

Development and Initial Validation of the Child Disgust Scale

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Although disgust sensitivity (DS) has been implicated in the development of anxiety disorders in children, the absence of a measure of DS specifically for children has not allowed for an adequate test of this claim. To fill this important gap in the literature, this investigation presents a series of studies on the development and examination of the psychometric properties (including reliability, validity, and factor structure) of scores on a newly developed Child Disgust Scale (CDS). Exploratory factor analysis in Study 1 ($N = 1,500$) found that a bifactor model, which allows for a “g” DS factor in addition to 2 distinct factors of Disgust Avoidance and Disgust Affect, was the best fit for the data. Study 2 ($N = 573$) confirmed a two-factor bifactor model above and beyond a 1-factor model that controlled for method effects due to reverse-worded items. Results from Study 3 ($N = 50$) provided support for convergent and discriminant validity such that scores on the CDS were significantly correlated with measures of anxiety and fear, but not depression. Finally, Study 4 ($N = 86$) found that the CDS differentiated children with a diagnosis of specific phobia ($n = 43$) from a matched nonclinical community sample of children ($n = 43$), such that those with a specific phobia reported greater DS compared with controls. Results from these studies suggest that the CDS is a developmentally appropriate measure with good psychometric properties that can aid research on the role of disgust sensitivity in anxiety-related disorders in children.

Keywords: Child Disgust Scale, disgust, factor analysis, measurement development, phobia

Disgust is a basic emotion marked by distinct behavioral, cognitive, and physiological properties (e.g., Ekman, 1992; Izard, 1992, 1993; Levenson, 1992) that originally evolved to protect

against oral incorporation of potential contaminants (Angyal, 1941; Tomkins, 1963). However, most contemporary accounts conceptualize disgust as a defensive response to a broad range of stimuli (Rozin, Haidt, & McCauley, 2000). Although disgust is considered a basic universal emotion (Ekman, 1992), individuals vary in the degree to which they experience disgust. Such individual differences have been captured by the construct of *disgust sensitivity*, defined as the propensity to experience disgust to a wide range of stimuli (de Jong & Merckelbach, 1998; Olatunji & Sawchuk, 2005). As such, disgust sensitivity is considered a dispositional personality trait that is stable and varies across individuals (Olatunji & McKay, 2006). It has been suggested that individual differences in disgust sensitivity may be acquired via social transmission during childhood (Rozin et al., 2000; Stevenson, Oaten, Case, Repacholi, & Wagland, 2010).

Programmatic research on the developmental origins of disgust sensitivity has informed current knowledge on the etiology of various forms of psychopathology. Indeed, disgust sensitivity has been implicated in the development and maintenance of anxiety-related disorders. For example, anxious individuals have been shown to report significantly greater disgust sensitivity than non-anxious controls (Koch, O'Neill, Sawchuk, & Connolly, 2002; Olatunji, Lohr, Smits, Sawchuk, & Patten, 2009). Measures of disgust sensitivity have also been found to correlate with measures

This article was published Online First April 6, 2015.

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This work was supported in part by Training Grant T32-MH18921 from the National Institute of Mental Health, Grant P30HD1505 from the National Institute of Child Health and Human Development, and Grant UL1TR000011 from the National Center for Advancing Translational Science/National Institutes of Health. We thank Bethany Ciesielski, Paul Rozin, Jon Haidt, and David Cole for their assistance throughout this project.

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of spider phobia (Mulken, de Jong, Merckelbach, 1996; Thorpe & Salkovskis, 1998), obsessive-compulsive disorder (OCD; David et al., 2009; McKay & Moretz, 2009; Thorpe, Patel, & Simonds, 2003), blood-injection-injury phobia (BII; Matchett & Davey, 1991; Olatunji, Williams, Sawchuk, & Lohr, 2006; Page & Tan, 2009), and posttraumatic stress disorder (Engelhard, Olatunji, & de Jong, 2011). Disgust sensitivity also appears to positively correlate with symptoms of hypochondriasis (Davey & Bond, 2006) and eating disorders (Davey, Buckland, Tantow, & Dallos, 1998). It is important that the association between disgust sensitivity and anxiety-related disorder symptoms remains significant even when controlling for negative affect (Muris, Merckelbach, Schmidt, & Tierney, 1999; Olatunji, 2006; Olatunji et al., 2009).

A very common observation in the literature is that females report more anxiety symptoms than males (see Craske, 2003, for review). Gender differences in anxiety among adults may emerge, in part, because of gender stereotyped reinforcement contingencies that have been established during childhood. For example, women are less likely than men to be reinforced for behaviors that are assertive and active, and more likely to be reinforced for avoidant and more passive behaviors (Chambless & Mason, 1986). Consequently, women may be more prone than men to experience fearful and anxious symptoms (Craske, 2003; Kelly, Forsyth, & Karekla, 2006). A similar pattern of differences between men and women has also been found for disgust sensitivity, with women reporting higher levels of disgust sensitivity than men (Haidt, McCauley, & Rozin, 1994). Subsequent research building on these findings has shown that the gender differences in some anxiety disorder symptoms in adults, including contamination-based OCD (Olatunji, Sawchuk, Arrindell, & Lohr, 2005), BII phobia (Olatunji, Arrindell, & Lohr, 2005), and spider phobia (Connolly, Olatunji, & Lohr, 2008), may be mediated by gender differences in disgust sensitivity.

Although previous research has implicated disgust sensitivity in anxiety-related disorders, most of the available studies have been undertaken with adults. Epidemiological data, however, suggest that the average age of onset for anxiety-related disorders is 11 years, with some disorders having even earlier onsets, such as spider phobia at 7 years and BII phobia at 5.5 years (Kessler, Ruscio, Shear, & Wittchen, 2009). The few available studies examining the relation between disgust sensitivity and clinical disorders in children have shown that disgust sensitivity is uniquely associated with symptoms of small animal phobias (de Jong, Andrea, & Muris, 1997; Muris et al., 1999; Muris, van der Heiden, & Rassin, 2008), social phobia, agoraphobia, OCD, and eating problems in school-age children (Muris et al., 1999; Muris, van der Heiden, & Rassin, 2008). A more recent extension of this correlational work with children has shown that disgust sensitivity may also be a causal factor in the development of anxiety symptoms. Specifically, Muris, Mayer, Huijding, and Konings (2008) found that providing children with disgust-related information about an unknown animal subsequently increased fear beliefs of the animal.

The available research suggests that disgust sensitivity is associated with anxiety symptoms in children. However, the number of available studies is limited in scope because of, in part, the absence of an age-appropriate measure of disgust sensitivity specifically designed for children. The handful of available studies that have examined disgust sensitivity among children have used simplified

or age-downward extensions of adult measures (see Muris et al., 1999; Muris, Mayer, et al., 2008; Muris, van der Heiden, & Rassin, 2008.). Although this approach has allowed for the assessment of disgust sensitivity and its association with fear and anxiety in children (Muris et al., 2012), it possesses significant limitations. For example, a simple downward extension of adult scales may not capture important developmental nuances that contribute to a more reliable and valid assessment of the disgust sensitivity construct in children. Furthermore, items on adult measures of disgust sensitivity may not be age appropriate and, thus, may be measuring a different construct in children than intended. Indeed, prior research has shown that there are developmental changes in the interpretation of instructions on measures of related constructs, such as negative affect (Campbell, Rapee, & Spence, 2001). Additionally, many of these studies have used a restricted range of children ages 9–13 years old. However, anxiety disorders often develop in even younger children (Kessler et al., 2005). Therefore, a developmentally sensitive measure of disgust sensitivity that can be employed in a wider age range of children is desired.

