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Sexual Orientation Self-Concept Ambiguity: Scale Adaptation and Validation

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Abstract

The current article describes the adaptation of a measure of sexual orientation self-concept ambiguity (SSA) from an existing measure of general self-concept clarity. Latent “trait” scores of SSA reflect the extent to which a person's beliefs about their own sexual orientation are perceived as inconsistent, unreliable, or incongruent. Sexual minority and heterosexual women ($n = 348$), ages 18 to 30, completed a cross-sectional survey. Categorical confirmatory factor analysis guided the selection of items to form a 10-item, self-report measure of SSA. In the current report, we also examine (a) reliability of the 10-item scale score, (b) measurement invariance based on respondents' sexual identity status and age group, and (c) correlations with preexisting surveys that purport to measure similar constructs and theoretical correlates. Evidence for internal reliability, measurement invariance (based on respondent sex), and convergent validity was also investigated in an independent, validation sample. The lowest SSA scores were reported by women who self-ascribed an *exclusively heterosexual* or *exclusively lesbian/gay* sexual identity, whereas those who reported a *bisexual*, *mostly lesbian/gay*, or *mostly heterosexual* identity, reported relatively higher SSA scores.

Keywords

scale adaptation; sexual self-concept; self-concept clarity; sexual minority

Winter (1988) defines the sexual self-concept as a person's evaluation of her or his own sexual feelings and actions. By extension, the *sexual orientation self-concept* can be described as a person's evaluation of her or his own sexual orientation. Individuals' sexual orientation self-concept is believed to be informed by various aspects of their sexuality, including their sexual attraction (i.e., degree of opposite- vs. same-sex sexual desire), sexual behavior (i.e., degree of opposite- vs. same-sex sexual activity), and sexual identity (i.e., labeling of self-identification; Galupo, Davis, Gryniewicz, & Mitchell, 2014; Laumann, Gagnon, Michael, & Michaels, 1994). Researchers who assess multiple facets of sexual

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orientation (e.g., sexual identity, sexual attraction, sexual behavior), find that these facets are often discordant within individuals across time (Burgard, Cochran, & Mays, 2005; Manley, Diamond, & van Anders, 2015; Smith, Rissel, Richters, Grulich, & Visser, 2003), or even within the same interview assessment (Brooks & Quina, 2009; Chandra, Mosher, Copen, & Sionean, 2011; Igartua, Thombs, Burgos, & Montoro, 2009; Narring, Huwiler, & Michaud, 2003; Pathela, Blank, Sell, & Schillinger, 2006; Ross, Essien, Williams, & Fernandez-Esquer, 2003). Facets of sexual orientation are often assessed independent of one another and as such, discordance among these facets is inherent to the current methodology used to assess aspects of sexuality in survey studies (e.g., sexual attraction vs. behavior, bipolar vs. unipolar ratings of attraction).

In studies that use separate items to assess facets of sexual orientation, women often endorse more discordance among and changes in facets of their sexual orientation than men (Diamond, 2000; Ott, Corliss, Wypij, Rosario, & Austin, 2011; Ott et al., 2013; Peplau & Garnets, 2000). Diamond (2000) posits that transitions, or “fluidity,” with regard to sexual identity and other facets of sexual orientation may be “an inevitable consequence of the fact that most sexual-minority women experience attractions for both sexes, although in widely varying degrees” (p. 242). Recent studies (Diamond, 2000, 2008, 2012; Galupo et al., 2014; Katz-Wise & Hyde, 2014; Ott et al., 2013; Savin-Williams, Joyner, & Rieger, 2012) provide evidence that women report high levels of variability over their lifetime with regard to both their sexual identity and the gender of their sexual and romantic partners. Recent work by Manley et al. (2015) suggests that sexual fluidity is also evident among men, although not to the same extent as in women.

When attempting to understand the impact of self-reported incongruence among lifetime, recent, and current assessments of sexual orientation, empirical research often relies on manipulations and transformations of mathematical associations among at least two or three facets of sexual orientation (e.g., Nield, Magnusson, Brooks, Chapman, & Lapane, 2015; Ott et al., 2011; Talley, Brown, Cukrowicz, & Bagge, 2015). Extant findings show that incongruence among independently assessed facets of sexual orientation may confer risk of negative health outcomes, including disordered alcohol use and elevated depression and anxiety symptoms (e.g., Bauer, Jairam, & Baidooobonso, 2010; Drabble, 2004; Gattis, 2013; Talley et al., 2015). It is believed that the dominant culture influences sexual identity development and transitions by promoting the assumption of a static, heterosexual identity. This creates potential challenges for women and men who are attempting to explore, update, and integrate same-sex sexual attractions and behaviors into their sexual self-concept (Tharinger & Wells, 2000). Ostensible incongruence or ambiguity with regard to self-reported facets of sexual orientation may, for some persons, reflect greater difficulty integrating aspects of a marginalized sexual identity into the self-concept. Nevertheless, it may also reflect individuals' attempts to amend previous characterizations of their sexual orientation self-concept, particularly during active identity explorations, the initiation or dissolution of romantic relationships, or increasing awareness of changes in sexual attraction. We believe it is reasonable to conjecture that for some individuals, greater sexual fluidity may, at times, relate to feelings of uncertainty regarding the sexual self-concept, but the two constructs are not isomorphic with one another, and sexual orientation sexual-

concept ambiguity (SSA) is not viewed as an inevitable consequence of changes or transitions in facets of sexual orientation or sexual expression.

Early theoretical models of sexual identity development (e.g., Cass, 1979; Troiden, 1989) described *identity confusion* as an exploratory stage that involved questioning cultural assumptions regarding one's own sexual orientation—for example, exploring same-sex attraction in environments with ubiquitous heterosexism. This stage was considered essential for integrating a nonexclusively heterosexual identity into the self-concept (Rosario, Schrimshaw, Hunter, & Braun, 2006). Although later models of sexual identity development acknowledge considerable variability in the process of identity formation and note that sexual identity is now considered to be more “fluid” than linear- or stage-process models of sexual identity represent (e.g., Diamond, 1998; Savin-Williams, 1998; Schneider, 2001), the concept of identity confusion remains central to understanding sexual identity development.

