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Social and Behavioral Consequences of the COVID-19 Pandemic: Validation of a Pandemic Disengagement Syndrome Scale (PDSS) in Four National Contexts

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The aim of the present study was to develop and validate a self-report measure that investigates people's general disengagement after the acute phases of the pandemic. Across three studies, we examined the psychometric features of the Pandemic Disengagement Syndrome Scale (PDSS) in four national contexts. In Study 1, we developed the instrument and investigated the factorial structure, internal consistency, measurement invariance across gender and countries (the United States and Italy), and discriminant validity. A bifactor model with two specific factors (Social Avoidance and Alienation) provided a better fit than the competing models. In Study 2, we tested the stability of the PDSS as well as its predictive validity. In Study 3, we conducted a quasi-experimental comparison between Norway and Sweden, to investigate whether scores on the PDSS are related to a markedly distinct approach to the pandemic in terms of mandatory lockdown. Overall, results from the three studies demonstrated that the PDSS is a valid and reliable measure of a syndrome of disengagement from others following a pandemic.

Public Significance Statement

This study is the first to validate a specific measure to assess a syndrome of disengagement from others following the acute phases of the pandemic. The findings of the study revealed that the Pandemic Disengagement Syndrome Scale is a psychometrically valid and reliable instrument. This measure could be used to help researchers and practitioners better understand the social and behavioral consequences of the COVID-19 pandemic among adults.

Keywords: COVID-19, quarantine, psychometrics, assessment, mental health

Supplemental materials: https://doi.org/10.1037/pas0001213.supp

The impact of COVID-19 on mental health has been the focus of extensive research during the acute phases of the pandemic (e.g., Prati & Mancini, 2021). However, less is known about the long-term and cumulative mental health consequences of the COVID-19 pandemic (Muehlschlegel et al., 2021). Since the beginning of the pandemic, the overall pattern of spread has been a series of COVID-19 waves. While during an acute COVID-19 wave we may expect a short-term amplification of worries about infection, over time persistently reactivated worries about infection could inhibit social behavior, reinforce avoidance, and inhibit reward seeking, irrespective of risk, a possibility consistent with a maladaptive activation of the behavioral inhibition system (Gray, 1982). Indeed, the effect of regional variation in COVID-19 risk on mental health suggests that worries about the pandemic can be decoupled from objective risk indices (Mancini & Prati, 2022). These potential negative consequences would be reinforced by wide-ranging public health measures that limit social behavior and rewarding activities (Banks & Xu, 2020), such as going to restaurants, movies, and large family gatherings, and by prolonged media exposure (Garfin et al., 2020). Together these effects of the pandemic could lead to a mutually reinforcing syndrome of disengagement from close others, social withdrawal, and reduced reward seeking (e.g., Corr & Cooper, 2016). These behavioral symptoms reflect a preference to maintain

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The first study (https://osf.io/23a7y) and the third study (https://osf.io/pk2y6) were preregistered. The second study was not preregistered. Data, study materials, and all analyses codes are freely available at https://osf.io/u2m5y/

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one's social life on the Internet (or telephone), to feel estranged from close others, and to feel less motivated to develop and maintain friendships and face-to-face relationships. In the present study, we investigated this possibility using a newly developed measure, the Pandemic Disengagement Syndrome Scale (PDSS).

Prior research on pandemics indicates that isolation and quarantine can have persistent psychological consequences. Several studies show that quarantine can lead people to feel stigmatized and rejected in their community and that a return to prequarantine functioning can be delayed for many months (Bai et al., 2004; Cava et al., 2005; Reynolds et al., 2008). In their study of severe acute respiratory syndrome in Toronto, Cava et al. (2005) found evidence of behavioral changes such as avoiding crowds after the quarantine period and staying in quarantine past the release date because people were not sure that the risk had passed. Reynolds et al. (2008) found similar avoidance behaviors (e.g., avoidance of crowded environments and public spaces) among a cohort of persons quarantined during the 2003 severe acute respiratory syndrome outbreak in Canada. One possible reason is fear of infecting others or being infected. In addition to such behavioral changes, there may be cognitive and emotional consequences. In a study during the severe acute respiratory syndrome outbreak, Bai et al. (2004) found that quarantined staff members were more likely to experience exhaustion and detachment from others than their nonquarantined counterparts.

Recent research indicates that the acute phase of the COVID-19 pandemic can result in persistent effects on mental health and feelings of social isolation. In a longitudinal study of the acute phase and 6 months later in Austria, Pieh et al. (2021) found that the negative psychological consequences of the COVID-19 pandemic (i.e., depressive, anxiety, and insomnia symptoms) remained largely unchanged, despite a small improvement in perceived stress and well-being at 6 months. Other studies have also observed enduring negative mental health and behavioral consequences after a COVID-19 wave or during a period of phased reopening (Beutel et al., 2021; Haucke et al., 2021; Sherman et al., 2020). Perceived daily hassles are also relatively stable both during and after an acute phase of the pandemic (Hargreaves et al., 2021). After restrictive measures were eased, the majority of people continued to value self-isolation (Zhou et al., 2020), and people continue to report elevated levels of loneliness, relative to prepandemic levels (Killgore et al., 2020). Other studies suggest that, after the acute phase, people express concerns about the safety of social interaction, fear the end of restrictions, and plan to continue to isolate themselves (Fisher et al., 2021) and reduce activities (Hood et al., 2021). Although some studies suggest people's loneliness returns to prepandemic levels (Beutel et al., 2021; Seifert & Hassler, 2020) or shows no change (Luchetti et al., 2020), these findings were observed in the early phase of the pandemic. Moreover, these mixed findings are consistent with heterogeneous effects of the pandemic (Mancini, 2020; Prati & Mancini, 2021; Seifert & Hassler, 2020; Zhou et al., 2020).

There are several possible explanations for these persistent forms of disengagement and avoidance behavior. One explanation is that public health measures to control spread (e.g., lockdowns that restrict normative activity and social behavior) are primarily responsible. If this is the case, we would expect to observe differences in disengagement and avoidance behavior when comparing countries that had different lockdown policies. A second explanation is that generalized concerns about the pandemic are primarily responsible, and given the global nature of the pandemic, these effects would be present in most contexts. In such a case, a wide variety of factors could contribute to disengagement and avoidance behavior, including the presence of global economic stressors, worries about another COVID-19 surge, learned helplessness, and persistent feelings of isolation and loneliness due to intentionally withdrawing from social interactions to reduce risk of infection (Haucke et al., 2021; Pieh et al., 2021).

