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The Assessment of Factorial Invariance in Need for Cognition Using Hispanic and Anglo Samples

SCOTT E. CULHANE and **OSVALDO F. MORERA**
Department of Psychology, The University of Texas at El Paso

P. J. WATSON
Department of Psychology, The University of Tennessee at Chattanooga

Abstract

The Need for Cognition Scale–Short Form (NCS–SF; J. T. Cacioppo, R. E. Petty, & C. F. Kao, 1984) is a commonly administered measure in the behavioral sciences, but little research has assessed its applicability across cultures. A sample of undergraduates in the southeastern United States and a sample of undergraduates at a southwestern U.S. university completed the NCS–SF. Hispanic respondents did not differ from Anglos in their mean NCS–SF scores. Confirmatory factor analysis revealed that factor parameter estimates and item intercepts were partially measurement invariant across samples.

Keywords

factorial invariance; need for cognition

The need for cognition scale–short form (NCS–SF; Cacioppo, Petty, & Kao, 1984) is an 18-item instrument for measuring “an individual’s tendency to engage in and enjoy thinking” (Cacioppo & Petty, 1982, p. 116). A person who scores high on need for cognition tends to generate more thoughts and to elaborate more on presented information, whereas a person who scores low tends to avoid cognitive effort. Findings by Cacioppo et al. (1984),Sanders, Gass, Wiserman, and Bruschke (1992), and Culhane, Morera, and Hosch (2004) estimated coefficient alphas of .90, .88, and .86, respectively, suggesting high internal consistency for the measure.

Need for cognition is a popular construct, as a PsycINFO search via the OVID platform revealed 30 references on the topic in 2003 alone. Need for cognition is also a construct with international appeal. The items in the NCS–SF have been translated into German (Bless, Waenke, Bohner, Fellhauer, & Schwarz, 1994), Turkish (Guelgoez & Sadowski, 1995), Spanish (Gutierrez, Bajen, Sintas, & Amat, 1993), French (Ginet & Py, 2000), Chinese (Kao, 1994), and Persian (Ghorbani, Watson, Bing, Davison, & LeBreton, 2003). Because need for cognition is frequently used across diverse cultural groups, it is important to ensure that the NCS–SF is reliable and valid for all samples.

Unfortunately, few researchers have made cross-cultural comparisons of the scale. Sanders et al. (1992) observed that Asian Americans had statistically smaller scores than did either Anglo or Hispanic respondents. Culhane et al. (2004) did not find mean differences between the

Address correspondence to Scott E. Culhane, who is now at the Department of Criminal Justice, University of Wyoming, P.O. Box 3197, Laramie, WY 82071.

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Hispanics from their sample and previously reported scores for Anglos (Haugtvedt & Petty, 1992). The absence of mean differences is not a sufficient condition for demonstrating a lack of ethnic differences, because such data are not necessarily indicative of measurement invariance (Thissen, Steinberg, & Gerrard, 1986). As Hispanics are one of the more under-represented samples in psychological research (Hall, Bansal, & Lopez, 1999) and the largest ethnic minority in the United States (U.S. Bureau of the Census, 2000), it is important to establish ethnic invariance of measures like the NCS–SF for this group of individuals.

Measurement Invariance

Many approaches can be used to determine whether mean differences (or lack of mean differences) are comparable across populations. Multigroup confirmatory factor analysis is one such method (Jöreskog, 1971;Meredith, 1993). For measured variables, the usual factor model can be expressed as

$$\Sigma_x = \Lambda_x \Phi \Lambda'_x + \Theta_\delta, \qquad (1)$$

where Σ_x is a p× p matrix of variances and covariances, Λ is a p × m matrix of factor pattern coefficients, φ is a m × m matrix of interfactor covariances, and Θ_δ is a p × p matrix of residuals among the manifest indicators. As Widaman and Reise (1997) explain, this equation can be modified by incorporating subscripts to denote group membership.

Meredith (1993) and Horn, McArdle, and Mason (1983) indicated that the first step in making group comparisons involves the establishment of a baseline model that has the same pattern for the Λ_x matrix across groups. The baseline model is called the *configural invariance model*. In other words, the configural invariance model assumes that the Λ_x matrix has the same pattern of zero and nonzero elements across groups.

