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Hierarchical Factor Structure of the Cognitive Assessment System: Variance Partitions From the Schmid–Leiman (1957) Procedure

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Orthogonal higher-order factor structure of the Cognitive Assessment System (CAS; Naglieri & Das, 1997a) for the 5–7 and 8–17 age groups in the CAS standardization sample is reported. Following the same procedure as recent studies of other prominent intelligence tests (Dombrowski, Watkins, & Brogan, 2009; Canivez, 2008; Canivez & Watkins, 2010a, 2010b; Nelson & Canivez, 2011; Nelson, Canivez, Lindstrom, & Hatt, 2007; Watkins, 2006; Watkins, Wilson, Kotz, Carbone, & Babula, 2006), three- and four-factor CAS exploratory factor extractions were analyzed with the Schmid and Leiman (1957) procedure using MacOrtho (Watkins, 2004) to assess the hierarchical factor structure by sequentially partitioning variance to the second- and first- order dimensions as recommended by Carroll (1993, 1995). Results showed that greater portions of total and common variance were accounted for by the second-order, global factor, but compared to other tests of intelligence CAS subtests measured less second-order variance and greater first-order Planning, Attention, Simultaneous, and Successive (PASS) factor variance.

Keywords: CAS, construct validity, hierarchical exploratory factor analysis, Schmid–Leiman higher-order analysis, structural validity

Recent revisions of major intelligence tests such as the Wechsler Intelligence Scale for Children–Fourth Edition (WISC–IV; Wechsler, 2003), Wechsler Adult Intelligence Scale–Fourth Edition (WAIS–IV; Wechsler, 2008), and the Stanford–Binet Intelligence Scales–Fifth Edition (SB–5; Roid, 2003) have utilized Carroll's (1993) model of the structure of cognitive abilities to facilitate subtest and factor selection and to aid in interpretation of scores and performance. Carroll's (1993, 2003) three-stratum theory of cognitive abilities is hierarchical in nature and proposes 50–60 narrow abilities (Stratum I), 8–10 broad ability factors (Stratum II), and Spearman's general ("g") ability factor (Spearman, 1904, 1927) at the top (Stratum III). Because the narrow abilities and broad ability factors are correlated, subtest performance on cognitive abilities tests reflect combinations of both first-order and second-order factors. Because of this Carroll argued that variance from the higher-order factor should be extracted first to residualize the lower-order factors, leaving them orthogonal to each other and the higher-order factor. Thus, variability associated with a higher-order factor is accounted for prior to interpreting variability associated with lower-order factors. Statistically, this apportioning of variance to the higher-order and lower-order dimensions may be conducted using the Schmid and Leiman (1957) procedure. This procedure was recommended by Carroll (1993, 1995, 1997, 2003); McClain (1996); Gustafsson and Snow (1997); Carretta and Ree (2001); Ree, Carretta, and Green (2003); and Thompson (2004). Specifically, Carroll (1995) noted:

> I argue, as many have done, that from the standpoint of analysis and ready interpretation, results should be shown on the basis of orthogonal factors, rather than oblique, correlated factors. I insist, however, that the

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> orthogonal factors should be those produced by the Schmid–Leiman, 1957, orthogonalization procedure, and thus include second-stratum and possibly third-stratum factors. p. 437.

To date, there have been several hierarchical factor structure investigations of modern editions of major intelligence tests where the Schmid and Leiman (1957) procedure was used to apportion variance. Four WISC–IV investigations (Bodin, Pardini, Burns, & Stevens, 2009; Watkins, 2006, 2010; Watkins, Wilson, Kotz, Carbone, & Babula, 2006) indicated that most common and total variance was associated with the higher-order general intelligence factor and substantially lesser amounts at the first-order factor level. Two investigations of the WAIS–IV also found that for the total standardization sample and the adolescent subsample (ages 16–19), the higher-order general intelligence factor accounted for substantially greater portions of common and total variance than the first-order factors (Canivez & Watkins, 2010a, 2010b). These studies concluded that interpretation of the WISC–IV and WAIS–IV should focus on the global FSIQ as the primary level of interpretation rather than the four first-order factor index scores as recommended in the respective manuals (Wechsler, 2003, 2008).

Similar results and even greater portions of total and common variance were apportioned to the higher-order general intelligence factor than lower-order factors in the SB–5 (Canivez, 2008). Results for the SB–5 showed no evidence for the purported five-factor model, a finding also reported by DiStefano and Dombrowski (2006). Higher-order factor structure studies of the Reynolds Intellectual Assessment Scales (RIAS; Reynolds & Kamphaus, 2003) using the Schmid and Leiman (1957) procedure also revealed primacy of general intelligence (Dombrowski, Watkins, & Brogan, 2009; Nelson & Canivez, 2011; Nelson, Canivez, Lindstrom, & Hatt, 2007) due to substantially larger portions of variance apportioned to the higher-order general intelligence factor than first-order factors. Reynolds and Kamphaus created the RIAS primarily to measure general intelligence in an efficient manner. A joint factor analysis investigation of the Wide Range Intelligence Test (WRIT; Glutting, Adams, & Sheslow, 2000) and the Wechsler Abbreviated Scale of Intelligence (WASI; Psychological Corporation, 1999) also found most total and common variability was associated with general intelligence and smaller portions of variance were apportioned to the first-order factors; supporting primary interpretation of the Full Scale IQ (FSIQ) and General IQ (GIQ) (Canivez, Konold, Collins, & Wilson, 2009).

