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Identity Status Measurement Across Contexts: Variations in Measurement Structure and Mean Levels Among White American, Hispanic American, and Swedish Emerging Adults

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We conducted this study to examine measurement equivalence and mean differences in identity status across 3 ethnic/cultural contexts: White American, Hispanic American, and Swedish. We used the Extended Objective Measure of Ego Identity Status II (EOM–EIS–II; Bennion & Adams, 1986), a commonly used instrument in the identity status literature. We conducted analyses to ascertain the extent to which the EOM–EIS–II functioned equivalently in 3 ethnically/culturally different samples. The internal structure of the measure was consistent across contexts. When we statistically controlled effects of age and gender, mean differences tended to be largely cross-cultural at the observed level of analysis but to be both cross-ethnic and cross-cultural at the latent level of analysis. This divergence in findings was found despite the limited age range represented in each of the samples. We therefore concluded that measurement error may have played a role in these differences and that data gathered using the EOM–EIS–II should be analyzed using latent variable methods. We discuss results in terms of using the EOM–EIS–II with diverse populations.

Identity has been the subject of widespread study for several decades. Erikson (1950) described identity as, among other things, a balance between self-knowledge and confusion about who one is. According to Erikson (1950), a coherent sense of identity lays the foundation for plotting a consistent and workable life course. A confused and haphazard sense of identity leaves one vulnerable to the pushes and pulls of social trends and fads such that one may employ a situational and disorganized approach to life (Côté & Levine, 2002; Erikson, 1968).

Although Erikson (1950, 1968) conceptualized identity as a developmental task of adolescence, recent social-structural changes in Western societies have resulted in a shift whereby many individuals address identity issues when they reach the age of majority (18 years old in most countries; Côté & Levine, 2002). Therefore, individuals in Western societies often address identity issues during emerging adulthood (Arnett, 2000). This life stage is characterized by partial, but not full, commitments to adult roles and responsibilities. In some contexts, the post-high-school emerging adult years may be an opportune time to study identity development and its measurement.

Erikson's (1950) conceptualization of identity was based on personal and clinical observations such that the theory is rich and elegant. However, Erikson's theory is somewhat lacking in precision and detail (Waterman, 1988). Marcia's (1966) attempt to adapt Erikson's conceptualization of identity for empirical research in the form of the identity status paradigm has received considerable theoretical and empirical attention.

THE IDENTITY STATUS PARADIGM

Since Marcia introduced it in 1966, the identity status paradigm has become one of the most influential frameworks for identity research (Archer, 1994). Marcia (1966) extracted from Erikson's (1950) theory the assumedly orthogonal dimensions of exploration and commitment. *Exploration* represents the process of sorting through potential life choices, whereas *commitment* represents selecting a set of choices to which one plans to adhere. Exploration and commitment pertain to values, beliefs, and goals in a number of life domains such as religion, occupation, and gender roles (Grotevant, 1993). Marcia (1966) crossed the two dimensions to create a 2 × 2 matrix where each quadrant within this matrix would correspond to an identity status. Each identity status represents a specific combination of a level (high or low) of exploration with a level (high or low) of commitment.

The statuses are achievement, moratorium, foreclosure, and diffusion. *Achievement* represents a commitment enacted following a previous period of exploration and is viewed as the successful consolidation of a sense of identity (Côté & Schwartz, 2002). *Moratorium* refers to ongoing exploration with only a partial or vague sense of commitment. It represents an ongoing, active struggle for a sense of identity. *Foreclosure* represents a commitment enacted without much prior exploration, and foreclosed individuals often adopt externally imposed roles and values without question or critical analysis. *Diffusion* represents a lack of concern about identity issues. Diffused individuals may or may not have engaged in prior identity exploration, but their exploration activity tends to be unfocused and haphazard. Diffusion is particularly distinguished by a strong absence of identity commitments (Schwartz, Côté, & Arnett, 2005).

Although the identity statuses are based on exploration and commitment, research has indicated that each status represents more than a simple permutation of exploration and commitment (Schwartz, 2002). It is important to note that statuses sharing a dimension in common (e.g., diffusion and moratorium are both characterized by a relative lack of commitment) may manifest this common dimension to differing extents. For example, the commitments held by foreclosed individuals are often more rigid and inflexible than are those held by achieved individuals (Schwartz, 2001), and lack of commitment is a much more prominent feature of diffusion than it is of moratorium (Schwartz et al., 2005).

To date, more than 500 published studies have drawn on the identity status paradigm. Each identity status has been empirically associated with a specific set of personality characteristics (see Berzonsky & Adams, 1999; Bourne, 1978; Kroger, 1993; Marcia, 1993; Waterman, 1999). For example, diffusion has been linked with apathy and disinterest, foreclosure with rigidity and authoritarianism, and moratorium with critical thinking and anxiety. Individuals with an achieved identity status generally score higher than the other three statuses on a number of positively valued qualities such as psychological well-being, cognitive complexity, academic motivation, and level of intimacy with friends. The identity statuses have been conceptualized within a number of content areas (e.g., political preferences, religious beliefs, dating styles). These content areas have been grouped into two broad categories: ideological (e.g., occupation, religion, politics) and interpersonal (e.g., friendships, dating, gender roles; for a review, see Schwartz, 2001).

Because the inclusion of some content areas (e.g., occupational choice and political preference within the ideological content-area category) within their assigned content-area categories may appear somewhat counterintuitive, it may be necessary to explain how these categorizations came about. In his original formulation of identity, Erikson (1950) focused on the content areas of ideology and occupation. When Erikson introduced identity status theory, Marcia (1966) redefined ideology as religion and politics. Several years later, Grotevant, Thorbecke, and Meyer (1982) extended the measurement of identity into interpersonal content areas (e.g., friendships, dating) that involved relational issues. The existing content areas were placed under the heading of ideological because they referred largely to intrapersonal issues. As additional content areas have been introduced, they have been added to either the ideological or interpersonal content-area category depending on whether or not they refer to relational issues.

Identity Status Measurement

A number of paper-and-pencil instruments have been developed to assess identity status (see Schwartz, 2001). Almost all of these instruments have been developed in North

America and have been normed on White samples. The Extended Objective Measure of Ego Identity Status II (EOM–EIS–II; Bennion & Adams, 1986) is the most commonly used paper-and-pencil identity status measure. The EOM–EIS–II measures identity status within four ideological content areas and four interpersonal content areas. The instrument contains items assessing endorsement of each of the four identity statuses. As a result, exploration and commitment are assessed together within each EOM–EIS–II item. An example of an item assessing achievement would be "After trying a lot of different recreational activities, I've found one or more I really enjoy doing by myself or with friends" (Adams, 1998, p. 84).

Evidence has been presented for the psychometric properties of the EOM–EIS–II with White Americans including internal consistency reliability (Jones & Streitmatter, 1987) and factorial, convergent, and discriminant validity (Bennion & Adams, 1986). The measure has been used to relate identity status to such constructs as separation anxiety (Bartle-Haring, Bruker, & Hock, 2002), substance use (Jones, Hartmann, Grochowski, & Glider, 1989), family environment (Adams, Ryan, & Keating, 2000), and academic autonomy (Berzonsky & Kuk, 2000).

IDENTITY STATUS ACROSS ETHNIC AND CULTURAL CONTEXTS

Researchers have long been interested in the cross-cultural applicability of Erikson's (1950) theory (e.g., Portes, Dunham, & del Castillo, 2000). However, a limitation of identity status measurement and of the identity status literature in general is the focus largely on White Americans (Sneed, Schwartz, & Cross, 2006). Although the United States has become increasingly ethnically diverse, identity status research has not attended to this diversity. Only a handful of studies (e.g., Schwartz et al., 2005) have used ethnically diverse samples and have examined cross-ethnic consistency in identity processes. Given the increase in ethnic minority populations, most notably Hispanics, in the United States (Day, 1996; Ramirez & de la Cruz, 2003), it is important to ascertain the appropriateness of identity status measures for ethnic minority populations. Moreover, although a growing number of identity status studies are being conducted outside of the United States, the extent to which identity status measures such as the EOM–EIS–II function equivalently (in terms of configural and metric invariance of factor structure) within and outside the United States has not been widely studied. Excluding research conducted in Canada, the majority of non-American identity status research has been conducted in Europe (e.g., Danielsen, Lorem, & Kroger, 2000; Goossens, 2001; Kumru & Thompson, 2003; Meeus, Iedema, Helsen, & Vollebergh, 1999). Comparatively less research has been conducted in other regions (e.g., Ohnishi, Ibrahim, & Owen, 2001, in Asia; Dwairy, 2004, in the Middle East).

