The psychometric properties of the Child PTSD Symptom Scale (CPSS) were examined in 2 samples. Sample 1 (N = 185, ages 6–17 years) consisted of children recruited from hospitals after accidental injury, assault, and road traffic trauma, and assessed 6 months posttrauma. Sample 2 (N = 68, ages 6–17 years) comprised treatment-seeking children who had experienced diverse traumas. In both samples psychometric properties were generally good to very good (internal reliability for total CPSS scores = .83 and .90, respectively). The point-biserial correlation of the CPSS with posttraumatic stress disorder (PTSD) diagnosis derived from structured clinical interview was .51, and children diagnosed with PTSD reported significantly higher symptoms than non-PTSD children. The CPSS demonstrated applicability to be used as a diagnostic measure, demonstrating sensitivity of 84% and specificity of 72%. The performance of the CPSS Symptom Severity Scale to accurately identify PTSD at varying cutoffs is reported in both samples, with a score of 16 or above suggested as a revised cutoff.

Keywords: CPSS, posttraumatic stress disorder, test–retest reliability, reliable change index, confirmatory factor analysis

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is acceptable (r ranging between .63 and .85 for the total score and subscales). In terms of validity, the original CPSS scores correlated highly with an established measure of child PTSD symptoms, the Child Posttraumatic Stress Reaction Index (CPTSD–RI; Pynoos et al., 1987), r = .80 (Foa et al., 2001). The CPSS can be used as a continuous measure of symptom severity (summation of Items 1–17 with possible scores ranging from 0 to 51), and in a sample of earthquake victims (N = 75) assessed 2 years posttrauma, a cutoff of 11 or greater was found to have sensitivity of 95% and specificity of 96%. Foa et al. (2001) also reported that the 17 symptom items could be scored dichotomously to generate a DSM–IV consistent diagnosis of PTSD.

While the CPSS has been used in a number of studies since its development, including child PTSD treatment research (e.g., Nixon, Sterk, & Pearce, 2012; Smith et al., 2007), surprisingly there has been only modest evaluation of its psychometric properties. In particular, whether the original recommended cutoff score of 11 to diagnose probable PTSD holds in other trauma types and samples is unknown. The CPSS was originally compared with another self-report measure (CPTSD–RI), and while one study has reported its correlation with clinical diagnosis (Rachamim, Helskog, Foa, Aderka, & Gilboa-Schechtman, 2011), the CPSS has yet to be validated against a structured clinical PTSD interview in terms of sensitivity and specificity metrics. Such a comparison not only has implications for the measure’s validity but could have significant cost-efficiency implications if it can be shown that the CPSS converges with structured interviews that are lengthy to administer and resource intensive (e.g., requiring interviewer training).

Rachamim et al. (2011) reported comparable psychometrics with a Hebrew version of the original measure in a sample of Israeli treatment-seeking youth (N = 156, ages 8–18), 78% of whom had been diagnosed PTSD. While the CPSS showed moderate correlation with PTSD diagnosis (.54), as measured with the Schedule of Affective Disorders and Schizophrenia for School Age Children, Revised for DSM–IV (K–SADS–R; Kaufman et al., 1997), the utility of the original cutoff score was not reported. Kohrt et al. (2011) examined the cross-cultural validity of the CPSS in Nepalese children exposed to war and displacement (N = 162). However, substantial modification to the original measure and validation of the measure against need for treatment rather than PTSD diagnosis per se precludes comparison of the psychometrics of this version.

Based on the consolidation of several data sets obtained from our published and unpublished work, we were in a position to examine the CPSS to answer some of the above questions. Accordingly, we were able to examine the psychometric properties of the CPSS in a sample of single-incident trauma survivors (Sample 1) and to test the sensitivity of different cutoff scores in PTSD-treatment-seeking children (Sample 2).

