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Validation of the Injustice experiences questionnaire in a heterogeneous trauma sample

Short title: **IEQ in Trauma**

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Abstract

Purpose/Objective: A recent study by Trost et al. (2015) investigated the influence of perceived injustice—reflecting appraisals of the severity and irreparability of loss following injury, blame, and unfairness—on physical and psychological outcomes in a sample of patients 12 months after sustaining traumatic injury. This brief report examines the psychometric properties of the Injustice Experiences Questionnaire (IEQ) using the previous sample from Trost et al. (2015) with added trauma patients (total N = 206).

Research Method/Design: Primary analyses included confirmatory and exploratory factor analyses to validate the measurement model of the IEQ in patients 12-months after traumatic injury. Reliability analyses were conducted and construct validity was assessed by examining associations between the IEQ and other pain-related, psychological, and health-related outcome variables of interest.

Results: Results replicated both one- and two-factor structures from past research, with a high factor correlation in CFA analyses and cross-loadings in EFA analyses. Item characteristics analysis demonstrated overall strong internal consistency ($\alpha = .95$). In addition, significant associations with psychosocial variables provide further support for the utility of measuring injustice perception using the IEQ in a trauma sample.

Conclusions/Implications: The IEQ shows strong psychometric properties and is suitable for use in a sample of diverse traumatic injury. However, results suggest the use of a one-factor model for the IEQ in this sample. Future trauma and rehabilitation research can use the IEQ to explore how injustice perceptions related to traumatic injury can prospectively influence physical and psychological outcomes.

Keywords: injustice perception, traumatic injury, trauma, outcomes, psychoassessment

Injustice Perception in a Trauma Sample: Validation and Psychometrics of the IEQ

Impact:

- Recent research has found that injustice perception as measured by the Injustice Experiences Questionnaire (IEQ) predicts physical and psychosocial outcomes one year after traumatic injury. This brief report presents a psychometric evaluation of the IEQ in a similar heterogeneous sample of patients with diverse traumatic injury.
- This report compares IEQ scores of patients with traumatic injury to one- and two-factor structures previously validated in patients with chronic pain and musculoskeletal injuries.
- Future trauma and rehabilitation research can explore how injustice perceptions related to traumatic injury can prospectively influence physical and psychological outcomes.

Recent research on traumatic injury has focused on psychosocial variables that may impact individuals' recovery, specifically those influencing the onset of depression and post-traumatic symptoms following injury (Bryant et al., 2010; Zatzick et al., 2013; deRoon-Cassini, Mancini, Rusch & Bonanno, 2010; Shih et al., 2012; Powers et al., 2014). One such variable of interest has been perceived injustice, which describes an appraisal reflecting the severity and irreparability of injury-related loss, blame, and unfairness (Sullivan et al., 2008).

To date, injustice perception has been assessed using the Injustice Experiences Questionnaire (IEQ); however, the measure is formally validated only in samples with musculoskeletal injuries to the back, neck, and knee, is primarily associated with pain-related constructs (Sullivan et al., 2009a, Sullivan et al., 2009b, Scott et al., 2013, Yakobov et al., 2014) and only recently with posttraumatic stress (Sullivan et al., 2009a; Scott & Sullivan, 2012). In Sullivan et al. (2008), exploratory factor analysis (EFA) of the IEQ revealed a two-factor structure determined to reflect appraisals of *(i) severity/irreparability of loss* and *(ii) blame/unfairness*. However, subsequent validation studies have shown that these factors can be highly correlated, with considerable cross-loadings among items in the IEQ (Kennedy & Dunstan, 2014; Rodero et al. 2012). Only one previous study has analyzed the IEQ using a confirmatory factor analysis (CFA) in a sample of individuals with musculoskeletal work injuries (Kennedy & Dunstan, 2014). The results of this study indicated that once post-hoc model fitting was applied via modification indices, both a two-factor and one-factor model demonstrate good fit. However, the authors ultimately supported the use of the two-factor model after subsequent differentially testing the subscales with measures of psychological distress (Kenney & Dunstan, 2014).

