A re-examination of the BIS/BAS scales: Evidence for BIS and BAS as unidimensional scales

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Abstract

Objectives: Carver and White's behavioral inhibition system and behavioral activation system (BIS/BAS) scales are the most widely used to assess constructs of the revised reinforcement sensitivity theory. This study provides a re-examination of the latent structure of the original BIS/BAS scales.

Methods: The interpretability of the three purported BAS subfactors relative to a "general behavioral activation" factor was assessed using Schmid-Leiman and standard confirmatory factor analysis. Regarding the BIS scale, comparisons were made between (a) Carver and White's unidimensional BIS model, (b) Johnson, Turner, and Iwata's 2-factor BIS model, (c) Heym, Ferguson, and Lawrence's alternative 2-factor BIS model, and (d) a modified Heym et al. model (unidimensional) controlling for method effects of reverse-scored items.

Results: Results revealed the majority of variance of individual BAS items was accounted for by a common, general BAS dimension. Additionally, for the BIS scale, results of the χ^2 difference statistical test supporting the 1-factor model, as well as the noted theoretical and psychometric difficulties in interpreting a multifactor BIS scale, provide converging support that BIS items actually represent a single, unidimensional factor.

Conclusions: The collective results suggested that the BIS and BAS scales should be conceptualized as separate unidimensional measures, which is consistent with theory behind the original development.

KEYWORDS

BIS/BAS, CFA, unidimensional scale

1 | INTRODUCTION

Reinforcement sensitivity theory (RST; Gray, 1990) is a theory of the neuropsychology of motivation, learning, and emotions (Smillie, Pickering, & Jackson, 2006) and is one of the most widely used theories of personality (Cogswell, Alloy, van Dulmen, & Fresco, 2006; Cooper, Gomez, & Aucote, 2007). The revised RST model (Gray & McNaughton, 2000) encompasses three constructs: the behavioral activation system (BAS), the fight, flight, freeze system (FFFS), and the behavioral inhibition system (BIS). The BAS is thought to function as a reward system mediating responses to all appetitive stimuli (Corr, 2004; Gray & McNaughton, 2000). The FFFS is considered part of the threat response system mediating the emotion of fear and motivating escape and avoidance (Corr, 2004). Finally, the BIS underlies the emotion of anxiety and is conceptualized as a conflict detection and resolution

device (Smillie et al., 2006). To date, the most popular measure of the RST constructs (from either conceptualization) are Carver and White's (1994) behavioral inhibition system and behavioral approach system (BIS/BAS) scales, which were developed to assess the constructs of BIS and BAS (Cooper et al., 2007; Levinson, Roodebaugh, & Frye, 2011), but not the FFFS.

The BIS/BAS scales comprise a self-report measure of avoidance and approach tendencies that contains a 7-item BIS scale and a 13-item BAS scale (Carver & White, 1994). The original developers also posited three BAS subfactors, corresponding to Reward Responsiveness (RR; five items), Drive (four items), and Fun-Seeking (FUN; four items) subscales. Numerous factor analytic studies employing both exploratory and confirmatory analyses have provided support for this proposed four-factor structure (i.e., BIS-1 factor; BAS-3 factor; Cogswell et al., 2006; Leone, Perugini, Bagozzi, Pierro, & Mannetti,

2001; Levinson et al., 2011; Ross, Millis, Bonebright, & Bailey, 2002), although the theoretical relevance of the three BAS subscales has also been questioned (Cogswell et al., 2006; Torrubia, Alvia, Molto, & Caseras, 2001). Campbell-Sills, Liverant, and Brown (2004), for example, demonstrated that although the four-factor structure of the BIS/BAS scales was supported, two BAS items did not load simply on their intended factors and other items were associated with nonsalient loadings. These researchers also found support for a general behavioral activation factor that may be more theoretically consistent and interpretable compared with the three-factor BAS model.

The unidimensional nature of the original BIS factor has also been questioned (Johnson, Turner, and Iwata's (2003)). Results from confirmatory and exploratory factor analyses have indicated two potentially problematic BIS items that appear to form a separate factor (Cogswell et al., 2006; Johnson et al., 2003). A recent study (Heym, Ferguson, & Lawrence's, 2008) extended these findings by separating the BIS scale into two factors: BIS-anxiety and FFFS-fear. The authors used confirmatory factor analysis to compare BIS models including Carver and White's (1994) single-factor BIS model, Johnson et al.'s (2003) 2-factor model, and their BIS-anxiety and FFFS-fear two-factor model. Results revealed that Heym et al.'s (2008) two-factor model (four-item BIS, three-item FFFS-fear) was associated with significantly better model fit compared with Carver and White's (1994) single-factor model and Johnson et al.'s (2003) two-factor model.