Method

Participants

Two samples were selected from several available data sets. Sample 1 consisted of children and adolescents (N = 185, age 6–17 years) recruited from the emergency department or pediatric inpatient ward of metropolitan hospitals following single-incident trauma. Approximately half this sample came from Australia and the other half the United Kingdom, with data from the present report representing their 6-month posttrauma assessment (Meiser-Stedman, Smith, Glucksman, Yule, & Dalgleish, 2008; Meiser-Stedman, Yule, Smith, Glucksman, & Dalgleish, 2005; Nixon, Ellis, Nehmy, & Ball, 2010). Children and adolescents who presented for treatment at the researchers’ traumatic stress clinics and were diagnosed with PTSD (N = 68, age 6–17 years) made up Sample 2, and these data were also derived evenly from Australia and the United Kingdom (Nixon et al., 2012; Smith et al., 2007). The published articles accompanying these data sets detail in full inclusion and exclusion criteria; briefly, children had to have been exposed to an event sufficient to cause PTSD, be able to complete assessments in English, and not be suspected of experiencing ongoing trauma (e.g., abuse). Key exclusion criteria included the presence of organic brain damage or learning difficulties, significant loss of consciousness at the time of the trauma, or unstable medication regimen (especially in the case of the treatment sample, Sample 2). The pooling of data from several studies has been effectively used in the child traumatic stress field (see, e.g., Dalgleish et al., 2008; Kassam-Adams et al., 2012; Meiser-Stedman et al., 2009) with such integrative data analysis approaches affording advantages of increased statistical power and sample heterogeneity (Curran & Hussong, 2009; Kassam-Adams et al., 2012). Table 1 shows the demographic and trauma-related characteristics of the sample.

Measures

The psychometric properties of the CPSS were detailed earlier in this article. PTSD diagnosis was established in the Australian sample with the Clinician Administered PTSD Scale for Children (CAPS–CA; Nader et al., 1998) and with the Anxiety Disorders Interview Schedule, Child and Parent Report version (ADIS–C/P; Silverman & Alban, 1996) in the United Kingdom sample. Structured clinical interviews are currently considered the gold standard for child PTSD assessment. While there is limited published information on the psychometric properties of the CAPS–CA, its format is a replication of the adult form, which has strong reliability and validity data (see Weathers, 2004, for review). In unpublished work (Ellis, 2008), internal reliability for the CAPS–CA scores was good (α = .88 for total symptom severity). We found interrater reliability performed on 24 randomly selected interviews across Samples 1 and 2 to be excellent (100% diagnostic agreement; Nixon et al., 2010, 2012). The ADIS–C/P is an established diagnostic tool for child anxiety disorders, including PTSD, and its scores have been shown to possess excellent test–retest reliability (Silverman, Saavedra, & Pina, 2001). Interrater reliability was similarly excellent (100% diagnostic agreement) for 21 interviews in Sample 1 (as reported in Meiser-Stedman et al., 2005). Smith et al. (2007) observed satisfactory interrater reliabil-

1 Consistent with the data pooling approach described earlier, samples were collapsed to increase power of the study, which is especially pertinent when examining sensitivity and specificity psychometrics where base rates of psychopathology may be low. Examination of the psychometric properties of the CPSS when divided by sample country of origin revealed negligible differences, including results where PTSD diagnosis was derived by either the CAPS–CA or ADIS; hence, data from the pooled samples is reported.
ity for PTSD diagnosis across 30 randomly selected interviews (κ = .82; Sample 2). Consistent with a general consensus in the child PTSD literature that current DSM–IV PTSD criteria are overly conservative when applied to children, children in the present study were coded as PTSD positive if they met either full DSM–IV criteria or subthreshold criteria (defined as meeting two of the possible three symptom clusters of re-experiencing, avoidance, and hyperarousal, as well as satisfying impairment criteria).2 All analyses reporting PTSD diagnosis include children with full or subthreshold PTSD.

Procedure

All children and caregivers provided written informed assent/consent at each research site. Children completed self-report questionnaires, with the CPSS indexed to the event that led to their hospital contact (Sample 1) or for which they were seeking treatment (Sample 2). Trained interviewers (clinical psychologists or advanced clinical psychology students) administered either the CAPS–CA or ADIS to derive diagnostic status in both samples.

Results and Discussion

Data were screened for normality and outliers. As expected for Sample 1 (a non-treatment-seeking sample with the majority of children demonstrating good adjustment), CPSS scores tended to be positively skewed. Analyses on transformed data did not alter the pattern of findings, with very little absolute change observed on correlations of interest. Similarly, extreme scores only just reached the conventions used to define outliers (e.g., Tabachnick & Fidell, 2001); thus, adjusting these had little effect. In particular, modifying outliers had no impact on cutoff findings. Data for Sample 2 were normally distributed without significant outliers. Accordingly, raw data are reported throughout.

