Article:

Neff, K., Tóth-Király, I., Yarnell, L. et al. (15 more authors) (2019) Examining the Self-Compassion Scale in 20 diverse samples: Support for use of a total score and six subscale scores. Psychological Assessment, 31 (1). pp. 27-45. ISSN 1040-3590

https://doi.org/10.1037/pas0000629

Running Head: SCS FACTOR STRUCTURE IN 20 DIVERSE SAMPLES

Examining the Factor Structure of the Self-Compassion Scale in 20 Diverse Samples: Support for Use of a Total Score and Six Subscale Scores

Kristin D. Neff,[1] István Tóth-Király,[2] Lisa M. Yarnell,[3] Kohki Arimitsu,[4] Paula Castilho,[5] Nima Ghorbani,[6] Hailan Xiaoxia Guo,[7] Jameson K. Hirsch,[8] Jörg Hupfeld,[9] Claudio S. Hutz,[10] Ilios Kotsou,[11] Woo Kyeong Lee,[12] Jesus Montero-Marin,[13] Fuschia M. Sirois,[14] Luciana K. de Souza,[10] Julie L. Svendsen,[15] Ross B. Wilkinson,[16] Michail Mantzios[17]

In press, *Psychological Assessment*

(Uncorrected version)

[1]Department of Educational Psychology, University of Texas at Austin, United States; [2]Doctoral School of Psychology, Department of Personality and Health Psychology, ELTE Eötvös Loránd University, Hungary; [3]American Institutes for Research, United States; [4]Department of Psychological Sciences, Kwansei Gakuin University, Japan; [5]Faculty of Psychology and Educational Sciences, University of Coimbra, Portugal; [6]Department of Psychology, University of Tehran, Iran; [7]Beijing Hailan Peer Education & Consultation Co., China; [8]Department of Psychology, East Tennessee State University, United States; [9]Department of Psychology, University of Bern, Switzerland; [10]Post-Graduate Program in Psychology, Universidade Federal do Rio Grande do Sul, Brazil; [11]Department of Psychology, Université Libre de Bruxelles, Belgium; [12]Department of Counseling Psychology, Seoul Cyber University, Korea; [13] Primary Care Prevention and Health Promotion Research Network (RedIAPP), Zaragoza, Spain;

[14]Department of Psychology, University of Sheffield, United Kingdom; [15]Department of Biological and Medical Psychology, University of Bergen, Norway; [16]School of Psychology, University of Newcastle, Australia; [17]Department of Psychology, Birmingham City University, United Kingdom

Correspondence concerning this article should be addressed to Kristin D. Neff, Department of Educational Psychology, University of Texas at Austin, 1912 Speedway, Stop D5800, Austin, TX, USA, 78712-1289. E-mail: kristin.neff@mail.utexas.edu

Author note: The clinical sample from the United Kingdom used in this study was drawn from the PREVENT trial, a project funded by the National Institute for Health Research Health Technology Assessment (NIHR HTA) Programme (project number 08/56/01). This trial is reported in full in the Lancet, DOI: http://dx.doi.org/10.1016/S0140-6736(14)62222-4. We are grateful to the trial team for allowing us to use the data.

Author contribution statement: KN conceived the paper and wrote the first draft. ITK conducted all statistical analyses. LY coordinated data collection and contributed her statistical expertise. The first three authors did the bulk of the writing. MM, the last author, coordinated putting together the translation information table in the supplementary materials with input from other authors. All other authors are listed in alphabetical order and contributed data as well as making comments on earlier drafts of the manuscript.

Public Significance Statement: This study examined the factor structure of the SCS in 20 diverse samples (N = 11,685), and excellent fit was found in every sample for an ESEM single-

bifactor model (with 95% of item variance explained by a general factor) and an ESEM six-factor correlated model. Results support use of a total SCS score or six subscale scores, but not two separate scores representing compassionate and uncompassionate self-responding.

Abstract

This study examined the factor structure of the Self-Compassion Scale (SCS) using secondary data drawn from 20 samples (*N* = 11,685) — 7 English and 13 non-English — including 10 community, 6 student, 1 mixed community/student, 1 meditator, and 2 clinical samples. Self-compassion is theorized to represent a system with six constituent components - self-kindness, common humanity, mindfulness and reduced self-judgment, isolation and over-identification. There has been controversy as to whether a total score on the SCS or if separate scores representing compassionate versus uncompassionate self-responding should be used. The current study examined the factor structure of the SCS using confirmatory factor analyses (CFA) and exploratory structural equation modeling (ESEM) to examine five distinct models: one-factor, two-factor correlated, six-factor correlated, single-bifactor (one general self-compassion factor and six group factors), and two-bifactor models (two correlated general factors each with three group factors representing compassionate or uncompassionate self-responding). Results indicated that a one- and two-factor solution to the SCS had inadequate fit in every sample examined using both CFA and ESEM, whereas fit was excellent using ESEM for the six-factor correlated, single-bifactor and correlated two-bifactor models. However, factor loadings for the correlated two-bifactor models indicated that two separate factors were not well specified. A general factor explained 95% of the reliable item variance in the single-bifactor model. Results support use of the SCS to examine six subscale scores (representing the constituent components of self-compassion) or a total score (representing overall self-compassion), but not separate scores representing compassionate and uncompassionate self-responding.

KEYWORDS: Self-Compassion Scale, self-compassion factor structure, bifactor analyses, Exploratory Structural Equation Modeling (ESEM)

Examining the Factor Structure of the Self-Compassion Scale in 20 Diverse Samples: Support for Use of a Total Score and Six Subscale Scores

The construct of self-compassion was first operationally defined and introduced into the psychological literature a decade and a half ago (Neff, 2003b). Theoretically, self-compassion is comprised of six components that combine and mutually interact to create a self-compassionate frame of mind when faced with personal inadequacy or life difficulties: self-kindness versus self-judgment, a sense of common humanity versus isolation, and mindfulness versus over-identification. Self-kindness entails being gentle, supportive, and understanding towards oneself. Rather than harshly judging oneself for shortcomings, the self is offered warmth and acceptance. Common humanity involves recognizing the shared human experience, understanding that all humans fail, make mistakes, and lead imperfect lives. Rather than feeling isolated by one's imperfection - egocentrically feeling as if "I" am the only one who has failed or am suffering - one takes a broader and more connected perspective with regard to personal shortcomings and individual difficulties. Mindfulness involves being aware of one’s present moment experience of suffering with clarity and balance, without running away with a dramatic storyline about negative aspects of oneself or one’s life experience - a process that is termed "over-identification." As Neff (2016a) writes, the various components of self-compassion are conceptually distinct and tap into different ways that individuals emotionally respond to pain and failure (with more kindness and less judgment), cognitively understand their predicament (as part of the human experience rather than as isolating), and pay attention to suffering (in a more mindful and less over-identified manner). The six elements of self-compassion are separable and do not co-vary in a lockstep manner, but they do mutually impact one another. Put another way, Neff (2016a, 2016b) proposes that self-compassion represents a dynamic system in which the various elements of self-compassion are in a state of synergistic interaction.

Over the last few years, research on self-compassion has grown at an exponential rate. There have been almost 1500 articles or dissertations written about self-compassion since 2003 (based on a Google Scholar search of entries with "self-compassion" in the title, May 2018), over half of which have been published in the last two years. The majority of research studies have utilized the Self-Compassion Scale (SCS; Neff, 2003a) to examine the construct of self-compassion. The SCS is intended to be used as a total score to measure self-compassion, or else as six subscale scores to assess its constituent elements: Neff (2016a, 2016b) proposes that the state of self-compassion entails more compassionate fewer uncompassionate responses to personal suffering, which is why the SCS measures both.

Neff's operationalization of the SCS was based on compassion for others as broadly conceptualized in Buddhist philosophy (2003b), although scores on the SCS have a relatively weak correlation with compassion for others (Neff & Pommier, 2013). This appears to be because most people have significantly more compassion for others than for themselves (Neff, 2003a; Neff & Pommier, 2013), meaning the two do not necessarily go hand in hand.

Research using the SCS suggests that self-compassion is a key indicator of wellbeing. For instance, cross-sectional research using the SCS shows that self-compassion has moderate to strong positive associations with outcomes such as happiness, optimism, life satisfaction, body appreciation and motivation and negative associations with outcomes such as depression, anxiety, maladaptive perfectionism and fear of failure – findings that are replicated using experimental methods such as interventions or mood manipulations (see Neff & Germer, 2017, for a review). While research suggests that self-compassion yields similar mental health benefits as other positive self-attitude constructs such as self-esteem (Neff, 2011), it does not appear to have the same pitfalls (Crocker & Park, 2004). For instance, Neff and Vonk (2009) found that while self-compassion and self-esteem were strongly correlated, simultaneous regressions indicated that self-compassion was associated with more stable and less contingent feelings of self-worth over time, and was associated with less social comparison, public self-consciousness, self-rumination, anger, closed-mindedness and narcissism than self-esteem. Similarly, an experience sampling study conducted by Krieger, Hermann, Zimmermann and Grosse Holtforth (2015) found that levels of self-compassion, but not self-esteem, predicted less negative affect when encountering stressful situations over a 14-day period.

The incremental predictive validity of SCS scores have been demonstrated with constructs such as neuroticism (Neff, Tóth-Király & Colosimo, in press; Stutts, Leary, Zeveney & Hufnagle, in press) and self-criticism (Neff, 2003a). Although a key feature of self-compassion is the lack of self-judgment, overall SCS scores still negatively predict anxiety and depression when controlling for self-criticism and negative affect (Neff, Kirkpatrick & Rude, 2007).

It should be mentioned that there are other models and measures of self-compassion, and that there is a lack of consensus in the field on how to define or measure compassion for self or others (Gilbert et al., 2017; Gilbert, Clarke, Hempel, Miles & Irons, 2004; Strauss et al., 2016). Given that the SCS is the most commonly used measure of self-compassion, however, the current study is aimed at examining the psychometric properties of the SCS in a way that is theoretically consistent with Neff's (2003b) operationalization of the construct.

The SCS was developed in a sample of U.S. college undergraduates (Neff, 2003a). Confirmatory factor analyses (CFA) were used to provide support that scale items fit as intended with the a priori theoretical model (Furr & Bacharach, 2008). An initial CFA found an adequate fit for a six-factor inter-correlated model and a higher-order factor model. Since that time at least 30 published studies have examined the factor structure of the SCS (see Table S1 in the supplemental materials for a summary). Multiple translations of the SCS have been published, most have which have replicated the six-factor structure of the SCS using CFA. While not all examined the higher-order model, those that did yielded inconsistent findings. For example, a higher-order factor was supported with a Czech (Benda & Reichová, 2016, Norwegian (Dundas et al., 2016), and two Portuguese samples (Castilho, Pinto-Gouveia, & Duarte, 2015; Cunha, Xavier & Castilho, 2016), but not with German (Hupfeld & Ruffieux, 2011), Italian (Petrocchi et al., 2013) or a third Portuguese sample (Costa et al., 2015).

Recently, there has been controversy over whether or not self-compassion should be measured as an overall construct, or if compassionate versus uncompassionate self-responding should be measured separately. Some have found that use of a total score is not justified through higher-order factor analyses, and have argued that two separate positive and negative factors demonstrate better fit (e.g., Costa et al., 2015; López et al., 2015; Montero-Marín et al., 2016). These researchers tend to use the term “self-compassion” to describe the positive factor and "self-criticism" or "self-coldness" to describe the negative factor (Costa et al., 2015; Gilbert et al., 2011; López et al., 2015). However, self-criticism and self-coldness primarily describe self-judgment, or how people emotionally respond to suffering, and do not describe isolation (a way of cognitively understanding suffering) or over-identification (a way of paying attention to suffering). Moreover, these terms may obscure the fact that items representing negative self-responding are reverse-coded to indicate their relative absence. Therefore, we prefer the terms compassionate self-responding (CS) to represent the three components of self-kindness, common humanity and mindfulness and reduced uncompassionate self-responding (RUS) to represent lessened self-judgment, isolation and over-identification measured by the SCS.

The question of whether the SCS can be used to measure self-compassion as a holistic state of being or if it should be used to measure two distinct states of being has important implications for our understanding of what self-compassion is. If self-compassion does not include RUS, the implication would be that how self-critical, isolated, and over-identified individuals are in times of struggle have little bearing on how self-compassionate they are. This, in turn, would have implications for researchers’ attempts to examine the link between self-compassion and well-being. For instance, Muris and Petrocchi (2017) conducted a meta-analysis of the link of the SCS subscales with psychopathology across 18 studies and found the three components representing RUS had a stronger association with psychopathology (e.g., depression, anxiety and stress) than the CS components. They argue that negative items “may inflate the relationship with psychopathology” (p. 734) and should therefore be excluded from the SCS. If, however, RUS is an integral part of self-compassion, then logically speaking it cannot “inflate” its own association with psychopathology. Rather, RUS could be interpreted to “explain” the link between self-compassion and psychopathology. Support for this point of view can be found in studies designed to examine self-compassion through mood induction (i.e., using writing prompts) or through intervention, which show that increasing self-compassion experimentally also leads to reduced negative outcomes such as depression, anxiety, shame, etc. (see Neff & Germer, 2017). Not including RUS subscales in the measurement of self-compassion, therefore, could potentially underestimate its relationship to psychopathology.

Some have argued that that the CS and RUS subscales should not be combined into a total self-compassion score because compassionate responding is associated with parasympathetic nervous system activity and uncompassionate responding with sympathetic activity (Gilbert, McEwan, Matos, & Rivis, 2011). However, research with the SCS shows that the CS and RUS subscales do not substantially differ in terms of their association with markers of sympathetic response (e.g., alpha-amylase, interleukin-6) after a stressful situation (Neff et al., in press), or vagally mediated heart-rate variability, a marker of parasympathetic response (Svendson et al, 2016). As Porges (2001) makes clear, the two types of autonomic nervous system responding themselves interact and co-vary as a system. The issue of whether self-compassion is best measured as a total score or if CS and RUS should be measured separately is largely a psychometric question, however, which has yet to be definitively established.

Alternative Models for Examining the Factor Structure of the SCS

It is important that the psychometric analyses used to examine psychological measures be consistent with the psychological theory underlying those measures (Morin, Arens, Tran & Caci, 2016b). Higher-order models are commonly employed to validate the simultaneous use of a total score and sub-scale scores in measures of multidimensional psychological constructs (e.g., Chen, West, & Sousa, 2006; Gignac, 2016). A higher-order model represents several first-order factors (representing sub-scale scores) and a higher-order factor (representing a total score) that explains their inter-correlation, but makes the strong assumption that the higher-order factor only influences individual item responses through the pathway of the first-order factors (appropriate for certain constructs like IQ). The original SCS publication (Neff, 2003a) used a higher-order model to justify use of a total and six subscale scores, but as mentioned above, support for a higher-order model has been inconsistent.

Williams, Dalgleish, Karl, and Kuyken (2014) did not find support for a higher-order factor in four different English samples (student, community, meditator, and clinical), but did find support for a six-factor correlated model. They suggested that the six subscales but not a total score be used. López et al. (2015) examined a Dutch community sample and did not find support for a higher-order factor, so conducted Exploratory Factor Analysis (EFA) and found that the positive items loaded on one factor and the negative items loaded on a second factor. No CFA was conducted to confirm this two-factor model, however. Costa et al. (2015) examined a Portuguese clinical sample and compared a higher-order model, a six-factor uncorrelated model, a two-factor uncorrelated model that separated positive and negative items, and a two-factor model that included correlated errors designed to improve model fit, and found that the two-factor model with correlated errors had the best fit. These latter two sets of researchers suggested that separate positive and negative scores be used rather than a total score.

The bifactor model is an increasingly popular way to model multidimensional constructs (Reise, 2012; Rodriguez, Reise, & Haviland, 2016a). A bifactor model does not assume that the general or group factors are higher or lower than the other but rather co-exist, and models the direct association of the general factor and group factors on individual item responses. This has the added benefit of enabling the calculation of omega values that represent the amount of reliable variance in item responding explained by the general factor. Note that with a bifactor model the group factors are not allowed to correlate. Although perhaps counter-intuitive, this improves interpretability. For instance, it models those aspects of an item (e.g., When something upsets me I try to keep my emotions in balance) that are shared by all items in the general factor (e.g., self-compassion), as well as those aspects that are only shared by other items in its group factor (e.g., mindfulness). Neff (2016a) argued that a bifactor model provides a better theoretical fit with her conceptualization of self-compassion than a higher-order model given that behaviors assessed by individual items are directly representative of self-compassion as a general construct in addition to its constituent group components.

Neff, Whittaker, and Karl (2017) examined the SCS using bifactor CFA analysis in four different U.S. populations: undergraduates, community adults, meditators, and a clinical population. While the one-factor, two-factor correlated, and higher-order models had poor fit across samples, the six-factor correlated and bifactor models had acceptable fit using liberal fit criteria in the undergraduate, community and meditator samples. Fit was inadequate in the clinical sample. Nonetheless, omega values revealed that over 90% of the reliable variance in scores could be explained by a general self-compassion factor in all four populations (including the clinical sample). Findings were interpreted as providing support for use of a total score as well as six subscale scores, but not as two positive and negative scores. Cleare, Gumley, Cleare and O'Conner (2018) independently replicated these findings in a Scottish sample: support was not found for a one-factor, two-factor, or higher-order model, but was found for a six-factor correlated and bifactor model, with 94% of the variance in item responding explained by a general factor.

