RESEARCH PAPER

# A NEW SHORT VERSION OF INTERNET GAMING DISORDER-20: AN EXPLORATORY STRUCTURAL EQUATION MODELING

Palmira Faraci, Rossella Bottaro, Giuliana Nasonte, Giusy Danila Valenti, and Giuseppe Craparo



## Abstract

*Objective*: The purpose of this paper was to contribute to the psychometric properties and dimensionality of the IGD-20.

*Method*: An online survey was completed by 392 Italian online gamers ($M_{age}$ = 29.2, SD = 11.3; 45.2% males). A battery of self-report questionnaires was administered to assess internet gaming disorder, internet addiction, loneliness, anxiety, depression, stress, social-interaction anxiety, self-esteem, and perceived social support. To test the factor structure of IGD-20, both traditional (i.e., EFA and CFA) and innovative (i.e., ESEM) techniques were applied. Convergent, concurrent, discriminant, and criterion-related validity were evaluated.

*Results*: Our study revealed the outperforming 3-factor ESEM model ($\chi^2$=39.951, $p$ = 0.0021; RMSEA = 0.056, 90% C.I. [0.032 - 0.079]; CFI = 0.986; TLI = 0.965; and SRMR = 0.017; ω = .76, .77, and .79, respectively) as a new short version (IGD-10SV) for the IGD-20. The validity of the IGD-10SV was supported by significant associations with theoretically related measures.

*Conclusions*: The current findings support the adoption of the analytic ESEM approach for complex multidimensional measures and the use of the IGD-10SV for the assessment of internet gaming disorder.

**Key words:** internet gaming disorder, behavioral addictions, exploratory structure equation modeling, factor analysis; psychometrics

Palmira Faraci[1], Rossella Bottaro[1], Giuliana Nasonte[1], Giusy Danila Valenti[2], and Giuseppe Craparo[1]
[1] Faculty of Human and Social Sciences - University of Enna "Kore"
[2] Department of Psychological, Educational Sciences and Human Movement – University of Palermo
Authors' ORCid:
Palmira Faraci: https://orcid.org/0000-0002-7377-8011
Rossella Bottaro: https://orcid.org/0000-0002-9232-164x
Giusy Danila Valenti: https://orcid.org/0000-0003-0529-8515
Giuseppe Craparo: https://orcid.org/ 0000-0001-5748-0592

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**Corresponding author**
Rossella Bottaro
Faculty of Human and Social Sciences, University of Enna "Kore", Cittadella Universitaria – 94100 Enna, Italy
Phone number: +39 0935 536536
E-mails: rossella.bottaro@unikorestudent.it; rossella.bottaro1@gmail.com

The internet has profoundly altered people's daily lives as well as their leisure time. It provides online spaces for players to play alone or with others. Gaming Disorder was included in the 11th edition of the International Classification of Diseases (ICD-11; World Health Organization, 2019) as «a pattern of gaming behavior ("digital-gaming" or "video-gaming") characterized by i) impaired control over gaming, ii) increasing priority given to gaming over other activities to the extent that gaming takes precedence over other interests and daily activities, and iii) continuation or escalation of gaming despite the occurrence of negative consequences» and a specification was included for its *predominantly online* form (ICD-11 code: 6C51.0). Instead, at this time, only pathological gaming is included among the "disorders of substance use and addiction" as behavioural addictions in the fifth edition of the *Diagnostic and Statistical Manual of Mental Disorders* (DSM-5; American Psychiatric Association, 2013). Indeed, Internet Gaming Disorder (IGD) is one of the clinical conditions introduced in the section, which included studies that require 'deeper further studies.' The IGD's DSM-5 provisional definition is «persistent and recurrent use of the Internet to participate in online games— often with other players— which leads to impairment or clinically significant distress» (p. 921) for a period of 12 months or more, endorsing at least five of the following criteria: preoccupation, withdrawal, tolerance, unsuccessful attempts to control, loss of interest, continued and excessive use despite awareness of psychosocial problems, deception, escape, and functional impairment. Beyond the evolving nature of the DSM-5 definition (Kuss et al., 2017a, 2017b; Schimmenti et al., 2014), the IGD was proposed as a specific sub-type of wider internet addiction (Griffiths, 2018), which also includes other pathological and unregulated consumption of social networks, chat rooms, online gambling, and so on (Wölfling et al., 2020). The lack of an official diagnostic classification for IGD, as well as the coexistence of various theoretical models for addictions, increases the need for knowledge on this topic.

The impairment on the online pathological players' lives was significant as was the impairment on people with drug addictions: risky decision-making in Internet gaming disorder, as well as an overabundance of choice, may have strong negative consequences (Dong & Potenza, 2016; Misuraca et al., 2016). Individuals with IGD always suffer from withdrawal symptoms (Yen et al., 2022) and continue to play despite adverse consequences on various domains of people's lives. However, there were some similarities and differences with drug addiction: unlike the concept of *tolerance*, which was essentially related to the need for an increasing dose to achieve the same level of response, tolerance in gaming disorder cannot be linked solely to the need for increased gaming time; in other words, the authors suggested that tolerance may not be the most representative term for these symptoms (Griffiths, 2018). Instead, a recent study (King et al., 2017) suggested that tolerance in gaming disorder was related to a need for the completion of increasingly intricate tasks and spending more time on difficult goals to achieve satisfaction and reduce the fear of missing out. Similar to drug addictions, the IGD is affected by *craving*, defined as an urgent and uncontrollable desire to take a substance or engage in a specific behaviour (Caretti et al., 2008). In line with the self-medication hypothesis (Khantzian, 1997), the craving's function is related to anaesthesia and psychological pain reduction: the target behavior would allow you not to feel the pain related to previous traumatic experiences. Lies and manipulation were consequences of minimizing the adverse life experiences for others or getting the means (e.g., money, internet access) to implement the additive behavior.

Furthermore, previous research has highlighted associations between pathological online gaming and mental health, personality disorders (Gervasi et al., 2017; Şalvarlı & Griffiths, 2021), body dissociation (Casale et al., 2021), emotional distress (Giardina et al., 2021), suicide risk (Pavarin et al., 2022), mood disturbances (Ostinelli et al., 2021), anxiety (Sundqvist & Wennberg, 2022), depression (González-Bueso et al., 2018; Wang et al., 2017), stress (Andreetta et al., 2020), and self-esteem (Paulus et al., 2018). It was also related to social responsiveness, resulting in a decrease in social relationships (Bum et al., 2018) and an increase in loneliness (Paulus et al., 2018; Zeliha, 2019): high levels of IGD, for example, were linked to social phobia, social interaction anxiety, and lower perceived social support (González-Bueso et al., 2018; Sioni et al., 2017; Uçur & Dönmez, 2021).

