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Alcohol Outcome Expectancies and Alcohol Use: A Latent Variable Cross-Lagged Panel Study

Kenneth J. Sher, Mark D. Wood, Phillip K. Wood, and Gail Raskin
University of Missouri—Columbia

The relation between alcohol outcome expectancies (EXP) and alcohol use was prospectively examined over 3 years in a mixed-gender sample of college students ($N = 465$) at low and high risk for the development of alcoholism. Alcohol use remained fairly stable over 4 years, but EXP decreased significantly over the course of the study. Structural equation modeling techniques were used to examine reciprocal relations between EXP and alcohol use over 1- and 3-year intervals. Reciprocal prospective effects were demonstrated, but the nature of these effects appears dependent on the interval between measurement periods. Conceptually, these findings indicate both an etiologic role for EXP in predicting future alcohol use, and the influence of alcohol consumption on the development and maintenance of EXP. Methodologically, they point to the importance of the consideration of measurement interval in longitudinal research.

In recent years, alcohol outcome expectancies, operationalized as beliefs that people have about the behavioral, cognitive, and emotional effects of drinking alcohol, have figured prominently in the search for psychosocial correlates and determinants of alcohol use and misuse. It appears that a variety of populations hold relatively specific beliefs about the effects of alcohol and these have been described in samples ranging from abstainers to alcoholics (Connors, O'Farrell, Cutter, & Thompson, 1986; Leigh, 1987). Outcome expectancies have been identified in children as young as 8 years (Miller, Smith, & Goldman, 1990) and have been found to be concurrent predictors of alcohol use in samples of adolescents, college students, and adults (Brown, Goldman, & Christiansen, 1985; Fromme, Stroot, & Kaplan, 1993; Leigh & Stacy, 1993; Mann, Chassin, & Sher, 1987; Wood, Nagoshi, & Dennis, 1992). Cross-sectional studies, such as these, although informative regarding the association between beliefs about the effects of alcohol and drinking behavior, are incapable of resolving the nature (i.e., direction) of the observed relations.

It is important to note that different theoretical perspectives have distinct implications with respect to the direction of effect in expectancy–alcohol use relations. As outlined by Stacy, Newcomb, and Bentler (1991), these relations can be described with reference to three general classes of theories. According to expectancy theory (e.g., Bolles, 1972), outcome expectancies are thought to be causally related to drinking behavior and thus should directly predict future alcohol use independent of previous use.[1] From a behavioral choice/self-perception framework (e.g., Bem, 1978; Vuchinich & Tucker, 1988), the opposite pattern is hypothesized. That is, alcohol use and its associated consequences should predict outcome expectancies. According to social learning theory (Bandura, 1977), both previous alcohol use and outcome expectancies should predict future drinking behavior, with the net result being a reciprocal influence process by which outcome expectancies influence drinking behavior, and drinking behavior results in altered expectations for alcohol's effects.

A hypothesized link between outcome expectancies and the initiation of alcohol use has received empirical support in several prospective studies. Consistent with expectancy theory, alcohol outcome expectancies predicted subsequent alcohol use in an adolescent sample (Christiansen, Smith, Roehling, & Goldman, 1989). Other prospective studies of early adolescents (Bauman, Fisher, Bryan, & Chenowith, 1985; Smith, Goldman, Greenbaum, & Christiansen, 1995) have observed reciprocal

Kenneth J. Sher, Mark D. Wood, Phillip K. Wood, and Gail Raskin, Department of Psychology, University of Missouri—Columbia.

Mark D. Wood is now at the Center for Alcohol and Addiction Studies, Brown University.

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Correspondence concerning this article should be addressed to Kenneth J. Sher, Psychology Department, 210 McAlester Hall, University of Missouri—Columbia, Columbia, Missouri 65211. Electronic mail may be sent via Internet to psycsher@mizzou1.missouri.edu.

[1] It is difficult to delineate the causal direction between stimuli, responses, and putative intervening constructs such as expectancies. Alcohol outcome expectancies, as cognitive representations of the anticipated consequences of drinking alcohol, are either directly or indirectly (in the case of vicarious learning) linked with previous drinking behavior, so that characterizing expectancy theory as unidirectional (from expectancy to behavior) is probably an oversimplification. It should be pointed out that expectancy theories (e.g., Bolles, 1972; Tolman, 1932), although emphasizing the influence of expectancies on behavior, do not preclude the possibility of behavior to expectancy or reciprocal effects. We would like to thank an anonymous reviewer for suggesting this point of clarification.

relationships between alcohol outcome expectancies and alcohol use that are consistent with social learning formulations. Bauman et al. (1985) observed significant cross-lagged effects between a measure of subjective expected utility of alcohol effects and alcohol use over a 1-year period in an early adolescent sample. Smith et al. (1995) examined a reciprocal influence model of outcome expectancy–alcohol use relations over a 2-year interval. They found significant prospective effects from outcome expectancy at Year 1 to alcohol use at Year 2, and from alcohol use at Year 1 to outcome expectancy at Year 2. The expectancy to alcohol use relation remained significant from Year 2 to 3, but alcohol use at Year 2 did not predict outcome expectancies at Year 3. These prospective effects are especially noteworthy in that they were demonstrated in a very conservative test (i.e., after controlling for the influence of autoregressive processes and cross-sectional covariances). Although these findings are important for examining the initiation of alcohol use, they are not as informative in describing the role that outcome expectancies may play later in development when drinking patterns have become more established and perhaps more autonomous. Therefore, prospective studies with older samples are needed to determine whether outcome expectancies play a maintaining or exacerbating role in the drinking behavior during later stages of psychosocial development. Such research is particularly important with respect to the feasibility of interventions that are based on expectancy modification (e.g., Darkes & Goldman, 1993).

One relevant study for investigating the role of outcome expectancies in older individuals was conducted by Stacy et al. (1991). In a 9-year prospective study of late adolescents/young adults, they examined relations between alcohol and drug expectancies and drug use and problems and noted, consistent with expectancy theory, significant prospective relations between drug expectancies and drug use and between alcohol expectancies and drug problems, but no evidence of a prospective effect from drug use to expectancies. Another prospective study of late adolescents/young adults (Stacy, Widaman, & Marlatt, 1990) also supported a direct influence from expectancies to alcohol use but did not investigate use to expectancy relations or reciprocal influence processes and therefore is not informative with respect to the other two classes of theories previously noted.

In contrasting the results of the three previous prospective studies that investigated reciprocal relations it is unclear whether observed differences in the pattern of results are more closely tied to sample differences or the length of the measurement intervals. Recall that both studies using early adolescent samples (Bauman et al., 1985; Smith et al., 1995) conducted their assessments at 1-year intervals, whereas Stacy et al. (1991) assessed their late adolescent/young adult participants over a 9-year period. Selection of the most "meaningful" measurement interval has been identified as a critical methodological issue in prospective research (Cohen, 1991; Gollob & Reichardt, 1987, 1991; Kessler & Greenberg, 1981). Unfortunately, it is difficult to discern, a priori, what constitutes the most appropriate measurement lag. If the interval is too long, critical periods of change may be missed. Alternatively, if the interval is too short, there may not have been sufficient time for a given variable to exert its effect (Gollob & Reichardt, 1987). Gollob and Reichardt (1991) demonstrate that autoregressive processes at least partially underlie the variations in effects observed over different time intervals. In addition to these methodological issues, considerations of age and developmental sequelae are also inextricably linked to the decision regarding which measurement intervals to include in a study. For example, in a study of adolescent drinking, one could argue that shorter measurement intervals would be more appropriate because early-to-mid adolescence is a period during which many individuals initiate alcohol use. Alternatively, in a study of late adolescents/young adults, with more stable alcohol use patterns, it may be that longer measurement intervals would more appropriately model alcohol use or its escalation to problematic levels. In any event, there appears to be some consensus among those discussing the issue that, given that one can never know the "true" causal interval, it is desirable to examine multiple measurement intervals (Kessler & Greenberg, 1981; Gollob & Reichardt, 1987).

