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The Factor Structure of Gambling-Related Cognitions in an Undergraduate University Sample

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Abstract Gambling is relatively common among university students, but few studies examine factors that contribute to gambling behavior in this cohort. Based on evidence that cognitive distortions may play a role in gambling behavior, this study examined the factor structure of gambling-related cognitive distortions using the Gambler's Beliefs Questionnaire (GBQ; Steenbergh et al., *Psychology of Addictive Behaviors*, 16:143–149, 2002) in a sample of 393 college undergraduates. Confirmatory factor analysis was used to test a previously reported two-factor model, comprising dimensions of Illusion of Control (IOC) and Luck/Perseverance (L/P). An oblique, but not orthogonal, two-factor model was confirmed but did not provide an incrementally better fit to the data than a one-factor model. However, multiple regression analyses showed that the L/P scale accounted for significant variance in the criterion when controlling for IOC items. This suggests that IOC items provide redundant information and that gambling-related cognitive distortions in this sample can be adequately assessed using solely the L/P factor.

Keywords Gambling · Cognitive distortions · University students · Factor analysis

These data were collected while the first four authors were at the State University of New York at Binghamton and the rest of the study was conducted at their subsequent institutions.

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Relative to research on gambling behavior in the general public, few studies have been conducted examining the factors associated with gambling behavior in university undergraduate students. However, undergraduate students frequently engage in a range of gambling behaviors. In a representative sample of the US university students, LaBrie et al. (2003) found that within a 1-year span, 42% of college students reported at least some gambling behavior, with 2.6% falling in the problematic range. It is therefore not surprising that some authors have urged that more research should examine the causal and maintaining variables underlying gambling behavior in this population (e.g., Stinchfield et al. 2006).

Recent research suggests that distorted cognitions may be an important variable in the maintenance of gambling behavior in general (for a review, see Toneatto 1999), and may specifically contribute to problems with gambling in college students (MacKillop et al. 2006a, b). Ladouceur and Walker 1996 noted two distinct types of gambling-related cognitive distortions that may lead individuals to miscalculate the economic utility of their gambling. One type, referred to as the illusion of control (IOC), comprises a collection of beliefs that one's behavior can influence the outcome of events determined by chance (Langer 1975). The second type, referred to as Luck/Perseverance (L/P), consists of interpretive biases that lead gamblers to overestimate their chances of winning. To directly assess IOC- and L/P-related distortions, Steenbergh et al. (2002) developed the Gambler's Beliefs Questionnaire (GBQ), which was designed to assess these constructs separately. In that study, a principal axis factor analysis yielded two primary factors, with the item factor loadings being consistent with the IOC and L/P constructs.

These data provided initial support for the two-factor model of gambling-related cognitive distortions. However, Steenbergh et al. (2002) used a mixed sample comprising two populations, undergraduates and community members, making it unclear whether these results applied to both samples separately. Indeed, there is evidence to suggest that, in contrast to adults, the determinants of gambling behavior may be different for university age individuals (e.g., Bergevin et al. 2006). It is therefore unclear whether gambling-related cognitive distortions in university students adhere to the two categories proposed by Steenbergh et al. (2002) and others. In addition, the IOC and L/P factors in Steenbergh et al. (2002) were obtained using a factor analytic method that examines the underlying factor model *a posteriori*. Although this is an important first step in measure development, the two-factor model for gambling beliefs has not been tested *a priori*. Taken together, the present investigation was designed to provide an *a priori* test of the GBQ's two-factor model of gambling-related cognitive distortions using Confirmatory Factor Analysis (CFA) with a sample composed solely of undergraduate college students. As the IOC and L/P cognitive distortions are not limited to those who engage in gambling behaviors only, but rather represent biases present in the general population, the foregoing hypotheses were tested using both individuals with and without a history of gambling behavior. This methodology also is consistent with the sampling methods used by Steenbergh et al. in the initial validation study.