Development of the Child Disgust Scale

This investigation sought to fill an important gap in the current literature by developing a reliable and valid self-report measure of disgust sensitivity that is age-appropriate for children. We first consulted the Disgust Scale—Revised (DS-R; Olatunji et al., 2007), a psychometrically sound and widely used adult measure of disgust sensitivity, as a reference for the range of stimuli that are generally considered to be disgusting. With this as an initial framework, we then consulted with developmental psychologists to get direction on the stimuli that would appropriately evoke disgust in children. This revealed that although some of the themes on the DS-R may be relevant to children, a majority of the items were generally not appropriate. With input from developmental psychologists, we then generated an initial set of items ourselves that was reviewed by developmental psychologists, experts in disgust, and parents of young children. Through this process that involved detailed consultation with these individuals, we were able to eliminate items and introduce new ones. It is through this iterative process that we were able to generate the necessary items for the scale.

The DS-R consists of three subscales: Core Disgust, Contamination Disgust, and Animal-Reminder Disgust. Thus, age-appropriate items that covered the three disgust domains were identified in developing the Child Disgust Scale (CDS). In addition to the age-relevance of the items, we also sought to expand the readability of the items, which was assessed by using the readability statistics offered by Microsoft Word 2010. The final version of the CDS was found to have a Flesch–Kincaid reading ease of 94.7 and a reading grade level of 2.9 (Flesch, 1951). For comparison, the DS-R possesses a reading ease of 75.6 and a reading grade level of 4.6.

Development of the CDS also included adopting a more age-appropriate rating scale. The DS-R currently has a 5-point Likert-type scale with 0 = *Strongly Disagree* and 4 = *Strongly Agree*, which may be too complex for children. Indeed, Ollendick (1983, p. 685) noted in his revision of the Fear Survey Schedule for Children that “young children, as well as impaired clinical-child populations, have difficulty understanding and discriminating re-

sponses on [a] 5-point scale". Accordingly, a 3-point response scale (0 = *Always*, 1 = *Sometimes*, 2 = *Never*) consistent with other self-report scales used with young children was used (e.g., Ebesutani et al., 2012).

The initial item pool of the CDS contained 22 items. Based on the feedback received from disgust researchers, developmental psychologists, and parents of young children, four items were dropped from the scale and several items were altered to increase readability and relevance to children (e.g., "I would share my drinks or snacks with my friends" was changed to "I would still drink my juicebox even if I saw another kid drink out of it").

The scale development process resulted in an 18-item CDS (see Appendix) that assesses sensitivity to *core disgust* (i.e., items related to oral incorporation of contaminants or contact with bodily waste or small animals), *contamination disgust* (i.e., items related to possible contamination by contagion of ill persons), and *animal-reminder disgust* (i.e., items related to threat to the body envelope, injury to the body, or death). The 18-item CDS was designed to have six exclusive items that were thematically related to each of the three disgust domains from the DS-R that have been used with adults. Given that previous studies have found that internalizing problems tend to be associated with less symptom differentiation during younger years of development (Brady & Kendall, 1992; De Bolle & Fruyt, 2010; Jacques & Mash, 2004; Price et al., 2013), we anticipated that the resultant factor structure may be simpler (and less differentiated) among this youth sample.

Scoring

In addition to the three hypothesized domains, the CDS yields a total (summed) score for which 0 = *Always*, 1 = *Sometimes*, and 2 = *Never* and for which higher numbers indicate greater disgust sensitivity. Seven items (Items 5, 6, 9, 14, 15, 16, 18) are reverse-scored.

Overview of the Present Investigation

The current investigation examines the factor structure and psychometric properties of the CDS scale scores across four studies. Study 1 examines the reliability of the CDS scores among elementary and middle school-age children. Bifactor exploratory factor analysis (EFA) was also used in Study 1 to examine the latent structure of the CDS and to evaluate whether disgust sensitivity in children can be conceptualized as unidimensional or multidimensional as it is in adult models. The bifactor analysis is preferable to traditional EFA procedures because it allows for a general "g" factor of disgust proneness, as well as specific disgust domains. Given that children, especially young children, may not have acquired reliable disgust responses to certain higher order factors (e.g., contamination), the bifactor model allows for a framework that can be applied to a wider developmental range. The bifactor model has also been found to fit psychological constructs well (Reise, Morizot, & Hays, 2007) thus making it a suitable model for examining disgust. Indeed, the bifactor model was recently found to be the best fit for the data in an adult measure of disgust propensity compared with exclusive unidimensional or multidimensional models (Olatunji, Ebesutani, & Reise, 2014). Confirmatory factor analysis (CFA) was then used in Study 2 to confirm the factor structure of the CDS in an independent

sample of elementary and middle schoolchildren based on the findings of Study 1. Study 3 then examined the convergent and discriminant validity of the CDS relative to self-report measures of fear, anxiety, and depression. Last, the "known-groups validity" of the CDS was examined in Study 4 by comparing differences between those with a diagnosis of a specific phobia and matched nonclinical children.

Study 1 Method

Participants

Participants were recruited from public schools in Mississippi. The final sample included 1,500 elementary and middle schoolchildren (778 boys and 722 girls) who completed the CDS. There were 186 (12%) children in 2nd grade, 198 (13%) in 3rd grade, 167 (11%) in 4th grade, 225 (15%) in 5th grade, 213 (14%) in 6th grade, 252 (17%) in 7th grade, 253 (17%) in 8th grade, and 6 (<1%) students that did not provide a grade. With respect to race, 1142 (76%) were White/Caucasian, 210 (14%) were Black/African American, 20 (1%) were Asian, 79 (6%) were Hispanic, and 49 (3%) self-identified as Other.

Among the 1,500 included youth, 1,365 (91.0%) had no missing data, 109 (7.3%) had one missing item, 16 (1.1%) had two missing items, six (0.4%) had three missing items, two (0.1%) had four missing items, one (0.1%) had five missing items, and one (0.1%) had six missing items.

Procedure

This study was reviewed and approved by the University of Mississippi Institutional Review Board. Passive consent was used, in which participants and their families were provided with the opportunity to decline participation one week prior to administration of the CDS. Families who chose to not participate in the study were asked to return a signed form to the school indicating their preference to be excluded from participation in the study. On the day of data collection, student assent forms and the CDS were distributed to the classrooms and administered by teachers. Students were given a second opportunity to decline participation prior to being given their forms. Administrators aided in distribution and collection of the scale, and a project research assistant was onsite to organize data collection, answer questions, and collect completed measures from each classroom. Children were given the instructions: "Each sentence below is a statement that might be disgusting. Choose how often you would do what the sentence says by circling: Always, Sometimes, or Never."

Data Analytic Strategy

Missing data. We used the recommended multiple imputation method available in Mplus (based on 10 imputed dataset) to handle missing data (Rubin, 1996).

Exploratory bifactor analysis (EFA) of the CDS. Using the PSYCH package available in the R statistical software (R Development Core Team, 2008), a Schmid-Leiman bifactor EFA using oblique rotation was used given that the factors were expected to be intercorrelated (see Reise, Moore, & Haviland, 2010, for a detailed description of the Schmid-Leiman bifactor EFA procedure).

ture). Data were treated as categorical (ordinal) because of the items being derived from a Likert-scale (Brown, 2006). We used the recommended procedures when conducting EFA on categorical data such that calculations were performed on polychoric correlation matrices (Holgado-Tello, Chacón-Moscó, Barbero-García, & Vila-Abad, 2010) with the robust weighted least squares estimator (WLSMV; Flora & Curran, 2004; Muthén, du Toit, & Spisic, 1997). The following metrics were used to evaluate the outcome of this analysis: (a) the number of eigenvalues greater than 1.0, (b) the scree plot, (c) the interpretability of each solution, and (d) the fit of each EFA solution according to the root-mean-square error of approximation (RMSEA) fit statistic. Additionally, given some criticism that the “eigenvalues greater than 1.0” criterion may yield too many factors (Velicer & Jackson, 1990), all criteria were considered when selecting the number of factors.