It has been suggested (e.g., Schneider, 2001; Yarhouse & Tan, 2004) that the process of sexual questioning, experiencing same-sex attraction, and engaging in behavior with same-sex partners will result in greater feelings of sexual identity uncertainty in American society, which stigmatizes same-sex sexual attraction and behavior. The *Identity Uncertainty* subscale of the Measure of Sexual Identity Exploration and Commitment (MoSIEC; Worthington, Navarro, Savoy, & Hampton, 2008) and the *Identity Uncertainty* subscale of the Lesbian, Gay, and Bisexual Identity Scale (LGBIS; Mohr & Fassinger, 2000) purport to assess a lack of clarity or certainty with regard to one's own sexual orientation. A significant limitation of the MoSIEC subscale is that one of the three items comprising the sub-scale is a simple negation of another item (i.e., “My sexual orientation is *not* clear to me” and “My sexual orientation is clear to me”), which may contribute to poor item criterion validity, content saturation, and inflated test–retest item stability (Holden & Fekken, 1990) as well as potentially create confusion among respondents (Weijters & Baumgartner, 2012). By contrast, because the LGBIS assesses six dimensions of lesbian, gay, and bisexual *identity*, the item content is primarily relevant to those who currently *identify* as bisexual, lesbian, or gay, but not to those who may be initiating the process of sexual questioning or who self-ascribe an exclusively heterosexual identity, despite that they may acknowledge same-sex attraction, behavior, or both (Diamond, 2000, 2005, 2008, 2012). Finally, an examination of the items from the MoSIEC and LGBIS subscales suggest they capture a diffuse sense of agnosticism regarding one's sexual identity, per se, yet do not assess perceived incongruence among facets of sexual orientation, which may be key for understanding why and under what conditions identity uncertainty contributes to negative psychosocial outcomes.

Thompson and Morgan (2008) found that women self-identifying as “mostly straight,” “bisexual,” and “lesbian” reported higher levels of *identity uncertainty* than “exclusively straight” women, as measured by the corresponding subscale of the MoSIEC (Worthington et al., 2008). Higher scores on the *identity uncertainty* subscale of the LGBIS (Mohr & Kendra, 2011) were shown to be associated with greater identity concealment, internalized homonegativity, and difficulty with the identity development process. Consistent with these findings, qualitative data (Thompson & Morgan, 2008) suggested that difficulties inherent in reconciling and labeling sexual attraction, behavior, and identity were common in narratives from persons expressing identity uncertainty. Taken together, findings show that persons

reporting a nonexclusively heterosexual identity may be at risk for reporting periods of relative uncertainty regarding their sexual orientation and these periods may confer risk of maladaptive health outcomes.

Work in this area would benefit from a valid and reliable scale to ensure researchers are adequately able to triangulate on levels of ambiguity that people perceive with regard to both their sexual orientation self-concept, in general, as well as among central facets of their sexual orientation. Adapting Campbell et al.'s (1996) definition of general self-concept clarity (SCC), SSA, by extension, can be defined as a person's awareness that facets of their own sexual orientation are perceived as inconsistent, unreliable, or uncertain. Again, prior work has shown that discrepancies across facets of sexual orientation are relatively common, especially among women (Diamond, 2003; Saewyc, Magee, & Pettingell, 2004; Savin-Williams, 2009; Savin-Williams & Vrangalova, 2013; Talley et al., 2015), yet no existing measure captures awareness of perceived discrepancies among sexual identity, attraction, and behavior. Despite that identity explorations, including sexual questioning, and discrepancies among various facets of sexual orientation are somewhat normative, particularly in young adulthood (e.g., Savin-Williams & Diamond, 2000; Schulenberg & Maggs, 2002), the extant literature is relatively silent about the role of inconsistencies among facets of sexual orientation in contributing to health-related disparities.

Recent theoretical work (Talley & Littlefield, 2014) outlined cognitive, affective, and behavioral pathways that lead to maladaptive health outcomes for persons who have adopted or are living with concealable stigmatized identities. The proposed framework highlighted the importance of identity-related characteristics, including sexual self-concept ambiguity, in contributing to the physical and mental health outcomes of sexual minority-identified persons. Notably, recent empirical work has supported hypothesized associations between the proposed SSA scale and alcohol misuse, risky sex, and suicidal ideation (e.g., Talley, Brown, Stevens, & Littlefield, 2014; Talley et al., 2015). Given previous findings in the extant literature, as well as pertinent theoretical models (see Meyer, 2003; Talley & Littlefield, 2014), we sought to provide evidence for criterion validity by showing that higher levels of SSA will be associated with greater alcohol misuse, depression symptoms, and anxiety symptoms.

The purpose of the current work is to adapt an existing measure of general SCC (Campbell et al., 1996) to the sexual orientation self-concept domain using a calibration sample of women who were recruited expressly for this purpose. We adapted items to reference a lack of clarity or certainty with regard to facets of sexual orientation. Categorical confirmatory factor analysis (CCFA) was used to guide scale adaptation and ensure a unidimensional factor structure that assessed the SSA construct efficiently and reliably (Cheung & Rensvold, 2002). After a final set of items was identified for the SSA scale, measurement invariance testing (Cheung & Rensvold, 2002; Kim & Yoon, 2011; Millsap & Olivera-Aguilar, 2013; Vandenberg & Lance, 2000) was conducted across age and sexual identity subgroups. We also examined regression models for expected positive associations between levels of SSA and self-reported severity of past-year alcohol use disorder (AUD) symptoms and distress caused by recent depression and anxiety symptoms, over and above associations with existing measures of sexual identity uncertainty. Subsequently, the factor structure of

the selected SSA items was validated in an independent sample of young adults in which measurement invariance testing was examined with regard to participants' biological sex. In both the calibration and validation samples, we examined the hypothesis (Mohr & Kendra, 2011) that SSA scores would be higher among individuals who identify as *primarily heterosexual*, *equally heterosexual and gay/lesbian* [e.g., *bisexual*], *primarily lesbian/gay*, and *exclusively lesbian/gay*, compared with their *exclusively heterosexual* counterparts, yet also posited that *exclusively lesbian/gay* individuals may be less likely to report heightened levels of ambiguity with regard to their sexual identity.

Method

Participants

The development of the SSA scale utilized data from a study examining how aspects of women's sexuality may influence their health risk behavior (e.g., Talley et al., 2014; Talley et al., 2015). Importantly, the calibration sample was recruited for the expressed purpose of adapting and validating this measure of SSA. For this reason, women who indicated any degree of same-sex sexuality in their lifetime were oversampled at a rate of 2 to 1. Specifically, we oversampled women who, in response to a phone screening, ever identified as “nonexclusively heterosexual,” ever acknowledged any degree of same-sex attraction, or ever acknowledged any same-sex sexual behavior.

