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Are We Certain about Which Measure of Intolerance of Uncertainty to Use Yet?

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Abstract

Intolerance of Uncertainty (IU) has been understood as a dispositional tendency to view the presence of negative events as unacceptable and threatening, regardless of the likelihood of those events occurring. The preference over the 12-item vs. 27-item of the IUS has been central to debate. The goals of the present study were to evaluate two competing models of measuring IU with model-fitting analyses and explore model invariance of gender (e.g., men vs. women). A sample of 980 individuals completed an online IUS survey. Results indicated that the two-factor short-form model provided better fit to the data compared to the full-length two-factor model. Results also indicated that the short-form IUS is gender invariant, suggesting acceptable use among men and women. These findings provide further support of a two-factor structure and suggest that the IUS is appropriate for men and women.

Keywords: intolerance of uncertainty, ambiguity, worry, confirmatory factor analysis

Intolerance of Uncertainty (IU) has been understood as a dispositional tendency to view the presence of negative events as unacceptable and threatening, regardless of the likelihood of those events occurring (Carleton, Norton, & Asmundson, 2007; Hong & Lee, 2015). IU has also been defined as "an individual's dispositional incapacity to endure the aversive response triggered by the perceived absence of salient, key, or sufficient information, and sustained by the associated perception of uncertainty" (Carleton, 2016; p. 31). Individuals with high IU tend to seek out additional information to increase their level of certainty about uncertain situations and try to avoid threatening information that leads to symptoms of anxiety (Ladouceur, Talbot, & Dugas, 1997). IU has previously been theorized to result from having a lower threshold for uncertainty perception, stronger responses to ambiguous situations (e.g., additional uncertainty, anxiety and worry), and anticipating future consequences of uncertainty as threatening (Krohne, 1993). A number of studies have measured aspects of IU to understand its relation to distressing symptoms, which has produced variations in how IU is measured.

While IU was originally conceptualized as a specific vulnerability for worry and generalized anxiety, recent research indicates that IU is a broad vulnerability across emotional disorders (Boswell, Thompson-Hollands, Farchione, & Barlow, 2013; Ciarrochi, Said, & Deane, 2005; Einstein, 2014; Gentes & Ruscio, 2011). IU may relate to rituals and compulsions as a function of reducing the distress resulting from uncertainty about a potentially fearful outcome (Steketee, Frost, & Cohen, 1998; Tolin, Abramowitz, Brigidi, & Foa, 2003). IU has been positively correlated with symptoms of social anxiety (Carleton, Collimore, & Asmundson, 2010), panic and agoraphobia (Carleton et al., 2014), and post-traumatic stress (Bardeen, Fergus, & Wu, 2013). Individuals who experience discomfort with uncertainty might opt to negatively evaluate uncertainty about future events, which may place them at higher risk for depression (Dupuy & Ladouceur, 2008; Yook, Kim, Suh, & Lee, 2010). In fact, associations between IU and higher depressive symptoms have been shown for clinical and subclinical samples (Berenbaum, Bredemeier, & Thompson, 2008; Brown & Naragon-Gainey, 2013; de Jong-Meyer, Beck, & Riede, 2009; Liao & Wei, 2011; Norton, Sexton, Walker, & Norton, 2005; Yook et al., 2010). Taken together, IU appears to be a cross-cutting cognitive vulnerability factor for developing emotional problems.

Given the importance of understanding IU across emotional disorders, the measurement of IU has produced variations in factorial structure that has been a focus of debate. The original full-length 27-item IU scale (IUS; Freeston, Rhéaume, Letarte, Dugas, & Ladouceur, 1994) was developed and later translated into an English version. Over the last decade, exploratory factor analytics studies have suggested two-, four-, and five-factor structures to measure IU. Early factorial research on the IUS extracted five interpretable factors (Freeston et al., 1994), while additional evidence suggests a four-factor structure (Berenbaum et al., 2008; Buhr & Dugas, 2002; Norton et al., 2005). Model fitting analyses have found poor support for unifactorial, four-, and five-factor structures as well as a difficulty with interpreting dimensions clearly (Carleton et al., 2007).

