Exploring biphobia among heterosexuals and LGBTQ+ individuals in Spain: Psychometric properties of the Biphobia Scale in a Spanish sample

Alberto González Robles[a*], Pablo Roca[b,c], María del Mar Sánchez-Fuentes[d], Yaiza Jiménez-Pérez[a], Daniel Ondé[e]

[a]*Department of Psychology and Sociology, University of Zaragoza, Teruel, Spain* [b]*Faculty of Health Sciences, Universidad Villanueva, Madrid, Spain.*

[c]*School of Health Sciences, Valencian International University, Valencia, Spain.*

[d]*Mind, Brain, and Behavior Research Center, University of Granada, Granada, Spain*

[e]*Department of Psychobiology & Behavioral Sciences Methods, Complutense University of Madrid, Madrid, Spain.*

***Corresponding author**: Alberto-González Robles (gonzaleza@unizar.es)

ORCID iD

Alberto González-Robles; https://orcid.org/0000-0002-3749-3756

Pablo Roca; https://orcid.org/0000-0001-9815-4902

María del Mar Sánchez-Fuentes; https://orcid.org/0000-0003-3657-4981

Yaiza Jiménez-Pérez; https://orcid.org/0009-0002-3104-2230

Daniel Ondé; https://orcid.org/0000-0002-1032-4506

Abstract

Self-reported questionnaires for assessing biphobia have been developed and validated in English-speaking countries. In Spain, despite considerable rates of biphobia, there is a lack of validated instruments to assess this phenomenon. The aim of this study was to explore the psychometric properties of the Spanish version of the Biphobia Scale (BS), a useful self-reported measure of biphobia. Four hundred and sixty-six participants from the general population in Spain completed the BS, the Ambivalent Sexism Inventory, the Modern Homophobia Scale, and the Internalized Homonegativity Scale. The factorial structure, internal consistency, and convergent validity were analyzed. The results of the factor analysis support a one-factor solution. Cronbach's alpha and Omega coefficient were .95 and .96, respectively. Evidence of convergent validity was provided, as shown by the significant and positive correlations between biphobia, sexism, homophobia, and internalized homonegativity. The results support the use of the BS in the Spanish population as a valid and reliable measure.

Key words: biphobia, lgbtq+, validity, reliability

Introduction

An extensive body of literature has shown associations between being lesbian, gay, bisexual, transgender, or queer (LGBTQ+) and negative social and psychological consequences, including bullying and cyberbullying (Berlan et al., 2010; Elipe et al., 2018), deterioration of mental health (Almeida et al., 2009; Wittgens et al., 2022), difficulties in social and family support (Hawthorne et al., 2020; Saewyc, 2011; Williams et al., 2005), physical aggression and sexual assault (Bender & Lauritsen, 2021; Effrig et al., 2011; Price et al., 2023), difficulties in access to health services (Alencar et al., 2016), and other subtle manifestations of LGBTQ+ discrimination, such as antigay jokes and slurs (Seelman et al., 2017).

Relative to other LGBTQ+ sexual orientations, bisexual individuals seem to be at a particularly high risk for developing social and psychological difficulties. For example, compared to gays and lesbians, bisexual individuals show more difficulties in obtaining and maintaining a social support network, higher rates of identity confusion, and lower disclosure and expression of their sexual identity (Balsam & Mohr, 2007; Dodge et al., 2012). These findings are consistent with the idea that people tend to categorize reality into binary or opposite categories, and each category has its mutually exclusive opposite. Men and women are seen as 'opposite categories,' and, as a result, bisexuality is commonly considered immature or confused compared to lesbians and gay individuals or even denied (Brewster & Moradi, 2010; Ochs, 1996). According to the literature, bisexual individuals also show a higher prevalence of psychopathology than gays and lesbians (Meyer, 2003; Persson & Pfaus, 2015). For instance, bisexuals have been shown to have a higher rate of suicidal ideation and attempts than heterosexuals, gays, and lesbians (Pompili et al., 2014). Moreover, meta-analytic studies have shown that, compared to lesbian/gay individuals, bisexuals are at a greater risk for mental disorders such as depression (Ross et al., 2018; Wittgens et al., 2022). Furthermore, biphobia has been associated with other forms of discrimination, such as homophobia directed towards gay men or lesbian women, sexism, transphobia, and internalized homophobia (Dierckx et al., 2017; Eliason, 1997; Garelick et al., 2017; Nagoshi et al., 2008, 2023), and these associations are stronger in those identified as heterosexuals than in those who consider themselves LGBTQ+ (e.g., López-Sáez et al., 2020; Mohr & Rochlen, 1999). These data should not be ignored, especially when considering that bisexuality is the most reported sexual orientation among LGBTQ+ individuals (COGAM, 2022; Jones, 2021).

Two common sociodemographic characteristics analyzed in the context of biphobia are sexual orientation and gender. Regarding sexual orientation, research conducted over the past decades supports the notion that bisexual individuals face 'double discrimination' from both heterosexual and LGBTQ+ communities (Doan Van et al., 2019; Friedman et al., 2014; Ochs, 1996). Furthermore, even though biphobia has been systematically found to be higher among heterosexuals than among LGBTQ+ individuals (e.g., Dodge et al., 2016; Friedman et al., 2014; Mulick & Wright, 2011), bisexuals also encounter discrimination from their monosexual (gays/lesbians) counterparts, although these forms of discrimination seem to be less overt and more challenging to detect (Brewster & Moradi, 2010; Weiss, 2011). Notably, bisexuals have been shown to have lower levels of biphobia compared to monosexual individuals, with the lowest levels of biphobia found among bisexual women (Friedman et al., 2014). When comparing negative attitudes toward bisexual men and women, the former are subject to higher levels of biphobia. Regarding gender, women generally report more positive attitudes toward bisexual individuals than men (e.g., Dodge et al., 2016; Eliason, 1997; Garelick et al., 2017; Mulick & Wright, 2011). It is important to note, however, that other studies have not reported significant differences in biphobia based on the gender of the participants surveyed (e.g., Hertlein et al., 2016; Mulick & Wright, 2002). Nevertheless, these studies did not use robust sampling procedures, and their samples were not representative of the general population. When more rigorous sampling methods are employed (e.g., probabilistic sampling), significant differences between men and women emerge (Dodge et al., 2016), indicating that these differences do indeed exist.

Assessing biphobia in Spain

In our country, it is not easy to find data that are representative of the general population regarding attitudes towards bisexuals. However, there are at least two sources of data that suggest the overall picture in terms of biphobia. Firstly, a few surveys about the attitudes of the population towards sexual minorities have been conducted over the past years. Nonetheless, these surveys primarily focused on LGBTQ+ individuals in general, although they do provide some noteworthy data regarding biphobia. For example, an analysis of LGBTQ+ prejudices among 1043 youth between 14 and 18 years old showed that between 14 and 22% of the sample had prejudices against bisexual individuals (SOMOS LGBT+, 2017). In a more recent study with a representative sample of 6256 adolescents, 42% of them expressed their unwillingness to have a bisexual partner (COGAM, 2022). In a related context, a meta-analysis that analyzed gender-based bullying against LGBTQ+ schoolchildren and adolescents in Spain reported that approximately half of LGBTQ+ minors self-identified as having experienced some form of bullying during their lives (Feijóo & Rodríguez-Fernández, 2021). According to these findings, many LGBTQ+ youth appear to face a hostile environment in schools. Secondly, most academic studies conducted in Spain that explore negative attitudes towards LGBTQ+ individuals have focused on college students (e.g., Amigo-Ventureira et al., 2022; López-Sáez et al., 2020). Overall, the results of these studies show a significant presence of negative attitudes towards LGBTQ+ individuals in Spain, particularly among heterosexuals, with women and LGBTQ+ individuals being the least biphobic, a pattern consistent with previous research (e.g., Anselmi et al., 2015; Dodge et al., 2016). However, research in Spain suggests that bisexual individuals are not immune to discrimination from monosexual individuals within the LGBTQ+ community (e.g., Mittal et al., 2022).

