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Article

Psychometric Properties of the Spanish Motives for Online Gaming Questionnaire in a Sample of College Students

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Abstract

This study investigates the psychometric properties of the Spanish version of the Motives for Online Gaming Questionnaire (MOGQ). We explored the factor structure and construct validity of the MOGQ through its relationships with gaming disorder symptoms (IGD–20) and impulsivity traits. We also analyzed if sociodemographic variables and gaming habits were related to gaming motives. An online cross-sectional survey was completed by 845 college students. Structure validity was examined using a combination of exploratory and confirmatory factor analyses, which supported a bifactor model composed of a general motivation factor and six uncorrelated factors (a mixed factor composed of escape and coping, competition, recreation, skill, social, and fantasy). Omega-hierarchical and omega coefficients were used to determine reliability of the MOGQ. The scale presented acceptable reliability for the general factor (ω_h = .79) and the specific factor scores (social ω = .79, escape/coping ω = .81, competition ω = .79, skill ω = .84, fantasy ω = .82, and recreation ω = .70). Positive associations were observed between the MOGQ and the IGD–20 symptoms, with escape/coping (r = .48) and fantasy (r =.40) showing the strongest ones. Null or low correlations were observed with impulsivity traits. Motives to play varied significantly across genders. These findings provide evidence that the Spanish version of the MOGQ is a reliable and valid tool to assess motives to play online games.

Keywords: gaming disorder; gender differences; loot boxes motives; online gaming

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Over the last forty years, the popularity of video games has significantly increased, making them an essential part of the global culture. At the beginning of the 21st century, the rise of the Internet and mobile phones launched video game use to new heights around the world, and, by 2023, the number of gamers is expected to reach 3.07 billion. In Spain, 18.2 million people played video games regularly in 2022 (Asociación Española de Videojuegos [AEVI], 2022; Newzoo International B.V., 2021). Video games are not only an important source of entertainment but also have a positive influence on several segments of our society. The video game industry plays a crucial role in the global economy (Tripp et al., 2020), the use of video games in educational settings has demonstrated practical value in supporting learning and facilitating the transmission of knowledge (Villani et al., 2018), and several studies have shown that video game-based interventions can have a positive impact in therapeutic and medical contexts (Halbrook et al., 2019; Xu et al., 2020). Video games can also provide a wide variety of psychological benefits (Verheijen et al., 2019; Wulansari et al., 2020), as enhancing intrinsic motivation, and fulfilling basic psychological needs like relatedness, autonomy, and competence (Przybylski et al., 2010). Despite this, negative consequences have also been associated with their excessive use, with their alleged addictive potential being a growing concern for mental health professionals and scientific communities (Chen et al., 2018; Reed et al., 2022).

In 2013, the *Diagnostic and Statistical Manual of Mental Disorders* (American Psychiatric Association, 2013) proposed internet gaming disorder (IGD) as a condition in need of further study; since then, numerous studies have investigated its psychological, social, and cultural correlates. More recently, but not without controversy, gaming disorder has been included as a new condition in the eleventh *International Classification of Diseases* (World Health Organization, 2019), eliminating the term 'Internet', as associating the disorder with Internet-related problems was not considered strictly necessary. A recent meta-analysis estimated that problematic gaming affects 1.4%–3.3% of the population worldwide (Kim, Son, et al., 2022). Notably, in Spain, the prevalence ranges from 1.7% to 8.3% among young adults (Beranuy et al., 2020; Buiza-Aguado et al., 2018) and from 1.8% (International Statistical Classification of Diseases and Related Health Problems, 11th Ed.; ICD–11 framework) to 3.1% (Diagnostic and Statistical Manual of Mental Disorders, 5th Ed.; DSM–5 framework) among adolescents (Nogueira-López et al., 2023). Also, problematic gaming is more prevalent in young men with family and interpersonal problems (Fumero et al., 2020; Mihara & Higuchi, 2017; Stevens et al., 2021). Regarding psychological variables, impulsivity (Ding et al., 2014; Ryu et al., 2018) and antisocial traits (Anderson et al., 2008; Müller et al., 2015; T'ng et al., 2020), among other individual differences

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factors, have been associated with problematic gaming symptoms. A large body of evidence has also shown that certain patterns of motives underlying gaming behaviors can predict gaming intensity (i.e., frequency and duration) and problems (i.e., addiction symptoms or negative consequences; Cheah et al., 2022; Király et al., 2022; Melodia et al., 2022).

Motives are the declarative goals people perceive as guiding their instrumental behaviors (Volkow et al., 2017). They encompass intrinsic, extrinsic, and experiential motivation (Banyte & Gadeikiene, 2015), and include pleasure seeking, curiosity, interest (Gómez-Maureira & Kniestedt, 2019; Gunnell & Gaudreau, 2015), experiencing intense emotions (Hemenover & Bowman, 2018), escaping from reality, dealing with negative feelings such as boredom or loneliness (Larche & Dixon, 2021), socializing with other people (Nebel & Ninaus, 2022), or obtaining economic rewards or prizes (Johnson et al., 2018). In the past, a large body of work has shown that the study of motives can help understand a variety of human behaviors, including substance and non-substance addictive behaviors (Barrada et al., 2019; Bennett & Holloway, 2017).

