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Bifactor Modeling of the Behavior Rating Inventory of Executive Function (BRIEF) in a Chilean Sample

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Abstract
The Behavior Rating Inventory of Executive Functions evaluates executive functioning through the observation of students' performance in real contexts. Most psychometric studies of the scale have only tested the first-order structure, despite the hierarchical configuration of its theoretical model. A bifactor model was conducted on a normative sample of 5- to 18-year-old Chileans (*M* age = 11.3 year, *SD* = 3.7) to test a hierarchical structure of three first-order factors and an independent second-order factor. Bifactor analyses showed best fit for the proposed hierarchical structure. Findings supported a method to evaluate executive functioning models that provides a general global factor score that may complement existing indices and thus help clinicians to make better inferences.

Keywords
bifactor modeling, executive functions, BRIEF

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Introduction

According to Chen, West, and Sousa (2006), bifactor and second-order confirmatory factor analysis (CFA) are two alternative approaches that allow representing general constructs comprised of several related factors. Furthermore, Reise, Moore, and Haviland (2010) emphasized that second-order CFA and bifactor models are the only choices to recognize multidimensionality and simultaneously retain the existence of a general factor. The latent model underlying the Behavior Rating Inventory of Executive Functions (BRIEF; Gioia, Isquith, Guy, & Kenworthy, 2000) is a good example of the previous idea, as it includes two first-orders factors (i.e., domain-specific factors) and a Global Executive Composite (GEC; i.e., general factor).

Recently, Roth, Lance, Isquith, Fischer, and Giancola (2013) implemented a second-order CFA to evaluate the latent structure of the BRIEF-A (Adult version), which is in line with the view of Chen et al. (2006) and Reise et al. (2010). Nonetheless, this has not been the case for most of the studies that evaluated the factorial structure underlying the BRIEF. Chen et al. (2006) asserted that bifactor modeling is less familiar than second-order CFA because it has been primarily used in the area of intelligence research. To the contrary, second-order CFA models are more widespread, as they have been used in a broader variety of research areas, including self-concept, psychological well-being, and personality (Chen et al., 2006). From Reise et al.'s (2010) perspective, bifactor modeling has not been well understood in personality research community and thus rarely used.

Regardless of why bifactor models are less known or used, this study purports to present some advantages of bifactor modeling over second-order CFA, analyzing the data of a Chilean sample of primary and secondary students who were evaluated using the BRIEF. Hence, some preliminary aspects of executive functioning and some studies analyzing the BRIEF's factorial structure will be briefly examined.

Executive functions are higher cognitive processes that account for a wide spectrum of mental abilities (Reynolds & Horton, 2008), which are considered essential to evoke efficient, creative, and socially adapted behaviors (Lezak, 1982). Most authors agree on its multidimensional nature, which includes planning, inhibition, flexibility, working memory, problem solving, and sequencing skills (Burgess, Veitch, de Lacy Costello, & Shallice, 2000; McCloskey & Perkins, 2012; Pennington & Ozonoff, 1996). Other authors (Bechara, Damasio, Damasio, & Lee, 1999; Damasio, 2005; Grafman & Litvan, 1999) also include personal abilities (e.g., regulation of social behavior and emotional control) as important components of executive functioning, which may be relevant for an adequate adaptation to everyday situations, at work, and to general social contexts. However, due to the complexity of executive functioning, a consensus definition among researchers, neuropsychologists, and cognitive scientists is still lacking.

Traditionally, executive functioning has been evaluated in laboratory or clinical settings using performance-based tests (e.g., Stroop, Tower of Hanoi, Wisconsin Card Sorting Test). On one hand, performance-based tests show high internal validity because they measure a limited but well-defined aspect of a particular behavior in a very controlled testing situation. But, on the other hand, these tests show low ecological validity and scarce representativeness of individual's functioning in real-life contexts (Bakar, Taner, Soysal, Karakas, & Turgay, 2011; García, González-Pienda, Rodríguez, Álvarez, & Álvarez, 2014; Gioia, Kenworthy, & Isquith, 2010).

Therefore, in order to improve ecological validity, some assessment scales evaluate executive functioning through the observation of person's behavior in specific situational environments (Chan, Shum, Toulopoulou, & Chen, 2008). This is the case of the BRIEF (Gioia et al., 2000), which addresses the ecological validity issue by asking parents, teachers, day care providers, or other informants about specific behaviors that can be displayed by the examinee in real-life situations (e.g., home, school, day care). The BRIEF has four versions: parent and teacher form, from 5 to 18 years old (BRIEF; Gioia et al., 2000); preschool version, from 2 to 5 years old (BRIEF-P; Gioia, Espy, & Isquith, 2003); self-report version, from 11 to 18 years old (BRIEF-SR; Guy, Isquith, & Gioia, 2004); and adult version, from 18 to 90 years old in informant or self-report forms (BRIEF-A; Roth, Isquith, & Gioia, 2005).