A recent study looking at physical and psychological outcomes among individuals admitted to a Level-I trauma center identified injustice perception as a significant predictor of greater pain intensity, depression, and post-traumatic symptoms 12 months after injury, even after controlling for relevant demographic and injury-related variables (Trost et al., 2015). These findings suggested that evaluating perceived injustice may be warranted in future studies of outcomes following traumatic injury. Initial checks of internal consistency in Trost et al. (2015) showed psychometric properties similar to those of previous studies; however, further validation of the IEQ in a heterogeneous trauma sample has yet to be formally conducted.

Accordingly, the current study had two primary aims: 1) to validate the measurement model of the IEQ proposed by Sullivan et al. (2008) and confirmed by Kennedy & Dunstan (2014) in a heterogeneous sample of patients admitted to a Level -1 trauma center, and 2) to test the reliability and validity of the IEQ in the same sample. A third aim was to examine the associations between IEQ scores and measures of pain, injury-appraisal measures, psychological distress, and health-related quality of life (HRQOL) to evaluate its clinical relevance. Greater understanding of the psychometric properties of the IEQ in a trauma population would further support the utility of this measure in research and clinical practice targeting individuals with a range of traumatic injuries.

Method

The current paper represents a secondary analysis of data examined in Trost et al. (2015). Recruitment and study procedures are comprehensively described in Trost et al. (2015). Study procedures were approved by the hospital's Institutional Review Board.

Participants

The original study by Trost et al. (2015) analyzed data from 158 participants admitted to a Level-I trauma center in the southwest United States between March 2012 and June 2014; participants completed measures 12-months following hospital discharge. In this study, an additional 48 participants completed 12-month follow-up assessment for a total sample of N = 206. Patients provided consent after initial hospitalization given the following inclusion criteria: admission to the trauma service > 24 hours, 18 years of age or older, and ability to provide at least one contact for follow-up. Exclusion criteria included inability to comprehend English or Spanish and/or the presence of cognitive deficits (i.e., screening for dementia and/or severe traumatic brain injury using the Cognistat screening tool; Kiernan, 1987) that impaired ability to provide informed consent. Immediately following enrollment, participants completed demographic and injury-related information. Missing demographic or injury-related information were gathered via chart review or the hospital's trauma registry.

Follow-up at 12-months

At 12 months (± 2 months) following admission, participants completed standardized measures administered by telephone. Reminders were sent one month prior to participants' 12-month follow-up date. During the 4-month follow-up window, researchers attempted to contact the participant a maximum of twelve times.

Measures at 12-month follow-up

Perceived Injustice. Perceived injustice was assessed using the Injustice Experience Questionnaires (IEQ), a 12-item measure that asks respondents to indicate the frequency with which they experience thoughts in relation to their injury using a 5-point Likert-type scale from 0 (never) to 4 (all the time). Representative items reflect elements of blame ("I am suffering because of someone else's negligence"), magnitude of loss ("I feel as if I have been robbed of something very precious"), irreparability of loss ("My life will never be the same"), and unfairness ("It all seems so unfair"). IEQ scores ranged from 0 to 48, with higher scores indicating higher perceptions of injustice. The IEQ has demonstrated good psychometric properties in samples with musculoskeletal injury (Scott, Trost, Milioto, et al., 2013; Sullivan et al., 2008).

Pain and injury-appraisal measures. A numeric rating scale *(NRS)* was used to measure participants' pain intensity; participants rated average pain from 0 (no pain at all) to 10 (worst possible pain) over the past two weeks. Higher scores indicated higher pain experience (Downie et al., 1978). Negative orientation toward pain experience was assessed using the *Pain Catastrophizing Scale (PCS)*, a 13-item measure of the degree to which individuals magnify, ruminate about, and feel helpless in the face of pain (Sullivan, Bishop, & Pivik, 1995). Lastly, fear of movement and (re)injury associated with pain was examined using the 13-item *Tampa Scale for Kinesiophobia (TSK*; Neblett et al., 2015*).*

Psychological distress measures. Depressive symptomatology was examined using the *Patient Health Questionnaire (PHQ-8)*, derived from the PHQ-9 with the question regarding suicide removed (Kroenke et al., 2009). Participants rated symptoms of depression over the previous two weeks from 0 (not at all) to 3 (nearly every day). The PHQ-8 has been validated in medical settings as a screening instrument for depression (Gilbody et al., 2007; Kroenke et al., 2010) as well as in samples of traumatic injury (Fann et al., 2005). Presence of posttraumatic symptomatology was assessed using the *Primary Care PTSD screen (PC-PTSD;* Prins et al., 2003). Presence of at least 3 out of 4 symptoms was considered a positive screen for clinical PTSD symptoms (Hanley, deRoon-Cassini, & Brasel, 2013).