Three additional studies on translations of the SCS have provided evidence for a model with six group factors and one general factor using a bifactor CFA approach: French (Kotsou & Leys, 2016) Brazilian Portuguese (Souza & Hutz, 2016), and Italian (Veneziani, Fuochi & Voci, 2017). However, Montero-Marín et al. (2016) did not find support for a CFA bifactor model in two Spanish and Brazilian-Portuguese samples of doctors, but did find support for two higher-order factors (CS and RUS) and six first-order factors. Moreover, Brenner, Health, Vogel and Credé (2017) found that a two-bifactor CFA model with six group factors and two uncorrelated general (CS and RUS) factors had better fit than a single-bifactor model in a sample of U.S. undergraduates, though findings for some indicators were poor and the choice of examining two uncorrelated general factors is not theoretically consistent with the construct of self-compassion. Thus, the dimensionality of the SCS is still in question. Also, the above-mentioned results suggest that the assumptions of CFA might be overly restrictive for the SCS, given the inconsistency of findings.

CFA makes the strict assumption that items can only load on their respective factors, and may fail to account for two main sources of construct-relevant dimensionality in complex scales like the SCS, potentially resulting in biased parameters (Morin, Arens & Marsh, 2016a; Morin, Arens et al., 2016b). These sources do not refer to random measurement error, but are related to the idea that items often present more than one source of true score variance and subsequently belong to more than one construct. The first source refers to the fact that individual items are expected to be associated with a global factor (e.g., self-compassion), in which specific factors are not differentiated, as well as specific group factors (e.g., self-kindness or reduced self-judgment), in which they are differentiated. As mentioned, the relation between specific and global factors can be modeled in a hierarchical or in a bifactor manner with the latter generally being preferred unless there are strong theoretical reasons for the application of the former.

The second source of dimensionality comes from the fact that the six components of the scale are conceptually close and interrelated as a system, meaning items within each subscale should be expected to have significant associations with items in other subscales. Indeed, a recent review of simulation studies (Asparouhov, Muthén & Morin, 2015) have shown that when cross-loadings between items and non-target factors are not expressed (i.e., cross-loadings are constrained to be zero), parameters are likely to be biased. Exploratory Structural Equation Modeling (ESEM) is specifically designed to model system level interactions (Marsh, Morin, Parker & Kaur, 2014; Morin, Marsh & Nagengast, 2013). In CFA, items are strictly allowed to load on one factor, and these additional associations between items and non-target factors are reflected in the form of modification indices and/or inflated inter-factor correlations, which are the only ways overlap can be expressed. In ESEM, these associations are expressed in the form of item cross-loadings. Unlike Exploratory Factor Analyses (EFA), in which no a priori hypotheses about models are advanced, ESEM with target rotation (Browne, 2001) can model a priori hypotheses and therefore be directly compared to CFA models (Marsh et al., 2014). ESEM has been suggested to result in substantially better fit and less strongly correlated factors than corresponding CFA solutions (Marsh, Liem, Martin, Morin, & Nagengast, 2011; Morin & Maïano, 2011; Tóth-Király, Orosz, et al., 2017).

ESEM has rarely been used to examine the SCS. However, Hupfeld and Ruffieux (2011) as well as Tóth-Király, Bőthe and Orosz (2017) applied ESEM to analyze the factor structure of the SCS and found that, compared to CFA, ESEM provided a better fit to the data. Moreover, to account for the two sources of construct-relevant dimensionality, Tóth-Király, Bőthe and Orosz (2017) also used the integrative bifactor ESEM framework (Morin, Arens, et al., 2016a, 2016b; Morin, Boudrias, Marsh, Madore, & Desrumeaux, 2016), and results strongly supported the presence of a global self-compassion factor as well as the six-specific factors. The overarching bifactor ESEM framework appears to be especially appropriate for the SCS because it can simultaneously model both the specific and overall relationship of items using a bifactor analytic approach as well as their interaction as a system with an ESEM approach.

The Current Study

In the current study, we examined the factor structure of the SCS using both CFA and ESEM analyses for five distinct models: a single factor, two-factor correlated, six-factor correlated, single-bifactor model (one general factor and six group factors), and a correlated two-bifactor model (a general factor representing CS with three group factors representing higher levels of self-kindness, common humanity and mindfulness, and a general factor representing RUS with three group factors representing lower levels of self-judgment, isolation and over-identification). Based on the existing literature, we expected that the one factor and two-factor correlated models would have poor fit, and the six factor-correlated, single-bifactor and two-bifactor models would have better fit. We also expected fit indices to be better in ESEM rather than CFA analyses given that it is more appropriate for modeling system-level interactions. Our overall goal was to determine the best-fitting solution that is also well-aligned with Neff's underlying model of self-compassion (2003b), given that this is the theoretical model that the SCS was created to measure.

We examined the SCS in 20 different samples. Because the SCS was developed in English we included 7 English samples, but also 13 samples from non-English speaking countries. We included student, community, meditator and clinical samples. The meditator and one of the clinical samples were the same as examined in Neff et al. (2017), and a second Portuguese clinical sample was also included (Castilho, Pinto-Gouveia, & Duarte, 2015). Given that the SCS is commonly used to assess outcomes of meditation-based and clinical interventions (e.g., Birnie, Speca, & Carlson, 2010; Kelly, Wisniewski, Martin-Wagar, & Hoffman, 2017), we felt it was important to include these populations. The comprehensiveness of this study was designed to try to find more definitive answers to questions regarding the factor structure of the SCS: Should a total score be used, or two separate scores representing CS and RUS?

Method

Procedure

This study was organized by the first three authors, who wanted to examine the factor structure of the SCS in a variety of international samples. SCS data for three samples from the United States (US) were contributed by the first and third authors, who originally collected the data for other research purposes. Appropriate Institutional Review Board (IRB) approval was received before collecting these data, which were de-identified for the current study before being statistically analyzed by the second author. To gather SCS data from samples outside of the US, researchers were contacted in other English and non-English-speaking countries. These researchers contributed SCS data for 17 additional samples, which had also been collected previously for other research purposes. (Information about the data source of each sample as well as participant recruitment procedures can be found in the supplementary materials). SCS data contributed from sources outside the US were received as secondary data and included no potential participant identifiers. IRB approval was not required for analyses of these de-identified secondary data, although researchers from outside the US also obtained local ethics committee approval before collecting their original data as appropriate.

Participants

The initial number of participants was 11,990 from 20 international samples drawn from the following counties: Australia, Brazil, Canada, China, France, Germany, Greece, Iran, Italy, Japan, South Korea, Norway, Portugal, Spain, United Kingdom, and United States. In total, we included 10 community, 6 student, 1 mixed community/student, 1 meditator, and 2 clinical samples. Participants were excluded if they were under age 18 or had more than 50% of their responses missing. Thus, the final sample included 11,685 respondents (3,296 males, 8,367 females, 22 unspecified), aged between 18 and 83 (M = 32.29, SD = 8.28). Specific sample characteristics can be seen in Table 1.

Measures

The SCS (Neff, 2003a) is a 26-item self-report questionnaire measuring the six components of self-compassion: Self-Kindness (5 items; e.g., “I try to be loving towards myself when I’m feeling emotional pain”), reduced Self-Judgment (5 items; e.g., “I’m disapproving and judgmental about my own flaws and inadequacies”), Common Humanity (4 items, e.g., “When things are going badly for me, I see the difficulties as part of life that everyone goes through”), reduced Isolation (4 items, e.g., “When I think about my inadequacies it tends to make me feel more separate and cut off from the rest of the world”), Mindfulness 4 items, e.g., (“When I’m feeling down I try to approach my feelings with curiosity and openness”), and reduced Over-Identification (4 items, e.g., "When something upsets me I get carried away with my feelings”). Responses are given on a scale from 1 (almost never) to 5 (almost always). Note that all items in the Self-Kindness, Common Humanity and Mindfulness subscales are positively-valenced, while all items in the Self-Judgment, Isolation and Over-Identification subscales are negatively valenced. Items representing uncompassionate self-responding are reverse-coded before calculating a total score to indicate their relative absence in a self-compassionate mindset. Means are calculated for each subscale, and a grand mean is calculated for a total self-compassion score. Neff (2003a) found that items forming a total SCS score evidenced good internal reliability (Cronbach's α =.92), as did the six subscales (Cronbach's α ranging from .75 to .81). Test-retest reliability over a three-week interval was good (total score, Cronbach's α =.93; six subscales, Cronbach's α ranging from .80 to .88). The current study also employed 12 SCS translations (out of 16 published): Brazilian Portuguese, Chinese, French, German, Greek, Persian, Italian, Japanese, Korean, Norwegian, Portuguese, and Spanish. A description of the psychometric properties of each SCS translation can be found in the supplementary materials.

Analyses

All statistical analyses were conducted with Mplus 7.4 (Muthén & Muthén, 1995-2015) with the weighted least squares mean- and variance-adjusted estimator (WLSMV) as it is more suitable for ordered-categorical items with five or less answer options than estimators based on maximum-likelihood (e.g., Bandalos, 2014; Finney & DiStefano, 2006). Prior to the main analyses, negative items were reverse-coded. In order to systematically investigate the potential sources of construct-relevant dimensionality of the SCS, five corresponding CFA and ESEM models were tested and subsequently contrasted: (1a, 1b) a one-factor model with a unitary self-compassion dimension; (2a, 2b) a two-factor correlated model with two unitary factors representing CS and RUS; (3a, 3b) a six-factor correlated model with six components of self-compassion; (4a, 4b) a single-bifactor model with a general self-compassion factor and six group factors; and (5a, 5b) a two-bifactor model including two correlated general factors representing CS and RUS, each with three group factors. As per typical model specifications, in the CFA-based models (1a-5a), items were only allowed to load on their a priori target factors with cross-loadings being constrained to zero. In the ESEM-based models (1b-5b), items were allowed to load on the non-target factors as well. ESEM was also estimated in a confirmatory manner with target rotation (Browne, 2001) as per prior suggestions (Asparouhov & Muthén, 2009) and applications (Tóth-Király, Bőthe, Rigó, & Orosz, 2017). In the correlated models (2a, 2b, 3a, and 3b), factors were allowed to correlate freely. In the case of the bifactor models (4a, 4b, 5a, and 5b), group factors were specified as orthogonal to the general factor, as is standard (e.g., Reise, 2012; Reise, Moore, & Haviland, 2010) but the two general factors were specified as correlated[1] (see also Tóth-Király, Morin, Bőthe, Orosz, & Rigó, 2018 for a similar application or Morin, Myers, & Lee, in press, for an overview). These models were tested in the total sample and individual samples.

In model assessment, instead of only relying on the chi-square test which is sensitive to sample-size (Marsh, Hau, & Grayson, 2005), commonly applied goodness-of-fit indices were

[1] In the two-bifactor ESEM model, the general factors were specified as CFA factors (i.e., no cross-loadings between them), while the six specific factors were specified as ESEM factors (i.e., cross-loadings between them were allowed).

examined with their respective thresholds (Hu & Bentler, 1999; Marsh et al., 2005): the Comparative Fit Index (CFI; ≥ .95 for good, ≥ .90 for acceptable), the Tucker–Lewis index (TLI; ≥.95 for good, ≥ .90 for acceptable), the Root-Mean-Square Error of Approximation (RMSEA; ≤ .06 for good, ≤ .08 for acceptable) with its 90% confidence interval, and the weighted root mean square residual (WRMR; ≤ 1.00 for acceptable). Note that we did not compare fit using AIC or BIC because these information criteria are only available for maximum likelihood-based estimations, which are less accurate for ordered categorical data. However, the primary purpose of these indices is to determine which models would be most likely to cross-validate in subsequent samples, and this study determines cross-validation directly by examining model fit in 20 different samples.

Analyses of data should not be based solely on fit indices, however. The close inspection of parameter estimates (e.g., factor loadings, cross-loadings and inter-factor correlations) and the theoretical conformity of the models may also reveal valuable information about measurement models (as proposed by Hu & Bentler, 1998; Marsh, Hau, & Wen, 2004; Marsh et al., 2011; Morin, Arens, et al., 2016a, 2016b; Morin, Boudrias, et al., 2016). When comparing first-order CFA and ESEM models, the emphasis should be on comparison of factor correlations and on the need to incorporate cross-loadings, assuming that both solutions have well-defined factors with strong target loadings. If there is a substantial difference in the size of correlations between CFA and ESEM, the latter results are preferred as they provide more exact estimates (Asparouhov et al., 2015). If differences are negligible, then CFA is preferred due to its greater parsimony. Relatively large cross-loadings in the ESEM model may suggest an unmodeled general factor, which can be tested with a bifactor model. The general factor should also be well-defined by strong and theoretically meaningful factor loadings. Additionally, reduced cross-loadings and some well-defined specific factors would also provide support for the bifactor representation[2]. A particularly important question relates to the inclusion of one or two general factors where, once

[2] Naturally, not all specific factors are well-defined in the bifactor model relative to the first-order model, given that the item-level covariance is disaggregated to two sources (general and specific factors) instead of one (e.g., Morin et al., 2016a; Tóth-Király et al., 2018).

again, the close examination of factor loadings is highly informative.

We also assessed the reliability of items in the models. In the case of the six-factor model we calculated composite reliability (Raykov, 1997) as opposed to Cronbach's alpha, which has been criticized as being less useful for determining the reliability of factors (Rodriguez et al., 2016a). It has the advantage of being model-based, taking into account factor loadings and item-specific measurement errors as well. Based on Bagozzi and Yi (1988), values above .60 are considered acceptable, whereas values above .70 are good. As bifactor models allow the partitioning of the different sources of variance into the global and specific factors, omega (ω) and omega hierarchical (ω_H) were also calculated for the best fitting models based on standardized estimates (Brunner, Nagy, & Wilhelm, 2012; Rodriguez, Reise, & Haviland, 2016b). Omega estimates the proportion of the variance in item responding that is attributed to both the global and specific factors. OmegaH estimates the proportion of variance that is attributed to the general factor only. Finally, we also compared the omega and omegaHs on the basis of Rodriguez et al. (2016b) to investigate the degree of reliable variance in item responding. For the variance attributed to the general factor, one should divide the value of omegaH by omega (i.e., ω_H / ω); for the remaining reliable variance attributable to the specific factors, one should subtract omegaH from omega (i.e., $\omega - \omega_H$). Reise, Bonifay, and Haviland (2013) suggest 75% or higher accounted for by the general factor as the ideal amount of variance to justify use of a total score in spite of the presence of multidimensionality of the data.

Results

Structural analyses

Because results were generally similar for the total sample and the individual samples, we mainly refer to results for the total sample for the sake of simplicity. We first examined the fit of the one-factor model for all samples (see the supplementary materials). The one-factor ESEM solution is fundamentally a one-factor CFA (using only different estimation routines in Mplus) as there are no cross-loadings in this model. In accord with our hypotheses, results clearly demonstrate the inadequacy of the unidimensional model (Total sample: CFI = .74, TLI = .73, RMSEA = .15 [90% CI .15-.15], WRMR = 14.44). Tables 2, 3, 4, and 5 present model fit indices for CFA and ESEM analyses for the two-factor, six-factor, single-bifactor and correlated two-bifactor models, respectively In the case of the two-factor correlated models (see Table 2), both the CFA (Total sample: CFI = .90, TLI = .89, RMSEA = .10 [90% CI .09-.10], WRMR = 7.48) and ESEM (Total sample: CFI = .88, TLI = .86, RMSEA = .11 [90% CI .11-.11], WRMR = 6.31) versions showed marginally acceptable fit indices in some samples, but the majority were not acceptable by commonly applied standards, hence we rejected these solutions. In the case of the six-factor correlated CFA and ESEM models (see Table 3), most CFA models had acceptable fit (Total sample: CFI = .95, TLI = .94, RMSEA = .07 [90% CI .07-.07], WRMR = 5.15). However, ESEM systematically outperformed these solutions as apparent with excellent fit indices (Total sample: CFI = .99, TLI = .97, RMSEA = .05 [90% CI .05-.05], WRMR = 1.75).

Following Morin et al. (2016a, 2016b), we also examined standardized item factor loadings for the corresponding CFA and ESEM solutions for the total sample to select the final models, presented in Tables 6, 7, and 8. When examining the six-factor correlated models (Table 6), all six factors were well-defined by their respective factor loadings (λ = .65 to .84, M_λ = .76) in CFA, but this solution also resulted in relatively high factor correlations (r = .38 to .91, Mr = .64), undermining the discriminant validity of interpretations of items in the six factors. In the ESEM model, factor loadings (λ = .26 to .97, M_λ = .56) as well as factor correlations (r = .16 to .66, Mr = .42) were systematically lower. These results are in line with previous studies (Morin et al., 2016a) showing that ESEM often provides a better representation of the inter-factor correlations. As expected, there were some cross-loadings ($|\lambda|$ = .00 to .42, M_λ = .10) between conceptually similar items ($\geq$ .32; Worthington & Whittaker, 2006). For example, the self-kindness item "I'm tolerant of my own flaws and inadequacies" cross-loaded on reduced self-judgment. Overall, cross-loadings were found for two self-kindness items on reduced self-judgment and one on mindfulness, one reduced self-judgment item on self-kindness, one mindfulness item on self-kindness and one on reduced over-identification, and two reduced over-identification items on reduced self-judgment.

