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Factor Structure of the Differential Ability Scales–Second Edition: Exploratory and Hierarchical Factor Analyses With the Core Subtests

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The present study examined the factor structure of the Differential Ability Scales–Second Edition (DAS–II; Elliott, 2007a) standardization sample using exploratory factor analyses, multiple factor extraction criteria, and hierarchical exploratory factor analyses (Schmid & Leiman, 1957) not included in the *DAS–II Introductory and Technical Handbook* (Elliott, 2007b). Exploratory factor analyses with multiple factor extraction criteria and hierarchical analyses with the Schmid and Leiman (1957) procedure were conducted with the 3 DAS–II standardization samples (Lower Early Years [Ages 2:6–3:5], Upper Early Years [Ages 3:6–6:11], School-Age [Ages 7:0–17:11]). All factor extraction criteria suggested 1 factor despite the author and publisher recommended and promoted 2 (Ages 2:6–3:5) or 3 (Ages 3:6–6:11, Ages 7:0–17:11) factors. Results indicated that most DAS–II subtests were properly associated with the theoretically proposed first-order factors. Hierarchical exploratory analyses with the Schmid and Leiman procedure, however, found that the hierarchical *g* factor accounted for large portions of total and common variance, while the 2 or 3 first-order factors accounted for small portions of total and common variance. It was concluded that the DAS–II provides strong measurement of general intelligence but clinical interpretation should be primarily at that level.

Keywords: DAS–II, exploratory factor analysis, factor extraction criteria, Schmid–Leiman higher-order analysis, structural validity

The Differential Ability Scales–Second Edition (DAS–II; Elliott, 2007a) is an individually administered battery of cognitive tests for children and adolescents ages 2–17 years. The DAS–II is a revised version of the DAS (Elliott, 1990), which originated from the British Ability Scales (Elliott, Murray, & Pearson, 1979), and was developed and standardized for use within the United States. It is divided into three levels: Lower Early Years (ages 2:6 through 3:5); Upper Early Years (3:6 through 6:11); and School Age (7:0 through 17:11). The three levels contain different mixtures of 10 core subtests that combine to yield a higher-order composite score called the General Conceptual Ability score (GCA), thought to measure psychometric *g* (Spearman, 1927), as well as three first-order composite scores called *cluster* scores (Verbal Ability, Nonverbal Reasoning Ability, and Spatial Ability) that are thought to reflect more specific and diverse aptitudes. The DAS–II also provides users with 10 supplemental diagnostic subtests, which can be combined to form three additional first-order cluster scores (School Readiness, Working Memory, and Processing Speed) across the age span. However, these measures do not contribute to the GCA or the three primary cluster scores, although it is suggested (e.g., Dumont, Willis, & Elliott, 2008) that they may provide users with additional information about cognitive strengths and weaknesses. It should be noted that not all of the aforementioned diagnostic clusters are available consistently throughout the entire DAS–II age span. For instance, the School Readiness cluster is only available from ages 4:6 through 8:0, whereas the Working Memory and Processing Speed clusters are available from ages 5:0 through 17:11. According to the *Introductory and Technical Handbook* (Elliott, 2007b), this is the result of an inability to adequately measure certain latent constructs (e.g., working memory and processing speed) at young ages.

In addition to deleting the Basic Number Skills, Spelling, and Word Reading subtests (eliminating achievement measures from the previous version of the DAS); combining Block Building and Pattern Construction into one subtest; creating and adding Phonological Processing, Recall of Digits Backward, Recall of Sequential Order, and Rapid Naming subtests; and increasing item coverage and range; the DAS–II theoretical foundation was updated. Although the *Introductory and Technical Handbook* indicates that the DAS–II development was not driven by a single theory of cognitive ability, the content and structure of the DAS–II was heavily influenced by the Cattell–Horn–Carroll model of cognitive abilities (CHC; Carroll, 1993, 2003; Cattell & Horn, 1978; Horn, 1991; Schneider & McGrew, 2012). The CHC or three-stratum model suggests that cognitive abilities are organized hierarchically according to their level of generality. Narrow abilities (Stratum I) serve as the foundation for the model, followed by broad ability factors (Stratum II), and at the apex (Stratum III), rests a general ability dimension. Consistent with other recently published intelligence tests, such as the Wechsler Intelligence Scale for Children–Fifth Edition (WISC–V; Wechsler, 2014), Wechsler Adult Intelli-

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gence Scale–Fourth Edition (WAIS–IV; Wechsler, 2008), Wechsler Preschool and Primary Scale of Intelligence–Fourth Edition (WPPSI–IV; Wechsler, 2012), Kaufman Assessment Battery for Children–Second Edition (KABC–II; Kaufman & Kaufman, 2004), Woodcock–Johnson IV Tests of Cognitive Abilities (WJ IV COG; Schrank, McGrew, & Mather, 2014), and the Stanford–Binet Intelligence Scales–Fifth Edition (SB5; Roid, 2003a); major elements of the CHC model were used to guide investigations of the structural validity of the DAS–II in the *Introductory and Technical Handbook* and serve as the primary method for interpretation of the scores provided by the DAS–II.

In terms of clinical interpretation, the *Introductory and Technical Handbook* suggests that users should interpret the scores obtained from the DAS–II in a stepwise fashion beginning with the GCA and then proceeding to more specific measures (e.g., clusters and subtests). Despite this recommendation, it is suggested that the profile of strengths and weaknesses generated at the cluster and subtest levels is of more value than the information provided by the GCA, especially in cases where considerable variability across the cluster scores is observed. According to Elliott (2007b), "the most satisfactory description of a child's abilities is nearly always at the level of profile analysis" (p. 87). However, such prescriptive statements are rarely justified in applied practice and require adherence to high standards of empirical evidence (Marley & Levin, 2011). Although profile analysis and the primary interpretation of part scores on intelligence tests such as the DAS–II are popular in clinical practice, empirical support for the validity of these practices has repeatedly been found wanting (e.g., Macmann & Barnett, 1997; McDermott, Fantuzzo, & Glutting, 1990; McDermott, Fantuzzo, Glutting, Watkins, & Baggaley, 1992; Glutting, Watkins, & Youngstrom, 2003; Miciak, Fletcher, Stuebing, Vaughn, & Tolar, 2014; Watkins, 2000a; Watkins, Glutting, & Lei, 2007).

