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The Learning Behaviors Scale: National standardization in Trinidad and Tobago

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ABSTRACT

This study reports on the national standardization and validation of the Learning Behaviors Scale (LBS) for use in Trinidad and Tobago. The LBS is a teacher rating scale centering on observable behaviors relevant to identifying childhood approaches to classroom learning. Teachers observed a stratified sample of 900 students across the islands' schools nationwide. Exploratory and confirmatory factor analyses yielded two reliable dimensions: Competence Motivation and Strategy/Flexibility. Scales were developed using IRT and Bayesian scoring methods and scores were used in HLM models to investigate variance accounted for in measures of academic achievement and psychopathology.

KEYWORDS

learning behaviors; item response theory; Trinidad and Tobago

The driving force behind most educational intervention is the desire to improve academic performance and emotional well-being. This necessarily requires understanding the foundational precursors of academic success and of behavioral adjustment. Conventional theories often link intellectual and cognitive capacities to successful academic performance (Deary, Strand, Smith, & Fernandes, 2007; Naglieri & Bornstein, 2003; Neisser et al., 1996; Watkins, Lei, & Canivez, 2007). Although traditional intelligence measures might be useful for decisions regarding general responsiveness to instruction (Braden & Shaw, 2009) or among exceptionally able students (Robertson, Smeets, Lubinski, & Benbow, 2010), they provide limited information nuanced enough to allow for assessment of specific behaviors and abilities that might prove useful in service of designing interventions for students struggling in the classroom (Burns et al., 2016; Elliott & Resing, 2015). Rather, development of effective interventions is alternatively linked to promotion of differential learning behaviors (Hyson, 2008; McDermott, 1984; Scarr, 1981; Schaefer & McDermott, 1999).

Learning behaviors (also called approaches-to-learning and learning-to-learn behaviors), are both positively *related* to desirable academic outcomes (Alexander, Entwisle, & Dauber, 1993; Hinshaw, 1992; Sattler, 1992;

Schaefer & McDermott, 1999; Stott, 1985; Yen, Konold, & McDermott, 2004) and demonstrably *teachable* through behavioral modification and modeling (Barnett, Bauer, Ehrhardt, Lentz, & Stollar, 1996; Engelmann, Granzin, & Severson, 1979; Stott, 1978, 1981; Stott & Albin, 1975; Weinberg, 1979), making them highly relevant for educators looking to design effective interventions. To the extent that positive learning behaviors can be taught, it may follow that interventions targeting them can have a direct impact on academic performance. Additionally, positive learning behaviors have been found to act as protective factors against behavioral maladjustment (McDermott, 1999; McDermott et al., 2001; Rikoon, McDermott, & Fantuzzo, 2012) and to reduce the risk of specific learning disabilities (McDermott, Goldberg, Watkins, Stanley, & Glutting, 2006).

Measurement of learning behaviors

Learning behaviors are defined as observable patterns of behavior exhibited by students as they respond to learning situations and react to academic tasks. They include indicators of effective effort and attitudes toward learning, strategic problem solving, flexibility, attentional persistence, reflectivity, and responses to novelty and error (Yen et al., 2004). The term is sometimes

interchanged with “learning style” in the research literature, its meaning thus pertaining to *how* a child learns (rather than *how well*, which is referred to as academic achievement) and to the child’s preferences when engaged in learning processes and when interacting with a learning environment (Yamazaki, 2005).

The learning behaviors construct evolved from early research on cognitive styles, temperament, reflectivity, and motivation. As these are based in concepts that are traditionally inferred from unobservable mediating psychological processes and thus fundamentally hypothetical, there are a number of challenges encountered in measuring them (Kagan, 1996; McClelland, 1973; Stott & Albin, 1975; Thomas & Chess, 1977; Zigler & Trickett, 1978). Seminal concepts of cognitive styles typically require inferences about students’ mediating psychological processes. Alternatively, the Learning Behaviors Scale (LBS; McDermott, Green, Francis, & Stott, 1999), rather than considering learning behaviors within the framework of these inferred concepts, bases assessment entirely on unobtrusive classroom observation by teachers. The rating scale content is comprised of entirely observable behavioral aspects, without reference to the student’s intentions, thoughts, or feelings, and the teacher interprets each behavior in regard to the student in an environment as compared to other children. The scale contains 29 items describing both positive and negative observable classroom behaviors that influence academic and social outcomes. Standardized and normed in the United States, it yields four subscale scores, representing behaviors related to expectation of success, positive classroom engagement, attentiveness, problem-solving strategies, and flexible approaches to tasks.

Based on large and representative U.S. samples, LBS assessments have been found to augment substantially the information afforded by measures of general intelligence in explaining the variability of students’ academic achievement, social adjustment, school attendance, and susceptibility to learning disabilities (McDermott, Rikoon, & Fantuzzo, 2014, 2016; McDermott et al., 2006; Yen et al., 2004). Moreover, LBS scores are not meaningfully related to demographic variables such as student gender or ethnicity (Schaefer & McDermott, 1999).

Cross-cultural studies on learning behaviors

Although the evidence on the LBS is generally supportive, it is understood that the properties of scores on a measure are derived according to the characteristics of a particular population and the conditions of a given study (Goodwin & Goodwin, 1999). There is a common misconception that norms and scoring routines can be readily

generalized across other populations without evaluation of normative data and evidence collected from the unique population of interest (Hu & Oakland, 1991; Oakland & Hu, 1993; van Widenfelt, Treffers, de Beurs, Sieblink, & Koudijs, 2005). Psychometric properties, including reliability and validity at the item and scale level, must be reassessed in every target population, especially for psychological and behavioral instruments where differences in cultural context might be expected to influence linguistic nuance and interpretation of the constructs in question (Hambleton & Patsula, 1998; K. H. Rubin, 1998; United Nations Children’s Fund, 2013).

