Psychometric Properties of the Body Checking Questionnaire (BCQ) and of the Body Checking Cognitions Scale (BCCS): A Bifactor-Exploratory Structural Equation Modeling approach

Christophe Maïano[a,c*], Alexandre J. S. Morin[b*], Annie Aimé [a,c], Geneviève Lepage[c], Stéphane Bouchard[a,c]

[a]Cyberpsychology Laboratory, Department of Psychoeducation and Psychology, Université du Québec en Outaouais (UQO), Saint-Jérôme, Canada.
[b] Substantive-Methodological Synergy Research Laboratory, Department of Psychology, Concordia University, Montreal, Canada.
[c]Department of Psychoeducation and Psychology, Université du Québec en Outaouais (UQO), Saint-Jérôme, Canada.

*The order of appearance of the first and second authors (C.M. and A.J.S.M) was determined at random: both should be considered first authors.

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Correspondence concerning this article should be addressed to: Christophe Maïano, Université du Québec en Outaouais, Campus de Saint-Jérôme, Département de Psychoéducation et de Psychologie, 5 rue Saint-Joseph, Saint-Jérôme (Québec) J7Z 0B7, email: christophe.maiano@uqo.ca

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Abstract

This research sought to assess the psychometric properties of the French versions of the Body Checking Questionnaire (BCQ) and the Body Checking Cognitions Scale (BCCS) among community samples. A total sample of 922 adolescents and adults was involved in a series of two studies. The results from the first study supported factor validity and reliability of responses obtained on these two measures, and showed that both measures were best represented by a bifactor-exploratory structural equation modeling (ESEM) representation of the data. The results from the second study replicated these conclusions, while also supporting the measurement invariance of the bifactor-ESEM solution and the equivalence of the correlations among the two measures (i.e., convergent validity) across samples. This second study also supported the criterion related validity of ratings on both measures with measures of global self-esteem, physical appearance, social physique anxiety, fear of negative appearance evaluation, and disturbed eating attitudes and behaviors. Finally, the results of this last study also supported the measurement invariance and lack of differential item functioning of both measures in relation to sex, age, diagnosis of eating disorders, and body mass index.

Keywords: Age, body mass index, differential item functioning, eating disorders, measurement invariance, multiple indicators multiple causes models, sex.

People struggling with disturbed eating attitudes and behaviors (DEAB) or a diagnosis of eating disorders (ED) tend to adopt a lifestyle allowing them to cope with their body image dissatisfaction (Reas, Whisenhunt, Netemeyer, & Williamson, 2002). To this end, some tend to check the size and weight of their overall body or specific parts of it in a ritualistically or compulsory manner (Mountford, Haase, & Waller, 2006; Reas et al. 2002). Body checking behaviors encompass, for example, repeated weighing using special clothes or jewelry (e.g., ring, watch) to estimate weight or shape, pinching specific body parts (e.g., cheeks, stomach, upper arms), examining oneself in the mirror, and scrutinizing fatness in specific body parts (Mountford et al., 2006; Reas et al. 2002).

Checking behaviors are often used as a way to prevent or reduce negative emotions, such as anxiety, feeling fat, depression, or loneliness (Fairburn, Cooper, & Shafran, 2003; Reas et al. 2002; Williamson, White, York-Crowe, & Stewart, 2004). Over time, these behaviors can become counterproductive and create a vicious circle whereby increasing attention toward one's body tends to magnify perceived body imperfections, in turn encouraging further body monitoring/checking and reinforcing body dissatisfaction (Fairburn et al., 2003; Reas et al. 2002; Williamson, White et al., 2004). Body checking behaviors tend to be reinforced by dysfunctional cognitions (Mountford et al., 2006) regarding the function of this type of behaviors (e.g., "*Body checking stops me from losing control of what I eat*").

Two instruments are available to assess either body checking behaviors or dysfunctional cognitions that may reinforce these behaviors: the Body Checking Questionnaire (BCQ; Reas et al., 2002) and the Body Checking Cognitions Scale (BCCS; Mountford et al., 2006). The BCQ includes 23 items on which participants are asked to indicate how often they currently engage in these behaviors using a five-point scale ranging from *never* (1) to *very often* (5). This instrument was validated by Reas et al. (2002) in two distinct community (n = 1,259) and clinical (diagnosed with anorexia nervosa, bulimia nervosa, and an ED not otherwise specified; n = 165) samples of female adults from the United States of America. The results from a principal component analysis (PCA) with oblique rotation provided a satisfactory solution encompassing three correlated factors: (a) overall appearance (OVAPP, 10 items), (b) specific body parts (SBP, 8 items), and (c) idiosyncratic checking (IDSC, 5 items). Additional results from a higher-order confirmatory factor analysis (CFA) performed among independent samples provided support for the first order factors (OVAPP, SBP, IDSC) and for a higher body-checking factor. The results also supported the convergent validity of the global score of the BCQ with the Body Shape Questionnaire (r = .86), the Eating Attitudes Test-26 (r = .70), and the Body Image Avoidance Questionnaire (r = .66). The criterion validity of the BCQ was finally supported by demonstrating its ability to differentiate various samples (i.e., clinical vs. community; very concerned vs. less concerned by their body shape; dieters vs. non-dieters).

The BCCS includes 19 items on which participants are asked to indicate how often each statement applies to them using a five-point scale ranging from *never* (1) to *very often* (5). This instrument was validated by Mountford et al. (2006) in two distinct community (n = 289) and clinical (diagnosed with anorexia nervosa, bulimia nervosa, and an ED not otherwise specified; n = 130) samples of female adults from the United Kingdom. The results from a PCA with oblique rotation provided a satisfactory solution encompassing four correlated factors: (a) objective evaluation (OBEVAL, 6 items), (b) reassurance (REASS, 4 items), (c) safety beliefs (SAFEB, 5 items), and (d) body control (BODC, 4 items). Additionally, a CFA performed among independent samples provided support for this four-factor structure (OBEVAL, REASS, SAFEB, BODC). The results also supported the convergent validity of the global score of the BCCS with the BCQ factors and with factors from the Eating Disorders Examination-Questionnaire (β = -.35 to .43). However, the validity of this more global composite score has yet to be established. Finally, the criterion validity of the BCCS was also supported based on a comparison of clinical and non-clinical samples.

Validation in Other Languages

The BCQ has been validated in other languages, such as German (Steinfiel et al., 2017; Vocks, Moswald, & Legenbauer, 2008), Norwegian (Reas, Von Soest, & Lask, 2009), Portuguese (Campana, Swami, Onodera, da Silva, & Fernandes, 2013) and Italian (Calugi, Dalle Grave, Ghisi, & Sanavio, 2006). The results from these studies replicated the higher-order CFA structure of the BCQ in samples of adolescents and/or adults. Additionally, several studies also supported the convergent validity of the BCQ factors with the BCCS factors (Haase, Mountford, & Waller, 2007; Mountford et al., 2006), and with measures of ED or disturbed eating attitudes and behaviors (Calugi et al., 2006; Mountford, Haase, & Waller, 2007; Reas, White, & Grilo, 2006; Reas et al., 2009; Steinfiel et al., 2017; Vocks et al., 2008), body uneasiness (Calugi et al., 2006), body image avoidance (Calugi et al., 2006; Campana et al., 2013), social anxiety (Campana et al., 2013), social physique anxiety (Haase et al., 2007), depression (Steinfiel et al., 2017), and self-esteem (Steinfiel et al., 2017).

In contrast, the BCCS has so far only been validated in Portuguese by Kachani et al. (2013).The results obtained by these authors provided support for the original four CFA structure of the BCCS in a sample of adult females. Additional studies have also supported the convergent validity of the BCCS scores with the BCQ (Mountford et al., 2007), measures of DEAB or ED (Kachani et al., 2013; Mountford et al., 2007), and body shape (Kachani et al., 2013).

Currently, no self-report questionnaires are available to assess either body checking behaviors or dysfunctional cognitions among French-speaking populations. Thus, information about these behaviors or the dysfunctional cognitions associated with them is currently lacking among French-speaking populations, and cross-cultural studies are impossible to conduct in a reliable and valid manner. The present study was designed to contribute to this body of research by proposing, and validating, a French version of these two instruments. The value of such a French version comes from the fact that French is one of the main first or second language spoken in Canada, across Europe and Africa, as well as one of the official languages of international institutions such as the Olympics or UNESCO. For practical purposes, the validation of a French-language version of the BCQ and BCCS will fill an important gap for researchers interested in assessing these constructs among French-speaking populations. For research and theory development, the validation of a French-language version of the BCQ and BCCS goes well beyond verifying the simple applicability of these instruments to French-speaking populations, but also provides an instrument on which to anchor cross-cultural comparisons necessary to the realization of large scale epidemiological or longitudinal studies (Tomás, Marsh, González-Romá, Valls, & Nagengast, 2014).

A BifactorRepresentation of the BCQ and BCCS

A key question that remains concerning the factor structure of both of these questionnaires is related to whether or not the specific dimensions covered in each of them can really be combined into a global score reflecting either body checking behaviors (BCQ) or dysfunctional cognitions (BCCS). Although there is some evidence that the specific dimensions of the BCQ can be combined into a global score (e.g., Calugi et al., 2006; Campana et al., 2013; Reas et al., 2002, 2009; Steinfiel et al., 2017; Vocks et al., 2008), this remains to be empirically verified for the BCCS. However, when considering this question, it is important to keep in mind that the first-order factor correlations observed among the specific dimensions assessed by these instruments are quite high. Indeed, correlations observed among the BCQ subscales range between .57 and .87 (Campana et al., 2013; Haase et al., 2007; Pellizzer, Tiggemann, Waller, & Wade, 2018; Reas et al., 2002), and those observed between the BCCS subscales range between .43 and .71 (Haase et al., 2007), thereby compromising the added value of these dimensions over and above that of the global construct itself. Clarifying the factorial structure of each instrument would inform researchers and clinicians about whether it is more appropriate to calculate a global score, to consider only the specific subscales, or to take both global and specific scores into account for the interpretation of the BCQ and BCCS.

The above considerations are expected to occur when, as is the case for the BCQ and BCCS, conceptually-related constructs are assessed within the same instrument. Morin and colleagues (Morin, Arens, & Marsh, 2016; Morin, Myers, & Lee, 2019) refer to this issue as construct-relevant psychometric multidimensionality. A first source of such multidimensionality relates to the assessment of coexisting global and specific constructs. When considering whether body checking behaviors or cognitions are best represented as global constructs or conceptually-distinct subscales, there is another possibility to consider, which is that they might exist as global entities coexisting with subscale specificities left unexplained by the global construct. Psychometrically, two approaches can be used to study this possibility. A first approach relies on higher-order factor models, where ratings are used to define first-order factors, themselves used to define a higher-order factor. However, hierarchical models rely on a stringent proportionality constraint according to which the ratio of variance explained by the global relative to the specific factors is exactly the same for all items associated with a first-order factor (for a more extensive discussion of this limitation of higher-order models, see Gignac, 2016). Bifactor models provide a more flexible alternative (Reise, 2012). In a bifactor model, one Global (G) factor reflects the global construct underlying all items and a series of orthogonal Specific

(S) factors explain the covariance among a set of items left explained by this G-Factor.

When considering a bifactor solution, it is important to keep in mind that the meaning of the S-factors is drastically different from the meaning of first-order factors estimated in a typical measurement model. Indeed, the latter reflect everything that is shared among a subset of items forming a subscale, and as such can directly be interpreted as providing a reflection of the construct covered in the subscale (e.g., SBP). By contrast, in a bifactor model, the S-factors reflect what remains, at the subscale level, once the global construct assessed by the whole questionnaire is taken into account. For instance, in the BCQ, the G-factor would provide a direct reflection of participants' global levels of body checking behaviors. However, their scores on the SPB S-factor would reflect the presence of a tendency to body check specific body parts to a greater or lower extent than their global tendencies to body check across dimensions. Sometimes, bifactor models might even result in the estimation of empty or weakly defined S-factors, reflecting the fact that participants levels on specific subscales seldom deviate from their global level across subscales. As such, the S-factors provide a direct test of specificity at the subscale level.