Despite initial support, the validity of this two-factor BIS model warrants further examination given ambiguity in the published literature. First, as noted by Campbell-Sills et al. (2004), the two *reverse-scored* BIS items are those that often load poorly on the BIS factor. This is important to note because reverse-scored items are known to be associated with method effects that convolute the dimensionality of measures when they are not appropriately accounted for in modeling procedures (see Brown, 2003; Marsh, 1996; Marsh, Scalas, & Nagengast, 2010). In addition, Carver and White (1994) directly stated that because the original fight-flight system was not the focus of Gray (1990), "this aspect of the theory has been disregarded in the work reported here" (p. 319). Because Carver and White (1994) did not construct the BIS/BAS scales to assess the FFFS-fear dimension, this lends less theoretical basis to separating the BIS factor into multiple subcomponents. In combination with the fact that previous investigations reporting a two-factor BIS structure did not control for method effects, it is possible that a closer examination that introduces this control would yield new insights as to the nature of the BIS scale's psychometric properties.

1.1 | The current study

The current study thus sought to clarify unresolved issues regarding the dimensionality of the BIS and BAS scales through the application of appropriate modeling procedures. Specifically, one aim of this study was to use exploratory bifactor analysis to examine the degree to which each of the three purported BAS subfactors (Drive, FUN, and RR) are substantively meaningful and interpretable relative to a "general BAS" dimension. The exploratory bifactor model has recently been shown to be effective at examining dimensionality—particularly unidimensionality—when multiple subscales are presumed to be present

and underlie a common general factor (e.g., Reise, Moore, & Haviland, 2010; Reise, Morizot, & Hays, 2007). The second aim was to use appropriate modeling procedures to account for method effects due to reverse-scored items (Brown, 2003; Marsh, 1996) to compare a unidimensional BIS model (accounting for the reverse-scored items via correlated error terms) to previously published one-factor and two-factor BIS models that did not account for such method effects (Carver & White, 1994; Heym et al., 2008; Johnson et al., 2003). We hypothesized that both behavioral approach and avoidance as measured by the BIS/BAS scales would load onto separate unidimensional factors.

2 | METHODS

2.1 | Participants

A total of 566 individuals completed an online survey via SONA systems, a cloud based participant management software, in the fall 2011 semester at a large University in the Southern United States. This survey consisted of a variety of self-report questionnaires across domains of social, experimental, and clinical psychology. Completion time of the survey averaged 45 min. Participants included both psychology and nonpsychology majors enrolled across a number of general psychology courses. The current study included participants with completed BIS/BAS forms with no missing data. Participants ($n = 29$) were excluded due to one ($n = 24$) or two ($n = 5$) missing items, leaving a final sample of 537. The mean age was 19.67 years ($SD = 2.92$; range 17–54), and 62% ($n = 333$) were female. Participant ethnicities were as follows: 70.2% White; 22.5% Black; 2.6% Asian; 2.2% Hispanic; 2.2% Other; and 0.2% no ethnicity reported.

2.2 | Measures

The BIS/BAS scales (Carver & White, 1994) comprise a 20-item self-report measure of trait sensitivity levels of the behavioral inhibition and activation systems. The original scale developers posited three BAS subscales that assess RR, Drive, and FUN. The same iteration of the measure posits BIS as a unitary scale measuring sensitivity to avoidance of potential punishment. Each statement is rated on a 4-point Likert-type scale ranging from 1 (*very false for me*) to 4 (*very true for me*).

2.3 | Procedure

All procedures were approved by the University's institutional review board. Participants were recruited through in-class announcements and email solicitations to complete an online survey. Following informed consent, individuals who chose to participate in the study completed a series of questionnaires including the BIS/BAS scales. Participants received course credit for their participation.

2.4 | Data analytic approach

Factor Structure of the BAS. The factor structure of the BAS was examined through Schmid-Leiman exploratory factor analysis (SL-EFA; Schmid & Leiman, 1957) using the Psych package (Revelle, 2016) in

the R statistical program (R Development Core Team, 2008). This technique is an exploratory bifactor model (Holzinger & Swineford, 1937) particularly useful for examining the extent to which variations in item responses are due to subfactors versus a general factor common to all the items. In addition to positing a common general factor (e.g., “general BAS,” as in the current study), a SL-EFA also posits multiple “group” factors that represent specific content domains. These “group” factors are orthogonal to each other as well as to the general factor, and they explain any additional item response variance above and beyond that accounted for by the general factor (see Reise et al., 2010 for full description of oblique [promax] rotation). Additionally, polychoric correlation matrices (Holgado-Tello, Chacón-Moscó, Barbero-García, & Vila-Abad, 2010) and the minimum residual (ordinary least squares) solution estimation method (Harman & Jones, 1966) were used due to the categorical nature of the data. Items were not reversed-coded for the factor analyses.