The next question that we addressed is whether the single-bifactor model with one general factor (representing self-compassion) or the correlated two-bifactor model with two general factors (representing CS and RUS) was able to provide an improved representation of the data. For the single-bifactor models (Table 4), CFA models were generally inadequate (Total sample: CFI = .85, TLI = .82, RMSEA = .12 [90% CI .12-.12], WRMR = 10.55), whereas ESEM models generally had much better fit (Total sample: CFI = .99, TLI = .98, RMSEA = .04 [90% CI .04-.04], WRMR = 1.42). Results for CFA and ESEM were less differentiated for the correlated two-bifactor models including two general factors (see Table 5), with generally adequate fit for the CFA models (Total sample: CFI = .96, TLI = .95, RMSEA = .06 [90% CI .06-.06], WRMR = 4.49), as well as for the ESEM models (Total sample: CFI = .99, TLI = .99, RMSEA = .04 [90% CI .03-.04], WRMR = 1.20). However, it should be noted that the correlated two-bifactor CFA model for the total sample had misspecifications, and almost half (9 out of 20) of the individual samples had negative residual variances, suggesting that the data did not support the hypothesized models. Therefore, we only compared the parameter estimates of the competing single- and correlated two-bifactor ESEM models.

The parameter estimates for the single-bifactor model (Table 7) revealed a well-defined general factor ($|\lambda|$ = .36 to .75, M = .62) reflecting a global level of self-compassion. As for the specific factors, common humanity retained a higher degree of specificity ($|\lambda|$ = .35 to .73, M = .53) once the general factor was extracted. By the same token, isolation ($|\lambda|$ = .24 to .58, M = .41) and mindfulness ($|\lambda|$ = .28 to .52, M = .41) had moderate degree of specificity, self-kindness ($|\lambda|$ = .06 to .56, M = .34) and overidentification ($|\lambda|$ = .19 to .50, M = .34) had a smaller degree of specificity, whereas self-judgment ($|\lambda|$ = .07 to .44, M = .22) retained almost no meaningful specificity. Finally, cross-loadings also slightly decreased in magnitude ($|r|$ = .01 to .34, M = .09) relative to the six-factor ESEM model. In the case of the correlated two-bifactor-ESEM model (see Table 8), while the correlation between the two factors were reduced (r = .09, p = .086), the two general factors were weakly defined by their respective factor loadings (Positive: $|\lambda|$ = .01 to .48, M = .22; Negative: $|\lambda|$ = .04 to .35, M = .17), arguing against the incorporation of a second general factor and supporting the superiority of the single-bifactor ESEM model with one general factor. Taking these results together, it appears that a six-factor correlated model (representing the six components of self-compassion) and a single-bifactor model (representing a general self-compassion factor and six specific factors) are supported, but a correlated two-bifactor model (representing CS and RUS) is not supported once parameter estimates are taken into account.

Reliability analyses

Finally, we estimated composite reliability indices for items in the six-factor model and the omega and omegaH indices for items the single-bifactor ESEM model in order to examine reliability. For the six-factor model (examining the sample as a whole), items in all factors had acceptable levels of composite reliability using Bagozzi and Yi's (1988) criteria of > .60 as adequate and > .70 as good: (self-kindness = .84, reduced self-judgment = .73, common humanity = .81, reduced isolation = .83, mindfulness = .67, and reduced over-identification = .70). (Composite reliability for items in the individual samples are available upon request from the first author). Reliability results for the single-bifactor model for all samples are presented in Table 9, although again we only discuss results for the total sample here. The single-bifactor model displayed high omega (.96) and omegaH (.91) values, demonstrating that a large majority of the variance in item responding can be attributed to the general factor. As per Rodriguez et al. (2016b), we compared the ratio of omega and omegaH to establish the amount of reliable variance of items attributable to the general factor (omegaH divided by omega) and that attributable to the multidimensionality caused by the specific factors (omegaH subtracted from omega). For the single-bifactor model, 95% of the reliable variance in item responding was attributed to the general self-compassion factor, whereas 5% was attributed to the group factors.

Discussion

Our analyses, which were designed to determine the best factor structure for the SCS, found that a one- and two-factor solution to the SCS had an inadequate fit using both CFA and ESEM. In contrast, a six-factor correlated solution had good fit using ESEM (CFA results for the six-factor solution were also acceptable) in every sample examined. The single-bifactor ESEM model (with one group and six specific factors) also had good fit in every sample. Moreover, inspection of factor loadings suggested good parameter estimates for a single general factor in ESEM. While the correlated two-bifactor ESEM model with two correlated general factors also had good model fit, factor loadings indicated poor specification of separate factors representing CS and RUS, so this model was rejected. Note that the single-bifactor ESEM model also had the highest level of theoretical conformity with Neff's (2003b) view that self-compassion is comprised of six components that interact as a global system. Results for our final selected models were remarkably similar across the 20 diverse populations examined - including student, community, clinical, and meditator samples in 13 different languages - providing strong support for the generalizability of the SCS to measure self-compassion.

Findings regarding cross-loadings in the ESEM models are also informative. In the six-factor model all factors were well defined, but eight cross-loadings were found (cross loadings were found equally within and across the CS and RUS dimensions). These cross-loadings highlight the importance of using models such as ESEM that can uncover this particular source of construct-relevant dimensionality. Use of a total SCS score was supported by the finding that 95% of the reliable variance in SCS item responding could be explained by a general factor for the total sample, ranging from 86% to 96% for the individual samples. This is well over the 75% or higher suggested by Reise et al. (2013) to justify use of a total score. All of the factors in the six-factor solution had adequate to good levels of composite reliability based on conventional thresholds (Bagozzi & Yi, 1988). In the single-bifactor model the general factor was well-defined and the specific factors were moderately well-defined. These observations give support for the idea that the specific factors assess relevant components over and above the general factors. They also support the system level interaction of components. We interpret these results as supporting use of a global score (representing self-compassion) or six subscale scores (representing self-kindness, common humanity, mindfulness and reduced self-judgment, isolation and over-identification), but not two separate CS and RUS scores.

The fact that the one- and two-factor solution had poor fit but a six-factor solution had good fit makes sense theoretically. It is potentially problematic to argue that self-compassion is a unitary construct (no theorists we are aware of have made this argument), or to argue that the three subscales representing CS versus RUS each form unitary constructs, as proposed by some (e.g., Costa et al., 2015; López et al., 2015). The three subscales within each of these dimensions are distinct, and tap into the way that people *emotionally* respond to suffering (with self-kindness or reduced self-judgment), *cognitively* understand their suffering (with common humanity or reduced isolation), and *pay attention* to their suffering (with mindfulness or reduced over-identification). Thus, within the dimensions of CS and RUS the three components are not thought to be identical.

Given that support was found for use of a total score and also six separate subscale scores, the question arises - when is use of a total score versus subscale scores warranted? The nomological network observed between the six subscales and important aspects of functioning indicates that there are areas of overlap but also difference between the subscales. For instance, Körner et al. (2015) found that it was mainly isolation that predicted depression, while Alda et al. (2016) found that common humanity had the strongest association with telomere length. Moreover, evidence from neuroimaging studies suggest the various components of self-compassion have distinct brain signatures. Longe et al. (2010) found that self-critical thinking (similar to self-judgment) and self-reassurance (similar to self-kindness) were associated with different regions of brain activity. Self-criticism was associated with activity in lateral prefrontal cortex (PFC) regions and dorsal anterior cingulate (dAC), linked to error processing and resolution, and also behavioral inhibition. Self-reassurance was associated with left temporal pole and insula activation, related to empathy. Mindfulness, on the other hand is linked to increased neural activation in the prefrontal cortex (PFC) and dorsal anterior cingulate cortex (dACC), associated with attentional control and emotion regulation (Young et al., 2017). These results suggest that the six components of self-compassion are not one unitary thing, nor are they two unitary things, but are six distinct but interrelated things.

Use of the subscales may have relevance for understanding the mechanisms by which self-compassion engenders well-being. Neff, Long, et al. (in press) recently explored the link of self-compassion and its components to psychological functioning in seven domains – psychopathology, positive psychological health, emotional intelligence, self-concept, body image, motivation, and interpersonal functioning. When examining the zero-order correlations between observed subscale scores and outcomes, they found that reduced self-judgment, isolation, and over-identification tended to have a stronger link to negative emotionality and self-evaluation than self-kindness, common humanity and mindfulness, while the latter tended to have a stronger association with outcomes like emotional awareness, goal re-engagement, compassion for others and perspective-taking. For many aspects of psychological functioning, however, such as happiness, wisdom, contingent self-esteem based on approval, body appreciation, or grit, all six subscales appeared to make an equal contribution to well-being. They interpreted findings to mean that although different elements of self-compassion may differentially explain its link with wellbeing, all are essential to the construct of self-compassion as a whole.

For most researchers, use of the SCS as a total score will be most appropriate given that self-compassion operates as a system. This view is supported by findings from intervention research indicating that self-compassion training changes all six components at the same time. The vast majority of intervention studies using a wide variety of methodologies that examined change in self-compassion have documented a simultaneous change in all six subscales of roughly the same magnitude: e.g. self-compassion meditation training (e. g., Toole & Craighead, 2016); online psycho-education (e.g., Krieger, Martig, van den Brink, & Berger, 2016); Compassion Focused Therapy (e.g., Beaumont, Irons, Rayner, & Dagnall, 2016); Compassionate Mind Training (e.g., Arimitsu, 2016) and the Mindful Self-Compassion program (e.g., Neff, 2016a). Not only do self-compassion interventions impact CS and RUS to the same degree, changes in both impact outcomes similarly. Krieger, Berger, and Holtforth (2016) used cross-lagged analyses to explore whether changes in self-compassion over the course of cognitive-behavioral psychotherapy led to changes in depression, and findings were the same whether a total score or two scores representing compassionate or uncompassionate responding were examined. They interpreted findings as evidence that self-compassion should be considered an overall construct rather than two separate constructs. Similarly, Neff (2016a) found that changes in SCS subscales representing CS and RUS after eight weeks of self-compassion training tended to be equally predictive of changes in happiness, life satisfaction, anxiety, depression and stress.

These findings suggest that self-compassion is experienced holistically. They also buttress current study findings supporting the use of a total SCS score to represent self-compassion as defined by Neff (2003b). Perhaps most importantly, they highlight why there is so much excitement about the construct of self-compassion in the field of psychology: It is a skill that can be learned (Neff & Germer, 2013). For researchers who are primarily interested in self-compassion as a trainable mind-state, therefore, use of a total score is probably most appropriate. For those more interested in unpacking the mechanisms of how self-compassion enhances well-being, however, it may be useful to examine the six constituent components themselves.

An important contribution of the present investigation is the finding that self-compassion is better represented with a single continuum rather than two distinct dimensions of CS and RUS. This notion was supported by the fact that the positively and negatively valenced items loaded on the general factor in a similar magnitude in the model including one general factor, whereas these loadings were weak in the model with two correlated general factors. It should be noted that the separation of positive and negative items sometimes results from a clustering effect where items with a similar valence load onto separate factors, basically forming method factors that mostly originate from the positive versus negative wording of the items (Crego & Widiger, 2014). This has been shown in research on self-esteem (Greenberger, Chen, Dmitrieva, & Farruggia, 2003; Marsh, 1996), for instance, where method factors emerged as a results of item wording. Generally, wording effects would be interpreted as substantively irrelevant artifacts, but in the case of the SCS, we do not believe that the separation of positively and negatively-valenced items are a result of item wording only. Rather, the distinction between compassionate and reduced uncompassionate responding toward oneself is conceptually meaningful and substantially contributes to the global self-compassion factor. Self-compassion can be conceptualized as a holistic state of mind representing the balance of CS and RUS along the three basic dimensions of emotional responding, cognitive understanding, and paying attention to personal distress.

Limitations and future directions

While this is one of the most comprehensive examinations of the factor structure of the SCS conducted to date, there were some limitations. For instance, the populations included were majority female and mainly community and student samples: only one meditator and two clinical samples were included. Fit in these latter samples was excellent, providing some confidence in use of the SCS with these populations. Still, it will be important to make sure the factor structure replicates in specific types of populations (anxious, eating disordered, etc.). Also, although findings support use of the SCS in different cultures, reliability coefficients and model fit did vary somewhat across samples (less so for our chosen models). Also, in some countries (e.g., China and Japan) multiple measurement models presented identification issues, and it should be investigated whether these issues relate to model misspecification or sampling-specific errors. Potential differences in the SCS structure should also be addressed with analyses of invariance across culture, population type, age and sex, as these may be additional sources of meaningful variation in the SCS that should be understood. (These analyses are being conducted for the current dataset and will be presented in a separate paper; Neff et al., 2018).

Given the superiority demonstrated by the ESEM models over CFA models, results suggest that future attempts to validate translations of the SCS or to examine the properties of the SCS in specific populations should use this approach (syntax files are available for interested readers in Appendix Five of the supplementary materials). Additional studies are also needed to examine the criterion-validity of test score interpretations using this improved representation in order to better capture the meaning of the subscales once the global level of self-compassion is accounted for. Use of the bifactor ESEM framework aligns with the proposition of Marsh and Hau (2007) who emphasized the need for the use of latent variable models which, compared to observed variables, more accurately define constructs with the explicit inclusion of measurement errors related to the imperfect items. Bifactor ESEM models are rather complex and sometimes difficult to incorporate into predictive models due to the relatively high number of estimated parameters, but one can construct separate measurement models and “translate” these measurement models into factor scores saved from these preliminary measurement models that are better at preserving the *a priori* nature of the constructs compared to observed variables (Morin, Meyer, Creusier, & Biétry, 2016; Gillet, Morin, Cougot, & Gagné, 2017). Neff, Tóth-Király, and Colosimo (2018) successfully used this approach to examine the incremental validity of self-compassion and neuroticism in predicting wellbeing. (Syntax for saving factor scores is also included in Appendix Five of the supplementary materials.)

Note that while the current study was aimed at examining the validity of test score interpretations on the SCS as a measure of Neff’s (2003b) conceptualization of self-compassion, in no way can it speak to the issue of whether this definition or measurement of self-compassion is superior to others. For example, Social Mentality Theory (SMT; Gilbert, 1989, 2005) posits that self-compassion is a state of mind that emerges from mammalian bio-social roles involving care-giving and care-seeking, while self-criticism emerges from evolved social roles that protect us from social threats. To this end Gilbert and colleagues developed the Forms of Self-Criticism and Self-Reassurance Scales (Gilbert, Clarke, Hempel, Miles & Irons, 2004) to measure these two ways of relating to oneself. More recently, Gilbert and colleagues (Gilbert et al., 2017) have developed a model of compassion for self, for others, and from others, based on the broadly used definition of compassion as sensitivity to suffering with a commitment to try to alleviate it (Goertz, Keltner, & Simon-Thomas, 2010). They developed the Compassion Engagement and Action Scales, including self-compassion and other compassion scales with items tapping into engagement with distress (e.g. tolerating and being sensitive to distress) and the motivation to alleviate that distress (e.g., thinking about and taking actions to help). Notably, the scales do not include kindness/concern or shared humanity as a feature of compassion. As with the SCS (Neff & Pommier, 2013), scale scores measuring compassion for self and others are only weakly correlated, with higher levels of compassion being reported for others than the self. It is unclear if the desire to alleviate distress operates the same way for self and others, however, given that the desire to alleviate personal distress overlaps with resistance to distress. Resistance can exacerbate psychopathology, which is why mindfulness-based clinical approaches such as Acceptance and Commitment Therapy (Hayes, Strosahl, & Wilson, 1999) and Mindfulness-Based Cognitive Therapy (Segal, Williams & Teasdale, 2012) are aimed at reducing resistance to personal distress.

Strauss et al. (2016) propose that measures of compassion should include five key elements: 1) Recognizing suffering; 2) Understanding the universality of suffering in human experience; 3) Feeling concern for the person suffering 4) Tolerating uncomfortable feelings in response to suffering, so remaining open to and accepting of the person suffering: and 5) Motivation to alleviate suffering. While the SCS taps into most of these elements, no items explicitly address the motivation to alleviate suffering. This is because the motivation to alleviate the self's suffering is easily conflated with resistance to personal distress (undermining the fourth element) in a way that is less problematic in measures of compassion for others. Still, future research might fruitfully explore whether adding items to the SCS that are focused on the motivation to help and support oneself in times of distress could strengthen the measurement of self-compassion.

