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Factor Structure of the Automatic Thoughts Questionnaire in a Clinical Sample

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None of the published factor analyses of the Automatic Thoughts Questionnaire— Negative (ATQ-N) [illegible] tions. To address this omission, we examined the factor structure of the ATQ-N in an adult sample (N = 178) seeking treatment for depression. A series of confirmatory factor analyses revealed poor fit indices with all previous models, sug- [illegible] cognitive dimensions than that of nonclinical samples. An exploratory analysis with our clinical sample identified five factors with eigenvalues >1 (demoralization, self-criticism, brooding, amotivation, and interpersonal disappointment) that accounted for 61% of the variance. Of these five factors, only the first two independently accounted for significant variability in levels of depression. Implications of the results and for further use of the ATQ-N within cognitive therapeutic research and practice are discussed.

An array of self-report instruments has been designed to assess cognitive outcome and possible mediating variables (Dozois, Covin, & Brinker, 2003) associated with cognitive therapy of depression (Beck, Rush, Shaw, & Emery, 1979). One of the first, and still one of the most widely used and respected measures (Nezu, Ronan, Meadows, & McClure, 2000), was developed by Hollon and Kendall (1980) to assess the frequency of 30 depressive automatic thoughts. Although their inventory initially was simply known as the Automatic Thoughts Questionnaire, it is now more widely referred to as the Automatic Thoughts Questionnaire – Negative (ATQ-N) to distinguish it from a similarly-formatted instrument (i.e., [illegible] and Wisnicki (1988) to assess the frequency of positive self-statements.

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By now the psychometric properties of the ATQ-N, including its internal (Chioqueta & Stiles, 2004; Deardorff, Hopkins, & Finch, 1984; Ghassemzadeh, Mojtabai, Karamghadiri, & Ebrahimkhani, 2006; Hollon & Kendall, 1980; Kazdin, 1990; Oei & Mukhtar, 2008; Sahin & Sahin, 1992) and temporal consistency (Chioqueta & Stiles, 2004; Ghassemzadeh et al., 2006; Sahin & Sahin, 1992), as well as convergent (Ghassemzadeh et al., 2006; Hollon & Kendall, 1980; Kazdin, 1990; Oei & Mukhtar, 2008; Sahin & Sahin, 1992) and discriminant validity (Chioqueta & Stiles, 2004; Hill, Oei, & Hill, 1989; Hollon & Kendall, 1980; Hollon, Kendall, & Lumry, 1986; Oei & Mukhtar, 2008; Sahin & Sahin, 1992) have been sufficiently documented within both nonclinical (Hollon & Kendall, 1980) as well as clinical samples ([illegible]).

However, with the exception of two possible studies (Kazdin, 1990; Netemeyer et al., 2002) all of the analyses of the factor structure of the ATQ-N to date have been conducted with nonclinical populations (Bryant & Baxter, 1997; Chioqueta & Stiles, 2006; Deardorff et al., 1984; Ghassemzadeh et al., 2006; Hollon & Kendall, 1980; Joseph, 1994; Oei & Mukhtar, 2008; Sahin & Sahin, 1992). Why there have been no factor analyses of the ATQ-N thus far with clinically depressed samples seems a bit unclear insofar as the Dysfunctional Attitudes Scale (Weissman & Beck, 1978) that has been in use as long as the ATQ-N as a frequent companion measure to it (e.g., Dobson & Breiter, 1983; Chioqueta & Stiles, 2004, 2006, 2007; Hill et al., 1989; Sahin & Sahin, 1992) has been factor analyzed with depressed outpatients (Imber et al., 1990).

Limited information about the dimensional properties of the ATQ-N with clinical samples unfortunately creates an inherent interpretational problem. In particular, it is unclear how what is known about the instrument's factor structure with nonclinical samples may generalize to clinical populations. While two previous factor analyses of the ATQ-N were with samples that could be characterized as clinical in nature (Kazdin, 1990; Netemeyer et al., 2002), neither was typical of cognitive therapy clients (i.e., adult outpatients). Kazdin's (1990) sample consisted of children admitted to an inpatient facility and only a minority of them (28%) received either a primary or secondary diagnosis of depression. Although Netemeyer et al. (2002) selected an adult sample, it was limited to self-identified problematic gamblers whose levels of depression were not systematically assessed.

Our major purpose in conducting this study was to address this gap in our understanding of the factor structure of the ATQ-N with a sample more representative of the clinical population to whom it is most often administered. To do so, we first determined the uniqueness of the factor structure of the ATQ-N when administered to a clinical sample by evaluating its fit with models derived from other populations. Poor fit with previously published factorial models would suggest the relevance to also conduct an exploratory analysis with our sample to identify the dimensions of automatic thinking that may be more specific to clinically depressed populations.

