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Psychometric Properties of the Depression Anxiety Stress Scales-21 (DASS-21) in Postgraduate Students from Southern Chile

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Abstract

The aim of this study was to examine the internal structure and internal consistency reliability of the 21-item Depression Anxiety Stress Scales (DASS-21) among postgraduate students from southern Chile. The sample comprised 520 master's and doctoral students from five state universities located in the regions of Biobío, La Araucanía, and Magallanes and the Chilean Antarctic; 53.9% were women and 46.1% were men, with a mean age of 32.80 years (SD = 7.87). The DASS-21 was administered via an online questionnaire. Its internal structure was examined using exploratory structural equation modelling (ESEM) with target rotation and weighted least squares mean and variance adjusted (WLSMV) estimation, while internal consistency was assessed using Cronbach's alpha. The three-factor model yielded $\chi^2(150) = 384.21$, CFI = 0.983, TLI = 0.976, RMSEA = 0.056, 90% CI [0.049, 0.063], and SRMR = 0.030, with $\alpha = 0.945$. The dimensions showed positive and statistically significant correlations ranging from 0.685 to 0.736. Most items displayed their primary loading on the expected factor, but Items 2, 6, 14, and 18 showed their strongest loadings on another dimension. These results describe a three-factor exploratory representation and high internal consistency.

Keywords: depression; anxiety; stress; psychometric properties; postgraduate students; DASS-21

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1. Introduction

The mental health of postgraduate students has become an area of growing interest within psychology and higher education institutions because of the demands associated with advanced academic training (Hazell et al., 2020; Pavlacic et al., 2024). Postgraduate students, particularly doctoral candidates, may simultaneously face pressure related to academic performance and research productivity, employment uncertainty, limited resources, social isolation, and difficulties balancing academic responsibilities with their personal and professional lives (Newlands et al., 2025; Zeeman et al., 2025). These conditions have

been associated with symptoms of anxiety and depression, stress, burnout, and reduced well-being. Their examination therefore requires consideration of both individual factors and the institutional and relational conditions that shape the postgraduate educational experience (Zeeman et al., 2025).

Empirical evidence supports the relevance of this issue. In a longitudinal study of doctoral students in Sweden, Bergvall et al. (2025) observed an increase in the use of medication associated with mental health problems over the course of doctoral training. Likewise, a systematic review and meta-analysis of 32 studies estimated that 24% of doctoral students experienced clinically significant depressive symptoms and 17% experienced anxiety symptoms, although substantial heterogeneity was found across studies (Satinsky et al., 2021). Hazell et al. (2020), in turn, identified elevated levels of stress and highlighted isolation, supervisory relationships, and social support as relevant factors. Taken together, these findings indicate that mental health during postgraduate education is an important concern, although its manifestation varies across contexts and educational trajectories.

Empirical support for bifactor representations of the DASS-21 has also been reported outside Chile. In the United Kingdom, Henry and Crawford (2005) examined 1794 non-clinical adults and found that the best-fitting model comprised a general psychological distress factor alongside three orthogonal specific factors representing depression, anxiety, and stress. Similarly, Osman et al. (2012), in two samples of university students from the United States ($N = 887$ and $N = 410$), identified a general factor accounting for the largest proportion of common variance and obtained good fit for a bifactor model in the follow-up study. More recently, Yılmaz Koğar and Koğar (2023) applied bifactor ESEM to 521 Turkish adults and identified a strong general negative affect factor, together with evidence of specific variance associated with depression. These findings suggest that the relative contributions of general distress and specific symptom dimensions may vary across populations, reinforcing the relevance of examining the instrument's internal structure in each target population.

In Chile, the available evidence has focused primarily on undergraduate university students. Álamo et al. (2020) documented an increase in symptoms of depression, anxiety, and stress among Chilean university students, while subsequent research described these manifestations in higher education settings following the pandemic (Martínez-Líbano et al., 2023). Research with Chilean undergraduate Social Work students has also examined associations between loneliness, academic organisation, and academic stress (Davinson-Pacheco et al., 2026). These studies contextualise student distress, but do not establish the measurement properties of the DASS-21 among postgraduate students. Evidence specifically concerning this population remains limited; properties observed in undergraduate samples cannot be assumed to generalise to master's and doctoral students.

