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THE TWENTY-ITEM TORONTO ALEXITHYMIA SCALE—I. ITEM SELECTION AND CROSS-VALIDATION OF THE FACTOR STRUCTURE

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Abstract—Addressing shortcomings of the self-report Toronto Alexithymia Scale (TAS), two studies were conducted to reconstruct the item domain of the scale. The first study resulted in the development of a new twenty-item version of the scale—the TAS-20. The TAS-20 demonstrated good internal consistency and test–retest reliability, and a three-factor structure theoretically congruent with the alexithymia construct. The stability and replicability of this three-factor structure were demonstrated in the second study with both clinical and nonclinical populations by the use of confirmatory factor analysis.

INTRODUCTION

FORMULATED during the early 1970s [1, 2], the alexithymia construct is generating increasing interest as a possible personality risk factor for a variety of psychiatric and psychosomatic disorders [3–8]. Investigations of alexithymia have been hampered, however, by problems with the various instruments designed to measure the construct [9, 10]. For example, several widely used self-report measures of alexithymia (viz., the Schalling Sifneos Personality Scale, the Revised Schalling Sifneos Personality Scale, and the MMPI Alexithymia Scale) have been shown to lack adequate reliability and/or validity [9–12].

Recognizing that these self-report alexithymia scales were developed without concern for construct validation and with virtually no consideration of contemporary standards of test construction [9, 10, 13, 14], we attempted to develop a new self-report measure of the construct using a combined empirical and rational method of scale construction. In our initial studies, we developed the twenty-six item Toronto Alexithymia Scale (TAS) [15], which demonstrated good internal consistency and test–retest reliability, and a four-factor structure theoretically congruent with the alexithymia construct—(F1) difficulty identifying and distinguishing between feelings and bodily sensations; (F2) difficulty describing feelings; (F3) reduced daydreaming; and (F4) externally-oriented thinking. Convergent, discriminant, and criterion validity of the TAS were also demonstrated [16–18].

Though the psychometric properties of the TAS are an improvement over those of other self-report measures of alexithymia [19], item development and scale validation are generally considered ongoing processes that must continue beyond the initial development of a scale [13, 20, 21]. In the course of further evaluating the

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TAS, several shortcomings of the scale became apparent that we have outlined elsewhere [22] and that prompted us to re-examine the compositional structure of the scale. For example, although the four-factor structure of the TAS was found to be replicable across clinical and nonclinical samples [23], the first two factors were highly correlated and had several items that cross-loaded significantly on both factors [15], thus questioning the independence of these two factors. In addition, on theoretical grounds, we expected that all of the TAS items and factors would be moderately and positively correlated; however, the items comprising the daydreaming factor had low magnitude corrected item-total correlations with the full TAS; this factor correlated negatively with the first factor (F1) [15]. Further, using causal modelling procedures, Haviland *et al.* [5] found negative correlations between the daydreaming factor and the factors assessing affect awareness and externally-oriented thinking, suggesting that the items assessing daydreaming have little theoretical coherency with the other facets of the alexithymia construct. Given that an externally-oriented cognitive style reflects, at least in part, an impoverished inner fantasy life, and that these facets of the alexithymia construct were linked by French psychosomaticists in the concept of *la pensée opératoire* [24, 25], we expected a higher correlation between F3 and F4. The experiences of other investigators using the TAS led to speculations that: (a) the self-reporting of daydreaming activity may be confounded by a social desirability response bias; and (b) the assessment of daydreaming alone may not capture adequately the capacity for imaginal activities.

With these observations in mind, we generated a new pool of items (the twenty-six items of the TAS plus seventeen new items that we wrote) and attempted to develop an improved and revised version of the TAS. Our initial work led to the elimination of all items assessing imaginal activity and to the construction of a twenty-three-item version of the scale—the TAS-R [22]. In contrast to the four-factor structure of the TAS, factor analysis of the TAS-R yielded a two-factor solution. Factor 1 comprised items assessing both the ability to distinguish between feelings and bodily sensations associated with emotional arousal and the ability to describe feelings to others; Factor 2 comprised items assessing externally-oriented thinking. However, subsequent testing of the factor structure of the TAS-R using confirmatory factor analyses indicated that a two-factor structure was not a good representation of the data, and that a three-factor structure was a much better match to the data. Given this finding, we decided to attempt a new revision of the TAS.