Overview of Pandemic Disengagement Syndrome

The aim of the current article is to develop and validate an instrument to investigate people's general disengagement after the acute phases of the pandemic (e.g., Fisher et al., 2021; Hood et al., 2021; Zhou et al., 2020). Such an instrument can greatly facilitate an assessment of the pandemic's long-term effects on mental health. Although social withdrawal may seem paradoxical given the restrictions on social behavior, there are sound reasons to propose that the resumption of social and other activities may be difficult for some people. Persistent activation of the behavioral inhibition system could reinforce and maintain vulnerabilities to specific broad dimensions of psychopathology (Gray, 1982). From the perspective of the hierarchical taxonomy of psychopathology (Kotov et al., 2017), for example, both the internalizing subfactor of fear, which involves social phobia, and the spectrum of detachment, which involves intimacy avoidance and avolition, could be at risk of being exacerbated by the pandemic. Together these concerns suggest that the pandemic may have unique effects on vulnerabilities to psychopathology, leading to detrimental outcomes with potential societal-level effects, given the worldwide impact of the pandemic.

On this basis, we created a new measure for the present study—the PDSS—that was designed specifically to assess disengagement from broad features of everyday life after an acute phase of the pandemic, including social withdrawal, feelings of Alienation in relation to close others, and general feelings of amotivation (Rubin & Barstead, 2018). In the first preregistered study (https://osf.io/23a7y), we developed an item pool based on published first-person reports of the pandemic's effects and investigated construct validity, internal consistency, measurement invariance, and convergent and discriminant validity of the instrument. To this end, we focused on two countries—Italy and the United States—that were among the most affected countries in the early phases of the pandemic. In the second study, we evaluated the test–retest reliability of the PDSS and predictive validity in a longitudinal sample of college students in the northeastern United States, an area that saw substantial cases during the early acute phase. In the third preregistered study (https://osf.io/pk2y6), we tested the link between disengagement syndrome and mandatory lockdowns, comparing two countries with broadly similar national cultures but substantially different lockdown policies (Sweden vs. Norway). This allowed us to assess whether disengagement syndrome is more likely attributable to prolonged exposure to the COVID-19 pandemic or whether lockdowns are primarily responsible for any observed effects. We also investigated the incremental validity of lockdown syndrome over and above social anxiety disorder, a theoretically relevant construct. Data, study materials, and all analyses codes are freely available at https://osf.io/u2m5y/.

Study 1

In Study 1, we focused on two countries, Italy and the United States, that were among the most affected countries during the first wave of the COVID-19 pandemic in 2020. Citizens' daily lives, including work, leisure activities, and social relationships substantially changed during the first wave of the pandemic in 2020. However, since the regions/states of the two countries were not equally affected, our sampling frame involved regions with different levels of objective COVID-19 risk in each country. The objective risk was based on regional differences in COVID-19 cases and death prevalence (high: New York and Lombardy; low: San Francisco and Campania). Given the lack of empirical findings to date, we did not make predictions related to the factorial structure of the PDSS. Instead, we used exploratory and confirmatory procedures to determine factor structure and then sought to establish reliability, convergent validity, and discriminant validity. We hypothesized that the PDSS would be moderately positively correlated with COVID-related worry, posttraumatic stress disorder symptoms (PTSD) symptoms, psychological distress, depression, anxiety ($r > .3$), social media use, and media exposure ($r > .2$), and moderately negatively correlated with assessments of well-being and social well-being ($r > -.3$), in-person social interaction frequency, and trust in institutions ($r > -.3$). We also expected that the PDSS would be relatively independent of political ideology and would show no differences across objective COVID-19 risk regions. Finally, we sought to investigate the measurement invariance of the PDSS across gender and countries. Previous studies suggest that both Italian and U.S. women consistently express higher concern, fear, and perceived risk of contracting COVID-19 and are more likely to adopt safety measures to protect themselves and others (e.g., the practice of social distancing) compared to men (Prati et al., 2021).

Method

Participants and Procedure

Approval from the Bioethics Committee of the University of Bologna and the Pace University institutional review board has been obtained for data collected in Italy and the United States, respectively. Based on the guidelines of MacCallum et al. (1999) regarding the minimum sample size in factor analysis, we determined that a sample size of 400 participants would obtain factor solutions that correspond closely to population factors and that are adequately stable even under conditions of weakly determined factors and low communalities. Using Monte Carlo data simulation techniques to evaluate sample size requirements for confirmatory factor analysis (CFA; Muthén & Muthén, 2002), we determined that a total of 400 participants is an appropriate sample (i.e., estimated power was greater than 80% with $\alpha = .05$) in one-, two-, three-, and four-factor models with factor loadings of .50 (Wolf et al., 2013). Since we planned to randomly split half of the sample, this required a total of 800 participants.

We recruited participants through survey firms in the United States and Italy. All surveys were self-administered and confidential. Two ongoing panel studies conducted by Demetra and Qualtrics were used to recruit and assess participants in Italy and the United States, respectively. To account for differences in objective COVID-19 risk, two regions that were differentially affected by the outbreak were selected within each country. In the United States, we selected a sample of people that resided in a zip code in one of the five boroughs of New York City (high-affected region) and was representative of New York City demographics. A matched demographic sample was chosen based on people living in a zip code in the County of San Francisco (low-affected region). In Italy, we used similar procedures to recruit participants in Lombardy (high-affected region) and Campania (low-affected region).

In the United States, the survey was available from July 29 to September 10. Five hundred and fifty-eight participants completed the surveys, and Qualtrics removed 143 participants because of unreliable survey completion times (25.6%). Thus, a sample of 415 complete surveys ($n = 210$, New York; $n = 205$, San Francisco) was used for the present study. The characteristics of the final sample in New York City approximated the 2019 census data (U.S. Census Bureau, 2019). Specifically, respondents were 50.7% female (U.S. census = 52.3%), 49.1% at least college-educated (U.S. census = 38.1%), 40% White (U.S. census = 42.7%), 23.3% Black (U.S. census = 24.3%), 19% Latino/Hispanic (U.S. census = 29.1%), 14.3% Asian (U.S. census = 14.1%), 2.9% multiracial (U.S. census = 3.6%), and reported a median income of $75,000–99,000 (U.S. census = $63,998). There were no differences in gender, age, non-White ethnicity, income, or education between the New York and the San Francisco sample ($p_s > .05$).

In Italy, the survey was available from July 31 to September 2, and 511 participants completed surveys. The survey firm Demetra removed 56 participants because of unreliable survey completion times (11.0%). Therefore, a sample of 455 complete surveys ($n = 231$, Lombardy; $n = 224$, Campania). Participants from Lombardy approximated 2019 population-level demographics in Italy (ISTAT, 2021). Specifically, respondents reported a mean age of 45.55 (Istituto Nazionale di Statistica [ISTAT] = 45.0), and were 51.1% female (ISTAT = 51.0%), 3.9% nonnative (ISTAT = 11.5%), and 88.3% at least college-educated (ISTAT = 84.9%). No differences in gender, age, education, or immigration status were found between the Lombardy and the Campania samples ($p_s > .05$). However, the Lombardy and the Campania sample differ in income ($p = .008$).

