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# Psychometric Properties of the Perceived Stress Scale in an Adult Psychiatric Inpatient Sample

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The data and the analysis script are publicly available at http://dx.doi.org/10.17632/k8zk3xhcjf.2. We have no known conflict of interest to disclose. The study is based on an unpublished Master's dissertation of Figalová (2019).

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**Abstract**

Despite the fact that the Perceived Stress Scale (PSS) is often used in the psychiatric population, only few studies reported its psychometric properties in such samples. This study aims to bridge this gap. We administered the 10-item version of the PSS to a sample of psychiatric inpatients ($n$ = 153) and evaluated its psychometric properties. Using the confirmatory factor analysis, we found that a bifactor model was the best fit. The scale showed excellent internal consistency ($\alpha$ = .91 and $\omega$ = .93 for the bifactor model). Item analysis discovered strong inter-item correlations, and indicated that item 9 had relatively low factor loading and item-total correlation. Women obtained a higher score of perceived stress than men. Our findings suggest that the scale works differently in the psychiatric sample than in the general population, and that the PSS might be omitting some of the important aspects of the perceived stress construct.

*Keywords:* Perceived Stress Scale, Reliability, Validity, Factor Structure, Psychiatric Inpatient

The *Perceived Stress Scale* (PSS), developed by Cohen, Kamarck, and Mermelstein (1983), is a self-report scale designed to measure the degree to which individuals appraise situations in their lives as stressful. The items of the PSS are developed to evaluate how unpredictable, uncontrollable, and overloaded respondents find their lives (Cohen & Williamson, 1988). Originally, the scale consists of 14 items. Two shorter versions, derived from the original scale, exist: the four-item *PSS-4*, and the 10-item *PSS-10*. The three forms of the PSS were previously compared (Andreou et al., 2011; Cohen & Williamson, 1988; Lesage, Berjot, & Deschamps, 2012; Leung, Lam, & Chan, 2010). Authors consensually report comparable or higher internal consistency, better factor structure, and higher sensitivity for the PSS-10 compared to both PSS-14 and PSS-10. This phenomenon was also confirmed in the Czech version of the scale (Figalová & Charvát, 2021a). Therefore, the PSS-10 is the most often used version of the scale.

Psychometric properties of the PSS were previously reported in a wide variety of samples, including general adult population (Andreou et al., 2011; Cohen & Janicki-Deverts, 2012; Figalova & Charvat, 2021a), students (Örücü & Demir, 2009; Ramírez & Hernández, 2007; Roberti, Harrington, & Storch, 2006), teachers (Reis, Hino, & Añez, 2010), policewomen (Wang et al., 2011), or adults with a physical illness (Golden-Kreutz, Browne, Frierson, & Andersen, 2004; Lee, Chung, Suh, & Jung, 2015; Leung, Lam, & Chan, 2010). Furthermore, two studies reporting data obtained in a psychiatric sample were identified. Authors of the first study (Jovanovic & Gavrilov-Jerkovic, 2015) administered the PSS-10 to a sample including 157 outpatients, diagnosed with depressive disorders (36.90%), mixed anxiety and depression (28.70%), anxiety disorders (26.80%), and other (7.60%). The authors suggested that the structure of the PSS was best represented by a bifactor model, comprising two specific factors

plus a general factor. The scale exhibited good internal consistency ($\alpha$ = .86), positive correlations with measures of depression ($r$ = .73), anxiety ($r$ = .61), stress ($r$ = .71), negative affect ($r$ = .78), and negative correlations with positive affect ($r$ = -.68) and life satisfaction ($r$ = -.55). In the second study, Hewitt, Flett, and Mosher (1992) administered the PSS-14 to a sample of 96 psychiatric patients (76 outpatients, 20 inpatients), diagnosed mostly with depressive disorders, schizophrenia, marital/family problems, alcoholism, and adjustment disorders. Authors suggested that a two-factor solution (which accounted for 46.60% of the variance) was most appropriate. The scale exhibited good internal consistency ($\alpha$ = .80) and positive correlation with the Beck Depression Inventory ($r$ = .57).