AIMS

The aims of this research were: (1) to extract a new set of items from the item pool used to create the TAS-R; (2) to cross-validate the factor structure of the revised scale with both clinical and non-clinical samples; and (3) to assess the convergent, discriminant, and concurrent validity of the revised scale. This first paper covers the first two aims. A second paper will describe the evaluation of the three types of validity.

STUDY 1: ITEM SELECTION AND SCALE REDEVELOPMENT

Method

Subjects. Subjects were 965 first- and second-year undergraduate students attending a large Canadian University. All students participated as volunteers. The sample (hereafter referred to as the derivation

sample) consisted of 389 males and 576 females with a mean age of 21.8 yr ($SD = 5.6$). A separate sample of 72 students (24 males, 48 females) was recruited at a later date to determine test–retest reliability of the revised scale. The mean age for this sample was 20.8 yr ($SD = 3.1$).

Procedure. In order to examine and explicate the relationships between the existing factors of the TAS, an additional seventeen items were written and added to the original twenty-six items. Seven of these new items were written to reflect the imaginal processing component of the alexithymia construct; two of these seven items specifically addressed daydreaming and five of them assessed other kinds of imaginal activity. Of the remaining ten new items, five items had content relevant to externally-oriented thinking and five items had content related to communicating feelings to others.

Four subscales were derived on the basis of previous factor analysis of the original twenty-six items [15] as well as rational–theoretical considerations and the content relevancy of the seventeen newly written items. In total, four hypothesized facets (i.e. content domains) of the alexithymia construct (identifying feelings, communicating feelings, imaginal processing, externally-oriented thinking) were represented by a pool of forty-three items each presented in a five-point Likert rating format. The relationship of social desirability with each of the forty-three items was examined by correlating these items with the Marlowe–Crowne Social Desirability Inventory (SDI) [26]. Any item that correlated ≥ 0.20 was deleted. The items were also examined for homogeneity of content with respect to their specific facet of the alexithymia construct. For this phase of the analysis, items were retained if they revealed adequate corrected item–total correlations (≥ 0.20) with the other items assessing the same facet of the construct and adequate corrected item–total correlations (≥ 0.20) with the remaining items on the other content domains. Finally, the items that remained following these two procedures were intercorrelated and subjected to a series of factor analyses (principal axis factoring, varimax rotation). Items were retained as meaningful to the construct if they loaded significantly (≥ 0.35) on one and only one of the rotated factors.

Results

Eight of the forty-three items had correlations ≥ 0.20 with the SDI; all of these were imaginal processing items. Further, analysis of the homogeneity of the imaginal processing items indicated that the seven items assessing daydreaming activity (five were from the original TAS, two were newly written items) clustered together (Cronbach's $\alpha = 0.76$, mean corrected item–total correlation = 0.47). However, the inclusion of the five new items assessing other types of imaginal activity decreased the overall alpha coefficient of the imaginal processing content scale to 0.69, and reduced the mean corrected item–total correlation to less than 0.20. Examination of the corrected item–total correlations with all forty-three items indicated that eight of the twelve imaginal processing items had corrected item–total correlations < 0.20 . In sum, only three of the twelve imaginal processing items met the pre-established statistical requirements. Given the failure to retain a sufficient number of items to measure this facet of the construct reliably and to show evidence of theoretical coherency with the other facets, we decided to discard this content domain. The corrected item–total correlations for each of the thirty-one items that comprised the three remaining facets were recalculated. Seven of these thirty-one items were subsequently eliminated because they had corrected item–facet or item–total correlations < 0.20 .

The remaining twenty-four items were then intercorrelated and the resulting matrix was subjected to principal axis factoring with iterations, squared multiple correlations in the diagonal.* Correspondent with the three homogeneous content domains identified and constructed in the previous analyses, three factors were

*Prior to principal axis factoring, the matrix was assessed for psychometric adequacy using Bartlett's test of sphericity [27], the Kaiser–Meyer–Olkin measure of sampling adequacy (MSA) [28], and inspection of the off-diagonal elements of the anti-image co-variants matrix [29], as recommended by Dziuban and Shirkey [30]. (Continued overleaf.)

rotated to a varimax solution. These three factors accounted for 27.9% of the total variance and each factor comprised sets of items that were interpretable and content relevant. In addition, the eigenvalue ≥ 1.00 criterion [28, 31] indicated that a three-factor solution was suitable for rotation, as only the first three unrotated factors were greater than unity. The scree test [32] also indicated a clear break between the third and fourth factors. In addition, all factors beyond the third produced trivial factors (i.e., factors with only one uniquely loading item) (see Gorsuch [33]). Thus, three factors were rotated to solution.