Scale Development

To identify scale content, we used an inductive process involving first-person accounts of responses to the lockdowns and the aftermath of the pandemic. Both authors read a series of blogs, social network posts, online articles, and scientific articles (e.g., Brooks et al., 2020; Cava et al., 2005; Reynolds et al., 2008; Zhou et al., 2020) to assess the phenomenological consequences after the acute phase of an epidemic. Based on these readings, our investigation was further informed by relevant dimensions in the hierarchical taxonomy of psychopathology, focusing on the subfactor of fear and the spectrum of detachment (Conway et al., 2019). Together, this material was used to develop relevant content and then to formalize specific item questions. In some cases, items were derived directly from first-person accounts. G.P. generated an initial set of items in English, and A.D.M suggested revisions and added additional items. In an iterative process, this resulted in a pool of 16 items (see Supplemental Table 1, for full item content)

Instrument

PDSS. Participants rated 16 items based on how accurately it described them from 1 (*not at all accurate*) to 5 (*very accurate*). Specifically, participants were asked "How accurately do each of the following sentences describe the way you feel right now?" For example, "Currently, I prefer to talk to friends using the Internet or the telephone rather than face-to-face." These items were subsequently factor analyzed to identify a final version of the scale.

General Distress. We measured general distress using the 21-item Depression Anxiety Stress Scales (DASS-21; (Bottesi et al., 2015; Lovibond & Lovibond, 1995). The DASS-21 consists of three seven-item subscales assessing depression, anxiety, and stress. Items are rated on a 4-point scale from *did not apply to me at all* to *applied to me very much, or most of the time* based on the past 2 weeks. Higher scores indicate more serious syndromes of depression, anxiety, and distress. Cronbach's alpha coefficients for distress ($\alpha = .91$), depression ($\alpha = .94$), and anxiety ($\alpha = .94$) were excellent.

Posttraumatic Stress Disorder Symptoms. We assessed PTSD symptoms using the eight-item PTSD-8 Inventory (Hansen et al., 2010). Participants rated how often they experienced PTSD symptoms in relation to the COVID-19 pandemic in the last month (e.g., "Sudden emotional or physical reactions when reminded of the COVID-19 pandemic.") These items were rated on a 4-point scale ranging from 1 (*not at all*) to 4 (*all the time*). Given that the PTSD-8 Inventory was viewed as unidimensional, we computed an average score for an estimate of overall PTSD symptom severity (higher scores indicate more severe PTSD symptoms). Internal consistency was satisfactory ($\alpha = .89$).

Well-Being. We used the five-item World Health Organization-Five Well-Being Index (Topp et al., 2015) as a short self-reported measure of current mental well-being. Participants rated how well each of the five statements applied to them in the last 14 days on a 5-point scale from 0 (*none of the time*) to 5 (*all of the time*). We computed an average score of mental well-being (higher scores represent greater well-being). Internal consistency was satisfactory ($\alpha = .93$).

Social Well-Being. The Social Well-being Scale of the mental health continuum–short form was used (Keyes, 2006; Petrillo et al., 2015). The Social Well-Being Scale consists of six items corresponding to feelings of social well-being. We asked American participants to rate the frequency of every feeling of social well-being in the past month on a 6-point scale ranging from 1 (*never*) to 6 (*always*), while Italian participants provided their responses on a 5-point scale ranging from 1 (*never*) to 5 (*always*). We computed a mean–item score by averaging responses and, to make comparable responses collected in Italy and the United States, we standardized the scores (the average score is given a value of zero). Cronbach's alpha was satisfactory for both the U.S. and Italian samples ($\alpha = .88$ and .81, respectively).

COVID-19 Worry. COVID-19 worry was assessed using a five-item scale derived from previous research on the pandemic influenza H1N1 2009 (Prati et al., 2011a, 2011b). Items were as follows: (a) "Do you think you are at risk of contracting coronavirus?" (b) "How often do you worry about contracting coronavirus?" (c) "How often do you worry about the future in relation to the Coronavirus pandemic?" (d) "How often do you worry about finances in relation to the Coronavirus pandemic?" and (e) "How often do you worry about school disruptions related to the Coronavirus pandemic?" Participants rated the extent to which each item applied to them over the past 2 weeks using a 7-point scale (1 = *not at all*, 7 = *extremely*). We derived a mean–item score by averaging responses. Higher scores correspond to greater worry about the epidemic of COVID-19. Cronbach's α was .77.

Institutional Trust. We assessed trust in institutions regarding the COVID-19 pandemic using a four-item scale (Prati, 2021). Specifically, we asked participants about their trust in local and regional/state authorities (two questions for each authority) to deal with the pandemic: "Do you think the local authorities/state government are competent in dealing with the COVID-19 pandemic?" and "Do you trust local authorities/state government to keep you safe during the pandemic?" Participants rated these questions on a 7-point scale (1 = *not at all*, 7 = *extremely*). Responses were averaged to derive a mean–item score. Higher scores correspond to greater institutional trust in the government's ability to address the challenges of the COVID-19 pandemic. Cronbach's α was satisfactory ($\alpha = .89$).

Media Exposure. To measure the consumption of media related to the COVID-19 pandemic, we asked participants three separate questions about how much time they have spent per day: (a) reading news articles, (b) watching news shows, or (c) watching online videos about the pandemic. Participants rated each item using a 7-point scale (from 1 = *not at all* to 7 = *3 hr or more*). We averaged responses to compute a mean–item score. Higher scores corresponded to more consumption of COVID-19-related media. Cronbach's α was adequate ($\alpha = .78$).

Social Media Use. Participants' social media use was measured using three separate questions about how much time they have spent per day on Facebook, Twitter, and Instagram/Snapchat over the past 2 weeks. For each question, participants were asked to indicate their response on a 7-point scale from 1 (*not at all*) to 7 (*extremely*). We averaged responses to derive a mean–item score. Higher scores represent greater social media use. Internal consistency was acceptable ($\alpha = .76$).

Social Interaction Frequency. Four items were used to measure social interaction frequency as follows: (a) Talking on the phone or video chats with family members per day? (b) Talking on the phone or video chats with friends per day? (c) Interacting in person with friends per day? (d) Interacting in-person with family per day? Participants were asked to rate the questions over the past 2 weeks on a 7-point scale from 1 (*not at all*) to 7 (*3 hr or more*). We averaged responses to derive a mean–item score. Higher scores correspond to greater social interaction frequency. Internal consistency was just below the conventional cut-off of .70 ($\alpha = .67$).

Results and Discussion

Preliminary Analyses

Data were first screened for normality, outliers, and missing data. No missing data were found. The median absolute deviation indicated the absence of influential outliers (Leys et al., 2019). Mardia (kurtosis and skewness), Henze–Zirkler, and Doornik–Hansen multivariate tests were statistically significant ($p < .001$), indicating the data were multivariate nonnormal. Visual examination revealed a floor effect for some items. Floor effects are not uncommon when investigating mental health outcomes. Therefore, analyses took into account multivariate nonnormality.