The theoretical model underlying the BRIEF defines executive functioning as a group of cognitive processes that are essential for problem solving (Gioia, Isquith, & Kenealy, 2008). Originally, the BRIEF included a group of eight scales: (a) Inhibit, (b) Shift, (c) Emotional Control, (d) Initiate, (e) Working Memory, (f) Plan/Organize, (g) Organization of Materials, and (h) Monitor. These scales were developed to assess different aspects of executive functioning, and allow obtaining a "Behavioral Regulation" index, and a "Metacognition" index, as well as a "Global Executive Composite" that accounts for a general executive functioning factor. Some studies (Alloway et al., 2009; Jarratt, Riccio, & Siekierski, 2005) showed high concurrent validity between the BRIEF, parent and teacher form, and other behavioral inventories such as Behavior Assessment System for Children parent and teacher form (Reynolds & Kamphaus, 2002) and Conners Teacher Rating Scale-Revised Short (Conners, 2005). Other investigations have reported only moderate correlations when estimating the association between executive functioning tests and the BRIEF's scales (Alloway et al., 2009; Bakar et al., 2011; Mahone et al., 2002; Toplak, Bucciarelli, Jain, & Tannock, 2009). The latter suggest that the above-mentioned measures may be targeting

> different levels of a hierarchically organized executive function system that can be conceptualized as a meta-construct. Each level gives rise to longer-term goals that require new abilities and skills so as to create increasingly more complex nested sets

> of goal-directed activities, organized and sustained across increasingly longer temporal durations and involving larger social networks to attain (Barkley & Fischer, 2011, p. 159).

The idea of a hierarchically organized structure for executive functioning has been well established by several theoretical models (Baddeley, 2012; Luria, 1984; Norman & Shallice, 1986; Posner & Petersen, 1990), which is in line with Barkley and Fischer's (2011) view. Assuming this conceptualization, the BRIEF's original model can also be seen as a hierarchical latent structure, composed by two first-order factors (i.e., domain-specific or indices) and a second-order factor (i.e., GEC).

In 2000, Gioia et al. first evaluated the construct validity of the BRIEF (teacher form) exploring the configuration of first-order factors, using a principal factor analysis approach. Nevertheless, the conducted analysis does not seem to be particularly suitable to evaluate the multidimensional nature of executive functioning (Chen et al., 2006; Reise et al., 2010), and even less so if only first-order factors were evaluated. As expected, results showed two first-order factors (i.e., Behavioral Regulation and Metacognition indices), which were moderately correlated ($r = .62$). Likewise, Gioia et al. (2000) found similar results for the parent form of the BRIEF, but the existence of a second-order general factor was not explored either. Later, Gioia, Isquith, Retzlaff, and Espy (2002) conducted a CFA in a clinical sample, using the BRIEF's parent form. Results did not replicate the two first-order correlated factors proposed by the BRIEF's original model. After these findings, Gioia et al. (2002) evaluated four different alternative models using the same analytical method. In the mentioned study, the authors reorganized the BRIEF's items into nine scales instead of the eight original ones. Results provided evidence for a three-factor model including: (a) Behavioral Regulation (i.e., Inhibit and Self Monitor scales), (b) Emotional Regulation (i.e., Emotional Control and Shift scales), and (c) Metacognition (i.e., Initiate, Working Memory, Plan/Organize, Organization of Materials, and Task Monitor scales). Again, the authors did not test the second-order general factor (i.e., GEC).

In 2008, Arango, Puerta, and Pineda conducted an exploratory factor analysis (EFA) to test the factorial structure of the BRIEF (parent and teacher form) in a sample of 128 adolescents. Findings revealed the existence of a single latent factor, which the authors named "Behavioral Supervision System" (Arango et al., 2008). This factor may be considered analogous to the "Global Executive Composite" originally proposed by Gioia et al. (2000). However, EFA may not be the most appropriate choice to test hierarchical models including first- and second-order factors either (Chen et al., 2006; Reise et al., 2010). Later on, Egeland and Fallmyr (2010) conducted a CFA in a sample of 158 children using both the parent and the teacher form as well as the nine scale's BRIEF version. Results also provided evidence for a three first-order factor latent structure for both BRIEF's forms. These results were also in line with Gioia et al.'s (2002) findings, which differentiated "Behavioral Regulation" from "Emotional Regulation" as related but separated first-order factors. The three first-order factor model (Gioia et al., 2002) showed better fit indices than the two first-order factors originally proposed (Gioia et al., 2000). More recently, Huizinga and Smidts (2011) obtained a two first-order factor structure underlying the eight original scales of the BRIEF (parent form). In 2014, García et al. observed the same underlying structure in a clinical sample of ADHD participants. Nevertheless, both studies used the exploratory method of principal component analysis, which also does not allow for testing hierarchical structure models.