Based on the findings of Steenbergh et al. (2002), it was hypothesized that a two-factor model, comprising orthogonal IOC and L/P dimensions, would provide an adequate fit to GBQ scores endorsed by undergraduate college students. However, as it is possible that the IOC and L/P are separate but overlapping dimensions of gambling-related cognition, an oblique two-factor model in which each factor was permitted to intercorrelate was also tested. As a control model, an undifferentiated one-factor model was also tested and compared to the two-factor models to determine incremental fit. As recommended by Bryant (personal communiqué), regression analyses were conducted using gambling behavior as a validation criterion to further test the adequacy of the two-factor structure. If the GBQ actually comprises two underlying factors, then the IOC and L/P scales should account for greater total variance in a criterion relative to an index that combines these scores. Furthermore, the IOC and L/P scales should account for unique variance in the criterion if they actually represent separate constructs. It was therefore hypothesized that the separate IOC and L/P dimensions would emerge as a better predictor of gambling behavior than would a single undifferentiated index that combines these dimensions, and that they would predict unique variance in the criterion.

Method

Participants and Procedures

The participants were recruited from a large research university in the Northeastern United States over the course of two semesters. Upon receiving approval from the University's Institutional Review Board, participants were recruited from undergraduate psychology classes via flyers and in-class announcements, and were compensated for their time with required research participation credits. Participants completed the GBQ and the South Oaks Gambling Screen (SOGS; Lesieur and Blume 1987) in a group administration, as well as an informed consent form and questionnaires unrelated to the current study. Presentation of the questionnaires was counterbalanced across participants. Three hundred and ninety-three participants provided complete data with 211 men and 182 women comprising the sample. The modal age was 18 for both men and women. The sample was predominantly Caucasian (n=265), with the remaining participants reporting the following racial and ethnic identifications: Asian (n=77), Hispanic (n=25), African American (n=12), Hawaiian or Pacific Islander (n=1), and "other" (n=13).

Measures

Gambler's Beliefs Questionnaire (GBQ; Steenbergh et al. 2002) The GBQ is a 21-item measure of cognitive distortions about gambling that yields the IOC and L/P subscales. The IOC subscale contains eight items and measures an individual's perception that chance occurrences are influenced by their behavior. The L/P subscale contains 13 items and purports to measure an individual's belief that chance occurrences will turn out in their favor. As indexed by Cronbach's α, the internal consistencies of the IOC subscale, L/P subscale, and combined GBQ score were 0.89, 0.94, and 0.93, respectively. Scores on the L/P dimension were significantly positively skewed and were normalized using a square root transformation.

South Oaks Gambling Screen (SOGS; Lesieur and Blume 1987) The SOGS, a self-report measure composed of 20 items that assesses pathological gambling based on DSM-III (American Psychiatric Association 1980) criteria, was used to assess the base rate of gambling behavior in the current sample. According to Lesieur and Blume (1987), the SOGS has demonstrated good reliability and validity in clinical samples, and discriminates pathological from non-pathological gamblers using a cut-off of score of five (cf., Ferris et al. 1999). SOGS scores were significantly positively skewed. A logarithmic transformation was then applied, which normalized the variable for use in the regression analyses.

Results

Preliminary Analyses

The mean SOGS score of all participants was 0.56 (SD=1.38), with a range of 0 to 13. The modal score was 0, with 76% of the sample not endorsing any problems with gambling (N=297), 21% receiving scores between 1 and 4 (N=83), and 3% obtaining scores of 5 or more (N=13), the commonly used criterion for probable pathological gambling. The most frequently endorsed gambling games of the participants were: card games for money (62.9%), raffles or fundraising tickets (46.5%), and scratch-and-win tickets (44.8%). The mean IOC and L/P scores were 16.98 (SD=8.66) and 20.78 (SD=9.82), respectively. The IOC and the L/P factors were highly significantly correlated (r=0.82, p<0.01). The IOC and L/P dimensions were significantly correlated with SOGS scores (r=0.22 and 0.29, respectively, ps<0.001). The combined GBQ index also significantly correlated with SOGS scores (r=0.26, p<0.001).