Examination of the factor loadings for the three-factor solution revealed that four items failed to load significantly (≥ 0.35) on any one of the three obtained factors; these four items were eliminated. The remaining twenty items, hereafter referred to as the twenty-item Toronto Alexithymia Scale (TAS-20), were subsequently factor analyzed (principal axis factoring, three factors rotated to a varimax solution). The results are presented in Table I.

TABLE I.—RESULTS OF THE EXPLORATORY FACTOR ANALYSIS OF THE TAS-20

Items	Factor 1	Factor 2	Factor 3
<i>Difficulty Identifying Feelings</i>			
1. I am often confused about what emotion I am feeling.	0.574	0.266	0.009
3. I have physical sensations that even doctors don't understand.	0.404	0.047	0.055
6. When I am upset, I don't know if I am sad, frightened, or angry.	0.592	0.187	0.051
7. I am often puzzled by sensations in my body.	0.570	0.066	0.068
9. I have feelings that I can't quite identify.	0.576	0.272	-0.022
13. I don't know what's going on inside me.	0.584	0.304	0.105
14. I often don't know why I am angry.	0.573	0.167	0.133
<i>Difficulty Describing Feelings</i>			
2. It is difficult for me to find the right words for my feelings.	0.281	0.618	0.103
4. I am able to describe my feelings easily.	0.149	0.641	0.142
11. I find it hard to describe how I feel about people.	0.341	0.568	0.223
12. People tell me to describe my feelings more.	0.261	0.470	0.078
17. It is difficult for me to reveal my innermost feelings, even to close friends.	0.137	0.479	0.154

(Continued).

All three measures of psychometric adequacy suggested that the correlation matrix was suitable for factor analysis: Bartlett's test of sphericity indicated that the items were interdependent [$\chi^2 = 4919.9$, $p < 0.001$]; the Kaiser-Meyer-Olkin measure of sampling adequacy was well above the 0.50 minimally accepted level (MSA = 0.86), indicating that the items belonged together psychometrically; and only 50 (9.1%) of the off-diagonal elements of the anti-image covariants matrix were greater than 0.09, signifying that the matrix of covariances of the individual items approach a diagonal.

TABLE I.—*continued*

Item	Factor 1	Factor 2	Factor 3
<i>Externally-Oriented Thinking</i>			
5. I prefer to analyze problems rather than just describe them.	-0.017	0.051	0.366
8. I prefer to just let things happen rather than to understand why they turned out that way.	0.151	0.094	0.352
10. Being in touch with emotions is essential.	0.069	0.024	0.474
15. I prefer talking to people about their daily activities rather than their feelings.	-0.017	0.148	0.438
16. I prefer to watch "light" entertainment shows rather than psychological dramas.	0.012	0.065	0.368
18. I can feel close to someone, even in moments of silence.	0.063	0.025	0.456
19. I find examination of my feelings useful in solving personal problems.	0.074	0.064	0.613
20. Looking for hidden meanings in movies or plays distracts from their enjoyment.	0.055	0.106	0.463
Eigenvalue	4.64	2.20	1.36
% Total Variance	12.60	9.63	8.75

Note: Items 4, 5, 10, 18, and 19 are negatively keyed.

The three-factor solution accounted for 31% of the total variance. Factor 1, which accounted for 12.60% of the total variance and 40.6% of the common variance, is comprised of items assessing the capacity to identify feelings and to distinguish between feelings and the bodily sensations of emotional arousal. Factor 2, accounting for 9.63% of the total variance and 31.1% of the common variance, reflects the inability to communicate feelings to other people. Factor 3, which accounted for 8.75% of the total variance and 28.2% of the common variance, is comprised entirely of items assessing externally-oriented thinking. The correlations among the unit-weighted subscales were 0.51 for F1 and F2; 0.18 for F1 and F3, and 0.29 for F2 and F3; all correlations were significant at the $p < 0.01$ level.