Process-based models of behavioral addiction have examined the dynamics of the relative utility attributed to positive and negative reinforcers in various stages of problematic gaming, including exposure, behavior consolidation, and dysregulated use (Perales et al., 2020). In this vein, a preponderant model such as the I-PACE (Interaction of Person-Affect-Cognition-Execution) includes motives among the moderating or mediating variables that explain the progression from recreational to dysregulated use of the Internet, including problematic video gaming (Brand et al., 2016). The onset of problematic gaming is more likely to occur in individuals who display a heightened motivation linked to gaming (Young & Brand, 2017). Notably, although the I-PACE model does not explicitly delineate specific motives, it considers the way the experience of positive and negative rewards during gaming aligns with distinct motives. This alignment is proposed to underlie the temporal progression of the disorder. In the initial stages, certain motives are postulated to be implicated in reward anticipation, while in later stages, other motives are hypothesized to be associated with coping with negative affect (Brand et al., 2016).

In view of this, numerous studies have investigated the motives why individuals engage in video game play (Király et al., 2022; Reid, 2012), and the circumstances in which these motives could be related to negative outcomes (Ng & Wiemer-Hastings, 2005; Wan & Chiou, 2006). Demetrovics et al. (2011) developed the most widely used measurement instrument to assess gaming motives: The Motives for Online Gaming Questionnaire (MOGQ). This instrument includes 27 items categorized into seven dimensions (social, escape, competition, coping, skill development, fantasy, and recreation motives). The use of this instrument has revealed that escape and fantasy motives are correlated with problematic gaming symptoms (Laconi et al., 2017). To date, although translations of the MOGQ exist in other languages, to our knowledge, it has only been systematically validated in Turkish, Italian, Chinese, Korean, and Persian. Table 1 summarizes the main findings of these studies.

Studies examining the psychometric properties of MOGQ across languages and samples have yielded slightly different factor structures, with specific motives such as coping, sometimes being identified as a unique factor, and sometimes being merged with other motives (Melodia et al., 2022). Most of these studies have however disregarded connections between motives, preferences, and gaming-related behaviors such as loot-box purchasing, which have been linked to gaming-related harms in the context of progressive gaming-gambling convergence (King & Delfabbro, 2020; Wardle & Zendle, 2021; Zendle & Cairns, 2019).

Additional research suggests that some cultural factors (e.g., individualistic vs. collectivist cultures, or culture-specific expressions of achievement motives) have an impact on gaming motivation and its association with problematic gaming (Snodgrass et al., 2013; Wang & Cheng, 2022), which reinforces the need to validate the scale in the language and culture in which it is used. In a similar vein, the role of motives in gaming behavior and their clinical implications can vary across genders (López-Fernández et al., 2019). Even though almost half of the population of gamers in Spain and other countries are females nowadays (AEVI, 2022; Newzoo International B. V., 2020), most MOGQ studies have focused on predominantly male samples, and studies in which females are fairly represented are much needed (López-Fernández et al., 2020b).

The main goal of this study is thus to test the psychometric properties of a Spanish version of the MOGQ (factorial structure, internal consistency, and validity) in a gender-balanced sample. Structure validity will be established through a combination of exploratory and confirmatory factor analyses (EFA/CFA). Convergent and divergent validity will be tested through the relationships between the MOGQ and problematic gaming symptoms, self-reported gaming time, impulsivity traits, and loot box consumption. Finally, we will explore gender differences among the previously cited variables.

Method

Participants and Procedure

Participants were recruited from four Spanish universities (the Catholic University of Murcia, the University of Granada, the University of Extremadura, and the University of Basque country). An email was sent to students inviting them to participate and complete an online survey. Participants were informed about the objectives of the study and the confidentiality of their responses. They gave consent to participate voluntarily. To encourage participation, five €15 gift cards were raffled at the end of the study. A total of 1,130 people participated in the study. The inclusion criteria were: (a) Being Spanish-speaking gamers, (b) reporting playing video games for at least 2 hours per week, and (c) being at least 18 years old. Participants were excluded if they did not meet the inclusion criteria, i.e., declaring they play less than two hours per week, being under 18 years old (n = 48), or providing invalid information (n = 104) (e.g., playing more than 7 days per week or more than 24 hours per day). The final sample consisted of 845 college students. The sample included 426 males (50.41%) and 417 females (49.35%), and participants were aged between 18 and 50 (M = 23.51, SD = 5.03) (please, see Table 2 for sample reported characteristics). All procedures performed in studies involving human participants were in accordance with the ethical standards of the institutional and/or national research committee and with the 1964 Helsinki Declaration and its later amendments or comparable ethical standards. The study was approved by the Catholic University of Murcia Ethics Committee (CE031905). This study is part of a larger project, and another study utilizing the same database has already been published (Infanti et al., 2023).

Measures

The Motives for Online Gaming Questionnaire (MOGQ) (Demetrovics et al., 2011), aims to measure seven gaming motives

Table 1. Validation Studies of the Motives for Online Gaming Questionnaire and Their Main Outcomes.

