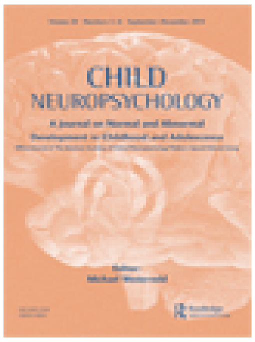

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Confirmatory factor analysis of the Behavior Rating Inventory of Executive Functioning (BRIEF) in children and adolescents with ADHD

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The Behavior Rating Inventory of Executive Functioning (BRIEF) is a parent report measure designed to assess executive skills in everyday life. The present study employed a confirmatory factor analysis (CFA) to evaluate three alternative models of the factor structure of the BRIEF. Given the executive functioning difficulties that commonly co-occur with attention-deficit/hyperactivity disorder (ADHD), the participants included 181 children and adolescents with a diagnosis of ADHD. The results indicated that an oblique two-factor model, in which the Monitor subscale loaded on both factors (i.e., Behavioral Regulation, Metacognition) and measurement errors for the Monitor and Inhibit subscales were allowed to correlate, provided an acceptable goodness-of-fit to the data. This two-factor model is consistent with previous research indicating that the Monitor subscale reflects two dimensions (i.e., monitoring of task-related activities and monitoring of personal behavioral activities) and thus loads on multiple factors. These findings support the clinical relevance of the BRIEF in children with ADHD, as well as the multidimensional nature of executive functioning.

Keywords: BRIEF; Confirmatory factor analysis; Factor structure; Executive functions; ADHD.

Executive functioning is an overarching term used to describe higher-order cognitive skills necessary for independent, goal-directed behavior (Lezak, 1995). These skills include initiation, inhibition, switching, working memory, attention, planning, problem-solving, self-regulation, and utilization of feedback, among others (Alvarez & Emory, 2006; Anderson, 2002; Barkley, 2000; Lezak, 2004). Executive functions have been conceptualized as both a unitary construct (Della Sala, Gray, Spinnler, & Trivelli, 1998; Shallice, 1990) and as a set of interrelated processes that work together (Alexander & Stuss, 2000). Debate around the unity or diversity of executive functions has led to various definitions of executive functioning and, subsequently, various methods for measuring this construct (Hughes & Graham, 2002).

It has been recognized by leaders in the field that the measurement of executive skills is inherently challenging for several reasons (Burgess, 1997; Gioia, Isquith,

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Kenworthy, & Barton, 2002). First, a true operational definition of executive functioning does not exist (Hughes & Graham, 2002), making it difficult to measure the multiple components it includes. Because of this problem, there is no single assessment measure which all individuals with executive dysfunction fail (Burgess, 1997). This is evident in the fact that, while many individuals with frontal lesions and observable problems with executive functioning in everyday life perform poorly on tests designed to be sensitive to executive skills, many do not (Cripe, 1996; Shallice & Burgess, 1991). Executive functions have also been considered one of the most difficult domains to measure using traditional laboratory tests (Salimpoor & Desrocher, 2006). Testing is typically conducted in a quiet room, without distractions and with a clinician coordinating test administration, explaining rules, setting goals, and initiating and stopping behaviors (Lezak, 1982). Additionally, multi-tasking and prioritizing are often not needed, since the clinician directs the tasks and determines the order in which they need to be completed (Manchester, Priestley, & Jackson, 2004). Therefore, core deficits may go unnoticed because of the structured nature of the tests and the non-distracting environment. Emotional arousal, a key component in decision-making and behavioral regulation, is carefully controlled in the testing environment and thus generally eliminated from the assessment process. Finally, patients with executive dysfunction may perform within the normal range on neuropsychological testing, but become exhausted in doing so, which is not reflected in performance.