The Cognitive Assessment System (CAS; Naglieri & Das, 1997a) is a standardized cognitive ability/intelligence measure based on the PASS theory (Das, Naglieri, & Kirby, 1994) that has direct links to Luria's neuropsychological theory (Luria, 1966a, 1966b, 1973, 1980, 1982). The theoretical PASS model of cognitive processes (viz., Planning, Attention, Successive, and Simultaneous) purports to measure Luria's three functional units of the brain (Unit 1–Attention, Unit 2–Successive & Simultaneous, Unit 3–Planning). There are 12 subtests that combine to measure four PASS factors (three for each PASS factor), and PASS scores were proposed as the primary interpretive level for the CAS. A Full Scale score, based on the sum of all 12 subtest scaled scores (or eight subtest scaled scores for the Basic Battery) is also provided. Although Naglieri (1999) denied an implied hierarchical CAS structure or measurement of a general intelligence with the Full Scale score, psychometrically the CAS seems similar to other measures of intelligence or cognitive abilities in that PASS scales are correlated and that implies a hierarchical model of measurement (Gorsuch, 1983; Thompson, 2004) with a general dimension at the apex and represented by the overall, global Full Scale score. The nature of that general dimension may, however, be different.

Naglieri and Das (1997b) provided mixed results regarding the factor structure of the CAS with the standardization sample. At different age levels, the CAS showed some empirical support for a four-factor model (PASS) while at other age levels empirical support for a three-factor model ([PA]SS; a combined Planning/Attention factor) was observed in both exploratory (EFA) and confirmatory (CFA) factor analyses. In the reported EFA the four-factor model appeared sufficient for the 5–7 and 14–17 age groups while the three-factor model appeared sufficient for the 8–10 and 11–13 age groups based on the maximum likelihood χ^2. It was reported in the CAS Interpretive Handbook that oblique and orthogonal rotations of extracted factors were performed, but only the orthogonal (varimax) coefficients were presented (and only for the three-factor extraction) and this illustrates relationships between subtests and factors that were forced to be uncorrelated. Also, by not presenting subtest factor coefficients for the four-factor extraction, it is not possible to determine simple structure or cross-loading subtests that might illustrate the lack of simple structure of the four-factor model. Tables A.1 through A.11 in the CAS Interpretive Manual present intercorrelations between the CAS subtests and PASS scales, and PASS scales were moderately correlated (as PASS theory posits). When factors are correlated it is quite informative to report factor pattern and structure coefficients from oblique rotations and resulting extracted factor intercorrelations (Thompson, 2004). When factors are correlated then a higher-order structure is implied and should be examined (Gorsuch, 1983; Thompson, 2004). Higher-order or hierarchical factor analyses (EFA or CFA) were also not reported in the CAS Interpretive Handbook (Naglieri & Das, 1997b). Thus, EFA results were insufficiently described and such explication would be informative.

Hierarchical structure of the 10 experimental PASS subtests that were precursors to the CAS subtests included in standardization was examined by Kranzler and Weng (1995). The sample for analyses included the 132 students in Grades 5–12 from an earlier study (Naglieri, Das, Stevens, & Ledbetter, 1991). Too few students in Grades K-2 ($N = 73$) who were administered only nine subtests precluded CFA analyses for the younger sample. In comparing three models, the theoretical four-factor PASS, a three-factor (PA)SS, and PASS + *g* (a hierarchical model) to the null model; all three showed satisfactory fit to the data and improvements over the null model based on reported fit statistics. While the root mean square error of approximation (RMSEA) fit statistic was not reported in this study, calculations using data (χ^2, *df*, and *n*) provided in Kranzler and Weng's Table 1 produced RMSEAs of .10 for the null model, .033 for the PASS + g, .017 for the PASS, and .013 for the (PA)SS. Thus, all three models provided close fit to the data and very small differences between them. Although a (PA)SS + g model was not tested against the other models, the (PA)SS Model was further examined using the Schmid and Leiman (1957) orthogonal transformation to apportion subtest variance to the first-order (PA)SS dimensions as well as the higher-order g dimension. While previously mentioned EFA based studies with Schmid and Leiman orthogonal transformations apportioned subtest variance to all first-order factors and the second-order factor, the coefficients presented in Kranzler and Weng's indicated that subtest variance was only apportioned to the theoretically assigned first-order factor (CFA restricted subtest association to only one first-order dimension) and the second-order factor (Kranzler & Weng, 1995). The resulting g loadings ranged from "poor" to "fair" using Kaufman's (1994) criteria and none met the .70 criterion for a "good" g loading. Results indicated that most variance was apportioned to the higher-order g factor as observed in studies of the WISC-IV, RIAS, SB-4, WASI, and WRIT. Schmid and Leiman orthogonal transformation was not provided for the PASS + g model. While interesting, there were significant changes to the experimental tasks, deletion of tasks, and addition of tasks, such that there are no direct implications for the standardized version of the CAS.

