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Original Citation

Sharratt, Kathryn, Foca, Liliana, Gavriluta, Cristina, Jones, Adele and Asiminei, Romeo (2014) Dimensionality and Construct Validity of the Romanian Self-Report Strengths and Difficulties Questionnaire (SDQ). Romanian Journal of Applied Psychology, 16 (2). pp. 33-39. ISSN 1454-8062

Romanian Journal of Applied Psychology
2014, Vol. 16, No. 2, 33-39

Dimensionality and Construct Validity of the Romanian Self-Report Strengths and Difficulties Questionnaire (SDQ)

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Received 02.12.2014; Accepted 22.12.2014
Available online 26.01.2015

The Strengths and Difficulties Questionnaire (SDQ) is one of the most widely utilised measures of behavioural and emotional difficulties among children and young people. Previous research has raised concerns about the psychometric properties of the measure, particularly the internal consistency of the CP and PP subscales. Confirmatory factor analysis (CFA) has generally supported a five-factor solution that is consistent with Goodman's (1997) original conceptualisation of the SDQ, but alternative factor structures have been validated including models with internalising and externalising factors, and a total difficulties factor. This was the first study to examine the dimensionality, construct validity and internal consistency of the Romanian self-report version of the SDQ. Based on data collected from 1,086 school children aged 9-17 years old, six alternative factor models were specified and tested using conventional CFA techniques and a confirmatory bifactor modelling approach. The five-factor model provided a better fit for the data than alternative factor structures, but was still unacceptable according to a range of overall model fit indices and individual item loadings. Model fit statistics for the five-factor solution were also notably poorer among boys than girls. Internal consistency was low for the CP, H and PP subscales among the total sample and girls only; and for the EP, CP, H and PP subscales among boys only. Results are discussed in terms of the appropriate interpretation of the Romanian SDQ.

Keywords: Strengths and Difficulties Questionnaire (SDQ); children; confirmatory factor analysis; bifactor modelling; internal consistency

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Acknowledgement. This paper was based on data collected as part of the COPING Project funded by the European Union Seventh Framework Programme (grant agreement number: 241988). The authors would like to thank Catalin Luca (Asociatia Alternative Sociale, Iasi, Romania) and Silviu Brebulet (University of Bucharest, Bucharest, Romania) for their assistance with data collection and preparation.

Introduction

Mental health disorders are largely unreported and therefore undiagnosed, but estimates suggest that almost 30,000 children and adolescents in Romania are registered with a condition (Open Society Institute, 2005). A history of large-scale institutionalisation of children and inter-country adoption (Vorria et al., 2006), disparities in social conditions between Roma and non-Roma children (Lee et al., 2014) and increases in parental migration (Brebuleţ, Gulei, Luca & Foca, 2012) are among several factors that have adversely impacted children's mental health. The social implications of mental health problems are inestimable but it is clear that the economic costs of treating disorders place a huge burden on health and social services (Smit et al., 2006), not least in Romania whose mental health infrastructure is being rebuilt following the Ceausescu regime (Gleason et al., 2011). Given that symptoms often have their onset in adolescence, and can present themselves up to four years before diagnosis, early intervention can decrease the risk of impairment across the lifespan (O'Connell, Boat & Warner, 2009). In any country, the development of effective prevention strategies is partly reliant on the identification of instruments that provide a reliable and valid assessment of the behavioural and emotional difficulties experienced by children and young people.

The Strengths and Difficulties Questionnaire (SDQ; Goodman, 1997) has been administered to various clinical and non-clinical samples of children and adolescents, including in Romania (e.g. Becker, Wolfgang, Hasselhorn, Banaschewski & Rothenberger, 2004; Brebuleţ et al., 2012; Pez, Boyd, Chritophe & Kovess-Masfety, 2013). The SDQ comprises of 25 items designed to load onto five separate subscales measuring Emotional Problems (EP), Peer Problems (PP), Conduct Problems (CP), Hyperactivity (H) and Prosocial Behaviour (PS). The first four subscales can be summed to provide a Total Difficulties Score (TDS), whereas the PS subscale assesses strengths and is considered independent of the difficulties subscales. The inclusion of strengths in the scale not only makes it more appealing to respondents but also provides a more balanced impression of children's mental health.

