

# One-Dimensionality, Reliability, Distribution, and Correlates of the Brief Interactive Optimism Scale among University Teachers

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## Abstract

The Brief Interactive Optimism Scale (BIOS-G), recently published, was developed to measure optimism based on the interactive personality styles theory in the general population of Mexican adults. It has been validated in both the general population in Mexico and other Latin American countries. This study aims to further consolidate the validity of the BIOS-G in a specific population. The objectives are to assess the psychometric properties of its items, validate its unidimensional structure, describe its distribution, verify its internal consistency, and examine its direct relationship with affective balance and positive affect, as well as its inverse relationship with emotional exhaustion and negative affect among Mexican university teachers. The BIOS-G, the CESQT emotional exhaustion subscale, and Scale of Positive and Negative Experience (SPANE) were administered to a non-probability sample of 213 Mexican university teachers. All items demonstrated discriminative power and internal consistency. The scale showed internal consistency (Green-Yang ordinal  $\omega = 0.74$  and McDonald ordinal  $\omega = 0.88$ ). The one-factor model presented good to acceptable fit to the data, with an average variance extracted of 0.65. Its distribution was negatively skewed and mesokurtic. The BIOS-G correlated directly with SPANE's affective balance and positive emotional experiences and inversely with negative emotional experiences and emotional exhaustion, with moderate strengths of association. Its mean score was higher in men than in women and correlated positively with age. It is concluded that the BIOS-G demonstrates internal consistency and validity evidence among Mexican university teachers, and its use in this specific population is recommended.

## Keywords

Optimism, Psychometric Properties, Affective Balance, Emotional Exhaustion, University Teachers

## 1. Introduction

Optimism, as a psychological construct, has played a central role in the field of positive psychology [1] [2], as it is a disposition associated with the physical [3], mental [4], and spiritual [5] well-being of individuals, as well as with their quality of life and happiness [6].

There are two main epistemological paradigms from which this construct has been studied: one derived from the theory of positive expectations and the other from the theory of attributional styles. The theory of positive expectations, which originated in the United States, represents the most influential model or conceptual framework at present. From this perspective, optimism is conceived as a personality trait characterized by a sense of confidence that favorable outcomes will occur in the future, satisfying the individual's expectations, and that decisions made will lead to positive consequences [7].

Regarding the theory of attributional styles, these can be conceptualized as patterns used to generate explanations for events—whether positive or negative—that occur in an individual's personal and/or social environment [8]. In other words, attributional style indicates how an individual habitually explains the causes of events that happen to them [9]. The optimistic attributional style is defined as a tendency to make external, specific, and transient attributions about previous negative events, whereas attributions for previous positive events tend to be internal, stable, and global [10].

The interactive personality styles theory, which originated in Mexico [11], represents a third conceptual framework for addressing the construct of optimism. From this perspective, optimism is conceived as the result of a historical and intricate web of changing interactions between the individual and their environment, others, and themselves. In this conceptualization, optimistic individuals view humanity as intrinsically good and believe that life is beautiful, generally positive, and conducive to good outcomes. In this model of optimism, both the past and the present play equally important roles and are inevitably linked [12].

García-Cadena *et al.* [13] point out that different sociocultural contexts, which have influenced the development of theories and measurement instruments, necessarily reflect the ethnocentric convictions of their creators. In this regard, the Brief Interactive Optimism Scale by García-Cadena *et al.* (BIOS-G) [12], which is intended to be validated in this study within a specific population (university teachers), was developed for the Mexican context based on the theory of interactive personality styles. It has been demonstrated that the scale is applicable to other Spanish-speaking contexts that share Latino culture [14].

## 2. Relevance of Optimism in Organizations

Currently, many organizations have developed programs to promote the well-being of their employees [15], as the literature has documented that a company with a workforce that enjoys a high level of well-being experiences higher organizational performance, greater labor retention, lower employee turnover, and a

stronger sense of community, organizational commitment, group cohesion, loyalty, and productivity [16] [17].

In the educational field, it is widely recognized that teachers play a valuable role in the formative process of students [18]. However, it is also well known—and there is consensus in the literature—that the teaching profession is a challenging activity that can lead to distress, depression, cognitive dissonance, decreased sleep hours, poor sleep quality, dissatisfaction, and job burnout syndrome [19] [20].

Consequently, it is important to develop programs that cultivate a high state of subjective well-being in teachers. This is necessary not only because it helps them feel better about themselves and satisfied with their teaching work, but also because it revitalizes them, improves their performance in their activities, and enhances their overall well-being [21], as well as students' performance, well-being, and educational satisfaction [18].

Focusing on the teaching profession in the Mexican context, Martínez-Ramírez *et al.* [22] identified relationships among optimism, affective balance, and emotional exhaustion, with moderate to high strengths of association. These Mexican researchers emphasized the need for educational institutions to develop and promote innovative programs aimed at supporting teachers' well-being. Optimism, in particular, could act as a protective factor against emotional exhaustion. Nevertheless, further studies with larger samples, including professors from public universities in both urban and rural areas, are necessary to confirm this hypothesis. Consequently, validating and assessing the psychometric properties of the BIOS-G, a measure of optimism, is essential within the professional context of Mexican university teachers.

In this regard, Vela *et al.* [23] argue that positive psychology provides a valuable framework for understanding teacher candidates' perceptions of mental health. In an empirical study conducted with teachers at a Hispanic-Serving Institution, they found that hope—which can be viewed either as a concrete form of optimism in response to specific challenging situations or as a manifestation of the trait of optimism in such contexts—and psychological grit were significant predictors of subjective well-being.

### 3. Study Objectives

As previously noted, García-Cadena *et al.* [12] published the Brief Interactive Optimism Scale (BIOS-G) to measure optimism based on the interactive personality styles theory. Their research involved a sample of 502 participants (263 women and 239 men) drawn from the general population of Mexican adults, yielding good reliability and validity results for a unidimensional four-item scale. In a subsequent study, García-Cadena *et al.* [24] expanded the sample to 3289 participants (2028 men and 1243 women) from the general Mexican population and obtained similar results.

Additionally, García-Cadena, Díaz *et al.* [13] [25] validated the BIOS-G in comparison with the Revised Life Orientation Test (LOT-R) [26] and the Optimism-

Pessimism Short Scale-2 (SOP2) [27] among Hispanic American social work students, finding that the BIOS-G exhibited the best reliability and validity properties. In a more recent study, the scale was validated in seven Latin American countries (Mexico, Colombia, Spain, Peru, Bolivia, Ecuador, and Chile), demonstrating good factorial invariance properties. The BIOS-G shows independence from, or a weak relationship with, gender and age. When an association is observed, optimism levels tend to be higher in men than in women and increase with age [14].

To strengthen the validity of this recently introduced yet underutilized scale in applied research, the objectives of the present study are as follows: 1) to describe the distribution of its four items and total score, interpreting the central tendency of these five distributions; 2) to assess the psychometric properties of discriminability and internal consistency for the four items, as well as the overall internal consistency of the scale; 3) to validate the proposed unidimensional structure of the scale; 4) to examine its direct relationships with positive affect and affective balance, as well as its inverse relationships with emotional exhaustion and negative affect; 5) to evaluate its independence from sex and age or identify significant relationships characterized by a weak association (specifically, greater optimism in men compared to women, and increasing optimism with age); and 6) to estimate its relationship with years of teaching experience among Mexican university teachers.

This research targets a specific population that has not yet been studied using the BIOS-G, at a time when aspects of positive psychology are receiving significant attention [28]. Therefore, this validation study is particularly relevant for measuring optimism in this population using the BIOS-G.

## 4. Method

### 4.1. Participants

The BIOS-G was administered to a non-probability sample of university teachers from the School of Medicine and Health Sciences at a private university with campuses in several Mexican states. A self-administered questionnaire was distributed via the Qualtrics digital platform. Data collection took place in March 2022. The teachers were from four different campuses: Monterrey, Mexico City, Guadalajara, and Chihuahua. Participation in the study was entirely voluntary. After eliminating the teachers who left any of the scales incomplete (five cases), the sample consisted of 213 participants (106 men and 107 women).