In what appears to be the only published independent factor analysis study with the CAS standardization sample, Kranzler and Keith (1999) reported a number of model comparisons to test the CAS theoretical model using CFA procedures. Kranzler and Keith compared various structural models in order to determine which model fit the CAS standardization sample data best and included uncorrelated (orthogonal) and correlated (oblique) four first-order PASS factor models as well as three- ([PA]SS), two- ([PA][SS]), and one-factor (g) models. Of these, Kranzler and Keith indicated that the correlated PASS models provided the best fit to the CAS standardization data.

Kranzler and Keith (1999) also examined hierarchical models for the CAS implied by the correlated PASS models. Results suggested that a general, higher-order factor (psychometric g) accounted for PASS variability and because of relations between Planning and Attention subtests, an intermediate factor related to a combination of Planning and Attention was proposed. Although their third-order hierarchical model produced improved model fit; Heywood cases indicated problems with the model such that a constrained model was necessary. Further, the factor correlations appear to be overestimated when compared to zero-order Pearson product-moment correlations for the PASS scale scores in Tables A.1 through A.11 in the CAS Interpretive Manual (Naglieri & Das, 1997b). Asparouhov and Muthen (2009) noted:

> When nonzero cross-loadings are specified as zero, the correlation between factor indicators representing different factors is forced to go through their main factors only, usually leading to overestimated factor correlations and subsequent distorted structural relations. (p. 398).

Thus, it is possible that some of the conclusions of Kranzler and Keith (1999), which Naglieri (1999) challenged on statistical and theoretical grounds, may be due, in part, to inherent problems in CFA procedures when cross-loadings are nonzero. It also seems that psychometric issues may exist with the interrelationships among subtests of the CAS, which further complicates matters. Inexplicably, Kranzler and Keith did not include a Schmid and Leiman (1957) orthogonal transformation to apportion subtest variance to the first- and second-order factors (or third-order) as was done in Kranzler and Weng (1995). Thus, an understanding of where subtest variance resides remains unclear.

CAS PASS factors are correlated; however, the zero-order factor intercorrelations are somewhat lower than those observed in other intelligence tests such as the WISC–IV, WAIS–IV, and especially the SB–5. This suggests that although a hierarchical measurement model appears present, as illustrated by Kranzler and Keith (1999), greater amounts of variability may be measured by the PASS first-order factors (what Carroll referred to as Stratum II dimensions) than the WISC–IV, WAIS–IV, and SB–5 factors. This seems necessary if one intends to use first-order (PASS) factor-based scores for interpretation and differential diagnosis as suggested in the CAS Interpretive Manual (Naglieri & Das, 1997b). Thus, as with other intelligence measures it is necessary to understand how subtest variance is apportioned to higher- versus lower-order dimensions so that interpretive weight of higher- versus lower-order dimensions is appropriately assigned and considered.

The purpose of the present study was to examine proportions of CAS subtest variance attributed to the higher-order dimension and to the first-order dimensions as insisted by Carroll (1995). Thus, this study is descriptive, rather than theory testing to determine which of various models fits data best. While other hierarchical investigations using CFA have been used to Test CAS "theory," the potentially complex nature of CAS subtest associations with more than one latent factors (cross-loadings) but restricting models permitting subtests loading on only one first-order factor (Kranzler & Keith, 1999; Kranzler & Weng, 1995) may have resulted in overestimation of relationships between latent factors. Using an EFA procedure to describe subtest variance apportions to all first- and second-order factors as has been previously reviewed would describe variance similarly to a CFA method that allowed subtest associations with all first-order factors (rather than restricting subtests to association to only one first-order factor). This is the first study to examine and describe the hierarchical structure of the CAS standardization sample using the Schmid and Leiman (1957) transformation procedure in order to apportion CAS subtest score variance to the first-order (PASS or [PA]SS) and second-order (psychometric g) factors and to identify proportions of common and total CAS variance attributed to a higher-order general factor and the lower-order (PASS or [PA]SS) scales. Because the CAS Interpretive Manual (Naglieri & Das, 1997b) and other studies (Kranzler & Keith, 1999; Kranzler & Weng, 1995) presented evidence for three and four first-order factor models, both were examined for variance apportions in the present study although only standard scores for four first-order factors are provided in the CAS for clinical use.

Method

Participants

Participants were members of the two age groups and subtest configurations from the CAS standardization sample (Naglieri & Das, 1997a) and included 801 5–7 year olds and 1,224 8–17 year olds who had complete subtest score data necessary for factor analytic procedures. Demographic characteristics are provided in detail in the CAS Interpretive Manual (Naglieri & Das, 1997b). The standardization sample was stratified and randomly sampled and closely matched the United States Census data on key demographic variables of geographic region, parent education level, race/ethnicity, and sex.