With the significant prevalence among the population and the potential negative impact of biphobia, an important area of research is the development and validation of assessment instruments with adequate psychometric properties to evaluate these attitudes. Several instruments have been developed for the assessment of biphobia. For the purposes of our study, we selected the Biphobia Scale (BS; Mulick & Wright, 2002), a 30-item instrument designed to assess negative cognitions, affects, and behaviors regarding bisexuality and bisexual individuals. Several factors were considered when choosing the BS. First, in line with previous systematic reviews on instruments for the measurement of binegativity (Bishop & Pynoo, 2022; Morrison et al., 2018), one of the most commonly used scales for assessing biphobia is the BS (Mulick & Wright, 2002, 2011). Second, according to Bishop and Pynoo (2022), the BS is one of the best available scales for assessing biphobia in terms of its psychometric properties. Another instrument analyzed in this review, the Stereotypes about Bisexual Men and Women Scales (Parent, 2012), was found to have better psychometric properties. However, we decided not to use it and instead use the BS for two reasons: a) its psychometric properties have not been published in a peer-reviewed journal, and b) it consists of one scale for each gender, resulting in a larger number of items, which can complicate participant recruitment. Third, the BS has shown adequate psychometric properties, including internal consistency, convergent and discriminant validity, and a unidimensional factor structure in English-speaking samples (Mulick & Wright, 2002, 2011). Finally, while the BS has been used in some studies with Spanish samples (e.g., López-Sáez et al., 2020), its psychometric properties have not yet been analyzed in Spanish individuals. To the authors' knowledge, there is no validated scale for assessing biphobia in Spain. Although some surveys have been conducted among adolescents as pinpointed above, a psychometrically sound measure is currently lacking to comprehensively evaluate biphobia in our country.

Current study

Given the distinct features of biphobia compared to other manifestations of LGBTQ+ negative attitudes (e.g., homophobia) and the absence of validated scales for biphobia in Spanish, the main goal of the current study was to examine the psychometric properties of the Spanish version of the BS (Mulick & Wright, 2002, 2011). The specific aims of this study were to analyze the factorial structure, the internal consistency, and the convergent validity of the instrument. In accordance with the results of the original validation (Mulick & Wright, 2002), we hypothesized high levels of internal consistency and a unidimensional factor structure. Furthermore, we expected significant positive correlations between biphobia and other measures (e.g., internalized homonegativity, sexism, and homophobia).

Method

Participants

A cross-sectional study was conducted. To participate in the study, individuals had to be 18 years old and sign a written informed consent. The final sample consisted of 466 participants from the general population (32.5% men and 67.5% women), with an age range of 18 to

65 years old (*M* = 27.04; *SD* = 8.4). Based on the recommendations of previous studies, the recruited sample was considered adequate for our study objectives (Rouquette & Falissard, 2011). Most of the participants were cisgender (91.6%), followed by unknown (5.8%), and only 2.6% of the participants considered themselves to be transgender. Regarding sexual orientation, most people considered themselves heterosexual (53.3%). Finally, most of the participants had university studies (64.5%). Table 1 shows the sample's socio-demographic characteristics in detail.

TABLE 1 ABOUT HERE

Measures

Sociodemographic questionnaire

Participants were asked about their sex, age, nationality, marital status, education level, and occupation.

Kinsey Scale

This scale was employed to assess the individuals with whom the participants have sexual relations. This scale uses eight response options for sexual practices, from "exclusively heterosexuals" (option 1) to "exclusively homosexuals" (option 7). An eighth option was included to account for "asexuality" (Kinsey et al., 1998).

Biphobia Scale (BS)

This is a scale consisting of 30 items rated in a 6-point Likert scale from 1 ("Strongly agree") to 6 ("Strongly disagree") that evaluates negative attitudes toward bisexuality and bisexual people. Higher scores indicate higher levels of biphobia. The total score is obtained summing up the scores on each item. The score on most items must be inverted before summing them up to obtain the total score (items 1, 2, 6, 7, 8, 9, 10, 12, 13, 14, 15, 17, 18, 19, 21, 22, 23, 24, 25, 26, 27, 28, and 30). Finally, 30 points are subtracted from the total score to obtain scores between 0 and 150 (Mulick & Wright, 2002).

Ambivalent Sexism Inventory (ASI)

We used the Spanish adaptation of the inventory (Expósito et al., 1998). The ASI consists of 22 items rated in a Likert scale from 0 ("Strongly disagree") to 5 ("Strongly agree") that evaluates ambivalence towards women and the perception of inferiority of women as a social group. The scale has two subscales: hostile sexism (ASI-HS) and benevolent sexism (ASI-BS). Higher scores on each of these factors indicate sexist attitudes (Glick & Fiske, 1996). In the present study, both the HS and BS subscales showed high levels of internal consistency (Cronbach's alpha: $\alpha_{ASI\text{-}HS} = .93$, $\alpha_{ASI\text{-}BS} = .85$; McDonald's omega: $\omega_{ASI\text{-}HS} = .95$, $\omega_{ASI\text{-}BS} = .89$).

Modern Homophobia Scale (MHS)

We used the Spanish validation of the scale (Rodríguez-Castro et al., 2013). This questionnaire is made up of two subscale that assess homophobia toward gay men and homophobia toward lesbian women. The *attitudes toward gay men* scale consists of 22 items rated in a 5-point Likert scale from 1 ("Strongly disagree") to 5 ("Strongly agree"). The scale has three subscales: a) personal discomfort with gay men (MHS-G-PD); b) deviance/changeability of male homosexuality (MHS-

G-D); and c) institutional homophobia toward gay men (MHS-G-IH). Higher scores indicate more favorable attitudes toward gay men. In the current study, all factors showed adequate internal consistency ($\alpha_{MHS-G-PD}$ = .75, $\alpha_{MHS-G-D}$ = .88, $\alpha_{MHS-G-IH}$ = .68; $\omega_{MHS-G-PD}$ = .83, $\omega_{MHS-G-D}$ = .92, $\omega_{MHS-G-IH}$ = .78). On the other hand, the *attitudes toward lesbians* subscale (MHS-L) contains 24 items rated in a 5-point Likert scale from ("Strongly disagree") to 5 ("Strongly agree"). The scale is also divided into the same three subscales: a) personal discomfort with lesbians (MHS-L-PD); b) deviance/changeability of female homosexuality (MHS-L-D); and c) institutional homophobia toward lesbians (MHS-L-IH:) (Raja & Stokes, 1998). Higher scores indicate more favorable attitudes toward lesbians. In the current study, all factors showed adequate internal consistency ($\alpha_{MHS-L-PD}$ = .72, $\alpha_{MHS-L-D}$ = .94, $\alpha_{MHS-L-IH}$ = .66; $\omega_{MHS-L-PD}$ = .78, $\omega_{MHS-L-D}$ = .94, $\omega_{MHS-L-IH}$ = .70*).*

Internalized Homonegativity Scale (IHS-16)

We used the Spanish adaptation of the scale (de la Rubia, 2013). It is a 16-item self-report questionnaire for the assessment of negative attitudes toward homosexuality. Each item is rated from 1 ("Completely disagree") to 9 ("Completely agree"). This scale has three subscales: a) acceptance of public manifestations of homosexuality (IHN-16-1); b) internal acceptance of homosexual feelings, desires, and identity (IHN-16-2); and c) promiscuity and incapacity for stable relationships (IHN-16-3). Higher scores on these factors indicate a higher level of internalized homophobia (Currie et al., 2004). Given the nature of the instrument, it was only completed by LGBTQ+ individuals (*n* = 196). In the present study, the three factors showed adequate internal consistency (α_{IHN-1} = .55, α_{IHN-2} = .69, α_{IHN-3} = .77; ω_{IHN-1} = .72, ω_{IHN-2} = .77, ω_{IHN-3} = .78).

Procedure

The present study was approved by the Community of Aragon Research Ethics Committee. Firstly, a research team composed of bilingual psychologists and experts in psychometrics conducted the translation and adaptation of the BS from English into Spanish. This initial translation was individually evaluated by a bilingual expert and one of the study's researchers with expertise in the field of sexuality. Following the initial translation, a bilingual expert performed a back translation. Guidelines from previous research (Elosua et al., 2014; Muñiz et al., 2013), and the standards of the American Educational Research Association (AERA, 2014), the American Psychological Association (APA), and the National Council on Measurement in Education were followed during the translation and adaptation process. Once translated, the Spanish adaptation was sent to five Spanish experts in sexuality research and psychological assessment to identify and suggest needed changes to some wording of the items. In this phase, no changes were made. Subsequently, we carried out a pilot study involving 10 individuals with similar sociodemographic characteristics to the final sample. They were asked to what extent they understood each item. If they found any ambiguous term or expression, they were asked to indicate which one and why. Because all the items achieved more than 80% agreement about their clarity, no changes were made. The Spanish version of the BS can be seen in Supplementary Material 1.