Unlike all previous studies, Roth et al. (2013) conducted a second-order CFA to examine the hierarchical latent structure of the BRIEF-A that, in Chen et al.'s (2006) and Reise et al.'s (2010) views, is one of the two proper ways to simultaneously test multidimensionality and general factors. In their study, Roth et al. (2013) provide evidence for three first-order factors (i.e., Metacognition, Behavioral Regulation, and Emotional Regulation) and a second-order general factor (i.e., GEC), supporting the multidimensional nature of executive functioning. In sum, aside from Roth et al.'s (2013) study, most mentioned studies used analytical approaches that would not be the most appropriate choices to evaluate the factor structure underlying the BRIEF (Chen et al., 2006; Reise et al., 2010). As one might expect, most findings regarding the factorial structure of the BRIEF led also to different results, suggesting the existence of a latent structure composed by either two or three first-order latent factors.

Hence, the present study aims to implement a bifactor analysis as an alternative method to test the latent structure of the BRIEF. Moreover, it is also of high importance to highlight the advantages of bifactor modeling over second-order CFA when concurrently evaluating both multidimensionality and the existence of a general factor.

Like second-order CFA models, bifactor models include a number of group factors (i.e., first-order or domain-specific factors), a general factor, and an explicit bifactor structure (Jennrich & Bentler, 2012). In 1999, Yung, Thissen, and McLeod demonstrated that second-order models are in fact nested within bifactor models. That is, for every bifactor model, there is an equivalent "full second-order model" with direct factor loadings from the second-order factor to every observed variable, over and above the second-order effect on the first-order factors. Therefore, a "standard second-order CFA model" is a constrained case in which the effects from the second-order factor to the observed variables (i.e., items or scales) are eliminated (Chen et al., 2006). Since bifactor models consider both effects (over and above the first-order factors), it becomes particularly useful when domain-specific and general factors are of focal interest (Chen et al., 2006). The canonical bifactor model assumes orthogonality (i.e., uncorrelated factors) among domain-specific factors and between general and domain-specific factors.

Chen et al. (2006) and also Reise et al., (2010) summarized several advantages of bifactor over second-order CFA models. The main advantage of bifactor over second-order CFA models is that bifactor analysis allows observing directly to which extent an item or scale (i.e., observed variable) reflects a common target trait (i.e., general factor) and, simultaneously, to which extent it may reflect a subtrait (i.e., domain-specific). That is, bifactor modeling allows retaining a single common latent factor but also controls for the variance that arises due to additional common factors (Reise et al., 2010).

A second advantage of bifactor models is that in second-order, CFA is not possible to observe the direct relationships between the observed variables and the general factor, but rather an "indirect effect" or a "mediated relationship" through the first-order factors. Therefore, to estimate the variance attributable to the general factor, the loading of the observed variable on the domain-specific factor must be multiplied by the loading of the domain-specific factor on the general factor (Chen et al., 2006). In contrast, bifactor modeling provides all factor loadings (over and above the general factor) and allows identifying whether a domain-specific factor makes a unique contribution to the prediction of external criteria (Chen et al., 2006). Because in a bifactor model, general and domain-specific factors are orthogonal, a simple inspection of the factor loadings on the second- and first-order factors is informative.

Since the second-order CFA model is nested into the more general bifactor model (i.e., less restricted), the latter can be used as a baseline model to compare the model fit as long as the model becomes more constrained, which constitutes a third advantage of bifactor models over second-order CFA. For instance, in a standard restricted second-order CFA model, it is assumed that correlations among first-order factors occur because they have a common cause (i.e., general factor). Hence, observing low factor loadings from a domain-specific factor and its related observed variables, and high factor loadings between the same observed variables and the general factor may suggest that these variables are better explained by a general factor and do not constitute a domain-specific factor. Then, "To the degree that the items reflect primarily the general factor and have low loadings on the group factors, subscales make little sense" (Reise et al., 2010, p. 555).

To show the benefits of using bifactor modeling, this study analyzes the same data set under a second-order CFA model and under a bifactor model. Fit indices and diagrams for each model are provided, together with a comparison between both nested models (i.e., chi-square difference). The discussion section will describe some implications for research and clinical practice. Attending to the complexity of executive functions and the multifactorial nature of its measures, this study proposes a bifactor model structure that previous studies have not yet proposed.

Method

Participants

Participants were primary and secondary students from 18 educational institutions in Chile. Students with mild specific learning or language difficulties were not excluded from the sample, to support the external validity of the study. However, students who were not very familiar to their teachers were excluded from assessment. The final sample included 300 students (155 males and 145 females), randomly selected from different class levels, ranging from 5 to 18 years old (M = 11.3, SD = 3.7) (Table 1). Written inform consent was obtained from each participant. This study was conducted in accordance with the Helsinki Declaration of the World Medical Association (Williams, 2008).

Measures

BRIEF—Teacher Form (Spanish version). The BRIEF consists of 86 items to which teachers respond whether the student exhibits problems with specific behaviors

Table 1. BRIEF Chilean sample distribution.

