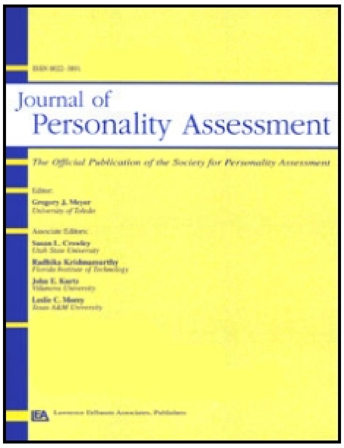

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Construct Validity of the Social Physique Anxiety Scale in a French Adolescent Sample

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We conducted a series of studies to investigate the psychometric properties of the Social Physique Anxiety Scale (SPAS; Hart, Leary, & Rejeski, 1989) among a total sample of 1,563 nonclinical French adolescents. The, first study provided support for the item content of the preliminary version of the French SPAS for adolescents. Then, the second study supported the convergent validity of the English and French version of the SPAS. Finally, Studies 3 to 6 provided support for the factor validity, measurement invariance (across genders, age groups and samples), latent mean structure invariance (across age category and samples), reliability, convergent validity, and criterion-related validity of a 7-item version of the SPAS.

Today, many adolescents are concerned with how their body shape appears to others (Thompson & Chad, 2002). Adolescents tend to have a preference for being perceived positively or as attractive to others. Many experience distress and negative thoughts when they perceive that their physicality is unlikely to be positively evaluated by others (Thompson & Chad, 2002). These adolescents can become anxious when displaying their physique in social settings laden with evaluation potential (Scott, Burke, Joyner, & Brand, 2004). Some adolescents may even feel compelled to avoid social situations and activities involving the prospect of body shape or physique evaluation. The term *social physique anxiety* (SPA) has been created to describe this phenomenon (Scott et al., 2004) and refers to a subtype of social anxiety representing the tendency to become apprehensive or nervous when others can observe or evaluate one's physique or body shape (Hart, Leary, & Rejeski, 1989). The extent to which physique-related anxiety among adolescents is associated with a variety of undesirable behaviors and attributes (e.g., low self-esteem, social avoidance, physical inactivity, or disordered eating) is unknown, but evidence from adult populations suggests that the possibility warrants investigation (Thompson & Chad, 2002).

Hart et al. (1989) did the seminal work in conceptualizing SPA and developing a measure of the degree to which individuals experience this anxiety by creating the 12-item unidimensional SPA scale (SPAS). The psychometric properties of this instrument were first investigated in a North American adult sample of 438 participants (43.8% males). Their results revealed that the SPAS presented: (a) acceptable internal consistency ($\alpha =$ .90); (b) moderate correlations with several convergent measures such as body satisfaction ($r = -.51$, $p < .01$), fear of negative evaluation ($r = .35$, $p < .01$), and interaction anxiousness ($r = .33$, $p < .01$); and (c) predictive validity by showing that higher results on the SPAS predicted higher levels of stress and more negative thoughts about their body for women during a physical fitness examination. Additional results regarding the convergent validity of the SPAS have been reported in additional studies of nonclinical samples of adults; they have reported moderate negative correlations between the SPAS and self-esteem (Diehl, Johnson, Rogers, & Petrie, 1998; Russell, 2002), body satisfaction (Crawford & Eklund, 1994; Martin, Rejeski, Leary, McAuley, & Bane, 1997; Petrie, Diehl, Rogers, & Johnson, 1996), physical self-perceptions (Kowalski, Crocker, & Kowalski, 2001; Petrie et al., 1996), social anxiety (Atalay & Gençöz, 2008), and disturbed eating attitudes and behaviors (Diehl et al., 1998; Haase & Prapavessis, 1998).

Although extensive research has not been conducted on SPA other than with adults, similar results have been found in samples of children and adolescents (aged 7–18; Baş, Asc1, Karabudak, & Kiziltan, 2004; Hausenblas & Mack, 1999; Smith, 2004; Thompson & Chad, 2002) including moderate of associations between the SPAS, body satisfaction, and disturbed eating attitudes and behaviors. Some studies have even suggested that the SPAS may represent one of the strongest predictors of disturbed eating attitudes and behaviors among adults (Diehl et al., 1998) and youths (Thompson & Chad, 2002). In fact, Mack, Strong, Kowalski, and Crocker (2007) recently reported that

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Because both Christophe Maïano and Alexandre J. S. Morin contributed equally to the preparation of this article, the order of appearance of the first and second authors (C. Maïano and A. J. S. Morin) was determined at random: Both should be considered first authors.

This manuscript was prepared while A. J. S. Morin was a visiting scholar at the University of Aix-Marseille II.

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SPA levels were significantly higher in adult females diagnosed to be suffering from clinical eating disorders (i.e., anorexia nervosa, bulimia, and not otherwise specified) or at risk of developing eating disorders (i.e., high levels of drive for thinness) than in adult females not included in these diagnostic categories.

The factor structure of the SPAS has been extensively investigated with confirmatory factor analyses (CFA) across English (Eklund, Kelley, & Wilson, 1997; Eklund, Mack, & Hart, 1996; Hagger et al., 2007; Martin et al., 1997; Motl & Conroy, 2000, 2001; Petrie et al., 1996) and non-English (Hagger et al., 2007; Isogai et al., 2001; Lindwall, 2004) samples of adults. Unfortunately, none of these studies have successfully replicated the 12-item unidimensional model of Hart et al. (1989). In fact, in the late 1990s, researchers (Eklund, 1998; Eklund et al., 1997, 1996) had noted that although the SPAS was conceptualized as unidimensional, the original 12-item version often proved multidimensional and yielded two 1st-order factors: Physique Presentation Comfort (PPC) and expectation of Negative Physical Evaluation (NPE). These authors (Eklund, 1998; Eklund et al., 1997, 1996) have argued that this was likely a methodological artifact stemming from presence of both positively (PPC) and negatively (NPE) worded items. To resolve this issue, Eklund (1998; Eklund et al., 1997, 1996) and Martin et al. (1997) have recommended three alternative strategies: (a) removing all of the positively worded items (i.e., 1, 2, 5, 8, and 11); (b) removing Items 1, 5, 8, and 11 and rephrasing Item 2 negatively; or (c) removing the three most conceptually misfitting or troublesome items (i.e., 1, 2, and 5). Subsequent analyses in samples of adults have confirmed the presence of a method effect in the original 12-item model (Hagger et al., 2007; Motl & Conroy, 2000; Motl, Conroy, & Horan, 2001) and the fact that this artifact could be resolved by excluding one or more of the positively worded items. To this end, Motl and Conroy (2000, 2001) have proposed to identify and remove the most troublesome items on the basis of their content and standardized residuals covariances (i.e., exceeding |2.58|).

