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The dimensionality and measurement properties of alcohol outcome expectancies across Hispanic national groups

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ABSTRACT

This study examines the psychometric properties of alcohol expectancies among Hispanic subgroups. Face-to-face interviews were conducted as part of the 2006 Hispanic Americans Baseline Alcohol Survey (HABLAS), which employed a multistage cluster sample design. A total of 5224 individuals (18 + years of age) representing four Hispanic national groups (Puerto Ricans, Cuban Americans, Mexican Americans, and South/Central Americans) were selected at random from the household population in five metropolitan areas (Miami, New York, Philadelphia, Houston, and Los Angeles). Alcohol expectancies included 18 items covering positive (e.g., laugh more, become more talkative) and negative dimensions (e.g., become aggressive, lose control) when alcohol is consumed. Confirmatory factor models replicated a previously proposed three-factor dimensional structure with a substantial majority of items exhibiting measurement invariance across Hispanic national group and gender. Items covering social extroversion were an exception, showing a lack of invariance for female Cuban and South/Central Americans. Latent mean differences across groups were detected for expectancies concerning emotional fluidity, and the pattern of differences largely mirrored known differences in alcohol consumption patterns. Results suggest that caution should be exercised in interpreting differences in expectancies concerning social extroversion across Hispanic groups, and additional work is needed to identify indices of this construct with invariant measurement properties. However, measures of emotional/behavioral impairment and emotional fluidity expectancies can be validly compared across gender and Hispanic national groups.

1. Introduction

Alcohol expectancies (AEs) reflect beliefs that individuals develop about the effects of alcohol on various aspects of behavior and cognition (Goldman, Brown, Christiansen, & Smith, 1991). These beliefs precede direct experience with alcohol and are influenced by social contexts and role models such as family, peers, and mass media (Christiansen, Goldman, & Inn, 1982). Once established, they influence various aspects of alcohol use (Goldman et al., 1991; Leigh & Stacy, 1991), including current drinking behaviors (for a review, see Baer, 2002; Critchlow, 1987; Leigh, 1987, 1989; Leigh & Stacy, 1993), future alcohol and drug use (Christiansen, Smith, Roehling, & Goldman, 1989; Stacy, Newcomb, & Bentler, 1991; Stacy, Widaman, & Marlatt, 1990), and alcohol dependence symptoms (Wood, Sher, & Strathman, 1996). Differences in AEs have been observed between subtypes of drinkers (e.g., lone vs. group problem drinkers, restrained and unrestrained drinkers), genders, and ethnicities (Bensley, 1991; Gustafson, 1993; Jones & McMahon, 1992; McMahon, Jones, & O'Donnell, 1994). The bulk of AE research has relied on predominantly White populations, although several studies have linked expectancies to drinking behaviors and alcohol-related problems among Hispanics (primarily of Mexican background; e.g., Corbett, Mora, & Ames, 1991; Gilbert, Mora, & Ferguson, 1994; Marín, 1996). Similarly, studies of the measurement properties of expectancy instruments have been examined primarily in non-Hispanic populations. For example, Leigh and Stacy (1993) studied the factor structure and psychometric properties of AEs and their relationship to self-reported alcohol use, finding evidence for a general two-factor structure with positive and negative dimensions. One exception can be found in Marín, Posner, and Kinyon (1993), who found evidence for a three-dimensional solution using exploratory factor analysis (with emotional and behavioral impairment, emotional fluidity, and social extroversion factors) in a Hispanic sample of mixed national origin.

Because there is pronounced heterogeneity in drinking levels and problems across Hispanics of different national origin (Ramisetty-Mikler, Caetano, & Rodriguez, in press; Vaeth, Caetano, Ramisetty-Mikler, & Rodriguez, 2009), important differences in any alcohol-related construct may be obscured by treating Hispanics as a homogeneous group. With respect to expectancy research, potential measurement issues complicate such comparisons further. To the extent that the measurement properties (e.g., loadings, intercepts) of expectancy items differ across these groups, the items (or composite scales) are not comparable (Byrne, Shavelson, &

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Muthen, 1989; Meredith, 1993; Muthén & Muthén, 2007). Consequently, there is a need for more formal evaluations of expectancy measures for use in Hispanic populations. In the present study, confirmatory factor analysis was used to compare two proposed dimensional structures of alcohol expectancies and to test for various forms of measurement non-invariance across four major Hispanic national groups.

2. Methods

Using a multistage cluster design, the 2006 HABLAS sampled Puerto Rican, Cuban, Mexican, and South/Central Americans (aged 18 or older) from five selected metropolitan areas of the U.S. The present analyses are restricted to respondents who were current drinkers ($N=2773$). Additional details concerning the survey methodology can be found in Caetano, Ramisetty-Mikler, and Rodriguez (2008).

2.1. Alcohol expectancy measure

In Marín et al. (1993), preliminary analyses reduced a large set of expectancy items to a pool of 18, and further factor analyses of this item set identified 11 items that loaded on three factors interpreted as emotional/behavioral impairment (EBI), emotional fluidity (EF), and social extroversion (SE). To compare Marin et al.'s three-factor solution with alternative structures proposed in the literature, the larger 18 item pool was used in the present study. Respondents rated how often alcohol would make them feel eight positive effects of alcohol consumption (laugh more^{SE}, become more talkative^{SE}, happier, relaxed^{EF}, romantic^{EF}, friendly^{EF}, sexually aroused^{EF}, and independent) and ten negative effects (become louder, aggressive, lose control^{EBI}, become careless^{EBI}, argumentative^{EBI}, lose coordination^{EBI}, become more emotional^{EBI}, sleepy, sad, and have difficulties thinking), where superscripts denote the 11 item configuration of Marín et al. (1993) three-factor solution. A Likert-type response format was used for each item, coded from 0 to 3 (almost never = 0, sometimes = 1, often = 2, almost always = 3).