Despite the fact that the PSS is the most widely used measure of perceived stress in the English-speaking countries and has been translated into more than 30 languages (Cohen & Janicki-Deverts, 2012), we found only two studies reporting the psychometric properties of the PSS in a psychiatric sample. Nevertheless, the PSS is often used with psychiatric patients. For example, Candrian et al. (2008) administered the PSS to patients with major medical depression who underwent a medical treatment with antidepressants, and Mugrabi et al. (2020) used the PSS to assess changes in perceived stress induced by psychiatric treatment in patients with psychiatric disorders. However, if an instrument is to be used in a very specific sample (e.g., psychiatric population), it is important to evaluate its psychometric properties in such samples. Otherwise, the external validity of an instrument may be low and the results biased. Item 9 of the PSS-10 may serve as an example. Respondents are asked: *„In the last month, how often have you been angered because of things that were outside of your control?* “. While a person from a general adult population usually controls the majority of things in his or her life, psychiatric inpatients are in a different position. They are often hospitalised against their will, live in a highly

controlled environment, have to follow the rules of the institution, and have even to share their personal space with other patients. Therefore, it is inappropriate to expect that this item would generate the same results under such different circumstances.

Even though the results of other psychometric studies seem promising, it might be misleading to generalise these results to a psychiatric population. Therefore, the goal of this study is to evaluate the psychometric properties of the Czech version of the PSS-10 in a population of psychiatric inpatients. We aim to assess its usability in psychiatric samples for other authors who consider employing this instrument in their studies, as well as for healthcare professionals who use the PSS in their practice.

## Method

### Sample

The participants had to meet the following inclusion criteria: 1. psychiatric hospital inpatient; 2. primary psychiatric diagnosis; 3. lucidity and vigilance on a level necessary for the administration of the PSS-10; 4. signed an informed consent prior to the administration.

We collected the data in two institutions, the Psychiatric Hospital Opava ($n$ = 79) and the Psychiatric Hospital Kroměříž ($n$ = 74), Czech Republic. The PSS-10 was administered to $N$ = 153 psychiatric hospital inpatients, out of which 56.2% were women ($n$ = 86). The age of the respondents ranged from 16 to 71 years ($M$ = 41.94, $SD$ = 12.49). The average time of hospitalization was 36 days ($SD$ = 27.16, five outliers with hospitalization ranging from 126 to 344 days were removed from the calculation). For future analyses, respondents were also divided into two groups according to the length of hospitalization. The first group consisted of 74 patients hospitalized for less than one month ($M$ = 14.51 days, $SD$ = 9.40), the second group consisted of 79 patients hospitalized for more than one month ($M$ = 68.39 days, $SD$ = 43.00). The

educational attainment and primary diagnosis of the respondents (according to the International Classification of Diseases-10; World Health Organisation, 1993) are presented in Table 1.

Table 1

*Characteristics of the sample by educational attainment and primary diagnosis*

| Variable | n (N=153) | % |
|---|---|---|
| Level of education | | |
| Secondary school or less (9 years or less) | 29 | 19.0 |
| Practical high school (12 years) | 56 | 36.6 |
| High school (13 years) | 41 | 26.8 |
| University (16 years or more) | 25 | 16.3 |
| Not specified | 2 | 1.3 |
| Primary diagnosis | | |
| F10–F19 Mental and behavioural disorders due to psychoactive substance use | 79 | 51.6 |
| F20-F29 Schizophrenia, schizotypal and delusional disorders | 11 | 7.2 |
| F30-F39 Mood [affective] disorders | 13 | 8.5 |
| F40-F48 Neurotic, stress-related and somatoform disorders | 34 | 22.2 |
| F60-F69 Disorders of adult personality and behaviour | 14 | 9.2 |
| Other | 2 | 1.3 |

This research was approved by the ethical committee of both institutions. All participants had to sign an informed consent prior to the test administration, and agree to share the necessary medical information with the researcher. All participants were informed that their participation is voluntary and can be terminated at any time. All data were anonymized and processed according to the General Data Protection Regulation (GDPR).

Data from Figalova and Charvat (2021b) representing the general adult population of the Czech Republic were used to compare the psychiatric sample presented in this paper with the general population. The sample of the general population consisted of 1725 adults aged 18–91 years (M = 44.32, SD = 12.83), of whom 56.9% were women, mostly with university education (70.7%). For further specification of the sample, see Figalova & Charvat (2021a).