It is clearly established that both the attributes of different learning styles and the frequencies with which they are subscribed to vary among cultures (Barmeyer, 2004; DeVita, 2001; Hofstede & McCrae, 2004; Pratt, 1991). Learning styles or behaviors are influenced by both the learner’s previous experiences and by the characteristics of the environment in which the learning takes place, and as such are context-dependent. As cultural expectations and philosophies generally set the norms for attitudes and motivations, it follows that both perception and expression of behaviors would differ across cultures.

In commercial management and higher education contexts, the Learning Style Inventory (LSI; Kolb, 1985), as well as more qualitative methods, are often used in cross-cultural research to assess differences in learning styles across cultures (Yamazaki, 2005). For instance, Barmeyer (2004) used the LSI to examine differences in learning style typologies with samples of French, German, and Quebecois business students, finding that German students emphasized “thinking” styles while French and Quebecois students emphasized “feeling” styles. However, Platsidou and Metallidou (2009) found that in the case of Greek students, the LSI resolved into an alternative two-factor structure that bore less resemblance to the original. Bowden, Abhayawansa, and Manzin (2015) examined approaches to learning in tertiary school students using the Study Process Questionnaire (SPQ; Biggs, 1987b), a self-report questionnaire originally designed in Australia that yields three approaches to learning as composites of motive and strategy: surface (minimal), deep (intrinsic interest) and achieving (driven by self-esteem or ego). Bowden et al. used the assessment to explore differentiation between learning approaches adopted by students from Confucian heritages, other Asian countries, and local English-speaking students at an Australian university. Their findings confirmed those of other literature indicating that Confucian heritage students and other Asian students are more likely to adopt deep learning approaches, or a mixed approach of deep and surface learning simultaneously. It has also been shown that,

while the SPQ's motive/strategy model is valid for use in Australia and Hong Kong, it is less so for Filipino and Dutch students (Stes, De Maeyer, Van Petegem, 2013; Watkins & Astilla, 1982). There has been much research into the posited dimensionality of the SPQ and its two-factor version, as it seems to be a cross-culturally sensitive instrument (Snelgrove & Slater, 2003; Stes et al., 2013).

There has, however, thus far been limited research into the cross-cultural implications of learning behaviors among school-age children, and developing countries have few nationally adapted and validated rating measures (Mpfu, Oakland, Ntinda, Seeco, & Maree, 2014). Wong, Lin, and Watkins (1996) applied the Learning Process Questionnaire (LPQ; Biggs, 1987a), a version of the SPQ for primary and secondary school students, to 10 samples of school children across Nigeria, Zimbabwe, Malaysia, Hong Kong, Beijing, and Canada, and found general support for the dimensions but with some variation in the position of the achieving dimension, where, like with the SPQ, it was sometimes associated with the surface approach or deeper approach rather than resolving as its own dimension. Previous cross-cultural LBS studies include Canivez and Beran (2011) in Canada, replicating the U.S. factor structure; Hamlet, Schaefer, Herrick, and Rai (2015) in Nepal; and Canivez et al. (2010) in China. Additionally, appropriate language versions of the LBS' preschool counterpart, the Preschool Learning Behaviors Scale (PLBS), have been standardized and validated with populations in Peru (Hahn, Schaefer, Merino, & Worrell, 2009) and in Greece (Penderi, Petrogiannis, & McDermott, 2014).

A small sample study ($N = 61$) of the LBS with Vincentian children in the Caribbean found that behaviors related to learning, attention, motivation, and anxiety accounted for more than 30% of variation in academic performance (Durbrow, Schaefer, & Jimerson, 2000). The follow-up study ($N = 65$) found that problem behaviors, particularly those related to attention and anxiety, were more predictive of academic achievement than cognitive ability scores, with LBS predictive power being about equal to that of cognitive ability scores (Durbrow, Schaefer, & Jimerson, 2001). These were small samples of a unique population, but suggested a focus on reducing attention problems might improve academic achievement for that group. Trinidad and Tobago, although one of the most developed nations in the Caribbean, is still classified as a developing economy by the International Monetary Fund (2015) and faces difficult economic circumstances, high crime rates, and distinctive parental disciplinary practices that compel the identification and mitigation of impediments to student learning as a means to improve academic, health, and social outcomes (Cappa & Khan, 2011; Greenberg &

Agozino, 2012; Krishnakumar, Narine, Roopnarine, & Logie, 2014; Roopnarine, Krishnakumar, Barine, Logie, & Lape, 2014; Williams, 2013).

The current study

This study examines the normative development, dimensionality, and validation of the LBS for national application in Trinidad and Tobago. The administration of the LBS was part of a larger national initiative undertaken by the Ministry of Education to identify children at risk for academic and behavior problems (Watkins, Hall, & Worrell, 2014). It builds upon a preliminary study conducted by Chomat-Mooney (2006) by taking advantage of modern methodological practices to account for inherent nonnormal item response distributions, as well as multiple imputation to avoid bias associated with missing data, IRT scoring methods to enhance reliability and generalizability, and multilevel modeling to provide more accurate estimates of external validity.

