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The Internet Gaming Disorder Scale

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Recently, the American Psychiatric Association included Internet gaming disorder (IGD) in the appendix of the 5th edition of the *Diagnostic and Statistical Manual of Mental Disorders* (*DSM–5*). The main aim of the current study was to test the reliability and validity of 4 survey instruments to measure IGD on the basis of the 9 criteria from the *DSM–5*: a long (27-item) and short (9-item) polytomous scale and a long (27-item) and short (9-item) dichotomous scale. The psychometric properties of these scales were tested among a representative sample of 2,444 Dutch adolescents and adults, ages 13–40 years. Confirmatory factor analyses demonstrated that the structural validity (i.e., the dimensional structure) of all scales was satisfactory. Both types of assessment (polytomous and dichotomous) were also reliable (i.e., internally consistent) and showed good criterion-related validity, as indicated by positive correlations with time spent playing games, loneliness, and aggression and negative correlations with self-esteem, prosocial behavior, and life satisfaction. The dichotomous 9-item IGD scale showed solid psychometric properties and was the most practical scale for diagnostic purposes. Latent class analysis of this dichotomous scale indicated that 3 groups could be discerned: normal gamers, risky gamers, and disordered gamers. On the basis of the number of people in this last group, the prevalence of IGD among 13- through 40-year-olds in the Netherlands is approximately 4%. If the *DSM–5* threshold for diagnosis (experiencing 5 or more criteria) is applied, the prevalence of disordered gamers is more than 5%.

Keywords: gaming disorder, game addiction, pathological gaming, Internet addiction, video games

The *Diagnostic and Statistical Manual of Mental Disorders* (*DSM*) is used by psychiatrists and psychologists in many countries as the main diagnostic tool for classifying psychiatric disorders. In the latest version of the *DSM* (5th ed. [*DSM–5*]), the American Psychiatric Association (APA) applied some changes to the descriptions and criteria for pathological behaviors and included *Internet gaming disorder* as a tentative disorder in the appendix of this manual (APA, 2013). The introduction of Internet gaming disorder in the *DSM–5* represents a major advance for the study, treatment, and prevention of problematic and pathological use of computer and video games. The crucial next step is to develop a survey instrument to measure Internet gaming disorder with solid psychometric properties that can be used for research and diagnostic purposes. Therefore, the main aim of the current study was to develop a valid and reliable survey instrument for Internet gaming disorder on the basis of the nine underlying criteria from the *DSM–5*. Because this instrument should be applicable to game-playing individuals of all ages, its properties were extensively tested among a representative sample of adolescents and adults (ages 13–40 years).

Over the last decade, many efforts have been made to define and measure the concept of pathological involvement with computer or video games. Although playing video games is not considered intrinsically pathologic or problematic, gaming can become pathological for some players when the activity becomes dysfunctional, harming an individual's social, occupational, family, school, and psychological functioning (Gentile et al., 2011). In general, *pathological gaming* can be described as persistent, recurrent, and excessive involvement with computer or video games that cannot be controlled, despite associated problems (Griffiths, 2005; Lemmens, Valkenburg, & Peter, 2009). Although *game addiction* is the most popular term to describe this disorder, this term is also considered ambiguous, because many players, developers, and reviewers use the term *addictive* as a positive adjective, indicating the enduring playability of a game and not destructive or pathological behavior (Adams, 2002). Most studies on game addiction or similar constructs have adapted the definition and criteria for pathological *gambling* from the *DSM–IV* (APA, 2000), and many have therefore applied the term pathological *gaming* to this type of behavior (e.g., Chiou & Wan, 2007; Gentile, 2009; Johansson & Götestam, 2004; Keepers, 1990; Lemmens, Valkenburg, & Peter, 2011a,2011b). After careful consideration by a multidisciplinary expert workgroup, the APA decided on the tentative term *Internet gaming disorder* in the *DSM–5* (Petry & O'Brien, 2013). Therefore, this term (or its abbreviation, IGD) and its underlying nine criteria will be used when addressing the measures under investigation in the current study.

The *DSM–5* states that only the use of Internet games must cause clinically significant impairment to constitute a diagnosis for IGD, not the use of sexual Internet sites, online gambling, or any other kind of Internet use (APA, 2013). Although the disorder is

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labeled "Internet" gaming disorder, the *DSM–5* states that "Internet gaming disorder most often involves specific Internet games, but it could involve non-Internet computerized games as well, although these have been less researched" (APA, 2013, p. 796). Indeed, several studies have indicated that pathological gaming is more prevalent among players of online games than it is among players of offline games (Ko, Yen, Liu, Huang, & Yen, 2009; Ng & Wiemer-Hastings, 2005; Seo, Kang, & Yom, 2009; Smyth, 2007; Tsitsika et al., 2009). However, pathological gaming is not limited to online games, as it has been associated with all sorts of offline computer and video games (e.g., Fisher, 1994; Griffiths & Hunt, 1995, 1998). In fact, several studies on pathological gaming were published years before online games became available to the general public (e.g., Shotton, 1989; Soper & Miller, 1983).

Regardless of whether IGD involves online games, offline games, or both, international efforts to assess this type of behavior have resulted in more than 18 tests, scales, and other measurement tools (for a review, see King, Haagsma, Delfabbro, Gradisar, & Griffiths, 2013). Although these instruments show various semantic and psychometric differences, many have adapted at least six of the criteria for pathological gambling (e.g., Gentile, 2009; Griffiths, 2005; Lemmens et al., 2009; Tejeiro Salguero & Bersabé Morán, 2002; Young, 1996, 2009) described in the *DSM–IV* (APA, 1994, 2000). These six criteria (*preoccupation, tolerance, withdrawal, persistence, escape*, and *problems*) can also be found among the proposed criteria for IGD in the *DSM–5*. In addition to these six criteria, the *DSM–5* included three other criteria that are diagnostic indicators of dysfunction: *deception* (e.g., Demetrovics et al., 2012; Gentile et al., 2011), *displacement* (e.g., Huang, Wang, Qian, Zhong, & Tao, 2007; Rehbein, Kleimann, & Mößle, 2010), and *conflict* (e.g., Lemmens et al., 2009; Young, 1996). In addition to the descriptions for the nine criteria provided by the *DSM*, we added labels to each criterion for clarity. The labels and adapted definitions for the nine *DSM–5* criteria for IGD can be found in Table 1. The words *Internet games* and *Internet gaming* have been replaced with *games* and *gaming* to reflect all types of computer and video games.

Development of the IGD Scales

Previous survey scales on pathological gaming generally applied either dichotomous or polytomous (i.e., continuous) response options when measuring the criteria for this behavior. Because the *DSM* adheres to dichotomous response options (i.e., *yes* or *no*) for diagnostic purposes, several researchers followed this dichotomous type of response in their measures of pathological gaming (e.g., Grüsser, Thalemann, & Griffiths, 2007; Tejeiro Salguero & Bersabé Morán, 2002). In an effort to increase the variance, and thereby the predictive power, of the measurement in survey research, other researchers have allowed for polytomous responses to the items, either through Likert-type response options (e.g., Gentile, 2009; Jap, Tiatri, Jaya, & Suteja, 2013) or ordinal frequency-based response options (e.g., Lemmens et al., 2009; Mehroof & Griffiths, 2010). Although both dichotomous and polytomous types of measurement generally contain identical sets of items and lead to highly similar results, the difference is reflected in the way both measurements are calculated. Dichotomous measures provide a cumulative score of affirmative answers, whereas polytomous answer options are often aggregated to reflect an individual's mean score. To obtain a valid and reliable instrument for IGD suited for diagnostic and research purposes, the current study aimed to statistically determine whether sum scores of affirmative answers on a dichotomous scale and mean scores on a polytomous measurement can both provide valid and reliable measurements.

