, the interfactor correlations were rather high (even compared to the CFA correlations), indicating potential conceptual overlap between these dimensions. New research needs to build on these prior results by (1) examining the SCS on a comprehensive sample and (2) incorporating a bifactor-ESEM framework which can be particularly useful to explore the psychometric dimensionality of an instrument (Morin et al. 2016; Morin et al. 2015). However, as noted by Marsh et al. (2009), when the ESEM solution does not fit the data better or does not result in smaller factor correlations, CFA is preferable due to its parsimony and it would be less advantage in performing an ESEM analysis.

Building on the existing literature on the Self-Compassion Scale, the present investigation examined its factor structure and construct validity by comparing both first-order and bifactor CFA and ESEM solutions. Subsequently, its measurement invariance was investigated in multiple groups in order to better understand the construct and to inspect the generalizability of the SCS.

Method

Participants

The final sample of 505 Hungarian respondents—who gave valid answers—was nationally representative among those who use Internet at least once a week in terms of gender (female = 265; 52.5%), age (M_{age} = 44.37 years; SD_{age} = 15.59 years; range 15–75 years), education (22.8% had primary level of education, 24.8% had vocational school degree, 31.5% graduated from high school, and 21.0% had higher education degree), and place of residence (18.8% in capital city, 19.6% in county capitals, 31.7% in cities, and 29.9% in villages).

Procedure

This research employed a nationally representative probability sample of Hungarians who used the Internet at least once a week. Participants were randomly selected from a pool with the help of a research market company in the summer of 2015. For the preparation of the sample, a multiple-step, proportionally stratified, probabilistic sampling method was employed, in which individuals were removed from the panel if they gave responses too quickly (i.e., without paying attention to their response) and/or had fake (unused) e-mail addresses. The study was conducted in accordance with the Declaration of Helsinki and with the approval of the Institutional Review Board of the Eötvös Loránd University. Participants were first informed about the aims and the content of the study. Then, they were assured about their anonymity and the confidentiality of their answers. They had to check a box if they were inclined to participate.

Measures

Self-Compassion Scale This questionnaire (Neff 2003) assesses six components of self-compassion: self-kindness (five items, e.g., "*When I'm going through a very hard time, I give myself the caring and tenderness I need.*"), self-judgment (five items, e.g., "*When times are really difficult, I tend to be tough on myself.*"), common humanity (four items, e.g., "*I try to see my failings as part of the human condition.*"), isolation (four items, e.g., "*When I fail at something that's important to me, I tend to feel alone in my failure.*"), mindfulness (four items, e.g., "*When something painful happens I try to take a balanced view of the situation.*"), and over-identification (4 items, e.g., "*When I'm feeling down I tend to obsess and fixate on everything that's wrong.*"). Participants can respond on a five-point scale (1 = almost never; 5 = almost always). The translation process followed the protocol of Beaton et al. (2000).

Data Analyses

All analyses were performed with Mplus 7.3 (Muthén and Muthén 1998-2015) with the weighted least squares mean- and variance-adjusted (WLSMV) estimator which was demonstrated to outperform maximum likelihood for ordered-categorical indicators with five or less answer categories (e.g., Bandalos 2014; Finney and DiStefano 2006). Negative factors were reversed prior to data analyses so that higher scores reflect higher levels of self-compassion. In the first phase of the analyses, by following previous recommendations (e.g., Morin et al. 2016), we first estimated a six-factor first-order CFA and ESEM models. In the first-order model, items were only allowed to load on their respective factors, while cross-loadings were constrained to zero. In the ESEM model, cross-loadings were allowed and target rotation was chosen (Asparouhov and Muthén 2009) due to the prior knowledge about the underlying theory. With this rotation, the main loadings were freely estimated, whereas the cross-loadings were "targeted" to be close to zero. This way, similarly to previous studies (e.g., Guay et al. 2015), ESEM was used as a confirmatory approach. These models were then contrasted with the bifactor solutions. In both bifactor models, items loaded on their respective factors and on a general self-compassion factor as well. Also, covariances between the factors were set to be zero for better interpretability. The only difference was that items were allowed to cross-load in the ESEM version, whereas they were not in the CFA version.

In the second phase of the analyses, measurement invariance was investigated across gender groups (male versus female) on the basis of the final model (Meredith 1993; Vandenberg 2002; Vandenberg and Lance 2000) to investigate the generalizability of the SCS. After establishing baseline models for the separate groups, the following sequential strategy—from the least restrictive model to the most restrictive one—was employed as outlined by previous studies (Guay et al. 2015; Meredith and Teresi 2006; Morin et al. 2013): (1) configural invariance (factor structure is the same, else is freely estimated), (2) weak invariance (factor loadings are invariant), (3) strong invariance (factor loadings and thresholds are invariant), (4) strict invariance (factor loadings, thresholds, and uniquenesses are invariant), (5) invariance of the variance–covariance matrix (factor loadings, thresholds, uniquenesses, and variance–covariances are invariant), and (6) latent mean invariance (factor loadings, thresholds, uniquenesses, variance–covariances, and latent means are invariant).

In order to assess the measurement models and the measurement invariance, commonly employed goodness of fit indices were examined. An often used index of model fit is the chi-square test (χ^2). However, it should not be considered reliable as it can be inflated by sample size and model misspecification (e.g., Brown 2015; Marsh et al. 2005). Therefore, as suggested by Brown (2015), multiple sample size-independent indices were observed with their good or acceptable cut-off values (Brown 2015; Hu and Bentler 1999): the Comparative Fit Index (CFI; ≥0.95 for good, ≥0.90 for acceptable), the Tucker–Lewis index (TLI; ≥0.95 for good, ≥0.90 for acceptable), and the Root-Mean-Square Error of Approximation (RMSEA; ≤0.06 for good, ≤0.08 for acceptable) with its 90% confidence interval. When comparing the nested models in the phase of measurement invariance, the chi-square difference test could be used via the DIFFTEST function of Mplus; however, it might be oversensitive (Marsh et al. 2005). Because of this, relative changes in fit indices were also inspected (Chen 2007; Cheung and Rensvold 2002): ΔCFI ≤0.010, ΔTLI ≤0.010, and ΔRMSEA ≤0.015. Parsimony-corrected indicators (TLI and RMSEA) might have great importance as the total number of estimated parameters is larger in ESEM than in CFA (Marsh et al. 2009; Morin et al. 2013). However, as Marsh et al. (2004) argued, these criteria should be seen as rough guidelines while also considering other important aspects such as parameter estimates or the underlying theory.