Following these recommendations, recent studies in samples of adults have succeeded in identifying cleaner, unidimensional models. However, they have mostly generated and tested truncated SPAS versions containing seven (Hagger et al., 2007; Isogai et al., 2001; Lindwall, 2004; Motl & Conroy, 2000, 2001), eight (Hagger et al., 2007), or nine items (Isogai et al., 2001; Martin et al., 1997). An examination of these studies also revealed that only three of them had evaluated the measurement invariance of the SPAS across gender, using the same set of items (i.e., seven). However, the results from these studies have been inconsistent and preclude clear conclusions regarding the possibility of interpreting SPAS results similarly across genders. Indeed, Motl and Conroy (2000, 2001) have found support for the gender-based measurement invariance of the SPAS and for latent means differences (with women scoring significantly higher than men), whereas Lindwall (2004) failed to replicate these results. However, it should be noted that the fact that Lindwall (2004) rely on a Swedish sample adds culture as a potential alternative explanation of their divergent results.

However, only two psychometric evaluations of the SPAS have been conducted with adolescents and both have relied on North American samples. The first of those studies (McAuley & Burman, 1993) tested the factor structure of the SPAS using an adolescent female sample (i.e., $n = 236$; 12–18 years old; $M_{\text{age}} = 14.44$, $SD_{\text{age}} = 1.80$) and failed to replicate Hart et al.'s (1989) original unidimensional structure. Recently, Smith (2004) reanalyzed the factor structure of the SPAS in a mixed-gender sample of 614 high school students ($M_{\text{age}} = 15.20$, $SD_{\text{age}} = 0.70$). Their results support the validity and the gender-based strict measurement invariance of a truncated nine-item version of the SPAS and showed significant latent mean differences across genders, with girls scoring significantly higher than boys.

In summary, substantial empirical support has been found for a variety of truncated unidimensional models in samples of adults. Preliminary results also suggest that these measurement models appear to be invariant across genders but that the latent means seem higher in women. Conversely, psychometric evidence regarding the psychometric properties of the SPAS in samples of adolescents is sorely lacking and relies on a reduced number of studies that present several limitations. First, none of them did replicate Hart et al.'s (1989) factor structure, and the results from the sole study that verified the psychometric properties of a truncated nine-item SPAS version have yet to be replicated in an independent sample. Second, these studies have been confined to English-speaking samples. Consequently, the generalizability of the SPAS adolescents from other cultures and from non-English-speaking countries remains an open question. The analysis of the factor structure of the SPAS in another country, such as France, would ensure that the SPA construct is not biased by colloquialisms and idiosyncrasies of its original language. Third, it is currently unknown whether the measurement and latent mean structure of the SPAS is invariant between boys and girls, preliminary evidence on this topic being limited to a single study. Fourth, none of the preceding studies have examined the measurement and latent mean invariance of the SPAS across age categories (i.e., early and late adolescence). This lack of research is surprising given the fact that this developmental period is characterized by important physical changes and by an increase in peers' influence that may both modify the meaning of the SPA construct. Fifth, current evaluations of the convergent validity of the SPAS were limited to measures of body satisfaction and disturbed eating attitudes and behaviors. Thus, evidence regarding the convergent validity of the SPAS with additional measures already tested in adults (i.e., self-esteem, fear of negative evaluation, social anxiety) remains unknown. Finally, despite the recent study conducted by Mack et al. (2007) on a sample of adults, the criterion-related validity of the SPAS (i.e., its ability to differentiate between community and clinical samples presenting an elevated level of body dissatisfaction) remains an open question.

The objectives of these series of studies were to: (a) develop a French version of the SPAS and test its applicability in youths, (b) examine the construct validity of this instrument (i.e., reliability, factor validity, measurement and latent mean structure invariance, convergent and criterion-related validity), and (c) cross-validate the results in an independent sample.

STUDY 1

The purpose of this study was to develop a preliminary version of the French SPAS and to verify its content clarity in a diversified sample of French youths.

TABLE 1.—Descriptive statistics for all samples.

Study and Sample	N	Age M	Age SD	BMI M	BMI SD
Study 1					
Total	30	10.80	1.56		
Boys	15	10.73	1.62		
Girls	15	10.87	1.55		
Study 2					
Total	23	33.78	13.49		
Males	12	37.25	13.20		
Females	11	30.00	13.37		
Study 3					
Total	678	15.30	1.84	20.24	2.79
Boys	297	15.16	1.84	20.72	2.62
Girls	381	15.41	1.84	19.87	2.64
Test–retest sample					
Total	23	16.57	0.95	20.01	2.38
Boys	12	16.50	0.90	19.60	2.29
Girls	11	16.64	1.03	20.47	2.50
Study 4					
Total	670	15.36	1.83	20.27	3.11
Boys	295	15.19	1.84	20.34	2.87
Girls	375	15.49	1.81	20.22	3.28
Study 5					
Total	119	14.55	2.28	20.26	2.98
Boys	70	14.04	2.21	20.37	3.36
Girls	49	14.90	2.27	20.09	2.24
Study 6					
Total	66	15.79	1.41	18.05	3.28
Anorexic	33	15.70	1.38	15.96	2.55
Nonclinical	33	15.88	1.45	20.14	2.52

Note. BMI = body mass index; BMI was calculated on the basis of the adolescents' self-reported weight and height and the following formula: weight/height × height (Cole, 1979).

TABLE 2.—Items of the French version of the social physique anxiety scale.