Factor Analysis

Initial data screening demonstrated that the majority of individual GBQ items were positively skewed. As such, the confirmatory factor analyses were conducted using techniques more appropriate for nonnormality. In particular, the current study used the Asymptotically Distribution Free (ADF) estimation procedure. The chi-square adjustment test statistic for small to medium samples (i.e., <2,500; Yuan and Bentler 1997), which yields values that are then compared to an F distribution, was used in the current analysis. In addition, goodness of fit was assessed using the incremental fit index (IFI; Bollen 1989), and the root mean square error of approximation (RMSEA; Brown and Cudeck 1993). The IFI was chosen because it performs well in small to medium samples compared to other existing fit indices (Bollen 1989). The RMSEA was selected because of its prevalent use in the factor analytic literature. Both the IFI and RMSEA were recalculated based on an adjusted ADF χ^2 estimate for small to medium samples (see Satorra and Bentler 1988). IFI estimates above 0.90 indicate good fitting models. RMSEA estimates of less than 0.05 indicate a close approximate fit, values ranging from 0.05 to 0.08 indicate a reasonable fit, and values of 0.10 or more are indicative of poor fitting models (Brown and Cudeck 1993). The CFA was conducted using AMOS 6.0.

Based on the adjusted ADF estimates, the orthogonal two-factor model was not a good fit to the data according to any of the fit indices, F_{ADF} (188, 205)=3.31, p<0.01, IFI=0.29, and RMSEA=0.12. The oblique two-factor model was identified as a good fit to the data; F_{ADF} (188, 205)=1.14, ns, IFI=0.91, and RMSEA=0.054. The one factor model was also identified as providing a good fit to the data based on the adjusted ADF estimator, F_{ADF} (189, 204)=1.16, ns, IFI=0.90, and RMSEA=0.059. Because these models were not nested, relative goodness of fit was assessed using Aikake's Information Criterion (AIC; Aikake 1987; Tabachnik and Fidell 2001). Similar to the estimates for the other fit indices, the AIC index was recalculated for both the one- and two-factor models using adjusted chi-square estimates for non-normal data. Compared to the orthogonal two-factor model, the one-factor model provided a significantly better fit to the data, $AIC_{difference}$ (1, n=393)=93.14, p<001. The oblique two-factor model did not provide an incrementally better fit to the data than the more parsimonious one-factor model, $AIC_{difference}$ (1, n=393)=1.00, ns. The factor loadings for the one-factor model are provided in Fig. 1.[1]

Regression Analysis

Two cases were identified as multivariate outliers using the Mahalonobis distance from the centroid (χ^2=13.82; p<0.001). These cases were not omitted from the sample as their removal had no impact on the obtained pattern of results. The SOGS was regressed on both the GBQ total score and the L/P and IOC dimensions in separate regressions. The squared multiple correlation (R^2) for the criterion measure was then examined to determine the proportion of variance that each regression analysis explained. The R^2 for the criterion measure should be higher when using the two separate GBQ subscales as predictors compared to using the GBQ total score if two factors are really more appropriate than one. The comparison of R^2 across models was conducted by first building 95% confidence intervals around (a) the estimate of R^2 from the regression using GBQ total score as a predictor and (b) the estimate of R^2 from the regression using the two GBQ subscales as separate predictors. The R^2s are considered to be significantly different if their confidence intervals are

[1] According to Bentler and Mooijaart (1989), the choice between two competing models that adequately fit the data should favor the less complex model on philosophical grounds (e.g., fewer theoretical assumptions; also see James et al. 1982) and given statistical considerations (e.g., more precise parameter estimates in the simpler model).