Confirmatory factor analytic (CFA) support for the DAS–II hierarchical structure was reported in the DAS–II *Introductory and Technical Handbook* (Elliott, 2007b), and Figures 8.1, 8.2, 8.3, and 8.4 illustrate the standardized validation models for the seven core and diagnostic subtests (ages 2:6–3:5) featuring two first-order factors, 11 core and diagnostic subtests (ages 4:0–5:11) with five first-order factors, 14 core and diagnostic subtests (ages 6:0–12:11) with seven first-order factors, and 12 core and diagnostic subtests featuring six first-order factors, respectively. In these models, several first-order factors not available in the actual DAS–II were specified (e.g., Auditory Processing, Visual–Verbal Memory, and Verbal Short-Term Memory). In addition, the Auditory Processing and Visual–Verbal Memory factors in the final validation models for ages 6–17 were each produced from a single indicator and reflect an underidentified dimension. Although the inclusion of single indicator variables is possible in CFA, variables assessed by a single measure should not be interpreted as factors due to the fact that they do not possess shared variance across multiple observed measures (Brown, 2015). It should also be noted that an independent CFA examination of the DAS–II core and diagnostic subtest structure by Keith, Low, Reynolds, Patel, and Ridley (2010) supported a six-factor hierarchical model that corresponded closely with CHC theory with general intelligence at the apex. However, their final validation model required the specification of a cross-loading for the Verbal Comprehension measure on crystallized ability and fluid reasoning factors. Although the latent structure of the core subtests was examined more directly in the *Introductory and Technical Handbook*, across ages 2–17, standardized solutions for these analyses were not provided. Inspection of the goodness-of-fit results in Table 8.4 indicated that a three-factor hierarchical model provided the most optimal solution for the core measures across the age span with fairly robust improvements in fit when compared to competing one- and two-factor hierarchical models. In general, CFA analyses have supported a hierarchical model with general intelligence at the apex and three first-order factors for the core subtests.

Unfortunately, despite the substantive structural and theoretical revisions to the DAS–II, Elliott (2007b) relied exclusively upon CFA to examine the structural validity. It was argued in the DAS–II *Introductory and Technical Handbook* (Elliott, 2007b) that the use of CFA procedures was justified due to the fact that development of the DAS–II was based on a previously established measurement instrument. However, overreliance on CFA procedures for examining the internal structure of intelligence tests can result in the retention of poorly defined factors, overfactoring of internal structure, and has been criticized within the technical literature (Canivez, 2013; Canivez & Watkins, 2010a, 2010b; Frazier & Youngstrom, 2007). Some methodologists have recommended that exploratory factor analytic (EFA) procedures should be used to compliment CFA procedures, especially when evaluating a new test or theory (Haig, 2005; Gerbing & Hamilton, 1996; Schmitt, 2011). EFA and CFA are considered to be complimentary procedures and Gorsuch (1983) noted that they provide answers to different empirical questions and that when the results from these procedures are in agreement, greater confidence can be placed in the internal structure of a test. Carroll (1998) argued that "CFA should derive its initial hypotheses from EFA results, rather than starting from scratch or from *a priori* hypotheses . . . [and] CFA analyses should be done to check my EFA analyses" (p. 8). Brown (2015) also noted that "in addition to a compelling substantive justification, CFA model specification is usually supported by prior (but less restrictive) exploratory analyses (i.e., EFA) that have established the appropriate number of factors, and pattern of indicator–factor relationships" (p. 141). Without the presentation of EFA procedures with the DAS–II norming sample data, clinicians are not able to consider the convergence or divergence of CFA and EFA results for the DAS–II or to consider alternate models EFA might suggest. This information is important for determining the relative importance of various scores for clinical interpretation.

Independent investigations of the factor structures of intelligence tests that have been completed using EFA procedures have challenged many of the latent structures reported in corresponding technical manuals obtained via CFA procedures alone. Both Dombrowski (2013) and Dombrowski and Watkins (2013), using data from the WJ III COG normative sample, obtained markedly different results for the WJ III COG than the CFA results reported in its technical manual (McGrew & Woodcock, 2001). These results supported a robust manifestation of general intelligence (*g*) and the additional presence of five to six first-order factors across the age span. No evidence for nine factors in the WJ III COG was found. Two investigations of the SB5 (Canivez, 2008; DiStefano & Dombrowski, 2006) indicated that the SB5 measured one fundamental dimension (*g*) with no support for the five first-order factors suggested in the SB5 technical manual (Roid, 2003b). Similarly, numerous investigations of various iterations of the Wechsler Scales (e.g., Canivez & Watkins, 2010a, 2010b, 2016; Canivez, Watkins, & Dombrowski, 2015; Dombrowski, Canivez, Watkins, & Beaujean, 2015; Watkins, 2006; Watkins & Beaujean, 2014) suggest that most of the reliable variance in those measures is associated with general intelligence and that interpretation should focus primarily on the Full Scale IQ (FSIQ) composite. As a consequence, it has been argued (e.g., Canivez, 2013; Glutting, Watkins, Konold, & McDermott, 2006; McGill, 2015) that the limited unique variance captured by first-order factors may be responsible for poor incremental validity of those scores in accounting for meaningful portions of achievement variance beyond that provided by the FSIQ in many contemporary intelligence test measures.

Also missing from the DAS–II *Introductory and Technical Handbook* were proportions of variance accounted for by the higher-order *g* factor and the proposed first-order factors, subtest *g* loadings, subtest specificity estimates, and incremental predictive validity estimates for the factors and subtest scores. Thus, clinicians do not have the necessary information for determining the relative importance of factor and subtest scores relative to the GCA score. If the factor or subtest scores fail to capture meaningful portions of true score variance they will likely be of limited clinical utility. The omission of incremental predictive validity results is especially troubling given that users are encouraged to interpret the DAS–II beyond the GCA level and an incremental validity investigation of the previous iteration of the DAS (Youngstrom, Kogos, & Glutting, 1999) found that interpretation beyond the GCA was not supported.