Results from laboratory tests present additional issues. It is difficult to translate executive skills into standardized tests; thus, critical executive functions may go unmeasured (Lezak, 1995; Sbordone, 2000). Performance-based tests measure individual components of executive functioning, rather than the integrated, multidimensional decision-making that is necessary during novel and/or complex situations (Goldberg & Podell, 2000; Shallice & Burgess, 1991). Therefore, reliance on performance-based types of tests alone may be inadequate in assessing executive functions because they attempt to separate integrated skills into component parts (Burgess, 1997) and can yield an incomplete and limited assessment (Bodnar, Prahme, Cutting, Denckla, & Mahone, 2007; Gioia & Isquith, 2004).

Given these limitations, the Behavior Rating Inventory of Executive Functioning (BRIEF) was developed using a multidimensional model of executive functioning based on the theoretical assumption that executive functions are distinct, yet related within an overarching executive system (Gioia et al., 2002). This measure attempts to overcome previous shortcomings by assessing parents' reports of their children's everyday executive behaviors in natural settings (Gioia et al., 2002).

The BRIEF has been submitted to factor analysis to examine its factor structure. Gioia, Isquith, Guy, and Kenworthy (2000) conducted an exploratory factor analysis on the eight scales comprising the BRIEF with parent and teacher ratings for both normative and clinical groups, and a two-factor structure was identified: a three-scale (Inhibit, Shift, Emotional Control) Behavioral Regulation factor and a five-scale (Initiate, Working Memory, Plan/Organize, Organization of Materials, Monitor) Metacognition factor. Slick, Lautzenhiser, Sherman, and Eyrl (2006) provided further evidence of a two-factor model of the eight scales comprising the BRIEF (i.e., Behavioral Recognition and Metacognition) using a sample of 80 children and adolescents with epilepsy.

However, it was proposed that the Monitor subscale may reflect both of these indices—monitoring of task-related activities and monitoring of personal behavioral activities. Gioia et al. (2002) investigated this hypothesis through a confirmatory factor analysis (CFA) using a sample of 374 children aged 5 to 18 years and with a variety of diagnoses, including attention-deficit/hyperactivity disorder (ADHD), learning disorders, autism spectrum disorder (ASD), Tourette's syndrome, affective disorders, and seizure disorders. Based on the current one-dimensional versus multidimensional models of executive functioning, four models of executive functioning were examined using nine subscales (with the Monitor subscale now divided into two subscales: Self-Monitor and Task-Monitor). The results of this study indicated that a three-factor model was the most appropriate structure for the nine scales: Emotional Regulation (Shift, Emotional Control), Behavioral Regulation (Self-Monitor, Inhibit), and Metacognition (Initiate, Working Memory, Plan/Organize, Organization of Materials, Task-Monitor). Egeland and Fallmyr (2010) more recently examined this issue using both the parent and teacher report of the Norwegian BRIEF in a mixed clinical group and healthy sample. Results indicated that a three-factor model provided the best fit to the data for both groups, indicating a distinction between emotional and behavioral regulation. A CFA of the adult version of the BRIEF (BRIEF-A) showed similar findings, with greater impairment on the Behavioral Regulation factor in adults with ADHD when using a three-factor model (Roth, Lance, Isquith, Fischer, & Giancola, 2013).

Further investigation of the BRIEF is needed, including the validity of the underlying structure of the BRIEF subscales (Gioia et al., 2002). The use of this measure with different clinical populations must also be examined to determine its validity, sensitivity, and specificity (Bodnar et al., 2007). Investigating the BRIEF within specific samples will help determine whether unique executive profiles arise from specific underlying executive function structures based on the disorder being examined (Gioia et al., 2002). Additionally, it may provide useful information on the generality and specificity of the model's executive functions.

Examining the factor structure of these measures in a sample of children and adolescents with ADHD is particularly important because of the significant deficits in executive functioning typically exhibited by this population. ADHD is characterized by persistent and developmentally inappropriate symptoms of inattention and/or hyperactivity and impulsivity (American Psychiatric Association, 2000). Executive skills difficulties are considered a crucial component of understanding ADHD (Barkley, 2006) and are often measured in clinical settings (Gioia et al., 2002; Heaton et al., 2001). Prior studies have examined the discriminant validity of the BRIEF for youth with ADHD (e.g., Gioia et al., 2002; Mahone et al., 2002; Reddy, Hale, & Brodzinsky, 2011; Toplak, Bucciarelli, Jain, & Tannock, 2009). Thus far, no studies have examined the factor structure of the BRIEF using only a sample of children and adolescents with ADHD.

