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Original article

French validation of the 7-item Game Addiction Scale for adolescents

Validation française de la Game Addiction Scale à 7-items pour adolescents

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ABSTRACT

Introduction. – The Game Addiction Scale (GAS: Lemmens, Valkenburg, & Peter, 2008, 2009) is a short instrument (7-item) for evaluating video game playing by adolescents.
Objective. – The aim of the current research was to investigate the psychometric properties of a French version of the 7-item Game Addiction Scale for adolescents.
Method. – Two studies were conducted with two samples of French adolescents between the ages of 10 and 18 (study 1: $n = 159$; study 2: $n = 306$). First, we examined the factor structure and internal consistency. Second, we added a concurrent validity analysis with estimation of the daily time spent playing video games and an assessment of depression and anxiety.
Results. – In both studies, the factor analysis revealed a one-factor structure that had good psychometric properties and fit the data well. The analysis also confirmed good internal consistency of the scale. Correlation analysis in the second study showed that the GAS score had significant positive relationships with the time spent playing video games, depression, anxiety, and the fact of being a boy, thereby supporting the concurrent validity of the scale.
Conclusion. – This French version of the GAS seems to be a reliable tool for identifying and assessing problematic use of video games.

Mots clés :
Game Addiction Scale
Propriétés psychométriques
Adolescents

RÉSUMÉ

Introduction. – La Game Addiction Scale (GAS : Lemmens, Valkenburg, & Peter, 2008, 2009) est un outil de mesure court (7 items) permettant d'évaluer spéciquement la pratique vidéo-ludique des adolescents, quel que soit le format de jeu vidéo utilisé.
Objectif. – L'objectif de ce travail est d'analyser les propriétés psychométriques de cette version française de la GAS.
Méthode. – Deux études ont été menées auprès de deux échantillons d'adolescents français, âgés de 10 à 18 ans (étude 1 : $n = 159$; étude 2 : $n = 306$). Les analyses de structure factorielle et de la consistance interne ont tout d'abord été menées dans la première étude. L'analyse de la validité concourante a ensuite été ajoutée dans la seconde étude avec l'estimation du temps de jeu quotidien, la mesure de la dépression et de l'anxiété.
Résultats. – Au sein des deux études, les analyses factorielles ont mis en avant une structure unidimensionnelle qui dispose de bonnes propriétés psychométriques et qui s'ajuste aux données. Les analyses ont également confirmée la consistance interne de l'outil. De plus, les analyses de corrélation effectuées dans la deuxième étude ont mis en avant des liens positifs significatifs entre le score de la GAS et le temps de jeu quotidien, le score de dépression, celui d'anxiété et le fait d'être un garçon. Ces éléments confirment la validité concourante de l'échelle.

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Conclusion. – Cette version française de la GAS 7-items pour adolescents dispose de bonnes propriétés psychométriques. Il semble qu'il s'agisse d'un outil fiable permettant de repérer et d'appréhender un usage problématique des jeux vidéo.

1. Introduction

Video game users represent approximately 40% of the French population (National Syndicate of Video Games, 2012). Adolescents are the primary users of video games (Fortin, Mora, & Trémel, 2005): 91% of 6- to 14-year-old and 95% of 15- to 24-year-old are video game players (French Videogame Agency, 2010). Yet excessive playing of video games has become a problem in modern society and is manifesting itself in treatment centers for adolescents.

Video game addiction can be defined as "excessive and compulsive use of video games that results in social and/or emotional problems; despite these problems, the gamer is unable to control this excessive use" (Lemmens, Valkenburg, & Peter, 2009, p. 78). The definition of the term "addiction" is still under debate, and many authors prefer to speak of "excessive" or "problematic" video game playing. Nonetheless, some studies have drawn a parallel between pathological gambling and playing behavior (Griffiths & Davies, 2005; Parker, Summerfeldt, Taylor, Kloosterman, & Keefer, 2013; Parker, Taylor, Eastabrook, Schell, & Wood, 2008). Some criteria used to describe video game addiction, such as cognitive salience, tolerance, and euphoria, are related to a high level of involvement and can be regarded as "peripheral" criteria (Wood, 2008). When these criteria indicate a level of involvement that has consequences on daily life, then the behavior can be seen as non-normal. To define addictive behavior, other criteria — the "core" of addiction (Wood, 2008) — must be present: conflict, withdrawal symptoms, relapse, and salience. Van Rooij, Schoenmakers, Vermulst, Van den Eijnden, and Van de Mheen (2010) made the distinction between addicted and non-addicted "heavy gamers" on the basis of the psychological, social, and/or physical implications of this involvement on daily life (King, Haagsma, Delfabbro, Gradisar, & Griffiths, 2013; Tejeiro Salguero & Bersabé Moran, 2002). The loss of control despite negative consequences is a sign of addictive use (Lemmens et al., 2009), which can be accounted for in terms of "dysfunctional preoccupations" (Parker et al., 2008, 2013).