Instrument

The CAS (Naglieri & Das, 1997a) is a cognitive assessment instrument based on the PASS theory and has a nationally representative standardization sample. A global, Full Scale score is provided in addition to four composite factor scores (Planning [P], Attention [A], Successive [SUC], and Simultaneous [SIM]) that represent dimensions from the PASS intelligence theory (Das, Naglieri, & Kirby, 1994). CAS subtest scores are scaled scores (mean [*M*] = 10, standard deviation [*SD*] = 3), whereas the PASS factor scores and Full Scale score are commonly scaled standard scores (*M* = 100, *SD* = 15). Extensive psychometric information regarding estimates of score reliability and validity are provided in the interpretive handbook (Naglieri & Das, 1997b).

Analyses

The CAS standardization samples (5–7 year olds and 8–17 year olds) were used in separate EFAs due to different subtest compositions. Following previous hierarchical EFA studies of prominent intelligence tests (Dombrowski et al., 2009; Canivez, 2008; Canivez et al., 2009; Canivez & Watkins, 2010a, 2010b; Nelson & Canivez, 2011; Nelson et al., 2007; Watkins, 2006; Watkins et al., 2006), the present study used principal axis extraction with two iterations in first-order factor extraction in estimating final communality estimates (Gorsuch, 2003) followed by oblique (Promax) rotations. The Promax rotated factors correlation matrix was then factor analyzed (second-order) to obtain factor coefficients and communality estimates for use in the Schmid and Leiman (1957) procedure as programmed in the MacOrtho computer program (Watkins, 2004). This transforms

> "an oblique factor analysis solution containing a hierarchy of higher-order factors into an orthogonal solution that not only preserves the desired interpretation characteristics of the oblique solution, but also discloses the hierarchical structuring of the variables"

(Schmid & Leiman, 1957, p. 53). Because the CAS manual presented mixed EFA and CFA evidence for both three- ([PA]SS) and four-factor (PASS) models, the present study also examined hierarchical structures and variance partitions for both the three- and four-factor models.

Results

Ages 5–7

Four-factor model—PASS. Table 1 presents factor structure coefficients and apportioned variance estimates in the four-factor model for 5–7 year olds. Results indicated that the second-order (general) factor accounted for 53.1% of the common variance and 23.9% of the total variance. The general factor accounted for between 10% and 39% (median [*Mdn*] = 24.5%) of individual subtest variability. The percent of common variance apportioned to the four first-order factors were as follows: Planning (17.2%), Successive (19.1%), Simultaneous (6.7%), and Attention (4.0%). The percent of total variance apportioned to the four first-order factors were: Planning (7.7%), Successive (8.6%), Simultaneous (3.0%), and Attention (1.8%). CAS subtests were generally associated with the theoretically consistent PASS factors. However, cross-loading was observed in that the Number Detection and Receptive Attention subtests had greater portions of variance apportioned to the Planning factor than to the Attention factor while the Expressive Attention subtest retained greater variance with the Attention factor. The Verbal-Spatial Relations subtest had equivalent variance apportioned to both the Successive and Simultaneous factors. The first- and second-order factors combined to measure 45.0% of the variance in CAS scores resulting in 55.0% unique variance (combination of specific and error variance).

Three-factor model—(PA)SS Table 2 presents factor structure coefficients and apportioned variance estimates for the three-factor model for 5–7 year olds. Results indicated that the second-order (general) factor accounted for 56.3% of the common variance and 28.4% of the total variance. The general factor accounted for between 12% and 50% (*Mdn* = 22%) of individual subtest variability. The percent of common variance apportioned to the three first-order factors were: Planning/Attention (25.5%), Successive (16.6%), and Simultaneous (1.5%). The percent of total variance apportioned to the four first-order factors were as follows: Planning/Attention

Table 1
Factor Structure Coefficients and Variance Sources for the CAS Standardization Sample Based on the Orthogonalized Higher-Order Factor Model From Four First-Order (PASS) Factors for Ages 5–7 (n = 801)

	General		Factor I (Planning)		Factor II (Successive)		Factor III (Simultaneous)		Factor IV (Attention)			
CAS subtest	b	$\%S^2$	b	$\%S^2$	b	$\%S^2$	b	$\%S^2$	b	$\%S^2$	h^2	u^2
Matching Numbers	.58	34	**.49**	**24**	.03	0	.05	0	–.12	2	0.59	0.41
Planned Codes	.56	31	**.44**	**19**	–.01	0	.07	1	–.09	1	0.52	0.48
Planned Connections	.51	26	**.40**	**16**	.12	2	–.06	0	–.01	0	0.44	0.56
Nonverbal Matrices	.48	23	.04	00	.02	0	**.37**	**14**	–.01	0	0.37	0.63
Verbal–Spatial Relations	.41	17	–.06	00	***.22***	***5***	**.22**	**5**	.12	1	0.28	0.72
Figure Memory	.52	27	.06	00	–.05	0	**.38**	**15**	.04	0	0.42	0.58
Expressive Attention	.37	14	.00	00	.03	0	.05	0	**.34**	**12**	0.26	0.74
Number Detection	.62	39	***.41***	***16***	–.10	1	.04	0	**.12**	**2**	0.58	0.42
Receptive Attention	.60	36	***.38***	***15***	–.02	0	–.05	0	**.20**	**4**	0.55	0.45
Word Series	.32	10	.00	00	**.60**	**36**	–.01	0	–.04	0	0.46	0.54
Sentence Repetition	.41	17	–.03	00	**.63**	**39**	.06	0	.01	0	0.56	0.44
Speech Rate	.37	14	.10	01	**.45**	**20**	–.08	1	.07	1	0.36	0.64
% Total S^2		23.9		7.7		8.6		3.0		1.8	45.0	55.0
% Common S^2		53.1		17.2		19.1		6.7		4.0	—	—