Within this context, instruments are needed that support defensible interpretations of depression, anxiety, and stress scores. The DASS-21 is widely used for dimensional assessment, but debate about its structure has direct consequences for the scores that researchers and practitioners calculate. Internal consistency, correspondence between items and factors, and the distinct information provided by conventional subscale sums are separate questions. Examining these distinctions is necessary before recommending interpretations or applications of DASS-21 scores among Chilean postgraduate students.

Depression Anxiety Stress Scales-21 (DASS-21): Description and Theoretical Structure

The 21-item Depression Anxiety Stress Scales (DASS-21) is the abbreviated version of the DASS-42 developed by Lovibond and Lovibond (1995). It comprises 21 items distributed across three seven-item subscales: depression, anxiety, and stress. In its original format, each item is rated using four ordinal response categories scored from 0 to 3 according to the

frequency or intensity with which the symptom was experienced during the previous week. The DASS-21 is a dimensional self-report measure, and its scores do not, in themselves, constitute clinical diagnoses (Lovibond & Lovibond, 1995; Xiong et al., 2025).

The depression dimension encompasses manifestations such as dysphoria, hopelessness, self-deprecation, loss of interest, anhedonia, and reduced motivation (Ahmed et al., 2024; Dwight et al., 2024; Guven et al., 2025). Conceptually, this factor is characterised by diminished positive affect and a reduced capacity to experience interest or satisfaction (Lovibond & Lovibond, 1995; Norambuena-Paredes et al., 2025; Paredes-Díaz et al., 2025). The anxiety dimension is primarily related to physiological arousal, fear, and situational anxiety (Alfonsson et al., 2017; Jacobsen et al., 2024; Lee et al., 2019). It includes manifestations such as breathing difficulties, trembling, feelings of panic, and apprehension in situations perceived as threatening (Alotiby, 2024; Katsikidou et al., 2025; Neyazi et al., 2025). Although anxiety is conceptually distinguished from depression by the predominance of physiological hyperarousal, it shares with stress components related to tension and heightened alertness (Lovibond & Lovibond, 1995; Yohannes et al., 2019). Finally, the stress dimension assesses a relatively persistent state of psychological tension and over-arousal, manifested through difficulty relaxing, agitation, irritability, impatience, and low tolerance of frustration. Lovibond and Lovibond (1995) proposed that stress and anxiety lie along a conceptual continuum, meaning that their empirical boundaries may be less clearly defined than those separating either dimension from depression (Nair et al., 2022; Richardson et al., 2017; Silva-Castillo et al., 2025).

The original formulation distinguishes related symptom domains. Depression encompasses dysphoria, hopelessness, self-deprecation, loss of interest, anhedonia, and reduced motivation. Anxiety concerns physiological arousal, fear, and situational anxiety, including breathing difficulties, trembling, and panic. Stress concerns persistent tension and over-arousal, manifested in difficulty relaxing, agitation, irritability, impatience, and low frustration tolerance (Lovibond & Lovibond, 1995; Antúnez & Vinet, 2012). These descriptions provide a theoretical basis for separate subscales, but shared symptom content means that their empirical differentiation must be assessed in the population in which the scores will be interpreted.

The original formulation proposes three distinguishable but correlated factors. However, accumulated evidence has sustained debate regarding three-factor, hierarchical, and bifactor structures. A systematic review and factorial meta-analysis documented support for three-factor solutions alongside evidence of substantial shared variance and favourable fit for bifactor representations (Yeung et al., 2020). Cross-cultural research has likewise examined the balance between general distress and specific dimensions (Zanon et al., 2021). The existence of three fitted factors does not, by itself, establish that the conventional sums distinguish the symptom domains sufficiently for their intended use (Gómez et al., 2020).