The calibration sample is composed of 348 women, ranging in age from 18 to 30 years ($M = 20.93$, $SD = 2.92$). Approximately 84% of participants identified as White or Caucasian, 10.8% as Black or African American, 4.8% as Asian, and 4.6% as Hispanic or Latina (choose all that apply). A majority of women in the calibration sample (58%, $n = 200$) self-identified as *exclusively heterosexual* at the time of the survey. Additional self-reported, sexual identity categories ($n = 147$) at the time of the survey included *exclusively lesbian/gay* (7.5%, $n = 26$), *primarily lesbian/gay* (2.9%, $n = 10$), *equally lesbian/gay and heterosexual* [e.g., *bisexual*] (6.3%, $n = 22$), *primarily heterosexual* (20.4%, $n = 71$), *questioning/unlabeled* (2.0%, $n = 7$), *queer* (2.6%, $n = 9$), and “other” ($n = 2$, i.e., “heterosomething”; “pansexual”). One participant did not respond to the item assessing sexual identity, but was retained in analyses because she completed all of the SSA items.

Procedure

Community and student participants were recruited into an online survey study from the fall of 2010 to the spring of 2013. Psychology students at a large Midwestern university were invited to enroll in a study examining “female sexuality and alcohol use.” Community participants were recruited through campuswide e-mail notices, advertisements in local newspapers, posted flyers in local businesses, and snowball sampling techniques. Women indicating interest in the online survey were administered a brief telephone screening to determine eligibility. Eligible persons (a) self-identified as female, (b) were between the ages of 18 and 30 years (inclusive), (c) were English-speaking, and (d) self-reported that they had never been in treatment for a substance use disorder. In addition, because another expressed goal of the larger study was to examine how identity-related aspects of women's sexuality may influence their alcohol use and consequences, only women who reported

consuming at least one alcoholic beverage in the prior 3 months were invited to participate. Eligible participants were e-mailed a link to a confidential, password-protected online survey, which allowed for consent and contained an embedded random-code identifier for reimbursement purposes. Students received course credit in return for their participation. Eligible community participants received a \$25 gift certificate to Amazon.com in exchange for their time. All study procedures and materials were reviewed and approved by the University of Missouri Institutional Review Board.

Measures

The SCC scale was developed by Campbell et al. (1996) to assess “the extent to which self-beliefs are clearly and confidently defined, internally consistent, and stable” (p. 141). The original scale included 12 items assessing individuals' level of SCC. To adapt the SCC scale to be relevant for the assessment of SSA, changes were made to the wording of question stems. For example, the question in the SCC scale, “Sometimes I feel that I am not really the person that I appear to be” was altered to “Sometimes I feel that my sexual orientation is not really what it appears to be.” In addition to altering the wording of the original 12 items of the SCC scale, three additional items were adapted from a list of clinically solicited indicators of general *identity diffusion* (Wilkinson-Ryan & Westen, 2000). These items, which are marked in Table 1, were adapted by the first author to ensure that items considered for inclusion in the adapted scale were sensitive to assessing perceived discrepancies among facets of sexual orientation, both within and across time. The original SCC scale was also included in the survey study to provide support for discriminate validity (Fiske, 1982).

Prior to completing the SSA scale–adapted items, participants read the following instructions: “Your sexual orientation is defined by your self-identification (e.g., lesbian, gay, bisexual, heterosexual) as well as your sexual attraction, sexual fantasies, and sexual behavior. Please consider all of these aspects of your sexual orientation when responding to the questions below. Please rate the extent to which you agree or disagree with each of the following statements.” The SSA scale items were administered using 4-point Likert-type response scales (1 = *strongly disagree*, 4 = *strongly agree*).

The MoSIEC (Worthington et al., 2008) was developed to assess hypothesized dimensions of sexual identity development. The scale contains 22 items and scores can be calculated for four primary subscales, namely *Commitment*, *Exploration*, *Uncertainty*, and *Synthesis*. Respondents are asked to indicate their endorsement of each MoSIEC item on a 6-point Likert-type response scale (1 = *very uncharacteristic of me*, 6 = *very characteristic of me*). Previous work by the scale authors indicated that scores demonstrated high internal reliability (Dillon, Worthington, Soth-McNett, & Schwartz, 2008; Worthington et al., 2008). An example item is, “My sexual orientation is not clear to me.” In the calibration sample, Cronbach's alpha for the three-item *Identity Uncertainty* subscale of the MoSIEC in the calibration sample was .67. The *Identity Uncertainty* subscale score of the MoSIEC (Worthington et al., 2008) was used in the current report to provide evidence for convergent validity of SSA scale scores to establish support that these two measures assess similar constructs (Fiske, 1982).

The LGBIS (Mohr & Kendra, 2011) was developed to assess a range of beliefs and feelings regarding one's sexual minority identity, which are believed to have clinical or theoretical relevance for psychosocial functioning. The scale contains 27 items and scores can be calculated for six primary subscales, namely *Need for Privacy*, *Need for Acceptance*, *Identity Uncertainty*, *Internalized Homonegativity/Binegativity*, *Difficult Process*, and *Superiority*. Respondents were asked to indicate their endorsement of each item on a 7-point Likert-type response scale (1 = *disagree strongly*, 7 = *agree strongly*). Four items from the *Identity Uncertainty* sub-scale of the LGBIS, which purports to measure “uncertainty about one's sexual orientation identity” (Mohr & Kendra, 2011, p. 239), were also used in the current report to provide evidence for convergent validity. Nevertheless, because the scale is worded to pertain to persons who currently identify as lesbian, gay, or bisexual, only women who indicated a nonexclusively heterosexual identity at the time they completed the survey were asked to respond to these items, of which a subset did ($n = 113$). An example item is, “I am not totally sure what my sexual orientation is.” Cronbach's alpha for *Identity Uncertainty* subscale scores in the calibration sample was good ($\alpha = .81$).

Measurement invariance testing (i.e., differential item functioning) was examined based on two respondent characteristics in multigroup CCFA structural equation models (SEM; Bollen, 1989; Cheung & Rensvold, 2002; Kim & Yoon, 2011). Testing the invariance of estimated SEM parameters based on age and sexual identity status provides evidence that the adapted SSA items have similar meaning and elicit consistent responding across these subgroups (Cheung & Rensvold, 2000). Participants were coded into two age groups based on a median age value of 21 in the calibration sample (i.e., younger [<21 years of age; $n = 210$] vs. older [$21\text{--}30$ years of age, $n = 138$]). A dichotomous indicator of self-reported sexual identity status (i.e., *exclusively heterosexual* = 0 [$n = 200$]; *primarily heterosexual–equally heterosexual and lesbian/gay–primarily lesbian/gay–exclusively lesbian/gay* = 1 [$n = 147$]), at the time of the survey, was also created for the purposes of testing measurement invariance. A chi-square test of independence suggested that women 21 years of age or older were more likely than those younger than 21 to identify as nonexclusively heterosexual, $\chi^2(1) = 25.17, p < .001$.