One of the most important features of a psychometrically sound survey instrument is its ability to assess the prevalence of IGD in a population. For diagnostic purposes, the *DSM–5* recommends a threshold of experiencing five or more criteria. The proposed diagnostic cutpoint of five criteria was conservatively chosen because lower thresholds will inflate diagnoses and could result in the classification of individuals who have not suffered significant clinical impairment. Overdiagnosis thus holds potential to undermine the importance of true psychiatric disorders, especially in the context of tentative disorders such as IGD (Petry et al., 2014). A recent study confirmed that experiencing five or more criteria is indeed an adequate cutoff point for diagnosing IGD (Ko et al.,

Table 1
Nine Criteria for Internet Gaming Disorder From the DSM–5

Criterion	Description
Preoccupation	Preoccupation relates to being all-absorbed by gaming and spending substantial amounts of time thinking or fantasizing about gaming during times of nonplay.
Tolerance	Tolerance is characterized by an increasing amount of time spent on games to feel their desired effects (e.g., excitement, satisfaction).
Withdrawal	Withdrawal refers to symptoms that emerge when unable to play or attempting to cut down or stop gaming. Symptoms typically involve feeling restless, irritated, angry, frustrated, anxious, or sad.
Persistence	Persistence entails an enduring desire for gaming or unsuccessful attempts to stop, control, or reduce gaming.
Escape	Escape relates to engaging in a behavior to escape from or relieve negative mood states, such as helplessness, guilt, anxiety, or depression.
Problems	This criterion refers to continued gaming despite being aware of negative consequences of this behavior for central areas of life.
Deception	Deception refers to individuals lying to others about, or covering up the extent of, their gaming behaviors.
Displacement	The gaming behavior dominates, with a resulting diminishment of other social and recreational activities.
Conflict	This reflects more substantial issues as a result of gaming, referring to losing, or nearly losing, an important relationship or opportunity related to schooling or employment.

Note. *DSM–5* = *Diagnostic and Statistical Manual of Mental Disorders* (5th ed.; American Psychiatric Association, 2013).

2014). To examine the aptness of the *DSM–5* diagnostic cutoff point of our scales, latent class analyses (LCAs) were performed. LCA estimates unconditional contingencies into conditional contingencies using maximum likelihood (Nylund, Asparouhov, & Muthén, 2007). LCA has been used in previous studies to identify empirically based subgroups (i.e., classes) among all sorts of pathological and addictive behaviors, including online gambling (Lloyd et al., 2010), Internet addiction (Rumpf et al., 2014), and online game addiction (Gentile et al., 2011; van Rooij, Schoenmakers, Vermulst, Van den Eijnden, & Van de Mheen, 2011). LCA may indicate whether there are more than two discernable groups of gamers (instead of just disordered gamers and nondisordered gamers) and may provide empirical validation for the *DSM–5* diagnostic threshold (≥5 criteria) for diagnosing disordered gamers.

LCA has an additional advantage: It provides an indication of the sensitivity and specificity of each criterion for diagnosis. There is considerable disagreement among researchers about certain indicators of IGD and, therefore, a need to determine the specificity and sensitivity of indicators used in survey instruments (Demetrovics et al., 2012). A recent study on the criteria for IGD by Ko et al. (2014) found that the escape criterion was very frequently endorsed by all gamers (i.e., low specificity) and that the deception criterion could be removed because of low sensitivity (i.e., people diagnosed with IGD did not often report deceiving or lying to others about their game use). Gentile (2009) found that the deception criterion did add to diagnostic accuracy but that the escape criterion was not specific enough to differentiate gamers from pathological gamers. Finally, Charlton and Danforth (2007) found that tolerance and preoccupation were of limited use in the classification of people as addicted to online computer games. Because LCA provides conditional probabilities of affirmatively answering each item in the latent classes, it can thereby be used to identify which of the nine criteria are best suited to distinguish disordered gamers from other gamers.

Validation of the IGD Scales

Several psychosocial constructs related to pathological gaming are included in the current survey study to assess the construct validity of the measures for IGD. *Criterion validity* refers to the degree in which a scale construct correlates with constructs to which it theoretically should be related. Thus, the relation between several psychosocial constructs and IGD should be similar to comparable measures used in previous research (e.g., game addiction, pathological gaming, problematic gaming) and these psychosocial constructs. To assess criterion-related validity of the IGD measures, respondents' sum scores on the dichotomous IGD scale and mean scores on the polytomous IGD scale should be correlated with the following variables in the expected direction: more time spent on games, more loneliness, less life satisfaction, lower self-esteem, less prosocial behavior, and more aggression.

Loneliness

Loneliness has been defined as an unpleasant experience that derives from important deficiencies in a person's network of social relationships (Peplau & Perlman, 1982). Cross-sectional studies have consistently confirmed the relation between loneliness and addiction to online games (e.g., Qin, Rao, & Zong, 2007). Loneliness has been found to be both a cause and a consequence of pathological online gaming, thereby indicating a reciprocal relation (J. Kim, LaRose, & Peng, 2009; Lemmens et al., 2011b). These studies indicate that although playing online games may temporarily provide an escape from the negative feelings associated with social deficiencies, pathological gaming does little to facilitate the development or maintenance of real-life contacts. In fact, the resulting displacement of real-world social interaction is likely to result in deterioration of existing relationships, thereby increasing loneliness. Regardless of the causal order between these constructs, a positive relation between IGD and loneliness is expected.

Satisfaction With Life

Life satisfaction refers to a general cognitive assessment of a person's subjective well-being (Diener, Emmons, Larsen, & Griffin, 1985). Studies have shown that lower satisfaction with daily life is related to game addiction (Ko, Yen, Chen, Chen, & Yen, 2005; Lemmens et al., 2009; Shapira et al., 2003). Compulsive use of online games often seems to stem from the motivation to relieve real-life dissatisfaction (Chiou & Wan, 2007). These studies signify that we can expect a negative relation between IGD and life satisfaction.

Self-Esteem

Self-esteem has been defined as an evaluation of one's self-concept, heavily dependent on reflected appraisals, social comparisons, and self-attributions (Rosenberg, Schooler, & Schoenbach, 1989). Lower self-esteem among male gamers is associated with more severe addiction to online games (Ko et al., 2005). Regardless of gender, a negative relation between self-esteem and game addiction has been found in several studies (Lemmens et al., 2011b; Schmit, Chauchard, Chabrol, & Sejourne, 2011), suggesting that a similar relation can be found between IGD and self-esteem.

Prosocial Behavior

Prosocial behavior has generally been defined as voluntary, intentional behavior that results in benefits for another, for which the actor's motives are unknown or unspecified (Eisenberg & Miller, 1987). An early study on excessive gaming among children by Wiegman and van Schie (1998) found that heavy players of video games showed significantly less prosocial behavior than either the nonplayers or moderate player groups. More recently, Cao and Su (2007) found a correlation between online game addiction and lower scores on prosocial behavior. The propensity to act prosocially toward peers has been regarded as the most defining indicator of children's social competence (Ladd & Profilet, 1996). Several researchers have suggested that online games particularly appeal to people who are socially unskilled, have an unmet need for sociability in their offline lives, and feel anxious over establishing real-life interpersonal relationships (Chak & Leung, 2004; Chiu, Lee, & Huang, 2004; Peters & Malesky, 2008). Indeed, other researchers have shown that lower social competence is an important antecedent of pathological gaming (Lemmens et al., 2011b).

Aggression

Numerous studies on the relation between game addiction and physical aggression (e.g., Grüsser et al., 2007; E. J. Kim, Namkoong, Ku, & Kim, 2008; Lemmens et al., 2009) have reported significantly more aggressive behavior among pathological gamers compared with nonpathological players. A recent longitudinal study found a cross-lagged effect of pathological gaming on physical aggression among adolescent boys (Lemmens et al., 2011a). The authors argued that the displacement of important activities such as school or homework may eventually cause problems at school and conflicts with parents. When attempts are made to stop this excessive behavior, withdrawal symptoms following from abstinence after prolonged use can lead to irritability and aggression (e.g., Young, 2009).

Method

Sample and Procedure

In June 2013, two versions of the IGD survey (polytomous and dichotomous) were equally distributed among a representative sample of 2,444 adults and adolescents age 13–40 years from the Netherlands. Half of them received the IGD survey with dichotomous response categories (i.e., *yes* or *no*; $n = 1{,}251$, $M_{\text{age}} = 24.8$ years, $SD = 8.1$); 50.4% of these respondents were female ($n = 630$). This type of measurement is labeled the *dichotomous IGD scale*. The other half of respondents received the survey with six polytomous response categories to the IGD items, which ranged from (0) *never* to (5) *every day or almost every day* over the last 12 months ($n = 1{,}193$, $M_{\text{age}} = 24.4$ years, $SD = 7.6$); 51.6% of these respondents were female ($n = 615$). This type of measurement is labeled the *polytomous IGD scale*. Apart from the different response options to the IGD items, the items on the two scales were identical. In our analyses, the sample was divided into three age groups: adolescents, young adults, and middle adults. The total sample consisted of 44.6% adolescents ages 13–20 years ($n = 1{,}091$, $M_{\text{age}} = 17.6$ years, $SD = 2.2$), 28.9% young adults ages 21–30 years ($n = 706$, $M_{\text{age}} = 25.1$ years, $SD = 2.8$), and 26.5% middle adults ages 31–40 years ($n = 647$, $M_{\text{age}} = 35.9$ years, $SD = 2.8$).