Meredith (1993) delineated three forms of factorial invariance: weak, strong, and strict. *Weak factorial invariance* occurs when the factor pattern coefficients in the Λ_x matrix are constrained to equality across groups. This form of measurement invariance has also been referred to as metric invariance (Hong, Malik, & Lee, 2003;Steenkamp & Baumgartner, 1998). Steenkamp and Baumgartner argued that observed differences among item scores can be meaningfully compared if metric invariance holds. In other words, item differences across groups reflect true differences across groups.

Strong factorial invariance requires an additional set of constraints on the confirmatory factor model. If the usual model is defined in terms of the observed item

$$x_i = \tau_i + \lambda_{ij}\xi + \delta_i, \qquad (2)$$

then any item score, x_i, is a linear combination of the latent variable, ξ, an intercept, τ_i, and an error term, δ_i. The λ_{ij} represents the factor pattern coefficient and can be viewed as the slope of the regression of the item score on the latent construct.

Strong factorial invariance is said to occur when both the factor pattern coefficients and the intercepts are constrained to equality across groups. This condition is also referred to as *scalar invariance* (Hong et al., 2003;Steenkamp & Baumgartner, 1998). The assessment of scalar invariance allows for the comparisons of means across groups. In other words, observed mean differences on the manifest variables imply that there are differences between groups on the latent means. Meredith (1995) refers to an additive bias when elements of the Λ matrix are constrained to equality across groups, but items have different intercepts across groups.

Therefore, the assessment of scalar invariance provides for a way of measuring "bias" across groups.

Partial metric and partial scalar invariance (Byrne, Shavelson, & Muthén, 1989;Steenkamp & Baumgartner, 1998) can be established by allowing a subset of the λ_{ij} and τ_i to vary freely across groups, while constraining the other λ_{ij}'s and τ_i's to equality. Whereas some have tested for measurement invariance with partial measurement invariance, such models may not be indicative of true invariance. Widaman and Reise (1997) indicated that different conclusions concerning group differences on latent means and variances may be generated because of the different ways in which the model would be identified.

Finally, Meredith (1993) describes *strict factorial invariance* as occurring when the error terms are constrained to equality across groups. If a researcher is interested in making mean comparisons among latent constructs, it is not necessary to establish strict factorial invariance. Therefore, latent means can be meaningfully interpreted if λ_{ij} and τ_i are fully invariant across groups. Other forms of invariance have been described elsewhere, in which factor covariances or factor variances can be constrained to equality across groups (Steenkamp & Baumgartner, 1998).

Prior research with the NCS–SF has indicated that a one-factor model best described the data. Sadowski (1993), for example, discovered only one component using principal components analyses. In a recently published article, Culhane et al. (2004) performed both exploratory and confirmatory factor analysis with separate samples of Hispanic participants. These authors demonstrated that a one-factor solution provided an adequate description for the NCS–SF.

The purpose of this article was to extend the research by Culhane et al. (2004) by testing whether the NCS–SF was measurement invariant in samples of U.S. Anglos and U.S. Hispanics. Again, Culhane et al. (2004) found no differences in Hispanic participant scores compared with previously reported Anglo participant scores. As noted, however, the absence of mean differences does not render a scale invariant across groups (Thissen et al., 1986). We hypothesized that the configural factorial invariance model would hold across groups. To compare competing models, we assessed whether the factor pattern coefficients of the NCS–SF would be invariant across ethnic groups. We then tested whether strict scalar invariance would hold. If scalar and metric invariance held, we hypothesized that the mean NCS–SF scores on the latent construct would not be different for the groups. Finally, we tested whether the error terms were invariant across groups.

Method

Participants

We recruited introductory psychology students for this project from two midsized state-run universities with fewer than 15,000 full-time undergraduate students. A predominantly self-identified Anglo sample (N = 367; 73.6% Anglo, 22.6% African American, 1.4% Hispanic, and 2.5% other ethnicities) was collected in a university in the southeastern United States. The gender of the sample was 63.8% female (133 men and 234 women), with a mean age of 19.16 (SD = 3.47). Participants were compensated with extra credit.

The other sample (N = 241) was collected in a university located on the border between the United States and Mexico and consisted of a self-identified Hispanic majority (78.4% Hispanic, 14.1% Anglo, 1.2% African American, and 6.2% other ethnicities). The mean age of this sample was 19.69 (SD = 3.53), and there were 93 men and 148 women who participated. The percentage of female students across samples did not differ, Z_{prop} = 0.60, p = *ns*.