Griffiths' descriptive approach to video game addiction (Griffiths & Davies, 2005) is based on Brown's clinical criteria of pathological gambling (Brown, 1993). Tejeiro Salguero and Bersabé Moran (2002) adapted these criteria to adolescent video game playing as follows:

- increased time spent playing or thinking about video games, scheduling next game, or remembering previous games;
- bad mood or irritability when it's impossible to play;
- increased time spent playing during difficult times;
- failed attempts to control game playing time;
- concealing the time spent playing from parents or friends;
- failing to do homework, or lying about it, in order to play video games;
- sleeping in, missed meals, and less time spent with family or friends in order to play more.

Lemmens, Valkenburg, and Peter (2008) and Lemmens et al. (2009) relied on these criteria to develop the 21-item and 7-item Game Addiction Scales [GAS], which are specifically aimed at assessing all kinds of video game playing and are well-suited to adolescents. Responses are scored on a five-point Likert scale, which affords greater sensitivity than dichotomous tools. The authors (Lemmens et al., 2009) conducted two surveys with two independent samples of adolescents (sample 1: $n = 352$, 67% boys; sample 2: $n = 369$, 68% boys). In both samples, the age varied between 12 and 18 years (sample 1: M = 14.80, SD = 1.64; sample 2: M = 15.20, SD = 1.35). Confirmatory factor analysis (CFA) of the 21-item GAS was performed and revealed a good structural fit (sample 1: $\chi^2(182) = 594.40$, $p < .001$; CFI = 0.90; RMSEA = 0.08 with a 90%CI [0.07, 0.09]). To determine whether their model held for the second sample, they performed a multigroup analysis. The model exhibited an adequate fit and its structure was very similar in the two samples. Cronbach's alpha indicated good internal consistency in each sample ($\alpha = 0.94$ and $\alpha = 0.92$). Exploratory factor analysis (EFA) was performed and revealed a single-factor structure that accounted for 47.40% of the variance. Loadings of the 21 items for the combined sample were computed (factor loadings varied between 0.48 and 0.85). The items with the highest overall loadings on each of the seven factors were selected in order to determine what items to use for the 7-item GAS. Multigroup analysis was performed. The model exhibited an adequate fit (unconstrained model for both samples: $\chi^2(28) = 69.90$, $p < .001$; CFI = 0.97; RMSEA = 0.05 with a 90%CI [0.03, 0.06]) and its structure was very similar in the two samples. The 7-item GAS showed good internal consistency in each sample ($\alpha = 0.86$ and $\alpha = 0.81$). Finally, the significant link between the GAS score and the time spent on games, loneliness, life satisfaction, aggression, and social competence indicated the concurrent validity of both versions of the GAS (7- and 21-item).

Even though the GAS was not validated clinically, the authors suggested using a cutoff point in line with the polythetic format applied in DSM-4 (American Psychiatric Association, 1996), that is, at least half of the criteria indicate that the subject's video game playing as problematic. In addition, the use of short, fast scales has proven necessary with adolescents (Maïano, Ninot, Morin, & Bilard, 2007) to prevent them from giving up or responding randomly (Coste, Guillemin, Pouchot, & Fermanian, 1997). Adolescence is known to be the period of immediacy, action, and impulsiveness. Moreover, attentional abilities are still under development, which causes adolescents to be easily distracted (Lacknera, Santessoa, Dywana, & Segalowitza, 2013; Monk et al., 2003; Steinberg, Dahl, Keating, Masten, & Pine, 2006). Thus, the use of a short scale seems appropriate. Using short scales could not only reduce the social desirability bias to which adolescents are particularly prone, but also help pinpoint the problematic use of video games by reducing the risk of drowning this problem in the adolescent process. Short forms are also necessary in the current clinical and scientific context, where "quick, preliminary evaluations" are needed in healthcare centers, and multivariate studies involving a battery of questionnaires must be conducted. Furthermore, in the field of addiction, both researchers and clinicians prefer short scales (Babor, Biddle-Higgins, Saunders, & Monteiro, 2001; Décamps, Battaglia, & Idier, 2010; Etter, Le Houezecn, & Perneger, 2003; Ewing, 1984; Heatherton, Kozlowski, Freckern, & Fagerstrom, 1991; O'Loughlin, Tarasuk, Difranzan, & Paradis, 2002). Moreover, there are very few validated scales in French for assessing the use of video games among adolescents (Romo, Bioulac, Kern, & Michel, 2012). To the best of our knowledge, there is only one exploratory study of the French version of Problem Video game Playing (Tejeiro

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Salguero & Bersabé Moran, 2002), which was conducted with ADHD children (Bioulac, Arfi, Michel, & Bouvard, 2010). It is important for French researchers and clinicians to have a reliable assessment tool for identifying and measuring the problematic use of video games among adolescents. All of these considerations prompted us to use the 7-item GAS.

