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The Structure of Perceived Emotional Control: Psychometric Properties of a Revised Anxiety Control Questionnaire

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The psychometric properties of the Anxiety Control Questionnaire (ACQ) were evaluated in 1,550 outpatients with *DSM-IV* anxiety and mood disorders and 360 nonclinical participants. Counter to prior findings, exploratory factor analyses produced a 3-factor solution (Emotion Control, Threat Control, Stress Control) based on 15 of the ACQ's original 30 items. Factor analyses in two independent clinical samples (e.g., confirmatory factor analysis, CFA) replicated the 3-factor solution. Multiple-groups CFAs indicated that the measurement properties of the ACQ were invariant in male and female patients, and that the ACQ was largely form and parameter equivalent in a clinical versus nonclinical sample. Hierarchical analysis supported the existence of a higher-order dimension of perceived control. Structural regression analyses indicated that each of the ACQ factors accounted for significant unique variance in one or both latent factors representing the dimensions of autonomic anxiety and depression. The results are discussed in regard to their conceptual and psychometric implications to the construct of perceived emotional control.

Contemporary theories propose that perceptions of diminished control over aversive events are vital to the etiology and maintenance of emotional disorders (Barlow, 2002; Barlow, Chorpita, & Turovsky, 1996). Similar to Bandura's (1986) theory of self-efficacy, Barlow (2002) describes anxiety as a cognitive-affective process in which an individual has a sense of unpredictability

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and uncontrollability over potentially negative or harmful events and emotions. This sense of unpredictability and uncontrollability is thought to be associated with physiological arousal, anxious apprehension, and uncertainty about one's ability to manage the threats. Research has consistently found that individuals who perceive limited control over aversive bodily or environmental events are at increased susceptibility to anxiety-related distress (Barlow et al., 1996; Chorpita & Barlow, 1998; Mineka & Zinbarg, 1996; Zvolensky, Eifert, Lejuez, Hopko, & Forsyth, 2000). For example, individuals with panic disorder report significantly less fear during laboratory panic provocations when they perceive the procedures as controllable (Rapee, Mattick, & Murrell, 1986; Sanderson, Rapee, & Barlow, 1989); similar laboratory-based findings have also been obtained in analog samples (e.g., Telch, Silverman, & Schmidt, 1996; Zvolensky, Eifert, Lejuez, & McNeil, 1999). Such results have augmented the theoretical and empirical significance of the construct in recent years; indeed, current conceptual models posit that low perceived control represents a shared psychosocial diathesis for anxiety and depression, helping to account for the overlap of these syndromes (Alloy, Kelly, Mineka, & Clements, 1990; Barlow, 2002; Mineka, Watson, & Clark, 1998; Zvolensky, Lejuez, & Eifert, 2000).

Although general measures of perceived control have been developed and evaluated in various populations (e.g., Craig, Franklin, & Andrews, 1984; Rotter, 1966; Shapiro et al., 1993), research has shown that more specific instruments are needed to evaluate predictions based on current theories of emotional disorders (Rapee, Craske, Brown, & Barlow, 1996). Despite the central role of perceived control in these theories, an instrument designed to specifically assess this construct has only recently been developed and tested. The Anxiety Control Questionnaire (ACQ; Rapee et al., 1996) is a 30-item instrument intended to assess an individual's perceived level of control over anxiety and anxiety-related events. Since its inception, the ACQ has been used in a variety of empirical contexts, such as studies of response to laboratory stressors (e.g., Feldner & Hekmat, 2001), interpretive biases (e.g., Zvolensky et al., 2001), psychopathology (e.g., White, Brown, Somers, & Barlow, 2003; Zebb & Moore, 2003), basic theoretical research (e.g., Stevens, 1997), and treatment outcome trials (e.g., Barlow, Rapee, & Brown, 1992). Though the ACQ has been frequently used as an indicator of the theoretically relevant dimension of perceived control, studies examining its latent structure have produced inconsistent findings.

The ACQ was designed to assess perceived control over a variety of potentially threatening internal and external events/situations that are directly relevant to the emotional disorders (Rapee et al., 1996). Items were developed to be face valid, and 18 of the items were reverse worded to avoid response bias. Using principal components analysis (PCA) with varimax rotation, Rapee et al. reported a two-factor solution representing perceived control over emotional reactions and external threats. The solution was obtained in a clinical sample of 250 patients with anxiety disorders and a nonclinical sample ($N = 236$).

After examination of internal consistency, factor structure, and items that double-loaded or that did not meet salient loading criteria (i.e., $\geq .30$), the final scale contained 30 items that formed two subscales labeled to reflect perceived control over external events (Events) and perceived control over internal reactions (Reactions). Rapee et al. subjected the ACQ to additional tests of reliability, validity, and sensitivity to change during treatment, and all evidence attested to the favorable psychometric properties of the scale. Although PCA resulted in a two-factor solution with good internal consistency, the authors recommended that the ACQ be scored using the 30-item total because "clear existence of a second factor is not strong" (Rapee et al., p. 289).

More recently, the latent structure of the ACQ was reevaluated by Zebb and Moore (1999) in a sample of 303 college students with exploratory analyses (PCA with varimax and oblique rotation). In contrast to the two-factor solution reported in Rapee et al. (1996), Zebb and Moore (1999) obtained a three-factor solution (two factors representing the absence of perceived control, one factor representing the presence of perceived control). In view of the numerous explanations that may have contributed to these inconsistent results (e.g., structural noninvariance between clinical and nonclinical samples), Zebb and Moore (1999) recommended that researchers continue to explore the ACQ's factor structure.

The importance of this recommendation is underscored by the numerous methodological shortcomings associated with extant structural analyses of the ACQ. For instance, although prior psychometric evaluations purported to conduct exploratory factor analyses (EFA), this in fact was not the case due to the exclusive use of PCA extraction. Unlike EFA, which is based on the common factor model (Thurstone, 1947), the objective of PCA is data reduction. Accordingly, PCA does not differentiate between common and unique variance (cf. Fabrigar, Wegener, MacCallum, & Strahan, 1999). Although the differing objectives of these extraction methods have been regarded by some as somewhat trivial, there are indeed many scenarios (e.g., low communalities, low item–factor ratios) where PCA and EFA can produce quite discordant results (Widaman, 1993).[1] Moreover, no study to date has examined the latent structure of the ACQ within the confirmatory factor analysis (CFA) framework. CFA affords a more comprehensive evaluation of latent structure, including significance tests of factor loadings and intercorrelations, presence/absence of negative residual variances, incorporation of an error theory (e.g., existence of minor factors), and direct statistical determination of the extent of measurement invariance of the solutions across relevant subpopulations (e.g., sexes, clinical versus nonclinical). Indeed, sample size limitations in Rapee et al. (1996) and Zebb and Moore (1999) precluded adequate cross-validation of the resulting solutions.

[1] Also, it is conceptually and mathematically inconsistent to conduct PCA when the stated objective is factor analysis (i.e., to reproduce the sample correlation matrix by estimating the pattern of relationships between common factors and indicators).