In summary, reciprocal influence models have received empirical support in two of the three prospective studies that have attempted to identify cross-lagged relations between outcome expectancies and alcohol use. It appears, at least for early adolescents, that outcome expectancies play an important role in the initiation of alcohol use, and alcohol use in turn strengthens outcome expectancies, thus contributing to a positive feedback process that may influence subsequent alcohol use (Smith et al., 1995).

Although the important prospective role of outcome expectancies in the early development of drinking patterns is not disputed, at this point in time, with the exception of the Stacy et al. (1991) study, the prospective role of expectancies in the development of drinking patterns in young adulthood is largely unexplored. It remains to be seen whether the pattern of reciprocal influences observed by Bauman et al. (1985) and Smith et al. (1995) extends beyond adolescence, an important period of initiation of alcohol use, to young adulthood where it could ostensibly play an important role in the maintenance of drinking and its escalation to problem levels.

This study seeks to extend and clarify what is known about alcohol outcome expectancies and alcohol use relations by prospectively assessing each in a multiwave, mixed-gender sample of college students at low and high risk for the development of alcoholism. The composition of this sample, with respect to being largely college students and of varying family history and gender, allows us to examine several relevant aspects of the alcohol outcome expectancy–drinking behavior literature from the context of a longitudinal design.

More specifically, college students are more likely than nonstudents of the same age to report heavy drinking, and, more generally, young adults report high levels of negative consequences related to alcohol use as compared with other age groups (Johnston, O'Malley, & Bachman, 1991). Moreover, the transition from high school to college is an important developmental stage that is often associated with increased opportunity for social interaction, which, for the first time, largely occurs outside the domain of parental or other supervisory control. Additionally, drinking in college, as opposed to in high school, tends to be fairly stable. Thus, examining reciprocal relations between alcohol outcome expectancies and alcohol use in the college years can potentially explicate the role of outcome expectancies in the maintenance and escalation of drinking in a sample already initiated into high levels of alcohol involvement.

Moreover, previous cross-sectional research on both our own college student sample (Sher, Walitzer, Wood, & Brent, 1991) and with other adolescent samples (Brown, Creamer, & Stetson, 1987; Mann et al., 1987) has found significant differences in the strength of alcohol outcome expectancies as a function of family history of alcoholism status. Both Mann et al. (1987) and Brown et al. (1987) found heightened expectancies for cognitive and motor functioning among individuals with a positive family history of alcoholism, suggesting that these types of expectancies may be one of the mechanisms by which risk is transmitted to behavior.

Finally, scant data exist regarding gender differences in alcohol outcome expectancies. Brown, Goldman, Inn, and Anderson (1980) found that women were more likely to endorse expectancies measuring "general positive social expectancies," and men were more likely to endorse expectancies measuring arousal and aggression. Alternatively, Rohsenow (1983) found that women expected fewer positive, global, social and physical pleasure and relaxation effects and had greater expectations of impairment effects than men. Leigh's (1987) study mirrored Rohsenow's with respect to increased impairment expectancies among women, and also found significantly greater expectations for "nastiness" among men compared with women. In previous cross-sectional analyses with our sample, Sher et al. (1991) found that men reported significantly stronger outcome expectancies than women for social lubrication, activity enhancement, and performance enhancement. Generalizations across studies are made difficult by the variety of different expectancy measures used, nonetheless, greater expectations of impairment among women has been a replicable finding.

Longitudinal studies examining gender and risk status differences in relations between alcohol outcome expectancies and alcohol use are virtually nonexistent, but researchers need them to investigate whether changes in outcome expectancies over time (and drinking experience) vary as a function of these factors. The sample composition of our study enables prospective examination of these unexplored potential differences in alcohol use and outcome expectancies as a function of gender and family history status over a period of time associated with relatively stable, high levels of alcohol use.

Accordingly, the present study was conducted to accomplish three major goals. First, it was hypothesized that findings of reciprocal influences between outcome expectancies and alcohol use from prospective studies with early adolescent samples would generalize to late adolescents/young adults who have greater personal experience with alcohol, as well as more stable alcohol use patterns. To test this hypothesis, nested structural equation models were specified to assess autoregressive and cross-lagged prospective relations between alcohol outcome expectancies and alcohol use. Second, models assessing prospective effects over both 1 and 3 years were specified to examine the importance of follow-up interval in the expression of functional relations between alcohol outcome expectancies and alcohol use. Third, we sought to examine whether prospective relations between outcome expectancies and alcohol use vary as a function of gender and family history of alcoholism.

Method

Participants

Baseline Sample

Participants at baseline, 489 freshmen at a large, midwestern university, were recruited from an initial screening sample of 3,156 entering, first-time freshmen (representing approximately 80% of all entering, first-time freshmen).[2] Participants who reported a family history of paternal alcoholism on a version of the Short Michigan Alcoholism Screening Test (SMAST; Selzer, Vinokur, & van Rooijen, 1975) adapted for assessing father's (F-SMAST) drinking problems (Crews & Sher, 1992) and on a Family History–Research Diagnostic Criteria (FH-RDC) interview (Endicott, Andreasen, & Spitzer, 1978) were classified children of alcoholics (COAs). A comparably sized sample of participants who reported an absence of substance use disorders in all first- and second-degree biological relatives and an absence of antisocial personality disorders in all first-degree biological relatives were classified as controls (non-COAs). (See Sher et al., 1991, for a more complete description of participant recruitment.)

Longitudinal Sample

Prospective analyses are based on 465 individuals (95% of those targeted for follow-up) including 458 participants who had complete data for all four waves and 7 participants who were missing data that could be estimated using means from data at surrounding waves. The prospective sample consisted of 109 male COAs, 127 female COAs, 111 male non-COAs, and 118 female non-COAs. To assess sample bias, 465 participants (i.e., the 458 with available data at all waves and the 7 for whom missing data could be estimated) were compared with the 22 subjects who did not have sufficient data for inclusion in the prospective analyses.[3] Specifically, these groups were contrasted on the four Wave 1 expectancy measures and the four Wave 1 consumption measures (see below) using t tests and on family history status and gender using chi-square tests of association. None of these analyses revealed significant between-groups differences (all $ps > .05$).

[2] Sample sizes differ slightly from those reported by Sher, Walitzer, Wood, and Brent (1991). Because of an ongoing system of checking data for unreliable or impossible values, some errors have been found and some unreliable data have been discovered. Two participants were deleted completely from this dataset after the Sher et al. (1991) article was published (1 was found to be adopted and the other provided questionable data on both the questionnaire and interview). Data from 1 participant were added late because the family history status was finalized after the Sher et al. article was published. Finally, 3 participants made errors in their self-reported coding of gender, and these errors were recently discovered and corrected. Note, however, that none of these changes (which affected approximately 1% of the total observations) alter the basic findings reported earlier.

[3] Data from only 22 (versus 24) could be compared with the prospective sample because 2 participants' Year 1 questionnaire data were deleted because their responses indicated a large number of impossible (i.e., out of range) values.

Measures

Alcohol Outcome Expectancies (EXP)

Forty-four items measuring positive expectations of alcohol's effects were used to assess the latent construct of alcohol outcome expectancies. On the basis of previous confirmatory and exploratory factor analyses (see Kushner, Sher, Wood, & Wood, 1994), four subscales were formed using unit-weighting scoring. Subscale scores were used as factor indicators and included (a) tension reduction (9 items; $\alpha = .89$), for example, "Drinking makes me feel less tense or nervous"; (b) social lubrication (8 items; $\alpha = .88$), for example, "Drinking makes me feel less shy"; (c) activity enhancement (9 items; $\alpha = .85$), for example, "Drinking makes many activities more enjoyable"; (d) performance enhancement (9 items; $\alpha = .81$), for example, "Drinking helps me have better ideas." To enhance item variability and yield a more suitable correlation matrix for factor analyses, response scales for each item ranged from 0 (*not at all*) to 4 (*a lot*).