2. The present research

The aim of the present research was to explore the psychometric properties of a French translation of the 7-item GAS for adolescents (Lemmens et al., 2008, 2009). We used a translation/back-translation procedure. Two researchers did the translation. They were clinical psychologists, so they had the scientific knowledge and clinical skills likely to be useful in producing a good translation. Each researcher translated the scale independently. Then the two French versions were back-translated into the original language by outside individuals who were unaware of the original version. We compared these versions in order to reach a consensus on the final version (the version we set out to validate in the present article). This process of scale validation is necessary for making comparisons between studies conducted in different languages, for developing scientific research among French speakers, and, with a clinical aim, for having a reliable tool to capture problematic use (Vallerand, 1989). A validated French version of the GAS specific to adolescents would therefore have scientific as well as clinical utility, and would allow for more epidemiological and empirical research on the use of video game playing by adolescents.

We conducted two studies with two different samples of French adolescents. The factor structure and internal consistency of the scale were evaluated. Concurrent validity was assessed by looking for significant links with several factors. The first was the time spent playing video games, considered as an initial but insufficient indicator of problematic playing (Parker et al., 2008, 2013; Schmit, Chauchard, Chabrol, & Sejourne, 2011; Tejeiro Salguero & Bersabé Moran, 2002). Indeed, even though the time spent playing is not a criterion of addiction, problematic gamers spend more time on video games than non-problematic gamers do (Parker et al., 2008, 2013; Schmit et al., 2011). In addition, the time spent on video games increases with involvement, which can lead to addiction (Van Rooij et al., 2010). The second and third factors assessed were depression and anxiety, both of which are associated with problematic use (Bonnaire & Varescon, 2009; Mehroof & Griffiths, 2010; Park et al., 2013; Schmit et al., 2011; Yen, Ko, Yen, Wu, & Yang, 2007). Cyberaddiction and problematic video game playing are accompanied by depressive affect and a high level of anxiety (Bonnaire & Varescon, 2009; Yen et al., 2007). Compared to non-addicted individuals, persons addicted to video games have a higher level of depression (Schmit et al., 2011) and anxiety (Mehroof & Griffiths, 2010). We therefore expected to find that the GAS score was positively related to the time spent on video games, depression, and anxiety. These relationships would then be considered evidence of concurrent validity. We also analyzed the gender distribution, even though gender differences in video game use/abuse are under debate. Despite current statistics indicating that as many girls as boys play video games, studies have shown that the majority of "non-players" are girls (Achab et al., 2011). Girls are generally casual players (TNS-Sofres, 2010), while regular and excessive players are mostly boys (Fortin et al., 2005; Griffiths & Hunt, 1995; French Videogame Agency, 2010). Finally, we conducted a third study that combined the samples from studies 1 and 2, in order to confirm that the factor structure was the same in each subgroup (i.e., boys and girls).

3. Study 1: factor analysis and reliability

The aim of the first study was to examine the factor structure and internal consistency of the 7-item GAS.

3.1. Method

3.1.1. Participants

The sample consisted of 159 French adolescents (52% boys). The participants' mean age was 14 years (range 10–18, M = 14.01, SD = 2, Md = 14.00).

3.1.2. Materials and procedure

The French translation of the 7-item GAS was used. All items were scored on a 5-point Likert scale ranging from 1 ("never") to 5 ("very often"). Four "validated" items — a response greater than or equal to 3 ("sometimes") validated the item — corresponded to problematic use of video games (Lemmens et al., 2008, 2009).

Following a briefing, a consent form was given to parents. The adolescents were also asked to give their written consent. The experiment took place in a classroom during school hours. A researcher was present during testing to answer questions and ensure confidentiality.