Items from the Young Adult Alcohol Problems Screening Test (Hurlbut & Sher, 1992) were used to calculate the frequency that participants had experienced nine AUD symptoms in the previous 12 months. The Young Adult Alcohol Problems Screening Test measure was adapted from one used in previous studies (Hurlbut & Sher, 1992) and included a response scale ranging from *never* to *40 or more times* in the past year. The Hopkins Symptom Checklist (Derogatis, Lipman, Rickels, Uhlenhuth & Covi, 1974) was used to assess distress caused by depression and anxiety symptoms in the previous 2 weeks, on a Likert-type scale (0 = *not at all*, 3 = *extremely*). Cronbach's alpha information is contained in Table 3.

Analytic Procedures

CFA SEMs with categorical latent factor indicators were used to adapt the SSA measure. CCFA was conducted in Mplus Version 7.11 (L. K. Muthén & Muthén, 1998–2013), and a weighted least squares means and variance adjusted (WLSMV) estimator was specified. In CCFA, item factor loadings and thresholds are estimated separately for each item, yet both

provide information on potential differential item functioning. Factor loadings represent slope estimates, indicating the rate at which the likelihood of item endorsement changes as SSA factor scores increase. Factor loadings reflect how efficiently an item, for example, can distinguish women characterized by low SSA factors scores from those characterized by moderate or high SSA factors scores. On the current scale, responding *strongly disagree* and *disagree* to SSA items would indicate lower levels of ambiguity and responding *agree* and *strongly agree* would reflect higher levels of ambiguity.

The final items selected for the SSA scale in the calibration sample were examined for metric invariance and residual variance invariance with respect to respondent age and sexual identity status (Cheung & Rensvold, 2002; Horn & McArdle, 1992; Kim & Yoon, 2011). The equivalence of factor loadings (λ) and thresholds (ν) across groups is referred to as metric and scalar invariance, respectively (Bollen, 1989; Cheung & Rensvold, 2002). We chose to fix the item with the highest factor loading on the overall latent factor at 1 (i.e., Item 15), specifying it as the marker indicator. Because selection of the marker variable can affect model results (Little, Lindenberger, & Nesselroade, 1999; Millsap, 2001), we also tested alternative models to ensure that specifying other items as marker variables did not alter final scale adaptation. Evidence for invariance of residual variances (Bollen, 1989) is provided by testing whether residual variances can be constrained across subgroups without a significant decrease in model fit. Residual variance invariance implies that the degree of measurement error is similar across groups. Because WLSMV was specified as the estimator in CCFA models, nested factor models were compared using the DIFFTEST function in *Mplus*.

Results

Prior to conducting CCFA, response frequencies (see Table 1) for the two reverse-scored items (SSA1R and SSA14R) were examined and distinct patterns of responding indicated that these items may have created confusion among participants. Evidence from three supplemental analyses suggested that these items should be removed. First, the two reverse-scored items had item–total correlations lower than .45, whereas all other items had item–total correlations above .71. Smaller relative item–total correlations suggested that the two reverse-scored items did not correlate highly with the overall scale score (Everitt & Skrondal, 2002). Second, when we examined the rotated exploratory factor analysis solution in SPSS Version 22 (specifying principal axis factor estimation, which entails no distributional assumptions; Fabrigar, Wegener, MacCallum, & Strahan, 1999), results showed that the two reversed-scored items had communalities below .30 (.28 and .26, respectively), whereas all other items had communalities above .54. Again, evidence from an examination of communalities suggested that these two items were not highly correlated with the extracted factor (Gorsuch, 1983). Third, estimated categorical exploratory factor analysis models (i.e., CEFA) in *Mplus* (with geomin rotation) that included the two reverse-scored items did not show adequate model fit. Specifically, both one- and two-factor solutions showed significant chi-square test values and root mean square of approximation (RMSEA) values above .05 (Hu & Bentler, 1999). Moreover, the second factor, in the two-factor CEFA model, was composed only of the reverse-scored items. These two items cross-loaded highly on both factors. Specifically, item SSA1R loaded at .49 on the primary scale

factor (with a majority of other items) and .38 on the second factor, whereas item SS14R loaded at .52 on the primary scale factor and .38 on the second factor. The two extracted latent factor scores in the CEFA model were only moderately correlated (.24). Item cross-loadings suggested complex relations between the only two reverse-scored items and all other adapted items. Seeking a unidimensional factor structure, we concluded the two reverse-scored items were problematic (Worthington & Whittaker, 2006). Notably, the coefficient alpha of the overall scale score, if these two items were removed, would remain the same as when both were included ($\alpha = .95$). At this stage, we elected to remove the two reverse-scored items.

Because the items were adapted from a previous scale in which the factor structure was known, a priori, to be unidimensional, CCFA was conducted on the remaining 13 items specifying a single latent factor and selecting the marker indicator (i.e., fixing the factor loading at 1) as the item with the highest factor loading on the overall latent factor (i.e., latent factor variance was freed). Results suggested a unidimensional factor structure adequately fit the data, $\chi^2(52) = 166.50$, $p < .001$, RMSEA = .067, (comparative fit index) CFI = .995, (Tucker–Lewis index) TLI = .994. As highlighted by Brown (2003), correlated errors among manifest indicators may reflect a number of potential method effects, including high content overlap, similar phrasings, differential susceptibility to demand characteristics, and carelessness or difficulty reading reverse-worded items (cf. Byrne, Shavelson, & Muthén, 1989; Cordery & Sevastos, 1993; Floyd & Widaman, 1995; Gerbing & Anderson, 1984; Schmitt & Stults, 1985; Spector, Van Katwyk, Brannick, & Chen, 1997). Given our expressed desire to create a unidimensional latent factor that adheres to the psychometric assumptions of local independence (Lord & Novick, 1968), while limiting participant burden, we elected to proceed with the removal of items, one-at-a-time, until model fit indices fell below recommended cutoff points (Hu & Bentler, 1999).