Study 1 Results

Descriptive Statistics

Descriptive statistics are presented in Table 1, including mean ratings for each of the CDS items and skew and kurtosis statistics. Examination of the skew and kurtosis of the CDS items revealed some significant z values, which suggested the presence of some nonnormal data. As noted above, we used the robust WLSMV estimator, which overcomes concerns related to potential biased parameter estimates caused by nonnormality (Muthén et al., 1997).

Exploratory Factor Analysis of the CDS

The bifactor EFA factor loadings associated with the three-factor and two-factor solutions appear in Table 2. Although a three-factor solution was originally hypothesized, Factor 1 did not have any items that met the cutoff criteria of .32 as identified by Tabachnick and Fidell (2001). Additionally, Factor 3 only contained two items with factor loadings above .32 and factors with fewer than three items are generally considered weak, unstable, and negligible (Costello & Osbourne, 2005). We therefore did not consider the three-factor model a viable solution. The two-factor

model was considered to be the most interpretable solution based on its strong model fit (RMSEA = .05) and interpretability. Three items were removed given that they did not load onto the general factor. The final measure therefore consisted of 15 items. Factor I consisted of 9 items that are largely characterized by avoidance of disgust eliciting stimuli (e.g., “If a dog licked my popsicle, I would still eat it”). We therefore labeled this first factor “Disgust Avoidance.” Factor II consisted of 6 items that were characterized by affective responses to disgust eliciting stimuli (e.g., “I feel sick if I see a dead animal on the side of the road”). We therefore labeled this second factor “Disgust Affect.”

Reliability of the CDS Score

The CDS total score (or general factor) was associated with adequate internal consistency reliability estimate ($\alpha = .78$). Table 1 displays alpha-if-deleted values for each of the 15 retained CDS items (with relation to the total score). These results do not reveal any items that need to be removed from the total score.

The Disgust Avoidance subscale scores ($\alpha = .78$) and Disgust Affect subscale scores ($\alpha = .69$) were also associated with adequate reliability estimates. We examined alpha-if-deleted values for each of the six items of the Disgust Affect subscale given that alpha fell just under the .70 reliability benchmark (Nunnally, 1978). All alpha-if-deleted values ranged from .65 to .66, suggesting that removal of items would not improve reliability. We therefore did not remove any items from this scale and we decided to retain the factor because its scale score reliability estimate fell extremely close to the .70 benchmark for adequate reliability. Item-total correlations for each item also appear in Table 1. All items moderately correlated with the total score.

Gender Differences

Significant sex differences were found for the CDS scores, $t(1,498) = 12.40, p < .001$, Cohen's $d = .55$, such that girls ($M = 19.39, SD = 4.79$) reported significantly greater disgust sensitivity than boys ($M = 16.27, SD = 4.95$). Compared with boys, girls were also found to report higher levels of Disgust Avoidance

Table 1
Study 1 Descriptive Statistics for Child Disgust Scale (CDS) Items

Item	<i>M</i> (0.00–2.00)	<i>SD</i>	Skewness	Kurtosis	Item–total correlation	Alpha-if-removed
CDS 2	1.73	0.59	–2.05	02.92	.34	.77
CDS 4	0.87	0.80	0.23	–1.41	.45	.76
CDS 5 ^a	0.42	0.66	1.29	0.38	.34	.77
CDS 6 ^a	0.47	0.71	1.16	–0.06	.32	.77
CDS 7	1.76	0.55	–2.17	3.63	.47	.76
CDS 8	1.68	0.60	–1.72	1.77	.39	.77
CDS 9 ^a	0.95	0.83	0.10	–1.56	.38	.77
CDS 10	1.43	0.67	–0.74	–0.55	.38	.77
CDS 11	1.14	0.81	–0.25	–1.43	.49	.76
CDS 12	1.76	0.53	–2.15	3.65	.34	.77
CDS 13	1.56	0.67	–1.24	0.26	.47	.76
CDS 15 ^a	0.60	0.76	0.80	–0.82	.41	.76
CDS 16 ^a	0.61	0.77	0.81	–0.87	.32	.77
CDS 17	1.63	0.61	–1.42	0.91	.35	.77
CDS 18 ^a	1.23	0.81	–0.50	–1.29	.33	.77

^a Item was reverse-scored.

Table 2
 Study 1 Exploratory Factor Analytic Results for the Two-Factor and Three-Factor Solutions

Item	2-factor			3-factor			
	General disgust	Disgust 1	Disgust 2	General disgust	Disgust 1	Disgust 2	Disgust 3
CDS 1	.24	.45		.47			
CDS 2	.34	.56		.67	.30		
CDS 3	.30	.52		.48			.70
CDS 4	.42	.46		.59	.20		
CDS 7	.49	.56		.68	.21		
CDS 8	.38	.53		.58			.20
CDS 10	.35	.46		.47			.48
CDS 11	.44	.39	.23	.56	.20	.28	
CDS 12	.35	.58		.60			.27
CDS 13	.45	.58		.66			.25
CDS 14 ^a	.31	.23	.21	.35		.25	
CDS 17	.32	.50		.59	.24		
CDS 5 ^a	.36		.54			.62	
CDS 6 ^a	.34		.48			.56	
CDS 9 ^a	.34		.43	.23		.50	
CDS 15 ^a	.42		.58	.26		.67	
CDS 16 ^a	.34		.49			.56	
CDS 18 ^a	.33		.48			.57	

Note. Factor loadings under .20 are not listed. CDS = Child Disgust Scale.

^a Item was reverse-scored.

($t[1,498] = 8.14, p < .001$, Cohen's $d = .42$; girls: $M = 1.59, SD = 0.35$; boys: $M = 1.43, SD = 0.41$) and Disgust Affect ($t[1,498] = 11.70, p < .001$, Cohen's $d = .60$; girls: $M = 0.93, SD = 0.45$; boys: $M = 0.67; SD = 0.41$).

Study 1 Discussion

The CDS showed adequate internal consistency in an initial sample of elementary and middle school-age children with girls reporting greater disgust sensitivity compared with boys. This finding is consistent with the adult research that finds that women report greater disgust sensitivity than men (Davey, 1994; Haidt et al., 1994; Schienle, Start, Walter, & Vaitl, 2003). Although the CDS was modeled after the three-factor structure of the adult DS-R (Core Disgust, Contamination Disgust, and Animal-Reminder Disgust), a bifactor EFA revealed a general disgust factor and only two interpretable factors (Disgust Avoidance and Disgust Affect). The two-factor bifactor solution suggests that the structure of disgust sensitivity among children (as assessed by the CDS) may be best characterized by responses to disgust eliciting stimuli rather than the nature of the stimuli themselves. However, the two-factor bifactor solution observed in Study 1 requires confirmation before definitive inferences can be made regarding the factor structure of the CDS. Accordingly, in Study 2, we used CFA in a new sample to examine the fit of the two-factor bifactor model relative to alternative models.

Study 2 Method

Participants

Participants were recruited from public schools in Mississippi. The final sample included an independent sample of 573 elementary and middle school children (262 boys and 311 girls). The mean age was 9.07 years old ($SD = 1.51$) with an age range of

6–13 years. With respect to race, 509 (89%) were White/Caucasian, 22 (4%) were Black/African American, 3 (1%) were Asian, 21 (4%) were Hispanic, and 18 (2%) self-identified as Other.