Students in this largely Hispanic sample participated as part of a research requirement of their course. Across samples, 25 students were removed because of errors in response or missing data. In addition, data from other ethnic groups were removed, as the comparison was between Anglos and Hispanics. The final sample consisted of 289 Anglos (59.8% women, *M* age = 19.33, *SD* = 3.45) and 175 Hispanics (64.8% women, *M* age = 19.61, *SD* = 3.60). A chi-square test of independence showed that the proportion of women across the two ethnicities did not differ statistically, $\chi^2(1, N = 464) = 1.18$, $p = ns$.

Power Analysis for the NCS–SF

Whereas Culhane et al. (2004) used 185 participants to estimate a confirmatory factor analytic model for the NCS–SF, others have suggested that a power analysis should be performed to assess necessary sample size in factor analysis applications (MacCallum, Widaman, Zhang, & Hong, 1999). Therefore, we used an a priori power analysis to determine the necessary sample size needed to estimate the one-factor model in each of the two ethnic groups. For this model to be identified, the factor variance of the latent variable was constrained to one while allowing the 18 factor pattern coefficients to vary freely. This model resulted in a factor model with 135 degrees of freedom for each group.

For the power analysis, the α level was set at .05, and β was set at .20. The power analysis was based on a test of "close fit" (MacCallum, Browne, & Sugawara, 1996). In a test of "close fit," it is customary to set $R_o = .05$, where R_o is the hypothesized value of the population value for the root mean square error of approximation (RMSEA) under the null hypothesis. The alternative hypothesis specifies that the population value of the RMSEA statistic equals .08. Using the Statistica software (Steiger, 1999), we determined that a minimum necessary sample of 108 participants was needed to estimate the model for each ethnicity. As there were 289 Anglos and 175 Hispanic respondents, we determined that model misfit would be more indicative of the inadequacy of the one-factor model, as opposed to the lack of a sufficient sample size to estimate the model. In addition, it should be noted that the sample size in these analyses was very similar to the sample size used in Culhane et al. (2004).

Materials

The questionnaire included a series of items to ascertain the participants' age, gender, and ethnicity. The ethnicity item instructed participants to use the following options: African American/Black, Caucasian/White, Hispanic, Middle Eastern, Oriental/Asian, or Other. The NCS–SF (Cacioppo et al., 1984) was scored on a 5-point Likert-type scale ranging from *extremely unlike me, unlike me, neutral, like me,* to *extremely like me*. Some examples of items on the NCS–SF are "I would prefer a task that is intellectual, difficult, and important to one that is somewhat important but does not require much thought" and "I try to anticipate and avoid situations where there is a likely chance I will have to think in depth about something" (reverse coded). Higher scores on the NCS–SF represent more favorable attitudes toward cognitive effort. The possible range of scores was from 0 to 72.

The English version of the test was used for both samples. The Southwest is a unique region that allows for data collection from many Hispanics who are capable of completing the English version. Data reported in a prior study (Culhane et al., 2004) indicated that Hispanic students in the United States are generally acculturated to United States norms. Acculturation data were not collected for this study. However, data were reexamined from the earlier project that used identical criteria for participant selection. The results showed that 68.8% of Hispanic respondents reported having spent their entire life in the United States, and only 7.7% spent more of their lifetime in Hispanic countries (e.g., Mexico, Peru) than in the United States. Also, the majority of participants (67.7%) were second-generation or later Americans (Culhane, Morera, & Hosch, 2003). There was no reason to believe that these participants were different.

Procedure

Participants responded to the questionnaires in large-group settings. Standardized answer sheets were used and were subsequently read by optical scanning equipment to transform the responses into a computer data file. The experimenter distributed informed consent forms that were completed prior to the administration of the questionnaire. The experimenter then read the instructions to the participants out loud, and research participants took as much time as they needed to complete the questionnaire.

Data Analysis

We conducted a one-way analysis of variance (ANOVA) using the ethnicities as the independent variable and NCS–SF scores as the dependent variable. We assessed measurement invariance using a multigroup confirmatory factor analysis. An independent clusters factor model, using maximum likelihood estimation, was estimated using the LISREL software version 8.30 (Jöreskog & Sörbom, 1999).

Model Fit

To evaluate the hypothesized factor models, we used a variety of fit indices. Hu and Bentler (1999) had recommended a strategy of reporting the standardized root mean square residual (SRMR) and one other index of model fit. An SRMR, which indicates poor model fit, should be close to .08 if the model describes the data. Of the other fit indices that can accompany SRMR, we found several. First, we looked at the RMSEA, which is a measure of badness of model fit. Hu and Bentler recommended that RMSEA values close to .06 are needed for the model to adequately describe the data. Widaman and Reise (1997) have also indicated that "reasonable" levels of model fit are said to occur when the RMSEA statistic is between 0.05 and 0.08. In addition, we also found a goodness of fit index (GFI). According to McDonald (1999), GFI indices close to 0.90 indicate adequate goodness of fit.