The online surveys were distributed through MSI, an international market research company. From their panel of more than 400,000 nationally representative Dutch participants, respondents were quasi-randomly selected using a combination of stratified sampling and quota sampling (half our sample consisted of 13–20-year-olds). Each respondent was randomly assigned to a questionnaire with either the polytomous or the dichotomous IGD scale. Respondents received points for participating that could later be redeemed for prizes. Most participants completed the survey within 25 min. If respondents had not played at least one computer or video game in the month prior to participation in the survey, they were exempt from filling in any game-related questions, including the IGD items.

In the adolescent sample, 84.5% ($n = 922$) had played a game. In the young adult sample, 80.5% ($n = 568$) had played a game, and among middle adults, 67.2% ($n = 435$) reported having played at least one game in the month prior to the survey. Only these 1,925 gamers (polytomous scale $n = 932$ [78.1%], and dichotomous scale $n = 993$ [79.4%]) were initially included in our analyses. If a response to an IGD item was missing, respondents received a message reminding them to fill in this item. Therefore, there were no missing data. However, 13 respondents indicated playing games for more than 140 hours per week (nine from the polytomous scale and four from the dichotomous scale), and their standard z scores on this variable were higher than 3. Furthermore, they did not provide clear responses to questions about the titles of games that they had played for more than 20 hours each day (e.g., *don't know*, *none*, *?*). Moreover, they generally provided extreme responses to all other items on the survey, including conflicting responses to some reverse-coded scale items. Because these responses convinced us that these 13 respondents were untrustworthy, they were removed from further analyses, resulting in a total of 1,912 gamers (polytomous IGD scale $n = 923$, and dichotomous IGD scale $n = 989$) who were included in our analyses.

Measures

IGD. The nine criteria for IGD are described in the *DSM–5* in very broad terms. For instance, the criterion *conflict* is described as "has jeopardized or lost a significant relationship, job, or educational or career opportunity because of participation in Internet games" (APA, 2013, p. 795). Several researchers have suggested that to distinguish specific aspects of the *DSM* criteria (e.g., relationship, job, education), items can be broken into discrete components (Petry et al., 2014). Indeed, literal adaptation of the nine criteria into nine survey items would not provide information about which specific aspect of a broadly defined criterion matches the concept of disordered gaming. Thus, each of the nine *DSM–5* definitions was measured with three items, either through separating core aspects of a criterion into different items or by applying slight changes in phrasing or synonyms. Furthermore, "Internet gaming" or "Internet games" as used in the descriptions of the *DSM–5* criteria does not match with what is usually meant by Internet games (i.e., browser games) and was therefore replaced with "gaming" or "games." This was also done to avoid exclusion of offline games. All 27 items were in Dutch and randomly distributed over the scale. Three items were created for each of the nine previously identified criteria—preoccupation, tolerance, withdrawal, persistence, escape, problems, deception, displacement, and conflict—resulting in a total of 27 items. The dichotomous and polytomous IGD scales both consisted of the same items, differing only in their response options (see the Appendix).

According to the *DSM–5*, gaming disorder is present when a person meets five (or more) of the nine criteria *during a period of 12 months* (APA, 2013). In accordance with this temporal rule, every item on both IGD scales was preceded by this statement: "During the last 12 months . . ." If respondents had received the dichotomous scale, they rated all items with either *no* (0) or *yes* (1). For analyses of the dichotomous scale, all individual *yes* answers (range: 0–27) were summed ($M = 4.20$, $SD = 5.37$). Respondents who received the polytomous IGD scale rated all items on a six-point ordinal-frequency scale: (0) *never*, (1) *one to four times in the last year*, (2) *five to 11 times in the last year*, (3) *about once to three times a month*, (4) *once or more a week*, and (5) *every day or almost every day*. Individual mean scores on the polytomous IGD scale were calculated ($M = 0.58$, $SD = 0.91$).

The 27-item polytomous IGD scale had a Cronbach's alpha of .94, and the dichotomous IGD scale had a Cronbach's alpha of .93.

Time spent on games. The weekly time spent on computer and video games and the specific time spent on consoles, computers, and handheld gaming devices (both offline and online) was measured by multiplying the days per week by the number of hours per day spent on these activities. The mean hours per week gamers spent on games was 8.85 ($SD = 12.80$). We also asked respondents for the title of the game that they had played the most in the last 12 months and, if they had played another game a lot, the game they had played second to most in the last 12 months. Respondents were then asked if they had played each of these games online. Response options ranged from *never* (0) to *always* (4) ($M = 2.83$, $SD = 1.47$).

Loneliness. Loneliness was measured using the five items with the highest item-total correlations from the 20-item UCLA Loneliness Scale (Russell, 1996). Sample items are "I feel alone" and "I feel like there is no one I can turn to." Response categories ranged from 1 (*totally disagree*) to 5 (*totally agree*). Confirmatory factor analysis (CFA) with maximum likelihood estimation was used to test the model of this five-item scale. None of the error terms were correlated. CFA indicated good model fits: $\chi^2(5, N = 2{,}444) = 25.72$, $p < .001$, comparative fit index (CFI) = .998, root mean square error of approximation (RMSEA) = .041 (90% confidence interval [CI]: .026, .058), indicating structural validity. The items were averaged to create the scale scores. This five-item scale was reliable, with a Cronbach's alpha of .93 ($M = 2.31$, $SD = 0.99$).

Life satisfaction. Respondents' degree of life satisfaction was measured using the five-item Satisfaction with Life Scale developed by Diener et al. (1985). Examples of items are "I am satisfied with my life" and "In most ways my life is close to my ideal." Response categories ranged from 1 (*totally disagree*) to 5 (*totally agree*). CFA with maximum likelihood estimation was used to test the model of this five-item scale. None of the error terms were correlated. CFA indicated good model fits: $\chi^2(5, N = 2{,}444) = 58.95$, $p < .001$, CFI = .994, RMSEA = .066 (90% CI: .052, .082), indicating structural validity. The items were averaged to create the scale scores. This scale was reliable, with a Cronbach's alpha of .92 ($M = 3.28$, $SD = 0.91$).

Self-esteem. The degree of self-esteem was measured using five items from the Self-Esteem Scale (Rosenberg et al., 1989). This measure detects feelings of self-acceptance, self-respect, and generally positive self-evaluation. Sample items are "I am able to do things at least as well as other people" and "I feel that I don't have much to be proud of" (reverse coded). Response categories ranged from 1 (*totally disagree*) to 5 (*totally agree*). CFA with maximum likelihood estimation was used to test this five-item model. None of the error terms were correlated. CFA indicated acceptable model fits: $\chi^2(5, N = 2{,}444) = 109.79$, $p < .001$, CFI = .981, RMSEA = .093 (90% CI: .078, .108), indicating adequate structural validity. The items were averaged to create the scale scores. This scale was reliable, with a Cronbach's alpha of .82 ($M = 3.66$, $SD = 0.72$).

Prosocial behavior. Respondents' degree of prosocial behavior was measured using the five prosocial items from the Strengths and Difficulties Questionnaire (Goodman, 1997). For some items, the wording was slightly altered to tap adults as well as youths. This measure detects voluntary, intentional behavior that results in benefits for another (Eisenberg, 2000). Sample items are "I am considerate of other people's feelings" and "I often volunteer to help others." Response categories ranged from 1 (*totally disagree*) to 5 (*totally agree*). CFA with maximum likelihood estimation was used to test this five-item model. None of the error terms were correlated. CFA indicated unacceptable model fits: $\chi^2(5, N = 2{,}444) = 192.99$, $p < .001$, CFI = .968, RMSEA = .124 (90% CI: .109, .132), indicating that this measure did not meet the RMSEA requirement for structural validity. However, when the residuals (errors) from two similar items (i.e., "I am considerate of other people's feelings" and "I am kind to other people") were correlated, the model fits improved considerably: $\chi^2(4, N = 2{,}444) = 7.66$, $p < .001$, CFI = .999, RMSEA = .019 (90% CI: .000, .040). This scale was reliable, with a Cronbach's alpha of .87 ($M = 3.78$, $SD = 0.68$).