A second source of construct-relevant psychometric multidimensionality likely to be present relates to the fact that items tend to present some degree of true score association with non-target constructs (Morin et al., 2016, 2018). This can be examined through exploratory factor analyses (EFA) by allowing free estimation of cross-loadings between items and conceptually-related constructs. EFA has recently been integrated with CFA and structural equation modeling into the exploratory structural equation modeling (ESEM) framework (Morin, Marsh, & Nagengast, 2013), making it possible to consider that items may present some degree of valid association with conceptually-related constructs (Morin et al., 2016, 2018). In particular, statistical research (for a review, see Asparouhov et al., 2015) shows that excluding cross-loadings as small as .100 tends to result in inflated estimates of factor correlations in CFA, and of the G-factor in bifactor-CFA, whereas unnecessary cross-loadings do not result in estimation biases. For these reasons, the present research relies on the bifactor-ESEM framework as a way to achieve a more accurate representation of BCQ and BCCS ratings.

Measurement Invariance and Latent Means Differences of the BCQ and BCCS

An important issue in the validation of any psychological measure relates to whether it can be used to compare individuals differing from one another on specific characteristics or attributes (e.g., age, body mass index, diagnosis, sex) in an unbiased, reliable and valid manner (Millsap, 2011). Such biases would occur when the factor structure of a questionnaire (i.e., configural invariance) differs across distinct subpopulations, but also when the degree of associations between items and latent constructs (i.e., factor loadings: weak invariance) differs across subpopulations, when average item ratings differ across subpopulations, when individuals have similar scores on the latent constructs (i.e., items intercepts: strong invariance), or when measurement errors differ across subpopulations (i.e., items uniquenesses: strict invariance). Whereas weak invariance is required to compare latent relations obtained in the different subpopulations and strong invariance is required to compare latent means, any comparison based on observed scale scores requires strict invariance.

Unfortunately, none of the original validation studies, nor any of the cross-linguistic validation studies, have reported evidence of measurement invariance for scores obtained on either the BCQ or the BCCS. Therefore, researchers or clinicians screening for body checking among participants with different characteristics cannot presently be assured that the BCQ and BCCS are psychometrically equivalent across different subpopulations. This lack of evidence of measurement invariance is surprising given that research generally shows statistically significant mean differences in BCQ and BCCS scores as a function of age, body mass index (BMI), diagnosis of ED, and sex. More specifically, studies using samples of women (Haase, Mountford, & Waller, 2011) or mixed-sex samples (Reas, White, & Grilo, 2006) have demonstrated that younger participants tend to have significantly higher scores on both the BCQ (global score and OVAPP and SBP subscales) and BCCS (all scales).Additional results from the same studies showed that participants with higher BMI tended to have significantly higher scores on the BCQ (global score and all subscales) and BCCS (REASS subscale). Moreover, other studies using women's samples have reported that participants with ED tend to have significantly higher scores on both the BCQ (global score and all subscales; Calugi et al., 2006; Mountford et al., 2006; Waller et al., 2008) and BCCS (all subscales: Kachani et al., 2013; Mountford et al., 2006; Waller et al., 2008). However, without evidence that these measures can operate without biases across various subpopulations, it remains impossible to clearly ascertain their validity.

So far, possible differential item functioning (DIF) on the BCQ has only been examined, in relation to sex, in a single study. In this study, Alfano Hildebrandt, Bannon, Walker and Walton (2011) hypothesized that, because they differ from one another in terms of body image, appearance concerns, and gendered socialization, men and women were likely to rely on distinct types of body checking behaviors and motives. Their results, limited to 10 items (nine from the OVAPP subscale and one from the IDSC subscale) from the BCQ, suggested that half of these items appeared to present some degree of DIF between men and women. For example, among women, the factor loadings appeared to be higher, and the response threshold lower on items focusing on the verification of whether ones' bones touched the floor while lying down, or checking one's bottom in the mirror (Alfano et al., 2011). Conversely, the factor loadings of items measuring wearing special clothing to assess the size of one's body, checking one's reflection in a reflecting surface, or comparing oneself to others seemed to be higher among men when compared to women (Alfano et al., 2011). However, this preliminary evidence came from the examination of unconstrained parameter estimates obtained in a model of configural invariance, rather than from systematic tests of measurement invariance, reinforcing the need to further investigate this possibility using proper methodologies. Yet, even though limited, this preliminary evidence suggests that body checking behaviors and cognitions are not immune from DIF. Therefore, these measures should not be used for comparing participants differing in terms of sex, age, BMI, diagnosis of ED, or any other characteristic likely to affect body shape, perceptions, cognitions, or behaviors, at least not before such evidence can be documented.

Overview of the Studies

The main objective of the present series of studies was to examine the psychometric properties of the French versions of the BCQ and BCCS among several community samples of French-speaking Canadian adolescents and adults. More specifically, the objective of Study 1 was to examine the factor validity and reliability of these two questionnaires using bifactor-ESEM. Study 2was then conducted to verify whether the best model identified for the BCQ and BCCS in study 1would replicate in a new independent community sample of adolescents and adults, and to assess the criterion-related validity of the best model identified for the BCQ and BCCS when considering measures of global self-esteem, physical appearance, social physique anxiety, fear of negative appearance evaluation, and disturbed eating attitudes and behaviors. The convergent validity of the BCQ and BCCS factors, and their equivalence across samples, were assessed using Samples from Studies 1 and 2. Finally, participants from Studies 1 and 2 were aggregated to obtain a larger sample in order to assess the measurement invariance or DIF of BCQ and BCCS ratings as a function of age, BMI, diagnosis of ED, and sex. All studies have been approved by the ethics committee of the Université du Québec en Outaouais.

Study 1: Factor Validity and Reliability

Method

Sample and Procedure. A convenience sample of 488 (411females and 77 males) French-speaking Canadian adolescents and adults, aged between 14 and 62 years old (M_{age} = 27 years, SD = 8.8), participated in this study. The participants were invited to participate via generic announcements sent via school messaging (letters or emails), in secondary schools, colleges and universities located in the Canadian Province of Quebec. Volunteers completed online informed consent form, and anonymously completed the questionnaires online via the LimeSurvey platform. The investigators and research team played no active role in selecting, or targeting, specific subpopulations of respondents, leading us to assume that the greater representation of females in the sample stems from their possible greater level of interest for this research thematic.

Measures. Participants were asked to self-report their age, sex, height, weight, and whether they had already received or not a diagnosis for ED. Their self-reported height and weight were then used to estimate their BMI. They also completed the French version of the BCQ and BCCS developed in a preliminary study (details on the translation process and the content validity of the translation are reported in section S1 in the online supplements), using the original response scale.

Analyses. Analyses were conducted using Mplus 7.4 (Muthén & Muthén, 2015) robust weighted least squares (WLSMV) estimator. To account for the limited level of missing data present at the item level (BCQ: M = 1.8%; BCCS: M = 2.9%), models were estimated using the full available information, based on algorithms implemented in Mplus for WLSMV estimation (Muthén & Muthén, 2015). First, the *a priori* original factor structure of the BCQ and BCCS were examined separately with CFA and ESEM. In CFA, it was hypothesized that participants' answers to the BCQ and BCCS would be explained, respectively, by three (OVAPP, SBP, IDSC) and four (OBEVAL, REASS, SAFEB, BODC) correlated factors, that each item would have a non-zero loading on the factor it was designed to measure and zero loadings on all other factors, and that uniquenesses would be uncorrelated. The *a priori* ESEM model was estimated using confirmatory target rotation (Asparouhov & Muthén, 2009; Browne, 2001) in which it was hypothesized that BCQ and BCCS would be explained, respectively, by three (OVAPP, SBP, IDSC)and four (OBEVAL, REASS, SAFEB, BODC)correlated factors, with all cross loadings "targeted" to be as close to zero as possible. In the second stage, bifactor-CFA and bifactor-ESEM representations of the BCQ and BCCS were examined separately. The bifactor solutions include one more factor than their CFA or ESEM counterparts. Indeed, in these solutions, all items were specified as having a main loading on both a global factor (G-factor) and on their *a priori* specific factors (BCQ: OVAPP, SBP, IDSC; BCCS: OBEVAL, REASS, SAFEB, BODC). As in typical bifactor representations (Morin et al., 2016; Reise, 2012), all factors were specified as orthogonal. In the bifactor-ESEM model, cross-loadings between the S-factors were targeted to be as close to zero as in the ESEM solution using anorthogonal bifactor target rotation (Reise, 2012). For all models, the composite reliability of the BCQ and BCCS factors was estimated using McDonald's (1970) omega (ω).

The goodness-of-fit of these models was assessed using a variety of fit indices (e.g., Marsh, Hau, & Grayson, 2005; Yu, 2002): the comparative fit index (CFI $\geq$ .90 or >. 95), the Tucker-Lewis index (TLI $\geq$ .90 or >. 95), the root mean square error of approximation (RMSEA $\leq$. 08 or <. 05), and the 90% confidence interval of the RMSEA. Changes in these fit indices were also used in the assessment of nested model comparisons, with changes in CFIs $\leq$ .01, TLIs $\leq$ .01 and RMSEAs $\leq$. 015 taken as being indicative of substantial change in model fit (Chen, 2007; Cheung & Rensvold, 2002). For purpose of complete disclosure, we also report the WLSMV chi-square test of exact fit (χ^2) and WLSMV chi-square difference tests ($\Delta W\chi^2$calculated using the Mplus DIFFTEST function). However, because of the well-established oversensitivity of these tests to sample size and minor misspecifications (e.g., Hu & Bentler, 1999; Marsh, Hau, & Grayson, 2005), we do not rely on those for model assessment and comparison purposes.

Morin et al. (2016) mentioned, and demonstrated using simulated data, that goodness-of-fit assessment was not sufficient to guide the selection of the optimal model among similarly fitting models. This observation led them to recommend a more complete strategy involving a careful examination of the parameter estimates (i.e., loadings, cross-loadings, latent correlations, composite reliability) obtained from these various models. More precisely, this examination should start with a comparison of the CFA and ESEM models, where the observation of reduced factor correlations in ESEM coupled with generally well-defined factors could be taken as evidence in favor of the ESEM solution over a similarly fitting CFA solution. Then, the retained model should be contrasted to its bifactor counterpart. In this second comparison, the observation of a well-defined G-factor coupled with at least a subset of well-defined S-factor could be taken as evidence supporting a bifactor solution over a similarly fitting first-order solution. The model providing the most optimal representation of the data on the basis of this examination was retained for subsequent analyses and for Study 2.

Results and Discussion

Factor Validity and Composite Reliability. The goodness-of-fit indices of the alternative measurement models estimated for the BCQ and BCCS are respectively reported in Tables 1 and 2. The CFA and bifactor-CFA solutions resulted in an excellent level of fit to the data for the BCQ (CFI-TLI > .95, RMSEA $\leq$.06), but failed to reach acceptability for the BCCS (CFI-TLI < .90, RMSEA $\geq$ .10). By contrast, the ESEM and bifactor-ESEM solutions resulted in an excellent fit to the data for both instruments (CFI-TLI > .95, RMSEA $\leq$.06), and in a level of fit that was substantially improved relative to their CFA (BCQ: ΔCFI = +.019, ΔTLI = +.015, ΔRMSEA = -.013; BCCS: ΔCFI = +.089, ΔTLI = +.094, ΔRMSEA = -.051) and bifactor-CFA (BCQ: ΔCFI = +.012, ΔTLI = +.012, ΔRMSEA = -.012; BCCS: ΔCFI = +.088, ΔTLI = +.104, ΔRMSEA = -.006) counterparts. The comparison of ESEM and bifactor-ESEM models suggested that the bifactor-ESEM provided a superior representation of the data than the ESEM for both the BCQ (ΔCFI = +.008, ΔTLI = +.010, ΔRMSEA = -.010) and the BCCS (ΔCFI = +.006, ΔTLI = +.006, ΔRMSEA = -.006). However, as mentioned by Morin et al. (2016), final model selection should always be conditioned on a careful examination of the parameter estimates, composite reliability and factor correlations of these models, starting with a comparison of CFA and ESEM solutions, before contrasting the most optimal of these representations with its bifactor counterpart.