In conventional independent-clusters confirmatory factor analysis (CFA), non-target loadings are fixed to zero. ESEM allows target loadings and cross-loadings to be estimated, which can help describe instruments with related symptom domains (Asparouhov & Muthén, 2009; Marsh et al., 2020). However, the conventional DASS-21 scoring rule continues to allocate each item exclusively to its original seven-item subscale. Allowing cross-loadings in the measurement model does not adjust those observed sums or make them equivalent to estimated latent factor scores. Interpretation must therefore consider the loading pattern, factor correlations, and evidence relevant to the actual scores being recommended.

In Chile, the DASS-21 was translated and adapted by Vinet et al. (2008). Antúnez and Vinet (2012) subsequently obtained a three-factor solution among university students, while also documenting cross-loadings and a weak loading for Item 2. Roman-Mella et al. (2014) and Román et al. (2016) provided evidence in adolescent and young populations.

Studies involving Chilean workers, adults, and healthcare personnel have also considered general distress and bifactor representations (Güilgüiruca et al., 2023; Cancino et al., 2024; Güilgüiruca-Retamal et al., 2026). These findings indicate that the relative differentiation of symptom dimensions requires evaluation in each population rather than an assumption based on a shared language or national setting.

Evidence from Spanish-speaking populations also raises concerns about score interpretation. González-Rivera et al. (2020) found poor support for the original three-factor CFA model and discriminant validity among 1073 Puerto Rican adults despite high internal consistency. In Mexico, Salinas-Rodríguez et al. (2023) compared several structures and favoured a bifactor representation. Castillo-Díaz and Henao Periañez (2025), however, supported a correlated three-factor solution among Honduran and Colombian university students. These differences require attention to what scores represent in each population, beyond their consistency or the fit of any one model.

The present study examined the internal structure and internal consistency of DASS-21 scores among master's and doctoral students from southern Chile. A three-factor ESEM specification with target rotation was estimated to describe target loadings, cross-loadings, and factor correlations. The study also examined how this pattern corresponds to the conventional item allocation. Only this ESEM specification was estimated; the analyses therefore cannot identify a preferred factorial structure or establish superiority over unidimensional, conventional three-factor CFA, hierarchical, or bifactor models. The internal consistency coefficients concern the original item sets and do not establish their distinct interpretation or clinical utility.

2. Materials and Methods

2.1. Participants

The analytical sample comprised 520 postgraduate students enrolled at five state universities located in the regions of Biobío, La Araucanía, and Magallanes and the Chilean Antarctic. Participants were selected using purposive non-probability sampling (Creswell & Creswell, 2023). The mean age was 32.80 years (SD = 7.87); 54.4% of the sample were women and 45.6% were men. Of the total sample, 63% were enrolled in master's programmes and 36.7% in doctoral programmes. Additional sociodemographic characteristics of the sample are presented in Table 1.

Table 1. Main Characteristics of the Sample.

Variable	Category	%
Sex	Men	45.6
	Women	54.4
Study level	Master's degree	63.3
	Doctoral degree	36.7
Marital status	Single	76.2
	Married	15.4
	Divorced	3.8
	Civil partnership	3.1
	Separated	1.3
	Widowed	0.2
Children	Yes	27.1
	No	72.9

Table 1. *Cont.*

Variable	Category	%
Region of residence	Biobío Region	14.8
	La Araucanía Region	68.8
	Magallanes and Chilean Antarctic	5.0
	Other Region	11.3
Cultural group	Indigenous	14.8
	Non-Indigenous	85.2

Note. Regional percentages may not sum exactly to 100% because of rounding.

2.2. Instruments

A sociodemographic questionnaire collected information on age, sex, postgraduate study level, marital status, presence of children, region of residence, and Indigenous identification, as reported in Table 1. Psychological symptoms were assessed with the DASS-21 (Lovibond & Lovibond, 1995), using the Spanish-language version validated among Chilean university students by Antúnez and Vinet (2012). The instrument comprises four ordered response categories referring to symptoms during the previous week; the analytical coding reported here ranges from 0 to 3. The conventional subscales assign Items 3, 5, 10, 13, 16, 17, and 21 to Depression; Items 2, 4, 7, 9, 15, 19, and 20 to Anxiety; and Items 1, 6, 8, 11, 12, 14, and 18 to Stress (Table 2). Each unmultiplied seven-item sum consequently ranges from 0 to 21. These conventional assignments were retained for descriptive comparability; they were not revised according to the estimated cross-loadings. Internal consistency estimates from the present sample are reported in the Results.