Health-Related Quality of Life. HRQL was assessed using the *Veterans RAND 12-Item Health Survey (VR-12)*, consisting of 12 items assessing physical and mental domains of health (i.e., VR-12 Physical and VR-12 Mental, respectively), with lower scores indicating diminished HRQOL (Selim et al., 2011).

Resilience. To assess the relationship between injustice perception and other variables of interest, a measure of resilience was included in the current analysis. Resilience, defined as the ability to maintain stable and healthy psychological function following exposure to an adverse event, was examined using the *Connor-Davidson Resilience Scale (CD-RISC;* Rainey et al., 2014), a 10-item measure assessing trait resilience on a 5-point scale (Connor & Davidson, 2003). Prior research in the trauma context suggests that resilience may reflect dispositional qualities (e.g., reduced negative affect and cognition) that may show a negative association with injustice perception (White, Driver, & Warren, 2010).

Data Analytic Plan

First, a CFA was conducted to assess goodness of model fit based on the factor structure established by Sullivan et al. (2008) and validated in Kennedy & Dunstan (2014). Data was screened for normality using an estimated asymptotic covariance matrix/polychoric correlation matrix (ACM/PCM) and then subjected to a two-factor model (Jöreskog,1990). Global indices of model fit and modification indices were evaluated to determine if the previous structural model was appropriate for the current sample (see Schreiber et al., 2006 for review). A one-factor structure was also explored to examine comparative fit. Based off of standard estimates determined by MacCallum, Brown and Sugwara (1996), power from these analyzed factor solutions would be between .608-.769 and adequate for comparison to past studies.

In the case of poor model fit, dimensionality of the IEQ at 12 months was re-assessed using an exploratory factor analysis (EFA). Due to latent variables believed to exist, Principal Axis Factoring was initially used. However, to compare past studies validating the IEQ via EFA, principal components analysis was also used in iterative analyses. Both methods were followed with oblique (direct oblimin) rotation to permit correlation between factors (Rodero et al., 2012; Sullivan et al., 2008).

We used several methods to examine validity and reliability of the IEQ in our trauma sample. Internal consistency of the IEQ was determined using Cronbach's alpha and item-total correlation coefficients. Convergent validity for the IEQ was examined by observing associations between the scale and pain and injury-appraisal questionnaires (i.e., NRS, TSK, PCS), and divergent validity with resilience (i.e. CD-RISC). Construct validity for the IEQ was examined by observing associations with HRQOL and psychological distress (i.e. VR-12, PHQ-8, and PC-PTSD). For all associations, bivariate correlations to continuous data and non-parametric tests for categorical data were used. Assumptions and matrix transformations were conducted using

PRELIS (Jöreskog,1990). Confirmatory factor analyses were conducted using LISREL (Jöreskog and Sörbom, 1993). All remaining analyses were conducted using SPSS software.

Results

Participant Characteristics

Demographic and injury-related variables of the sample (N = 206) are reported in Table 1. At the time of analysis, the retention rate was 41%. Participants who completed the study at 12 months tended to be significantly older ($M_{\text{diff}} = 5.47$, $p < .001$), have a college degree (χ^2 (8)= 20.97, $p = .007$), a higher income (i.e., > $50K, χ^2 (4) = 13.16, $p = .011$), and possess private insurance (χ^2 (3) = 25.42, $p < .001$), and sustained injury from falls, motorcycle collisions (MCC), automobile vs. pedestrian collisions, and animal attacks; by contrast, those who did complete the study tended to sustain injuries from motor vehicle collisions (MVC) and gunshot wounds (GSW; χ^2 (11) = 23.14, p =.017). All other demographic and injury-related variables showed nonsignificant differences between those who completed the study and those who did not.

IEQ Scores and Assumptions

To conduct the 2-factor CFA, a Pearson-r correlation matrix was produced containing all twelve items (see Table 2). Means and standard deviations are also reported. The overall mean score on the IEQ was 16.74 (*SD* =14.92), comparable (i.e., within 1 SD) to IEQ total scores obtained in previous studies (e.g., *M* = 19.86; Scott et al. 2013). The mean score on the subscale *severity/irreparability* was 9.46 (*SD* = 7.74) and the mean score on subscale *blame/unfairness* was 5.82 (*SD* = 6.45).