Modification indices were examined, along with estimated factor loadings, to identify items for removal and improve overall model fit. The removal of items ceased once a desirable model fit was achieved (i.e., chi-square p value greater than .05 and RMSEA values below .05). For the 13-item CCFA, the highest modification index value suggested correlated residuals between Items 3 and 15. Because Item 15 had a higher factor loading on the overall factor, Item 3 was removed and fit was reassessed. Overall fit increased with the removal of Item 3, $\chi^2(54) = 122.25$, $p < .001$, RMSEA = .06, (95% [confidence interval] CI [.05, .07]), CFI = .996, TLI = .995. The next largest modification index showed correlated residuals between Items 8 and 9. The estimated factor loading of Item 9 was higher than that of Item 8, indicating that Item 9 was contributing more to the overall fit than Item 8. After the deletion of Item 8, model fit improved, $\chi^2(44) = 85.57$, $p < .001$, RMSEA = .05, (95% CI [.05, .07]), CFI = .998, TLI = .997. The next largest modification index showed a correlated residual between Items 13 and 5, and an additional modification index suggested a correlated residual between Items 13 and 2. Given the complex, correlated residuals with Item 13 and other items in the scale, Item 13 was deleted. Following deletion of Item 13, no further modifications to improve overall model fit were displayed, at which point the overall model fit fell below specified cutoff values.

Ultimately, five items (SSA1R, SSA8, SSA3, SSA13, SSA14R) were removed before the model indicated excellent fit to the data, $\chi^2(35) = 47.02$, $p = .08$, RMSEA = .03, (95% CI [.00, .05]), CFI = .999, TLI = .999. As shown in Figure 1, the final 10-item, single-factor accounted for over 70% of the variance in item responding and few correlations (<4%) among item residuals (observed vs. estimated) exceeded |.04|. We considered the 10-item SSA scale to be initially adapted at this stage. The final SSA scale score demonstrated excellent reliability in the calibration sample ($\alpha = .95$).

Measurement invariance testing was conducted in *Mplus* using the DIFFTEST command in nested model comparisons (estimator = WLSMV; B. Muthén & Asparouhov, 2002; Zumbo, 2007). Theta parameterization was specified to allow for residual variances as parameters so that strict invariance tests could be conducted. We examined support for metric invariance in the calibration sample by constraining factor loadings to be equivalent across women in younger and older emerging adulthood and testing whether model fit worsened (Horn & McArdle, 1992). Metric invariance testing indicated that both younger and older women endorsed equivalent rates of change in response probabilities across equivalent levels of SSA factor scores, $\chi^2(9) = 5.12$, $p = .82$. Separate metric invariance tests were also conducted with sexual identity status as the grouping variable, $\chi^2(9) = 6.12$, $p = .73$. Nested model comparisons suggested equivalent factor loadings, and thus provided support for metric invariance, regardless of women's age in emerging adulthood or their sexual identity.

We examined support for scalar invariance by constraining item thresholds to be equivalent across subgroups (assuming metric invariance) and testing whether model fit worsened (Meredith, 1993). Scalar invariance testing indicated that both younger and older women endorsed equivalent response option probabilities at equivalent levels of SSA factor scores, $\chi^2(28) = 20.40$, $p = .85$. Separate scalar invariance tests were conducted with sexual identity status as the grouping variable, $\chi^2(28) = 24.51$, $p = .65$. Nested model comparisons revealed equivalent item thresholds, and thus provided support for scalar invariance, regardless of age or sexual identity status. Taken together, findings suggested SSA is measured equally well with the adapted 10-item scale in various age and sexual orientation subpopulations.

To establish support for strict factorial invariance, we examined evidence for invariance of residual variances by constraining residual variances to a value of one across both subgroups (assuming metric and scalar invariance) and testing whether model fit worsened with this specification (B. Muthén & Asparouhov, 2002). Multigroup models indicated that residual variances were not equivalent across younger and older women, $\chi^2(10) = 20.77$, $p = .02$. An examination of residuals for Items 2, 11, and 12 suggested that older women, compared with younger women, reported more variance for Item 2, but less for Items 11 and 12. Relaxing constraints on residual variances for these three items suggested that remaining error variances were statistically equivalent across age groups, $\chi^2(7) = 4.04$, $p = .77$. Tests of invariance of residual variances were also conducted with regard to sexual identity status. This analysis also revealed that residual variances were not invariant across sexual identity subgroups, $\chi^2(10) = 31.09$, $p < .001$. Results indicated that 5 of 10 SSA scale items elicited similar degrees of measurement error, regardless of age and sexual identity status, whereas Items 4, 5, 9, 11, and 12, tended to elicit greater variation in random responding among

women who identified as exclusively heterosexual at the time of the survey. Relaxing constraints on residual variances for those five items showed that remaining error variances were equivalent across sexual identity subgroups, $\chi^2(5) = 10.06$, $p = .07$. Taken together, we assert that noninvariance has been demonstrated by the current invariance tests, and thus feel confident that interpretable structural models can be estimated using the 10-item SSA scale (B. Muthén & Asparouhov, 2002).

Evidence for SSA Construct Validity in the Calibration Sample

Construct validity is supported by expected mean-level differences in SSA scores, based on respondents' sexual identity. A one-way analysis of variance, conducted in SPSS Version 22, showed significant differences in mean-levels of SSA scores across the Kinsey-based, 5-point sexual identity scale, Welch $F(4) = 15.12$, $p < .001$ (see Table 2). Noting unequal variances in SSA scores across sexual identity groups, we utilized Games Howell post hoc comparisons (Milligan, Wong, & Thompson, 1987), and showed that SSA scores were equivalent and low between *exclusively heterosexual* women and *exclusively lesbian/gay* women. *Primarily heterosexual*, *bisexual*, and *primarily lesbian/gay* women had higher SSA scores than *exclusively heterosexual* women. However, only *primarily heterosexual* women reported higher levels of SSA than *exclusively lesbian/gay* women.

We also sought to establish initial evidence for convergent and discriminate validity. Table 3 contains Spearman correlations between SSA total scores and assessment scores that are purported to provide support for convergent, discriminate, and criterion-related validity. An examination of correlations contained in Table 3 shows that SSA total scores correlated most highly, as expected, with existing subscale scores purporting to measure similar constructs. Specifically, the Spearman correlation between SSA total scores and MoSIEC's *Identity Uncertainty* subscale scores was high, $\rho = .62$. The Spearman correlation between SSA total scores and *Identity Uncertainty* subscale scores, for the subset of women who identified as nonexclusively heterosexual and completed the LGBIS, was also high, $\rho = .70$, $n = 113$. Taken together, these bivariate associations provide support that the SSA scale measures beliefs that aspects of one's sexual orientation are perceived as inconsistent or uncertain. Notably, the Spearman correlation between SSA scores and SCC scores was moderate, $\rho = -.39$, suggesting the content of the 10-item SSA scale is not isomorphic with the content of Campbell et al.'s (1996) SCC scale, from which SSA items were adapted.