Among the 573 included youth, 523 (91.3%) had no missing data, 40 (7.0%) had one missing item, 7 (1.2%) had two missing items, and 3 (0.5%) had three missing items.

Procedure

The procedure was the same as described in Study 1.

Data Analytic Strategy

Confirmatory factor analyses. We conducted CFA on the CDS items using Mplus 7.11 (Muthén & Muthén, 2010). Because of the CDS data being categorical (ordinal) in nature, we used polychoric correlations (Holgado-Tello et al., 2010) and the robust weighted least-squares with mean and variance adjustment (WLSMV) estimator, Flora et al., 2004; Muthén et al., 1997). We used full-information maximum likelihood (FIML) to impute missing data given that FIML has been recommended as one of the best methods for handling missing data in many contexts (Allison, 2003; Arbuckle, 1996; Schafer & Graham, 2002); that said, bifactor modeling has only recently begun to be applied in psychology and so relatively less is known about how it performs in these contexts, such as when data are imputed via these methods. We examined model fit via the chi-square statistic; the RMSEA (Steiger, 1990), for which smaller values (e.g., less than .08) are indicative of good fit (Hu & Bentler, 1999) and the comparative fit index (CFI; Bentler, 1990), for which larger values (e.g., greater than .95) are considered to indicate good model fit. We used the chi-square difference test (i.e., $\Delta\chi^2_{diff}$) to examine the significance of modifications to the original model.

Measurement invariance across gender. We evaluated measurement invariance of the derived, best-fitting model across males

($n = 262$) and females ($n = 311$) using multigroup CFA. The recommended steps of this process have been outlined by Brown (2006). Specifically, we first examined fit of the single-sample solutions in the male-only and female-only subsamples, separately. If both single-sample solutions evidenced good model fit (based on the fit statistic benchmarks noted in the section above), configural invariance (i.e., “equal form”) is then examined in the combined full sample. Configural invariance examines whether the data from both groups are associated with the same number of factors and item-to-factor loading patterns. Configural invariance is considered supported if the fit indices meet the previously mentioned benchmarks of good model fit (cf. Brown, 2006).

If configural invariance is supported, then metric invariance (i.e., “equal factor loadings”) and scalar invariance (i.e., “equal item thresholds”) can be tested, in successive order. Metric invariance is tested by constraining all *factor loadings* to be the same across groups, and scalar invariance is tested by constraining all *item thresholds* to be the same across groups. For both metric and scalar invariance, we used the Δ CFI difference test to determine whether the invariance model is supported (Chen, 2007). If the difference in the CFI fit index between the constrained and non-constrained model is less than .01 (Δ CFI < .01), then invariance (at the constrained model level) is supported (Chen, 2007; Cheung & Rensvold, 2002). For example, if the equality constraint of equal item thresholds across groups did not lead to a substantial degradation in model fit, then scalar invariance is supported. Scalar invariance is important to examine given that this is the test of differential item functioning (McDonald, 1999). If scale scores are associated with differential item functioning, then individuals who fall on the same level of the underlying latent trait provide systematically different observed scores on that measure’s items.

Without establishing scalar invariance (or the lack of differential item functioning) it has been said that the comparison of mean scores across subgroup is ambiguous because “the effects of a between-groups difference in the latent means are confounded with differences in the scale and origin of the latent variable” (see Cheung & Rensvold, 2002, p. 238).

Study 2 Results

Confirmatory Factor Analyses

We first examined the fit of the two-factor bifactor model resulting from the exploratory bifactor analyses in Study 1. In this bifactor model, all items from the CDS loaded on the general factor (Items 1, 3, and 14, which were dropped from the measure entirely in Study 1 because of insignificant loading on the general factor). Items 2, 4, 7, 8, 10, 11, 12, 13, and 17 loaded on the Factor 1 called (labeled “Disgust Avoidance”) and Items 5, 6, 9, 15, 16, and 18 loaded on Factor 2 called (labeled “Disgust Affect”). This two-factor bifactor model revealed an excellent fit to the data based on the full sample (i.e., RMSEA = .048; CFI = .995). One item did not load significantly on the General Factor (Item 9). We therefore eliminated this item and reran the two-factor bifactor model with Item 9 removed. This resulted in a 14-item bifactor model that was also associated with excellent fit (i.e., RMSEA = .053; CFI = .995). All items then loaded significantly on both the general disgust dimension and their respective content sub-domains. The final two-factor bifactor model with 14 items is presented in Figure 1. It is worth noting that Items 4 (“I would pick up a worm with my hand”) and 18 (“I feel sick if I see someone throw up”) had significant and positive loadings on their respective

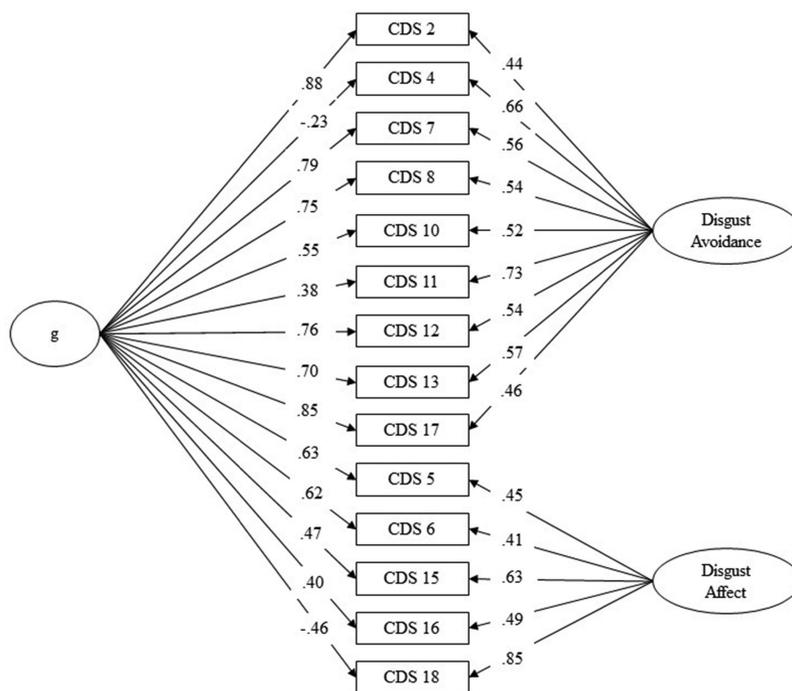


Figure 1. Two-factor bifactor model of the final 14-item Child Disgust Scale (CDS).

subdomains, and significant loadings on the general factor (indicating that they are significantly relevant and pertinent to this bifactor model); however, the loadings on the general factor were negative. Thus, the bifactor model was respecified with the general factor path to Items 4 and 18 removed. The removal of these items was associated with very poor model fit based on some indices, $\chi^2(91) = 20,487.40$, $p < .001$; RMSEA = .113. A chi-square difference test showed that a model that included the negative paths demonstrated significantly better fit than a model that removed those pathways, $\chi^2_{\text{diff}}(28) = 20,321.86$, $p < .001$. Given these analyses, the two negative pathways were retained for the remainder of analyses.

We then compared this (14-item) two-factor bifactor model (that included two negatively loaded items on the general factor) against (a) standard (correlated traits) two-factor model, (b) a unidimensional model, and (c) a unidimensional model that controlled for method effects due to the reverse-worded items. The fit of all competing models appear in Table 3. The two-factor (correlated traits) model was first tested. This two-factor model consisted of Disgust Avoidance and Disgust Affect, without a general factor and was associated with relatively poor model fit based on some fit indices, $\chi^2(76) = 887.74$, $p < .001$; RMSEA = .137, although acceptable model fit based on others (CFI = .960; TLI = .952). However, the chi-square difference test showed that the two-factor bifactor model fit significantly better than this (correlated traits) two-factor model, $\chi^2_{\text{diff}}(13) = 496.70$, $p < .001$.