Secondly, we employed two strategies to recruit the study sample: a snowball sampling technique and posts on social media platforms promoting the study (i.e., Twitter, Instagram). The recruitment phase lasted for 4 months, taking place between October 2022 and January 2023. When participants clicked on the link, they were granted access to the study information and informed consent, which outlined the study's purpose. None of the questions were mandatory, except for agreeing to the informed consent. Participation was anonymous, and all participants were volunteers who did not receive any compensation for taking part in this research. To mitigate the potential influence of uncooperative participants, three items aimed at assessing respondents' attentiveness were included in the survey (i.e., *1. Please select the option 'completely agree'*; 2. *Please select the option 'completely disagree'*; and *3. Please select the option 'moderately agree'*). Eight participants did not respond correctly to these three items, so they were excluded from the analyzed sample (*n* = 8)."

Statistical analyses

The data analysis procedure follows a similar structure to that of the authors of the original validation (Mulick & Wright, 2002). All analyses were performed using the R Studio software (R Core Team, 2021).

First, preliminary analyses of the BS items were conducted using descriptive statistics (missing values, mean, standard deviation, and response distribution). The results of these analyses were crucial for determining the estimation methods that would be used in subsequent analyses (i.e., the use of robust estimation methods in the case of strong asymmetry) (Curran et al., 1996).

Second, Exploratory Factor Analysis (EFA) was conducted to evaluate the dimensionality or factorial structure of the BS. Before applying EFA, we conducted an initial evaluation of sample adequacy or factorability of the data using the Kaiser-Meyer-Olkin index and Bartlett’s test. Several techniques for assessing dimensionality were employed to compare different clusters of items, including parallel analysis (PA), eigenvalues, and the consistency of factor loadings of factor solutions.

To achieve convergence in the results of the PA, two random eigenvalue extraction methods were employed: Principal Components (PA-PC) and Unweighted Least Squares (PA-ULS), using, in each case, the mean criterion and the 95th percentile criterion (see the study by Ledesma and Valero-Mora for technical details) (Ledesma & Valero-Mora, 2019). Eigenvalues were used to determine whether the first factor extracted from a factor solution clearly explained more variance than the second factor. In this case, the correlation matrix is mostly determined by a single source of variation (i.e., a single common factor).

Additionally, to assess the consistency of the factorial solution of the Spanish version, we inspected the factor loading values and compared them with those obtained by Mulick and Wright (2002) for the one-factor model, using the Congruence Coefficient (Ck) (Tucker & Lewis, 1973) to assess the similarity between two factorial solutions that share the same internal structure. Ck calculates the discrepancy between the factor loadings of two factor solutions ($\lambda 1$ and $\lambda 2$) for each item i of factor k using the following expression (1): Ck ranges between −1 and 1, and values close to 1 reflect high similarity between the two factorial solutions compared.

$$Ck = \frac{\sum_{ik} \lambda 1_{ik} \lambda 2_{ik}}{\sqrt{(\sum_{ik} \lambda 1_{ik}^2)(\sum_{ik} \lambda 2_{ik}^2)}} \quad (1)$$

Third, the reliability of the BS was calculated using Cronbach's Alpha coefficient (α) and the Omega coefficient (ω). Although Mulick and Wright (2002) argued that the instrument can be conceptualized as a single dimension or factor, they also suggested that there could be up to four facets of content represented by the factorial solution of a single factor: cognitive (13 items), affective (8 items), behavioral avoidance (5 items), and behavioral acting-out (4 items). This issue raises the need to evaluate the reliability of the one-factor model in the presence of a certain degree of multidimensionality (i.e., if the measure can be considered essentially unidimensional) (Reise, 2012). To address this issue, we estimated reliability using the Hierarchical Omega coefficient (ω_H) and the Explained Common Variance index (ECV) by applying an Exploratory Bifactor Model (Reise, 2012). Values of $\omega_H > .70$ and ECV $> .60$ indicate that the measure is essentially unidimensional (Reise et al., 2013).

Fourth, a two-way independent-measures ANOVA was conducted to assess the effects of sex and sexual orientation (independent variables) on the total score of the BS (dependent variable), estimating the effect size using partial eta squared (η^2) and post-hoc comparisons corrected by Bonferroni's method. To assess sexual orientation, we used the Kinsey Scale (Kinsey et al., 1998), grouping individuals as heterosexual if they chose the options "exclusively heterosexual" and "mostly heterosexual, slightly homosexual" (n = 246), and as LGBTQ+ if they selected any of the other options (n = 210) (except for the "asexual" option).

Finally, Pearson correlations were calculated between the BS total score and the other scales (i.e., ASI, MSH-G, MSH-L, and IHS-16) to assess the convergent validity of the instrument. The correlations were interpreted following Cohen's convention (Cohen, 1988): values between .10 and .30 are considered small, between .30 and .50 are considered medium, and above .50 are considered large. Based on previous studies (Mulick & Wright, 2002), we expected positive correlations between the BS and the ASI and IHS-16 scales, as well as negative correlations between the BS and the MHS-G and MHS-L scales (higher scores indicate more favorable attitudes toward gays and lesbians).

Results

Preliminary analyses

The mean of the responses to the BS items ranges between 1.04 and 1.71, with standard deviations between .38 and 1.57 (see Table 2). The distribution of the answers shows a positive asymmetry, with more than 90% of the answers concentrated in category 1 ("Strongly disagree", after the corresponding items have been inverted). Consequently, the assumption of normality was not fulfilled. For this reason, robust estimation methods were used in the application of EFA. Because there was only 1.4% of missing data, and the pattern of missing data was completely at random (Little MCAR test: $\chi^2_{(983)} = 983.29$; $p > .05$), only complete responses to all the items were used for the analysis (the final sample included 458 participants). The mean score for the total sample ($N = 458$) in the BS was 5,5 ($SD = 13.02$).

TABLE 2 ABOUT HERE

Exploratory factor analysis (EFA)

Sample adequacy tests indicated that the correlation matrix was adequate to apply EFA (Bartlett's test: $\chi^2 = 9988.5$; $p < .0005$; Kasier-Meyer-Olkin = .96). PA suggested the presence of three factors (both with PA-PC and PA-ULS). The first four eigenvalues were 13.8, 2.2, 1.5, and 1.1. The difference between the eigenvalue 1 and 2 was quite high (13.8 > 2.2), which suggests the existence of a single underlying dimension, in line with the original proposal by Mulick and Wright (2002). However, the eigenvalues 2 and 3 (2.2 and 1.5, respectively) may indicate the presence of specific variance in the data. The eigenvalue 4 (1.1) is borderline if we compare it with Kaiser's rule ($k > 1$), and it does not seem to contribute much to explain the communality of the items.

Based on these results, EFA was applied using Unweighted Least Squares (ULS) and Oblimin rotation (under the hypothesis that the multifactorial models would be correlated). The factorial solutions were forced to one factor (following the original proposal by Mulick and Wright), three factors (from the result of the PA) and four factors (to explore the four content facets suggested by Mulick and Wright). Table

3 shows the factor loadings obtained for each factorial solution.

TABLE 3 ABOUT HERE

In the 1-factor solution, the factor loadings (λ_{ik}*) are above .30 in most items, except for items 3, 16, and 29. Table 3 also shows the loadings obtained by Mulick and Wright (2002) for the 1-factor model. A congruence coefficient value Ck = .95 was obtained (values between .92 and .98 indicate an acceptable similarity between both versions of the scale) (MacCallum et al., 1999).

In the 3-factor solution, more than half of the items yielded greater λ_{ik}* values in factor 1 (F1: 19 items); items 3, 4, 5, 11, 16, 20 and 29 showed greater loadings in factor 2 (F2); and items 2, 15, 17 and 23 showed higher loadings in factor 3 (F3). Item 19 did not show loadings greater than .30 in any factor, so this item was excluded from subsequent analyses. In the 3-factor solution, the following correlations between factors were obtained: $\phi_{(1,2)} = .42$, $\phi_{(1,3)} = .56$, $\phi_{(2,3)} = .37$.