Exploratory Factor Analysis

To address the factorial structure of the scale, we first conducted an exploratory factor analysis on a randomly split half of the sample. Ordinal factor analysis (i.e., using raw-data matrix of polychoric correlations; Basto & Pereira, 2012) was conducted using the principal axis factor followed by an oblique method of rotation (i.e., oblimin). Bartlett's test of sphericity values justified the application of exploratory factor analysis, $\chi^2(2,120) = 5,087.375$, $p < .001$. The Kaiser–Meyer–Olkin measure of sampling adequacy was .934, indicating that the sampling was adequate. Estimates of communalities were above .35. Both parallel analysis and minimum average partials indicated two factors. In addition, Kaiser's rule revealed two factors. The first and the second factors explained 51.30% and 9.15% of the variance, respectively, for a total of 60.44 variance explained.

Supplemental Table 1 displays the results from the factor analysis for all PDSS items. One item (No. 8) was dropped because it did not load saliently on only one factor (crossloading). Factor 1 (Social Avoidance) encompassed eight items with factor loadings ranging from .83 to .56, while Factor 2 (Alienation) comprised seven items with factor loadings ranging from .87 to .60. The term *Alienation* was chosen because these items encompass feelings of isolation from others, powerlessness, and amotivational states. These subjective states constitute major dimensions of a classical formulation of Alienation (Kanungo, 1979; Seeman, 1991) and are consistent with the detachment spectra of the hierarchical taxonomy of psychopathology (HITOP; Conway et al., 2019). For the second factor, we chose the term *Social Avoidance* because the items indicate a motivated reduction in social interaction, fear of direct in-person interaction because of COVID-19, generalized fears of public places, and loss of social connections (Cumming & Henry, 1961). These subjective states are consistent with classical formulations of Social Avoidance in which a subtype is social withdrawal (Asendorpf, 1990), and they are also consistent with the fear spectra of HITOP and its associated subfactors of agoraphobia and specific phobia, here in relation to COVID-19 (Conway et al., 2019).

Confirmatory Factor Analysis

Based on the resulting factor structure (i.e., two factors), we then conducted a CFA on the other half of the sample. We used a robust weighted least squares estimator using a diagonal weight matrix in Mplus Version 8.7. The fit of a two-correlated factor model using the other half of the sample was satisfactory, $\chi^2(89) = 519.892$, $p < .001$; Tucker–Lewis index (TLI) = .94; comparative fit index (CFI) = .95; standardized root-mean-square residual (SRMR) = .053. A series of confirmatory factor analyses using the whole sample was performed to determine whether the instrument is best represented by a higher order factor model (one superordinate factor and two subordinate factors), a bifactor model, or a two-correlated factor model (two latent variables which were allowed to correlate). The fit of the higher order factor model was not satisfactory, $\chi^2(90) = 6,240.979$, $p < .001$; TLI = .64; CFI = .59; SRMR = .140. The fit of the two-correlated factor model (see Supplemental Table 2) was satisfactory, $\chi^2(89) = 826.415$, $p < .001$; TLI = .95; CFI = .96; SRMR = .045. Compared to that of the two-correlated factor model, the fit of the bifactor model (see Supplemental Table 2) was slightly better, $\chi^2(75) = 573.458$, $p < .001$; TLI = .96; CFI = .97; SRMR = .029. Therefore, we selected a bifactor model as the optimal factor structure of the PDSS. We then computed bifactor indices such as explained common variance = .70 (general factor), omega = .95, omegaS = .93 (Social Avoidance) and .91 (Alienation), hierarchical omega (omegaH) = .80, percentage of uncontaminated correlations = .53, factor determinacy = .94 (general factor), construct replicability = .93 (general factor), and average relative parameter bias = .15. The indices revealed an acceptable level of parameter bias and the presence of some multidimensionality that was not serious enough to preclude the use of the total score.

In sum, these analyses provided evidence of factorial coherence for a novel construct in response to the COVID-19 pandemic. This overarching construct encapsulated social withdrawal, amotivation, and feelings of Alienation in relation to close others. Exploratory factor analyses indicated that this measure was best characterized by two factors, while confirmatory factor analyses revealed that the two-correlated factors model had a good fit and a bifactor solution resulted in an improved fit.

Measurement Invariance Across Gender and Countries

We used multiple group analyses to test measurement invariance across countries and gender. To this end, we first specified an unconstrained model in which the same items load on the same factors in both groups (configural model). We then compared the fit of the configural model to a model which constrains corresponding factor loadings to be comparable (equivalent) across groups (metric model). Finally, we tested the metric model fit to one in which the factor loadings and items are comparable (equivalent) across groups (scalar model). Based on Monte Carlo studies (Chen, 2007), we determined that a cutoff of $\Delta CFI \geq -0.01$ supplemented by a criterion requiring a change of $\geq .010$ in SRMR can be used to conclude whether a measure can be considered invariant.

The configural model for gender had an adequate fit across men and women, $\chi^2(158) = 630.453$, CFI = .973, SRMR = .034. Subsequently, we tested a metric invariance model which had adequate fit, $\chi^2(176) = 580.837$, CFI = .977, SRMR = .034. The small change-in-model fit indices, $\Delta CFI = -.004$, $\Delta SRMR < .001$, supported metric invariance. Next, a threshold invariance model was tested and had a satisfactory fit, $\chi^2(219) = 620.930$, CFI = .977, SRMR = .036. The change in model fit indices was very small, $\Delta CFI < -.001$, $\Delta SRMR = .002$, indicating the fit of the threshold model did not significantly differ from that of the metric model. Thus, we found support for threshold invariance across gender.

To test measurement invariance across countries, we first established configural invariance by examining model fit. The configural model had adequate, albeit borderline, fit across countries, $\chi^2(158) = 644.966$, CFI = .973, SRMR = .038. The metric model was then fitted and compared to the configural model. The metric model indicated acceptable fit, $\chi^2(176) = 823.644$, CFI = .964, SRMR = .042. Change in fit indices provided evidence for metric invariance across countries, $\Delta CFI = -.009$, $\Delta SRMR = .004$. Finally, a threshold model invariance was evaluated and compared to the metric model; the fit of this model was acceptable, $\chi^2(219) = 1,118.477$, CFI = .950, SRMR = .051. However, the change did not provide support for scalar invariance across countries, $\Delta CFI = -.014$, $\Delta SRMR = .009$. We tested partial invariance by freeing the threshold of Item 12, which accounted for a large source of misfit. After freely estimating the threshold of Item 12, partial scalar invariance was supported, $\Delta CFI = -.010$, $\Delta SRMR = .007$. In addition, we tested threshold invariance across gender after freeing the threshold of Item 12. The threshold invariance model with threshold of Item 12 freely estimated was tested and had a satisfactory fit, $\chi^2(215) = 622.235$, CFI = .977, SRMR = .035. The change in model fit indices was very small, ΔCFI < .001, ΔSRMR = .001, indicating support for threshold invariance across gender after freeing the threshold of Item 12.