The SDQ is highly regarded for its clinical utility in distinguishing children and young people most vulnerable to mental health problems (e.g. Becker et al., 2004; Bourdon, Goodman, Rae, Simpson & Koretz, 2005; Muris, Meesters & van den Berg, 2003). However, evidence concerning the psychometric properties of the SDQ is less convincing. Several studies have reported low internal consistencies for the CP and PP subscales (e.g. Di Riso et al., 2010; Du, Kou & Coghill, 2008; Goodman, 2001), suggesting that they might be measuring more heterogeneous constructs than intended.

Given that a hypothesised factor structure already exists for the SDQ, the most appropriate technique for testing the construct validity and dimensionality is confirmatory factor analysis (CFA). Unlike exploratory techniques, CFA has the added advantage of providing a robust indication of whether items load onto the anticipated latent constructs in the absence of measurement error (Bollen, 1989). CFA has confirmed the existence of five separate subscales in various samples of children across Europe (e.g. Giannakopoulos et al., 2009; Goodman, 2001; Rønning, Handegaard, Sourander & Mørch, 2004). However, some of these studies have used model fit statistics that are vulnerable to distortions of the normal distribution or have relied on the correlation of error terms to achieve acceptable model fit. In the general population, SDQ scores are often positively skewed to reflect the low occurrence of behavioural and emotional difficulties, whereas the opposite is true of clinical samples (see Bourdon et al., 2005). As demonstrated by Boduszek, Hyland, Dinghra and Mallet (2013) the correlation of error terms should be avoided as this would imply the presence of an additional unspecified latent variable and can also over-complicate the interpretation of models.

Exploratory studies observed a tendency for the positively worded items to load onto the same factor suggesting the possibility of a "positive construal" or method effect (e.g. Du et al., 2008; Niclasenet al., 2012). The application of CFA has produced equivocal results regarding the presence of a method effect; whilst some studies have found that the inclusion of a factor for the positively phrased items significantly enhances the fit of the five-factor model (e.g. Dickey & Blumberg, 2004), others have observed little improvement (e.g. Van Roy, Veenstra & Clench-Aas, 2008).

Alternative theoretically justifiable model structures have been validated elsewhere in the literature. Consistent with Goodman's (1997) original conceptualisation of the SDQ, Yao and colleagues found that a hierarchical model with a total difficulties factor underlying the four problem subscales provided acceptable model fit among adolescents in China (Yao et al., 2009). In the only study to apply bifactor modelling techniques to the SDQ, Kobór, Takács and Urbán (2013) found support for a total difficulties factor in addition to five grouping factors akin to the original subscales (see Hyland, Boduszek, Dhingra, Shevlin & Egan, 2014 for an explanation of bifactor modelling).

Other investigators have tested a model where the EP and PP subscales are replaced with an internalising problems factor, and the CP and H subscales with an externalising problems factor, whilst still retaining a separate PS factor. This three-factor model provided acceptable model fit for children in Italy (Di Riso et al., 2010) and in Belgium was superior to the five-factor model (Van Leeuwen, Meerschaert, Bosmans, De Medts & Braet, 2006). In the UK, Goodman, Lamping & Ploubidis (2010) provided little support for replacing the subscales with internalizing and externalizing factors. Instead a hierarchical model with higher-order internalising and externalising factors (and a separate PS factor) achieved acceptable model fit indices for the self-report and informant versions of the scale.

In summary, CFA has most often supported a five-factor solution to the SDQ that is consistent with the intended subscales, but alternative model structures have been validated including those with a total difficulties factor and internalising and externalising factors. Contradictory evidence concerning the appropriate factor structure and low internal consistency scores for certain subscales (most notably CP and PP) have continued to raise concerns about the efficacy of the SDQ. The aim of the present study was to provide the first examination of the dimensionality, construct validity and internal consistency of the Romanian self-report version of the SDQ.