For the identity status literature to be justified in branching out beyond White Americans and into American ethnic minority and non-American contexts, the extent to which identity measures function similarly within (across ethnic/subcultural variations) and outside of the United States (i.e., across cultures) needs to be ascertained. The question of similarity refers to two separate issues—measurement equivalence and mean differences. Measurement equivalence refers to configural and metric invariance—that is, whether the hypothesized measurement structure, which specifies which items are assigned to which subscales for a given instrument, functions equivalently across contexts (cf. Vandenberg & Lance, 2000). Measurement equivalence is a necessary prerequisite for assuming that the same phenomena are being measured in each ethnic or cultural context (van de Vijver & Leung, 2001). As it concerns the EOM–EIS–II, lack of measurement equivalence would indicate that the identity content areas participants are asked to rate should vary by culture or by ethnic subculture (cf. Marcia, 2001). Mean differences would not preclude the use of the EOM–EIS–II across ethnic or cultural groups, but they would suggest that the statuses may have different valences within each group.

Cross-Cultural Variation

The EOM–EIS–II is the only identity status measure that has been used to explore cross-cultural differences between North American and other contexts. A search of the PsycINFO psychological literature database from January 1985 through October 2004, covering the life span of the EOM–EIS–II, yielded four cross-cultural identity studies using the EOM–EIS–II. To test for measurement equivalence, Ohnishi et al. (2001) examined the measurement structure of the EOM–EIS–II ideological items in samples of adult women from the United States and Japan. The foreclosure and achievement items had the highest internal consistency estimates in both the U.S. and Japanese samples. Relative to the U.S. sample, items measuring diffusion and moratorium had higher internal consistency estimates in the Japanese sample. In the U.S. sample, factor-analytic procedures failed to differentiate diffusion items from moratorium items in most of the content areas surveyed. In a sample of Palestinian adolescents, Dwairy (2004) found that the structure of the EOM–EIS–II ideological identity status subscales also was different from the structure that had been normed on American adolescents.

Regarding the question of mean differences in identity status across cultural contexts, two Norwegian studies (Jensen, Kristiansen, Sandbekk, & Kroger, 1998; Stegarud, Solheim, Karlsen, & Kroger, 1999) have found that mean identity scores for all four identity statuses were lower in Norwegian samples relative to published means from White American samples. In contrast, Dwairy (2004) found that Palestinian adolescents scored significantly higher on foreclosure relative to published means from White American adolescent samples. Taken together, the extant literature suggests that both the measurement structure (e.g., Dwairy, 2004; Ohnishi et al., 2001) and mean levels (e.g., Dwairy, 2004; Jensen et al., 1998; Stegarud et al., 1999) of identity status may differ between American and non-American samples.

Although the aforementioned studies provide guidance regarding cross-cultural variation in measurement structure and mean levels of identity status, two primary areas remain unaddressed. First, whereas Dwairy (2004) and Ohnishi et al. (2001) examined cross-cultural differences in the structure of identity, they did so only in ideological identity content areas (e.g., politics, religion) and not in interpersonal identity content areas (e.g., friendships, dating). Second, whereas Dwairy (2004), Jensen et al. (1998), and Stegarud et al. (1999) reported differences in mean identity scores between American and non-American samples, the American means used in these studies were taken from the EOM–EIS–II reference manual. Although the raw data used to generate the American means presumably could have been obtained and analyzed, this step was not taken in these studies. Cross-sample equivalence tests and internal consistency analyses, both of which require raw data from both samples, were not conducted. As a result, it is not known whether the constructs assessed in the non-American samples were identical to, or somewhat different from, the constructs (i.e., identity status) that the EOM–EIS–II was designed to measure.

Cross-Ethnic Variations

A search of the PsycINFO psychological literature database between January 1985 and October 2004 revealed no studies in which the measurement structure or mean levels of identity status has been compared between White and ethnic minority American samples. As a result, little is known about how identity status measures may function in minority American samples. We designed this study, in part, to address this research question.

Issues in Interpreting Cross-Cultural and Cross-Ethnic Variation

Those studies that have found differences between American and non-American samples in terms of identity measurement structure and/or mean levels have ascribed these differences to cultural factors (i.e., differences between American culture and Norwegian, Japanese, or Palestinian culture). However, we note that all of these studies used translated versions of the EOM–EIS–II. It is possible that at least some of the variations found in these studies can be attributed to the measurement error associated with the creation and use of translated measures. For example, idiosyncratic expressions in one language may not have precise equivalents in other languages. The resulting differences in meaning may produce measurement error. Similar issues may occur when using the instrument with ethnic minority individuals in the United States. For example, because many Hispanic and Asian individuals in the United States may speak English as a second language, difficulty in reading or interpreting items may produce differences in response patterns. Cultural issues (e.g., cultural proscriptions against identity exploration in areas such as dating and gender roles) also may affect the ways in which ethnic minority individuals interpret and respond to the items. As a result, with respect to both cross-cultural and cross-ethnic use of the EOM–EIS–II and other identity measures, differences in measurement structure or mean levels of identity status may be attributable to language issues, cultural differences, or both. It is therefore important to examine such differences using latent variables (e.g., in structural equation modeling [SEM] analyses) that attempt to uncouple measurement error from the effects of more substantive linguistic or cultural issues. For example, differences in dating patterns, religious observations, political histories, and traditional gender roles in one country (or cultural group) versus another may increase or reduce the salience of some content areas. These differences may emerge as differences in measurement structure, as mean differences, or both.

THIS STUDY

We designed this study to empirically evaluate the appropriateness of the EOM–EIS–II for use in diverse populations. Our objective in this study was to examine variations in the measurement structure and mean levels of identity status scores in samples drawn from three different contexts. We used a White American sample because it represents the population on which most previous identity status research has been conducted (Sneed et al., 2006). We used a Hispanic American sample because (a) ethnic minority individuals will continue to comprise increasing proportions of the U.S. population (Day, 1996) and (b) Hispanics are the largest minority group in the United States (Ramirez & de la Cruz, 2003). We selected a European (Swedish) sample because the majority of non-North American identity research has been conducted in Europe (e.g., Danielsen et al., 2000; Kumru & Thompson, 2003; Meeus et al., 1999). No published research has examined identity status or its measurement in Sweden, and we wanted to help address this knowledge gap.

We guided this study by two specific aims. Our first aim was to examine the measurement equivalence (i.e., configural and metric invariance; see Vandenberg & Lance, 2000) of the EOM–EIS–II across these three culturally dissimilar samples. Provided that measurement equivalence could be demonstrated, our second aim was to examine variation in mean levels of endorsement of the four identity statuses across samples. We conducted these mean comparisons at the observed level using traditional analysis of variance methods used in past research (e.g., Dwairy, 2004; Jensen et al., 1998; Stegarud et al., 1999) and at the latent level using SEM methods. Differences in findings between the two approaches may provide an estimate of the extent to which measurement error contributes to the differences observed in the analysis of variance results.

METHOD

Participants and Procedures

White American Sample

The White American sample consisted of 223 students (36 men, 187 women) attending a public university in central Florida with a largely White student population. The mean age of the sample was 21.2 years (SD = 3.86 years), with 89% of participants between the ages of 17 and 23. Of those participants who provided socioeconomic data (78% of White American participants), 19% reported annual household incomes below $30,000; 17% reported annual household incomes between $30,000 and $50,000; 34% reported annual household incomes between $50,000 and $100,000; and 25% reported annual household incomes above $100,000. Participants completed the EOM–EIS–II on the Internet and received course credit for their participation.