Finally, in the 4-factor solution, more than half of the items loaded in factor 1 (F1: 16 items); items 15, 25 and 26 showed greater loadings in factor 2 (F2); in factor 3 (F3) items 3, 4, 5, 16, 20 and 29; and in factor 4 (F4) items 2, 11, 17 and 23. Again, item 19 obtained loadings under .30 in all the factors and was not included in the following analyses. In the 4-factor solution, the following correlations between factors were obtained: $\phi_{(1,2)} = .33$, $\phi_{(1,3)} = .68$, $\phi_{(1,4)} = .48$, $\phi_{(2,3)} = .39$, $\phi_{(2,4)} = .29$, $\phi_{(3,4)} = .39$.

Reliability

Coefficients α and ω were calculated for all the items of the BS (1-factor model) and for the group of items of each of the subscales identified by EFA (multidimensional models of 3 and 4 factors). For the 1-factor model, values of $\alpha = .95$ and $\omega = .96$ were obtained. Removing items with λ_{ik}*< .30 (items 3, 16 y 29), these values were slightly higher ($\alpha = .96$, $\omega = .97$). For the 3-factor model, the following values were obtained: F1: $\alpha = .96$ and $\omega = .97$; F2: $\alpha = .72$ and $\omega = .80$; F3: $\alpha = .76$ and $\omega = .78$. For the 4-factor model, the following results were obtained: F1: $\alpha = .96$ and $\omega = .97$; F2: $\alpha = .84$ and $\omega = .85$; F3: $\alpha = .68$ and $\omega = .79$; F4: $\alpha = .73$ and $\omega = .75$.

The results on internal consistency indicate adequate levels for the three solutions evaluated. The results obtained show the existence of a general factor that explain the scores on the scale (1-factor model). However, there is also some degree of multidimensionality, as shown by the EFA solutions and the eigenvalues of factors 2 and 3. Therefore, to explore the unidimensionality of the scale, ω_H and ECV were calculated. The values obtained for the BS were $\omega_H = .71$ and ECV = .63 ($\omega_H = .81$ and ECV = .74 if items 3, 16 and 29 are removed). These values support the essential unidimensionality of the scale (i.e., $\omega_H > .70$ and ECV > .60).

Convergent validity

Table 4 shows the correlations between the BS and the rest of the measures to explore the convergent validity (i.e., ASI, MSH-G, MSH-L and IHS-16). As hypothesized, the results showed significant positive correlations between the BS and the ASI, both with the total score and with the two subscales (i.e., hostile sexism and benevolent sexism), with a medium-large size. No significant correlations were found between the BS and the MHS-G. However, significant negative correlations were observed between the BS and the MHS-L, both with the total score and with the three subscales (i.e., personal discomfort, deviance/changeability, and institutional homophobia), with a medium-large size. Likewise, there was a significant positive correlation between the BS and the IHS-16, both with the total score and with the three subscales (i.e., acceptance of public manifestations of homosexuality; internal acceptance of homosexual feelings, desires, and identity; and promiscuity and incapacity for stable relationships), with a medium size.

TABLE 4 ABOUT HERE

Associations between sex and sexual orientation with biphobia

To assess the potential effects of sex and sexual orientation on the BS total score (following the logic outlined above about the essential unidimensionality of the data), a two-way independent-measures ANOVA was computed. There were three missing values in the sex variable (i.e., non-response), so the analyzed sample included 455 participants. The interaction effect (sexual orientation x sex) was not significant ($F_{(1,451)} = 3.3$; $p = .072$). However, the two main effects analyzed were statistically significant: sexual orientation ($F_{(1,451)} = 18.0$,

$p < .0005$; $\eta^2 = 0.038$) and sex ($F_{(1,451)} = 6.1$, $p = .014$; $\eta 2 = .013$). There were significant differences between the heterosexual sample ($M = 7.5$; $SD = 16.0$) and the LGBTQ+ sample ($M = 3.0$; $SD = 7.7$). Regarding sex, men ($M = 7.6$; $SD = 18.2$) showed a greater score than women ($M = 4.4$; $SD = 9.6$). However, these results must be considered with caution because the statistical power associated with the effect of sex on the BS was less than .80 ($1 - \beta = .696$). The statistical power associated with the effect of sexual orientation was high ($1 - \beta = .989$).

Discussion

Main findings

The main objective of this study was to explore the psychometric properties of the Spanish version of the Biphobia Scale (BS) (Mulick & Wright, 2002). Specifically, our goals were to evaluate the internal structure of the instrument, estimate internal consistency, and explore the convergent validity between the BS and other related measures. Furthermore, we aimed to assess the degree of biphobia among Spanish individuals. In the following paragraphs, the main findings of the study are discussed.

First, the results showed that individuals in our sample exhibited a degree of biphobia in the mild range, according to the score interpretation proposed by Mulick and Wright (2002) (scores lower than 30). Moreover, in line with the original validation, heterosexuals were found to be significantly more biphobic than LGBTQ+ individuals. These results are consistent with previous studies showing that heterosexual individuals tend to hold more negative attitudes towards bisexuals compared to LGBTQ+ individuals (Dodge et al., 2016; Friedman et al., 2014; Mulick & Wright, 2002, 2011). A difference with the original validation is that we did find a significantly higher degree of biphobia among men than women, in line with previous studies showing sex differences in negative attitudes toward sexual minorities (Dodge et al., 2016; Ross et al., 2018). In any case, all the results on biphobia, regardless of sex or sexual orientation, were close to 0 (between 3 and 7.6). In this context, it is important to remember that the maximum score on the BS is 150, and that the original authors reported a mean score of 32.28 for heterosexuals and 13.33 for monosexual individuals. These differences might be partly explained by the context in which the measures were obtained (e.g., Spain vs. United States, early 2000s vs. 2022). In fact, according to an international survey conducted in 2019, Spain has been ranked as one of the least homophobic countries worldwide (Poushter & Kent, 2020). Other factors should be considered when interpreting these differences, such as the capacity of the instrument to capture common manifestations of biphobia in our country versus other countries or cultures (e.g., United States). Nevertheless, the results suggest that, overall, the individuals in our sample have a mild degree of biphobia.

Second, to assess the internal structure of the instrument, we replicated the analyses performed by Mulick and Wright (2002) and examined different factorial solutions. As in the original validation, the results of the different analyses performed suggest that the most suitable model is the one-factor solution because the factor loadings are greater than .30 in most of the items. Additionally, the results of the Hierarchical Omega coefficient support that the scale can be considered essentially unidimensional, hence supporting the recommendation of using a single global score by the authors of the original version (Mulick & Wright, 2002). Regarding internal consistency, a value of $\alpha = .95$ was obtained for the one-factor model, and this value increased ($\alpha = .96$) when removing the items with factor loadings < .30 (items 3, 16 and 29), showing excellent reliability.

Regarding convergent validity, the medium to large significant positive correlations between the BS and the ASI, and between the BS and the IHS-16, support previous studies showing the relationship between internalized homophobia, sexism, and other related constructs such as biphobia (Galupo, 2006; Makwana et al., 2018; Szymanski et al., 2005; Warriner et al., 2013). In fact, high levels of sexism and internalized homophobia have been identified as predictors of negative attitudes toward LGBTQ+ individuals (Barnes & Meyer, 2012; Davies, 2004; Glick et al., 2015). Also as predicted, a significant negative correlation was observed between the BS and the attitudes toward the lesbian subscale of the Modern Homophobia Scale (MHS). These results suggest that individuals hold similar prejudices against the monosexual and bisexual population. Research has shown that homophobia is associated with several factors that might be influencing other forms of LGBTQ+ phobia, such as biphobia. For example, some individuals may have conservative and rigid beliefs about sexuality, and others may have personality traits that facilitate the emergence of prejudices (Ficarrotto, 1990). However, no correlation was found between the attitudes toward gays subscale of the MHS. According to the literature, lesbian women are more exposed to discrimination just because they are women and, consequently, they also face sexist prejudices in addition to homophobic prejudices (González & Kokozos, 2019). Therefore, a potential explanation for the absence of a correlation between biphobia and homophobia towards gays and, at the same time, the positive correlation between biphobia and homophobia towards lesbians is that bisexuals and lesbians are not as accepted as gays. People might be more accepting towards gays but less open to accepting other (more invisible) forms of non-heterosexual orientation, such as bisexuality and lesbianism.