The current study examined the construct and assessment of executive functioning with a sample of youth diagnosed with ADHD. A maximum likelihood CFA via LISREL 8 (Jöreskog & Sörbom, 1996) was used to test hypotheses about the structure of executive functioning using the BRIEF data. Specifically, it was hypothesized that a measurement model consisting of two correlated factors—namely, Behavioral Regulation (Inhibit, Shift, Emotional Control) and Metacognition (Initiate, Working Memory, Plan/Organize, Organization of Materials, Monitor)—would (a) fit the data better than a model consisting of two orthogonal factors, (b) fit the data better than a one-factor model, and (c) provide an acceptable goodness-of-fit to the data. Although we hypothesized that the two-factor model would fit the BRIEF data well, we realized that it might require modifications in order to achieve a fully acceptable goodness-of-fit. We thus anticipated the potential need to modify the a-priori model slightly to include cross-loadings or correlated measurement errors based on the relevant literature on executive functioning in order to achieve an acceptable overall model fit.

METHOD

Subjects

This study was part of a larger, ongoing data collection of clinically-referred children seen for a neuropsychological evaluation in a university-based outpatient neuropsychology clinic in a large urban city. Participants included children and adolescents diagnosed with ADHD (aged 6 to 16 years) and their parents. Youth were diagnosed with ADHD following a comprehensive neuropsychological battery by a licensed clinical psychologist and board certified neuropsychologist. Diagnosis was established using objective test data, subjective parent report (e.g., Conners Rating Scales), and corroboration from a third party (e.g., teachers). Exclusionary criteria included a full scale IQ (FSIQ) of less than 75 or a diagnosis of a neurological condition (e.g., seizures), as these could impact neuropsychological test performance. A total of 181 youth were included in the current study. None of the participants declined to have their data used anonymously. Demographic data for the sample are presented in Table 1. The majority of the sample was male (73%), Caucasian (56%), and had a least one other comorbid diagnosis (53%). Comorbid conditions included learning disorder (39%), mood disorder (14%), anxiety

Table 1 Sample Characteristics (n = 181).

	Mean	*SD*
Age	10.32	2.67
	n	Percentage of Sample
Gender		
Male	132	72.9
Female	49	27.1
Race/Ethnicity		
Caucasian	102	56.4
African American	34	18.8
Latino/Latina	16	8.8
Biracial	14	7.7
Asian	2	1.1
Missing	13	7.2
ADHD Subtype		
Combined	87	48.1
Inattentive	79	43.6
Hyperactive/Impulsive	3	1.7
NOS	1	0.6
Unspecified	11	6.1
Comorbid Diagnosis		
Learning Disorder	70	38.7
Mood Disorder	26	14.4
Anxiety Disorder	18	9.9
Disruptive Behavior Disorder	7	3.9
Other	4	2.2

disorder (10%), and disruptive behavior disorder (4%). Approximately half of the sample was diagnosed with the ADHD Combined type (48%) and 44% were diagnosed with the Inattentive type. Medication use was not reported, although all participants were newly diagnosed with ADHD at the time of the study.

Procedure

Data collection took place over the course of 4 years. Parents provided informed consent for assessment and to have their child's clinical data de-identified and used for research purposes. Children aged 12 and over provided assent in addition to parent consent. Demographic information was collected from parents through their completion of a child neuropsychology history questionnaire and a clinical interview conducted by a licensed clinical psychologist. Children and adolescents completed a neuropsychological test battery that included the assessment of intelligence, and parents of the participants completed the BRIEF. All procedures were supervised by a licensed clinical neuropsychologist and approved by the university's institutional review board.