Hispanic American Sample

The Hispanic American sample was recruited from an university in Miami, a city in which Hispanic individuals comprise the majority of the population. The Hispanic American sample consisted of 362 students (83 men, 279 women). Of these participants, 62% were born in the United States, and all were raised by at least one immigrant parent. The most common countries of origin for these participants and/or their parents were Cuba (47%), Colombia (10%), Nicaragua (6%), and Puerto Rico (3%). The mean age of the sample was 20.4 years (SD = 2.36 years), with 90% of participants between the ages of 18 and 23. Of those participants who provided socioeconomic data (75% of Hispanic American participants), 22% reported annual household incomes below $30,000; 33% reported annual household incomes between $30,000 and $50,000; 30% reported annual household incomes between $50,000 and $100,000; and 15% reported annual household incomes above $100,000. Participants completed the EOM–EIS–II at home over the weekend and returned it to their instructor the following week. All participants completed an English version of the measure. Participants received course credit for their participation.

Swedish Sample

The Swedish sample consisted of 517 students of European descent (284 men, 233 women) from five gymnasiums in a mid-size Swedish city. The mean age of the sample was 17.2 years (SD = 0.67 years), with 99% of participants between the ages of 17 and 21. Although socioeconomic data were not collected on the Swedish sample, in Sweden socioeconomic status can be estimated roughly according to whether students are on track toward university (higher socioeconomic status) or vocational education (lower to middle socioeconomic status). In the Swedish sample, 73% of participants were on track for university, whereas 27% were on track for vocational education. Participants completed the EOM–EIS–II in class and returned it to their instructor.

The Swedish National Agency for Education (2004) described the gymnasium as being similar to American community colleges and universities; it is described as

> The Swedish gymnasium is a free and non-compulsory form of school which is similiar to American community colleges and universities. According to The Swedish National Agency for Education (2005), "every municipality in Sweden is required by law to offer all students who have completed compulsory school an upper secondary education … . The right to begin an upper secondary school program applies up to and including the calendar year in which the student turns 20." (p. 65)

As is the case in American community colleges, students live at home with their parents but are viewed and treated as young adults. Students choose their own program and specialization, and their progress is monitored through grades at the end of each course. The Swedish sample was significantly younger than either the White American sample, $t(737) = 23.09$, $p < .001$, $d = 1.70$; or the Hispanic American sample, $t(877) = 29.75$, $p < .001$, $d = 2.01$. In addition, the Swedish sample was characterized by a significantly more balanced gender distribution than was either the White American sample, $\chi^2(1, N = 740) = 95.51$, $p < .001$, $\phi = .36$, or the Hispanic American sample, $\chi^2(1, N = 879) = 89.67$, $p < .001$, $\phi = .32$.

Organization of the Samples for Data Analysis

Data collection for the White American and Swedish samples included some ethnic minority participants, and data collection for the Hispanic American sample included some White Americans as well as members of multiple ethnic minority groups (81% of these minority participants were of Hispanic origin). As a strategy to minimize the confounding of ethnicity with culture and nationality, only non-Hispanic White participants were retained in the White American sample. Only Hispanic participants were retained in the Hispanic American sample. Only White participants (i.e., those of European descent) were retained in the Swedish sample. Inclusion of only one ethnic group within each sample minimizes the extent to which ethnic heterogeneity could account for observed similarities and differences among the three samples. In total, 107 participants (26 non-Hispanic Black, 44 Hispanic, 13 Asian, and 24 Other) were dropped from the White American sample; 193 participants (98 non-Hispanic White, 60 non-Hispanic Black, 22 Asian, and 13 Other) were dropped from the Hispanic American sample; and 69 participants (1 non-Hispanic Black, 9 Hispanic, 8 Asian, and 51 Other) were dropped from the Swedish sample. The sample sizes reported in the Participants and Procedures section refer to the sample compositions after potentially overlapping individuals were dropped.

Measure

The EOM–EIS–II (Bennion & Adams, 1986) assesses identity status in four ideological content areas (politics, religion, occupation, and philosophical lifestyle) and in four interpersonal content areas (friendships, dating, gender roles, and recreation/leisure). The EOM–EIS–II contains 64 items, 16 measuring each of the 4 statuses. Of the 16 items targeting each identity status, 2 assess each content area. The EOM–EIS–II generates continuous scores for each status both in overall terms and within the ideological and interpersonal content-area clusters. The measure uses a 6-point Likert-type response scale ranging from 1 (*strongly disagree*) to 6 (*strongly agree*). In past research, internal consistency reliability estimates for scores on the EOM–EIS–II scales have ranged from .60 to .80 for White American samples (Bennion & Adams, 1986) and from .49 to .74 for ethnic minority American samples (Schwartz, 2002). To our knowledge, our study is the first to use the EOM–EIS–II with a Swedish sample.

Creation of the Swedish Translation of the EOM–EIS–II

To create the Swedish version of the EOM–EIS–II, a professional translator translated the original English version into Swedish. A second bilingual individual then translated the Swedish version back into English. The two English versions were then compared, with the two translators discussing and resolving discrepancies between the original and back-translated English versions. The Swedish translation was then pilot tested for clarity on a small group (n = 10) of Swedish gymnasium students. Some items were slightly modified based on feedback from these students. Data from these pilot participants were not used in the analyses for this report.

RESULTS

Data Analytic Strategy

The data analyses for this study proceeded in three general steps: (a) confirmatory factor analyses evaluating the equivalence of measurement structure across samples, (b) comparisons of mean identity status scores across samples, and (c) computation of internal consistency reliability (alpha) coefficients. We conducted both the alpha coefficient analyses and mean comparisons at both the observed and latent levels of analysis. We conducted both the confirmatory factor analyses and latent mean comparisons using SEM techniques (cf. Byrne, 2001; Hancock, 1997). These SEM analyses were conducted using AMOS Release 5.0 (Arbuckle & Wothke, 2004; Byrne, 2001).

Confirmatory Factor Analyses

We conducted confirmatory factor analyses to test the measurement equivalence of the EOM–EIS–II across samples. We entered The 64 EOM–EIS–II items into four, 3-group confirmatory factor analysis models (with culture as the grouping variable), one model for each identity status. We estimated the models separately by identity status because including all four statuses in a single model would have resulted in a sample size to parameter estimation ratio of considerably less than 10:1 (less than 2:1 in the White American sample), potentially causing model instability or misspecification (cf. Kline, 1998). Moreover, one of our goals in this study was to examine the adequacy of the EOM–EIS–II hypothesized measurement structure, which specifies four separate sets of items (one set of 16 items for each identity status), across ethnic/cultural groups. In each confirmatory factor analysis, we created separate latent variables for the ideological and interpersonal content-area clusters. We drew a bidirectional path between the ideological and interpersonal latent variables because the ideological and interpersonal subscales for each status are often correlated with one another (Bennion & Adams, 1986; Schwartz & Montgomery, 2002).

In tests of measurement equivalence using confirmatory factor analysis methodology, we estimated the hypothesized measurement structure (i.e., determining which items load on which subscales) on each of the cultural groups examined (Raju, Laffitte, & Byrne, 2002). This is accomplished by way of a five-step process suggested by Byrne (2001). We added additional steps in this study to (a) examine the contributions of age and gender to measurement equivalence and (b) identify items that do not adequately represent the subscales to which they are assigned in the hypothesized measurement structure. First, we applied the hypothesized measurement structure separately to each of the groups under study. This step facilitates examination of the extent to which the algorithm fits adequately within each sample. As part of this step, we consulted modification indexes to identify error terms that should be correlated to improve model fit. Only those error terms for which a correlation can be theoretically justified are then correlated. In this case, for each status, we correlated only error terms within the ideological content areas and within the interpersonal content areas. To ensure that the analyses were congruent with the theoretical separation of identity content areas into ideological and interpersonal clusters, we did not correlate error terms across content-area clusters (i.e., an error term for an ideological item with an error term for an interpersonal item). Second, we estimated a multigroup model in which the pattern coefficients linking the observed questionnaire item scores to their latent subscale factors were free to vary across samples (i.e., the unconstrained model). This unconstrained model represents the aggregation of the three models examined in the first step. Third, we estimated a model in which the pattern coefficients linking the observed questionnaire item scores to their latent subscale factors were constrained equal across the three samples (i.e., the constrained model). Fourth, we compared the fit indexes for the unconstrained and constrained models. If the unconstrained model fit the data adequately and if the difference in fit indexes between the unconstrained and constrained models was judged to be trivial, then measurement equivalence was assumed (Little, 1997; Muthén & Muthén, 2004).