To summarize, in the 20 diverse samples we examined, the excellent fit of single-bifactor ESEM and six-factor correlated ESEM models found across samples strongly supports the conclusion that self-compassion as measured by the SCS can be viewed as a general construct (explaining 95% of the reliable variance in item responding), comprised of six separate components. While the constituent elements of self-compassion are distinct and can be measured separately, they operate in tandem, as suggested by the large body of research examining self-compassion interventions. Hopefully these findings can help put some of the controversy over the factor structure of the SCS to rest: A total score rather than two separate scores should be used.

References

Alda, M., Puebla-Guedea, M., Rodero, B., Demarzo, M., Montero-Marin, J., Roca, M., & Garcia-Campayo, J. (2016). Zen meditation, length of telomeres, and the role of experiential avoidance and compassion. *Mindfulness, 7*(3), 651-659.

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

Arimitsu, K. (2016). The effects of a program to enhance self-compassion in Japanese individuals: A randomized controlled pilot study. *The Journal of Positive Psychology*, *11*(6), 559-571.

Arimitsu, K., Aoki, Y., Furukita, M., Tada, A., & Togashi, R. (2016). Construction and Validation of a Short Form of the Japanese version of the Self-Compassion Scale. *Komazawa Annual Reports of Psychology, 18,* 1-8.

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

Asparouhov, T., Muthén, B., & Morin, A. J. S. (2015). Bayesian structural equation modeling with cross-loadings and residual covariances: Comments on Stromeyer et al. *Journal of Management, 41*(6), 1561-1577.

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.

Bagozzi, R. P., & Yi, Y. (1988). On the evaluation of structural equation models. *Journal of the Academy of Marketing Science, 16*(1), 74–94.

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.

Beaumont, E., Irons, C., Rayner, G., & Dagnall, N. (2016). Does Compassion-Focused Therapy training for health care educators and providers increase self-compassion and reduce selfpersecution and self-criticism? *Journal of Continuing Education in the Health Professions, 36*(1), 4-10.

Benda, J., & Reichová, A. (2016). [Psychometric characteristics of the Czech version of the Self-Compassion Scale] Psychometrické charakteristiky české verze Self-Compassion Scale (SCS-CZ). *Československá psychologie. 60*(2), 20-36.

Birnie, K., Speca, M., & Carlson, L. E. (2010). Exploring self-compassion and empathy in the context of Mindfulness-based Stress Reduction (MBSR). *Stress and Health, 26*(5), 359-371.

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

Brenner, R. E., Heath, P. J., Vogel, D. L., & Credé, M. (2017). Two is more valid than one: Examining the factor structure of the Self-Compassion Scale (SCS). *Journal of Counseling Psychology, 64(6),* 696-707.

Browne, M. (2001). An overview of analytic rotation in exploratory factor analysis. *Multivariate Behavioral Research, 36*, 111-150.

Brunner, M., Nagy, G., & Wilhelm, O. (2012). A tutorial on hierarchically structured constructs. *Journal of Personality, 80*(4), 796–846.

Castilho, P., & Pinto-Gouveia, J. (2011). Self-Compassion: Validation of the Portuguese version of the Self-Compassion Scale and its relation with early negative experiences, social comparison and psychopathology. *Psychologica, 54,* 203-231.

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.

Chang, E. C., Yu, T., Jilani, Z., Fowler, E. E., Yu, E. A., Lin, J., & Hirsch, J. K. (2015). Under Assault: Understanding The Impact Of Sexual Assault On The Relation Between Hope And Suicidal Risk In College Students. *Journal of Social and Clinical Psychology*, *34*(3), 221-238.

Chen, F. F., West, S. G., & Sousa, K. H. (2006). A comparison of bifactor and second-order models of quality of life. *Multivariate Behavioral Research*, *41*(2), 189-225.

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.

Cleare, S., Gumley, A., Cleare, C. J., & O'Conner, J. C. (2018). An investigation of the factor structure of the Self-Compassion Scale. *Mindfulness, 9*(2), 618-628.

Coroiu, A., Kwakkenbos, L., Moran, C., Thombs, B., Albani, C., Bourkas, S., ... & Körner, A. (2018). Structural validation of the Self-Compassion Scale with a German general population sample. *PloS one*, *13*(2), e0190771.

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.

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.

Costa, J., Marôco, J., Pinto-Gouveia, J., Ferreira, C., & Castilho, P. (2015). Validation of the psychometric properties of the Self-Compassion Scale. *Clinical Psychology & Psychotherapy, 23,* 460-468.

Crego, C. & Widiger, T. A. (2014). Psychopathy, *DSM-5*, and a caution. *Personality Disorders: Theory, Research, and Treatment, 5*(4)*,* 335–347.

Crocker, J., & Park, L. E. (2004). The costly pursuit of self-esteem. *Psychological Bulletin, 130*, 392-414.

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.

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*, 58–72.

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.

Furr, R. M., & Bacharach, V. R. (2008). *Psychometrics: An Introduction*. Thousand Oaks: Sage.

Garcia-Campayo, J., Navarro-Gil, M., Andrés, E., Montero-Marin, J., López-Artal, L., & Demarzo, M. M. (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*(1), 4.

Ghorbani, N., Chen, Z., Saeedi, Z., Behjati, Z.,Watson,P.J. (2013). Sakhtare Ameli Meghyase Shafeghat e Khod dar Iran (Factorial Structure of Self-Compassion Scale in Iran) *Pazhohesh Haye Karbordi Ravanshenakhti, 4*(3), 29-41. (In Persian)

Gignac, G. E. (2016). The higher-order model imposes a proportionality constraint: That is why the bifactor model tends to fit better. *Intelligence*, *55*, 57-68.

Gilbert, P. (1989). *Human nature and suffering*. London: Routledge.

Gilbert, P. (2005). *Compassion: Conceptualizations, research and use in psychotherapy.* London: Brunner-Routledge.

Gilbert, P., Catarino, F., Duarte, C., Matos, M., Kolts, R., Stubbs, J…..& Basran, J. (2017). The development of compassionate engagement and action scales for self and others**.** *Journal of Compassionate Health Care,* DOI 10.1186/s40639-017-0033-3

Gilbert, P., Clarke, M., Hempel, S., Miles, J. N., & Irons, C. (2004). Criticizing and reassuring oneself: An exploration of forms, styles and reasons in female students. *British Journal of Clinical Psychology*, *43*(1), 31-50.

Gilbert, P., McEwan, K., Matos, M., & Rivis, A. (2011). Fears of compassion: Development of three self-report measures. *Psychology and Psychotherapy, 84,* 239–255.

Gillet, N., Morin, A. J. S., Cougot, B., & Gagné, M. (2017). Workaholism profiles: Associations with determinants, correlates, and outcomes. *Journal of Occupational and Organizational Psychology*, *90*(4), 559-586.

Goertz, J. L., Keltner, D., & Simon-Thomas, E. (2010). Compassion: An evolutionary analysis and empirical review. *Psychological Bulletin, 136*, 351–374.

Greenberger, E., Chen, C., Dmitrieva, J., & Farruggia, S. P. (2003). Item-wording and the dimensionality of the Rosenberg Self-Esteem Scale: do they matter? *Personality and Individual Differences*, *35*(6), 1241-1254.

Halamová, J., Kanovský, M., & Pacúchová, M. (2018). Self-Compassion Scale: IRT Psychometric Analysis, Validation, and Factor Structure–Slovak Translation. *Psychologica Belgica*, *57*(4), 190-209.

Hayes, S. C., Strosahl, K. D., & Wilson, K. G. (1999). *Acceptance and commitment therapy* (p. 6). New York: Guilford Press.

Hu, L. T., & Bentler, P. M. (1998). Fit indices in covariance structure modeling: Sensitivity to underparameterized model misspecification. *Psychological Methods*, *3*(4), 424-453.

Hu, L., & Bentler, P. M. (1999). Cutoff criteria for fit indexes in covariance structure analysis: Conventional criteria versus new alternatives. *Structural Equation Modeling, 6*, 1-55.

Hupfeld, J., & Ruffieux, N. (2011). Validation of a German version of the self-compassion scale (SCS-D). *Zeitschrift für Klinische Psychologie und Psychotherapie, 40*(2), 115–123.

Karakasidou, E., Pezirkianidis, C., Galanakis, M., & Stalikas, A. (2017). Validity, Reliability and Factorial Structure of the Self Compassion Scale in the Greek Population. *Journal of Psychology and Psychotherapy*, *7*, 313.

Kelly, A. C., Wisniewski, L., Martin-Wagar, C., & Hoffman, E. (2017). Group-based Compassion-Focused Therapy as an adjunct to outpatient treatment for eating disorders: A pilot randomized controlled trial. *Clinical Psychology & Psychotherapy*, *24*(2), 475-487.

Kim, K. E., Yi, G. D., Cho, Y. R., Chai, S. H., & Lee, W. K. (2008). The validation study of the Korean version of the self-compassion scale. *Korean Journal of Health Psychology, 13*, 1023–1044.

Körner, A., Coroiu, A., Copeland, L., Gomez-Garibello, C., Albani, C., Zenger, M., & Brähler, E. (2015). The role of self-compassion in buffering symptoms of depression in the general population. *PloS One, 10*(10). doi: 10.1371/journal.pone.0136598

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). doi: 10.1371/journal.pone.0152880

Krieger, T., Berger, T., & grosse Holtforth, M. (2016). The relationship of self-compassion and depression: Cross-lagged panel analyses in depressed patients after outpatient therapy. *Journal of affective disorders*, *202*, 39-45.

Krieger, T., Hermann, H., Zimmermann, J., & Grosse Holtforth, M. (2015). Associations of self-compassion and global self-esteem with positive and negative affect and stress reactivity in daily life: Findings from a smart phone study. *Personality and individual differences*, *87*, 288-292.

Krieger, T., Martig, D. S., van den Brink, E., & Berger, T. (2016). Working on self-compassion online: A proof of concept and feasibility study. *Internet Interventions, 6*, 64-70.

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.

Longe, O., Maratos, F. A., Gilbert, P. Evans, G., Volker, F., Rockliff, H., et al. (2009). Having a word with yourself: Neural correlates of self-criticism and self-reassurance. *Neuroimage, 49*, 1849–1856.

López, A., Sanderman, R., Smink, A., Zhang, Y., van Sonderen, E., Ranchor, A., & Schroevers, M. J. (2015). A reconsideration of the Self-Compassion Scale's total score: self-compassion versus self-criticism. *PloS One, 10*(7). doi: 10.1371/journal.pone.0132940

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, 123-132.

Marsh, H. W. (1996). Positive and negative global self-esteem: A substantively meaningful distinction or artifactors?. *Journal of Personality and Social Psychology*, *70*(4), 810-819.

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). NJ: Erlbaum.

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.

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.

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.

Montero-Marín, J., Gaete, J., Demarzo, M., Rodero, B., Lopez, L. C. S., & García-Campayo, J. (2016). Self-criticism: A measure of uncompassionate behaviors toward the self, based on the negative components of the self-compassion scale. *Frontiers in Psychology*, *7:*1281. doi: 10.3389/fpsyg.2016.01281

Morin, A. J. S., Arens, A. K., & Marsh, H. W. (2016a). 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.

Morin, A. J. S., Arens, A. K., Tran, A., & Caci, H. (2016b). 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*, *25*(4), 277-288.

Morin, A. J. S., Boudrias, J. S., Marsh, H.W., Madore, I., & Desrumeaux, P. (2016). Further reflections on disentangling shape and level effects in person-centered analyses: An illustration exploring the dimensionality of psychological health. *Structural Equation Modeling, 23*, 438-454.

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.

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., Meyer, J. P., Creusier, J., & Biétry, F. (2016). Multiple-group analysis of similarity in latent profile solutions. *Organizational Research Methods*, *19*(2), 231-254.

Morin, A. J. S., Myers, N. D., & Lee, S. (in press). Modern factor analytic techniques: Bifactor models, exploratory structural equation modeling (ESEM) and bifactor-ESEM. In G. Tenenbaum, & Eklund, R.C. (Eds.), *Handbook of Sport Psychology, 4th Edition*. Wiley.

Muris, P., & Petrocchi, N. (2017). Protection or vulnerability? A meta-analysis of the relations between the positive and negative components of self-compassion and psychopathology. *Clinical Psychology & Psychotherapy, 24*(2), 373-383.

Muthén, L. K., & Muthén, B. O. (1998-2017). *Mplus users guide (8th ed.).* Muthén & Muthén.

Neff, K. D. (2003a). Development and validation of a scale to measure self-compassion. *Self and Identity, 2*, 223-250.

Neff, K. D. (2003b). Self-compassion: An alternative conceptualization of a healthy attitude toward oneself. *Self and Identity, 2*, 85-102.

Neff, K. D. (2011). Self-compassion, self-esteem, and well-being. *Social and Personality Compass, 5,* 1-12.

Neff, K. D. (2016a). The Self-Compassion Scale is a valid and theoretically coherent measure of self-compassion. *Mindfulness, 7*(1), 264-274.

Neff, K. D. (2016b). Does self-compassion entail reduced self-judgment, isolation, and over-identification? *Mindfulness, 7,* 791-797.

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

Neff, K. D. & Germer, C. (2017). Self-compassion and psychological wellbeing. In J. Doty (Ed.) *Oxford Handbook of Compassion Science (pp. 371-386)*. Oxford University Press.

Neff, K. D., Kirkpatrick, K. & Rude, S. S. (2007). Self-compassion and its link to adaptive psychological functioning. *Journal of Research in Personality, 41,* 139-154.

Neff, K. D., Long, P., Knox, M. C., Davidson, O., Kuchar, A., Costigan, A…., & Breines (in press). The forest and the trees: Examining the association of self-compassion and its positive and negative components with psychological functioning. *Self and Identity.*

Neff, K. D., Pommier, E. (2013). The relationship between self-compassion and other- focused concern among college undergraduates, community adults, and practicing meditators. *Self and Identity, 12*(2)*,*160-176.

Neff, K. D., Tóth-Király, I., & Colisomo, K. (in press). Self-compassion is best measured as a global construct and is overlapping with but distinct from neuroticism: A response to Pfattheicher, Geiger, Hartung, Weiss, and Schindler (2017). *European Journal of Personality.*

Neff, K. D., Tóth-Király, I., Yarnell, L. M., Arimitsu, K., Castilho, P. Ghorbani, N., . . . Mantzios, M. (2018). *Examining invariance in the factor structure of the Self-Compassion Scale in 20 international samples.* Manuscript in preparation.

Neff, K. D., Whittaker, T., & Karl, A. (2017). Evaluating the factor structure of the Self-Compassion Scale in four distinct populations: Is the use of a total self-compassion score justified? *Journal of Personality Assessment, 99,* 596-607.

Neff, K. D. & Vonk, R. (2009). Self-compassion versus global self-esteem: Two different ways of relating to oneself. *Journal of Personality, 77,* 23-50.

Petrocchi, N., Ottaviani, C., & Couyoumdjian, A. (2013). Dimensionality of self-compassion: translation and construct validation of the self-compassion scale in an Italian sample. *Journal of Mental Health*, (0), 1-6.

Porges, S. W. (2001). The polyvagal theory: Phylogenetic substrates of a social nervous system.

International Journal of Psychophysiology, *42*(2), 123-146.

Raykov, T. (1997). Estimation of composite reliability for congeneric measures. *Applied Psychological Measurement, 21*(2), 173–184.

Reise, S. P. (2012). The rediscovery of bifactor measurement models. *Multivariate Behavioral Research, 47*(5), 667-696.

Reise, S. P., Bonifay, W. E., & Haviland, M. G. (2013). Scoring and modeling psychological measures in the presence of multidimensionality. *Journal of Personality Assessment, 95*(2), 129-140.

Reise, S. P., Moore, T. M., & Haviland, M. G. (2010). Bifactor models and rotations: Exploring the extent to which multidimensional data yield univocal scale scores. *Journal of Personality Assessment, 92*(6), 544-559.

Rodriguez, A., Reise, S. P., & Haviland, M. G. (2016a). Applying bifactor statistical indices in the evaluation of psychological measures. *Journal of Personality Assessment, 98*, 223-237

Rodriguez, A., Reise, S. P., & Haviland, M. G. (2016b). Evaluating bifactor models: Calculating and interpreting statistical indices. *Psychological Methods, 21*(2), 137-150.

Segal, Z. V., Williams, J. M. G., & Teasdale, J. D. (2012). *Mindfulness-based cognitive therapy for depression*. Guilford Press.

Souza, L. K. d., & 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.

Strauss, C., Taylor, B. L., Gu, J., Kuyken, W., Baer, R., Jones, F., & Cavanagh, K. (2016). What is compassion and how can we measure it? A review of definitions and measures. *Clinical psychology review*, *47*, 15-27.

Stutts, L., Leary, M., Zeveney, A., & Hufnagle, A. (in press). A longitudinal analysis of the relationship between self-compassion and the psychological effects of perceived stress. *Self and Identity*.

Svendsen, J. L., Osnes, B., Binder, P. E., Dundas, I., Visted, E., Nordby, H., ... & Sørensen, L.

(2016). Trait Self-Compassion Reflects Emotional Flexibility Through an Association with High Vagally Mediated Heart Rate Variability. *Mindfulness*, 1-11.

Toole, A. M., & Craighead, L. W. (2016). Brief self-compassion meditation training for body image distress in young adult women. *Body Image, 19*, 104-112.