According to Ferraro et al. (2020), the prevalence of IGD in Italy is 43%, or nearly one out of every two online gamers. However, to avoid clinical stigmatization, it is critical, as with all behavioral addictions, to draw a clear line between normal and pathological internet gaming (Sampogna et al., 2018). Indeed, an inconsistency between the time actually spent in the game and that perceived was discovered (Jin et al., 2022). For this reason, a generic instrument for Internet Addiction (e.g., Internet Addiction Test; IAT; Young, 1996) is insufficiently useful. Besides, unlike other psychometric instruments for this topic (e.g., IGDS9-SF; Pontes & Griffiths, 2015), a multidimensional tool is required for a more articulated understanding of pathology, especially for clinical purposes.

## The Internet Gaming Disorder-20

The Internet Gaming Disorder-20 (IGD-20; Pontes et al., 2014) is one of the most commonly used psychometric instruments to assess IGD according to DSM-5 diagnostic criteria (American Psychiatric Association, 2013). It was developed by recruiting a sample of 1003 English-speaking gamers from 57 countries. Its 20 items use a five-point Likert-scale ranging from 1 (*strongly disagree*) to 5 (*strongly agree*). According to the DSM-5 proposed criteria, a confirmative factorial analysis (CFA) supported the IGD-20's six-factor structure (i.e., salience, mood modification, tolerance, withdrawal, conflict, and relapse). Furthermore, the original study showed good validity and reliability scores (Pontes et al., 2014). The main problem of the original version was the high estimate correlation between the factors: for example, the correlation between Salience and Tolerance was extremely high ($r$ = .94) as was the correlation between Tolerance and Withdrawal ($r$ = .77), Salience and Conflict ($r$ = .74), and Tolerance and Conflict ($r$ = .74). Generally, all the correlations between factors were too high resulting in multicollinearity issues and poorly defined factors.

Until today, the IGD-20 has been validated in six languages: Spanish (Fuster et al., 2016), Arabic (Hawi & Samaha, 2017), Korean (Kim, 2019), Chinese (Shu et al., 2019), Polish (Grajewski & Dragan, 2021), and Turkish (Çakıroğlu & Soylu, 2019). Most versions showed some good reliability (e.g., Arabic, Korean, and Turkish total Cronbach's α = .92, .85, .86, respectively). However, despite the fact that it is frequently presented as a multidimensional instrument (with the exception of the Arabic validation), the Cronbach's α values for each subscale have been omitted (i.e., French and Korean validations; see **table 1**). Moreover, some psychometric issues were found in all versions: although the 6-factor structure is very common in the different published

**Table 1.** *Reliability (Cronbach's α) of IGD-20 test for the available validations*

| Validation | Tot | Salience | Tolerance | Mood Modification | Withdrawal | Conflict | Relapse |
|---|---|---|---|---|---|---|---|
| English (original study) | .88 | .64 | .63 | .78 | .80 | .74 | .63 |
| Spanish (Fuster at al., 2016) | U | .68 | .61 | .79 | .85 | .76 | .66 |
| Arabic (Hawi & Samaha, 2017) | .91 | N.A. | N.A. | N.A. | N.A. | N.A. | N.A. |
| Chinese (Shu et al., 2019) | .90 | | .84 | .59 | .82 | .52 | .82 |
| French (Plessis et al., 2021) | U | U | U | U | U | U | U |
| Korean | .85 | U | U | U | U | U | U |

*Note.* U = Undetectable; N.A. = Not applicable.

validations, the estimated correlations among the factors often remain high ($r \geq .90$) as well as in the original study. In the Spanish validation, for example, the estimated correlations between Salience and Tolerance ($r = .95$), Withdrawal and Conflict ($r = .92$), and Tolerance and Withdrawal ($r = .89$) were all high. Additionally, the Korean validation found the same correlational results between Salience and Tolerance ($r = .99$), Salience and Conflict ($r = .90$), Tolerance and Withdrawal ($r = .88$), Withdrawal and Conflict ($r = .88$), and Conflict and Relapse ($r = .88$) (see the factor correlation matrix in the original studies). To address this issue, various solutions were proposed, and different studies did not aim for the same factorial structure. The French validation inserted the correlations between errors as modification indices to reduce the correlation between latent factors (i.e., items 2 and 8, 8 and 14 for Mood Modification, 5 and 20, 19 and 20 for Conflict, 6 and 12 for Relapse) which remained very high (e.g., correlation between Tolerance and Withdrawal =.91; Salience and Tolerance = .89). Given the high correlation between Salience and Tolerance reported in previous studies, the Chinese and Polish versions proposed a 5-factor structure that combined Salience and Tolerance into a single factor without solving the problem (Salience/Tolerance and Withdrawal = .96, Salience/Tolerance and Conflict = .94, Salience/Tolerance and Relapse = .92), while the Arabic version fit well with a one-factor model. Further concerns emerge nonetheless at item-level analysis: Some items often displayed problematic factor loadings: e.g., item 2 (French $\lambda = .12$; Korean $\lambda = .29$, Chinese = .24); item 19 (French $\lambda = .23$, Chinese $\lambda = .01$).

Ultimately, as with other commonly used self-report measures (Faraci et al., 2013a; Faraci & Tirrito, 2013; Triscari et al., 2011) testing the tool's factorial model in different languages may be beneficial in controlling cultural distortions and improving diagnostic capacities. However, addressing the highlighted psychometric challenges would be valuable in order to make the instrument more valid and reliable while preserving its multidimensionality for more expendable clinical applications. Indeed, the inflated correlations lead to a poor distinction between factors and biased relationships with other constructs (Marsh et al., 2011, 2014; Perera, 2015). In this regard, it also appears critical to provide new evidence about various validity types (e.g., discriminant and convergent validity) to improve our understanding of the clinical path of IGD.

## Methodological focus: Exploratory Structural Equation Modeling (ESEM)

Many psychological instruments have an apparently adequate factor structure based on the traditional Confirmatory Factor Analysis (CFA) approach. However, while returning adequate goodness-of-fit indices, a more careful analysis from an eclectic perspective (Marsh et al., 2010), which considers the integration of a variety of different indices, a detailed evaluation of the estimated parameters, and a comparison between viable alternative models, can arouse some psychometric issues. The advantages of the CFA approach were clearly distinguishable in the IGD-20 study: the DSM-5 diagnostic criteria suggested a very informative theory-driven starting point to test the dimensionality of the tool. The misspecification of zero factor loadings, though, may lead to an estimate of 'pure factors' that is parsimonious but often unrealistic (Asparouhov & Muthén, 2009; Marsh et al., 2009, 2010, 2011).

Exploratory Structural Equation Modeling (ESEM) incorporates new parts of variance in the model, permitting the estimation of the cross-loadings. This allows a better distinction of latent factors that are less related, improving the ability to differentiate between multiple factors (Marsh et al., 2011; Perera, 2015). As a consequence of the inflated correlations in the CFA model, the evidence about discriminant validity, relationships with other constructs, and the tenability of the higher-order representation could result in erroneous inferences (Marsh et al., 2014). Here, in light of the limitations highlighted in IGD-20's previous research, we described an application of the ESEM, an evolving psychometric approach. We combined the advantages of Exploratory Factor Analysis (EFA), a data-driven approach, and CFA, a theory-driven approach, that could offer a new insight on the dimensionality of the tool.