A unidimensional model of disgust sensitivity was then tested with all 14 CDS items as indicator variables. This model was also associated with relatively poor fit to the data based on some fit indices, $\chi^2(77) = 1,103.77$, $p < .001$; RMSEA = .153, but acceptable model fit based on others (CFI = .950; TLI = .941). The chi-square difference test however showed that the two-factor bifactor model fit significantly better than this one-factor model, $\chi^2_{\text{diff}}(14) = 591.17$, $p < .001$.

Last, we examined the one-factor model that controlled for method effects due to the reverse-worded items. In this model, we set all error terms among all the negatively worded items to be correlated, based on the correlated uniqueness model (cf. Brown, 2003; Marsh, 1996). This one-factor model (controlling for wording method effects) was also associated with relatively poor fit based on some fit indices, that is, $\chi^2(41) = 516.43$, $p < .001$; RMSEA = .142, but acceptable fit based on others (CFI = .977; TLI = .948). This model was not nested under the bifactor model and so the chi-square difference test could not be used to compare model fit. Although Cronbach's alpha is widely used as a measure of reliability, it can sometimes yield misleading results, especially

when data are multidimensional, given that coefficient alpha reflects the reliability of *all* sources of systematic variance, including variance of the presence of the general factor, content group factors, and specific factors (Cortina, 1993). Omega provides a better estimate of reliability as it assumes that items on congeneric rather than tau equivalent (Graham, 2006). Omega hierarchical computed for the total score composite (OmegaH; Zinbarg, Barlow, & Brown, 1997; Zinbarg, Revelle, Yovel, & Li, 2005) provides an estimate of the proportion of variance in scores that is due to the general factor (e.g., general disgust sensitivity). Omega hierarchical for each subscale composite provides an index of the degree to which the subscale scores provide reliable variance after accounting for the general factor. Based on the CFA bifactor loadings in Figure 1, omega for the total scale was .96, and omega hierarchical for the total score was .62. This reveals the presence of a relatively strong general factor, whereby 62% of the variance of this total composite could be attributable to variance on the general factor. Omega hierarchical for the Disgust Avoidance and Disgust Affect subscale composites were .45 and .63, respectively. These results suggest that the Disgust Affect subscale scores provide a high degree of reliable variance after accounting for the general factor; the Disgust Avoidance subscale scores, however, provide a much lower degree of reliable variance after accounting for the general factor.

Measurement Invariance Across Gender

The single-sample solutions evidenced good model fit in the male-only sample (RMSEA = .046; CFI = .998; TLI = .997) and also in the female-only sample (RMSEA = .050; CFI = .994; TLI = .991). All items also loaded significantly on the general factor and their respective subdomain in both groups. The tests of configural invariance was also supported, as evidenced by strong fit indices, RMSEA = .049; CFI = .996, TLI = .995; $\chi^2(126) = 210.94$. Because of the nature of the bifactor model (whereby items load on both the general and a specific factor), the Mplus multigroup confirmatory factor analysis procedures did not allow the specific test of metric invariance by itself; Mplus only allowed the test of configural invariance, and then scalar invariance (constraining both factor loadings and item thresholds, simultaneously). We were thus forced to skip the specific test of metric invariance, and proceed to the test of scalar invariance. The fit indices associated with the scalar invariance model were also strong, RMSEA = .044; CFI .996, TLI = .996; $\chi^2(165) = 256.66$. The Δ CFI test revealed that scalar invariance was supported given that Δ CFI between the configural and scalar model was less than

Table 3
Study 2 Confirmatory Factor Analysis of the 14-Item Child Disgust Scale (CDS) Model Fit Indices

Model tested	χ^2	df	χ^2/df	RMSEA	TLI	CFI
1-Factor model	1,103.77	77	14.33	.153	.941	.950
2-Factor model	887.74	76	11.68	.960	.952	.960
1-Factor model w/ method effects	516.43	41	12.60	.142	.977	.948
2-Factor bifactor model	165.54	63	2.63	.053	.993	.995

Note. $N = 574$. The best fitting model is indicated in bold. χ^2/df = a ratio of chi square, divided by the degrees of freedom (see Kline, 2005); RMSEA = root-mean-square error of approximation; TLI = Tucker-Lewis Index (Tucker & Lewis, 1973); CFI = comparative fit index.

.01. Because the test for scalar invariance also includes constraining factor loadings to be equal across groups, this test simultaneously provided support for metric invariance across gender. Based on these results supporting measurement invariance all the way to the scalar invariance level, we then were able to proceed with comparing mean scores across gender.

Internal Consistency and Gender Differences

The overall Cronbach's alpha for the 14-item CDS scale was an acceptable .87. The Cronbach's alpha estimates for the two subscales of the CDS were Disgust Avoidance = .93 and Disgust Affect = .64. Contrary to predictions, there were no gender differences in disgust sensitivity for the total score ($t[572] = .38, p > .05$; girls: $M = 21.02, SD = 7.80$; boys: $M = 21.28, SD = 8.44$), Disgust Avoidance ($t[572] = .28, p > .05$; girls: $M = 1.33, SD = 0.59$; boys: $M = 1.35, SD = 0.69$), or Disgust Affect ($t[572] = .57, p > .05$; girls: $M = 1.08, SD = 0.51$; boys: $M = 1.10, SD = 0.51$).

Study 2 Discussion

Although the CDS contained items that were intended to sample distinct disgust domains identified in previous research (Olatunji et al., 2007), EFA of CDS items in Study 1 indicated that a bifactor model that allows for measurement of the two identified factors as well as a general factor provided the best fit to the data. CFA in Study 2 confirmed that the bifactor model was a better fit to the data above and beyond competing models including a model that controlled for method effects due to reverse-worded items. This suggests that although the Disgust Affect factor contains all reverse-worded items (see Table 1), there is a true factor apart from the method effects. Examination of the bifactor model showed that one item did not load onto the general factor. Removal of this item resulted in a final 14-item scale. Additionally, two items (Item 4 and 18) loaded negatively onto the general disgust factor. However, the items were retained given that they loaded positively onto the intended subfactors. Further, removal of these negative pathways resulted in poor model fit. The negative loadings of the two items on the general disgust factor, despite having positive loadings on the intended subfactors, is unexpected and may reflect a methodological artifact akin to statistical suppression. Further research is needed to explore the bifactor analytical method in more detail to delineate the origins of such effects. Consistent with Study 1, the findings of Study 2 also suggest that the CDS total score has good reliability among youth. However, it is not yet clear the extent to which the CDS items correlated with measures of fear and anxiety in children. Accordingly, Study 3 was conducted to examine the convergent validity of the scale and its two factors in relation to measures of anxiety and fear. Further, to assess whether the correlation between disgust, anxiety, and the convergent measures was a true correlation and not simply an artifact of negative affect, a measure of depression was also included to examine discriminant validity.

Study 3 Method

Participants

Participants included 50 children who were recruited through an online participant recruitment system, the Vanderbilt University

Kennedy Center Study Finder. The children ranged in age from 5 to 12 years ($M = 7.62, SD = 2.18$; 52% boys) and were mostly Caucasian (70%).