The detailed parameter estimates from the CFA and ESEM solutions for the BCQ and BCCS are respectively presented in Tables S3 and S6 of the online supplements, while those of the bifactor-CFA and bifactor-ESEM are respectively presented in Tables S4 and S7. In the CFA models, the factors loadings of the BCQ (λ = .630-.887, M_λ = .751) and BCCS (λ = .612-.869, M_λ = .748) were high and associated with satisfactory levels of composite reliability (BCQ: ω = .882-.905, M_ω = .908; BCCS: ω = .826-.889, M_ω = .863). However, some of the latent correlations observed among the BCQ (r = .806-.869, M_r = .829) and BCCS (r = .382-.905, M_r = .657) factors appeared to be high enough to call into question the distinctiveness of some factors, suggesting the need to account for possible sources of construct-relevant psychometric multidimensionality present at the items level through the incorporation of cross-loadings, and/or of a global factor to the models (Morin et al., 2016). Factor correlations were substantially reduced in ESEM for both the BCQ (r = .583-.784, M_r = .652) and the BCCS (r = .177-.680, M_r = .386). Similarly, the ESEM factors themselves remained well-defined by high factor loadings (λ = .342-.850, M_λ = .594), small cross-loadings ($|\lambda|$ = .001-.478, $M_{|\lambda|}$ = .163), and satisfactory estimates of composite reliability for the BCCS (ω = .723-.871, M_ω = .815). However, two items associated with the OBEVAL factor (BCCS11 and BCCS13) presented a weak pattern of association with two factors (i.e., high cross-loadings, .351 to .474), suggesting that these items might be stronger indicators of a global body checking cognition factor than of any specific type of cognition. The BCQ ESEM solution resulted in a very similar pattern of results (λ = .193-1.10, M_λ = .636; cross-loadings $|\lambda|$ = .001-.390, $M_{|\lambda|}$ = .148; ω = .823-.918, M_ω = .877), but revealed a greater number of items presenting an inconsistent and weak pattern of association to multiple factors (suggesting again, correspondence to a broadband construct rather than to specific behaviors). However, this model was not fully proper statistically as it revealed two out-of-bound factor loadings (> 1.0). Taken together, these results support the superiority of an ESEM representation of the data relative to a CFA representation for both instruments. However, the out-of-bound parameter estimates associated with the BCQ solution, and the items presenting a weak and inconsistent pattern of loadings and cross-loadings suggest that the ESEM solution might not be optimal and supported the need to consider alternative models.

The bifactor-ESEM solutions revealed a well-defined and reliable G-factor for both the BCQ (λ = .524-.824, M_λ = .683; ω = .966) and the BCCS (λ = .240-.830, M_λ = .608; ω = .951), suggesting that all items used in these questionnaires do indeed tap into a global overarching construct of body checking or body checking cognitions, respectively, over and above their associations with more specific facets of these global constructs. The S-factors from these two models generally appeared to be much more weakly defined when the variance attributed to the G-factor was considered. Still, the results did show that the S-factors of the BCQ retain some meaningful specificity over and above participants' global levels of body checking (OVAPP: λ = .025-.486, M_λ = .284, ω = .646; SBP: λ = .120-.574, M_λ = .312, ω = .712; IDSC: λ = .091-.546, M_λ = .345, ω = .632). By contrast, for the BCCS, the results revealed two generally very weak S-factors related to OBEVAL (λ = .100-.527, M_λ= .262, ω = .516) and SAFEB (λ = .075-.339, M_λ = .206, ω = .322) and two reasonably well-defined S-factors related to BODC (λ = .267-.568, M_λ = .396, ω = .700) and REASS (λ = .622-.791, M_λ = .710, ω = .870). Altogether, these results support the superiority of the bifactor-ESEM solution for both instruments.

Study 2: Cross-validation, Convergent Validity, Criterion-Related Validity, and Measurement Invariance

Method

Sample and Procedure. A new sample of 434 (378females and 56males) French-speaking Canadian adolescents and adults, aged between 15 and 68 years old (M_{age} = 26.5 years, SD = 9.6), participated in this study. Recruitment was conducted in secondary schools, colleges, universities, community organizations for eating disorders, and a private clinic specialized in eating disorders all located in the Canadian Province of Quebec. The participants were invited to participate by generic announcements sent via school messaging (letters or emails) bulletin board or community organizations websites, and completed an online informed consent form, before anonymously completing the questionnaires online using the LimeSurvey platform.

Measures. Participants completed the same questions as in Study 1, as well as measures of:

Self-Conceptions. Participants levels of Global Self-Esteem (GSE) and Physical Appearance (PA) were measured using three items from the GSE (α = .857) and PA (α = .743) subscales from the Short Physical Self-Inventory (PSI-S; Maïano et al., 2008; Morin & Maïano, 2011). Each item was rated on a six-point response scale ranging from *not at all* (1) to *entirely* (6). The PSI-S has been validated and cross-validated (Maïano et al., 2008; Morin & Maïano, 2011) among large samples (combined N = 3,047) of French-speaking adolescents. Findings from these studies provide support for the factor validity, reliability (α = .64-.73 to .75-.87 across three independent samples), temporal stability (r = .74-.84) and convergent validity of this instrument.

Social Physique Anxiety. Participants completed the French version (Maïano, Morin, Eklund et al., 2010) of the Social Physique Anxiety Scale (SPAS; Hart et al., 1989). The seven items (α = .942) from this instrument were rated on a five-point response scale ranging from *not at all* (1) to *extremely* (5). The French SPAS has been validated in a series of six studies (Maïano, Morin, Eklund et al., 2010) conducted among a large sample of adolescents (N = 1,563). Findings from these studies provide support for the factor validity, reliability (α = .81 to .87 across two independent samples), temporal stability (r = .78), and convergent validity of this instrument.

Fear of Negative Appearance Evaluation. Participants completed the French version (Maïano, Morin, Monthuy-Blanc et al., 2010) of the Fear of Negative Appearance Evaluation Scale (FNAES; Lundgren et al., 2004). This scale includes five items (α = .948) rated on a five-point scale ranging from *not at all* (1) to *extremely* (5).The French FNAES has been validated in a series of three studies (Maïano, Morin, Monthuy-Blanc et al., 2010) conducted among a sample of 684 adolescents. Findings from these studies support the factor validity, reliability (α = .83 across two independent samples), temporal stability (r = .77), and convergent validity of this instrument.

Disturbed Eating Attitudes and Behaviors. Participants completed the French version (Leichner, Steiger, Puentes-Neuman, Perreault, & Gottheil, 1994) of the Eating Attitudes Test-26 (EAT-26; Garner, Olmsted, Bohr, & Garfinkel, 1982). This instrument comprises 26 items and assesses three dimensions: dieting (α = .930), bulimia-food preoccupation (α = .875), and oral control (α = .822). Participants answer each item using a six-point response scale ranging from *always* (6) to *never* (1). The French EAT-26 has been validated among samples of clinical and non-clinical adolescents and adults (Leichner et al., 1994). Findings from this study dsupport the factor validity, reliability (α = .54 to .86), and criterion-related validity of this instrument.

Analyses. The model retained as the most optimal in the second study was first replicated using Mplus 7.4 WLSMV estimator (Muthén & Muthén, 2015) using the full available information to account for the limited amount of missing data present at the item level (BCQ: M = 1.2%; BCCS: M = 2.3%). In Study 2, model fit assessment and comparisons followed the same procedures outlined in Study 1.The measurement invariance of the best model identified for the BCQ and BCCS was then examined across samples from the first and second studies. These tests were performed in the following sequence adapted to WLSMV (Guay, Morin, Litalien, Valois, & Vallerand, 2015; Morin, Arens, Tran, & Caci, 2016; Morin et al., 2011): (i) configural invariance; (ii) invariance of the factor loadings (weak invariance); (iii) invariance of the thresholds (strong invariance); (iv) invariance of the uniquenesses (strict invariance); (v) invariance of the latent variances/covariances; and (vi) invariance of thelatent means.

In a second stage, we examined the convergent validity the BCQ and BCCS with one another using latent variable correlations separately in both samples, before assessing the equivalence of these correlations across samples. This test started from the most invariant measurement model previously identified for both instruments, and involved adding invariance constraints on the covariance between the BCQ and BCCS factors. Then, in a third stage, the criterion-related validity of the BCQ and BCCS with measures of GSE, PA, SPAS, FNAES, and EAT-26 was examined using latent factor correlations.

In a fourth stage, and assuming that the measurement invariance of the bifactor-ESEM solution would be demonstrated, samples from Studies 1 and 2 were combined in order to conduct further tests of measurement invariance as a function of participants' sex and ED diagnosis. Because tests of measurement invariance require the formation of groups, we relied on tests of differential item functioning (DIF) to assess the impact of continuous scores of age and BMI on items ratings (item thresholds with ordinal indicators and the WLSMV estimator). These tests were conducted using multiple indicators multiple causes (MIMIC) in which linear and curvilinear effects of age (age, age^2)

and BMI (BMI, BMI^2) were used as predictors of BCQ and BCCS ratings. To facilitate interpretations, age and BMI were standardized prior to the analyses. These MIMIC tests of DIF were performed in the following sequences (Marsh, Nagengast, & Morin, 2013; Morin, Marsh, & Nagengast, 2013): (a) null effects model (the paths from age and BMI to the latent factors and item responses [thresholds] are constrained to be zero); (b) saturated model (the paths from age and BMI to the item responses are freely estimated, while the paths from age and BMI to the latent factors are constrained to be zero); and (c) invariant model (the paths from age and BMI to the latent factors are freely estimated, while the paths from age and BMI to the item responses thresholds are constrained to be zero).

Results and Discussion

Cross-validation. The bifactor-ESEM solutions of the BCQ and BCCS were replicated in this second independent sample. Tables 1 and 2, respectively present the goodness-of-fit indices of these models and reveal an excellent level of fit to the data for both questionnaires (CFI-TLI > .95, RMSEA ≤.06). The parameter estimates from this model are presented in Tables S5 and S8 of the online supplements, and are highly similar to those reported in Study 1.

Measurement Invariance across samples. Results from the tests of measurement invariance of the BCQ and BCCS conducted across samples are respectively reported in Tables 1 and 2 (see models 3-1 to 3-6). These results support the full invariance (loadings, thresholds, uniquenesses, latent variances-covariances) of the bifactor-ESEM models for both questionnaires (ΔCFI/ΔTLI < .010; ΔRMSEA < .015). However, for the BCQ, the ΔCFI-TLI (.01) and ΔRMSEA (> .015) revealed the presence of latent means differences across samples. These results showed that the sample from the second study tended to present higher latent means on the G-factor (.334, p< .001), on the OVAPP S-factor (.199, p = .047), and on the IDSC S-factors (.299, p = .014). No significant difference was found for the SBP S-factor (.110, p = .225). Similarly, for the BCCS, the ΔRMSEA (+.015) suggested the presence of latent mean differences across samples (while ΔCFI-TLI remained marginal at -.008). These results showed that the sample from the second study tended to present higher mean levels on the G-factor (.345, p< .001), and lower means levels on the REASS S-factor (-.157, p = .038), and on the SAFEB S-factors (-.297, p = .014). No significant difference was found for the OBEVAL S-factor (-.076, p = .474), and the BODC S-factor (-.041, p = .684).

Convergent validity. The bifactor-ESEM solution was estimated to examine the convergent validity of the BCQ and BCCS factors across samples from Studies1 and 2. The model in which correlations between these factors were freely estimated [χ^2(1670) = 2047.50, CFI = .990, TLI = .990, RMSEA = .022, 90%CI = .019-.025] and the model in which these were constrained to be equal across samples [χ^2(1690) = 1967.09, CFI = .993, TLI = .993, RMSEA = .019, 90%CI = .015-.022] were both excellent and equivalent (ΔWχ^2 = 39.96, Δdf = 20, p .06, ΔCFI = +0.003, ΔTLI +0.003, ΔRMSEA -0.003). These results thus support the equivalence of the latent covariances between the BCQ and BCCS factors across samples, which are reported in Table 3. These results show that the G-factor of the BCQ is significantly and (a) positively related to the SAFEB and BODC S-factors and the G-factor of the BCCS; and (b) negatively related to the REASS S-factor of the BCCS. The OVAPP S-factor of the BCQ is significantly and (a) positively related to the G-factor of the BCCS, and (b) negatively related to the BODC S-factor of the BCCS. The SBP S-factor of the BCQ is significantly and negatively related to the REASS and SAFEB S-factors of the BCCS. Finally, the IDSC S-factor of the BCQ is significantly and negatively related to the REASS S-factor of the BCCS.