Tóth-Király, I., Bőthe, B., & Orosz, G. (2017). Exploratory structural equation modeling analysis of the Self-Compassion Scale. *Mindfulness*, *8*(4), 881-892.

Tóth-Király, I., Bőthe, B., Rigó, A., & Orosz, G. (2017). An illustration of the exploratory structural equation modeling (ESEM) framework on the Passion Scale. *Frontiers in Psychology*, *8*:1968. doi: 10.3389/fpsyg.2017.01968

Tóth-Király, I., Morin, A. J. S., Bőthe, B., Orosz, G., & Rigó, A. (2018). Investigating the multidimensionality of need fulfillment: A bifactor exploratory structural equation modeling representation. *Structural Equation Modeling, 25*(2)*, 267-286.*

Tóth-Király, I., Orosz, G., Dombi, E., Jagodics, B., Farkas, D., & Amoura, C. (2017). Cross-cultural comparative examination of the Academic Motivation Scale using exploratory structural equation modeling. *Personality and Individual Differences, 106*, 130-135.

Veneziani, C. A., Fuochi, G., & Voci, A. (2017). Self-compassion as a healthy attitude toward the self: Factorial and construct validity in an Italian sample. *Personality and Individual Differences*, *119*, 60-68.

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.

Worthington, R. L., & Whittaker, T. A. (2006). Scale development research: A content analysis and recommendations for best practices. *The Counseling Psychologist, 34*(6), 806-838.

Young, K. S., van der Velden, A. M., Craske, M. G., Pallesen, K. J., Fjorback, L., Roepstorff, A., & Parsons, C. (2017). The impact of mindfulness-based interventions on brain activity: A systematic review of functional magnetic resonance imaging studies. *Neuroscience & Biobehavioral Reviews.*

Table 1

Characteristics of the Total and Individual Samples

Country	Language	Type	Initial *N*	Final *N*	Females	*M* Age (*SD*)
Total		Combined	11990	11685	8367 (71.6%)	32.29 (8.28)
AUS	English	Community	316	316	240 (75.9%)	37.20 (14.67)
BRA	Brazilian Portuguese	Community	312	312	241 (77.2%)	30.36 (10.76)
CAN	English	Student	395	362	308 (85.1%)	21.23 (4.02)
CHI	Chinese	Community	262	261	255 (97.7%)	37.02 (7.68)
FRA	French	Community	1554	1545	1362 (88.2%)	43.07 (12.48)
GER	German	Community	396	380	303 (79.7%)	29.43 (10.15)
GRE	Greek	Community	981	974	612 (62.8%)	21.99 (6.09)
IRA	Persian	Student	575	448	239 (53.3%)	25.33 (7.38)
ITA	Italian	Community	384	380	257 (67.6%)	33.56 (10.46)
JAP	Japanese	Student	718	718	291 (40.5%)	19.42 (1.16)
KOR	Korean	Student	353	343	180 (52.5%)	38.80 (9.22)
NOR	Norwegian	Student	327	318	189 (59.4%)	23.03 (3.40)
POR 1	Portuguese	Mixed	1128	1101	824 (74.8%)	24.71 (8.01)
POR 2	Portuguese	Clinical	314	297	236 (79.5%)	29.37 (8.43)
SPA	Spanish	Community	434	434	306 (70.5%)	49.71 (10.83)
UK 1	English	Community	1108	1085	969 (89.3%)	21.38 (5.69)
UK 2	English	Clinical	390	390	300 (76.9%)	50.16 (11.08)
US 1	English	Community	984	974	619 (63.6%)	38.17 (12.88)
US 2	English	Student	844	833	486 (58.3%)	21.22 (3.53)
US 3	English	Meditator	215	214	150 (70.1%)	47.36 (11.62)

Note. AUS = Australia; BRA = Brazil; CAN = Canada; CHI = China; FRA = France; GER = Germany; GRE = Greece; IRA = Iran; ITA = Italy; JAP = Japan; KOR = South Korea; NOR = Norway; POR = Portugal; SPA = Spain; UK = United Kingdom; US = United States

Table 2

Goodness-of-Fit indices for the Total and Individual Samples: Two-Factor Correlated Models

	Two-Factor Correlated CFA					Two-Factor Correlated ESEM				
	CFI	TLI	RMSEA	90% CI	WRMR	CFI	TLI	RMSEA	90% CI	WRMR
Total	.90	.89	.10	.09-.10	7.48	.88	.86	.11	.11-.11	6.31
AUS	.93	.92	.10	.09-.10	1.51	.92	.91	.10	.10-.11	1.25
BRA	.94	.93	.08	.08-.09	1.36	.94	.93	.08	.07-.09	1.13
CAN	.89	.88	.09	.09-.10	1.70	.89	.87	.10	.09-.10	1.38
CHI	.96	.96	.10	.10-.11	1.41	.96	.95	.11	.10-.12	1.20
FRA	.89	.88	.11	.11-.11	3.18	.89	.87	.12	.11-.12	2.68
GER	.81	.80	.11	.10-.12	1.89	.84	.81	.10	.10-.11	1.51
GRE	.92	.92	.08	.08-.09	2.19	.90	.88	.10	.10-.10	1.91
IRA	.81	.79	.09	.08-.09	1.88	.90	.88	.07	.06-.07	1.19
ITA	.85	.84	.12	.12-.13	2.15	.87	.84	.12	.11-.12	1.62
JAP	.86	.84	.10	.09-.10	2.64	.78	.73	.13	.12-.13	2.52
KOR	.91	.90	.09	.08-.10	1.79	.94	.93	.08	.07-.09	1.08
NOR	.89	.88	.10	.09-.10	1.68	.88	.86	.10	.10-.11	1.47
POR 1	.92	.92	.10	.10-.10	2.75	.90	.89	.12	.11-.12	2.22
POR 2	.89	.88	.10	.09-.10	1.57	.89	.87	.10	.09-.11	1.37
SPA	.82	.80	.11	.11-.11	2.25	.88	.85	.09	.09-.10	1.43
UK 1	.88	.87	.11	.10-.11	2.68	.88	.85	.11	.11-.11	2.27
UK 2	.89	.88	.09	.08-.09	1.68	.88	.85	.10	.09-.10	1.40
US 1	.91	.90	.10	.09-.10	2.29	.92	.90	.10	.09-.10	1.86
US 2	.83	.81	.11	.10-.11	2.61	.86	.83	.10	.10-.10	1.93
US 3	.92	.91	.10	.09-.11	1.37	.92	.91	.10	.10-.11	1.16

Note. AUS = Australia; BRA = Brazil; CAN = Canada; CHI = China; FRA = France; GER = Germany; GRE = Greece; IRA = Iran; ITA = Italy; JAP = Japan; KOR = South Korea; NOR = Norway; POR = Portugal; SPA = Spain; UK = United Kingdom; US = United States

Table 3

Goodness-of-Fit Indices for the Total and Individual Samples: Six-Factor Correlated Models

	Six-Factor Correlated CFA					Six-Factor Correlated ESEM				
	CFI	TLI	RMSEA	90% CI	WRMR	CFI	TLI	RMSEA	90% CI	WRMR
Total	.95	.94	.07	.07-.07	5.15	.99	.97	.05	.05-.05	1.75
AUS	.94†	.93	.09	.08-.10	1.27	.98	.97	.06	.06-.07	0.52
BRA	.96	.96	.06	.06-.07	1.05	.99	.98	.05	.04-.06	0.51
CAN	.93	.92	.08	.07-.08	1.30	.97	.94	.06	.06-.07	0.60
CHI	.97	.97	.09	.08-.10	1.17	.99†	.99	.06	.05-.07	0.46
FRA	.92	.91	.09	.09-.10	2.54	.98	.96	.06	.06-.06	0.89
GER	.87†	.85	.09	.09-.10	1.53	.98	.97	.05	.04-.05	0.50
GRE	.97	.96	.06	.05-.06	1.40	.98	.97	.05	.05-.06	0.68
IRA	.85†	.83	.08	.07-.08	1.62	.96	.93	.05	.04-.06	0.66
ITA	.91	.90	.10	.09-.10	1.60	.98	.97	.06	.05-.06	0.53
JAP	.93	.92	.07	.07-.07	1.75	.96	.93	.06	.06-.07	0.84
KOR	.92	.91	.09	.08-.09	1.60	.98	.96	.06	.05-.06	0.53
NOR	.93	.92	.08	.07-.08	1.26	.98	.97	.05	.04-.06	0.48
POR 1	.94†	.94	.09	.08-.09	2.20	.99	.97	.06	.05-.06	0.71
POR 2	.92†	.91	.08	.08-.09	1.30	.97	.95	.06	.05-.07	0.56
SPA	.86†	.84	.10	.09-.10	1.90	.97	.95	.05	.05-.06	0.58
UK 1	.94	.93	.08	.07-.08	1.80	.98	.97	.05	.04-.05	0.67
UK 2	.92	.90	.08	.07-.08	1.41	.98	.96	.05	.04-.06	0.55
US 1	.96	.95	.07	.07-.07	1.51	.99	.98	.04	.04-.05	0.57
US 2	.92	.91	.07	.07-.08	1.73	.98	.96	.05	.05-.06	0.67
US 3	.95	.95	.08	.07-.09	1.05	.99	.98	.05	.04-.06	0.43

Note. AUS = Australia; BRA = Brazil; CAN = Canada; CHI = China; FRA = France; GER = Germany; GRE = Greece; IRA = Iran; ITA = Italy; JAP = Japan; KOR = South Korea; NOR = Norway; POR = Portugal; SPA = Spain; UK = United Kingdom; US = United States; † These solutions had model identification issues, suggesting overparameterization.

Table 4

Goodness-of-Fit Indices for the Total and Individual Samples: Bifactor Models

	Bifactor CFA					Bifactor ESEM				
	CFI	TLI	RMSEA	90% CI	WRMR	CFI	TLI	RMSEA	90% CI	WRMR
Total	.85	.82	.12	.12-.12	10.55	.99	.98	.04	.04-.04	1.42
AUS	.92	.90	.11	.10-.11	1.53	.99	.98	.05	.04-.06	0.40
BRA	.93	.92	.09	.08-.09	1.34	.99	.98	.05	.04-.06	0.45
CAN	.85	.82	.12	.11-.12	2.31	.97	.95	.06	.05-.07	0.51
CHI	.95†	.94	.12	.11-.13	1.57	.99†	.99	.05	.04-.06	0.39
FRA	.89	.87	.11	.11-.12	3.23	.99	.98	.05	.05-.05	0.69
GER	.88	.85	.09	.09-.10	1.53	.99	.97	.04	.03-.05	0.43
GRE	.83	.80	.13	.13-.13	3.51	.99	.98	.04	.04-.05	0.53
IRA	.67	.61	.12	.12-.13	2.34	.97	.94	.05	.04-.06	0.57
ITA	.89	.87	.11	.10-.11	1.88	.99	.97	.05	.04-.06	0.45
JAP			no identification			.97†	.95	.06	.05-.06	0.68
KOR	.63	.56	.19	.19-.20	3.96	.98	.97	.05	.04-.06	0.45
NOR	.87	.85	.11	.10-.11	1.70	.99	.97	.05	.04-.06	0.43
POR 1	.83	.80	.15	.15-.15	4.61	.99	.98	.05	.05-.06	0.62
POR 2	.85	.82	.12	.11-.12	1.79	.98	.95	.06	.05-.07	0.48
SPA	.74	.69	.14	.13-.14	2.84	.98	.96	.05	.04-.06	0.48
UK 1	.89	.87	.11	.10-.11	2.73	.99	.98	.04	.04-.05	0.55
UK 2	.82†	.79	.11	.11-.12	2.04	.98	.97	.05	.04-.05	0.48
US 1	.90	.88	.11	.10-.11	2.54	.99	.99	.04	.03-.04	0.49
US 2	.80	.76	.12	.11-.12	3.20	.98	.96	.05	.04-.05	0.57
US 3	.92	.90	.11	.10-.11	1.32	.99	.99	.04	.02-.05	0.36

Note. AUS = Australia; BRA = Brazil; CAN = Canada; CHI = China; FRA = France; GER = Germany; GRE = Greece; IRA = Iran; ITA = Italy; JAP = Japan; KOR = South Korea; NOR = Norway; POR = Portugal; SPA = Spain; UK = United Kingdom; US = United States; † These solutions had model identification issues, suggesting overparameterization.

Table 5

Goodness-of-Fit Indices for the Total and Individual Samples: Correlated Two-Bifactor Models

	Correlated Two-Bifactor CFA					Correlated Two-Bifactor ESEM				
	CFI	TLI	RMSEA	90% CI	WRMR	CFI	TLI	RMSEA	90% CI	WRMR
Total	.96†	.95	.06	.06-.06	4.49	.99	.99	.04	.03-.04	1.20
AUS	.96	.95	.08	.07-.08	1.09	.99	.98	.04	.03-.05	0.36
BRA	.97†	.96	.06	.06-.07	1.02	.99	.98	.04	.03-.05	0.41
CAN			no identification			.98	.97	.05	.04-.06	0.44
CHI			no identification			.99†	.99	.05	.04-.06	0.36
FRA	.95†	.94	.08	.07-.08	2.06	.99	.98	.04	.04-.05	0.59
GER	.90	.88	.08	.08-.09	1.37	.99†	.97	.04	.03-.05	0.42
GRE			no identification			.99	.99	.04	.03-.04	0.46
IRA			no identification			.98	.96	.04	.03-.05	0.51
ITA	.94	.93	.08	.07-.08	1.31	.99	.98	.04	.03-.05	0.40
JAP			no identification			.99†	.97	.04	.04-.05	0.55
KOR			no identification			.99	.97	.05	.04-.06	0.40
NOR	.93	.91	.08	.08-.09	1.31	.99	.98	.04	.03-.05	0.40
POR 1	.96	.96	.07	.07-.08	1.88	.99	.98	.05	.04-.05	0.54
POR 2			no identification			.98	.96	.06	.05-.06	0.44
SPA			no identification			.99	.97	.04	.03-.05	0.44
UK 1	.94	.93	.08	.07-.08	1.74	.99	.99	.03	.03-.04	0.48
UK 2	.93†	.92	.07	.06-.08	1.29	.99	.97	.04	.03-.05	0.44
US 1	.96	.95	.07	.07-.08	1.56	.99	.99	.03	.02-.04	0.42
US 2	.91	.90	.08	.08-.08	1.83	.99	.97	.04	.04-.05	0.52
US 3			no identification			.99	.99	.04	.02-.05	0.33

Note. AUS = Australia; BRA = Brazil; CAN = Canada; CHI = China; FRA = France; GER = Germany; GRE = Greece; IRA = Iran; ITA = Italy; JAP = Japan; KOR = South Korea; NOR = Norway; POR = Portugal; SPA = Spain; UK = United Kingdom; US = United States; † These solutions had model identification issues, suggesting overparameterization.

Table 6

Standardized Factor Loadings for the Six-Factor Correlated CFA and ESEM Solutions of the Self-Compassion Scale for the Total Sample

	CFA	ESEM					
	SF (λ)[1]	SK (λ)	SJ (λ)	CH (λ)	IS (λ)	MI (λ)	OI (λ)
Self-kindness							
sk5	**.74**	**.69**	.04	.12	*.01*	.02	*.01*
sk12	**.82**	**.84**	.03	.02	*.01*	*.01*	.04
sk19	**.84**	**.80**	.03	*.00*	*.00*	.08	*.01*
sk23	**.75**	**.26**	.42	.08	.05	.31	.16
sk26	**.79**	**.36**	.34	.11	*.00*	.34	.18
Self-judgment							
sj1	**.76**	.09	**.61**	.05	.06	.05	.18
sj8	**.74**	.25	**.43**	*.00*	.13	.16	.22
sj11	**.74**	.09	**.51**	*.02*	.13	.05	.13
sj16	**.83**	.05	**.49**	.02	.23	.03	.20
sj21	**.73**	.33	**.33**	*.01*	.14	.14	.21
Common humanity							
ch3	**.70**	.04	.11	**.45**	.15	.19	*.01*
ch7	**.65**	.08	.06	**.97**	.04	.15	.04
ch10	**.73**	*.00*	.02	**.87**	.08	.06	.03
ch15	**.84**	.09	.05	**.43**	.07	.29	.06
Isolation							
is4	**.79**	*.01*	.28	.08	**.43**	*.01*	.13
is13	**.79**	*.01*	.10	.03	**.97**	.02	.06
is18	**.72**	.04	.10	.02	**.90**	*.00*	.03
is25	**.79**	*.00*	.20	.13	**.37**	.04	.26
Mindfulness							
mi9	**.66**	.13	.14	.09	.08	**.53**	.33
mi14	**.79**	.16	.15	.12	.09	**.58**	.17
mi17	**.77**	.14	*.01*	.16	.10	**.49**	.05
mi22	**.72**	.38	.02	.13	.08	**.29**	.06
Over-identification							
oi2	**.82**	.02	.34	.05	.20	.04	**.38**
oi6	**.78**	.03	.40	.08	.16	*.01*	**.31**
oi20	**.68**	.06	.07	*.01*	*.01*	.20	**.69**
oi24	**.69**	.05	.03	.02	.14	.21	**.58**

Note. CFA = confirmatory factor analysis; ESEM = exploratory structural equation modeling; SF = specific factor; [1] = Each item loaded on its respective specific factor, while cross-loadings were constrained to zero; SK = self-kindness; SJ = self-judgment; CH = common humanity; IS = isolation; MI = mindfulness; OI = over-identification; λ = standardized factor loadings. Target factor loadings are in bold. Non-significant parameters ($p \geq .05$) are italicized.