Measures

The Screen for Child Anxiety Related Emotional Disorders—Revised (SCARED-R; Muris, Merckelbach, Schmidt, & Mayer, 1998) is a 66-item measure of seven domains of anxiety disorder symptoms. Severity of symptoms are rated using a 0- to 2-point rating scale, with 0 = *not true or hardly ever true*, 1 = *sometimes true*, and 2 = *true or often true*. The present study excluded the Separation Anxiety and School Phobia and Traumatic Stress Disorder scales because of poor factor loading of the construct leaving 50 items. The SCARED-R demonstrated good internal consistency ($\alpha = .93$) in the current sample.

The Fear Survey Schedule for Children—Revised (FSSC-R; Ollendick, 1983) is an 80-item measure designed to assess common childhood fears. The present study excluded the “Fear of Failure and Criticism” factor because of length as well as its failure to map onto our construct of interest. Therefore, this study used a modified FSSC-R that consists of 51 items rated on a 3-point Likert-type scale (*None, Some, or A lot*). The FSSC-R demonstrated good internal consistency ($\alpha = .90$) in the current sample.

The Child Depression Inventory (CDI; Kovacs, 1992) is a 27-item measure designed to assess depression symptoms in children. Each item has three statements, and the child is asked to select the one answer that best describes his or her feelings over the past two weeks. The CDI demonstrated acceptable internal consistency ($\alpha = .79$) in the current study.

Procedure

Participants were recruited through Vanderbilt University's online recruitment system and provided verbal informed consent over the phone after hearing the details of the study. Participants were then e-mailed a unique link to the study survey that also included a hard copy of the assent which required participants to agree prior to being presented with the survey. The study survey data were collected and managed using REDCap electronic data capture, a secure, web-based application designed to support data capture for research studies, hosted at Vanderbilt University (Harris et al., 2009). Each questionnaire measure was presented individually and the option to skip items was included for each question (i.e., “I would prefer not to answer this question”). Parents were told they could help children if necessary and a question was included at the end of each questionnaire to determine whether parental assistance was used. Examination of this question revealed that children needed help slightly over half the time depending on the questionnaire. Specifically, parents helped children 50% of the time for the CDS, 48% of the time for the SCARED-R, 60% of the time for the FSSC-R, and 74% of the time for the CDI. Further, parental assistance on the questionnaires was significantly correlated with age ($r_s = .39-.54, p_s < .001$), such that younger children needed more help from parents to complete the questionnaires.

Study 3 Results

Internal Consistency and Gender Differences

Descriptive statistics and correlations for each measure in Study 3 are presented in Table 4. The Cronbach's alpha estimate for the CDS 14-item total score was adequate at .76 with an average interitem correlation of .20. Further, the two CDS subscales also demonstrated acceptable internal consistency estimates: Disgust Avoidance, $\alpha = .73$, and Disgust Affect, $\alpha = .60$. Girls ($n = 24$) reported significantly greater disgust sensitivity compared with boys ($n = 26$) for the total CDS score ($t[48] = 4.44, p < .001, d = 1.26$; girls: $M = 18.29, SD = 3.84$, boys: $M = 12.88, SD = 4.68$). Girls also reported greater Disgust Avoidance ($t[48] = 3.27, p = .002, d = .93$; girls: $M = 1.50, SD = 0.28$, boys: $M = 1.18, SD = 0.40$) and Disgust Affect ($t[48] = 3.95, p < .001, d = 1.11$; girls: $M = 0.96, SD = 0.43$, boys: $M = 0.47, SD = 0.45$) than boys.

Convergent and Divergent Validity

Correlational analyses were used to examine the convergent validity of the CDS scores in relation to measures of fear (FSSC-R) and anxiety (SCARED-R). As demonstrated in Table 4, the CDS total score was significantly related to anxiety-related disorder symptoms and common childhood fears. The pattern of correlations that emerged between the CDS total score and anxiety measures also supported the convergent and discriminant validity of the scale. For example, the correlation between the CDS scores and specific phobia ($r = .55$) was larger than the correlation between the CDS and social anxiety ($r = .25$); this difference however did not reach statistical significance ($z = 1.70, p = .09$), likely due to a small sample size. Table 4 also shows that the pattern of correlations with measures of fear and anxiety were

generally weaker with the Disgust Affect factor consisting of reversed scored items. In contrast, discriminant validity of the CDS scores was demonstrated by a nonsignificant relationship with depression on the CDI (see Table 4).

Study 3 Discussion

Results of Study 3 provide initial support of the convergent and discriminant validity of the CDS. The CDS total score was significantly correlated with measures of fear and anxiety but not depression. The findings of Study 3 also suggest that disgust sensitivity in children may be more strongly associated with some anxiety disorders (specific phobia) than others (social anxiety disorder). The absence of an association with depression suggests that the CDS measures a distinct vulnerability that is not simply an artifact of negative affect. Having provided initial evidence in support of the convergent and divergent validity of the CDS, Study 4 was conducted to examine the known-groups validity of scores on the CDS. Examination of the extent to which the CDS yields different scores for groups known to vary on disgust sensitivity would speak well to the validity of the scale as well as its clinical utility. Consistent with previous research (Muris et al., 1999; Muris, van der Heiden, & Rassin, 2008), it was predicted that children with a diagnosis of a specific phobia would report significantly greater disgust sensitivity than an age-, gender-, and ethnicity-matched nonclinical sample (NCS).

Study 4 Method

Participants

Forty-three children with a primary diagnosis of specific phobia (42% female; 93% Caucasian; $M_{\text{age}} = 9.16$ years, $SD = 1.90$ years) were recruited for the present study. The clinical sample was recruited in the United States from contacts with mental health treatment clinics, pediatricians, family practice physicians, and school systems, as well as newspaper articles and TV and radio advertisements. The following specific phobias were included: being alone/Darkness (46.5%), storms (16.3%), dogs (14.0%), costumes (7.0%), loud noises (4.7%), bees/insects (4.7%), spiders (4.7%), and BII (2.3%). An NCS of 43 children who were matched for age, gender, and ethnicity (42% female; 93% Caucasian; $M_{\text{age}} = 9.16$ years, $SD = 1.90$ years) was also recruited.

Measures

The Anxiety Disorders Interview Schedule for DSM-IV—Child and Parent Versions (ADIS-IV-C/P; Silverman & Albano, 1996). The ADIS-IV-C (child version) and ADIS-IV-P (parent version) are reliable and well validated semistructured diagnostic interviews designed to facilitate diagnoses of anxiety and mood disorders and other disorders in children and adolescents between 6 and 17 years old.

Procedure

Children in the clinical sample completed the CDS as part of the diagnostic intake prior to undergoing treatment through the Child Study Center at Virginia Polytechnic Institute and State University

Table 4
Associations Between the Child Disgust Scale (CDS) and Measures of Convergent and Discriminant Validity

Measure	CDS total	Disgust avoidance	Disgust affect	<i>M</i>	<i>SD</i>
Convergent measures					
FSSC-R	0.47**	0.49**	0.26	98.60	15.89
Unknown	0.48**	0.49**	0.29*	32.86	6.51
Injury/Small Animal	0.50**	0.54**	0.26	30.78	5.76
Danger/Death	0.19	0.22	0.04	27.44	5.38
Medical Fears	0.23	0.17	0.26	7.52	2.08
SCARED-R	0.49**	0.44**	0.39*	86.48	18.51
Panic Disorder	0.33*	0.28*	0.29*	19.92	5.34
GAD	0.34*	0.34*	0.22	15.02	5.09
Social Phobia	0.25	0.28*	0.13	8.88	2.04
Specific Phobia	0.55**	0.44**	0.50**	25.42	6.03
OCD	0.40**	0.37**	0.29*	16.50	4.61
Discriminant measure					
CDI	-0.21	-0.24	-0.07	52.86	3.77
<i>M</i>	15.48	1.34	0.71		
<i>SD</i>	05.06	0.38	0.50		

Note. FSSC-R = Fear Survey Schedule—Revised; GAD = generalized anxiety disorder; OCD = obsessive-compulsive disorder; SCARED-R = Screen for Child Anxiety Related Emotional Disorders—Revised; CDI = Child Depression Inventory.