### 4.2. Instruments of Measurement

*Brief Interactive Optimism Scale* (BIOS-G) [12]. This self-report instrument consists of four items. Participants indicate their degree of agreement or disagreement with each statement using a four-point Likert-type scale, with response options ranging from 0 to 3: 0 = of course not, 1 = I think not, 2 = I think so, and 3 = of course yes (Appendix 1). The overall level of optimism is calculated by summing the scores of the four items, with item 1 being reverse-scored. Higher total

scores indicate a greater degree of optimism. In a sample of 502 Mexican adults (263 women and 239 men), García-Cadena *et al.* (2019) [12] reported a single-factor structure with a good fit to the data using Maximum Likelihood:  $\chi^2(2, n = 502) = 1.22, p = 0.544; \chi^2/df = 0.609, GFI = 0.999, NFI = 0.999, CFI = 1, RMSEA = 0.001$  (90% CI: 0.0001, 0.077), and  $SRMR = 0.007$ . Additionally, the scale demonstrated factorial invariance between genders and equivalence of means. The internal consistency of the scale was high ( $\omega = 0.87, \alpha = 0.86$ , and ordinal  $\alpha = 0.91$ ). Furthermore, the correlations with concurrent validity criteria were strong: positive with self-esteem ( $r = 0.78, p < 0.001, 95\% CI [0.82, 0.74]$ ) and negative with depression ( $r = 0.81, p < 0.01, 95\% CI [0.84, 0.77]$ ) and psychopathy ( $r = 0.67, p < 0.001, 95\% CI [0.72, 0.61]$ ).

*Positive and Negative Experience Scale (SPANE)* [29]. This self-report instrument consists of 12 items: six related to positive affect (PA subscale) and six related to negative affect (NA subscale). The items are rated on a five-point Likert-type scale ranging from 0 to 4 (0 = never, 1 = almost never, 2 = sometimes, 3 = almost always, and 4 = always). See Appendix 1. In the original study, the scale demonstrated good overall internal consistency ( $\alpha = 0.89$ ) and in both subscales ( $\alpha = 0.87$  for PA and 0.81 for NA). It was recently validated in Mexico with a sample of 600 university students, with a mean age of 20.24 years ( $sd = 2.34$ ), of whom 57.3% were women [30]. The Mexican validation showed good internal consistency ( $\omega = 0.93$  for PA and 0.87 for NA), convergent validity ( $AVE = 0.68$  for PA and  $AVE = 0.54$  for NA), and discriminant validity ( $r^2 [AP, AN] = 0.47, p < 0.01$ ). The goodness-of-fit of the correlated two-factor model ranged from acceptable ( $\chi^2/df = 4.99, GFI = 0.929, AGFI = 0.895, NFI = 0.946, NNFI = 0.945$ ) to good ( $CFI = 0.956$  and  $SRMR = 0.040$ ) using the Maximum Likelihood method.

*Emotional Exhaustion (PE)* [31]. This is a subscale of the Questionnaire for the Evaluation of Occupational Burnout Syndrome (CESQT, in Spanish), consisting of four Likert-type items with five response options. Participants indicate how often (0 = never, 1 = seldom, 2 = sometimes, 3 = most of the time, and 4 = always) they have felt overloaded, overwhelmed, tired, or emotionally drained as a result of their work (Appendix 1). The CESQT was validated in Mexico with a sample of 698 elementary school teachers [32]. The PE subscale demonstrated skewness and kurtosis values within the normal range, and its internal consistency was high ( $\alpha = 0.82$ ). In this study, an abbreviated version with three items was used, based on the findings of a previous study [33]. The item related to emotional exhaustion was eliminated.

### 4.3. Procedure and Ethic Aspects

A descriptive-correlational study was conducted using an ex post facto cross-sectional design. Informed consent was obtained for participation, ensuring anonymity and confidentiality of the information in accordance with the ethical research standards of the American Psychological Association [34] and the Mexican Society of Psychology [35].

#### 4.4. Data Analysis

In relation to the first objective, which was to describe the distributions of the items (ordinal variables) and the BIOS-G total score (quantitative variable), measures of central tendency were employed (median and mode for the items, and the arithmetic mean for the BIOS-G total score). Additionally, measures of variation were used (semi-interquartile range and quartile coefficient of variation for the items, and sample standard deviation and Pearson's coefficient of variation for the BIOS-G total score). Measures of shape were also considered, including the quartile coefficient of skewness and percentile coefficient of kurtosis for the items, and skewness and kurtosis coefficients based on Pearson's standardized central moments for the BIOS-G total score. Symmetry was tested using D'Agostino's test [36], and mesokurtosis was assessed using Anscombe and Glynn's test [37] for the BIOS-G total score. The quantiles of the item distributions were calculated following Rule 2 (rounding) in R [38]. Bootstrap confidence intervals at the 95% level for the quartile coefficient of skewness and percentile coefficient of kurtosis were obtained using the percentile method with replacement from 1000 bootstrap samples [39]. Bootstrap confidence intervals, using the Bias-Corrected and Accelerated (BCa) percentile method, were calculated for Pearson's shape measures.

The fit of the BIOS-G distribution to a normal curve was tested using the Shapiro-Wilk  $W$  test [40] with Royston's standardization [41] and the K-squared test [42]. To interpret the scores, as well as the average, median, and mode of the BIOS-G (which is the sum of the items, with item 1 inverted, divided by the number of items), four class intervals of constant amplitude ( $a = (3 - 0)/4 = 0.75$ ) were defined between the minimum value of 0 and the maximum value of 3. These intervals correspond to the four values of the items, allowing the use of the response labels to determine the level of optimism.

In relation to the second objective, two psychometric properties of the items were estimated: discriminability and internal consistency [43]. Discriminative power was assessed by the capacity to statistically differentiate between the high-score group (scores greater than or equal to the 73rd percentile) and the low-score group (scores less than or equal to the 27th percentile) on the sum of the four items constituting the total BIOS-G score [44]. This was tested using the Mann-Whitney  $U$  test.

The internal consistency of items was assessed using three indicators: 1) a polyserial correlation of at least 0.5 between each item and the sum of the remaining items making up the scale ( $r_{i-1,i}$ ), 2) no significant decrease in the overall reliability coefficient after excluding the item, and 3) an explained variance of the item by the single-factor model of at least one-quarter, preferably at least one-half [44].

The polyserial correlation was estimated using the two-step method [45]. Reliability was assessed using ordinal omega (ordinal  $\omega_{i-1}$ ) coefficients, according to Green-Yang [46] and McDonald [47] formulas, due to non-compliance with the assumption of tau equivalence, that is, because of a congeneric measurement

model with heterogeneous measurement weights among its four indicators. Tau equivalence was tested by fitting a one-factor model with equivalent measurement weights, estimating its parameters using the Diagonally Weighted Least Squares method, also known as Weighted Least Squares with Mean and Variance adjusted (DWLS/WLSMV), which utilizes polychoric correlations [48]. To evaluate the significant decrease in reliability, bootstrap confidence intervals were calculated using the BCa method for the ordinal omega coefficients. The decrease was considered significant when the upper limit of the interval excluding the item was below the lower limit of the interval that included it.

Six indices were considered to assess model fit: chi-square goodness-of-fit statistic ( $\chi^2$ ), relative chi-square ( $\chi^2/df$ ), Bentler's Non-Normalized Fit Index (*NNFI*), Bentler-Bonnet Comparative Fit Index (*CFI*), Steiger-Lind Root Mean Square Error of Approximation (*RMSEA*), and Standardized Root Mean Square Residual (*SRMR*). The values for good, acceptable, and poor fit are shown in **Table 1** [48]-[50]. The single-factor model with four indicators was estimated using the DWLS/WLSMV method, where the standardized squared measurement weights indicate the variances of the items explained by the factor:  $(\lambda_i)^2$ .

**Table 1.** Goodness-of-fit indices for the specified model.

Indices	Null hypothesis	Good	Acceptable	Poor
<i>p</i> -value for the $\chi^2$ statistic	One-factor with four indicators	>0.05	<0.01	≤0.01
$\chi^2/df$		≤2	<5	≥5
<i>NNFI</i>		≥0.95	≥0.90	<0.90
<i>CFI</i>		≥0.95	≥0.90	<0.90
<i>p</i> -value for <i>RMSEA</i>	<i>RMSEA</i> ≤ 0.05	≥0.1	≥0.05	<0.05
	<i>RMSEA</i> ≥ 0.08	<0.1	≥0.10	>0.10
<i>SRMR</i>		<0.08	<0.10	>0.10

Internal consistency, as evaluated using McDonald's or Green-Yang ordinal omega coefficients [46] [47], was interpreted as excellent for values in the range (0.95, 0.90], good for (0.90, 0.80], acceptable for (0.80, 0.70], questionable for (0.70, 0.60], poor for (0.60, 0.50], and unacceptable for values below 0.50. Values equal to or greater than 0.95 were deemed undesirable, as they suggest item redundancy [43] [51].