Results

As anticipated, the observed mean NCS–SF scores for the Anglos and Hispanics were practically equivalent, 40.30 ($SD = 9.35$) and 41.85 ($SD = 9.03$), respectively, $F(1, 462) = 3.05$, $p = .08$, $\eta^2 = .007$. This finding suggested that no substantive differences existed between groups. Cronbach's alpha of the NCS–SF was equal to .87 for Anglos (95% CI = .84 to .89) and .85 for Hispanics (95% CI = .82 to .88). These confidence intervals indicated that the point estimates of internal consistency reliability did not statistically differ across groups. Moreover, the lower bound of each confidence interval indicated that internal consistency reliability was adequate. Tables 1 and 2 contain the means and standard deviations and the interitem correlations among the NCS–SF items across the two groups.

We tested the configural invariance model, and the final factor pattern coefficients for each group are presented in Table 3. For the most part, a visual inspection of the items suggested that the two groups were not substantially different, with the exception of Items 4 and 5. Item 4 reads, "I would rather do something that requires little thought than something that is sure to challenge my thinking abilities." Item 5 reads, "I try to anticipate and avoid situations where there is a likely chance I will have to think in depth about something." For both of these items, Anglos had higher factor pattern coefficients.

To determine whether the configural invariance model provided an adequate description to the data, we used the model fit indices described earlier. As Table 4 demonstrates, both the point estimate of the RMSEA and SRMR value indicated adequate fit, whereas the GFI almost but did not quite meet its heuristic level of adequate fit. The configural invariance model, therefore, appeared to provide an acceptably good description of the data. However, caution must be exercised, as the GFI did not meet its purported cut-off value.

To assess weak factorial invariance (metric invariance), we constrained the elements of the λ matrix to equality across groups. The value of the chi-square statistic increased by 14.80, and 18 degrees of freedom were added to the model. As the weak factorial invariance model was nested within the configural invariance model, we performed a chi-square difference test, which indicated that the difference in fit between these two models was not statistically significant. It was also worth noting that the point estimates of RMSEA and SRMR were still acceptable, whereas the GFI did not dramatically decrease.

In addition, we examined individual modification indices for the 18 factor pattern coefficients. A modification index informs the researcher of the reduction in the value of the chi-square statistic when any parameter estimate is allowed to freely vary. After allowing the factor pattern coefficients for Items 4 and 5 to vary freely across the two groups, the fit of the model statistically improved, $\chi^2(2, N = 464) = 8.13$, $p < .02$. We concluded that the factor pattern coefficients were partially measurement invariant. In other words, most of the factor pattern coefficients on the NCS–SF did not statistically differ across groups. Only Items 4 and 5 had different factor pattern coefficients.

Next, strong factorial invariance (scalar invariance) was tested by constraining the values of the τ_i vector across groups. The value of the chi-square statistic increased by 38.51 with 18 degrees of freedom added. As the strong factorial invariance model was nested within the weak factorial invariance model, we performed another chi-square difference test, and the model fit statistically worsened. In other words, at least one of the intercepts differed across groups. It should be worth noting that the modification indices for Items 1, 7, 8, and 12 were in excess of 3.84, which would indicate that freeing any one of these particular estimates would statistically improve model fit.

Although strict factorial invariance requires the factor pattern coefficients, the intercepts, and the error terms to be invariant across populations, we tested the invariance of error terms for completeness. As the strong factorial invariance model did not hold, we compared a model that constrained error terms and factor pattern coefficients to the weak partially factorial invariant model. Although the value of the chi-square statistic increased by 21.08 in constraining the error terms, this did not represent a statistical decrease in model fit. Of further note was the fact that the value of the RMSEA statistic was still acceptable, whereas the value of the GFI statistic remained the same. Freeing the error term for the fourth item would have improved model fit, $\chi^2(1, N = 464) = 9.50$, $p < .01$. All of these models are summarized in Table 4.

Discussion

In this project, we assessed the measurement equivalence of the NCS–SF across samples of Anglo and Hispanic participants. The results revealed that the NCS–SF was partially measurement invariant, as two of the factor pattern coefficients and at least one intercept differed across groups. As partial invariance may not be indicative of true invariance (Widaman & Reise, 1997), we did not compare the latent means across the two groups.