According to Carroll (2003), all cognitive measures are composed of reliable variance that is attributable to a higher-order general factor, reliable variance that is attributable to first-order group factors, and error variance. Because of this, Carroll argued that variance from the higher-order factor must be extracted first to residualize the lower order factors, leaving them orthogonal to the higher-order dimension. Thus, variability associated with a higher-order factor is accounted for before interpreting variability associated with lower-order factors, resulting in variance being apportioned correctly to higher-order and lower-order dimensions. To accomplish this task, Carroll (1993, 1995) recommended second-order factor analysis of first-order factor correlations followed by a Schmid–Leiman transformation (SL; Schmid & Leiman, 1957). The Schmid–Leiman technique allows for the orthogonalization of higher-order variance from lower-order factors. According to Carroll (1995):

> 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 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).

Although Keith et al. (2010) provided the results of residualized subtest factor loadings in their DAS–II CFA analyses, the clinical utility of these results are limited due to the fact they were derived from a hypothesized first-order latent structure that deviates significantly from the structure suggested in the *Introductory and Technical Handbook* (Elliott, 2007b).

As noted by Reise (2012), the SL procedure is an exploratory bifactor method (approximate bifactor) and the most dominant method used to date, although there are two other, less examined, exploratory bifactor methods: target bifactor rotation (Reise, Moore, & Maydeu–Oliveres, 2011) and analytic bifactor rotation (Jennrich & Bentler, 2011). Because the SL has been the dominant method used and because of its application with Wechsler scales (Canivez & Watkins, 2010a, 2010b, 2016; Canivez et al., 2015; Golay & Lecerf, 2011; Watkins, 2006), SB5 (Roid, 2003a) (Canivez, 2008), Wechsler Abbreviated Scales of Intelligence (WASI; Psychological Corporation, 1999) and Wide Range Intelligence Test (WRIT; Glutting, Adams, & Sheslow, 2000; Canivez, Konold, Collins, & Wilson, 2009), Reynolds Intellectual Assessment Scales (RIAS; Reynolds & Kamphaus, 2003; Dombrowski, Watkins, & Brogan, 2009; Nelson & Canivez, 2012; Nelson, Canivez, Lindstrom, & Hatt, 2007), the Cognitive Assessment System (CAS; Naglieri & Das, 1997; Canivez, 2011), and the Woodcock–Johnson Psychoeducational Battery III (WJ III; Woodcock, McGrew, & Mather, 2001; McGrew & Woodcock, 2001; Dombrowski, 2013, 2014a, 2014b; Dombrowski & Watkins, 2013), use of the SL procedure allows comparison of DAS–II results to these other studies. Until now, the variance decomposition procedures described by Carroll (1995) have yet to be applied to current or previous versions of the DAS.

Purpose of the Current Study

To address this gap in the literature the present study used subtest correlation matrices from the DAS–II normative sample published in the DAS–II *Introductory and Technical Handbook* (Elliott, 2007b) to independently examine the factor structure using EFA procedures that allow the data to "speak for itself" (Carroll, 1985, p. 26) in order to examine the following research questions: a) Using multiple extraction criteria, how many factors should be extracted and retained for the DAS–II normative sample across the three test levels (i.e., Lower Early Years, Upper Early Years, and School Age)? and b) When forced extracting the number of first-order factors suggested by the test publisher across the different DAS–II test levels and applying the Schmid and Leiman (1957) procedure, what portions of variance are attributed to general intelligence (*g*) and the first-order broad ability clusters? Use of the SL procedure allows comparison to similarly obtained results from studies of other intelligence tests previously reported. It is believed that results from the current study provide practitioners with important information regarding the correct interpretation of the DAS–II within clinical practice. If the interpretive procedures recommended in the DAS–II *Introductory and Technical Handbook* are utilized by clinicians, it is imperative that they know how variability is apportioned across first-and second-order dimensions.

Method

Participants

Participants were members of the DAS–II standardization sample and included a total of 3,480 individuals ranging in age from 2–17 years. Demographic characteristics are provided in detail in the DAS–II *Introductory and Technical Handbook* (Elliott, 2007b). The standardization sample was obtained using stratified proportional sampling across demographic variables of age, sex, race/ethnicity, parent educational level, and geographic region. Examination of the tables in the *Introductory and Technical Handbook* revealed a close correspondence to the October 2002 U.S. census estimates across the stratification variables.

Instrument

The DAS–II is an individual test of general intelligence for ages 2–17 that is a decedent of the British Ability Scales (Elliott et al., 1979). Consistent with other contemporary measures of intellectual ability (e.g., Wechsler Scales), the DAS–II measures general intelligence through the administration of numerous subtests, each of which is a unique indicator of psychometric *g*. According to Elliott (2007b), psychometric *g* is defined as "the ability of an individual to perform complex mental processing that involves conceptualization and the transformation of information" (p. 17). The DAS–II uses different combinations of the 10 core subtests to produce the GCA at different points in the age span. Whereas four subtests combine to form the GCA for ages 2:6 through 3:5, six core subtests are needed from ages 3:6 through 17:11. The core subtests also combine to form three primary cognitive clusters at the first-order level, each composed of two subtests. Although the Verbal Ability and Nonverbal Reasoning Ability clusters are provided throughout the age span, an additional Spatial Ability cluster is only available from ages 3:6 through 17:11. As previously discussed, different combinations of supplemental diagnostic subtests are provided throughout the age range, which can be combined to yield additional first-order clusters (e.g., Working Memory, Processing Speed, and School Readiness) however, these measures are not utilized to calculate the higher-order GCA composite or its lower-order cognitive clusters. Additionally, it should also be noted that the diagnostic measures cannot be used to substitute for core subtests at any point in the age range.

Procedure and Analyses

DAS–II core subtest correlation matrices for the three standardization sample subgroups (2:6–3:5, 3:6–6:11, 7:0–17:11) were obtained from Tables 8.1, 8.2, and 8.3, respectively, in the *Introductory and Technical Handbook* (Elliott, 2007b) to conduct hierarchical exploratory factor analyses. Analyses in the current study were limited to the core subtests in order to provide a consistent examination of the DAS–II structure across the age span. Multiple criteria (Gorsuch, 1983) were examined to determine how many latent factors were suggested and included eigenvalues >1 (Kaiser, 1960), the scree test (Cattell, 1966), standard error of scree (SE_{Scree}; Zoski & Jurs, 1996), Horn's parallel analysis (HPA; Horn, 1965), and minimum average partials (MAP; Velicer, 1976). HPA and MAP have been found to be the most accurate empirical criteria with scree sometimes a useful adjunct according to simulation studies (Velicer, Eaton, & Fava, 2000; Zwick & Velicer, 1986). Random data and resulting eigenvalues for HPA were produced using the Monte Carlo PCA for Parallel Analysis computer program (Watkins, 2000b) with 100 replications to provide stable eigenvalue estimates. HPA performance, however, has the tendency to underfactor in the presence of a strong general factor (Crawford et al., 2010). The scree test is a subjective criterion so the SE_{Scree} as programmed by Watkins (2007) was used because it was reportedly the most accurate objective scree method (Nasser, Benson, & Wisenbaker, 2002).