As for reliability in terms of internal consistency, Cronbach's alpha value was observed (Nunnally 1978) with the usual guidelines (0.70 is acceptable; 0.80 is good). Although this reliability index provides a comparative information in relation to previous studies, it tends to be less reliable (Rodriguez et al. 2016; Sijtsma 2009). Therefore, two additional indices were calculated that are adequate in the case of bifactor models (Brunner et al. 2012; Rodriguez et al. 2016) and can help in understanding the complex relationship patterns of the examined constructs. First is coefficient omega (ω) which is a model-based reliability index and estimates the proportion of variance that can be attributed to the blend of the global construct and the specific factors. Second is omega hierarchical (ω_H) which indicates the proportion of variance that can only be attributed to the specific factor.

Results

Model fit indices of all four models can be seen in Table 2. The six-factor CFA model showed unacceptable fit to the data (CFI = 0.845; TLI = 0.823; RMSEA = 0.101 [90% CI 0.097–0.106]). Interestingly, the bifactor CFA model demonstrated a worse fit compared to the six-factor one (CFI = 0.711; TLI = 0.656; RMSEA = 0.141 [90% CI 0.137–0.146]). Compared to these, the six-factor ESEM solution had superior fit (CFI = 0.966; TLI = 0.940; RMSEA = 0.059 [90% CI 0.053–0.065]), while the bifactor ESEM model had slightly better model fit indices than the ESEM solution (CFI = 0.972; TLI = 0.945; RMSEA = 0.057 [90% CI 0.050–0.063]). Based on these results, the bifactor ESEM solution appeared to be the most adequate model and the best representation of the data. However, it has to be noted that model selection should not only be based on model fit indices but on the examination of parameter estimates and the underlying theory as well (Marsh et al. 2004; Morin et al. 2016). First, the CFA and ESEM solutions were compared, and then the ESEM and the bifactor ESEM models as suggested by Morin et al. (2015).

CFA Versus Bifactor CFA Versus ESEM

Table 3 demonstrates the standardized parameter estimates for all models. In the six-factor CFA model, factor loadings were high ($\lambda = 0.48$ to 0.87, $M = 0.70$) and factor correlations were medium ($r = 0.07$ to 0.96, $M = 0.45$). However, the three positive ($M_r = 0.75$) and the three negative factors ($M_r = 0.79$) had high intercorrelations with the adjacent factors, while they had weaker links with the other three, resulting in an overall medium level of 0.45. Despite the high factor loadings, the model did not fit the data adequately, indicating that the hypothesized structure might not be well represented.

The fit of the bifactor CFA model was unacceptable, a result possibly attributable to the rather weakly defined general self-compassion factor ($\lambda = 0.07$ to 0.81, $M = 0.48$) and the specific factors which had acceptable factor loadings: self-kindness ($\lambda = 0.39$ to 0.79, $M = 0.54$), self-judgment ($\lambda = 0.30$ to 0.71, $M = 0.53$), common humanity ($\lambda = 0.52$ to 0.80, $M = 0.66$), isolation ($\lambda = 0.18$ to 0.64, $M = 0.42$), mindfulness ($\lambda = 0.48$ to 0.55, $M = 0.50$), and over-identification ($\lambda = 0.19$ to 0.78, $M = 0.36$).

Table 2 Goodness-of-fit statistics and information criteria for the estimated models on the Self-Compassion Scale

Model	WLSMV χ^2 (df)	CFI	TLI	RMSEA	90% CI	Comparison	$\Delta\chi^2$ (df)	ΔCFI	ΔTLI	ΔRMSEA
CFA 6-factor first-order model	1755.661* (284)	0.845	0.823	0.101	0.097–0.106	–	–	–	–	–
CFA bifactor model	3025.583* (273)	0.711	0.656	0.141	0.137–0.146	–	–	–	–	–
ESEM 6-factor first-order model	509.921* (184)	0.966	0.940	0.059	0.053–0.065	–	–	–	–	–
ESEM bifactor model	482.847* (164)	0.972	0.945	0.057	0.050–0.063	–	–	–	–	–
Gender invariance										
Baseline male	273.237* (164)	0.972	0.944	0.053	0.041–0.064	–	–	–	–	–
Baseline female	343.542* (164)	0.967	0.935	0.064	0.055–0.074	–	–	–	–	–
M1. Configural	613.888* (328)	0.970	0.940	0.059	0.052–0.066	–	–	–	–	–
M2. Weak (loadings)	835.555* (461)	0.960	0.944	0.057	0.051–0.063	M2-M1	310.170* (133)	-0.010	+0.004	−0.002
M3. Strong (loadings, thresholds)	892.150* (532)	0.962	0.953	0.052	0.046–0.058	M3-M2	107.132* (71)	+0.002	+0.009	−0.005
M4. Strict (loadings, thresholds, uniquenesses)	935.110* (558)	0.960	0.953	0.052	0.046–0.057	M4-M3	58.341* (26)	−0.002	0.000	0.000
M5. Latent variance–covariance	**982.510* (586)**	**0.958**	**0.953**	**0.052**	**0.046–0.057**	**M5-M4**	**102.600* (28)**	**−0.002**	**0.000**	**0.000**
M6. Latent means	1151.874* (593)	0.941	0.935	0.061	0.056–0.066	M6-M5	82.896* (7)	−0.017	−0.018	+0.009

Bold entries indicate the final levels of invariance that were achieved

CFA confirmatory factor analysis, *ESEM* exploratory structural equation modeling, *WLSMV* weighted least squares mean- and variance-adjusted estimator, χ^2 chi-square, *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, $\Delta\chi^2$ chi-square difference test based on the Mplus DIFFTEST function for WLSMV estimator, *ΔCFI* change in CFI value compared to the preceding model, *ΔTLI* change in the TLI value compared to the preceding model, *ΔRMSEA* change in the RMSEA value compared to the preceding model