Next, for each identity status, we identified items that loaded on their respective subscales at .30 or below in at least two of the three samples. Loadings below .30 in confirmatory factor analyses may indicate that the item does not represent the construct adequately (Comrey & Lee, 1992). Therefore, identifying such items may be a necessary prerequisite to future revisions of the EOM–EIS–II.

Last, we entered age and gender into the model as covariates to ascertain whether age and gender differences among the samples would affect the extent to which measurement equivalence could be assumed. We estimated paths between each covariate and each of the indicator variables in the model. A nontrivial difference in model fit with versus without covariates indicates that the covariates affected the extent of measurement equivalence observed. A finding of measurement equivalence, coupled with the absence of age and gender differences in the extent of equivalence observed, would suggest that the EOM–EIS–II operated equivalently across ethnic/cultural contexts, genders, and age groups.

In evaluating the fit of the unconstrained model to the data, we used the comparative fit index (CFI), which compares the specified model to a null model with no paths or latent variables, and the root mean square error of approximation (RMSEA), which represents the extent to which the covariance structure specified in the model deviates from the covariance structure observed in the data. Good model fit is represented by CFI values of .95 or greater and by RMSEA values of .05 or less (Bentler & Bonnett, 1980; Byrne, 2001), with .90 representing the lower bound for an acceptable CFI value and .08 representing the upper bound for an acceptable RMSEA value (Quintana & Maxwell, 1999). We report the chi-square statistic, but it was not used to evaluate model fit because it often indicates significant differences between the model and observed covariances even when such deviations are quite small (Kline, 1998). In the comparisons between the unconstrained and constrained models, we followed the recommendations of Vandenberg and Lance (2000) who suggested reporting the difference in chi-square values as well as the difference in at least one adjunctive fit index. Following Little's (1997) recommendations for measurement equivalence analyses in cross-cultural measurement research, as an adjunctive fit index, we reported the difference in the nonnormed fit index (NNFI). We evaluated measurement equivalence across samples using the method recommended by Little (1997), in which the NNFI estimates from the two models are compared. A difference of .05 or less is considered trivial. Little's method, rather than the chi-square difference test, was used in the formal invariance tests because the chi-square difference test often indicates significant differences in model fit when the fit indexes themselves are only trivially different (Little, 1997).

Mean Identity Status Scores Across Ethnic/Cultural Groups

The second step of analysis was to compare the mean identity status scores across the three samples. The vast majority of identity status research has used analyses conducted on observed scores. In this study, we used latent variable methods to ascertain the extent to which these comparisons might better be conducted at the latent level. We report observed-level multivariate analyses of variance (MANOVAs) in the Discussion for comparison purposes.

We used SEM to compare the mean scores across samples. This procedure requires measurement equivalence across samples and uses the factor structure (including the same correlated error terms) emerging from the invariance analyses. In mean comparisons using SEM methodology, the latent subscale means are compared, and measurement error is factored out of these latent means when they are compared (Hancock, 1997; Hancock, Lawrence, & Nevitt, 2000). We conducted the mean comparisons in a manner similar to that proposed by Byrne (2001). First, we constrained the intercepts for all items equal across samples, one item at a time, and we freed items for which the assumption of scalar invariance was statistically rejected (using the difference in NNFI values between constrained and unconstrained models; see Little, 1997; Vandenberg & Lance, 2000). Second, because three groups were being compared, we used an omnibus test to test for the presence of differences among the groups (and to control for Type I error). Similar to the procedure used to establish measurement equivalence, a model with the latent means constrained equal was compared to a model in which the means were free to vary (Little, 1997; Vandenberg & Lance, 2000). Because comparisons of latent means across groups are error free and therefore precise, Little recommended that only the chi-square statistic be used to compare the constrained and unconstrained mean-difference models. A significant chi-square difference is akin to a significant omnibus effect in an analysis of variance, indicating that pairwise comparisons should be conducted among the latent group means. In these pairwise comparisons, the mean for a reference group is constrained to zero, and the ratio of the other latent mean to the square root of the pooled (i.e., averaged between the groups being compared and weighted according to the respective sample sizes) variance estimate indicates the number of pooled standard deviations by which the nonreference group mean differs from the reference group mean (Hancock, 2004). That is, in each comparison, the ratio of the nonreference group's latent mean to the pooled standard deviation represents the effect size (i.e., Cohen's *d*) for the mean comparison between that group and the reference group. We evaluated statistical significance for each of these mean comparisons as a *z* test, which was evaluated against the *t* distribution with infinite degrees of freedom (Hancock, 1997). For each identity status for which a significant omnibus mean difference emerged, we first compared the Hispanic American and Swedish latent means to the White American latent mean, and we then compared the Swedish latent mean to the Hispanic American latent mean. In keeping with recent recommendations regarding reporting of results, we report effect sizes as well as statistical significance. In the pairwise latent mean comparisons, effect size is reported as Cohen's *d* because *d* represents a focused, directional, and precise effect size index (McCartney & Rosenthal, 2000).

For each identity status, we then investigated the contribution of age and gender to the mean differences obtained. As was done in the measurement invariance analyses, we estimated paths between each covariate and each of the indicator variables in the model. Meaningful contributions of age and gender to mean differences would affirm the importance of these variables in identity research (Lewis, 2003; Schwartz & Montgomery, 2002).

Internal Consistency Reliability Analyses

The third step of analysis was to conduct internal consistency analyses on each of the EOM–EIS–II subscales (i.e., for each status, both overall and within the ideological and interpersonal content areas) within each sample. We computed internal consistency analyses at both the observed (Cronbach, 2004) and latent (Fornell & Lacker, 1981) levels. The latent internal consistency reliability formula posits reliability as the ratio of variance explained by the construct to the total variance among the indicators, and it controls for measurement error within the indicators. Because the latent variable represents the variability shared among the indicators, unreliability (i.e., alpha coefficients below 1.0) is due to nonperfect correlations among the indicators. Put another way, to the extent to which each indicator possesses unique variability that is not shared with the others, the alpha coefficient will be decreased to a corresponding degree.

Organization of the Presentation of Results

Because we conducted both the measurement equivalence analyses and the structural mean comparisons separately by identity status, we present these two sets of analyses together for each status. Such a format is consistent with Byrne (2001), Little (1997), and Vandenberg and Lance (2000), all of whom conceptualized measurement equivalence tests and structural mean comparisons under the rubric of multigroup comparisons. For ease of presentation and because latent mean comparisons are inherently more powerful and accurate than comparisons of observed means (Hancock et al., 2000), we focus the reports of mean comparisons on latent-level analyses. We briefly present results of observed level analyses in the Discussion section followed by comparisons of latent and observed level results. We present internal reliability analyses under a single heading in the Results section because they are reported across statuses.

Measurement Equivalence and Mean Differences Across Samples

Diffusion

Measurement equivalence. Following Byrne (2001), we first evaluated the fit of the EOM–EIS–II Diffusion subscale to the data separately within each of the three samples. These models produced an adequate fit to the data: χ^2 values (91 *df* each) ranged from 131.73 to 188.00 (all ps < .001), CFI values ranged from .92 to .94, and RMSEA values ranged from .04 to .05. When we compared the constrained and unconstrained models, this indicated that the model fit equivalently across the three samples, $\Delta\chi^2(32) = 70.37$, $p < .001$; ΔNNFI = .010. Adding age and gender as covariates did not significantly alter the model fit, $\Delta\chi^2(3) = 0.14$, *ns;* ΔNNFI = .015. As a result, we assumed measurement equivalence for the EOM–EIS–II Diffusion items, and this equivalence did not appear to be affected by age and gender differences across samples.