Severities of PTSD Symptoms and PTSD Status

Table 1 summarizes children’s self-reported PTSD symptom severity on the CPSS. Scores in Sample 1 (e.g., CPSS total score, M = 10.06, SD = 9.80) were only marginally higher than that reported by Foa et al. (2001; M = 7.60, SD = 8.10). Not surprisingly, scores in the treatment-seeking sample (Sample 2) were considerably higher (CPSS total score, M = 27.44, SD = 9.33). Using structured clinical interviews, 10% (n = 19) of children in Sample 1 were diagnosed with PTSD, and there were no significant differences between genders or ethnicities in PTSD rates. All children in Sample 2 had PTSD.

Age and Gender3

Age differences were observed only on the Re-experiencing subscale in Sample 1, where younger children (6–11 years old; 2 Subthreshold or alternative diagnostic criteria used in the child PTSD literature has included the definition adopted in the present study (e.g., Meiser-Stedman et al., 2005): a requirement for at least one symptom each of re-experiencing, avoidance, and hyperarousal (e.g., Kassam-Adams & Winston, 2004; Kenardy, Spence, & Macleod, 2006), or one symptom of re-experiencing, one of avoidance, and two of hyperarousal (Meiser-Stedman et al., 2008; Scheeringa, Wright, Hunt, & Zeanah, 2006). We found prevalence rates were comparable regardless of which criteria were used.

3 At the suggestion of one reviewer, we examined the psychometric properties of the CPSS in Sample 1 for two age groups: children (6- to 11-year-olds) and adolescents (12- to 17-year-olds). As can be seen in the supplemental materials available online, the psychometric properties for the two age groups were remarkably similar and consistent with findings when the sample was considered as a whole. The only minor impact of this age segregation appeared to be in relation to cutoff scores, where some differences were observed relative to when the entire Sample 1 was examined. However, this is more likely to be due small cell sizes for PTSD positive cases (where percentages can be artificially skewed) rather than age effects per se.
M = 2.95, SD = 3.20, n = 82) reported significantly more re-experiencing symptoms than middle-school children (12–14 years old; M = 1.12, SD = 1.94, n = 55, p < .001). In the treatment-seeking sample (Sample 2), older children (15–17 years old, n = 24) reported significantly higher total severity scores (M = 31.53, SD = 8.01) than younger children did (M = 24.31, SD = 8.96, n = 32, p = .003). These older children also reported higher avoidance and impairment scores (avoidance: M = 12.00, SD = 3.86, vs. M = 8.45, SD = 3.80, p < .002; impairment: M = 3.87, SD = 1.73, vs. M = 2.16, SD = 1.93, p < .001). No gender differences were observed on CPSS scores in Sample 1 or Sample 2.

**Internal Reliability**

Internal reliability (Cronbach’s alpha) was very good in Sample 1 (total score, .90; re-experiencing, .84; avoidance, .78; and arousal, .79). Intercorrelations between the total CPSS score and subscales were also good (re-experiencing, .73; avoidance, .89; arousal, .88) and subscale intercorrelations ranged from .58 to .69. Internal reliability was lower but within acceptable limits in Sample 2 (total score, .83; re-experiencing, .78; avoidance, .65; arousal, .67). The total CPSS score correlated highly with each subscale (re-experiencing, .75; avoidance, .87; arousal, .83), and subscale intercorrelations ranged from .42 to .64.

**Test–Retest Reliability**

For a subset of Sample 1 the CPSS had been administered 3 months and 6 months posttrauma (n = 113). On the one hand, knowing the test–retest reliability over such an interval is useful when one might wish to calculate reliability of change indices (Jacobson, Roberts, Berns, & McGlinchey, 1999). For example, this interval matches the length of a treatment intervention, as opposed to the relatively short test–retest interval (i.e., 1–2 weeks) that is typically reported in psychometric studies. On the other hand, it should be recognized that there can be a great deal of change in symptoms in the first 6 months following trauma in children (Le Brocque, Hendrikz, & Kenardy, 2010), and the sample was non-treatment-seeking; thus, one might expect lower stability relative to a short test–retest interval. Consistent with this, test–retest reliability coefficients were good for the total score (although lower than the .84 reported by Foa et al., 2001) and modest for subscales (total score: .75; re-experiencing, .50; avoidance, .62; arousal, .70).