**Instrument**

Participants were asked to provide basic demographic characteristics. Furthermore, we administered the 10-item Czech version of the Perceived Stress Scale (Figalová & Charvát, 2021a). The questions of the scale ask about feelings and thoughts during the last month. Respondents report how often they felt in a certain way on a 5-point scale (from 0 = *never* to 4 = *very often*). The PSS consists of both negatively stated items (measuring perceived distress) and positively stated items (measuring perceived self-efficacy). The total score is obtained by reversing the scores for the positively stated items and then summing all items across the scale (Cohen, Kamarck, & Mermelstein, 1983). The data collection was performed either in group or individually. In some cases, participants required assistance, such as reading the items aloud by the researcher. We did not administer any further instruments alongside the PSS, as this was evaluated as demanding and potentially stressful for the participants.

**Analytic strategy**

We asked participants to fill the PSS-10 in a paper-pencil form. We did not include protocols with three or more missing items in the analysis (in total 33 protocols were omitted). The confirmatory factor analysis (CFA) and reliability analyses were performed in the *lavaan* package (Rosseel, 2012) in the RStudio. For the CFA, we used the standard settings of the function '*cfa'* with standardized estimates. We set the items as ordered; hence the weighted least square mean and variance adjusted (WLSMV) estimator was used. We report robust (scaled) test statistics. We compared the one-factor model, two-factor model, and bifactor model of the PSS-10. To evaluate the internal consistency, we computed both Cronbach's $\alpha$ and McDonald's $\omega$. The item analysis was performed with a focus on an item's mean, inter-item correlation, item-total correlation, and Cronbach's α if the item was deleted. Descriptive analyses and known-

group difference analyses were performed using SPSS, version 26. As the data of different subsamples do not show normal distribution and/or homoscedasticity, we used non-parametric methods to compare these subsamples.

## Results

### Confirmatory factor analysis

Results of the CFA are presented in Table 2. We compared the one-factor, two-factor, and bifactor model. The *AGFI* statistic, as well as the incremental fit indices (*NFI, TLI, CFI, IFI*), were satisfactory in all three models (>.90, Hooper, Coughlan, & Mullen, 2008). The *SRMR* values should be below .05 to indicate a good fit, however, values as high as .08 are deemed acceptable (Hooper, Coughlan, & Mullen, 2008). The observed *SRMR* values suggested that the bifactor model was the best fit for the data, and the two-factor model was an acceptable fit. Moreover, the threshold of an *RMSEA* suggesting good fit is generally considered to be below .07 (Steiger, 2007). In this sample, only the bifactor model was close to this threshold.

Table 2

*Results of the CFA comparing the one-factor, two-factor, and bifactor model*

| | Absolute Fit Indices | | | | | | Incremental Fit Indices | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| Model | $\chi 2$ | *df* | *p* | *AGFI* | *SRMR* | *RMSEA* | *NFI* | *TLI* | *CFI* | *IFI* |
| one-factor | 170.153 | 35 | .000 | .954 | .086 | .159 | .912 | .908 | .928 | .929 |
| two-factor | 77.406 | 34 | .000 | .984 | .051 | .092 | .960 | .970 | .977 | .977 |
| bifactor | 44.588 | 25 | .009 | .990 | .033 | .072 | .977 | *.981* | .990 | .990 |

*Note*. $\chi 2$ = chi-square; df = degrees of freedom; AGFI = Adjusted good fit index; SRMR = Standardized root mean square of residuals; RMSEA = Root mean square of approximation, NFI = Normed fit index; TLI = Tucker-Lewis index; CFI = Comparative fit index, IFI = Incremental fit index.

Table 3 presents the factor loading of the PSS-10 items in all three models we compared. The observed factor loading was acceptable in case of the one-factor and two factor models, only the factor loading of item 9 was slightly below the recommended threshold (.60, Awang, 2014). For the bifactor model, we observed low factor loading of the items on the group factors, but the factor loadings on the general factor was satisfactory. The low loading of the items on the group factors is not problematic if the scores for these factors are not reported individually (DeMars, 2013). The correlation between factors in the two-factor model was $r = .75$. The factors in the bifactor model were set as orthogonal.

Table 3

*Factor loading of the PSS-10 items*

| | One-factor model | Two-factor model | | Bifactor model | | |
|---|---|---|---|---|---|---|
| Item | *PS* | *PI* | *NI* | *PI* | *NI* | *PS* |
| PSS_1 | .680 | | .715 | | .319 | .634 |
| PSS_2 | .707 | | .735 | | .193 | .705 |
| PSS_3 | .729 | | .762 | | .288 | .699 |
| PSS_4r | .764 | .814 | | .335 | | .718 |
| PSS_5r | .786 | .835 | | .359 | | .730 |
| PSS_6 | .585 | | .625 | | .304 | .551 |
| PSS_7r | .754 | .798 | | .527 | | .641 |
| PSS_8r | .737 | .781 | | .543 | | .619 |
| PSS_9 | .532 | | .571 | | .799 | .364 |
| PSS_10 | .865 | | .932 | | .327 | .854 |

*Note*. PS = Perceived Stress; PI = Positively Stated Items; NI = Negatively Stated Items

## Reliability

We computed both Cronbach's $\alpha$ and the McDonald's $\omega$ to assess internal consistency. The total values for the full scale are presented in Table 4. The internal consistency was very good in all three models.