Maximum likelihood (ML) was used to conduct the CFA; all required assumptions, including adequate sample size, indicator measurement, and multivariate normality were addressed (Brown, 2006). Because of the Likert-scale nature of the data, a polychoric correlation matrix was produced. Skewness (z = 21.05) and kurtosis (z = 13.00) values for multivariate normality indicated significant non-normality (both $p < .001$; Brown, 2006), therefore analyses were conducted using an ACM derived from PRELIS.

Analysis 1. Validating the factor structure and construct validity of IEQ in a Level-I trauma sample.

2-factor model CFA. Standard global indices to determine overall model fit are shown in Table 3 (see Schreiber et al., 2006 for review). Although CFI and SRMR fit indices met acceptable standards (>.95 and < .08, respectively), values for GFI (>.95) and RMSEA (< .05) indicated an overall poor fit for the structural model. Figure 1 shows the initial path diagram for the two-factor structure. Though each item loaded onto its respective factor sufficiently ($\lambda > .40$; Tabachnick & Fidell, 2001), the high covariance between factors *severity/irreparability* and *blame/unfairness* (r = .93) indicated that the model could be more parsimonious.

Further examination of modification indices showed large standardized residuals, confirming significant variance left unexplained by the model. In addition, modification indices suggested loading specific items on different factors— specifically, IEQ1 (“Most people don’t understand how severe my condition is”) and IEQ8 (“I worry that my condition is not being taken seriously”) loading onto the blame factor would create a large expected change in factor loading (λ = -.489 and .838, respectively).

One-factor model CFA. Table 3 also shows the fit indices for the re-evaluated one-factor model. The overall fit of the one-factor model indicated similar issues as the two factor model: CFI and SRMR continued to meet acceptable standards, while values for GFI and RMSEA still indicated a poor fit.

Figure 2 shows the path diagram for the one-factor structure. Examination of residuals and modification indices suggested correlating error variances for IEQ7-IEQ9, IEQ6-IEQ7, and IEQ6-IEQ2. Interestingly, the high modification index for IEQ6-IEQ2 was also observed in the CFA conducted by Kennedy & Dunstan (2014) and ultimately added to the model, as both items directly address the perceived permanence of participants' injury. In addition, inclusion of this specific error covariance resulted in better fit.

Analysis 2. Re-assessing factor structure through exploratory factor analysis.

Due to issues of fit, an initial EFA was conducted by requesting a two-factor structure as suggested by Sullivan et al. (2008). Principal axis factoring was used with oblique rotation due to the high correlation between the latent factors in the previous analysis. An initial EFA requesting a two-factor structure yielded 66.37% variance explained, with coefficients detailed in Table 4. Significant cross-loadings existed for IEQ4, IEQ8, IEQ10, IEQ11, and IEQ12, departing slightly from the initial cross-loadings described in Sullivan et al. (2008). A one-factor EFA model was then run for comparison, yielding 65.23% variance explained and showing high loadings and communality coefficients for all items, with the exception of IEQ3 (see Table 4). A CFA was run without the IEQ3 to observe any improvement in fit; little change was observed in global fit indices (e.g. ΔGFI = .080, ΔRMSEA < .004).

Analysis 3. Reliability and construct validity of the IEQ.

The IEQ showed strong internal consistency (α = .95). Table 4 shows the item-total correlation coefficients for the scale. Values for all item-deleted alphas remained stable; however, IEQ3 ("I am suffering because of someone else's negligence") was the only item showing a higher value. Associations between the IEQ, NRS, PCS, TSK, and CDRISC were examined to demonstrate convergent or divergent validity, while associations with HRQoL and psychological distress were computed to demonstrate construct validity. Table 5 shows associations between IEQ, pain-related measures and other psychosocial measures at 12-month post-injury. As expected, higher IEQ scores were significantly correlated with higher pain-intensity, higher TSK and PCS scores, as well as higher scores for the PHQ and PC-PTSD. Also as expected, higher IEQ scores were significantly associated with lower scores on the CD-RISC, and both VR-12 Physical and Mental subscales.