Educational level	N	Girls	Boys	Ages M (SD)	Learning disability
Preschool[a]	23	13	10	5.35 (.49)	3
Elementary school (Primary)					
1st	23	16	7	6.65 (.71)	6
2nd	32	13	19	7.59 (.61)	14
3rd	26	15	11	9.04 (.92)	7
4th	24	10	14	9.58 (.72)	7
5th	25	12	13	10.64 (.70)	7
Middle school (Secondary)					
6th	22	13	9	11.64 (.66)	5
7th	22	8	14	12.64 (.66)	4
8th	24	15	9	13.67 (.87)	6
High school (Secondary)					
1st	21	7	14	14.76 (.77)	3
2nd	17	8	9	15.71 (.59)	1
3rd	20	12	8	16.45 (.51)	2
4th	21	13	8	17.19 (.51)	5
Total	300	155	145		70

[a]The equivalent for the pre-school level in the Chilean educational system is "Kindergarten".
BRIEF: Behavior Rating Inventory of Executive Function.

at school. Items are rated as 1 (*Never*), 2 (*Sometimes*), or 3 (*Often*). The BRIEF's items aggregate to eight clinical scales: (a) Inhibit, the ability to suppress impulses and to stop one's own behavior at the appropriate times; (b) Shift, the ability to adjust behavior flexibly to the changing demands of a situation; (c) Emotional Control, the capacity to modulate emotional responses; (d) Initiate, the capacity to initiate tasks or activities and independent generation of ideas, strategies, or responses; (e) Working Memory, the ability to keep information in mind while completing a task; (f) Plan/Organize, the capacity to manage current and future-oriented task demands; (g) Organization of Materials, the orderliness of work, play, and storage spaces; and (h) Monitor, the ability to check work and performance during and immediately after finishing a task (Isquith, Gioia, & PAR-Staff, 2008).

These scales allow calculating two composite scores: A Behavioral Regulation Index and a Metacognition Index, which sum allows obtaining a summary score called GEC (Isquith et al., 2008). In 2002, Gioia et al. proposed a new organization of the BRIEF items into nine clinical scales, parceling the Monitor subscale into two aspects: Social Behavior Regulation (Self Monitor) and Task Regulation (Task Monitor). The nine scale scoring system is aggregated into three composites: (a) Behavioral Regulation, (b) Emotional Regulation, and (c) Metacognition. The present study used Gioia et al.'s (2002) nine scale configuration.

Procedure

Psychological Assessment Resources provided the required authorization to conduct a validation study of the BRIEF test in a Chilean normative sample.

A total of 50 schools were contacted and 23 agreed to collaborate. School Directors gave authorization to contact teachers. After informed consent was obtained from the primary teachers of each class, one student from each class list was randomly selected and his teacher was asked to evaluate him/her with the BRIEF teacher form. Therefore, every teacher evaluated only one student. This was done to preserve the independence of the evaluations. Six research assistants, who were properly trained to apply the BRIEF helped during the assessment period. Research assistants provided the instructions to teachers, received inform consents, and solved in situ questions regarding the questionnaire.

Data Analysis

Descriptive statistics for the BRIEF's Chilean normative sample (Table 2) and internal consistency using Cronbach's alpha (Table 3) were estimated. Missing data were imputed using the expectation-maximization algorithm. These analyses were carried out using the IBM SPSS STATISTIC (2011) program, version 19.

Table 2. BRIEF descriptive statistics for the Chilean normative sample.

BRIEF scale	Min	Max	Mean	SD
Shift	10.00	30.00	15.94	4.87
Emotional Control	9.00	27.00	13.83	5.15
Initiate	7.00	21.00	12.70	3.98
Working Memory	10.00	30.00	18.00	6.07
Plan/Organize	10.00	30.00	17.96	5.56
Organization of Materials	7.00	21.00	10.74	4.08
Task Monitor	4.00	12.00	7.64	2.48
Inhibit	10.00	30.00	16.70	6.50
Self Monitor	6.00	18.00	10.28	3.65

BRIEF: Behavior Rating Inventory of Executive Function.

Table 3. BRIEF internal consistency.

BRIEF scale	Corrected item total correlation Min—Max	Number of items
Shift	.574–.704	10
Emotional Control	.653–.835	9
Initiate	.540–.723	7
Working Memory	.632–.828	10
Plan/Organize	.506–.735	10
Organization of Materials	.646–.828	7
Task Monitor	.525–.803	6
Inhibit	.710–.851	10
Self Monitor	.671–.712	4

BRIEF: Behavior Rating Inventory of Executive Function.

Second-order CFA and Bifactor Modeling. This analysis was conducted to evaluate the data fit to the theoretical model of the BRIEF proposed by Gioia et al. (2002), and recently tested by Roth et al. (2013). Consequently, a second-order CFA was implemented in order to identify whether the proposed BRIEF's factorial structure may be replicated for the Chilean normative sample. Mean raw score ratings from each scale (range 1–3) were used as input for all observed variables, using Robust Maximum Likelihood estimation method. All analyses were conducted using using MPlus 7.11 software (Muthén & Muthén, 2013).