Aggressive behavior. Aggressive behavior was measured using seven items from the Physical Aggression subscale from Buss and Perry's (1992) Aggression Questionnaire. All seven items measured acts of physical aggression toward others (e.g., fighting, punching). Respondents were asked to reflect on the past 6 months when responding to items such as "There are people that pushed me so far that we came to blows" and "Once in a while I can't control the urge to strike another person." Response categories ranged from 1 (*totally disagree*) to 5 (*totally agree*). CFA with maximum likelihood estimation was used to test this seven-item model. None of the error terms were correlated. CFA indicated good model fits: $\chi^2(14, N = 2{,}444) = 122.60$, $p < .001$, CFI = .994, RMSEA = .056 (90% CI: .047, .066), indicating structural validity. The items were averaged to create the scale scores. This scale was reliable, with a Cronbach's alpha of .96 ($M = 1.48$, $SD = 0.81$).

Results

Differences Across Gender and Age

Out of 2,444 representative Dutch adolescents and adults between 13 and 40 years of age, 1,912 respondents (78.2%) reported having played a game in the last month. Male respondents were more likely to report having played a game in the last month ($n = 1{,}018$ [85.7%]) than female respondents ($n = 894$ [71.9%]), $\chi^2(1) = 798.21$, $p < .001$. Because the sizes of the three age groups were not equal, the harmonic mean sample size (773.47) was used when performing Scheffé post hoc comparison of age group differences. Adolescents ($n = 918$ [84.5%]) and young adults ($n = 564$ [80.3%]) were both more likely to play games than middle adults ($n = 430$ [67.0%]; $p < .001$). Among the respondents who indicated having played a game in the last month (i.e., the "gamers"; $n = 1{,}912$), male respondents spent more than twice as much time on games per week ($M = 12.45$, $SD = 18.09$) than females did ($M = 6.08$, $SD = 12.06$), $t(1{,}548) = 8.07$, $p < .001$. Similar significant gender differences in time spent on games were found within all three age groups. Although adolescent gamers spent slightly more time on games per week ($M = 10.20$, $SD = 15.19$) than young adult gamers ($M = 8.54$, $SD = 16.22$) and middle adult gamers ($M = 8.95$, $SD = 16.65$), these differences were not significant.

Regarding IGD, differences similar to time spent on games were found across age and gender: Men scored higher on the polytomous IGD scale ($M = 0.74$, $SD = 1.02$) than women ($M = 0.39$, $SD = 0.72$), $t(925) = 6.13$, $p < .001$. Men also scored higher on the dichotomous IGD scale ($M = 4.87$, $SD = 5.88$) than women ($M = 3.44$, $SD = 4.60$), $t(981.18) = 4.31$, $p < .001$. Analysis of variance (ANOVA) with a Scheffé post hoc test indicated that middle adults (ages 31–40 years) scored significantly lower on the dichotomous IGD scale ($M = 3.22$, $SD = 5.07$) than young adults ($M = 4.63$, $SD = 5.90$) and adolescents ($M = 4.48$, $SD = 5.14$; $p < .01$). The polytomous IGD scale showed a similar pattern of lower scores among adolescents and middle adults, but these scores did not differ significantly (adolescents: $M = 1.57$, $SD = 0.87$; young adults: $M = 1.61$, $SD = 0.97$; middle adults: $M = 1.51$, $SD = 0.91$).

Dimensional Structure of the IGD Scales

Our first aim was to investigate whether the nine *DSM* criteria for IGD can be accounted for by one higher order factor: IGD. We used structural equation modeling (SEM) with weighted least squares estimators to test these second-order factor models using CFA in MPlus (Asparouhov & Muthén, 2009). Although maximum likelihood is the most common estimation method in CFA, this method assumes that observed variables are continuous and normally distributed in the population (Lubke & Muthén, 2004). Because this assumption was not met with our skewed distribution of IGD and ordinal levels of measurement, a weighted least squares approach was applied to our data, allowing any combination of dichotomous, ordered categorical, or continuous observed variables (Flora & Curran, 2004). Three items were used as indicators for each of the nine latent criteria. The error terms (residuals) associated with each observed item were all uncorrelated (Byrne, 2001). The correlations among the nine latent criteria can be entirely explained by one higher order factor: IGD. The goodness of fit was evaluated using the chi-square value, the CFI, the weighted root mean square residual (WRMR), and the RMSEA and its 90% CI. Particularly when dealing with large samples, the CFI, WRMR, and RMSEA indices are considered informative fit criteria in SEM. A good fit is expressed by a CFI greater than .95, a WRMR value below 1.0, and an RMSEA value less than .08 (Byrne, 2001; Hu & Bentler, 1999).

The 27-item second-order factor model resulted in an acceptable model fit for the polytomous IGD scale, $\chi^2(315, N = 923) = 959.420$, $p < .001$, CFI = .991, WRMR = 1.005, RMSEA = .047 (90% CI: .043, .050). The same model for the dichotomous IGD scale also showed an acceptable model fit, $\chi^2(315, N = 989) = 486.825$, $p < .001$, CFI = .989, WRMR = .966, RMSEA = .019 (90% CI: .019, .027). Although some measurement and structural loadings differed between the two models, Table 2 shows that these differences were consistently small. Despite a minor transgression of the WRMR value for the polytomous IGD scale (0.005 above 1.0), the overall structure of the second-order factor model showed a good fit that was similar across the two measurement types. The second-order factor model, including the standardized mean measurement loadings of the 27 observed items on the nine first-order factors (i.e., the mean loadings of items across dichotomous and polytomous scales) is shown in Figure 1. It also shows the standardized loadings of the first-order factors on the second-

Table 2
Means and Confirmatory Factor Loadings for IGD Polytomous and Dichotomous Scales

No.	Criterion label	Polytomous IGD (*M*; range: 0–4)	Dichotomous IGD (% *yes*)	Polytomous measurement loadings (β)	Dichotomous measurement loadings (β)
1	Preoccupation 1	0.80	22	.840	.788
2	Preoccupation 2	0.47	10	.928	.934
3	Preoccupation 3	0.47	13	.915	.858
4	Tolerance 1	1.25	41	.768	.730
5	Tolerance 2	0.82	25	.874	.814
6	Tolerance 3	0.47	12	.982	.930
7	Withdrawal 1	0.52	12	.921	.855
8	Withdrawal 2	0.48	12	.941	.914
9	Withdrawal 3	0.46	12	.954	.941
10	Persistence 1	0.63	15	.853	.900
11	Persistence 2	0.49	11	.929	.920
12	Persistence 3	0.41	10	.984	.974
13	Escape 1	0.95	28	.899	.937
14	Escape 2	0.89	31	.947	.970
15	Escape 3	0.82	27	.939	.916
16	Problems 1	0.36	6	.887	.809
17	Problems 2	0.66	24	.778	.525
18	Problems 3	0.41	10	.934	.813
19	Deception 1	0.49	13	.910	.825
20	Deception 2	0.42	12	.953	.913
21	Deception 3	0.48	16	.908	.849
22	Displacement 1	0.71	14	.896	.846
23	Displacement 2	0.60	15	.917	.891
24	Displacement 3	0.51	13	.940	.851
25	Conflict 1	0.38	7	.915	.840
26	Conflict 2	0.34	6	.937	.958
27	Conflict 3	0.31	5	.964	.913

Note. IGD = Internet gaming disorder; No. = number.