1. Je suis satisfait(e) de mon apparence physique[a] (*I am comfortable with the appearance of my physique/figure*)	
2. Cela me gênerait de porter des vêtements qui m'aminciraient ou qui me grossiraient (*I would worry about wearing clothes that might make me look too thin or overweight*)	SPAS–7
3. J'aimerais ne pas être aussi nerveux(se), au sujet de mon apparence physique (*I wish I wasn't so uptight about my physique/figure*)	SPAS–7
4. Parfois cela m'inquiète que d'autres personnes puissent évaluer négativement mon poids ou mon développement musculaire (*There are times when I am bothered by thoughts that other people are evaluating my weight or muscular development negatively*)	SPAS–7
5. Lorsque je me regarde dans le miroir, je me sens bien dans mon corps[a] (*When I look in the mirror, I feel good about my physique/figure*)	
6. Certaines parties de mon corps que je n'aime pas, me rendent nerveux(se) dans certains contextes sociaux (*Unattractive features of my physique/figure make me nervous in certain social settings*)	SPAS–7
7. En présence d'autres personnes, je m'inquiète au sujet de mon apparence physique (*In the presence of others, I feel apprehensive about my physique/figure*)	SPAS–7
8. Je suis sûr de paraître en forme aux yeux des autres[a] (*I am comfortable with how fit my body appears to others*)	
9. Je me sentirais mal à l'aise de savoir que les autres jugent mon corps (*It would make me uncomfortable to know others were evaluating my physique/figure*)	SPAS–7
10. Je suis une personne timide et je n'aime pas montrer mon corps aux autres (*When it comes to displaying my physique/figure to others, I am a shy person*)	
11. Je suis détendu(e), lorsqu'il est évident que les autres regardent mon corps[a] (*I usually feel relaxed when it is obvious that others are looking at my physique/figure*)	
12. Lorsque je porte un maillot de bain, je me sens souvent nerveux(se) à propos de ma silhouette (*When in a bathing suit, I often feel nervous about the shape of my body*)	SPAS–7

Note. SPAS = social physique anxiety scale; SPAS–7 = items from the SPAS that were retained in this French version for adolescents.
[a]Reversed score.

Method

Sample. Boys ($n = 15$) and girls ($n = 15$) aged between 9 and 13 years were recruited from one elementary and one middle school located in Southern France. This age bracket was chosen to develop a questionnaire that is easily understandable and accessible to a wider audience comprising older children, "normal" adolescents, and adolescents with reading/learning difficulties (who sometimes present a comprehension level similar to older children; Harter, 1999). Although the objective of these series of studies was to develop and validate an adolescent version of the SPAS, we also decided to include younger children to ensure that future researchers interested in the SPAS validity in older children or in adolescents with reading/learning difficulties will be able to start from the same set of items. The descriptive statistics of this sample and of the samples used in the following studies are presented in Table 1.

Measure. The original version of the SPAS was translated into French following the standardized back-translation techniques widely described in the literature (Brislin, 1986; Van de Vrijver & Hambleton, 1996). Initially, translation from English into French was done separately by two bilingual researchers (C. Maïano, A. J. S. Morin) and a bilingual translator. Translation discrepancies between the three translated forms were subsequently discussed to develop an initial French version of the inventory. A second bilingual translator whose native language was English and who had not seen the original English version translated this initial French version back into English. The back-translated version was then compared with the original English version and any inconsistencies and incongruence were highlighted. The translation process was repeated until the back-translated versions were equivalent to the original English version. As suggested by Eklund et al. (1997), Item 2 was also changed from a positively worded item to a negatively worded item. The 12 items (see Table 2) are answered on a 5-point Likert scale ranging from 1 (*not at all*) to 5 (*extremely*).

Procedure. After agreeing to participate and returning the informed consent forms signed by their parents, the participants completed the preliminary version of the SPAS in standardized conditions (i.e., isolation, quiet classroom conditions, groups of up to 10 students, and help for reading if necessary). The original response format was replaced by a 5-point Likert scale assessing the clarity of the items and ranging from 1 (*not at all clear*) to 5 (*completely clear*). The participants were invited to respond as honestly as possible to the questions. Following the completion of the questionnaire, we used individual interviews to investigate how unclear items could be clarified. All of the studies reported in this manuscript were approved by the local ethics committee.

Results and Discussion

Analyses of the clarity of the items from the preliminary French version were performed following Gonzales-Reigosa (1976) and Vallerand's (1989) suggestions including the recommendation that an item clarity score of less than 4 (on a 5-point scale) should be considered unsatisfactory. In this study, we considered all items satisfactory, with observed scores ranging from a low of 4.50 ($SD = 0.51$) for Item 2 through a high of 4.83 ($SD = 0.38$) for Item 1. These results thus support to the appropriateness of the translated items for older children and adolescents.

Study 2

The objective of this study was to test the convergent validity of the resulting French version with the original English version in a bilingual sample of adults (see Behling & Law, 2000; Brislin, 1986).

Method

Sample. Adult males ($n = 12$) and females ($n = 11$) were recruited at the University of Nice Sophia-Antipolis and Montpellier in Southern France (see Table 1) for this study because in France, bilingual youths are notably hard to recruit, whereas a bilingual university sample could be conveniently obtained. As suggested by Gonzales-Reigosa (1976) and Vallerand (1989), the degree of participants' bilingualism was evaluated through four items rated on a 4-point Likert scale ranging from 1 (*not at all true*) to 4 (*completely true*) on which they were asked to indicate if they could (in English and French separately): (a) read, (b) write, (c) understand a conversation, and (d) express themselves. The degree of bilingualism of the participants was satisfactory ($M_{\text{French}} = 14.04$, $SD = 1.82$; $M_{\text{English}} = 14.30$, $SD = 1.64$; all individuals scores > 12 for both languages), and no differences were found between languages, $t(22) = -.514$, $p = .61$. All participants were thus characterized by a homogeneous and sufficient degree of bilingualism to complete the questionnaires.

Measures and procedure. We used the original English version and the aforementioned preliminary French version of the SPAS in this study. The first 12 participants completed the English version and then, half an hour later, the preliminary French version. The other participants followed the opposite order. The items were presented in a different order across the two versions. Items were rated on the original SPAS 5-point Likert scale ranging from 1 (*not at all*) to 5 (*extremely*).

Results and Discussion

We performed correlational analyses to evaluate the relationships between the scores on the original and translated items as well as the full-scale scores. We observed correlation coefficients that were very strong (rs ranging from .81–.96) and statistically significant (ps $< .001$). Student t tests for matched samples also revealed no significant differences between English and French responses to individual all items and the full scale score ($p > .05$). Consequently, we considered the French and English versions comparable.

Study 3

The objectives of this study were to examine the factor validity and reliability of the preliminary version of the French SPAS in adolescent males and females through the examination of the relative fit of the alternative models reported in the literature and to assess the measurement and latent mean structure invariance of the SPAS across gender and age.