Table 2. Distribution of DASS-21 Items.

n°	Subdimension	Item
1	Stress	I found it very difficult to release tension.
2	Anxiety	I noticed that my mouth was dry.
3	Depression	I could not experience any positive feelings.
4	Anxiety	I found it difficult to breathe.
5	Depression	I found it difficult to take the initiative to do things.
6	Stress	I overreacted in certain situations.
7	Anxiety	I felt my hands trembling.
8	Stress	I felt that I was expending a great deal of energy.
9	Anxiety	I was worried about situations in which I might panic or make a fool of myself.
10	Depression	I felt that there was nothing to look forward to.
11	Stress	I felt restless.
12	Stress	I found it difficult to relax.
13	Depression	I felt sad and depressed.
14	Stress	I could not tolerate anything that prevented me from continuing what I was doing.
15	Anxiety	I felt that I was on the verge of panic.
16	Depression	I could not get enthusiastic about anything.
17	Depression	I felt that I had little worth as a person.
18	Stress	I tended to get angry easily.
19	Anxiety	I felt my heart beating even though I had not exerted myself physically.
20	Anxiety	I felt afraid for no reason.
21	Depression	I felt that life had no meaning.

2.3. Procedure

Formal contact was established with the academic authorities of the participating universities, who granted the corresponding institutional approvals. The study was approved by the Scientific Ethics Committee of Universidad de La Frontera (UFRO File No. 145/23), in accordance with the ethical principles applicable to research involving human participants.

Data were collected using an online questionnaire hosted on QuestionPro. A total of 550 postgraduate students were invited through institutional email. The invitation included a questionnaire link and an informed consent form describing the study aims, voluntary participation, confidentiality and anonymity, and the right to withdraw without academic or personal consequences. The questionnaire was self-administered.

The analytical dataset was obtained by filtering the QuestionPro export to include complete responses from eligible participants. This yielded 520 complete questionnaires, corresponding to 94.5% of the 550 invitations. The count of 520 refers to records included in the analysis; it does not identify how many of the remaining invitees did not initiate the survey, submitted an incomplete questionnaire, or did not meet the eligibility criteria. A completeness filter alone does not establish that all retained records are unique or otherwise valid.

2.4. Data Analysis

A total of 520 complete responses were collected. Descriptive statistics and Spearman's rank-order correlations among the three dimensions were subsequently calculated, taking into account the ordinal nature of the responses and the absence of normality. These analyses were conducted using SPSS v.23. The reported subscale-level descriptive statistics and Spearman correlations were interpreted separately from the latent ESEM solution and were not treated as evidence confirming the original item assignments.

To examine the internal structure of the instrument, a three-factor exploratory structural equation modelling (ESEM) model with target rotation was estimated. This approach allows both primary and cross-loadings to be freely estimated (Asparouhov & Muthén, 2009; Marsh et al., 2020). The model was estimated in Mplus v.8.10 (Muthén & Muthén, 2017) using the weighted least squares mean and variance adjusted (WLSMV) estimator, which is appropriate for ordinal indicators. With 21 items and three exploratory factors, this measurement model has 150 degrees of freedom, as does an unrestricted three-factor exploratory solution. Target rotation guides the interpretation of the loading matrix without changing global model fit; it does not impose the zero cross-loadings required by a simple-structure confirmatory model (Asparouhov & Muthén, 2009).

Model fit was described using χ^2, the comparative fit index (CFI), Tucker–Lewis index (TLI), root mean square error of approximation (RMSEA) with its 90% confidence interval, and standardised root mean square residual (SRMR). Conventional cutoffs developed for continuous indicators may be too permissive for ordinal responses (McNeish, 2023). No model-specific simulation-based cutoffs were derived; therefore, the indices are reported descriptively rather than used to classify fit as satisfactory, good, or acceptable. Interpretation also considered primary loadings, cross-loadings, and factor correlations. Internal consistency was assessed using Cronbach's alpha for the total score and each conventional subscale; these coefficients were not interpreted as evidence of unidimensionality. Model-based omega coefficients for the ESEM factors and measurement invariance across sex and programme level were not evaluated.