*$p < .01$

Examining the ESEM solution revealed reduced factor correlations (r = |.02| to |.54|, M = 0.21) compared to the CFA solution. The pattern was still similar in that the positive (r = .34) and the negative factors (r = .34) had higher intercorrelation with the conceptually similar factors and weaker ones with the opposing factors, resulting in an overall coefficient of 0.21. With ESEM, the theoretically similar factors are much more differentiated. Similarly to the correlations, target factor loadings were also reduced (λ = 0.26 to .81, M = 0.53). As expected, although most of the non-target item loadings were small, some had cross-loading to a high degree (>0.32). These non-target cross-loadings could indicate conceptual overlap between the different facets of self-compassion. The decrease in target loadings and increase in non-target loadings could be attributed to the statistical procedure (ESEM) which shows that there might be a more general construct along with the six components of self-compassion. These results support the usefulness and appropriateness of both the ESEM and the bifactorial procedures.

ESEM versus Bifactor ESEM

The bifactor ESEM solution provided the best fit to the data as it was slightly better than the six-factor ESEM model (see Table 2). The examination of the parameter estimates (Table 3) revealed that a well-defined general self-compassion factor with mostly strong and significant factor loadings (λ = 0.04 to 0.78, M = 0.46). Most of the specific factors were well defined as well: self-kindness (λ = 0.41 to 0.65, M = 0.51), self-judgment (λ = 0.16 to 0.69, M = 0.46), and common humanity (λ = 0.46 to 0.86, M = 0.62). These factors retained a certain degree of specificity in the presence of the general self-compassion dimension. While the isolation factor was borderline in terms of specific factor loadings (λ = 0.17 to 0.50, M = 0.41), the mindfulness (λ = 0.11 to 0.25, M = 0.19) and over-identification (λ = 0.21 to 0.40, M = 0.32) factors were rather weakly defined, indicating that these specific items might be indicators of self-compassion rather than the specific dimensions.

Measurement Invariance

In the second phase of the analyses, measurement invariance was tested across gender groups (male vs. female) on the final bifactor ESEM solution (see Table 2). Configural models were successfully estimated in all groups, and then constraints were gradually imposed on the models. Although almost all χ^2 and $\Delta\chi^2$ test were significant, other model fit indices (ΔCFI,

Table 3 Parameter estimates for the CFA and ESEM solutions of the Self-Compassion Scale

	CFA	Bifactor CFA		ESEM						Bifactor ESEM						
	SF (λ)[a]	SC (λ)	SF (λ)[a]	SK (λ)	SJ (λ)	CH (λ)	IS (λ)	MI (λ)	OI (λ)	SC (λ)	SK (λ)	SJ (λ)	CH (λ)	IS (λ)	MI (λ)	OI (λ)
Self-kindness																
sk5	**0.72**	**0.48**	**0.39**	**0.44**	*−0.07*	0.16	0.17	0.22	*−0.02*	**0.46**	**0.41**	−0.17	0.21	*0.05*	*0.13*	*−0.08*
sk12	**.68**	**0.40**	**0.52**	**0.65**	*−0.03*	0.11	0.10	*0.01*	*−0.04*	**0.35**	**0.60**	−0.08	0.20	*0.06*	*0.02*	*−0.02*
sk19	**0.74**	**0.39**	**0.79**	**0.71**	0.20	*0.02*	−0.10	0.18	−0.09	**0.40**	**0.65**	*0.07*	0.12	−0.16	*0.06*	−0.13
sk23	**0.48**	**0.21**	**0.52**	**0.34**	0.57	0.27	−0.22	*−0.01*	−0.13	**0.18**	**0.44**	0.48	0.22	*−0.15*	*0.13*	−0.16
sk26	**0.69**	**0.44**	**0.49**	**0.50**	0.14	0.08	−0.16	0.25	0.12	**0.44**	**0.45**	*0.04*	0.15	−0.21	*0.11*	*0.00*
Self-judgment																
sj1	**0.51**	**0.21**	**0.66**	*−0.02*	**0.66**	*0.07*	−0.11	−0.20	0.22	***0.17***	*0.04*	**0.69**	*−0.08*	*0.07*	0.07	*0.15*
sj8	**0.66**	**0.42**	**0.54**	0.19	**0.65**	−0.16	0.28	*0.05*	0.16	**0.50**	*0.05*	**0.49**	−0.21	*−0.03*	*−0.22*	−0.24
sj11	**0.81**	**0.63**	**0.30**	0.20	**0.29**	−0.25	0.25	0.22	0.22	**0.68**	*−0.02*	***0.16***	−0.25	*−0.04*	*−0.19*	*0.06*
sj16	**0.54**	**0.21**	**0.71**	0.11	**0.62**	*−0.01*	−0.15	−0.30	0.33	***0.18***	0.11	**0.69**	*−0.11*	*0.05*	*−0.05*	0.28
sj21	**0.72**	**0.53**	**0.43**	0.33	**0.39**	−0.17	0.38	*−0.01*	*−0.05*	**0.56**	*0.13*	***0.27***	−0.17	*0.09*	−0.28	*−0.08*
Common humanity																
ch3	**0.70**	**0.40**	**0.53**	−0.09	*−0.03*	**0.57**	0.19	0.30	*−0.02*	**0.38**	*0.05*	*−0.14*	**0.47**	*0.09*	0.37	−0.16
ch7	**0.64**	***0.07***	**0.80**	0.15	−0.17	**0.80**	0.12	−0.32	*−0.07*	***0.04***	0.24	−0.14	**0.86**	*0.05*	−0.20	*0.05*
ch10	**0.70**	**0.14**	**0.78**	0.17	−0.11	**0.69**	*−0.07*	−0.10	*0.03*	***0.12***	0.26	−0.13	**0.68**	−0.15	*−0.01*	*0.03*
ch15	**0.84**	**0.54**	**0.52**	*−0.05*	0.11	**0.57**	0.10	0.48	*−0.04*	**0.52**	*0.09*	*−0.09*	**0.46**	*−0.11*	*0.34*	−0.24
Isolation																
is4	**0.87**	**0.81**	**0.18**	*−0.00*	0.17	*0.03*	**0.36**	0.20	0.45	**0.78**	−0.13	*0.11*	*−0.05*	**0.17**	*0.03*	0.22
is13	**0.84**	**0.70**	**0.50**	−0.08	0.18	0.13	**0.81**	*0.02*	*0.04*	**0.72**	−0.17	*0.09*	*−0.04*	**0.49**	*−0.03*	*−0.04*
is18	**0.76**	**0.60**	**0.64**	−0.09	*0.06*	0.09	**0.74**	*−0.01*	0.12	**0.62**	−0.17	*0.01*	*−0.06*	**0.50**	*−0.01*	*0.05*
is25	**0.77**	**0.67**	**0.34**	0.11	*−0.00*	*0.04*	**0.55**	−0.22	0.41	**0.58**	*−0.04*	*0.06*	*−0.03*	**0.46**	*−0.15*	0.36
Mindfulness																
mi9	**0.65**	**0.46**	**0.49**	0.22	−0.25	0.13	*0.06*	**0.41**	0.24	**0.49**	0.19	−0.34	0.18	*−0.05*	***0.21***	*0.06*
mi14	**0.70**	**0.47**	**0.55**	0.22	−0.08	0.21	*−0.02*	**0.47**	0.11	**0.50**	0.23	−0.23	0.23	*−0.16*	***0.24***	−0.08
mi17	**0.63**	**0.40**	**0.48**	0.24	*−0.06*	0.22	*−0.03*	**0.37**	0.09	**0.45**	0.21	−0.20	0.26	−0.23	***0.11***	*−0.07*
mi22	**0.71**	**0.47**	**0.48**	0.51	−0.19	0.12	*0.02*	**0.26**	0.11	**0.42**	0.48	−0.26	0.20	−0.01	***0.20***	*0.02*
Over-identification																
oi2	**0.84**	**0.80**	**0.19**	*−0.03*	0.27	0.07	0.33	0.10	**0.45**	**0.71**	−0.11	0.24	−0.06	0.22	*0.08*	**0.25**
oi6	**0.82**	**0.79**	**0.16**	*−0.02*	0.27	*0.02*	0.21	0.26	**0.48**	**0.78**	−0.17	*0.19*	*−0.04*	*−0.01*	*−0.01*	**0.21**
oi20	**0.57**	**0.44**	**0.78**	−0.12	0.11	−0.08	0.09	*0.04*	**0.59**	**0.44**	−0.21	0.17	−0.15	*0.12*	*0.01*	**0.40**
oi24	**0.71**	**0.64**	**0.32**	*0.04*	*0.05*	−0.09	0.34	*−0.01*	**0.54**	**0.58**	−0.13	*−0.01*	−0.12	0.27	*−0.08*	**0.40**