Method

Setting

Located about seven miles off the north coast of Venezuela in the Caribbean, Trinidad and Tobago is a twin-island nation that had been earlier colonized by several different European countries before gaining its independence from Great Britain in 1962. It remains an English-speaking, Commonwealth nation and maintains a free and obligatory education system for all children aged 5 to 16 years. Respect for authority and obedience are highly valued in children, with corporeal punishment accepted as a norm within the home, and parent-child interactions commonly characterized by discipline and order (Barrow, 2008; Cappa & Khan, 2011; Gopaul-McNicol, 1999; Roopnarine et al., 2014).

Participants

The study included a probability sample of 900 children aged 4–15 years ($M = 8$, $SD = 2$), drawn from 75 government and assisted schools across the islands. The sample was representative of current ethnic distributions in the islands, with 50.3% female, 39.9% of African descent, 38.3% of East Indian descent, and 21.7% of mixed descent (Central Intelligence Agency, 2014). Schools were stratified by region, with regions with a lower proportion of school-aged children in the population being less represented than those with higher proportions. The sample was further blocked by grade level and gender to create a representative national normative sample ($n = 700$) and a supplemental validity

oversample ($n = 200$). The normative sample was used for scale calibration and the oversample was included in structural and validity analyses.

Instruments

Classroom learning behavior

The Learning Behaviors Scale is an objective teacher response scale designed to measure differential learning behaviors of school-aged children. It was originally standardized and validated in the United States, with a nationally representative sample of children aged 5–17 years stratified by age, gender, and grade level ($N = 1,500$; McDermott, 1999). The instrument contains 29 items with responses on a 3-point Likert scale (2 = *Most often applies*, 1 = *Sometimes applies*, 0 = *Does not apply*) for teachers to describe the frequency with which a student manifests a given behavior. Teachers complete the form after at least 50 days of observation, to ensure that they have adequate familiarity with the child. For U.S. application, the measure yields a total score as well as four subscores assessing distinct dimensions of learning behaviors, including Competence Motivation, Attitude Toward Learning, Attention-Persistence, and Strategy/Flexibility.

Factorial invariance has been reported in a variety of contexts (Canivez & Beran, 2011; Canivez, Willenborg, & Kearney, 2006; Worrell, Vandiver, & Watkins, 2001), and the measure has been validated for Head Start alumni (Rikoon et al., 2012) and academically talented students (Worrell & Schaefer, 2004). Interobserver agreement was established by Buchanan, McDermott, and Schaefer (1998) and Worrell et al. (2001), internal consistency (coefficient $\alpha = .75$ to $.85$) by McDermott (1999), and short-term temporal stability by Canivez, Willenborg, and Kearney (2006). The incremental validity for augmenting standardized intelligence test scores with LBS scores in order to predict academic achievement was investigated by Lakebrink (2014) for special education populations and established by Yen et al. (2004) for the U.S. normative sample. Convergent and divergent validity evidence for LBS have been demonstrated with measures of intellectual functioning using the Differential Ability Scales (Elliott, 1990), academic enablers using the Academic Competence Evaluation Scales (Smith, 2015), classroom behavior using the Adjustment Scales for Children and Adolescents (ASCA; McDermott, Stott, & Marston, 1993), and academic performance using the Basic Achievement Skills Individual Screener (The Psychological Corporation, 1983).

Classroom social-emotional adjustment

The Adjustment Scales for Children and Adolescents (ASCA) is a teacher rating scale that enables a teacher to

describe a student's classroom behaviors using 156 binary behavioral indicators (McDermott et al., 1993). The instrument is designed to investigate phenological types of problem behaviors and, based on a Trinidad and Tobago normative sample, yields scores on two reliable broad-band behavioral dimensions (coefficient $\alpha > .70$): Overactivity (i.e., externalizing) and Underactivity (i.e., internalizing) problems (McDermott et al., 2015). Originally developed, normed, and validated in the United States, these same two dimensions also emerged in analyses of several other populations, including Hispanic/Latino, Native American, and Canadian (Canivez & Beran, 2009; Canivez & Bohan, 2006; Canivez & Sprouls, 2005, 2010; McDermott, 1993). Substantial evidence of convergent and divergent validity, factorial validity, and internal consistency has been provided by McDermott, Steinberg, and Angelo (2005).

In addition to identifying broad-band phenological syndromes, the ASCA analyzes problem behaviors within 29 specific classroom situational contexts. Teachers observe students over a two-month period and endorse the behaviors associated with each context. Contexts encompass a wide range of situations relevant to the classroom, from interaction with peers to playing fairly to coping with new learning tasks. Each context contains between three and eight different behavioral indicators, including one item with a positive behavioral description intended to reduce the response bias associated with measures composed only of negative descriptions (LeBoeuf, Fantuzzo, & Lopez, 2010). The Trinidad and Tobago normative sample yields scores on three reliable situational context dimensions ($\alpha \geq .75$); that is to say, Peer, Learning, and Teacher Context Problems (McDermott, et al., 2016). These are the same three dimensions found in the original U.S. standardization sample (McDermott et al., 2005).

Classroom clinical behavior

The Disruptive Behavior Disorders Rating Scale (DBDRS; Pelham Evans, Gnagy, & Greenslade, 1992; Pelham Greenslade, Gnagy, & Milich, 1992) draws its items from symptoms described in the *Diagnostic and Statistical Manual of Mental Disorders* (3rd edition, revised [DSM-III-R]); disruptive behavior categories of attention deficit hyperactivity disorder, oppositional-defiant disorder, and conduct disorder. The symptoms listed for these categories in the DSM-III-R correspond well to more recent versions of the DSM, though diagnostic criteria vary (Pelham, Fabiano, & Masseti, 2005). Teachers record the frequency of each symptom exhibited by a student using a 4-point Likert scale (*Not at all*, *Just a little*, *Pretty much*, and *Very much*). The scale yields scores for Inattention, Oppositional/Defiant, and Impulsivity/

Overactivity, and has proven suitable for detecting behavioral and pharmacological effects (Pelham et al., 2005). Internal consistency ($\alpha = .91$ to $.96$) and predictive and discriminant validity have been established with various populations, including clinical cohorts, males in regular classrooms (Pelham, Gnagy et al., 1992), and special education settings (Pelham, Evans et al., 1992). For the current sample, coefficient alpha ranges from $.87$ (Impulsivity/Overactivity) to $.91$ (Inattention).