In addition, whereas both studies addressed the issue of possible ACQ sex differences, these analyses were performed in a cursory fashion (*t* tests of coarse composite scores) in the absence of CFA-based evaluation of the many possible sources of noninvariance (e.g., differing factor structures or factor loadings). The strengths of CFA can also be capitalized on to address the important question of the extent to which the discrepant structures obtained by Rapee et al. (1996) and Zebb and Moore (1999) were due to the use of clinical versus nonclinical samples. Because the ACQ is used frequently in clinical and nonclinical research alike, inconsistencies in its measurement properties between these populations would have important implications (e.g., complications with clinical and nonclinical group comparisons; possible need for different scoring methods).

Thus, a key aim of the present study was to provide a thorough evaluation of the latent structure of the ACQ. In addition to conducting large-sample replications of the factor solutions, we examined the degree of measurement equivalence of the ACQ in selected subpopulations (male versus female patients; clinical versus nonclinical samples). We also tested a higher-order model of perceived control where the correlations among first-order factors were explained by a single second-order factor. Finally, we evaluated the extent to which the ACQ factors differentially predicted (and added to the prediction of) core dimensions of anxious arousal and depression.

Method

Participants

The clinical sample consisted of 1,550 patients who presented for assessment and treatment at the Center for Anxiety and Related Disorders between the periods of October 1996 and November 2001. Women constituted the larger portion of the sample (60.5%); average age was 33.22 (*SD* = 11.23, range = 18 to 75). Most patients were Caucasian (88.6%), with smaller numbers of individuals identifying as Asian (4.3%), African American (3.5%), Hispanic (3.0%), and other (0.6%). Diagnoses were established with the Anxiety Disorders Interview Schedule for *DSM-IV*—Lifetime version (ADIS-IV-L; Di Nardo, Brown, & Barlow, 1994), a semistructured interview designed to ascertain reliable diagnosis of the *DSM-IV* anxiety, mood, somatoform, and substance use disorders, and to screen for the presence of other conditions (e.g., psychotic disorders). A reliability study of a subset of the current sample (n = 362) who had two independent administrations of the ADIS-IV-L indicated good-to-excellent interrater agreement for current principal disorders (range of κs = .67 to .86), except dysthymia (e.g., κ = .22; Brown, Di Nardo, Lehman, & Campbell, 2001). For each diagnosis, interviewers assign a 0-to-8 clinical severity rating (CSR) that indicates the degree of distress and impairment associated with the disorder (0 = *none* to 8 = *very severely disturbing/disabling*). In patients with two or more current diagnoses, the "principal" diagnosis is the one receiving the highest CSR. For current and lifetime

disorders that meet or surpass the threshold for a formal *DSM-IV* diagnosis, CSRs of 4 (*definitely disturbing/disabling*) or higher are assigned ("clinical" diagnoses). Current clinical diagnoses not deemed to be the principal diagnosis are referred to as "additional" diagnoses. The breakdown of current clinical disorders (collapsing across principal and additional diagnoses) was as follows: panic disorder with or without agoraphobia (n = 594), generalized anxiety disorder (n = 318), social phobia (n = 639), specific phobia (n = 318), obsessive-compulsive disorder (n = 172), major depression (n = 408), dysthymic disorder (n = 122), other anxiety/mood disorder (e.g., posttraumatic stress disorder, anxiety or depressive disorder NOS; n = 261).

The nonclinical sample consisted of 360 undergraduates who completed the ACQ during their participation in an assessment study (for course credit) conducted at the University of Albany, State University of New York (190 males, 169 females, 1 missing). The average age of this sample was 18.68 (SD = 0.98, range = 17 to 22). Most nonclinical participants were Caucasian (65.8%), with smaller numbers of individuals identifying as African American (10.3%), Hispanic (10.0%), Asian (8.1%), and other (5.9%).

Measures

As noted earlier, the ACQ is a 30-item instrument designed to measure perceived control over emotional reactions and external threats (Rapee et al., 1996; see Appendix). Participants respond to the items using a 0-to-5 scale (0 = *strongly disagree*, 5 = *strongly agree*). After reverse scoring 18 items, two subscales are formed by summation (Reactions, 14 items; Events, 16 items), as well as a 30-item total score.

In addition to the ACQ, the following questionnaires were administered to the clinical sample: Beck Depression Inventory (BDI; Beck & Steer, 1987), Beck Anxiety Inventory (BAI; Beck & Steer, 1990), and Depression Anxiety Stress Scales (DASS; Lovibond & Lovibond, 1995). These measures were used in validity analyses involving the latent factors of Depression (BDI, DASS-Depression) and Anxious Arousal (BAI, DASS-Anxiety).

Procedure

The clinical sample completed the ACQ as part of their initial intake evaluation (consisting of the ADIS-IV-L and a questionnaire battery) at the Center for Anxiety and Related Disorders. In the nonpatient sample, the ACQ was one of several measures administered in small-group testing sessions (roughly 30 participants per group) as part of a questionnaire study on the psychological risk factors for nonclinical panic (Forsyth, Karekla, & Zvolensky, 2002). Because of the psychometric nature of the present study, only cases in both samples who completed all 30 ACQ items were included in the analyses.

To thoroughly replicate the ACQ factor structure, the clinical sample (N = 1,550) was randomly divided into three subsamples. Sample 1 (n = 450; 184 males, 266 females) was used to conduct initial EFAs of the original 30-item ACQ and revised solutions. The inconsistent prior findings regarding latent

structure (Rapee et al., 1996; Zebb & Moore, 1999) and the methodological shortcomings noted earlier required that initial analyses of the ACQ's latent structure be conducted in an exploratory fashion (i.e., a strong empirical basis meriting CFA was lacking). Sample 2 (n = 400; 151 males, 249 females) and Sample 3 (n = 700; 277 males, 423 females) were used to replicate the final ACQ structure obtained in Sample 1. To provide a maximal N, analyses of measurement invariance between male and female patients were conducted using both replication samples (n = 1,100). Patients in the replication samples who were age 22 or below (n = 272) were also used in multiple-groups CFAs that evaluated the form and parameter equivalence in comparison to a nonclinical, undergraduate sample (n = 360).

Data Analysis

The sample variance-covariance matrices were analyzed using latent variable software programs and maximum-likelihood minimization functions (LISREL 8.53, Jöreskog & Sörbom, 2002; Mplus 2.13, Muthén & Muthén, 1998). Goodness of fit was evaluated using the Root Mean Square Error of Approximation (RMSEA; Steiger, 1990), p value for test of close fit (RMSEA < 0.05) (CFit), Comparative Fit Index (CFI; Bentler, 1990), and the Tucker-Lewis Index (TLI; Tucker & Lewis, 1973). Multiple indices were selected because they provide different information for evaluating model fit (i.e., absolute fit, fit adjusting for model parsimony, fit relative to a null model); used together, these indices provide a more conservative and reliable evaluation of the model fit (cf. Jaccard & Wan, 1996). In instances where competing models were nested, comparative fit was evaluated with nested χ^2 tests.