Alcohol Use (AU)

The following measures were used to assess the latent construct of alcohol use: (a) total quantity/frequency (QF) of alcohol consumption, calculated from individual QF estimates for beer, wine, wine coolers, and hard liquor during the past 30 days; (b) frequency of alcohol consumption per week based on the past year; (c) quantity of alcohol consumption per drinking occasion, based on the past year; (d) number of heavy drinking occasions (5 or more drinks on a single occasion) per week, based on the past 30 days. We were interested in using both past month and past year estimates because the former are less likely to be affected by recall bias, whereas the latter are less likely to be affected by transient fluctuations in drinking patterns. Each of these items were assessed with continuous response scales and were converted to weekly equivalents.[4] Coefficients alpha for the four items constituting the AU latent variables ranged from .82 to .88 over the four measurement occasions.

Results

Changes in Alcohol Use and Alcohol Outcome Expectancies Over Time

Prior to constructing our latent variable models, we conducted 2 × 2 × 4 (Family History × Gender × Time) analyses of variance (ANOVAs) on our primary measures of interest, the alcohol consumption variables and alcohol outcome expectancies. As seen in Figure 1, COAs reported higher levels on frequency of alcohol consumption per week, $F(1, 461) = 8.48, p < .01$. Men reported higher levels on each of the four consumption measures ($ps < .0001$). For quantity of alcohol per drinking occasion, there was a significant Family History × Gender × Time interaction, $F(3, 1383) = 3.50, p < .02$, with use for men remaining fairly stable over time, whereas patterns of consumption for COA and non-COA women tended to converge over time.[5]

COAs reported higher levels of tension reduction, social lubrication, and performance enhancement alcohol outcome expectancies ($ps < .01$), and men reported higher levels of each of the four outcome expectancy subscales ($ps < .01$). As can be seen in Figure 2, alcohol outcome expectancies decreased over time for each of the subscales ($ps < .0001$). Profile contrast analyses indicated, generally, that outcome expectancies remained relatively stable over the first 2 years of college, but then decreased significantly over the remaining 2 years. Omnibus tests revealed no significant higher order interactions between gender and family history status and time.[6]

Latent Variable Models

Four-Wave Models

Measurement model specification. Our structural equation modeling approach was based on the two-step method proposed by Anderson and Gerbing (1988). The initial measurement model specified associations between measured and latent variables, and bidirectional associations (covariances) among latent and manifest study variables (i.e., gender). It was expected that measurement errors of like indicators (e.g., quantity–frequency) would be associated over time, therefore error covariances of lags one, two, and three were estimated (Anderson & Gerbing, 1988). Additionally, because invariance of factor loadings over time is conceptually desirable and has been suggested as an important criteria for meaningful comparisons in longitudinal data (Hoyle & Smith, 1994), we constrained like indicators to be the same across measurement occasions. The overall fit of this initial measurement model was acceptable, $\chi^2(454, N = 465) = 1{,}284.10, p < .0001$, Comparative Fit Index (CFI) = .93, Nonnormed Fit Index (NNFI) = .92, Normed Fit Index (NFI) = .90, indicating that the specified model closely fit the data. All factor loadings from measured to latent variables were significant, with standardized path estimates ranging from .63 to .88 for expectancy indicators and from .50 to .92 for alcohol use indicators. To examine whether the assumption of invariance at the level of factor loadings was tenable, the measurement model was respecified with factor loadings free to vary across measurement occasions. The overall fit of this model was also acceptable, $\chi^2(436, N = 465) = 1{,}203.03, p < .0001$, CFI = .94, NNFI = .92, NFI = .91. Next

[4] Copies of the specific items, their response options, and associated scoring programs are available from the first author.

[5] Subsequent trend analyses indicated several higher order interactions for the alcohol consumption variables. For frequency of alcohol consumption per week, there was a significant Family History × Time (Linear) interaction, $F(1, 461) = 5.49, p < .02$, with COA's scores tending to converge with non-COA's scores over time. For quantity of alcohol per drinking occasion, there was a significant Family History × Time (quadratic) interaction, $F(1, 461) = 10.91, p < .001$; non-COAs demonstrated more of a quadratic effect than COAs, with a maximum at Time 3. For heavy drinking in the past 30 days, there was a significant Gender × Time (linear) interaction, $F(1, 461) = 5.04, p < .05$, with heavy drinking occasions showing a slight increase over time for men and a slight decrease over time for women.

[6] More fine-grained trend analyses indicated a significant Family History × Time (linear) interaction for performance enhancement outcome expectancies, $F(1, 461) = 5.17, p < .05$, in which COAs demonstrated a steeper decrease in these expectancies over time than did non-COAs. Additionally, for tension reduction outcome expectancies, there was a significant Family History × Gender × Time (quadratic) interaction, $F(1, 461) = 5.72, p < .02$, in which COA men demonstrated more of a quadratic effect over time, while COA women's expectancy scores tended to parallel those of non-COA men and women over time.

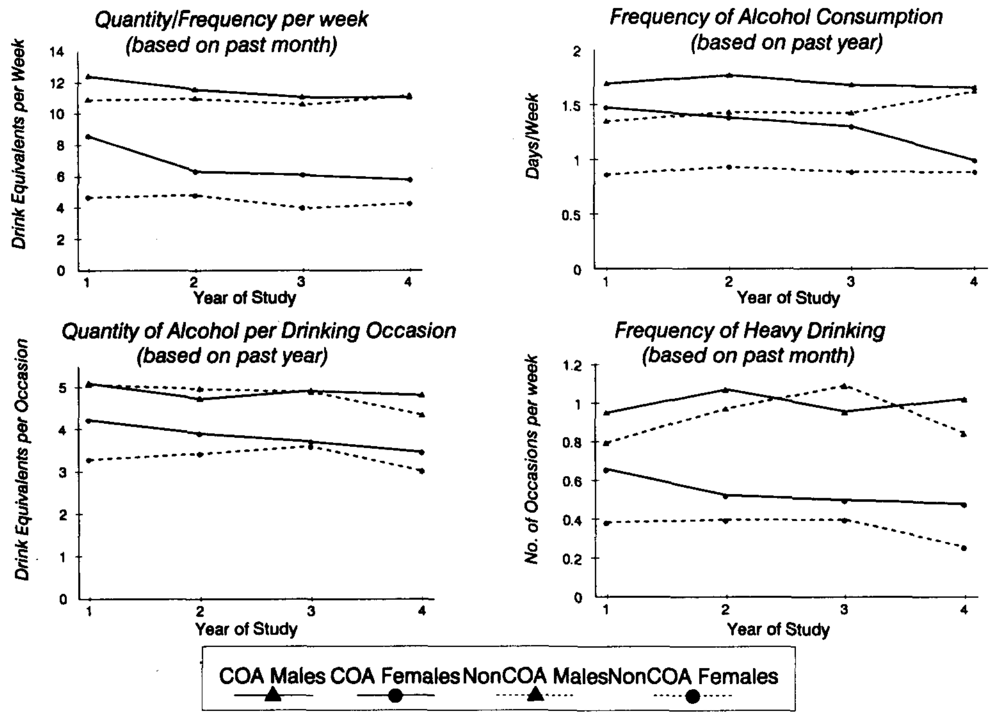

Figure 1. Course of alcohol consumption over four annual waves. COA = children of alcoholics; NonCOA = not children of alcoholics.

we compared these two models on the basis of chi-square difference scores and indexes of relative (incremental) fit (Marsh, Balla, & McDonald, 1988). Relaxing the constraints for factor loadings over time resulted in a significant improvement in model fit based on both the chi-square difference score, $\chi^2(18) = 81.07$, $p < .01$, and more conservative estimates of incremental fit (relative Type II NNFI = .04, relative Type II Akaike's Information Criterion [AIC] = .58). Note that these types of fit indexes indicate the increment in fit of nested hierarchical models with respect to plausible alternative models and should not be compared with traditional fit measures that are based on comparison with an implausible "independence" null model. Although the unconstrained model demonstrated significant increments in model fit over the constrained model, examination of the pattern of standardized factor loadings indicated minimal differences across the two models. Therefore in the interest of parsimony and measurement invariance over time, we elected to estimate our base and comparison structural models with factor loadings for like indicators constrained to be equal over time.