In addition to the exploratory bifactor model, standard three-factor EFA solutions of the BAS items were examined using oblique rotation, polychoric correlation matrices, and ordinary least squares with the R statistical package. Factor loadings $>.30$ were considered to be strong loadings on the general factor (in the SL-EFA solution) and on the specific subfactors (in both the SL-EFA and standard EFA solutions). The degree to which each of the three purported BAS subfactors (Drive, FUN, and RR) are substantively meaningful and interpretable (in addition to the “general BAS” factor) was assessed based on whether the following criteria were met: (a) at least three items—the minimum number of items needed to just identify a factor—loaded on the subfactor in both the standard EFA and SL-EFA solutions with loadings $>.30$; (b) the pattern of factor loadings of the items were consistent with theory-based predictions (and not related simply due to nuisance factors such as the manner an item is worded; i.e., reversed scored statements); (c) these items did not load (or cross-load) on factors inconsistent with the originally posited three-factor (Drive, FUN, and RR) BAS structure; and (d) the factor loading of a given item in the standard EFA solution did not evidence a substantial drop relative to the item's factor loading when subjected to a SL-EFA solution (given that such a drop would suggest that it would be difficult to obtain meaningful variation from the item/subscale once accounting for the “general BAS” factor; cf. Reise et al., 2010).

In addition to EFA, we also conducted confirmatory factor analysis (CFA) using Mplus Version 7.2 (Muthén & Muthén, 2012) to examine the fit of the bifactor model in a CFA context. As with the EFA bifactor model, all factors were set as orthogonal to each other and data were treated as ordinal (categorical). The weighted least-squares with mean and variance (WLSMV) estimator was also used for these analyses. The fit indices used to evaluate the model fit are described below.

2.4.1 | Factor structure of the BIS

A CFA was conducted in Mplus version 7.2 (Muthén & Muthén, 2012) to evaluate and compare competing BIS models. CFA was used for these analyses to examine competing BIS factor structures given that CFA provides advantages over EFA when comparing competing (nested) models that involving modeling method effects (Brown, 2003; Marsh, 1996). Computations were based on polychoric

correlations and the robust WLSMV adjustment estimator, as has been recommended for use with categorical (ordinal) data (Flora & Curran, 2004; Muthén, du Toit, & Spisic, 1997). Given that a unidimensional model that controls for method effects by specifying correlated error terms among the two reverse-scored BIS items (cf. Brown, 2003) have not yet been examined, CFA was used to compare our newly posited competing unidimensional model whereby method effects (due to the two reverse-scored items) were accounted for by specifying correlated error terms (also referred to as correlated uniqueness; Brown, 2003; Marsh, 1996) to (a) Carver and White's (1994) unidimensional BIS model (not accounting for method effects due to the reverse-scored items), (b) Johnson et al.'s (2003) two-factor correlated BIS model (which relegated the two reverse-scored items into their own factor, called the BIS-anxiety and the FFFS-fear subscales), (c) Heym et al.'s (2008) two-factor correlated BIS model (which includes a third item as part of the FFFS subscale—yet without accounting for method effects between the two reverse-worded items), and (d) a modified Heym et al. (2008) two-factor correlated BIS model whereby we included correlated error terms between the two reverse worded items of the FFFS-fear subscale to account for method effects.¹

2.4.2 | Examining model fit

The following fit indices were used to evaluate these competing models: the comparative fit index (CFI), Tucker–Lewis index (TLI), and the root mean square error of approximation (RMSEA). Values of CFI and TLI greater than .95 (Hu & Bentler, 1999) and RMSEA values less than .08 (Browne & Cudeck, 1993) suggest good model fit. These cut-offs were used, however, as only rough guidelines for interpreting good model fit.

2.4.3 | Comparing the fit of nested models

Due to the correlated error terms included in our newly specified one-factor model, this one-factor model was not nested relative to the competing models b and c (noted above)—thereby precluding our ability to conduct χ^2 difference tests to statistically compare their relative model fit. Our newly specified one-factor model (with correlated error terms included between the two reverse-worded items), however, was nested within one-factor model without correlated error terms as well as the modified Heym et al. (2008) two-factor correlated BIS model that included correlated error terms between the two reverse-worded items of the three-item FFFS-fear subscale (to account for method effects of these two reverse-worded items). This nesting thus afforded our ability to statistically test whether relegating the three FFFS-fear items to comprise a separate second factor (as in the Heym et al., 2008 model) was associated with significantly better model fit relative to the more parsimonious one-factor model (with included error terms between the two reverse-worded items). Given that these analyses were based on the WLSMV limited information estimator, we

¹Although it has been suggested that method effects could also be modeled by correlating error terms among the *nonreversed-worded* items of a measure, researchers more recently have reported that correlating error terms among the *reversed-worded* items more appropriately accounts for method effects (e.g., Corwyn, 2000; Marsh, 1996; Tomas & Oliver, 1999). We thus examined method effects by correlating error terms among the reversed-worded (and not the nonreversed-worded) items in the present study.

conducted the χ^2 difference test using the “difftest” procedure available in Mplus, which calculates the appropriate χ^2 difference test statistics when using limited information estimators (see Asparouhov & Muthen, 2006; Muthen & Muthen, 2012).