Results

Exploratory Factor Analysis

Using Sample 1 (n = 450), the 30 ACQ items were submitted to an exploratory factor analysis (EFA; maximum likelihood estimation, promax rotation). Acceptability of the factor models (e.g., factor selection) was evaluated by goodness of model fit (RMSEA < .08, CFit p > .05),[2] the interpretability of the solution, and the strength of the parameter estimates (e.g., primary factor loadings ≥.30, absence of salient cross-loadings). Counter to the conclusions of Rapee et al. (1996), a 2-factor solution did not fit the data well, $\chi^2(376) = 961.83$, $p < .001$, RMSEA = .059, CFit = .001 (eigenvalues ≥ 1.0 for the unreduced correlation matrix were: 6.97, 1.97, 1.74, 1.24, 1.23, 1.12, 1.05). Conversely, acceptable model fit was obtained for a 3-factor solution,

[2] In addition to practical considerations (Mplus provides only a handful of goodness-of-fit indices for EFA), RMSEA and its significance test (CFit) are emphasized as indices of EFA model fit for reasons discussed in MacCallum, Browne, and Sugawara (1996; e.g., χ^2 evaluates the overly stringent hypothesis of exact fit; RMSEA has favorable statistical properties including its penalization of poor model parsimony and its known distributional properties that allow for the generation of confidence intervals and significance tests).

$\chi^2(348) = 738.99, p < .001$, RMSEA = .050, CFit = .497, and a 4-factor solution, $\chi^2(321) = 611.54, p < .001$, RMSEA = .045, CFit = .942 (a 5-factor model did not converge). However, 7 items did not have clearly salient loadings on any factor in these solutions (items 1, 4, 6, 9, 21, 25, 30). Thus, another EFA was conducted without these items. As before, a 2-factor solution was poor-fitting, $\chi^2(208) = 640.79, p < .001$, RMSEA = .068, CFit < .001 (eigenvalues $\geq$ 1.0 = 5.79, 1.79, 1.59, 1.22, 1.09, 1.00). Whereas the 3- and 4-factor models provided an adequate fit to the data (e.g., 3-factor model: $\chi^2[187] = 442.16, p < .001$, RMSEA = .055, CFit = .102), a few problematic items remained that either had low primary factor loadings or had salient loadings on more than one factor (items 11, 12, 27, 28, 29). Moreover, both the 3- and 4-factor solutions included a factor defined by 3 items (items 7, 19, 23) pertaining to social impression management (e.g., "There is little I can do to influence people's judgment of me," "I can usually make sure people like me if I work at it"). Because this dimension lacked a strong conceptual basis to the construct of perceived control (cf. Barlow, 2002), and was relatively unrelated to other factors in the solution (*r*s = .19 and .21 in the 3-factor solution), this factor was dropped in subsequent analyses.

Therefore, a third EFA was conducted using the remaining 15 items. Once again, a 2-factor solution was poor-fitting, $\chi^2(76) = 250.15, p < .001$, RMSEA = .071, CFit < .001 (eigenvalues $\geq$ 1.0 = 4.67, 1.41, 1.28). However, the 3-factor solution provided a good fit to the data, $\chi^2(63) = 125.46$, $p < .001$, RMSEA = .047, CFit = .648 (49% variance explained). Whereas the 4-factor solution also fit the data well, $\chi^2(51) = 80.56, p < .01$, RMSEA = .036, CFit = .946, the fourth factor was poorly defined (i.e., all potential items on this factor had salient cross-loadings on other factors). The promax-rotated pattern matrix of the 3-factor solution is presented in Table 1. The first factor, labeled "Emotion Control" (EC), consists of 5 of 14 items that belonged to the Reactions subscale reported in Rapee et al. (1996); although 1 item is reverse scored (#26), these items appear to reflect one's ability to effectively control one's emotions (e.g., "I am able to control my level of anxiety"). The second factor, labeled "Threat Control" (TC), consists of 6 of the original 16 items of the Events subscale (cf. Rapee et al.) whose content bears on the belief that the occurrence of or escape from frightening events is out of one's control (e.g., "There is little I can do to change frightening events"). The third factor, "Stress Control" (SC), is comprised of 4 items (2 each from the Reaction and Events subscales) measuring perceived difficulty coping and regulating one's emotions in stressful situations (e.g., "When I am put under stress, I am likely to lose control"). As seen in Table 1, all 15 items had salient loadings on their primary factor (range = .36 to .86) and no salient cross-loadings were evident. The factors were moderately intercorrelated: EC with TC and SC (both *r*s = .54); TC with SC = .52.

Although reported infrequently in applied psychometric research, factor determinacy data are important in the evaluation of factor analytic findings (e.g., a highly indeterminate factor can produce radically different factor

TABLE 1
LATENT STRUCTURE OF THE ANXIETY CONTROL QUESTIONNAIRE: EXPLORATORY FACTOR ANALYSIS (n = 450) AND EXPLORATORY FACTOR ANALYSIS CONDUCTED WITHIN THE CONFIRMATORY FACTOR ANALYSIS FRAMEWORK (n = 400)

	Factor					
	Emotion Control		Threat Control		Stress Control	
ACQ Item	EFA	E/CFA	EFA	E/CFA	EFA	E/CFA
10	**.557**	**.574**	.026	.088	.007	.136
13	**.645**	**.664**	.044	.099	.113	.074
17*	**.633**	**.675**	.044	.000	.027	.000
22	**.364**	**.567**	.166	.058	.053	.014
26	**.453**	**.463**	.110	.162	.060	.043
5	.174	.162	**.608**	**.523**	.019	.062
8	.054	.027	**.587**	**.618**	.172	.134
14*	.012	.000	**.637**	**.703**	.049	.000
15	.001	.104	**.526**	**.635**	.024	.057
16	.046	.013	**.458**	**.583**	.120	.055
20	.087	.072	**.543**	**.469**	.133	.106
2	.118	.091	.071	.059	**.499**	**.493**
3*	.061	.000	.016	.000	**.863**	**.667**
18	.011	.003	.195	.109	**.483**	**.431**
24	.163	.053	.039	.018	**.534**	**.662**
Determinacy	.855	.881	.885	.896	.904	.855

Note. ACQ = Anxiety Control Questionnaire; EFA = exploratory factor analysis (maximum likelihood extraction, promax rotation); E/CFA = exploratory factor analysis within the confirmatory factor analysis framework (maximum likelihood). Items with asterisks were used as anchor indicators in the E/CFA analysis. Factor loadings ≥ .30 are in bold.

scores that are nonetheless equally consistent with the obtained factor loadings). Accordingly, factor determinacies (i.e., validity coefficients: correlation between factor score estimates and their respective factors) were computed on the refined factor scores (least squares regression approach; Thurstone, 1947) using the SAS PROC IML routines provided by Grice (2001). As shown in Table 1, the 3 factors evidenced favorable levels of determinacy (range = .86 to .90), per the recommendations of Gorsuch (1983; i.e., ≥.80).[3]

[3] Grice (2001) has also developed SAS PROC IML methods for computing the correlational accuracy (equivalence of correlations among factor scores to correlations among factors) and univocality of factor scores (extent that factor scores are excessively or insufficiently correlated with other factors). Although these results were favorable in the current data set (e.g., absolute discrepancies between correlations among factors and refined factor scores were .12, .13, and .14), they are not reported in detail because of space considerations and because factor scores were not used as variables in subsequent analyses.