It should be noted that the reliability (i.e., internal consistency) of the scores obtained through these instruments is a key element in accurately estimating the correlations between scores. In this regard, the scores obtained from the IHN-1 and the IHN-2 subscales showed Cronbach's alpha values lower than 0.70. However, the Omega values exceeded this standard cut-off point for internal consistency. Therefore, we concluded that the three IHS-16 subscales exhibit good reliability, given that Omega is a more precise estimator of internal consistency than Cronbach's alpha in the absence of tau-equivalence (Trizano-Hermosilla & Alvarado, 2016), as observed in the data obtained in this study.

Limitations, conclusions, and future directions

This study has some limitations that should be mentioned. First, we used non-probabilistic sampling methods, so the results obtained cannot be generalized to the Spanish population. Second, most participants were female and had university studies, which may affect the representativeness of the results. Future research would benefit from recruiting a more heterogeneous sample (e.g., regarding sex, sexual orientation, and education). Relatedly, we found a pattern of extreme responses because more than 90% of the responses were concentrated in category 1 ("Strongly disagree"). Therefore, the assumption of normality was not fulfilled. This may be interpreted as a positive outcome from a social perspective, as individuals in our sample showed low levels of biphobia, at least as measured with the BS. However, some potential sources of bias should be considered when interpreting these results, such as the effects of social desirability in responding to the scale, or the difficulty of the instrument to identify subtle manifestations of biphobia because most items contain statements on overt biphobia. For example, item 2: "I do not like bisexual individuals," item 6: "I discriminate against bisexual people," item 14: "Bisexual people should not get married," or item 22: "I avoid bisexual people." In this vein, there is some published research that supports the existence of microaggressions like microinsults or microinvalidations that may be overlooked in many of the scales for the assessment of LGBTQ+ prejudices (Sue, 2010). An example of a microinsult is "a gay man being told he is gay because his mother was too overbearing" (Woodford et al., 2015, p. 1662). A microinvalidation entails the trivialization of experiences of oppression by marginalized groups. "We are all just people. Your sexuality doesn't matter" (Woodford et al., 2015, p. 1662), represents an example of microinvalidation. As can be seen in Supplementary Material 1, the BS evaluates generally overt manifestations of cognitive, affective, and behavioral biphobia, but it does not evaluate these subtle experiences. In any case, these types of instruments tend to elicit highly polarized responses without affecting the sensitivity of the item content. Finally, the results of the EFA 1-factor solution suggest that items 3, 16, and 29 should be reviewed to achieve a better fit because they showed loadings below the minimum recommended value of .30. However, to remove these items from the final version, it would be necessary to perform a confirmatory factor analysis with a larger sample of participants (since the data show a high degree of asymmetry). A future study with a larger sample is warranted to confirm the factorial structure found in this study.

To conclude, the Spanish version of the BS showed adequate internal consistency, convergent validity, and a 1-factor structure consistent with the original validation. As the results showed, the sample examined exhibited a low degree of biphobia. To the authors' knowledge, this is the first study to translate and examine the psychometric properties of a scale for the assessment of biphobia among Spanish individuals from the general population. Although some surveys have been conducted to assess the degree of biphobia among Spanish individuals, a more comprehensive and validated measure is lacking, and so we expect this study to contribute to filling this gap. Also, biphobia prevention programs conducted in Spain may benefit from a validated measure of biphobia for both the heterosexual and LGBTQ+

populations. For future research, we suggest that more heterogeneous samples should be recruited to obtain more representative outcomes, especially in terms of sexual orientation and education because most participants were heterosexual and highly educated. Moreover, a larger sample should be recruited to perform a confirmatory factor analysis. On the other hand, social desirability biases are more likely when socially unaccepted attitudes are to be evaluated (Krumpal, 2013), and therefore, future studies should consider controlling its influence by including a social desirability measure. Research focused on microaggressions toward bisexual individuals should be studied in future research. As the results showed, the degree of biphobia among our sample was considered mild. While this is a positive outcome, it might be reflecting the inability of the BS to capture subtle manifestations of biphobia that also have deleterious consequences on bisexual individuals. Additionally, future research should analyze whether the Spanish version of the BS is invariant depending on sex and sexual orientation. Likewise, cross-cultural studies are important because there are cultural differences, for example, in the sexual double standard (Sánchez-Fuentes et al., 2020). Finally, the validation of the BS in other Spanish-speaking countries could stimulate intercultural research and address the cultural bias present in this area of research.

Acknowledgements

This work was supported by Fundación Antonio Gargallo (University of Zaragoza) under Grant number 2021/B008.

Disclosure statement

The authors declare no conflict of interest.

References

Alencar Albuquerque, G., de Lima, C., da Silva, G., Alves, M. J. H., Belém, J. M., dos Santos Figueiredo, F. W., ... Adami, F. (2016). Access to health services by lesbian, gay, bisexual, and transgender persons: systematic literature review. *BMC International Health and Human Rights*, *16*, 1-10. https://doi.org/10.1186/s12914-015-0072-9

Almeida, J., Johnson, R. M., Corliss, H. L., Molnar, B. E., & Azrael, D. (2009). Emotional distress among LGBT youth: The influence of perceived discrimination based on sexual orientation. *Journal of Youth and Adolescence*, *38*, 1001-1014. https://doi.org/10.1007/s10964-009-9397-9

American Educational Research Association, American Psychological Association, and National Council on Measurement in Education (2014). *Standards for educational and psychological testing*. American Educational Research Association.

Amigo-Ventureira, A. M., Depalma, R., & Durán-Bouza, M. (2022). Homophobia and transphobia among Spanish practicing and future teachers. *American Journal of Sexuality Education*, *17*(3), 277-303. https://doi.org/10.1080/15546128.2022.2035290

Balsam, K. F., & Mohr, J. J. (2007). Adaptation to sexual orientation stigma: a comparison of bisexual and lesbian/gay adults. *Journal* *of Counseling Psychology*, *54*, 306-219. https://doi.org/10.1037/0022-0167.54.3.306

Barnes, D. M., & Meyer, I. H. (2012). Religious affiliation, internalized homophobia, and mental health in lesbians, gay men, and bisexuals. *American Journal of Orthopsychiatry*, *82*, 505-515. https://doi.org/10.1111/j.1939-0025.2012.01185.x

Bender, A. K., & Lauritsen, J. L. (2021). Violent victimization among lesbian, gay, and bisexual populations in the United States: Findings from the National Crime Victimization Survey, 2017–2018. *American Journal of Public Health*, *111*(2), 318-326. https://doi.org/10.2105/AJPH.2020.306017

Berlan, E. D., Corliss, H. L., Field, A. E., Goodman, E., & Austin, S. B. (2010). Sexual orientation and bullying among adolescents in the growing up today study. *Journal of Adolescent Health*, *46*(4), 366-371. https://doi.org/10.1016/j.jadohealth.2009.10.015

Bishop, C. J., & Pynoo, E. (2022). How well are outgroup attitudes and behaviours toward bisexual individuals measured? A systematic review of the psychometric properties of binegativity measures. *Psychology & Sexuality*, *13*, 182-200. https://doi.org/10.1080/19419899.2020.1756390

Brewster, M. E., & Moradi, B. (2010). Perceived experiences of anti-bisexual prejudice: Instrument development and evaluation. *Journal of Counseling Psychology*, *57*, 451-468. https://doi.org/10.1037/a0021116

COGAM Asociación LGTBI+ (2022). *LGTBFobia en las aulas*

2021/2022. Cogam. https://cogam.es/informe-lgtbifobia-en-las-aulas-2021-22/

Cohen, J. (1988). *Statistical power analysis for the behavioral sciences. 2nd ed*. Erlbaum Associates.