2.2. Statistical analyses

Measurement and structural models were fit with M*plus* 5.2 (Muthén & Muthén, 2007) using a robust maximum likelihood estimator. The complex sampling design used in the HABLAS was accounted for. Model fit was assessed with the comparative fit index (CFI; Bentler, 1990), root mean square error of approximation (RMSEA; Steiger, 1990), the standardized root mean square residual (SRMR), and the Tucker–Lewis Index (TLI; Bentler & Bonett, 1980; Tucker & Lewis, 1973). We have not relied on model χ^2 values (which were typically significant) because of the statistic's sensitivity to trivial misfits in large sample sizes (Cudeck & Browne, 1983). Following the specification of an acceptable configural invariance model, metric (loading), scalar (intercept), and strict measurement invariance (Steenkamp & Baumgartner, 1998; Vanderberg & Lance, 2000) in the factor structure of AEs was assessed across eight groups defined by the four Hispanic national identities and gender. Nested models were compared with scaled chi-square difference tests (Satorra & Bentler, 2001), which were computed using model log likelihoods and scaling correction factors following procedures described on the M*plus* website (http://www.statmodel.com/chidiff.shtml).

3. Results

Demographic characteristics of the HABLAS sample can be found in Caetano et al. (2008). The average inter-item correlation was .38 with a range of $r=.12$ to $r=.79$. Preliminary correlated two- and three-factor CFA models fit to the overall sample and to individual subgroups consistently showed extreme modification indices (>10; the standard 3.84 value was not used to avoid capitalization on chance) for three residual covariances: "more romantic" with "sexually aroused," "become argumentative" with "more emotional," and "lose self control" with "become careless." Because it is sensible that these pairs of items would each share common influences beyond the presently modeled expectancy factors, these residual covariances were modeled and all others were constrained to zero for subsequent models.

For the reduced 11 item set of Marín et al. (1993), a two-factor solution (involving positive and negative dimensions; Leigh & Stacy, 1993) is nested within the Marin et al. three-factor solution (the EF and SE factors consist of exclusively positive items, whereas EBI items are exclusively negative). Consequently, we first compared the three-factor solution with a two-factor solution by constraining the correlation between the latent EF and SE factors to 1. Imposing this constraint significantly degraded model fit (χ^2_{diff} $(1)=122.5$, $p<.0001$; two-factor: CFI = .92, TLI = .90, RMSEA = .06, SRMR = .06; three-factor: CFI = .95, TLI = .93, RMSEA = .05, SRMR = .06). This finding was corroborated by examination of fit indices for group-specific models. Across all subgroups, indices consistently suggested better fit for the 11-item three-factor model. For example, CFI values ranged from .92 to .99 for the three-factor model (Table 1) and from .89 to .97 for the two-factor model. Other group-specific fit indices showed the same pattern, reflecting a small but consistently better fit for the three-factor model. To ensure the result was not an artifact of using a reduced item set, we examined group-specific fit indices for the two-factor model using the full 18 item set (the 11-item three-factor solution is not defined for all 18 items, but the two-factor solution, which was derived from a larger item set, is). CFI values ranged from .77 to .92 for this model and other indices behaved similarly, again suggesting support for the 11-item three-factor solution.

Invariance testing began by specifying a three-factor configural invariance model with unrestricted parameters across all subgroups (model 1). Model fit statistics and hierarchical testing results from all invariance analyses are presented in Table 2. Fit indices for non-rejected models in Table 2 are within ranges generally considered "acceptable" (CFI/TLI ≥ .90, RMSEA/SRMR < .10; Hu & Bentler, 1999; Kline, 2005). Subsequent discussion will therefore be directed toward discussion of scaled difference tests. Although constraining loadings to equality (Table 2; model 2 vs. model 1) did not significantly degrade model fit, constraining intercepts across groups (model 3 vs. model 2) did. Consequently, modification indices from model 3 were consulted for extreme values to identify the source of the invariance. Of 88 (8 groups × 11 items) possible group-specific intercepts, these suggested that three should be freely estimated: "more relaxed" for Puerto Rican males, and "laugh more" and "talk more" for Cuban and South/Central females. Puerto Rican males were less likely to endorse "more relaxed" than other groups, and Cuban American females and (to a slightly lesser extent) South/Central females were each less likely to endorse "laugh more" and "talk more" than other groups. The resulting model (model 4) did not differ significantly from model 2, and this model was retained. Note that because the SE factor is made up of only these two indicators, measurement invariance on this factor cannot be assumed for Cuban American and South/Central females. Consequently, latent SE means for these two groups were left constrained at zero for subsequent testing.

Stricter forms of invariance were examined next. Constraining residual variances to equality (model 5 vs. model 4) did not degrade model fit, but constraining the three estimated residual covariances to equality (model 6 vs. model 5) did. Although there was some variation across ethnic groups in extent, the three residual covariances ("lose self-control" with "become careless," "become argumentative" with "become emotional," and "more romantic" with "sexually aroused") were stronger (more positive) tended to be stronger for females than for males. Finally, constraints on factor variances and covariances (model 7 vs. model 5) did not significantly degrade model fit. Parameter estimates for this final model appear on the right of Table 1. All estimates were significant ($p<.001$).