Discussion

This study is the first to examine the psychometric properties of the IEQ in a heterogeneous sample of trauma patients 12-months after injury. Using a series of CFA and EFA modeling, we replicated results from the original model proposed by Sullivan et al. (2008) as well as the measurement models confirmed by Kennedy & Dunstan (2014) in support of both one- and two-factor measurement models. Though neither model initially provided good fit, the error covariance suggested by Kennedy & Dunstan (2014; between IEQ2 and IEQ6—both of which address belief that an injury is not adequately acknowledged by others) improved model fit for both the one- and two-factor structures. Although the two-factor model structure was supported, the high correlation between the two factors (i.e. severity/irreparability of loss and blame/unfairness) suggests that a single factor model may better represent the relationship between items in this sample, with findings from subscales interpreted with caution.

Additional analyses using EFA found significant departures from loadings proposed by Sullivan et al. (2008). Specifically, in the current sample items IEQ4, IEQ5 and IEQ8 showed high cross-loadings not demonstrated in past research, which further suggests that the single-factor model is more appropriate (Sullivan et al., 2008). In addition, IEQ3 showed poorer item characteristics (e.g., communality coefficients, item-deleted internal consistency values) than other questionnaire items. Given the content of IEQ3 (“I have been robbed of something very precious”), this item may reflect a more explicit blame attribution in comparison to the more indirectly/implicitly-worded items in the rest of the measure.

In addition, modification indices in Analysis 1 suggested covariances between numerous items across factors, indicating redundancies in the IEQ. For example, high covariances between IEQ7 with IEQ4 and IEQ6 in the 2-factor model, and between IEQ6 and IEQ9 in the one-factor model, indicate the possibility of item exclusions that could improve model fit without affecting the strong reliability of the measure as a whole (see Table 4 for item-deleted alpha coefficients) Especially for clinical samples, a shorter/more concise measure may be advantageous in that perceptions of injustice could be screened quickly within routine clinical assessment and care.

This study was limited by several factors. One concern is the possibility of an underpowered analysis, despite the suggested power estimates determined by McCallum et al. (1996); the ratio between sample size and estimated parameters was 17:1, below the recommended 20:1 for strong SEM analyses (Jackson et al., 2003). However, analyses by Sullivan et al. (2008) and Kennedy and Dunstan (2014) were conducted with either similar or significantly smaller sample sizes, making these findings comparable to previous psychometric evaluation. Further, analysis of data was cross-sectional in nature, 12 months after the injury occurred. A longitudinal design coupled with path analytic modeling could assess injustice perception before related outcomes, providing the predictive validity absent in this study. Moreover, assessing injustice before related outcomes can provide stronger findings regarding the stability of the IEQ factor structure. In addition, this is the first sample reflecting diverse sources of injury (e.g., lacerations, blunt trauma, penetrating injury) without a single etiology (e.g., injuries resulted from falls, motor vehicles, violent crime etc.). The heterogeneity of injury type and etiology in the current sample may have contributed to differential model fit and factor loadings as compared to previous studies which have examined more homogenous patient populations. Accordingly, further analyses might examine the factor structure of the IEQ in subsamples of patients with more similar injury characteristics.

Overall, the findings of this study support the use of the IEQ to measure perceived injustice in a diverse trauma sample. In particular, this study supports existing evidence that injustice perception is a unique correlate of physical and mental well-being during recovery from trauma (Trost et al., 2015, Monden et al., 2016). As such, this suggests that the IEQ could be a useful patient tool for assessing individual perception of overall adjustment to physical injury; future clinical research should empirically examine the influence of injustice perception immediately after injury in relation to longer-term follow-up outcomes. If future research identifies the IEQ as a prospective predictor of poor recovery in diverse trauma sample, then the IEQ could be used to identify at-risk individuals. Future research will also need to develop and evaluate interventions to facilitate recovery for individuals with high levels of perceived injustice.