Principal axis exploratory factor analyses (Fabrigar, Wegener, MacCallum, & Strahan, 1999) were used to analyze the DAS–II standardization samples core subtest correlation matrices using SPSS 21 for Macintosh OSX. Promax (oblique) rotation ($k = 4$; Gorsuch, 1983) was applied to extracted factors. As per Child (2006) salient factor pattern coefficients were defined as those ≥.30. Following Carroll's (1995) guidance, the Schmid and Leiman (1957) procedure was used to orthogonalize first-order factors by removing all variance associated with the second-order dimension using as programmed in the MacOrtho program (Watkins, 2004). This transforms "an oblique factor analysis solution containing a hierarchy of higher-order factors into an orthogonal solution which 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). Thus, Thompson (2004) noted "This allows the researcher to determine what, if any, variance is unique to a given level of analysis or perspective" (p. 74) and he further argued (like Carroll, 1995), "Whenever oblique rotation is performed, higher-order factors are implied and should be derived and interpreted. The Schmid–Leiman solution is particularly useful in such cases" (p. 81). The Schmid–Leiman procedure also allows one to "interpret the second-order factors in terms of the measured variables, rather than as a manifestation of the factors of the measured variables" (p. 74).

The Schmid–Leiman (SL) orthogonalization procedure produces an approximate exploratory bifactor (Holzinger & Swineford, 1937) solution (Canivez, in press; Reise, 2012), but may be constrained by proportionality (Yung, Thissen, & McLeod, 1999), and may be problematic with nonzero cross-loadings (Reise, 2012). Reise presented the analytic bifactor (Jennrich & Bentler, 2011) and target bifactor (Reise et al., 2011) as more recent alternative exploratory bifactor methods that do not include proportionality constraints. The present application of the SL orthogonalization procedure was selected because there are numerous studies of its application with other intelligence tests (Canivez, 2008, 2011; Canivez et al., 2009; Canivez & Watkins, 2010a, 2010b, 2016; Canivez et al., 2015; Dombrowski, 2013, 2014a, 2014b; Dombrowski & Watkins, 2013; Dombrowski et al., 2009; Golay & Lecerf, 2011; Nelson & Canivez, 2012; Nelson et al., 2007; Watkins, 2006), which facilitates comparison of DAS–II results to these other studies. This method is referred to and labeled SL bifactor (Reise, 2012).

Omega-hierarchical and omega-subscale (Reise, 2012) were estimated as model-based reliability estimates of the latent factors (Gignac & Watkins, 2013) from the higher-order solutions to estimate the unique portions of true score variance apportioned to the different latent factors. Chen, Hayes, Carver, Laurenceau, and Zhang (2012) stressed that "for multidimensional constructs, the alpha coefficient is complexly determined, and McDonald's omega-hierarchical (ω_h; McDonald, 1999) provides a better estimate for the composite score and thus should be used" (p. 228). This is also an inherent problem with other internal consistency estimates such as split-half or KR-20. ω_h is the model-based reliability estimate for the hierarchical general intelligence factor independent of the variance of group factors. Omega-subscale (ω_s)

is the model-based reliability estimate of a group factor with all other group and general factor(s) removed (Reise, 2012). Omega estimates (ω_h and ω_s) may be obtained from EFA SL bifactor solutions and were produced using the Omega program (Watkins, 2013), which is based on the tutorial by Brunner, Nagy, and Wilhelm (2012) and the work of Zinbarg, Revelle, Yovel, and Li (2005) and Zinbarg, Yovel, Revelle, and McDonald (2006). Omega coefficients should at a minimum exceed .50, but .75 is preferred (Reise, 2012; Reise, Bonifay, & Haviland, 2013).

Results

Factor Extraction Criteria Comparisons

Table 1 summarizes results from the six factor extraction criteria for the three DAS–II age groups (2:6–3:5, 3:6–6:11, 7:0–17:11). All criteria suggested the extraction of one factor except the author/publisher recommended theory, which suggested two factors for the 2:6–3:5-year-old age group and three factors for the 3:6–6:11 and 7:0–17:11-year-old age groups. It has been argued that it is better to overextract than underextract (Gorsuch, 1997; Wood, Tataryn, & Gorsuch, 1996) so two factors were extracted for the 2:6–3:5-year-old age group, while three and two factors were extracted for the for the 3:6–6:11 and 7:0–17:11-year-old age groups to examine subtest associations based on the author/publisher suggested structure and to allow examination of the performance of smaller factors.

Exploratory Factor Analyses: Ages 2:6–3:5

Table 2 presents results from extracting two DAS–II factors with promax ($k = 4$) rotation. The g loadings (factor structure coefficients from first unrotated factor) ranged from .565 (Pattern Construction) to .836 (Naming Vocabulary) and all were within the fair to good range based on Kaufman's (1994) criteria (≥.70 = good, .50–.69 = fair, <.50 = poor). A broad general intelligence factor appears for all subtests. The Picture Similarities subtest demonstrated cross-loading on both the Verbal and Nonverbal factors and had a slightly higher factor pattern coefficient on the Verbal factor; demonstrating a complex structure. The high first-order factor correlation of .782 implies a higher-order (general intelligence) factor structure requiring explication (Carroll, 1993, 1995, 1998; Gorsuch, 1983, 1988, 2003; Thompson, 2004).