Examination of structural relations. With respect to structural relations between EXP and AU over time, we initially estimated a base model with autoregressive but no cross-lagged paths as a basis for comparison with later structural models. The base and subsequent structural models all had (a) gender as an exogenous manifest variable with paths estimated to each of the EXP and AU latent variables;[7] (b) family history as an exogenous factor with the variance set equal to 1, the path estimate for the error of the observed family history variable constrained to .2861 as a correction for attenuation due to measurement unreliability[8] (Cohen, Cohen, Teresi, Marchi, & Velez, 1990), and paths estimated to EXP and AU at Year 1[9]; (c) EXP and AU as fully endogenous autoregressive processes of lag 1 (e.g., Year 1 to Year 2, Year 2 to Year 3); (d) covariances estimated between contemporaneous disturbances; (e) one factor loading for each of the AU and EXP latent variables constrained to be equal to 1; (f) factor loadings for like indicators

[7] We estimated paths between gender and each of the AU and EXP latent constructs because we were interested in examining whether the associations observed at baseline would remain similar or change over the course of the study. In the four-wave models, for gender to AU relations we found significant but decreasing associations over the 4 years of the study. Men demonstrated higher levels of alcohol involvement; standardized path coefficients were −.22 at baseline, −.10 at Year 2, −.08 at Year 3, and −.09 at Year 4. Relations between gender and AU were also significant at both measurement occasions in the two-wave model. For gender to EXP in the four-wave model, the relations at baseline were significant (−.14, denoting stronger endorsement of expectancy items among men), but decreased to nonsignificant levels at Years 2, 3, and 4. In the two-wave model, both the gender to EXP Year 1 and gender to EXP Year 4 paths were significant.

[8] This value represents the square root of the variance in Family History status that is accounted for by error.

[9] Initial models estimated relations between Family History and AU and EXP at each of the four waves. Because there were no significant family history effects beyond those demonstrated at Year 1 for both EXP and AU, we elected not to estimate these paths in analyses presented in text.

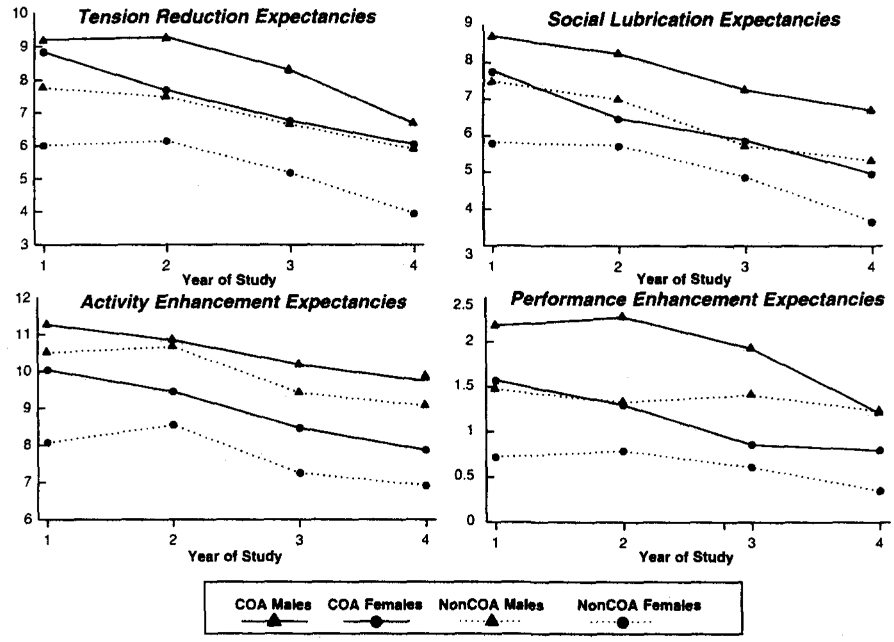

Figure 2. Course of alcohol outcome expectancies over four annual waves. COA = children of alcoholics; NonCOA = not children of alcoholics.

constrained to be equal over time; (g) covariances of lags 1, 2, and 3, estimated between error variances of like indicators; and (h) although close to zero by design, the covariance between Family History and gender was estimated. As can be seen in Table 1, the base model fit the data adequately, $\chi^2(478, N = 465) = 1{,}400.84, p < .0001$, CFI = .93, NNFI = .91, NFI = .89.

Separate structural models estimated paths from the Alcohol Use latent variables to the Expectancy latent variables (AU → EXP) to examine the effect of alcohol use on beliefs about the effects of drinking over 1-year intervals and from the Expectancy latent factors to the Alcohol Use latent factors (EXP → AU) to examine the effect of beliefs on drinking behavior over the same interval. As displayed in Table 2, the inclusion of the three AU → EXP paths provided slight but significant increments in model fit over that of the base model, $\Delta\chi^2(3) = 23.61, p < .01$, relative Type II NNFI = .016, relative Type II AIC = .531. Moreover, there were significant prospective effects from AU to EXP for each of the three paths, with standardized path coefficients ranging from .08 to .15. The EXP → AU model also demonstrated a significant but modest increment in model fit over the base model, $\Delta\chi^2(3) = 9.47, p < .05$, relative Type II NNFI = .001, relative Type II AIC = .105, and significant prospective associations between EXP and AU for two of the three specified paths (for the Year 1 EXP to Year 2 AU path, $p < .05$, one tailed). Standardized path coefficients ranged from .02 to .09.

Table 1
Traditional Fit Measures for the Four-Wave Model

Base or structural model	Traditional fit measure				
	χ^2	*df*	NFI	NNFI	CFI
Base	1,400.84	478	.89	.91	.93
AU → EXP	1,377.23	475	.90	.92	.93
EXP → AU	1,391.37	475	.90	.91	.93
FCLP	1,368.19	472	.90	.92	.93

Note. NFI = Bentler-Bonett Normed Fit Index; NNFI = Bentler-Bonett Nonnormed Fit Index; AU = Alcohol Use; EXP = Outcome Expectancy; FCLP = Full Cross-Lagged Panel; CFI = comparative fit index.

Table 2
Chi-Square Difference Tests and Type II Relative Nonnormed Fit Index for the Four-Wave Model

Base or structural model	Target model $\Delta\chi^2$ (*df*)		
	AU → EXP	EXP → AU	FCLP
	Hierarchical χ^2 (*df*) difference tests		
Base	23.61 (3)**	9.47 (3)*	32.65 (6)**
AU → EXP		na	9.04 (3)*
EXP → AU			23.18 (3)**
	Relative Type II NNFI fit measures		
Base	.016	.001	.017
AU → EXP		na	.004
EXP → AU			.016

Note. AU = Alcohol Use; EXP = Outcome Expectancy; FCLP = Full Cross-Lagged Panel; NNFI = Bentler-Bonett Nonnormed Fit Index; na = cannot be compared because the models are not nested.
* $p < .05$. ** $p < .01$.