## The present study

Substantively, this paper aimed to address issues relevant to the psychometric properties of IGD-20. Previous studies (Pontes & Griffiths, 2015) have shown that a one-dimensional structure frequently fails to adequately describe the complexities of psychological constructs and is therefore ineffective in clinical and research settings. As a result, the current study aimed to provide a multidimensional psychometric tool that has already been used in different cultures to evaluate the IGD and improve the instrument's theory. A comparison between CFA and ESEM models contributed to the literature supporting the application of ESEM in psychological research.

In particular, the present study was designed to: (a) define a stable factor structure that overcomes the psychometric limits discovered in previous research; and (b) examine the instrument's convergent, concurrent, and discriminant validity based on the following hypotheses:

Hypothesis 1 (H1): high levels of IGD as measured by IGD-20 should be associated with high levels of IGD as measured by a different psychometric instrument designed for the same construct, such as the Internet Gaming Disorder Scale;

Hypothesis 2 (H2): high levels of IGD should be associated with high levels of internet addiction, anxiety, depression, stress, loneliness, and social interaction anxiety;

Hypothesis 3 (H3): high levels of IGD should be associated with low levels of self-esteem and perceived social support from family, friends, and significant others.

## Method

### *Participants*

Participants (N = 392) were Italian adults (i.e., age > 18; $M_{age} = 29.2$, $SD = 11.3$, range = 13–75) who play online games using various informatics software and hardware. We excluded children and adolescents (i.e., under the age of 18) as well as people who do not play online. The overall sample was nearly gender balanced (45.2% males and 54.8% females) and began playing online at an average age of 14.5 ($SD = 10.9$, range = 0-69).

The total sample was randomly divided into two subsamples: Group I (n = 208, $M_{age} = 30$, $SD = 11.6$)

was used for the Exploratory Factor Analysis (EFA). They were 44.7% males and 55.3% females, and they began playing online at $M_{age}$ = 15.9 (*SD* = 14.1). Group II (n = 184, $M_{age}$ = 28.3, *SD* = 10.9) was used for the Confirmatory Factor Analysis (CFA) and Exploratory Structural Equation Modeling (ESEM). They were 45.7% males and 54.3% females, and they began playing online at $M_{age}$ = 15.7 (*SD* = 16). See **table 2** for more details.

**Table 2.** *Sample socio-demographic characteristics (N = 392)*

| | Total Sample (n=392) | | Group I (n=208) | | Group II (n=184) | |
|---|---|---|---|---|---|---|
| | N | % | N | % | N | % |
| Have a relationship | | | | | | |
| Yes | 222 | 56.6 | 113 | 54.3 | 109 | 59.2 |
| Not | 148 | 37.8 | 82 | 39.4 | 66 | 35.9 |
| I prefer not to specify | 22 | 5.6 | 13 | 6.3 | 9 | 4.9 |
| Devices used to play online | | | | | | |
| Smartphone/Tablet | 285 | 72.7 | 157 | 75.5 | 128 | 69.6 |
| Console (e.g., playstation, Xbox, etc.) | 162 | 41.3 | 84 | 40.4 | 78 | 42.4 |
| Computer | 192 | 49 | 112 | 53.8 | 80 | 43.5 |

*Note*. Group I = Group I for the Exploratory Factor Analysis; Group II = Group II for the Confirmatory Factor Analysis and the Exploratory Structural Equation Model.

## *Measures*

Two Italian native speakers independently translated the English version of IGD-20 into Italian. The two translations were then compared, and no substantial differences were found. One bilingual speaker who was familiar with the psychological topic back-translated the first final version into English. A minor revision was required after comparing the back-translation with the original version.

In addition to the IGD-20, participants completed other self-report measures to provide evidence based on relationships to other variables. Ad hoc items were developed to detect the socio-demographic sample's characteristics and online gambling habits.

*Internet Addiction*. The *Internet Addiction Test* (IAT; Young, 1996) is a pioneering psychological instrument for assessing Internet Addiction that was developed in an 8-item first version in accordance with the DSM-IV (Diagnostic and Statistical Manual of Mental Disorders-4th edition) criteria for pathological gaming. Next, the author (Young, 1998) proposed a 20-item extension (e.g., "Do people around you complain about the amount of time you spend online?", "Are your performance at work or your productivity affected negatively by the Internet?") that assesses the severity of the disorder on a 5-point Likert scale ranging from 0 (*not at all*) to 5 (*always*). In the present study, we used the Italian version (Faraci et al., 2013a), which demonstrated good psychometric properties, including for the two-factor solution (Cronbach's α =.88 and Cronbach α =.79 for the Italian version, respectively, and Cronbach's α =.93 for the present sample).

*Internet Gaming Disorder*. The *Internet Gaming Disorder Scale-Short-Form* (IGDS9-SF; Pontes & Griffiths, 2015) is a brief instrument (i.e., a unidimensional scale with nine items) for assessing online internet gaming disorder in accordance with the nine DSM-5 criteria (e.g., "Have you lost interests in previous hobbies and other entertainment activities as a result of your engagement with the game?" or "Have you continued your gaming activity despite knowing it was causing problems between you and other people?"). It shows good validity and reliability in both the original study (Pontes & Griffiths, 2015; Cronbach's α = .87) and the present sample (Cronbach's α = .91). It has also been validated in 15 different languages. In this paper, we used the Italian version (Monacis et al., 2016).

*Self-esteem*. To assess self-esteem levels, we administered the *Rosemberg Self-Esteem Scale* (RSES; Rosemberg, 1989). It is a 10-item self-report scale (e.g., "I think I have a number of qualities", "I guess I don't have much to be proud") with a 4-point Likert scale ranging from 1 (*strongly agree*) to 4 (*strongly disagree*). We used the Italian version (Prezza et al., 1997), which showed good internal consistency (Cronbach's α =.84). Cronbach's α for the present sample was .91.

*Stress, anxiety, and depression*. To assess stress, anxiety, and depression with a unique psychometric instrument, we used the Depression Anxiety Stress Scales-21 (DASS-21; Lovibond & Lovibond, 1995) in its Italian version (Bottesi et al., 2015). It is a 21-item self-report scale (e.g., "I felt a lot of tension and I had difficulty recovering a state of calm", "I just couldn't feel any positive emotions", "I felt stressed out") with a 4-point Likert scale ranging from 0 (*never happened to me*) to 3 (*it happened to me almost always*) with good internal consistency and temporal stability in both the original version (anxiety Cronbach's α = .74; depression Cronbach's α = .82; stress Cronbach's α = .85; total Cronbach's α = .90; Lovibond & Lovibond, 1995) and the present sample (anxiety Cronbach's α = .81; depression Cronbach's α = .91; stress Cronbach's α = .90; total Cronbach's α = .95).