In relation to the third objective, the validation of the dimensional structure was carried out in two ways. First, the number of factors was determined using the matrix of polychoric correlations, estimated via the two-step method. Seven different analyses were employed to determine the number of factors: parallel analysis of the eigenvalues [52], comparison data method [53], minimum average partial correlation squared [54] and raised to the fourth power [55], the Kaiser empirical criterion [56], and the standard error of predicted eigenvalues in the scree test, where values greater than the inverse of the number of items were

considered significant [57]. The value on which the analyses converged was sought [43] [58]. Second, confirmatory factor analysis was conducted. The method used to minimize the discrepancy function was DWLS/WLSMV. The goodness-of-fit indices and their interpretation are shown in **Table 1**.

From the standardized measurement weights, the Average Variance Extracted (*AVE*) [59] was calculated. The overall ordinal omega coefficient, computed using the Green-Yang approach based on equivalent parallel forms (Green-Yang ordinal  $\omega$ ) and the McDonald approach based on standardized measurement weights (McDonald ordinal  $\omega$ ), is considered good when it is equal to or greater than 0.80, indicating that 80% of the empirical scores obtained through the measurement instrument are free from error, and acceptable when it is equal to or greater than 0.70 [43]. The *AVE* was regarded as evidence of the convergent validity of a latent variable, as it measures the amount of variance a latent variable explains in the items [59]. A combination of  $AVE > 0.50$  and ordinal  $\omega \geq 0.70$  reflects construct validity in the measurement model [60].

In relation to the fourth objective, the correlations between BIOS-G total score and the concurrent validity criteria (affective balance, AP and AN of the SPANE, and DES) were calculated using Spearman's rank-order correlation coefficient ( $r_s$ ). Asymptotic confidence intervals were derived via z-transformation [61]. The strength of the association was interpreted as follows: trivial for  $r_s$  values less than 0.10, small for  $r_s$  values between 0.10 and 0.29, medium for values between 0.30 and 0.49, large for values between 0.50 and 0.69, very large for values between 0.70 and 0.89, nearly perfect for values between 0.90 and 0.99, and perfect for a value of 1 [62].

To test the differences in means between sexes for the fifth objective, Student's t-test for independent samples was used, along with a 95% BCa confidence interval for the mean difference. The assumption of homogeneity of variance was tested with Levene's test. Effect size was calculated using Cohen's d statistic, with the following interpretation: values of d less than 0.10 were considered to reflect a trivial effect size, values within the interval [0.10, 0.20) a very small effect, [0.20, 0.50) a small effect, [0.50, 0.80) a medium effect, and values greater than or equal to 0.80 a large effect [63]. However, new cutoff points (0.15, 0.36, and 0.65) have been proposed following a review of 134 meta-analyses [64].

Continuing with the fifth objective, the correlation between the BIOS-G total score and age was computed using Pearson's product-moment correlation coefficient, with confidence intervals estimated using the Bias-Corrected and Accelerated (BCa) percentile method, drawing 1000 random samples with replacement (bootstrap). The same approach was applied to examine the correlation between the BIOS-G total score and years of teaching experience, addressing the sixth objective. The significance level was set at 0.05.

To compute the polychoric and polyserial correlations, the Green-Yang ordinal omega, McDonald ordinal omega coefficients, and their bootstrap confidence intervals; calculate alpha ordinal and its asymptotic confidence interval; determine the number of factors; and perform factor analysis, R version 4.4.1 was used with

the following libraries: polycor, misty, MBESS [65], sjpierce/piercer [66], EFA, dimensions [57], psych, lavaan [48], and semPlot. Excel 2021 for Windows was used to calculate the Fornell-Larcker *AVE* [59] and the asymptotic confidence intervals for the Spearman correlations [61]. All other calculations were performed using SPSS version 29.

## 5. Results

### 5.1. Description of the Sample

The sample consisted of 106 men (49.8%) and 107 women (50.2%), with no significant difference between sexes according to the binomial test (exact two-tailed  $p = 1$ ). Regarding education level, 2 out of the 213 participants (0.9%) held a bachelor's degree, 151 (70.9%) held a master's degree or medical specialty, and 60 (28.2%) held a doctoral degree.

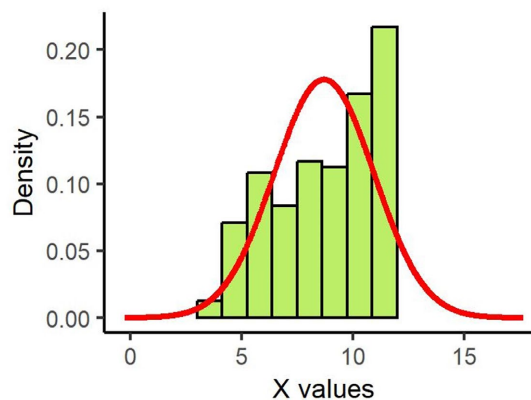
The mean age of the 213 participants was 44.39 years (95% BCa *CI* [43.10, 45.99]), and the median was 43 years (95% BCa *CI* [40, 44]). The mode, according to Grenander's procedure for unimodal continuous variables, was 38.51 years (95% BCa *CI* [36.55, 39.95]), and the most frequent value was 38 years ( $f_{max} = 0.06$ ), which was only modal value. Age ranges from 25 to 74 years, with a sample standard deviation of 10.34. The Pearson's coefficient of variation was 23.3, indicating small dispersion. The age distribution was positively skewed ( $\sqrt{b_1} = 0.51$ , 95% BCa *CI* [0.28, 0.74]; D'Agostino skewness test:  $z(\sqrt{b_1}) = 2.97$ , two-tailed  $p = 0.003$ ) and platykurtic, or with thinned tails ( $b_2 = 2.33$ , 95% BCa *CI* [1.96, 2.87]; Anscombe-Glynn kurtosis test:  $z(b_2) = -2.84$ , two-tailed  $p = 0.004$ ). It deviated from normality according to the Shapiro-Wilk test with Royston's standardization ( $w = 0.95$ ,  $z_w = 4.59$ , right-tailed  $p < 0.001$ ) and the D'Agostino-Berlanger-D'Agostino test ( $k^2 = 16.93$ , right-tailed  $p < 0.001$ ).

The mean number of years of teaching experience was 12.66 years (95% BCa *CI* [11.39, 14.11]), and the median was 10 years (95% BCa *CI* [7.79, 10]). The Grenander's mode was 5.51 years (95% BCa *CI* [4.50, 6.01]), and the most frequent value was 2 years ( $f_{max} = 0.08$ ), which was the only modal value. The number of years of teaching experience ranged from 0 to 54 years, with a sample standard deviation of 9.73. The Pearson coefficient of variation was 76.8, indicating a wide dispersion. The distribution was positively skewed ( $\sqrt{b_1} = 1.01$ , 95% BCa *CI* [0.69, 1.58]; D'Agostino skewness test:  $z(\sqrt{b_1}) = 5.30$ , two-tailed  $p < 0.001$ ) and leptokurtic, or with heavy tails ( $b_2 = 3.84$ , 95% BCa *CI* [2.56, 6.55]), as indicated by Anscombe-Glynn kurtosis test ( $z(b_2) = 2.16$ , two-tailed  $p = 0.031$ ), deviating from normality according to the Shapiro-Wilk test with Royston's standardization ( $w = 0.91$ ,  $z_w = 6.11$ , right-tailed  $p < 0.001$ ) and the D'Agostino-Berlanger-D'Agostino test ( $k^2 = 32.74$ , right-tailed  $p < 0.001$ ).