Note. CAS = Cognitive Assessment System; b = subtest factor structure coefficient (loading); $\%S^2$ = percent of variance explained in the subtest; h^2 = communality; u^2 = uniqueness (specific and error variance). Coefficients and variance percents for subtests within their theoretically assigned factor are highlighted in bold. Coefficients and variance percents in bold italic represent cross-loading on alternate factor.

(12.9%), Successive (8.4%), and Simultaneous (0.8%). CAS subtests were generally associated with the proposed theoretical factors but cross-loading was again observed with the Verbal–Spatial Relations subtest in the three-factor model having a greater proportion of first-order variance apportioned to the Successive factor. The first- and second-order factors combined to measure 50.4% of the variance in CAS scores resulting in 49.6% unique variance (combination of specific and error variance).

Ages 8–17

Four-factor model—PASS Table 3 presents factor structure coefficients and apportioned variance estimates for the four-factor model for 8–17 year olds. Results indicated that the second-order (general) factor accounted for 57.2% of the common variance and 28.9% of the total variance. The general factor accounted for between 12% and 43% (*Mdn* = 28.5%) of individual subtest variability. The percent of common variance apportioned to the four first-order factors were: Planning (21.4%), Successive (13.9%), Simultaneous (5.1%), and Attention (2.3%). The percent of total variance apportioned to the four first-order factors were as follows: Planning (10.8%), Successive (7.0%), Simultaneous (2.6%), and Attention (1.2%). As with the 5–7 age group, cross-loading was observed with Expressive Attention, Number Detection, and Receptive Attention all having greater portions of variance apportioned to the Planning factor than the Attention factor. No other subtests demonstrated substantial cross-loadings. The first- and second-order factors combined to measure 50.4% of the variance in CAS scores resulting in 49.6% unique variance (combination of specific and error variance).

Three-factor model—(PA)SS Table 4 presents factor structure coefficients and apportioned variance estimates for the three-factor model for 8–17 year olds. Results indicated that the second-order (general) factor accounted for 57.8% of the common variance and 28.5% of the total variance of CAS subtests. The general factor accounted for between 12% and 41% (*Mdn* = 27.5%) of individual subtest variance. The percent of common variance apportioned to the three first-order factors were as follows: Planning/

Table 2
Factor Structure Coefficients and Variance Sources for the CAS Standardization Sample Based on the Orthogonalized Higher-Order Factor Model From Three First-Order (PA)SS Factors for Ages 5–7 (n = 801)

	General		Factor I (Planning/ Attention)		Factor II (Successive)		Factor III (Simultaneous)			
CAS subtest	*b*	$\%S^2$	*b*	$\%S^2$	*b*	$\%S^2$	*b*	$\%S^2$	h^2	u^2
Matching Numbers	.69	47	**.53**	**28**	.01	0	.02	0	0.76	0.24
Planned Codes	.64	41	**.49**	**24**	–.02	0	.03	0	0.66	0.34
Planned Connections	.65	42	**.50**	**25**	.12	1	–.03	0	0.69	0.31
Nonverbal Matrices	.35	13	.03	00	.02	0	**.19**	**4**	0.16	0.84
Verbal–Spatial Relations	.34	12	–.02	00	***.23***	***5***	**.12**	**1**	0.19	0.81
Figure Memory	.39	15	.09	01	–.05	0	**.19**	**4**	0.20	0.80
Expressive Attention	.35	12	**.19**	**04**	.06	0	.04	0	0.16	0.84
Number Detection	.69	48	**.58**	**34**	–.08	1	.03	0	0.83	0.17
Receptive Attention	.71	50	**.60**	**36**	.00	0	–.02	0	0.85	0.15
Word Series	.40	16	–.03	00	**.59**	**34**	–.01	0	0.50	0.50
Sentence Repetition	.46	21	–.05	00	**.62**	**38**	.03	0	0.60	0.40
Speech Rate	.48	23	.17	03	**.45**	**20**	–.04	1	0.46	0.54
% Total S^2		28.4		12.9		8.4		0.8	50.4	49.6
% Common S^2		56.3		25.5		16.6		1.5	—	—

Note. CAS = Cognitive Assessment System; *b* = subtest factor structure coefficient (loading); $\%S^2$ = percent of variance explained in the subtest; h^2 = communality; u^2 = uniqueness (specific and error variance). Coefficients and variance percents for subtests within their theoretically assigned factor are highlighted in bold. Coefficients and variance percents in bold italic represent cross-loading on alternate factor.