*Loneliness*. The *UCLA Loneliness-Scale- Version 3* (UCLA-LS; Russell, 1996) was used to measure one's subjective feelings of loneliness as well as feelings of social isolation. For the purpose of the current study, we administered its Italian version (Boffo et al., 2012). It is a 20-item scale with a 4-point Likert scale ranging from 1 (*I never fell this way*) to 4 (*I often feel this way*) (e.g., "I am unhappy doing so many things alone", "I feel isolated from others"). This measure has a high level of internal consistency in both its original version (Russell, 1996; Cronbach's α = .96) and the present sample (Cronbach's α = .92).

*Social interaction anxiety*. The *Social Interaction Anxiety Scale* (SIAS; Mattick & Clarke, 1998) is a 20-item scale used to assess anxiety when engaging in social interactions (e.g., "It makes me uncomfortable to meet an acquaintance on the street", "It's hard for me to chat with other people"). Respondents are asked to rate their social interaction anxiety on a 4-point Likert-type scale; the total score reflects the severity of the anxiety (Cronbach's α = .86 for the original version; Mattick & Clarke, 1998; Cronbach's α = .92 for the present sample). In this study, the SIAS Italian version was used (Sica et al., 2007).

*Perceived Social Support.* The *Multidimensional Scale of Perceived Social Support* (MSPSS; Zimet et al., 1988) is composed of three subscales that can differentiate the sources of perceived social support: family, friends, and significant others. Its 12 items (e.g., "There is a special person who is around when I am in need", "My family really tries to help me", "I can count on my friends when things go wrong") are evaluated on a 7-point Likert scale ranging from 1 (*very much disagree*) to 7 (*very much agree*). The MSPSS's Italian validation (Di Fabio & Palazzeschi, 2015) showed good reliability for each subscale (Cronbach's α for Significant other = .91, Cronbach's α for Family = .87, and Cronbach's α for Friends = .85), as well as the present sample (Cronbach's α for Significant other = .91, Cronbach's α for Family = .91, and Cronbach's α for Friends = .94).

## Procedure and ethics

Data were collected through an online survey via SNS (e.g., Facebook, Instagram, and WhatsApp) using a Google Form. The research project proposal was carried out in accordance with the Declaration of Helsinki and was approved by the Internal Review Board of the psychological research of the **** University.

The measures were administered in compliance with the privacy guarantee regulations outlined in Legislative Decree n. 196/2003 and the GDPR (EU Regulation n.2016/679). All participants agreed to provide informed consent.

## Data analyses

First, we compared Confirmatory Factor Analysis (CFA), Full and Bifactor Exploratory Structural Equation Modeling (ESEM and B-ESEM, respectively) to test the most recurring factor structure (i.e., the 6-factor structure) presented in previous research according to the common accepted criteria (i.e., RMSEA and SRMR [0.06-0.08] marginally acceptable or [0.01-0.05] excellent; CFI and TLI [0.90-0.95] marginally acceptable or [0.96-0.99] excellent) (Hu & Bentler, 1999). After finding several psychometric limits, according to Alamer and Marsh (2022), we performed an Exploratory Factor Analysis (EFA) on the first randomly extracted subsample (Group I) to determine the most performant factor structure. CFA and ESEM were applied to the second randomly selected subsample (Group II) in order to find the best factorial solution based on the data-driven results and the theoretical model proposed in previous literature (Caretti et al., 2008). All models' estimations were performed using Mplus 7 software (Muthén & Muthén, 1998-2012). Finally, we evaluated the convergent, concurrent, discriminant, and criterion-related validity using IBM SPSS version 25.

## Results

### Preliminary data processing

The minimum required sample size for CFA and ESEM estimation was a priori calculated (Christopher Westland, 2010; Cohen, 2013; Soperg, 2022). We estimated it for a medium effect size (i.e., 0.3), a desired statistical power level of 80%, a confidence interval of 95%, and the most expensive model among those provided (i.e., 20 observed variables and 6 latent factors). The minimum sample size recommended was 161. Therefore, our sample largely meets the requirements.

In the first phase, the entire dataset was screened for potential issues. Thanks to a mandatory response format, no missing data were found. The multivariate normality of the data was checked by the computation of Mahalanobis' distance, which revealed that the data were approximately normal (i.e., Mardia's multivariate Kurtosis coefficient = 183.97; critical value = 143, chi-square critical value = 31.264 when $p < .001$).

### Comparison between CFA and ESEM 6-factors models

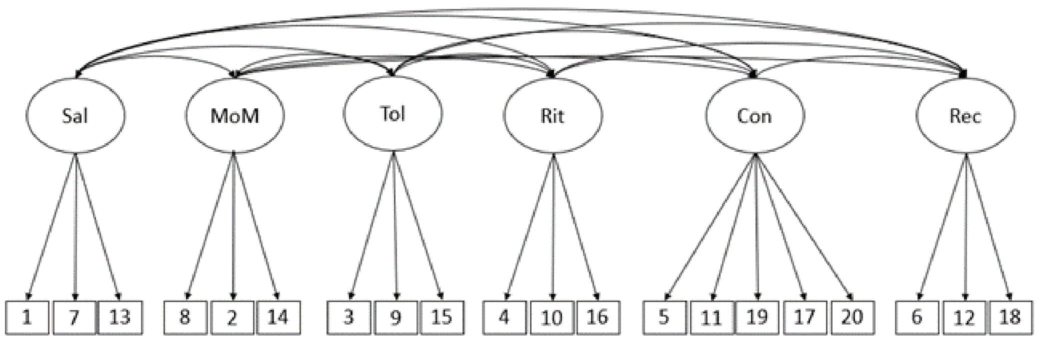
Based on the preliminary results, we used the Maximum Likelihood (ML) estimator which assumed the normality of the data distribution. A CFA was performed to test the most common structure presented in previous studies (i.e., 6-factors model; **figure 1**). Overall, the model fit indices appeared to be well-fitting, except for the CFI and TLI that were really on the verge of acceptability (i.e., chi-square test = 460.833, df=155, p-value = 0.00; RMSEA = 0.071, 90% C.I. [0.064 - 0.078]; CFI = 0.910; TLI = 0.890; and SRMR = 0.048; **table 3**). However, other psychometric criticisms have emerged, similar to those found in previous validations. First and foremost, the estimated factor correlation showed very high values and highlighted an inadmissible model (e.g., three