Academic achievement

Oral Reading Fluency (ORF) is a curriculum-based assessment that measures a child's reading fluency based on the number of words accurately read aloud in one minute from grade-level text (Hasbrouck & Tindal, 1992). Two ORF passages were selected from local grade-level reading texts for each test occasion (fall, winter, and spring) and determined to be of appropriate difficulty via the Flesh-Kincaid readability measure. Passages were administered in counterbalanced order to control for order effects. Scores reflect the number of words read correctly, averaged over two passages, and are collected from fall ($M = 59.1$, $SE = 38.2$), winter ($M = 65.3$, $SE = 39.2$), and spring ($M = 59.1$, $SE = 38.2$) totals. The average correlation between the two passages for the study sample was $.85$ and the average correlation of scores across time (fall vs. winter vs. spring) was $.92$.

ORF has been nationally normed in the United States (Hasbrouck & Tindal, 2006) and has demonstrated validity as a screening tool and progress monitoring measure of student reading proficiency (Fuchs, Fuchs, Hosp, & Jenkins, 2001; Good, Simmons, & Kame'enui, 2001), as well as identifying students at later risk for reading problems (Good, Simmons, Kame'enui, Kaminski, & Wallin, 2002). Fuchs et al. (2001) found evidence for a steeper growth curve for fluency in the primary grades, with negative acceleration in later years. Convergent validity evidence for ORF has been substantiated with other curriculum-based measures and state reading assessments (Deno, Fuchs, Marston, & Shin, 2001; Stage & Jacobsen, 2001; Wood, 2006). Research has also supported the predictive validity and clinical utility of ORF scores (Hart et al., 2013; Petscher & Kim, 2011).

Home social-emotional behavior

The Adjustment Scales for Children and Adolescents-Home Edition (ASCA-H; Watkins & McDermott, 2002) is a parent rating scale containing 202 items relevant to 28 situations associated with behavior in the home. The ASCA-H is similar to the ASCA, but the behavioral indicators represent behaviors in the home rather than the classroom, and children are observed by the parent

rather than the teacher. Parents observe children over a two-month period, and endorse any behavioral descriptions that reflect the child's typical behavior (Watkins & McDermott, 2002). Preliminary evidence of structural validity and scale reliability was substantiated by Coffey (2006), finding four first-order factors (Aggressive-Oppositional, Attention-Seeking Impulsive, Detached, and Diffident) with internal reliability coefficients ranging from $.65$ to $.92$ in a pilot study of 426 U.S. children. Another pilot study with 314 U.S. children found a stable three-factor structure (Unsocialized, Avoidant, and Restless-Impulsive) and established convergent and divergent validity with other parent rating scales (Mordell, 2001). Although the scale is still in development, three factors (ADH, Conduct Problems, and Overactivity) were derived from the current Trinidad and Tobago standardization sample, with internal consistency reliability coefficients ranging from $.74$ – $.82$.

Procedure

Data for the study were collected over the course of one academic year as part of a consultation project agreement between research teams based at a university in the United States and the Trinidad and Tobago Ministry of Education (Watkins et al., 2014). Guidance and Special Education Officers from the Ministry of Education received training from the consulting team and were assigned to gather data from the educational divisions in which they already worked. They, as well as teachers and parents, were paid an honorarium for participating.

Exploratory analysis

There were 799 children for whom all LBS items had teacher reports, with an additional 101 cases with incomplete responses for some items. Missing values for the latter cases were considered to be missing at random, and were imputed using a Markov Chain Monte Carlo (MCMC) multiple imputation method as recommended by D. B. Rubin (1987) and Schafer (1997). The full sample of 900 was then randomly partitioned into an exploratory factor analytic subsample ($n = 500$) and confirmatory subsample ($n = 400$). Inasmuch as item-level data were effectively ordinal rather than continuous, two-stage maximum-likelihood estimation was applied to produce a smoothed polychoric correlation matrix (Olsson, 1979). The matrix was submitted to minimum average partialing (MAP; Velicer, 1976) to suggest the optimal number of retained factors. Iterated common factoring with squared multiple correlations on the principal diagonal of the correlation matrix was conducted, as rotated toward simple structure via varimax and promax

criteria. Solution criteria included approximate simple structure as reflected by a maximized hyperplane count (Yates, 1987) and coverage of items, at least four salient items (loadings $\geq .40$) per factor, reliable factors (i.e., $\alpha \geq .70$), and theoretical plausibility, parsimony, and concordance with leading research (Fabrigar, Wegener, MacCallum, & Strahan, 1999).

Confirmatory analysis

Salient marker items from the exploratory solution were submitted to maximum-likelihood estimation under the Satorra–Bentler scaled difference chi-square for non-normal data (Satorra & Bentler, 2001). An alternative model consisted of the structure found with the U.S. standardization sample. Acceptable fit was indicated by a root mean squared error of approximation (RMSEA) $\leq .08$ and comparative fit index (CFI) $\geq .90$ (Marsh, Liem, Martin, Morin, & Nagengast, 2011).