Inspection of the confirmatory factor pattern coefficients revealed that four Ideological Diffusion items and two Interpersonal Diffusion items loaded below .30 on their respective subscales in at least two of the three samples. These included both items in the occupational domain, one item in the political domain, one item in the lifestyle domain, one item in the dating domain, and one item in the leisure domain.

Mean differences. When we constrained the item intercepts equal across samples, the assumption of scalar invariance was rejected for two items—one in the religion content area and another in the gender roles content area. The structural mean comparison revealed mean differences in latent Diffusion scores among the three samples, $\Delta\chi^2(4) = 276.34$, $p < .001$. Significant differences in Ideological Diffusion emerged between the Swedish sample and the two American samples, with the Swedish sample having the highest mean. The latent means for the two American samples did not differ significantly from one another. For Interpersonal Diffusion, the Swedish sample scored significantly lower than the two American samples, which were not significantly different from one another (see Tables 1 and 2).

We then added age and gender to the mean difference comparison model as covariates. For Ideological Diffusion, the three samples were all significantly different from one another, with the Hispanic American sample scoring highest and the Swedish sample scoring lowest. On Interpersonal Diffusion, the three samples were again all significantly dif-

TABLE 1
Sample Means and Standard Deviations With and Without Covariates

	Without Covariates						*With Covariates*					
	White American[a]		*Hispanic American*		*Swedish*		*White American*[a]		*Hispanic American*		*Swedish*	
Subscale	*M*	*SD*	*M*	*SD*	*M*	*SD*	*M*	*SD*	*M*	*SD*	*M*	*SD*
Ideological Diffusion												
Observed	24.20	6.45	25.04	5.43	29.04	7.30	24.69	7.05	25.39	6.87	28.50	7.61
Latent	0.00	0.25	–0.01	0.18	0.31	0.28	0.00	0.67	0.45	0.45	–0.29	0.80
Interpersonal Diffusion												
Observed	21.51	5.65	21.58	5.71	20.67	6.22	22.38	6.22	22.20	6.06	19.56	6.72
Latent	0.00	0.51	0.03	0.50	–0.29	0.58	0.00	0.39	0.17	0.39	–0.33	0.39
Ideological Foreclosure												
Observed	19.33	6.51	20.09	6.68	18.45	6.72	20.47	7.14	20.82	6.96	17.31	7.72
Latent	0.00	0.41	0.03	0.43	–0.12	0.41	0.00	0.75	0.27	0.79	0.49	0.76
Interpersonal Foreclosure												
Observed	17.31	6.06	17.85	6.58	16.65	6.32	18.29	6.77	18.52	6.59	15.66	7.30
Latent	0.00	0.65	0.11	0.72	–0.14	0.68	0.00	0.67	–0.18	0.76	0.80	0.69
Ideological Moratorium												
Observed	24.07	6.13	24.74	6.63	26.22	6.55	25.12	7.00	25.09	6.81	25.62	7.55
Latent	0.00	0.44	0.07	0.48	0.00	0.47	0.00	0.09	–0.19	0.09	0.26	0.11
Interpersonal Moratorium												
Observed	26.40	5.46	25.88	5.79	21.78	6.16	27.22	6.39	26.38	6.22	21.11	6.90
Latent	0.00	0.74	–0.01	0.74	–0.16	0.71	0.00	0.74	1.42	0.72	0.53	0.73
Ideological Achievement												
Observed	32.10	5.46	33.63	5.91	27.51	6.96	31.79	6.82	33.38	6.65	28.20	7.36
Latent	0.00	0.17	0.02	0.16	–0.18	0.18	0.00	0.19	–0.04	0.18	–1.19	0.21
Interpersonal Achievement												
Observed	33.70	6.11	34.91	6.03	28.60	6.51	33.47	6.58	34.38	6.52	29.53	7.22
Latent	0.00	0.21	0.07	0.20	–0.24	0.21	0.00	0.44	–0.25	0.40	–2.34	0.41

[a]Latent mean constrained to zero (cf. Hancock, 2004).

ferent from one another, with the Hispanic American sample scoring highest and the Swedish sample scoring lowest (see Tables 1 and 2).

Foreclosure

Measurement equivalence. When we evaluated the fit of the Foreclosure subscale to the data within each sample, the models fit the data adequately. Chi-square values (83 *df* each) ranged from 180.04 to 250.27 (all *p*s < .001), CFI values ranged from .93 to .94, and RMSEA values ranged from .06 to .07. When we compared the constrained and unconstrained models, results indicated that the Foreclosure items fit equivalently across the three samples, $\Delta\chi^2(32) = 84.32$, $p < .001$; ΔNNFI = .001. Adding age and gender did not significantly change the model fit, $\Delta\chi^2(3) = 3.74$, *ns;* ΔNNFI = .012. As a result, we assumed measurement equivalence for these items irrespective of age and gender differences across samples. No foreclosure items loaded on their respective subscales at .30 or below.

Mean differences. When we constrained the item intercepts equal across samples, the assumption of scalar invariance was statistically rejected for three items—one in the lifestyle content area, one in the friendships content area, and one in the gender roles content area. The structural mean comparison revealed mean differences in latent Foreclosure scores among the three samples, $\Delta\chi^2(4) = 24.49$, $p < .001$. Significant pairwise differences in Ideological Foreclosure emerged between the Swedish sample and the two American samples, with the Swedish sample scoring lowest. The two American samples were not significantly different from one another. For Interpersonal Foreclosure, significant pairwise differences emerged between the Swedish sample and the two American samples, with the Swedish sample scoring lowest. Again, the two American samples were not significantly different from one another (see Tables 1 and 2).

When we controlled for age and gender, Ideological Foreclosure scores differed significantly among all three samples, with the Swedish sample scoring highest and the White American sample scoring lowest. Interpersonal Foreclosure scores also differed significantly among all three samples, with the Swedish sample scoring highest and the White American sample scoring lowest (see Tables 1 and 2).

Moratorium

Measurement equivalence. Evaluating the fit of the Moratorium subscale to the data yielded an adequate fit: χ^2 values (81 *df* each) ranged from 137.74 to 150.95 (all *p*s < .001), CFI values ranged from .92 to .94, and RMSEA values ranged from .04 to .06. Comparing the constrained and unconstrained models indicated that the EOM–EIS–II Moratorium subscales fit equivalently across the three samples, $\Delta\chi^2$

TABLE 2
Pairwise Mean Differences in Identity Status Scores by Subscale

	White American Vs. Hispanic American				*White American Vs. Swedish*				*Hispanic American Vs. Swedish*			
	Without Covariates		*With Covariates*		*Without Covariates*		*With Covariates*		*Without Covariates*		*With Covariates*	
Subscale	*M*	*d*	*M*	*d*	*M*	*d*	*M*	*d*	*M*	*d*	*M*	*d*
Ideological Diffusion												
Observed	1.69	0.14	1.21	0.10	9.37***	0.69	6.93***	0.51	9.03***	0.61	6.37***	0.43
Latent	0.60	0.05	10.02***	0.83	15.48***	1.14	5.16***	0.38	19.40***	1.31	16.14***	1.09
Interpersonal Diffusion												
Observed	0.12	0.01	0.36	0.03	1.90	0.14	5.84***	0.43	2.21*	0.15	6.02***	0.41
Latent	0.72	0.06	5.31***	0.44	7.06***	0.52	11.54***	0.85	8.59***	0.58	18.95***	1.28
Ideological Foreclosure												
Observed	1.33	0.11	0.58	0.60	1.77	0.13	5.88***	0.43	3.70***	0.25	7.11***	0.48
Latent	0.85	0.07	4.22***	0.35	3.93***	0.29	8.83***	0.65	5.33***	0.36	4.29***	0.29
Interpersonal Foreclosure												
Observed	0.99	0.08	0.36	0.03	1.49	0.11	5.03***	0.37	2.81**	0.19	6.07***	0.41
Latent	1.86	0.16	2.90**	0.25	2.85**	0.21	15.89***	1.17	5.28***	0.36	20.14***	1.36
Ideological Moratorium												
Observed	1.22	0.10	0.05	0.00	4.48***	0.33	0.95	0.07	3.28**	0.22	1.01	0.07
Latent	1.77	0.15	23.41***	2.11	0.08	0.00	33.82***	2.49	2.22*	0.15	65.30***	4.41
Interpersonal Moratorium												
Observed	1.09	0.09	1.57	0.13	9.68***	0.78	11.30***	0.91	10.07***	0.68	11.85***	0.80
Latent	0.16	0.01	23.54***	1.95	2.78**	0.22	9.02***	0.72	3.11**	0.21	18.21***	1.23
Ideological Achievement												
Observed	3.26**	0.27	2.78**	0.24	8.75***	0.70	6.22***	0.50	13.77***	0.93	10.80***	0.73
Latent	1.43	0.12	2.56*	0.22	12.69***	1.02	72.74***	5.84	16.96***	1.16	84.66***	5.81
Interpersonal Achievement												
Observed	2.35*	0.20	1.63	0.14	9.96***	0.80	6.99***	0.56	14.81***	1.00	10.36***	0.70
Latent	3.00**	0.34	7.20***	0.60	14.17***	1.14	69.67***	5.59	22.36***	1.51	76.26***	5.15