**Table 3.** *Fit indices of examined models*

| Model | $\chi^2$ | *Df* | *p* | CFI | TLI | RMSEA [90% C.I.] | SRMR | AIC | BIC | aBIC | Meets the criteria |
|---|---|---|---|---|---|---|---|---|---|---|---|
| **6-factors** | | | | | | | | | | | |
| CFA | 460.833 | 155 | 0.00 | 0.910 | 0.890 | 0.071 [0.064 – 0.078] | 0.048 | 20018.918 | 20316.762 | 20078.790 | No |
| ESEM | 479.258 | 151 | 0.00 | 0.904 | 0.879 | 0.074 [0.067 – 0.082] | 0.043 | 20045.342 | 20359.071 | 20108.407 | No |
| **3-factors short version** | | | | | | | | | | | |
| CFA 3-factors | **739.452** | **45** | **0.002** | **0.959** | **0.942** | **0.070 [0.042- 0.096]** | **0.041** | **4708.641** | **4814.734** | **4710.215** | **Yes** |
| ESEM 3-factors | **24.086** | **18** | **0.152** | **0.991** | **0.978** | **0.043 [0.000-0.083]** | **0.022** | **4699.990** | **4851.092** | **4702.231** | **Yes** |

*Note.* In bold models that meet criteria; *df*= degree of freedom; CFI=comparative fit index; TLI=Tucker-Lewis index; RMSEA= root-mean-square error of approximation; 90% CI=90%confidance interval for RMSEA; SRMR= Standardized-root-mean square residual.

**Figure 1.** *Confirmatory factor analysis of the 6-Factor Model*




*Note.* Sal = Salience; Mom = Mood Modification; Tol = Tolerance; Rit = Withdrawal; Con = Conflict; Rec = Relapse

correlations >.99; **table 4**). Furthermore, Cronbach's values for two subscales were well below acceptable levels (i.e., Mood Modification and Conflict α = .05 and .51, respectively). The Cronbach's alpha values for the other subscales were modest (i.e., Tolerance, Salience, and Relapse α =.62, .67, and .68, respectively). Only the withdrawal subscale had a good value (i.e., Cronbach's α = .79). We also compared these CFA model with the correspondent ESEM (**table 3**) with target rotation and with Bifactor ESEM because this approach is recognized as decreasing the correlation between factors. Not even the ESEM model showed fully acceptable fit indices (i.e., chi-square test = 479.258, df=151, p-value = .00; RMSEA = .074, 90% C.I. [.067 - .082]; CFI = .904; TLI = .879; and SRMR = .044) and the bifactor model with target orthogonal rotation did not converge. As a result of these findings, the most common factorial structure was disconfirmed.

**Table 4.** *Factor loadings, estimated correlation matrix and Cronbach's α for the latent variables in CFA 6-Factors Model*

| | 1. Salience | 2.Mood Modification | 3. Tolerance | 4.Withdrawal | 5. Conflict | 6. Relapse |
|---|---|---|---|---|---|---|
| Item 1 | .54 | | | | | |
| Item 7 | .65 | | | | | |
| Item 13 | .70 | | | | | |
| Item 8 | | .81 | | | | |
| Item 2 | | .33 | | | | |
| Item 14 | | .79 | | | | |
| Item 3 | | | .47 | | | |
| Item 9 | | | .80 | | | |
| Item 15 | | | .64 | | | |
| Item 4 | | | | .65 | | |
| Item 10 | | | | .84 | | |
| Item 16 | | | | .78 | | |
| Item 5 | | | | | .73 | |
| Item 11 | | | | | .66 | |
| Item 19 | | | | | -.14 | |
| Item 17 | | | | | .56 | |
| Item 20 | | | | | .68 | |
| Item 6 | | | | | | .69 |
| Item 12 | | | | | | .57 |
| Item 18 | | | | | | .76 |
| 1. Salience | - | | | | | |
| 2. Mood Modification | .76 | - | | | | |
| 3. Tolerance | .99 | .76 | - | | | |
| 4. Withdrawal | .87 | .68 | .87 | - | | |
| 5. Conflict | 1.05 | .68 | .99 | .87 | - | |
| 6. Relapse | .99 | .67 | .94 | .89 | 1.01 | - |
| Cronbach's α | .67 | .05 | .62 | .79 | .51 | .68 |

## *Exploratory factor analysis*

Following, we performed an EFA on Group I. Bartlett's test of sphericity ($\chi^2$ = 914.609; *df* = 45) was significant ($p$ < .001), and the KMO measure of sampling adequacy was .881, indicating that the questionnaire items were suitable for factor analysis.

Parallel analysis determined that four factors had to be extracted. The factor correlation matrix, indicating a prominent inter-correlation among factor scales, supported the use of the oblique rotation procedures (promax criterion). Based on the resultant pattern matrix, item 7 "I usually think about my next gaming session when I am not playing" and item 3 "I have significantly increased the amount of time I play games over last year" that loaded simultaneously on two factors, without a difference of at least .30 between loading on the primary factor and loading on other factors, were not retained (item 7 loaded on F1 at .339, and on F3 at .333; item 8 loaded on F2 at .444, and on F3 at .361). At this point, item 19 "I know my main daily activity (i.e., occupation, education, homemaker, etc.) has not been negatively affected by my gaming" failed to load at least .30 on the extracted three factors. Therefore, item 19 was removed. Further, item 5 "I have lost interest in other hobbies because of my gaming", that loaded simultaneously on two factors, was not retained (item 5 loaded on F1 at .414, and on F2 at .451). Additionally, item 1 "I often lose sleep because of long gaming sessions", which failed to load at least .30 on the extracted three factors, was removed and item 18 "I often try to play games less but find I cannot", that loaded simultaneously on two factors, was not retained (item 18 loaded on F1 at .341, and on F2 at .499). Item 2 "I never play games in order to feel better" was removed due to communality equal to .179. Moreover, item 17 "I think my gaming has jeopardized the relationship with my partner", which loaded simultaneously on two factors, was not retained (item 5 loaded on F1 at .491, and on F2 at .314). At this point, the resulting number of factors was evidently over-defined, with a factor comprised by only one indicator (i.e., item 8 "I play games to help me cope with any bad feelings I might have" that was, thus, removed). Finally, item 9 "I need to spend increasing amounts of time engaged in playing games" loaded simultaneously on two factors (i.e., item 9 loaded on F2 at .375, and on F3 at .307) and, thus, it was removed.

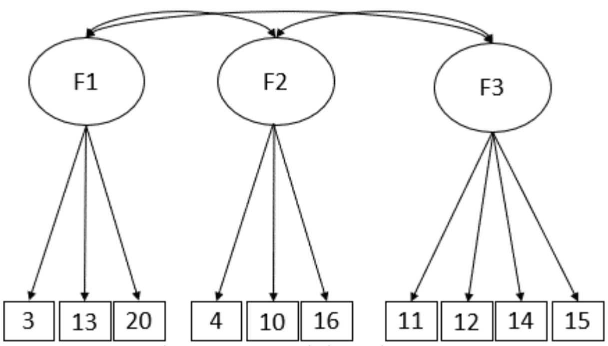
Ultimately, in accordance with previous research on behavioral addiction (Caretti et al., 2008; King et al., 2017; Yen et al., 2022) and the parallelism with the labels of the original factors, we named the three factors as follows: F1) Prevalence (i.e., items 6, 13, and 20), because this latent factor described the impact of pathological games on users' lives in terms of time and context, according to the Cambridge Dictionary definition of «the fact that something is very common or happens often»; F2) Withdrawal (i.e., items 4, 10, and 16), because it included the same items of the original factor with the same name; F3) Craving (i.e., items 11, 12, 14, and 15), because these items reflected the essential characteristics of this longing desire and its impact on emotional and social experiences in daily life (e.g., lies and worries) in line with the Caretti et al. (2008) theoretical definition of craving.