3. Results

Using the analytical coding scheme ranging from 0 to 3, the mean scores were 1.22 (SD = 0.66) for stress, 0.85 (SD = 0.72) for depression, and 0.78 (SD = 0.70) for anxiety.

The three dimensions showed positive and statistically significant associations. The strongest correlation was observed between stress and anxiety ($r_s = 0.753$, $p < 0.001$), followed by stress and depression ($r_s = 0.707$, $p < 0.001$), and depression and anxiety ($r_s = 0.703$, $p < 0.001$). The magnitude of these associations indicates substantial overlap among the dimensions; however, the correlations alone do not provide evidence for the existence of a general factor or demonstrate the independence of the factors (see Table 3).

The ESEM model showed positive correlations among the three latent factors, ranging from 0.598 to 0.777. The strongest correlation was between Depression and Anxiety (r = 0.777, SE = 0.022), followed by Stress and Anxiety (r = 0.635, SE = 0.037) and Stress and Depression (r = 0.598, SE = 0.036) (see Table 4).

Table 3. Spearman's Correlations Among Stress, Depression, and Anxiety.

Variable	1	2
1. Stress	—	
2. Depression	0.707 ***	—
3. Anxiety	0.753 ***	0.703 ***

Note. Values are Spearman's rank-order correlations (r_s) among observed subscale scores, not latent ESEM correlations. All tests were two-tailed. *** $p < 0.001$.

Table 4. Latent Factor Correlations for the ESEM Model.

Latent Factors	Correlation	SE
Stress–Depression	0.598	0.036
Stress–Anxiety	0.635	0.037
Depression–Anxiety	0.777	0.022

Exploratory Structural Equation Modelling (ESEM)

The three-factor ESEM model yielded $\chi^2(150) = 407.170$, CFI = 0.983, TLI = 0.976, RMSEA = 0.057, 90% CI [0.051, 0.064] (see Table 5). These indices describe the discrepancy between the observed item associations and a three-factor exploratory representation; they do not directly test the conventional item-to-subscale allocation. Overall internal consistency was high ($\alpha = 0.945$), with coefficients of 0.855 for stress, 0.898 for depression, and 0.873 for anxiety. Most items showed their highest loading on the theoretically expected factor; however, relevant cross-loadings were observed, and several items loaded most strongly on another dimension (see the factor-loading table). The rotated loading pattern therefore shows partial correspondence with the theoretical allocation, together with departures that limit a straightforward interpretation of the conventional subscales.

Table 5. Factor Loadings for the Three-Factor ESEM Model of the DASS-21.

Item	Factor 1: Stress	Factor 2: Depression	Factor 3: Anxiety	Residual Variance
Item 1	0.712 (0.045)	−0.046 (0.056)	0.112 (0.063)	0.424
Item 2	0.380 (0.056)	−0.005 (0.080)	0.301 (0.090)	0.625
Item 3	0.194 (0.051)	0.594 (0.055)	0.099 (0.064)	0.346
Item 4	0.186 (0.057)	0.084 (0.070)	0.571 (0.077)	0.403
Item 5	0.262 (0.052)	0.550 (0.069)	−0.147 (0.076)	0.609
Item 6	0.177 (0.057)	0.145 (0.064)	0.435 (0.069)	0.532
Item 7	0.192 (0.059)	0.051 (0.074)	0.552 (0.082)	0.466
Item 8	0.573 (0.051)	0.057 (0.067)	0.144 (0.072)	0.491
Item 9	0.294 (0.047)	0.087 (0.058)	0.504 (0.068)	0.366
Item 10	0.071 (0.041)	1.039 (0.043)	−0.315 (0.053)	0.264
Item 11	0.652 (0.043)	0.073 (0.054)	0.152 (0.061)	0.345
Item 12	0.848 (0.047)	0.122 (0.043)	−0.012 (0.056)	0.159
Item 13	0.175 (0.041)	0.642 (0.049)	0.096 (0.056)	0.297
Item 14	0.137 (0.059)	0.413 (0.059)	0.280 (0.073)	0.435
Item 15	0.038 (0.042)	0.239 (0.046)	0.674 (0.054)	0.193
Item 16	−0.025 (0.041)	0.961 (0.037)	−0.030 (0.045)	0.147
Item 17	−0.089 (0.043)	0.776 (0.050)	0.177 (0.054)	0.248
Item 18	0.100 (0.054)	0.190 (0.063)	0.480 (0.063)	0.498
Item 19	0.058 (0.044)	−0.033 (0.062)	0.809 (0.065)	0.325
Item 20	0.025 (0.047)	0.069 (0.052)	0.795 (0.057)	0.249
Item 21	−0.248 (0.049)	0.852 (0.047)	0.183 (0.054)	0.248