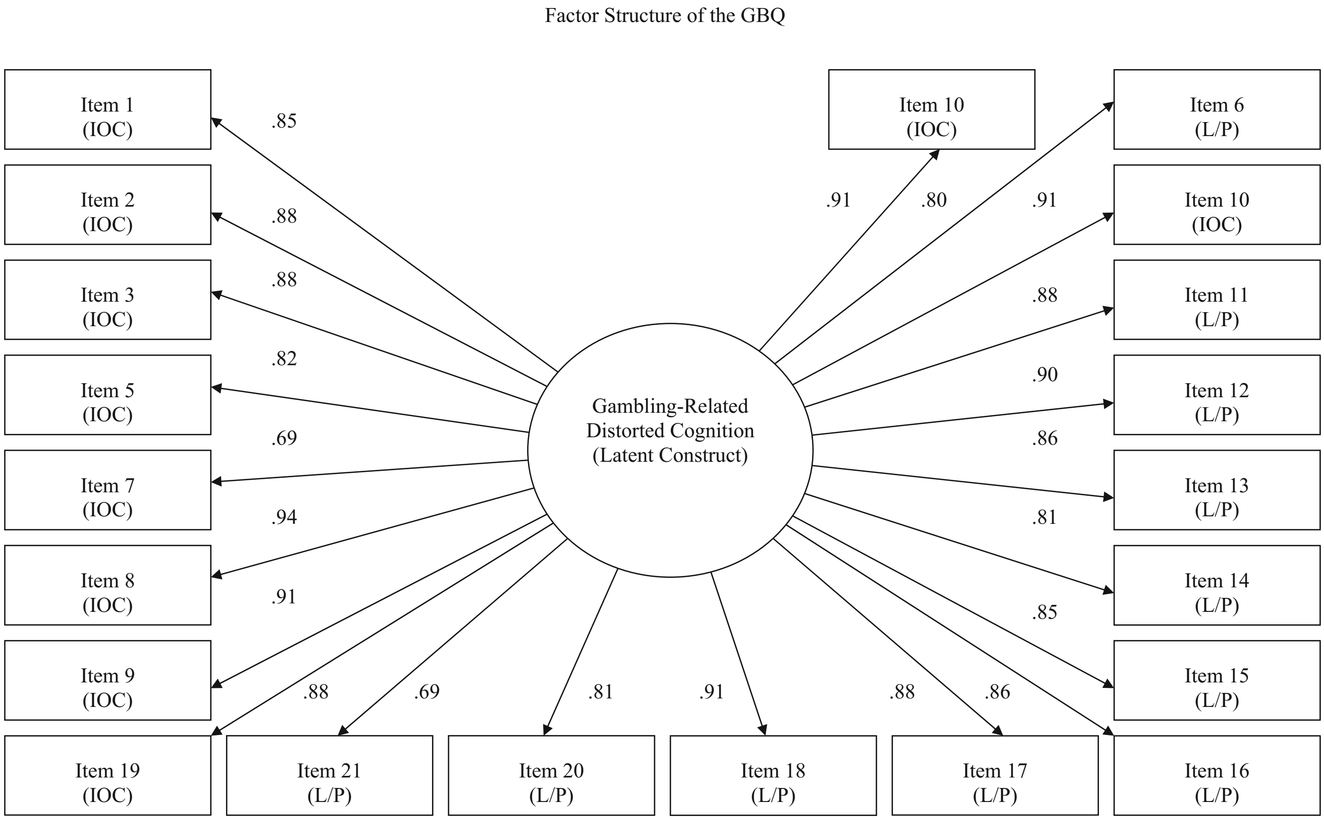

Fig. 1 One-factor model of the GBQ items in an undergraduate sample with standardized regression weights

non-overlapping (for an additional example of this approach, see Bryant et al. 1996).

It was hypothesized that the SOGS variance accounted for would be greater in the regression equation that modeled the GBQ dimensions separately as opposed to combining them into one undifferentiated measure of gambling-related cognitive distortions. Both regressions analyses were significant, with R^2 values of 0.07 ($p<0.001$; CI=$0.02<R^2<0.13$) for the GBQ total score and 0.08 ($p<0.001$; CI=$0.01<R^2<0.12$) for the model comprising the separate IOC and L/P dimensions. The confidence intervals for the two R^2s almost completely overlap, indicating that the difference in variance explained across the two equations is not significant.

The unique variance in SOGS accounted for by the IOC and L/P dimensions was evaluated using the significance test for the t value associated with each parameter and the magnitude of the effect was based on their respective squared semipartial correlations (si_j^2). The hypothesis that the separate GBQ dimensions would account for unique variance in SOGS scores was partially supported. In particular, the L/P dimension significantly accounted for unique variance in SOGS, t=3.76; $p<0.001$ (sr_i^2=0.18, when controlling for the IOC dimension). The IOC dimension, however, did not significantly predict SOGS scores, t=−0.59; *ns* (sr_i^2=−0.03) when controlling for the L/P dimension.

Discussion

The purpose of the present study was to further examine the factor structure of gambling-related cognitive distortions in an undergraduate university sample using the GBQ. Contrary to what was predicted, the data did not support a two-factor structure to the GBQ comprising the IOC and L/P dimensions proposed by Steenbergh et al. (2002). Although an oblique two-factor model adequately fit the data, conceptualizing the IOC and L/P factors as two overlapping constructs did not provide an incrementally better fit than a one-factor control model. This indicates that the two-factor model carries additional parameters unnecessary for model fit (Anderson and Gerbing 1988).