The TAS-20 demonstrated acceptable internal consistency (Cronbach's $\alpha = 0.81$) with the derivation sample, as did each of the three factors ($F1 = 0.78$, $F2 = 0.75$, $F3 = 0.66$). There was a small but statistically significant difference between the mean TAS-20 score for men (mean = 51.14, $SD = 10.40$) and the mean score for women (mean = 48.99, $SD = 11.48$) [$t(964) = 2.96$, $p < 0.01$]. In addition, there was a low magnitude correlation between TAS-20 scores and age ($r = -0.13$, $p < 0.01$). The test-retest reliability of the TAS-20 with the separate sample of 72 students, who were requested to complete the scale at the beginning of class on two occasions 3 wk apart was 0.77 ($p < 0.01$). Examination of the endorsement pattern frequency for each of the twenty items revealed no evidence of skewness for any of the items. Five of the items are negatively-keyed.

STUDY 2: CROSS-VALIDATION OF THE THREE-FACTOR STRUCTURE

Method

Subjects. This study used separate undergraduate university student and psychiatric out-patient samples. The student sample consisted of 401 subjects (159 males, 242 females) with a mean age of 21.1 yr ($sd = 4.2$). The psychiatric out-patient sample consisted of 218 patients (94 males, 124 females) referred to the psychiatric out-patient department of a large metropolitan teaching hospital. This was a diagnostically heterogeneous sample of patients referred by general practitioners and medical specialists associated with the hospital. Preliminary admission intake psychiatric diagnoses were given according to DSM-III-R criteria—15.6% were diagnosed as anxiety disorder, 15.6% somatoform disorder, 11.1% personality disorder, 10.6% dysthymia, 10.1% major depression or bipolar disorder, 8.7% adjustment disorder, 6.4% schizophrenia or other psychotic disorders, 2.3% substance use disorder, 1.8% eating disorder, and 17.9% were undiagnosed due to incomplete assessments. None of the patients were impaired to the degree that they required hospitalization. The mean age of the psychiatric sample was 35.2 yr ($sd = 11.5$).

Procedure. The student subjects completed the forty-three-item pilot scale in a classroom setting; the psychiatric patients were administered this scale as part of the general out-patient admission procedure.

To test the replicability of the theoretical model represented by the three factor structure of the TAS-20, the correlation matrices for the student and psychiatric out-patient samples were analyzed with LISREL 7.16 [34]. The model that we constructed was based on the factor loadings reported in Table 1. Each item was considered to be a measure of only a single latent factor as determined by a factor loading ≥ 0.35 derived from the exploratory factor analysis of the derivation sample. Given the hypothesized association among the three facets of the construct, an oblique model was tested. Following the recommendation of Cole [35], the goodness-of-fit was evaluated using four criteria: chi-square goodness-of-fit, the goodness-of-fit (GFI) [36], the adjusted goodness-of-fit (AGFI) [36], and the root mean-square residual (RMS). Multiple criteria were used as each of these indices has different strengths and weaknesses in assessing the goodness-of-fit between the hypothetical model and the actual data [35, 37]. Although the chi-square test of goodness-of-fit is not a good indicator of fit when sample sizes are large, typically producing a significant χ^2 statistic with larger samples [37, 38], we retained this criterion because it still provides meaningful information (e.g. ratio of degrees of freedom to chi-square), and because the chi-square difference can be used to identify which models are better able to reproduce the data [39]. The following criteria standards [38] were used to indicate the goodness-of-fit of the model to the data: for the chi-square test, a nonsignificant χ^2 (a significant chi-square indicates that a significant amount of covariance remains unexplained); $GFI > 0.85$; $AGFI > 0.80$; $RMS < 0.10$.

Results

The parameter estimates for each of the items for both samples are presented in Table II. For the student sample, the chi-square goodness-of-fit was significant [χ^2 (167, $N = 401$) = 502.85, $p < 0.001$]. However, the GFI (0.886), AGFI (0.856) and RMS (0.069) all met the criteria standards, thus signifying adequacy of the fit. A similar pattern of results was found with the psychiatric out-patient sample—the chi-square goodness-of-fit was significant [χ^2 (167, $N = 218$) = 358.07, $p < 0.001$], but the GFI (0.864), AGFI (0.829) and RMS (0.070) met the criteria standards.