Table 1
Number of Factors Suggested for Extraction Across Six Different Criteria by Age Group

Extraction criterion	DAS–II age group 2:6–3:5	3:6–6:11	7:0–17:11
Eigenvalue >1	1	1	1
Scree test	1	1	1
Standard Error of Scree (SEScree)	1	1	1
Horn's Parallel Analysis (HPA)	1	1	1
Minimum Average Partials (MAP)	1	1	1
Theory proposed	2	3	3

Table 2
DAS–II Exploratory Factor Analysis Results for Standardization Sample Ages 2:6–3:5 (N = 352) Two Oblique Factor Solution

	General	Verbal		Nonverbal		
	S	*P*	*S*	*P*	*S*	h^2
DAS–II subtest						
Verbal comprehension	.793	**.629**	.784	.199	.690	.630
Naming vocabulary	.836	**.890**	.858	−.041	.655	.737
Picture similarities	.624	***.363***	.598	***.301***	.585	.393
Pattern construction	.565	.039	.502	**.591**	.622	.388
Eigenvalue		2.46		.66		
% Variance		50.88		2.80		

Note. S = Structure coefficient, *P* = Pattern coefficient, h^2 = Communality. General structure coefficients are based on the first unrotated factor coefficients (g-loadings). Salient factor pattern coefficients are presented in bold (pattern coefficient ≥.30). The promax based correlation between the two factors was .782.

Hierarchical EFA: SL Bifactor Model Ages 2:6–3:5

The first-order oblique solution was transformed with the Schmid–Leiman (SL) orthogonalization procedure. Results for the Schmid and Leiman orthogonalization of the higher-order factor analysis of two DAS–II first-order factors (see Table 2) are presented in Table 3. All subtests were properly associated (higher residual variance) with their theoretically proposed factor after removing g variance except Picture Similarities, which had residual variance approximately evenly split between the Verbal factor and its theoretically related Nonverbal factor. The hierarchical g factor accounted for 43.9% of the total variance and 81.7% of the common variance.

The general factor also accounted for between 31.0% (Pattern Construction) and 56.4% (Naming Vocabulary) of individual subtest variability. At the first-order level, the Verbal factor accounted for an additional 7.2% of the total variance and 13.4% of the common variance and the Nonverbal factor accounted for an additional 2.6% of the total variance and 14.9% of the common variance. The general and group factors combined to measure 53.7% of the variance in DAS–II scores resulting in 46.3% unique variance (combination of specific & error variance).

Omega-hierarchical and omega-subscale coefficients were estimated based on the SL results in Table 3. To examine latent factor reliability of the theoretically based scores Picture Similarities was included in the Nonverbal factor for estimation of ω_s. The ω_h coefficient for general intelligence (.729) was high and sufficient for scale interpretation; however, the ω_s coefficients for the two specific DAS–II group factors (Verbal and Nonverbal) were considerably lower (.151 and .064, respectively). Thus, the two specific DAS–II group factors likely possess too little unique true score variance for confident clinical interpretation (Reise, 2012; Reise et al., 2013).

Exploratory Factor Analyses: Ages 3:6–6:11

Table 4 presents results from extracting three DAS–II factors with promax ($k = 4$) rotation. The three-factor EFA solution appeared to be the most reasonable solution as extraction of two-factors resulted in the Matrices subtest failing to have a salient factor pattern coefficient on any factor. The g loadings (factor

Table 3
Sources of Variance in the Differential Abilities Scale-Second Edition (DAS–II) Normative Sample Ages 2:6–3:5 (N = 352) According to an Exploratory Bifactor Model (Orthogonalized Higher-Order Factor Model)

	General		Verbal		Nonverbal			
	b	S^2	b	S^2	b	S^2	h^2	u^2
DAS–II subtest								
Verbal comprehension	.732	.536	**.294**	**.086**	.093	.009	.631	.369
Naming vocabulary	.751	.564	**.416**	**.173**	−.019	.000	.737	.263
Picture similarities	.587	.345	*.170*	*.029*	**.141**	**.020**	.393	.607
Pattern construction	.557	.310	.018	.000	**.277**	**.077**	.387	.613
Total variance		.439		.072		.026	.537	.463
Common variance		.817		.134		.149		
	ω_h = .729		ω_s = .151		ω_s = .064			

Note. b = loading of subtest on factor, S^2 = variance explained, h^2 = communality, u^2 = uniqueness, ω_h = Omega hierarchical, ω_s = Omega subscale. Bold type indicates coefficients and variance estimates consistent with the theoretically proposed factor. Italic type indicates coefficients and variance estimate larger for alternate factor.

structure coefficients from first unrotated factor) ranged from .568 (Matrices) to .724 (Naming Vocabulary) and all were within the fair to good range based on Kaufman's (1994) criteria (≥.70 = good, .50–.69 = fair, <.50 = poor). A broad general intelligence factor appears for all subtests. All DAS–II subtests were saliently and properly associated with their theoretical factor and none of the subtests demonstrated cross-loading on multiple factors, demonstrating desirable simple structure. The moderate to high factor correlations presented in Table 4 (.664 to .705) imply a higher-order (general intelligence) factor structure requiring explication (Carroll, 1993, 1995, 1998; Gorsuch, 1983, 1988, 2003; Thompson, 2004).

Hierarchical EFA: SL Bifactor Model Ages 3:6–6:11

As the three-factor EFA solution appeared to be the most reasonable solution; that first-order oblique solution was transformed with the Schmid–Leiman (SL) orthogonalization procedure. Results for the Schmid and Leiman orthogonalization of the higher-order factor analysis of three first-order DAS–II factors (see Table 4) are presented in Table 5. All subtests were properly associated (higher residual variance) with their theoretically proposed factor after removing *g* variance. The hierarchical *g* factor accounted for 31.5% of the total variance and 66.3% of the common variance.

The general factor also accounted for between 27.5% (Pattern Construction) and 37.3% (Verbal Comprehension) of individual subtest variability. At the first-order level, the Verbal factor accounted for an additional 7.0% of the total variance and 14.8% of the common variance, the Nonverbal factor accounted for an additional 3.8% of the total variance and 8.1% of the common variance, and the Spatial factor accounted for an additional 5.2% of the total variance and 10.9% of the common variance. The

Table 4
DAS–II Exploratory Factor Analysis Results for Standardization Sample Ages 3:6–6:11 (N = 928) Three Oblique Factor Solution

	General	Verbal		Spatial		Nonverbal		
	S	P	S	P	S	P	S	h^2
DAS–II subtest								
Verbal comprehension	.664	**.547**	.679	.066	.531	.130	.540	.477
Naming vocabulary	.724	**.921**	.839	−.040	.530	−.082	.501	.710
Picture similarities	.617	.180	.531	.187	.547	**.335**	.587	.395
Matrices	.568	−.042	.432	−.052	.460	**.768**	.703	.497
Pattern construction	.719	.054	.561	**.687**	.765	.058	.578	.589
Copying	.605	−.042	.442	**.777**	.699	−.070	.450	.493
Eigenvalue		3.02		.78		.74		
% Variance		42.54		5.99		4.16		
Promax-based factor correlations		Verbal		Spatial		Nonverbal		
Verbal		—						
Spatial		.683		—				
Nonverbal		.664		.705		—		

Note. S = Structure Coefficient, P = Pattern Coefficient, h^2 = Communality. General structure coefficients are based on the first unrotated factor coefficients (*g*-loadings). Salient factor pattern coefficients are presented in bold (pattern coefficient ≥.30).