The separate IOC and L/P factors also did not predict SOGS scores beyond what was accounted for by a combined GBQ total score. There is therefore little incremental criterion-related validity associated with assessing the IOC and L/P dimensions separately. However, when the L/P and IOC dimensions were modeled together, the former accounted for unique variance in SOGS scores whereas the latter accounted for almost no unique variance in the criterion. Considered alongside the CFA analyses, this pattern of results suggests that (a) information about gambling-related cognitive distortions assessed by IOC items overlapped almost entirely with L/P items, and (b) the lack of incremental fit for the two-factor model likely resulted from the redundancy of information provided by the IOC scale.

The finding that the L/P dimension accounts for additional variance in the criterion beyond what was accounted for by the IOC dimension bears further consideration. It is possible that IOC items overlapped with a subset of L/P items, but that a different subset of non-overlapping L/P items accounted for additional variance in SOGS scores. This might indicate that the L/P dimension is not by itself a unitary factor. This is, however, an unlikely possibility given that the L/P scale was highly internally consistent and obtained an alpha coefficient even greater than the combined GBQ score. It is more likely that the increased explained variance in SOGS resulted from the greater number of items on the L/P scale ($n=5$). Alternatively, it is possible that L/P items accounted for unique variance in SOGS scores because they more precisely measure gambling-related cognitive distortions relative to the IOC items in this sample. In either case, these findings suggest that gambling-related cognitive distortions are usefully construed as a unitary construct that can be adequately assessed solely using the L/P dimension.

The discrepancy between the current findings and those of Steenbergh et al. (2002) might have emerged as a result of differences between the samples. As the current sample solely comprised college-aged individuals, these findings might indicate that gambling-related cognitive distortions are less differentiated in this population. In addition, although the base rate of gambling behavior in this sample was similar to what might be expected for the population (LaBrie et al. 2003), participants in this study reported less frequent gambling behavior than did Steenbergh et al.'s participants. It is possible that distinct notions about gambling did not emerge because the university students comprising this sample are young adults potentially with less gambling experience. Also, it should be noted that the current sample was a general university population sample and these findings might not generalize to clinical samples of pathological gamblers from university populations. In such individuals, it is possible that distinct differences between domains of cognitive distortions may conform to the IOC and L/P dimensions proposed by Steenbergh et al. While the sample differences limit the comparability of the present study to make comparisons across different levels of gambling, future research should explore the factor structure of gambling cognitions across differing levels of gambling frequency and severity in the population.

Importantly, the fact that the two-factor structure was not confirmed in this general college sample does not necessarily vitiate the utility of the GBQ. Rather, it contraindicates heavy emphasis of the two different factors when used in a general university sample. In that context, the current findings suggest that either the L/P scale or the combined GBQ measure may well be useful in characterizing the presence and magnitude of cognitive distortions and the association of these beliefs with gambling behavior and related criteria in college students. Equally, if the GBQ is applied in clinical contexts using cognitive approaches with college-age individuals, the measure may well be useful, but the utility of assessing the two separate dimensions would be questionable.

Finally, two additional considerations of the current study bear consideration. First, it should be noted that the proportion of variance in gambling involvement accounted for by the GBQ, although highly statistically significant, was relatively modest in absolute magnitude, accounting for about 8% of the variance. This clearly suggests that cognitive attributions about gambling, as assessed by the GBQ, represent but one variable contributing to gambling in this sample and further research characterizing other variables that influence college gambling is warranted. Second, it should also be noted that caution should be exercised in the generalization of these data to ethnic or racial minority populations because the sample was predominantly Caucasian. More specific study of the GBQ in minority samples is therefore warranted.

In summary, these data supported a one-factor structure of the GBQ that comprises the degree of gambling-related cognitive distortions in university students in the United States. Furthermore, these data indicate that the IOC and L/P subscales are more usefully construed as overlapping parts of a unitary construct in which the greater number or precision of items on the L/P confer a distinct advantage in accounting for variability in gambling behavior. Despite diverging from the previously reported two-factor structure, the data suggest that GBQ may still be useful in research and clinical contexts as a single factor measure of cognitive distortions in an undergraduate population.

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