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Table 1. Sample Characteristics (N = 206)

Variable	**N (%) or Mean (SD)**
Demographic Variables	
Age	47.48 (17.11), 18-68
Gender	
Female	83 (40.3%)
Male	123 (59.%)
Ethnicity	
Non-Hispanic	90 (56.9%)
Hispanic	33 (20.9%)
Other	23 (14.6%)
Missing/Unobtainable	11 (6.9%)
	1 (.6%)
Race	
Caucasian	150(72.8%)
Black	43(20.9%)
Native American/Alaskan Native	3(1.5%)
Asian	1(.5%)
Multiracial	8(3.9%)
Unobtainable	1(.5)
Marital Status	
Never Married	68 (33.0%)
Married	77 (37.4%)
Divorced	37 (18.0%)
Separated	7 (3.4%)
Widowed	15 (7.3%)
Missing/Unobtainable	2 (1.0%)
Education Level	
Less than HS	35 (17.0%)
HS Diploma	82 (39.8%)
Associate’s	25 (12.1%)
Bachelor’s	37 (18.0%)
Graduate/Professional	24 (13.2%)
Missing/Unobtainable	3 (1.5%)
Income	
<25K	49 (23.8%)
25-49K	23 (11.2%)
50-74K	35 (17.0%)
75+K	48 (23.3%)
Missing/Unobtainable	51 (24.8%)
Insurance	
No Insurance	41 (19.9%)
Public	29 (14.1%)
Private	68 (33.0%)
Public/Private	21 (10.2%)
Missing/Unobtainable	47 (22.8%)

Cause of injury	
Fall	57 (27.7%)
Machine	3 (1.5%)
Stab	3 (1.5%)
Motor Vehicle Collision (MVC)	42 (20.4%)
Bicycle	7 (3.4%)
Gun Shot Wound	13 (6.3%)
Aggravated Assault	13 (6.3%)
Motorcycle Collision (MCC)	32 (15.5%)
Automobile vs Pedestrian	16 (7.8%)
Dive	1 (<0.5%)
Animal	8 (3.9%)
Other	4 (1.9%)

Table 2. Correlation matrix between IEQ items.

	Mean	SD	IEQ1	IEQ2	IEQ3	IEQ4	IEQ5	IEQ6	IEQ7	IEQ8	IEQ9	IEQ10	IEQ11	IEQ12
IEQ1	1.44	1.39	1											
IEQ2	1.69	1.57	.695**	1										
IEQ3	1.07	1.51	.428**	.408**	1									
IEQ4	1.26	1.59	.617**	.591**	.533**	1								
IEQ5	1.88	1.66	.680**	.635**	.357**	.644**	1							
IEQ6	2.13	1.63	.642**	.716**	.432**	.617**	.671**	1						
IEQ7	1.16	1.51	.628**	.597**	.559**	.715**	.672**	.569**	1					
IEQ8	1.06	1.44	.681**	.591**	.397**	.587**	.585**	.540**	.727**	1				
IEQ9	1.20	1.56	.593**	.593**	.519**	.623**	.658**	.593**	.810**	.735**	1			
IEQ10	1.43	1.59	.679**	.682**	.565**	.653**	.702**	.696**	.762**	.682**	.825**	1		
IEQ11	.96	1.45	.612**	.597**	.369**	.542**	.593**	.546**	.654**	.627**	.651**	.688**	1	
IEQ12	1.47	1.60	.618**	.610**	.476**	.573**	.677**	.628**	.758**	.613**	.661**	.717**	.605**	1

Table 3. Fit indices for all-tested CFA models.

Model Name	CFI	GFI	RMSEA	90% RMSEA	SRMR	χ^2	df
2-Factor CFA	.975	.869	.111	.094; .128	.040	186.74	53
2-Factor CFA w/ error covariance	**.971**	**.881**	**.106**	**.089; .124**	**.038**	**172.53**	**52**
1-Factor CFA	.968	.843	.124	.107; .141	.044	224.05	54
1-Factor CFA w/ error covariance	.972	.866	.115	.098; .132	.041	197.22	53

Table 4. Reliability Analyses and EFA (2- and 1-Factor) Coefficients

Item	Reliability Analyses		2-Factor Structure			1-Factor Structure	
	Corrected Item-Total R	Alpha w/ Item Deleted	Comm.	Blame	Severity	Comm.	Injustice
IEQ1	.775	.946	.676	.441	**.694**	.664	**.815**
IEQ2	.756	.947	.706	.363	**.758**	.637	**.798**
IEQ3	.556	.953	.346	**.513**	.289	.372	**.610**
IEQ4	.753	.947	.589	.564	**.520**	.630	**.794**
IEQ5	.776	.946	.659	.484	**.652**	.669	**.818**
IEQ6	.749	.947	.691	.358	**.751**	.625	**.791**
IEQ7	.846	.944	.875	**.858**	.374	.768	**.876**
IEQ8	.760	.947	.624	.637	**.467**	.650	**.806**
IEQ9	.822	.944	.795	**.798**	.398	.737	**.858**
IEQ10	.872	.943	.800	**.691**	.568	.805	**.897**
IEQ11	.727	.947	.558	**.548**	.507	.601	**.775**
IEQ12	.783	.946	.644	**.605**	.527	.677	**.823**

Table 5. Associations between IEQ scores, pain-related and quality of life variables at 12-month follow-up.