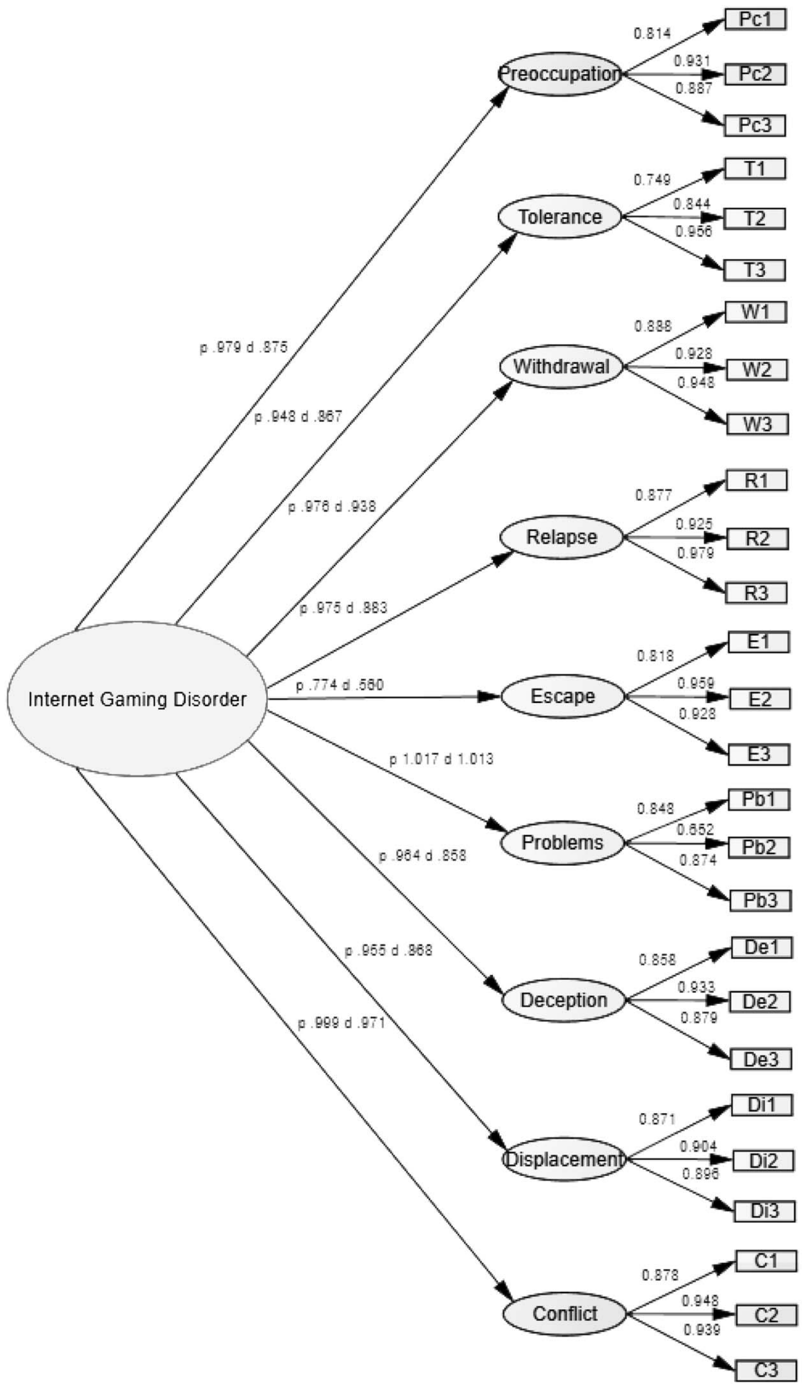

Figure 1. Second-order factor model with standardized mean measurement loadings and structural loadings for the polytomous scale (p) and the dichotomous scale (d).

order factor (the structural loadings) for the polytomous (p) and dichotomous (d) scales.

As Figure 1 shows, both standardized structural loadings of the problem criterion were slightly higher than 1, also known as a Heywood case (Bollen, 1987). A plausible explanation is that influential outliers (i.e., disordered gamers) were causing this Heywood case (Kolenikov & Bollen, 2012). A solution to this problem, as suggested by Bollen, is to remove influential outliers. Although removing univariate outliers on IGD from our sample did solve the Heywood case for the standardized structural loadings of problems (p = .996, d = .994), it severely affected all other loadings. More important, because these outliers on IGD were likely among the few disordered gamers, removing them from the data was not a good solution for theoretical reasons. Others have suggested that contrary to conventional wisdom, it is possible for standardized coefficients to be larger than 1 in absolute value as a function of both the degree of multicollinearity present among the set of independent variables and the correlation of each of the independent variables with the dependent variable (Deegan, 1978). Thus, we argue that the standardized loadings higher than 1 for problems could be explained by the similarities with other criteria (e.g., deception, conflict) and are, therefore, not necessarily a reason for concern. The mean scores on the IGD scale items, the percentage of affirmative answers on the items from the dichotomous IGD scale, and the standardized measurement loadings (β) for both scale models are shown in Table 2.

To determine the reliability of the factor loadings and model fits, the samples were split in half so as to assess whether the proposed higher order model could be generalized to both halves of the data sets. For this split-sample reliability, both the dichotomous and polytomous data sets were randomly split in two halves, and model fits were calculated for all four halves. The first half of the polytomous IGD scale sample provided an acceptable fit, $\chi^2(315, N = 466) = 800.698$, $p < .001$, CFI = .984, WRMR = 1.010, RMSEA = .058 (90% CI: .053, .062). The second half of the polytomous sample also provided an acceptable model fit, $\chi^2(315, N = 463) = 564.043$, $p < .001$, CFI = .997, WRMR = 0.994, RMSEA = .041 (90% CI: .036, .047). The dichotomous scale provided good model fits for the first half, $\chi^2(315, N = 496) = 398.939$, $p < .001$, CFI = .991, WRMR = 0.858, RMSEA = .023 (90% CI: .015, .030), and second half of the sample, $\chi^2(315, N = 497) = 378.966$, $p < .001$, CFI = .986, WRMR = 0.876, RMSEA = .020 (90% CI: .011, .027). Thus, all four models provided reliable model fits across split samples.

Because of gender differences in both mean and sum IGD scores, we examined whether the factor structure held for male and female respondents. To do so, we performed multigroup analysis (Jaccard & Wan, 1996) to test whether gender differences in the structural loadings were statistically significant. First, we estimated a model in which we did not pose any cross-group constraints—that is, we allowed the nine structural paths to vary between male and female respondents. In a subsequent model, we constrained, one at a time, each structural path to be equal across genders. Finally, we used both the polytomous and the dichotomous sample to test whether the fit of the constrained model differed from the fit of the unconstrained model for all nine structural loadings. A significant change in model fit would indicate that the constrained path differed between male and female respondents. Neither the polytomous nor the dichotomous sample showed significant chi-square changes in structural loadings across gender, indicating that the structural loadings of the nine criteria shown in Figure 1 provided similar reflections of the higher order construct across genders.

Constructing Short IGD Scales

To facilitate incorporation of both IGD scales into space-limited surveys, an additional aim of this study was to investigate whether a model with fewer items would provide an equal or even better description of the data. To create short versions of the scales, the highest loading item from each criterion was selected to create nine-item versions of the scales that encompassed all criteria. Because measurement loadings slightly differed between the polytomous and dichotomous models (the criteria displacement and conflict provided different items with the highest measurement loading), we used the combined mean measurement loadings (last column from Table 2) to select one set of nine items with the highest overall loadings from each of the nine first-order factors. These short versions of the scales were then tested as first-order structural models using CFA in Mplus.

The unconstrained model for the short, nine-item polytomous IGD scale yielded an acceptable model fit, $\chi^2(16, N = 932) = 112.240$, $p = < .001$, CFI = .988, WRMR = 0.810, RMSEA = .080. The unconstrained model for the short dichotomous IGD scale yielded a good model fit, $\chi^2(23, N = 993) = 28.96$, $p = < .001$, CFI = .997, WRMR = 0.639, RMSEA = .016. Items for the short, nine-item IGD scales are displayed in Table 3. The short, nine-item polytomous IGD scale was strongly correlated with the 27-item polytomous IGD scale ($r = .98$, $p < .001$) and showed good reliability, with a Cronbach's alpha of .95 ($M = 0.50$, $SD = 0.93$). Similarly, the short, nine-item dichotomous IGD scale was strongly correlated with the dichotomous 27-item scale ($r = .93$, $p < .001$) and proved reliable, with a Cronbach's alpha of .83 ($M = 1.18$, $SD = 1.92$). At the level of measurement, therefore, all versions of the scale yielded good results.