Criterion-related validity. The goodness of fit of the bifactor-ESEM solution examining the convergent validity of the BCQ and the BCCS with criterion measures resulted in an excellent level of fit to the data (see models 2-2 in Tables 1 and 2). The results from these analyses are reported in Table 4, and show that the BCQ G-factor is significantly and (a) negatively related to GSE and PA; and (b) positively related to the FNAES, SPAS, and all subscales of the EAT-26. The OVAPP S-factor is significantly and (a) negatively related to GSE; and (b) positively related to the FNAES, SPAS, and oral control subscale of the EAT-26. Moreover, the SBP S-factor is significantly and positively related to GSE and PA. Finally, the IDSC S-factor is significantly and (a) positively related to GSE and PA, and the oral control subscale of the EAT-26; and (b) negatively related to the FNAES and SPAS.

Concerning the BCCS, the G-factor is significantly and (a) negatively related to GSE and PA; and (b) positively related to the FNAES, SPAS, and all subscales of the EAT-26. The OBEVAL S-factor is also significantly and negatively related to the bulimia-food preoccupation and oral control subscales of the EAT-26. Furthermore, the REASS S-factor is significantly and (a) positively related to GSE and

PA; and (b) negatively related to the FNAES, SPAS, and the dieting, and bulimia-food preoccupation subscales of the EAT-26. The SAFEB S-factor is significantly and (a) positively related to GSE and PA; and (b) negatively related to the FNAES, SPAS, and the dieting subscale of the EAT-26. Finally, the BODC specific factor is positively related to all subscales of the EAT-26.

For interested readers, we provide comparable results taken from a first-order ESEM solution in TablesS9 and S10 of the online supplements. As noted in the introduction, the S-factors from a bifactor solution can be directly interpreted as providing a test of the added-value of considering subscale-specific scores over and above global construct scores. In this context, it is not surprising that the results reported in Table 4 reveal associations with criterion measures that are dominated by the global factors, while also providing evidence of the added-value of considering some specific subscales in a way that differ across criterion measures (e.g., the REASS scale from the BCCS appeared particularly important to consider in relation with global self-esteem and physical appearance). By failing to partial out global constructs from subscale-specific scores, the ESEM solutions resulted in a far less differentiated pattern of associations with the criterion measures, showing statistically significant associations between almost all factors and criterion measures.

Measurement Invariance: Sex and ED Diagnosis. The results from the tests of the measurement invariance of the BCQ and BCCS as a function of participants' sex and diagnosis of ED conducted on the combined sample are reported in Tables 1 and 2 (see models 4-1 to 4-6 and 5-1 to 5-6). These results support the full invariance (loadings, thresholds, uniquenesses, latent variances and covariances) of the bifactor-ESEM representation of the BCQ and BCCS as a function of sex and diagnosis of ED (ΔCFI/ΔTLI < .010; ΔRMSEA < .015). However, for both questionnaires, the ΔCFI-TLI (> .01) and ΔRMSEA (> .015) suggest the presence of latent means differences as a function of sex and diagnosis of ED.

For the BCQ, the results showed that males tended to score significantly lower on the G-factor (-.780, $p < .001$) and the OVAPP (-.819, $p < .001$) and SBP (-1.189, $p < .001$) S-factors when compared to females, but significantly higher on the IDSC (.835, $p = .012$) S-factor. Participants without a diagnosis of ED tended to score significantly lower on the G-factor (-1.066, $p < .001$) and higher on the IDSC S-factor (.667, $p < .001$) than those with a diagnosis of ED. No significant differences were found for the OVAPP (.241, $p = .197$) and SBP (-.256, $p = .078$) S-factors.

For the BCCS, the results similarly showed that boys/men tended to score significantly lower on the G-factor (-.931, $p < .001$) and higher on the SAFEB S-factor (.564, $p < .001$) relative to females, but no significant differences for the OBEVAL (.074, $p = .654$), REASS (.120, $p = .251$), and BODC (.310, $p = .090$) S-factors. Finally, the results showed that participants without a diagnosis of ED tended to score significantly lower on the G-factor (-1.166, $p < .001$) and higher on the REASS (.515, $p < .001$) and SAFEB (.684, $p = .001$) S-factors, with no significant differences noted for the OBEVAL (-.029, $p = .873$) and BODC (-.124, $p = .463$) S-factors.

DIF: Age and BMI. The results from the MIMIC tests of DIF conducted as a function of age and BMI levels on the combined sample are reported in Tables 1 and 2 (see models 6-1 to 6-3 and 7-1 to 7-3). For the BCQ, the results showed that none of the models allowing for age and BMI to influence scores on the BCQ/BCCS responses (saturated model) or factors (invariant model) resulted in any meaningful improvement in model fit when compared to the null effects model. These results are consistent with the complete equivalence of the BCQ/BCCS responses as a function of age and BMI, and a lack of effects of age and BMI on scores on the BCQ/BCCS latent factors.

General Discussion

The first study provided support for the superiority of the bifactor-ESEM solution (when compared to alternative CFA, bifactor-CFA and ESEM solutions) for both questionnaires. This solution reveals a well-defined and reliable G-factor for both the BCQ and the BCCS, suggesting that all items used in these questionnaires do indeed tap into a global overarching construct of body checking or body checking cognitions, over and above their associations with more specific facets of these global constructs. Interestingly, the S-factors from these two questionnaires generally appear to be much more weakly defined when the variance attributed to the G-factor is considered. Still, the results do show that the BCQ S-factors retain some meaningful specificity over and above participants' global levels of body checking (OVAPP, SBP, IDSC). For the BCCS, the results reveal two generally very weak S-factors related to OBEVAL and SAFEB and two reasonably well-defined S-factors related to BODC and REASS.

The second study successfully replicated the bifactor-ESEM solutions of the BCQ and BCCS on a new independent sample, and supported the full invariance of this solution across samples for both questionnaires. However, the analyses also showed that the sample from the second study tended to present (a) higher latent mean levels on the G-factor of the BCQ, on the OVAPP S-factor, and on the IDIOC S-factors; and (b) higher latent mean levels on the G-factor of the BCCS, and lower latent means levels on the REASS and SAFEB S-factors. This observation may be explained by the fact that nearly three times more participants reported having received a diagnosis for ED in the sample from the second study than in the sample from the first study.

Additionally, the results supported the equivalence of the covariances between the BCQ and the BCCS across samples. Findings obtained from these analyses of convergent validity revealed that the G-factors of both questionnaires are highly positively correlated and that the REASS and BODC S-factors of the BCCS are moderately correlated with most of the S-factors of the BCQ. Interestingly, this research is the first to report associations involving global scores on these instruments. In addition, given that the present research was the first to rely on bifactor modeling, it is perhaps not so surprising to note that the results obtained for the S-factors do not perfectly match those obtained in previous studies at the subscale levels, which tended to show positive correlations between the BCQ and BCCS subscales (Haase et al., 2007; Mountford et al., 2006).Indeed, in bifactor models, the S-factors reflect the covariance shared among the items forming these subscales but left unexplained by the G-factor. By contrast, subscale correlations reported in these earlier studies relied on a conflation of the variance attributed to participants' global levels of body checking behaviors and cognitions with that more specifically related to the subscales. For instance, in the BCQ, scores on the S-factors reflect a tendency to over- or under-check one's overall appearance, specific body parts, or idiosyncratic characteristics over and above one's global body checking tendencies, i.e., some degree of imbalance in the extent to which one tends to focus on specific dimensions relative to the others. Similarly, for the BCCS, scores on the S-factors also likely reflect some degree or over- or under-investment in specific types of cognitions relative to one's global body checking cognitions.

Additional analyses supported the criterion-related validity of both questionnaires in relation to a variety of criterion measures. More specifically, findings showed that the G-factors from both questionnaires were highly correlated (positively or negatively) with the criterion measures in a direction that matched the results from previous studies (Calugi et al., 2006; Campana et al., 2013; Kachani et al., 2013; Reas et al., 2002, 2006, 2009; Steinfiel et al., 2017; Vocks et al., 2008). Additionally, moderate correlations were found for the S-factors of the BCQ and BCCS and criterion measures. More specifically, results related the OVAPP S-factor of the BCQ and of the BODC S-factor of the BCCS were consistent with those found in previous studies (Campana et al., 2013; Haase et al., 2007; Mountford et al., 2006; Reas et al., 2006). However, the correlations related to the remaining S-factors were largely inconsistent with those obtained in previous studies (Campana et al., 2013; Haase et al., 2007; Mountford et al., 2006; Reas et al., 2006). Once again, this is likely related to the specific nature of the bifactor representation adopted in the present research.

The results also supported the ability to use these questionnaires to conduct unbiased comparisons as a function of participants' sex, ED diagnoses, age, and BMI based on evidence of measurement invariance or lack of DIF as a function of these individual characteristics. Whereas the results showed that age and BMI had no effects whatsoever on BCQ and BCCS ratings, they also revealed significant latent mean differences related to participants' sex and ED diagnosis. More specifically, the results showed that males, compared to females, tend to present significantly lower levels on the G-factors of the BCQ and BCCS. Likewise, participants without a diagnosis of ED, compared to those with a diagnosis of ED, tended to present significantly lower levels on the G-factors of the BCQ and BCCS. These findings are very similar to those found in previous studies in relation to sex (Reas et al., 2006; Vocks et al., 2008) or ED diagnosis (Calugi et al., 2006; Kachani et al., 2013;Mountford et al., 2006;Pellizzer et al., 2018;Steinfiel et al., 2017; Vocks et al., 2008; Waller, Sines, Meyer, & Mountford, 2008). However, significantly higher levels on the OVAPP and SBPS-factors of the BCQ, and on the SAFEB S-factor of the BCCS were found for males relative to females. Similarly, higher levels on the IDSC S-factor of the BCQ and on REASS and SAFEB S-factors of the BCCS were found for participants without a diagnosis of ED. Once again, this is likely to be related to the specific nature of the bifactor representation adopted in the present research.

A possible critical implication of these results has to do with the scoring of measures known to follow a bifactor structure. For research purposes, the recommendation is simple and involves relying on latent variable models similar to those used in the present study. For practical purposes, however, the solution is not as straightforward. Indeed, typical scale scores obtained via the addition or averaging of the items forming a scale are not able to accommodate a bifactor structure or cross loadings (Brown, Finney, & France, 2011). Despite the apparent complexity of this question, the solution is in fact relatively simple. As noted by Perreira et al. (2018, p.79), "*scoring should be computerized and based on algorithms similar to those used to generate factor scores. Interestingly, the Mplus statistical package can be used in such a manner, relying on an estimation of factor scores conditioned on the exact parameter estimates of the final bifactor-ESEM solution*."

Four main limitations of the current series of studies should be considered when interpreting the present findings. First, the information about the present or previous diagnosis of ED was self-reported. It is thus probable that the percentage of respondents suffering from clinical levels of ED has been over- or under-estimated. Consequently, future studies should investigate whether the bifactor-ESEM structure of the BCQ and BCCS will hold in samples with a formal diagnosis of ED. Second, it is important to use caution when considering some of these results, particularly those related to measurement invariance, given the smaller sample size of some of the groups of participants (males in particular). Therefore, it is important for future studies to replicate the present results on larger, and more representative samples, and particularly on samples including a greater proportion of males. The present study relied on a convenience mode of sampling based on voluntary participation. The greater proportion of females volunteering for participation suggests that males might be less interested in this research theme. It would thus appear important, for future research, to consider ways in which to build male interest for participation, possibly via the addition of different research themes (e.g., muscular verification behaviors) into the questionnaire.

Third, we did not examine the temporal stability of the BCQ and BCCS. Therefore, the test-retest reliability of these questionnaires should be assessed in the future to obtain a fuller picture of their psychometric properties. Fourth, the criterion validity of the BCQ and BCCS related to their ability to distinguish clinical and non-clinical samples has not been assessed in the present research. It is thus unknown whether these questionnaires can differentiate between individuals with and without a formal diagnosis of ED, or between subgroups of individuals with anorexia nervosa, bulimia nervosa or ED not otherwise specified.

In conclusion, the present results validate and cross-validate the French versions of the BCQ and BCCS, and support the idea that these questionnaires are best represented according to a bifactor-ESEM solution. These questionnaires can thus be used to assess body checking behaviors and body checking cognitions among community samples of French-speaking adolescents and adults with similar characteristics (sex, age, and BMI).

References

Alfano, L., Hildebrandt, T., Bannon, K., Walker, C., & Walton, K. E. (2011). The impact of gender on the assessment of body checking behavior. *Body Image*, *8*(1), 20-25.

Asparouhov, T., & Muthén, B.O. (2009). Exploratory structural equation modeling. *Structural Equation Modeling*, *16*, 397-438.