Following poor support for a multifactorial and unifactorial structure, a two-factor structure was conceptualized and has frequently been tested. Carleton et al. (2007) developed a 12-item short-form version of the two-factor structure from the original 27 items with factors labeled as *prospective* and *inhibitory anxiety* and shared a strong inter-item correlation (r = .73). Despite attempts to reduce item redundancy, Sexton and Dugas (2009) argued that the short version (12 item) of IUS did not fully represent the construct of IU and inconsistencies between proposed structures were likely due to unreliable and oversampling of factor structures. Therefore, Sexton and Dugas (2009) proposed a full-length two-factor structure of IU having excellent model fit with the following dimensions: (1) *uncertainty has negative behavioral and self-referent implications* and (2) *uncertainty is unfair and spoils everything* (Sexton & Dugas, 2009).

While Sexton and Dugas (2009) have criticized the methodology for developing the IUS short-form, more recent studies with student samples have demonstrated stronger model fit for the short-form compared to the full-length two-factor model (Fergus & Wu, 2012; Helsen, Van den Bussche, Vlaeyen, & Goubert, 2013; Hong, & Lee, 2015). Although Hong and Lee's (2015) results ultimately found greater convergent validity for an alternative 18-item version (two-factor) using a sample from Singapore, support for the two-factor short-form has been consistent in clinical samples as well (Carleton et al., 2012; Jacoby, Fabricant, Leonard, Riemann, & Abramowitz, 2013; McEvoy & Mahoney, 2011). Despite growing support for the two-factor short-form version, direct comparison of the two competing factor structures has only been completed in one western study using a student sample (Fergus & Wu, 2012). Khawaja and Yu (2010) compared both 27- and 12-item versions in an Australian sample and concluded that both measures were psychometrically comparable. Although these findings are encouraging, these results would be strengthened further within a structural equation modeling framework. Therefore, the next appropriate step is for a direct comparison of the 12-item and 27-item measures of IU using a broader demographic sample (non-student) together with modeling fitting analyses.

Given the higher endorsements of emotional symptoms among women compared to men, gender differences in IU may be present (Eaton et al., 2012). To our knowledge, a few studies have found that there were no factorial differences in IU with regard to gender (Carleton et al., 2012; Helsen et al., 2013; Khawaja & Yu, 2010). While Robichaud, Dugas, and Conway (2003) found no gender score differences, findings based on factor scores might mask potential item-level differences that may reveal patterns of responding based on gender. Additionally, factorial gender invariance of the two-factor version of IU has only been evaluated in one non-western sample (Carleton et al., 2012). Therefore, demonstrating item-level gender invariance in a nonclinical and community-based sample would strengthen the debate on the utility of the two-factor IUS.

Although recent research has begun to support the use of the short-form IUS measure, few research studies have compared the short-form measure directly with the full-length measure in western samples (see Einstein, 2014; Fergus & Wu, 2012). Given the debate between the utilization of the short-form vs. full-length IUS measure, psychometric evaluation of these two models in a single study is a needed step to help resolve this debate. Additionally, establishing confirming evidence for factorial gender invariance will help establish the IUS as a valid and unbiased measure in a variety of settings. The present study was designed to: (1) directly compare the IUS-12 (Carleton et al., 2007) to the full-length IUS (Sexton & Dugas, 2009) for the best-fitting model and (2) examine model invariance of gender (e.g., men and women) in a community-based sample.

Method

Participants

The sample of 980 participants was drawn from a larger online sample as described in the Procedures section. Participants' age ranged from 21 to 75 (M = 36.5, SD = 12.25). See Table 1 for additional characteristics.