Curran, P. J., West, S. G., & Finch, J. F. (1996). The robustness of test statistics to nonnormality and specification error in confirmatory factor analysis. *Psychological methods*, 1(1), 16-29. https://doi.org/10.1037/1082-989X.1.1.16

Currie, M. R., Cunningham, E. G., & Findlay, B. M. (2004). The Short Internalized Homonegativity Scale: Examination of the factorial structure of a new measure of internalized homophobia. *Educational and Psychological Measurement*, *64*, 1053- 1067. https://doi.org/10.1177/0013164404264845

Davies, M. (2004). Correlates of negative attitudes toward gay men: Sexism, male role norms, and male sexuality. *Journal of Sex Research*, *41*, 259-266. https://doi.org/10.1080/00224490409552233

de la Rubia, J. M. (2013). Dimensionalidad, consistencia interna y distribución de la Escala Homonegatividad Internalizada en estudiantes mexicanos de Ciencias de la Salud [Dimensionality, internal consistency and distribution of the Internalized Homonegativity Scale in Mexican students of Health Sciences] *Acta de Investigación Psicológica*, *3*, 986-1004. https://doi.org/10.1016/S2007-4719(13)70947-5

Dierckx, M., Meier, P., & Motmans, J. (2017). "Beyond the box": A comprehensive study of sexist, homophobic, and transphobic attitudes among the Belgian population. *DiGeSt. Journal of Diversity and Gender Studies*, *4*(1), 5-34. https://doi.org/10.11116/digest.4.1.1

Doan Van, E. E., Mereish, E. H., Woulfe, J. M., & Katz-Wise, S. L. (2019). Perceived discrimination, coping mechanisms, and effects on health in bisexual and other non-monosexual adults. *Archives of Sexual Behavior*, *48*, 159-174. https://doi.org/10.1007/s10508-018-1254-z

Dodge, B., Herbenick, D., Friedman, M. R., Schick, V., Fu, T. C., Bostwick, W., ... Sandfort, T. G. (2016). Attitudes toward bisexual men and women among a nationally representative probability sample of adults in the United States. *PloS One*, *11*(10), e0164430. https://doi.org/10.1371/journal.pone.0164430

Dodge, B., Schnarrs, P. W., Reece, M., Martinez, O., Goncalves, G., Malebranche, D., Van Der Pol, B., Nix, R., & Fortenberry, J. D. (2012). Individual and social factors related to mental health concerns among bisexual men in the Midwestern United States. *Journal of Bisexuality*, *12*(2), 223-245. https://doi.org/10.1080/15299716.2012.674862

Effrig, J. C., Bieschke, K. J., & Locke, B. D. (2011). Examining victimization and psychological distress in transgender college students. *Journal of College Counseling*, *14*(2), 143-157. https://doi.org/10.1002/j.2161-1882.2011.tb00269.x

Eliason, M. J. (1997). The prevalence and nature of biphobia in heterosexual undergraduate students. *Archives of Sexual Behavior*, *26*, 317-326. https://doi.org/10.1023/A:1024527032040

Elipe, P., de la Oliva Muñoz, M., & Del Rey, R. (2018). Homophobic bullying and cyberbullying: Study of a silenced problem. *Journal of Homosexuality*, *65*(5), 672-686. https://doi.org/10.1080/00918369.2017.1333809

Elosua, P., Mujika, J., Almeida, L. S., & Hermosilla, D. (2014). Procedimientos analítico-racionales en la adaptación de tests. Adaptación al español de la batería de pruebas de razonamiento [Analytical-rational procedures in the adaptation of tests. Adaptation to Spanish of the battery of reasoning test]. *Revista Latinoamericana de Psicología*, *46*(2), 117–126. https://doi.org/10.1016/S0120-0534(14)70015-9

Expósito, F., Moya, M. C., & Glick, P. (1998). Sexismo ambivalente: medición y correlatos. [Ambivalent sexism: Measurement and correlates]. *International Journal of Social Psychology, 13*, 159-169. https://doi.org/10.1174/021347498760350641

Feijóo, S., & Rodríguez-Fernández, R. (2021). A meta-analytical review of gender-based school bullying in Spain. *International Journal of Environmental Research and Public Health*, *18*(23), 12687. https://doi.org/10.3390/ijerph182312687

Ficarrotto, T. J. (1990). Racism, sexism, and erotophobia: Attitudes of heterosexuals toward homosexuals. *Journal of*

Homosexuality, *19*, 111-116. https://doi.org/10.1300/J082v19n01_07

Friedman, M. R., Dodge, B., Schick, V., Herbenick, D., Hubach, R. D., Bowling, J., Goncalvez, G., Krier, S., & Reece, M. (2014). From bias to bisexual health disparities: Attitudes toward bisexual men and women in the United States. *LGBT Health*, *1*, 309-318. https://doi.org/10.1089/lgbt.2014.0005

Galupo, M. P. (2006). Sexism, heterosexism, and biphobia: The framing of bisexual women's friendships. *Journal of Bisexuality*, *6*, 35-45. https://doi.org/10.1300/J159v06n03_03

Garelick, A. S., Filip-Crawford, G., Varley, A. H., Nagoshi, C. T., Nagoshi, J. L., & Evans, R. (2017). Beyond the binary: Exploring the role of ambiguity in biphobia and transphobia. *Journal of Bisexuality*, *17*, 172-189. https://doi.org/10.1080/15299716.2017.1319890

Glick, P., & Fiske, S. T. (1996). The Ambivalent Sexism Inventory: Differentiating hostile and benevolent sexism. *Journal of Personality and Social Psychology, 70,* 491-512. https://doi.org/10.1037/0022-3514.70.3.491

Glick, P., Wilkerson, M., & Cuffe, M. (2015). Masculine identity, ambivalent sexism, and attitudes toward gender subtypes. *Social Psychology*, 46, https://doi.org/10.1027/1864-9335/a000228

González, M., & Kokozos, M. J. (2019). Prejudice reduction in public schools: A dialogic approach. *Journal of Educational Research and Practice*, *9*, 24. https://doi.org/10.5590/JERAP.2019.09.1.24

Hawthorne, O., Camic, P. M., & Rimes, K. A. (2020). Understanding the structure, experiences and challenges of social support for older lesbian, gay and bisexual people: a systematic review. *Ageing & Society*, *40*(2), 282-305. https://doi.org/10.1017/S0144686X18000910

Hertlein, K. M., Hartwell, E. E., & Munns, M. E. (2016). Attitudes toward bisexuality according to sexual orientation and gender. *Journal of Bisexuality*, *16*(3), 339-360. https://doi.org/10.1080/15299716.2016.1200510

Jones, J. M. (2021, February 24). *LGBT Identification Rises to 5.6% in Latest U.S. Estimate*. Optumhealtheducation. https://www.optumhealtheducation.com/sites/default/files/LGBT%20Identification%20Rises%20to%205.6%25%20in%20Latest%20U.S.%20Estimate.pdf

Kinsey, A. C., Pomeroy, W. B., & Martin, C. E. (1998). *Sexual behavior in the human male.* Indiana University Press (Original published 1948).

Krumpal, I. (2013). Determinants of social desirability bias in sensitive surveys: a literature review. *Quality & quantity*, *47*, 2025-2047. https://doi.org/10.1007/s11135-011-9640-9

Ledesma, R. D., & Valero-Mora, P. (2019). Determining the number of factors to retain in EFA: An easy-to-use computer program for carrying out parallel analysis. *Practical Assessment, Research, and Evaluation*, 12(1), 2. https://doi.org/10.7275/wjnc-nm63

López-Sáez, M. Á., & García-Dauder, D. (2020). Los test de masculinidad/feminidad como tecnologías psicológicas de control de género [Masculinity/femininity tests as psychological technologies of gender control]. *Athenea digital: revista de pensamiento e investigación social*, *20*, 00012. https://doi.org/10.5565/rev/athenea.2521

MacCallum, R. C., Widaman, K. F., Zhang, S., & Hong, S. (1999). Sample size in factor analysis. *Psychological Methods*, 4(1), 84-99. https://doi.org/10.1037/1082-989X.4.1.84

Makwana, A. P., Dhont, K., De keersmaecker, J., Akhlaghi-Ghaffarokh, P., Masure, M., & Roets, A. (2018). The motivated cognitive basis of transphobia: The roles of right-wing ideologies and gender role beliefs. *Sex Roles*, *79*, 206-217. https://doi.org/10.1007/s11199-017-0860-x

Meyer, I. H. (2003). Prejudice, social stress, and mental health in lesbian, gay, and bisexual populations: conceptual issues and research evidence. *Psychological Bulletin*, *129*, 674-697. https://doi.org/10.1037/0033-2909.129.5.674

Mittal, I., Ronzón-Tirado, R., Charak, R., & Cano-Gonzalez, I. (2022). Experiences of Microaggressions, Discrimination, & Exclusion within the LGBTQ+ Community in Spain. *MEDI 9331 Scholarly Activities Clinical Years*, *65*.