Scaling

The IRT graded response model for ordinal data was used for scaling. Scores were computed through expected a posteriori (EAP) Bayesian estimation. The normative sample parameters were applied to produce score estimates for the oversample ($n = 200$). Scaled scores were subjected to a linear conversion to center at $M = 50$ and $SD = 10$ for purpose of interpretation. Reliability was assessed for dimensions using Cronbach's α , and examination of the relationships between estimated subscale information functions (the reciprocal of squared measurement error) and measurement error.

External validity

Product–moment correlations were computed to determine the magnitude and direction of relationships between scores on each LBS subscale and external criterion variable. As data were nested within teachers, hierarchical linear modeling (HLM) was applied, where each LBS subscale served as a group-mean centered predictor in a two-level conditional HLM model, indicating the percentage of between-children within-teacher variance accounted for by respective LBS subscales.

Results

MCMC imputation for missing data was effective, with a relative efficiency mean of .9982 and a lowest relative efficiency value of .9949.

Dimensionality

MAP for the 29 items suggested that three factors might be extracted. Thus, 1- through 5-factor models were

tested against the stated criteria, where the 2-factor, promax-rotated ($k = 2$) model emerged as the optimal solution. Models extracting more than two factors contained underidentified and unreliable dimensions, and the 1-factor model compressed the 2-factor model into an uninterpretable composite factor. Waller's (2001) goodness-of-fit index = .98 and root mean squared residual = .07 for the 2-factor model. Five items failed to load saliently on either factor, and the remaining 24 items were retained. Based on item content and the patterns of descending loadings, the scales were named Competence Motivation (14 items) and Strategy/Flexibility (10 items). Item descriptions and factor loadings for each scale are shown in Table 1. These scales have an interfactor correlation of .44.

The two-factor structure was assessed for the confirmatory subsample, as represented by 24 items. The Satorra–Bentler $\chi^2(274) = 677.06$, CFI = .90, and RMSEA = .061 (90% CI = .055/.066) provided evidence of adequate model fit. The U.S. structure was inadequate (CFI = .83, RMSEA = .078 with CI of .074–.082), indicating the misfit of the island data to the U.S. model.

Scaling and reliability

The graded response model threshold parameter for Competence Motivation exhibited an average of -1.09 ($SD = 0.51$) with M slope of 1.07 ($SD = 0.12$); Strategy/Flexibility showed M threshold and slope at -1.69 ($SD = 0.81$) and 0.87 ($SD = 0.16$), respectively. Maximum information for Competence Motivation is 5.42 at $\theta = -.3$ and for Strategy-Flexibility is 2.88 at $\theta = -1.1$.

Expected a posteriori (EAP; Thissen, Pommerich, Billeaud, & Williams, 1995) scaled scores were produced for students, with the normative sample $M = 50$ and $SD = 10$. Figure 1 illustrates the overlap of measurement error and total test information for the dimensions. The figure highlights the practical score range for making discriminations among students, which is approximately from ~ 2.5 SDs below the population mean to ~ 1.3 SDs above the population mean for Competence Motivation and from ~ 3 SDs below the population mean to ~ 1 SD above the population mean for Strategy/Flexibility. Scale scores were internally consistent with Competence Motivation yielding an α coefficient of .89 and Strategy/Flexibility an α of .80.

External validity

Table 2 displays the concurrent relationships between LBS scores and independent criterion measures. All correlations are in the expected direction, with LBS scores exhibiting negative correlations with measures of maladjusted behavior and moderately positive

Table 1. Dimensional structure and properties of the Learning Behaviors Scale.

Item description ^a	Scale pattern loadings ^b		Communality	Item/scale r^c	Item/scale polyserial r^d
	I	II			
Scale I: Competence motivation (coefficient $\alpha = .89^e$)					
Lively interest in learning activities	.76	-.01	.57	.61	.71
Cooperates sensibly in class activities	.74	-.01	.54	.59	.68
Accepts new tasks, no fear	.74	-.05	.52	.58	.70
Takes refuge in dullness/incompetence	.72	.28	.78	.71	.84
Hesitant about giving answer	.71	-.05	.48	.56	.69
Gives up easily	.70	.28	.75	.71	.82
Don't-care attitude toward success	.64	.37	.76	.69	.81
Little desire to please teacher	.64	.21	.57	.60	.72
Reluctant to tackle new tasks	.63	.21	.56	.63	.74
Responds in manner showing attention	.63	-.06	.36	.47	.59
Too unenergetic for interest/effort	.59	.11	.42	.43	.63
Delays answer, wait for hint	.48	.21	.36	.48	.63
Says task too hard without effort	.47	.33	.47	.55	.68
Sticks to task with minor distractions	.41	.05	.19	.34	.48
Scale II: Strategy/ Flexibility (coefficient $\alpha = .80^e$)					
Invents silly ways to do tasks	.05	.78	.64	.58	.73
Hostile when frustrated or corrected	.09	.67	.52	.50	.69
Enterprising ideas that don't work	-.22	.64	.33	.36	.55
Performs own not accepted way	.05	.59	.38	.46	.64
Peculiar/inflexible procedures	.11	.58	.40	.46	.63
Charms others to do work	.12	.57	.40	.41	.60
Fidgets, squirms, leaves seat	.25	.56	.49	.53	.72
Headaches and pains to avoid learning	.22	.56	.47	.44	.64
Doesn't work well in bad mood	.07	.53	.32	.44	.64
Distracted easily or seeks distraction	.38	.48	.54	.51	.75

^a Descriptions are abbreviated for convenient presentation.