Criterion Validity of the IGD Scales

To determine the validity of our construct, respondents' mean scores on the polytomous IGD scales and sum scores on the dichotomous IGD scales should be correlated with time spent on games, loneliness, life satisfaction, self-esteem, prosocial behavior, and aggression in the expected directions. In general, criterion validity is quantified by significant correlation coefficients between measurements. As Table 4 shows, both the long (27-item) and short (nine-item) versions of the polytomous and dichotomous IGD scales generally showed significant small-to-moderate correlations with time spent on games, loneliness, self-esteem, and prosocial behavior. The correlations with aggressive behavior were remarkably strong across scales. With one exception (life satisfac-

Table 3
Items for the Short, Nine-Item IGD Scales

Criterion	During the last year . . .
Preoccupation	have there been periods when all you could think of was the moment that you could play a game?
Tolerance	have you felt unsatisfied because you wanted to play more?
Withdrawal	have you been feeling miserable when you were unable to play a game?
Persistence	were you unable to reduce your time playing games, after others had repeatedly told you to play less?
Escape	have you played games so that you would not have to think about annoying things?
Problems	have you had arguments with others about the consequences of your gaming behavior?
Deception	have you hidden the time you spend on games from others?
Displacement	have you lost interest in hobbies or other activities because gaming is all you wanted to do?
Conflict	have you experienced serious conflicts with family, friends or partner because of gaming?

Note. IGD = Internet gaming disorder.

Table 4
Correlations Between IGD and Measures of Criterion Validity

Variable	27-item polytomous	9-item polytomous	27-item dichotomous	9-item dichotomous
Time on games	.23*	.23*	.22*	.21*
Loneliness	.41*	.39*	.31*	.30*
Prosocial	−.11*	−.12*	−.17*	−.18*
Life satisfaction	−.05[a]	−.04[b]	−.16*	−.14*
Self-esteem	−.11*	−.10[c]	−.21*	−.20*
Aggression	.60*	.61*	.43*	.43*

[a] $p = .10$. [b] $p = .28$. [c] $p = .002$.
* $p < .001$.

tion), all correlations were significant at least at $p < .001$ and in the expected directions. The correlations with the criterion variables were highly comparable across the long and short scales (*r*s difference ≤ .02), with one exception: The correlations between life satisfaction and polytomous scales were nonsignificant ($r = -.05$, $p = .10$, and $r = -.04$, $p = .28$), whereas the dichotomous scales did show significant negative correlations ($r = -.16$, $p < .001$, and $r = -.15$, $p < .001$). Because of these scale differences in correlations with life satisfaction, the dichotomous scales showed more criterion-related validity than the polytomous scales. The correlations of the nine-item scales were not much different from those of the 27-item versions, indicating that the nine-item versions of the scales are just as valid as the longer versions.

Diagnosing IGD

The nine-item dichotomous IGD scale was used to assess the prevalence of gaming disorder among 989 adolescent, young adult, and middle adult gamers. The majority of gamers ($n = 532$ [53.8%]) did not report any signs of IGD whatsoever (i.e., scored 0 on the nine-item scale). Combined with the respondents who did not report playing games ($n = 262$), we found that 63% of all 13–40-year-olds did not experience any signs of IGD. To assess the prevalence of IGD, two types of assessments of disordered gamers were applied: (a) the *DSM–5* cutoff point of five or more criteria and (b) LCA of a disordered group. First, according to the threshold applied by the *DSM–5*, endorsement of at least five of the nine criteria is required for a positive diagnosis. Among gamers who filled in the dichotomous IGD scale ($N = 989$), we found that 67 players had responded *yes* to five or more of the criteria (6.8%). Compared with the whole sample of respondents, regardless of whether they played computer or video games ($N = 1{,}247$), this corresponds with 5.4% of all respondents (ages 13–40 years) who could be considered disordered gamers according to the proposed *DSM–5* threshold (APA, 2013). A stricter cutoff point for IGD (≥6 criteria) yielded 51 disordered gamers, 4% of all respondents ages 13–40 years (5.2% of all gamers). The prevalence of disordered gaming was slightly higher among male respondents (6.8%) than among female respondents (4.0%), $\chi^2(N = 1{,}247) = 4.94$, $p = .026$. However, comparing gender differences only among gamers, the difference in disordered gamers among male (8.0%) and female respondents (5.4%) was found to be not significant, $\chi^2(N = 989) = 2.50$, $p = .112$. Although the prevalence of gaming disorder among young adults (ages 21–30 years) was higher (6.7% [8.5% of young adult gamers]) than the prevalence of IGD among adolescents (ages 13–20 year; 5.5% [6.4% of adolescent gamers]) and middle adults (ages 31–40 years; 3.9% [5.6% of middle adult gamers]), these prevalence differences between age groups did not differ significantly. Similarly, differences between male and female respondents within age groups (adolescents: 7.1% and 3.9%; young adults: 8.9% and 5.1%; middle adults: 5.5% and 2.0%, respectively) did not differ significantly.

The second type of assessment of the prevalence of IGD came from an LCA on the short, nine-item IGD scale. LCA provided an estimate of the number of latent classes within IGD, the prevalence of players in those classes, and the conditional probabilities of affirmative answers on each item within these classes. There are two preferred methods of determining the ideal number of classes: The bootstrap likelihood ratio test (BLRT) and the Bayesian information criterion (BIC). The BLRT is a parametric bootstrap method described in McLachlan and Peel (2000). By using bootstrap samples to estimate the distribution of the log likelihood difference test statistic, the BLRT compares the improvement in fit between the current model and a model with one less class. A significant *p* value would indicate a statistically significant better fit for the current model (Nylund et al., 2007). The second common method of selecting the number of latent classes is the BIC, with lower scores indicating better solutions. For categorical LCA models, the BIC consistently provides an accurate identification of the number of classes across all model types and sample sizes (Jedidi, Jagpal, & DeSarbo, 1997; Nylund et al., 2007; Yang, 2006). Out of convention, we also report the log likelihood value and Akaike information criterion (AIC).

The comparison of six latent classes is shown in Table 5. Testing for additional classes provided no relevant information. The lowest BIC score was provided by a three-class solution. The BLRT also showed no significant *p* values after the third class, indicating that no significant information was added after a three-class solution. Thus, a three-class solution fit the data better than a two-class solution, as proposed in the *DSM–5*. On the basis of the properties of the classes, the groups were labeled as *normal gamers* ($n = 784$ [79%]), *risky gamers* ($n = 157$ [16%]), and *disordered gamers* ($n = 48$ [4.9%]). Normal gamers scored between 0 and 2 on the dichotomous 9-item scale ($M = 0.36$, $SD = 0.56$). Risky gamers' scores ranged between 1 and 6 ($M = 3.08$, $SD = 1.08$). The last group of 48 disordered gamers' scores ranged between 6 and 9 ($M = 7.30$, $SD = 1.23$) on the nine-item IGD

Table 5
Comparison of the Number of Latent Classes

Latent classes	Log likelihood	AIC	BIC	BLRT p value for $k - 1$
One	−3,271.1	6,560.27	6,604.4	—
Two	−2,622.0	5,281.98	5,375.1	<.001
Three	−2,546.5	5,150.99	5,293.1	<.001
Four	−2,530.3	5,138.58	5,329.7	.27
Five	−2,516.6	5,131.27	5,371.4	.35
Six	−2,506.6	5,131.14	5,420.3	.41

Note. AIC = Akaike information criterion; BIC = Bayesian information criterion; BLRT = bootstrap likelihood ratio test.

scale. On the basis of the properties of this third class, 4.9% of gamers, or approximately 3.8% of all Dutch 13–40-year-olds, could be labeled as disordered gamers. It is interesting to note that the percentage of male respondents among these disordered gamers (63%) was identical to the percentage of male respondents among disordered gamers when using the *DSM* assessment (≥5 criteria).

The criterion validity of both types of assessment (*DSM* and LCA) was examined by comparing scores on the criterion-related constructs between disordered gamers and other groups. For the LCA assessment, an ANOVA with a Scheffé post hoc test was performed with the three classes (normal gamers, risky gamers, and disordered gamers). For the *DSM* assessment, *t* tests were performed between disordered and nondisordered gamers. The differences between groups for both assessment types are shown in Table 6. Overall, both assessments of disordered gamers indicated good criterion-related validity. The three groups from the LCA showed significant differences between normal gamers, risky gamers and disordered gamers on loneliness and aggression, with disordered gamers showing significantly more aggression and loneliness than the other two groups. Furthermore, disordered gamers displayed significantly less self-esteem and spent almost 10 hours more on games per week than normal gamers did. Similarly, *t* tests between disordered and nondisordered gamers using the *DSM* assessment showed significant differences in time spent on games, loneliness, self-esteem, prosocial behavior, and aggression. Life satisfaction did not differ significantly between any of the groups.