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

Brown, A., Finney, S., & France, M. (2011). Using the bifactor model to assess the dimensionality of the Hong Psychological Reactance Scale. *Educational and Psychological Measurement*, *71*, 170-185.

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

Calugi, S., Dalle Grave, R., Ghisi, M., & Sanavio, E. (2006). Validation of the Body Checking Questionnaire (BCQ) in an eating disorders population. *Behavioural and Cognitive Psychotherapy*, *34*, 233-242.

Campana, A. N. N. B., Swami, V., Onodera, C. M. K., da Silva, D., & Fernandes, M. D. C. G. C. (2013). An initial psychometric evaluation and exploratory cross-sectional study of the Body Checking Questionnaire among Brazilian women. *PlosOne*, *8*, e74649.

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

Cheung, G. W., & Rensvold, R. B. (2002). Evaluating goodness-of fit indexes for testing measurement

invariance. *Structural Equation Modeling*, *9*, 233-255. doi: 10.1207/S15328007SEM0902_5
Fairburn, C. G., Cooper, Z., & Shafran, R. (2003). Cognitive behaviour therapy for eating disorders: A "transdiagnostic" theory and treatment. *Behaviour Research and Therapy*, *41*, 509-528.
Garner, D. M., Olmstead, M. P., Bohr, Y., & Garfinkel, P. E. (1982). The Eating Attitude Test: Psychometric features and clinical correlates. *Psychological Medicine*, *12*, 871-878.
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.
Guay, F., Morin, A. J. S., Litalien, D., Valois, P., & Vallerand, R. J. (2015). An application of exploratory structural equation modeling to evaluate the Academic Motivation Scale. *The Journal of Experimental Education*, *83*, 51–82. doi: 10.1080/00220973.2013.876231
Haase, A. M., Mountford, V., & Waller, G. (2007). Understanding the link between body checking cognitions and behaviors: The role of social physique anxiety. *International Journal of Eating Disorders*, *40*, 241-246.
Haase, A. M., Mountford, V., & Waller, G. (2011). Associations between body checking and disordered eating behaviors in nonclinical women. *International Journal of Eating Disorders*, *44*(5), 465-468.
Hart, E.H., Leary, M.R., & Rejeski, W.J. (1989). The measurement of social physique anxiety. *Journal of Sport & Exercise Psychology*, *11*, 94-104.
Hu, L., & Bentler, P. M. (1999). Cutoff criteria for fit indexes in covariance structure analysis. *Structural Equation Modeling, 6,* 1–55. doi: 10.1080/10705519909540118
Kachani, A. T., Barroso, L. P., Brasiliano, S., Hochgraf, P. B., Cordás, T. A., & Conti, M. A. (2013). Psychometric evaluation of the body checking cognitions scale (BCCS) Portuguese version. *Perceptual and Motor Skills*, *116*, 175-186.
Leichner, P., Steiger, H., Puentes-Neuman, G., Perreault, M., & Gottheil N. (1994). Validation d'une échelle d'attitudes alimentaires auprès d'une population québécoise francophone. *Revue Canadienne de Psychiatrie*, *39*, 49-54.
Lundgren, J. D., Anderson, D. A., & Thompson, J. K. (2004). Fear of negative appearance evaluation: Development and evaluation of a new construct for risk factor work in the field of eating disorders. *Eating Behaviors*, *5*, 75–84. doi:10.1016/S1471-0153(03)00055-2
Maïano, C., Morin, A. J. S., Eklund, R. C., Monthuy-Blanc, J., Garbarino, J.-M., & Stephan, Y.(2010). Construct validity of the Social Physique Anxiety Scale in a French adolescent Sample. *Journal of Personality Assessment*, *92*, 1-10.
Maïano, C., Morin, A. J. S., Monthuy-Blanc, J., & Garbarino, J. M. (2010). Construct validity of the Fear of Negative Appearance Evaluation Scale in a community sample of French adolescents. *European Journal of Psychological Assessment*, *26*, 19–27. doi:10.1027/1015-5759/a000004
Maïano, C., Morin, A. J. S., Ninot, G., Monthuy-Blanc, J., Stephan, Y., Florent, J.-F., & Vallée, P. (2008). A short and very short form of the Physical Self-Inventory for adolescents: Development and factor validity. *Psychology of Sport & Exercise*, *9*, 830-847.
Marsh, H. W., Hau, K-T., & Grayson, D. (2005). Goodness of fit evaluation in structural equation modeling. In A. Maydeu-Olivares, & J. McArdle (Eds.), *Contemporary psychometrics* (pp. 275-340). Hillsdale, NJ: Erlbaum.
Marsh, H. W., Nagengast, B., & Morin, A. J. S. (2013). Measurement invariance of big-five factors over the life span: ESEM tests of gender, age, plasticity, maturity, and la dolce vita effects. *Developmental Psychology*, *49*, 1194-1218. doi: 10.1037/a0026913
McDonald, R. P. (1970). Theoretical foundations of principal factor analysis, canonical factor analysis, and alpha factor analysis. *British Journal of Mathematical and Statistical Psychology*, *23*, 1–21. doi: 10.1111/j.2044-8317.1970.tb00432.x
Millsap, R. E. (2011). *Statistical approaches to measurement invariance*. New York, NY: Routledge.
Morin A. J. S., Moullec G., Maïano C., Layet L., Just J.-L., Ninot G. (2011). Psychometric properties of the Center for Epidemiologic Studies Depression Scale (CES-D) in French clinical and nonclinical adults. *Epidemiology and Public Health [Revue d'Épidémiologie et de Santé Publique]*, *59*, 327–340. doi: 10.1016/j.respe.2011.03.061
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. *Psychology of Sport & Exercise*, *12*, 540-554.

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

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

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* (2nd ed., pp. 395-438). Charlotte, NC: Information Age.

Morin, A.J.S., Myers, N.D., & Lee, S. (2019). Modern factor analytic techniques: Bifactor models, exploratory structural equation modeling (ESEM) and bifactor-ESEM. In G. Tenenbaum & R.C. Eklund (Eds.), *Handbook of sport psychology*, 4th Edition. London, UK: Wiley.

Mountford, V., Haase, A., & Waller, G. (2006). Body checking in the eating disorders: Associations between cognitions and behaviors. *International Journal of Eating Disorders*, *39*, 708-715.

Mountford, V., Haase, A. M., & Waller, G. (2007). Is body checking in the eating disorders more closely related to diagnosis or to symptom presentation? *Behaviour Research & Therapy*, *45*, 2704-2711.

Muthén, L. K., & Muthén, B. (2015). *Mplus user's guide (7thed.).* Los Angeles, CA: Muthén & Muthén.

Pellizzer, M. L., Tiggemann, M., Waller, G., & Wade, T. D. (2018). Measures of body image: Confirmatory factor analysis and association with disordered eating. *Psychological Assessment*, *30*, 143-153. doi: 10.1037/pas0000461

Perreira, T.A.,Morin, A.J.S., Hebert, M., Gillet, N., Houle, S.A., & Berta, W. (2018). The short form of the Workplace Affective Commitment Multidimensional Questionnaire (WACMQ-S): A bifactor-ESEM approach among healthcare professionals. *Journal of Vocational Behavior, 106,* 62-83.

Reas, D. L., Von Soest, T., & Lask, B. (2009). Reliability and validity of the Norwegian version of the Body Checking Questionnaire. *Tidsskrift for Norsk Psykologforening*, *46*, 2.

Reas, D. L., Whisenhunt, B. L., Netemeyer, R., & Williamson, D. A. (2002). Development of the Body Checking Questionnaire: A self-report measure of body checking behaviors. *International Journal of Eating Disorders*, *31*, 324-333.

Reas, D. L., White, M. A., & Grilo, C. M. (2006). Body checking questionnaire: Psychometric properties and clinical correlates in obese men and women with binge eating disorder. *International Journal of Eating Disorders*, *39*, 326-331.

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

Steinfeld, B., Bauer, A., Waldorf, M., Engel, N., Braks, K., Huber, T. J., & Vocks, S. (2017). Validierung einer deutschsprachigen Fassung des Body Checking Questionnaire (BCQ) an Jugendlichen mit Anorexia und Bulimia Nervosa. *PPmP-Psychotherapie· Psychosomatik· Medizinische Psychologie*, *67*, 38-46.

Tomás, I., Marsh, H. W., González-Romá, V., Valls, V., & Nagengast, B. (2014). Testing measurement invariance across Spanish and English versions of the Physical Self-Description Questionnaire: An application of exploratory structural equation modeling. *Journal of Sport and Exercise Psychology*, *36*, 179-188. doi: 10.1123/jsep.2013-0070

Vocks, S., Moswald, C., & Legenbauer, T. (2008). Psychometrische Überprüfung einer deutschsprachigen Fassung des Body Checking Questionnaire (BCQ). *Zeitschrift für Klinische Psychologie und Psychotherapie*, *37*, 131-140.

Waller, G., Sines, J., Meyer, C., & Mountford, V. (2008). Body checking in the eating disorders: Association with narcissistic characteristics. *Eating Behaviors*, *9*, 163-169.

Williamson, D. A., White, M. A., York-Crowe, E., & Stewart, T. M. (2004). Cognitive-behavioral theories of eating disorders. *Behavior Modification*, *28*, 711-738.

Yu, C. Y. (2002). *Evaluating cutoff criteria of model fit indices for latent variable models with binary and continuous outcomes*. Los Angeles, CA: University of California.

Table 1
Goodness-of-Fit Statistics of the Alternative Models Estimated for the BCQ

Models	Study	N°	Description	χ^2(df)	CFI	TLI	RMSEA	90% CI	CM	$\Delta W\chi^2$(df)	ΔCFI	ΔTLI	ΔRMSEA
CFA[a]	1	1-1	CFA	679.899(227)*	.961	.957	.064	.059-.070	1-3	230.57(40)*	-.019	-.015	+.013
		1-2	Bifactor-CFA	489.678(207)*	.976	.970	.053	.047-.059	1-4	172.79(40)*	-.012	-.012	+.012
ESEM[a]	1	1-3	ESEM	425.719(187)*	.980	.972	.051	.045-.058	1-4	106.29(20)*	-.008	-.010	+.010
		1-4	Bifactor-ESEM	303.303(167)*	.988	.982	.041	.034-.048					
ESEM[b]	2	2-1	Bifactor-ESEM	358.903(167)*	.983	.975	.053	.046-.061	-	-	-	-	-
		2-2	Convergent analysis with criterion measures	4132.271(2032)*	.961	.957	.050	.048-.053	-	-	-	-	-
Bifactor-	1-2	3-1	Configural invariance	672.122(334)*	.985	.978	.048	.043-.053	-	-	-	-	-
ESEM:		3-2	Weak (λ) invariance	755.569(410)*	.985	.982	.044	.039-.048	3-1	136.05(76)*	.000	+.004	-.004
samples[c]		3-3	Strong (λ, ν) invariance	819.700(475)*	.985	.984	.040	.036-.045	3-2	108.42(65)*	.000	+.002	-.004
		3-4	Strict (λ, ν, δ) invariance	786.889(498)*	.988	.987	.036	.031-.041	3-3	32.20(23)	+.003	+.003	-.004
		3-5	Full (λ, ν, δ, ξ/φ) invariance	629.576(508)*	.995	.995	.023	.017-.029	3-4	10.78(10)	+.007	+.008	-.013
		3-6	Latent mean (λ, ν, δ, ξ/φ, η) invariance	860.116(512)*	.985	.985	.039	.035-.044	3-5	59.44(4)*	-.010	-.010	+.016
Bifactor-	2	4-1	Configural invariance	631.581(334)*	.986	.979	.045	.039-.050	-	-	-	-	-
ESEM:		4-2	λ invariance	706.348(410)*	.986	.983	.040	.035-.045	4-1	153.55(76)*	.000	+.004	-.005
Sex[d]		4-3	λ,ν invariance[☆]	706.102(467)*	.989	.988	.034	.029-.039	4-2	56.00(57)	+.003	+.005	-.006
		4-4	λ, ν, δ invariance	757.330(490)*	.987	.987	.035	.030-.040	4-3	66.18(23)*	-.002	-.001	+.001
		4-5	λ, ν, δ, ξ/φ invariance	693.470(500)*	.991	.991	.030	.024-.035	4-4	15.34(10)	+.004	+.004	-.005
		4-6	λ, ν, δ, ξ/φ, η invariance	1440.695(504)*	.956	.956	.065	.061-.069	4-5	237.68(4)*	-.035	-.035	+.035
Bifactor-	2	5-1	Configural invariance	633.825(334)*	.982	.972	.047	.041-.052	-	-	-	-	-
ESEM: ED		5-2	λ invariance	678.802(410)*	.984	.980	.040	.034-.045	5-1	128.65(76)*	+.002	+.008	-.007
diagnosis[e]		5-3	λ,ν invariance	751.537(475)*	.983	.982	.038	.032-.043	5-2	117.95(65)*	-.001	+.002	-.002
		5-4	λ, ν, δ invariance	746.956(498)*	.985	.985	.035	.030-.040	5-3	34.51(23)	+.002	+.003	-.003
		5-5	λ, ν, δ, ξ/φ invariance	636.276(508)*	.992	.992	.025	.018-.031	5-4	12.47(10)	+.007	+.007	-.010
		5-6	λ, ν, δ, ξ/φ, η invariance	1334.369(512)*	.949	.950	.063	.059-.067	5-5	195.457(4)*	-.043	-.042	+.038
DIF: Age[f],	2	6-1	MIMIC Null model	510.534(213)*	.988	.983	.040	.035-.044	-	-	-		-
Age²		6-2	MIMIC Saturated	587.955(167)*	.983	.970	.053	.049-.058	6-1	116.59(46)*	-.005	-.013	+.013
		6-3	MIMIC Invariant	684.271(205)*	.981	.972	.051	.047-.056	6-2	128.298(38)*	-.002	+.002	-.002
DIF: BMI[g],	2	7-1	MIMIC Null model	509.908(213)*	.988	.983	.040	.036-.044	-	-	-		-
BMI²		7-2	MIMIC Saturated	571.310(167)*	.984	.971	.053	.048-.057	7-1	120.45(46)*	-.004	-.012	+.013
		7-3	MIMIC Invariant	699.650(205)*	.980	.971	.053	.048-.057	7-2	153.30(38)*	-.004	.000	.000