* $p < .05$. ** $p < .01$.

in Blacksburg, VA. During the child's intake interview, parents completed several questionnaires about themselves and their families and a structured diagnostic interview regarding their child. Presence of a specific phobia was determined during a clinical consensus meeting, based solely on the child and parent diagnostic interviews. Based on independent raters, kappa for this sample was .91. The matched NCS of children completed the CDS in a classroom setting using the procedure and sample from Study 1 and Study 2.

Study 4 Results

Internal Consistency and Gender Differences

The Cronbach's alpha estimate for the CDS 14-item total score was questionable at .64 and an average interitem correlation at .13. However, both the Disgust Avoidance subscale and the Disgust Affect subscale demonstrated acceptable internal consistency ($\alpha = .88$ and $.75$, respectively). Regarding differences between girls ($n = 18$) and boys ($n = 25$) among the specific phobia group, Disgust Affect was greater among girls ($M = 0.92$, $SD = 0.48$) compared with boys ($M = 0.55$, $SD = 0.58$), $t(41) = 2.00$, $p < .05$, $d = .70$. However, Disgust Avoidance did not significantly differ between girls ($M = 1.56$, $SD = 0.43$) and boys ($M = 1.50$, $SD = 0.35$), $t(41) = .54$, $p = .60$, $d = .15$). General disgust sensitivity also did not significantly differ between girls ($M = 18.67$, $SD = 4.59$) and boys ($M = 16.24$, $SD = 4.56$), $t(41) = 1.66$, $p = .11$, $d = .53$. There were no gender differences found among the NCS group based on the CDS total score or the two subscales ($ps > .05$).

Group Differences

A univariate analysis of variance was conducted to examine whether the specific phobia group and NCS group differed on the CDS total score. As shown in Table 5, significant group differences were found on the CDS total score, with the specific phobia group reporting significantly greater disgust sensitivity compared with the NCS group, $F(1, 84) = 11.42$, $p = .001$, $\eta_p^2 = .12$. Table 5 also shows significant group differences for Disgust Avoidance with the specific phobia group scoring higher than the NCS group, $F(1, 84) = 11.08$, $p = .001$, $\eta_p^2 = .12$. However, no significant

group differences in Disgust Affect was observed, $F(1, 84) = .09$, $p = .77$, $\eta_p^2 = .001$.

Study 4 Discussion

Results of Study 4 support the known-group's validity of the CDS among a sample of children with specific phobia, compared with a nonclinically referred, community sample matched for relevant demographic characteristics. Consistent with previous research (Muris et al., 1999), children with a diagnosis of specific phobia reported greater disgust sensitivity compared with nonclinical controls. The findings of Study 4 also showed that the two groups did not significantly differ on the Disgust Affect factor, which consists of the reverse-worded items.

General Discussion

The present investigation examined the psychometric properties of the newly developed CDS in four independent samples of children ages 5 to 13 years. Across these four samples, the CDS total score demonstrated an average Cronbach's alpha of .77. Although the internal reliability of the CDS total score was acceptable in the present investigation (Nunnally, 1978), it was not exceptional. However, this may not be a limitation of the measure per se but a byproduct of assessing complex emotional processes in children. Research has shown that young children often have difficulty labeling and communicating emotion, and they may not be able to provide fully accurate reports of their own emotions until 8–9 years of age, when their understanding of emotions becomes based on internal mental cues (see Schniering, Hudson, & Rapee, 2000, for review). Therefore, some young children may experience difficulty recognizing and describing the features of disgust, which may contribute to a less internally consistent assessment of their sensitivity to disgust. This highlights the importance of a multimodal approach to the assessment of disgust sensitivity in children that is sensitive to developmental considerations. Although development of the CDS represents an initial step in facilitating such research, subsequent revision of the scale items and response options may prove valuable in improving the reliability of the measure.

Although an EFA in Study 1 revealed a two-factor bifactor solution which includes a general factor that accounts for the covariance among all the CDS items and two independent subfactors corresponding to Disgust Avoidance and Disgust Affect best represented the data, there were concerns given that the Disgust Affect factor contained only reverse-worded items. In an initial attempt to determine whether the reverse-worded items emerged as an artifact of a method effect or if they represent a distinct and theoretically meaningful factor, a CFA of the two-factor bifactor model of the CDS was examined in Study 2. Results showed that the bifactor model fit the data significantly better than a one-factor model that controlled for method effects due to reverse-scoring of the negatively worded items. This suggests that the CDS measures a general disgust sensitivity response as well as two distinct subfactors (Disgust Affect and Disgust Avoidance) that cannot be explained by method effects alone. A bifactor model of the CDS scores has important implications for the study of disgust sensitivity and its association with anxiety disorder symptoms among youth. For example, the bifactor structure of the scale offers

Table 5
Study 4: Descriptive Statistics and Group Differences on the Child Disgust Scale (CDS) Total Score and Subscales Among Children With a Specific Phobia Diagnosis and Nonanxious Controls

CDS	Specific phobia participants ($n = 43$)		Nonanxious control participants ($n = 43$)		$F(1, 84)$	η_p^2
	M	SD	M	SD		
CDS total score	17.26	4.83	13.86	4.48	11.42**	.12
Disgust Avoidance	1.52	0.38	1.15	0.62	11.08**	.12
Disgust Affect	0.71	0.57	0.74	0.57	0.09	.00

Note. η_p^2 represents partial eta squared.

** $p < .01$.

researchers the flexibility to examine how general and specific disgust responses simultaneously confer risk for anxiety disorders. Similarly, examination of profile scores on general and specific disgust facets of the scale may offer greater incremental predictive validity than either the total score on the CDS or the two subfactors alone.

The present findings suggest that the CDS consists of a general disgust sensitivity factor in addition to two distinct component factors. However, the present findings did reveal lower reliability estimates for the Disgust Affect factor relative to the Disgust Avoidance factor. The relatively lower reliability of the Disgust Affect factor may be partially due to reverse-worded items that can be more difficult for children to understand. The lower reliability may also be partially accounted for by fewer items for the Disgust Affect factor relative to the Disgust Avoidance factor. Despite the lower reliability, the present findings revealed that the Disgust Affect subscale scores provided a higher degree of reliable variance than the Disgust Avoidance subscale scores after accounting for the general disgust sensitivity factor. This suggests that the Disgust Affect factor may offer some incremental utility above and beyond the general disgust sensitivity factor of the CDS. Assuming that the Disgust Avoidance and Disgust Affect factors of the CDS represent distinct processes, the lower reliability estimates for the Disgust Affect factor may reflect a more complex process to assess in children. Early in development, children learn through intrafamilial modeling with facial expressions and social referencing (Stevenson et al., 2010) to avoid disgust elicitors (e.g., “don’t put that in your mouth!”). Behavioral avoidance of disgusting stimuli (Disgust Avoidance) may be what is actively taught to young children through such learning processes which may facilitate a more internally consistent response than affective labeling of disgust responses (Disgust Affect) that require cognitive resources that are underdeveloped in young children (Rozin & Fallon, 1987).