Exploratory Factor Analysis Within the Confirmatory Factor Analysis Framework

The next two clinical samples were used to evaluate the replicability of the 4-factor solution obtained in the aforementioned EFA of the 15 ACQ items. As an intermediate step between EFA and confirmatory factor analysis (CFA), the ACQ data from Sample 2 (n = 400) were analyzed using the EFA within the CFA framework approach (E/CFA; Jöreskog, 1969). Although underutilized in applied factor analysis research, E/CFA is a very helpful precursor to CFA. In this strategy, the CFA applies only to the identifying restrictions used in EFA (e.g., fix factor variances to 1.0 for m restrictions; for each factor, select an anchor item to freely load onto it but fix its cross-loadings to 0; allow nonanchor items to be freely estimated). Whereas this specification uses the same number of restrictions (m^2) and produces the same overall fit as EFA, the CFA maximum likelihood estimation provides a great deal of additional information that can be important in developing realistic CFA models in subsequent analyses (e.g., standard errors to determine the significance of factor loadings and factor correlations; modification indices reflecting possible residual covariances).

Using items 3, 14, and 17 as anchor items,[4] the 3-factor E/CFA model provided a good fit to the data, $\chi^2(63) = 109.48, p < .001$, RMSEA = .043, CFit = .799, CFI = .967, TLI = .945. As shown in Table 1, the magnitudes of primary loadings were strong (range = .43 to .70) and statistically significant (range of zs = 5.35 to 14.19). All factor determinacies (calculated by the Mplus software) were above .80 (range = .86 to .90). The factor intercorrelations were as follows: EC with TC and SC = .42 and .51, respectively; TC with SC = .44 (ps < .001). More importantly, E/CFA provides modification indices for indicator error covariances that might point to the existence of minor factors or a more complex error structure that would be undetected by EFA. In this regard, only one noteworthy modification index was observed (MI = 17.73) that involved correlated error between items 15 and 16 [standardized expected parameter change (EPC) = .17].

Collectively, the results of EFA and E/CFA indicate that the latent structure of the 15 retained ACQ items is defined by 3 moderately intercorrelated factors (EC, TC, SC), each associated with a favorable level of determinacy. In addition to upholding the 3-factor structure, E/CFA findings suggest that subsequent CFA models can be pursued under the assumption of random measurement error, with the possible exception of items 15 and 16 (both from the TC factor).

Confirmatory Factor Analysis

On the basis of the solutions obtained in Samples 1 and 2, a 3-factor CFA model was fit to the Sample 3 data (n = 700). This model provided an acceptable

[4] Anchor items are selected by identifying one item per factor that has a strong primary loading and very small cross-loadings. To achieve m^2 restrictions (in addition to fixing factor variances to unity), all loadings are freely estimated except for anchor item cross-loadings which are fixed to zero (cf. Jöreskog, 1969).

fit to the data, $\chi^2(87) = 243.81, p < .001$, RMSEA = .051, CFit = .423, CFI = .934, TLI = .920. Fit diagnostics indicated no salient strains in the solution except for the residual covariance of items 15 and 16 (MI = 44.45, standardized EPC = .20). Given the consistency of this outcome with the E/CFA results and the frequent necessity to model nonrandom error in test instruments due to method effects (cf. Brown, 2003; Byrne, Shavelson, & Muthén, 1989; Gerbing & Anderson, 1984; Marsh, 1996), the solution was respecified correlating the residuals of items 15 and 16. In addition to improving fit over the previous solution, $\chi^2_{diff}(1) = 43.57, p < .001$, this model fit the data well, $\chi^2(86) = 200.24, p < .001$, RMSEA = .044, CFit = .908, CFI = .952, TLI = .941. As shown in Table 2, the magnitudes of the loadings were strong (range = .43 to .71) and the factor determinacies were quite satisfactory (all >.87). Factor intercorrelations were: EC with TC and SC = .51 and .62, respectively; TC with SC = .63 ($ps < .001$).

The scale reliabilities of the 3 factors were calculated within the CFA

TABLE 2
LATENT STRUCTURE OF THE ANXIETY CONTROL QUESTIONNAIRE: CONFIRMATORY FACTOR ANALYSIS USING SAMPLE 3 (n = 700) AND MALE (n = 428) AND FEMALE (n = 672) PATIENTS

	Factor								
	Emotion Control			Threat Control			Stress Control		
ACQ Item	S3	Males	Females	S3	Males	Females	S3	Males	Females
10	.553	.541	.580						
13	.644	.573	.697						
17	.574	.578	.612						
22	.618	.686	.580						
26	.567	.556	.583						
5				.708	.668	.700			
8				.624	.691	.645			
14				.620	.616	.661			
15				.551	.488	.592			
16				.495	.503	.535			
20				.428	.400	.443			
2							.485	.523	.452
3							.711	.622	.737
18							.563	.515	.564
24							.686	.710	.667
Determinacy	.873	.871	.885	.884	.880	.894	.879	.859	.881
Reliability	.728	.724	.749	.730	.721	.756	.706	.684	.700
Mean	6.57	6.71	6.33	16.86	16.89	16.52	10.15	10.25	9.83
SD	4.88	4.75	4.91	6.09	5.81	6.43	4.61	4.31	4.70

Note. S3 = Sample 3; ACQ = Anxiety Control Questionnaire. Means and *SD*s are based on coarse factor scores (i.e., raw score composites).

model using the approach developed by Raykov (2001). This method reconciles the problems with Cronbach's α which is a misestimator of scale reliability except in the rare instance when all elements of a multiple-item measure are tau-equivalent and possess random measurement error (Lord & Novick, 1968; McDonald, 1999; Raykov, 2001). In LISREL, the procedure entails specifying three dummy latent variables whose variances are constrained to equal the numerator (true score variance), denominator (total variance), and corresponding ratio of true score variance to total score variance, per the classic formula for scale reliability estimation (Lord & Novick, 1968). As shown in Table 2, although likely attenuated to some degree by the small number of items per factor, the scale reliabilities were acceptable (i.e., $\geq .65$; ρs = .71, .73, and .73 for SC, EC, and TC, respectively).