Note. χ^2 = robust weighed least square (WLSMV) chi-square ; BCQ = Body Checking Questionnaire; BMI = body mass index; CFA = confirmatory factor analytic model; CFI = comparative fit index; CM = comparison model; df = degrees of freedom; DIF = differential item functioning;ED = eating disorders; ESEM = exploratory structural equation modeling; MIMIC = multiple indicators multiple cause; TLI = Tucker-Lewis index; RMSEA = root mean square error of approximation; 90% CI = 90% confidence interval of the RMSEA; λ = factor loadings; ν = thresholds; δ = Uniquenesses; ξ = factor variances; φ = factor covariances; η = factor means; $\Delta W\chi^2$ = WLSMV chi square difference test (calculated with the Mplus DIFFTEST function); Δ = change from the previous model.The fact that WLSMV χ^2 values are not exact, but "estimated" as the closest integer necessary to obtain a correct *p* value explains the fact that the χ^2 and the resulting CFI values can be non-monotonic with model complexity.*p <.01; [a] *N*= 482; [b]*N*= 406;[c]*N*= 888; [d] females: *N*= 761, males: *N*= 127; [e] No-diagnosis: *N*= 720, diagnosis: *N*= 105; [f]*N*= 886; [g]*N*= 873.[☆]Because all response categories were not used in at least one of the subsamples for some of the items, the responses to these items were recoded into fewer categories.

Table 2
Goodness-of-Fit Statistics of the Alternative Models Estimated for the BCCS

Models	Study	N°	Description	χ^2(df)	CFI	TLI	RMSEA	90% CI	CM	ΔWχ^2(df)	ΔCFI	ΔTLI	ΔRMSEA
CFA [a]	1	1-1	CFA	974.003(146)*	.890	.871	.108	.102-.114	1-3	551.04(45)*	-.089	-.094	+.051
		1-2	Bifactor-CFA	910.964(133)*	.897	.867	.110	.103-.116	1-4	552.93(47)*	-.088	-.104	+.059
ESEM [a]	1	1-3	ESEM	258.650(101)*	.979	.965	.057	.048-.065	1-4	65.34(15)*	-.006	-.006	+.006
		1-4	Bifactor-ESEM	196.475(86)*	.985	.971	.051	.042-.061	-	-	-	-	-
ESEM [b]	2	2-1	Bifactor-ESEM	147.678(86)*	.993	.986	.041	.029-.052	-	-	-	-	-
		2-2	Convergent analysis with criterion measures	4049.640(1768)*	.957	.953	.055	.052-.057	-	-	-	-	-
Bifactor-ESEM: Samples[c]	1-2	3-1	Configural invariance	370.128(172)*	.988	.976	.050	.043-.057	-	-	-	-	-
		3-2	Weak (λ) invariance	448.098(242)*	.988	.982	.043	.037-.049	3-1	124.37(70)*	.000	+.006	-.007
		3-3	Strong (λ, ν) invariance	432.228(294)*	.992	.990	.032	.025-.038	3-2	32.53(52)	+.004	+.008	-.011
		3-4	Strict (λ, ν, δ) invariance	473.753(313)*	.990	.989	.033	.027-.039	3-3	45.35(19)*	-.002	-.001	+.001
		3-5	Full (λ, ν, δ, ξ/φ) invariance	397.344(328)*	.996	.996	.021	.012-.029	3-4	19.05(15)	+.006	+.007	-.012
		3-6	Latent mean (λ, ν, δ, ξ/φ, η) invariance	531.211(333)*	.988	.988	.036	.030-.042	3-5	45.74(5)*	-.008	-.008	+.015
Bifactor-ESEM: Sex[d]	2	4-1	Configural invariance	295.672(172)*	.992	.985	.040	.032-.047	-	-	-	-	-
		4-2	λ invariance	358.795(242)*	.993	.990	.032	.025-.039	4-1	105.65(70)*	+.001	+.005	-.008
		4-3	λ,ν invariance	386.744(294)*	.994	.993	.026	.018-.033	4-2	63.22(52)	+.001	+.003	-.006
		4-4	λ, ν, δ invariance	405.706(313)*	.994	.994	.025	.018-.032	4-3	28.78(19)	.000	+.001	-.001
		4-5	λ, ν, δ, ξ/φ invariance	433.240(328)*	.994	.993	.026	.019-.033	4-4	33.06(15)*	.000	-.001	+.001
		4-6	λ, ν, δ, ξ/φ, η invariance	797.282(333)*	.972	.971	.055	.050-.060	4-5	129.37(5)*	-.022	-.022	+.029
Bifactor-ESEM: ED diagnosis[e]	2	5-1	Configural invariance	277.399(172)*	.992	.984	.038	.029-.046	-	-	-	-	-
		5-2	λ invariance	352.457(242)*	.991	.988	.033	.025-.040	5-1	110.25(70)*	-.001	+.004	-.005
		5-3	λ,ν invariance	433.042(294)*	.989	.987	.033	.026-.040	5-2	94.43(52)*	-.002	-.001	.000
		5-4	λ, ν, δ invariance	504.853(313)*	.985	.984	.038	.032-.044	5-3	71.04(19)*	-.004	-.003	+.005
		5-5	λ, ν, δ, ξ/φ invariance	453.341(328)*	.990	.990	.030	.023-.036	5-4	34.48(15)*	+.005	+.006	-.008
		5-6	λ, ν, δ, ξ/φ, η invariance	1148.445(333)*	.937	.935	.076	.071-.080	5-5	230.18(5)*	-.053	-.055	+.046
DIF: Age[f], Age²	2	6-1	MIMIC Null model	168.931(124)*	.997	.996	.020	.011-.027	-	-	-	-	-
		6-2	MIMIC Saturated	238.154(86)*	.991	.979	.044	.037-.051	6-1	42.64(38)	-.006	-.017	+.024
		6-3	MIMIC Invariant	260.749(114)*	.992	.985	.037	.031-.043	6-2	46.24(28)	+.001	+.006	-.007
DIF: BMI[g], BMI²	2	7-1	MIMIC Null model	330.227(124)*	.988	.979	.043	.037-.049	-	-	-	-	-
		7-2	MIMIC Saturated	224.278(86)*	.992	.980	.042	.035-.049	7-1	134.83(38)*	+.004	+.001	-.001
		7-3	MIMIC Invariant	315.179(114)*	.988	.978	.044	.038-.050	7-2	100.74(28)*	+.004	-.002	+.002

Note. χ^2 = robust weighed least square (WLSMV) chi-square ; B = bifactor; BCCS = Body Checking Cognitions Scale; BMI = body mass index; CFA = confirmatory factor analytic model; CFI = comparative fit index; CM = comparison model; df = degrees of freedom; DIF = differential item functioning; ED = eating disorders; ESEM = exploratory structural equation modeling; MIMIC = multiple indicators multiple cause; TLI = Tucker-Lewis index; RMSEA = root mean square error of approximation; 90% CI = 90% confidence interval of the RMSEA; λ = factor loadings; ν = thresholds; δ = Uniquenesses; ξ = factor variances; φ = factor covariances; η = factor means; ΔWχ^2 = WLSMV chi square difference test (calculated with the Mplus DIFFTEST function); Δ = change from the previous model. The fact that WLSMV χ^2 values are not exact, but "estimated" as the closest integer necessary to obtain a correct *p* value explains the fact that the χ^2 and the resulting CFI values can be non-monotonic with model complexity. * $p < .01$.[a]*N*= 487; [b]*N*= 434; [c]*N*= 921;[d] females: *N*= 789, males: *N*= 132;[e] No-diagnosis: *N*= 749, diagnosis: *N*= 109; [f]*N*= 919; [g]*N*= 905.

Table 3.
Latent Factor Correlations from the Convergent Validity Analyses of the BCQ and BCCS[a]

BCQ	BCCS				
	OBEVAL	REASS	SAFEB	BODC	G-factor
OVAPP	-.001	.073	.034	-.449**	.193**
SBP	-.018	-.207**	-.308**	-.019	.033
IDSC	-.005	-.568**	-.020	.101	-.015
G-factor	-.053	-.154**	.073*	.082**	.865**

Note. BCCS = Body Checking Cognitions Scale; BCQ = Body Checking Questionnaire; BODC = body control; G-factor = global factor from a bifactor model; IDSC = idiosyncratic checking; OBEVAL =objective evaluation; OVAPP = overall appearance; REASS = reassurance; SAFEB = safety beliefs; SBP = specific body parts. **p*< .05; ***p* < .01; [a] *N* = 922.

Table 4.
Latent Factor Correlations from the Criterion-Related Validity Analyses of the BCQ and BCCS

Measures	Subscales	BCQ[a]				BCCS[b]				
		OVAPP	SBP	IDSC	G-factor	OBEVAL	REASS	SAFEB	BODC	G-factor
PSI-S	Global self-esteem	-.244*	.165*	.233*	-.680*	-.054	.549*	.352*	.094	-.642*
	Physical appearance	-.040	.236*	.227*	-.599*	.003	.641*	.418*	.044	-.521*
FNAES	-	.350*	-.044	-.234*	.730*	-.148	-.335*	-.299*	-.015	.701*
SPAS	-	.285*	-.085	-.220*	.767*	-.155	-.385*	-.422*	-.089	.758*
EAT-26	Dieting	-.041	-.038	.007	.846*	-.060	-.273*	-.175*	.327*	.833*
	Bulimia-food preoccupation	.054	-.075	.083	.721*	-.152*	-.327*	-.080	.208*	.699*
	Oral control	.232*	.025	.345*	.643*	-.187*	.023	-.027	.356*	.707*

Note. BCCS = Body Checking Cognitions Scale; BCQ = Body Checking Questionnaire; BODC = body control; EAT = Eating Attitudes Test; FNAES = Fear of Negative Appearance Evaluation Scale; G-factor = global factor from a bifactor model; IDSC = idiosyncratic checking; OBEVAL =objective evaluation; OVAPP = overall appearance; PSI-S= short form of the Physical Self-Inventory; REASS = reassurance; SAFEB = safety beliefs; SBP = specific body parts; SPAS = Social Physique Anxiety Scale. **p*< .01; [a] *N* = 406; [b] *N* = 434.