A limitation to these findings involves the value of the GFI. Although the RMSEA and SRMR pointed toward acceptably fitting models, the value of the GFI contradicted this conclusion. It is worth noting that the largest modification index always occurred for the freeing of a correlated error term, and although this study looked at modification indices to assess partial measurement invariance, reliance on modification indices to improve model fit capitalizes on chance characteristics of the data set (MacCallum, Roznowski, & Necowitz, 1992). Therefore, we refrained from using modification indices merely to improve model fit and increase the GFI to an acceptable level.

A second potential limitation of this research was that the Spanish version of the NCS–SF was not made available to the Hispanic participants. These participants were recruited from classes where the medium of instruction was English. Future researchers may nevertheless want to offer the Spanish version to participants of Hispanic ethnicity. Yet another interesting project might involve a within-subject examination of the scale presented in two languages.

At the same time, however, it is important to remember that Hispanics are the largest and fastest growing ethnic minority in the United States (U.S. Bureau of the Census, 2000). This means that researchers throughout the United States will likely work with samples that include larger numbers of Hispanics. Even near the southeastern university where the present Anglo sample was obtained, communities now have substantial populations of Hispanic adults working in local industries and Hispanic children taking courses in both Spanish and English in local school systems. How Hispanic individuals, and indeed members of other ethnic minorities, respond to English questionnaires is an important empirical question that researchers in a pluralistic society may increasingly need to answer.

Another potential limitation was the use of linear confirmatory factor analysis with items from a 5-point Likert scale. Item scores from ordered categories may not meet the strict assumptions of linearity in a linear confirmatory factor analysis model (McDonald, 1999;Panter, Swygert, Dahlstrom, & Tanaka, 1997). In addition, the use of a single item to assess ethnicity perhaps resulted in questionable construct validity (Tanaka, Ebreo, Linn, & Morera, 1998). For example, Hispanic respondents who use a "rule-out" strategy in assessing ethnicity (i.e., I'm not Anglo, I'm not African American . . . Therefore, I'm left with Hispanic) may differ in how they identify themselves ethnically from respondents who clearly acknowledge Hispanic as their ethnicity. Moreover, great within-group heterogeneity exists for individuals who identify themselves as Hispanic. Within the Hispanic label are individuals who recently immigrated from Mexico and other Latin American countries, whereas other Hispanic individuals may have lived in the United States for many generations.

Overall, the results of this investigation suggested that the NCS–SF was configurally invariant across the two ethnic samples that we examined. The results of the one-way ANOVA were the first to directly compare mean NCS–SF scores of U.S. Hispanics and Anglos. However, the factor pattern coefficients and the intercepts were not strictly invariant for the two ethnic groups, which did not allow for a comparison of the latent means. At present, we can conclude that observed NCS–SF scores do not statistically differ across U.S. Hispanic and U.S. Anglo groups. We can also conclude that the NCS–SF is partially measurement invariant across U.S. Hispanic and U.S. Anglo samples.

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TABLE 1

Means (0–4 Possible), Standard Deviations, and Interitem Correlations of the Need for Cognition Scale–Short Form (Anglos)

Item	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15	16	17	18	*M*	*SD*
1	—																		1.70	1.05
2	.43	—																	2.35	0.85
3	.39	.41	—																2.62	0.92
4	.40	.45	.50	—															2.45	0.92
5	.30	.40	.47	.50	—														2.77	0.82
6	.26	.31	.34	.37	.32	—													1.67	0.90
7	.32	.19	.31	.39	.35	.17	—												2.20	0.98
8	.20	.27	.20	.25	.25	.21	.20	—											1.85	0.99
9	.29	.28	.24	.40	.30	.20	.29	.25	—										1.81	0.90
10	.21	.38	.36	.37	.33	.22	.12	.07	.21	—									2.65	0.83
11	.36	.44	.34	.44	.36	.20	.22	.20	.30	.44	—								2.51	0.92
12	.26	.28	.32	.44	.28	.25	.20	.08	.21	.40	.50	—							2.54	0.90
13	.33	.36	.30	.39	.34	.32	.24	.13	.29	.29	.42	.40	—						1.80	0.96
14	.23	.35	.41	.39	.33	.35	.24	.18	.18	.41	.46	.43	.40	—					2.22	0.99
15	.16	.21	.22	.32	.37	.23	.27	.15	.18	.17	.22	.29	.29	.26	—				2.21	0.90
16	.29	.19	.09	.26	.23	.17	.21	.12	.07	.11	.29	.18	.14	.12	.09	—			2.03	1.08
17	.34	.16	.21	.40	.35	.19	.30	.27	.20	.09	.25	.27	.16	.17	.20	.33	—		2.45	1.04
18	.10	.24	.17	.23	.20	.21	.27	.06	.12	.14	.15	.18	.18	.23	.15	.10	.09	—	2.46	0.98