Authors	Country	Confirmatory Factor Analysis	Item analysis	Outcomes
Ballabio et al. (2017)	Italy	Seven–factor structure (corresponding to the original model).	The similar meaning of some items explains the need for error covariances. Item 8 (How often do you lose track of time when gaming?) was closely related to Item 13 (How often do you feel time stops while gaming?), and Item 6 (How often do you fail to meet up with a friend because you were gaming?) was similar to Item 16 (How often do you choose gaming over going out with someone?).	Reliability: Ranging from .75 to .88. Convergent validity: Positive correlations with some dimensions of the Online Gaming Motivation Scale, Gaming Motivation Scale and The Player Experience of Need Satisfaction. Criterion validity: Psychiatric symptoms were directly and indirectly associated with problematic online gaming (via escape and fantasy motivations).
Wu et al. (2017)	China	Bifactor structure with seven specific factors and a general motivation factor subsuming all items.	Item 18 (originally from the coping factor) was not associated with any factor and was removed. Item 4 (originally from the coping factor) was associated with recreation and coping factors.	Reliability: Ranging from .83 to .90. Criterion validity: Positive correlation with psychological need satisfaction and time spent to play. General motivation factor, high escape, and low skill development emerged as motives related to problematic gaming.
Evren et al. (2020)	Turkey	Six–factor structure (coping and escape merged as single factor).	Item 18 (originally from the coping factor) showed a negative loading for recreation factor and a negative loading for coping/escape factor. The authors decided to remove this item.	Reliability: Ranging from .87 to .92. Convergent validity: Positive correlations with Electronic Gaming Motives and symptoms of problematic gaming and its severity. Criterion validity: Correlations with time spent gaming.
Marino et al. (2020)	Italy	Seven–factor structure (corresponding to the original model).	The items all load on their respective factors, and no cross–loading problems were identified.	Reliability: Ranging from .87 to .92. Criterion validity: Escape, coping and fantasy were positively related to social anxiety. Recreation showed a negative relationship with social anxiety. Escape dimensions payed a mediating role between social anxiety and problematic gaming symptoms.
Kim & Kang (2021)	South Korea	Six–factor structure (coping factor disappeared with respect the original model).	Item 4 (originally from the coping factor) and Item 11 cross–load on the escape and recreation factors and were removed. Item 25 (originally from the coping factor) is cross–loading on escape and skill development factors and was removed. Item 17 (originally from the competition factor) is cross–loading on skill development and competition and was removed. Item 18 (originally from the coping factor) loaded on the fantasy factor instead of the coping factor.	Reliability: Ranging from .80 to .87. Incremental validity: Escape and fantasy predict problematic gaming symptoms above and beyond what was explained by extraversion, neuroticism and impulsivity personality variables.
Hamzehzadeh et al. (2022)	Iran	Seven factor structure (corresponding to the original model).	The items all load on their respective factors, and no cross–loading problems were identified.	Reliability: Ranging from .68 to .87. Test–retest reliability: Intraclass correlation coeficients for the entire questionnaire was .85. Convergent validity: Positive correlations with some dimensions of the Gaming Motivation Scale and the Player Experience of Need Satisfaction.
Dowran et al. (2022)	Iran	Seven factor structure (corresponding to the original model).	The items all load on their respective factors, and no cross–loading problems were identified.	Reliability: Ranging from .79 to .85. Convergent validity: Positive correlations with problematic gaming symptoms.

(social, escape, competition, skill development, coping, fantasy, and recreation). A description of the seven dimensions can be found in Table 3. The questionnaire had 27 items scored on a five-point scale (1 = *Almost never/Never* to 5 = *Almost always/Always*). The original version showed good internal consistency and Cronbach's alpha values range from .79 to .90. To develop the Spanish version, a native Spanish psychologist translated the questionnaire from English to Spanish. Next, an English philologist with Spanish linguistics skills translated the questionnaire back from Spanish to English. Finally, a group of three psychologists (two English and one Spanish), fluent in both languages, reviewed the final version. They concluded that the language and expressions were clear and adjusted to the linguistic context.

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Table 2. Sociodemographic and Game-Related Variables.

	Mean (*SD*)	*N* (%)
Age	23.51 (5.03)	
Hours of gaming per day	2.02 (1.79)	
Gender		
Male		426 (50.41)
Female		417 (49.35)
Other		2 (0.24)
Educational level		
Primary education		4 (0.47)
Secondary education		15 (1.78)
Vocational education and training		13 (1.54)
Middle school education		58 (6.86)
High school education		157 (18.58)
University degree		439 (51.95)
Master's degree		139 (16.45)
Doctorate		20 (2.37)
Loot boxes consumption		
Yes		194 (22.96)
No		651 (77.04)
Playing on PC		
Male		182
Female		87
Other		1
Total		270 (31.95)
Playing on Console		
Male		159
Female		61
Other		0
Total		220 (26.04)
Playing on portable/tablet		
Male		85
Female		269
Other		1
Total		355 (42.01)
Online		
Yes		666 (78.82)
No		179 (21.18)

The Internet Gaming Disorder test (IGD–20) (Spanish version by Fuster et al. [2016], original English version by Pontes et al. [2014]) evaluates the presence of IGD symptoms based on the DSM–5 framework and the component model of addiction by Griffiths (2005). The scale consists of 20 items that are answered using a 5-point Likert scale (1 = *Strongly disagree* to 5 = *Strongly agree*). The Spanish version of the IGD–20 showed good psychometric properties and Cronbach's alpha was .87 (Fuster et al., 2016).

Table 3. Questionnaire Variables of the Motives for Online Gaming Questionnaire and the Impulsivity Scale.