Retained items produced consistent and satisfactory loadings on each factor, meeting minimum requirements for inclusion. **Table 4** displays the scale's items and factor loadings. Intercorrelations between subscale scores were $r$ = .564 ($p$ < .01) for withdrawal and prevalence, $r$ = .637 ($p$ < .01) for withdrawal and craving, and $r$ = .723 ($p$ < .01) for prevalence and craving. As expected, the dimensions correlated significantly with each other, indicating that the questionnaire subscales measured the IGD construct. Cronbach's subscales' alpha coefficients were satisfying (prevalence, $\alpha$ = .82; withdrawal, $\alpha$ = .82; craving, $\alpha$ = .75). For a more practical presentation, this new short version is referred to as *Internet Gaming Disorder-10 Short Version* (IGD-10SV; Appendix 1).

## *Comparison between CFA and ESEM 3-Factors Model (short version)*

The feasibility of the emerging three-factor solution was examined through a CFA on Group II and compared with the corresponding ESEM solution. As presented in **table 3**, the fit indices of the three-factor solution met the criteria for adequacy for both CFA (i.e., chi-square test = 739.452, p-value = .0016; RMSEA = .070, 90% C.I. [.042 - .096]; CFI = .959; TLI = .942; and SRMR = .041) and ESEM models (i.e., chi-square test = 24.086, p-value = .1522; RMSEA = .043, 90% C.I. [.000 - .083]; CFI = 0.991; TLI = .978; and SRMR = .022; **table 5-6-7**; **figure 3**) with a slight and expected improvement in the latter. Despite the ESEM model had a slightly high BIC than the CFA model, recent research (Cao & Liang, 2022) has revealed that this information criterion was biased in the ESEM technique, favouring the more parsimonious model, as well as in our results. However, the correlation between the factors was once again quite high ($r$ > .70) in the CFA model, while the ESEM 3-factor model showed better defined factors and significantly reducing the factor correlations ($r$ < .46) to the corresponding CFA model (**table 6**). At item level, all items loaded significantly ($p$ < 0.05) on their hypothesized latent factors in CFA. The standardized parameter estimates are shown in **figure 2**. On the other hand, the ESEM model produced well-defined factors (with all items $\lambda$ > .30, with the exception of item 12, which was statistically significant for $p$ <.01). However, according to van Zyl and ten Klooster (2022), these results should be considered in the context of the theory, and thresholds should not be taken rigidly. Indeed, the item 12 "I do not think I could stop gaming" as well as the item 6 "I would like to cut down my gaming time but it's difficult to do" showed statistically significant cross-loadings on all non-target factors. These items are theoretically related to the time spent playing online and the difficulties related to stopping playing. According to previous research (Griffiths, 2018; King et al., 2017), game time dynamics are crucial elements in IGD diagnosis, with withdrawal, craving, and prevalence implications as reflected in the present factorial structure. Moreover, item 11 "I have lied to

**Figure 2.** *Confirmatory factor analysis of the 3-Factor Model of IGD-10SV*




*Note*. F1 = Prevalence; F2 = Withdrawal; F3 = Craving.

**Figure 3.** *Exploratory structure equation model of the 3-Factor Model of IGD-10SV*

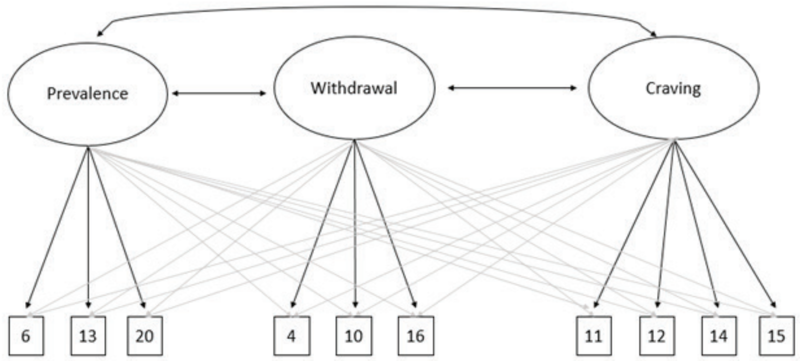



**Table 5.** *Factor loadings of the scale items (pattern matrix) for exploratory factor analysis (n = 208)*

| Item | λ F1 | λ F2 | λ F3 |
|---|---|---|---|
| 20.I believe my gaming is negatively impacting on important areas of my life. | .865 | -.166 | .144 |
| 6.I would like to cut down my gaming time but it's difficult to do. | .783 | .238 | -.251 |
| 13.I think gaming has become the most time-consuming activity in my life. | .647 | -.002 | .180 |
| 16.I tend to get anxious if I can't play games for any reason. | .039 | .738 | .073 |
| 4.When I am not gaming I feel more irritable. | .012 | .694 | -.004 |
| 10.I feel sad if I am not able to play games. | -.024 | .633 | .276 |
| 12.I do not think I could stop gaming. | -.115 | .091 | .678 |
| 11.I have lied to my family members because the amount of gaming I do. | .045 | .045 | .676 |
| 15.I often think that a whole day is not enough to do everything I need to do in-game. | .077 | .080 | .532 |
| 14.I play games to forget about whatever's bothering me. | .283 | .026 | .402 |

**Table 6.** *Correlations among latent factors of confirmative factor analysis and exploratory structural equation model of IGD-10SV in group II (N = 184)*

| | 1. | 2. | 3. |
|---|---|---|---|
| 1. P | - | .735*** | .791*** |
| 2. W | .459*** | - | .771*** |
| 3. C | .373*** | .368*** | - |

*Note.* CFA above diagonal, ESEM below
P = Prevalence; W = Withdrawal; C = Craving.
**p < .05; **p < .01; ***p < .001*

my family members because the amount of gaming I do" called up "greed" for the game at whatever cost typical of craving but also the impact on the loved ones of players. Ultimately, the item 16 "I tend to get anxious if I can't play games for any reason" showed significant target loading ($\lambda = 0.607$) on Withdrawal for its social implication but also on Craving for the need to implement online gambling behavior to escape painful emotional states, in line with the self-medication hypothesis (Khantzian, 1997).