Table 5
Sources of Variance in the Differential Abilities Scale–Second Edition (DAS−II) Normative Sample Ages 3:6–6:11 (N = 928) According to an Exploratory Bifactor Model (Orthogonalized Higher-Order Factor Model)

	General		Verbal		Spatial		Nonverbal			
	b	S^2	b	S^2	b	S^2	b	S^2	h^2	u^2
DAS−II subtest										
Verbal comprehension	.611	.373	**.326**	**.106**	.035	.001	.073	.005	.486	.514
Naming vocabulary	.532	.283	**.550**	**.303**	−.021	.000	−.046	.002	.588	.412
Picture similarities	.600	.360	.107	.011	.098	.010	**.188**	**.035**	.416	.584
Matrices	.532	.283	−.025	.001	−.027	.001	**.430**	**.185**	.469	.531
Pattern construction	.524	.275	.032	.001	**.362**	**.131**	.033	.001	.408	.592
Copying	.564	.318	−.025	.001	**.409**	**.167**	−.039	.002	.488	.512
Total variance		.315		.070		.052		.038	.476	.524
Common variance		.663		.148		.109		.081		
	ω_h = .697		ω_s = .255		ω_s = .206		ω_s = .136			

Note. b = loading of subtest on factor, S^2 = variance explained, h^2 = communality, u^2 = uniqueness, ω_h = Omega hierarchical, ω_s = Omega subscale. Bold type indicates coefficients and variance estimates consistent with the theoretically proposed factor.

general and group factors combined to measure 47.6% of the variance in DAS–II scores resulting in 52.4% unique variance (combination of specific & error variance).

Omega-hierarchical and omega-subscale coefficients were estimated based on the SL results in Table 5. The ω_h coefficient for general intelligence (.697) was moderate and likely sufficient for scale interpretation; however, the ω_s coefficients for the three specific DAS–II group factors (Verbal, Spatial, Nonverbal) were considerably lower (.136–.255). Thus, the three specific DAS–II group factors likely possess too little unique true score variance for confident clinical interpretation (Reise, 2012; Reise et al., 2013).

Exploratory Factor Analyses: Ages 7:0–17:11

Three first-order factors. Table 6 presents results from extracting three DAS–II factors with promax (k = 4) rotation. The g loadings (factor structure coefficients from first unrotated factor) ranged from .668 (Recall of Designs) to .809 (Sequential and Quantitative Reasoning) and all were within the fair to good range based on Kaufman's (1994) criteria (≥.70 = good, .50–.69 = fair, <.50 = poor). A broad general intelligence factor appears for all subtests. All DAS–II subtests were saliently and properly associated with their theoretical factor and none of the subtests demonstrated cross-loading on multiple factors, demonstrating desirable simple structure. The moderate to high factor correlations presented in Table 6 (.689 to .812) imply a higher-order (general intelligence) factor structure requiring explication (Carroll, 1993, 1998; Gorsuch, 1983, 1988, 2003; Thompson, 2004).

Two first-order factors. Table 7 presents results from extracting two DAS–II factors with promax (k = 4) rotation. The g loadings (factor structure coefficients from first unrotated factor)

Table 6
DAS−II Exploratory Factor Analysis Results for Standardization Sample Ages 7−17 (N = 2,200) Three Oblique Factor Solution

	General	Verbal		Spatial		Nonverbal		
	S	P	S	P	S	P	S	h^2
DAS−II subtest								
Pattern construction	.722	−.006	.555	**.578**	.751	.218	.683	.579
Matrices	.741	.074	.606	.210	.682	**.519**	.744	.574
Recall of designs	.668	.042	.520	**.741**	.735	−.044	.589	.540
Word definitions	.737	**.779**	.810	.000	.570	.041	.623	.657
Verbal similarities	.745	**.776**	.814	.038	.586	.015	.626	.664
Sequential and quantitative reasoning	.809	.045	.659	.017	.703	**.807**	.854	.731
Eigenvalue		3.64		.72		.52		
% Variance		54.51		5.69		2.21		
Promax-based factor correlations		Verbal		Spatial		Nonverbal		
Verbal		—						
Spatial		.689		—				
Nonverbal		.747		.812		—		

Note. S = Structure Coefficient, P = Pattern Coefficient, h^2 = Communality. General structure coefficients are based on the first unrotated factor coefficients (g-loadings). Salient factor pattern coefficients are presented in bold (pattern coefficient ≥.30).

Table 7
DAS−II Exploratory Factor Analysis Results for Standardization Sample Ages 7−17 (N = 2,200) Two Oblique Factor Solution

	General	Nonverbal		Verbal		
	S	*P*	*S*	*P*	*S*	h^2
DAS−II subtest						
Pattern construction	.726	**.798**	.764	−.045	.558	.585
Matrices	.743	**.662**	.751	.118	.618	.571
Recall of designs	.655	**.667**	.679	.016	.520	.462
Word definitions	.741	.016	.620	**.799**	.811	.659
Verbal similarities	.748	.043	.631	**.779**	.811	.658
Sequential and quantitative reasoning	.789	**.655**	.788	.177	.672	.635
Eigenvalue		3.64		.72		
% Variance		54.02		5.47		

Note. *S* = Structure Coefficient, *P* = Pattern Coefficient, h^2 = Communality. General structure coefficients are based on the first unrotated factor coefficients (*g*-loadings). Salient factor pattern coefficients are presented in bold (pattern coefficient ≥.30). The promax based correlation between the two factors was .755.

ranged from .655 (Recall of Designs) to .789 (Sequential and Quantitative Reasoning) and all were within the fair to good range based on Kaufman's (1994) criteria (≥.70 = good, .50–.69 = fair, <.50 = poor). A broad general intelligence factor appears for all subtests. All DAS–II subtests were saliently and properly associated with theoretically oriented factors and none of the subtests produced cross-loading on multiple factors, demonstrating desirable simple structure. The high first-order factor correlation of .755 implies a higher-order (general intelligence) factor structure requiring explication (Carroll, 1993, 1998; Gorsuch, 1983, 1988, 2003; Thompson, 2004).