Attention (24.5%), Successive (15.2%), and Simultaneous (2.4%). The percent of total variance apportioned to the three first-order factors were: Planning/Attention (12.1%), Successive (7.5%), and Simultaneous (1.2%). Subtests were associated with their theoretically consistent factors and no cross-loadings were observed. The first- and second-order factors combined to measure 49.3% of the variance in CAS scores resulting in 50.7% unique variance (combination of specific and error variance).

Discussion

As observed in the orthogonal higher-order structure investigations of the WISC–IV (Bodin et al., 2009; Watkins, 2006; Watkins et al., 2006), WAIS–IV (Canivez & Watkins, 2010a, 2010b), RIAS (Dombrowski et al., 2009; Nelson & Canivez, 2011; Nelson et al., 2007), SB–5 (Canivez, 2008), and the WASI and WRIT (Canivez et al., 2009); and, consistent with Jensen's (1998) observations, the present study also found that most of the total and common CAS variance was associated with a general, second-order (g) factor and interpretation of CAS performance at this level is supported, despite Naglieri's (1999) rejection of a hierarchical model and intentions for CAS interpretation to be at the PASS level. However, compared to the WISC–IV, WAIS–IV, SB–5, RIAS, WASI, and WRIT, the CAS subtests had less variance apportioned to the higher-order general factor (g) and greater proportions of variance apportioned to first-order (PASS or [PA]SS) factors. This is consistent with the subtest selection and construction in an attempt to measure PASS dimensions linked to PASS theory (Das, Naglieri, & Kirby, 1994) and neuropsychological theory (Luria, 1966a, 1966b, 1973, 1980, 1982), but which were not specifically linked to a theory of general intelligence. CAS subtests also had lower g-loadings (first unrotated factor structure coefficients from EFA) which allowed for capturing greater portions of first-order variance. Within the 5–7 age group, only the Number Detection subtest had a "good" g–loading and within the 8–17 age group, only the Planned Connections subtest had a "good" g–loading according to Kaufman's (1994) criteria (≥.70). There also ap-

Table 3
Factor Structure Coefficients and Variance Sources for the CAS Standardization Sample Based on the Orthogonalized Higher-Order Factor Model From Four First-Order (PASS) Factors for Ages 8–17 (n = 1,224)

	General		Factor I (Planning)		Factor II (Successive)		Factor III (Simultaneous)		Factor IV (Attention)			
CAS subtest	*b*	$\%S^2$	*b*	$\%S^2$	*b*	$\%S^2$	*b*	$\%S^2$	*b*	$\%S^2$	h^2	u^2
Matching Numbers	.48	23	**.54**	**29**	.02	0	.00	0	.00	0	0.52	0.48
Planned Codes	.35	12	**.45**	**20**	.04	0	.01	0	–.11	1	0.34	0.66
Planned Connections	.58	34	**.45**	**20**	.03	0	.08	1	.03	0	0.55	0.45
Nonverbal Matrices	.57	32	.02	0	.04	0	**.34**	**11**	–.08	1	0.44	0.56
Verbal–Spatial Relations	.61	37	.00	0	.04	0	**.23**	**5**	.14	2	0.45	0.55
Figure Memory	.59	35	–.01	0	–.04	0	**.36**	**13**	.01	0	0.48	0.52
Expressive Attention	.48	23	***.28***	***8***	.02	0	–.02	0	**.24**	**6**	0.36	0.64
Number Detection	.50	25	***.44***	***20***	–.04	0	.00	0	**.14**	**2**	0.47	0.53
Receptive Attention	.48	23	***.58***	***33***	–.04	0	–.04	0	**.08**	**1**	0.57	0.43
Word Series	.48	23	.00	0	**.56**	**31**	–.02	0	–.07	0	0.55	0.45
Sentence Repetition	.61	37	.01	0	**.58**	**33**	–.01	0	.02	0	0.71	0.29
Sentence Questions	.65	43	–.01	0	**.43**	**18**	.05	0	.12	1	0.63	0.37
% Total S^2		28.9		10.8		7.0		2.6		1.2	50.4	49.6
% Common S^2		57.2		21.4		13.9		5.1		2.3	—	—

Note. CAS = Cognitive Assessment System; *b* = subtest factor structure coefficient (loading); $\%S^2$ = percent of variance explained in the subtest; h^2 = communality; u^2 = uniqueness (specific and error variance). Coefficients and variance percents for subtests within their theoretically assigned factor are highlighted in bold. Coefficients and variance percents in bold italic represent cross-loading on alternate factor.

peared to be greater unique variance measured by the CAS (particularly for Nonverbal Matrices, Verbal–Spatial Relations, Figure Memory, and Expressive Attention in the 5–7 age group) when compared to the WISC–IV, WAIS–IV, SB–5, RIAS, WASI, and WRIT (see Tables 1-4). Within the CAS, the Planning and Successive processing scales provided the greatest levels of apportioned variance suggesting greater interpretability while the Simultaneous and Attention processing scales provided the lowest levels of apportioned variance and less interpretability when four factors were extracted. In the three-factor models, the Planning/Attention and Successive processing scales provided the greatest levels of apportioned variance with Simultaneous processing providing substantially less variance.