Note. ESEM = exploratory structural equation modelling; DASS-21 = Depression Anxiety Stress Scales-21. Boldface indicates the loading on the factor to which each item is theoretically assigned. Cross-loadings were freely estimated under target rotation.

Items 2, 6, 14, and 18 showed patterns that warrant particular attention. Item 2, theoretically assigned to anxiety, had a loading of 0.301 on that factor and a higher loading on stress (0.380). Item 6, theoretically assigned to stress, showed a loading of 0.177 on stress and a higher loading on anxiety (0.435). These findings indicate that, in this sample, both items were more strongly associated with a dimension other than the one theoretically expected.

Item 14, theoretically corresponding to stress, showed its highest loading on depression (0.413), followed by anxiety (0.280) and stress (0.137). Item 18, also assigned to stress, had a loading of 0.100 on its theoretical factor and a higher loading on anxiety (0.480). In both cases, the predominant loading did not correspond to the dimension originally proposed.

Additionally, Item 10 showed a loading of 1.039 (SE = 0.043) on depression and a secondary loading of −0.315 (SE = 0.053) on anxiety. Examination of the standardized solution showed an R^2 of 0.736 and a positive residual variance of 0.264.

4. Discussion

The study provides preliminary evidence about a specified three-factor ESEM representation of DASS-21 responses and the internal consistency of the conventional item sets among postgraduate students from southern Chile. These findings address different measurement questions. Global fit describes the reproduction of item associations; the rotated loadings describe correspondence with the proposed factors; and alpha describes consistency within the scores that were calculated. Their combination does not establish that the original Depression, Anxiety, and Stress sums provide sufficiently differentiated information for every proposed use. The principal scoring concern is the conventional Stress composite, for which three of seven indicators loaded weakly on Stress and most strongly on another factor.

The distinction between observed and latent associations matters for this interpretation. Stress and Anxiety had the strongest observed-score correlation (0.753), but Depression and

Anxiety had the strongest latent ESEM correlation (0.777). Thus, the latent solution does not simply reproduce the ranking of associations between conventional sums. The correlations are consistent with overlapping symptom domains, but do not by themselves demonstrate either a general distress factor or sufficient discriminant validity. In particular, the fact that correlations are below unity is insufficient to validate distinct individual profiles.

ESEM makes the overlap visible by estimating cross-loadings, but does not resolve its implications for conventional scoring automatically. Each original subscale sums seven items with equal weights and allocates each item to only one domain. That score therefore retains contributions associated with non-target factors when items cross-load. The estimated latent factors account for these associations within the measurement model; the conventional observed sums do not make the same distinctions. Consequently, evidence about a rotated factor cannot be transferred directly to its similarly named subscale sum. Factor scores would also require their own reliability and determinacy evaluation before being proposed as an alternative (Asparouhov & Muthén, 2009; van Zyl & ten Klooster, 2022).

González-Rivera et al. (2020) illustrate why reliability and score differentiation must be considered separately. In their Puerto Rican sample, latent correlations were high and both maximum and average shared variance exceeded average variance extracted, despite high consistency. The present ESEM represents cross-loadings, but does not demonstrate that the conventional sums overcome this overlap. The weak Stress loadings make that distinction especially relevant. Their results cannot be transferred directly to Chilean postgraduate students: the populations differ, and their maximum-likelihood CFA with Satorra–Bentler correction differs from the present ordinal WLSMV ESEM. Our contribution is to document the current loading pattern and its scoring implications; it is not to establish either universal equivalence or universal failure of the three subscales.