Data analysis was performed using the Statistical Package for the Social Sciences (SPSS 16.0, SPSS Inc., Chicago) and the Analysis of Moment Structures (AMOS 6, AMOS Development Corporation). An exploratory factor analysis (EFA) was performed using the Varimax rotation method, in order to determine the dimensionality of the questionnaire. Confirmatory factor analyses (CFA) were also computed using the maximum likelihood estimator. We calculated the Chi^2 value (χ^2), which should be nonsignificant in order to consider the fit as acceptable (Hu & Bentler, 1999). The following fit indexes were used to evaluate the model's goodness of fit. The Comparative Fit Index (CFI) and the Tucker-Lewis Index (TLI) indicate a very good fit if they are close to 1, and an acceptable fit if they are greater than 0.90. The Root Mean Square Error of Approximation (RMSEA) with a 90% confidence interval (90%CI) and the Standardized Root Mean Square Residual (SRMR) indicate an acceptable fit if less than 0.08 and a good fit if less than 0.05 (Hu & Bentler, 1999; Schermelleh-Engel, 2003). Cronbach's alpha was used to measure internal consistency as an indicator of scale reliability; it should be greater than 0.70.

3.2. Results

The EFA revealed a factor loading for every item that was greater than 0.50 (Table 1), and a single-factor structure that accounted for 46.50% of the variance. The eigenvalue of the first factor was above 1, while the others were below 1 (Table 2). So the scale was considered one-dimensional (Carmines & Zeller, 1979). In the CFA, we computed a one-factor model in which the seven items on the scale were hypothesized to be a unique latent factor representing video game addiction (Table 3). The χ^2 statistic of the one-factor model was nonsignificant, $\chi^2(14) = 21.66$, $p = 0.09$, which corresponds to an acceptable fit. The combination of the four fit indexes indicated an acceptable fit (CFI = 0.97, TLI = 0.96, RMSEA = 0.06 with a 90%CI [0.0, 0.11], and SRMR = 0.04). Cronbach's alpha was 0.80, which indicates good internal consistency of the scale.

Based on the GAS scores of the 159 adolescents, 20 were problematic video game players (75% boys) and 139 were non-problematic video game players (48% boys). In addition, 18% of the boys and 7% of the girls were problematic gamers ($\chi^2 = 4.90$, $p < .05$). There was no significant correlation between age and problematic videogame playing, nor a significant difference

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Table 1
Means, standard deviations and factor loadings.

How often during the last six month...	Mean (SD)			Factor loading		
	Study 1	Study 2	Original	Study 1	Study 2	Original
Item 1: did you feel addicted to a video game? *(T'es-tu senti addicté à un jeu vidéo ?)*	1.88 (1.05)	2.48 (1.24)	1.59 (0.95)	.79	.78	.70
Item 2: did you spend more and more time on video games? *(As-tu augmenté ton temps passé à jouer aux jeux vidéo ?)*	1.55 (0.86)	1.84 (1.04)	1.64 (0.95)	.51	.55	.79
Item 3: did you play video games to escape reality? *(As-tu joué aux jeux vidéo pour fuir/oublier la réalité ?)*	1.47 (0.94)	1.90 (1.20)	1.52 (0.86)	.55	.51	.76
Item 4: others have unsuccessfully tried to reduce your playing time? *(D'autres ont-ils essayé sans succès de réduire ton temps de jeu ?)*	1.75 (1.07)	2.27 (1.28)	1.66 (1.07)	.65	.61	.61
Item 5: did you feel bad when unable to play? *(T'es-tu senti mal quand tu ne pouvais pas jouer ?)*	1.45 (0.83)	1.87 (1.10)	1.32 (0.72)	.80	.79	.85
Item 6: did you have hard arguments with others on your time spent on games? *(As-tu eu des conflits avec les autres au sujet de ton temps passé à jouer aux jeux vidéo ?)*	1.81 (1.03)	2.24 (1.28)	1.29 (0.65)	.64	.70	.74
Item 7: your playing time caused sleep deprivation? *(Ton temps de jeu a-t-il causé des manques de sommeil ?)*	1.74 (1.02)	2.19 (1.18)	1.50 (0.91)	.78	.72	.67

No missing data.

Table 2
Eigenvalues.

	Eigenvalues	
	Study 1	Study 2
Factor 1	3.3	3.2
Factor 2	0.9	0.9
Factor 3	0.8	0.7
Factor 4	0.7	0.6
Factor 5	0.5	0.5
Factor 6	0.4	0.5
Factor 7	0.3	0.4

No missing data.

between the average ages of problematic gamers (M = 13.49, SD = 2.5) and non-problematic gamers (M = 14.18, SD = 2) (t = −1.38, p > .05). We divided the sample into two age groups at the median (Md = 14, so [10–13] and [14–18]); no significant difference between the two groups was found concerning the problematic use of video games (t = 1.47, p > .05).

The results indicated that the one-factor model of the 7-item GAS had good psychometric properties and fit the data well. These results are similar to those obtained by Lemmens et al. (2009) in their original study. However, the number of participants was relatively small, barely above the generally accepted value of at least 10 subjects per item needed to validate a scale (Tabachnik & Fidell, 2007). Our sample size may therefore limit the generality of our results regarding the prevalence of addiction.