3 | RESULTS

3.1 | Factor structure of the BAS

The factor loadings associated with the 13 BAS items from the three-factor SL-EFA, standard three-factor EFA, and unidimensional solutions appear in Table 1. The factor loadings of the unidimensional model were also added to Table 1 to compare those loadings to the loadings of the general factor of the bifactor model. The intercorrelations among the three BAS factors in the standard EFA solutions were .55, .50, and .48. The SL-EFA solution revealed that all 13 BAS items loaded strongly on the general BAS factor (all loadings >.40, range .43–.70). The loadings of the general BAS factor were also comparably high compared with the loadings of the unidimensional model. These results demonstrate that all 13 BAS items are good indicators of the general BAS dimension. Further, only one subscale (the BAS RR) met all criteria for being a meaningful and interpretable subscale. Both the BAS Drive and FUN subscales failed to yield three items that met criteria for subscale retention. Together, these results reveal that the majority of the variance of the individual BAS items may be accounted for by a common, general BAS dimension.

The confirmatory factor analytic results yielded similar findings in support of the unidimensionality of the BAS items. Specifically, when fitting a confirmatory bifactor model, all items loaded significantly on the general BAS factor (see factor loadings in parentheses in Table 1).

Due to the FUN items only loading significantly on the general BAS factor but not on the specific FUN domain, these items were specified to only load on the general BAS factor. All other items loaded significantly on both the general BAS factor and also on their respective RR and Drive specific domains (also shown in Table 1). This confirmatory bifactor model also fit the data well (i.e., CFI = .96, TLI = .95, RMSEA = .072).

As a result, the general BAS factor (i.e., the BAS total score) appears to be the most meaningful and interpretable BAS dimension, consistent with the underlying theory that drove its initial development. It is important to note, however, that several of the loadings on the specific factors are somewhat high (and in a few cases, higher than the loading on the general factor), such as some of the factor loadings of the RR items based on the EFA model. The specific content domains underlying the BAS scale are thus likely not trivial and may provide unique and meaningful contributions to item variation that warrants further research. Results based on the CFA bifactor model, however, suggest that the FUN items may best represent the general BAS domain as opposed to also the specific FUN domain, and so more research is needed to better understand the degree to which these items provide specific information above and beyond the overarching general BAS dimension.

3.2 | Factor structure of the BIS

The fit statistics associated with the various competing BIS models appear in Table 2. As hypothesized, Carver and White's (1994) one-factor model (not accounting for method effects) was associated with poor model fit (e.g., CFI = .73, RMSEA = .208), consistent with previous findings (Heym et al., 2008; Johnson et al., 2003). Also as expected, Johnson et al.'s (2003) two-factor correlated model was associated

TABLE 1 SL- and standard EFA three-factor solutions for the 13 BAS items

Scale/item	SL-EFA solution (CFA solution)				Three-factor EFA solution			Unidimensional EFA solution	Items
	g	Drive	RR	FUN	Drive	RR	FUN		
Drive 1	.52 (.46)	.50 (.75)			.74			.54	I go out of my way to get things I want.
Drive 2	.70 (.62)	.61 (.51)			.88			.69	When I want something, I usually go all out to get it.
Drive 3	.62 (.59)	— (.24)			.38			.70	If I see a chance to get something I want, I move on it right away.
Drive 4	.63 (.68)	— (.17)		.46			.67	.63	When I go after something, I use a “no holds barred” approach.
RR 1	.45 (.52)		.53 (.41)		.70			.59	When I am doing well at something, I love to keep at it.
RR 2	.52 (.57)		.61 (.56)		.79			.67	When I get something I want, I feel excited and energized.
RR 3	.55 (.63)		.44 (.36)		.57			.69	When I see an opportunity for something I like, I get excited right away.
RR 4	.46 (.49)		.53 (.53)		.70			.59	When good things happen to me, it affects me strongly.
RR 5	.43 (.48)		.44 (.40)		.59			.56	It would excite me to win a contest.
FUN 1	.54 (.69)		.31	(—)	.40	.43		.65	I am always willing to try something new if I think it will be fun.
FUN 2	.56 (.69)			.42 (—)			.66	.63	I will often do things for no other reason than that they might be fun.
FUN 3	.50 (.60)			.44 (—)			.71	.53	I often act on the spur of the moment.
FUN 4	.60 (.73)			.35 (—)	.30	.53		.69	I crave excitement and new sensations.

Note. BAS = behavioral activation system; CFA = confirmatory factor analysis; EFA = exploratory factor analysis; FUN = Fun Seeking; g = general BAS factor; RR = Reward Responsiveness; SL = Schmid-Leiman; Loadings <.30 were suppressed.