Target factor loadings are in bold. Non-significant parameters ($p \geq .05$) are italicized

CFA confirmatory factor analysis, *ESEM* exploratory structural equation modeling, *SC* general self-compassion factor, *SK* self-kindness, *SJ* self-judgment, *CH* common humanity, *IS* isolation, *MI* mindfulness, *OI* over-identification, λ standardized factor loadings, *SF* specific factors of the Self-Compassion Scale

[a] Each item loaded on their respective specific factor, while cross-loadings were constrained to zero

ΔTLI, ΔRMSEA) did not decrease more than the recommended cut-off values, indicating gender invariance on the level of latent variance–covariance matrix. However, latent mean invariance could not be achieved in either group. The inspection of the latent means revealed that—when the latent means of the males were set to be zero for the purpose of identification—females' latent means were significantly lower on self-judgment ($M = -0.38$, $p < .01$) dimension while having higher scores on self-kindness ($M = +0.65$, $p < .001$). The general self-compassion means were higher among males than females ($M = -0.55$, $p < .001$).

Reliability Indices

In the last phase, the reliability of the questionnaire was assessed (see Table 4). Regarding internal consistency, Cronbach's alpha values were satisfactory for all scales ($\alpha = 0.73$ to 0.88). However, as it can be less reliable

Table 4 Reliability indices and descriptive statistics of the Self-Compassion Scale

Scales	*N* of items	α	ω	ω_H	Range	M	SD
0. Self-compassion	26	0.88	0.96	0.53	1–5	3.08	0.46
1. Self-kindness	5	0.76	0.81	0.53	1–5	2.88	0.78
2. Self-judgment	5	0.74	0.82	0.45	1–5	3.23	0.74
3. Common humanity	4	0.74	0.85	0.72	1–5	3.12	0.84
4. Isolation	4	0.85	0.90	0.24	1–5	2.94	1.03
5. Mindfulness	4	0.73	0.66	0.09	1–5	3.29	0.79
6. Over-identification	4	0.80	0.84	0.17	1–5	3.06	0.94

N of items number of items on the factors, α Cronbach's alpha, ω coefficient omega, ω_H omega hierarchical, *M* mean, *SD* standard deviation

(Rodriguez et al. 2016; Sijtsma 2009), other indices were also considered that are especially useful in the case of bifactor models. Coefficient omega levels were generally high (ω = 0.66 to 0.96, M = 0.83), indicating that the scale scores adequately represent the blend of self-compassion and the six specific components. In contrast, omega hierarchical values had high variance (ω_H = 0.09 to 0.72, M = 0.83). Although there are no specific cut-off values for omega coefficients, over-identification (ω_H = 0.17) and isolation (ω_H = 0.24) had relatively small omega hierarchical values, while mindfulness had the lowest value (ω_H = 0.09). Regarding these three factors—especially in the case of mindfulness—only a small proportion of variance can be attributed to the specific factor, while a larger proportion can be explained by the general self-compassion factor.

Finally, as suggested by Rodriguez et al. (2016), we compared the omega (0.96) and the omegaH (0.53) of the general self-compassion factor. The ratio of the two indices (0.53/0.96 = 0.55) revealed that approximately half of the variance in the total scores can be attributed to the general self-compassion dimension, while about 43% (0.96–0.53) of the variance in the total score can be attributed to the multidimensionality caused by the specific factors.