Online Supplements for:

Psychometric Properties of the Body-Checking Questionnaire (BCQ) and of the Body Checking Cognitions Scale (BCCS): A Bifactor-ESEM Approach

S1. Translation and Content Validity of the Translation

The objective of this preliminary study was to examine the content validity of the translated Body Checking Questionnaire (BCQ) and Body Checking Cognitions Scale (BCCS) items in a sample of community adolescents.

Method

Sample and Procedure. A sample of 18 (14 girls and 4 boys) French-speaking Canadian adolescents, aged between 14 and 18 years old (M_{age} = 16.7 years, *SD* = .96) participated in this study. Participants were recruited by the research team in school settings or community organizations located in the Canadian Province of Quebec through a letter of information summarizing the study. Participants completed a written informed consent form in which they also provided their email. Those who consented were then invited by email to answer an online questionnaire (via the LimeSurvey platform).

Measures. The permission to translate and validate the BCQ and BCCS was respectively obtained from Dr. Deborah L. Reas and Dr. Victoria Mountford. The adaptation of the original English version of the BCQ and BCCS into French was done following standardized translation back-translation techniques (Van de Vrijver & Hambleton, 1996). First, the original English items were translated into French by a professional bilingual translator. Second, the translated items were discussed in committee by four authors of this manuscript familiar with cross-linguistic psychometric adaptation procedures. Perceived discrepancies between the translated items and the original items were discussed and resolved by consensus. Third, the approved French items were back-translated into English by a second independent professional bilingual translator. Fourth, the back-translated items were compared with the original English items in committee. Inconsistencies were discussed and resolved by consensus. The final French versions of the BCQ and the BCSS are respectively presented in Tables S1 and S2 of these online supplements. In the current study, the original response scale (BCQ/BCCS: *never* = 1 to *very often* = 5) was replaced by a five-point scale assessing the clarity of the items (*not at all clear* = 1 to *completely clear* = 5). Average clarity scores for items were considered to be unsatisfactory if they fell under four (Gonzales-Reigosa, 1976; Vallerand, 1989), leading to a re-examination of these items.

Results

Average clarity ratings were satisfactory for all BCQ items, with observed scores ranging from 4.50 (*SD* = .62) for Item 6 to 4.78 (*SD* = 0.43) for Item 5. The global mean across the 23 items was 4.65 (*SD* = .09). Similarly, average clarity ratings were also satisfactory for all BCCS items, ranging from 4.00 (*SD* = 1.19) for Item 1 to 4.61 (*SD* = 0.50) for Item 11. The global mean across the 19 items was 4.43 (*SD* = 0.18). These results support the clarity of the translated items.

References

Gonzalez-Reigosa, F. (1976). The anxiety arousing effect of taboo words in bilinguals. In Spielberger C. D., & Diaz-Guerrero R. (Eds.), *Cross-cultural anxiety* (pp. 89-105). Washington, DC: Hemisphere.

Vallerand, R. J. (1989). Vers une méthodologie de validation transculturelle de questionnaires psychologiques : Implications pour la recherche en langue française. *Canadian Psychology, 4,* 662-680.

Van de Vrijver, F.J.R., & Hambleton, R.K. (1996). Translating tests: Some practical guidelines. *European Psychologist*, *1*, 89–99.

Table S1.
French Back-Translated Items from the BCQ[a]

Items	Subscales	French Items
1	SBP	Je vérifie si mes cuisses s'écrasent et prennent plus de place lorsque je suis assis.
2	SBP	Je pince ou tire la peau de mon ventre pour en mesurer le gras.
3	OVAPP	J'essaie certains vêtements en particulier pour m'assurer qu'ils me font encore.
4	IDSC	Je vérifie le tour (diamètre) de mon poignet pour m'assurer qu'il n'a pas changé de taille.
5	OVAPP	Je vérifie mon reflet dans les portes vitrées ou les fenêtres d'auto pour voir de quoi j'ai l'air.
6	SBP	Je pince ou tire la peau du haut de mes bras pour en mesurer le gras.
7	IDSC	Je touche le dessous de mon menton pour m'assurer que je n'ai pas de « double-menton ».
8	OVAPP	Je regarde d'autres personnes pour comparer ma silhouette à la leur.
9	SBP	Je frotte (ou touche) mes cuisses lorsque je suis assis pour vérifier si elles sont grosses.
10	SBP	Je vérifie le tour de mes jambes pour m'assurer qu'elles n'ont pas changé de taille.
11	OVAPP	Je demande leur poids ou leur taille de vêtement aux autres afin de me comparer à eux.
12	OVAPP	Je vérifie l'apparence de mes fesses dans le miroir.
13	OVAPP	Je me pratique à m'asseoir et à rester debout dans différentes positions pour voir de quoi j'ai l'air dans chacune de ces positions.
14	SBP	Je vérifie si mes cuisses se frottent entre elles.
15	OVAPP	J'essaie de susciter des commentaires de la part des autres au sujet de ma grosseur.
16	SBP	Je vérifie si ma graisse ballotte.
17	OVAPP	Je rentre mon ventre pour voir de quoi j'ai l'air quand mon ventre est complètement plat.
18	IDSC	Je vérifie que mes bagues/joncs me font aussi bien qu'avant.
19	SBP	Je vérifie si j'ai de la cellulite sur mes cuisses quand je suis assis.
20	IDSC	Je me couche par terre pour voir si je peux sentir mes os toucher le plancher.
21	OVAPP	Je serre mes vêtements le plus possible autour de moi pour voir à quoi je ressemble.
22	OVAPP	Je me compare aux top-modèles de la télévision ou des magazines.
23	IDSC	Je pince ou tire la peau de mes joues pour en mesurer le gras.
Answer Scale		1 = Jamais 2 = Rarement 3 = Parfois 4 = Souvent 5 = Très souvent

Note. BCQ = Body Checking Questionnaire; IDSC = idiosyncratic checking; OVAPP = overall appearance; SBP = specific body parts.
[a] *for the English items see Reas, D. L., Whisenhunt, B. L., Netemeyer, R., & Williamson, D. A. (2002). Development of the Body Checking Questionnaire: A self-report measure of body checking behaviors. International Journal of Eating Disorders, 31, 324-333.*

Table S2.
French Back-Translated Items from the BCCS[a]

Items	Subscales	French Items
1	BODC	Examiner mon corps aujourd'hui me permet de décider de la quantité de nourriture que je pourrai manger demain.
2	REASS	Examiner mon corps me rassure à propos de ma silhouette.
3	REASS	Examiner mon corps m'aide à me calmer quand je suis anxieux à propos de ma silhouette ou de mon poids.
4	BODC	Examiner mon corps m'aide à contrôler mon poids.
5	REASS	C'est bon pour moi d'examiner mon corps.
6	BODC	Examiner mon corps m'empêche de perdre le contrôle de ce que je mange.
7	REASS	Examiner mon corps m'aide à me sentir mieux.
8	OBEVAL	En examinant mon corps, je peux savoir combien de poids j'ai pris.
9	OBEVAL	Examiner mon corps m'aide à confirmer le poids indiqué par la balance (pèse-personne).
10	OBEVAL	Il faut que j'examine mon corps pour voir comment mon poids se répartit.
11	OBEVAL	Je continue d'examiner mon corps dans l'espoir d'être un jour heureux de mon apparence.
12	BODC	Si j'arrête d'examiner mon corps, mon poids montera en flèche.
13	OBEVAL	Le moyen le plus fiable de juger de mon apparence est d'examiner mon corps.
14	SAFEB	Avant de quitter la maison, je dois m'assurer que mon corps est caché (par mes vêtements) de la manière qui me convient.
15	SAFEB	Si je résiste à la tentation d'examiner mon corps, je me sentirai moins bien.
16	SAFEB	Examiner mon corps m'aide à savoir comment je me sens.
17	SAFEB	J'oublie de quoi j'ai l'air si je ne vérifie pas.
18	SAFEB	Le fait d'examiner mon corps me met plus à l'aise face aux autres.
19	OBEVAL	Examiner mon corps me permet de savoir quand j'ai besoin de faire plus d'exercice.
Answer Scale		1 = Jamais 2 = Rarement 3 = Parfois 4 = Souvent 5 = Très souvent

Note. BCCS = Body Checking Cognitions Scale; BODC = body control; OBEVAL =objective evaluation; REASS = reassurance; SAFEB = safety beliefs.
[a] *for the English items see Mountford, V., Haase, A., & Waller, G. (2006). Body checking in the eating disorders: Associations between cognitions and behaviors. International Journal of Eating Disorders, 39, 708-715.*

Table S3.
Standardized Parameters Estimates from the Confirmatory Factor Analytic and Exploratory Structural Equation Models of the BCQ – Sample 1

	Confirmatory factor analysis				Exploratory structural equation modeling			
Standardized factor loadings (λ) and uniquenesses (δ)								
Items	OVAPP(λ)	SBP(λ)	IDSC(λ)	δ	OVAPP(λ)	SBP(λ)	IDSC(λ)	δ
BCQ3	.665			.558	.193	.230	.333	.564
BCQ5	.667			.556	.766	*.048*	-.162	.481
BCQ8	.841			.293	.832	*.100*	*-.084*	.251
BCQ11	.647			.581	.702	-.191	.202	.519
BCQ12	.660			.564	.657	.197	-.226	.503
BCQ13	.773			.403	.560	*.103*	.175	.419
BCQ15	.673			.547	.686	*-.075*	*.110*	.513
BCQ17	.739			.453	.482	.161	.155	.479
BCQ21	.686			.530	.448	*-.047*	.390	.495
BCQ22	.630			.603	.802	*-.006*	-.191	.506
BCQ1		.868		.247	-.129	1.103	-.143	.133
BCQ2		.723		.477	.330	.356	*.088*	.502
BCQ6		.775		.399	*.044*	.580	.231	.400
BCQ9		.887		.213	-.130	1.027	*-.023*	.160
BCQ10		.838		.297	*-.015*	.682	.254	.288
BCQ14		.785		.383	*.098*	.720	*-.013*	.374
BCQ16		.841		.292	.239	.464	.223	.320
BCQ19		.708		.499	.280	.538	*-.107*	.487
BCQ4			.755	.430	*.057*	*.156*	.629	.406
BCQ7			.780	.391	*.097*	.266	.475	.452
BCQ18			.713	.492	.188	*-.001*	.615	.452
BCQ20			.766	.413	*.107*	*.115*	.633	.391
BCQ23			.855	.270	*.100*	.173	.673	.265
ω	.905	.936	.882		.888	.918	.823	
Factor correlations								
Factor	SBP	IDSC			SBP	IDSC		
OVAPP	-				-			
SBP	.869	-			.784	-		
IDSC	.813	.806	-		.588	.583	-	

Note. BCQ = Body Checking Questionnaire; IDSC = idiosyncratic checking; OVAPP = overall appearance; SBP = specific body parts; Greyscale = main loadings; non-significant parameters are in italics; *ω* = McDonald's omega coefficient of composite reliability. All correlations are statistically significant ($p \leq .001$).

Table S4.
Standardized Parameters Estimates from the Bi-Factor Confirmatory Factor Analytic and Exploratory Structural Equation Models of the BCQ – Sample 1

	Bifactor confirmatory factor analysis					Bifactor exploratory structural equation modeling				
Standardized factor loadings (λ) and uniquenesses (δ)										
Items	OVAPP(λ)	SBP(λ)	IDSC(λ)	G-factor	δ	OVAPP(λ)	SBP(λ)	IDSC(λ)	G-factor	δ
BCQ3	*-.101*			.673	.537	*.025*	*.038*	.155	.641	.563
BCQ5	.389			.601	.488	.241	*-.079*	-.283	.650	.433
BCQ8	.419			.771	.230	.390	*.056*	-.111	.762	.251
BCQ11	.293			.598	.557	.440	*.010*	.224	.551	.452
BCQ12	.302			.608	.540	.216	*.031*	-.305	.623	.471
BCQ13	.161			.741	.425	.253	*.037*	*.078*	.716	.416
BCQ15	.315			.619	.517	.447	*.065*	.149	.572	.447
BCQ17	*.047*			.724	.474	*.096*	*-.037*	*-.039*	.729	.456
BCQ21	*.079*			.668	.548	.247	*-.001*	.293	.621	.467
BCQ22	.450			.552	.493	.486	.101	*-.092*	.524	.471
BCQ1		.584		.732	.123	*.007*	.574	-.090	.729	.131
BCQ2		*.046*		.717	.483	*.005*	*.052*	*-.107*	.723	.464
BCQ6		.175		.747	.412	-.111	.169	*.014*	.765	.373
BCQ9		.503		.762	.167	*-.022*	.508	*-.027*	.761	.162
BCQ10		.288		.784	.302	*.083*	.394	.250	.735	.235
BCQ14		.348		.716	.367	.119	.394	*.013*	.685	.360
BCQ16		*.091*		.827	.307	*.001*	.120	*.012*	.824	.307
BCQ19		.219		.668	.506	.158	.283	*-.095*	.633	.485
BCQ4			.307	.661	.468	*.020*	*.045*	.444	.647	.382
BCQ7			.357	.676	.415	-.186	-.096	*.091*	.756	.377
BCQ18			.240	.631	.544	*.023*	*-.076*	.368	.637	.453
BCQ20			.268	.679	.467	*.141*	*.116*	.546	.612	.294
BCQ23			.612	.729	.094	-.153	-.135	.275	.811	.225
ω	.576	.656	.616	.964		.646	.712	.632	.966	

Note. BCQ = Body Checking Questionnaire; G-factor = global factor from a bifactor model; Greyscale = main loadings; IDSC = idiosyncratic checking; OVAPP = overall appearance; SBP = specific body parts; non-significant parameters are in italics; *ω* = McDonald's omega coefficient of composite reliability.