Method

Sample, measure, and procedure. Adolescents ($n = 678$) aged 11 to 18 years (see Table 1) were recruited from four middle and high schools located in Southern France. Inclusion criteria for recruitment required: (a) no self-reported history of eating disorders and obesity and were neither underweight, overweight, or obese at the time of the study (according to body mass index cut-off scores for boys and girls provided by Cole, Bellizzi, Flegal, & Dietz, 2000, and Cole, Flegal, Nicholls, & Jackson, 2007); (b) being schooled in regular classes and thus presented no intellectual, motor, or sensory disability (according to the French education policies); and (c) never having repeated a school year according to self-reports. We retested a subgroup of 23 of these adolescents (12 boys and 11 girls) after a 2-week period to obtain data for evaluation of instrument test–retest reliability. After agreeing to participate and returning signed parental informed consent forms, the participants completed the preliminary version of the SPAS developed in Study 1 in standardized conditions.

Data analysis. In this study, we performed the CFAs using bootstrapped maximum likelihood (ML) estimation with AMOS 7.0 (Arbuckle, 2006), given the significant non-normality of the data (normalized coefficients values for kurtosis: 29.57). Consequently, all fit indexes were based on the Bollen–Stine bootstrap (Fouladi, 2000; Yuan & Hayashi, 2003) p value and bootstrap adjusted chi-square and goodness-of-fit indexes. We identified models by fixing the loading of a single item for a factor to 1. Assessment of fit of the CFA models was based on multiple indicators: the chi-square statistic, the comparative fit index (CFI), the Tucker-Lewis Index (TLI; Tucker & Lewis, 1973), the standardized root mean square residual (SRMR), the root mean square error of approximation (RMSEA), and the 90% confidence interval (CI) of the RMSEA (RMSEA 90% CI). Values greater than .90 for CFI and TLI were considered to be indicative of adequate model fit, although values greater than .95 are preferable (Byrne, 2005; Hu & Bentler, 1999). Values smaller than .08 or .05 for the RMSEA and smaller than .10 and .08 for the SRMR support, respectively, acceptable and good model fit (Hu & Bentler, 1999). Concerning the RMSEA 90% CI, values less than .05 for the lower bound (left side) and less than .08 for the upper bounds (right side) or containing 0 for the lower bound and less than .05 for the upper bounds (right side) provide, respectively, acceptable and good model fit (MacCallum, Browne, & Sugawara, 1996). We evaluated critical values for the tests of gender and age measurement and latent mean structure invariance in CFAs relative to several criteria: χ^2 difference tests, and CFI and RMSEA changes resulting from the application of additional invariance constraints (Chen, 2007; Cheung & Rensvold, 2002). A CFI difference of .01 or less and RMSEA differences of .015 or less between a reference model and the resulting model indicate that the measurement invariance hypothesis should not be rejected. Finally, we computed reliability from the model's standardized parameters estimates using the formula provided by McDonald (1999): $\omega = (\Sigma\lambda i)^2/([\Sigma\lambda i]^2 + \Sigma\delta ii)$, where λi are the factor loading and δii the error variances.

Results

We examined seven a priori CFA measurement models from the extant literature: (a) Model 1: the original 12-item model postulated by Hart et al. (1989) with the factor loading of Item 1 fixed to 1.0; (b) Model 2: the 9-item model identified by Martin et al. (1997) and examined by Isogai et al. (2001) and Smith (2004) containing Items 3, 4, 6, 7, 8, 9, 10, 11, and 12; (c) Model 3: the 8-item model identified by Hagger et al. (2007) for the British, Estonian, and Swedish samples containing Items 2, 3, 4, 6, 7, 9, 10, and 12; (d) Model 4: the 7-item model employed by Hagger et al. (2007; for a Spanish sample) and Isogai et al. (2001) containing Items 3, 4, 6, 7, 9, 10, and 12; (e) Model 5: the 7-item model of that has been identified by Motl and Conroy (2000, 2001) and employed by Lindwall (2004) containing Items 3, 4, 6, 7, 8, 9, and 10; (f) Model 6: the 7-item model identified by Hagger et al. (2007) for a Turkish sample containing Items 2, 3, 4, 6, 9, 10, and 12; and (g) Model 7: the 12-item, two-factor model evaluated by Eklund et al. (1996) and Petrie et al. (1996), with the Items 1, 5, 8, and 11 assigned to the PPC factor and Items 2, 3, 4, 6, 7, 9, 10, and 12 assigned to the NPE factor.

The goodness-of-fit statistics and parameters estimates of the various CFA models tested for the SPAS are displayed in Table 3. They showed that four of the estimated models (1, 2, 6, and 7) exhibited significant bootstrapped χ^2 values, CFIs, and/or TLIs under .90; RMSEAs exceeding the .08 criterion; and SRMR indicator between .05 and .11. For the other three CFA models (3, 4, and 5), the results also showed significant bootstrapped χ^2 values and RMSEAs values over .08 but CFIs and TLIs values exceeding .90 and SRMRs under .06. In addition, we did not observe satisfactory fits for any of four method effects measurement models (with correlated uniquenesses or method factors underlying the positively or negatively worded items) tested to evaluate Eklund et al.'s (1997) hypothesis that methodological artifacts might influence SPAS answers (these results are available on request from C. Maïano).

Given the unsatisfactory fit of the various a priori models, we deemed them all inadequate for French adolescents. Consequently, we used the original SPAS model (Hart et al., 1989) as a starting point and modified it via an iterative process following Motl and Conroy (2000) recommendations, which involve removing the problematic items responsible for the misspecification in the models and then rerunning the analyses. Following these guidelines, problematic items should be removed based on large standardized residuals covariances (i.e., exceeding |2.58|) and substantive arguments concerning item content (i.e., redundancy and salience). Examination of the standardized residuals covariances showed that five items (1, 5, 8, 10, and 11) were problematic. We performed a final CFA (Model 8) on the remaining seven items, and this measurement model was supported by all of the fit indexes. These items formed a scale with an average score of 18.10 ($SD = 6.78$) and an acceptable level of internal consistency ($\omega = .87$). In addition, the temporal stability of this seven-item version, estimated using test–retest reliability correlations uncorrected for measurement errors on the subsample that was reevaluated after 2 weeks, also proved satisfactory ($r = .78$, $p < .05$).