Bifactor Modeling. Bifactor modeling with CFA (Harman, 1976; Holzinger & Swineford, 1937) was conducted to evaluate whether domain-specific factors and a general factor best represented the latent structure underlying the BRIEF. As Reise (2012) stated

> A bifactor structural model specifies that the covariance among a set of item responses can be accounted for by a single general factor that reflects the common variance running among all scale items and group factors that reflect additional common variance among clusters of items, typically, with highly similar content. It is assumed that the general and group factors all are orthogonal. (p. 667)

As previously stated, the difference between bifactor modeling and a second-order factor analysis is that, in the former, domain-specific factors are not used as indicators of the more general factor but are measured variables themselves. For that reason, bifactor models allow examining the role of the domain-specific factors in parallel to the role of a more general factor (Chen et al., 2006).

In this study, the bifactor model permitted each scale to load onto its respective first-order factor, and simultaneously onto a second-order factor. Correlations between first-order and second-order factors were set to zero to allow proper model identification. Following the canonical definition of bifactor modeling, correlations among first-order factors were also set to zero (i.e., orthogonal). Bifactor modeling allowed the examination of the BRIEF's clinical scales, regardless of the common aspects associated with a more general executive functioning factor (i.e., GEC). Bifactor modeling analysis was also conducted using MPlus 7.11 software (Muthén & Muthén, 2013).

Second-order CFA vs. Bifactor Analysis. In order to demonstrate the advantages of bifactor over second-order CFA, fit indices for each model were estimated and then compared. The difference between the two models' fit indices was estimated conducting a chi-squared difference test. The difference between these nested models was obtained (i.e., $\Delta\chi^2$) as well as the difference of each model's degrees of freedom (i.e., Δdf). If the $\Delta\chi^2$ difference value is significant, the model with more freely estimated parameters (i.e., larger model) fits the data better than the model in which the parameters in question are fixed (i.e., smaller model) (Kline, 2005). Chi-squared difference tests applied to nested models have essentially the same strengths and weaknesses as χ^2 tests applied to any model: They are directly influenced by sample size, and for large samples, even trivial differences may become significant (Kline, 2005).

Finally, the effect size of the $\Delta\chi^2$ difference between both models was also estimated using Cohen's ω coefficient, which values can be interpreted in terms of Cohen's (1988) suggested standards for effect size interpretation (i.e., small = .1; medium = .3; large = .5).

Results

Table 2 shows the main central tendency and dispersion statistics obtained for the Chilean normative sample. Missing data were less than 8% across data and less than 4.6% within cases.

A reliability analysis was performed for each clinical scale. Internal consistency coefficients for each scale are (a) Shift, $\alpha = .89$; (b) Emotional Control, $\alpha = .94$; (c) Initiate, $\alpha = .87$; (d) Working Memory, $\alpha = .94$; (e) Plan/Organize, $\alpha = .91$; (f) Organization of Materials, $\alpha = .92$; (g) Task Monitor $\alpha = .88$; (h) Inhibit, $\alpha = .95$; and (i) Self Monitor, $\alpha = .85$). Table 3 shows the corrected item-total correlations for each scale (i.e., *ai* parameter in Item Response Theory) and indicated similar discrimination levels among items.