METHOD

PARTICIPANTS

[illegible] archival data were available. The remainder were [illegible] with a primary diagnosis of unipolar depression from our departmental training clinic who had been administered the ATQ-N and assessed at intake for level of depression, or participants in another, but as of yet unpublished, research project on treatment of depression. The majority of participants were female (147 or [illegible]) and White ([illegible]) with [illegible] of 40.48 [illegible] (SD = 11.64).

MEASURES

Automatic Thoughts Questionnaire—Negative. Participants were asked to separately [illegible] on a 5-point scale (1 = not at all, 5 = all the time) how frequently, if at all, 30 depressing thoughts occurred over the last week (see Table 3 for a list of the items). Total ATQ-N scores range from 30–150, with higher scores reflective of increased rates of negative self-statements. As already discussed, [illegible] ATQ-N score (M = 93.42, SD = 24.18) was comparable to that reported for other depressed outpatient samples (e.g., Harrell & Ryon, 1983; Hill et al., 1989).

[illegible] Beck, Steer, & Brown, 1996). The BDI-II is a widely used, 21-item, self-report measure of depression. Total scores range from 0–63, with higher scores [illegible] of the BDI-II, including acceptable levels of reliability as well as evidence of its concurrent and discriminant validity, are well established (Beck et al., 1996).

Because the BDI-II had not yet been developed at the time, participants within the clinical sample culled from the earlier depression studies had been assessed with the original Beck Depression Inventory (BDI; Beck, Ward, Mendelson, Mock, & Erbaugh, 1961). The BDI, like its successor, has exhibited strong psychometric properties, with acceptable levels of reliability with clinical populations (Beck, Steer, & Garbin, 1988; Steer, Beck, & Garrison, 1986) as well as concurrent and construct validity (Beck & Steer, 1987). Combining data from the two versions of the BDI seemed appropriate, given that both are psychometrically [illegible] in further analyzing BDI scores, but to simply verify that the [illegible] (M = 29.25, SD = 7.16) fell within the moderate to severe ranges (Beck et al., 1996; Kendall, Hollon, Beck, Hammen, & Ingram, 1987).

RESULTS

We first analyzed our data set for possible gender differences. None were detected for age or ATQ-N scores, although female participants reported higher levels of depression than their male counterparts (M = 29.82 vs. 26.55), $t(176) = 2.54$, p = .02, d = .45. However, because both means fell within the [illegible] range of depression as defined by Beck et al. (1996), [illegible] gender difference to be of any practical importance.

INTERNAL CONSISTENCY

Both alpha (r = .95) and split-half reliability coefficients (r = .94) further substantiated the internal consistency of the ATQ-N documented by other researchers (e.g., Chioqueta & Stiles, 2004; Deardorff et al., 1984; Hollon & Kendall, 1980; Sahin & Sahin, 1992). Corrected item-total correlations for the 30 items of the scale ranged from .42 (item 4, [illegible]) to .77 ([illegible] failure) with a mean of .62. The average item-total correlation and their [illegible] are comparable to those reported in other studies (e.g., Deardorff et al., 1984; Ghassemzadeh et al., 2006; Kazdin, 1990), thereby supporting the integrity of our overall findings.

EVALUATION OF DIMENSIONAL INVARIANCE

The ATQ-N's high level of internal consistency suggested that it was comprised of a limited number of factors. To determine how many, we used an SPSS syntax program developed by O'Connor (2000) for conducting a minimum average partial test (MAP; Velicer, 1976). The MAP computes the residual covariance matrix rescaled to the variance of each variable, giving the partial correlations after each factor is extracted. The average partial r^2 (old criterion) and r^4 (new criterion) are computed after each factor extraction and decrease until all common variance has been extracted and then start increasing. At this point, factor extraction ceases and the number of factors before the increase is used.