Questionnaire	Subscale	Subscale description
Motives for Online Gaming Questionnaire		
	Social	Enjoyment from social interaction, meeting new people, and collaborating towards gaming goals.
	Escape	Desire to leaving reality and real–life problems.
	Competition	Seeking challenges with others and striving to obtain or win something.
	Skill development	Improvement of cognitive skills or abilities related to gaming.
	Coping	Enhancement of mood, managing negative emotions and aggressive behaviors.
	Fantasy	Fantasizing new identities, stories, and atmospheres that are impossible in the real world.
	Recreation	Intention to have fun, relax, or experience enjoyment.
UPPS Impulsivity Scale		
	Negative urgency	Tendency to act rashly in negative emotional contexts (e.g., sadness, fear, anger…).
	Lack of premeditation	Tendency to make decisions without considering the potential consequences.
	Lack of perseverance	Tendency to quit of boring or difficult tasks.
	Sensation seeking	Tendency to seek out novel experiences and preferences for thrilling activities.
	Positive urgency	Tendency to act rashly in positive emotional contexts (e.g., happiness, enjoyment, euphoria…).

Note. UPPS-P = Urgency (negative), Premeditation (lack of), Perseverance (lack of), Sensation Seeking, Urgency (positive), Impulsive Behavior Scale.

In the present sample, Cronbach's alpha (computed using Spearman correlations) was .91 for the total score.

The short Spanish version of the Urgency (negative), Premeditation (lack of), Perseverance (lack of), Sensation Seeking, Urgency (positive), Impulsive Behavior Scale (UPPS-P Impulsivity Scale) (Spanish version by Cándido et al. [2012], original French version by Billieux et al., [2012]) assesses five impulsivity traits (negative urgency, lack of premeditation, lack of perseverance, sensation seeking, and positive urgency). Table 3 provides a description of these dimensions. The scale consists of 20 items scored on a Likert scale (1 = *Strongly agree* to 4 = *Strongly disagree*). The original version showed good internal reliability, with Cronbach's alpha ranging from .61 and .82. For our current sample, a Cronbach's

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alpha for each dimension was computed using Spearman correlations and were .79 for lack of perseverance, .76 for lack of premeditation, .81 for sensation seeking, .66 for positive urgency, and .82 for negative urgency. In the current study, we merged positive and negative urgency as recent studies suggest that they belong to a single construct (Billieux et al., 2021). The Cronbach's alpha of general urgency was .81.

Participants were asked about different gaming habits, including: (a) How many days per week they dedicated to gaming, (b) how many hours per day they dedicated to gaming, and (c) how much money they spent on loot boxes in the last month.

Data Analytic Strategy

In the present paper, we separated the whole dataset into two different samples. One sample was used for the purpose of an exploratory factor analysis (EFA) ($n = 272$) while the other sample was used to test the structure validity through a confirmatory factor analysis (CFA) ($n = 573$). The decision to combine EFA and CFA obeyed to several reasons. First, previous studies have favored slightly different factor structures (Evren et al., 2020; Kim & Kang, 2021; Wu et al., 2017), so it is unclear which of them should be favored *a priori* in a strictly theory-driven CFA. And second, the original seven-factor model was derived from samples that primarily consisted of males (Ballabio et al., 2017; Dowran et al., 2022; Hamzehzadeh et al., 2022; Marino et al., 2020), which could compromise its generalizability. The combination of EFA and CFA prevents the structure selection from being influenced by theoretical preconceptions, and it has been successfully applied in previous psychometric MOGQ studies (Evren et al., 2020; Kim & Kang, 2021;).

We followed Hair et al.'s (2014) guidelines for EFA and CFA analyses. The separation of the dataset was done to have a ratio of ten observations per variable (e.g., a minimum of 270 observations) for the EFA. The two samples did not present significant differences in demographic data (education level, age, gaming hours per day). Before performing the EFA, we assessed sample adequacy using the Kaiser-Meyer-Olkin (KMO) and Bartlett's sphericity test. To select the number of factors to retain in our model, we used scree plot, eigenvalue, and parallel analysis methods. We then performed an EFA using the maximum likelihood method with Promax rotation (as the gaming motives might be related and are thus not presumed to be orthogonal) and 100 iterations. These options were also used in the Korean and Turkish validation studies of the MOGQ Scale (Evren et al., 2020; Kim & Kang, 2021). As the EFA sample size is larger than 250 and smaller than 350, items having a factor loading value higher than .35 were identified as significant (Hair et al., 2014). When an item had more than one significant factor loading or no significant factor loadings, the item was discarded. For the CFA, we used a robust variant of maximum likelihood as an estimator (MLR). Several CFAs were computed to test different competing models and were then compared using a series of fit measures. Models tested included the original seven-factor model, a second-order model, a model created using the EFA factor structure, a bifactor model consisting of a general motivation and the seven original factors (uncorrelated), and a bifactor model consisting of a general motivation and the factors identified by EFA (uncorrelated). Bifactor models have been included following evidence from the Chinese validation of the MOGQ (Wu et al., 2017).