TABLE 2 Fit statistics for confirmatory factor analytic BIS models

Model	CFI	TLI	RMSEA	χ^2	df
Carver and White's (1994) 1-factor BIS (unidimensional) model	.73	.60	0.21	339.08	14 ^a
Johnson et al.'s (2003) 2-factor BIS (anxiety/FFFS) model (not controlling for method effects)	.95	.93	0.09	72.69	14 ^a
Heym et al.'s (2008) 2-factor BIS (anxiety/FFFS-fear) model (not controlling for method effects)	.78	.64	0.20	286.77	13 ^a
Heym et al.'s (2008) 2-factor BIS (anxiety/FFFS-fear) model (controlling for method effects)	.96	.93	0.09	58.80	12 ^a
Alternate 1-factor BIS (unidimensional) model (controlling for method effects)	.96	.94	0.08	56.67	13 ^a

Note. BIS = behavior inhibition scale; CFI = comparative fit index; FFFS = fight, flight, freeze system; RMSEA = root mean square error of approximation; TLI = Tucker–Lewis index.

^aDegrees of freedom estimated using the special Mplus procedure.

with good model fit (e.g., CFI = .95), and fit significantly better than Carver and White's (1994) one-factor model [χ^2 difference test (1) = 151.95, $p < .001$].² This model however did not meet the benchmark for good fit based on RMSEA (.09). Heym et al.'s (2008) two-factor model interestingly also demonstrated poor model fit according to RMSEA (.20) and CFI (.78) but exhibited significantly better model fit than Carver and White's (1994) one-factor model [χ^2 difference test (1) = 37.14, $p < .001$].

Specifying correlated error terms in Heym et al.'s (2008) two-factor model substantially improved model fit relative to the original Heym et al. (2008) two-factor model that did not include correlated error terms to account for the method effects of the two reverse-scored items of the FFFS-fear subscale (i.e., CFI improved from .78 to .96, RMSEA improved from .20 to .09). The χ^2 difference test also supported these models as being significantly different: χ^2 difference test (1) = 121.56, $p < .001$.

Tests of the more parsimonious one-factor model (which included correlated error terms to account for method effects of the two reverse-scored items) revealed good model fit (e.g., CFI = .96, RMSEA = .08). As noted above, this one-factor model is not nested relative to the competing models above, which precluded our ability to conduct χ^2 difference tests to compare their relative model fit. The one exception was the comparison with the one-factor model (with no correlated error terms), which is nested within the one-factor unidimensional model with correlated error terms. The χ^2 difference test supported these models as being significantly different: χ^2 difference test (1) = 162.09, $p < .001$.

Nonetheless, the newly proposed one-factor model was associated with the best fit indices relative to those models and was the only model that met both the .95 and .08 cutoff for good model fit based on CFI and RMSEA, respectively. It is important to note, however, that when considering the cutoff point as general guidelines rather than rigid cutoff points, Heym et al.'s (2008) two-factor BIS (anxiety/FFFS-fear) model (controlling for method effects) also had relatively supportive model fit indices overall (e.g., CFI = .96, TLI = .93, RMSEA = .09).

Because our newly posited one-factor model (with correlated error terms between the two reverse-scored items) was nested within

the modified Heym et al. (2008) two-factor model (with correlated error terms between the two reverse-scored items of the FFFS-fear scale), we used the χ^2 difference to examine differences between these factor structures. Results revealed that the modified Heym et al. (2008) two-factor model was not associated with significant better model fit relative to our newly posited one-factor model [$\chi^2_{diff}(1) = .193$, $p = .661$], thereby providing support for our more parsimonious one-factor model. Further, although the modified Heym et al. (2008) two-factor model was associated with good fit based on CFI (.96), it was associated with slightly less good fit based on RMSEA (.09), although not significantly so. It is also important to note that the two factors under the modified Heym et al.'s (2008) model (which corrected for correlated error terms to account for method effects between the reverse-scored items; e.g., Brown, 2006) almost perfectly correlated with each other (albeit negatively) at a correlation of $-.96$. When two factors correlate this highly, this almost always suggests the presence of unidimensionality (i.e., that items on both “factors” in fact belong to the same factor). The near -1.0 correlation between the two “factors” in the modified Heym et al.'s (2008) model thus also suggests that those two “factors” may represent a single unidimensional factor (albeit with opposite positive/negative valence). We also conducted a chi-square difference test to statistically test the assumption that unidimensionality. Specifically, we compared the two-factor Heym et al. (2008) model with and without the correlation between the two factors set to 1.0. The chi-square difference test was nonsignificant, providing additional support for the model: χ^2 difference test (1) = .19, $p = .66$ (ns).

Further, a one-factor model (with correlated error terms), as seems supported in the present study, is more interpretable given the parsimony of its one-factor structure.

Regarding Johnson et al.'s (2003) two-factor model, it is also notable that the second factor in Johnson et al.'s (2003) two-factor model comprises only two items (i.e., the two reverse-scored items). This is problematic given that factors with fewer than three items are considered weak and often negligible (Costello & Osborne, 2005). Given the results of the χ^2 difference statistical test supporting the one-factor model, as well as the noted theoretical and psychometric difficulties in interpreting a multifactor BIS scale, the current results provide converging support that the BIS items actually represent a single, unidimensional factor and should be interpreted as such. All five models (including factor loadings, correlation between factors, and correlations between correlated residuals) appear in Figures 1–5.