The revised MAP test (Velicer, Eaton, & Fava, 2000) revealed five factors with eigenvalues ranging from 12.88 to 1.21 that accounted for 61% of the variance. Only Sahin and Sahin (1992) also reported five factors in their analysis of a Turkish version of the ATQ-N administered to college students, with the other models consisting of a single (Kazdin, 1990) to four factors (Hollon & Kendall, 1980). Ostensibly the varying number of dimensions identified primarily with nonclinical samples strongly suggests that the factor structure of the ATQ-N cannot be generalized to clinical populations. This interpretation, however, must be tempered somewhat by the acknowledgment that other researchers have used varied methods, from principal components (Deardorff et al., 1984; Hollon &

TABLE 1. Goodness-of-Fit Statistics for Various Factor Models of the ATQ-N

				Measures of Relative Fit			
Factor Model	Factors[a]	Items[b]	χ^2	*df*	*NC*[c]	*RMSEA*	*GFI*
Single	1	30	1097.60	405	2.71	.098	.689
Chioqueta & Stiles (2006)	2	30	976.61	375	2.60	.095	.723
≥ .50 Loadings	2	12	152.03	53	2.87	.103	.874
Oei & Mukhtar (2008)	2	17	310.10	118	2.63	.096	.870
≥ .50 Loadings	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]
Deardorff et al. (1984)	3	15	202.37	87	2.33	.087	.874
≥ .50 Loadings	2	14	187.92	74	2.54	.093	.874
Hollon & Kendall (1980)	4	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]
≥ .50 Loadings	2	11	164.90	43	3.83	.127	.850
[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	.757
≥ .50 Loadings	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]	[illegible]

Notes. [a]Values refer to the number of factors; [b]Values refer to the number of items; [c]NC, the normed chi-square, denotes the chi-square value divided by degrees of freedom.

Kendall, 1980; Sahin & Sahin, 1992) to principal axis (Oei & Mukhtar, 2008) in extracting factors. There has been more consensus in the means used to determine the number of extracted factors to retain as all of the previous studies have primarily, if not exclusively, used the Kaiser (1960) rule of eigenvalues [illegible] to retain, although as noted, the MAP identified the same number of factors as the Kaiser rule.

Because at least some of the apparent dimensional variance of the ATQ-N might be attributable to different ways in which the number-of-factors problem has been approached (Gregorich, 2006), we opted to err on the side of caution by conducting a series of confirmatory factor analyses (CFA). Specifically, in doing so [illegible] located during our literature search (Chioqueta & Stiles, 2006; Deardorff et al., 1984; Hollon & Kendall, 1980; Joseph, 1994; Kazdin, 1990; Netemeyer et al., 2002; Oei & Mukhtar, 2008; Sahin & Sahin, 1992). Factor loadings for ATQ-N items were included within each publication with the exception of Kazdin (1990), Joseph (1994), and Netemeyer et al. (2002). [illegible] reported by Kazdin by using all 30 items within the CFA in manner consistent with [illegible] latent construct. [illegible] but both reported that they were no longer available.

Table 1 presents the results from a series of CFAs conducted using the Analysis of Moment Structures (AMOS 5.0) program (Arbuckle, 2003) to test the fit of various factorial models to our clinical sample. In order to provide a more exhaustive evaluation of fit, we tested two iterations for each of the models with the exception of the single factor model. The first iteration listed in Table 1 for each

TABLE 2. Correlation Matrices of ATQ-N Factors

	1.	2.	3.	4.	5.
1. Demoralization	.89				
2. [illegible]	.67	.91			
3. Brooding	.66	.71	.87		
4. Amotivation	.66	.64	.66	.81	
5. Interpersonal Disappointment	.30	.36	.37	.31	.70

Notes. [a]Diagonal entries are alpha coefficients.

of the models included all reported factor loadings with the restriction that only the highest value was included for items that loaded on more than one factor. The second iteration followed this same restrictive rule, but was limited to items with loadings ≥.50 in order to increase the likelihood of determining an adequate fit. As can be seen in Table 1, this resulted in fewer items for all of the models tested as well as a reduction in the number of factors for those original models containing four or more factors (i.e., Hollon & Kendall, 1980; Sahin & Sahin, 1992). The most profound impact of this second round of testing was on the model of Sahin and Sahin (1992), resulting in a loss of four factors and 22 items.

In both iterations, we tested oblique versions of the models even for those that were originally based on an orthogonal rotation (Bouvard et al., 1994; Hollon & Kendall, 1990; Sahin & Sahin, 1992). Table 1 reports three different measures that we used to assess goodness of fit. Because the chi-square (χ^2) statistic may overestimate the lack of model fit (Bollen, 1989) due to its sensitivity to sample size, we divided it by the degrees of freedom (χ^2/df) to calculate a normed chi-square (*NC*) that approaches one as model fit improves. The value of the root mean square error of approximation (*RMSEA*; Joreskog & Sorbom, 1997) is also inversely related to model fit, while the third measure reported in Table 1, the goodness-of-fit index (*GFI*; Joreskog & Sorbom, 1997) increases as the fit of a given model improves. Following the guidelines of Bollen (1989) and Hu and Bentler (1998), we regarded *NC* values of ≤ 3, RMSEA values of ≤ .06, and *GFI* values of ≥.90 as indicative of good model fit.