The present study established full measurement invariance across gender and partial measurement invariance across countries. Full measurement invariance across gender implies that the PDSS can be interpreted in a conceptually similar manner by men and women. Therefore, practitioners and researchers employing PDSS can compare scores meaningfully across gender. Invariance across countries revealed some systematic variability regarding one item (Item 12). Although scores on general distress and COVID-19 worry were not significantly different between Italy and the United States, people from the United States reported more PTSD symptoms, institutional trust, and media exposure during the pandemic (Mancini & Prati, 2022). Therefore, such systematic variability regarding Item 12 ("I worry that people don't take the pandemic as seriously as I do") may reflect a small difference in the experience of the pandemic among people from Italy and the United States. Researchers can employ the correctly specified partial invariance model to compare countries on latent means or variances (Luong & Flake, 2022). Taken together, these results showed that the PDSS demonstrated a similar meaning, interpretation, and level, regardless of gender and country.

Discriminant Validity of the PDSS

We used the Confidence intervals in confirmatory factor analysis(sys) technique to assess discriminant validity (Rönkkö & Cho, 2020). Specifically, we computed correlation coefficients (Spearman's rho) and the upper and lower limits of the 95% confidence intervals (CIs) of the standardized factor correlations using CFA (ρ_{CFA}) between PDSS and its subscales and study variable. Table 1 reports correlation coefficients (Spearman's rho) among PDSS and its subscales and study variables. The absolute values of the 95% upper/lower limit did not exceed the recommended cutoff of .80, thereby indicating no evidence of a problem with discriminant validity (Rönkkö & Cho, 2020). According to the guidelines of Cohen (1988), correlation coefficients in the order of .10, .30, and .50 are considered "small," "medium," and "large," respectively, in terms of the magnitude of effect sizes.

Table 1

Correlations (Spearman's Rho) Among Pandemic Disengagement Syndrome Scale (PDSS) and Its Subscales and Study Variables

	PDSS	Social avoidance	Alienation
Variable	r_s	r_s	r_s
Well-being	−.10**	−.03	−.16***
Social well-being	.00	.01	−.02
Stress (DASS-21)	.32***	.13***	.49***
Anxiety (DASS-21)	.33***	.13***	.51***
Depression (DASS-21)	.34***	.13***	.52***
PTSD	.57***	.44***	.59***
COVID worry	.53***	.49***	.46***
Trust	.18***	.22***	.08*
Social media use	.18***	.09**	.26***
Media exposure	.29***	.24***	.29***
Social interaction frequency	−.08*	−.15***	.03
PDSS	—	—	—
Social avoidance	.90***	—	—
Alienation	.88***	.60***	—

Note. CI = confidence interval; DASS-21 = Depression Anxiety Stress Scales–21; PTSD = posttraumatic stress disorder symptoms.
[a] Upper limit (for positive correlations) or lower limit (for negative correlations) of the 95% CIs of the estimated factor correlations.
* $p < .05$. ** $p < .01$. *** $p < .001$.

Mental Health Symptoms. Correlation coefficients (Spearman's rho) analyses were largely consistent with our preregistered hypotheses. The PDSS and its dimensions were significantly and positively related to COVID-related worry, PTSD symptoms, psychological stress, depression, and anxiety. The size of these correlation coefficients was medium to large except for those involving Social Avoidance. The size of the correlation coefficients between Social Avoidance and stress, anxiety, and depression were small, albeit significant.

Positive Functioning. Except for Social Avoidance, the PDSS and Alienation were significantly and negatively related to well-being. As expected, the size of the correlation coefficients was small. Unexpectedly, Social Avoidance was not related to social well-being. In addition, contrary to our expectations, social well-being was not related to the PDSS and its dimensions.

Media Use. The PDSS and its dimensions were significantly and positively related to trust, social media use, and media exposure. The size of the correlation coefficients involving trust and social media use was relatively small, whereas that of media exposure was small to medium.

Social Functioning. Consistent with the theoretical construct, social interaction frequency was negatively related to the overall PDSS, to Social Avoidance but not to Alienation.

COVID-19 Prevalence. We used a robust independent samples t test (i.e., Yuen's t test, the trim portion was set to 0.2) to investigate whether the scores of the PDSS were different between regions with high and low objective COVID-19 risk. Results revealed nonsignificant differences between high and low affected regions on PDSS, Yuen's $t(517.9) = 0.949$, $p = .343$, $\xi = .05$, Social Avoidance, Yuen's $t(515.8) = 1.436$, $p = .152$, $\xi = .07$, and Alienation, Yuen's $t(519.9) = 0.072$, $p = .942$, $\xi = .01$. Together, these findings indicate that disengagement after an acute pandemic phase is not a function of objective indicators of risk. Instead, fear, worries, perceptions, and beliefs related to the pandemic may influence subsequent disengagement irrespective of objective indicators of risk.

Political Identification. The PDSS showed an unexpected relation to political ideology, as found in a comparison of means across levels of ideology (Supplemental Figure 1), $F(4, 704) = 5.780$, $p < .001$, $\omega^2 = .03$. Post hoc tests revealed that liberal people reported higher scores on PDSS than slightly liberal ($p = .035$) and slightly conservative people ($p < .001$). Moreover, slightly conservative people reported lower scores on PDSS than conservative people ($p = .006$) and moderate people ($p = .007$). Political party affiliation may help to explain risk perception and differential engagement in COVID-related health behaviors, such that liberal people are more likely than conservative people to report higher risk perception of COVID-19, mask use, and social distancing (e.g., Barbieri & Bonini, 2021; Bruine de Bruin et al., 2020; Calvillo et al., 2020). The present findings suggest that political identification has a complex relationship to PDSS.

In sum, the dimensions of the PDSS showed expectable positive relations with anxiety, stress, depression, COVID-19 worry, PTSD symptoms, social media, and media exposure, as well as negative links with well-being. We also found preliminary evidence for the factor structure, reliability, and construct validity of the pandemic disengagement syndrome. However, we were unable to assess test–retest reliability and predictive validity because of the cross-sectional data.

Study 2

In Study 2, we used an undergraduate sample in the fall of 2020 from the United States to further examine the psychometric properties of the PDSS. Specifically, to assess test–retest reliability, we examined whether PDSS showed stability over two time points. We also assessed predictive validity in a longitudinal analysis. We anticipated that PDSS would predict subsequent loneliness, depression, and social support 1 month later, holding constant baseline levels of each outcome variable.

Method

Participants and Procedure

Participants were 193 undergraduate and master's psychology students from the United States who completed online surveys for course credit using the Qualtrics platform at two time points spaced approximately 1 month apart during the fall semester of 2020 (Mdn_{baseline} = October 30th; Mdn_{wave1} = November 24th). Participants completed surveys on pandemic social withdrawal syndrome, mental health functioning, social functioning, personality, interpersonal trust, and other variables. All surveys were approved by the Pace University institutional review board (No. 1592051). All participants had to be 18 years of age or older.