Table 4

*Internal consistency*

| | One-factor | Two-factor | Bifactor |
|---|---|---|---|
| Cronbach's $\alpha$ | .91 | .91 | .91 |
| McDonald's $\omega$ | .89 | .90 | .93 |

Furthermore, we carried an item analysis to evaluate the quality of the items and how these behave in a sample consisting of psychiatric inpatients. Table 5 shows the results. No attenuation effect was observed. All items had satisfactory item-total correlation and the overall internal consistency of the scale would not increase by removal of any of the items.

Table 5

*Item analysis*

| Item | *M* | *SD* | Response option | | | | | Item-Total Correlation | $\alpha$ if Item Deleted |
|---|---|---|---|---|---|---|---|---|---|
| | | | 0 | 1 | 2 | 3 | 4 | | |
| PSS_1 | 2.49 | 1.13 | .05 | .15 | .29 | .29 | .22 | .61 | .88 |
| PSS_2 | 2.44 | 1.12 | .06 | .13 | .31 | .31 | .19 | .64 | .88 |
| PSS_3 | 2.71 | 1.08 | .04 | .08 | .29 | .31 | .27 | .65 | .88 |
| PSS_4r | 1.75 | 1.19 | .18 | .24 | .32 | .18 | .08 | .66 | .88 |
| PSS_5r | 1.95 | 1.14 | .12 | .23 | .33 | .24 | .09 | .68 | .88 |
| PSS_6 | 2.16 | 1.13 | .07 | .23 | .33 | .24 | .14 | .54 | .89 |
| PSS_7r | 1.83 | .99 | .09 | .27 | .40 | .20 | .04 | .62 | .88 |
| PSS_8r | 1.93 | 1.04 | .09 | .23 | .40 | .22 | .07 | .60 | .88 |
| PSS_9 | 2.33 | 1.19 | .08 | .16 | .30 | .26 | .20 | .48 | .89 |
| PSS_10 | 2.28 | 1.39 | .13 | .18 | .27 | .13 | .29 | .79 | .87 |

*Note*. *M* = Mean; *SD* = Standard Deviation; Item-Total Correlation = Corrected Item-Total Correlation; $\alpha$ if Item Deleted = Cronbach's Alpha if item deleted

Table 6 presents the inter-item correlation. The upper part of the table shows Pearson correlations, the bottom part of the table shows polychoric correlations. The observed values

ranged from $r$ = .15 to $r$ =.62 for Pearson correlation, and from $\rho$ = .15 to $\rho$ = .67 for polychoric correlation. Cohen, Swerdlik, and Phillips (1996) recommends that the ideal range of inter-item correlation is between .20 and .40. With an exception of the correlation between items 8r and 9, which was under the suggested threshold, all items fit within or above this threshold. The high observed correlations are further discussed in the Discussion section.

Table 6

*Inter-item correlation*

| | PSS_1 | PSS_2 | PSS_3 | PSS_4r | PSS_5r | PSS_6 | PSS_7r | PSS_8r | PSS_9 | PSS_10 |
|---|---|---|---|---|---|---|---|---|---|---|
| PSS_1 | | .49 | .56 | .41 | .38 | .37 | .38 | .36 | .44 | .53 |
| PSS_2 | .53 | | .49 | .41 | .49 | .41 | .41 | .42 | .36 | .61 |
| PSS_3 | .61 | .54 | | .48 | .46 | .34 | .44 | .37 | .45 | .58 |
| PSS_4r | .45 | .44 | .53 | | .58 | .34 | .61 | .56 | .32 | .57 |
| PSS_5r | .42 | .52 | .50 | .62 | | .38 | .59 | .62 | .27 | .58 |
| PSS_6 | .41 | .45 | .39 | .36 | .41 | | .28 | .34 | .39 | .59 |
| PSS_7r | .43 | .45 | .49 | .67 | .64 | .30 | | .62 | .18 | .47 |
| PSS_8r | .40 | .46 | .41 | .61 | .67 | .37 | .68 | | .15 | .47 |
| PSS_9 | .48 | .40 | .50 | .34 | .29 | .43 | .20 | .15 | | .53 |
| PSS_10 | .58 | .67 | .64 | .63 | .64 | .65 | .53 | .53 | .58 | |