Table 7

Standardized Factor Loadings for the Bifactor CFA and ESEM Solutions of the Self-Compassion Scale for the Total Sample

	Bifactor-CFA		Bifactor-ESEM						
	GF (λ)	SF (λ)[1]	GF (λ)	SK (λ)	SJ (λ)	CH (λ)	IS (λ)	MI (λ)	OI (λ)
Self-kindness									
sk5	**.59**	**.50**	**.58**	**.47**	.04	.17	.05	.12	.06
sk12	**.66**	**.54**	**.64**	**.56**	*.01*	.11	.03	.12	.03
sk19	**.67**	**.53**	**.68**	**.50**	.03	.08	.07	.12	.07
sk23	**.66**	**.27**	**.72**	**.06**	.04	*.01*	.15	.08	.24
sk26	**.67**	**.33**	**.73**	**.13**	.13	.06	.19	.12	.26
Self-judgment									
sj1	**.65**	**.42**	**.67**	.06	**.44**	.12	.02	.12	.05
sj8	**.62**	**.44**	**.66**	.04	**.20**	.13	.06	.25	.13
sj11	**.64**	**.39**	**.70**	.09	**.15**	.12	*.01*	.14	.03
sj16	**.72**	**.39**	**.75**	.10	**.23**	.13	.11	.12	.09
sj21	**.63**	**.36**	**.67**	.11	**.07**	.10	.05	.25	.13
Common humanity									
ch3	**.51**	**.39**	**.46**	.09	.11	**.38**	.05	.24	.03
ch7	**.38**	**.72**	**.36**	.08	.07	**.73**	.05	.04	.02
ch10	**.48**	**.64**	**.44**	.11	.03	**.65**	.07	.10	*.01*
ch15	**.63**	**.36**	**.58**	.08	.07	**.35**	.05	.27	.12
Isolation									
is4	**.69**	**.26**	**.66**	.08	.20	.05	**.26**	.06	.10
is13	**.64**	**.58**	**.64**	.06	.02	.06	**.58**	*.00*	.05
is18	**.56**	**.57**	**.57**	.07	.03	.06	**.55**	.02	.07
is25	**.69**	**.25**	**.67**	.07	.11	*.00*	**.24**	.10	.20
Mindfulness									
mi9	**.53**	**.45**	**.50**	.09	.15	.15	.07	**.43**	.16
mi14	**.65**	**.55**	**.59**	.12	.10	.19	*.00*	**.52**	.04
mi17	**.64**	**.39**	**.61**	.08	.07	.17	.03	**.40**	.06
mi22	**.61**	**.27**	**.55**	.25	.07	.17	.04	**.28**	.12
Over-identification									

oi2	**.75**	**.23**	**.69**	.05	.34	.07	.17	.04	**.27**
oi6	**.71**	**.17**	**.68**	.07	.25	.07	.11	.08	**.19**
oi20	**.58**	**.57**	**.59**	.12	.06	.05	.05	.03	**.50**
oi24	**.60**	**.42**	**.60**	.11	.08	.03	.12	.05	**.41**

Note. CFA = confirmatory factor analysis; ESEM = exploratory structural equation modeling; GF = general factor of self-compassion; SF = specific factor; [1] = Each item loaded on its respective specific factor, while cross-loadings were constrained to zero; SK = self-kindness; SJ = self-judgment; CH = common humanity; IS = isolation; MI = mindfulness; OI = over-identification; λ = standardized factor loadings. Target factor loadings are in bold. Non-significant parameters ($p \geq .05$) are italicized.

Table 8

Standardized Factor Loadings for the Two-Bifactor CFA and ESEM Solutions of the Self-Compassion Scale for the Total Sample

	Two-bifactor-CFA			Two-bifactor-ESEM							
	CS (λ)	RUS (λ)	SF (λ)[1]	CS (λ)	RUS (λ)	SK (λ)	SJ (λ)	CH (λ)	IS (λ)	MI (λ)	OI (λ)
Self-kindness											
sk5	**.70**		**.32**	**.43**		**.24**	.31	.34	.10	.37	.06
sk12	**.78**		**.39**	**.48**		**.32**	.39	.30	.13	.40	.09
sk19	**.80**		**.31**	**.46**		**.25**	.42	.28	.11	.44	.07
sk23	**.77**		**.21**	**.31**		**.22**	.49	.26	.14	.40	.05
sk26	**.79**		**.08**	**.36**		***.17***	.45	.32	.09	.47	*.03*
Self-judgment											
sj1		**.72**	**.28**		**.16**	*.00*	**.71**	.07	.30	.06	.26
sj8		**.69**	**.33**		**.15**	*.11*	**.67**	.06	.21	.07	.23
sj11		**.70**	**.24**		**.09**	*.08*	**.64**	.09	.22	.17	.22
sj16		**.79**	**.22**		***.06***	*.04*	**.66**	.08	.34	.15	.31
sj21		**.69**	**.23**		**.35**	.18	**.67**	.09	.09	.14	.19
Common humanity											
ch3	**.56**		**.31**	***.07***		*.06*	.09	**.48**	.18	.39	.09
ch7	**.44**		**.70**	***.03***		*.07*	.04	**.79**	.04	.16	.06
ch10	**.54**		**.59**	***.08***		*.07*	.10	**.75**	.05	.23	.11
ch15	**.68**		**.27**	***.09***		*.00*	.22	**.50**	.13	.49	.05
Isolation											
is4		**.74**	**.14**		**.10**	*.01*	.50	.11	**.45**	.13	.28
is13		**.69**	**.52**		**.32**	*.03*	.31	.09	**.68**	.20	.21
is18		**.62**	**.52**		**.32**	*.02*	.26	.07	**.63**	.15	.21
is25		**.74**	**.12**		**.18**	*.02*	.46	.17	**.39**	.11	.37
Mindfulness											
mi9	**.59**		**.36**	***.09***		*.06*	.08	.26	.07	**.58**	.29
mi14	**.71**		**.48**	***.01***		*.13*	.15	.29	.16	**.76**	.15
mi17	**.70**		**.26**	***.12***		*.01*	.21	.32	.17	**.60**	.13
mi22	**.67**		**.12**	**.27**		*.07*	.22	.32	.14	**.49**	.04

Over-identification										
oi2	**.80**	**.05**	***.07***	.10	.52	.08	.41	.15	**.47**	
oi6	**.76**	***.02***	***.04***	*.02*	.56	.11	.31	.11	**.37**	
oi20	**.63**	**.86**	**.19**	*.01*	.28	.08	.16	.20	**.68**	
oi24	**.66**	**.21**	**.21**	*.01*	.28	.11	.23	.24	**.57**	

Note. CFA = confirmatory factor analysis; ESEM = exploratory structural equation modeling; CS = general factor representing Compassionate Self-Responding; RUS = general factor representing Reduced Uncompassionate Self-responding; SF = specific factor; [1] = Each item loaded on its respective specific factor, while cross-loadings were constrained to zero; SK = self-kindness; SJ = self-judgment; CH = common humanity; IS = isolation; MI = mindfulness; OI = over-identification; λ = standardized factor loadings. Target factor loadings are in bold. Non-significant parameters ($p \geq .05$) are italicized.

Table 9

Reliability Estimates for the Bifactor ESEM Model for the Total and Individual Samples

	Bifactor			
	ω	ω_H	GF	SF
Total	.96	.91	.95	.05
AUS	.98	.93	.95	.05
BRA	.97	.91	.94	.06
CAN	.96	.88	.92	.08
CHI	Negative Residual Variance			
FRA	.97	.92	.95	.05
GER	.96	.88	.92	.08
GRE	.97	.91	.94	.06
IRA	.93	.85	.91	.08
ITA	.96	.89	.93	.07
JAP	Negative Residual Variance			
KOR	.95	.82	.86	.13
NOR	.96	.89	.93	.07
POR 1	.97	.90	.93	.07
POR 2	.96	.90	.94	.06
SPA	.94	.83	.88	.11
UK 1	.97	.92	.95	.05
UK 2	.96	.89	.93	.07
US 1	.97	.93	.96	.04
US 2	.95	.87	.92	.08
US 3	.98	.93	.95	.05

Note. ω = Omega; ω_H = Omega Hierarchical; GF = Reliable variance explained by the general factor; SF = Reliable variance explained by the specific factors; AUS = Australia; BRA = Brazil; CAN = Canada; CHI = China; FRA = France; GER = Germany; GRE = Greece; IRA = Iran; ITA = Italy; JAP = Japan; KOR = South Korea; NOR = Norway; POR = Portugal; SPA = Spain; UK = United Kingdom; US = United States

Supplementary Materials

Examining the Factor Structure of the Self-Compassion Scale in 20 Diverse Samples: Support for Use of a Total Score and Six Subscale Scores

Table of Contents:

Appendix 1 Table S1

Previous studies examining the factor structure of the Self-Compassion Scale†

Original paper	Nation (Sample type)	Sample Size	Examined alternative models	Final chosen model
Arimitsu (2014)	Japan	N = 366 M_{age} = 19.6	(1) one-factor first-order CFA (2) six-factor first-order CFA (3) six-factor higher-order CFA (with one G-factor) (4) six-factor higher-order CFA (with two correlated G-factors)	(2)
Azizi et al. (2013)	Iran (student)	N = 265 M_{age} = 22.1	(1) six-factor first-order CFA	(1)
Benda & Reichová (2016)	Czech Republic (community)	N = 5368 M_{age} = —	(1) six-factor first-order CFA (2) six-factor higher-order CFA (with one G-factor)	both
Bento et al. (2016)	Portugal (pregnant women)	N = 417 M_{age} = 33	(1) six-factor first-order CFA	(1)
Brenner et al. (2017)	USA (student)	N = 1115 M_{age} = 19.4	(1) one-factor first-order CFA (2) two-factor first-order CFA (3) three-factor first-order CFA (4) six-factor first-order CFA (5) six-factor higher-order CFA (with one G-factor) (6) six-factor higher-order CFA (with two G-factors) (7) bifactor CFA (with one G-factor and six S-factors) (8) bifactor CFA (with two uncorrelated G-factors and six S-factors)	(8)
Castilho et al. (2015)	Portugal (community)	N = 1128 M_{age} = 24.5	(1) six-factor first-order CFA (2) six-factor higher-order CFA (with one G-factor)	both
	Portugal (clinical)	N = 316 M_{age} = 28.69	(1) six-factor first-order CFA (2) six-factor higher-order CFA (with one G-factor)	both
Chen et al. (2011)	China (student)	N = 660 M_{age} = —	(1) EFA (six factors were extracted) (2) six-factor first-order CFA	both
Cleare et al. (2018)	United Kingdom (community)	N = 526 M_{age} = 23	(1) EFA (five factors were extracted) (2) six-factor first-order CFA (3) six-factor higher-order CFA (with one G-factor) (4) one-factor CFA (5) two-factor first-order CFA (6) bifactor CFA (with one G-factor and six S-factors) (7) five-factor CFA (based on EFA)	(6)
Coroiu et al. (2018)	Germany (community)	N = 2510 M_{age} = 50.23	(1) one-factor first-order CFA (2) two-factor first-order CFA (3) three-factor first-order CFA (4) six-factor first-order CFA (5) six-factor higher-order CFA (with two correlated G-factors) (6) bifactor CFA (with one G-factor and two S-factors) (7) bifactor CFA (with two uncorrelated G-factors and two S-factors)	(7)

			(8) bifactor CFA (with one G-factor and 6 S-factors)	
Costa et al. (2015)	Portugal (clinical)	N = 361 M_{age} = 25.19	(1) six-factor first-order CFA (2) six-factor higher-order CFA (with one G-factor) (3) two-factor first-order CFA	(3)
Cunha et al. (2016)	Portugal (adolescent)	N = 3165 M_{age} = 15.49	(1) six-factor first-order CFA (2) six-factor higher-order CFA (with one G-factor)	both
de Souza & Hutz (2016)	Brazil (community)	N = 432 M_{age} = 32.5	(1) six-factor first-order CFA (2) six-factor higher-order CFA (with one G-factor) (3) bifactor CFA (with one G-factor and six S-factors)	(3)
Deniz et al. (2008)	Turkey (student)	N = 341 M_{age} = 19.81	(1) six-factor first-order CFA (2) EFA (one G-factor being extracted)	(2)
Dundas et al. (2016)	Norway (student)	N = 277 M_{age} = 22.9	(1) one-factor first-order CFA (2) three-factor first-order CFA (3) six-factor higher-order CFA (with one G-factor)	(3)
Garcia-Campayo et al. (2014)	Spain (student)	N = 268 M_{age} = 20.54	(1) six-factor first-order CFA	(1)
Halamová et al. (2018)	Slovakia (community)	N = 1181 M_{age} = 30.30 N = 676 M_{age} = 29.90	(1) six-factor first-order IRT CFA (2) bifactor IRT CFA (with one G-factor and six S-factors) (3) two-bifactor IRT CFA (with two correlated G-factors and six S-factors)	(3)
Hupfeld & Ruffieux (2011)	Germany (community)	N = 561 M_{age} = 26.04	(1) six-factor first-order ESEM	(1)
Karakasidou et al. (2017)	Greek (community)	N = 642 M_{age} = 36.83	(1) six-factor first-order CFA (2) two-factor first-order CFA (3) six-factor higher-order CFA (with one G-factor)	(1)
Kotsou & Leys (2016)	France (community)	N = 1554 M_{age} = 42.92	(1) six-factor first-order CFA (2) six-factor higher-order CFA (with one G-factor) (3) bifactor CFA (with one G-factor and six S-factors)	(3)
Lee & Lee (2010)	Korea (community females)	N = 405 M_{age} = —	(1) six-factor first-order CFA	(1)
López et al. (2015)	Netherlands† (community)	N = 1643 M_{age} = 54.9	(1) six-factor first-order CFA (2) six-factor higher-order CFA (with one G-factor) (3) two-factor EFA	(3)
Mantzios et al. (2015)	Greece (community)	N = 556 M_{age} = 24.43	(1) six-factor EFA	(1)

Montero-Marin et al. (2016)	Brazil (doctors)	N = 406 M_{age} = 41.09	(1) one-factor first-order CFA (2) two-factor first-order CFA (3) six-factor first-order CFA (4) six-factor higher-order CFA (with *one* G-factor) (5) six-factor higher-order CFA (with *two* correlated G-factors) (6) six-factor higher-order CFA (with *three* correlated G-factors) (7) six-factor third-order CFA (with one third order G-factor, three uncorrelated second order G-factors) (8) bifactor CFA (with one G-factor and six S-factors) (9) three-factor first-order CFA (positive items only) (10) three-factor higher-order CFA (with one G-factor, positive items only) (11) three-factor first-order CFA (negative items only) (12) three-factor higher-order CFA (with one G-factor, negative items only)	(12)
	Spain (doctors)	N = 416 M_{age} = 49.71		
Neff (2003)	USA (student)	N = 391 M_{age} = 20.91	(1) six-factor first-order CFA (2) six-factor higher-order CFA (with one G-factor)	(2)
Neff et al. (2008)	Taiwan† (student)	N = 164 M_{age} = 20.5	(1) six-factor first-order CFA	(1)
	Thailand† (student)	N = 223 M_{age} = 19.8	(1) six-factor first-order CFA	(1)
Neff et al. (2017)	USA (student)	N = 222 M_{age} = 20.94	(1) one-factor first-order CFA (2) two-factor first-order CFA (3) six-factor first-order CFA (4) six-factor higher-order CFA (with one G-factor) (5) bifactor CFA (with one G-factor and six S-factors)	(5)
	USA (community)	N = 1394 M_{age} = 36.01		
	USA (mediator)	N = 215 M_{age} = 47.40		
	USA (clinical)	N = 390 M_{age} = 50.16		
Neff et al. (2018)	USA (community)	N = 576 M_{age} = 37.21	(1) one-factor first-order CFA (2) one-factor first-order ESEM (3) two-factor first-order CFA (4) two-factor first-order ESEM (5) six-factor first-order CFA (6) six-factor first-order ESEM (7) bifactor CFA (with one G-factor and six S-factors) (8) bifactor ESEM (with one G-factor and six S-factors) (9) two-bifactor CFA (with two correlated G-factors and six S-factors)	(8)
	USA (community)	N = 581 M_{age} = 36.40		

			(10) two-bifactor ESEM (with two correlated G-factors and six S-factors)	
Petrocchi et al. (2014)	Italy (community)	N = 424 M_{age} = 36.53	(1) six-factor first-order CFA (2) six-factor higher-order CFA (with one G-factor) (3) one-factor first-order CFA (4) two-factor first-order CFA	(1)
Pfattheicher et al. (2017)	USA (community)	N = 576 M_{age} = 37.21	(1) one-factor first-order CFA (2) two-factor first-order CFA (3) six-factor higher-order CFA (with one G-factor) (4) six-factor higher-order CFA (with two G-factors) (5) six-factor first-order CFA	(4)
Tóth-Király et al. (2017)	Hungary (community)	N = 505 M_{age} = 44.37	(1) six-factor first-order CFA (2) six-factor first-order ESEM (3) bifactor CFA (with one G-factor and six S-factors) (4) bifactor ESEM (with one G-factor and six S-factors)	(4)
Veneziani et al. (2017)	Italy (community)	N = 522 M_{age} = 30.05	(1) one-factor first-order CFA (2) two-factor first-order CFA (3) six-factor first-order CFA (4) six-factor higher-order CFA (with one G-factor) (5) bifactor CFA (with one G-factor and two S-factors) (6) bifactor CFA (with one G-factor and six S-factors)	(6)
Williams et al. (2014)	UK (community)	N = 821 M_{age} = —	(1) one-factor first-order CFA (2) six-factor first-order CFA (3) six-factor higher-order CFA (with one G-factor)	(2)
	UK (meditator)	N = 211 M_{age} = —		
	UK (clinical)	N = 390 M_{age} = —		
Zeng et al. (2016)	China (Buddhist community)	N = 179 M_{age} = 35.5	(1) six-factor first-order CFA (2) two-factor first-order CFA (3) three-factor first-order CFA (separate for positive and negative items)	(3)
	China (non-Buddhist community)	N = 232 M_{age} = 30.1		

Note. Literature search was performed on April 10, 2018.; † = Examined an unpublished/unvalidated translation of the SCS; M_{age} = mean age; N = number of participants; CFA = confirmatory factor analysis; EFA = exploratory factor analysis; ESEM = exploratory structural equation modeling; IRT = item response theory; G-factor = general factor; S-factor = specific factor.