### 5.2. Description of the Distribution of BIOS-G and Its Items and Interpretation of Their Scores

The mean score on the BIOS-G was 8.71 (95% BCa *CI* [8.40, 9]), bootstrap standard

error ( $bse$ ) = 0.16), and the median was 9 (95% BCa  $CI$  [8, 9],  $bse$  = 0.28). The Grenander's mode was 9.61 (95% BCa  $CI$  [8.58, 10.50],  $bse$  = 0.47), and the most frequent value was 10 ( $f_{\max}$  = 0.19), which was the only modal value. The BIOS scores ranged from 3 to 12, with a standard deviation of 2.24. The Pearson coefficient of variation was 25.7 (95% BCa  $CI$  [23.44, 28.10],  $bse$  = 1.17). The distribution was symmetrical by the D'Agostino skewness test with a significance level of 0.05, but negatively skewed with a significance level of 0.1 by this test ( $z(\sqrt{b_1})$  = -1.69, two-tailed  $p$  = 0.090). Additionally, the 95% BCa bootstrap confidence interval revealed negative skewness ( $\sqrt{b_1}$  = -0.28, 95% BCa  $CI$  [-0.51, -0.06],  $bse$  = 0.11). On the other hand, it was platykurtic, or with heavy tails ( $b_2$  = 2.10, 95% BCa  $CI$  [1.86, 2.52],  $bse$  = 0.16); Anscombe-Glynn kurtosis test ( $z(b_2)$  = -4.84, two-tailed  $p$  < 0.001), deviating from normality according to the Shapiro-Wilk test with Royston's standardization ( $w$  = 0.94,  $z_w$  = 5.05, right-tailed  $p$  < 0.001) and the D'Agostino-Berlanger-D'Agostino test ( $k^2$  = 26.27, right-tailed  $p$  = 0.001). See **Figure 1**.



**Figure 1.** Friedman-Diaconis histogram with overlaid normal curve.

By considering four intervals of constant amplitude ( $a = (max - min)/\text{number of discrete values} = (3 - 0)/4 = 0.75$ ), the scores, mean, median, and mode of the BIOS-G can be interpreted using the item response labels. An average in the interval [0, 0.75) corresponds to a discrete value of 0 (clearly not or very pessimistic), in the interval [0.75, 1.5) corresponds to 1 (I think not or pessimistic), in the interval [1.5, 2.25) corresponds to 2 (I think so or optimistic), and in the interval [2.25, 3] corresponds to 3 (clearly yes or very optimistic). The average response on the scale indicates an optimistic level, since the mean ( $m = 2.18$ ) falling within the range [1.5, 2.25), while the median and modal responses (2.25 and 2.40, respectively) correspond to a very optimistic level [2.25, 3].

The mode of reverse item 1 ("Life is NOT ugly") was 3, which is the maximum value and corresponds to a very optimistic level of response. The median was 2, reflecting an optimistic level of response. Although the modal value at the upper end of the distribution accounted for 30.5% of the frequency, the item cannot be considered to exhibit a ceiling effect, as the null hypothesis of a uniform distribution

was not rejected, assuming a probability of 1/4 for each ordered response category ( $\chi^2 [3] = 3.99$ , right-tailed  $p = 0.260$ ). The scores ranged from 0 to 3, with a semi-interquartile range of 1. The quartile coefficient of variation was 0.5. The frequency distribution exhibited symmetry ( $qcs = 0$ , 90% percentile  $CI [-0.33, 0.33]$ ,  $bse = 0.24$ ). A two-tailed bootstrap probability value for the null hypothesis of symmetry ( $H_0: QCS = 0$ ) was 1, supporting the null hypothesis. The profile of the bootstrap sampling distribution was platykurtic or shortened-tailed ( $pck_0 = 0.07$ , 90% percentile  $CI [0.07, 0.24]$ ,  $bse = 0.08$ ). The two-tailed bootstrap probability value for the null hypothesis of mesokurtosis ( $H_0: PCK_0 = 0$ ) was below the significance level (two-tailed  $p = 0.028 < \alpha = 0.1$ ), leading to the rejection of the null hypothesis. See **Table 2**.

**Table 2.** Frequency and descriptive statistics of the items.

Item	Relative frequency				Descriptive statistics					
	1	2	3	4	<i>mo</i>	<i>mdn</i>	<i>sir</i>	<i>qvc</i>	<i>qcs</i>	<i>pck<sub>0</sub></i>
1	0.244	0.211	0.239	0.305	3	2	1	0.5	0 [-0.333, 0.333]	0.070 [0.070, 0.237]
2	0	0.014	0.408	0.577	3	3	0.5	0.2	-1 [-1, -1]	0.237 <i>bse</i> = 0
3	0	0.033	0.413	0.554	3	3	0.5	0.2	-1 [-1, 1]	0.237 <i>bse</i> = 0
4	.01	0.014	0.183	0.573	2	2	0	0	0 [0, 1]	-0.263 [-0.263, -0.013]

Note. Items: 1 = life is NOT ugly (being reverse-scored), 2 = life is beautiful, 3 = life is good, and 4 = most people are good. Descriptive statistics: *mo* = mode, *mdn* = median, *sir* = semi-quartile range, *qvc* = quartile coefficient of variation, *qcs* = Bowley's quartile coefficient of skewness [percentile bootstrap confidence interval at 90%], and *pck<sub>0</sub>* = Kelley's percentile coefficient of kurtosis centered at 0 [percentile bootstrap confidence interval at 90%].

The median and mode of item 2 ("Life is beautiful") coincided at 3, indicating that the average response corresponds to a very optimistic level. The item presented a ceiling effect, as the maximum value of 3 concentrated 57.7% of the frequency distribution ( $>1/4 \times 100$ ). The scores ranged from 1 to 3, with a semi-interquartile range of 0.5. No participant opted for the first response option. The quartile coefficient of variation was 0.2. The frequency distribution exhibited negative skewness ( $qcs = -1$ , 90% percentile  $[-1, -1]$ ,  $bse = 0.26$ ), indicating that the left side of the distribution is less concentrated or longer than the right (50th percentile - 25th percentile =  $3 - 2 = 1$  versus 75th percentile - 50th percentile =  $3 - 3 = 0$ ). Thus, when the left-tailed bootstrap probability was calculated for the null hypothesis of extreme negative skewness ( $H_0: QCS \leq -1$ ), the null hypothesis was upheld at the 10% significance level ( $p = 0.983 > \alpha = 0.1$ ). Consistently, two-tailed bootstrap probability value for null hypothesis of symmetry ( $H_0: QCS = 0$ ) was 0.034 ( $< \alpha = 0.1$ ). The profile of the bootstrap sampling distribution was platykurtic or shortened-tailed ( $pck_0 = 0.24$ ,  $bse = 0$ ). Accordingly, the right-tailed bootstrap

probability value for the null hypothesis of extreme platykurtosis ( $H_0: PCK_0 \geq 0.24$ ) was 1, while the two-tailed bootstrap probability value for hypothesis of mesokurtosis ( $H_0: PCK_0 = 0$ ) was 0. See **Table 2**.

The median and mode of item 3 (“Life is good”) are 3, indicating that the average response corresponds to a very optimistic level. This item exhibited a ceiling effect, as the maximum value of 3 accounted for more than a quarter of the frequency distribution. Scores ranged from 1 to 3, with a semi-interquartile range of 0.5. As in item 2, the first response option received no responses. The quartile coefficient of variation was 0.2. The frequency distribution exhibited negative skewness ( $qcs = -1$ ), indicating that the left side is less concentrated or shorter than the right side (50th percentile - 25th percentile =  $2 - 3 = 1$ , compared to 75th percentile - 50th percentile =  $3 - 3 = 0$ ). It is worth noting that the 90% bootstrap confidence interval spanned the entire potential range of the coefficient  $[-1, 1]$ , due to the high bootstrap standard error ( $bse = 0.46$ ). This occurred because the bootstrap sampling distribution of the QCS statistic included only two values: -1 ( $f = 0.94$ ) and 1 ( $f = 0.06$ ). Consequently, the one-tailed bootstrap probability for the null hypothesis of extreme negative skewness ( $H_0: QCS \leq -1$ ) supported the null hypothesis at the 10% significance level (one-tailed  $p = 0.943 > \alpha = 0.1$ ), whereas the two-tailed bootstrap probability for the null hypothesis of symmetry ( $H_0: QCS = 0$ ) was 0.114, a much lower value. The distribution profile was platykurtic or shortened-tailed ( $pck_0 = 0.24$ ,  $bse = 0$ ). Accordingly, the right-tailed bootstrap probability for the null hypothesis of extreme platykurtosis ( $H_0: PCK_0 \geq 0.24$ ) was 1, while the two-tailed bootstrap probability for the null hypothesis of mesokurtosis ( $H_0: PCK_0 = 0$ ) was 0. See **Table 2**.