The Mexican comparison by Salinas-Rodríguez et al. (2023), which favoured a bifactor structure during the first COVID-19 lockdown, further illustrates the heterogeneity of Hispanic evidence. Conversely, Castillo-Díaz and Henao Periañez (2025) supported a correlated three-factor representation among Honduran and Colombian university students. Neither result identifies the preferred structure for Chilean postgraduate students. Together with the Chilean cross-loading findings, this literature motivates explicit evaluation of dimensional representation and the information retained in conventional scores, while considering the population and estimation method.

The contribution of ESEM relative to simpler or more complex structures remains untested here. Gómez et al. (2020) compared conventional CFA, ESEM, bifactor CFA, and bifactor ESEM in 738 community adults. Although the models had favourable global fit, some ESEM and bifactor specific factors were poorly defined; the authors preferred the parsimonious first-order CFA model because its factors were well defined and supported by reliability and validity evidence. Their findings illustrate why a model should not be selected from fit indices alone. In the present sample, comparisons with a unidimensional model and the conventional three-factor CFA model would clarify the contribution of freely estimated cross-loadings. A theoretically justified bifactor model could additionally examine general and specific variance, provided that its estimates are admissible and its factors meaningfully defined. No such comparison was performed, so no preferred structure is identified.

Fit indices also require consideration of the ordinal response categories and the model specification. McNeish (2023) showed that conventional cutoffs may not transfer adequately to categorical factor models and proposed model-specific calibration. Such calibration was not undertaken here. The indices are therefore presented descriptively and considered alongside the weak target loadings and cross-loadings. The interpretation of Stress does not depend on assigning a favourable verbal label to the global fit: those item-level departures remain relevant to the score that users would calculate.

The item findings are partly consistent with earlier Chilean evidence, although their implications should not be reduced to a general statement that the dimensions are related. Antúnez and Vinet (2012) reported cross-loadings and a weak Anxiety loading for Item 2 among undergraduate students. In the present sample, Items 6 and 18 were more strongly associated with Anxiety than Stress, while Item 14 was more strongly associated with Depression. The largest-loading pattern identifies a substantive problem for the theoretical allocation even though differences between loadings were not formally tested. Item 10 also requires scrutiny because its target loading exceeds 1 and it has a negative Anxiety cross-loading; its positive residual variance does not establish overall model stability.

The available evidence is insufficient to interpret the conventional seven-item Stress sum as a sufficiently specific measure of Stress in this population. This conclusion concerns the support for that interpretation, rather than proof that the score contains no stress-related information. Items 1, 8, 11, and 12 retain their largest loading on Stress, whereas Items 6, 14, and 18 contribute responses more strongly associated with other dimensions in this solution. If the conventional sum is retained for descriptive comparison with previous research, it should be identified as the original seven-item composite and accompanied by this limitation. Deleting these items, assigning them to another subscale, or counting them in multiple subscales would create a different scoring rule requiring independent evaluation. Neither a four-item Stress score nor any automatic reassignment is supported by the present analyses. Depression and Anxiety also require score-specific evidence before being treated as fully differentiated measures.

The alpha coefficients (0.855–0.898 for the conventional subscales) indicate internal consistency, but do not quantify how much reliable variance is specific to each intended construct. Shared variance associated with other dimensions can also contribute to consistency. Model-based reliability should therefore be defined for the intended score and evaluated under a measurement model that accommodates its multidimensionality. If a bifactor representation is justified and admissible, omega hierarchical and omega hierarchical subscale coefficients, together with factor definition and relevant external associations, could help assess the interpretation of general and specific scores (Gómez et al., 2020; Trizano-Hermosilla et al., 2021). These quantities were not estimated here. A high total alpha also does not establish that replacing the subscales with a single general distress score is appropriate.

Measurement invariance across sex and programme level was not examined. The sample size and alpha coefficients do not establish equivalence of item loadings, thresholds, or score interpretations across these groups. Group comparisons therefore require additional evidence. Temporal invariance and repeated-measures data are also needed before differences across occasions can be interpreted as changes in the same construct.