4. Study 2: factor analysis, reliability, and concurrent validity

The aim of the second study was to confirm the factor structure and internal consistency of the 7-item GAS, and to examine its concurrent validity with a larger sample of adolescents.

4.1. Method

4.1.1. Participants

The sample consisted of 306 French adolescents (65% boys). The participants' mean age was 14 years 9 months (range 10–18 years, M = 14.75, SD = 2.4, Md = 14.40).

4.1.2. Materials and procedure

The French translation of the 7-item GAS was used. The French version (Moor & Mack, 1982) of the Child Depression Inventory or CDI (Kovacs & Beck, 1977) was used to assess depression (α = 0.86). The CDI contains 27 items, each consisting of three statements. For each item, the adolescent is asked to select the statement

Table 3
Confirmatory factor analysis: fit indexes.

Model	χ^2	CFI	TLI	RMSEA	SRMR
Study 1	21.66 p = 0.086	.97	.96	.06 (90%CI: .0; .11)	.04
Study 2	41.88 p < .05	.95	.92	.08 (90%CI: .05; .11)	.05
Study 3					
Total sample	32.74 p < .05	.98	.97	.05 (90%CI: .03; .08)	.04
Boys	32.30 p < .01	.96	.94	.06 (90%CI: .04; .09)	.04
Girls	17.13 p = .25	.99	.98	.04 (90%CI: .0; .08)	.04
Under 14 years	23.02 p = .06	.98	.97	.06 (90%CI: .0; .09)	.04
14 or older	24.57 p < .01	.98	.96	.06 (90%CI: .01; .09)	.04
Original	69.90 p < .01	.97	/	.05 (90%CI: .03; .06)	/

No missing data. CFI: Comparative Fit Index; TLI: Tucker-Lewis Index; RMSEA: Standardized Root Mean Square Residual; SRMR: Standardized Root Mean Square Residual.

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that best describes his/her feelings. The French version (Turgeon & Chartrand, 2003) of the Revised Children's Manifest Anxiety Scale or RCMAS (Reynolds & Richmond, 1979) was used to assess anxiety. The RCMAS consists of 37 yes/no items that rate the level of anxiety (total anxiety score), and the nature of the anxiety on three dimensions: physiological anxiety, worry, and social anxiety (α = 0.78, 0.57, 0.82, and 0.70, respectively). Another dimension of this scale corresponds to defensiveness. It assesses the tendency to deny common negative behaviors. High scores on this dimension reflect ideal behavior (which is not characteristic of anyone), and may be indicative of social desirability (desirability scale, α = 0.73). The participants were asked to assess the time spent daily on video games during weekdays, weekends, and holidays. They were also asked about their weekly frequency of video game playing (less than once per week, 1–3 times per week, more than three times a week, and every day).

Participants were recruited through advertising in their schools and on social network sites. An information leaflet for adolescents and their parents containing the link to the protocol was circulated. Measures were performed with an online questionnaire containing a consent form (Google documents: https://www.docs.google.com).

Data analysis was performed using SPSS 16.0 (SPSS Inc., Chicago) and AMOS 6 (AMOS Development Corporation). Exploratory and confirmatory factor analyses were computed (using Varimax rotation method for the EFA). For the CFA, we used the maximum likelihood estimator and computed five fit indexes (χ^2, CFI, TLI, RMSEA with a 90%CI, and SRMR). Internal consistency (α) was used to assess reliability. Correlation analysis (Bravais-Pearson's correlation coefficient r) was performed to determinate concurrent validity. We looked at how the GAS score was related to the time spent on video games, depression, anxiety, and gender (boys were labeled +1 and girls −1). A strong correlation between the GAS and these variables was considered as evidence of its concurrent validity. We also analyzed the distribution of subjects by gender (χ^2) to check for consistency with the literature.

4.2. Results and discussion

The EFA revealed a factor loading greater than 0.50 on every item (Table 1) and a single-factor structure that accounted for 44.90% of the variance. The eigenvalue of the first factor was above 1, while the others were below 1 (Table 2). So the scale was considered one-dimensional (Carmines & Zeller, 1979). In the CFA, we computed a one-factor model (Table 3). The Chi^2 value was significant. However, it is known that χ^2 increases with sample size and that it is unusual to obtain a nonsignificant χ^2 when performing CFA on self-report questionnaires (Byrne, 2010). Moreover, the combination of the four indexes indicated an acceptable fit (CFI = 0.95, TLI = 0.92, RMSEA = 0.08 with a 90%CI [0.05, 0.11], and SRMR = 0.05). Cronbach's alpha was 0.79, which confirms the good internal reliability of the scale.