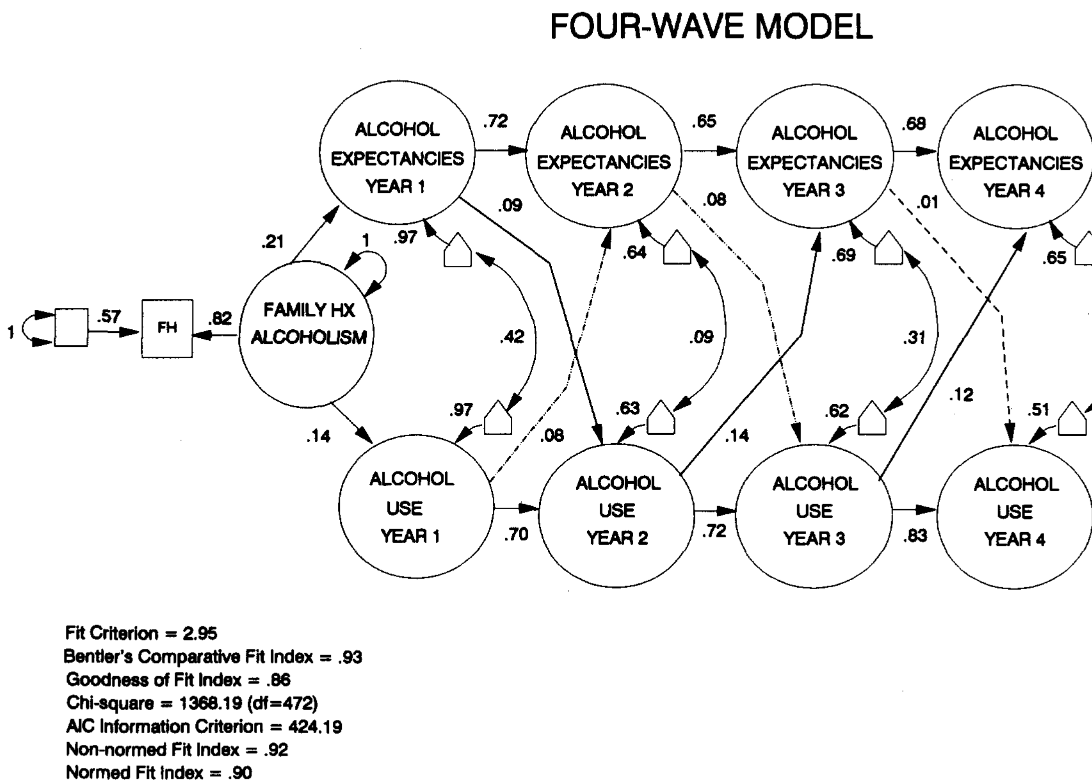

Figure 3. Standardized solution for the four-wave model. Solid lines represent paths $p < .05$ for a two-tailed test. Dash-dotted lines represent $p < .05$ for a one-tailed test. Dashed lines represent $p > .05$. Estimates from gender to each of the Outcome Expectancy and Alcohol Use latent variables are not shown. Family History (HX) latent construct variance was set to 1, and the path for the error of the observed family history variable was set to .2861 in the unstandardized solution (Crews & Sher, 1992). One factor loading for each of the Alcohol Use and Outcome Expectancy latent variables was constrained to be equal to 1, and factor loadings for like indicators (e.g., Quantity/Frequency [QF] Year 1, QF Year 2, QF Year 3, QF Year 4) were constrained to be equal across measurement occasions. Error variances were free to vary across measurement intervals and covariances of lag 1, lag 2, and lag 3 from measurement errors of like indicators were estimated.

The final four-wave structural model included cross-lagged paths of lag 1 in order to examine possible reciprocal relations between expectancies and alcohol use over time. As with the base and previous structural models, the overall fit of the full cross-lagged panel (FCLP) model was acceptable, $\chi^2(472, N = 465) = 1{,}368.19$, $p < .0001$, CFI = .93, NNFI = .92, NFI = .90. Comparison of the FCLP with the base model indicated a significant increment in model fit, $\Delta\chi^2(6) = 32.65$, $p < .01$, with modest increments on more conservative indexes of relative fit (relative Type II NNFI = .017, relative Type II AIC = .62). As can be seen in Figure 3, examination of the specific cross-lagged path coefficients indicated statistically significant, but modest, prospective effects for each of the three AU to EXP paths and for two of the three EXP to AU paths (for the Year 1 EXP to Year 2 AU path, $p < .05$, one tailed).

Two-Wave Models

Although the four-wave model takes into account all the data, it is important to note, as discussed in the introduction, that previous research differs with respect to the time course over which cross-lagged effects have been examined. To evaluate the hypothesis that one might see a different pattern of prediction over longer intervals, we constructed a series of two-wave models to examine relations between outcome expectancies and alcohol use over a 3-year interval.

Measurement model specification. As with the four-wave model, we used a two-step approach to examine relations between EXP and AU over the longer (3-year) interval (Anderson & Gerbing, 1988). Again, the initial measurement model specified covariances among all latent factors and manifest study variables (i.e., gender). Covariance estimates from measurement errors of like indicators were estimated and factor loadings for like indicators were constrained to be equal over both measurement occasions. Consistent with the four-wave measurement model, overall model fit for the two-wave measurement model was acceptable, $\chi^2(120, N = 465) = 390.20$, $p < .0001$, CFI = .95, NNFI = .92, NFI = .93. All factor loadings from measured to latent variables were significant with loadings that very closely approximated those from the four-wave model. Relaxing the equality constraints for like factor loadings resulted in a nonsignificant increment in model fit, $\Delta\chi^2(6) = 5.59$, *ns*; therefore, consistent with the four-wave model, we elected to estimate structural models with like indicators constrained to be equal over time.

Examination of structural relations. The specification of the two-wave base model was identical to that of the four-wave model, except that because there were only two measurement occasions, only error covariances of lag 3 were estimated. As can be seen in Table 3, traditional fit indexes for the two-wave base model indicated a reasonable fit of the data to the model, $\chi^2(124, N = 465) = 406.17$, CFI = .94, NNFI = .93, NFI = .92.

Separate structural models were estimated to examine the AU → EXP relation, the EXP → AU relation, and reciprocal relations (FCLP). As can be seen in Table 4, the addition of the AU → EXP path did not result in a significant increment in model fit, $\Delta\chi^2(1) = 1.77$, *ns*, relative Type II NNFI = .005, relative Type II AIC = .007, whereas the addition of the EXP → AU path did significantly increment model fit, $\Delta\chi^2(1) = 15.92$, $p < .01$, relative Type II NNFI = .045, relative Type II AIC = .407. Comparison of the FCLP model with the EXP → AU model resulted in a nonsignificant increment in model fit, $\Delta\chi^2(1) = 0.43$, relative Type II NNFI = .010, relative Type II AIC = .074. Consistent with the four-wave model, the two-wave model demonstrated significant (and larger magnitude) prospective prediction from EXP to AU. In contrast to the four-wave model, no prospective effects were found from AU to EXP (see Figure 4).

Robustness testing. To evaluate a number of potential threats to the validity of our model testing procedures we recalculated our final cross-lagged models (described above and referred to in Table 5 as the standard analyses) in a variety of ways. First, because of concerns about the effects of departures from normality on study findings, we recomputed our models: (a) after subjecting the data to normalizing transformation procedures[10] and (b) using a robust estimation option designed to provide more conservative standard error estimates for tests of the significance of factor loadings. Second, in order to assess the possible influence of multivariate outliers on our findings, we recomputed our analyses after eliminating the five participants with the largest values on Mardia's coefficient. Third, we excluded those participants who were Wave 1 abstainers (n = 31; defined conservatively as reporting no alcohol consumption during the previous year). We undertook these analyses because individuals who were abstainers during their freshman year would be relatively unlikely to show much variability in use over the course of the study. Finally, we recalculated both final models using only participants who remained either full- or part-time students ($N = 275$) at the same university throughout the course of the study. Our reason for these final analyses was to examine the effects of interest in the context of continuous college enrollment at the same university. We anticipated (and indeed found) that restricting the sample in this way would lead to reduced cross-lagged prediction for two reasons: (a) the homogeneity of the college environment over time would be expected to increase autoregressive stability and consequently decrease unexplained criterion variance; and (b) eliminating drop outs, stop outs, and transfers has the effect of decreasing the proportion of excessive and problematic drinkers (Wood, DeBord, & Sher, June 1994). Nevertheless, we felt that these ancillary analyses would be useful in demonstrating biases that might be introduced by failing to track individuals who left the university.