Finally, the McDonald's Omega showed an optimal value for each subscale ($\omega$ = .76, .77, and .79, respectively) — $\omega$ is to be preferred over the Cronbach's alpha because it does not assume equal loadings— according to suggested reliability criteria (i.e., $\omega$ > .70; McDonald, 1999). As a result, and according to the *eclectic approach* (Marsh et al., 2010, p. 488), we supported the ESEM 3-factors model for the present short version as a new short version for the IGD-20 (that we called the IGD-10SV). Based on our findings, we performed the subsequent analysis using the suggested model.

## *Convergent, concurrent, discriminant, and criterion-related validity of the IGD-10SV*

To verify the convergent, concurrent, discriminant, and criterion-related validity, the correlation between our version of IGD-10SV and existing psychometric instruments was investigated. We computed the scores of the new three factors (i.e., prevalence, withdrawal, and craving) and explored their associations with other variables. Convergent validity was detected in relation

**Table 7.** *Item level descriptive statistics, standardized loadings of confirmative factor analysis and exploratory structural equation model of IGD-10SV in group II (n = 184)*

| Factors | Item | Mean | SD | Skew-ness | Kurtosis | CFA | | | | | | | | ESEM | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | | P | | W | | C | | | | P | | W | | C | | | |
| | | | | | | λ | *S.E.* | λ | *S.E.* | λ | *S.E.* | δ | $R^2$ | λ | *S.E.* | λ | *S.E.* | λ | *S.E.* | δ | $R^2$ |
| **P** | 6 | 1.95 | 1.05 | 0.89 | -0.19 | **0.636** | 0.054 | | | | | 0.595 | 0.405 | **0.588** | 0.093 | 0.322 | 0.098 | -0.214 | 0.088 | 0.476 | 0.524 |
| | 13 | 1.79 | 1.06 | 1.41 | 1.25 | **0.768** | 0.768 | | | | | 0.411 | 0.589 | **0.649** | 0.076 | 0.069 | 0.080 | 0.166 | 0.087 | 0.416 | 0.584 |
| | 20 | 2.10 | 1.18 | 0.81 | -0.41 | **0.732** | 0.732 | | | | | 0.464 | 0.536 | **0.628** | 0.082 | 0.143 | 0.071 | 0.031 | 0.031 | 0.484 | 0.516 |
| **W** | 10 | 1.79 | 1.00 | 1.19 | 0.62 | | | **0.823** | 0.823 | | | 0.322 | 0.678 | 0.109 | 0.084 | **0.724** | 0.084 | 0.101 | 0.090 | 0.318 | 0.682 |
| | 16 | 1.54 | 0.75 | 1.31 | 1.18 | | | **0.771** | 0.771 | | | 0.406 | 0.594 | 0.070 | 0.079 | **0.607** | 0.083 | 0.244 | 0.084 | 0.407 | 0.593 |
| | 4 | 1.68 | 0.92 | 1.40 | 1.48 | | | **0.576** | 0.576 | | | 0.668 | 0.332 | 0.099 | 0.079 | **0.322** | 0.092 | -0.016 | 0.058 | 0.640 | 0.360 |
| **C** | 11 | 1.62 | 0.99 | 1.64 | 1.92 | | | | | **0.723** | 0.044 | 0.477 | 0.523 | 0.444 | 0.084 | 0.017 | 0.094 | **0.429** | 0.094 | 0.463 | 0.537 |
| | 12 | 2.08 | 1.91 | 0.72 | -0.77 | | | | | **0.616** | 0.054 | 0.621 | 0.379 | 0.249 | 0.087 | 0.244 | 0.081 | **0.271** | 0.081 | 0.650 | 0.350 |
| | 14 | 2.25 | 1.26 | 0.50 | -1.09 | | | | | **0.696** | 0.696 | 0.516 | 0.484 | 0.385 | 0.084 | 0.084 | 0.081 | **0.371** | 0.081 | 0.457 | 0.453 |
| | 15 | 1.74 | 1.00 | 1.41 | 1.42 | | | | | **0.737** | 0.044 | 0.456 | 0.544 | 0.025 | 0.088 | 0.192 | 0.097 | **0.781** | 0.097 | 0.223 | 0.777 |

*Note.* P = Prevalence; W = Withdrawal; C = Craving. Bold items= Significant target loadings (p < 0.05); Underlined items indicate cross-loading items; S.E. Standard Error; λ = Standardized Factor loadings; δ=Item Uniqueness.

to the IGDS: all subscales ($r$ = .70, .62, .71, respectively, $p$ < .001) showed significant high positive associations with another measure of Internet Gaming Disorder. In terms of convergent validity, the all IGD-10SV subscales' score showed significant positive correlations with measures of internet addiction, anxiety, depression, stress, loneliness, and social interaction anxiety ($p$ <.001 for all measures, except for anxiety, which showed $p$ < .01; **table 8**). Additionally, with regard to discriminant validity, self-esteem ($r$ = -.36, -.28, -.36, $p$ < .001) and MSPSS subscales (i.e., perceived social support from family: $r$ = -.35, -.28, -.41; friends: $r$ = -.28, -.30, -.21; and significant others: $r$ = -.28, -.19, -.27) showed significant negative associations with all IGD-10SV subscales ($p$ < .001). Finally, to investigate criterion-related validity, we conducted correlations between the IGD-10SV subscales and play hours (i.e., daily, weekly, and monthly). All associations were statistically significant and positive ($p$ < .001; **table 8**).

# Discussion

The present paper investigated the psychometric properties of the IGD-20 and offered a contribution to the literature supporting the application of ESEM. In this regard, our results offered new knowledge to enhance instrument theory and a broader conceptualization of the IGD, which is still designated as needing further study.

In contrast to previous research (Çakıroğlu & Soylu, 2019; Fuster et al., 2016; Hawi & Samaha, 2017; Kim, 2019; Grajewski & Dragan, 2021; Pontes et al., 2014; Shu et al., 2019), which highlighted poorly factor loadings or high correlations between latent factors that resolved in different factor structures or in the use of modification indices, we purposed a new methodological perspective comparing both traditional (i.e., EFA and CFA) and innovative (i.e., ESEM) methods. First, we used the CFA method to test the most common IGD-20 structure (i.e., 6-factors), which resulted in the same problems as previous validations. As a new methodological point of view, we compared it with ESEM and Bifactor ESEM models using target rotation in a confirmatory way—indeed in recent research (Marsh et al., 2010, 2014; Perera, 2015), target rotation was suggested as the preferable choice for a strong theoretical model because it freely estimates the cross-loadings but closes them near zero—with poor results. Second, we applied an EFA that suggested a new brief version with a three-factor structure, which we named IGD-10SV. As a practical implication, a shortened version may be an easy tool for the assessment of addictions, especially within large batteries. Therefore, we compared the CFA and ESEM techniques to better represent the IGD-10SV. Indeed, the cross-loadings unforced to zero capture much of the correlation between the latent factors, allowing us to solve the multicollinearity issues that plagued previous versions of the instrument. Finally, we proposed the 3-factor ESEM model as the best factor solution for the IGD-10SV short version according to an eclectic approach (Marsh et al., 2010) which simultaneously considered goodness-of-fit indices, correlations among factors, and item level parameters. Despite the availability of other brief psychometric tools for Internet Gaming Disorder, such as the IGDS-9SF (Pontes & Griffiths, 2015), the present short form also maintained the dimensionality originally proposed for the assessment of the IGD, which allows for the detection of addictions in their widely recognized (Caretti et al., 2008; King et al.,