^b Values are promaxian pattern loadings at $k = 2$, where hyperplane count is maximized. Salient pattern loadings ($\geq .40$) are italicized. $N = 500$ comprising the random exploratory analysis subsample.

^c Each correlation reflects the relationship between an item and the sum of the other items composing a given scale, where item distributions were standardized to unit-normal form.

^d Values are correlations between each item and the continuous sum of all items comprising a scale.

^e Reliability is based on the exploratory subsample ($N = 500$).

correlations with the direct assessment of reading proficiency. As the data are nested (children within classrooms), the last column in Table 2 lists the percentage of criterion measure variance that reflects children's actual individual differences, with parenthetical values indicating how much of that variance is accounted for by a given LBS scale. These results from HLM generally comport with interpretations based on ordinary correlations, but are more precise in estimating relative criterion-related validity for LBS scores.

Of particular note, both LBS dimensions have the expected moderately high negative correlation with behavior problems within a learning context. Table 2's last column entry for the ASCA Learning Context Problems scale indicates that, whereas 85.2% of score variance stems from children's individual differences (rather than teacher or classroom characteristics), 52.0% of that variance is predictable from children's Competence Motivation scores and 33.5% from Strategy/Flexibility scores, suggesting good correspondence. As many attentive behavior items are found under Competence Motivation, it is unsurprising that this dimension has a high negative correlation with the DBDRS Inattention scale at $-.70$, and predicts markedly more of the

explainable variance of that scale than of the others, at 63.8%. In contrast, only 5.6% of the explainable variance in the ASCA-H ADH scale can be predicted from Strategy/Flexibility scores, suggesting a smaller amount of association than even the modest .22 correlation. The low negative and occasionally nonsignificant relationships between LBS scores and parent behavior ratings were anticipated based on past research discoveries of behavior rating differences between teachers and parents (Dinnebeil et al., 2013; Hartley, Zakriski, Wright, 2011; Lane, Paynter, & Sharman, 2013) and the fact that parents and teachers view different types of behaviors, and the LBS is specifically focused on classroom learning behaviors (Hartman, Rhee, Willcutt, & Pennington, 2007).

Demographic trends

Table 3 displays the mean population distribution of Competence Motivation and Strategy/Flexibility by student gender and grade level in Trinidad and Tobago, whereas Table 4 shows the distribution by student gender and ethnicity. Multivariate analysis of variance (MANOVA), with grade level, gender, and ethnicity as independent variables and the two LBS dimensions as

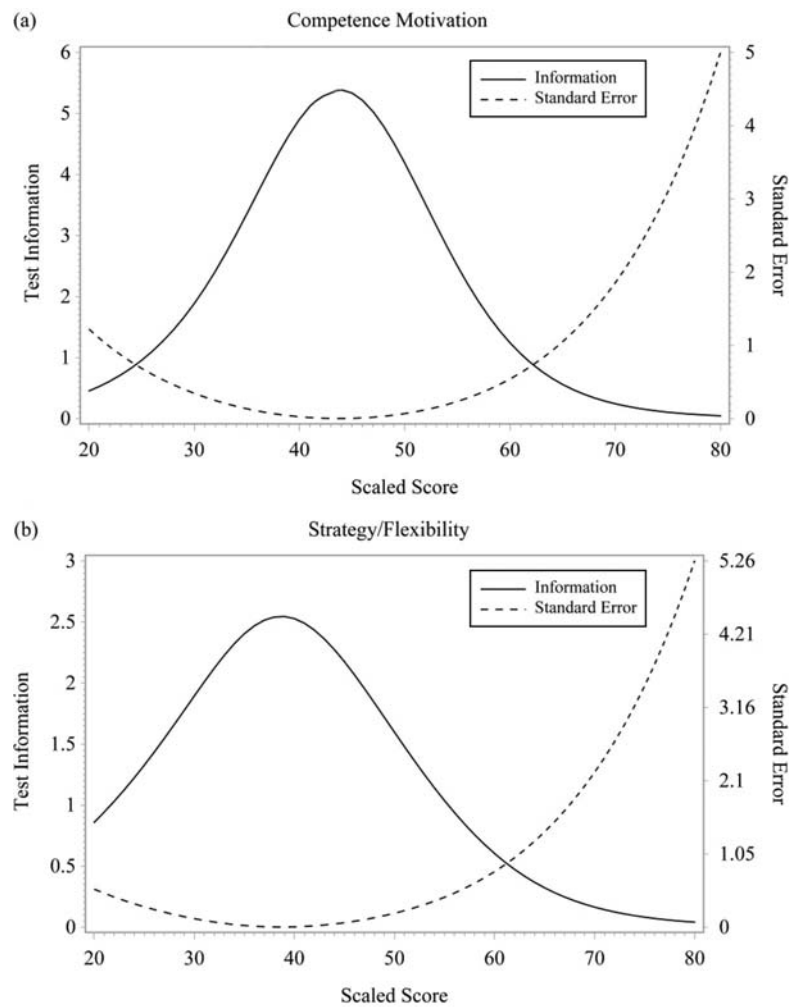


Figure 1. Distributions of estimated information functions and standard errors for LBS scales.

dependent variables, was used to indicate whether there were mean differences in dimensional scores across demographic groups. Subsequent univariate analyses of variance (ANOVAs) suggest the means of LBS Strategy/Flexibility scores vary significantly at $p < .05$ where scores are higher for children in Standard 2 than in Standard 5, for East Indian children compared to African children, and for female children compared to male children. The latter effect is in line with findings that Caribbean girls generally perform better academically than boys (Bowe, 2015; De Lisle, Smith, & Jules, 2005; Kutnick, Jules, & Layne, 1997), with Evans (1999) at least partially attributing this to girls being more likely to pay attention and boys engaging in more disruptive behaviors. No significant effects were found among groups on the Competence Motivation dimension.