Note. Numbers presented for latent mean comparisons are z values. Numbers presented for observed mean comparisons are t values. Degrees of freedom for t values are 583 for the comparison between the White American and Hispanic American samples, 738 for the comparison between the White American and Swedish samples, and 877 for the comparison between the Hispanic American and Swedish samples. Cohen's d = effect sizes. t and z values are a function of both effect size and the sample size for each group (Kline, 2004). As the sample size increases, the z value associated with a given effect size increases.
$^{*}p < .05$. $^{**}p < .01$. $^{***}p < .001$.

(32) = 80.28, $p < .001$; ΔNNFI = .012. Adding age and gender to the model did not significantly change the model fit, $\Delta\chi^2(3) = 5.49$, *ns;* ΔNNFI = .022. As a result, we assumed measurement equivalence for the EOM–EIS–II Moratorium items, and this equivalence did not appear to be affected by age and gender differences across samples.

Inspection of pattern coefficients between the items and their respective latent constructs revealed that one Ideological Moratorium item and three Interpersonal Moratorium items loaded below .30 in at least two of the three samples. These included one lifestyle item, one gender role item, and both recreation items.

Mean differences. When we constrained the item intercepts equal across samples, the assumption of scalar invariance was rejected for six items—both items in the political content area and one item apiece in the occupational, gender roles, friendship, and recreation content areas. The structural mean comparison did not reveal significant mean differences in Moratorium scores among the three samples, $\Delta\chi^2(4) = 6.67$, *ns.*

When we added age and gender to the mean difference model as covariates, the structural mean comparison revealed significant mean differences in Moratorium scores among the three samples, $\Delta\chi^2(4) = 83.95$, $p < .001$. Ideological Moratorium scores differed significantly among all three pairs of samples, with the Swedish sample scoring highest and the Hispanic American sample scoring lowest. On Interpersonal Moratorium, latent mean scores again differed significantly among all three pairs of samples, with the Hispanic American sample scoring highest and the White American sample scoring lowest (see Tables 1 and 2).

Achievement

Measurement equivalence. When we examined the fit of the EOM–EIS–II Achievement subscales to the data, the model provided a good fit: χ^2 values (83 *df* each) ranged from 118.57 to 147.76 (all *p*s < .001), CFI values ranged from .94 to .96, and all three RMSEA values were .04. When we compared the constrained and unconstrained models, results indicated that the EOM–EIS–II Achievement subscales fit equivalently across the three samples, $\Delta\chi^2(32) = 47.85$, $p < .025$; ΔNNFI = .001. Adding age and gender to the model did not significantly change the fit, $\Delta\chi^2(3) = 0.96$, *ns;* ΔNNFI = .014. As a result, we assumed measurement equivalence for the EOM–EIS–II Achievement items, and this equivalence did not appear to be qualified by age and gender differences between samples.

Inspecting the confirmatory factor pattern coefficients indicated that three Ideological Achievement items and one Interpersonal Achievement item loaded below .30 in at least two of the three samples. Two of these items were in the political content area, one was in the lifestyle content area, and one was in the recreational content area.

Mean differences. When we constrained the item intercepts equal across samples, the assumption of scalar invariance was rejected for three items—both items in the religion content area and one item in the gender roles content area. The structural mean comparison revealed mean differences in latent Achievement scores among the three samples, $\Delta\chi^2(4) = 266.43$, $p < .001$. Pairwise comparisons indicated that Ideological Achievement scores differed significantly between the Swedish sample and the two American samples, with the Swedish sample scoring lowest. The two American samples were not significantly different from one another. For Interpersonal Achievement, pairwise comparisons indicated that latent Ideological Achievement scores differed significantly between all three pairs of samples, with the Hispanic American sample highest and the Swedish sample lowest (see Tables 1 and 2).

When we added age and gender to the mean difference model as covariates, pairwise comparisons indicated that Ideological Achievement scores differed significantly between all three pairs of samples, with the White American sample highest and the Swedish sample lowest. For Interpersonal Achievement, pairwise comparisons indicated that latent Interpersonal Achievement scores differed significantly between all three pairs of samples, again with the White American sample highest and the Swedish sample lowest (see Tables 1 and 2).

Internal Consistency Reliability Analyses

We examined internal consistency in two ways. First, we computed Cronbach's alpha coefficients for the each of EOM–EIS–II Identity Status scales within each sample. Second, we computed reliability (alpha) coefficients for the latent identity status constructs (Fornell & Lacker, 1981). For purposes of consistency and simplicity, we refer to these coefficients as "latent alpha coefficients." Cronbach's alpha values and latent alpha coefficients were below the acceptable range (i.e., $\alpha < .70$) for Diffusion, Moratorium, and Achievement (see Table 3). However, when items loading on their respective subscales at .30 or below in at least two of the three samples were removed from consideration, only 14% (5 of 36) of the latent alpha coefficients were below .70. Moreover, whereas the mean alpha coefficient was equivalent between the observed (.68) and latent (.66) levels with no items removed, when problematic items were removed, the mean alpha coefficient was markedly higher at the latent level (.80) than at the observed level (.69). Four of the six latent alpha coefficients that were below .70 after problematic items were removed were for indexes of Diffusion, and in two of these cases, the latent alpha coefficient decreased after low-loading items were removed. Cronbach's alpha coefficients for observed scores remained below .70 for Diffusion in all three samples and for Moratorium and Achievement in the Hispanic American and

TABLE 3
Internal Consistency Estimates for American and Swedish Samples

	White American		*Hispanic American*		*Swedish*	
Status Measure[a]	*Observed*	*Latent*	*Observed*	*Latent*	*Observed*	*Latent*
Diffusion						
Ideological (4 items removed)	.66, .56	.59, .72	.45, .40	.57, .63	.63, .62	.75, .75
Interpersonal (2 items removed)	.63, .55	.55, .69	.58, .56	.73, .68	.56, .53	.76, .67
Overall (6 items removed)	.72, .62	.73, .85	.65, .61	.60, .86	.64, .59	.84, .83
Foreclosure[b]						
Ideological (0 items removed)	.69	.78	.68	.74	.60	.71
Interpersonal (0 items removed)	.82	.81	.81	.79	.75	.76
Overall (0 items removed)	.86	.89	.84	.87	.80	.85
Moratorium						
Ideological (1 item removed)	.69, .71	.66, .86	.70, .72	.67, .87	.60, .55	.55, .75
Interpersonal (3 items removed)	.61, .61	.61, .75	.61, .61	.54, .75	.58, .54	.50, .66
Overall (4 items removed)	.76, .76	.78, .93	.77, .79	.76, .94	.70, .66	.68, .90
Achievement						
Ideological (3 items removed)	.64, .58	.59, .69	.61, .64	.53, .74	.60, .62	.48, .71
Interpersonal (1 item removed)	.74, .73	.74, .87	.66, .66	.63, .83	.54, .58	.53, .78
Overall (4 items removed)	.79, .77	.81, .92	.75, .76	.74, .93	.69, .70	.67, .87

Note. For both the observed and latent level analyses, the first number represents the alpha coefficient before problematic items were removed, and the second number represents the alpha coefficient after problematic items were removed.
[a]The parenthetical note in this column indicates the number of problematic items (those loading below .30 on their assigned subscales) that were removed to increase internal consistency. [b]No Foreclosure factor pattern coefficients were below .30; therefore, no Foreclosure items were removed from analysis.