Appendix 2: Sample recruitment information

Australia

This sample was not part of a previously published study. Participants were recruited for this community sample (*N* = 316) through announcements posted online and surveys were completed online. Participants could choose to enter a lottery draw for the chance to win one of four $50 gift card prizes. Participants were 75.9% females, *M* age = 37.20 (*SD* = 14.67). Education information is not available.

Brazil

See Souza, Ávila-Souza and Gauer (2016) for a full description of recruitment procedures. Participants for this community sample (*N* = 312) were recruited through online announcements, along with the link for a widely used research platform. Participants were 77.2% females, *M* age = 30.36 (*SD* = 10.76). 70% of participants had a university degree.

Canada

See Sirois, Kitner and Hirsch (2015) for a full description of recruitment procedures. Participants for this student sample (*N* = 395) were randomly selected from a psychology subject pool at a large Canadian university. Participants were 85.1% females, *M* age = 21.23 (*SD* = 4.02).

China

This sample was not part of a previously published study. Participants in this community sample (*N* = 262) signed up for a Mindful Self-Compassion course, and took the SCS at pre-test. Participants were 97.7% females, *M* age = 37.02 (*SD* = 7.68). 8% had a high school education only, 64% had a college degree, and 23% had a graduate degree.

France

See Kotsou and Leys (2016) for a full description of recruitment procedures. Participants for this community sample (*N* = 1554) were recruited through an announcement posted online, and surveys were completed online. Participants were 88.2% females, *M* age = 43.07 (*SD* = 12.48). 64% of participants had at least an undergraduate level of education.

Germany

See Hupfield and Ruffieux (2011) for a full description of recruitment procedures. Participants for this community sample (*N* = 396) were recruited via online portals for German-speaking people with interest in psychological research... Participants were 79.7% females, *M* age = 29.43 (*SD* = 10.15). 71% of participants had at least an undergraduate level of education.

Greece

This sample was not part of a previously published study. Participants were recruited for this community sample (*N* = 981) by sending an email to employees of four Greek Universities inviting them to participate in an online survey, with instructions to pass the link on to other individuals or groups who might be interested in taking part, but avoid sharing with students. Following the same procedure, the survey link was also disseminated to employees of two Greek Hospitals - with instructions to avoid sending the link to patients. The aim was to obtain a nonclinical and a nonstudent sample, which would be broadly representative of the general population in Greece. Participants were 62.8% females, *M* age = 21.99 (*SD* = 6.09). Education information is not available.

Iran

This sample was not part of a previously published study. Undergraduate and graduate seminary students were recruited for this student sample (*N* = 575). Participants were 53.3% females, *M* age = 25.33 (*SD* = 7.38). 24% were graduate students holding a Bachelor's degree or higher.

Italy

See Petrocchi et al. (2014) for a full description of recruitment procedures. Participants for this community sample (*N* = 384) were recruited via several professional mailing lists and completed on online survey. Participants were 67.6% females, *M* age = 33.56 (*SD* = 10.46). Most respondents had finished high school (38.2%), 19.6% had a Bachelor’s degree, 42.2% had a Master’s degree or higher.

Japan

See Arimitsu, Aoki, Furukita, Tada & Togashi (2016) for a full description of recruitment procedures. Participants for this student sample (*N* = 718) were recruited from two metropolitan Japanese universities. Participants were 40.5% females, *M* age = 19.42 (*SD* = 1.16).

Korea

See Woo Kyeong (2013) for a full description of recruitment procedures. Participants for this student sample (*N* = 315) were randomly recruited from a counseling psychology subject pool at an online university. Participants were 52.5% females, *M* age = 38.80 (*SD* = 9.22).

Norway

See Dundas et al. (2016) for a full description of recruitment procedures. Participants for this undergraduate student sample (*N* = 327) were recruited from a Norwegian university (medical and psychology students) and a university college (engineering students). Participants were 59.4% females, *M* age = 23.03 (*SD* = 3.40).

Portugal 1

See Castilho, Pinto-Gouveia, and Duarte (2015) for a full description of recruitment procedures. Participants for this mixed student-community sample (*N* = 1128) were recruited from two large universities in Portugal and also from community groups in Portugal using nonrandom methods. Students were informed of the study by announcements made at the end of lectures, and participants from the community sample were recruited in several Portuguese institutions. Participants were 74.8% females, *M* age = 24.71 (*SD* = 8.01). 78% of participants were students.

Portugal 2

See Castilho, Pinto-Gouveia, and Duarte (2015) for a full description of recruitment procedures. Participants for this clinical sample (*N* = 316) were recruited from the outpatient psychiatric services of different public hospitals in Portugal and were referred by the psychologists and psychiatrists in charge. A trained therapist clinically assessed all participants using diagnostic structured interviews: Structured Clinical Interview for DSM-IV Axis I Disorders, Anxiety Disorders Interview Schedule for DSM-IV, Structured Clinical Interview for

DSM-IV Axis II Personality Disorders, and Borderline Personality Disorder Severity Index. Only patients with Axis I and II disorders participated in the study. Participants were 80% females. *M* age = 28.69 (*SD* = 8.74). 40% of participants were students.

Spain

See Montero-Marín, Zubiaga, et al. (2016) for a full description of recruitment procedures. Participants for this community sample (*N* = 434) were health care professionals who were randomly recruited from the mailing list of the Aragon Health Service. Participants were 70.5% females, *M* age = 49.71 (*SD* = 10.83). 49% subjects were physicians, 42% were nurses, and 10% were residents.

United Kingdom 1

This sample was not part of a previously published study. Participants for this community sample (*N* = 1108) were recruited by sending an invitation email to employees of two British Universities to take an online survey, with instructions to pass the invitation on to other individuals or groups who might be interested in taking part, but to avoid sharing the online link with students. The aim was to obtain a nonstudent sample, which would be broadly representative of a community sample in the United Kingdom. Participants were 89.3% females, *M* age = 21.38 (*SD* = 5.69). Education information is not available.

United Kingdom 2

See Williams et al. (2014) for a full description of recruitment procedures. Participants for this clinical sample (*N* = 405) were recruited through primary care settings in the United Kingdom. Criteria for this group included having a diagnosis of recurrent major depressive disorder in full or partial remission according to the Diagnostic and Statistical Manual of Mental Disorders (4th ed.; DSM–IV), having three or more previous major depressive episodes, and being 18 or older. Participants were 76.6% females, *M* age = 50.16 (*SD* = 11.8). 22% of participants had some education, 41% of participants had a high school or vocational education, and 32% had a university degree or other professional qualification.

United States 1

This sample was not part of a previously published study. Participants were recruited for this community sample (*N* = 984) from Mechanical Turk. Participants were directed to Survey Monkey in order to take the study, and were paid 30 cents for completing it. Participants were 63.6% females, *M* age = 38.17 (*SD* = 12.88). 37% of participants reported having a 4-year college degree, 21% completed some college, 14% had a two-year degree, 18% had pursued graduate school, and 9% had a high school degree or less.

United States 2

See Chang et al. (2015) and Yarnell & Neff (2013) for a full description of recruitment procedures for this student sample. Data were combined from two studies in order to increase sample size (N = 844), given that the SCS was developed in a US student sample. Participants were randomly selected from subject pools at two large southern American universities. Participants were 58.3% females, *M* age = 21.22 (*SD* = 3.53).

United States 3

See Neff et al. (2017) for a full description of recruitment procedures. Participants for this sample of meditators (*N* = 215) were recruited via an e-mail that invited them to complete an online questionnaire via Survey Monkey. E-mails were sent to individuals affiliated with Seattle Insight Meditation Society, Spirit Rock, the Insight Meditation Society, and similar groups. Participants reported a wide range in meditation experience from beginner to advanced (1 to 20 years of meditation practice). The average length of meditation practice for the sample was 6.67 years (*SD* = 3.86). Participants were 70% females, *M* age = 47.40 (*SD* = 12.88). Education information not available.

Appendix 3: SCS translation information

Brazilian Portuguese (Souza & Hutz, 2016)

Translation procedure: Two forward translations, two focus groups (research team and licensed psychologists), community and major undergraduate samples, two bilingual experts in psychometrics, back-translation by a bilingual Buddhist, followed by a final check by K. Neff, were the steps taken to ensure cultural validation.

Internal Consistency: Cronbach's $\alpha = .92$ for the total score, ranging from .66 to .81 for the subscales.

Test-retest reliability: N/A

Factor structures tested: Both single- and six-factor correlated CFA model displayed a good fit.

Chinese (Chen, Yan & Zhou, 2011)

Translation procedure: A standard forward-backward translation method was utilized.

Internal Consistency: Cronbach's $\alpha = .84$

Test-retest reliability: $r = .89$

Factor structures tested: A six-factor EFA and CFA model displayed a good fit.

French (Kotsou & Leys, 2016)

Translation procedure: A standard forward-backward translation method was utilized.

Internal Consistency: Cronbach's $\alpha = .94$

Test-retest reliability: $r = .85$

Factor structures tested: A higher-order one-factor, bi-factor, and a six-factor correlated model were calculated. The six-factor correlated CFA model displayed a good fit, while the higher-order one-factor model displayed a weaker fit and the bi-factor showing an acceptable fit.

German (Hupfeld & Ruffieux, 2011)

Translation procedure: A standard forward-backward translation method was utilized.

Internal Consistency: Cronbach's α = .91

Test-retest reliability: *r* = .92

Factor structures tested: A six-factor correlated ESEM model displayed a good fit.

Greek (Mantzios, Wilxon & Giannou, 2015)

Translation procedure: A standard forward-backward translation method was utilized.

Internal Consistency: Cronbach's α = .87

Test-retest reliability: *r* = .89

Factor structures tested: N/A

Italian (Petrocchi, Ottaviani & Couyoumdjian, 2013)

Translation procedure: A standard forward-backward translation method was utilized.

Internal Consistency: Cronbach's α = .90

Test-retest reliability: *r* = .85

Factor structures tested: A six-factor correlated CFA model displayed an adequate fit after removing two items. A one-factor, higher order one-factor model, and two-factor models were displayed a poor fit.

Japanese (Arimitsu, 2014)

Translation procedure: A standard forward-backward translation method was utilized.

Internal Consistency: Cronbach's α = .84

Test-retest reliability: *r* = .83

Factor structures tested: A one-factor, higher order one-factor model, higher two-factor, and a six-factor correlated model were calculated. In comparison, the six-factor correlated CFA model displayed the best fit.

Korean (Kim, Yi, Cho, Chai & Lee, 2008)

Translation procedure: A standard forward-backward translation method was utilized.

Internal Consistency: Cronbach's α = .90

Test-retest reliability: r = .85

Factor structures tested: A six-factor correlated CFA model displayed a good fit.

Norwegian (Dundas, Svendsen, Wiker, Granli & Schanche, 2016)

Translation procedure: A standard forward-backward translation method was utilized.

Internal Consistency: Cronbach's α = .89

Test-retest reliability: N/A

Factor structures tested: A one-factor, two-factor, three-factor, and a six-factor correlated model were calculated. In comparison, the six-factor CFA model displayed the best fit.

Persian (Ghorbani, Chen, Saeedi, Behjati, Watson, 2013)

Translation procedure: A standard forward-backward translation method was utilized, but psychometric analyses were not conducted. However, information for a similar Persian translation (Azizi, Mohammadkhani, Lotfi & Bahramkhani, 2013), reported below.

Internal Consistency: Cronbach's α = .78

Test-retest reliability: N/A

Factor structures tested: CFA found a marginally good fit for a six-factor correlated model.

Portuguese (Castilho & Pinto-Gouveia, 2011)

Translation procedure: A standard forward-backward translation method was utilized.

Internal Consistency: Cronbach's α = .89

Test-retest reliability: r = .78

Factor structures tested: Both single and six-factor correlated CFA model displayed a good fit.

Spanish (Garcia-Campayo, Navarro-Gil, Andrés, Montero-Marin, López-Artal, & Demarzo, 2014)

Translation procedure: A standard forward-backward translation method was utilized.

Internal Consistency: Cronbach's $\alpha = .87$

Test-retest reliability: ICC = .92

Factor structures tested: A six-factor CFA model displayed a good fit.

Appendix 4 Table S2

Goodness-of-Fit Indices for the Total and Individual Samples: One-Factor Models

	One-Factor CFA					One-Factor ESEM				
	CFI	TLI	RMSEA	90% CI	WRMR	CFI	TLI	RMSEA	90% CI	WRMR
Total	.74	.73	.15	.15-.15	14.44			N/A		
AUS	.86	.85	.13	.13-.14	2.11			N/A		
BRA	.88	.87	.11	.11-.12	1.83			N/A		
CAN	.76	.75	.14	.13-.14	2.97			N/A		
CHI	.93	.92	.14	.14-.15	2.10			N/A		
FRA	.80	.79	.15	.14-.15	4.67			N/A		
GER	.75	.73	.13	.12-.13	2.23			N/A		
GRE	.74	.72	.16	.15-.16	4.68			N/A		
IRA	.60	.56	.13	.12-.13	2.69			N/A		
ITA	.76	.74	.15	.15-.16	2.93			N/A		
JAP	.49	.45	.18	.18-.18	5.56			N/A		
KOR	.55	.52	.20	.20-.21	4.51			N/A		
NOR	.78	.77	.13	.13-.14	2.30			N/A		
POR 1	.76	.74	.17	.17-.17	5.89			N/A		
POR 2	.77	.75	.14	.13-.14	2.34			N/A		
SPA	.64	.61	.15	.15-.16	3.49			N/A		
UK 1	.78	.77	.14	.14-.14	3.98			N/A		
UK 2	.73	.70	.14	.13-.14	2.61			N/A		
US 1	.83	.81	.13	.13-.14	3.54			N/A		
US 2	.69	.66	.14	.14-.15	4.28			N/A		
US 3	.84	.83	.14.	.14-.15	1.94			N/A		

Note. AUS = Australia; BRA = Brazil; CAN = Canada; CHI = China; FRA = France; GER = Germany; GRE = Greece; IRA = Iran; ITA = Italy; JAP = Japan; KOR = South Korea; NOR = Norway; POR = Portugal; SPA = Spain; UK = United Kingdom; US = United States

Appendix 5: Syntax files for the examined models

Model 1a: One-factor CFA

```
! Commands preceded by ! sign are comments that Mplus ignores.

DATA:
   FILE IS C:\Users\scs.dat;
! Path to and name of data file changes per study.

VARIABLE:
   MISSING ARE ALL (9999);
   NAMES ARE
      scsj1 scoi2 scch3 scis4 scsk5 scoi6 scch7 scsj8 scmi9 scch10 scsj11
      scsk12 scis13 scmi14 scch15 scsj16 scmi17 scis18 scsk19 scoi20
      scsj21 scmi22 scsk23 scoi24 scis25 scsk26;
   USEVARIABLES ARE
      scsj1 scoi2 scch3 scis4 scsk5 scoi6 scch7 scsj8 scmi9
      scch10 scsj11 scsk12 scis13 scmi14 scch15 scsj16 scmi17
      scis18 scsk19 scoi20 scsj21 scmi22 scsk23 scoi24 scis25
      scsk26;
! Specifying that we're treating the variables as categorical.
   CATEGORICAL ARE all;

ANALYSIS:
! Requesting the weighted least squares mean- and variance-adjusted estimator
   estimator = wlsmv;

MODEL:
! Specifying the latent self-compassion factor with the 'BY' statement
sc BY
      scsk5* scsk12 scsk19 scsk23 scsk26
      scsj1 scsj8 scsj11 scsj16 scsj21
      scch3 scch7 scch10 scch15
      scis4 scis13 scis18 scis25
      scmi9 scmi14 scmi17 scmi22
      scoi2 scoi6 scoi20 scoi24;

      sc@1;

! Requesting standardized parameter estimates
OUTPUT: stdyx;
```