Model Chi-Square, comparative fit index (CFI), Tucker Lewis index (TLI), root mean square error of approximation (RMSEA), standardized root mean square residual (SRMR), and Akaike information criteria (AIC) were used for the models' evaluation. Good fit model references are a normed χ^2 lower than 5.0, a CFI value higher than .95, a TLI value higher than .95, a SRMR value lower than .08, a RMSEA value lower than .08, and an AIC smaller value (Hooper et al., 2008). To assess reliability, we computed the Omega and Cronbach's alpha coefficients using the CFA sample. Invariance across genders was tested in the CFA sample using multiple group CFA (MGCFA) and chi-square difference tests to compare configural and metric models, followed by a comparison between metric and scalar models. The criterion was the non-significance of the chi-square difference test. We then assessed the convergent and divergent validity by observing the partial correlations between the scores of the MOGQ Scale, the IGD–20 Scale, the short UPPS-P, and the estimated gaming hours per day while controlling for gender. Finally, we observed if there were significant differences in MOGQ Scale scores as a function of gender and loot box consumption. These analyses were computed using R (v4.0.3) and the following packages: haven (Wickham et al., 2022), tidyverse (Wickham et al., 2019), corrplot (Wei & Simko, 2021), psych (Revelle, 2022), lavaan (Rosseel, 2012), and alookr (Ryu, 2022). The dataset and the code are available via the following Open Science Framework (OSF).[1]

Results

Shapiro-Wilk normality tests showed non-normal distribution for the MOGQ total score and its factors.

Exploratory Factor Analyses

Premise analyses confirmed the sampling adequacy of the EFA sample, with good KMO (measure of sampling adequacy, MSA = .9) and a significant Bartlett's sphericity test ($\chi^2 = 5{,}270.655$, $df = 351$, $p < .001$). Eigenvalues rule (value greater than 1), scree plot, and parallel analyses suggested the presence of six factors. EFA results are reported in Table 4. Factor 1 (Items 2, 4, 9, 11, 16, 23, 25) is a mix of escape and coping dimensions present in the original scale. Factor 2 (Items 3, 10, 17, 24) corresponds to the competition dimension. Factor 3 (Items 7, 14, 21) corresponds to the recreation dimension. Factor 4 (Items 5, 12, 19, 26) corresponds to the skill dimension. Factor 5 (Items 1, 8, 15) corresponds to the social dimension. Finally, Factor 6 (Items 6, 13, 20) corresponds to the fantasy dimension. Items 18 ("*…because it helps me to channel my aggression*") and 22 (*"…because playing gives me company"*) did not load on any retained factors and were thus deleted. Item 27 ("*…because I can be in another world*") cross-loaded onto Factors 1 and 6 and was also deleted from the EFA model.

Confirmatory Factor Analyses

Several CFAs were conducted to compare the goodness of fit of the different models. In total, five models were tested. Model 1 is derived from the original MOGQ paper (Demetrovics et al., 2011) and consists of seven factors correlated to each other without the presence of a general factor. Model 2 is a second-order model where the items are related to the seven factors which are themselves related to a general factor. Model 3 consists

[1] https://osf.io/jk94v/?view_only=118f5cee309a4d9aa48fdf1dde1392e4

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Table 4. Exploratory Factor Analysis of the MOGQ (EFA sample, n = 272).

Item	Factor loadings 1.	Factor 2	Factor 3	Factor 4	Factor 5	Factor 6	Item–total correlation
2	0.90						.66**
4	0.45						.75**
9	1.08						.67**
11	0.62						.73**
16	1.01						.70**
23	0.88						.64**
25	0.53						.74**
3		0.42					.52**
10		0.69					.54**
17		0.89					.54**
24		1.02					.53**
7			0.55				.45**
14			1.08				.51**
21			1.02				.56**
5				0.97			.69**
12				0.97			.71**
19				0.79			.66**
26				0.99			.69**
1					0.95		.38**
8					0.95		.50**
15					0.42		.61**
6						0.48	.60**
13						0.94	.62**
20						0.87	.58**
27	0.42					0.40	.72**
18	/	/	/	/	/	/	.52**
22	/	/	/	/	/	/	.65**
Range	7–35	4–20	3–15	4–20	3–15	3–15	
M	17.25	9.54	11.71	9.02	5.61	5.17	
M / *n* items	2.46	2.39	3.90	2.26	1.87	1.72	
SD	7.66	4.25	3.25	4.54	2.85	2.93	

Note. EFA = exploratory factor analysis.
**Significant at $p < .001$ level. * Significant at $p < .05$ level.

of the six factors derived from the EFA without any general factor. Model 4 is a bifactor model where all the items are related to a general motivation and to seven uncorrelated factors. Finally, Model 5 is a bifactor model where all the items are related to a general motivation and to six uncorrelated factors (as identified by EFA).

Globally, bifactor models (Model 4 and 5) presented the best fit, suggesting the presence of a general factor related to all the items and several uncorrelated factors. Among them, Model 5 (EFA bifactor model, a bifactor model consisting of a general motivation and six uncorrelated factors as identified in EFA) satisfied most of the fit indices (χ^2/df = 3.267, CFI = .938, TLI = .925, RMSEA = .068 [.062, .073], SRMR = .065, AIC = 35,981.793) (please, see Table 5). Even though Model 5 had incremental fit values (CFI and TLI) lower than .95, it is important to note that a threshold of .90 was used in the Korean and Turkish validation for these values (Evren et al., 2020; Kim & Kang, 2021). We thus conclude that the CFI and TLI values could be acceptable. Invariance analysis across genders (*n* male = 290; *n* female = 281) yielded no significant chi-square difference tests. No differences have been found when comparing configural invariance and metric invariance models ($\Delta\chi^2$ = 53.212, *df* = 41, *p* = .096), and for metric and scalar invariance models ($\Delta\chi^2$ = 22.103, *df* = 17, *p* = .181).