In contrast, Kranzler and Weng (1995) found greater apportioned variance to the Successive scale after apportioning variance to the higher-order factor in the experimental CAS tasks but the Matrices subtest was among the higher g-loading subtests here in the 8–17 age group within the three-factor model and in Kranzler and Weng as is commonly reported (Jensen, 1998). Given the significant changes to CAS tasks, addition and deletion of tasks/subtests, CFA restriction of subtest loading on only one first-order factor, and the very small sample size in the Kranzler and Weng study; direct comparisons of the present results to those of Kranzler and Weng are difficult at best.

A problem articulated by Kranzler and Keith (1999) was their observation that the CAS Planning and Attention scales were "virtually indistinguishable" (p. 139) in their CFA. In the present study, the Number Detection and Receptive Attention subtests had substantially greater variance apportioned to the CAS Planning factor than the Attention factor after extracting variance due to a general intelligence factor in both the 5–7 and 8–17 age groups (see Tables 1 & 3). In the 8–17 age group, the Expressive Attention had roughly equivalent portions of variance apportioned to Planning and Attention. This complicates CAS interpretation, as these three subtests do not appear to uniquely measure Attention. This relationship between Planning and Attention subtests and factors is why a three-factor solution appears viable in several analyses and illustrated in Table 4.14 in the CAS Interpretive Handbook (Naglieri & Das, 1997b, p. 58) and elsewhere (Kranzler & Keith,

Table 4
Factor Structure Coefficients and Variance Sources for the CAS Standardization Sample Based on the Orthogonalized Higher-Order Factor Model From Three First-Order (PA)SS Factors for Ages 8–17 (n = 1,224)

	General		Factor I (Planning/ Attention)		Factor II (Successive)		Factor III (Simultaneous)			
CAS subtest	b	$\%S^2$	b	$\%S^2$	b	$\%S^2$	b	$\%S^2$	h^2	u^2
Matching Numbers	.46	21	**.55**	**30**	.00	0	–.01	0	0.51	0.49
Planned Codes	.35	12	**.42**	**17**	.00	0	–.01	0	0.29	0.71
Planned Connections	.58	33	**.47**	**22**	.02	0	.05	0	0.55	0.45
Nonverbal Matrices	.62	39	–.02	0	.01	0	**.21**	**5**	0.44	0.56
Verbal–Spatial Relations	.62	39	.04	0	.08	1	**.17**	**3**	0.42	0.58
Figure Memory	.64	41	–.02	0	–.05	0	**.24**	**6**	0.48	0.52
Expressive Attention	.43	19	**.36**	**13**	.07	0	.01	0	0.32	0.68
Number Detection	.47	22	**.50**	**25**	–.02	0	.01	0	0.47	0.53
Receptive Attention	.45	20	**.62**	**38**	–.04	0	–.02	0	0.58	0.42
Word Series	.47	22	–.03	0	**.56**	**31**	–.02	0	0.53	0.47
Sentence Repetition	.59	35	.01	0	**.60**	**36**	–.01	0	0.70	0.30
Sentence Questions	.63	40	.02	0	**.47**	**22**	.05	0	0.62	0.38
% Total S^2		28.5		12.1		7.5		1.2	49.3	50.7
% Common S^2		57.8		24.5		15.2		2.4	—	—

Note. CAS = Cognitive Assessment System; b = subtest factor structure coefficient (loading); $\%S^2$ = percent of variance explained in the subtest; h^2 = communality; u^2 = uniqueness (specific and error variance). Coefficients and variance percents for subtests within their theoretically assigned factor are highlighted in bold.

1999), but there are no composite scores available for this configuration. Planning and Attention were also postulated to have higher correlation than they would with either the Simultaneous or Successive factors based on theory and anatomical relationships between Luria's first and third functional units (Luria, 1966a; Naglieri, 1999). Zero-order Pearson product–moment correlations between the three Planning subtests and the three Attention subtests ranged from .173 to .573 (Mdn_r = .473) for the 5–7 age group and ranged from .269 to .540 (Mdn_r = .421) for the 8–17 age group. The zero-order Pearson product–moment correlations between Planning and Attention scores were .627 (39% shared variance) for the 5–7 age group, and .672 (45% shared variance) for the 8–17 age group. The fact that CAS Attention factor subtests had nonzero cross-loadings suggests that the CFA results of Kranzler and Keith (1999) implying Planning and Attention were indistinguishable may be a result of overestimating factor correlations and subsequent distortion of structural relations due to cross-loading subtests being set to zero (Asparouhov & Muthen, 2009). This may also have affected results reported by Kranzler and Weng (1995).