	IEQ	Age	TSK	PCS	Pain	CDRISC	VR-12 P	VR-12 M	PHQ	PC-PTSD
IEQ	1									
Age	-.201**	1								
TSK	.704**	-.117	1							
PCS	.667**	-.075	.577**	1						
Pain	.624**	-.014	.607**	.540**	1					
CDRISC	-.413**	-.001	-.375**	-.387**	-.380**	1				
VR-12 Physical	-.404**	-.192*	-.532**	-.347**	-.610**	.327**	1			
VR-12 Mental	-.617**	.070	-.497**	-.567**	-.511**	.509**	.199*	1		
PHQ	.681**	-.086	.561**	.548**	.588**	-.532**	-.445**	-.786**	1	
PCPTSD	.602**	-.182**	.505**	.513**	.497**	-.466**	-.256**	-.642**	.670**	1

Figure 1. Path diagram for the IEQ two-factor structural model with error covariance.

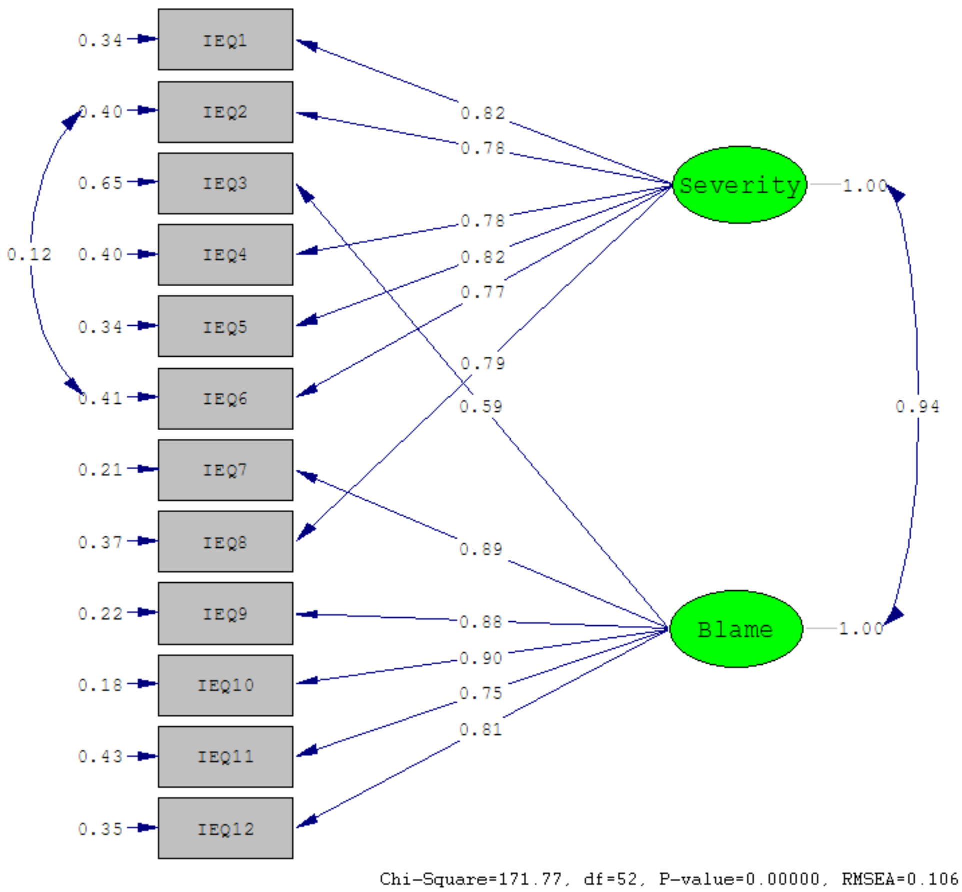

Figure 2. Path diagram for the IEQ one-factor structural model with error covariance.