**Descriptive statistics and known-group differences**

The total score of the scale ranged from 5 to 38, the average score of the whole sample was $M$ = 21.88 ($SD$ = 8.10). The average score of women was $M$ = 24.41 ($SD$ = 6.89), the average score of men was $M$ = 18.63 ($SD$ = 8.43). No relationship between age and total score was observed ($r$ = -.05, $p$ = .532). A weak negative correlation was observed between the time of hospitalization and total score ($r$ = -.318, $p$ < .001).

A Mann-Whitney test indicated that the score of perceived stress was higher for women ($Mdn$ = 24) than for men ($Mdn$ = 18), $U$ ($n_{women}$ = 86, $n_{men}$ = 67) = 1750.00, $z$ = -4.16, $p$ < .001. A large effect size was observed ($\eta^2$= .11, $d$ = 0.71). Furthermore, respondents with a primary

diagnosis from the group of Neurotic, stress-related and somatoform disorders (*Mdn* = 29) had a higher score of perceived stress compared to respondents with diagnoses from other categories (*Mdn* = 21). A Mann-Whitney test indicated that this difference was statistically significant, $U$ ($n_{stress}$ = 34, $n_{other}$ = 119) = 919.00, $z$ = -4.85, $p$ <.001. A large effect size was observed ($\eta^2$ = .15, $d$ = 0.85).

To compare the sample of psychiatric inpatients with the general adult population, we used data shared by Figalova & Charvat (2021b). Respondents recruited from the psychiatric inpatient population reported a higher score of perceived stress (*Mdn* = 22) than respondents recruited from the general adult population (*Mdn* = 18). A Mann-Whitney test indicated that this difference was statistically significant, $U$ ($n_{psychiatric}$ = 153, $n_{general}$ = 1725) = 94969.50, $z$ = -5.76, $p$ < .001. A small effect size was observed ($\eta^2$ = .02, $d$ = 0.27).

## Discussion

The aim of this study was to evaluate the psychometric properties of the Czech translation of the 10-item Perceived Stress Scale (PSS-10) in a population of psychiatric inpatients. The sample consists of $N$ = 153 respondents. The respondents were hospitalised in a psychiatric hospital with various mental and behavioural disorders, classified accordingly to the International Classification of Diseases-10 (World Health Organisation, 1993). The average time of hospitalisation was 36 days.

The construct validity of the PSS was examined using confirmatory factor analysis (CFA). The fit of a one-factor, two-factor, and bifactor model was compared and several absolute and incremental fit indices were reported. The $\chi 2$ statistic indicated a poor fit for all three compared models. However, the $\chi 2$ statistic is very sensitive to sample size and is no longer used as a basis for an acceptance or rejection of a model (Vandenberg, 2006). Considering the

*RMSEA* absolute fit indices, only the value of the bifactor model reached the threshold for a good fit (*RMSEA* = .072). However, values below .10 are generally considered to be tolerable (Steiger, 2007), suggesting that the two-factor model could also be a good fit. Furthermore, the *SRMR* values indicating good fit should be below .05, however, values as high as .08 are deemed acceptable (Hooper, Coughlan, & Mullen, 2008). The *SRMR* value observed in the present study suggests the bifactor model was the best fit for the data (*SRMR* = .033), and that the two-factor model had an acceptable fit (*SRMR* = .051). Overall, the one-factor model did not fit the data well. The two-factor model could be considered acceptable, although it was clearly inferior to the bifactor model in all observed indices. The bifactor model fit data the best. This finding is in line with the reports of other authors who compared the one-factor, two-factor, and bifactor model (Figalova & Charvat, 2021a; Jovanovic & Gavrilov-Jerkovic, 2015).