Table S5.

Standardized Parameters Estimates from the Bifactor Exploratory Structural Equation Model of the BCQ – Sample 2

Items	OVAPP(λ)	SBP(λ)	IDSC(λ)	G-factor	δ
BCQ3	*-.053*	*-.050*	*-.046*	.734	.454
BCQ5	.349	*.016*	*.042*	.631	.478
BCQ8	.304	.144	*.068*	.758	.307
BCQ11	.375	.141	*.055*	.656	.407
BCQ12	.335	.291	*-.036*	.566	.481
BCQ13	.258	*.001*	.113	.651	.497
BCQ15	.223	-.105	-.163	.711	.407
BCQ17	-.138	*-.066*	*-.048*	.777	.371
BCQ21	.149	*.044*	*.059*	.727	.444
BCQ22	.330	.126	*.010*	.610	.503
BCQ1	*.058*	.578	*-.009*	.690	.186
BCQ2	-.186	*.033*	*.068*	.776	.358
BCQ6	-.083	.099	.216	.767	.349
BCQ9	*.047*	.544	.120	.778	.082
BCQ10	.188	.235	*.019*	.765	.323
BCQ14	*.062*	.222	-.159	.774	.323
BCQ16	-.133	*-.050*	-.142	.816	.294
BCQ19	*.059*	.354	-.178	.626	.448
BCQ4	*.089*	*.033*	.160	.724	.440
BCQ7	*-.048*	*-.024*	.364	.665	.422
BCQ18	*-.107*	-.297	*.006*	.694	.419
BCQ20	*-.010*	*.081*	*.098*	.805	.335
BCQ23	.093	*-.008*	.673	.648	.118
ω	.592	.654	.494	.969	

Note. BCQ = Body Checking Questionnaire; G-factor = global factor from a bifactor model; Greyscale = main loadings; IDSC = idiosyncratic checking; OVAPP = overall appearance; SBP = specific body parts; non-significant parameters are in italics; *ω* = McDonald's omega coefficient of composite reliability.

Table S6.

Standardized Parameters Estimates from the Confirmatory Factor Analytic and Exploratory Structural Equation Models of the BCCS – Sample 1

	Confirmatory factor analysis					Exploratory structural equation modeling				
Standardized factor loadings (λ) and uniquenesses (δ)										
Items	OBEVAL(λ)	REASS(λ)	SAFEB(λ)	BODC(λ)	δ	OBEVAL(λ)	REASS(λ)	SAFEB(λ)	BODC(λ)	δ
BCCS8	.709				.498	.813	*.034*	-.183	.128	.311
BCCS9	.629				.605	.673	*.038*	-.160	.141	.494
BCSS10	.815				.336	.512	*-.049*	.384	*.079*	.314
BCSS11	.720				.481	.427	-.238	.474	*.061*	.347
BCSS13	.750				.437	.351	.209	.386	*.052*	.438
BCSS19	.714				.490	.342	*.063*	.101	.337	.508
BCCS2		.804			.354	-.190	.770	.118	.214	.298
BCCS3		.834			.304	-.208	.605	.290	.258	.398
BCCS5		.747			.442	.272	.733	*-.032*	-.119	.404
BCCS7		.774			.401	.115	.850	*-.014*	*-.078*	.274
BCSS14			.721		.480	.335	-.120	.463	*.104*	.434
BCSS15			.864		.254	.214	.071	.606	.162	.251
BCSS16			.795		.367	.249	.098	.505	.129	.385
BCSS17			.464		.784	*-.049*	*.001*	.363	.240	.752
BCSS18			.612		.626	.119	.478	.501	-.148	.454
BCCS1				.792	.373	*.072*	*-.068*	.162	.673	.351
BCCS4				.797	.365	.139	.331	-.091	.631	.306
BCCS6				.804	.353	*.081*	.094	-.127	.843	.247
BCSS12				.869	.244	.244	-.160	.169	.617	.226
ω	.869	.869	.826	.889		.801	.864	.723	.871	
Factor correlations										
Factor	REASS	SAFEB	BODC			REASS	SAFEB	BODC		
OBEVAL	-					-	-			
REASS	.382	-				.177	-			
SAFEB	.905	.517	-			.506	.181	-		
BODC	.873	.478	.788	-		.680	.255	.516	-	

Note. BCCS = Body Checking Cognitions Scale; BODC = body control; OBEVAL =objective evaluation; REASS = reassurance; SAFEB = safety beliefs; Greyscale = main loadings; non-significant parameters are in italics; *ω* = McDonald's omega coefficient of composite reliability. All correlations are statistically significant ($p \leq .001$).

Table S7.
Standardized Parameters Estimates from the Bifactor Confirmatory Factor Analytic and Exploratory Structural Equation Models of the BCCS– Sample 1

	Bifactor confirmatory factor analysis						Bifactor exploratory structural equation modeling					
Standardized factor loadings (λ) and uniquenesses (δ)												
Items	OBEVAL(λ)	REASS(λ)	SAFEB(λ)	BODC(λ)	G-factor	δ	OBEVAL(λ)	REASS(λ)	SAFEB(λ)	BODC(λ)	G-factor	δ
BCCS8	.485				.662	.326	.527	*-.008*	*-.039*	.154	.609	.326
BCCS9	.497				.575	.422	.476	*.030*	-.197	*.077*	.550	.425
BCSS10	*.004*				.803	.355	.186	-.095	*.070*	*-.072*	.798	.309
BCSS11	*.066*				.714	.486	*.100*	-.269	*.024*	-.152	.750	.332
BCSS13	*.055*				.748	.437	.101	.142	.225	*-.014*	.699	.431
BCSS19	*.098*				.699	.501	.184	*.017*	.100	.213	.641	.499
BCCS2		.722			.375	.338	-.084	.759	*-.068*	*.045*	.353	.286
BCCS3		.597			.470	.422	-.213	.622	-.193	*-.063*	.517	.259
BCCS5		.645			.359	.455	.181	.666	.266	.101	.265	.372
BCCS7		.790			.309	.280	.075	.791	.157	*.069*	.240	.282
BCSS14			.126		.686	.513	*.053*	-.157	*.075*	-.091	.726	.431
BCSS15			.467		.788	.162	*-.060*	*.000*	.256	*-.017*	.830	.242
BCSS16			.330		.728	.360	*-.004*	*.031*	.242	*.001*	.751	.376
BCSS17			.192		.432	.777	-.145	*-.025*	*.117*	*.053*	.465	.746
BCSS18			.162		.568	.652	*-.084*	.405	.339	-.124	.517	.431
BCCS1				.273	.718	.410	*-.013*	.066	-.140	.287	.748	.334
BCCS4				.318	.704	.403	.122	.298	*.063*	.463	.624	.288
BCCS6				.673	.682	.082	.088	.075	*-.025*	.568	.661	.226
BCSS12				.241	.797	.307	.092	-.169	-.080	.267	.811	.227
ω	.365	.835	.398	.653	.948		.516	.870	.322	.700	.951	

Note. BCCS = Body Checking Cognitions Scale; BODC = body control; G-factor = global factor from a bifactor model; OBEVAL =objective evaluation; REASS = reassurance; SAFEB = safety beliefs; Greyscale = main loadings; non-significant parameters are in italics; *ω* = McDonald's omega coefficient of composite reliability.

Table S8.
Standardized Parameters Estimates from the Bifactor Exploratory Structural Equation Models of the BCCS – Sample 2

Items	OBEVAL(λ)	REASS(λ)	SAFEB(λ)	BODC (λ)	G-factor	δ
BCCS8	.543	*.022*	*-.020*	.138	.675	.230
BCCS9	.408	*.030*	-.143	*.051*	.644	.395
BCSS10	*-.083*	*-.038*	*-.021*	*.032*	.800	.350
BCSS11	-.133	-.244	*-.014*	-.087	.776	.313
BCSS13	*.032*	*.020*	.188	-.148	.723	.419
BCSS19	*-.051*	*.033*	*-.025*	.207	.685	.484
BCCS2	*-.080*	.783	*-.046*	*.046*	.231	.323
BCCS3	*-.050*	.774	-.240	*-.026*	.429	.156
BCCS5	.171	.519	.317	.121	.299	.497
BCCS7	.087	.826	.229	*.077*	.214	.207
BCSS14	*-.012*	-.284	*-.089*	-.127	.703	.401
BCSS15	*-.004*	*-.047*	.184	*.037*	.838	.260
BCSS16	*-.006*	*.039*	.158	*-.039*	.818	.302
BCSS17	-.203	*.049*	*.166*	*-.049*	.453	.721
BCSS18	-.109	.451	.222	-.127	.462	.506
BCCS1	*-.017*	-.072	*-.080*	.334	.782	.265
BCCS4	.137	.279	*-.041*	.470	.658	.248
BCCS6	.080	.167	*.013*	.405	.742	.251
BCSS12	*-.009*	-.151	*.077*	.213	.830	.237
ω	.416	.877	.234	.669	.955	

Note. BCCS = Body Checking Cognitions Scale; BODC = body control; G-factor = global factor from a bifactor model; OBEVAL =objective evaluation; REASS = reassurance; SAFEB = safety beliefs; Greyscale = main loadings; non-significant parameters are in italics; *ω* = McDonald's omega coefficient of composite reliability.

Table S9.
Latent Factor Correlations from the Criterion-Related Validity Analyses of the BCQ with a First Order ESEM Solution

Measures	Subscales	BCQ		
		OVAPP	SBP	IDSC
PSI-S	Global self-worth	-.753*	-.483*	-.481*
	Physical appearance	-.606*	-.366*	-.509*
FNAES	-	.834*	.597*	.442*
SPAS	-	.844*	.603*	.512*
EAT-26	Dieting	.763*	.686*	.723*
	Bulimia & Food preoccupation	.670*	.572*	.626*
	Oral control	.628*	.585*	.550*

Note. BCQ = Body Checking Questionnaire; EAT = Eating Attitudes Test; FNAES = Fear of Negative Appearance Evaluation Scale; IDSC = idiosyncratic checking; OVAPP = overall appearance; PSI-S = short form of the Physical Self-Inventory; SBP = specific body parts; SPAS = Social Physique Anxiety Scale. *$p < .01$.

Table S10.
Latent Factor Correlations from the Criterion-Related Validity Analyses of the BCCS with a First Order ESEM Solution

Measures	Subscales	BCCS			
		OBEVAL	REASS	SAFEB	BODC
PSI-S	Global self-worth	-.591*	.396*	-.626*	-.490*
	Physical appearance	-.456*	.527*	-.567*	-.411*
FNAES	-	.559*	-.171*	700*	.565*
SPAS	-	.601*	-.225*	.797*	.572*
EAT-26	Dieting	.720*	-.051	.635*	.866*
	Bulimia & Food preoccupation	.564*	-.122	.617*	.680*
	Oral control	.554*	.218*	.531*	.766*

Note. BCCS = Body Checking Cognitions Scale; BODC = body control; EAT = Eating Attitudes Test; FNAES = Fear of Negative Appearance Evaluation Scale; OBEVAL =objective evaluation; PSI-S= short form of the Physical Self-Inventory; REASS = reassurance; SAFEB = safety beliefs; SPAS = Social Physique Anxiety Scale. *p < .01.