Sensitivity and Specificity of the IGD Items

Both types of assessment of disordered gamers—(a) five or more *DSM* criteria and (b) the disordered group from the LCA—were examined for the sensitivity and specificity of the nine indicators for diagnosis. Sensitivity is demonstrated by the proportion of disordered gamers who answered positively on an indicator. Specificity is indicated by the proportion of nondisordered gamers who reported a negative answer to an indicator. Ideally, both sensitivity and specificity of an item should be high to discriminate false positives and false negatives (Glaros & Kline, 1988). As Table 7 shows, both assessment types indicated adequate sensitivity for the items and high specificity for most items. In general, the LCA assessment showed higher sensitivity and slightly lower specificity than the *DSM* assessment, as would be expected considering the higher mean score on the nine-item IGD scale for the LCA assessment of the 48 disordered gamers (M = 7.33, SD = 1.23) compared with the *DSM* mean score on the nine-item scale for the 67 disordered gamers (M = 6.72, SD = 1.44). The diagnostic accuracy, as indicated by the proportion of all true positives and true negatives, was almost identical across assessment types.

The sensitivity of conflict was the lowest of all items across both assessment types (.66 for *DSM*, .73 for LCA), indicating that between 66% and 73% of all disordered gamers had experienced serious conflicts with friends, family, or partners because of games. Conversely, the specificity of conflict was highest across assessment types (.99 for *DSM*, .98 for LCA), indicating that only one or two players who had experienced this type of conflict were not among the disordered gamers. This resulted in conflict acquiring the highest diagnostic accuracy (97% across assessments). The item with the lowest specificity was escape (.73 for *DSM*, .72 for LCA), which indicated that 27%–28% of all nondisordered gamers had played games to forget about annoying things. Escape also had the lowest diagnostic accuracy (.74 for *DSM*, .73 for LCA), much lower than all other indicators (≥.90). To determine whether this relatively low specificity and diagnostic accuracy of escape was caused by the item that was selected for the nine-item scale, the specificity and diagnostic accuracy of the two other items for escape from the 27-item scale were examined for both assessments of disordered gamers. The specificity of escape item 1 (played games to forget about problems) showed low specificity across assessments (.75 for *DSM*, .74 for LCA), with similar diagnostic

Table 6
Mean Differences Between Groups With LCA and DSM–5 Assessment Types for IGD

Variable	Light gamers (n = 784)	Heavy gamers (n = 157)	Disordered gamers (n = 48)	*DSM–5* <5 criteria (n = 922)	*DSM–5* ≥5 criteria (n = 67)
IGD short score	0.36[a]	3.08[a]	7.33[a]	0.73[x]	6.72[x]
Time on games	7.64[a,b]	18.09[a]	17.60[b]	9.26[x]	16.03[x]
Loneliness	2.25[a]	2.61[a]	3.12[a]	2.29[x]	3.13[x]
Self-esteem	3.72[a,d]	3.57[d]	3.26[a,d]	3.70[x]	3.29[x]
Life satisfaction	3.36	3.18	3.12	3.33	3.16
Prosocial	3.83[a,d]	3.67[d]	3.36[a,d]	3.81[x]	3.44[x]
Aggression	1.36[a]	1.81[a]	2.70[a]	1.42[x]	2.65[x]

Note. Coefficients in rows with identical superscripts a, b, and x differ significantly at $p < .001$. Coefficients in rows with identical the superscript d differ significantly at least at $p < .05$. LCA = latent class analysis; *DSM–5* = *Diagnostic and Statistical Manual of Mental Disorders* (5th ed.; American Psychiatric Association, 2013); IGD = Internet gaming disorder.

Table 7
Sensitivity (Sens) and Specificity (Spec) of the Nine Criteria for Assessing Disordered Gamers

Criterion	All gamers (N = 993) *Yes*	Disordered gamers LCA (n = 48) *Sens/Spec*	LCA diagnostic Accuracy	Disordered gamers *DSM* (n = 67) *Sens/Spec*	*DSM* diagnostic Accuracy
Preoccupation	.10	.81/.94	.93	.67/.95	.93
Tolerance	.12	.77/.91	.91	.78/.93	.92
Withdrawal	.12	.92/.93	.93	.82/.94	.93
Persistence	.10	.92/.95	.95	.78/.96	.94
Displacement	.13	.73/.90	.90	.70/.92	.90
Problems	.10	.73/.93	.92	.70/.95	.93
Deception	.12	.85/.92	.92	.75/.93	.92
Escape	.31	.88/.72	.73	.87/.73	.74
Conflict	.06	.73/.98	.97	.66/.99	.97

Note. LCA = latent class analysis; *DSM* = *Diagnostic and Statistical Manual of Mental Disorders.*

accuracy (.75 for *DSM*, .74 for LCA). Escape item 3 (played games to escape negative feelings) also showed low specificity (.75 for *DSM*, .75 for LCA) and comparable diagnostic accuracy (.75 for *DSM*, .75 for LCA). Thus, regardless of assessment type or item formulation, the escape criterion shows low specificity when distinguishing disordered gamers from nondisordered gamers.

Discussion

Numerous studies over the last few decades have demonstrated that the concept of pathological involvement with computer or video games is valid enough for the APA to include IGD in the *DSM–5* as an issue worth further study. Because measurement tools are needed for research and diagnostic purposes, the main aim of the current study was to test the reliability and validity of four survey measures for IGD based on the nine criteria from the *DSM–5* (APA, 2013). The properties of a 27-item IGD scale with either polytomous or dichotomous response options were tested among a representative sample of 2,444 Dutch adolescents and adults aged 13–40 years. CFAs indicated that both scales had good psychometric properties, as indicated by the fit indices of the models. On the basis of the factor loadings, the most suited item from each of the nine criteria was selected for short, nine-item versions of the scales, which also showed good model fits, indicating solid structural validity.

Criterion-related validity of both long and short versions of the scales was indicated by the significant correlations with time spent on games, loneliness, self-esteem, life satisfaction, prosocial behavior, and aggression. In general, higher mean scores on the polytomous scales and higher sum scores on the dichotomous scales indicated more loneliness, lower self-esteem, less life satisfaction (although this correlation was not significant for the polytomous scale), and more aggressive behavior. Thereby, these players showed a pattern that would be predicted on the basis of what is known about pathological gaming, substance and gambling addictions, and what would theoretically be expected if IGD is a legitimate health issue. In particular, the relatively strong correlations with loneliness and aggressive behavior are considerable reasons for concern, especially because these two social characteristics are known to increase as a result of prolonged pathological involvement with games (Lemmens et al., 2011a, 2011b).

Although several researchers have argued the relative merits of either a dichotomous or continuous approach to measuring pathological use of games, no studies have empirically compared these types of measurements. Sum scores of affirmative answers on the dichotomous scales showed correlations with all criterion-related measures, whereas the polytomous scales showed significant correlations with all criterion-related measures except life satisfaction. More important, sum scores of dichotomous answers on the nine-item scale could best be used to assess the prevalence of IGD when the *DSM–5* cutoff point of five or more (out of nine) was administered. Although the IGD items for the scales were identical across measurement types, direct comparison of prevalence rates between polytomous and dichotomous scales was not possible, because the items on the polytomous scale did not provide a clear cutoff point for affirmative answers, whereas the sum score on the dichotomous scale provided a straightforward indication of prevalence. When the *DSM* cutoff point of five or more was applied to determine the prevalence of IGD among the Dutch sample of 13–40-year-olds, we found that more than 5% of this population met the criteria for a positive diagnosis. Although most of the disordered gamers were male (63%), the gender difference in IGD was considerably smaller than that reported in previous studies on similar constructs (Gentile, 2009; Ko et al., 2009). This study lends support to the *DSM* style of dichotomous measurement for screening purposes, although we do not support diagnosis based on this measurement alone. A clinical interview in addition to information from family members, friends, or partners would probably be the most effective method to determine with the highest degree of accuracy if a person is suffering from gaming disorder.

To investigate the aptness of the *DSM–5* threshold for diagnosis, an LCA was conducted to assess empirically based subgroups (i.e., classes) of gamers on the basis of endorsement of the nine dichotomous IGD criteria. Three subgroups were identified: normal gamers, risky gamers, and disordered gamers. The relative prevalence of this last group was slightly lower (3.8%) than when prevalence was assessed using the *DSM–5* threshold (5.0%). However, all disordered gamers in the LCA group had experienced at least six of the nine criteria, thus indicating that raising the threshold for diagnosis to six out of nine may be appropriate so as to avoid overdiagnosing disordered gamers (Petry, 2005). Furthermore, when the cutoff point was raised to six or more criteria for positive diagnosis, the prevalence of IGD came closer (4.0%) to the prevalence of the disordered group indicated by the LCA. Nonetheless, because every endorsed item documents some damage to functioning, five symptoms may be a reasonable threshold, because it demonstrates potentially significant clinical impairment. Regardless of the type of assessment, disordered gamers showed significantly less self-esteem, less prosocial behavior, more loneliness, and more aggression than nondisordered gamers.