²In order to conduct this chi-square difference test between Carver and White's (1994) one-factor model and Johnson et al.'s (2003) two-factor model, we modified Carver and White's (1994) one-factor model slightly to make this model nested within Johnson et al.'s (2003) two-factor model. Specifically, we specified the loadings from both reverse-scored items to the target factor to be equal (as this is the case in Johnson et al.'s two-factor model).

FIGURE 1 Carver and White's (1994) one-factor BIS (unidimensional) model. BAS = behavioral activation system; BIS = behavior inhibition scale

FIGURE 2 Johnson et al.'s (2003) two-factor BIS (anxiety/FFFS) model. BAS = behavioral activation system; BIS = behavior inhibition scale; FFFS = fight, flight, freeze system

4 | DISCUSSION

The main goal of the present study was to re-examine the factor structure of the BIS/BAS scales. Specifically, the BAS scales (both a general BAS factor and the three BAS subscales model) were examined to determine the extent to which they were substantively meaningful and interpretable. In addition, competing BIS models (both unidimensional and two-factor) were compared with determine best model fit.

Results from this study indicated that all 13 BAS items were good indicators of the general BAS dimension. In addition, with the

exception of the BAS-RR subscale, neither the BAS-Drive nor the BAS-FUN subscales demonstrated adequate psychometric support to be considered meaningful and interpretable subscales. Specifically, some items cross loaded, and many subscale factor loadings in the standard EFA evidenced substantial drop relative to the factor loadings when subjected to an exploratory bifactor model. This lends further support to the assertion that once accounting for the "general BAS" dimension, very little meaningful item variation due to the subscales remains. More research is needed to confirm these findings, particularly given previous support for both the presence of the "general

FIGURE 3 Heym et al.'s (2008) two-factor BIS (anxiety/FFFS-fear) model. BAS = behavioral activation system; BIS = behavior inhibition scale; FFFS = fight, flight, freeze system

FIGURE 4 Alternate one-factor BIS (unidimensional) model (controlling for method effects). Note the factor loading of $-.17$ is small, but significant ($z = 3.58, p < .01$). BAS = behavioral activation system; BIS = behavior inhibition scale

BAS" factor and meaningful loadings on the purported RR, FS, and Drive subdomains above and beyond the general BAS dimension (Campbell-Sills et al., 2004). Across both studies, however, results suggest the general BAS factor is the most meaningful and should be relied on for interpretation. Results supporting the BAS scale as primarily a unidimensional construct are also more theoretically

consistent with both the original (Gray, 1990) and revised RST (Gray & McNaughton, 2000). This finding may also aid in understanding previous inconsistencies among subscales noted in the literature (e.g., Campbell-Sills et al., 2004; Smillie et al., 2006) by providing additional evidence for use of the more parsimonious and theoretically guided unidimensional scales. For example, previously published subscale

FIGURE 5 Modified Heym et al.'s (2008) two-factor BIS (anxiety/FFFS-fear) model accounting for method effects. BAS = behavioral activation system; BIS = behavior inhibition scale; FFFS = fight, flight, freeze system

associations (i.e., RR, FS, and Drive) with other variables have been difficult to interpret in terms of how each subscale related theoretically to general BAS functioning (Cogswell et al., 2006; Jackson & Smillie, 2004). Re-examining these associations while conceptualizing the BAS primarily as a unidimensional factor may elucidate new and more theoretically consistent conclusions.

The present findings also provide strong evidence that covariances among the seven items of the BIS scale are best explained by a single underlying construct. Although refuting previous assertions that the latent structure of the BIS consists of two potentially meaningful factors (BIS-anxiety and FFFS-fear; Heym et al., 2008), these results are consistent with the literature detailing potential method effects in scales using both positively and negatively worded items (e.g., Marsh, 1996). If method effects are not modeled in the factor analyses of the BIS scale, the resulting latent structure of the scale may be confounded or masked by method effects (Marsh, 1996). Specifically, the unidimensional BIS model evidenced better fit than the two-factor models once controlling for method effects. Although Johnson et al.'s (2003) two-factor model (not controlling for method effects) achieved substantially similar fit as the one-factor model (with included method effects), the current one-factor model is still favorable given more parsimony and consistency with theory. Additionally, Johnson et al.'s (2003) two-factor model is problematic given that one of the factors comprises only two items (i.e., the only two reversed-scored items in the scale). Similarly, it is notable that the modified Heym et al. (2008) two-factor model that allowed correlated error terms among the two reversed-worded items of the FFFS-fear subscale to control for method effects was not associated with better model fit relative to our newly proposed one-factor model (also controlling for method effects), despite Heym et al.'s (2008) two-factor model being composed of more factors than our one-factor model.