**Convergent Validity**

A point-biserial correlation was calculated between the CPSS total score and PTSD diagnosis in Sample 1, showing a significant relationship (r = .51, p < .001), which was comparable to that reported by Rachamim et al. (2011; r = .54). The CPSS can be scored according to DSM–IV criteria for use as a diagnostic measure. Accordingly, we examined how a diagnosis derived from the CPSS using the full and subthreshold criteria outlined earlier compared against PTSD diagnosis obtained via structured clinical interview. Whether requiring impairment items to be endorsed impacted on this was also examined. Similarly, we tested whether the stringency of the criteria required for a symptom to be considered endorsed affected diagnostic rates. That is, we examined CPSS frequency scores of 0 versus 1–3 (as done by Foa et al., 2001) as well as dichotomizing scores of 0–1 versus 2–3. When we used the 0 versus 1–3 criterion without requiring an impairment item to be endorsed, the sensitivity of the CPSS was high in Sample 1, identifying 95% (n = 18) of the children diagnosed with PTSD by interview. Specificity, however, was low (51%, n = 85). Requiring at least one impairment item to be endorsed improved overall performance, with sensitivity slightly lower (84%, n = 16) but specificity increasing to 72% (n = 118). Adopting a stricter approach to symptom scoring (0–1 vs. 2–3) did not improve performance of the CPSS. For example, this stricter scoring without the requirement of an impairment item to be endorsed resulted in sensitivity of 68% (n = 13) and specificity of 84% (n = 139). Requiring impairment to be endorsed further reduced sensitivity (58%, n = 11), whereas specificity further improved (88%, n = 145).

**Discriminant Validity**

CPSS severity was examined against clinical diagnosis using receiver–operator characteristic (ROC) curves, with the area under the curve (AUC) significant in Sample 1 (AUC = .89, p < .001). ROC analysis was inappropriate for Sample 2 as all children were PTSD positive. Table 2 details the performance of varying cutoff scores for both samples, including the original published cutoff of 11 (Foa et al., 2001). Given that Sample 1 contained a significant subset of children who presented with accidental injury that varied in severity, the performance of the CPSS was also examined in just those who had suffered assault or road traffic trauma.

Foa et al. (2001) reported achieving high sensitivity (95%) and specificity (96%) using a cutoff score of 11. As is apparent from Table 2, such levels were not achieved in the current samples. However the data suggest that a higher cutoff score (e.g., 16) could be used while still obtaining an optimal balance of sensitivity and specificity (sensitivity between .84 and .93 across Sample 1 and 2; specificity of .83 in Sample 1). It should also be remembered that Foa et al. compared the CPSS against another self-report measure, whereas the present data included diagnostic status derived from structured clinical interviews.

Severity scores on the CPSS clearly discriminated between those diagnosed with PTSD and those who were non-PTSD. As summarized in Table 3, children in Sample 1 with PTSD had significantly higher scores on all subscales and the total CPSS score than non-PTSD children.

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4 The same reviewer suggested conducting factor analysis for Sample 1. A comprehensive report on the factor analytic structure of the CPSS has been previously published (see Kassam-Adams, Marsac, & Cirilli, 2010) in a similar, albeit larger, sample, and Sample 1 was modest for the purposes of factor analysis. However, confirmatory factor analysis was conducted. As reported in the supplemental materials (see Table S5), none of the models tested reached adequate fit criteria. However the pattern of fit indices across tested models were similar to those observed by Kassam-Adams et al. (2010; e.g., the four-factor models commonly labeled as “numbing” and “dysphoria” in the literature were significantly better than the DSM–IV factor model). This suggested the most likely explanation of less than adequate fit was due to sample size limitations (see supplemental materials for further details).
### Functional Impairment

Table 1 details the descriptive data for severity of impairment for both samples. Internal reliability for both samples was acceptable (.80 for Sample 1; .75 for Sample 2). Test–retest reliability over a 3-month interval for a subset of participants in Sample 1 (n = 113) was acceptable (r = .60, p < .001).

As expected, in Sample 1 total impairment scores were significantly correlated with total CPSS severity (r = .58, p < .001), and each of the CPSS subscales (re-experiencing, r = .30, p < .001; avoidance, r = .51, p < .001; and arousal, r = .57, p < .001). Similar patterns were observed in Sample 2 with impairment scores correlated with total CPSS severity (r = .58, p < .001), and each of the CPSS subscales (re-experiencing, r = .26, p = .04; avoidance, r = .61, p < .001; and arousal, r = .54, p < .001).