Discussion

The present study applied ESEM, an advanced statistical method, to investigate the construct validity of the SCS—the most widely used instrument to measure self-compassion—in an attempt to contribute to the ongoing discussion about its dimensionality. The bifactor ESEM framework proved to be superior and useful in uncovering the underlying dimensionality of the SCS and demonstrated a more realistic representation of the relationship between the items and the factors. Both the general self-compassion dimension and the six components were confirmed; however, not all can be considered fully reliable. Also, this study contributes to the steadily increasing literature on ESEM (e.g., Caci et al. 2015; Joshanloo and Lamers 2016; Maïano et al. 2013).

As self-compassion is an important research area, it has also become increasingly important to measure this multifaceted construct with an appropriate instrument, the Self-Compassion Scale. Although the psychometric properties of the questionnaire have been investigated in many countries and samples (for an overview, see Table 1), there is still no accordance regarding its factor structure (e.g., Muris et al. 2016) which was mainly investigated with confirmatory factor analysis. Even though CFA and bifactor CFA models had well-defined factors in the current research, these solutions still had unsatisfactory model fit. These results are in line with previous CFA studies detailed in Table 1 (e.g., Arimitsu 2014; Costa et al. 2015; Williams et al. 2014) which would suggest that CFA might be too restrictive for the SCS. The application of ESEM—similar to the study of Hupfeld and Ruffieux (2011)—resulted in better fit indices and smaller factor correlations as it has been hypothesized based on previous studies (e.g., Arens and Morin 2016; Guay et al. 2015; Morin et al. 2015). It also confirmed that while items of the scale formed separate factors, they still exhibited cross-loadings, suggesting a certain degree of conceptual overlap between the factors. These non-target loadings could also indicate the presence of a global construct in the data.

By employing ESEM with a bifactorial framework, important information has been discovered about the scale. The six specific factors and the general self-compassion factor had been partially confirmed. Goodness-of-fit indices showed that the bifactor solution was the most adequate (as it has been recently suggested by Neff 2016a, 2016b). However, in the presence of an overarching construct, some of the small ones became less well-defined. Morin et al. (2015) argued that as long as the global factor is well-defined, not all specific factors are required to be equally well-defined. The present results clearly support this argument with the global factor and some facets of self-compassion being well-defined. The results further confirm the partial reliability of the scale as not all indices were satisfactory. More precisely, mindfulness was not salient when accounting for the global factor, suggesting that it does not contribute to the theory over and above the general factor. In bifactor models, it is possible that specific factor loadings become non-significant, resulting in less-defined factors which could indicate that the general factor can explain a larger proportion of the variance (Reise et al. 2007). However, in the case of mindfulness, it cannot be easily distinguished from self-compassion. Prior research argued that mindfulness is related to the internal experience as feelings, emotions, or thoughts, whereas self-compassion focuses more on the experiencer than the experience itself (Germer 2009; Neff and Germer 2013). Following this notion, if the wording of the mindfulness items focused more on the experience, while the wording of the other self-compassion items focused on the experiencer, it could be expected that omega hierarchicals of mindfulness and its factor loadings could also increase. This step could also help in the differentiation between self-compassion and mindfulness.

High levels of measurement invariance were also supported across gender groups (i.e., invariance on the level of latent variance–covariance matrix). The factor structure could be generalized across subgroups of participants, which allows the comparison of the groups. Additionally, gender-based mean differences were also discovered. These results are consistent with previous studies in that females had lower levels of overall self-compassion than did males (e.g., Raes 2010; Yarnell et al. 2015). However, they also scored lower on self-judgment and higher on self-kindness, indicating that gender differences should also be investigated on the level of the six factors, not just the overall self-compassion dimension.

Although the current study has many strengths (such as the diverse and online representative sample and the exhaustive statistical analyses), several limitations need to be acknowledged. First, the research was cross-sectional and questionnaire-based, implying possible biases. A longitudinal research could be fruitful in uncovering potential effects of life events (e.g., failure or criticism) that can influence one's self-compassion. Second, the present results need to be replicated in other countries in order to draw a more solid conclusion about the nature of self-compassion. Temporal stability could also be investigated, possibly with the inclusion of non-clinical and clinical samples as well, to better understand the psychometric properties of the instrument. It would also be useful in future studies to examine its convergent and discriminant validity.

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Compliance with Ethical Standards

Ethics Statement The study was conducted in accordance with the Declaration of Helsinki and with the approval of the Institutional Review Board of the Eötvös Loránd University.

Conflict of Interest The authors declare that they have no conflict of interest.

References

Arens, A. K., & Morin, A. J. (2016). Examination of the structure and grade-related differentiation of multidimensional self-concept instruments for children using ESEM. *The Journal of Experimental Education, 84*(2), 330–355. doi:10.1080/00220973.2014.999187.

Arimitsu, K. (2014). Development and validation of the Japanese version of the self-compassion scale. *The Japanese Journal of Psychology, 85*(1), 50–59.

Asparouhov, T., & Muthén, B. (2009). Exploratory structural equation modeling. *Structural Equation Modeling: A Multidisciplinary Journal, 16*(3), 397–438. doi:10.1080/10705510903008204.

Azizi, A., Mohammadkhani, P., Lotfi, S., & Bahramkhani, M. (2013). The validity and reliability of the Iranian version of the self-compassion scale. *Iranian Journal of Clinical Psychology, 2*(3), 17–23.

Bandalos, D. L. (2014). Relative performance of categorical diagonally weighted least squares and robust maximum likelihood estimation. *Structural Equation Modeling: A Multidisciplinary Journal, 21*(1), 102–116. doi:10.1080/10705511.2014.859510.

Beaton, D. E., Bombardier, C., Guillemin, F., & Ferraz, M. B. (2000). Guidelines for the process of cross-cultural adaptation of self-report measures. *Spine, 25*(24), 3186–3191.

Benda, J., & Reichová, A. (2016). [Psychometric characteritics of the Czech version of the self-compassion scale] Psychometrické charakteristiky české verze self-compassion scale (SCS-CZ). *Československá psychologie., 60*(2), 20–36.