Discussion

This study investigated the factor structure of the LBS for a normative sample of schoolchildren in Trinidad and

Tobago. Exploratory and confirmatory analyses revealed two unique and reliable learning behavior scales (Competence Motivation and Strategy/Flexibility). This result did not support the construct generalization of the four scales from the U.S. standardization sample, replicating only two of them, unlike the dimensions observed in other cultural populations. Generally when a scale is not replicated across cultures, it may be due to either lack of configural invariance (number of factors underlying the construct), or lack of loading invariance (which items comprise the factors). The former can be attributed to attempts to simply import a construct from one culture to another, where it may be more nuanced or differentiated; the latter may be explained by incomplete overlap in concepts and definitions such that item content is less appropriate (Chen, 2008).

In the Trinidad and Tobago normative sample, the items comprising the other two U.S. dimensions (Attitude Toward Learning and Attention-Persistence) were mostly subsumed under Competence Motivation, essentially combining to form one factor related to

Table 2. Relationships between LBS scores and concurrent criterion measures.

Criterion measure	Competence motivation	Strategy & flexibility	% Explainable variance ^a
Adjustment Scales for Children and Adolescents, Phenotype Scales (teacher rating)			
Overactivity (<i>n</i> = 845)	-.44 (27.0)	-.61 (43.2)	86.2
Underactivity (<i>n</i> = 845)	-.53 (40.3)	-.24 (8.3)	87.3
Adjustment Scales for Children and Adolescents, Context Scales (teacher rating)			
Peer context problems (<i>n</i> = 845)	-.43 (18.4)	-.58 (41.6)	73.1
Teacher context problems (<i>n</i> = 845)	-.55 (31.3)	-.38 (8.8)	86.8
Learning context problems (<i>n</i> = 845)	-.69 (52.0)	-.62 (33.5)	85.2
Disruptive Behavior Disorder Rating Scale (teacher rating)			
Inattention (<i>n</i> = 682)	-.70 (63.8)	-.63 (58.3)	87.9
Oppositional/Defiant (<i>n</i> = 640)	-.44 (45.7)	-.58 (52.9)	93.6
Impulsivity/Overactivity (<i>n</i> = 638)	-.31 (33.9)	-.59 (56.0)	81.9
Oral Reading Fluency (direct assessment)			
Fall ORF Mean of A & B passages (<i>n</i> = 682)	.35 (37.5)	.23 (17.2)	59.6
Winter ORF Mean of A & B passages (<i>n</i> = 711)	.38 (29.3)	.24 (7.7)	58.0
Spring ORF Mean of A & B passages (<i>n</i> = 682)	.35 (37.5)	.23 (17.2)	59.6
Adjustment Scales for Children and Adolescents—Home (parent rating)			
ADH (<i>n</i> = 712)	-.15 (9.5)	-.22 (5.6)	90.3
Conduct Problems (<i>n</i> = 711)	-.19 (17.3)	-.18 (18.9)	100.0
Underactivity (<i>n</i> = 711)	-.09 [†] (2.8)	-.05 [†] (-0.2)	87.0

Note. Nonparenthetical entries are Pearson product moment correlations. Parenthetical entries indicate the percentage of variance in the respective criterion measure scores between children within classrooms that is accounted for by a given LBS scale score. Values equal 1 – reduction in the intraclass correlation (100) as estimated via hierarchical linear modeling. Each two-level random coefficients model entered a given LBS scale as the covariate. All correlations and fixed effects associated with LBS scales are significant statistically at $p < .01$ unless indicated [†](nonsignificant). LBS = Learning Behaviors Scale, ORF = Oral Reading Fluency, ADH = Attention Deficit and Hyperactivity disorder.

^aTotal percentage of potentially explainable variance between children within classrooms. Values equal 1 – intraclass correlation (100), where the intraclass correlation was estimated via hierarchical linear modeling. Each two-level, unconditional means model applied random intercepts for classrooms, where the random effect was significant at $p < .001$.

motivation while the other was related to strategy. This finding illustrates the necessity of confirming cross-cultural equivalence of measures, as it may suggest either that children in the islands demonstrated insufficient behavioral variability in regard to motivation or teachers were insensitive to such variability that would permit

Table 3. Mean population distribution of competence motivation and strategy/flexibility by gender and grade level in Trinidad and Tobago.

Gender	Competence motivation	Strategy/Flexibility
	<i>M</i> (<i>SD</i>)	<i>M</i> (<i>SD</i>)
Infant 1		
Male (<i>n</i> = 50)	49.9 (9.4)	51.6 (8.2)
Female (<i>n</i> = 50)	49.4 (9.2)	50.4 (9.3)
Infant 2		
Male (<i>n</i> = 50)	50.4 (11.6)	50.0 (11.5)
Female (<i>n</i> = 50)	50.9 (8.9)	51.5 (8.4)
Standard 1		
Male (<i>n</i> = 50)	49.0 (11.6)	48.7 (10.6)
Female (<i>n</i> = 50)	49.3 (8.3)	48.5 (10.5)
Standard 2		
Male (<i>n</i> = 50)	49.8 (9.6)	50.2 (9.3)
Female (<i>n</i> = 50)	54.5 (10.5)	53.7 (10.0)
Standard 3		
Male (<i>n</i> = 50)	49.7 (12.7)	46.6 (10.2)
Female (<i>n</i> = 50)	48.7 (8.8)	49.5 (9.9)
Standard 4		
Male (<i>n</i> = 50)	50.1 (9.6)	48.9 (9.5)
Female (<i>n</i> = 50)	51.1 (8.5)	53.4 (8.7)
Standard 5		
Male (<i>n</i> = 50)	48.0 (9.8)	47.0 (10.8)
Female (<i>n</i> = 50)	48.8 (9.9)	48.0 (11.2)
Total		
Male (<i>n</i> = 350)	49.5 (10.6)	49.0 (10.1)
Female (<i>n</i> = 350)	50.4 (9.3)	50.7 (9.9)