**Table 8.** *Convergent, concurrent, and discriminant validity of the 3-Factors Model.*

| | F1 Prevalence | F2 Withdrawal | F3 Craving |
|---|---|---|---|
| IGDS | .70*** | .62*** | .71*** |
| IAT | .59*** | .51*** | .64*** |
| DASS_Anxiety | .16** | .17** | .19** |
| DASS_Depression | .34*** | .29*** | .35*** |
| DASS_Stress | .23*** | .23*** | .26*** |
| LS | .35*** | .32*** | .32*** |
| SIAS | .39*** | .32*** | .35*** |
| RSES | -.36*** | -.28*** | -.36*** |
| MSPSS_Family | -.35*** | -.28*** | -.41*** |
| MSPSS_Friends | -.28*** | -.30*** | -.21*** |
| MSPSS_Significant Other | -.28*** | -.19*** | -.27*** |
| Hours of play_Daily | .40*** | .36*** | .47*** |
| Hours of play_Weekly | .38*** | .31*** | .46*** |
| Hours of play_Monthly | .38*** | .33*** | .50*** |

*Note.* *** $p < .001$, ** $p < .01$, * $p < .05$

IGDS= Internet Gaming Disorder Scale; IAT = Internet Addiction Test; DASS = Depression, Anxiety, Stress Scale- 21; LS = Ucla Loneliness Scale v.3; SIAS = Social Interaction Anxiety Scale; RSES = Rosemberg Self-Esteem Scale; MSPSS = Multidimensional Scale of Perceived Social Support; IGD-10SV = Internet Gaming Disorder-10 Short Version.

2017; Yen et al., 2022) components of prevalence, withdrawal, and craving.

The IGD-10SV's score validity was supported by significant associations with theoretically related measures. According to our H1, the IGD-10SV subscales showed strong positive associations with another measure of Internet Gaming Disorder, such as the IGDS, demonstrating that the test measures what it was designed to measure. Furthermore, they were strongly associated with the IAT; previous research (Wölfling et al., 2020) supports the framing of the IGD as a subset of the larger internet addiction. In fact, online gaming is only one of many activities available on the internet. The associations with measures of psychological symptoms such as anxiety, depression, stress, loneliness, and social interaction anxiety were used to investigate convergent validity. According to our H2 and previous research (Andreetta et al., 2020; Bum et al., 2018; González-Bueso et al., 2018; Ostinelli et al., 2021), these results support the negative effects of IGD on users' mental health and social skills: high levels of IGD (i.e., Prevalence, Withdrawal, and Craving) are positively correlated with high levels of psychological symptoms. Moreover, according to previous research (Paulus et al., 2018; Sioni et al., 2017; Uçur & Dönmez, 2021; Zeliha, 2019) and our initial H3, our findings about discriminant validity support the harmful impact on mental health of the moderate negative associations with self-esteem and different sources of perceived social support (i.e., family, friends, and significant others). As a clinical implication, the assessment of IGD should also include measures of psychiatric symptoms, self-perception, and social responsiveness in order to provide a multi-assessment that can account for the complexities of this addiction behavior. Furthermore, criterion-related validity suggests that spending more time playing online is associated with higher levels of all IGD-10SV subscales. Thus, as a practical implication, it is critical to supplement the assessment of this disorder with information on the time of use, as this is an important risk factor for the development of this behavioral addiction. Finally, the IGD-10SV has been shown to be a valid psychometric tool in terms of concurrent, convergent, and discriminant validity scores.

## Limitations and further implications

The present study contributed to the dimensionality of a new short version of IGD-20 using an innovative ESEM approach, which was able to resolve the strong multicollinearity problems in previous validations. A convenience small sample and cross-sectional design, however, were used. Further studies in Italian-speaking samples will be required to improve the generalizability of the results. It would also be useful for clinical implications in determining a clinical cut-off. These analyses were not possible due to the characteristics of our sample (i.e., there was no distinction between clinical and general population).

However, one advantage of this study was that it provided additional knowledge about addictive behaviours in clinical and research settings and further methodological contributions to the application of ESEM in psychological research. Anyway, at a psychometric level, we used the previous suggested golden rules for goodness-of-fit indices shared by the scientific community for SEM models (Hu & Bentler, 1999). Despite the fact that this problem is currently blocked by an eclectic approach (Marsh et al., 2010), further research is necessary to test the adequacy of these thresholds also in ESEM models. Furthermore, our outcomes may also contribute to a better understanding of how to distinguish online gaming behaviour from all other online behaviors, with the goal of accurately distinguishing Internet addictions (Schimmenti et al., 2014). Further studies should be conducted to investigate the relationship between IGD and other psychological and social implications for users' well-being in order to improve the multidimensional assessment of this complex pathological behaviour. Finally, additional research is needed to further examine this composite and developing pathological behavior.

## Conclusion

The purpose of this paper was to develop a psychometric model for assessing IGD using an evolving methodological approach (i.e., ESEM). On a theoretical level, the present paper helps to understand the complex construct of internet gaming disorder and its potential nomothetic definition in future diagnostic manuals. As a result of the psychometric implications, the ESEM method may provide a more faithful (i.e., adherent to the complex reality of psychological constructs and far from the stringent pure factors typical of the CFA approach) reading of cross-loadings than traditional methods. Furthermore, the possibility of using a new, brief version of a psychometric tool while retaining its multidimensionality is advantageous for practical assessment while also providing adequate depth for clinical practice. In fact, in accordance with current knowledge about behavioural addictions (Caretti et al., 2008; King et al., 2017; Yen et al., 2022), being able to assess the persistence of the disorder in space and time, the presence of withdrawal symptoms, and craving syndrome represents a practical advantage in both assessment and clinical practice. In conclusion, based on the current findings and the psychometric issues raised by previous studies, it appears that additional research addressing the complexity of the IGD construct would be beneficial.

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## Appendix 1

*Items for IGD-10SV*

4. When I am not gaming I feel more irritable.

6. I would like to cut down my gaming time but it's difficult to do.

10. I feel sad if I am not able to play games.

11. I have lied to my family members because the amount of gaming I do.

12. I do not think I could stop gaming.

13. I think gaming has become the most time-consuming activity in my life.

14. I play games to forget about whatever's bothering me.

15. I often think that a whole day is not enough to do everything I need to do in-game.

16. I tend to get anxious if I can't play games for any reason.

20. I believe my gaming is negatively impacting on important areas of my life.