Those diagnosed with PTSD reported significantly more impairment than children without PTSD (Table 3). There were no age or gender differences for impairment scores in Sample 1 or Sample 2.

### General Comments

The aim of the present article was to examine some of the psychometric properties of the CPSS, which was developed over 10 years ago. The present data show that the CPSS continues to perform well as a measure of self-reported PTSD symptoms. A strength of the current data is its comparison against diagnostic status obtained via structured clinical interview. In addition to CPSS severity being correlated with diagnostic status and children with PTSD demonstrating significantly higher CPSS scores than non-PTSD children, the utility of the CPSS as a quasidiagnostic instrument was also illustrated. This has significant practical implications, suggesting that when time or resource issues preclude formal diagnostic assessment with lengthy clinical interviews, the CPSS could be used as a proxy.

Most of the psychometric properties observed in the original development report of the CPSS were replicated in the present study. Some slight differences were observed. For example, test–retest correlations in the present data were lower; however, this is likely to reflect the difference in sample and methods. Foa et al. (2001) used a 1- to 2-week test interval, 2 years posttrauma. Children in Sample 1 in our study were assessed over a 3-month interval, and at 3- and 6-month posttrauma assessments. There typically is further remission of symptoms in the first 6 months following trauma in children, with relatively little change from 6 months to 2 years (Le Brocque et al., 2010), which likely accounts for the reduced stability observed in our data. Other differences observed in the current report included lack of gender effects but some age effects, whereas the reverse was observed by Foa et al.

While sensitivity and specificity metrics in the current data were good, they were not as high as those obtained by Foa et al. (2001), especially in Sample 1. As mentioned previously, the present report measured discriminant validity against a structured clinical interview, arguably a more rigorous assessment than the self-report

### Table 2

<table>
<thead>
<tr>
<th>CPSS cutoff (≥)</th>
<th>Sample 1&lt;sup&gt;a&lt;/sup&gt; Sensitivity</th>
<th>Sample 1&lt;sup&gt;b&lt;/sup&gt; Sensitivity</th>
<th>Sample 2&lt;sup&gt;c,d&lt;/sup&gt; Sensitivity</th>
</tr>
</thead>
<tbody>
<tr>
<td>10</td>
<td>89% (17/19)</td>
<td>88% (15/17)</td>
<td>99% (67/68)</td>
</tr>
<tr>
<td>11</td>
<td>89% (17/19)</td>
<td>88% (15/17)</td>
<td>99% (67/68)</td>
</tr>
<tr>
<td>12</td>
<td>84% (16/19)</td>
<td>94% (16/17)</td>
<td>97% (66/68)</td>
</tr>
<tr>
<td>13</td>
<td>84% (16/19)</td>
<td>94% (16/17)</td>
<td>97% (66/68)</td>
</tr>
<tr>
<td>14</td>
<td>84% (16/19)</td>
<td>94% (16/17)</td>
<td>96% (65/68)</td>
</tr>
<tr>
<td>15</td>
<td>84% (16/19)</td>
<td>80% (132/166)</td>
<td>85% (58/68)</td>
</tr>
<tr>
<td>16</td>
<td>84% (16/19)</td>
<td>83% (137/166)</td>
<td>85% (58/68)</td>
</tr>
<tr>
<td>17</td>
<td>84% (16/19)</td>
<td>83% (138/166)</td>
<td>85% (58/68)</td>
</tr>
<tr>
<td>18</td>
<td>79% (15/19)</td>
<td>94% (16/17)</td>
<td>81% (99/122)</td>
</tr>
<tr>
<td>19</td>
<td>79% (15/19)</td>
<td>87% (145/166)</td>
<td>88% (60/68)</td>
</tr>
<tr>
<td>20</td>
<td>79% (15/19)</td>
<td>89% (148/166)</td>
<td>93% (63/68)</td>
</tr>
</tbody>
</table>

Note. CPSS = Child PTSD Symptom Scale; PTSD = posttraumatic stress disorder.  
<sup>a</sup> Full hospital sample (N = 185; 19 PTSD, 166 non-PTSD).  
<sup>b</sup> Hospital sample (road traffic trauma and assault only, N = 139; 17 PTSD, 122 non-PTSD).  
<sup>c</sup> Treatment sample (N = 68, all PTSD).  
<sup>d</sup> Specificity not applicable for treatment sample, as all participants had PTSD.