Bento, E., Xavier, S., Azevedo, J., Marques, M., Freitas, V., Soares, M. J., et al. (2016). Validation of the self-compassion scale in a community sample of Portuguese pregnant women. *European Psychiatry, 33*, S238. doi:10.1016/j.eurpsy.2016.01.598.

Brown, T. A. (2015). *Confirmatory factor analysis for applied research* (2nd ed.). New York: Guilford Press.

Brunner, M., Nagy, G., & Wilhelm, O. (2012). A tutorial on hierarchically structured constructs. *Journal of Personality, 80*(4), 796–846. doi:10.1111/j.1467-6494.2011.00749.x.

Caci, H., Morin, A. J., & Tran, A. (2015). Investigation of a bifactor model of the strengths and difficulties questionnaire. *European Child & Adolescent Psychiatry, 24*(10), 1291–1301. doi:10.1007/s00787-015-0679-3.

Castilho, P., Pinto-Gouveia, J., & Duarte, J. (2015). Evaluating the multifactor structure of the long and short versions of the self-compassion scale in a clinical sample. *Journal of Clinical Psychology, 71*(9), 856–870. doi:10.1002/jclp.22187.

Chen, F. F. (2007). Sensitivity of goodness of fit indexes to lack of measurement invariance. *Structural Equation Modeling, 14*(3), 464–504. doi:10.1080/10705510701301834.

Chen, J., Yan, L., & Zhou, L. (2011). Reliability and validity of Chinese version of self-compassion scale. *Chinese Journal of Clinical Psychology, 19*(6), 734–736.

Chen, F. F., Hayes, A., Carver, C. S., Laurenceau, J. P., & Zhang, Z. (2012). Modeling general and specific variance in multifaceted constructs: a comparison of the bifactor model to other approaches. *Journal of Personality, 80*(1), 219–251. doi:10.1111/j.1467-6494.2011.00739.x.

Cheung, G. W., & Rensvold, R. B. (2002). Evaluating goodness-of-fit indexes for testing measurement invariance. *Structural Equation Modeling, 9*(2), 233–255. doi:10.1207/S15328007SEM0902_5.

Chiorri, C., Marsh, H. W., Ubbiali, A., & Donati, D. (2016). Testing the factor structure and measurement invariance across gender of the big five inventory through exploratory structural equation modeling. *Journal of Personality Assessment, 98*(1), 88–99. doi:10.1080/00223891.2015.1035381.

Costa, J., Marôco, J., Pinto-Gouveia, J., Ferreira, C., & Castilho, P. (2015). Validation of the psychometric properties of the self-compassion scale. Testing the factorial validity and factorial invariance of the measure among borderline personality disorder, anxiety disorder, eating disorder and general populations. *Clinical Psychology & Psychotherapy*. doi:10.1002/cpp.1974.

Cunha, M., Xavier, A., & Castilho, P. (2016). Understanding self-compassion in adolescents: validation study of the self-compassion scale. *Personality and Individual Differences, 93*, 56–62. doi:10.1016/j.paid.2015.09.023.

de Souza, L. K., & Hutz, C. S. (2016). Adaptation of the self-compassion scale for use in Brazil: evidences of construct validity. *Trends in Psychology, 24*(1), 159–172. doi:10.9788/TP2016.1-11.

Deniz, M., Kesici, Ş., & Sümer, A. S. (2008). The validity and reliability of the Turkish version of the self-compassion scale. *Social Behavior and Personality, 36*(9), 1151–1160. doi:10.2224/sbp.2008.36.9.1151.

Dundas, I., Svendsen, J. L., Wiker, A. S., Granli, K. V., & Schanche, E. (2016). Self-compassion and depressive symptoms in a Norwegian student sample. *Nordic Psychology, 68*(1), 58–72. doi:10.1080/19012276.2015.1071203.

Ferreira, C., Pinto-Gouveia, J., & Duarte, C. (2013). Self-compassion in the face of shame and body image dissatisfaction: implications for eating disorders. *Eating Behaviors, 14*(2), 207–210. doi:10.1016/j.eatbeh.2013.01.005.

Finney, S. J., & DiStefano, C. (2006). Non-normal and categorical data in structural equation modeling. In G. R. Hancock & R. D. Mueller (Eds.), *Structural equation modeling: a second course* (pp. 269–314). Charlotte: Information Age Publishing.

Friis, A. M., Johnson, M. H., Cutfield, R. G., & Consedine, N. S. (2016). Kindness matters: a randomized controlled trial of a mindful self-compassion intervention improves depression, distress, and HbA1c among patients with diabetes. *Diabetes Care, 39*(11), 1963–1971. doi:10.2337/dc16-0416.

Garcia-Campayo, J., Navarro-Gil, M., Andrés, E., Montero-Marin, J., López-Artal, L., & Demarzo, M. M. P. (2014). Validation of the Spanish versions of the long (26 items) and short (12 items) forms of the self-compassion scale (SCS). *Health and Quality of Life Outcomes, 12*(4), 1–9. doi:10.1186/1477-7525-12-4.

Germer, C. K. (2009). *The mindful path to self-compassion: freeing yourself from destructive thoughts and emotions*. New York: Guilford Press.

Guay, F., Morin, A. J. S., Litalien, D., Valois, P., & Vallerand, R. J. (2015). Application of exploratory structural equation modeling to evaluate the academic motivation scale. *The Journal of Experimental Education, 83*(1), 51–82. doi:10.1080/00220973.2013.876231.

Hu, L., & Bentler, P. M. (1999). Cutoff criteria for fit indexes in covariance structure analysis: conventional criteria versus new alternatives. *Structural Equation Modeling, 6*(1), 1–55. doi:10.1080/10705519909540118.

Hupfeld, J., & Ruffieux, N. (2011). Validierung einer deutschen version der self-compassion scale (SCS-D) [validation of a German version of the self-compassion scale (SCS-D)]. *Zeitschrift für Klinische Psychologie und Psychotherapie, 40*(2), 115–123. doi:10.1026/1616-3443/a000088.

Joshanloo, M., & Lamers, S. M. (2016). Reinvestigation of the factor structure of the MHC-SF in the Netherlands: contributions of exploratory structural equation modeling. *Personality and Individual Differences, 97*, 8–12. doi:10.1016/j.paid.2016.02.089.