Thus, the more restrictive CFA solution (e.g., unlike EFA, all item cross-loadings and all but one error covariance were fixed to 0) provided further support for the 3-factor ACQ solution. In addition to results indicating acceptable degrees of overall and localized fit, the suitability of the 3 factors was evidenced by favorable magnitudes of the factor loadings (range = .43 to .71), factor determinacies ($>.87$), and scale reliabilities ($\geq .71$).

Measurement Invariance and Population Heterogeneity

Male versus female patients. The degree of measurement invariance (e.g., equal factor loadings and indicator intercepts) and population heterogeneity (e.g., equal factor *M*s) of the 15-item ACQ were examined in male and female patients using multiple-groups CFA. To maximize available sample size, the two replication samples (total n = 1,100) were combined (males = 428, females = 672). Prior to the multiple-groups solutions, CFAs were conducted separately to verify adequate fit in the male and female subsamples. Next, a two-group CFA was conducted to test equal ACQ form between sexes. As shown in Table 3, these models fit the data well. The factor loadings and factor determinacies for males and females are provided in Table 2 (correlations of item 15 and item 16 residuals were .23 and .16 for males and females, respectively, ps $< .001$). Given evidence of equal form, the parameter equivalence of the ACQ was evaluated in a series of two-group CFAs that entailed increasingly restrictive constraints. The first analysis indicated that the factor loadings were equivalent in the male and female samples, $\chi^2_{diff}(12) = 18.90$, *ns*. The next analysis, which constrained the indicator intercepts to equality, produced a significant degradation in model fit, $\chi^2_{diff}(12) = 25.84$, $p < .05$. Fit diagnostics indicated that the decrease in fit was due to a difference on the intercept of item 2. This intercept was freely estimated in subsequent models in order for the parameter equivalence analysis to proceed (partial intercept invariance; cf. Byrne et al., 1989). As shown in Table 3, models that imposed additional equality constraints on the solution did not have deleterious effects on model fit (equal factor variances and covariances; no sex differences on the 4 factor *M*s). Thus, the collective findings suggest that the ACQ measurement model is invariant between male and female patients, with

TABLE 3
CONFIRMATORY FACTOR ANALYSES OF THE ANXIETY CONTROL QUESTIONNAIRE: TESTS OF MEASUREMENT INVARIANCE AND POPULATION HETEROGENEITY

	χ^2	*df*	χ^2_{diff}	Δdf	RMSEA	CFit (*p*)	CFI	TLI
Clinical Sample: Males and Females								
Males (n = 428)	147.37	86			.041	.912	.954	.944
Females (n = 672)	169.51	86			.038	.992	.966	.958
Equal form (n = 1,100)	316.88	172			.039	1.00	.962	.953
Equal factor loadings	335.78	184	18.90	12	.039	1.00	.960	.954
Equal intercepts	361.62	196	25.84*	12	.039	1.00	.956	.953
Partial intercept invariance[a]	350.97	195	15.19	11	.038	1.00	.959	.955
Equal factor variances/ covariances	360.11	201	9.14	6	.038	1.00	.958	.956
Equal factor means	361.88	204	1.77	3	.038	1.00	.958	.957
Clinical and Nonclinical Sample (Age ≤ 22)								
Clinical sample (n = 272)	146.08	86			.051	.452	.934	.919
Nonclinical sample (n = 360)	174.23	86			.053	.300	.909	.889
Nonclinical sample (item 26 load on Stress factor)	156.54	86			.048	.609	.927	.911
Partial form equivalence (n = 632)	302.62	172			.049	1.00	.930	.915
Equal factor loadings[b]	317.12	183	14.50	11	.048	1.00	.929	.918
Equal intercepts	382.66	194	65.54***	11	.055	1.00	.899	.891
Partial intercept invariance[c]	329.89	190	12.77	7	.048	1.00	.925	.918
Equal factor variances/ covariances	348.61	196	18.72**	6	.050	1.00	.925	.913
Partial factor variance invariance[d]	336.32	194	6.43	4	.048	1.00	.924	.918
Equal factor means	679.25	197	342.93***	3	.088	.093	.743	.726

Note: χ^2_{diff} = nested χ^2 difference; RMSEA = root mean square error of approximation; CFit = *p* value for test of close fit (RMSEA < .05); CFI = comparative fit index; TLI = Tucker-Lewis Index.

[a] All intercepts held invariant except for item 2.

[b] Item 26 freely estimated (loading, intercept) in all tests.

[c] All intercepts held invariant except for items 5, 13, 15, 20.

[d] All variances/covariances held invariant except for variances of Threat Control and Stress Control.

*p < .05; **p < .01; ***p < .001.

the exception of the item 2 intercept (a difference that could be due to chance in view of the number of constraints that were imposed on the models).

In addition, Table 2 provides the ACQ scale reliabilities for male and female patients. In addition to point estimates of scale reliabilities (Raykov, 2001), Raykov (2002) has recently developed covariance structure analytic procedures to statistically test for differences in scale reliabilities across groups. This approach entails fixing the paths from the error terms into the indicators to equal the squared sum of factor loadings. The resulting error terms reflect

reciprocals of the right-hand side of the classic reliability equation (Lord & Novick, 1968): theta divided by total true score variance (this parameterization produces re-scaled error terms that, when summed, indicate the ratio of total error variance to total true score variance; i.e., the product of these summed terms and true score variance = total error variance). The null hypothesis of equal scale reliabilities is tested by a nested comparison of two models: (1) re-scaled error terms freely estimated in each group; and (2) the sums of the re-scaled error terms are constrained to equality across groups. Thus, the method avoids difficulties associated with imposing nonlinear constraints on the reliability coefficients themselves (Raykov, 2002). Because the two-group model without constraints on the error terms is a reparameterization of a structural equivalence model, goodness of fit was identical to the equal form solution; e.g., $\chi^2(172) = 316.88$ (see Table 3). The second model, where the sum of males' and females' re-scaled error terms were constrained to equality, produced a $\chi^2(175) = 319.41$, RMSEA = .039, CFit = 1.00, CFI = .98, TLI = .98. As reflected by the nonsignificant nested χ^2 test, $\chi^2_{diff}(3) = 2.53$, *ns*, these findings indicate the scale reliabilities of the 3 ACQ factors are equivalent in the male and female patients.

Clinical versus nonclinical sample. To determine the equivalence of the 3-factor solution in a nonclinical sample ($n = 360$), multiple-group CFA solutions were conducted using clinical and nonclinical participants. Because the nonclinical sample was drawn from a college student population, the clinical comparison sample was limited to patients age 22 or below ($n = 272$), to foster the demographic similarity of the groups.