The practical contribution is to identify how the specified ESEM solution behaves in an understudied postgraduate population and where conventional scoring requires further evaluation. The study does not establish screening accuracy, appropriate severity cutoffs, or the validity of individual profiles across the three subscales. Screening requires evidence against a suitable external clinical criterion, including sensitivity and specificity in the target setting. Symptom monitoring requires evidence of temporal comparability, measurement error, and responsiveness to meaningful change. Stating that the questionnaire does not provide a diagnosis does not establish these other properties. Accordingly, the present findings support cautious descriptive research use with explicit scoring limitations, but do not independently justify screening, clinical decision rules, or longitudinal monitoring applications.

Limitations and Future Research

First, the cross-sectional design allowed the internal structure of the DASS-21 to be examined at a single point in time, but it did not permit assessment of the temporal stability of its scores, sensitivity to change, or the evolution of psychological distress throughout postgraduate training. Nor does this design allow temporal relationships between academic demands and symptomatology to be established. Longitudinal studies could examine factorial stability and within-person changes across different stages of postgraduate training, particularly given that previous evidence has shown variations in mental health over the course of doctoral education (Bergvall et al., 2025, Gálvez-Nieto et al., 2025; Norambuena-Paredes et al., 2026).

Second, purposive non-probability recruitment from five state universities limits generalisability. The complete-response sample represented 520 of 550 invitees, but this proportion does not establish representativeness or eliminate self-selection and complete-case selection bias. The reported region of residence is not a substitute for university affiliation or disciplinary composition. Without a documented participant flow and distributions by institution and discipline, participation across these strata cannot be assessed. Indigenous identification was reported by 14.8% of the sample, but equivalence of measurement across Indigenous and non-Indigenous participants was not examined. Future studies should broaden institutional and disciplinary coverage and evaluate invariance across relevant groups (Bibi et al., 2020; Castillo-Díaz & Henao Periañez, 2025).

Third, responses came from a single online self-report administration and may be affected by recall, social desirability, and common-method variance. Filtering QuestionPro exports for complete responses defines the analysed records but does not establish the absence of partial submissions, duplicate participation, or invalid response patterns. These issues require the recruitment and survey records and an explicit account of the checks actually applied. The sensitivity of the results to any exclusions based on extreme symptom scores also remains unestablished; high symptom scores should not, by themselves, be treated as invalid responses. External symptom and clinical measures are needed to examine convergent, discriminant, incremental, and criterion-related validity.

Finally, only one factorial specification was estimated. Comparisons with a unidimensional model and the conventional correlated three-factor CFA model, followed by theoretically justified bifactor alternatives where appropriate, are needed to examine plausible explanations of the overlap. Evaluation should consider admissibility, factor definition, interpretable scores, and model-based reliability alongside fit and parsimony. Several of these questions could be addressed through reanalysis of the existing item-level dataset; they are unresolved analyses, rather than properties established by the present interpretation. Replication, invariance testing, and external validation would then be required to assess the stability and usefulness of the intended scores. Academic stage, workload, supervision, and social support may also help contextualise symptom reports in future studies (Hazell et al., 2020; Newlands et al., 2025; Zeeman et al., 2025).

5. Conclusions

The present study describes the loading pattern, factor correlations, and internal consistency associated with a three-factor ESEM specification of the DASS-21 among postgraduate students from southern Chile. The fit indices concern this specification; they do not establish a preferred factorial structure or validate the original scoring system. High internal consistency does not, by itself, demonstrate that the three conventional sums provide sufficiently distinct information.

The conventional seven-item Stress composite requires particular caution because Items 6, 14, and 18 had weak loadings on Stress and their largest loading on another factor.

The study extends evidence to master's and doctoral students from five state universities in southern Chile. Its contribution is an initial examination of the specified ESEM solution and the scoring questions revealed by its cross-loadings. Comparisons of plausible models, reliability analyses appropriate to the intended scores, and tests of discriminant and incremental validity remain necessary before stronger score interpretations can be supported.

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