Next, we evaluated the measurement and latent mean structure invariance of this seven-item CFA model across gender and age groups. In several age categories, the sample sizes were small, so age was "dichotomized" in two groups: 11 to 14 years old and 15 to 18 years old. We first estimated CFA models separately in all gender- and age-related subsamples (Models 8A and 8C), and then we performed measurement invariance tests across gender and age groups (Models 8B and 8D) in the sequential order recommended by Byrne (2004): (a) configural invariance (reference model), (b) factor loadings invariance, (c) factor variance invariance, and (d) item uniquenesses invariance. Finally, we performed the invariance of the latent mean structure of the CFA model across gender and age groups (Models 8E and 8F) in the sequential order recommended by Byrne, Shavelson, and Muthén (1989): (a) factor loadings invariance

TABLE 3.—Goodness-of-fit statistics of SPAS models.[a]

Study	No. of Items	Model	Number of Factors	χ^2(B–S)	*df*	CFI	TLI	SRMR	RMSEA	RMSEA 90% CI	$\lambda(\delta)$
Study 3[b]	12	1	1	66.585*	54	.747	.690	.113	.139	.131–.148	.229 (.05) to .753 (.57)
	9	2	1	32.339*	27	.924	.899.	.054	.092	.080–.105	.171 (.03) to .775 (.60)
	8	3	1	24.523*	20	.945	.923	.045	.089	.074–.104	.331 (.11) to .781 (.61)
	7	4	1	17.882*	14	.946	.918	.048	.103	.086–.121	.536 (.29) to .785 (.62)
	7	5	1	16.656*	14	.948	.923	.046	.091	.074–.109	.138 (.02) to .797 (.64)
	7	6	1	16.820*	14	.932	.898	.050	.103	.086–.121	.328 (.11) to .779 (.61)
	12	7	2	63.949*	53	.904	.881	.088	.086	.078–.096	PPC: .531 (.28) to .799 (.64)
											NPE: .325 (.11) to .777 (.60)
	7	8	1	16.669*	14	.970	.954	.037	.072	.055–.091	328 (.11) to .802 (.64)
Study 4[c]	7	8	1	17.498*	14	.973	.960	.035	.071	.053–.090	.364 (.13) to .819 (.67)

Note. SPAS = social physique anxiety scale; χ^2(B–S) = Bollen–Stine chi-square; CFI = comparative fit index; TLI = Tucker–Lewis index; SRMR = standardized root mean square residual; RMSEA = root mean square error of approximation; RMSEA 90% CI = 90% confidence interval (CI) for the RMSEA point estimate; λ = factor loading; δ = uniquenesses; PPC = physique presentation comfort; NPE = negative physique evaluation. Model 1: Hart, Leary, and Rejeski (1989) and McAuley and Burman (1993); Model 2: Martin, Rejeski, Leary, McAuley, and Bane (1997), Isogai et al. (2001), and Smith (2004); Model 3: Hagger et al. (2007) for British, Estonian, and Swedish samples; Model 4: Isogai et al. (2001), and Hagger et al. (2007) for Spanish sample; Model 5: Motl and Conroy (2000, 2001), and Lindwall (2004); Model 6: Hagger et al. (2007) for Turkish sample; Model 7: Eklund, Mack, and Hart (1996), and Petrie, Diehl, Rogers, and Johnson (1996); Model 8: alternative model.

[a]Bootstrapped goodness-of-fit indexes are reported in this table because of the significant multivariate non-normality within these data. [b]$n = 678$. [c]$n = 670$.

*$p < .05$.

TABLE 4.—Goodness-of-fit statistics of invariance tests for alternative SPAS models.[a]

Study	No. of Items	Model	Description	χ^2(B–S)	*df*	CFI	TLI	SRMR	RMSEA	RMSEA 90% CI	$\Delta\chi^2$	Δdf	|ΔCFI|	|Δ RMSEA|
Study 3[b]	7	8A	Boys[c]	16.765*	14	.937	.905	.054	.071	.046–.097				
			Girls[d]	16.405*	14	.972	.958	.037	.062	.036–.089				
		8B	Configural invariance	34.388*	28	.953	.929	.054	.064	.051–.077				
			Factor loading invariance	40.576*	34	.950	.939	.070	.059	.048–.072	6.19	6	.003	.005
			Factor variance invariance	41.544*	35	.951	.941	.073	.058	.047–.072	7.16	7	.002	.006
			Uniquenesses invariance	50.286*	42	.927	.931	.089	.065	.054–.075	15.90	14	.026	.001
	7	8C	11–14[e]	17.483*	14	.975	.963	.044	.065	.027–.100				
			15–18[f]	17.665*	14	.953	.929	.048	.091	.070–.114				
		8D	Configural invariance	34.957*	28	.960	.940	.044	.059	.046–.072				
			Factor loading invariance	41.399*	34	.962	.953	.046	.052	.040–.065	6.44	6	.002	.007
			Factor variance invariance	42.120*	35	962	.955	.046	.051	.039–.063	7.16	7	.002	.008
			Uniquenesses invariance	50.097*	42	.961	.961	.049	.047	.036–.059	15.14	14	.001	.012
		8E	Factor loading invariance	40.576*	34	.953	.929	.070	.064	.042–.069				
			Intercept invariance	46.881*	40	.927	.923	.068	.066	.032–.055	6.31	6	.026	.002
			Intercept partial invariance ($\tau_{7,\ 9,\ 12\ \text{free}}$)	43.430*	37	.949	.942	.069	.058	.047–.069	2.67	3	.004	.006
			Factor mean invariance ($\tau_{7,\ 9,\ 12\ \text{free}}$)	44.285*	38	.940	.934	.067	.062	.051–.073	3.71	4	.013	.002
		8F	Factor loading invariance	41.399*	34	.962	.953	.046	.052	.040–.065				
			Intercept invariance	47.392*	40	.961	.959	.046	.048	.037–.060	6.07	6	.001	.004
			Factor mean invariance	48.394*	41	.961	.960	.046	.048	.037–.059	7.06	7	.001	.004
Study 4[g]	7	8A	Boys[h]	18.454*	14	.974	.961	.043	.075	.046–.104				
			Girls[i]	17.255*	14	.970	.955	.038	.069	.044–.095				
		8B	Configural invariance	35.363*	28	.972	.958	.043	.050	.037–.064				
			Factor loading invariance	41.403*	34	.970	.964	.061	.047	.035–.060	6.04	6	.002	.003
			Factor variance invariance	42.303*	35	.971	.965	.065	.046	.034–.059	6.94	7	.001	.004
			Uniquenesses invariance	52.564*	42	.961	.961	.078	.049	.038–.060	17.20	14	.011	.001
	7	8C	11–14[j]	16.654*	14	.970	.956	.045	.075	.040–.110				
			15–18[k]	18.084*	14	.977	.965	.034	.065	.042–.089				
		8D	Configural invariance	35.701*	28	.975	.962	.045	.049	.035–.062				
			Factor loading invariance	42.124*	34	.977	.972	.047	.042	.029–.055	6.42	6	.002	.007
			Factor variance invariance	42.966*	35	.977	.973	.050	.041	.028–.054	7.27	7	.002	.008
			Uniquenesses invariance	51.954*	42	.979	.979	.049	.036	.023–.048	16.25	14	.004	.013
		8E	Factor loading invariance	41.403*	34	.972	.958	.061	.050	.037–.064				
			Intercept invariance	47.408*	40	.943	.940	.059	.060	.049–.071	6.01	6	.029	.010
			Intercept partial invariance ($\tau_{7,9,12\ \text{free}}$)	44.355*	37	.966	.961	.059	.049	.037–.061	2.95	3	.006	.001
			Factor mean invariance ($\tau_{7,9,12\ \text{free}}$)	45.476*	38	.957	.953	.058	.054	.042–.065	4.07	4	.015	.004
		8F	Factor loading invariance	42.124*	34	.977	.972	.047	.042	.029–.055				
			Intercept invariance	48.837*	40	.976	.975	.047	.040	.027–.052	6.71	6	.001	.002
			Factor mean invariance	49.218*	41	.974	.974	.047	.040	.028–.052	7.09	7	.003	.002