Mohr, J. J., & Rochlen, A. B. (1999). Measuring attitudes regarding bisexuality in lesbian, gay male, and heterosexual populations. *Journal of Counseling Psychology*, *46*, 353-369. https://doi.org/10.1037/0022-0167.46.3.353

Morrison, M. A., Bishop, CJ, & Morrison, T. G. (2018). A systematic review of the psychometric properties of composite LGBT prejudice and discrimination scales. J*ournal of Homosexuality*, *66*(4), 549-570. https://doi.org/10.1080/00918369.2017.1422935

Mulick, P. S., & Wright Jr, L. W. (2002). Examining the existence of biphobia in the heterosexual and homosexual populations. *Journal of Bisexuality*, *2*, 45-64. https://doi.org/10.1300/J159v02n04_03

Mulick, P. S., & Wright Jr, L. W. (2011). The biphobia scale a decade later: Reflections and additions. *Journal of Bisexuality*, *11*, 453-457. https://doi.org/10.1080/15299716.2011.620486

Muñiz, J., Elosua, P., & Hambleton, R. K. (2013). Directrices para la traducción y adaptación de los tests: segunda edición [Guidelines for the translation and adaptation of tests: second edition]. *Psicothema*, *25*, 151-157. https://doi.org/ 10.7334/psicothema2013.24

Nagoshi, J. L., Adams, K. A., Terrell, H. K., Hill, E. D., Brzuzy, S., & Nagoshi, C. T. (2008). Gender differences in correlates of homophobia and transphobia. *Sex Roles*, *59*, 521-531. https://doi.org/10.1007/s11199-008-9458-7

Nagoshi, C. T., Bechky, A., Nagoshi, J., & Pillai, V. K. (2023). For Straight Persons, Correlates of Biphobia Are Mostly the Same as for Homophobia. *Journal of Bisexuality*, *23*(2), 115-125. https://doi.org/10.1080/15299716.2022.2106607

Ochs, R. (1996). Biphobia: It goes more than two ways. In B. A.

Firestein (Ed.), *Bisexuality: The psychology and politics of an invisible minority* (pp. 217-239). Sage Publications.

Parent, M. C. (2012). *Content mapping and operationalization of stereotypes about lesbian women, gay men, bisexual women, and bisexual men* (Doctoral dissertation). ProQuest Dissertations and Theses database. (UMI No. 3586631)

Persson, T. J., & Pfaus, J. G. (2015). Bisexuality and mental health: Future research directions. *Journal of Bisexuality*, *15*(1), 82-98. https://doi.org/10.1080/15299716.2014.994694

Pompili, M., Lester, D., Forte, A., Seretti, M. E., Erbuto, D., Lamis, D. A., Amore, M., & Girardi, P. (2014). Bisexuality and suicide: a systematic review of the current literature. *The Journal of Sexual Medicine*, *11*(8), 1903-1913. https://doi.org/10.1111/jsm.12581

Poushter, J., Kent, N. O., & Center, P. R. (2020, August 3). The global divide on homosexuality persists but increasing acceptance in many countries over past two decades. Pew Research Center. https://www.pewresearch.org/global/2020/06/25/global-divide-on-homosexualitypersists

Price, M. N., Green, A. E., DeChants, J. P., & Davis, C. K. (2023). Physical dating violence victimization among LGBTQ youth: disclosure and association with mental health. *Journal of Interpersonal Vi*olence, *38*(15-16), 9059-9085. https://doi.org/10.1177/08862605231162655

R Core Team (2021). R: A language and environment for statistical computing. R Foundation for Statistical Computing, Vienna, Austria. https://www.R- project.org/

Raja, S., & Stokes, J. P. (1998). Assessing attitudes toward lesbians and gay men: The modern homophobia scale. *International Journal of Sexuality and Gender Studies*, *3*, 113-134. https://doi.org/10.1023/A:1023244427281

Reise, S. P. (2012). The rediscovery of bifactor measurement models. *Multivariate Behavioral Research*, *47*, 667-696. https://doi.org/10.1080/00273171.2012.715555

Reise, S. P., Scheines, R., Widaman, K. F., & Haviland, M. G. (2013). Multidimensionality and structural coefficient bias in structural equation modeling: A bifactor perspective. *Educational and Psychological Measurement*, *73*, 5-26. https://doi.org/10.1177/00131644124498

Rodríguez-Castro, Y., Lameiras-Fernández, M., Carrera-Fernández, V., & Vallejo- Medina, P. (2013). Validación de la escala de homofobia moderna en una muestra de adolescentes [Validation of the modern homophobia scale in a sample of adolescents]. *Anales de Psicología, 29,* 523-533. https://dx.doi.org/10.6018/analesps.29.2.137931

Ross, L. E., Salway, T., Tarasoff, L. A., MacKay, J. M., Hawkins, B. W., & Fehr, C. P. (2018). Prevalence of depression and anxiety among bisexual people compared to gay, lesbian, and heterosexual individuals: A systematic review and meta-analysis. *The Journal of Sex Research*, *55*(4-5), 435-456. https://doi.org/10.1080/00224499.2017.1387755

Rouquette, A., & Falissard, B. (2011). Sample size requirements for the internal validation of psychiatric scales. *International Journal of Methods in Psychiatric Research*, *20*(4), 235-249. https://doi.org/10.1002/mpr.352

Saewyc, E. M. (2011). Research on adolescent sexual orientation: Development, health disparities, stigma, and resilience. *Journal of research on adolescence*, *21*(1), 256-272. https://doi.org/10.1111/j.1532-7795.2010.00727.x

Sánchez-Fuentes, M. M., Moyano, N., Gómez-Berrocal, C., & Sierra, J. C. (2020). Invariance of the Sexual Double Standard Scale: A cross-cultural study. *International Journal of Environmental Research and Public Health, 17*, 1569. https://doi.org/10.3390/ijerph17051569

Seelman, K. L., Woodford, M. R., & Nicolazzo, Z. (2017). Victimization and microaggressions targeting LGBTQ college students: Gender identity as a moderator of psychological distress. *Journal of Ethnic & Cultural Diversity in Social Work*, *26*, 112-125. https://doi.org/10.1080/15313204.2016.1263816

Somos LGBT+. (n.d.). *Informe sobre homofobia, lesbofobia, bifobia y transfobia en las aulas aragonesas*. FELGTB. https://www.felgtb.com/ciberbullying/materiales/INFORME%20LGTBIFOBIA%20EN%20LAS%20AULAS%20ARAGON.pdf

Sue, D. W. (2010). *Microaggressions in everyday life: Race, gender, and sexual orientation*. John Wiley & Sons.