Prior to the multiple-group CFA, CFAs were conducted separately for the clinical and nonclinical samples. The 3-factor measurement model fit the clinical sample data well (see Table 3). As seen in Table 4, the magnitude of primary factor loadings (range = .46 to .76) and factor determinacies (all >.87) were quite favorable, and no significant areas of ill fit were noted. However, this was not the case for the nonclinical sample (e.g., TLI = .89, see Table 3). Fit diagnostics indicated that a considerable point of strain in this model involved item 26. Although item 26 had a primary loading on the EC factor in the clinical sample (and no evidence of cross-loadings), model fit could be improved in the nonclinical sample by allowing this item to load on the SC factor (modification index = 23.34, standardized EPC = .64). Respecification of this solution, allowing item 26 to load on SC instead of EC, produced an acceptable fitting model, $\chi^2(86) = 156.54$, RMSEA = .048, CFit = .609, CFI = .927, TLI = .911 (no evidence of a salient cross-loading on the EC factor was noted; e.g., EPC = .21). The factor loadings and factor determinacies of this model are presented in Table 4. Finally, a multiple-group CFA was conducted to simultaneously estimate the parameters in both groups (partial form equivalence). This model provided an acceptable fit to the data, $\chi^2(172) = 302.62$, RMSEA = .049, CFit = 1.00, CFI = .930, TLI = .915.

As seen in Table 3, equality constraints to the factor loadings (except item 26) did not significantly degrade the fit of the model, $\chi^2_{diff}(11) = 14.50$, *ns*,

TABLE 4
LATENT STRUCTURE OF THE ANXIETY CONTROL QUESTIONNAIRE: CONFIRMATORY FACTOR ANALYSIS OF CLINICAL (n = 272) AND NONCLINICAL SAMPLE (n = 360)

	Factor					
	Emotion Control		Threat Control		Stress Control	
ACQ Item	CLIN	NC	CLIN	NC	CLIN	NC
10	.542	.487				
13	.584	.499				
17	.522	.577				
22	.649	.671				
26	.606					.603
5			.755	.587		
8			.660	.567		
14			.646	.568		
15			.414	.460		
16			.478	.532		
20			.471	.490		
2					.456	.349
3					.711	.602
18					.513	.399
24					.716	.710
Determinacy	.874	.846	.896	.861	.878	.882
Reliability	.717	.649	.742	.691	.692	.670
Mean	5.89	11.96	16.15	18.25	9.31	13.75
SD	4.44	3.65	5.99	5.05	4.30	4.48

Note. ACQ = Anxiety Control Questionnaire; CLIN = clinical sample; NC = nonclinical sample; Means and *SD*s are based on coarse factor scores (e.g., Emotion Control *M*s are summations of 4 and 5 items for nonclinical and clinical groups, respectively).

indicating that 15 ACQ items were related to the factors in comparable ways in both groups (i.e., a factor loading is a regression slope relating an item to its corresponding latent variable). However, imposing these constraints on the item intercepts produced a significant χ^2 decrease, $\chi^2_{diff}(11) = 65.54$, $p < .001$. Sequential testing to determine the nature of the noninvariance (cf. Byrne et al., 1989) revealed that the intercepts for items 5, 13, 15, and 20 were noninvariant. A subsequent model that allowed these four intercepts to be freely estimated was equivalent to the equal factor loadings model, $\chi^2(7) = 12.77$, *ns*. Similarly, a model that held equal all factor variances and covariances led to a significant decline in goodness of fit, $\chi^2_{diff}(6) = 18.72, p < .01$. Sequential testing revealed that the nature of this noninvariance pertained to group differences in the TC and SC factor variances. Indeed, a respecified model that held all factor variance-covariances invariant except for these two elements did not degrade overall fit, $\chi^2(4) = 6.43$, *ns*. Consistent with the expectation of clinical versus nonclinical group differences on levels of

perceived control, a model that constrained the latent factor means to equality was poor fitting (e.g., CFI = .743), $\chi^2_{diff}(3) = 342.93$, $p < .001$. Post hoc evaluation of the sources of this noninvariance revealed significant group differences ($ps < .01$, with the clinical group scoring in the direction of less perceived control) on the means of all ACQ factors (z tests reflecting group differences in latent means were 10.19, 2.92, and 5.73 for EC, TC, and SC, respectively).

Finally, the scale reliabilities of the ACQ factors were calculated for the clinical and nonclinical samples using the Raykov (2001) method. As shown in Table 4, acceptable reliabilities were found for all scales in the clinical and nonclinical samples for each of the three ACQ factors (range = .65 to .74).

In summary, whereas the majority of the aspects of measurement invariance were equivalent in the clinical and nonclinical samples (e.g., the same number of factors; except for item 26, the same pattern and magnitude of factor loadings), group differences were obtained in some tests of population heterogeneity (i.e., 2 noninvariant factor variances, 3 noninvariant factor means). Such differences are consistent with the notion that the levels and dispersion of the underlying perceived control constructs differ between groups (e.g., clinical participants evidence lower levels of perceived control and a greater range on the construct continua).

Hierarchical Model of Perceived Control

Next, the viability of a higher-order "Perceived Control" (PC) dimension was considered using Sample 3 ($n = 700$). Because the structural component of the higher-order model was just-identified (i.e., a single higher-order PC factor, 3 lower-order factors of EC, TC, and SC), the appropriateness of hierarchical structure was assessed in terms of the magnitude of factor loadings of the 3 ACQ factors and individual ACQ items. Using factor correlations obtained from the first-order solution, results indicated that EC, TC, and SC loaded strongly on the higher-order PC factor (loadings = .70, .72, and .86, respectively, $ps < .001$). A Schmid-Leiman (1957) transformation was then conducted to obtain the loadings of the 15 ACQ items on the higher-order PC factor, and residualized loadings of the ACQ items on the 3 lower-order factors. As shown in Table 5, every ACQ item had a salient loading on the higher-order PC factor (range = .31 to .62). Moreover, with the exception of 3 items (2, 18, 20), all of the residualized first-order loadings were above .30, suggesting that the component dimensions of EC, TC, and SC account for unique salient variance in their constituent indicators, above the variance explained by the higher-order PC factor (as shown in Table 5, residualized primary loadings ranged from .39 to .46 for EC, from .30 to .49 for TC, and from .23 to .34 for SC).

The scale reliability of the higher-order PC factor was estimated using the LISREL parameterization developed by Raykov and Shrout (2002) for composites with general structure. Indeed, the scale reliability estimate of the general PC factor was quite favorable ($\rho = .848$). Collectively, these results

TABLE 5
HIGHER-ORDER FACTOR LOADINGS AND RESIDUALIZED PRIMARY LOADINGS FOR THE 15-ITEM ANXIETY CONTROL QUESTIONNAIRE (n = 700)

ACQ Item	Higher Order Factor Loading	Residualized Primary Loading
	Perceived Control	Emotion Control
10	.389	.393
13	.453	.457
17	.404	.408
22	.435	.439
26	.399	.403
	Perceived Control	Threat Control
5	.508	.493
8	.448	.435
14	.445	.432
15	.395	.384
16	.355	.345
20	.307	.298
	Perceived Control	Stress Control
2	.424	.235
3	.622	.345
18	.493	.272
24	.600	.332

Note. ACQ = Anxiety Control Questionnaire. Loadings transformed using the Schmid-Leiman procedure.

support the suitability of an ACQ latent structure defined by 3 primary factors, and a single higher-order factor of perceived control that has favorable scale reliability and accounts for salient variance in the 3 lower-order factors and 15 items.