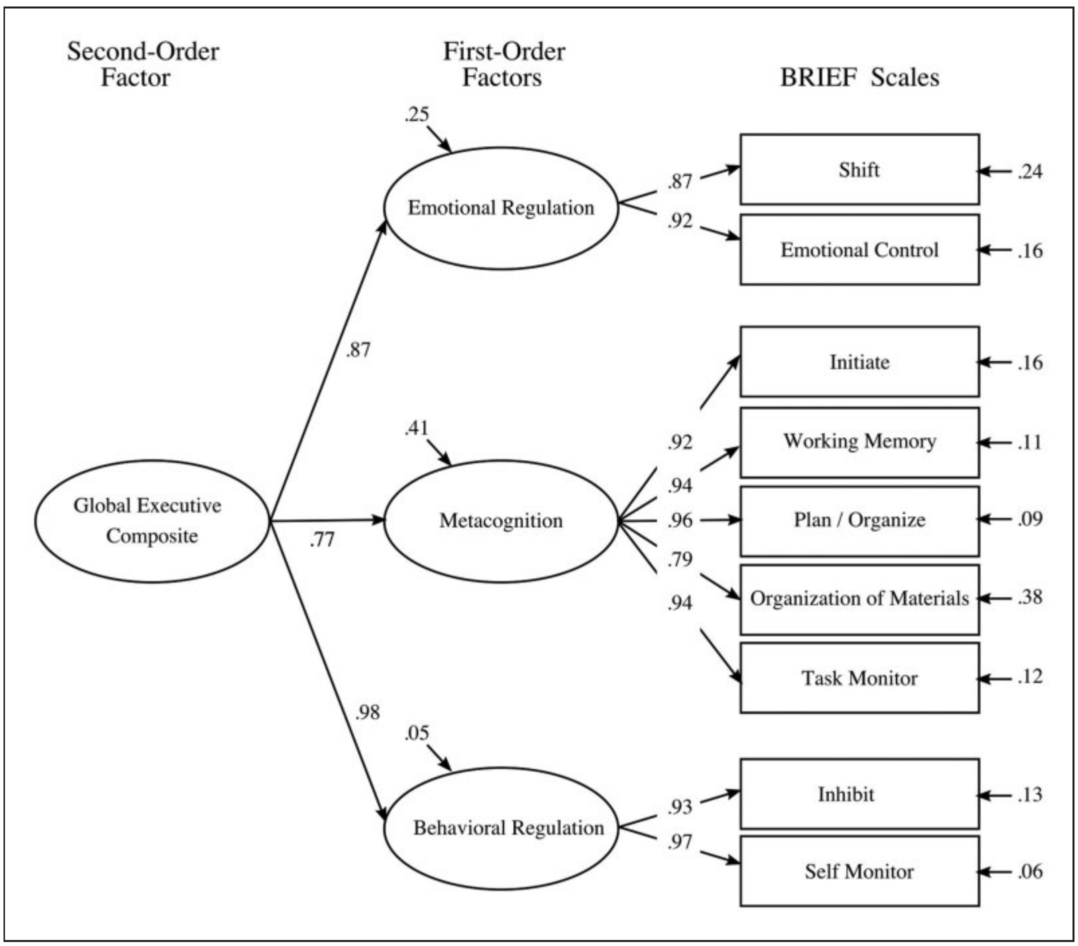

Figure 1. Diagram of the Second-Order Factor Structure underlying the BRIEF Scales. First-order factors: (a) Emotional Regulation; (b) metacognition; (c) Behavioral Regulation. Second-order factor: (a) global executive composite.

Table 4. Factor loadings for the BRIEF's first and second-order factors.

First-order factors						
BRIEF scale/factor	Emotional Regulation		Metacognition		Behavioral Regulation	
Shift	1.000	(.872)[a]				
Emotional Control	1.235	(.916)				
Initiate			1.000	(.917)		
Working Memory			1.097	(.942)		
Plan/Organize			1.022	(.957)		
Organization of Materials			0.879	(.785)		
Task Monitor			1.115	(.938)		
Inhibit					1.000	(.932)
Self Monitor					0.975	(.968)
Global executive composite (Second-order factor)	1.000	(.865)	1.092	(.771)	1.608	(.976)

Note: $N = 300$. BRIEF: Behavior Rating Inventory of Executive Function.
[a]Standardized solution is included in parentheses.

Second-order CFA: Model Fit Indices

Indices for the second-order model showed good fit to the data ($\chi^2 = 157.11$, $df = 24$, $p < .01$; CFI = .945; RMSEA = .136; SRMR = .043). Although RMSEA value was higher than recommended standards (i.e., RMSEA = .06 (Hu & Bentler, 1999), RMSEA = .07 (Steiger, 2007), factor loadings ranged between .771 and .968 (Figure 1 and Table 4).

Bifactor modeling: Model Fit Indices

Like second-order CFA analysis, the bifactor model showed good general fit indices ($\chi^2 = 128.64$, $df = 18$, $p < .01$; CFI = .954; RMSEA = .143; SRMR = .039). Likewise, the bifactor model showed a RMSEA value that was also higher than recommended standards (RMSEA = .06 –.07). On one hand, all direct factor loadings from the "General Factor" to the observed variables (i.e., BRIEF Scales) were significant and higher than those of the domain-specific factors to the observed variables (Figure 2 and Table 5). On the other hand, only two of the three domain-specific factors had significant factor loadings toward the BRIEF scales (i.e., Emotional Regulation = (a) Shift, and (b) = Emotional Control; and Metacognition = (a) Initiate, (b) Working Memory, (c) Plan/Organize, (d) Organization of Materials), and (e) Task Monitor (Figure 2). The factor loadings from Behavioral Regulation (i.e., (a) Inhibit and (b) Self Monitor) were low and not significant (Table 5).