Table 1. Demographic data for confirmatory sample

	(N = 980)	
Characteristic	*n*	%
Region		
South	350	35.7
Midwest	209	21.3
West	228	23.3
Northeast	182	18.5
Not Reported	11	1.2
Gender		
Men	378	39
Women	593	60
Other Gender	9	1
Sexual Orientation		
Heterosexual	871	89
Bisexual	64	7
Gay or lesbian	29	3
Other Reported	13	1
Not Reported	3	.3
Ethnicity		
European American	745	76
African-American	80	8
Asian American or Pacific Islander	52	5
Multiethnic/other	59	6
Latino	39	4
Education		
Associates Degree or Higher	571	59
Some College or Vocational School	281	28
High School Equivalency or Lower	128	13
Relationship		
Single	473	48
Married, Living Together	437	44
Other	82	8
Employment Hours Weekly		
More than 40	408	42
Fewer than 40	306	31
Not Employed	265	27

Measure

The IU scale (IUS; Carleton, Gosselin, & Asmundson, 2010) is a 27-item scale using a five-point Likert-type response scale of one (*not at all characteristic of me*) to five (*very characteristic of me*). The IUS assesses emotional, cognitive, and behavioral reactions to ambiguous situations, implications of being uncertain, and attempts to control the future. Higher scores indicate greater IU. Sample items include “uncertainty stops me from having a strong opinion” and “uncertainty makes life intolerable.” Buhr and Dugas (2002) reported strong internal consistency for the IUS (r = .94), which was consistent with reliability analysis from the present study (r = .91).

Procedures

A total of 1302 responses were received, but approximately 322 responses were eliminated due to duplicate data (*N* = 176), not currently living in the United States (*N* = 39), younger than age 19 (age of majority in Nebraska; *N* = 13), and responses that fail a "Turing test" (i.e., random responding designed to catch nonhuman or illogical response patterns; "If you are reading this, answer with number 3;" *N* = 94).

Participants were recruited through Amazon's Mechanical Turk system. Mechanical Turk is an online market for labor requests such that requestors post jobs and workers choose jobs to complete for varying pay rates. Research indicates that data collection via Mechanical Turk is at least as reliable as traditional methods and compensation rates do not affect data quality (Buhrmester, Kwang, & Gosling, 2011). All data were collected via online survey. Participants were required to affirm that they were at least 19 years old, not currently in college, and that they had read and electronically signed the informed consent form before beginning the survey. After the survey was completed, participants were presented with a debriefing form and instructed to enter a specific code in order to receive compensation (e.g., $0.10 US). Participation was limited to those who were not currently in college in an attempt to reflect a more community sample given much of the previous work has used college samples. All procedures were approved by the university Institutional Review Board.

Analyses

Confirmatory factor analyses (CFA) were used to: (a) assess the fit of the two-factor models suggested by Sexton and Dugas (2009) and Carleton et al. (2007) and (b) and to assess if the factor structure of the most adequately fitting model, measurements weights (i.e., the relationship between the measured variables and their latent variables), and structural covariances (i.e., the covariances among the latent variables). To assess whether the factor structure differed for men and women, a multiple-group CFA procedure in AMOS as described by Byrne (2001, 2004) was utilized to conduct invariance analysis. Multiple group analysis in structural equation modeling allows comparisons of the same construct across samples for any identified structural equation model simultaneously. Widely used methodology for invariance testing often involves testing a baseline model, which is appropriate because these models do not involve between-group constraints (Byrne, 2004). With a multiple-group approach, invariance testing imposes equality constraints on particular parameters and the data for all groups must be analyzed simultaneously to obtain efficient estimates. Additionally, the pattern of free and fixed parameters remains consistent with the baseline model specification for each group (Byrne, 2004). AMOS (Arbuckle, 1999) allows testing of whether the groups meet an assumption of equality by examining whether different sets of path coefficients are invariant. Statistically significant differences in measurement weights or structural covariances would suggest that the measurement of the respective construct is not comparable across groups. Overall, the procedure for testing multigroup invariance is based on analysis of covariance structures.