Hierarchical EFA: SL Bifactor Models Ages 7:0–17:11

Three first-order factors. The three first-order oblique EFA factor solution was transformed with the Schmid–Leiman (SL) orthogonalization procedure. Results for the Schmid and Leiman orthogonalization of the higher-order factor analysis of three first-order DAS–II factors (see Table 6) are presented in Table 8. All subtests were properly associated (higher residual variance) with their theoretically proposed factor after removing *g* variance. The hierarchical *g* factor accounted for 48.6% of the total variance and 78.6% of the common variance.

The general factor also accounted for between 40.3% (Recall of Designs) and 63.2% (Sequential and Quantitative Reasoning) of individual subtest variability. At the first-order level, the Verbal factor accounted for an additional 7.4% of the total variance and 12.0% of the common variance, the Spatial factor accounted for an additional 3.9% of the total variance and 6.2% of the common variance, and the Nonverbal factor accounted for an additional 2.0% of the total variance and 3.2% of the common variance. The general and group factors combined to measure 61.9% of the variance in DAS–II scores resulting in 38.1% unique variance (combination of specific and error variance).

Omega-hierarchical and omega-subscale coefficients were estimated based on the SL results in Table 8. The ω_h coefficient for general intelligence (.819) was high and likely sufficient for scale interpretation; however, the ω_s coefficients for the three specific DAS–II group factors (Verbal, Spatial, Nonverbal) were consid-

Table 8
Sources of Variance in the Differential Abilities Scale–Second Edition (DAS−II) Normative Sample Ages 7−17 (N = 2,200) According to an Exploratory Bifactor Model (Orthogonalized Higher-Order Factor Model)

	General		Verbal		Spatial		Nonverbal			
	b	S^2	*b*	S^2	*b*	S^2	*b*	S^2	h^2	u^2
DAS−II subtest										
Pattern construction	.697	.486	−.004	.000	**.289**	**.084**	.076	.006	.575	.425
Matrices	.720	.518	.045	.002	.105	.011	**.181**	**.033**	.564	.436
Recall of designs	.635	.403	.025	.001	**.370**	**.137**	−.015	.000	.541	.459
Word definitions	.659	.434	**.471**	**.222**	.000	.000	.014	.000	.656	.344
Verbal similarities	.665	.442	**.469**	**.220**	.019	.000	.005	.000	.663	.337
Sequential and quantitative reasoning	.795	.632	.027	.001	.009	.000	**.282**	**.080**	.712	.288
Total variance		.486		.074		.039		.020	.619	.381
Common variance		.786		.120		.062		.032		
	ω_h = .819		ω_s = .266		ω_s = .140		ω_s = .066			

Note. *b* = loading of subtest on factor, S^2 = variance explained, h^2 = communality, u^2 = uniqueness, ω_h = Omega hierarchical, ω_s = Omega subscale. Bold type indicates coefficients and variance estimates consistent with the theoretically proposed factor.

erably lower (.266–.066). Thus, the three specific DAS–II group factors likely possess too little unique true score variance for confident clinical interpretation (Reise, 2012; Reise et al., 2013).

Two first-order factors. The first-order oblique solution was transformed with the Schmid–Leiman (SL) orthogonalization procedure. Results for the Schmid and Leiman orthogonalization of the higher-order factor analysis of two DAS–II first-order factors (see Table 7) are presented in Table 9. All subtests were properly associated (higher residual variance) with their theoretically related factor after removing *g* variance. The hierarchical *g* factor accounted for 46.1% of the total variance and 77.6% of the common variance.

The general factor also accounted for between 35.2% (Recall of Designs) and 52.1% (Sequential and Quantitative Reasoning) of individual subtest variability. At the first-order level, the Nonverbal factor accounted for an additional 8.0% of the total variance and 13.5% of the common variance and the Verbal factor accounted for an additional 5.3% of the total variance and 8.9% of the common variance. The general and group factors combined to measure 59.4% of the variance in DAS–II scores resulting in 40.6% unique variance (combination of specific and error variance).

Omega-hierarchical and omega-subscale coefficients were estimated based on the SL results in Table 9. The ω_h coefficient for general intelligence (.769) was high and sufficient for scale interpretation; however, the ω_s coefficients for the two specific DAS–II group factors (Nonverbal and Verbal) were considerably lower (.179 and .185, respectively). Thus, the two specific DAS–II group factors likely possess too little unique true score variance for confident clinical interpretation (Reise, 2012; Reise et al., 2013).

Discussion

The present study was conducted to independently examine the DAS–II (Elliott, 2007a) factor structure using EFA procedures not included in the DAS–II *Introductory and Technical Handbook* and that included the Schmid and Leiman (1957) procedure to determine what portions of DAS–II subtest variance are attributed to general intelligence (*g*) and the first-order broad ability clusters. Understanding the convergence or divergence of results from EFA and CFA as well as understanding how variability is apportioned across first- and second-order DAS–II dimensions is important for clinicians to decide if sufficient variability is present among the DAS–II scores they interpret.

While all of the DAS–II subtests (except Picture Similarities for ages 2:6–3:5) were properly associated with their theoretically proposed latent first-order factor (Elliott, 2007b) for all age groups, examination of variance apportions to the hierarchical *g* factor and the group factors found substantially greater total and common variance associated with the hierarchical *g* factor. This is a result observed in numerous other studies examining the latent factor structure of intelligence or cognitive ability tests using both EFA with SL bifactor procedure and CFA procedures using either higher-order or bifactor models (Bodin, Pardini, Burns, & Stevens, 2009; Canivez, 2008, 2011, 2014; Canivez et al., 2009; Canivez & Watkins, 2010a, 2010b, 2016; Canivez et al., 2015; DiStefano & Dombrowski, 2006; Dombrowski, 2013, 2014a, 2014b; Dombrowski & Watkins, 2013; Dombrowski et al., 2009; Gignac & Watkins, 2013; Nelson & Canivez, 2012; Nelson et al., 2007; Nelson, Canivez, & Watkins, 2013; Watkins, 2006, 2010; Watkins & Beaujean, 2014; Watkins, Canivez, James, James, & Good, 2013; Watkins, Wilson, Kotz, Carbone, & Babula, 2006). These results are also consonant with the literature regarding the importance of general intelligence (Deary, 2013; Gottfredson, 2008; Jensen, 1998; Lubinski, 2000; Ree, Carretta, & Green, 2003). As such, the principal interpretation of DAS–II should be of the GCA, the estimate of *g*. The dominance of *g* variance captured by the DAS–II subtests is a likely reason that methods to determine how many factors to extract and retain such as HPA and MAP indicated only one factor (Crawford et al., 2010).