Participants were 18.72 years of age on average (SD = 1.51), 64.2% female, and 53.4% reported their ethnicity as White, 23.8% as Latino, 13.5% as Black/African American, 5.2% as Asian American, .5% as Native American, and 3.6% as multiethnic or other. The median family income was $80–99,000. No eligible participants were excluded from the analysis. Missing data at Wave 1 (W1) were 35.4% (available n = 124). The order of surveys was randomized for each participant. However, due to an error in specifying the number of scales for randomization (there were 17 total scales but only 14 were identified for randomization), some scales were inadvertently omitted by design in each wave. Although this reduced the sample by 30 additional participants, these cases were missing completely at random (MCAR), because of the randomized omission of scales. Specifically, Little's MCAR test was not significant, $\chi^2(1{,}188)$ = 1,151.793, p = .769. Thus, 94 participants had complete data for the PDSS at baseline and W1. A fully conditional specification multiple imputation procedure (n = 10) was used to handle missing data. Our sample size was based on the available participant pool for the fall semester, but the samples were sufficient to detect small to medium-sized effects in our correlational (r = .20) and regression analyses (f^2 = .10) at adequate power (>80%), according to conventional methods (G^*Power Version 3.1; Faul et al., 2009). The data file for the present study will be made available.

Measures

PDSS. The same scale was used for Study 2. Cronbach's alpha for the scale and its subscales are reported in the Supplemental Table 3.

Depression Symptoms. We used the short 11-item Center for Epidemiological Studies Depression Scale (Radloff, 1977). Scores range from 0 to 33 with respondents indicating how frequently they have experienced symptoms during the "past week or so" on a 4-point scale ranging from *not at all or less than 1 day* to *most or nearly every day*. Higher scores indicate greater depressive symptom severity. Internal consistency was satisfactory at W1 (α = .81) and Wave 2 (W2; α = .85).

Loneliness. We used a 13-item version of the UCLA Loneliness scale (Russell et al., 1980). The scale measures global feelings of social isolation and lack of companionship (such as feeling "that there is one you can turn to"). Six items were reverse-scored (such as feeling "that you are part of a group of friends"). Items were rated on a 4-point Likert-type scale ranging from 1 (*never*) to 4 (*often*). Reliability was high at W1 (α = .92) and W2 (α = .90).

Social Support. We used the 12-item Multidimensional Perceived Social Support Scale (MSPSS) to measure individuals' perceptions of social support from three different sources: friends, family, and significant others (Zimet et al., 1988). The MSPSS is rated on a 5-point Likert-type scale ranging from 1 (*strongly disagree*) to 5 (*strongly agree*). Reliability was high at W1 (α = .91) and W2 (α = .92).

Five-Factor Model of Personality. We used a brief 10-item scale of the five-factor model personality (Gosling et al., 2003). Two-item scales were used to measure agreeableness, openness, extraversion, conscientiousness, and neuroticism.

Results and Discussion

Stability

First, we calculated correlations between two assessments of the same measure taken at two separate time-points as an indication of the relative agreement between the two measurements. Test–retest correlations (Supplemental Table 3) ranged from .76 (Social Avoidance) to .82 (total scale). Correlation coefficients assess the linearity of the relation between the two measurements but cannot establish the equality of individual values between the two measurements. Therefore, a large correlation between two measurement points is a necessary but insufficient condition to establish an agreement.

To demonstrate agreement, we calculated intraclass correlation coefficients (ICCs) using a two-way random effects model for absolute agreement that corresponds to the ICC (2, k) in the nomenclature proposed by Shrout and Fleiss (1979). The test–retest reliability coefficients (ICC) for the full and its subscales were as follows: .90, 95% CI [.86, .92], for the Full Total scale; .87, 95% CI [.83, .90], for the Social Avoidance subscale; and.87, 95% CI [.83, .90], for the Alienation subscale. According to Cicchetti (1994), ICC values between. 75 and 1.00 can be interpreted as excellent. Therefore, the PDSS exhibited adequate stability over time.

Discriminant Validity

Table 2 displays the correlation coefficients between PDSS and its subscales and the five-factor model personality, loneliness, social support, and depression at W1. Except for a significant (small to medium) correlation between neuroticism and Alienation, the PDSS and its subscales were not significantly associated with extraversion, agreeableness, conscientiousness, openness, and neuroticism. The

Table 2

Correlations (Spearman's Rho) and Upper Limits (or Lower Limits for Negative Correlations) of the 95% CIs of the Estimated Factor Correlations Among Pandemic Disengagement Syndrome Scale (PDSS) and Its Subscales, Five Factor Model Personality, Loneliness, Social Support, and Depression at Wave 1 (W1)

	PDSS		Social avoidance		Alienation	
Variable	r_s	95%L[a]	r_s	95%L[a]	r_s	95%L[a]
Extraversion	−.12	−.45	−.11	.39	−.13	.46
Agreeableness	−.10	−.35	−.07	−.42	−.11	−.52
Conscientiousness	−.09	−.49	−.03	−.28	−.16	−.55
Openness	.17	.57	.11	.26	.21*	.40
Neuroticism	.15	.53	.08	.35	.24*	.70
Loneliness	.45*	.62	.32*	.43	.52*	.70
Social support	−.20*	−.33	−.10	−.13	−.31*	−.45
Depression	.45*	.65	.34*	.43	.51*	.58

Note. CI = confidence interval.
[a] Upper limit (for positive correlations) or lower limit (for negative correlations) of the 95% CIs of the estimated factor correlations.
* $p < .05$.

magnitude of the correlation between PDSS and its subscales and social support was small to medium. Medium to large effect sizes were found in the correlation coefficients between PDSS and its subscales and loneliness and depression.

To assess discriminant validity using the Confidence intervals in confirmatory factor analysis(sys) technique (Rönkkö & Cho, 2020), we calculated the upper and lower limits of the 95% CIs of the standardized factor correlations between PDSS and its subscales and loneliness, social support, and depression at W1. The 95% upper limit for positive correlation and the lower limit for negative correlation pairs were <.80, that is the cutoffs proposed by Rönkkö and Cho (2020). These results provided evidence of discriminant validity.

Predictive Validity

We used four ordered polytomous regression analyses to examine whether PDSS longitudinally at W1 predicted relevant outcomes (i.e., loneliness, social support, and depression) at W2 holding constant baseline (W1) values of the outcome. Results revealed that PDSS longitudinally predicted social support, $b = -0.21$, $SE = 0.05$, $p < .001$, 95% CI [−0.32, −0.11], $f^2 = .07$, loneliness, $b = 0.21$, $SE = 0.09$, $p = .021$, 95% CI [0.03, 0.38], $f^2 = .05$, and depression, $b = 0.34$, $SE = 0.08$, $p < .001$, 95% CI [0.18, 0.51], $f^2 = .13$, at W2 holding constant baseline (W1) values of the outcome.