Based on the GAS scores of the 306 adolescents, 86 were problematic video game players (82.50% boys) and 220 were non-problematic video game players (58.50% boys). In addition, 35.50% of the boys and 15% of the girls were problematic gamers (χ^2 = 15.63, p < .05). As in Study 1, there was no significant correlation between age and problematic video game playing and no significant difference between the average age of problematic gamers (M = 14.94, SD = 2.3) and non-problematic ones (M = 14.67, SD = 2.4) (t = 0.90, p > .05). No significant difference between the two age groups in the problematic use of video games was found (under 14 and 14 or over; t = 0.58, p > .05).

During weekdays, adolescents classified as problematic gamers did not play more than the others (M = 2.6 hours, SD = 2.1, and M = 2.2 hours, SD = 2, respectively, p > .05). However, 53.40% of the problematic gamers played every day vs. 21.70% for non-problematic ones, and 19.80% played less than once a week vs. 43.80% for non-problematic gamers (χ^2 = 33.49, p < .05). During weekends and holidays, problematic gamers played more (M = 5.4, SD = 3.1, and M = 5.9, SD = 3.3, respectively) than did non-problematic ones (M = 3.8, SD = 3.1, and M = 4.3, SD = 3.5, respectively) (t = 3.43, p < .05, and t = 3.18, p < .05).

Table 4
Concurrent validity: correlation analysis.

	GAS
Time spent daily on video games during	
Weekdays	.22**
Weekends	.40**
Holidays	.28**
Depression	.31**
Anxiety	
Total	.15*
Physiological	.07
Worry	.12
Social-anxiety	.20**
Gender (boy = +1; girl = −1)	.30**

No missing data. *p < .05; **p < .01. GAS: Game Addiction Scale.

Correlation analyses (Table 4) showed that the GAS score was significantly and positively correlated with the time spent on video games daily, on weekdays, on weekends, and on holidays, and also with depression, anxiety (especially social anxiety), and the fact of being a boy.

This second study was conducted to confirm the results of the first with a larger sample, and to assess the concurrent validity of the 7-item GAS. The factor analysis indicated a one-factor model with good psychometric properties that fit the data well. Internal reliability was good. These elements confirm the construct validity of this scale and are consistent with those obtained by Lemmens et al. (2009) in the original study.

Furthermore, in line with the literature (Parker et al., 2008, 2013; Schmit et al., 2011), there was a significant positive correlation between the GAS score and time spent on video games, and adolescents classified as problematic gamers played more frequently and longer than did non-problematic gamers. Our results confirmed that problematic use of video games is more prevalent among boys. The general population, as well as the players themselves, consider video games as a male activity (Fox & Tang, 2014). They are played more often by male adolescents (Bioulac & Michel, 2012) and the majority of regular and problematic players are boys (French Videogame Agency, 2010; Fortin et al., 2005; Griffiths & Hunt, 1995). In addition, we found significant positive correlations between the GAS score and depression, anxiety, and especially social anxiety. These results are consistent with the literature, insofar as depression and anxiety are considered comorbidities of cyberaddiction (Bonnaire & Varescon, 2009) and Internet addiction (Yen et al., 2007). In the case of online video game addiction, adolescents and young adults have been shown to have a depressive symptomatology (Schmit et al., 2011). Finally, a significant relationship between anxiety (trait and state) and online video game addiction was found, suggesting that anxiety promotes excessive online playing (Mehroof & Griffiths, 2010). Thus, the concurrent validity of this French version of the 7-item GAS seems satisfactory.

5. Study 3: multigroup factor analysis

The aim of the third study was to confirm the invariant factorial structure of the 7-item GAS in each of four adolescent subgroups: boys, girls, under 14 years of age, and 14 or older. We tested the equality of the factor loadings for each group, in order to determine whether the same model was applicable across groups.

Table 5
Multigroup confirmatory factor analysis: fit indexes.