Kotsou, I., & Leys, C. (2016). Self-compassion scale (SCS): psychometric properties of the French translation and its relations with psychological well-being, affect and depression. *PloS One, 11*(4), e0152880. doi:10.1371/journal.pone.0152880.

Krieger, T., Altenstein, D., Baettig, I., Doerig, N., & Holtforth, M. G. (2013). Self-compassion in depression: associations with depressive symptoms, rumination, and avoidance in depressed outpatients. *Behavior Therapy, 44*(3), 501–513. doi:10.1016/j.beth.2013.04.004.

Leary, M. R., Tate, E. B., Adams, C. E., Allen, A. B., & Hancock, J. (2007). Self-compassion and reactions to unpleasant self-relevant events: the implications of treating oneself kindly. *Journal of Personality and Social Psychology, 92*(5), 887–904. doi:10.1037/0022-3514.92.5.887.

Lee, W. K., & Lee, K. (2010). The validation study of the Korean version of the self-compassion scale with adult women in the community. *Journal of the Korean Neuropsychiatric Association, 49*, 193–200.

López, A., Sanderman, R., Smink, A., Zhang, Y., van Sonderen, E., Ranchor, A., et al. (2015). A reconsideration of the self-compassion scale's total score: self-compassion versus self-criticism. *PloS One, 10*(7), e0132940. doi:10.1371/journal.pone.0132940.

MacBeth, A., & Gumley, A. (2012). Exploring compassion: a meta-analysis of the association between self-compassion and psychopathology. *Clinical Psychology Review, 32*(6), 545–552. doi:10.1016/j.cpr.2012.06.003.

Maïano, C., Morin, A. J., Lanfranchi, M. C., & Therme, P. (2013). The eating attitudes test-26 revisited using exploratory structural equation modeling. *Journal of Abnormal Child Psychology, 41*(5), 775–788. doi:10.1007/s10802-013-9718-z.

Mantzios, M., Wilson, J. C., & Giannou, K. (2015). Psychometric properties of the Greek versions of the self-compassion and mindful attention and awareness scales. *Mindfulness, 6*(1), 123–132. doi:10.1007/s12671-013-0237-3.

Marsh, H. W., Hau, K. T., & Wen, Z. (2004). In search of golden rules: comment on hypothesis-testing approaches to setting cutoff values for fit indexes and dangers in overgeneralizing Hu and Bentler's (1999) findings. *Structural Equation Modeling, 11*(3), 320–341. doi:10.1207/s15328007sem1103_2.

Marsh, H. W., Hau, K.-T., & Grayson, D. (2005). Goodness of fit evaluation. In A. Maydeu-Olivares & J. McArdle (Eds.), *Contemporary psychometrics* (pp. 275–340). Mahwah NJ: Erlbaum.

Marsh, H. W., Muthén, B., Asparouhov, T., Lüdtke, O., Robitzsch, A., Morin, A. J. S., & Trautwein, U. (2009). Exploratory structural equation modeling, integrating CFA and EFA: application to students' evaluations of university teaching. *Structural Equation Modeling: A Multidisciplinary Journal, 16*(3), 439–476. doi:10.1080/10705510903008220.

Marsh, H. W., Lüdtke, O., Muthén, B., Asparouhov, T., Morin, A. J. S., Trautwein, U., & Nagengast, B. (2010). A new look at the big five factor structure through exploratory structural equation modeling. *Psychological Assessment, 22*(3), 471–491. doi:10.1037/a0019227.

Marsh, H. W., Liem, G. A. D., Martin, A. J., Morin, A. J. S., & Nagengast, B. (2011). Methodological measurement fruitfulness of exploratory structural equation modeling (ESEM): new approaches to key substantive issues in motivation and engagement. *Journal of Psychoeducational Assessment, 29*(4), 322–346. doi:10.1177/0734282911406657.

Marsh, H. W., Nagengast, B., & Morin, A. J. S. (2013). Measurement invariance of big-five factors over the life span: ESEM tests of gender, age, plasticity, maturity, and la dolce vita effects. *Developmental Psychology, 49*(6), 1194–1218. doi:10.1037/a0026913.

Marsh, H. W., Morin, A. J. S., Parker, P. D., & Kaur, G. (2014). Exploratory structural equation modeling: an integration of the best features of exploratory and confirmatory factor analysis. *Annual Review of Clinical Psychology, 10*, 85–110. doi:10.1146/annurev-clinpsy-032813-153700.

Meredith, W. (1993). Measurement invariance, factor analysis and factorial invariance. *Psychometrika, 58*(4), 525–543. doi:10.1007/BF02294825.

Meredith, W., & Teresi, J. A. (2006). An essay on measurement and factorial invariance. *Medical Care, 44*(11), S69–S77. doi:10.1097/01.mlr.0000245438.73837.89.

Morin, A. J. S., & Maïano, C. (2011). Cross-validation of the short form of the physical self-inventory (PSI-S) using exploratory structural equation modeling (ESEM). *Psychology of Sport and Exercise, 12*(5), 540–554. doi:10.1016/j.psychsport.2011.04.003.

Morin, A. J. S., Marsh, H. W., & Nagengast, B. (2013). Exploratory structural equation modeling. In G. R. Hancock & R. O. Mueller (Eds.), *Structural equation modeling: a second course* (pp. 395–436). Charlotte: Information Age Publishing, Inc..

Morin, A. J. S., Arens, A. K., Tran, A., & Caci, H. (2015). Exploring sources of construct-relevant multidimensionality in psychiatric measurement: a tutorial and illustration using the composite scale of morningness. *International Journal of Methods in Psychiatric Research.* doi:10.1002/mpr.1485.

Morin, A. J. S., Arens, A. K., & Marsh, H. W. (2016). A bifactor exploratory structural equation modeling framework for the identification of distinct sources of construct-relevant psychometric multidimensionality. *Structural Equation Modeling: A Multidisciplinary Journal, 23*(1), 116–139. doi:10.1080/10705511.2014.961800.