emergence of nuanced dimensionality. This sensitivity to cross-cultural differences has been noted in other scales that measure approaches to learning (e.g., LSI, LPQ), and is not necessarily unexpected when dealing with concepts relying heavily on environmental and behavioral factors that would be heavily influenced by cultural context. There has been much exploration around the possible consequences of using behavioral assessment instruments designed for use in Western cultures in Eastern contexts, or simply translating instruments into other languages without assessing construct validity in that other population; this study serves to add to the literature emphasizing the risk of automatically assuming psychometric generalizability across all cultures even when the

Table 4. Mean population distribution of competence motivation and strategy/flexibility by gender and ethnicity in Trinidad and Tobago.

Gender	Competence motivation	Strategy/Flexibility
	<i>M</i> (<i>SD</i>)	<i>M</i> (<i>SD</i>)
African descent		
Male (<i>n</i> = 137)	47.8 (11.0)	47.9 (10.3)
Female (<i>n</i> = 132)	50.1 (9.3)	49.7 (10.6)
East Indian descent		
Male (<i>n</i> = 126)	51.0 (10.5)	50.0 (9.6)
Female (<i>n</i> = 131)	50.5 (9.3)	52.1 (8.8)
Mixed descent		
Male (<i>n</i> = 76)	50.4 (9.9)	49.1 (10.1)
Female (<i>n</i> = 73)	51.2 (8.8)	49.9 (10.1)
Total		
Male (<i>n</i> = 350)	49.5 (10.6)	49.0 (10.1)
Female (<i>n</i> = 350)	50.4 (9.3)	50.7 (9.9)

language is the same, and assuming that all scales and subscales contain the same components.

Nevertheless, the present study suggests an alternative model for Trinidad and Tobago that, while not the same as the U.S. model, measures similar constructs, affords locally valid and reliable scores, and will yield information that can be used in formulating treatment plans for children with academic difficulties. In the United States, many early education programs, including most state departments of education, find it useful to support curricula meant to build cognitive and socio-behavioral skills with initiatives intended to foster the development of learning behaviors (McDermott et al., 2016; Scott-Little, Kagan, & Frelow, 2005). As something behavioral in nature, learning behaviors are theoretically teachable, and improving them should help to improve the cognitive and sociobehavioral skills that are linked to them. By identifying which learning behaviors are potentially in need of improvement, teachers can better determine which behaviors that contribute to them might be worth targeting. For example, in practice, a low score on the Strategy/Flexibility scale suggests that a child might have some deficit in orderly problem solving. This concern could be mitigated by a program to help the child sequentially improve understanding of cause and effect and recognize possible consequences of an action, translating into an increased ability to develop a plan for a multistage activity while considering possible consequences. A low score on the competence motivation scale might suggest that efforts should be made to encourage the child's engagement in activities that were previously challenging, and increase frustration tolerance and persistence.

Limitations and future directions

The inability of teachers to discern the two other constructs found in the U.S. standardization sample warrants further investigation as to cause. Is this something that can be rectified by training, or is it simply reflecting a more well-disciplined population where nuances are less sharp than in the United States? If the latter, more item development might be justified to tailor the assessment to more closely match population characteristics and needs. Norms for classroom behavior may be demonstrably different in U.S. classrooms as compared to Trinidad and Tobago, necessitating further research into local classroom learning dynamics.

Future work should also include additional standardized academic achievement measures to strengthen the criterion validity assessment for the LBS. Validity analyses of academic achievement were limited to oral reading fluency in this study because no other academic

assessments were administered in Trinidad and Tobago during the initial data collection. As well, since the inclusion of secondary schools in the study was logistically unfeasible, the normative sample was limited to elementary school students. Future work might expand the measure to include secondary school students.

As the advantage of learning behaviors is predicated on the idea that they can be modified, there is value to learning how they might change over time, so that interventions might be better targeted temporally. Thus, further work might include studies over short intervals within an academic year (e.g., see McDermott et al., 2011, on the Learning-to-Learn Scales), and the developmental evolution of learning behaviors as children age and transition through areas of schooling (e.g., see McDermott et al., 2014, on longitudinally equating the PLBS and LBS across prekindergarten, kindergarten, and first grade with a sample of Head Start children). Questions also remain as to whether efforts to improve learning behaviors should be orchestrated through behavioral modeling programs that target learning behaviors themselves, or a more holistic approach in carefully integrating them within the central curricular content (Fantuzzo, Gadsden, & McDermott, 2011).

Conclusion

The present study provides a foundation for studying learning behaviors in Trinidad and Tobago, though more research is necessary into the environmental factors unique to classrooms in the island schools. However, the information provided by these scales will aid in determining patterns of learning behaviors and be useful in targeting interventions and instructional approaches for struggling students. Combined with information on social and emotional adjustment, cognitive ability, and academic achievement, information on learning behaviors can help educators develop a better understanding of opportunities for instructional modifications. Additionally, the notion of positive learning behaviors serving as protective factors against negative academic and behavioral outcomes is worthy of study, as learning behaviors, unlike other established protective factors, can be improved through intervention.

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