In sum, we found that the PDSS and its subscales demonstrated excellent test–retest reliability and discriminant validity. Moreover, pandemic disengagement syndrome longitudinally predicted increases in depression and loneliness and decreases in social support, controlling for baseline values of those variables. This finding further supported the predictive validity of the PDSS.

Study 3

Study 3 examined whether disengagement syndrome was specifically linked with lockdowns and whether it had incremental validity over and above social anxiety. To address this possibility, we conducted a quasi-experimental comparison between Norway and Sweden, two countries with similar national cultures but with markedly distinct approaches to the pandemic. In Sweden, no national lockdowns were imposed, schools remained open, and masks remained optional. These policies were implemented at the beginning of the pandemic and remained in place throughout. By contrast, Norway implemented a full lockdown early in the pandemic, closed schools, and had mandatory mask requirements. Thus, we recruited demographically similar participants in Norway and Sweden and compared their levels of pandemic disengagement syndrome, depression, and social anxiety. Our preregistered hypothesis was that lockdown syndrome would be higher in Norway than in Sweden, controlling for depression and social anxiety (see https://osf.io/pk2y6). We also explored whether PDSS incrementally predicted depression over and above social anxiety symptoms, a theoretically similar construct.

Method

Participants and Procedure

Participants were recruited through the Prolifics online survey platform. We specified Norway or Sweden as the country of origin and asked participants to identify their country of residence. Data were collected from March 31 to April 29, 2021 (*Mdn* = April 2). Participants were paid the equivalent of $10 per hour for a survey that took approximately 5 min per person. According to our preregistered power analysis, we determined the sample size based on an anticipated small-sized effect for a general linear model ($f^2 = .05$) with three independent variables and the following specifications: statistical power = 80% and $p < .05$. We conducted power analysis using the *R* package "pwr." The required sample size was 220 participants. We sought to exceed this sample somewhat to increase power and to account for data lost because of exclusions. We recruited a sample of 327 participants. However, some participants indicated they were currently residing in a country other than Norway and Sweden (80 participants), leaving a final sample of 237 participants (n = 143 Sweden participants; n = 137 Norway participants). Norway participants were mostly male (65%), and were aged 18–25 (50.4%), 26–34 (29.2%), 35–44 (14.6%), 45–54 (2.9%), and 65–74 (0.7%). Median income was 426,000–635,000 Krona. Full-time employment was 41.6%, part-time 16.8%, and unemployed 39.4% of the sample. Sweden participants were mostly male (60.8%), and aged 18–25 (37.8%), 26–34 (39.9%), 35–44 (12.6%), 45–54 (7.0%), and 65–74 (2.8%). Median income was 211,000–425,000 Krona. Full-time employment was 37.8%, part-time was 21.0%, and unemployed were 41.3% of the sample. Across the two countries, there were no differences in gender ($\chi^2 = 2.5$, $p = .28$) or employment status ($\chi^2 = 2.5$, $p = .63$), but Swedish participants were older than Norwegian participants ($t = 2.13$, $p = .03$), and Norwegian participants reported more income than Swedish participants ($t = -4.81$, $p < .001$). Therefore, as preregistered, we controlled for age and income in all analyses. Because most people in Sweden and Norway speak English fluently, surveys were administered in English. All surveys were approved by the Pace University institutional review board (No. 1592051-2). All participants had to be 18 years of age or older.

Measures

PDSS. The same scale was used. Reliability was high for both Norway ($\alpha = .81$) and Sweden ($\alpha = .86$).

Social Anxiety. We used a short six-item version of the Social Interaction Anxiety Scale (Le Blanc et al., 2014). The scale measures anxiety related to social interaction and fears of being scrutinized in social or performance situations (e.g., "I have difficulty making eye contact with others" and "I am tense mixing in a group"). Items are scored on a 5-point Likert scale from 0 (*not at all a characteristic or true of me*) to 4 (*extremely characteristic or true of me*). Internal consistency was high for both Norway ($\alpha = .87$) and Sweden ($\alpha = .88$).

Depression Symptoms. We used the four-item Patient Health Questionnaire-4 to measure depression symptoms (Löwe et al., 2010). The scale measures dysphoric emotion ("I am downhearted and blue") and anhedonia ("I was unable to become enthusiastic about anything"). Items are scored on a 4-point Likert scale from 0 (*not at all*) to 4 (*applied to me very much or most of the time*). Internal consistency was high for both Norway ($\alpha = .91$) and Sweden ($\alpha = .88$).

Results and Discussion

Preliminary Analyses

We first screened data for normality, outliers, and missing data. We found missing data for three participants (1.1%). Given the low proportion of missing data, we used pairwise deletion (Newman, 2014). No influential outliers were observed based on the median absolute deviation (Leys et al., 2019). Mardia's Skewness and Kurtosis test as well as the Doornik–Hansen test were significant ($p < .001$), while the Henze–Zirkler multivariate test was not statistically significant ($p = .490$). Visual examination indicated some floor effects. Taken together, these findings suggested that the data could be multivariate nonnormal. Therefore, we decided to test country differences in latent means using a robust weighted least squares estimator using a diagonal weight matrix in Mplus 8.7.

Comparison of the PDSS Scores Between Norway and Sweden

We used CFA and multiple group analysis. First, we conducted a series of increasingly stringent nested model comparisons to test configural, metric, and scalar invariance. Once we obtained partial or full measurement invariance based on the criteria provided by Chen (2007), we could conclude that scores are comparable across groups, and, therefore, we evaluated latent mean differences. To assess latent mean differences, we used the value of the critical ratio (CR) which is calculated by dividing the parameter estimate by its standard error (values larger than 1.96 represent statistically significant differences).

In evaluating measurement invariance based on countries (i.e., Norway or Sweden), the configural model was a good fit to the data, $\chi^2(158) = 305.435$, CFI = .949, SRMR = .068. Subsequently, we compared the fit indices of the metric model to that of the configural model. The changes in approximate fit indices revealed no meaningful decrement in fit, ΔCFI $\leq -.001$, ΔSRMR = .004. Comparing the fit of a threshold model to the metric model, we found evidence of full scalar invariance, ΔCFI = .003, ΔSRMR = .004. Measurement invariance indicated that the PDSS could be interpreted in a conceptually similar manner by respondents in Norway and Sweden. Measurement invariance is a necessary condition to test differences in latent means across the two countries. Indeed, the establishment of measurement invariance indicated that the mean differences of the PDSS can be compared directly. No significant country differences on latent means were found for PDSS, CR = -0.363, $p = .717$, Social Avoidance, CR = 0.645, $p = .519$, and Alienation, CR = -0.958, $p = .338$. Thus, Norway did not exhibit higher scores on PDSS than Sweden participants.