Model 1b: One-factor ESEM

```
MODEL:
  sc BY
    scsk5 scsk12 scsk19 scsk23 scsk26
    scsj1 scsj8 scsj11 scsj16 scsj21
    scch3 scch7 scch10 scch15
    scis4 scis13 scis18 scis25
    scmi9 scmi14 scmi17 scmi22
    scoi2 scoi6 scoi20 scoi24 (*1);
```

Model 2a: Two-factor CFA

```
! All syntax preceding this model are the same as in the previous CFA models unless indicated
! otherwise.
MODEL:
  pos BY scsk5* scsk12 scsk19 scsk23 scsk26
      scch3 scch7 scch10 scch15
      scmi9 scmi14 scmi17 scmi22;
  neg BY scsj1* scsj8 scsj11 scsj16 scsj21
      scis4 scis13 scis18 scis25
      scoi2 scoi6 scoi20 scoi24;

  pos@1; neg@1;
```

Model 2b: Two-factor ESEM

```
ANALYSIS:
  estimator = wlsmv;
  rotation = target;
! Target rotation was used in all models in conjunction with the (~) sign for all ESEM models.
! Cross-loadings are targeted to be as close to zero as possible

MODEL:
  pos BY scsk5 scsk12 scsk19 scsk23 scsk26
      scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
      scch3 scch7 scch10 scch15
      scis4~0 scis13~0 scis18~0 scis25~0
      scmi9 scmi14 scmi17 scmi22
      scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
  neg BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
      scsj1 scsj8 scsj11 scsj16 scsj21
      scch3~0 scch7~0 scch10~0 scch15~0
      scis4 scis13 scis18 scis25
      scmi9~0 scmi14~0 scmi17~0 scmi22~0
      scoi2 scoi6 scoi20 scoi24 (*1);
```

Model 3a: Six-factor CFA

```
! All syntax preceding this model are the same as in the previous CFA models unless indicated
! otherwise.
MODEL:
! self-kindness
   sk BY scsk5* scsk12 scsk19 scsk23 scsk26;
! self-judgment
   sj BY scsj1* scsj8 scsj11 scsj16 scsj21;
! common humanity
   ch BY scch3* scch7 scch10 scch15;
! isolation
   is BY scis4* scis13 scis18 scis25;
! mindfulness
   mi BY scmi9* scmi14 scmi17 scmi22;
! overidentification
   oi BY scoi2* scoi6 scoi20 scoi24;

   sj@1; oi@1; ch@1; sk@1; mi@1; is@1;
```

Model 3b: Six-factor ESEM

```
! All syntax preceding this model are the same as in the previous ESEM models unless indicated
! otherwise.
MODEL:
   sk BY scsk5 scsk12 scsk19 scsk23 scsk26
        scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
        scch3~0 scch7~0 scch10~0 scch15~0
        scis4~0 scis13~0 scis18~0 scis25~0
        scmi9~0 scmi14~0 scmi17~0 scmi22~0
        scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
   sj BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
        scsj1 scsj8 scsj11 scsj16 scsj21
        scch3~0 scch7~0 scch10~0 scch15~0
        scis4~0 scis13~0 scis18~0 scis25~0
        scmi9~0 scmi14~0 scmi17~0 scmi22~0
        scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
   ch BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
        scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
        scch3 scch7 scch10 scch15
        scis4~0 scis13~0 scis18~0 scis25~0
        scmi9~0 scmi14~0 scmi17~0 scmi22~0
        scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
   is BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
        scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
        scch3~0 scch7~0 scch10~0 scch15~0
        scis4 scis13 scis18 scis25
```

```
        scmi9~0 scmi14~0 scmi17~0 scmi22~0
        scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
   mi BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
        scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
        scch3~0 scch7~0 scch10~0 scch15~0
        scis4~0 scis13~0 scis18~0 scis25~0
        scmi9 scmi14 scmi17 scmi22
        scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
   oi BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
        scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
        scch3~0 scch7~0 scch10~0 scch15~0
        scis4~0 scis13~0 scis18~0 scis25~0
        scmi9~0 scmi14~0 scmi17~0 scmi22~0
        scoi2 scoi6 scoi20 scoi24 (*1);
```

Model 4a: Bifactor-CFA (1 G- and 6 S-factors)

```
! All syntax preceding this model are the same as in the previous CFA models unless indicated
! otherwise.
MODEL:
   sc BY scsk5* scsk12 scsk19 scsk23 scsk26 scsj1 scsj8 scsj11 scsj16
        scsj21 scch3 scch7 scch10 scch15 scis4 scis13 scis18 scis25
        scmi9 scmi14 scmi17 scmi22 scoi2 scoi6 scoi20 scoi24;

   sk BY scsk5* scsk12 scsk19 scsk23 scsk26;
   sj BY scsj1* scsj8 scsj11 scsj16 scsj21;
   ch BY scch3* scch7 scch10 scch15;
   is BY scis4* scis13 scis18 scis25;
   mi BY scmi9* scmi14 scmi17 scmi22;
   oi BY scoi2* scoi6 scoi20 scoi24;

   sc@1; sj@1; oi@1; ch@1; sk@1; mi@1; is@1;

   sc WITH sk-oi@0;
   sk WITH sj-oi@0;
   sj WITH ch-oi@0;
   ch WITH is-oi@0;
   is WITH mi-oi@0;
   mi WITH oi@0;
```

Model 4b: Bifactor-ESEM (1 G- and 6 S-factors)

```
! All syntax preceding this model are the same as in the previous ESEM models unless indicated
! otherwise.
ANALYSIS:
   estimator = wlsmv;
```

```
   rotation = target (orthogonal);
! Factors are specified as orthogonal to each other.
MODEL:
   sc BY scsk5 scsk12 scsk19 scsk23 scsk26 scsj1 scsj8 scsj11 scsj16
       scsj21 scch3 scch7 scch10 scch15 scis4 scis13 scis18 scis25
       scmi9 scmi14 scmi17 scmi22 scoi2 scoi6 scoi20 scoi24 (*1);
MODEL:
   sk BY scsk5 scsk12 scsk19 scsk23 scsk26
       scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
       scch3~0 scch7~0 scch10~0 scch15~0
       scis4~0 scis13~0 scis18~0 scis25~0
       scmi9~0 scmi14~0 scmi17~0 scmi22~0
       scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
   sj BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
       scsj1 scsj8 scsj11 scsj16 scsj21
       scch3~0 scch7~0 scch10~0 scch15~0
       scis4~0 scis13~0 scis18~0 scis25~0
       scmi9~0 scmi14~0 scmi17~0 scmi22~0
       scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
   ch BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
       scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
       scch3 scch7 scch10 scch15
       scis4~0 scis13~0 scis18~0 scis25~0
       scmi9~0 scmi14~0 scmi17~0 scmi22~0
       scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
   is BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
       scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
       scch3~0 scch7~0 scch10~0 scch15~0
       scis4 scis13 scis18 scis25
       scmi9~0 scmi14~0 scmi17~0 scmi22~0
       scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
   mi BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
       scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
       scch3~0 scch7~0 scch10~0 scch15~0
       scis4~0 scis13~0 scis18~0 scis25~0
       scmi9 scmi14 scmi17 scmi22
       scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
   oi BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
       scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
       scch3~0 scch7~0 scch10~0 scch15~0
       scis4~0 scis13~0 scis18~0 scis25~0
       scmi9~0 scmi14~0 scmi17~0 scmi22~0
       scoi2 scoi6 scoi20 scoi24 (*1);
```

Model 5a: Two-bifactor (two-tier) CFA model (2 G- and 6 S-factors)

```
! All syntax preceding this model are the same as in the previous CFA models unless indicated
! otherwise.
! positive items
 po BY scsk5* scsk12 scsk19 scsk23 scsk26
      scch3 scch7 scch10 scch15
      scmi9 scmi14 scmi17 scmi22;
! negative items
  ne BY scsj1* scsj8 scsj11 scsj16 scsj21
      scis4 scis13 scis18 scis25
      scoi2 scoi6 scoi20 scoi24;

  sk BY scsk5* scsk12 scsk19 scsk23 scsk26;
  sj BY scsj1* scsj8 scsj11 scsj16 scsj21;
  ch BY scch3* scch7 scch10 scch15;
  is BY scis4* scis13 scis18 scis25;
  mi BY scmi9* scmi14 scmi17 scmi22;
  oi BY scoi2* scoi6 scoi20 scoi24;

  po@1; ne@1; sj@1; oi@1; ch@1; sk@1; mi@1; is@1;
! general factors are allowed to correlate with each other
  po WITH sk-oi@0;
  ne WITH sk-oi@0;
  sk WITH sj-oi@0;
  sj WITH ch-oi@0;
  ch WITH is-oi@0;
  is WITH mi-oi@0;
  mi WITH oi@0;
```

Model 5b: Two-bifactor (two-tier) ESEM model (2 G- and 6 S-factors)

```
! All syntax preceding this model are the same as in the previous ESEM models unless indicated
! otherwise.
ANALYSIS:
  estimator = wlsmv;
  rotation = target (orthogonal);

MODEL:
  sk BY scsk5 scsk12 scsk19 scsk23 scsk26
      scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
      scch3~0 scch7~0 scch10~0 scch15~0
      scis4~0 scis13~0 scis18~0 scis25~0
      scmi9~0 scmi14~0 scmi17~0 scmi22~0
      scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
  sj BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
      scsj1 scsj8 scsj11 scsj16 scsj21
```

```
      scch3~0 scch7~0 scch10~0 scch15~0
      scis4~0 scis13~0 scis18~0 scis25~0
      scmi9~0 scmi14~0 scmi17~0 scmi22~0
      scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
ch BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
      scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
      scch3 scch7 scch10 scch15
      scis4~0 scis13~0 scis18~0 scis25~0
      scmi9~0 scmi14~0 scmi17~0 scmi22~0
      scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
is BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
      scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
      scch3~0 scch7~0 scch10~0 scch15~0
      scis4 scis13 scis18 scis25
      scmi9~0 scmi14~0 scmi17~0 scmi22~0
      scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
mi BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
      scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
      scch3~0 scch7~0 scch10~0 scch15~0
      scis4~0 scis13~0 scis18~0 scis25~0
      scmi9 scmi14 scmi17 scmi22
      scoi2~0 scoi6~0 scoi20~0 scoi24~0 (*1);
oi BY scsk5~0 scsk12~0 scsk19~0 scsk23~0 scsk26~0
      scsj1~0 scsj8~0 scsj11~0 scsj16~0 scsj21~0
      scch3~0 scch7~0 scch10~0 scch15~0
      scis4~0 scis13~0 scis18~0 scis25~0
      scmi9~0 scmi14~0 scmi17~0 scmi22~0
      scoi2 scoi6 scoi20 scoi24 (*1);

po BY scsk5* scsk12 scsk19 scsk23 scsk26
      scch3 scch7 scch10 scch15
      scmi9 scmi14 scmi17 scmi22;
ne BY scsj1* scsj8 scsj11 scsj16 scsj21
      scis4 scis13 scis18 scis25
      scoi2 scoi6 scoi20 scoi24;

po@1; ne@1;

po WITH sk-oi@0;
ne WITH sk-oi@0;
```

Supplementary Material References

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

Arimitsu, K., Aoki, Y., Furukita, M., Tada, A., & Togashi, R. (2016). Construction and Validation of a Short Form of the Japanese version of the Self-Compassion Scale. *Komazawa Annual Reports of Psychology, 18,* 1-8.

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.

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

Brenner, R. E., Heath, P. J., Vogel, D. L., & Credé, M. (2017). Two is more valid than one: Examining the factor structure of the Self-Compassion Scale (SCS). *Journal of counseling psychology*, *64*(6), 696-707.

Castilho, P., & Pinto-Gouveia, J. (2011). Self-Compassion: Validation of the Portuguese version of the Self-Compassion Scale and its relation with early negative experiences, social comparison and psychopathology. *Psychologica, 54,* 203-231.

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.

Chang, E. C., Yu, T., Jilani, Z., Fowler, E. E., Yu, E. A., Lin, J., & Hirsch, J. K. (2015). Under Assault: Understanding The Impact Of Sexual Assault On The Relation Between Hope And Suicidal Risk In College Students. *Journal of Social and Clinical Psychology*, *34*(3), 221-238.

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.

Cleare, S., Gumley, A., Cleare, C. J., & O'Connor, R. C. (2018). An Investigation of the Factor Structure of the Self-Compassion Scale. *Mindfulness*, *9*(2), 618-628.

Coroiu, A., Kwakkenbos, L., Moran, C., Thombs, B., Albani, C., Bourkas, S., ... & Körner, A. (2018). Structural validation of the Self-Compassion Scale with a German general population sample. *PloS one*, *13*(2), e0190771.

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.

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.

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.

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*, 58–72.

Garcia-Campayo, J., Navarro-Gil, M., Andrés, E., Montero-Marin, J., López-Artal, L., & Demarzo, M. M. (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*(1), 4.

Ghorbani, N., Chen, Z., Saeedi, Z., Behjati, Z.,Watson,P.J. (2013). Sakhtare Ameli Meghyase

Shafeghat e Khod dar Iran (Factorial Structure of Self-Compassion Scale in Iran) *Pazhohesh Haye Karbordi Ravanshenakhti, 4*(3), 29-41. (In Persian)

Halamová, J., Kanovský, M., & Pacúchová, M. (2018). Self-Compassion Scale: IRT Psychometric Analysis, Validation, and Factor Structure–Slovak Translation. *Psychologica Belgica*, *57*(4), 190-209.

Hupfeld, J., & Ruffieux, N. (2011). Validation of a German version of the self-compassion scale (SCS-D). *Zeitschrift für Klinische Psychologie und Psychotherapie, 40*(2), 115–123.

Karakasidou, E., Pezirkianidis, C., Galanakis, M., & Stalikas, A. (2017). Validity, Reliability and Factorial Structure of the Self Compassion Scale in the Greek Population. *Journal of Psychology and Psychotherapy*, *7*, 313.

Kim, K. E., Yi, G. D., Cho, Y. R., Chai, S. H., & Lee, W. K. (2008). The validation study of the Korean version of the self-compassion scale. *Korean Journal of Health Psychology, 13*, 1023–1044.

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

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

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, 123-132.

Montero-Marin, J., Zubiaga, F., Cereceda, M., Demarzo, M. M. P., Trenc, P., & Garcia-Campayo, J. (2016). Burnout subtypes and absence of self-compassion in primary healthcare professionals: A cross-sectional study. *PLoS One, 11*(6), e0157499

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

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.

Neff, K., Tóth-Király, I., & Colosimo, K. (in press). Self-compassion is best measured as a global construct and is overlapping with but distinct from neuroticism: A response to Pfattheicher, Geiger, Hartung, Weiss, and Schindler (2017). *European Journal of Psychology.* Early view doi: 10.1002/per.2148

Neff, K. D., Whittaker, T. & Karl, A. (2017). Evaluating the factor structure of the Self-Compassion Scale in four distinct populations: Is the use of a total self-compassion score justified? *Journal of Personality Assessment, 1,* 1-12.

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.

Pfattheicher, S., Geiger, M., Hartung, J., Weiss, S., & Schindler, S. (2017). Old Wine in New Bottles? The Case of Self-compassion and Neuroticism. *European Journal of Personality*, *31*, 160-169.

Sirois, F. M., Kitner, R., & Hirsch, J. K. (2015). Self-compassion, affect, and health-promoting behaviors. *Health Psychology, 34,* 661-669.

Souza, L. K. de, Ávila-Souza, J., & Gauer, G. (2016). Escala de Autocompaixão. In: C. S. Hutz (Ed.), *Avaliação em Psicologia Positiva* (pp. 169-177). São Paulo: Hogrefe/CETEPP.

Souza, L. K. D., & 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.

Tóth-Király, I., Bőthe, B., & Orosz, G. (2017). Exploratory Structural Equation Modeling Analysis of the Self-Compassion Scale. *Mindfulness*, *8*, 881-892.

Veneziani, C. A., Fuochi, G., & Voci, A. (2017). Self-compassion as a healthy attitude toward the self: Factorial and construct validity in an Italian sample. *Personality and Individual Differences*, *119*, 60-68.

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.

Woo Kyeong, L. (2013). Self-compassion as a moderator of the relationship between academic burn-out and psychological health in Korean cyber university students. *Personality and Individual Differences*, *54*(8), 899-902.

Yarnell, L. M., Neff, K. D. (2013). Self-compassion, interpersonal conflict resolutions, and well-being. *Self and Identity.* 2:2, 146-159.

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.