EFA Bifactor Model Reliability

Because the EFA bifactor model presents the best fit, we computed the omega-hierarchical (ω_h) for the general factor (general motivation) (Flora, 2020). The omega-hierarchical is especially relevant when it comes to measuring the reliability of the general factor of a bifactor model. Indeed, it reflects the proportion of the total-score variance as a result of a general factor, regardless of the multidimensional aspect of the scale (Flora, 2020). For the specific factors, we reported the omega (ω) which represents the stability of the specific factors but also the general factor (Reise, 2012). The omega-hierarchical for the general factor was .79 (α = .93). Regarding the specific factors, omegas were .79 for social (α = .80), .81 for escape/coping (α = .91), .79 for competition (α = .85), .84 for skill (α = .92), .82 for fantasy (α = .85), and .70 for recreation (α = .82). Thus, the scale presents an acceptable reliability for the general factor and the specific factor scores.

EFA Bifactor Model and External Variables

Using the EFA bifactor model, we explored how impulsivity traits, IGD symptoms, and gaming hours per day correlated with the MOGQ's dimensions and general motivation. Results are displayed in Table 6. All MOGQ's subscales showed significant correlations with daily gaming hours (correlations between r_s = .14, $p < .001$, with competition motivation, to r_s = .35, $p < .001$, with general motivation). All gaming motives positively correlated with IGD–20 total score, except recreation motives. In general, correlations between impulsivity traits and motives to play were small and non-significant. Significant correlations ranged from –.07 (p = .041, between lack of premeditation and recreation motivation) to .16 ($p < .001$, between urgency and fantasy motivation). Finally, loot box consumption was linked to all specific gaming motives, the highest relationship being observed with the competition dimension (r_{pb} = .25, $p < .001$), and the lowest with fantasy dimension (r_{pb} = .08, p = .021).

Regarding the gender comparison (please, see Table 7), we found males to score significantly higher on general motivation but also in the competition, recreation, and skill dimensions. No significant differences were found for the escape/coping, social, and fantasy dimensions.

Discussion

The main aim of this study was to establish the factor structure of the Spanish version of the MOGQ in a large convenience sample of college students regularly participating in video gaming activities. The combination of EFA and CFA results supports a bifactor model

Table 5. Model Fit of the Measurement Models for MOGQ Items (n = 573).

			Absolute fit			Parsimony fit	Incremental fit	
Model	*df*	χ^2	χ^2/df	RMSEA	SRMR	AIC	CFI	TLI
Model 1: Original 7 Factors model	303	1,360.652	4,491	.086 [.081, .090]	.106	41,015.724	.883	.865
Model 2: Second Order Factors (7 original Factors)	317	1,515.096	4,779	.089 [.084, .093]	.116	41,172.985	.868	.853
Model 3: 6 Factors model (EFA)	237	1,185.819	5,003	.090 [.085, .095]	.092	36,481.286	.886	.867
Model 4: Bifactor (general and 7 specifics factors)	297	1,174.722	3,955	.074 [.070, .079]	.064	40,652.348	.913	.898
Model 5: Bifactor (general and 6 specifics factors)	228	744.925	3,267	.068 [.062, .073]	.065	35,981.793	.938	.925

Note. CFA = confirmatory factor analysis; RMSEA = root-mean-square error of approximation; SRMR = standardized root mean square residual; AIC = akaike information criteria; CFI = comparative fit index; TLI = Tucker Lewis index; EFA = exploratory factor analysis.

Table 6. EFA Model Spearman's Partial Correlations with Impulsivity Traits, Gaming Disorder Symptoms as Measured by the IGD–20, Gaming Hours per day, Loot Boxes[a]

			MOGQ						
			Escape/Coping	Competition	Recreation	Skill	Social	Fantasy	General motivation
Gaming hours/day		r_s	.30	.14	.17	.30	.27	.27	.35
		p–value	< .001	< .001	< .001	< .001	< .001	< .001	< .001
IGD	Total score	r_s	.48	.28	.06	.30	.33	.40	.46
		p–value	< .001	< .001	.071	< .001	< .001	< .001	< .001
UPPS–P	Urgency	r_s	.13	.11	–.09	.05	.09	.16	.11
		p–value	< .001	.001	.007	.133	.013	< .001	.002
	Sensation seeking	r_s	.05	.13	–.08	.07	.06	.12	.08
		p–value	.132	< .001	.024	.040	.059	< .001	.014
	Lack of premeditation	r_s	.02	.02	–.07	–.05	–.01	.06	–.01
		p–value	.497	.506	.041	.157	.719	.088	.857
	Lack of perseverance	r_s	.12	.03	–.02	–.04	.03	.15	.07
		p–value	< .001	.407	.614	.200	.387	< .001	.032
Loot Boxes[a]		r_{pb}	.10	.25	.18	.22	.14	.08	.22
		p–value	.003	< .001	< .001	< .001	< .001	.021	< .001

Note. Conditioned on Gender. *p*-value adjustment method: Benjamini & Hochberg (1995). EFA = exploratory factor analysis; MOGQ = Motives for Online Gaming Questionnaire; IGD = Internet Gaming Disorder; UPPS-P = Urgency (negative), Premeditation (lack of), Perseverance (lack of), Sensation Seeking, Urgency (positive), Impulsive Behavior Scale.
[a]Point-Biserial method instead of Spearman.

consisting of a general motivation factor and six uncorrelated subfactors (a merged escape/coping motives subfactor, and separated subfactors for competition, recreation, skill, social, and fantasy motives); the final scale showed satisfactory psychometric properties. Omega and Cronbach's alpha coefficients showed acceptable or good reliability and the correlations with time spent on video games and IGD symptoms support the construct validity of the scale.