Instruments

The BRIEF (Gioia et al., 2000) is an 86-item parent report questionnaire designed to assess executive functioning in children aged 5 to 18 years. Parents rate if their child's behavior is "never", "sometimes", or "often" a problem, with higher ratings indicative of greater perceived impairment. The BRIEF is composed of eight clinical scales (Initiate, Working Memory, Plan/Organize, Organization of Materials, Monitor, Inhibit, Shift, Emotional Control) that generate two broad indices: Metacognition Index and Behavior Regulation Index. An overall score is obtained (the Global Executive Composite) from the raw scores of the Metacognition Index and the Behavioral Regulation Index. The Behavioral Regulation Index includes the Inhibit, Shift, and Emotional Control subscales (Gioia et al., 2000). The Metacognition Index is comprised of the following subscales: Initiate, Working Memory, Plan/Organize, Organization of Materials, and Monitor (Gioia et al., 2000). It also has two validity scales to identify the informants' response styles. The BRIEF was normed on 1419 control children and 852 children from referred clinical groups. Adequate test-retest reliability, internal consistency, content and construct validity, and convergent and discriminate validity have been demonstrated. Specifically, test-retest reliability statistics range from .79 to .88 during a 2-week period and internal consistency is reported as ranging from .80 to .98 (Gioia et al., 2000).

The Wechsler Intelligence Scale for Children – Fourth Edition (WISC-IV; Wechsler, 2003) is a commonly used, well-normed measure of intellectual functioning that was used to determine IQ for exclusionary purposes in the current study (i.e., participants with an IQ of less than 75 were excluded). It consists of ten subtests which yield four domain scores (Verbal Comprehension, Perceptual Reasoning, Working Memory, Processing Speed) and an overall measure of intellectual functioning (FSIQ). The WISC-IV was nationally standardized with a representative sample of 2200 children aged 6 to 16 years, and it has demonstrated good psychometric properties.

Data Analyses

The factor structure of the BRIEF was examined using CFA via LISREL 8.80. As required by CFA, the model specified which scales were expected to load on which factors, how these factors intercorrelate, and the relations among unique-error terms for the observed indicators. In these models, scales were forced to have a single loading, factors were standardized (i.e., variances fixed at one), and unique errors were considered independent. To establish fit, χ^2 values and four measures of goodness-of-fit were used to assess CFA models in the study: (1) the root mean square error of approximation (RMSEA), (2) the standardized root mean square residual (SRMR), (3) the non-normed fit index (NNFI), and (4) the comparative fit index (CFI). According to Hu and Bentler (1998), the RMSEA measure of absolute fit should be no greater than .10 and the SRMR value should be less than .08. For measures of relative fit, Bentler and Bonett (1980) suggest that values above .90 are indicative of a good fit for the NNFI and the CFI. In addition to conventional cutoff values, the fit of a model is also interpreted relative to competing models. In the current study, a second-order CFA was used to examine three competing models for the BRIEF.

RESULTS

Descriptive data for the sample is presented in Table 2. WISC-IV FSIQ ranged from below average to superior and was overall in the average range (mean standard score = 97.88, *SD* = 12.17). Of the scores on the BRIEF subscales, 25 to 62% were in the clinical range (*T*-scores of 65 or greater), with the greatest problems reported on the Working Memory subscale (62% in the clinical range). Intercorrelations of the BRIEF are presented in Table 3.

The first hypothesis stated that a measurement model of the BRIEF consisting of two correlated factors (Behavioral Regulation, Metacognition) would provide a better fit to the data than a global one-factor model for this sample of children with ADHD. A test of the one-factor model of executive functioning provided a poor absolute fit [$\chi^2(20, n = 181) = 200.27$, RMSEA = .24, SRMR = .12] as well as a poor relative fit (NNFI = .78, CFI = .84; see Table 4). Supporting the a-priori

Table 2 Descriptive Characteristics ($n = 181$).