As can be seen in Table 1, all of the 11 factor models we tested failed to display sufficient goodness-of-fit with our clinical sample. None of the models met criterion on all three of our fit indices and only that of Oei and Mukhtar (2008) met at least two of them. *NC* was met for at least one of the iterations for all of the models except that of Hollon and Kendall (1980). However, none of the models displayed an adequate fit according to *RMSEA* under either iteration, and *GFI* only rose to an acceptable level for two of the models (Oei & Mukhtar, 2008; Sahin & Sahin, 1992) when items were limited to those with loadings ≥.50. As indicated in Table 1, the general impact of this second iteration was to marginally inflate all three fit indices, and in the case of Sahin and Sahin's (1992) model, increase *NC* (5.98) to an unacceptably high level.

TABLE 3. Factor Loadings for Exploratory Factor Analysis With Promax Rotation of ATQ-N

Factor	Loading	Item
1. Demoralization	.84	12. I can't stand this anymore.
	.79	6. I don't think I can go on.
	.69	19. Wish I could just disappear.
	.61	29. It's just not worth it.
	.51	28. My future is bleak.
	.51	25. I feel so helpless.
	.47	15. I wish I were somewhere else.
	[illegible]	[illegible]
	.43	4. No one understands me.[a]
	.40	26. Something has to change.[b]
	.34	24. I'll never make it.[c]
2. Self-Criticism	.97	18. I'm worthless.
	.78	17. I hate myself.
	.71	21. I'm a loser.
	.62	23. I'm a failure.
	.56	2. I'm no good.[d]
	.45	7. I wish I were a better person.
	.40	24. I'll never make it.
	.36	3. Why can't I ever succeed?
3. Brooding	.93	20. What's the matter with me?
	[illegible]	[illegible]
[illegible]	[illegible]	[illegible]
	.46	26. Something has to change.
	.34	10. I'm so disappointed in myself.
[illegible]	.90	30. I can't finish anything.
	.73	16. I can't get things together.
	.72	13. I can't get started.
5. Interpersonal Disappointment	.50	5. I've let people down.
	.48	2. I'm no good.
	.34	4. No one understands me.
	.34	8. I'm so weak.

Notes. [a]Also loads .34 on Factor 5; [b]Also loads .46 on Factor 3; [c]Also loads .40 on Factor 2; [d]Also loads .48 on Factor 5.

EXPLORATORY FACTOR ANALYSIS

The failure to obtain an adequate fit with any of the previous models, underscored the need to conduct a separate EFA of the ATQ-N with our clinical sample to identify its unique dimensional structure. We chose an oblique (i.e., Promax with a Kaiser normalization) rather than orthogonal [illegible] because we had no a priori reason to anticipate that they would be unrelated to each other

TABLE 4. Regression Analysis Predicting Depression From ATQ-N Factor Scores

Factor Score	B	SE B	β	t	p
1. Demoralization	2.34	.79	.31	2.96	.004
2. Self-Criticism	2.79	.81	.38	3.45	.001
3. Brooding	−1.43	.83	−.19	−1.72	.088
4. Amotivation	.34	.79	.04	.43	.668
5. Interpersonal Disappointment	.73	.60	.09	1.22	.224

Notes: $R^2 = .33$, $p < .001$.

based upon the most recent analyses of the ATQ-N (Chioqueta & Stiles, 2006; Joseph, 1994; Netemeyer et al., 2002; Oei & Mukhtar, 2008). As can be seen in Table 2, the five factors [illegible] with each other.

Table 3 presents a summary of the loadings for each factor using ≥.32 as salient for inclusion (Tabachnick & Fidell, 2007). Using this criterion, 90% of the items (27 of 30) loaded on at least one factor and only four items (items 2, 4, 24, and 26) loaded on two. We have referred to Factor 1 as Demoralization. It appears to encompass thoughts that one lacks the wherewithal, stamina, and personal attributes to persevere in meeting [illegible] (e.g., [illegible] I can't stand this anymore). As such, it seems to generally parallel the type of negative thoughts about the world that represent the second component of the cognitive triad (Beck et al., 1979). We have termed Factor 2 Self-Criticism as the eight items that load on it can be seen as comprising negative judgments about the self (e.g., item 17: I hate myself). Such thinking has been linked to suicidal risk (Morrison & O'Connor , 2008) and seems to reflect the first component of Beck's negative cognitive triad. Another pattern of thinking that may itself account for the relationship between self-criticism and suicidality (O'Connor & Noyce, 2008) appears to be represented by Factor 3. We named it Brooding to reflect a type of ruminative thinking (Treynor, Gonzalez, & Nolen-Hoeksema, 2003) focused on why one has fallen short of some comparative standard (e.g., item 14: What's wrong with me?). We have identified Factor 4 as Amotivation as the three items that load on it reflect difficulties in initiating (e.g., item 12: I can't get started) and completing (e.g., item 30: I can't finish anything) goal-directed activities. The fifth and final factor in our view reflects Interpersonal Disappointment as it appears to encompass [illegible] (e.g., item 4: No one understands me).