The median and mode of item 4 (“Most people are good”) coincided at 2, indicating that the average response corresponds to an optimistic level. The item did not present a ceiling effect, as the maximum value concentrated less than a quarter of the frequency distribution, although the median covered 57.3% of the distribution. The semi-interquartile range and the quartile coefficient of variation were null, suggesting that there is no variability. However, scores ranged from 0 to 3, and when the semi-percentile rank ( $spr = (90\text{th percentile} - 10\text{th percentile})/2 = 1$ ) and the percentile coefficient of variation ( $pcv = (90\text{th percentile} - 10\text{th percentile})/(10\text{th percentile} + 90\text{th percentile}) = 0.5$ ) were calculated, a variability similar to the previous items was observed. The frequency distribution showed symmetry ( $qcs = 0$ , 90% percentile CI  $[0, 1]$ ,  $bse = 0.45$ ), revealing that the two sides of the distribution are equal (50th percentile - 25th percentile =  $2 - 2 = 0$  versus 75th percentile - 50th percentile =  $2 - 2 = 0$ ). Two-tailed bootstrap probability value for null hypothesis of symmetry was 1. The profile of the bootstrap sampling distribution was leptokurtic or elongated-tailed ( $pck_0 = -0.26$ , 90% percentile  $[-0.26, -0.01]$ ,  $bse = 0.11$ ). The right-tailed bootstrap probability value for the null hypothesis of extreme leptokurtosis ( $H_0: PCK_0 \leq -0.26$ ) was 0.76, while the two-tailed bootstrap probability value for null hypothesis of mesokurtosis ( $H_0: PCK_0 = 0$ ) was 0. See **Table 2**.

### 5.3. Discriminability and Internal Consistency

The four items that make up BIOS-G demonstrated discriminative power, as each item was able to significantly differentiate between the high ( $\leq 73$ rd percentile = 10) and low ( $\geq 27$ th percentile = 7) scoring groups, the items 1 and 2 with a mean absolute difference greater than 1 and items 3 and 4 greater than 0.91 (Table 3).

**Table 3.** Discriminability and internal consistency of items.

It.	<i>mdn</i>	Mann-Whitney			Polyserial correlation		H <sub>0</sub> : $\rho = 0$		H <sub>0</sub> : $\rho \geq 0.5$	
		<i>U</i>	<i>z</i>	<i>p</i>	<i>r<sub>i,t-i</sub></i> [95% <i>CI</i> ]	<i>se</i>	<i>z</i>	<i>p</i>	<i>z</i>	<i>p</i>
1	2.050	292	-10.141	<0.001	0.455 [0.345, 0.553]	0.053	8.559	<0.001	-0.874	0.191
2	1.030	46	-12.290	<0.001	0.756 [0.701, 0.801]	0.026	29.489	<0.001	7.310	>0.999
3	0.920	502	-10.357	<0.001	0.572 [0.484, 0.649]	0.042	13.602	<0.001	1.621	0.948
4	0.913	1072	-7.485	<0.001	0.429 [0.315, 0.531]	0.055	7.774	<0.001	-1.336	0.091

Note. It. = Items: 1 = Life is NOT ugly (being reverse-scored), 2 = Life is beautiful, 3 = Life is good, and 4 = Most people are good. |*mdn*| = absolute mean difference between groups of high and low scores on the scale. Mann-Whitney test: *U* = test statistic for the difference in central tendency between the two groups, *z* = standardized test statistic, and *p* = two-tailed asymptotic probability value. Polyserial correlation of the item with the rest of the scale:  $r_{i,t-i}$  [95% *CI*] =  $[\tanh(\operatorname{atanh}(r) \mp z_{1-\alpha/2} \times se/(1 - r^2))] =$  point estimate and 95 % asymptotic confidence interval, *se* = asymptotic standard error,  $z = (1 - r^2) \times (\operatorname{atanh}(r) - \operatorname{atanh}(0.5))/se =$  test statistic, *p* = probability value for null hypothesis of  $\rho = 0$ ,  $z = (1 - r^2) \times (\operatorname{atanh}(r) - \operatorname{atanh}(0.5))/se =$  test statistic for the hypothesis for H<sub>0</sub>:  $\rho \geq 0.5$ , *p* = left-tailed asymptotic probability value.

The polyserial correlations of each item with the sum of the remaining items ranged from 0.43 to 0.86 with a mean of 0.55. Although items 1 and 4 had correlations below 0.5, the null hypothesis of a correlation greater than or equal to 0.5 for all four items was retained (Table 3).

The tau-equivalence assumption among the four items was tested by fitting a one-factor model with all measurement weights set as equivalent. The null hypothesis of goodness-of-fit was rejected:  $\chi^2$  [5, *n* = 213] = 90.58, *p* < 0.001;  $\chi^2/df = 90.58/5 = 18.12$ ; *RMSEA* = 0.284 (90% *CI* [0.235, 0.337], *p*-value < 0.001 for H<sub>0</sub>: *RMSEA*  $\leq$  0.05), and *SRMR* = 0.179 > 0.10.

These results indicate that the measurement model is congeneric, with heterogeneous factor loadings. Consequently, the ordinal omega coefficient is chosen for calculating reliability.

The overall internal consistency was acceptable, with a Green-Yang ordinal  $\omega$  of 0.74 (95% BCa *CI* [0.68, 0.79]). The removal of item 1 resulted in a non-significant increase in this value (Green-Yang ordinal  $\omega = 0.77$ , 95% BCa *CI* [0.70, 0.82]), while the removal of item 2 led to a significant decrease (Green-Yang ordinal  $\omega = 0.56$ , 95% BCa *CI* [0.47, 0.65]). The removal of the other two items also resulted in a non-significant decrease (Table 4).

**Table 4.** Internal consistency of items.

Scale	Green-Yang ordinal $\omega$	McDonald ordinal $\omega$	Ordinal alfa	$\lambda^2$
Without item 1	0.771 [0.700, 0.821]	0.878 [0.821, 0.906]	0.864 [0.829, 0.893]	0.4994
Without item 2	0.562 [0.466, 0.646]	0.877 [0.833, 0.899]	0.709 [0.634, 0.771]	0.9999
Without item 3	0.651 [0.541, 0.745]	0.877 [0.833, 0.899]	0.794 [0.741, 0.838]	0.7406
Without item 4	0.690 [0.599, 0.755]	0.904 [0.866, 0.921]	0.868 [0.834, 0.896]	0.3575
BIOS-G total score	0.737 [0.676, 0.786]	0.877 [0.833, 0.899]	0.826 [0.784, 0.861]	

Note. Ordinal  $\omega$  = ordinal omega coefficients by Green and Yang [46] and McDonald [47]; their confidence intervals were obtained by bootstrap, using the BCa percentile method with the extraction of 1000 samples. Ordinal alfa and its approximate confidence interval using the Feldt method [67].  $\lambda^2$  = explained variance of the item by the single factor estimated using the WLSMV method (with two measurement weight specified for the solution was admissible): Optimism  $\approx -0.60 * O1 + 0.849 * O2 + O3 + O4$ .

When the ordinal omega coefficient was calculated using the measurement weights from the one-factor model estimated with the DWLS/WLSMV method and the formula provided by McDonald [47], good overall consistency was obtained (McDonald ordinal  $\omega = 0.88$ , 95% BCa *CI*[0.83, 0.90]). This value increased non-significantly with the removal of item 4 (McDonald ordinal  $\omega = 0.90$ , 95% BCa *CI*[0.87, 0.92]) and of item 1 (McDonald ordinal  $\omega = 0.88$ , 95% BCa *CI*[0.82, 0.91]). The removal of the other two items did not result in any change. See **Table 4** for details.

On the other hand, if the ordinal alpha coefficient were used [51], the overall reliability would also be good (ordinal  $\alpha = 0.864$ , 95% Feldt-type [67] approximate *CI* [0.829, 0.893]). Removing items 1 or 4 would result in a non-significant increase in this value, while removing items 2 and 3 would decrease it—item 2 significantly (0.709, 95% Feldt-type approximate *CI* [0.634, 0.771]) and item 3 non-significantly. See **Table 4**. However, due to non-compliance with the tau-equivalence assumption, it would be inadequate to estimate the proportion of true variance explained (reliability) by this coefficient.