Swedish samples. As we noted in the confirmatory factor analysis results, across statuses, problematic ideological items tended to be in the political and lifestyle content areas (three items apiece), and problematic interpersonal items tended to be in the recreational content area (four items). The alpha reliability coefficients, at both the observed and latent levels, suggest that the EOM–EIS–II Diffusion items may be most in need of revision.

As we noted previously, some of the reliability coefficients decreased when potentially problematic items were removed from consideration. As a result, as an additional check to further examine the extent to which these items were indeed problematic, for each potentially problematic item, we examined the item-total correlations (i.e., the correlation between item responses and the score for the subscale to which the item was assigned). Similar to our criterion for identification of potentially problematic items in the confirmatory factor analyses, we considered items problematic if their item-total correlations were below .30 in at least two of the three samples.

For Ideological Diffusion, three of the four items with factor pattern coefficients below .30 had item-total correlations below .30 in at least two of the three samples. The fourth item, which assesses political diffusion, had an item-total correlation below .30 in only one of the three samples. For Interpersonal Diffusion, the two items with factor pattern coefficients below .30 in at least two of the samples also had item-total correlations below .30 in both the Hispanic American and Swedish samples. For Ideological Moratorium, the one item with a factor pattern coefficient below .30 in at least two of the samples had item-total correlations of .08 and .10 in the White American and Hispanic American samples, respectively, and an item-total correlation of .23 in the Swedish sample. For Interpersonal Moratorium, two of the three items with factor pattern coefficients below .30 in at least two of the samples also had item-total correlations below .30 in all three samples. The third item, which assesses recreational moratorium, had item-total correlations above .30 in both the White American and Hispanic American samples. For Ideological Achievement, all three of the items with factor pattern coefficients below .30 in at least two of the samples also had item-total correlations below .30 in at least two of the three samples. For Interpersonal Achievement, the one item with a factor pattern coefficient below .30 in at least two of the samples had item-total correlations of .22 and .06 in the Hispanic American and Swedish samples, respectively, and an item-total correlation of .38 in the White American sample. In the cases in which internal consistency estimates decreased when these items were removed from consideration, these decreases may have been due to decreasing the number of items in the reliability analysis, particularly at the observed level of analysis.

DISCUSSION

We undertook this study to ascertain how a commonly used measure of identity status, the EOM–EIS–II, may function across three dissimilar samples. This study is the first cross-cultural evaluation of an identity measure in which (a) full data sets were compared between American and non-American samples and (b) both the ideological and interpersonal content-area clusters were used in analysis. Moreover, this study included U.S. samples from White and Hispanic groups and examined these groups separately. As a result, these results can be brought to bear on the applicability of the EOM–EIS–II in both cross-cultural and cross-ethnic contexts. Given the increasingly international scope of identity research and in light of the rapidly growing ethnic minority (particularly Hispanic) population of the United States, this is a critical research direction.

The measurement structure of the EOM–EIS–II appeared equivalent across samples, but latent variable mean endorsement of each identity status (except for Moratorium when covariates were not included) differed across samples. When they were evaluated at both the latent and observed levels, internal consistency estimates were generally below acceptable levels. Only when items not adequately representing their corresponding subscales were eliminated from analysis were latent internal consistency estimates acceptable (and comparable) for Foreclosure, Moratorium, and Achievement across the three samples. Internal consistency reliability estimates for observed scores, however, remained below acceptable levels even when these low-loading items were removed from analysis. It is possible that removing low-loading items from the EOM–EIS–II is more beneficial at the latent level of analysis than at the observed level of analysis.

These findings suggest that the EOM–EIS–II is appropriate for use in diverse ethnic and cultural contexts—but that some items (especially those assessing diffusion in the occupational domain, moratorium in the recreational domain, and achievement in the political domain) may need to be revised or deleted to improve the internal consistency of scores generated using the instrument. It is also possible that the grouping of content areas into ideological and interpersonal categories may create internal consistency problems because the content areas classified into each category are not strongly conceptually or empirically related (cf. Schwartz, 2001). Moreover, the differences in latent versus observed internal consistency estimates once these low-loading items were removed suggests that analyses using EOM–EIS–II data may best be conducted at the latent level of analysis. Analyses conducted using only observed scores may be somewhat less reliable (and may have less certain meaning) than may analyses conducted using latent factor pattern coefficients and mean scores.

Table 4 summarizes the patterns of mean differences at the observed versus latent levels, and with and without controls for age and gender. Because the overall MANOVA of observed scores across the three samples yielded a significant result, $F(16, 2076) = 171.76, p < .001$, and because all univariate results were significant at $p < .05$, we were justified in reporting pairwise differences at the observed level. Overall, one noteworthy parallel and three noteworthy differences in findings emerged between the observed and latent levels of analysis. In terms of parallels, at the observed level (both with and without covariates) and at the latent level (without covariates), the Swedish sample tended to be characterized by the lowest scores and was significantly different from the Hispanic American sample in all of the comparisons conducted. Differences between the Swedish and White American samples were inconsistent at both the observed and latent levels of analysis when covariates were not controlled. This pattern of findings appears to suggests that although the EOM–EIS–II is somewhat sensitive to cross-cultural differences, it is most sensitive to differences that are both cross-cultural and cross-ethnic. The Swedish and Hispanic American samples, which were significantly different from one another on all of the comparisons conducted, are dissimilar in terms of both nationality and majority/minority status.

In terms of differences between observed and latent analyses, one clear pattern that emerged from an inspection of Table 2 is that in nearly all of the mean comparisons conducted, effect sizes were larger at the latent level of analysis. This pattern is consistent with Hancock et al. (2000) who demonstrated that latent mean comparisons tend to be more statistically powerful and precise than observed-level mean comparisons because of controls for measurement error. Second, the White American and Hispanic American samples differed significantly from one another at the latent level when covariates were controlled, but observed-level comparisons generally did not differentiate the White American and Hispanic American samples regardless of whether covariates were controlled. This pattern suggests that latent-level analyses, with appropriate controls for demographic differences between samples, appear more suited to detect cross-ethnic differences in EOM–EIS–II identity status means than do observed-level analyses. Third, latent-level analyses on EOM–EIS–II data may be more sensitive to age and gender effects than may observed-level analyses. Whereas the patterns of pairwise latent mean differences changed when age and gender differences were controlled, the observed-level patterns generally did not change (or changed less appreciably) when age and gender were taken into account. All three of these findings suggest that, at least with respect to the EOM–EIS–II, mean-difference analyses may best be conducted at the latent level of analysis. Given the comparative insensitivity of observed-level analyses to removal of problematic items, age and gender effects, and cross-ethnic differences, the accuracy of mean-difference analyses conducted and interpreted at the observed level may be somewhat uncertain.