To establish initial evidence for criterion and incremental validity, a series of bivariate and multivariate linear regression models were conducted in SPSS Version 22.0 (Hunsley & Meyer, 2003). SSA scale scores that remain significantly associated with pertinent criterion variables when *Identity Uncertainty* subscale scores from preexisting measures are included in multivariate regression models would provide evidence for incremental validity of the SSA scale. In multivariate models that included both SSA and MoSIEC's *Identity Uncertainty* scores, only SSA scores were significantly related to frequency of reported AUD symptoms in the past year, $\beta = .14$, (standard error) $SE = .05$, $p = .01$ ($R^2 = .02$, $F(1, 313) = 7.11$, $p < .01$). Despite that the bivariate association between MoSIEC's *Identity Uncertainty* scores and depressive symptoms was high, $\beta = .29$, $SE = .03$, $p < .001$, in multivariate models that included both SSA and MoSIEC's *Identity Uncertainty* scores, only

SSA scores were significantly related to the reported severity of recent depressive symptoms, $\beta = .30$, $SE = .05$, $p < .001$ ($R^2 = .02$, $F[1, 313] = 6.45$, $p = .01$). Similarly, despite that the bivariate association between MoSIEC's *Identity Uncertainty* subscale scores and anxiety symptoms was also high, $\beta = .26$, $SE = .03$, $p < .001$, in multivariate regression models that included both SSA and MoSIEC's *Identity Uncertainty* scores, only SSA scores were significantly related to the reported severity of anxiety symptoms felt in the previous 2 weeks, $\beta = .29$, $SE = .05$, $p < .001$ ($R^2 = .03$, $F[1, 314] = 9.15$, $p < .01$).

In multivariate models that included both SSA and the LGBIS's *Identity Uncertainty* subscale scores, only SSA scores were shown to relate to frequency of reported AUD symptoms in the past year, $\beta = .35$, $SE = .13$, $p = .01$ ($R^2 = .06$, $F[1, 106] = 7.24$, $p < .01$). Despite that, in bivariate models, LGBIS's *Identity Uncertainty* scores and recent depressive symptoms were related, $\beta = .27$, $SE = .04$, $p < .01$, in multivariate regression models neither SSA, $\beta = .04$, $SE = .12$ ($R^2 = .00$, $F[1, 106] = 0.09$, $p = .76$), nor *Identity Uncertainty* scores were significantly associated with depressive symptoms felt in the previous 2 weeks. Finally, although the LGBIS's *Identity Uncertainty* subscale score was related to reports of anxiety symptoms in bivariate models, $\beta = .23$, $SE = .04$, $p = .01$, neither SSA, $\beta = .13$, $SE = .12$ ($R^2 = .01$, $F[1, 109] = 0.96$, $p = .33$), nor *Identity Uncertainty* scores were significantly associated with recent anxiety symptoms in multivariate regression models.

Thus, we found limited evidence for incremental validity, above and beyond LGBIS's *Identity Uncertainty* sub-scale scores, for SSA scores in predicting internalizing symptoms. Nevertheless, in light of significant bivariate associations between levels of SSA and hypothesized criterion variables (see Table 3), there is reason to believe that the SSA scale measures beliefs that one's sexual orientation is perceived as inconsistent or uncertain and, as expected, is associated with alcohol misuse and psychological distress.

CCFA Results for Validation Sample

CCFA for the final 10-item SSA scale was conducted in an independent validation sample ($n = 1,046$) of introductory psychology students. The validation sample was composed of 474 men and 572 women, ranging in age from 17 to 47 years ($M = 18.98$, $SD = 1.84$). Approximately 83% ($n = 871$) of participants identified as White or Caucasian, 11.6% ($n = 121$) as Black or African American, 6.0% ($n = 63$) as Asian, and 3.2% ($n = 33$) as Hispanic or Latina (choose all that apply). In the validation sample, 91.0% of respondents ($n = 948$) self-identified as *exclusively heterosexual* at the time of the survey. Additional sexual identity categories at the time of the survey included *exclusively gay/lesbian* (1.0%, $n = 10$), *primarily gay/lesbian* (0.4%, $n = 4$), *equally gay/lesbian and heterosexual* (1.0%, $n = 10$), *primarily heterosexual* (5.1%, $n = 53$), *questioning/unlabeled* (0.3%, $n = 3$), *queer* (0.3%, $n = 3$), and "other" ($n = 3$, e.g., "pansexual"). A subset of individuals chose not to answer this item (1.0%, $n = 10$).

The 10-item SSA CCFA model indicated excellent fit to the data, $\chi^2(35) = 78.29$, $p < .001$, RMSEA = .03, (95% CI [.02, .05]), CFI = .999, TLI = .998. Moreover, the single-factor model explained 92% of the variance in item responding and no problematic residual correlations among error indicators were apparent. The final 10-item SSA scale scores also demonstrated excellent reliability in the validation sample ($\alpha = .95$).

We examined metric invariance with respondent biological sex as the grouping variable by constraining factor loadings to be equivalent across both women ($n = 572$) and men ($n = 474$; constraining thresholds to be equivalent across groups) and testing changes in model fit. Metric invariance testing indicated that both women and men endorsed equivalent rates of change in response probabilities across equivalent levels of SSA factor scores, $\chi^2(9) = 6.83, p = .65$. Scalar invariance testing indicated that both women and men endorsed equivalent response probabilities across equivalent levels of SSA factor scores, $\chi^2(28) = 28.59, p = .43$. As a final step, we examined evidence for residual variance invariance by constraining residual errors to be equivalent across subgroups (assuming metric and scalar invariance) and testing change in model fit (B. Muthén & Asparouhov, 2002). Models indicated that residual variances were equivalent across women and men, $\chi^2(10) = 14.03, p = .17$. Findings support that the SSA construct is measured equally well across both men and women with the adapted 10-item scale.

Further Evidence for SSA Construct Validity in the Validation Sample

Construct validity was partially supported by expected mean-level differences in SSA scores, based on sexual identity status. Specifically, SSA total scores were compared among young adults who identified as *exclusively heterosexual* ($n = 948$), *primarily heterosexual* ($n = 53$), *equally lesbian/gay and heterosexual* ($n = 10$), *primarily lesbian/gay* ($n = 4$), and *exclusively lesbian/gay* ($n = 10$) respondents. A one-way analysis of variance confirmed significant differences in mean-level SSA scores based on sexual identity, Welch $F(4) = 13.01, p < .001$. Noting unequal variances in SSA scores across sexual identity subgroups, Games Howell post hoc comparisons revealed that SSA scores were lower among *exclusively heterosexual* persons compared with *bisexual* and *primarily heterosexual* persons. No other post hoc comparisons were significant. Although the validation sample did not replicate differences between the SSA scores of various sexual identity subgroups found in the calibration sample, we note that the pattern of means is similar.