Factor analyses of measures of disgust sensitivity in adults has consistently produced multidimensional solutions, suggesting that disgust sensitivity is not a unitary construct (de Jong & Merckelbach, 1998; Olatunji et al., 2007). However, the present findings suggest that emergence of distinct disgust domains may be moderated by development. The CDS was modeled after the DS–R, the most commonly used measure of disgust sensitivity in adults. The DS–R consists of three disgust domains including Core Disgust, Contamination Disgust, and Animal-Reminder Disgust (Olatunji et al., 2007). However, the present findings suggest that distinct disgust domains of this sort that are thematically driven may not be readily observed early in childhood. That is, older children and adults may be more sensitive to gradations in the *content* of disgust stimuli whereas younger children are less cognizant of such nuances. The acquisition of disgust is thought to develop in stages, starting with basic taste and smell aversions in infancy and early childhood, followed by an understanding of contagion in late childhood, and more complex responses such as sociomoral disgust appearing in later childhood and adolescence (see Sawchuk, 2009, for review). The adaptation of the disgust response to more complex stimuli across development may be a byproduct of increasing cognitive maturity. Therefore, developmental limitations in young children may prevent observation of complex disgust domains found among adult samples. Recent research examining the factor structure of disgust responses, as assessed by the DS–R, in adolescents found three factors corresponding to Contagion,

Mortality, and Contact Disgust (Kim, Ebessutani, Young, & Olatunji, 2013). This finding is consistent with the view that the nature of disgust domains may evolve over development. With increasing cognitive capacity, more complex disgust domains that are characterized by differences in content and contagion potency may be more readily observed. This pattern of findings also highlights the importance of future research examining the factor structure of disgust responses across the developmental continuum. In addition to the psychometric implications, such an approach may inform knowledge on how disgust responses are acquired over time and how they are extended to various domains.

The CDS also demonstrated good convergent and known-groups validity in the present investigation. As predicted, scores on the total CDS were significantly correlated with scores on measures of anxiety and fear. This finding is consistent with previous research among adults (Matchett & Davey, 1991; Mulkens et al., 1996), as well as research using adult measures of disgust sensitivity in children (Muris, van der Heiden, & Rassin, 2008). The present study also found that youth with a diagnosis of a specific phobia reported greater disgust sensitivity on the CDS compared with a nonclinical youth sample. This finding is consistent with prior research implicating disgust sensitivity in the development and maintenance of specific phobias (Matchett & Davey, 1991; Mulkens et al., 1996; Olatunji et al., 2006; Page & Tan, 2009). The findings further suggest that the disgust sensitivity-specific phobia association is readily observed even in young children. Of note is that the Disgust Affect factor did not significantly differentiate youth with a diagnosis of a specific phobia from controls. This finding suggests that the Disgust Affect factor may have limited utility in differentiating those with a wide range of phobias from those that do not. Future research is needed to examine if the CDS and its factors have greater utility in differentiating samples with more homogenous phobias from those that do not have such phobias. The CDS also demonstrated good discriminant validity in the present investigation, as scores on the scale did not significantly correlate with scores on a measure of depression. This finding is consistent with a previous study that found no association between disgust sensitivity and depression symptoms in a sample of adults (Muris et al., 2000). Although there is some evidence that disgust experienced toward the self may confer risk for depression (Overton, Markland, Taggart, Bagshaw, & Simpson, 2008), these findings suggest that disgust experienced toward stimuli in one’s environment may play less of a role in the development of depression.

Findings with adults have consistently shown that women report higher disgust sensitivity compared with men (Davey, 1994; Haidt et al., 1994; Schienle et al., 2003). The present investigation is largely consistent with this notion as girls generally scored higher on the CDS than boys in three samples. However, no gender differences on the CDS were found in two samples. One interpretation of these findings is that gender differences in sensitivity to disgust may be less reliable in young children. That is, the gender stereotyped reinforcement contingencies that may give rise to gender differences in disgust sensitivity may not yet be firmly established until later in development. Given that prior research with adults suggests that gender differences in some anxiety disorder symptoms can be accounted for by gender differences in disgust sensitivity (Olatunji, Arrindell, & Lohr, 2005; Olatunji, Sawchuk, et al., 2005), future research examining the processes, which may include socialization or parental transmission, that may

facilitate heightened disgust sensitivity in young girls may reveal unique opportunities for early intervention.

Although the present series of studies suggests that the CDS has good psychometric properties, several limitations of the current investigation should be considered. First, although the bifactor model was found to be the best fit for the data, inclusion of a relatively large number of reverse-worded items indicates that wording of the items may warrant further consideration. Indeed, young children may have some difficulty understanding reverse-worded items, thereby reducing the consistency of responses. Future research should assess whether these items can be reworded to improve reliability. This would also provide an opportunity to reexamine the factor structure of the scale. Second, the assessment of the reliability of the scale is limited to internal consistency, and examination of the test–retest reliability of the CDS scale scores in future research will prove valuable in determining the stability of the scale. Although this is a psychometric issue, such research may also inform understanding of the nature of disgust sensitivity itself, especially in regard to whether and how disgust sensitivity changes over time and how such change is moderated by development. Third, the CDS was not directly compared with downward versions of disgust sensitivity measures that are commonly used with children in other studies. Such comparisons may prove valuable in providing researchers with some guidance as to which scale to use and when. A likely outcome from such work is that the CDS may be more appropriate for younger children whereas currently available downward measures of disgust sensitivity may be more appropriate for older children. However, this remains an empirical question. Last, the present investigation is also limited by the exclusive use of self-report in examining the convergent and discriminant validity of the CDS. Linking scores on the CDS with disgust responding through other assessment modalities will be an important aim for future research. For example, disgust-related behavioral approach tasks have been used with young children (Steven-son et al., 2010) and performance on such tasks would be predicted to correspond with scores on the CDS. Despite these limitations, the initial studies presented here suggest that the CDS is an age-appropriate measure of disgust sensitivity in children. The CDS is thus likely to provide utility in further illuminating the role of disgust sensitivity in anxiety-related disorders among youth.

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(Appendix follows)

Appendix

Child Disgust Scale

Each sentence below is a statement that might be disgusting. Choose how often you would do what the sentence says by circling: *Always*, *Sometimes*, or *Never*.

1. If I saw my favorite toy in the garbage I would take it out and play with it.	Always	Sometimes	Never
2. If a dog licked my popsicle I would still eat it.	Always	Sometimes	Never
3. I would sit next to someone even if they wore the same underwear all week.	Always	Sometimes	Never
4. I would pick up a worm with my hand.	Always	Sometimes	Never
5. When I see blood I feel dizzy.*	Always	Sometimes	Never
6. I feel gross when I touch raw meat to help cook dinner.*	Always	Sometimes	Never
7. I would touch a sandwich with green mold on it.	Always	Sometimes	Never
8. I would still eat my soup if I saw a hair in it.	Always	Sometimes	Never
9. I would feel gross if I accidentally touched someone's bloody cut.*	Always	Sometimes	Never
10. I would sit next to a sweaty kid at lunch.	Always	Sometimes	Never
11. I would watch a TV show that showed people's guts.	Always	Sometimes	Never
12. I would use the toilet even if there was poop still in it.	Always	Sometimes	Never
13. I would share markers with someone that had touched a dead bird.	Always	Sometimes	Never
14. I won't eat unless I can wash my hands.*	Always	Sometimes	Never
15. I feel sick if I see a dead animal on the side of the road.*	Always	Sometimes	Never
16. I don't like seeing the blood in meat at the grocery store.*	Always	Sometimes	Never
17. I would still drink my juice box even if I saw another kid drink out of it.	Always	Sometimes	Never
18. I feel sick if I see someone throw up.*	Always	Sometimes	Never

Note. Items in bold indicate retained items. * Indicates reverse-scored item.

Received March 3, 2013
Revision received November 13, 2014
Accepted January 9, 2015 ■