Without question, the CAS appears to have a complex four-factor structure that does not satisfy simple structure with subtests having variance apportioned to more than one first-order factor and this could have affected the Kranzler and Keith (1999) CFA results. This is not unique to the CAS as other measures have been found to show satisfactory EFA results, but questionable or unsatisfactory CFA results when significant cross-loading of subtests was observed (Marsh et al., 2010). Exploratory structural equation modeling (E-CFA; Asparouhov & Muthen, 2009) was proposed as one solution to address such issues within a CFA approach. The present study illustrates well the apportioned variance of subtests to each of the test levels and summarizes CAS variance at both the higher-order and lower-order levels and provides an expanded examination of EFA results not included in the CAS Interpretive Handbook or extant literature. The present study also illustrates why it is important to consider both EFA and CFA results when investigating the structure of a test. The present results help describe where subtest variance is apportioned and has implications for interpretation of the Full Scale score as well as the PASS factors, but was not conducted to test which model is theoretically better. The present results should help to inform future revisions of the CAS where attempts to modify existing subtests or creating new ones may help to achieve simple structure where subtests are associated with only one of the first-order factors. If Planning and Attention are dimensions that are highly interdependent this may prove difficult to achieve and it may thus be necessary to apply E-CFA methods when examining the theoretical structure as suggested by Asparouhov and Muthen. Certainly greater explication of EFA and CFA results should be provided in the technical manual.

Investigations of the structural validity of tests are important but they are insufficient. Canivez et al. (2009) and others (Carroll, 1997; Kline, 1994; Lubinski & Dawis, 1992) pointed out that factor analytic methods (CFA and EFA) cannot fully answer questions of test score validity or utility. Also, because the latent constructs from CFA are not directly observable, and latent construct scores are difficult to calculate and not readily available, they offer no direct practical clinical application (Oh, Glutting, Watkins, Youngstrom, & McDermott, 2004). Additional methods are also required to assess the relative importance of higher-order versus lower-order interpretations.

Given the greater portions of CAS variance observed at the first-order factor level than the WISC–IV, WAIS–IV, and SB–5, there is perhaps an improved chance for incremental validity of CAS PASS scores over and above the CAS Full Sale score, particularly for the Planning (or Planning/Attention) and Successive processing scales. Establishment of incremental validity (Haynes & Lench, 2003; Hunsley, 2003; Hunsley & Meyer, 2003) is essential when considering interpretation of scores representing different levels of a test. In this way the relative importance of the PASS ([PA]SS) factor scores versus the global Full Scale score in predicting academic achievement may be assessed. If lower-order scores (PASS factors) do not account for meaningful portions of external criteria variance (i.e., academic achievement) after accounting for prediction from the higher-order score (Full Scale), the lower-order scores may be unimportant or simply redundant. Such incremental validity of first-order factor index scores of intelligence tests has yet to achieve sufficient support (Freberg, Vandiver, Watkins, & Canivez, 2008; Glutting, Watkins, Konold, & McDermott, 2006; Glutting, Youngstrom, Ward, Ward, & Hale, 1997; Kahana, Youngstrom, & Glutting, 2002; Ryan, Kreiner, & Burton, 2002; Watkins, Glutting, & Lei, 2007; Youngstrom, Kogos, & Glutting, 1999). Incremental validity is an important feature that must also be examined for the CAS PASS scores.

Validity, however, should not be confused with diagnostic utility (Meehl, 1959; Mullins–Sweatt & Widiger, 2009; Wiggins, 1988) as the latter is concerned with the application of test score interpretation to the individual. Diagnostic use of the CAS PASS scores is another area for further investigation. For example, PASS theory (Das, Naglieri, & Kirby, 1994; Naglieri & Das, 1997b) proposes that children with attention deficit hyperactivity disorder (ADHD) would, as Barkley (2003, 2006) suggests, be more impulsive (and less reflective) in their cognitive processing, which in turn would negatively impact planning processing. Attention difficulties would be expected to negatively affect attention processing. Studies of CAS performance of children with ADHD typically show lowest performance on Planning with concurrent deficits on Attention but normal Simultaneous and Successive processing scores (Crawford, 2002; Naglieri & Das, 1997b; Naglieri, Goldstein, Iseman, & Schwebach, 2003; Naglieri, Salter, & Edwards, 2004; Paolitto, 1999; Pottinger, 2002; Van Luit, Kroesbergen, & Naglieri, 2005). Such group differences studies provide support for the construct validity of the CAS via distinct group differences; however, such support is inadequate for determining the utility of the CAS in individual diagnostic decision-making (Mullins–Sweatt & Widiger, 2009). Distinct group differences are necessary but not sufficient; so if the portions of first-order PASS variance presently observed are sufficient, diagnostic utility of PASS scores may be possible and should be examined. Future research will help determine the extent to which CAS PASS scores possess acceptable incremental validity and diagnostic utility.

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