Note. SPAS = social physique anxiety scale; χ^2(B–S) = Bollen–Stine chi-square; CFI = comparative fit index; TLI = Tucker–Lewis Index; SRMR = standardized root mean square residual; RMSEA = root mean square error of approximation; RMSEA 90% CI: RMSEA 90% confidence interval (CI); τ = factor intercepts; $\Delta\chi^2$ = change in goodness-of-fit χ^2 relative to baseline model; Δdf = change in *df* relative to baseline model; ΔCFI = change in CFI relative to baseline model; ΔRMSEA = Change in RMSEA relative to baseline model. Model 8A: CFA for boys and girls subsamples; Model 8B: CFA, gender invariance tests; Model 8C: CFA for 11 to 14 and 15 to 18 years old subsamples; Model 8D: CFA, age invariance tests; Model 8E: CFA, gender latent mean invariance tests; Model 8F: CFA, age latent mean invariance tests.

[a]Bootstrapped goodness-of-fit indexes are reported in this table because of the significant multivariate non-normality within these data. [b]$N = 678$. [c]$n = 297$. [d]$n = 381$. [e]$n = 231$. [f]$n = 447$. [g]$N = 670$. [h]$n = 295$. [i]$n = 375$. [j]$n = 217$. [k]$n = 453$.

*$p < .01$.

(reference model), (b) item intercepts invariance, and (c) latent means invariance.

The results from the Models 8A and 8C are reported in Table 4 and show that the seven-item model performed relatively well in the separate gender and age subsamples (i.e., CFI and TLI > .90; SRMR < .06; and RMSEA < .10). Moreover, we found acceptable and equivalent reliability coefficients in all subsamples ($\omega = .87$ for boys, girls, and 11–14 years old; and $\omega = .81$ for 15–18 years old). Measurement and latent mean structure invariance tests (Models 8B, 8D, 8E and 8F) suggested that: (a) the measurement model fully overlapped between genders and age groups, although the uniquenesses of Items 4 and 12 may not be invariant across gender (ΔCFI = .026); (b) item intercepts were partially invariant across gender (i.e., equality constraints on item intercepts for Items 7, 9, and 12 had to be relaxed) but fully invariant across age groups; and (c) the latent means significantly differed across genders but not age groups, with boys (latent mean fixed to zero) presenting significantly lower scores than girls [latent $M = -.16$, $t(676) = -4.26$, $p < .001$, $d = .33$].

Discussion

The objectives of this study were to investigate the psychometric properties of a version of the French SPAS in adolescents and to assess its measurement and mean structure invariance across gender and age. These findings indicate that none of the models identified in previous studies of adults (e.g., Eklund et al., 1997, 1996; Hagger et al., 2007; Hart et al., 1989; Isogai et al., 2001; Lindwall, 2004; Martin et al., 1997; Motl & Conroy, 2000, 2001; Petrie et al., 1996) or adolescents (McAuley & Burman, 1993; Smith, 2004) provided a satisfactory level of fit to the data in this French adolescent sample. Nevertheless, a truncated alternative seven-item model, composed exclusively of negatively worded items, provided satisfactory goodness-of-fit indexes, internal consistency coefficients, and test–retest reliability coefficients. Although many of the preceding studies also converge on a seven-item model, none of them converged on the same seven-item model. This observation is important because it provided the impetus for our consideration of all of the alternative models in this study. A core of common items is evident across studies (including this study), but an item or two differs in each instance. However, it should be noted that none of the seven-item models emerging in earlier studies has been replicated in a second independent sample, which makes it possible that the data-driven modifications to arrive at the final models in each instance provided a better model fit for the data set being analyzed instead of a more generalizable model.

The group-based tests of measurement and latent mean structure invariance conducted in this study represent a first step in the evaluation of the potential generalizability of the results. These analyses support the measurement invariance of the seven-item model identified in this study across gender and age groups. These results are consistent with gender invariance testing in earlier studies involving adolescent (Smith, 2004) and adult samples (Motl & Conroy, 2000, 2001), and the age invariance findings are promising in that this is the first study to examine this measurement issue with SPAS data. These findings provide evidence of SPAS construct measurement equivalence that affords the prospect of meaningful comparisons of SPAS data across gender and age groups. Regarding SPAS latent mean structure, these results partially support the mean structure equivalence of the seven-item model across gender even though the intercepts of Items 7, 9, and 12 were not observed to be invariant. On average, it appears that boys' and girls' answers on Items 7, 9, and 12 differed in a manner that is independent from the observed differences on the overall SPAS factor score. Nonetheless, these results provide evidence of the construct similarity of a four-item version of the SPAS across gender and suggest that meaningful comparisons can be made across gender, at least on the basis of the four invariant items. Moreover, as would be expected given earlier findings with adolescents using a nine-item version (i.e., Smith, 2004), higher latent mean scores were observed for girls than for boys. This finding was consistent across analyses relying on the four invariant items (i.e., 2, 3, 4, and 6) and analyses involving on all of the seven items (i.e., 2, 3, 4, 6, 7, 9, and 12) in the preferred seven-item model. Thus, the finding of gender-based latent mean differences is replicable across studies and appears quite robust. Finally, the results provide evidence of the complete latent mean structure invariance across age groups.