The overlap between coping and escape motives is one of the most noteworthy results of the present study and converges with part of the existing literature (Evren et al., 2020; Kim & Kang, 2021). Relatedly, some items included in the coping factor in previous studies present questionable levels of robustness and representativity (Evren et al., 2020; Kim & Kang, 2021; Wu et al., 2017). For instance, Item 18 *("…because it helps me to channel my aggression")*, did not load significantly on any factor and was subsequently removed from the final version in the present study. This item was also removed from the coping factor in the Chinese version (Wu et al., 2017), in which it aligned with skill development and recreational motives. In the Turkish version (Evren et al., 2020), the same item negatively loaded onto the recreation factor, and positively onto the coping/escape factor. And finally, in the Korean version (Kim & Kang, 2021) it loaded onto the fantasy factor. We did not find any other unexpected loadings for items within the coping factor, although the Korean version (Kim & Kang, 2021) excluded Items 4 and 11, and the Chinese version (Wu et al., 2017) found Item 4 to cross-load onto the coping and recreation factors.

The superiority of six-factor models relative to their seven-factor counterparts –with coping and escape motives loading on a single factor– supports the theoretical model proposed by Demetrovics et al. (2011) for the original version of the MOGQ and closely replicates the findings from the Turkish version (Evren et al., 2020). Although a six-factor structure was also reported for the Korean version of the MOGQ this was due to the disappearance of the coping factor (Kim & Kang, 2021). In contrast, the Italian, Hungarian, and Persian validations leaned towards the seven-factor structure (Ballabio et al., 2017; Demetrovics et al., 2011; Dowran et al., 2022; Hamzehzadeh et al., 2022). Although general East-West

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Table 7. Wilcoxon Mann Whitney Test between Female and Male on MOGQ, IGD–20, UPPS-P, and Gaming Hours per Day.

	Female		Male			
	M (*SD*)	*Mdn*	*M* (*SD*)	*Mdn*	*W* test	*p*-value
MOGQ						
Escape/Coping	17.27 (7.28)	17	17.38 (7.54)	16	88,727	.979
Competition	8.52 (3.83)	8	10.58 (4.40)	10	113,413	< .001
Recreation	11.17 (3.27)	12	12.16 (3.03)	13	105,649	< .001
Skill	8.95 (4.43)	8	10.12 (4.80)	9.5	101,293	< .001
Social	5.58 (2.84)	5	5.83 (2.84)	5	94,844	.083
Fantasy	5.30 (3.06)	4	5.42 (3.10)	4	90,957	.526
General motivation	56.79 (17.99)	56	61.49 (17.78)	61	100,776	< .001
IGD–20						
Total score	34.65 (10.65)	31	38.82 (12.61)	37	105,701	< .001
UPPS–P						
Urgency	19.22 (4.93)	19	18.07 (4.92)	18	76,312	< .001
Sensation seeking	9.87 (3.03)	10	10.06 (3.02)	10	92,398	.309
Lack of premeditation	7.17 (2.33)	7	7.09 (2.41)	7	86,410	.491
Lack of perseverance	7.31 (2.51)	7	7.67 (2.76)	7	94,553	.102
Gaming hours/day	1.58 (1.30)	1	2.44 (2.08)	2	115,088	< .001

Note. MOGQ = *Motives for Online Gaming Questionnaire*; IGD = *Internet Gaming Disorder Test*; UPPS-P = Urgency (negative), Premeditation (lack of), Perseverance (lack of), Sensation Seeking, Urgency (positive), Impulsive Behavior Scale.

cultural differences have been proposed as an explanation for this divergence (Wu et al., 2017), our results are closer to those of the Chinese validation than to those from European and Middle East countries, and thus fail to support this explanation.

Complementary evidence points to the possible role of the distribution of individual differences in the samples used for different studies. Motives are tightly related to gaming preferences, and these to other individual characteristics and sociodemographic variables (Gómez-Gonzalvo et al., 2020, Kim, Nam, et al., 2022; Rehbein et al., 2016; Veltri et al., 2014). More specifically, some of these studies report differences in gaming preferences between males and females (Gómez-Gonzalvo et al., 2020; Phan et al., 2012; Veltri et al., 2014), with male video gamers tending to be more competitive than women and showing a preference for games that require navigation and attentional skills (Buser et al., 2021; Tungodden & Willén, 2022). Although we are aware that these results must be interpreted with caution (as video game studies is a male-dominated area with a high presence of gender stereotypes and women are often invisibilized (Lange et al., 2021, Wohn et al., 2020), male participants in our study scored significantly higher than female ones in the competition, recreation, and skill motives.

With regard to the possible role of gender, an important coincidence between the Chinese validation and ours, as well as a methodological advantage relative to other studies, is the almost gender-balanced sample composition in both (54.6% males to 45.4% females and 50.41% males to 49.35% females, respectively). This makes our results more representative of the factor structure of motives in the general population of gamers. Although, in the present study, we found measurement invariance across genders, additional factor invariance analyses, as well as gender-informed analyses for other individual characteristics and demographics, are still needed.