We also examined support for convergent validity in the validation sample. Nevertheless, participants were not asked to complete all items contained in the LGBIS and MoSIEC assessments and, instead, only responded to items from the relevant subscale of each respective measure. The alpha reliability for the three-item *Identity Uncertainty* sub-scale of the MoSIEC in the validation sample was unacceptable ($\alpha = .48, n = 1,041$). The Spearman correlation between SSA total scores and the MoSIEC's *Identity Uncertainty* subscale scores was moderate, $\rho = .55$, in the validation sample; however, after dropping the single, reverse-scored item ("My sexual orientation is certain to me."), the alpha coefficient improved greatly ($\alpha = .89$) and the correlation was moderate, $\rho = .64$. Among those who self-identified as nonexclusively heterosexual in the validation sample, the Spearman correlation between their SSA total score and the LGBIS *Identity Uncertainty* subscale scale ($\alpha = .88$) scores was very high, $\rho = .81, n = 77$. Findings provide evidence that the SSA scale measures ambiguity related to individuals' sexual orientation self-concept among both men and women in emerging adulthood.

Discussion

The adapted SSA scale contains 10 items, rated on a 1-(*strongly disagree*) to 4-point (*strongly agree*) response scale. The scale is easily administered and scored. Items may be averaged or summed to create an index capturing a person's current level of sexual orientation self-concept ambiguity. Convergent validity of the SSA measure was supported by strong correlations with available measures that include subscales assessing *identity uncertainty* (Mohr & Kendra, 2011; Worthington et al., 2008). Even still, the SSA scale has a number of benefits over currently available subscales that purport to assess similar constructs (e.g., MoSIEC, LGBIS). First, unlike Mohr and Kendra's (2011) LGBIS measure, the SSA scale is worded to pertain to all individuals, regardless of their self-ascribed sexual identity. Second, psychometric properties of SSA scores demonstrate robust internal reliability and initial support for criterion and discriminate validity of the SSA scale was promising. Third, invariance testing supported that the adapted SSA scale is a robust measure for men and women across early-young adulthood (between the ages of 18 and 20 years) and later-young adulthood (between the ages of 21 and 30 years) and regardless of sexual identity status. Consistent with Meyer's (2003) minority stress theory, Talley and Littlefield (2014) theorized that identity-related characteristics of sexual minority individuals, including reported levels of SSA, may account, in part, for disparities in substance misuse and other negative mental health outcomes. Thus, we sought to provide initial support for these hypothesized relations in predicting alcohol misuse, depression symptoms, and anxiety symptoms, over and above currently available assessments that purport to measure similar constructs. Findings suggested that incremental validity was somewhat supported by showing that the SSA total score was more strongly related to AUD symptoms in the past year than either of the two existing subscales, and also more strongly related to recent depressive and anxiety symptoms than the MoSIEC's *Identity Uncertainty* subscale (Worthington et al., 2008). Despite limited evidence for the incremental validity of SSA scores to demonstrate reliable associations with internalizing symptoms over and above LGBIS *Identity Uncertainty* subscale scores, we are encouraged by bivariate associations that provided support for criterion-related validity.

The SSA scale provides an efficient and reliable means to assess the extent to which aspects of a person's sexual orientation are inconsistent, unreliable, or ambiguous. Because this is a self-report scale, accurate assessment relies on reporters' acknowledgment of a lack of clarity or consistency with regard to their sexual identity, attraction, and behavior within and across time. Hypotheses that levels of SSA would be elevated among *primarily lesbian/gay, bisexual*, and *primarily heterosexual* individuals, compared with their *exclusively heterosexual* peers, were supported and consistent with previous findings from Thompson and Morgan (2008). In both the calibration and validation samples, *exclusively lesbian/gay* individuals reported levels of SSA that were not statistically different from their *exclusively heterosexual* counterparts. Thus, individuals who report a monosexual sexual identity, attraction, or behavior may be less likely to report heightened SSA. Criterion validity was supported in the calibration sample and showed hypothesized relations between higher levels of SSA and a greater frequency of AUD symptoms in the past year as well as greater levels of distress stemming from recent depression and anxiety symptoms. Talley and Littlefield

(2014) hypothesize that intervening cognitive, affective, and behavioral processes relevant for persons with concealable stigmatized identities will influence relations between identity-related characteristics (e.g., SSA) and mental health outcomes (see also Meyer, 2003). For example, bisexual-identified individuals, who were found in the current study to report higher levels of SSA, may perceive more identity-related stress due to anticipated stigma, from both heterosexual or lesbian and gay persons, that characterizations of their previous sexual orientation experience were “just a phase” (e.g., Tabatabai & Linders, 2011).

Although the current study provides evidence to support the validity and reliability of the adapted SSA scale, findings should be interpreted with the following limitations in mind. First, the current study relied on convenience samples of women in young adulthood and students enrolled in an introductory psychology class who largely self-identified as *exclusively heterosexual* at the time of the surveys. Moreover, women in the calibration sample self-selected into the study, which purported to examine “female sexuality and alcohol use.” Although phone screenings excluded potential participants who reported receiving substance use treatment at any point in their lifetime, the recruitment wording may have resulted in a greater number of participants with higher proclivities toward or engagement in substance use behavior. Thus, convenience sampling may have contributed to bias in responding. Although preliminary evidence for criterion validity is presented from data available in the calibration sample, we were unable to examine parallel associations in the validation sample, which was composed of young adults in emerging adulthood who largely identified as *exclusively heterosexual*. Thus, future work is needed to further validate the current measure of SSA (e.g., among adults in middle and older adulthood or persons with diverse ethnic and racial identities) and inform the development of a nomological network pertinent to the SSA construct (Cronbach & Meehl, 1995).

Despite the aforementioned limitations, we believe that the 10-item SSA scale will provide researchers with a measure that will be useful for guiding future theoretical and empirical work that attempts to better understand *why* and *under what conditions* nonexclusively heterosexual persons may be a risk for various health-related disparities. Future work would benefit from attempts to identify personal or environmental factors that mediate associations between sexual self-concept ambiguity and mental health outcomes, including substance misuse and internalization symptoms. Alternatively, work might attempt to identify protective factors that buffer against negative health outcomes among individuals who report greater variation in or changes among facets of their sexual orientation.

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Figure 1.