Study 4

The objectives of the fourth study were to cross-validate the factor structure of the seven-item French SPAS and its measurement and latent mean structure invariance (across gender and age groups) in a new independent sample of adolescents.

Method

Sample, measure, and procedure. Adolescents ($n = 670$) aged between 11 and 18 years (see Table 1) were recruited from four middle and high schools located in Southern France. This sample met the same inclusion criteria as that in Study 3 and completed the seven-item French SPAS developed in Studies 1 through 3 in the same aforementioned standardized conditions and after agreeing to participate and returning the informed consent forms signed by their parents.

Data analysis. In this study, we performed the CFA on the final model identified in Study 3 (Model 8) using bootstrapped ML estimation with AMOS 7.0 given the significant non-normality of the data (normalized coefficients values for kurtosis = 15.40). We based assessment of fit and reliability of this model on the indicators and formula described in Study 3.

Results

The results from the CFA model estimated on the Study 4 full sample were essentially the same as those found in Study 3 and were indicative of acceptable goodness of fit (see Table 3). Examination of measurement and latent mean structure invariance of the model across gender and age groups occurred in the same sequence as in Study 3 (see Table 4). The results from the Models 8A and 8C indicate that the seven-item model performed relatively well and slightly better than in Study 3 (i.e., CFI and TLI $> .95$; SRMR $< .05$; and RMSEA $< .08$) in the gender and age subsamples. Moreover, acceptable and equivalent reliability coefficients were found in all subsamples ($\omega =$ .87 for boys, girls, 11–14 years old, and 15–18 years old). The measurement and latent mean structure (Models 8B, 8D, 8E and 8F) invariance tests were also replicated and hence strengthen the results found in Study 3: (a) the SPAS measurement model was found to fully overlap between genders and age groups, although Item 12 uniquenesses may slightly differ across gender (ΔCFI = .011); (b) the intercepts were partially invariant across gender (i.e., equality constraints on Intercepts 7, 9, and 12 again had to be relaxed) but fully invariant across age groups; and (c) the latent means significantly differed across genders but not age groups, with boys presenting significantly lower scores than girls [latent $M = -.20$, $t(668) = -4.454$, $p < .001$, $d = .35$].

Discussion

The objectives of the fourth study were to cross-validate the factor structure and the measurement and the mean structure invariance of the seven-item French SPAS in a new independent sample of adolescents to ensure that the final model identified in Study 3 was not the result of capitalization of chance but rather a generalizable model. The results obtained using data from a new independent sample indicate that the seven-item model emerging in this investigation does exhibit generalizability among French adolescents. This effort amounts to several steps beyond earlier studies of adults SPAS psychometric properties (e.g., Hagger et al., 2007; Isogai et al., 2001; Lindwall, 2004) because those studies had not featured an attempt to cross-validate the identified seven-item model across an additional sample. The results observed for the optimal identified model of the French version of the SPAS for adolescents were found to be completely replicable across studies and across subsamples (i.e., gender and age groups) within those studies. Nonetheless, although the results from these series of studies brought strong evidence in favor of one specific seven-item SPAS model, this model was obtained on a French adolescent sample. For this reason, care should be taken to avoid using these results as a proof that the SPAS measurement debates have found a solution. This solution may rather have been found for French adolescents only, and similar methods should be used to clarify these issues further with additional populations.

Study 5

The objective of this study was to verify the convergent validity of the seven-item French SPAS with measures of social anxiety, self-esteem, fear of negative evaluation, social physique anxiety, disturbed eating attitudes and behaviors, and body image disturbance.

Method

Sample and procedure. Adolescents (70 boys and 49 girls) aged between 11 and 18 years (see Table 1) were recruited from two middle and high schools located in Southern France. All adolescents agreed to participate; returned consent forms signed by their parents; met the same inclusion criteria as that in Studies 3 and 4; and completed, in the same aforementioned standardized conditions, a booklet of instruments that were presented to them in the order listed in the Measures section.

Measures. We used the 7-item French SPAS developed in Studies 1 through 4 to measure adolescents' SPA. We also employed five widely employed French versions of: (a) the 24-item anxiety subscale of the Liebowitz Social Anxiety Scale (LSAS–AN; Liebowitz, 1987; Yao et al., 1999) in which participants to rate their anxiety on a Likert scale ranging from 0 (*none*) to 3 (*severe*); and (b) the 26-item EAT (EAT–26; Garner, Olmstead, Bohr, & Garfinkel, 1982; Leichner, Steiger, Puentes-Neuman, Perreault, & Gottheil, 1994) on which the global score provides an index of the presence of disturbed eating attitudes and behaviors. Participant responses are obtained on a 6-point scale ranging from 5 (*always*) to 0 (*never*) but consistent with Garner et al.'s (1982) recommendations, and this answer scale is recoded on a 4-point scale ranging from 0 to 3, with zero combining the least symptomatic answers (1–2–3) from the original rating scale; (c) the 10-item Rosenberg Self-Esteem Inventory (RSEI; Rosenberg, 1965; Vallières & Vallerand, 1990), which measures feelings of self-worth or self-acceptance using a 4-point Likert response format ranging from 4 (*strongly agree*) to 1 (*strongly disagree*); (d) 30-item Fear of Negative Evaluation Scale (FNES; Musa, Kostogianni, & Lépine, 2004; Watson & Friend, 1969), which measures "apprehension about others' evaluations, distress over their negative evaluations, avoidance of evaluative situations, and the expectation that others would evaluate oneself negatively"(Watson & Friend, 1969, p. 449) using 30 items rated on a 1 (*true*) to 0 (*false*) response scale; and (e) the 19-item Body Image Avoidance Questionnaire (BIAQ; Rosen, Srebnik, Saltzberg, & Wendt, 1991; Maïano, Morin, Monthuy-Blanc, & Garbarino, 2009) on which the global score provides an index of behavioral manifestations of body image disturbance on a 6-point Likert scale ranging from 0 (*never*) to 5 (*always*).