The overlap between coping and escape motives is also supported by their consistent and shared associations with gaming-related problems (Griffiths et al., 2016; Kuss et al., 2012) and other problematic behaviors (Estévez et al., 2021; Froushani & Akrami, 2017; Jauregui et al., 2017; Lee-Winn et al., 2018). This convergence of internal and external validity results has led some authors to suggest that coping and escape behaviors are underpinned by a common mechanism (see Barrada et al., 2019). According to this view, problematic gaming would work as an overt emotion regulation strategy and would be maintained by negative reinforcement (Kardefelt-Winther, 2017). Still, an open question remains regarding the potential distinction between the healthy and unhealthy use of video gaming to deal with stress and other aversive states. On the one hand, as noted earlier, coping/escape motives are consistently associated with the severity of a range of potentially problematic activities (see also Chang & Lin, 2019; de Hesselle et al., 2021; Kircaburun et al., 2020; Laconi et al., 2017; López-Fernández et al., 2020a; Melodia et al., 2022; Rafiemanesh et al., 2022). On the other hand, many individuals report using video gaming as a buffer against life stress, with video gaming positively contributing to their well-being as a consequence of this (Bourgonjon et al., 2016; Ceranoglu, 2010). In the words of Gee (2007) "*there are escapes that lead nowhere, like hard drugs, and escapes like […] gaming that can lead to the imagination of new worlds, new possibilities to deal with those perils and pitfalls, new possibilities for better lives for everyone*" (p. 1741). In this direction, Giardina et al. (2023) proposed to distinguish between *escapism* (a temporal and adaptive process with positive effects for the individuals) and *escape* (a maladaptive avoidant coping strategy used to disconnect from reality).

In any case, further research (and maybe a less biased approach; see Granic et al., 2014) is needed to understand the connection between coping/escape and video gaming problems. Tentatively, some factors beyond motives could explain why negatively reinforced video gaming acquires a disproportionate salience and value in some individuals, relative to other activities, in a way that ends up being harmful (see Perales et al., 2020). Among these factors, there is a pre-existing impoverished or not sufficiently rewarding environment, lack of skills to satisfy the same motives using other activities or strategies, emotion regulation problems, or impulsivity. The importance of the latter is indeed supported by the observed significant association of urgency with the specific motives most strongly related to IGD–20 scores. However, urgency-motives correlations remain weak (see also Zsila et al., 2018), which supports the divergent validity of the motives scale and suggests that impulsivity could play a mediating or moderating role in problematic video game use, as it seems to occur in other problematic behaviors (Adams et al., 2012; Marzilli et al., 2020).

With regard to the rest of the motives, fantasy was distinct enough from coping/escape to confirm a separated motivational dimension, but the two motive types are aligned in terms of their relationship with urgency and gaming-related problems. The most relevant difference between the two is the specific significant (but weak) association between fantasy motives and sensation seeking. This general parallelism could indicate that fantasy and coping/escape motives can have a similar negative function. At the other end, recreation motives are least strongly related with video gaming problems, replicating previous results (Bäcklund et al., 2022; Kircaburun et al., 2020; Moudiab & Spada, 2019). Again, this converges with evidence from other behavioral domains in which fun, thrill, or entertainment motives are unrelated to problems derived from those activities, or even protect against them (see Barrada et al., 2019).

Finally, and concerning gaming habits, we found loot box consumption to be significantly (but weakly) related to all MOGQ dimensions, with the two highest correlations observed for competition (r_{pb} = .25, p < .001) and skill (r_{pb} = .22, p < .001), and the two lowest for fantasy (r_{pb} = .08, p = .021) and coping/escape (r_{pb} = .10, p = .03). In other words, there is little overlap between predictors of video gaming problems and predictors of loot box use. Loot box use seems to be motivated by the desire to perform better or to be more competitive in the game, which makes sense in view that loot box contents in some popular games are often not only cosmetic but also impactful on gaming performance (e.g., player packs in FIFA Ultimate Game; Lemmens, 2022). Consequently, problematic loot box use might be due to other mechanisms than those linking problematic gaming with negative reinforcement motives (see, for example, Li et al., 2019, for a depiction of the specific contribution of loot box consumption to the comorbidity between gambling and gaming problems).

The present study is not without limitations. First, participants were recruited from university colleges, which limit the representativeness of the sample and the generalizability of results. Second, the cross-sectional design of the study does not allow establishing the directionality of links between the variables of interest. Third, data were collected using self-report questionnaires, meaning they could be affected by a desirability bias or question misinterpretation. Fourth, the participants in this study are heterogeneous and belong to different segments within the gaming community in terms of gaming intensity. Therefore, generalizing the results to specific player profiles (e.g., high-intensity gamers) is not warranted. Finally, most of the correlations found to be significant were so because of the large sample used, but they are mostly small-sized and range in a narrow interval. In view of that, their practical and theoretical significance should be interpreted cautiously, especially in the case of tentative explanations for comparisons between correlations (i.e., differential correlation patterns of motives with different outcomes).

The main strengths of the study include the large sample, which allows a robust set of associations, while the combined use of EFA and CFA ensures that factor structure selection is not biased by theoretical preconceptions. Theory-driven models are contrasted against data-driven counterparts, where the latter are cross-validated across subsamples. The superiority of six-factor solutions over seven-factor ones can thus be considered robust for the population of reference; the MOGQ is a reliable and valid instrument to assess motives to play online games. In theoretical terms, our study also confirms that gaming motives are related to gaming disorder symptoms and highlights the comparative importance of coping/escape and fantasy motives. Furthermore, this study confirms that gender and urgency could be relevant to understand how motives impact video gaming and its outcomes.

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