Domain	Mean	SD	Median	Range	% Clinically elevated*
WISC-IV FSIQ (Standard Score)	97.88	12.17	96	75–126	N/A
BRIEF (*T*-scores)					
Behavioral Regulation Index					
Inhibit	58.97	13.72	57	36–100	33
Shift	58.18	13.58	57	36–95	33
Emotional Control	55.65	12.73	56	35–91	25
Metacognition Index					
Initiate	60.55	11.71	62	35–87	39
Working Memory	68.16	10.82	70	39–93	62
Plan/Organize	64.73	11.61	65	33–89	55
Organization of Materials	59.01	10.03	61	33–76	37
Monitor	62.37	10.90	64	31–91	50

Note. *Clinically elevated scales have *T*-scores ≥ 65.

Table 3 Means, *SD*s, and Intercorrelations for BRIEF Subscales (n = 181).

	1	2	3	4	5	6	7	8
1 Inhibit	1.00							
2 Shift	.530*	1.00						
3 Emotional Control	.537*	.648*	1.00					
4 Initiate	.288*	.492*	.387*	1.00				
5 Working Memory	.236*	.335*	.231*	.658*	1.00			
6 Planning/Organization	.238*	.402*	.262*	.650*	.718*	1.00		
7 Organization of Materials	.277*	.261*	.216*	.459*	.540*	.635*	1.00	
8 Monitor	.533*	.469*	.417*	.575*	.510*	.679*	.576*	1.00
Mean	58.97	58.18	55.65	60.55	68.16	64.73	59.01	62.37
Standard deviation	13.72	13.58	12.73	11.71	10.82	11.61	10.03	10.90

Note. *Correlation is significant at the .01 level (2-tailed).

Table 4 Goodness-of-Fit Statistics for BRIEF Factor Models (n = 181).

						Measures of fit			
Factor model	χ^2	*df*	$\Delta\chi^2$	Δdf	p <	RMSEA	SRMR	NNFI	CFI
1. One global factor	200.27	20	-	-	.001	.24	.12	.78	.84
2. Two oblique factors	89.48	19	110.79	1	.001	.14	.07	.91	.94
3. Two orthogonal factors	133.12	20	24.06	1	.001	.16	.23	.86	.90
4. Two oblique factors with Monitor subscale loading on both factors	69.56	18	131.83	2	.001	.12	.06	.93	.95
5. Two oblique factors with Monitor subscale loading on both factors and Monitor measurement error correlated with Inhibit subscale	**48.93**	**17**	**153.06**	**3**	**.001**	**.09**	**.05**	**.95**	**.97**

Note. χ^2 = chi-square test statistic; *df* = degrees of freedom; $\Delta\chi^2$ = change in chi-square test statistic; Δdf = change in degrees of freedom; RMSEA = root mean square error of approximation; SRMR = standardized root mean square residual; NNFI = non-normed fit index; CFI = comparative fit index. Model 5 (indicated in bold) provides the best fit to the data.

hypothesis, the oblique two-factor model fit the data better than the one-factor model [$\Delta\chi^2(1)$ = 110.79, p < .001]. Also supporting the a-priori hypothesis, an oblique version of this two-factor model provided a better fit to the data than did an orthogonal version [$\Delta\chi^2(1)$ = 24.06, p < .001]. However, contrary to predictions, the overall goodness-of-fit of the oblique two-factor model was unacceptable. Although the model's SRMR was less than .08 and its NNFI and CFI values were greater than .90, indicating a good fit, its RMSEA value was greater than .10, suggesting a poor model fit.

Due to the poor overall fit of the two-factor model and the previous literature suggesting that the Monitor subscale may be related to both factors (e.g., Gioia et al., 2002), a modified version of the two-factor model that allowed the Monitor subscale to load on both the Behavioral Regulation and Metacognition factors was tested next. This modified model fit the data better than the original two-factor model [$\Delta\chi^2(1) = 19.92$, $p < .001$], but still provided a poor absolute fit (RMSEA = .12) despite an adequate relative fit (NNFI = .93, CFI = .95).