Relationships of the ACQ Factors with Latent Dimensions of Anxious Arousal and Depression

In view of theoretical arguments that perceived control represents a psychosocial diathesis for both anxiety and depression (e.g., Barlow, 2002), additional models were developed using the Sample 3 data to evaluate the associations of the 3 ACQ factors with latent factors corresponding to the dimensions of anxious arousal and depression (cf. Brown, Chorpita, & Barlow, 1998; Clark, Watson, & Mineka, 1994). Along the lines of Brown et al. (1998), the BAI and DASS-Anxiety scales were specified as indicators of an Anxious Arousal (AA) latent factor, and the BDI and DASS-Depression scales were used as indicators of a Depression (DEP) factor (a correlated residual was estimated for the DASS-Anxiety and DASS-Depression indicators to reflect the nonrandom error due to method effects associated with subscales of the same

instrument). Ten patients in Sample 3 had missing data for at least one of these 4 indicators, resulting in an $n = 690$ for these analyses.

Because more factors emerged from latent structure analyses than originally proposed by the ACQ's developers, the question arises as to the meaning and importance of these dimensions or whether this result, albeit consistent across replications, is a relatively trivial artifact of scale development. Thus, of particular interest was whether the ACQ factors were differentially related to AA and DEP and whether each dimension explained significant unique variance in these outcomes. Drawing from the initial substantive interpretation of the factors, it was tentatively expected that the EC and TC factors would have greater relevance to the prediction of AA, and the SC factor of roughly equal relevance to AA and DEP.

Prior to interpretation of structural coefficients, the fit of the 5-factor measurement model was examined (EC, TC, SC, AA, DEP). This model provided a good fit to the data, $\chi^2(140) = 307.51, p < .001$, RMSEA = .042, CFit = .977, CFI = .982, TFI = .979. Fit diagnostics revealed no localized points of strain in the model and factor loadings for the AA and DEP latent variables indicated that the anxiety and depression measures were reasonable indicators of these constructs (AA: factor loadings = .90 and .89 for BAI and DASS-Anxiety, respectively; DEP: factor loadings = .97 and .81 for BDI and DASS-Depression, respectively; AA-DEP factor intercorrelation = .58).

The factor correlations (ϕs) of the ACQ factors with the AA and DEP factors are provided in Table 6 (all ps $< .001$). The differential magnitude of the correlations of each ACQ factor with the AA and DEP factors was evaluated

TABLE 6
DIFFERENTIAL RELATIONSHIPS OF THE ANXIETY CONTROL QUESTIONNAIRE FACTORS WITH LATENT FACTORS OF ANXIOUS AROUSAL AND DEPRESSION ($n = 690$)

	Criterion		
	Arousal	Depression	
ACQ-Emotion Control			
ϕ	−.582	−.487	$z = 3.36, p < .01$
γ	−.373***	−.165**	
ACQ-Threat Control			
ϕ	−.499	−.441	$z = 1.98, p < .05$
γ	−.200**	−.057	
ACQ-Stress Control			
ϕ	−.532	−.615	$z = 3.06, p < .01$
γ	−.177*	−.478***	

Note. ACQ = Anxiety Control Questionnaire; ϕ = zero-order correlations among latent factors; γ = completely standardized regression coefficients; z = z-test of the differential magnitude of factor intercorrelations.
*$p < .05$; **$p < .01$; ***$p < .001$.

using the z-test procedure of Meng, Rosenthal, and Rubin (1992). As shown in Table 6, both the EC and TC factors were more strongly correlated with the AA factor than the DEP factor. In contrast, the SC factor evidenced a significantly stronger zero-order relationship with the DEP factor than with the AA factor.

Next, the AA and DEP factors were regressed onto the ACQ factors to examine if each ACQ dimension accounted for significant unique variance in anxious arousal and depression.[5] Because the ACQ factors were not expected to account for all the covariance of the AA and DEP factors, correlated disturbances were specified for these outcomes (thus, this model was structurally just-identified and model fit was the same as the measurement model). The completely standardized regression coefficients (γs) from this model are presented in Table 6 (R^2 = .41 and .40 for AA and DEP, respectively). All ACQ factors significantly contributed to the prediction of AA. On the other hand, the EC and SC factors, but not the TC factor, accounted for significant unique variance in DEP.

Discussion

Owing partly to substantial methodological differences (e.g., large N analyses based on the common factor model and CFA), the present findings were largely inconsistent with the results of latent structure analyses from previous studies (Rapee et al., 1996; Zebb & Moore, 1999). Initial analyses led to the elimination of many original ACQ items due to modest or ambiguous primary factor loadings. Post hoc evaluation indicated that, in many instances, the poor psychometric performance of these deleted items may have been due to content that was largely unrelated to the purported trait of emotional control (e.g., state-like symptoms: "I often shake uncontrollably"; situational/emotional avoidance: "I am usually able to avoid threat quite easily"; general worry/hypochondriacal cognition: "When I hear someone has a serious illness, I worry that I am next"). After the instrument was refined to 15 items, the resulting 3-factor structure (EC, TC, SC) was quite consistent across clinical replication samples and between male and female patients. In addition, multiple-group CFAs indicated that the measurement properties of the revised ACQ were invariant in male and female patients (e.g., factor loadings, factor variances/covariances, latent Ms, scale reliabilities). With the exception of one item (#26), the ACQ evidenced form equivalence in a clinical and nonclinical sample (age 22 and below). Analyses in which factor loadings were held invariant did not degrade the multiple-group model, suggesting that the ACQ items were related to the factors in comparable ways in clinical and nonclinical participants (metric invariance). Not surprisingly, however, clinical versus nonclinical group differences were obtained on many aspects of population

[5] It is noteworthy that the ACQ factors are not multicollinear. In Sample 3 analyses where each ACQ factor was regressed onto the remaining 2 factors, range of R^2 = .40 to .51 (R = .63 to .72).

heterogeneity (two heterogeneous factor variances, clinical group scored in the direction of less perceived control on all three latent factor *M*s).

Whereas these results provide encouraging evidence of the structural and measurement invariance of the factor solutions, additional development and psychometric evaluation of the ACQ may be warranted. For example, while factor determinacies were adequate in all replications (indicating that the ACQ's factor scores could serve as suitable substitutes for the factors themselves in scenarios where latent structural analyses are not feasible), in some instances the estimates of scale reliability were somewhat modest (e.g., ρs ranged from .68 to .76 in the clinical samples). This was most evident in the estimates for SC, reflecting the well-known positive association between scale length and scale reliability (e.g., SC = 4 items; cf. Lord & Novick, 1968). Although methodologists have concluded that as few as three measured elements are sufficient to represent common factors (e.g., MacCallum, Widaman, Zhang, & Hong, 1999; Velicer & Fava, 1998), it would nonetheless be useful to expand the items of some ACQ factors to improve the reliability of their sum scores. In addition, evidence of the long-term temporal stability of the ACQ would be helpful in view of the conceptualization of perceived emotional control as a psychosocial diathesis to emotional disorders (Barlow, 2002). As the current study's analyses were limited to predominantly Caucasian outpatient and college student samples, additional research on the generalizability of the ACQ's measurement properties is necessary (e.g., community and varied ethnic and racial samples).