Results

Confirmatory factor analysis

Two separate measurement models of IU were examined using CFA in IBM SPSS Amos 19 (Tables 2 and 3). The first model tested the full-length two-factor structure by Sexton and Dugas (2009). This model consisted of 15 items on factor 1 (i.e., 1, 2, 3, 9, 12, 13, 14, 15, 16, 17, 20, 22, 23, 24, 25) and 12 items on factor 1 (i.e., 4, 5, 6, 7, 8, 10, 11, 18, 19, 21, 26, 27). Measures of internal consistency for factor 1 (α = .93) and factor 2 (α = .91) from Sexton and Dugas (2009) were excellent. The second model was from Carleton et al.'s (2007) two-factor short form of the IUS. The model from Carleton et al. (2007) consisted of seven items on factor 1 (i.e., 7, 8, 10, 11, 18, 19, 21) and five items on factor 2 (i.e., 9, 12, 15, 20, 25). Measures of internal consistency for factor 1 (α = .85) and factor 2 (α = .88) from Carleton et al. (2007) were in the acceptable range. For both models, the two factors were allowed to covary given that they were highly correlated and subscales were direct facets of the same construct. Good model fit is indicated by values close to .95 for CFI, GFI, and values equal or less than .06 for RMSEA (Hu & Bentler, 1999; Meyers, Gamst, & Gaurino, 2013; Tabachnick & Fidell, 2001). Although Meyers and colleagues (2013) suggest that RMSEA values equal or less than .08 are considered to have adequate fit.

Table 2. IUS factor structure and item loadings from confirmatory factor models

Sexton and Dugas (2009)			Carleton et al. (2007)		
Item	Factor 1	Factor 2	Item	Factor 1	Factor 2
17	.81		1	.77	
15	.80		6	.74	
9	.79		5	.69	
12	.78		2	.66	
14	.77		7	.62	
25	.73		4	.60	
20	.72		3	.55	
13	.72		10		.80
22	.71		8		.79
24	.69		9		.77
23	.64		12		.74
16	.63		11		.73
3	.58				
2	.44				
1	.43				
26		.81			
6		.80			
7		.78			
27		.77			
5		.74			
19		.69			
8		.64			
18		.64			
4		.60			
21		.56			
11		.55			
10		.48			

Table 3. Univariate summary statistics and item-total correlations of the IUS

Item	M	SD	r_{corr}	Item	M	SD	r_{corr}
1	2.68	1.11	.36	15	2.54	1.16	.77
2	2.08	.96	.36	16	2.71	1.28	.59
3	2.38	1.10	.56	17	2.61	1.18	.80
4	2.30	1.16	.61	18	3.06	1.21	.54
5	2.53	1.25	.68	19	2.72	1.16	.65
6	2.99	1.20	.72	20	2.37	1.17	.67
7	2.80	1.15	.71	21	3.15	1.10	.50
8	3.42	1.08	.58	22	2.38	1.18	.66
9	2.35	1.17	.77	23	2.27	1.21	.60
10	3.37	1.00	.44	24	2.45	1.26	.66
11	2.95	1.15	.53	25	2.22	1.15	.70
12	2.11	1.11	.69	26	2.69	1.19	.78
13	2.00	1.09	.70	27	2.78	1.26	.75
14	2.17	1.09	.70				

Note: IUS = intolerance of uncertainty scale

The full-length IUS model from Sexton and Dugas (2009) produced model fit indices below recommendations, χ^2 (323) = 2043.96, p < .001, GFI = .85, CFI = .89, RMSEA = .07. However, chi-square as a measure of good fit can be unreliable, especially in large samples (Brown & Moore, 2006). Carleton et al.'s (2007) short-form two-factor model produced the same model fit indices near recommended values, χ^2 (53) = 403.27, p < .001, GFI = .94, CFI = .94, RMSEA = .08. While modification indices could be considered to further improve model fit, they were not utilized in order to maintain consistency with previous methodology (Carleton et al., 2007; Sexton & Dugas, 2009).