The reason for this is likely due to the fact that Heym et al.'s (2008) "two factors" were nearly perfectly correlated at .96, thereby essentially making that "two factor" model a single-factor model. This lack of improved fit of Heym et al.'s (2008) two-factor model supports our more parsimonious one-factor model—particularly in light of the good fit evidenced by the one-factor model (with included method effects). As noted above, researchers have often interpreted the second factor as having substantive meaning (e.g., a factor of fear that is related to FFFS; Heym et al., 2008). As demonstrated in the present study, however, it is more likely that the appearance of a second factor is due to lack of modeling method effects as opposed to an actual, distinct construct related to FFFS. That said, factor analysis alone cannot determine whether correlated error is a "method" factor or a factor with substantive meaning. Future validity studies are thus needed to substantiate this claim.

Results highlight potential problems of having reverse-scored items in a scale (e.g., Brown, 2003). In fact, Marsh (1996) questioned whether or not the advantages of including reverse-scored items (e.g., reducing acquiescence reporting biases) outweigh the problems related to method effects, such as complexity of scoring reverse-scored items (Brown, 2003). More research is needed to determine how to best handle reverse-scored items, and whether or not they should be retained in the BIS scale. Although the current results are interesting and provide the basis for future research, it is important to note that this is a single study. Replication of the methods used in this study would be beneficial to add support to the results and resultant conceptualization of BIS/BAS factor structure. Despite limitations, the collective results provide a strong psychometric basis for the use of the original BIS/BAS scales (controlling for method effects in the BIS scale) to be used in future research. Results also caution against the use of a two-factor BIS scale to separate and assess the constructs

of behavioral inhibition and FFFS (cf. Heym et al., 2008). Although the FFFS construct is important in RST research (Walker & Jackson, 2017), and recent instrumentation allows assessment of the FFFS independent of other RST constructs (Maack, Buchanan, & Young, 2015), it remains the task of future studies to develop a means of more accurately measuring, conceptualizing, and integrating the FFFS construct into RST research more broadly.

DECLARATION OF INTEREST STATEMENT

The authors have no conflicts of interest to declare.