Method

Participants

A total of 1,086 children aged 9 to 17 completed the self-report version of the SDQ. The County Centres for Resources in Educational Assistance facilitated access to children attending school in four counties in Romania (Iasi, Botosani, Vaslui, and Bacau). A convenience sampling strategy was adopted, and the SDQ was administered by teachers during classes. School counsellors were available to provide support during questionnaire completion, and a short debriefing took place after SDQ administration. The sample included slightly more girls than boys (57.3% and 42.7% respectively) with a mean age of 13.14 (SD = 2.41). Less than 1% of the data was missing in a non-random fashion and therefore these cases were omitted from the analysis.

Measure

The self-report version of the SDQ comprises of 25 items which are responded to on a 3-point ordinal scale (0 = not true; 1 = somewhat true; 2 = certainly true). Item scores can be summed to provide scores on five subscales measuring Emotional Problems (EP), Peer Problems (PP), Conduct Problems (CP), Hyperactivity (H) and Prosocial Behaviour (PS) as well as a Total Difficulties Score (TDS). The Romanian translation of the scale used in this study and the scoring procedures can be found at: http://sdqinfo.org.

Analysis

The dimensionality of the SDQ was investigated using confirmatory factor analytic (CFA) techniques with weighted least squares means and variance adjusted (WLSMV) estimation in Mplus version 6.0 (Muthen & Muthen, 1998–2010). The WLSMV statistic is regarded as the most appropriate estimator for ordinal-level data, especially when the scale has fewer than five response options (see Moshagen & Musch, 2014). Six alternative model conceptualisations were specified and tested including (i) a 25-item unidimensional model; (ii) a three-factor model consisting of internalising, externalising and prosocial factors; (iii) a five-factor model reflecting the original subscales; (iv) a bifactor model consisting of a general total difficulties factor and five grouping factors analogous to the original subscales; v) a hierarchical model with a higher-order total difficulties factor and a separate PS factor; and (vi) a hierarchical model with higher-order internalising and externalising factors and a separate PS factor. In the bifactor model items were allowed to load onto the general factor as well the grouping factors, and the grouping factors were restricted to be uncorrelated with each other and the general factor (see Reise, Moore, & Haviland, 2010; Reise, Morizot, & Hays, 2007). In all models, measurement error terms remained uncorrelated as suggested in previous research (Boduszek et al, 2013).

Overall model fit was assessed using a range of goodness-of-fit statistics and the appropriateness of the model parameters. The chi-square ($\chi 2$) statistic assessed the sample and implied covariance matrix; a good fitting model is indicated by a non-significant result. The chi-square statistic is, however, strongly associated with sample size, and as such good models tend to be over-rejected. The Comparative Fit Index (CFI; Bentler, 1990) and the Tucker Lewis Index (TLI; Tucker & Lewis, 1973) are measures of how much better the model fits the data compared to a baseline model where all variables are uncorrelated. For these indices values above 0.9 are considered acceptable (Bentler, 1990; Hu & Bentler, 1999). The standardized root mean-square residual (SRMR; Joreskog & Sorbom, 1981) and the root mean-square error of approximation (RMSEA; Steiger, 1990) are also presented; values less than .08 are considered acceptable for these indices. The Akaike Information Criterion (AIC; Akaike, 1974) was used to compare the alternative models, with the smaller value indicating the best fitting model.

Results

Descriptive statistics, including Cronbach's alpha are presented in Table 1. Internal consistency was low for the CP, H and PP subscales within the total sample, girls only and boys only. Internal consistency was also low for the EP subscale among boys only.