The parameter estimates of the relationships among the three factors for the student and psychiatric samples are presented in Table III. With the exception of the estimate between F1 and F3 for the student sample, all estimates are statistically significant. Given these significant parameter estimates, we sought to examine whether the twenty items were best represented as a unidimensional construct (one factor solution) or as a construct comprised of three interrelated facets (three factor solution). These models can be compared by testing the difference in the chi-squares, in the present instance the one-factor model versus the three-factor model. Results indicated that the three factor solution was better able to reproduce the data than the unidimensional model for both the university student sample [χ^2 difference = 451.49, ($df = 3$), $p < 0.001$] and the psychiatric out-patient sample [χ^2 difference = 163.98, ($df = 3$), $p < 0.001$].

TABLE II.—PARAMETER ESTIMATES FROM THE RESULTS OF THE CONFIRMATORY FACTOR ANALYSIS OF THE TAS-20 FOR THE STUDENT AND PSYCHIATRIC OUT-PATIENT SAMPLES

Item	University students			Psychiatric out-patients		
	F1	F2	F3	F1	F2	F3
<i>Difficulty Identifying Feelings</i>						
1	0.55*	0.00	0.00	0.65*	0.00	0.00
3	0.43*	0.00	0.00	0.36*	0.00	0.00
6	0.60*	0.00	0.00	0.65*	0.00	0.00
7	0.58*	0.00	0.00	0.52*	0.00	0.00
9	0.70*	0.00	0.00	0.73*	0.00	0.00
13	0.68*	0.00	0.00	0.71*	0.00	0.00
14	0.60*	0.00	0.00	0.70*	0.00	0.00
<i>Difficulty Describing Feelings</i>						
2	0.00	0.72*	0.00	0.00	0.68*	0.00
4	0.00	0.71*	0.00	0.00	0.70*	0.00
11	0.00	0.54*	0.00	0.00	0.47*	0.00
12	0.00	0.51*	0.00	0.00	0.58*	0.00
17	0.00	0.72*	0.00	0.00	0.68*	0.00
<i>Externally-Oriented Thinking</i>						
5	0.00	0.00	0.45*	0.00	0.00	0.45*
8	0.00	0.00	0.28*	0.00	0.00	0.42*
10	0.00	0.00	0.62*	0.00	0.00	0.35*
15	0.00	0.00	0.39*	0.00	0.00	0.46*
16	0.00	0.00	0.30*	0.00	0.00	0.39*
18	0.00	0.00	0.47*	0.00	0.00	0.44*
19	0.00	0.00	0.70*	0.00	0.00	0.58*
20	0.00	0.00	0.39*	0.00	0.00	0.38*

* $p < 0.05$

TABLE III.—PARAMETER ESTIMATES FOR THE RELATIONSHIPS AMONG THE THREE FACTORS OF THE TAS-20 FOR THE UNIVERSITY STUDENT AND PSYCHIATRIC OUT-PATIENT SAMPLES

	Factor 1	Factor 2	Factor 3
Factor 1	—	0.65*	0.10
Factor 2	0.72*	—	0.36*
Factor 3	0.32*	0.50*	—

* $p < 0.05$.

Note: The university student sample is above the diagonal and the psychiatric out-patient sample is below the diagonal.

Factor 1 = difficulty identifying feelings; Factor 2 = difficulty describing feelings; Factor 3 = externally-oriented thinking.

In addition to comparing the unidimensional model against the three factor solution, we also tested whether a three-factor solution produced a better fit than a two factor solution. A two-factor solution model was constructed by collapsing Factor 1 and Factor 2 from the TAS-20 into a single factor, while Factor 3 constituted the second factor. Results indicated that the three factor solution was better able to reproduce the data than the two factor solution model for both the university student sample [χ^2 difference = 155.48, (df = 2), $p < 0.001$] and the psychiatric outpatient sample [χ^2 difference = 68.67, (df = 2), $p < 0.01$].