	χ^2	df	CFI	RMSEA	Base	CMIN	Δ - CFI	Δ - RMSEA
Boys/girls								
Model a: unconstrained - Configural invariance	34.49	28	0.949	0.02	–	–	–	–
Model b: measurement weights - Metric invariance	46.44	34	0.931	0.03	ma	11.96	0.018	0.01
Model c: measurement intercepts - Strong invariance	97.27	41	0.627	0.06	mb	50.83*	0.322	0.03
Model c2: partial strong invariance (i1, i4, i6 & i7 free)	48.78	36	0.915	0.03	mb	2.34	0.016	0.00
Model d: structural means (i1, i4, i6 & i7 free)	51.59	37	0.903	0.03	mc2	2.81	0.012	0.00
Model e: structural covariance	52.85	38	0.902	0.03	md	1.26	0.001	0.00
Model f: measurement residuals - Partial strict invariance (e1, e4, e6 & e7 free)	64.75	41	0.843	0.04	me	11.90*	0.059	0.01
Model f2: partial strict invariance (e1, e4, e5, e6 & e7 free)	53.40	40	0.911	0.03	me	0.55	0.009	0.00
Under 14 years/14 years or older								
Model a: unconstrained Configural invariance	47.57	28	0.978	0.03	–	–	–	–
Model b: measurement weights metric invariance	51.35	34	0.980	0.03	ma	3.78	0.002	0.00
Model c: measurement intercepts - Strong invariance	63.38	40	0.973	0.03	mb	12.07	0.007	0.00
Model d: structural means	63.43	41	0.974	0.03	mc	0.05	0.001	0.00
Model e: structural covariance	64.69	42	0.974	0.03	md	1.26	0.000	0.00
Model f: measurement residuals - Strict invariance	73.23	49	0.972	0.03	me	8.54	0.002	0.00

No missing data. *$p < .05$. CFI: Comparative Fit Index; RMSEA: Standardized Root Mean Square Residual.

5.1. Method

We combined the samples from studies 1 and 2 ($n = 465$; 60% boys and 40% girls; 47% under age 14, and 53% 14 or older). First, we performed a confirmatory factor analysis (CFA) for the total sample and for each subgroup. Before testing measurement invariance across groups, the properties of the model were assessed using the whole sample. We calculated five fit indexes: χ^2, CFI, TLI, RMSEA, and SRMR. If the model appeared to fit the data reasonable well for the individual sub samples, assessment of invariance between the groups continued by sequentially testing a series of progressively restrictive and nested models in the following order:

- model a: unconstrained – configural invariance;
- model b: measurement weights (equal factor loadings) – metric invariance;
- model c: measurement intercepts (equal indicator intercepts) – strong invariance;
- model d: structural means (equal latent means);
- model e: structural covariance (equal factor co-variances);
- model f: measurement residuals (equal factor variances) – strict invariance.

Our rationale for the model testing sequence was that the models a, b and c specifically assess configural, metric and scalar measurement invariance respectively. The models d, e and f compare subgroup distributional properties of the constructs. We calculated the Chi^2 and the CFI and RMSEA difference statistics measuring the difference between the original and constrained models. A non-significant Chi^2 difference (CMIN, $p > .05$) indicates that the model exhibits measurement invariance across groups (Byrne, 2010). A delta-CFI value less than or equal to 0.01 indicates that the null hypothesis of invariance should not be rejected (Cheung & Rensvold, 2002); a value between 0.01 and 0.02 means that differences may exist; a value above 0.02 indicates the non-invariance of the model (Vandenberg & Lance, 2000). A very small delta-RMSEA represents substantively insignificant variations in the model's fit (Marsh, Ellis, Parada, Richards, & Heubeck, 2005). Data analysis was performed using AMOS 6 software (AMOS Development Corporation).

5.2. Results and discussion

CFA was performed for the whole sample and for each subgroup. We computed a one-factor model (Table 3). For the whole sample, for boys, and for adolescents under 14 years of age, the Chi^2 value was significant. However, it is known that χ^2 increases with sample size, and that it is unusual to obtain a nonsignificant χ^2 when performing CFA on self-report questionnaires (Byrne, 2010). The combination of the four indexes indicated a good fit for the total sample and for each subgroup (Table 3).

Multigroup CFA was performed on the gender and age groups (Table 5).

Gender invariance: The two first models fit suggested acceptable evidences for configural and metric invariances (Model a: $\chi^2(28) = 34.49$, $p > .05$; CFI = 0.95, RMSEA = 0.02; Model b: CMIN = 11.96, $p > .05$; Δ-CFI < 0.02, Δ-RMSEA = 0.01). The third model (Model c: Measurement intercepts) testing strong factorial invariance was rejected (CMIN = 50.83, $p < .05$; Δ-CFI > 0.02, Δ-RMSEA = 0.03). Modification indices suggested that the cross-group equality constraint on the intercept for items 1, 4, 6 and 7 contributed most strongly to the lack of fit. The fourth model freely estimated these parameters in both groups and provided evidence of partial strong factorial invariance (Gregorich, 2006) (model c2: CMIN = 2.34, $p > .05$; Δ-CFI < 0.02, Δ-RMSEA = 0.00). The fifth and sixth models fit suggested acceptable evidences for structural means and structural covariance invariances (model d: CMIN = 2.81, $p > .05$; Δ-CFI < 0.02, Δ-RMSEA = 0.00; and model e: CMIN = 1.26, $p > .05$; Δ-CFI < 0.01, Δ-RMSEA = 0.00). The seventh model tested partial strict factorial invariance by imposing additional cross-group equality constraints on all corresponding item residual variances except those for items 1, 4, 6 and 7. This model was rejected (CMIN = 11.90, $p < .05$; Δ-CFI > 0.02, Δ-RMSEA = 0.01). The modification indices suggested freely estimating the residual variance for item 5, resulting in a well-fitting partial strict factorial invariance model (Gregorich, 2006) (model f2: CMIN = 0.55 $p > .05$; Δ-CFI < 0.01, Δ-RMSEA = 0.00). These results allow us to consider the partial factorial invariance across gender, which implies to avoid or restrict group comparison to those items with appropriate invariant parameters (Gregorich, 2006).