The communality, or the variance explained by the factor, was greater than one-third for each item and exceeded seven-tenths for items 2 and 3 (**Table 4**).

#### 5.4. Concurrent Validity

The BIOS-G showed one-dimensionality by the Horn's parallel analysis, the Ruscio-Roche comparison data analysis, the Braeken-van Assen empirical Kaiser criterion, the Velicer's minimum average of the partial correlations squared and to the fourth power, as well as by the criterion of a standard error in the predicted eigenvalue in the sedimentation plot greater than 0.25 or inverse of the number of items.

The average polychoric correlations among the four ordinal items was 0.59, ranging from 0.34 (between items 1 and 4) to 0.87 (between items 2 and 3). All

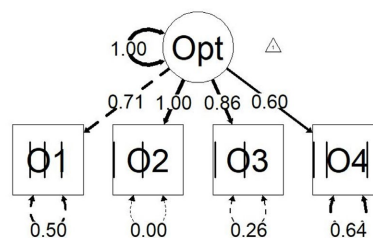
correlations were significant (Table 5). The one-factor model with four indicators, estimated by the WLSMV method using polychoric correlations, showed a good fit according to six indices:  $\chi^2(2, n = 213) = 2.82, p = 0.244 > 0.05; \chi^2/df = 1.41 < 2; NNFI = 0.999$  and  $CFI = 1 > 0.95; SRMR = 0.035 < 0.05$ ; and  $RMSEA = 0.044$  (90% CI [0, 0.151],  $p = 0.411 > 0.1$  for  $H_0: RMSEA = 0.05; p = 0.385$  for  $H_0: RMSEA \geq 0.080$ ). The average variance extracted was greater than 0.5 ( $AVE = 0.65$ ) and composite reliability was greater than 0.8 (McDonald  $\omega = 0.87$ , 95% BCa CI [0.83, 0.90]). However, the solution was not admissible since the standardized measurement weight of item 2 was greater than one.

**Table 5.** Polychoric correlations and estimated model parameters.

It.	Polychoric correlations			CFA					
	1	2	3	<i>L</i>	<i>se</i>	<i>z</i>	<i>p</i>	<i>l</i>	<i>s</i> <sup>2</sup> (res)
1				0.600				0.707	0.501
2	0.732*** [0.049, 0.620]			0.849				0.99996	<0.001
3	0.459*** [0.070, 0.311]	0.870*** [0.030, 0.797]		0.731	0.057	12.740	<0.001	0.861	0.259
4	0.336*** [0.073, 0.187]	0.617*** [0.064, 0.477]	0.549*** [0.066, 0.406]	0.508	0.054	9.401	<0.001	0.598	0.643

Note. Polychoric correlations: point estimate and 95% asymptotic confidence interval and z-test for significance.: ns  $p > 0.05$ , \*  $p < 0.01$ , \*\*  $p < 0.01$ , \*\*\*  $p < 0.001$ . Confirmatory factor analysis: *L* = unstandardized measurement weight (value of 1 indicates that the weight value was identified or constrained to 1 to make the model identifiable), *se* = standard error of measurement weight, *z* = test statistic, *p* = two-tailed probability, *l* = standardized measurement weight, *s*<sup>2</sup>(res) = 1 - *f*<sup>2</sup> = variance of the standardized residual. The polychoric correlations were estimated using the two-step method.

To overcome this problem, due to the collinearity between items 2 and 3, a second parameter was specified. Apart from the measurement weight of item 1, specified as 0.6, the measurement weight of item 2 was specified with a value of 0.849. The factor loadings or standardized measure weights ranged from 0.60 to 1 (Figure 2). All the parameters were significant (Table 5), including factor variance:  $s^2(\text{Factor}) = 1.39, se = 0.11, z = 12.90, p < 0.001$ . Refer to Appendix 2 for the computation of the last model using R. The script allows displaying the graph of the factor model in the console or saving it as a JPEG file.



**Figure 2.** Estimates in the standardized model. The dotted line indicates which parameter was identified in the estimation by the WLSMV method, using polychoric correlations.

The average variance extracted was greater than 0.5 ( $AVE = 0.65$ ) and composite reliability was greater than 0.8 (McDonald  $\omega = 0.88$ , 95% BCa  $CI [0.83, 0.90]$ ). After reviewing the goodness-of-fit indices, the scale showed a good fit according to three of six indices:  $\chi^2 (3, n = 213) = 6.67, p = 0.083 > 0.05$ ;  $NNFI = 0.998$  and  $CFI = 0.996 > 0.95$ . It was acceptable according to three indices:  $\chi^2/df = 2.22 < 3$ ;  $SRMR = 0.056 < 0.08$ ; and  $RMSEA = 0.076$  (90%  $CI [0, 0.155]$ ,  $p = 0.225 > 0.1$  for  $H_0$ :  $RMSEA = 0.05$ ;  $p = 0.545$  for  $H_0$ :  $RMSEA \geq 0.080$ ). Although the parsimony of the model is low ( $PR = 0.3$ ), the model is not overidentified, as the relative chi-square is greater than one.

### 5.5. Concurrent Validity

The BIOS-G total score showed direct correlations with the SPANE affective balance (AB) and SPANE positive affect (PA) subscale, the former with a medium strength of association and latter large. Likewise, the BIOS-G had negative correlations with the SPANE negative affect (NA) subscale and the CESQT emotional exhaustion subscale, both a medium strength of association (Table 6). Thus, evidence of convergent and divergent validity is provided.

**Table 6.** Correlates of BIOS-G.

Statistics	AB	PA	NA	PW
$r_s$	0.482	0.453	-0.401	-0.371
95% $CI$	[0.325, 0.585]	[0.299, 0.560]	[-0.497, -0.273]	[-0.475, -0.240]
$z$	7.110	6.608	-5.747	-5.270
$p$	<0.001	<0.001	<0.001	<0.001

Note.  $r_s$  = Spearman's rank-order correlation coefficient, 95%  $CI$  = asymptotic 95% confidence interval calculated via z-transformation. Significance test:  $z$  = test statistic, and  $p$  = two-tailed probability value. AB = SPANE affective balance, PA = SPANE positive affect subscale, and NA = SPANE negative affect subscale. PW = CESQT emotional exhaustion subscale.

### 5.6. Relationship of BIOS-G with Sex, Age, and Teaching Experience

The mean BIOS-G total score for females ( $m = 8.39$ , 95% BCa  $CI [7.94, 8.84]$ ) was significantly lower (mean difference:  $md = 0.64$ , 95% BCa  $CI [0.06, 1.17]$ ;  $t [211] = 2.09, p = 0.038$ ) than that for males ( $m = 9.03$ , 95% BCa  $CI [8.62, 9.41]$ ). However, their variances were statistically equivalent (Levene:  $F = 0.17, p = 0.677$ ). The effect size of sex on optimism was small (Cohen's  $d = 0.286$ , 95%  $CI [0.016, 0.556]$  and Hedges-Olkin  $g = 0.285$ , 95%  $CI [0.015, 0.555]$ ) according to both Cohen [63] and Lovakov-Agadullina [64] criteria.

The BIOS-G total score was significantly correlated with age ( $r = 0.17$ , 95% BCa  $CI [0.04, 0.31]$ ) and years of teaching experience ( $r = 0.14$ , 95% BCa  $CI [0.003, 0.29]$ ). When partializing the effect of age, the correlation between optimism and psychological burnout remains significant (Fisher's  $r_p = -0.34$ , 95% BCa  $CI [-0.45, -0.22]$ ). The same holds true when partializing for teaching experience

(Fisher's  $r_p = -0.36$ , 95% BCa CI  $[-0.47, -0.25]$ ) or for both experiential variables simultaneously (Fisher's  $r_p = -0.35$ , 95% BCa CI  $[-0.47, -0.21]$ ).