Muris, P., Otgaar, H., & Petrocchi, N. (2016). Protection as the mirror image of psychopathology: further critical notes on the self-compassion scale. *Mindfulness, 7*(3), 787–790. doi:10.1007/s12671-016-0509-9.

Muthén, L. K., & Muthén, B. O. (1998-2015). *Mplus user's guide* (7th ed.). Los Angeles: Muthén & Muthén.

Neff, K. D. (2003). The development and validation of a scale to measure self-compassion. *Self and Identity, 2*(3), 223–250. doi:10.1080/15298860309027.

Neff, K. D. (2011). Self-compassion, self-esteem, and well-being. *Social and Personality Psychology Compass, 5*(1), 1–12. doi:10.1111/j.1751-9004.2010.00330.x.

Neff, K. D. (2016a). The self-compassion scale is a valid and theoretically coherent measure of self-compassion. *Mindfulness, 7*(1), 264–274. doi:10.1007/s12671-015-0479-3.

Neff, K. D. (2016b). Does self-compassion entail reduced self-judgment, isolation, and over-identification? A response to Muris, Otgaar, and Petrocchi (2016). *Mindfulness, 7*(3), 791–797. doi:10.1007/s12671-016-0531-y.

Neff, K. D., & Germer, C. K. (2013). A pilot study and randomized controlled trial of the mindful self-compassion program. *Journal of Clinical Psychology, 69*(1), 28–44. doi:10.1002/jclp.21923.

Neff, K. D., Pisitsungkagarn, K., & Hsieh, Y. P. (2008). Self-compassion and self-construal in the United States, Thailand, and Taiwan. *Journal of Cross-Cultural Psychology, 39*(3), 267–285. doi:10.1177/0022022108314544.

Nunnally, J. C. (1978). *Psychometric theory* (2nd ed.). New York: McGraw-Hill.

Petrocchi, N., Ottaviani, C., & Couyoumdjian, A. (2014). Dimensionality of self-compassion: translation and construct validation of the self-compassion scale in an Italian sample. *Journal of Mental Health, 23*(2), 72–77. doi:10.3109/09638237.2013.841869.

Raes, F. (2010). Rumination and worry as mediators of the relationship between self-compassion and depression and anxiety. *Personality and Individual Differences, 48*(6), 757–761. doi:10.1016/j.paid.2010.01.023.

Reise, S. P. (2012). The rediscovery of bifactor measurement models. *Multivariate Behavioral Research, 47*(5), 667–696. doi:10.1080/00273171.2012.715555.

Reise, S. P., Morizot, J., & Hays, R. D. (2007). The role of the bifactor model in resolving dimensionality issues in health outcomes measures. *Quality of Life Research, 16*(1), 19–31. doi:10.1007/s11136-007-9183-7.

Rodriguez, A., Reise, S. P., & Haviland, M. G. (2016). Evaluating bifactor models: calculating and interpreting statistical indices. *Psychological Methods, 21*(2), 137–150. doi:10.1037/met0000045.

Sijtsma, K. (2009). On the use, the misuse, and the very limited usefulness of Cronbach's alpha. *Psychometrika, 74*(1), 107–120. doi:10.1007/s11336-008-9101-0.

Vandenberg, R. J. (2002). Toward a further understanding of and improvement in measurement invariance methods and procedures. *Organizational Research Methods, 5*(2), 139–158. doi:10.1177/1094428102005002001.

Vandenberg, R. J., & Lance, C. E. (2000). A review and synthesis of the measurement invariance literature: suggestions, practices, and recommendations for organizational research. *Organizational Research Methods, 3*(1), 4–70. doi:10.1177/109442810031002.

Williams, M. J., Dalgleish, T., Karl, A., & Kuyken, W. (2014). Examining the factor structures of the five facet mindfulness questionnaire and the self-compassion scale. *Psychological Assessment, 26*(2), 407–418. doi:10.1037/a0035566.

Yarnell, L. M., Stafford, R. E., Neff, K. D., Reilly, E. D., Knox, M. C., & Mullarkey, M. (2015). Meta-analysis of gender differences in self-compassion. *Self and Identity, 14*(5), 499–520. doi:10.1080/15298868.2015.1029966.

Zeng, X., Wei, J., Oei, T. P., & Liu, X. (2016). The self-compassion scale is not validated in a Buddhist sample. *Journal of Religion and Health, 55*(1996), 1–14. doi:10.1007/s10943-016-0205-z.

Zessin, U., Dickhäuser, O., & Garbade, S. (2015). The relationship between self-compassion and well-being: a meta-analysis. *Applied Psychology: Health and Well-Being, 7*(3), 340–364. doi:10.1111/aphw.12051.

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ERRATUM

Erratum to: Exploratory Structural Equation Modeling Analysis of the Self-Compassion Scale

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The authors would like to acknowledge an error in the calculation of the omega and omega hierarchical values for the general self-compassion factor in the above-mentioned paper. These values for the six subscales are correct. The error is on page 9 and the corresponding Table 4 (see the updated version on the right).

The correct omega and omega hierarchical values for the general self-compassion factor would be 0.94 and 0.81, respectively (instead of 0.96 and 0.53). As an additional correction, the ratio of the two indices (0.81/0.94 = 0.86) revealed that 86% of the variance in the total score can be attributed to the general self-compassion factor, while 13% (0.94-0.81) of the variance in the total score can be attributed to the multidimensionality caused by the specific factors.

Table 4 Reliability indices and descriptive statistics of the Self-Compassion Scale

Scales	*N* of items	α	ω	ω_H	Range	M	SD
0. Self-compassion	26	.88	.94	.81	1-5	3.08	0.46
1. Self-kindness	5	.76	.81	.53	1-5	2.88	0.78
2. Self-judgment	5	.74	.82	.45	1-5	3.23	0.74
3. Common humanity	4	.74	.85	.72	1-5	3.12	0.84
4. Isolation	4	.85	.90	.24	1-5	2.94	1.03
5. Mindfulness	4	.73	.66	.09	1-5	3.29	0.79
6. Over-identification	4	.80	.84	.17	1-5	3.06	0.94

N of items number of items on the factors, α Cronbach's alpha, ω coefficient omega, ω_H omega hierarchical, *M* mean, *SD* standard deviation

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