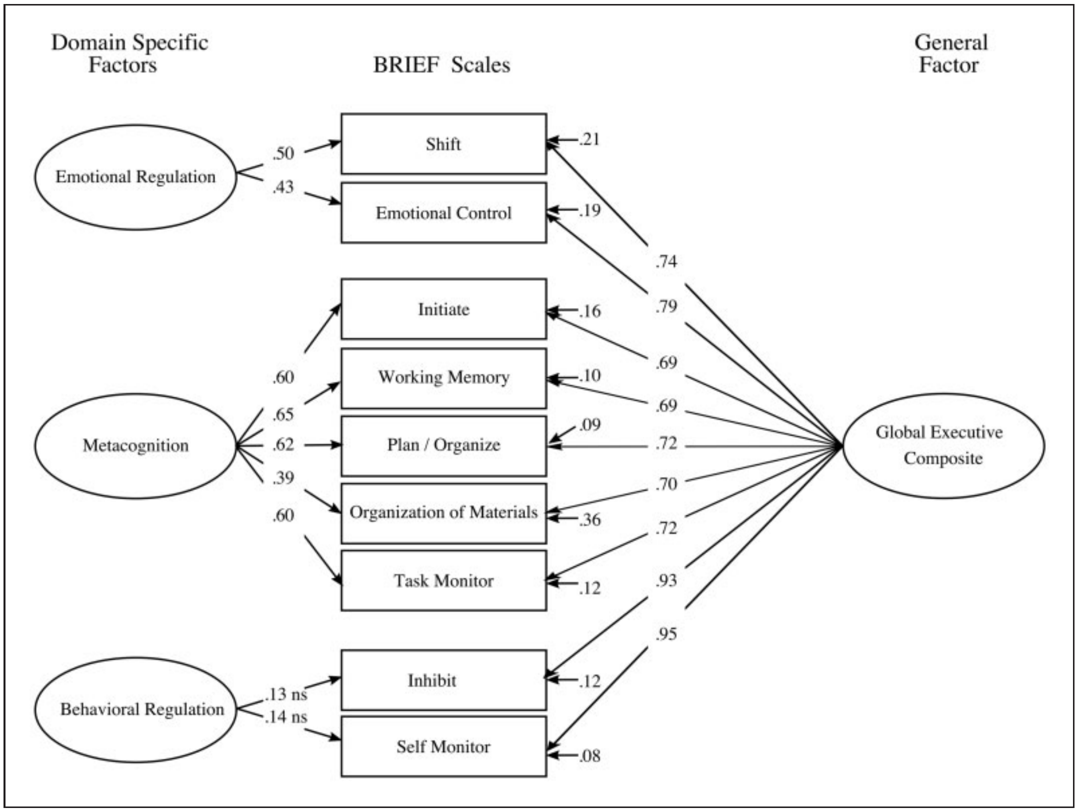

Figure 2. Diagram of the bifactor structure underlying the BRIEF scales.

Table 5. Factor loadings for the BRIEF's first and general and domain-specific factors.

BRIEF scale	Global executive composite	Emotional Regulation	Metacognition	Behavioral Regulation
Shift	1.000 (.738)[a]	1.000 (.500)		
Emotional Control	1.263 (.793)	1.004 (.427)		
Initiate	1.095 (.693)		1.000 (.601)	
Working Memory	1.157 (.686)		1.158 (.652)	
Plan/Organize	1.121 (.724)		1.016 (.624)	
Organization of Materials	1.132 (.698)		0.662 (.388)	
Task Monitor	1.241 (.721)		1.084 (.598)	
Inhibit	1.681 (.931)			1.000 (.129)[b]
Self Monitor	1.610 (.949)			1.006 (.139)[b]

Note: N = 300. BRIEF: Behavior Rating Inventory of Executive Function.
[a]Standardized solution is included in parentheses.
[b]Factor loading was not significant.

Table 6. Chi-square difference between second-order and bifactor model (nested comparison).

						Second order vs. bifactor		
	x^2	*df*	RMSEA (90% CI)	SRMR	CFI	Δx^2	Δdf	ω
Second-order model (Figure 1)	157.11*	24	.136 (.12–.16) *(ns)*	.043	.945	28.47*	6	.125
Bifactor model (Figure 2)	128.64*	18	.143 (.12–.17) *(ns)*	.039	.954			

Notes: N = 300; *df* = model degrees of freedom; RMSEA = root mean square error of approximation; 90% CI = 90% confidence interval for RMSEA; SRMR = standardized root mean residual; CFI = comparative fit index; Δx^2 = Chi-Square value for nested comparison; Δdf = degree freedom for nested comparison; ω = Cohen's ω (effect size).
*$p < .01$

Second-order CFA vs. Bifactor Modeling Comparison

The bifactor model solution showed a better fit to the data than the second-order factor solution. The Chi-square difference test was significant ($\Delta\chi^2 = 28.47$, $df = 6$, $p < .01$), which indicates that the larger model (i.e., bifactor modeling) had a better fit to the data than the more "restricted" second-order CFA model (Table 6).

Regarding the magnitude of the chi-squared difference, the Cohen's ω indicated a rather "small effect-size" ($\omega = .13$) between both models. However, since both models had good fit indices this finding might have been expected. Moreover, as previously mentioned, in relatively larger samples even slight differences may become significant.

Discussion

The multidimensional nature of executive functioning makes its evaluation a complex endeavor and a permanent challenge. Although Gioia et al. (2000) addressed the ecological validity issue in executive functioning assessment and proposed a hierarchical structure underlying the BRIEF, reported studies had neither tested a "domain-specific" and "general" factor structure nor implemented bifactor modeling as an analytical strategy. Except for Roth et al. (2013), most studies have conducted CFA and EFA to test only the structural model of the BRIEF (i.e., first-order factors) and few of them have explored a hierarchical and mediational structure (i.e., a hierarchical factor structure with first and second order factors) as proposed by the BRIEF's authors.