Szymanski, D. M. (2005). Heterosexism and sexism as correlates of psychological distress in lesbians. *Journal of Counseling & Development*, *83*, 355-360. https://doi.org/10.1002/j.1556-6678.2005.tb00355.x

Trizano-Hermosilla, I., & Alvarado, J. M. (2016). Best alternatives to Cronbach's alpha reliability in realistic conditions: congeneric and asymmetrical measurements. *Frontiers in Psychology*, *7*, 769. https://doi.org/10.3389/fpsyg.2016.00769

Tucker, L.R., & Lewis, C. (1973). A reliability coefficient for maximum likelihood factor analysis. *Psychometrika*, *38*, 1-10. https://doi.org/10.1007/BF02291170

Warriner, K., Nagoshi, C. T., & Nagoshi, J. L. (2013). Correlates of homophobia, transphobia, and internalized homophobia in gay or lesbian and heterosexual samples. *Journal of Homosexuality*, *60*, 1297-1314. https://doi.org/10.1080/00918369.2013.806177

Weiss, J. (2011). Reflective paper: GL versus BT: The archaeology of biphobia and transphobia within the US gay and lesbian community. *Journal of Bisexuality*, *11*(4), 498-502. https://doi.org/10.1080/15299716.2011.620848

Williams, T., Connolly, J., Pepler, D., & Craig, W. (2005). Peer victimization, social support, and psychosocial adjustment of sexual minority adolescents. *Journal of Youth and Adolescence*, *34*, 471-482. https://doi.org/10.1007/s10964-005-7264-x

Wittgens, C., Fischer, M. M., Buspavanich, P., Theobald, S., Schweizer, K., & Trautmann, S. (2022). Mental health in people with minority sexual orientations: A meta-analysis of population-based studies. *Acta Psychiatrica Scandinavica*, *145*(4), 357-372. https://doi.org/10.1111/acps.13405

Woodford, M. R., Chonody, J. M., Kulick, A., Brennan, D. J., & Renn, K. (2015). The LGBQ microaggressions on campus scale: A scale development and validation study. *Journal of Homosexuality*, *62*, 1660-1687. https://doi.org/10.1080/00918369.2015.1078205

Table 1. Sociodemographic characteristics.

Variables	*N*	%
Sexual orientation		
Exclusively heterosexual	248	53.3
Heterosexual, with sporadic homosexual contacts	26	5.6
Heterosexual, with frequent homosexual contacts	6	1.3
Bisexual	90	19.4
Homosexual, with frequent heterosexual contacts	7	1.5
Homosexual, with sporadic heterosexual contacts	19	4.1
Exclusively homosexual	62	13.3
Asexual	7	1.5
Marital Status		
Single	265	56.9
Married or partnered	189	40.6
Divorced or widowed	12	2.5
Educational Level		
Basic studies	10	2.2
Secondary studies	155	33.3
University studies	301	64.5
Occupation		
Student	200	43
Housekeeper	3	.6
Employed	215	46.1
Unemployed	41	8.8
Off work	5	1.1
Retired	2	.4

Table 2. Means (M), standard deviations (SD) for the BS items, and % of response for category 1 ("Strongly disagree").

Item	*M*	*SD*	*% response category 1*
1	1.2	.82	93.9%
2	1.2	.76	93.7%
3	1.7	1.15	64.6%
4	1.2	.81	95.4%
5	1.3	1.04	91.3%
6	1.1	.55	96.1%
7	1.1	.42	96.5%
8	1.1	.55	94.3%
9	1.1	.61	95.4%
10	1.0	.38	98.7%
11	1.3	.93	90.2%
12	1.1	.49	96.9%
13	1.1	.58	96.3%
14	1.1	.53	96.7%
15	1.3	.88	90.2%
16	1.6	1.57	83.4%
17	1.2	.87	91.0%
18	1.1	.43	97.4%
19	1.3	.91	87.1%
20	1.2	.91	94.3%
21	1.1	.56	95.6%
22	1.1	.47	97.4%
23	1.1	.61	95.4%
24	1.1	.45	97.4%
25	1.2	.69	92.4%
26	1.2	.70	93.9%
27	1.1	.71	96.3%
28	1.1	.46	98.0%
29	1.4	1.27	90.8%
30	1.1	.42	98.0%

Table 3. Estimated factor loadings for each factorial solution.

Item	Factorial loadings (λ_{ik}*)								
	1 factor	1 factor (Mulick & Wright, 2002)	3 factors			4 factors			
			F1	F2	F3	F1	F2	F3	F4
1	**,672**	**,77**	**,395**	,172	,273	**,356**	,125	,142	,263
2	**,713**	**,68**	,254	,234	**,487**	,216	,170	,185	**,468**
3	,283	**,52**	-,074	**,415**	,170	-,095	,106	**,370**	,184
4	**,410**	**,59**	,106	**,621**	-,064	,149	-,058	**,659**	,005
5	**,311**	**,57**	,145	**,535**	-,194	,183	-,076	**,567**	-,123
6	**,651**	**,63**	**,624**	,130	-,045	**,632**	-,033	,162	-,015
7	**,842**	**,34**	**,944**	-,006	-,101	**,869**	,095	-,007	-,104
8	**,689**	**,50**	**,717**	,002	-,014	**,626**	,153	-,026	-,037
9	**,666**	**,63**	**,601**	,082	,047	**,549**	,097	,070	,041
10	**,910**	**,61**	**,979**	-,060	-,006	**,976**	-,032	-,016	,014
11	**,388**	**,56**	-,136	**,430**	**,404**	-,152	,126	**,374**	**,425**
12	**,754**	**,56**	**,657**	,079	,098	**,569**	,175	,041	,074
13	**,747**	**,65**	**,524**	-,038	**,377**	**,518**	,056	-,038	**,367**
14	**,782**	**,48**	**,528**	-,071	**,454**	**,546**	,015	-,059	**,457**
15	**,585**	**,48**	,276	,181	**,306**	,036	**,524**	,014	,208
16	,248	**,55**	-,096	**,562**	,031	-,180	,213	**,465**	,046
17	**,516**	**,39**	,212	,125	**,346**	,187	,114	,094	**,327**
18	**,895**	**,62**	**,809**	,043	,119	**,704**	,207	,002	,084
19	**,530**	**,61**	,185	,291	,267	,094	,247	,211	,233
20	**,332**	**,54**	,164	**,367**	-,055	,118	,097	**,324**	-,025
21	**,681**	**,55**	**,719**	,067	-,082	**,667**	,067	,067	-,076
22	**,858**	**,78**	**,835**	,074	,004	**,845**	-,041	,113	,038
23	**,708**	**,67**	**,398**	,053	**,423**	**,384**	,093	,038	**,407**
24	**,854**	**,59**	**,950**	-,070	-,038	**,956**	-,056	-,020	-,012
25	**,749**	**,58**	**,449**	,293	,207	,073	**,782**	,066	,054
26	**,689**	**,63**	**,487**	,220	,126	,068	**,874**	-,030	-,075

27	**,626**	**,79**	**,586**	,122	-,022	**,480**	,198	,073	-,049
28	**,918**	**,72**	**,921**	,005	,029	**,853**	,105	,000	,023
29	,276	**,54**	-,026	**,489**	,032	-,121	,228	**,396**	,033
30	**,805**	**,69**	**,849**	-,044	,008	**,849**	-,032	-,006	,030

Table 4. Correlations of the BS with other measures

	BS	ASI	ASI-HS	ASI-BS	MHS-G	MHS-G-PD	MHS-G-D	MHS-G-IH	MHS-L	MHS-L-PD	MHS-L-D	MHS-L-IH	IHS-16	IHS-16-1	IHS-16-2	IHS-16-3
BS	1	.49**	.42**	.45**	–.06	–.06	–.04	–.04	–.62**	–.49**	–.46**	–.48**	.44**	.30**	.40**	.36**
ASI		1	.93**	.83**	–.03	–.01	–.01	–.03	–.51**	–.30**	–.28**	–.47**	.51**	.43**	.31**	.46**
ASI-HS			1	.56**	.01	–.01	.01	.01	–.51**	–.27**	–.27**	-.50**	.45**	.37**	.27**	.44**
ASI-BS				1	–.06	–.02	–.03	–.07	–.37**	–.25**	–.23**	–.31**	.42**	.37**	.27**	.36**
MHS-G					1	.77*	.76**	.90**	.07	.08	–.01	.05	–.19**	–.17**	–.21**	–.06
MHS-G-PD						1	.35**	.45**	.08	.12*	–.04	.06	–.03	–.03	–.08	–.03
MHS-G-D							1	.68**	–.01	–.03	–.02	.02	–.32**	–.31**	–.32**	–.13
MHS-G-IH								1	.06	.07	.03	.05	–.14	–.12	–.15*	–.06
MHS-L									1	.64**	.57**	.90**	–.40**	–.35**	–.25**	–.35**
MHS-L-PD										1	.24**	.28**	–.13	–.06	–.08	–.17*
MHS-L-D											1	.39**	–.31**	–.37**	–.13	–.21**
MHS-L-IH												1	–.38**	–.34**	–.25**	–.31**
IHS-16													1	.82**	.81**	.75**
IHS-16-1														1	.54**	.37**
IHS-16-2															1	.43**
IHS-																1