Another aim of the LCA was to examine sensitivity and specificity of each criterion for diagnosis. Sensitivity was demonstrated by the proportion of disordered gamers who answered positively on an indicator, whereas specificity was indicated by the proportion of negative responses among nondisordered gamers. Using the short, nine-item dichotomous IGD scale, both types of assessment (LCA and *DSM*) showed high specificity and adequate sensitivity, resulting in good diagnostic accuracy across assessment types. Although the specificity of both assessment types was almost identical, the LCA assessment showed higher sensitivity than the *DSM* assessment, indicating that the LCA assessment of disordered gamers included more gamers who had experienced each criterion. High specificity may be prioritized when identifying clinical cases for administration of treatment in an inpatient setting, whereas prioritizing sensitivity may be more appropriate in epidemiological research (King et al., 2013).

Regarding the appropriateness of the nine IGD criteria, gamers in general are much less likely to experience conflict (6%) than escape (31%), indicating that not all criteria are equally prevalent and that some items may provide better discriminative power when diagnosing this disorder than others. Although escape may have shown good sensitivity (more than 87% of disordered gamers had played games to escape problems), it showed low specificity (more than a quarter of nondisordered gamers also played to escape problems). This finding supports conclusions from Gentile (2009) and Ko et al. (2014), who also found that escape did not add to diagnostic accuracy because of the lack in specificity. Because this criterion is also experienced by the vast majority of disordered gamers, there is no immediate reason to remove it from the scale. However, its low specificity does provide an explanation of why the implied LCA threshold of six or more criteria for positive diagnosis seems appropriate. Contrary to suggestions from previous studies regarding the low specificity of tolerance and preoccupation (Charlton & Danforth, 2007) or the low sensitivity of deception (Ko et al., 2014), both types of assessment indicated that these criteria were useful in distinguishing disordered gamers from nondisordered gamers.

When studying any disorder with a low prevalence rate, larger samples are necessary for increased confidence in the measures and prevalence rates among the group. This study used a large, nationally representative sample with a much wider age range than has typically been included in gaming disorder research. The prevalence of IGD among Dutch 13–40-year-olds (i.e., 5.4% using the *DSM* assessment or 3.8% using the LCA assessment) was slightly lower than the prevalence of pathological gaming, or similar constructs, reported in other countries, such as the United States (8.5%; Gentile, 2009), Germany (11.9%; Grüsser et al., 2007), 7.5% in Taiwan (Ko, Yen, Yen, Lin, & Yang, 2007) and Singapore (8.7%; Choo et al., 2010). It is possible that the slightly lower prevalence rate is related to the inclusion in our sample of a substantial group of middle adults (ages 31–40 years), who showed less signs of IGD than the adolescents or young adults did. The dichotomous short IGD scale and *DSM* cutoff point seem to support, or at least not invalidate, all previously published measures, albeit with a slightly more conservative estimate of the prevalence of gaming disorder. However, despite the high reliability and validity of both the long and short versions of our IGD scale, the short versions are derived from specific aspects of the broad *DSM* criteria. Therefore, prevalence rates based on these scales should be interpreted with caution until their items are validated in a clinical or diagnostic setting.

Future studies may examine how increasing the threshold for diagnosing disordered gamers to six or more criteria will affect prevalence estimates and relations with criterion-related variables. In addition to the dichotomous scale, the frequency in which the criteria are met (i.e., the polytomous IGD scale) may provide more information on the severity of the addiction than dichotomous answers alone can provide. It is also important to note that several of the items on the short scale (i.e., persistence, problems, deception, and conflict) cannot exist within a social vacuum, which means that some form of social struggle surrounding games is necessary for these criteria to occur. It seems plausible that players who live without a partner, roommate, or family may be underdiagnosed with this scale because they simply do not experience the social surveillance required for these criteria. Future studies could include measures of home, family, or other social circumstances that may provide additional information on the social aspects of some criteria for IGD.

Several shortcomings of our study should be addressed. First, the use of an online survey excludes people who do not have access to the Internet. Although 98% of the Dutch population has home access to the Internet, specifically sampling some respondents who do not use the Internet would make the sample more representative and possibly less biased toward use of online games and computers. Another shortcoming is that we did not examine convergent validity of our IGD measure (i.e., we did not empirically examine the relation between our IGD measures and similar measures). Including multiple measures for IGD, game addiction, or problematic gaming in a survey would allow for comparison of the psychometric properties of these instruments and their underlying criteria. Some of the items selected for the short scale covered a specific aspect of a criterion, as suggested by Petry & O'Brien (2013), thereby excluding other aspects. Although the specific items selected for the nine-item scale may have adequately reflected the whole sample of 13–40-year-old men and women, the most suitable item from each criterion may differ between genders and age groups. Future studies may indicate whether certain aspects are more suitable for diagnosing IGD among specific demographic groups. Finally, the two assessments for diagnosing disordered gamers should have indicated a significant difference in life satisfaction between disordered gamers and nondisordered gamers (Ko et al., 2005; Lemmens et al., 2009), whereas we found no significant differences. The relation between IGD and life satisfaction needs further examination to determine whether this inconsistency was the result of our assessment of IGD or the negative relation with life satisfaction is simply not as stable as expected.

In conclusion, this study is among the first to begin empirical examination of the new definition of IGD from the *DSM–5*. We investigated its underlying criteria and how this disorder can be measured reliably and validly. The short dichotomous version of the scale provided a valid and reliable measure of IGD with good diagnostic accuracy that can be used for research and diagnostic purposes among male and female gamers of all ages. Having developed a survey instrument that is suitable for all gamers, we hope that this psychometrically robust scale will facilitate research that may answer important questions about the time course of IGD, its causes and consequences among different age groups, and the outcomes with or without treatment. Regardless of the terminology that is applied to this disorder, disordered involvement with video games has become a serious health concern for a sizable group of players around the world, and this phenomenon deserves further scientific attention. This study thereby represents a major step forward in the measurement, validation, and general understanding of IGD.

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Appendix

Twenty-Seven Items for the Internet Gaming Disorder Scale

Preoccupation

During the last year . . .
. . . have there been periods when you were constantly thinking about a game while at school or work?
. . . have there been periods when all you could think of was the moment that you could play a game?
. . . have there been periods when you were constantly fretting about a game?

Tolerance

During the last year . . .
. . . have you felt the need to continue playing for longer periods of time?
. . . have you felt the need to play more often?
. . . have you felt unsatisfied because you wanted to play more?

Withdrawal

During the last year . . .
. . . have you been feeling tense or restless when you were unable to play games?
. . . have you been feeling angry or frustrated when you were unable to play games?
. . . have you been feeling miserable when you were unable to play a game?

Persistence

During the last year . . .
. . . did you want to play less, but couldn't?
. . . did you try to play less, but couldn't?
. . . were you unable to reduce your time playing games, after others had repeatedly told you to play less?

Escape

During the last year . . .
. . . have you played games to forget about your problems?
. . . have you played games so that you would not have to think about annoying things?
. . . have you played games to escape negative feelings?

Problems

During the last year . . .
. . . have you skipped work or school so that you could play games?
. . . have you played throughout the night, or almost the whole night?
. . . have you had arguments with others about the consequences of your gaming behavior?

Deception

During the last year . . .
. . . have you lied to your parents or partner about the time you spent playing games?
. . . have you hidden the time you spend on games from others?

(*Appendix continues*)

Appendix (*continued*)

. . . have you played games secretively?

Displacement

During the last year . . .
. . . have you been spending less time with friends, partner or family in order to play games?
. . . have you lost interest in hobbies or other activities because gaming is all you wanted to do?
. . . have you neglected other activities (e.g., hanging out with friends, hobbies or sports) so that you could play games?

Conflict

During the last year . . .
. . . have you experienced serious problems at work or school because of gaming?
. . . have you experienced serious conflicts with family, friends or partner because of gaming?
. . . have you lost or jeopardized an important friendship or relationship because of gaming?

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