Given that the two factors are highly correlated, both models above were compared to unidimensional models. As expected, the full-length unidimensional model produced model fit indices lower than recommendations, χ^2 (324) = 2574.58, p < .001, GFI = .78, CFI = .85, RMSEA = .08. The relative fit of the Sexton and Dugas (2009) model was not subjected to comparison to a unidimensional model because of initial model fit indices that were lower than the recommended values for model fit (Hu & Bentler, 1999; Meyers et al., 2013; Tabachnick & Fidell, 2001). The short-form unidimensional model also produced values lower than recommended for adequate fit to the data, χ^2 (54) = 765.84, p < .001, GFI = .85, CFI = .87, RMSEA = .11. In comparing the relative fit to the data between the short-form two-factor model and related unidimensional model, a significant chi-square difference test indicated that the two-factor short-form provided more adequate fit to the data, χ^2 (1) = 158.7, p < .001, ϕ = .40. Between the full-length and short-form two-factor models, relative model fit indices based on AIC and BIC values were compared to evaluate which model fit the data more adequately. As evidenced by smaller values, the short-form model (AIC = 455.27; BIC = 577.46) fit the data more adequately compared to the full-length model (AIC = 2153.96; BIC = 2422.78).

Testing model invariance of IUS

Given stronger evidence for good model fit to the data of the short-form two-factor model of the IUS (Carleton et al., 2007), the short-form was reassessed in IBM Amos 19 for whether or not the confirmatory factor structure was invariant across gender. Self-identified men (n = 378) and women (n = 593) were included in the analysis. Participants (n = 8) who self-identified as other gender were removed from the present analysis due an inadequate sample size for this group.

The multiple group analysis evaluates the difference between an unconstrained model, which assumes that the groups are yielding different estimates of the parameters and a constrained model, which assumes the groups are yielding equivalent estimates of the parameters when the model is applied to the data. Two model comparisons were completed. The first comparison that included only the measurement weights (i.e., pattern coefficients) was not statistically significant, χ^2 (10) = 13.83, $p > .05$. The second comparison including the structural covariances (i.e., combined factors of path coefficients and variance/covariance of the factors) was also not statistically significant, χ^2 (13) = 21.81, $p > .05$. In summary, the two-factor short-form model of the IUS appeared to be invariant between males and females.

Discussion

The first aim of this study was to compare the two-factor structures from the short-form IUS (Carleton et al., 2007) and full-length IUS (Sexton & Dugas, 2009). Greater psychometric support was found for the short-form vs. the full-length IUS using model, which was consistent with Fergus and Wu (2012). Both unidimensional models of the 12-item and 27-item measures did not result in adequate fit to the data. In fact, the two-factor 12-item model fit the data significantly better than the 12-item unidimensional structure. Within the second aim of the study, we found that the 12-item two-factor model (Carleton et al., 2007) was factorially invariant with regard to gender. Evidence from multiple group invariance findings further supports the use of the IUS-12, suggesting no particular dimension bias based on gender.

Evidence for supporting the two-factor structure of the IUS has implications, given the debate between the utility of a short-form or full-length measure (see Birrell, Meares, Wilkinson, & Freeston, 2011; Einstein, 2014). Results from Khawaja and Yu (2010) indicated that the IUS-27 had slightly better reliability, but that both measures were equally effective. While these findings are encouraging, reliability estimates alone are not an accurate measure of internal consistency, as additional information is needed (see Sijtsma, 2009). Our results are consistent with Fergus and Wu (2012) that indicate that the IUS-12 significantly yielded better fit to the data compared to the IUS-27 in a western sample. Research using non-western samples have also supported the use of the IUS-12 over the full-length measure (Helsen et al., 2013; Hong & Lee, 2015). The present study attempted to address concerns about the development of the IUS-12 as articulated by Sexton and Dugas (2009). That is, we intentionally recruited a sufficient sample size in order to achieve greater reliability in our factor solutions. Interestingly, some items on the IUS-27 appear to have greater specificity toward generalized anxiety disorder and worry, although these items were removed on the IUS-12 (Gentes & Ruscio, 2011). Overall, the present findings contribute to the literature by directly evaluating two competing IUS measures and providing support for the IUS-12.