Thus, the two-factor oblique model underwent additional modifications in order to improve overall model fit. Previous research investigated the idea that monitoring one's behavior and behavioral inhibition share the need to regulate one's actions and its impact on others, and the results indicated that the Monitor and Inhibit subscales were statistically related (Gioia et al., 2002). Therefore, it was expected that these two subscales may share variance over and above what was reflected in the common factors in the model. Thus, a two-factor model that allowed the Monitor subscale to load on both factors (i.e., Behavioral Regulation and Metacognition), as well as allowing the measurement errors for the Monitor and Inhibit subscales to correlate, was tested next. This two-factor oblique model provided a good absolute fit [$\chi^2(17, n = 181) = 48.93$, RMSEA = .09, SRMR = .05] and a good relative fit (NNFI = .95, CFI = .97; see Figure 1). Additionally, it provided a better fit than a one-factor model [$\Delta\chi^2(3) = 153.06$, $p < .001$].

As expected given the large sample size, chi-square statistics for all models were significant; however, the chi-square value was the lowest for the final model. Inspection

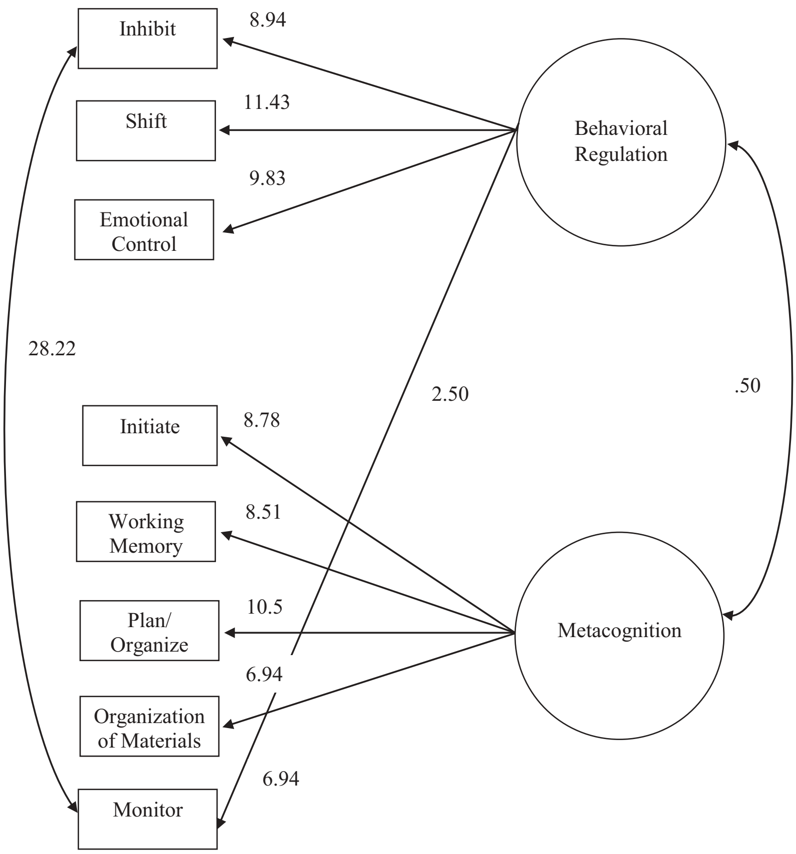

Figure 1 Completely standardized parameter estimates for the final two-factor CFA model (n = 181).

of the factor intercorrelation from this two-factor model demonstrated that the factors were moderately intercorrelated (r = .50, p < .001).

DISCUSSION

A lack of research on the validity of recent measures of executive functioning, as well as continued debate surrounding the unity or diversity of executive functions, underscored the importance of investigating the factor structure of the BRIEF (Hughes & Graham, 2002; Silver, 2000). The constructs underpinning the BRIEF required further investigation through statistical scrutiny (i.e., CFA) with an ADHD sample given the executive skills difficulties that typically exist within this population, as well as the frequent use of executive functioning measurement in clinical settings (Gioia et al., 2002; Heaton et al., 2001). Thus, analyses were completed to assess the factor structure of the BRIEF using a sample of children and adolescents diagnosed with ADHD.