The key question also arises as to whether the multiple factors that emerged from these analyses, albeit consistent across replications, are clinically and theoretically meaningful. Though the current findings might ultimately contribute to a reconceptualization of perceived control (e.g., the construct is more complex than originally believed), it is also possible that the multidimensionality of the ACQ is essentially artifactorial (e.g., differential clustering of similarly worded items; cf. Marsh, 1996). This competing explanation is particularly salient in view of the fact that the ACQ was originally developed under the premise that it assessed a unidimensional construct (Rapee et al., 1996). Nonetheless, it may be that the psychometric evolution of the ACQ will prove to be similar to that of many other measures of trait vulnerability and psychopathology. For example, although the construct of anxiety sensitivity was initially considered to be unidimensional, psychometric research on the Anxiety Sensitivity Index (ASI; Peterson & Reiss, 1992) has produced consistent findings of multifactorial structure (e.g., Mohlman & Zinbarg, 2001; Zinbarg, Barlow, & Brown, 1997). In addition to prompting new conceptualizations of this construct (cf. Taylor, 1999), these findings have led to research showing that the various ASI dimensions (e.g., physical concerns, social concerns) are differentially relevant to the prediction of outcomes in information processing and biological challenge paradigms (e.g., Carter, Suchday, & Gore, 2001; McNally, Hornig, Hoffman, & Han, 1999; Zinbarg, Brown, Barlow, & Rapee, 2001; Zvolensky & Forsyth, 2002).

Nevertheless, the validity and utility of the three lower-order ACQ dimensions identified in this study await future research. Initial analyses of this nature indicated that each of the ACQ dimensions accounted for significant unique variance in one or both latent factors representing autonomic anxiety and depression, and that these differential relationships were often in the expected direction (e.g., EC and TC were more strongly related to anxiety than depression). However, research is needed to evaluate the concurrent and predictive validity of these dimensions in context of a wide array of disorder-specific outcomes (e.g., latent factors corresponding to the *DSM-IV* anxiety and mood disorders; cf. Brown et al., 1998), laboratory provocations (e.g., predict response to biological panic challenge), and other dimensions of temperament and psychosocial vulnerability. Although subsequent research may show that the ACQ factors are differentially related to the various emotional disorders (e.g., EC and SC to generalized anxiety disorder), it is equally important to demonstrate the discriminant and predictive validity of perceived emotional control in relation to broader dimensions of vulnerability (e.g., neuroticism/negative affect, behavioral inhibition; cf. Brown & Barlow, 2002; Clark et al., 1994; Watson, Clark, & Harkness, 1994) and other closely related, but more experiential, parameters such as perceived predictability and loss of control (cf. Foa, Zinbarg, & Rothbaum, 1992; Mineka & Kelly, 1989; Zvolensky et al., 1999; Zvolensky, Eifert, et al., 2000).

However, the analysis of the higher-order structure of the ACQ supports the use of a broader dimension of perceived control. Indeed, the present findings indicated that each ACQ item had a salient loading on the higher-order perceived control dimension. Indeed, the scale reliability estimate of this higher-order dimension was quite favorable ($\rho = .85$) and in fact was higher than all estimates obtained for the lower-order ACQ dimensions. These results suggest that while it is appropriate to use the ACQ subscale scores in clinical research, a total ACQ score may also be utilized as an indicator of the broader domain of perceived control.

Although important questions remain, the present findings indicate that the revised ACQ's latent structure is defined by three moderately intercorrelated lower-order factors and a single higher-order dimension. Thus, investigators are encouraged to use this revised 15-item version of the ACQ entailing these three, noncollinear factors, as well as a total score reflecting a broader dimension of perceived control.[6] In addition to reconciling measurement difficulties, this strategy should advance our understanding of the meaning and relevance of the theoretically important construct of perceived emotional control.

[6] In future research using the revised 15-item ACQ, it would be helpful to validate the current 3-factor structure to rule out the possibility that the input item covariances used in the present analyses were influenced by the 15 excluded items.

Appendix
Anxiety Control Questionnaire

Listed below are a number of statements describing a set of beliefs. Please read each statement carefully and, on the 0–5 scale below, indicate how much you think *each* statement is typical of *you*.

0	1	2	3	4	5
Strongly Disagree	Moderately Disagree	Slightly Disagree	Slightly Agree	Moderately Agree	Strongly Agree

1. I am usually able to avoid threat quite easily.
2. **How well I cope with difficult situations depends on whether I have outside help. (R)**
3. **When I am put under stress, I am likely to lose control. (R)**
4. I can usually stop my anxiety from showing.
5. **When I am frightened by something, there is generally nothing I can do. (R)**
6. My emotions seem to have a life of their own. (R)
7. There is little I can do to influence people's judgments of me. (R)
8. **Whether I can successfully escape a frightening situation is always a matter of chance with me. (R)**
9. I often shake uncontrollably. (R)
10. **I can usually put worrisome thoughts out of my mind easily.**
11. When I am in a stressful situation, I am able to stop myself from breathing too hard.
12. I can usually influence the degree to which a situation is potentially threatening to me.
13. **I am able to control my level of anxiety.**
14. **There is little I can do to change frightening events. (R)**
15. **The extent to which a difficult situation resolves itself has nothing to do with my actions. (R)**
16. **If something is going to hurt me, it will happen no matter what I do. (R)**
17. **I can usually relax when I want.**
18. **When I am under stress, I am not always sure how I will react. (R)**
19. I can usually make sure people like me if I work at it.
20. **Most events that make me anxious are outside my control. (R)**
21. I always know exactly how I will react to difficult situations.
22. **I am unconcerned if I become anxious in a difficult situation, because I am confident in my ability to cope with my symptoms.**
23. What people think of me is largely outside of my control. (R)
24. **I usually find it hard to deal with difficult problems. (R)**
25. When I hear someone has a serious illness, I worry that I am next. (R)
26. **When I am anxious, I find it hard to focus on anything other than my anxiety. (R)**
27. I am able to cope as effectively with unexpected anxiety as I am with anxiety that I expect to occur.
28. I sometimes think, "Why even bother to try coping with my anxiety when nothing I do seems to affect how frequently or intensely I experience it?" (R)
29. I often have the ability to get along with "difficult" people.
30. I will avoid conflict due to my inability to successfully resolve it. (R)

Note. The 15 items in bold were retained in the final factor models; Emotion Control = items 10, 13, 17, 22, 26; Threat Control = items 5, 8, 14, 15, 16, 20; Stress Control = items 2, 3, 18, 24; (R) = reverse-worded item.

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