Table 1: Descriptive statistics and internal consistency for SDQ subscales

	M	SD	Min	Max	α
Children (total)					
Emotional Problems	4.266	2.498	0	10	.629
Conduct Problems	2.751	1.961	0	10	.474
Hyperactivity	3.519	2.169	0	10	.557
Peer Problems	2.842	1.874	0	10	.380
Prosocial Behaviour	8.077	1.968	1	10	.684
Total Difficulties Score	13.378	5.863	0	35	.730
Girls					
Emotional Problems	4.783	2.511	0	10	.631
Conduct Problems	2.588	1.915	0	10	.489
Hyperactivity	3.445	2.193	0	10	.587
Peer Problems	2.643	1.828	0	9	.404
Prosocial Behaviour	8.368	1.831	1	10	.674
Total Difficulties Score	13.460	5.862	0	31	.741
Boys					
Emotional Problems	3.573	2.309	0	10	.575
Conduct Problems	2.968	2.003	0	9	.457
Hyperactivity	3.619	2.133	0	10	.514
Peer Problems	3.108	1.903	0	9	.343
Prosocial Behaviour	7.688	2.076	1	10	.674
Total Difficulties Score	13.267	5.870	1	35	.725

Independent sample t-tests indicated a significant difference between girls and boys on the EP (t[1084]= 8.127, $p < .001$, Mdiff = 1.210, 95%CI = .918/1.502, Cohen's d = .501), CP (t[1084]= -3.165, $p < .05$, Mdiff = -.379, 95%CI = -.614/-.616, Cohen's d = .193), PP (t[1084]= -4.072, $p < .001$, Mdiff = -.465, 95%CI = -.689/-.241, Cohen's d = .249) and PS (t[1084]= 5.618, $p < 0.001$, Mdiff = .681, 95%CI = .443/.918, Cohen's d = .347) subscales. There was no statistically significant difference on the H subscale (t[1084]= -1.302, $p > .05$) or Total Difficulties Score (t[1084]= .535, $p > .05$).

Correlations between the subscales are presented in Table 2. As expected, all four difficulties subscales displayed moderate-strong positive correlations with each other (ranging from 0.42-0.79) within all groups of children (total sample, girls only, boys only). The PS subscale also displayed moderate-strong negative correlations with the CP, H, PP and PS subscales (ranging from -0.40-0.62) within all groups of children. Contrary to expectations, the PS and EP subscales displayed a very weak correlation that failed to reach statistical significance and in girls was in the opposite direction to that expected (total sample = -0.03; boys only = -0.05; girls only = 0.03).

Table 3 presents the fit indices and comparative fit indices for the six alternative models of the SDQ within the total sample, girls only and boys only. None of the models provided an acceptable fit of the data based on CFI, TLI, RMSEA or SRMR statistics, but the five-factor model displayed comparatively better model fit indices than alternative factor structures among all samples of children. The five-factor model also demonstrated a lower AIC value providing some support for its statistical superiority relative to other models. According to model fit statistics the five-factor model provided a better fit for the data collected from girls than it did boys.

Table 2: Correlation between SDQ factors for total sample, girls only and boys only

SDQ Factors	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15
1. EP(T)	-														
2. CP(T)	.42	-													
3. H(T)	.44	.77	-												
4. PP(T)	.47	.79	.58	-											
5. PS(T)	.03*	-.61	-.51	-.58	-										
6. EP(G)						-									
7. CP(G)						.50	-								
8. H(G)						.48	.79	-							
9. PP(G)						.53	.78	.56	-						
10. PS(G)						-.03*	-.62	-.57	-.58	-					
11. EP(B)											-				
12. CP(B)											.58	-			
13. H(B)											.48	.77	-		
14. PP(B)											.68	.73	.60	-	
15. PS(B)											-.05*	-.53	-.40	-.47	-

Note: EP = Emotional Problems; CP = Conduct Problems; H = Hyperactivity; PP = Peer Problems; PS = Prosocial; (T) = total sample; (G) = girls only; (B) = boys only. All correlations were significant at $p < .001$ except coefficients with * which failed to reach significance at $p < 0.05$.

The adequacy of the five-factor model can also be determined in relation to its parameter estimates. As can be seen in Table 4, all items displayed statistically significant factor loadings on their repsective latent factors (at $p < 0.01$) within the total sample, girls only and boys only. In addition, all factor loadings were in the expected direction but five items (7, 11, 16, 17, 20) failed to exceed 0.4 among all samples of children.