As can be seen in Table 5, for the two-wave model, the prospective effects from EXP to AU were robust with respect to all of the threats to validity that we tested. The magnitude of the EXP to AU effect in the model estimated using only participants who were continuously enrolled at the same university was reduced but still significant (standardized path coefficient = .10, $p < .05$, one tailed). Likewise, the prospective effects from AU to EXP observed in the four-wave model were largely unaffected by most of the procedures described above. In only two instances did significant AU to EXP effects observed in the standard analyses fail to emerge in the robustness testing analyses. Specifically, when models were estimated with the five largest multivariate outliers deleted and with only those individuals who remained enrolled at the same university for the 4 years of the study, the Year 1 AU → Year 2 EXP path became nonsignificant. Thus, on the basis of these ancillary analyses, we conclude that there is clear evidence for a prospective effect of alcohol expectancies on alcohol consumption over an extended time period (i.e., 3 years). However, in both the standard and robustness testing analyses, the evidence for this effect over shorter time intervals (i.e., 1 year) is weak. On the other hand, there is also evidence for prospective effects from use to expectancies, but only at the shorter time interval (i.e., 1 year). We also note that if one considers only those participants continuously enrolled at the same university for 4 years, the magnitude of the AU → EXP link is slightly decreased, thus pointing to the importance of following participants who leave campus.

Invariance analyses. To evaluate the generalizability of our models across gender and family history status, we conducted multigroup analyses by gender and again by family history status with both the two-wave and four-wave models. Given the large

Table 3
Traditional Fit Measures for the Two-Wave Model

Base or structural model	Traditional fit measures				
	χ^2	*df*	NFI	NNFI	CFI
Base	406.17	124	.92	.93	.94
AU → EXP	404.40	123	.92	.93	.94
EXP → AU	390.25	123	.92	.93	.95
FCLP	389.83	122	.92	.93	.95

Note. NFI = Bentler-Bonett Normed Fit Index; NNFI = Bentler-Bonett Nonnormed Fit Index; AU = Alcohol Use; EXP = Outcome Expectancy; FCLP = Full Cross-Lagged Panel; CFI = comparative fit index.

[10] Data were normalized using the BLOM procedure (SAS Institute, 1989). Briefly, each variable is rank transformed and the resulting rank (i.e., percentile) is then replaced by its corresponding standard normal equivalent.

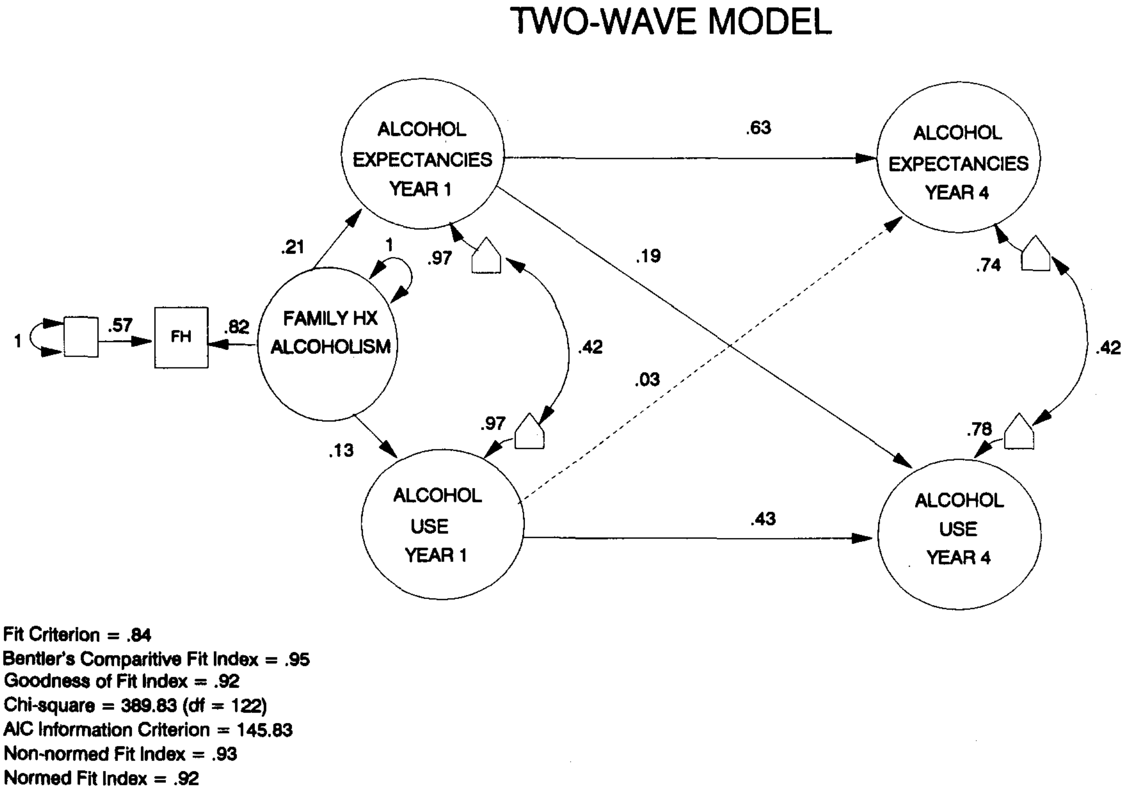

Figure 4. Standardized solution for the two-wave model. Solid lines represent paths $p < .05$ for a two-tailed test. Dashed lines represent paths $p > .05$. Estimates from gender to each of the Outcome Expectancy and Alcohol Use latent variables are not shown. Family HX latent construct variance was set to 1, and the path for the error of the observed family history variable was set to .2861 in the unstandardized solution (Crews & Sher, 1992). One factor loading for each of the Alcohol Use and Outcome Expectancy latent variables was constrained to be equal to 1, and factor loadings for like indicators (e.g., Quantity/Frequency [QF] Year 1, QF Year 4) were constrained to be equal across measurement occasions. Error variances were free to vary across measurement intervals and covariances of lag 3 from measurement errors of like indicators were estimated.

Table 4
Chi-Square Difference Tests and Type II Relative Nonnormed Fit Measures for the Two-Wave Model

Base or structural model	Target model $\Delta\chi^2$ (*df*)		
	AU → EXP	EXP → AU	FCLP
	Hierarchical χ^2 (*df*) difference tests		
Base	1.77 (1)	15.92 (1)**	16.35 (2)**
AU → EXP		na	14.57 (1)**
EXP → AU			0.43 (1)
	Relative Type II NNFI fit measures		
Base	.005	.045	.035
AU → EXP		na	.040
EXP → AU			.010

Note. AU = Alcohol Use; EXP = Outcome Expectancy; FCLP = Full Cross-Lagged Panel; NNFI = Bentler-Bonett Nonnormed Fit Index, na = cannot be compared because the models are not nested.
** $p < .01$.