## 6. Discussion

In relation to the first objective of the study, which was to describe the distribution of the scale and its four constituent items and interpret the central tendency of these five distributions, the scale has a steeply ascending staircase distribution, *i.e.*, a profile with negative skewness and platykurtosis, deviating from normality. This distribution is characteristic of personal or socially desirable traits [68]. The distributions of items 2 and 3 are more similar to that of the scale. These items exhibit a ceiling effect, with high concentration at the upper values, very low concentration at the lower values (negative skewness), zero frequency for the first ordered category of response, reduced variability, and shortened tails, lacking a right tail. Both items are very similar in content, as they request a global assessment of one's life, and are positively keyed. Items 1 and 4 have distinct distributions. Neither shows a ceiling effect; the frequency of the first ordered category is not zero, and both are symmetrical. Item 1, which is the negatively keyed item, has a uniform distribution (platykurtic), probably indicating some confusion in the answers, while item 4, which differs in content from the other three by asking for an overall assessment of the majority of persons, has an elongated-tailed distribution (leptokurtic), indicating a more measured or less extreme assessment.

Again, the central tendency of the scale, assessed using the median and mode, is more similar to items 2 and 3 than to items 4 and 1. The central tendency for items 2 and 3 corresponds to a very optimistic view. These items request an overall and direct evaluation in aesthetic terms (beautiful) or hedonic terms (good). The central tendency of item 1 is between very optimistic (mode) and optimistic (median). This item is negatively keyed and requests an overall evaluation in terms of aesthetics (ugly). The central tendency of item 4 corresponds to optimistic. This item uses a moral adjective (good) to qualify life in most people. On the one hand, overall assessments elevate the central tendency [44]. On the other hand, aesthetic or hedonic adjectives, which connote desire, facilitate a more optimistic evaluation than moral adjectives, which connote what "ought to be" [69] [70].

Notably, these non-normal distributions—characterized by positive skewness and platykurtosis for the scale and items 2 and 3, uniformity for item 1, and symmetry with leptokurtosis for item 4—are also observed in the general population [12] [14] [24] and among university social work students [13]. The central tendency of the BIOS-G, assessed by the median and mode as very optimistic and by the mean as optimistic, is consistent with findings from the aforementioned studies conducted across various populations.

The identification of a ceiling effect in items 2 and 3 is a finding that has not been emphasized in previous studies on the BIOS-G. This distributional characteristic limits the variability of the total score but does not invalidate it. Moreover, as evidenced by data from this and prior research, it does not significantly impact

the scale's reliability.

In relation to the second objective of testing the discriminability and internal consistency of the items, it was found that all four items discriminate effectively for the total scale score, even though three of them exhibit a ceiling effect. Although this effect is typically seen as a negative trait in items, particularly in performance and ability tests [44], it is not detrimental in this trait scale. The overall internal consistency of the scale is good, ranging from 0.80 to 0.89 across McDonald ordinal omega and ordinal alpha, consistent with prior studies [12]-[14] [24], and from 0.70 to 0.79 using Green-Yang ordinal omega, suggesting that the item content is ultimately not redundant despite its high similarity [51]. Given that the assumption of tau-equivalence among ordinal variables is not met, ordinal omega is preferable to ordinal alpha, as well as to hierarchical omega and Cronbach's alpha coefficients [43] [60] [65].

The internal consistency of the four items is satisfactory, as each item's correlation with the sum of the other items is statistically equal to or greater than 0.5, and the removal of any item does not significantly increase the Green-Yang ordinal omega coefficient. Additionally, a single factor accounts for more than one-third of the item variances [44]. However, according to the McDonald ordinal omega, the removal of item 4 does increase overall internal consistency. This is because item 4 has the lowest correlations with the other items, as it differs most in content, although it contributes significantly to the one-factor model with a standardized measurement weight of 0.6. Consequently, no item warrants elimination.

In relation to the third objective of validating the unidimensional structure, the hypothesis is confirmed [12] [14]. The empirical criteria for determining the number of factors converge on 1, even in the Ruscio-Roche comparative analysis, which typically overestimates the number of factors when the number of indicators is low [57]. Factor loadings are significant and equal to or above 0.6. Additionally, the goodness-of-fit indices range from good to acceptable, with the model supported by the chi-square test, and the relative chi-square greater than 1, showing the model is not overdetermined [49]. Furthermore, indicators of convergent validity ( $AVE > 0.50$ ) and composite reliability (Green-Yang ordinal  $\omega > 0.70$  and McDonald ordinal  $\omega > 0.80$ ) reach desirable values, indicating construct validity when considered in combination [60]. These results are consistent with previous studies [14]. It should be noted that the factor loadings are not homogeneous, so the measurement model is congeneric [65].

In this sample of university teachers, the one-factor model reveals a collinearity problem between items 1 and 2, which are opposites of each other. However, item 1 is likely more confusing in its response due to being a negatively keyed item. This collinearity issue was satisfactorily resolved by specifying measurement weights (other than 1) for these two items—0.6 for item 1 and 0.849 for item 2. This problem could be seen as a weakness of the scale, suggesting the need to revise item 1. In its revised version, item 1 could request an overall positive

assessment, similar to items 2 and 3. Thus, its new wording could be: “Life always offers opportunities to move forward”. Notably, the content of this proposed item aligns with the interactive personality styles theory, which suggests that optimistic individuals perceive humanity as intrinsically good and view life as beautiful, generally positive, and conducive to favorable outcomes [12]. It is worth noting that the issue of collinearity between items 1 and 2 emerged as a finding of this study conducted among teachers; it was neither anticipated nor an established fact. This observation has led to the proposed rewording of the first item of the BIOS-G.

In relation to the fourth objective concerning concurrent validity, the inverse relationship with emotional exhaustion and negative affect, as well as the direct relationship with affective balance and positive affect, were confirmed. Correlations with positive traits are higher than those with negative traits, although these differences are not significant, as their confidence intervals overlap. Notably, the inverse and moderate association with emotional exhaustion represent a novel contribution to studies using the BIOS-G. Higher optimism is associated with lower psychological fatigue, with a shared variance of one-seventh. This result is consistent with findings of an inverse relationship between the LOT-R and burn-out among healthcare personnel in non-Latino cultural contexts [71] [72].

In relation to the fifth objective, the relationship of the BIOS-G with gender and age was significant in this sample of university teachers. The level of optimism is greater in men than women and increases with age as expected [14]. In addition, the strength of the association with both sociodemographic variables was weak, consistent with expectations. The lower optimism observed in women may be linked to their greater predisposition toward negative affect (e.g., anxiety and depression) compared to men [8].

Considering a cultural factor, traditional Latin American culture has historically emphasized greater dependence and submission of women to the extended family of origin, the immediate nuclear family, and male authority. This culturally expected behavior may contribute to a less optimistic attitude among women compared to men, particularly in the professional sphere. However, it is important to note that cultural norms are now shifting toward greater female empowerment, independence, and leadership. Consequently, this cultural influence may have less impact on younger women than on older women [73].

On the other hand, the increase in optimism with age may stem from improved social and personal adaptation, as well as greater wisdom acquired over the years [74]. In light of the theory of interactive personality styles [11], personal interaction experiences further reinforce the meaning attributed to optimism within the individual’s cultural context, which helps explain this positive relationship. The weak strength of the association may be attributed to the fact that internalization occurs primarily during childhood and adolescence.

Regarding the sixth objective, years of teaching experience were also positively related to optimism, with a weak strength of association. When partializing the effect of age and years of teaching experience, the relationship between optimism and emotional exhaustion remains significant and a moderate strength of

association. Consequently, optimism is an antecedent trait that acts as a protective factor against emotional exhaustion in university teachers. This finding is particularly relevant to the study of this population from a positive psychology perspective [2] [28].

### 6.1. Limitations

As limitations of the study, it should be noted that non-probabilistic sampling was used, and the sample exhibited a slight female bias, with women comprising half of the participants despite representing only two-fifths of the population. Although the sample size is not small, it exceeds the recommended minimum of 200 for structural equation modeling [50] and represents approximately one-fifth of the teaching population at the School of Medicine and Health Sciences of the private university studied. Given that factor loadings were high ( $\geq 0.60$ ) on a dimension with four indicators, the exploratory factor analysis is well-powered [75]. Moreover, there are more than 20 participants for each parameter to be estimated, which is considered a good participant-to-parameter ratio for confirmatory factor analysis [50]. Additionally, statistical analyses were conducted in accordance with the ordinal nature of the measurement scale and the lack of compliance with the tau-equivalence assumption [45] [76].