Age invariance (under 14 vs. 14 or older): the six models fit suggested acceptable evidences for configural, metric, strong, strict and structural invariances (Table 5: model a: $\chi^2(28) = 47.57$, $p > .05$; CFI = 0.98, RMSEA = 0.03; model b: CMIN = 3.78, $p > .05$; Δ-CFI < 0.01, Δ-RMSEA = 0.00; model c: CMIN = 12.07, $p > .05$; Δ-CFI < 0.01, Δ-RMSEA = 0.00; model d: CMIN = 0.05, $p > .05$; Δ-CFI < 0.01, Δ-RMSEA = 0.00; model e: CMIN = 1.26, $p > .05$; Δ-CFI < 0.01, Δ-RMSEA = 0.00; model f: CMIN = 8.54, $p > .05$; Δ-CFI < 0.01, Δ-RMSEA = 0.00). These results allow us to consider the factorial invariance across age groups.

Thus, the hypothesis of the inconsistency of the model between boys and girls, on the one hand, and between adolescents under

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14 and 14 or older, on the other, can be rejected. Study's 3 results confirm that the factor structure of the scale did not vary across subgroups (boys vs. girls, under 14 vs. 14 or older), thereby confirming the validity of the results obtained in studies 1 and 2.

6. General discussion

Over the past decade, excessive/problematic use of video games has received much empirical and theoretical attention. There is increasing interest in this issue, particularly for adolescents. However, much more research is needed to better detect and understand this problem, and to allow for the development of targeted treatment programs (Collins, Freeman, & Chamarro-Premuzic, 2012). The GAS was introduced as a self-report measure of the problematic use of video games (Lemmens et al., 2008, 2009). The main aim of the current research was to examine the psychometric properties of the 7-item GAS in two French samples.

Due to the absence of a consensus on the concept of addiction to video games (Wood, 2008), and according to recent studies, the prevalence of problematic use varies between 3% and 12% (Bonnaire & Varescon, 2009; Gentile, 2009). In the present studies where we applied the polythetic format, 12.60% (study 1) and 28.10% (study 2) of the adolescents were classified as problematic video game players. With a monothetic format (validation of all criteria) only 0.6 and 1% of adolescents were classified as problematic video game players, which could be an underestimation (Van Rooij et al., 2010). However, we noted two items that had very low factor loadings in both samples: Item 2 ("How often during the last six month did you spend more and more time on video games?"), a measure of tolerance (.51 and .55, in studies 1 and 2, respectively) and Item 3 ("How often during the last six month did you play video games to escape reality?"), a measure of mood changes (.55 and .51, respectively). It should be noted that in the original study (Lemmens et al., 2008, 2009) the samples were composed only of adolescents who had played video games within the last month. In our studies, the samples were not selective for video game playing. This could explain differences in the factor loadings. In video game playing, these two items correspond to "peripheral criteria" for addiction (Wood, 2008). As such, they are related to a high level of involvement and not necessarily to the problematic use of video games. Due to the interactive and immersion qualities of any video game, it is not necessary to spend a lot of time playing to reach the desired state; simply playing might be enough. So in this context, one cannot assume that playing longer is a question of tolerance. We *can* assume, however, that these items correspond to a high level of involvement, but are not enough to consider the behavior problematic.

Finally, these considerations may raise the question of the cutoff points to apply on the 7-item GAS. The polythetic format applied by DSM-4 could be supplemented by a recommendation, such as having to validate at least two or three items referring to the core criteria of addiction, out of the required four.

To conclude, this French version of the 7-item GAS exhibits acceptable psychometric properties that confirm its theoretical and concurrent validity. Although a great many studies are needed to assess the clinical validity of any scale, the French version of the 7-item GAS seems to be a reliable assessment tool for identifying and capturing the problematic use of video games and should be useful in future research.

Disclosure of interest

The authors declare that they have no conflicts of interest concerning this article.

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