In this study, Roth et al.'s (2013) study was first replicated, using a second-order CFA, obtaining good fit indices for a solution including three first-order factors with a second-order general factor. In this case, obtained results were in line with both Gioia et al.'s (2002) and Roth et al.'s (2013) findings. It is important to emphasize that the proposed structural model also assumed the same three first-order domains of executive functioning (i.e., Emotional Regulation, Metacognition, and Behavioral Regulation) observed by Gioia et al. (2002) and also replicated by other researchers (e.g., Egeland & Fallmyr, 2010; Gioia & Isquith, 2002; Roth et al., 2013). In sum, implementing a second-order CFA approach, this article replicated Roth et al.'s (2013) findings in a Chilean normative sample.

The second analysis implemented bifactor modeling as a method that may better account for the underlying structure of the BRIEF. Furthermore, it is proposed that bifactor modeling might also be useful to evaluate the pertinence and interpretability of having domain-specific factors and, simultaneously, a general—but not hierarchical—factor. When implementing second-order factor analysis, higher-order factors (i.e., General) derivate from first-order factors (i.e., domain-specific). Hence, lower-order factors partially account for the second-order factor variance. Conversely, in a bifactor analysis, both first- and second-order factors simultaneously account for the direct variance of the whole measurement model (i.e., test scales and observed variables).

Results obtained from bifactor modeling showed good fit indices and significantly better to those observed when conducting second-order CFA. Outcomes provide evidence for an underlying common general factor (i.e., GEC) as has been reported by Gioia et al. (2000) and Arango et al. (2008). More importantly, results also provide evidence for the existence of two latent domain-specific factors (i.e., Metacognition and Emotional Regulation) explaining variance over and above the general factor, as first proposed by Gioia et al. (2000).

Regarding the Behavioral Regulation Factor, reported findings suggest that there is explained variance that is not uniquely attributable to this domain-specific factor, since there is some variance that also accounts for a more General Factor (i.e., GEC). Moreover, Behavioral Regulation does not show significant loadings for its related scales (i.e., Inhibit and Self Monitor). This finding could imply that Behavioral Regulation does not provide evidence for the existence of an independent domain-specific trait of executive functioning for the tested model. However, the Emotional Regulation factor showed significant factor loadings on its observed variables (i.e., Shift and Emotional control).

This finding might suggest that the so-called Emotional Regulation factor may be better represented by what Gioia et al. (2000) originally called Behavioral Regulation factor (excepting from the Inhibit scale, which had a non-significant factor loading). Nevertheless, this interpretation is not consistent with Gioia et al.'s (2002) findings, which differentiated "Behavioral Regulation" from "Emotional Regulation" as related but separated factors. In this line, it may be necessary to replicate a bifactor modeling approach in different samples to clarify this sensitive aspect.

From a psychometric perspective, reliability analysis showed excellent internal consistency of the BRIEF's scales, supporting the possibility of interpreting the BRIEF results in terms of both, a single global score (i.e., GEC) and more specific first-order composites (i.e., domain-specific factors). In this case, only the two first-order factors that showed significant factor loadings (i.e., Emotional Regulation, Metacognition).

In order to obtain interpretable scale scores that reflect the variability attributable to their latent contributors, it can be suggested to calculate weighted scores by multiplying each scale's raw score by its respective non-standardized regression coefficient. These weighted scores may allow for a good interpretation of both single and global scores. However, it must be stated that these regression coefficients must be estimated for each normative sample, since they will vary depending on the sample's characteristics.

As previously stated, interpreting a single global score as well as domain-specific scores has research and clinical implications. An important implication for research is that a single global score may allow the independent testing of predictive relationships between domain-specific factors and external criteria over and above the general second-order factor (Chen et al., 2006). This issue is relevant for the BRIEF questionnaire because, even though some of its item-behavior descriptions mainly indicate to use or not a particular well-delimited cognitive domain, most item-behavior descriptions are more likely to account for one or more aspects of many cognitive domains, which have the potential of being related to executive functioning, and thus integrating a cluster (McCloskey & Perkins, 2012).

From a clinical standpoint, the existence of a general factor accounting for executive functioning has been widely proposed (e.g., Baddeley, 2012; Luria, 1984; Norman & Shallice, 1986; Posner & Petersen, 1990) as a mechanism that processes, or integrates, many other sub-functions, which make possible the resolution a novel problem or situation (Detterman, 1982). For that reason, a clinical advantage of interpreting a global score might be precisely the ability of assessing the underlying common mechanism that explains some variations in frontal functioning (Miyake et al., 2000). The latter idea would allow for evaluating executive functioning impairments in a broadly defined manner, which might not have been detected separately using single scale scores.

A limitation of this study is the relatively small sample size used and the lack of geographical representativeness. However, the sample did not exclude children with learning disabilities, making the sample more heterogeneous in this aspect. On the other hand, a clear advantage of using bifactor modeling is to open the possibility of measuring a single latent trait, while controlling for the variance explained by an additional common factor. In this way, the researcher would be able to explore the extent to which items or scales represent a common target trait and the extent to which they represent subtraits (Reise et al., 2010).

Finally, the results encourage cognitive science researchers and test developers to explore the potential benefits of using bifactor modeling as a way to evaluate the latent dimensions underlying cognitive assessment instruments.

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