TABLE 2

Means (0–4 Possible), Standard Deviations, and Interitem Correlations of the Need for Cognition Scale–Short Form (Hispanics)

Item	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15	16	17	18	*M*	*SD*
1	—																		1.97	0.92
2	.48	—																	2.38	0.93
3	.20	.39	—																2.66	0.94
4	.23	.27	.40	—															2.62	0.91
5	.25	.29	.33	.46	—														2.67	0.85
6	.23	.35	.27	.32	.27	—													1.73	1.01
7	.22	.28	.23	.22	.28	.13	—												2.10	0.99
8	.19	.18	.19	.25	.11	.29	.11	—											2.07	1.06
9	.22	.30	.31	.41	.29	.27	.32	.27	—										1.97	0.98
10	.24	.37	.24	.25	.19	.23	.31	.10	.25	—									2.70	0.86
11	.31	.50	.28	.33	.22	.37	.14	.28	.19	.33	—								2.63	0.87
12	.25	.35	.30	.26	.21	.34	.17	.26	.23	.27	.40	—							2.82	0.88
13	.38	.36	.21	.24	.22	.19	.24	.09	.33	.21	.32	.27	—						1.97	0.90
14	.43	.41	.19	.22	.23	.18	.23	.21	.18	.37	.39	.26	.31	—					2.17	0.86
15	.38	.36	.23	.20	.14	.20	.15	.16	.21	.27	.31	.15	.12	.34	—				2.23	0.94
16	.30	.21	.12	.23	.27	.13	.14	.09	.21	.25	.30	.25	.22	.25	.15	—			2.05	1.10
17	.23	.27	.29	.33	.17	.17	.31	.24	.32	.41	.26	.45	.31	.23	.19	.20	—		2.63	0.98
18	.06	.13	.05	.14	.13	.19	.15	.11	.04	.33	.05	.07	.13	.18	.18	.00	.12	—	2.47	0.95

TABLE 3

Configural Invariance: Final Factor Pattern Coefficients for the Need for Cognition Scale–Short Form

	Anglo		Hispanic	
Item	**Coefficient**	***SE***	**Coefficient**	***SE***
1	.55	.06	.56	.07
2	.62	.06	.67	.07
3	.62	.06	.51	.08
4	.75	.05	.55	.07
5	.64	.06	.47	.08
6	.48	.06	.48	.08
7	.46	.06	.42	.08
8	.34	.06	.36	.08
9	.46	.06	.53	.08
10	.52	.06	.54	.07
11	.64	.06	.61	.07
12	.58	.06	.50	.07
13	.57	.06	.54	.08
14	.59	.06	.55	.07
15	.42	.06	.45	.08
16	.33	.06	.40	.08
17	.43	.06	.53	.08
18	.31	.06	.23	.08

TABLE 4

Fit Indices for Modified Structural Models

Model	χ^2	**df**	**SRMR**	**RMSEA**	**90% CI**	**GFI**
1. Baseline	521.03	270	.063	.065	.057–.073	.88
2. LX Inv	535.83	288	.074	.063	.055–.071	.87
2a. Partial LX	527.70	286	.068	.062	.054–.070	.87
3. Partial LX, TX Inv	566.21	304	.071	.063	.055–.070	.87
4. Partial LX, TD Inv	545.13	304	.070	.061	.053–.069	.87
4a. Partial LX, Partial TD	535.63	303	.068	.059	.052–.067	.87
Model comparison	$\Delta\chi^2$	Δdf	ΔSRMR	ΔRMSEA		ΔGFI
1. Model 2–Model 1	14.80	18	.011	−.002		−.01
2. Model 2a–Model 2	8.13*	2	−.006	−.001		.00
3. Model 3–Model 2a	38.51*	18	.003	.001		.00
4. Model 4–Model 2a	21.08	18	−.001	−.002		.00
5. Model 4a–Model 4	9.50	1	−.002	−.002		.00

Note. LX = λ_{ij}s, TX = τ_is, and TD = Θ_δs.

* $p < .05$.