TABLE IV.—INTERNAL RELIABILITY COEFFICIENTS AND MEAN INTERITEM CORRELATION COEFFICIENTS FOR THE DERIVATION, UNIVERSITY STUDENT, AND PSYCHIATRIC OUT-PATIENT SAMPLES

	Derivation sample	Student sample	Psychiatric sample
<i>Alpha Coefficients</i>			
TAS-20	0.81	0.80	0.83
Factor 1	0.78	0.79	0.81
Factor 2	0.75	0.75	0.75
Factor 3	0.66	0.66	0.64
<i>Mean Interitem Correlations</i>			
TAS-20	0.17	0.16	0.19
Factor 1	0.34	0.35	0.39
Factor 2	0.38	0.38	0.38
Factor 3	0.20	0.20	0.19

Note: Factor 1 = difficulty identifying feelings; Factor 2 = difficulty describing feelings; Factor 3 = externally-oriented thinking.

The internal reliability coefficients and mean interitem correlation coefficients of the TAS-20 for the derivation sample, second student sample, and psychiatric out-patient sample for the full TAS-20 and each of the three factors are displayed in Table IV.

There were no significant gender differences in mean TAS-20 scores in the student sample (mean for males = 47.40, SD = 9.77; mean for females = 47.38, SD = 10.96) and the psychiatric sample (mean for males = 55.27, SD = 12.24, mean for females = 54.45, SD = 13.48) and the scale showed no relationship with age in these two samples.

DISCUSSION

The TAS-20 was developed in an attempt to surmount the shortcomings of the original TAS and to improve upon an earlier revision of the scale (the TAS-R). In Study 1, twenty items were selected from a pool of forty-three items on the basis of theoretical considerations and predetermined statistical criteria. As occurred during item selection for the TAS-R [22], all items directly assessing daydreaming and other imaginal activity were eliminated because of low corrected item-total correlations and/or high correlations with a measure of social desirability. Several other items were eliminated because they failed to meet the criterion for item-factor loadings, which was set higher than it was in the development of both the TAS and TAS-R. Factor analysis of the twenty-item scale yielded three intercorrelated factors that are congruent with the theoretical construct of alexithymia. The first two factors correspond to Factors 1 and 2 of the original TAS; the third factor corresponds to Factor 4 of the TAS. In contrast to the original scale, there is no significant cross-loading of items on the factors of the TAS-20.

The results of the confirmatory factor analyses in Study 2 indicate that the three-factor structure of the TAS-20 is stable and replicable across clinical and nonclinical populations. Factor 1 correlates strongly with Factor 2, which is expected since the ability to communicate feelings is obviously contingent on an ability to recognize

one's own affects. Similarly, Factors 2 and 3 are correlated, as an externally-oriented cognitive style contains little or no reference to a person's inner feelings. Together Factors 2 and 3 appear to reflect the *pensée opératoire* aspect of the alexithymia construct, namely, a cognitive style that shows a preference for the external details of everyday life rather than thought content related to feelings, fantasies, and other aspects of a person's inner experience [2, 24, 25].

Although Factors 1 and 3 of the TAS-20 were unrelated in the student sample in Study 2, they were significantly related to each other in the psychiatric outpatient sample and in the derivation sample in Study 1. Nonetheless, the validity of the three factor solution of the TAS-20 was confirmed by the results of the LISREL analyses comparing one-factor and two-factor models with the three-factor model.

The high coefficient alphas obtained for the full TAS-20 across samples indicate excellent internal consistency of the revised scale. The homogeneity of the TAS-20 was confirmed also by the values of the mean interitem correlation coefficients, which are within the range recommended by Briggs and Cheek [40] for multifactor scales. The results represent an improvement over the estimates of internal consistency and item homogeneity for the original TAS [15, 23, 41]. Although the coefficient alphas for the three factors of the TAS-20 are lower than for the full scale, they still indicate acceptable levels of internal consistency. The mean interitem correlation coefficients for the three factors also indicate that each factor has an optimal level of homogeneity [40]. Like the TAS [15, 41], the TAS-20 demonstrated good test-retest reliability.

In the derivation sample, male university students scored significantly higher on the TAS-20 than female students, and there was a weak relationship between TAS-20 scores and age. Further testing of the scale with diverse samples is needed to determine whether these findings indicate true effects or are explainable by the large sample size.

While the initial evidence of reliability and factorial validity of the TAS-20 is encouraging, other kinds of evidence are needed to show that the scale measures the alexithymia construct in an adequate way. Some of this evidence is presented in the second paper.

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