A total of five factor analyses were completed for the BRIEF. It was hypothesized that a two-factor oblique model of executive functioning would provide a good fit to the data compared to a one-factor model and a two-factor orthogonal model. A one-factor model and a two-factor orthogonal model were both rejected. A two-factor oblique model provided a better fit than the former two models; however, it did not provide an adequate fit overall (i.e., a good relative fit but a poor absolute fit).

Additional model modifications were conducted, and a two-factor oblique model in which the Monitor subscale loaded equally on both factors (i.e., Behavioral Regulation, Metacognition) and the measurement errors for the Monitor and Inhibit subscales were allowed to correlate provided a good fit for the data. This two-factor model is consistent with previous research indicating that the Monitor subscale reflects two dimensions (i.e., monitoring of task-related activities and monitoring of personal behavioral activities) and thus loads on multiple factors (Gioia & Isquith, 2004; Slick et al., 2006). Gioia et al. (2002) previously explored this hypothesis through CFA with the Monitor subscale separated into two components, and the results indicated that a three-factor model provided the best fit to the data for a mixed clinical group. Nonetheless, the two-factor model with modifications as indicated in the current study is the most appropriate factor structure since proper administration and scoring of the BRIEF in clinical settings only produces eight subscales. Additionally, parsimony would dictate that the two-factor model is superior to a three-factor model. In the current model, the measurement errors for the Monitor and Inhibit subscales were allowed to correlate, which is consistent with previous research indicating that these two subscales were statistically related (Gioia et al., 2002) given their shared ability to regulate one's actions and its impact on others.

Previous literature indicated a need for factor replication of the BRIEF within specific clinical groups (e.g., ADHD) in order to elucidate the generality or specificity of executive functioning models (Gioia et al., 2002). The current study provides support for the use of a two-factor model of executive functions for the BRIEF for youth diagnosed with ADHD. In addition, the Behavioral Regulation and Metacognition factors within this model were moderately correlated, suggesting that these factors are related but separate, which is consistent with a multidimensional theory of executive functioning.

Limitations

Although this study provides important information about the measurement of executive functioning, there are several limitations that need to be acknowledged. First, the sample was heterogeneous in terms of ADHD subtype, with approximately half of the sample being diagnosed as the Inattentive type and the other half diagnosed as the Combined type. Future research should obtain adequate sample sizes to compare findings between subtypes to determine whether the factor structure of measures of executive functioning is similar across groups. Another limitation of the sample was the rate of comorbidity, with 53% having been diagnosed with at least one other disorder. Although this limits the specificity of the results for ADHD, it increases generalizability. In fact, this rate of comorbidity is consistent with community samples that demonstrate comorbidity rates of up to 44% (Szatmari, Offord, & Boyle, 1989) and rates near 87% in clinic-referred children (Kadesjo & Gillberg, 2001). Comparing samples of youth diagnosed with ADHD only with comorbid samples may be helpful in future research to clarify the specific impact of different diagnoses on executive functions. Finally, a three-factor model was not examined in the current study given prior research demonstrating increased support for a three-factor model with the Monitor subscale separating into two. Nonetheless, the two-factor model has been maintained on the current version of the BRIEF and is thus being used in clinical practice, offering applied value to our study.

Conclusions

Despite the aforementioned limitations, the present study makes several important contributions to the literature. It provides support for the existing factor structure of the BRIEF through its replication of the factor structure using CFA. This finding also supports a model of executive functioning that is comprised of multiple but interrelated components. Finally, it provides evidence for the use of the existing factor structure of the BRIEF with a sample of children and adolescents with ADHD. This is particularly important given the characteristic executive functioning impairments that often co-occur with ADHD and the need to adequately assess executive functions in neuropsychological assessments with this population.

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