Incremental Validity

A final analysis examined whether disengagement syndrome incrementally predicted depression symptoms, over and above social anxiety. The PDSS correlated significantly with both depression ($r_s = .50$, $p < .001$) and social anxiety ($r_s = .39$, $p < .001$). We conducted a hierarchical multiple regression analysis, entering control variables for age and income on the first step, social anxiety on the second step, and disengagement syndrome on the third step. The first step of this analysis was significant, $F(2, 271) = 9.33$, $p \leq .001$, $R^2 = .06$, and both age ($b = -.17$, $SE = .05$, $p < .001$) and income ($b = -.07$, $SE = .05$, $p < .001$) were inverse predictors of depression. The second step also contributed additional significant variance to depression, $F_{\text{change}}(1, 270) = 67.30$, $p \leq .001$, $R^{2\text{change}} = .19$, indicating social anxiety was also a strong predictor of depression. More important, in a third step, forced entry of disengagement syndrome significantly improved model fit, $F_{\text{change}}(1, 269) = 50.30$, $p \leq .001$, $R^{2\text{change}} = .12$, indicating that disengagement syndrome uniquely explained 12% of the variance in depression, net of age, income, and social anxiety. The inclusion of country as a control variable did not alter the findings.

The absence of significant country-level differences in pandemic disengagement provides an important conceptual contribution to our understanding of the PDSS and the underlying construct. Specifically, the evidence suggested that pandemic disengagement syndrome is embedded in core psychological responses to the pandemic and does not depend on mandatory restrictions on social behavior. It should be noted that in the first wave of the pandemic, the level of voluntary protective avoidant behavior was high among European countries (including Sweden), independent of the stringency and severity of policies (Jørgensen et al., 2021). Therefore, Social Avoidance behaviors may have occurred to a similar degree in the two countries, irrespective of the mandatory or voluntary nature of lockdowns.

In sum, these findings imply that disengagement syndrome is likely to be present in a wide variety of contexts, not merely those subject to lockdown orders. This further supports the idea that disengagement syndrome is attributable to the pandemic's broader consequences for economic, political, social, and health functioning (Haucke et al., 2021; Pieh et al., 2021). We also found that disengagement syndrome contributes unique variance to functioning, independent of a related syndrome of social anxiety disorder, providing evidence of incremental validity.

General Discussion

In the present studies, we evaluated the psychometric properties of a novel instrument that assessed general disengagement after the acute phases of the pandemic. We used three studies and five samples from four countries. Our results suggest that the PDSS is a psychometrically valid and reliable instrument. Specifically, we found that the PDSS exhibited acceptable levels of internal consistency, test–retest reliability, discriminant validity, predictive validity, incremental validity, factorial validity, and measurement invariance across gender and countries. We found that factor analyses supported a general factor of disengagement and two subfactors, which we described as Social Avoidance and Alienation. Given that both findings of one general factor model (bifactor) and two-correlated factors model of the PDSS were supported by fit indices, we argue that researchers and practitioners can use either model depending on their interests and analytic purposes.

There may be multiple reasons for withdrawing from the social milieu. After the end of a pandemic phase, people may still fear infections, report concerns about interaction in person with others, and, consequently, plan to continue to isolate, and reduce activities (Fisher et al., 2021; Hood et al., 2021). Therefore, some people may still report loneliness during the reopening phase (Killgore et al., 2020). It should be noted that previous research did not show that COVID-19 lockdowns increase loneliness to a large extent (Prati & Mancini, 2021). It seems likely that only a portion of the population may experience higher social disengagement and loneliness following the acute phase of a pandemic. In addition, stay-at-home and social-distancing policies that have been enforced or recommended during the acute pandemic phases may interfere with the development and maintenance of social skills that occur across the lifespan (Beauchamp & Anderson, 2010). We suspect that temporarily limited access to stimulating social environments may also detract from our willingness to engage in social behavior, because of the absence of positive reinforcement. Future research is needed to understand the underlying causes of pandemic disengagement after the acute phase of a pandemic.

Although many self-report measures have been validated or adapted to investigate the psychological consequences of the COVID-19 pandemic (e.g., Sawicki et al., 2022), these assessments were designed to assess worries or fears during acute phases of the COVID-19 pandemic. By contrast, the present study seeks to assess the persistent psychological impairments that may have resulted from the pandemic. Our self-report measure can help to document the long-term psychological costs of the pandemic, assist in the diagnosis of COVID-19-related mental health problems, and identify the appropriate treatment for those afflicted. We believe that this instrument will continue to be useful as the consequences of the pandemic persist, and a subset of people will continue to demonstrate these symptoms. Moreover, because future national and local outbreaks of variants of severe acute respiratory syndrome coronavirus 2 or other pathogens cannot be excluded, the current instrument could be adapted for use in future outbreaks and pandemics.

It is important to emphasize that pandemic disengagement is not proposed to be a new diagnosis but rather a pathogenic–pathoplastic expression of mental illness in response to the unique stresses of the pandemic (Lilienfeld, 2017; McNally, 2012). In such a perspective, disengagement syndrome involves core dimensions of psychopathology that are refracted through the lens of the pandemic. This lens involves basic threats activated by the pandemic (e.g., fear of infection, harming others, contributing to spread) that would activate narrowly specific symptoms, the origins of which lie simultaneously in broad dimensional vulnerability to psychopathology (Conway et al., 2019; Kotov et al., 2017) and situational triggers of the pandemic. In this way, disengagement syndrome is conceptually similar to culturally bound expressions of psychopathology such as kayak angst, a form of agoraphobia that afflicts Inuit seal hunters in Greenland (Amering & Katschnig, 1990; Lilienfeld, 2017), and taijin kyofusho, an expression of social anxiety in Japan in which sufferers worry about offending others through their comments, appearance, or body odor (Kirmayer, 1991).

Limitations and Strengths

Although this study fills an important gap in the extant literature on the psychological consequences of acute phases of a pandemic, several limitations should be acknowledged. First, the recruitment of samples from Western countries limits generalizability across cultures. Second, our study focused only on adults. It is particularly important to investigate child and adolescent adaptation because of the potential for the pandemic to impair social skills. Future research may wish to examine the psychometric properties of our instrument among adolescents and children. Fourth, the current research is based on online surveys; therefore, people without internet access did not have the opportunity to participate. A final limitation is the lack of independent informant ratings of disengagement syndrome. Nevertheless, we note that our findings were based on adequately powered samples in four countries and five samples. Moreover, we found evidence that the PDSS possesses a coherent factor structure, measurement invariance across gender and country, construct validity and reliability, longitudinal stability, and predictive, discriminant, and incremental validity.

Conclusion

As the world adapts to the changes wrought by the COVID-19 pandemic, an important concern is its long-term psychological costs. The present research finds evidence that one such cost may be a specific psychopathological syndrome resulting from the pandemic. A better understanding of this syndrome is a critical task, and the PDSS can assist researchers to investigate it and clinicians to treat it.

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