Discussion

The aim of this study was to examine the dimensionality, construct validity and internal consistency of the Romanian self-report version of the SDQ within a sample of school children. Six competing models were specified and tested using conventional CFA techniques and a confirmatory bifactor modelling approach. None of the models provided an acceptable fit for the data based on a range of goodness-of-fit statistics, but the five-factor model performed comparatively better than alternative factor structures. This finding is consistent with Goodman's (1997) original conceptualisation of the instrument and is also the model that is most frequently supported in the empirical literature (e.g. Giannakopoulos et al., 2009; Goodman, 2001; Rønning et al., 2004). Even less support was found for alternative model structures, including those comprising of internalising and externalising factors or a total difficulties factor despite them having been occasionally validated elsewhere in the literature (e.g. Goodman et al., 2010; Yao et al., 2009). In the present study it was also observed that the five-factor model provided a somewhat better fit of the data collected from girls than boys.

Table 3: Fit indices for six alternative models of SDQ

Models	$\chi 2$ (df)	CFI/TLI	SRMR	RMSEA	AIC
Children (total sample)					
1 factor	1949.27 (275)	.54/.50	.08	.08	54996.67
3 factors	1324.47 (272)	.71/.68	.07	.06	54377.86
5 factors	1090.64 (265)	.77/.74	.06	.05	54157.43
Bifactor (5 grouping + 1 general)	3001.67 (261)	.25/.14	.31	.10	56011.06
5 factors + 1 hierarchical with separate PS	1181.30 (270)	.75/.72	.06	.06	54238.69
5 factors + 2 hierarchical with separate PS	1165.93 (268)	.75/.73	.06	.06	54227.32
Girls					
1 factor	1214.52 (275)	.57/.53	.08	.07	30398.43
3 factors	873.45 (272)	.72/.70	.07	.06	30063.37
5 factors	746.86 (265)	.78/.75	.06	.05	29950.79
Bifactor (5 grouping + 1 method)	1954.48 (261)	.22/.10	.39	.10	31166.39
5 factors + 1 hierarchical with separate PS	805.41 (270)	.75/.73	.06	.06	29999.33
5 factors + 2 hierarchical with separate PS	793.15 (268)	.76/.73	.06	.06	29991.06
Boys					
1 factor	1031.25 (275)	.51/.46	.08	.08	24117.04
3 factors	770.31 (272)	.67/.64	.07	.06	23922.10
5 factors	732.77 (265)	.69/.65	.07	.06	23898.56
Bifactor (5 grouping + 1 method)	1365.58 (261)	.28/.17	.26	.10	24539.37
5 factors + 1 hierarchical with separate PS	761.40 (270)	.68/.64	.07	.06	23917.19
5 factors + 2 hierarchical with separate PS	745.01 (268)	.68/.64	.07	.06	23904.79

On further inspection it was apparent that a number of the item loadings were unacceptably low, suggesting that the items might not be a good representation of the intended latent construct. This was particularly true for the PP subscale where three out of the five items failed to reach acceptable values for the total sample and girls only; rising to four items among boys only. Overall, item loadings were comparatively poorer among boys and slightly more items displayed unacceptably low item loadings than among girls (nine compared to seven).

Table 4: Standardized factor loadings for five factor SDQ model

Items	Children	Girls	Boys
Emotional Problems			
1.I get a lot of headaches, stomach-aches or sickness	.408	.414	.402
2.I worry a lot	.523	.519	.439
3.I am often unhappy, down-hearted or tearful	.677	.672	.647
4.I am nervous in new situations, I easily lose confidence	.453	.488	.378
5.I have many fears, I am easily scared	.482	.449	.490
Conduct Problems			
6.I get very angry and often lose my temper	.386	.424	.389
7.I usually do as I am told	.213	.230	.158*
8.I fight a lot. I can make other people do what I want	.453	.430	.481
9.I am often accused of lying or cheating	.519	.556	.464
10.I take things that are not mine from home, school or elsewhere	.461	.391	.511
Hyperactivity			
11.I am restless, I cannot stay still for long	.364	.391	.342
12.I am constantly fidgeting or squirming	.454	.451	.477
13.I am easily distracted, I find it difficult to concentrate	.571	.600	.527
14.I think before I do things	.382	.429	.312
15.I finish the work I am doing, my attention is good	.488	.493	.461
Peer Problems			
16.I am usually on my own, I generally play alone or keep to myself	.322	.324	.393
17.I have one good friend or more	.301	.293	.258
18.Other people my age generally like me	.440	.435	.391
19.Other children or young people pick on me or bully me	.432	.498	.413
20.I get on better with adults than with people my own age	.218	.240	.186
Prosocial Behaviour			
21.I try to be nice to other people, I care about their feelings	.502	.399	.597
22.I usually share with others (food, games, pens etc.)	.523	.536	.480
23.I am helpful is someone is hurt, upset or feeling ill	.596	.581	.608
24.I am kind to younger children	.580	.613	.533
25.I often volunteer to help others (parents, teachers, other children)	.562	.578	.521*