number of parameters to participants in some of the analyses, invariance hypotheses were grouped according to conceptual criteria outlined by Bentler (1989; see also Horn, McArdle, & Mason, 1983). Specifically, we tested the hypotheses that all structural parameters were identical across populations (strict invariance) against four alternative models. First, models in which all parameters except for error variances and covariances were constrained to be equal across groups were estimated (factor covariance, pattern and structural path invariance models). Second, in addition to the error components described in the factor covariance, pattern, and structural path invariance models, disturbance variances and contemporaneous disturbance covariances for the endogenous latent variables of AU and EXP were allowed to vary across groups (factor pattern and structural path invariance models). Third, we specified models in which all parameters of the model were different across groups except factor loadings (factor pattern invariance models). In the fourth set of models, all parameters were free to vary across the two groups (no invariance models). Note that these models are nested within each other. Evaluation of the best-fitting invariance model for the data was done by examination of standard fit sta-

Table 5
Cross-Lagged Standardized Path Coefficients From Standard Analyses and Robustness Testing

Model	Standard analyses	Robust standard error estimation	Multivariate outliers eliminated	Abstainers eliminated	Transformed data	Continuous enrollment
			Four-wave models			
EXP → AU						
Year 1 → 2	.09*	.09	.05	.10*	.09*	.06
Year 2 → 3	.08†	.08	.00	.09*	.08†	.00
Year 3 → 4	.01	.01	.02	.01	.01	.03
AU → EXP						
Year 1 → 2	.08†	.08†	.06	.08*	.08†	−.03
Year 2 → 3	.14*	.14*	.13*	.11*	.14*	.12*
Year 3 → 4	.12*	.12†	.12*	.14*	.12*	.10*
			Two-wave models			
EXP → AU	.19*	.19*	.18*	.21*	.19*	.10†
AU → EXP	.03	.03	.04	.00	.03	.06

Note. Robust standard error estimation refers to the statistical significance patterns of the maximum likelihood estimates under "standard analyses" using the robust standard error estimate described by Bentler (1989). EXP = Outcome Expectancy; AU = Alcohol Use.
† $p < .05$, one tailed. * $p < .05$, two tailed.

tistics such as the goodness-of-fit index (GFI), CFI, NNFI and NFI, chi-square difference tests between models, and relative fit indexes (relative Type II NNFI and AIC).

For the gender invariance analyses of the four-wave model, we concluded that the model of factor pattern and structural path invariance best represented the data, $\chi^2(917, N = 465) = 2{,}161.95, p < .01$, CFI = .90, NFI = .84, NNFI = .89. Although a statistically significant chi-square difference score resulted from comparison of the factor pattern and structural path invariance model with the factor pattern invariance model, $\Delta\chi^2(14) = 67.08, p < .01$, little gain was found for either traditional or relative measures of fit (e.g., CFI went from .90 to .91, relative fit indexes were incremented < .1). Examination of the differences in disturbance variances from the factor covariance, pattern, and path invariance model indicated substantially higher disturbance terms for the men than for the women in this study. For these reasons, we conclude that these data can be taken as support for the conclusion that the structural paths between alcohol use and alcohol expectancies in this model are roughly the same for men and women, although parameters related to measurement error and the variability in the Alcohol Use and Alcohol Expectancies latent variables differ substantially for men and women. The invariance results from the two-wave models by gender basically paralleled that of the four-wave analyses. That is, the chi-square difference score comparing the factor pattern and structural path invariance model with the factor pattern invariance model was significant, with little evidence for model differences with respect to either traditional or relative indexes of fit (CFI's were .91 and .93, respectively, and relative fit indexes were <.2).[11]

We were unable to conduct similar invariance analyses for the four-wave model by family history due to the failure of some of the models to converge, even if start values were provided at the overall solution and if the iterations specified were generous. For that reason, our discussion of invariance by family history is limited to the two-wave model. For the two-wave model, the model of factor pattern and structural path invariance was,

[11] Because there was some evidence against gender invariance at the level of structural paths, we conducted six additional invariance tests for the four-wave model and two for the two-wave model. In each of these analyses, first, all parameters were free to vary by gender except factor patterns (loadings). Then we constrained one cross-lagged effect at a time to be equal across gender to examine whether Lagrange multiplier tests would indicate whether relaxing these constraints would improve model fit. For the four-wave models, in only one case did Lagrange tests indicate that the constraints were not tenable in our sample. Specifically, for the four-wave model, the equality constraint for the path from Year 3 AU to Year 4 EXP was not supported by the data, $\chi^2(1) = 11.67$, $p < .001$. When not constrained to be equal by gender, the unstandardized path coefficient from Year 3 AU to Year 4 EXP was .16, $p < .001$, for women, and .02, *ns*, for men, indicating that the main effect noted in the overall model appears to characterize women more than men at this wave. Note that gender differences were not found for any of the other significant cross-lagged effects in the overall model. For the two-wave models, Lagrange tests indicated that neither the AU to EXP nor the EXP to AU equality constraints were supported by the data, $\chi^2(1) = 8.69, p < .01$, and $\chi^2(1) = 4.52, p < .05$, respectively. Of most importance, the EXP to AU effect was significant for both men and women (unstandardized path coefficients from EXP to AU were .56, $p < .001$, for men and .30, $p < .01$, for women), but stronger for men than women in the unstandardized values. Note that the standardized values were .23 for both men and women. With respect to the AU to EXP path in the two-wave model, there was no evidence for an effect for men (unstandardized path coefficient = −.017, *ns*), but there was a significant association for women (unstandardized path coefficient = .10, $p < .05$), indicating that some caution is warranted in generalizing across gender. These differences should also be interpreted in the context of smaller autoregressive effects for women compared with men.

again, the most reasonable interpretation. The chi-square difference score comparing the factor pattern and structural path invariance model with the factor pattern invariance model was not statistically significant and both traditional and relative fit measures were relatively unaffected (e.g., CFI did not differ across the two models, relative fit indexes < .1). Overall fit for the model of factor pattern and structural path invariance was adequate, $\chi^2(230) = 556.33$, $p < .0001$, CFI = .94, NNFI = .92, NFI = .90.

Clearly, as previously noted, the less-than-favorable ratio of parameters to participants in the study and the number of invariance hypotheses tested are grounds for caution in the interpretation of these invariance models. The slight gains in fit for the models relaxing constraints for the structural paths may represent ceiling or floor effects in the manifest variables, differential censoring of the male and female samples, nonlinearity in the measurement model, or some combination thereof. Nevertheless, the general pattern of results (i.e., that the structural relationships between AU and EXP are the same for men and women and for family history positive and negative individuals, but that men and family history positive individuals are more variable in AU and EXP than women and family history negative individuals) seems reasonable and addresses the concern that the proposed model is an amalgam of completely different structural models.

Discussion

Previous studies of the relation between alcohol use and alcohol outcome expectancies have revealed robust cross-sectional associations between these two constructs, but the direction of effect has not been systematically studied in prospective investigations. Although existing longitudinal data indicate prospective, bidirectional influences between alcohol outcome expectancies and alcohol use during *adolescence* (when most individuals are beginning their initiation into alcohol use; Bauman et al., 1985; Smith et al., 1995), the extent that similar effects occur in older subjects who have already established relatively stable alcohol use patterns is unclear. Stacy et al. (1991) found outcome expectancies during adolescence predicted drug use 9 years later, but they observed no evidence for a reciprocal effect (i.e., adolescent drug use predicting later expectancies).

As noted by Stacy et al. (1991), the direction of influence between outcome expectancies and alcohol use has important implications for theories of the relation between cognitive variables and substance use. Thus, the seemingly different patterns of findings resulting from early-to-mid adolescent samples (Bauman et al., 1985; Smith et al., 1995) and late adolescent to young adult samples (Stacy et al., 1991) could indicate that the relation between outcome expectancies and alcohol use differs as a function of stage of development. Consequently, different theories of the relation between outcome expectancies and alcohol use might be needed to explain the role of expectancies in the initiation of alcohol use versus the maintenance (and further development) of alcohol use patterns in experienced drinkers.