Results

The internal consistencies of the different instruments used in this study were all in the acceptable range (EAT–26: $\alpha =$ .83; BIAQ: $\alpha =$ .70; LSAS–AN: $\alpha =$ .92; FNES: $\alpha =$ .86; RSEI: $\alpha =$ .84; and seven-item SPAS: $\alpha =$ .74). We employed a Bonferroni correction to manage error rate inflation and thus set the alpha error at .01 (i.e., .05/5). The seven-item French SPAS was significantly and positively correlated with the global score of the EAT–26 ($r = .50$, $p < .001$), the LSAS–SA ($r = .47$, $p < .001$), the BIAQ ($r = .34$, $p < .001$), and the FNES ($r = .45$, $p < .001$) but negatively correlated with the RSEI ($r = -.60$, $p < .001$).

Discussion

The purpose of the fifth study was to examine the convergent validity of the seven-item SPAS. Moderate correlations between the French SPAS and these measures were observed. These findings were in the expected directions and were very similar to those found in previous validation studies in adolescents (Baş et al., 2004; Hausenblas & Mack, 1999; Smith, 2004; Thompson & Chad, 2002) and adults (e.g., Atalay & Gençöz, 2008; Crawford & Eklund, 1994; Haase & Prapavessis, 1998; Hart et al., 1989). Accordingly, they offer preliminary support for the convergent validity of the seven-item version of the SPAS.

Study 6

The objective of this study was to test the criterion-related validity of the seven-item French SPAS in adolescent girls with anorexia nervosa and without eating disorders (ED).

Method

Sample. Adolescent females ($n = 33$) without ED and adolescent female patients ($n = 33$) suffering from anorexia nervosa (AN) according to *Diagnostic and Statistical Manual of Mental Disorders* (4th ed. [*DSM–IV*]; American Psychiatric Association, 1994) and to International Classification of Diseases (ICD–10) criteria (World Health Organization, 1994) participated in this study (see Table 1). The anorectic patients were recruited from an inpatient psychiatric unit and the diagnostic and history of AN was obtained with the fifth French version of the Mini International Neuropsychiatric Interview (MINI; Sheehan et al., 1998). The adolescents without ED were recruited in a high school in southern France and met the same inclusion criteria as that of adolescents from Studies 3 through 5 except that the MINI was employed to confirm a lack of diagnostic and history of ED. All adolescents agreed to participate and returned consent forms signed by their parents.

Measures. We used the seven-item French SPAS developed in Studies 1 through 5 to measure adolescents' SPA. The fifth French version of the MINI was used for diagnostic purposes and to obtain a patient history of AN and bulimia nervosa (BN). This instrument is a short, structured, diagnostic interview that can be used as a tool to diagnose 16 Axis I psychiatric disorders according to the *DSM–IV* and ICD-10 criteria. The MINI has 16 separate modules (e.g., major depressive episode, anxiety, etc.) each involving standardized, structured, close-ended questions. Interviewers read these questions verbatim to the interviewees. We made psychiatric diagnosis and history in a specific module according to the number of affirmative replies to each question. In this study, we used only the AN and BN modules.

Procedure. The MINI was completed by a psychiatrist in the psychiatric unit during the first interview with the patient. The interview with the adolescents without ED was conducted by J. Monthuy-Blanc of this study. The SPAS was completed after the interview.

Results

We compared the scale score from the seven-item French SPAS between the clinical and nonclinical samples using an independent-samples t test (one-tailed) and Cohen's (1992) measure of effect size. The independent-samples t tests were significant, $t(64) = 11.16$, $p < .001$, $d = 1.69$, and reflect a large effect size. These results revealed that the clinical group ($M = 25.58$, $SD = 5.06$) had higher scores than the nonclinical group ($M = 17.00$, $SD = 2.38$).

Discussion

In this study, we sought to evaluate the criterion-related validity of the SPAS using a clinical sample of anorectic girls. The results were consistent with Mack et al.'s (2007) findings in showing that the anorectic girls did present significantly higher scores than the nonclinical sample. This finding supports the criterion-related validity of the French SPAS.

General Discussion

In conclusion, we evaluated the psychometric properties of the French SPAS in a series of six studies. The obtained results suggest that the seven-item version may be confidently used in research measuring SPA in French adolescents because it: (a) was found to be applicable to French adolescents; (b) was equivalent to the original English version; (c) demonstrated satisfactory factor validity and measurement invariance across gender and age groups, a result that was replicated in two large and independent samples, (d) revealed higher latent mean scores of SPA for females and latent mean structure invariance across age categories, a result that was replicated in two large and independent samples; and (e) had satisfactory convergent and criterion-related validity. These results are encouraging.

Nevertheless, three limitations of these studies must be taken into account when interpreting these findings. First, the factor structure and measurement invariance analyses of the French SPAS were based on a mixed (boys and girls) sample of nonclinical and normally achieving early and late adolescents, which might not be considered representative of the French adolescent population. This indicates that the use of this instrument should be limited to samples with these same characteristics. Therefore, examining the factor structure and measurement invariance of the seven-item French version of the SPAS across a more diverse sample of adolescents should be a future research priority. Such research could be performed using samples comprised of adolescents from various clinical populations (i.e., bulimic, obese), socioeconomic status, activity levels, and cultural or linguistic groups.

Second, the reliance on a cross-sectional sample also precludes the verification of the developmental stability or change of the SPAS for adolescents. Although this study allowed for the verification of the 2-week test–retest reliability of the instruments, a complete test of the construct validity of the French SPAS would involve testing the developmental change of SPA during the early to late adolescent years. This issue should clearly be addressed in the context of longitudinal studies with different age groups and as well as with different Age × Gender groups.

Third, we performed the criterion-related validity analysis using data obtained from relatively small samples of nonclinical and clinically anorectic adolescent girls. The possibility exists that this result may be sample specific and hence of limited generalizability. Therefore, examining the criterion-related validity of this instrument in boys and across both clinical and nonclinical samples of adolescents should be a future research priority.

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