REGRESSION ANALYSIS

As expected, we obtained a significant correlation between ATQ-N and BDI scores ($r = .53$, $p < .01$). To better understand the relationship between specific dimensions of negative thinking and clinical depression, we conducted a regression analysis using the five factor scores to predict variability in BDI scores. As indicated in Table 4, a significant model [illegible]

[illegible] 33% of variability in BDI scores. However, only the first two factors dealing [illegible] made significant contributions, with the latter accounting for the highest proportion of variance in levels of depression. Particularly surprising was the negative, albeit insignificant, relationship between brooding and depression, especially in light of other research that has found this [illegible] of rumination to be predictive of self-reported depression in at least community samples (Treynor et al., 2003).

DISCUSSION

Our primary purpose in conducting this study was to examine the factor structure of the ATQ-N in a sample more representative of those to whom it is administered in clinical research and practice. Our findings strongly suggest that the dimensions of automatic negative thinking identified thus far in nonclinical populations are not representative of the categories of self-statements endorsed by those seeking treatment of clinical depression. Specifically, the number of common factors (five) identified within our clinical sample differed from all previous factor analyses except that of Sahin and Sahin (1992) with Turkish college students. Moreover, we [illegible] of dimensional invariance for the ATQ-N when CFAs that would maximize goodness-of-fit were conducted. It seems worth reiterating that all of the previous models that we examined for fit were from nonclinical samples, with the exception of Kazdin (1990) that evaluated child inpatients. Unfortunately, [illegible] to obtain item loadings from Netemeyer et al. (2002) of the only previous factor analysis of the ATQ-N conducted with what might be construed as at least an adult quasi-clinical population (i.e., problem gamblers). While it would obviously be desirable for others to further substantiate [illegible] model of the ATQ-N by replication with a similar sample, we believe that our overall results justify the need to be cautious in generalizing [illegible] the ATQ-N with nonclinical populations to those who struggle with clinical depression.

We deemed the five factors of the ATQ-N identified within our clinical sample as demoralization, self-criticism, brooding, amotivation, and interpersonal disappointment. The degree to which the first three dimensions, in particular, correspond to similarly designated constructs and variables within the literature at this point is unclear, but could be resolved empirically by correlating ATQ-N factor scores with existing measures of demoralization (Tellegen et al., 2003), [illegible] 1997) and the brooding subtype of rumination (Treynor et al., 2003). Until this additional research is conducted, any interpretation of our regression analysis should be held lightly. Nevertheless, it may be informative to at least tentatively relate our findings to other research, in [illegible] examined the relationship between rumination, on the one hand, and depression, suicidality, and related clinical phenomena, on the other.

Clearly the most unexpected finding of our regression analysis was that rumination typified by a negative self-focus (e.g., What's the matter with me?) did not significantly contribute to variability in BDI scores. Instead brooding was negatively, although not significantly, related to levels of self-reported depression. It seems worthy of further research to examine whether this apparent disconnect may in part be yet just another reflection of a discontinuity between nonclinical and clinical samples in studying ostensibly similar processes. Unlike the reflective [illegible] rumination brood- [illegible] ing has been shown to be more predictive of BDI scores [illegible] community and clinical samples (Morrison & O'Connor, 2008). While rumination in general has been implicated in both the onset (Robinson & Alloy, 2003) [illegible] of depression (Nolen-Hoeksema, McBride, & Larson, 1997), more recent research that found the distinction between the two types of rumination to be blurred in currently depressed populations (Whitmer & Gotlib, 2011). When viewed in the aggregate, the literature to date on rumination in combination with our findings suggests that the possibility that brooding may contribute to the development of both suicidality and depression in both nonclinical and clinical populations, but independently account [illegible] variability in levels of depression in clinical samples seems worthy of further exploration. Our hope is that [illegible] negative automatic thoughts identified in our factor analysis might play at least some small role in this larger endeavor.

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