Note: All factor loadings were significant at $p < .001$ except items marked with * which were significant at $p < 0.01$.

Low internal consistency scores further illustrated the problematic nature of the subscales. Three of the subscales had unacceptably low Cronbach's alpha scores among girls (CP, H and PP), and four among boys (EP, CP, H and PP). This might simply reflect the fact that the subscales contain a limited number of items, but combined with poor overall model fit and weak item loadings it might be more likely that the subscales are measuring more disparate constructs than intended. Previous studies have typically revealed low internal consistency scores of the CP and PP subscales (e.g. Di Riso et al., 2010; Goodman, 2001) so the findings of the present study are somewhat more disappointing.

The desire to identify differential psychological outcomes for boys versus girls is evident in the literature surrounding the SDQ, and it is frequently reported that girls score higher on the EP and PP subscales, whereas boys score higher on the CP and H subscales (e.g. Di Riso et al., 2010; Muris et al., 2004). The pattern of results in the present study was slightly different, with girls scoring significantly higher on the EP subscale but boys scoring significantly higher on the CP and PP subscales. Given the aforementioned problems with the PP subscale in particular, it would be inadvisable to regard this as conclusive evidence that boys in Romania display uncharacteristically poorer peer relations than girls.

Overall, the strength and direction of correlations among subscales was comparable among boys and girls. As anticipated, the four difficulties subscales all correlated positively with one another, and the PS subscale correlated negatively with the CP, H and PP subscales. However, the PS subscale displayed little correlation with the EP subscale and was also in an opposite direction among boys and girls (albeit very weak and non-significant). Further qualitative research might offer valuable insights into this unexpected finding.

This was the first study to test the psychometric properties of the Romanian SDQ and therefore provides novel information about the usefulness of the instrument within Romanian samples, but the study is not without limitations. A notable advantage of the SDQ is the availability of parallel versions of the instrument for completion by self-report, parents and teachers, which enables the triangulation of results. However, the present study was not able to extend findings to the informant versions of the instrument or comment on the comparability of self-report and informant versions. Further research including both self-report and informant versions of the scale within Romanian samples might provide further insights into why the self-report version possesses weak psychometric properties. The present study was also unable to comment on the ability of the SDQ to accurately distinguish children at heightened risk of mental health problems.

In conclusion, none of the models tested provided an acceptable fit for the data, but the five-factor model was preferable to alternative factor structures. Problems with the SDQ were further reflected in low item loadings, particularly for the PP subscale, and low internal consistency scores for all but one of the subscales (PS). Findings also suggested that the SDQ performed worse among boys than girls; evidenced by poorer overall model fit indices, item loadings and internal consistency scores. Overall, this implies that in the current format the Romanian translation of the SDQ provides a poor representation of the emotional and behavioural difficulties experienced by children, especially among boys. Further qualitative research to inform the potential rephrasing or reorganisation of items in order to better capture the intended latent constructs (and therefore improve construct validity and internal consistency) would be beneficial. It is important to note that usefulness of the SDQ as a screening tool does not depend entirely on its factor structure (see for example, Becker et al., 2004), and it would be premature to dismiss the SDQ as unhelpful for clinical purposes in Romanian based solely on these findings.

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