### Table 3

<table>
<thead>
<tr>
<th>CPSS scale</th>
<th>PTSD</th>
<th>Non-PTSD</th>
<th>F</th>
<th>Cohen’s d</th>
</tr>
</thead>
<tbody>
<tr>
<td>Re-experiencing</td>
<td>5.47 (4.38)</td>
<td>1.78 (2.32)</td>
<td>34.35&lt;sup&gt;∗&lt;/sup&gt;</td>
<td>1.05</td>
</tr>
<tr>
<td>Avoidance</td>
<td>9.32 (5.21)</td>
<td>2.93 (3.53)</td>
<td>50.05&lt;sup&gt;∗&lt;/sup&gt;</td>
<td>1.44</td>
</tr>
<tr>
<td>Arousal</td>
<td>8.58 (3.31)</td>
<td>3.18 (3.23)</td>
<td>47.42&lt;sup&gt;∗&lt;/sup&gt;</td>
<td>1.65</td>
</tr>
<tr>
<td>Total</td>
<td>24.73 (10.40)</td>
<td>8.38 (8.23)</td>
<td>63.54&lt;sup&gt;∗&lt;/sup&gt;</td>
<td>1.74</td>
</tr>
<tr>
<td>CPSS impairment</td>
<td>2.44 (1.80)</td>
<td>0.86 (1.48)</td>
<td>18.56&lt;sup&gt;∗&lt;/sup&gt;</td>
<td>0.96</td>
</tr>
</tbody>
</table>

Note. PTSD, n = 19; non-PTSD, n = 166. CPSS = Child PTSD Symptom Scale; PTSD = posttraumatic stress disorder.  
<sup>∗</sup> p < .001.
measure used by Foa et al. Important sample differences were also apparent. Sample 1 in particular contained fewer Caucasian children than the original normative sample (61% vs. 89%) and more boys (58% vs. 41%), and both Sample 1 and 2 included children ages 6–7 as well as 16–17, compared with children in the 8- to 15-year-old age range. Foa et al.’s sample also comprised children who had experienced the same event (earthquake) 2 years previously, whereas in the present study mixed types of traumatic experiences constituted the stressful events, typically 6 months prior to assessment. It is likely that some of these differences contributed to the lower sensitivity and specificity observed in the present data. Of course, the clinician’s goal for using the CPSS would guide the selection of an appropriate cutoff. If it is used for screening purposes where there are the resources to manage potential false positives, a score of 16 seems sufficient, and we suggest that in most cases clinicians would err on the side of identifying children with probable PTSD. While near perfect sensitivity (99%) was observed in our treatment seeking sample using Foa et al.’s cutoff of 11, we lacked specificity data, and it is worth noting that sensitivity for a cutoff of 16 was very good (93%).

Taking into account these factors, the fact that we used a diagnostic clinical interview for comparison purposes, and our clinical experience with using the CPSS in a variety of contexts, we argue that a cutoff of 11 is probably too low; thus, we recommend considering a score of at least 16 as the cutoff. Considering all the above issues, the discrepancy in the other achieved psychometrics (internal reliability, intercorrelations between subscales, etc.) between the current data and the original development study is relatively small, underscoring the robustness of the CPSS as a measure of PTSD.

Several limitations are acknowledged. We did not have consistent measures across data sets for other psychopathology (depression, general anxiety); thus, we could not test the divergent validity of the CPSS. Prior studies of the CPSS (e.g., Foa et al., 2001; Rachamim et al., 2011) have found good support for its divergent validity, and prior publication of the current data sets (e.g., Meiser-Stedman et al., 2005, 2008; Nixon et al., 2010, 2012; Smith et al., 2007) has documented levels of anxiety and depression in the samples. Data were collected prior to proposed DSM-5 criteria for PTSD; thus, we were not in a position to examine how these additional symptoms might impact on performance of the CPSS. Prior research has examined child–caregiver agreement on symptoms and diagnosis, especially in younger children (Meiser-Stedman et al., 2008). Given that the CPSS was completed by children, it remains to be seen whether optimal identification of child PTSD would be assisted by incorporating parent or caregiver responses. Despite these limitations, the current study indicates that the CPSS is a useful measure of child PTSD. It is short, easy for children to understand, easily scored, and can be used as a reasonable proxy for clinical diagnosis.

References