Ten-item categorical confirmatory factor model for the calibration and validation samples.

Note. SSA = sexual orientation self-concept ambiguity. Values before the slash are culled from the calibration sample model, whereas values following the slash are culled from the validation sample model. Values inside the circles are residual variances for each item.

Table 1

SSA Item Distributions in the Calibration Sample ($n = 348$).

Item	Item wording	Strongly disagree (%)	Disagree (%)	Agree (%)	Strongly agree (%)	<i>M</i> (<i>SD</i>)
SSA1R ($n = 348$)	I seldom experience conflict between different aspects of my sexual orientation.	177 (50.9)	107 (30.7)	43 (12.4)	21 (6.0)	1.74 (0.90)
*SSA2 ($n = 348$)	On one day I might have one opinion of my sexual orientation and on another day I might have a different opinion.	210 (60.3)	60 (17.2)	60 (17.5)	17 (4.9)	1.67 (0.93)
SSA3 ($n = 347$)	I spend a lot of time wondering about what my sexual orientation really is.	233 (67.0)	62 (17.8)	42 (12.1)	10 (2.9)	1.51 (0.82)
*SSA4 ($n = 348$)	I feel as though my sexual orientation is different depending on whom I am with. ^a	227 (65.2)	54 (15.5)	51 (14.7)	16 (4.6)	1.59 (0.90)
*SSA5 ($n = 347$)	My views about my sexual orientation change rapidly or unpredictably. ^a	263 (75.6)	58 (16.7)	23 (6.6)	3 (0.9)	1.33 (0.64)
*SSA6 ($n = 347$)	Sometimes I feel that my sexual orientation is not what it appears to be.	216 (62.1)	68 (19.5)	51 (14.7)	12 (3.4)	1.59 (0.86)
*SSA7 ($n = 348$)	When I think about my sexuality in the past, I'm not sure what my sexual orientation was.	246 (70.7)	57 (16.4)	37 (10.6)	8 (2.3)	1.45 (0.77)
SSA8 ($n = 348$)	Sometimes I think I know other people's sexual orientation better than I know my own.	228 (65.5)	57 (16.4)	45 (12.9)	18 (5.2)	1.58 (0.90)
*SSA9 ($n = 348$)	My beliefs and actions regarding my sexual orientation often seem contradictory. ^a	227 (65.2)	71 (20.4)	40 (11.5)	10 (2.9)	1.52 (0.81)
*SSA10 ($n = 348$)	If I were asked to describe my sexual orientation, my description might end up being different from one day to another day.	248 (71.3)	59 (17.0)	35 (10.1)	6 (1.7)	1.42 (0.74)
*SSA11 ($n = 348$)	My beliefs about my sexual orientation often conflict with one another.	262 (75.3)	51 (14.7)	30 (8.6)	5 (1.4)	1.36 (0.70)
*SSA12 ($n = 348$)	Even if I wanted to, I don't think I could tell someone what my sexual orientation is really like.	251 (72.1)	53 (15.2)	33 (9.5)	11 (3.2)	1.44 (0.79)
SSA13 ($n = 345$)	My beliefs about my sexual orientation seem to change very frequently.	265 (76.1)	55 (15.8)	19 (5.5)	6 (1.7)	1.32 (0.66)
SSA14R ($n = 347$)	In general, I have a clear sense of who I am with regard to my sexual orientation.	190 (54.6)	95 (27.3)	28 (8.0)	34 (9.8)	1.73 (0.97)
*SSA15 ($n = 347$)	It is often hard for me to make up my mind about my sexual orientation because I don't really know.	255 (73.3)	49 (14.1)	36 (10.3)	7 (2.0)	1.41 (0.76)

Note. R indicates a reverse-scored item.

* Denotes items retained in the final, 10-item Sexual Orientation Self-Concept Ambiguity (SSA) scale. Three items were adapted from a list of clinically solicited indicators of general *identity diffusion* (Wilkinson-Ryan & Westen, 2000).

Table 2
SSA Total Score Means in the Calibration and Validation Samples, Split by Sexual Identity Status.

Sexual identity status	Calibration sample			Validation sample		
	<i>n</i>	<i>M</i>	<i>SD</i>	<i>n</i>	<i>M</i>	<i>SD</i>
Exclusively heterosexual	200	1.22	0.46 ^{ac}	948	1.21	0.01 ^a
Primarily heterosexual	71	1.89	0.71 ^b	53	1.64	0.05 ^b
Bisexual ^f	22	1.70	0.70 ^{bc}	10	1.97	0.12 ^b
Primarily lesbian/gay	10	2.06	0.63 ^{bc}	4	2.00	0.18
Exclusively lesbian/gay	26	1.39	0.63 ^{ac}	10	1.42	0.12

Note. SSA = sexual orientation self-concept ambiguity. Current sexual identity status was reported at the time of the survey.
^f Response option: “equally heterosexual and lesbian/gay (e.g., bisexual).” Means sharing the same superscript are not significantly different from each other (Games Howell, *p* < .05).

Table 3

Spearman Correlations Between SSA Total Scores and Convergent and Criterion Assessment Total Scores and Subscale Scores in the Calibration Sample.

Scale score	α	ρ	n^a
MoSIEC total score	.78	.21 *	333
MoSIEC commitment	.75	-.28 *	339
MoSIEC exploration	.90	.40 *	338
MoSIEC identity uncertainty	.70	.64 *	339
MoSIEC synthesis	.94	-.48 *	328
LGBIS total score	.95	.60 *	105
LGBIS need for privacy	.85	.41 *	91
LGBIS need for acceptance	.69	.37 *	118
LGBIS internalized homonegativity	.77	.46 *	102
LGBIS difficult process	.78	.37 *	110
LGBIS identity uncertainty	.81	.70 *	131
LGBIS identity superiority	.34	.06	144
Self-concept clarity	.92	-.39 *	320
Frequency of alcohol use disorder symptoms—Past 12 months	.82	.09 *	325
Anxiety symptoms—Past 2 weeks	.80	.22 *	323
Depressive symptoms—Past 2 weeks	.89	.25 *	322

Note. SSA = sexual orientation self-concept ambiguity; MoSIEC = Measure of Sexual Identity Exploration and Commitment (Worthington et al., 2008); LGBIS = Lesbian, Gay, and Bisexual Identity Scale (Mohr & Kendra, 2011); Self-Concept Clarity (Campbell et al., 1996).

^aSample sizes differ according to skip patterns established in the interview assessment (i.e., only women who indicated a nonexclusively heterosexual identity were asked to respond to items on the LGBIS).

* $p < .05$.