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REFERENCES

- Asparouhov, T., & Muthén, B. (2006). Robust chi square difference testing with mean and variance adjusted test statistics. Mplus Web Notes: No. 10. May 26, 2006. Retrieved September 26, 2010, from <http://www.statmodel.com/download/webnotes/webnote10.pdf>
- Brown, T. A. (2003). Confirmatory factor analysis of the Penn State worry questionnaire: Multiple factors or method effects? *Behaviour Research and Therapy*, 41, 1411–1426.
- Brown, T. A. (2006). *Confirmatory factor analysis for applied research*. New York: Guilford Press.
- Browne, M., & Cudeck, R. (1993). Alternative ways of assessing model fit. In K. A. Bollen, & J. S. Long (Eds.), *Testing structural equation models* (pp. 136–159). Newbury Park, CA: Sage.
- Campbell-Sills, L., Liverant, G., & Brown, T. (2004). Psychometric evaluation of the behavioral inhibition/behavioral activation scales in a large sample of outpatients with anxiety and mood disorders. *Psychological Assessment*, 16, 244–254.
- Carver, C. S., & White, T. (1994). Behavioral inhibition, behavioral activation, and affective responses to impending reward and punishment: The BIS/BAS scales. *Journal of Personality and Social Psychology*, 67, 319–333.
- Cogswell, A., Alloy, L. B., van Dulmen, M. H. M., & Fresco, D. M. (2006). A psychometric evaluation of behavioural inhibition and approach self report measures. *Personality and Individual Differences*, 40, 1649–1658.
- Cooper, A., Gomez, R., & Aucote, H. (2007). The behavioural inhibition system and behavioural approach system (BIS/BAS) scales: Measurement and structural invariance across adults and adolescents. *Personality and Individual Differences*, 43, 295–305.
- Corr, P. J. (2004). Reinforcement sensitivity theory and personality. *Neuroscience and Biobehavioral Reviews*, 28, 317–332.
- Corwyn, R. F. (2000). The factor structure of global self-esteem among adolescents and adults. *Journal of Research in Personality*, 34, 357–379.
- Costello, A. B., & Osborne, J. W. (2005). Best practices in exploratory factor analysis: Four recommendations for getting the most from your analysis. *Practical Assessment, Research and Evaluation*, 10, 1–9.
- Flora, D., & Curran, P. (2004). An empirical evaluation of alternative methods of estimation for confirmatory factor analysis with ordinal data. *Psychological Methods*, 9, 466–491.
- Gray, J. A. (1990). Brain systems that mediate both emotion and cognition. *Cognition & Emotion*, 4(3), 269–288.
- Gray, J. A., & McNaughton, N. (2000). *The neuropsychology of anxiety*. Oxford, England: Oxford University Press.
- Harman, H., & Jones, W. (1966). Factor analysis by minimizing residuals (minres). *Psychometrika*, 31(3), 351–368.
- Heym, N., Ferguson, E., & Lawrence, C. (2008). An evaluation of the relationship between Gray's revised RST and Eysenck's PEN: Distinguishing BIS and FFS in Carver and White's BIS/BAS scales. *Personality and Individual Differences*, 45, 709–715.
- Holgado-Tello, F., Chacón-Moscoso, S., Barbero-García, I., & Vila-Abad, E. (2010). Polychoric versus Pearson correlations in exploratory and confirmatory factor analysis of ordinal variables. *Quality and Quantity*, 44(1), 153–166.
- Holzinger, K. J., & Swineford, R. (1937). The bifactor method. *Psychometrika*, 2, 41–54.
- Hu, L. T., & Bentler, P. M. (1999). Cutoff criteria for fit indexes in covariance structure analysis: Conventional criteria versus new alternatives. *Structural Equation Modeling*, 6, 1–55.
- Jackson, C. J., & Smillie, L. D. (2004). Appetitive motivation predicts the majority of personality and an ability measure: A comparison of BAS measures and a re-evaluation of the importance of RST. *Personality and Individual Differences*, 36(7), 1627–1636.
- Johnson, S., Turner, R., & Iwata, N. (2003). BIS/BAS levels and 783 psychiatric disorder: An epidemiological study. *Journal of Psychopathology and Behavioral Assessment*, 25, 25–36.
- Leone, L., Perugini, M., Bagozzi, R. P., Pierro, A., & Mannetti, L. (2001). Construct validity and generalizability of the Carver-White behavioural inhibition system/behavioural activation system scales. *European Journal of Personality*, 15(5), 373–390.
- Levinson, C. A., Roodebaugh, T. L., & Frye, T. (2011). An examination of the factor, convergent, and discriminant validity of the behavioral inhibition system and behavioral activation system scales. *Journal of Psychopathology and Behavioural Assessment*, 33, 87–100.
- Maack, D. J., Buchanan, E., & Young, J. (2015). Development and psychometric investigation of an inventory to assess fight, flight, and freeze tendencies: The fight, flight, freeze questionnaire. *Cognitive Behaviour Therapy*, 44, 117–127.
- Marsh, H. W. (1996). Positive and negative global self-esteem: A substantively meaningful distinction or artifact? *Journal of Personality and Social Psychology*, 70, 810–819.
- Marsh, H. W., Scalas, L. F., & Nagengast, B. (2010). Longitudinal tests of competing factor structures for the Rosenberg self-esteem scale: Traits, ephemeral artifacts, and stable response styles. *Psychological Assessment*, 22, 366–381.
- Muthén, B., du Toit, S. H. C., & Spisic, D. (1997). Robust inference using weighted least squares and quadratic estimating equations in latent variable modeling with categorical and continuous outcomes. *Psychometrika*.
- Muthén, B., & Muthén, L. (2012). In L. K. Muthén, & B. O. Muthén (Eds.), *Mplus user's guide* (7th ed.). Los Angeles, CA: Muthén & Muthén. 1998–2012
- R Development Core Team (2008). R: A language and environment for statistical computing. In *R foundation for statistical computing*. Vienna: Austria. URL: <http://www.r-project.org>
- Reise, S., Moore, T., & Haviland, M. (2010). Bifactor models and rotations: Exploring the extent to which multidimensional data yield univocal scale scores. *Journal of Personality Assessment*, 92, 544–559.
- Reise, S., Morizot, J., & Hays, R. (2007). The role of the bifactor model in resolving dimensionality issues in health outcomes measures. *Quality of Life Research*, 16, 19–31.
- Revelle, W. (2016). Psych: Procedures for personality and psychological research. Northwestern University, Evanston, <http://cran.r-project.org/web/packages/psych/>. R package version 1.6.4.
- Ross, S. R., Millis, S. R., Bonebright, T. L., & Bailey, S. E. (2002). Confirmatory factor analysis of the behavioral inhibition and activation scales. *Personality and Individual Differences*, 33, 861–865.
- Schmid, J., & Leiman, J. M. (1957). The development of hierarchical factor solutions. *Psychometrika*, 22, 53–61.
- Smillie, L. D., Pickering, A. D., & Jackson, C. J. (2006). The new reinforcement sensitivity theory: Implications for personality measurement. *Personality and Social Psychology Review*, 10(4), 320–335.

- Tomas, J., & Oliver, A. (1999). Rosenberg's self esteem scale: Two factors or method effects. *Structural Equation Modeling*, 6, 84–98.
- Torrubia, R., Alvia, C., Molto, J., & Caseras, X. (2001). The sensitivity to punishment and sensitivity to reward questionnaire (SPSRQ) as a measure of Gray's anxiety and impulsivity dimensions. *Personality and Individual Differences*, 31, 837–862.
- Walker, B. R., & Jackson, C. J. (2017). Examining the validity of the revised reinforcement sensitivity theory scales. *Personality and Individual Differences*, 106, 90–94.

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