The factor loading of the items was overall satisfactory. However, the factor loading of item 9 (.532 and .571 for the one-factor and two-factor model, respectively) was slightly below the recommended threshold (.60; Awang, 2014). This was also observed previously in a sample of Czech general population (Figalova & Charvat, 2021a). We believe that the reason might be an imperfect translation of this item, or cultural differences in responses to this item. Moreover, bias could also arise from the fact that this item asks participants how often they have been angered because of things that were outside of their control. This question might provide misleading results if asked respondents living in a highly controlled environment of a psychiatric hospital. Furthermore, we found low factor loading of most of the items on group factors in the bifactor model. Therefore, reporting individual scores for the separate “subscales” formed by positively and negatively stated items is not recommended (DeMars, 2013). This finding

supports the recommendation of Cohen and Williamson (1988), who claim that only a single total score of perceived stress should be obtained while using the PSS.

We used both Cronbach's α and McDonald's ω to evaluate the internal consistency of the PSS-10. While Cronbach's α is the most commonly used measure of internal consistency, McDonald's ω is more appropriate for multidimensional data and should therefore be preferred (Dunn et al., 2014). All the observed values suggested very good internal consistency. This agrees with other authors, who also generally reported good or very good internal consistency of the scale.

We have also conducted an item analysis in order to evaluate how each item of the PSS-10 behaves in the studied sample. We found that all items are of a good quality. The only potential problem arose when we analysed the inter-item correlation. Cohen et al. (1996) recommends that the ideal range of inter-item correlation is between .20 and .40, suggesting that while the items are reasonably homogenous, they contain sufficiently unique variance. However, a considerable number of inter-item correlation values reported in this study exceeded the recommended upper threshold. It could suggest that the items might be too homogenous, not capturing the entire bandwidth of the construct. This finding might be also interesting in relation to convergent validity of the PSS. A number of previous authors reported strong and very strong correlations between the PSS and measures of anxiety (Figalova & Charvat, 2021a; Pbert et al., 1992; Remor, 2006; Roberti et al., 2006). We presume that the PSS might omit some of the important aspects of the perceived stress, and this might be accented even more in a sample of psychiatric inpatients. We recommend that this issue be addressed directly in a future study, and potentially a new method is developed to address the full complexity of perceived stress.

We observed higher scores of perceived stress amongst women compared to men in the psychiatric inpatient sample. Similar gender effect was previously observed in several studies on different populations (Andreou et al., 2011; Lesage et al., 2012; Leung et al., 2010; Remor, 2006). The study by Hewitt et al. (1992) on psychiatric population also reported higher levels of perceived stress amongst women. Moreover, gender differences were found in the Czech general adult population, signifying that women have a higher score of perceived stress than men (Figalova & Charvat, 2021a). Contrary, Jovanovic and Gavrilov-Jerkovic (2015) did not observe any gender differences in perceived stress in a Serbian psychiatric sample. We believe that the results of the present study support the good quality of the Czech version of the PSS, as the translation seems to work in a similar manner as the original, English language version. However, the gender difference effect size between the Czech general adult population ($d$ = 0.34) and Czech psychiatric inpatient population ($d$ = 0.71) is relatively large. This difference may be accounted for by the clinical nature of the present sample. The observed difference in total score between the Czech general adult population and the Czech psychiatric sample support this thought. Overall, our findings suggest that the PSS indeed works differently for the general adult population and for the psychiatric inpatient population.

Interestingly, we observed a large effect size in differences between respondents with a primary diagnosis from the group of *Neurotic, stress-related and somatoform disorders* and respondents with diagnoses from other categories ($d$ = 0.85). We believe this finding supports the idea that the PSS really measures, at least up to a point, stress-related symptoms.

The present study has several limitations. First, the sample size was relatively small and a convenience sampling method was used. Second, a relatively small number of psychiatric diagnoses is represented in the sample. In order to generalise the findings of this study, we

recommend using larger sample size and quota sampling to obtain a representative sample of psychiatric inpatient population. Third, no further self-report measures were administered alongside the PSS. Although this was necessary in order not to expose participants to a potentially stressful situation, collecting more data to assess convergent validity would be extremely useful. Nevertheless, the diagnosis of the patient can be, up to a certain point, linked to criterion validity.

To summarize, our findings suggest that the scale works differently in the psychiatric inpatient sample and general adult sample. Moreover, our results suggest that the PSS might be omitting some of the important aspects of the perceived stress construct. Therefore, we recommend caution when using the PSS to assess the level of perceived stress, especially when used in the very specific context of psychiatric inpatients. A new tool to assess perceived stress, that would include more aspects of the complex and multifaceted perceived stress phenomena, would be beneficiary not only for psychiatric inpatient samples, but presumably for researchers studying all populations.

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