Before examining the issue of direction of influence between alcohol use and outcome expectancies, we need to consider the major differences in the mean trajectory of alcohol use and alcohol expectancies between those reported in adolescent samples and our late adolescent/young adult sample. Previous research has suggested that outcome expectancies increase and become more homogenous from childhood through adolescence (Christiansen, Goldman, & Inn, 1982; Miller et al., 1990), and Smith et al. (1995) reported an increase in outcome expectancies for social facilitation in their 3-year study. In contrast, we note a significant decrease in outcome expectancies over time in our older sample. The picture that results from piecing our findings together with those of adolescence researchers is that expectancies for (at least certain forms) of reinforcement from alcohol increase over the course of adolescence, then plateau and begin to moderate in early adulthood. Perhaps outcome expectancies, which are though to be formed through vicarious learning, are initially strengthened by direct associative experience with alcohol (as well as myriad other social learning factors such as interaction with peers, media portrayals, etc.), but over the course of time and repeated exposure are subsequently tempered by such experiences.

It is also of note that in contrast to the mean increases in alcohol use found in adolescent samples (e.g., Smith et al., 1995), our sample was characterized by a relatively stable pattern of alcohol use over the college years. Consequently, it permits an examination of the direction of effect in individuals who have established (and often heavy) drinking patterns. Thus in our study we were able to examine the extent to which expectancies play a role in the maintenance of drinking patterns.

Our findings are also generally consistent with the small body of research that has examined the relation between alcohol use and outcome expectancies prospectively. Of particular note is the demonstration of reciprocal prospective effects between alcohol outcome expectancies and alcohol use. The cross-lagged effects we observed were not of large magnitude, nor were they particularly robust in the case of predicting use from expectancies in the four-wave model. However, it is important to note that they were obtained in conservative analyses (i.e., after controlling for autoregressive effects and cross-sectional covariances).[12] Moreover, they were obtained in the context of relative stability of drinking patterns in young adulthood.

Although the hypothesized pattern of reciprocal influences between alcohol outcome expectancies and alcohol use was observed, the nature of the prospective effects appears to be determined by the interval between measurement periods. Over 1-year intervals, there is evidence of a "reverse" effect, that is, alcohol outcome expectancies appear to vary as a function of

Although we report these numbers in the interest of thoroughness, given the sample size of the present study the estimates are less precise than would be obtained in studies with larger samples.

[12] It is important to control for cross-sectional covariances because these associations may be due to either contemporaneous direct effects or other causal factors not included in the model. Moreover, not estimating cross-sectional covariances when they exist has been shown to have a biasing effect on autoregressive and cross-lagged estimates (Anderson & Williams, 1992).

prior alcohol use (a pattern consistent with behavioral choice/self-perception approaches; Bem, 1978; Vuchinich & Tucker, 1988).[13] However, over longer measurement intervals (i.e., 3 years), there was clear evidence for a robust, prospective effect of alcohol outcome expectancies on subsequent drinking behavior (a pattern consistent with expectancy theory; Bolles, 1972). These effects were obtained after we controlled for variance associated with autoregressive effects and cross-sectional covariances. Furthermore, although these participants were drinking fairly heavily, and, as the analysis of variance findings show, relatively consistently over this period of time, we were still able to demonstrate prospective effects from outcome expectancy to alcohol use. Correlational studies such as this one cannot be used to infer causality; nevertheless, our findings are very consistent with the view that outcome expectancies play an important role in the maintenance of alcohol use, and perhaps in the escalation of drinking to problematic levels.

Our findings may seem paradoxical, because, a priori, one might expect better prospective effects over shorter rather than longer measurement intervals. Therefore, our finding that the relation between outcome expectancies and alcohol use was more robust over 3 years than over 1 year deserves comment. It may be more difficult to demonstrate prospective relations to behavior from more traitlike aspects of the individual, such as outcome expectancies, over shorter intervals because, particularly in late adolescence and early adulthood, there is a great deal of stability in the behavior to be predicted. Thus, over shorter measurement intervals, more of the variance in alcohol use is accounted for by previous alcohol use, leaving relatively lower levels of variance to be predicted by potential etiologic and maintaining factors. In fact, our results are quite consistent with this explanation. Recall that in the two-wave models, the autoregressive path for AU is much smaller (standardized path coefficient = .44) than the shorter, 1-year interval path coefficients in the four-wave model (M = .72).[14]

Alternatively, if one is interested in discovering the consequences of alcohol consumption, shorter time intervals can often be more sensitive because the effects of drinking on outcome expectancies might be relatively transient in young adults (or other individuals with substantial drinking experience). It also seems reasonable to speculate that in studies examining expectancy influences on drinking behavior among early adolescents, the most meaningful causal lag is likely to be one of shorter interval, given that adolescence is a time when many individuals make the transition into regular drinking. This notion is consistent with the findings reported by Smith et al. (1995), who noted significant associations between expectancy and drinking over both of the 1-year intervals they examined in their study of adolescents.

Returning to the issue raised by Stacy et al. (1991) of whether the pattern of directional effects can inform us to the relevance of three broad classes of explanations between cognitive variables and behavior, our findings suggest that the issue is more complex than it appears on the surface. That is, the nature of the prospective, functional relations between cognitive and behavioral variables cannot be determined independently of the time intervals between measurement occasions. Unfortunately, theories of the relation between alcohol outcome expectancies and alcohol use have not yet developed sufficient precision to provide clear guidelines on what these might be.

Of particular interest in this study is the relation between family history of alcoholism and alcohol expectancies, and the extent that alcohol expectancies mediate part of the risk associated with familial alcoholism. The finding that COAs differed significantly from non-COAs on measures of outcome expectancies is consistent with some previous research (Brown et al., 1987; Mann et al., 1987) and extends our previous analysis of baseline relations (Sher et al., 1991) to an extended observation period. As before, our findings suggest that outcome expectancies mediate part of the risk associated with a family history of alcoholism, albeit the magnitude of this mediation is small at any measurement occasion (e.g., see Figures 3 and 4). Note, however, in the current study the focus has been on alcohol use, not alcohol abuse or dependence, and other data suggest stronger family history relations with more problematic outcomes.

A related question concerns the relations among family history of alcoholism, alcohol outcome expectancies, and drinking behavior. Some authors (e.g., Penick, Read, Crowley, & Powell, 1978) have suggested that family history can be used to subtype alcoholics with respect to etiological processes. If different etiological processes are implicated for family history positive versus family history negative processes, it is reasonable to examine whether the structural relations among the constructs of alcohol expectancy and alcohol use differ as a function of family history. Consistent with recent research (Molina, Chassin, & Curran, 1994), we found no evidence to implicate different etiologic relations between family history positive and family history negative groups. However, given the limitations of our invariance analyses (i.e., less than optimal participant/parameter ratio, inability to estimate the family history four-wave models), our findings are in need of replication in larger samples.

Our invariance analyses also suggested that the structural relations we noted between alcohol outcome expectancies and alcohol use are basically consistent across groups defined on the basis of gender. However, these findings also should be interpreted with caution because of the low participant/parameter ratio and some evidence against structural invariance (although neither the increments in model fit nor the observed differences in the magnitude of effects were large).

In summary, the results of this study are important for both substantive and methodological reasons. Conceptually, these findings indicate both an etiologic and maintaining role for alcohol outcome expectancies in predicting future alcohol use in individuals with substantial drinking experience and the influence of alcohol consumption on the development and mainte-

[13] As noted, reciprocal effects between outcome expectancies and alcohol use were observed in the four-wave model (see Figure 3). However, in ancillary analyses we found that the expectancy to alcohol use paths *were not* particularly robust, while the alcohol use to expectancy paths *were* consistent across most of our robustness testing analyses (see Table 3).

[14] We thank an anonymous reviewer for noting the relevance of our findings to this point.

nance of alcohol outcome expectancies. However, these reciprocal effects appear to have different time-bound functional relations. Methodologically, they point to the importance of the consideration of measurement interval in longitudinal research.

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