The age and gender differences may themselves be important to interpret, inasmuch as age and gender are considered as important variables in identity development (Lewis, 2003; Schwartz & Montgomery, 2002). The patterns of mean differences obtained with age and gender uncontrolled may reflect the contributions of these two variables. The Swedish sample was significantly younger and had a greater representation of males than did the American samples. At both the observed and latent levels, the Swedish sample scored significantly higher on indexes of Ideological Diffusion and Moratorium than did either of the American samples. Diffusion and Moratorium both represent low levels of commitment, which appears to differentiate more from less successful identity development (Schwartz et al., 2005). In the Schwartz et al. study, both men and younger emerging adults were overrepresented among participants characterized as less successful in their identity development. The fact that this difference emerged in the ideological domains may reflect cultural differences in the salience of some of these domains. For example, American emerging adults have generally been characterized as disinterested in politics as evidenced by the low election turnout among this group (Kaplan & Venezky, 1994). Similarly, religion may be more salient to Americans than to northern Europeans (cf. Goossens, 2001). The contributions of age and gender in this study are consistent with all

TABLE 4
Patterns of Significant Differences for Observed and Latent Analyses With and Without Controls for Age and Gender

	Without Covariates		*With Covariates*	
Subscale	*Observed*	*Latent*	*Observed*	*Latent*
Ideological Diffusion	S > HA = WA	S > WA = HA	S > HA = WA	HA > WA > S
Interpersonal Diffusion	HA > S	HA = WA > S	WA = HA > S	HA > WA > S
Ideological Foreclosure	HA > S	HA = WA > S	HA = WA > S	S > HA > WA
Interpersonal Foreclosure	HA > S	HA = WA > S	HA = WA > S	S > WA > HA
Ideological Moratorium	S > HA = WA	NA[a]	NS[b]	S > WA > HA
Interpersonal Moratorium	WA = HA > S	NA[a]	WA = HA > S	HA > S > WA
Ideological Achievement	HA > WA > S	HA = WA > S	HA > WA > S	WA > HA > S
Interpersonal Achievement	HA > WA > S	HA > WA > S	HA > WA > S	WA > HA > S

Note. An equals sign (=) indicates that the sample means in question did not differ on that subscale at $p < .05$. In cases where the relationship between only two sample means is indicated, the third sample mean did not differ significantly from either of the other sample means. S = Swedish; HA = Hispanic American; WA = White American.

[a]The omnibus test for latent mean differences was not statistically significant. [b]None of the contrasts between samples were significant at $p < .05$.

of these prior findings. It should be noted, however, that the age ranges within each of the samples (especially the Swedish sample) were limited. The findings might have been somewhat different had a wider range of ages been represented in each sample.

These findings are consistent with those of Jensen et al. (1998) and Stegarud et al. (1999) who reported that Scandinavians scored significantly lower than Americans on all of the EOM–EIS–II scales only when measurement error and age and gender differences were not statistically controlled. Although our observed-level results are largely consistent with those of the Norwegian studies, these latent-level results suggest that when age differences, gender differences, and measurement error were taken into account, the Swedish sample scored highest on both Foreclosure scales and Ideological Moratorium (and intermediate on Interpersonal Moratorium). These findings highlight the importance of controlling for demographic differences and for measurement error when conducting cross-cultural mean comparisons of identity status scores. Previous studies (Dwairy, 2004; Jensen et al., 1998; Stegarud et al., 1999) addressing this topic, which have used only published American means, have not been able to implement such controls.

It is noteworthy that at the latent level and with age and gender differences controlled, the Hispanic American sample scored significantly higher than the White American sample on Diffusion in both sets of domains, Foreclosure in the interpersonal domains, and Moratorium in the ideological domains. At first glance, this pattern of results appears implausible given one of the basic assumptions of the identity status model—that foreclosure and moratorium are incompatible with one another. However, one should consider that the majority of Hispanic American participants were United States born but raised by immigrant parents. The incompatibilities between American culture and many Hispanic cultures may have important implications for identity development in second-generation Hispanic immigrants (cf. Schwartz & Montgomery, 2002). It is possible that the Hispanic American participants in this study were attempting to straddle two cultures—making identity commitments as per Hispanic culture while at the same time “keeping their options open” as per American culture. It is noteworthy, however, that White Americans scored significantly higher than Hispanic Americans on achievement in both sets of domains, perhaps suggesting that straddling two sets of cultural expectations interferes with identity consolidation.

The fact that the Swedish participants completed a Swedish-language version of the measure may also been partially responsible for some of the differences found in the comparisons of observed means. For example, Goossens (2001) observed that concepts such as dating and religion have different meanings in European languages than in English. The consistency of factor structure suggests that the same basic set of content areas can be applied across ethnic and cultural groups. However, mean differences in endorsement between the Swedish sample and the American samples (who completed an English version of the measure) were the only significant differences that emerged when measurement error was left uncontrolled. This may suggest that when the EOM–EIS–II and other identity measures are translated into other languages, content areas and item wording may need to be altered slightly so as to be culturally and linguistically syntonic. Literal translation of terms from one language to another may inadvertently increase measurement error.

It is noteworthy that the factor structure of the EOM–EIS was consistent across the three samples studied despite the fact that (a) the Swedish participants completed a translated version of the measure and (b) English was likely a second language for many of the Hispanic American participants. Mean differences in item endorsement and subscale means emerging from these analyses may be a function of cultural differences, language or translation issues, or both. Although measurement error related to translation/language issues can be controlled in SEM, there may be more substantive issues related to language and translation that do not affect the factor structure of the instrument but are associated with higher or lower endorsement of items in specific content areas. Future cross-cultural and cross-ethnic research using identity measures should attempt to uncouple the effects of translation and language issues from the effects of ethnic and/or cultural differences. Asking bilingual ethnic minority and international participants to complete the measure both in English and in their native languages may help to identify differences that are due to translation or language issues. It may also be advantageous to administer the measure to immigrant as well as second generation, ethnic minority participants to more systematically determine the effects of acculturation on responses.

Limitations and Future Directions

This study is limited in at least two important ways that should be considered when interpreting the results. First, significant age and gender differences emerged between the Swedish sample and the two American samples. Compared to the two American samples, the Swedish sample was significantly younger and was characterized by a more balanced gender distribution. It should be noted that the White American and Hispanic American samples included small numbers of men (White American, n = 36; Hispanic American, n = 83). Men represented only 16% of the White American sample and 23% of the Hispanic American sample. Although it is possible to statistically control for these age and gender differences, the fact that the Swedish and American participants were at somewhat different points in their respective educational systems, along with the possibility of gender differences in identity status (Lewis, 2003; Schwartz & Montgomery, 2002), may have affected the results in ways that cannot be statistically controlled.

Second, the EOM–EIS–II was administered differently within each sample. White American participants completed the measure on the Internet, Hispanic American participants completed the measure at home, and Swedish participants completed the measure in class. Although the effects of such methodological variations are beyond the scope of this study, it is reasonable to expect that differences in the setting in which participants completed the EOM–EIS–II may have impacted their responses (e.g., because of potential differences in social desirability; see Schwartz, Mullis, & Dunham, 1998). The fact that the factor structure of the EOM–EIS–II was equivalent across samples may temper this limitation to some extent. However, the different administration methods may have been somewhat responsible for the mean differences we observed in this study. The effects of such methodological variations should be examined in future empirical investigations.

Despite these limitations, this study has generated potentially important results concerning the cross-cultural and cross-ethnic properties of a widely used identity measure. The measurement structure of the EOM–EIS–II was consistent across three diverse contexts and administration methods, which suggests that the measure is appropriate for use with culturally and ethnically diverse groups. It is also important to specify exactly how latent-level, mean-difference analyses could be conducted in new research (or with existing data) using the EOM–EIS–II and how such methods can advance identity research. The technical aspects of latent mean comparisons are described in Hancock (1997) and Hancock et al. (2000). Substantively, it is important to identify the extent to which the correlates of each identity status differ depending on whether identity status itself is measured at the latent versus observed level. For example, one of van Hoof's (1999) criticisms of identity status theory is that the four statuses do not each map onto a distinct set of correlates. A cleaner and more theoretically consistent pattern of correlates of each status might have emerged in prior research had measurement unreliability been statistically controlled. To further evaluate the theoretical and empirical benefits of analyzing identity status data using latent-variable methods, it is important for research to ascertain the correlates of each identity status at the latent level of analysis.

These results also suggest that some EOM–EIS–II items did not pattern well onto their assigned subscales. As a result, future research should further examine the extent to which revising or removing items that load poorly on their assigned subscales as well as modifying certain content areas may improve the appropriateness of the measure both overall and for specific cultural and ethnic groups. Through such a process, advances in identity measurement may facilitate more cross-ethnic and cross-cultural identity research as well as research on mean endorsement levels for specific identity content areas within and between cultures.

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