Given the small portions of total and common variance uniquely attributed to the DAS–II Verbal, Nonverbal, and Spatial factors and the low portions of true score variance in these factors (as estimated by ω_s coefficients), there appears to be little variance apart from *g* in these factor scores to warrant clinical interpretation (Reise, 2012; Reise et al., 2013), or if interpreted, done with extreme caution. Relatedly, the confidence intervals provided for

Table 9

Sources of Variance in the Differential Abilities Scale–Second Edition (DAS−II) Normative Sample Ages 7−17 (N = 2,200) According to an Exploratory Bifactor Model (Orthogonalized Higher-Order Factor Model)

	General		Nonverbal		Verbal			
	b	S^2	b	S^2	b	S^2	h^2	u^2
DAS−II subtest								
Pattern construction	.654	.428	**.396**	**.157**	−.022	.000	.585	.415
Matrices	.677	.458	**.328**	**.108**	.059	.003	.569	.431
Recall of designs	.593	.352	**.331**	**.110**	.008	.000	.461	.539
Word definitions	.707	.500	.008	.000	**.396**	**.157**	.657	.343
Verbal similarities	.713	.508	.021	.000	**.386**	**.149**	.658	.342
Sequential and quantitative reasoning	.722	.521	**.325**	**.106**	.088	.008	.635	.365
Total Variance		.461		.080		.053	.594	.406
Common Variance		.776		.135		.089		
	ω_h = .769		ω_s = .179		ω_s = .185			

Note. b = loading of subtest on factor, S^2 = variance explained, h^2 = communality, u^2 = uniqueness, ω_h = Omega hierarchical, ω_s = Omega subscale. Bold type indicates coefficients and variance estimates consistent with the theoretically proposed factor.

the DAS–II factor scores are considerably smaller (due to conflated general intelligence variance) than they would be if only the unique true score variance of the group factor scores was used with latent factor scores.

Validity and interpretation of latent factors cannot be solely determined by factor analytic approaches because there may be many plausible models that adequately fit. Ultimately a test's structure must be assessed against external criteria to answer questions of validity or diagnostic utility/efficiency (Canivez & Gaboury, 2016; Canivez et al., 2009; Carroll, 2012; Kline, 1998; Lubinski & Dawis, 1992). In the case of intelligence tests such as the DAS–II, academic achievement is one such external criterion of interest. Due to the hierarchical nature of tests of intelligence the use of incremental validity is particularly necessary to determine the relative contribution of higher-order versus lower-order intelligence constructs in accounting for variability in academic achievement (Haynes & Lench, 2003; Hunsley, 2003; Hunsley & Meyer, 2003).

The limited portions of unique variance captured by first-order factors in intelligence tests may be responsible for poor incremental validity of such scores in accounting for meaningful portions of achievement variance beyond that provided by the omnibus composite IQ score in many contemporary intelligence tests (e.g., Canivez, 2013; Canivez, Watkins, James, Good, & James, 2014; Freberg, Vandiver, Watkins, & Canivez, 2008; Glutting et al., 2006; Glutting, Youngstrom, Ward, Ward, & Hale, 1997; Kahana, Youngstrom, & Glutting, 2002; McGill & Busse, 2015). Assessment of incremental validity of the DAS factor scores (Youngstrom et al., 1999) found that interpretation beyond the GCA as predictors of achievement was not supported. While incremental validity of DAS–II factor scores above and beyond the GCA has not yet been reported, but should be investigated, it is hard to imagine these group factors would provide useful incremental information when predicting performance in academic achievement or relations with other external criteria.

Another problem for DAS–II interpretation is the recommended practice of identification of factor based cognitive strengths and weaknesses (ipsative comparisons) because analyses of DAS–II factor score differences at the observed score level conflate *g* variance and specific group factor (Verbal, Nonverbal, Spatial) variance. The same is true of analyses of subtest based processing strengths and weaknesses. Because it is not possible to disaggregate these sources of variance for individuals there is no way to know how much of the variance in performance is due to the hierarchical general factor, specific group factor, or the narrow subtest ability. Add to this the general problems identified in ipsative score comparisons (McDermott et al., 1990, 1992; McDermott & Glutting, 1997) and it is easy to see why such practices have been eschewed.

Limitations

A limitation of the present study is that it is based on EFA methods and procedures that may produce results different from those of CFA-based bifactor methods. Differences may also be present if exploratory bifactor rotation methods (i.e., target bifactor rotation [Reise et al., 2011] or analytic bifactor rotation [Jennrich & Bentler, 2011] were used (Dombrowski et al., 2015). As such it is important that further examinations of the latent factor structure of the DAS–II use CFA methods and procedures that include comparison of rival bifactor models (see Canivez in press; Reise, 2012) to the higher-order models provided in the DAS–II *Introductory and Technical Handbook* (Elliott, 2007b). Bifactor-based CFA models are absent from the DAS–II *Introductory and Technical Handbook* and may provide better fit and representation of DAS–II structure. Additionally, it is possible that results of the present analyses may not apply to populations quite different from the normative sample such as those with extremely high intellectual abilities who may exhibit important cognitive profiles (Robertson, Smeets, Lubinski, & Benbow, 2010).

Conclusion

The present study provides clinicians with important information substantially qualifying interpretive recommendations of the DAS–II (Elliott, 2007b). As "the ultimate responsibility for appropriate test use and interpretation lies predominantly with the test user" (AERA, APA, & NCME, 2014, p. 141), clinicians using the DAS–II in clinical evaluations must seriously consider the present information to make informed decisions about which DAS–II scores have satisfactory reliability, validity, and utility. Clinical decision making with scores or score comparisons that lack sufficient reliability and validity have implications for individual clients and the ethical clinician must "know what their tests can do and act accordingly" (Weiner, 1989, p. 829).

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