One might ask why a non-probabilistic sampling method (“accidental or convenience sampling”) was chosen for this study rather than a probabilistic sampling method (“simple random sampling”). At the inferential level, it is crucial to obtain a random and representative sample of the population. Probability sampling, in which the likelihood of any population element being included in the sample can be determined a priori, provides stronger assurances of meeting this requirement compared to non-probability sampling. However, in research involving human participants, factors such as informed consent, voluntary participation, and the personal investment of time and cognitive effort to reliably respond to questions or perform tasks make this condition insufficient. This contrasts with research involving plant organisms or mechanical objects, where such constraints do not apply [77].

Indeed, it could be argued that the ideal conditions assumed by many inferential tests, which typically rely on probability sampling, are rarely met in studies involving human participants, as probability sampling can be very costly and time-consuming. Conversely, a convenience sample of voluntary participants recruited through a means accessible to the entire population can yield comparable results at a significantly lower cost in terms of resources and time [78]. Therefore, this approach was deemed a more practical alternative without substantially compromising the study’s quality or inferential validity. Nonetheless, employing quota sampling (by sex) could have resulted in a more representative sample, without the slight bias of female participation.

### 6.2. Future Directions

The use of this scale is recommended in studies adopting a positive psychology

perspective within this population, as the findings indicate that optimism acts as a protective antecedent factor against emotional exhaustion. This effect is independent of years of teaching experience and age, aligning with the trait-based nature of the concept measured [12]. Maintaining a well-adjusted and satisfied teaching staff, free from emotional exhaustion, enhances student learning outcomes [18]. Furthermore, when evaluating internal consistency with this scale, the use of ordinal omega is essential, as the measurement model is congeneric—meaning the factor loadings of its ordinal indicators (Likert-type items) are not uniform [43] [65]. Finally, it is suggested to test a new item (“life always offers opportunities to move forward”) as a potential substitute for item 1 to mitigate collinearity issues with item 2 and reliability problems arising from the potential confusion caused by its negative wording. Furthermore, this new item aligns with the interactive personality styles theory on which the scale is based.

## 7. Conclusion

The items demonstrate discriminative power, although items 2 and 3 exhibit a ceiling effect, and can be considered reliable. The internal consistency of the scale is acceptable according to Green and Yang’s ordinal omega and good according to McDonald’s ordinal omega, with no evidence of content redundancy, as values exceed 0.95. However, there is a collinearity issue between items 1 and 2, requiring the specification of their measurement weights for the factorial solution to be admissible. The fit of the one-factor model ranges from good to acceptable, and the model demonstrates construct validity, representing a congeneric measurement model. Hypotheses for convergent validity were confirmed, showing a direct relationship with affective balance and positive affect and an inverse relationship with emotional exhaustion and negative affect. The BIOS-G was weakly associated with male gender, older age, and greater years of teaching experience within the sample of professors from the Faculty of Medicine and Health Sciences of a private university. Notably, the scale’s association with emotional exhaustion was independent of age and teaching experience.

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## Conflicts of Interest

The authors declare no conflicts of interest regarding the publication of this paper.

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## Appendix 1. Scales Applied

*BIOS-G.* Please answer by selecting one option for each statement. There are no right or wrong answers; each person is different and responds based on their personal opinion or perception.

Ítem	Claro que no	Creo que no	Creo que sí	Claro que sí
1. La vida es fea.				
2. La vida es bella.				
3. La vida es buena.				
4. La mayoría de las personas son buenas				

Note. Items: 1. Life is ugly, 2. Life is beautiful, 3. Life is good, 4. Most people are good. Response labels: 0 = Of course not, 1 = I think not, 2 = I think so, and 3 = Of course yes. Item 1 is reverse-scored.

*SPANE.* Please reflect on your activities and experiences from the past four weeks. Then, indicate your degree of agreement with the feelings expressed below using the responses provided.

0	1	2	3	4
Nunca	Pocas veces	Algunas veces	La mayoría de las veces	Siempre

Enunciado	Respuesta
1. En las últimas cuatro semanas me he sentido positivo(a)	
2. En las últimas cuatro semanas me he sentido negativo(a)	
3. En las últimas cuatro semanas me he sentido bien	
4. En las últimas cuatro semanas me he sentido mal	
5. En las últimas cuatro semanas he tenido sentimientos agradables	
6. En las últimas cuatro semanas he tenido sentimientos desagradables	
7. En las últimas cuatro semanas me he sentido feliz	
8. En las últimas cuatro semanas me he sentido triste	
9. En las últimas cuatro semanas he sentido miedo/temor	
10. En las últimas cuatro semanas me he sentido alegre	
11. En las últimas cuatro semanas me he sentido enojado(a)	
12. En las últimas cuatro semanas me he sentido satisfecho(a)	

Note. Statement of items: 1. In the last four weeks, I have been feeling positive; 2. In the last four weeks, I have been feeling negative; 3. In the last four weeks, I have felt good; 4. In the last four weeks, I have felt bad; 5. In the last four weeks, I have had pleasant feelings; 6. In the last four weeks, I have had unpleasant feelings; 7. In the last four weeks, I have felt happy; 8. In the last four weeks, I have felt sad; 9. In the last four weeks, I have felt afraid/fearful; 10. In the last four weeks, I have felt joyful; 11. In the last four weeks, I have felt angry; and 12. In the last four weeks, I have felt content. Answer labels: 0 = Never, 1 = Rarely, 2 = Sometimes, 3 = Most of the time, and 5 = Always.



```

3, 3, 3, 3, 2, 3, 3, 3, 3, 3, 3, 3, 2, 3, 3, 2, 3, 3, 2, 3, 2, 2, 2, 3, 2, 3, 2, 3, 2, 3, 3,
3, 1, 3, 3, 3, 3, 3),
O4 = c(2, 2, 2, 1, 2, 2, 2, 2, 2, 2, 2, 2, 2, 2, 2, 2, 3, 2, 2, 1, 3, 2, 2, 2, 2, 1, 2, 2, 2, 3,
3, 3, 3, 2, 3, 1, 1, 3, 2, 1, 3, 3, 2, 1, 3, 1, 2, 3, 2, 3, 2, 3, 3, 2, 3, 3, 2, 3, 3, 2, 2, 2, 2, 2,
2, 2, 2, 1, 2, 3, 3, 2, 3, 1, 2, 2, 2, 2, 2, 2, 1, 1, 3, 2, 2, 1, 2, 0, 2, 2, 0, 2, 2, 2, 2, 2, 1, 2,
2, 1, 2, 2, 2, 2, 2, 2, 1, 2, 3, 2, 2, 2, 2, 2, 1, 1, 2, 1, 0, 2, 2, 3, 2, 1, 2, 2, 1, 2, 2, 2, 2,
2, 2, 2, 1, 1, 3, 2, 2, 2, 1, 2, 3, 2, 1, 3, 3, 2, 3, 2, 2, 3, 1, 1, 1, 2, 2, 2, 2, 2, 2, 3, 2, 3, 3,
2, 3, 3, 2, 3, 1, 2, 3, 2, 2, 3, 3, 3, 3, 3, 1, 1, 2, 2, 2, 2, 1, 1, 3, 2, 2, 1, 3, 2, 3, 1, 1, 1, 2, 2,
1, 1, 3, 2, 3, 2, 2))
data$O1 <- ordered(data$O1)
data$O2 <- ordered(data$O2)
data$O3 <- ordered(data$O3)
data$O4 <- ordered(data$O4)

# Estimate and test the model by WLSMV using polychoric correlations.
model <- "Optimism =~ 0.60 * O1 + 0.849 * O2 + O3 + O4"
fit <- cfa(model, data = data, estimator = "WLSMV", ordered = c("O1", "O2",
"O3", "O4"), start = "simple")

# Results of the factor model fit and its graph, which can be displayed in the console or saved as a JPEG file. In the latter case, remove the blue hashtag symbol.
summary(fit, fit.measures = TRUE, standardized = TRUE)
# jpeg("Figure_CFA_plot.jpeg", width = 800, height = 600, units = "px", res = 300)
semPaths(fit, "std", whatLabels = "std", layout = "tree", edge.label.cex = 2,
edge.color = "black", label.color = "black", fade = FALSE, node.width = 1.4,
sizeMan = 12, sizeLat = 12, node.label.cex = 1.4, label.cex = 1.4)
dev.off()

```