Given the disproportionate level of emotional problems endorsed by women (Eaton et al., 2012), the potential difference in gender was examined in the measurement of IU. Our findings add to the literature by demonstrating that the IUS-12 was invariant for men and women (Carleton et al., 2012). Previous research has either not tested item-level invariance in a confirmatory analytic framework (Robichaud et al., 2003) or has tested gender invariance in non-western samples (Helsen et al., 2013; Khawaja & Yu, 2010). Item-level invariance testing may reveal patterns of responding that is represented by gender, which might otherwise be ignored by compared average scores. While the present study was limited by inadequate sample sizes for testing invariance across other demographic variables (i.e., ethnicity, sexual orientation, and education), future research should explore these relationships. Findings from invariance analysis suggest that researchers and clinicians can be more confident that patterns of responding will be similar with regard to gender. Therefore, the present study further supports the use of the IUS-12 without specific concern for measurement bias with regard to gender.

The present study has particular strengths that improved generalizability of the findings. First, utilizing Amazon's Mechanical Turk as vehicle for research solicitation allowed recruiting beyond a specific geographic area. Second, excluding college students results in a community sample with a broader range on demographic variables than in the typical college student sample in much of the nonclinical literature on IU. However, the sample was also largely Euro-American, as is typical of samples collected via the Internet (Buhrmester et al., 2011). Limited representation of ethnic groups limits generalizability to other cultural groups. Future research should include samples that are more representative of multicultural society, especially because there is already known effects for Euro-Americans and African-Americans (e.g., Fergus & Wu, 2012).

One limitation of the present study was the inability to present evidence of construct validity as another measure of the appropriateness of the IUS-12 and IUS-27. Given the established literature on construct validity for IU (see Birrell et al., 2011; Carleton, 2016), data collection for the present study did not take this into consideration. We also acknowledge limitations in the use of "Turing tests" for eliminating responses that could decrease the quality of the data. Some researchers note that screening methods that flag inattentive responses may have high measurement error as such methods rely on the "questionable assumption that measured attentiveness is constant throughout the task and may tap into other correlated traits" (Paolacci & Chandler, 2014, p. 186). Unfortunately, it is difficult determine the nature of Turing test responses (i.e., correlated traits rather than state-level differences in attentiveness) in the absence of additional information.

Interest in the measurement of IU has grown considerably, as evidenced by the proliferation of factor analytic studies and other specific measures of IU (see Einstein, 2014). For example, the IU inventory (Gosselin et al., 2008), disorder-specific IU (Thibodeau et al., 2015), and a situation-specific measure of IU (McEvoy & Mahoney, 2011) have been recent developments in the measurement of IU. The next step is to understand the incremental explanatory power of these measures, although some research has already begun to this end (Fergus, 2013). The current pragmatic shift from diagnosis-driven research to an emphasis on cross-cutting constructs such as IU will benefit from precise measurement of constructs of interests. However, some attempts to measure IU as a transdiagnostic measure have taken on a pragmatic approach to investigating the explanatory power of IU by splitting aspects of IU to specific nosology (Thibodeau et al., 2015). Given evidence of shared pathology across many emotional disorders, disorder-specific measurement of IU may inherently discount aspects of IU that are important across emotional disorders. The present study supports the measurement of IU as a transdiagnostic construct across emotional disorders, especially the strength of the IUS-12 in accomplishing this aim.

Taken together, the present study contributes to the growing support of the IUS-12 measure and its appropriateness with regard to gender, especially as preference over the full-length IUS has been central for debate. Researchers and clinicians can be more confident when using the IUS to further understand the role of IU across emotional disorders and, ultimately, further develop interventions to help remediate those disorders.

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