Validation of the Metacognitions about Online Gaming Scale (MOGS)

among Chinese gamers

Le DANG1,2,3

Juliet Honglei CHEN1,2

Hui ZHOU1,2

Marcantonio M. SPADA4

Anise M. S. WU1,2\*

1 Department of Psychology, Faculty of Social Sciences, University of Macau, Macao, China.

2 Centre for Cognitive and Brain Sciences, Institute of Collaborative Innovation, University of Macau, Macao, China.

3 Faculty of Teacher Education, Pingdingshan University, Pingdingshan, China.

4 School of Applied Sciences, London South Bank University, London, UK.

\*Corresponding author:

Prof. Anise M. S. Wu

Department of Psychology, Faculty of Social Sciences, University of Macau, Avenida da Universidade, Taipa, Macao, China.

E-mail: [anisewu@um.edu.mo](mailto:anisewu@um.edu.mo)

Phone: +(853) 8822 8377

Fax: +853-8822-2337

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**Abstract**

With the largest online gamer population worldwide and a heightened rate of Internet Gaming Disorder (IGD), China has a long-lasting need to identify salient correlates of IGD and provide corresponding assessment tools to support cost-effective IGD screening and interventions. To respond to such a need, the present study aimed to validate the Metacognitions about Online Gaming Scale (MOGS) among Chinese gamers to provide an additional tool for promoting studies investigating metacognition, a promising and newly emerged correlate of IGD, in China. To evaluate the psychometric properties of MOGS, we acquired, through a survey method, the data of 1340 lifetime gamers (59.3% female, *M*age = 19.84 years) from universities in China, in which 262 gamers also participated in the one-month retest. Our results indicated that the Chinese version of MOGS has a two-factor structure and satisfactory reliabilities (*α* = .90 and.92, ICC = .60 and.64, AVE = .56 and .70, ρc = .88 and .92). Moreover, MOGS’s convergent validity was evidenced by the expected, positive associations with generic metacognitions, stress, and IGD tendency (*r* (1338) = .29-.55, *p* < .001) as well as significant MOGS differences between probable IGD and non-IGD gamers (*p* < .001). A 6-item, short-form MOGS, which displayed equivalent psychometric soundness as its full-scale counterpart, was also developed. As the first study to validate MOGS among Chinese gamers, the present study attested to the readiness of this measure in facilitating further studies of gaming-specific metacognitions for early identification of and tailored interventions for high-risk gamers in China.

***Keywords*:** Chinese, gaming disorder, generic metacognitions, Internet Gaming Disorder, metacognitions about online gaming, psychometric properties

**1. Introduction**

Internet Gaming Disorder (IGD) or gaming disorder (GD) is a persistent and recurrent pattern of excessive and uncontrollable Internet gaming, which results in a cluster of cognitive and behavioral symptoms, impaired daily functioning, and significant psychological distress (American Psychiatric Association [APA], 2013; World Health Organization [WHO], 2019). The prevalence of IGD, formerly known as problematic online gaming and gaming addiction, ranges from 3.5% to 17.0% in China (Long et al., 2018), higher than the pooled estimation of its global prevalence of 2.47- 3.05% (Pan et al., 2020; Stevens et al., 2021). According to the latest statistics of gamers worldwide (Lai, 2021), 20.5% of the overall 3243.3 million gamers were from China. This alarmingly large-sized Chinese gamer population, accompanied with the heightened IGD rate, underscores the long-lasting need of identifying risk factors of IGD and providing reliable and valid assessment tools for these corresponding factors in the early identification of high-risk gamers and timely interventions. Responding to such a need, the present study aimed to validate the Metacognitions about Online Gaming Scale (MOGS; Spada & Caselli, 2017) among Chinese people.

Metacognition refers to thinking about one’s thinking. When Wells and Matthews (1994) introduced the first metacognitive model of psychopathology, the Self-Regulatory Executive Function (S-REF) model, they defined metacognition as beliefs about cognitive-affective experiences and ways of controlling these experiences. According to the S-REF model, metacognitions exert top-down or strategic influences on processing bias and underpinning coping styles that lead to psychological distress and addictive behaviors (Spada et al., 2015). A growing body of literature, based on cross-sectional studies and systematic reviews, has shown that metacognitive beliefs could be associated with not only anxiety and depression (Sun et al., 2017) but also psychosis (Sellers et al., 2017), eating disorders (Palmieri et al., 2021), stress-related disorders (Wells et al., 2012), and behavioral addictions (Zhang et al., 2020). As metacognitions may vary across disorders (Casale et al., 2021), Spada and Caselli (2017) drew researchers’ attention from generic metacognitions (i.e., generic beliefs about cognitive-affective experiences, such as “I need to control my mind at all times”) to specific metacognitions involved in IGD, by developing MOGS. Metacognitions about online gaming are theorized to guide cognitive appraisal and coping style and (dis)regulate behaviors during the pre-, during-, and post-engagement phases towards external triggers (e.g., exposure to online gaming). As shown by subsequent studies, these specific metacognitions about online gaming have stronger associations with IGD (e.g., .45-.75; Akbari et al., 2021; Nazligül, & Süsen, 2021) compared to generic metacognitions (e.g., .16-.33; Aydın et al., 2020; Zhang et al., 2020). Therefore, these studies suggest a good fit of the S-REF model for IGD and are in line with what has been observed for other addictive behaviors, including alcohol misuse and nicotine dependence (Caselli et al., 2016, 2018; Poormahdy et al., 2022).

In China, the application of the metacognitive tenet of psychopathology in addiction research and interventions, as outlined by Spada et al. (2015), is still in its infancy. According to two recent systematic reviews on metacognitions in addictive behaviors, there was only one Chinese study (i.e., Zhang et al., 2020) on generic metacognitions and IGD undertaken so far, compared to a total of 38 and 13 eligible studies included by Hamonniere & Varescon (2018) and Casale et al. (2021), respectively, on addictive behaviors across worldwide regions. This lack of relevant studies in China can be attributed, at least in part, to the limited number of easy-administrable and validated metacognitive assessment tools for Chinese people; hence, the present study aimed to validate MOGS (Spada & Caselli, 2017), a self-reported, specific assessment tool for metacognitions about online gaming, among Chinese gamers to address this missing link and lay the groundwork for integrating metacognition in screening and treating IGD in China.

According to the original conceptualization (Spada & Caselli, 2017), MOGS contains two primary domains: one evaluates one’s negative metacognitions about online gaming (N-MOG; i.e., beliefs about uncontrollability and dangers of online gaming) and the other gauges one’s positive metacognitions about online gaming (P-MOG; i.e., beliefs about the benefits of online gaming as a cognitive and affective self-regulation strategy). This two-factor structure has received empirical support and applications among Italian (Caselli et al., 2021; Marino et al., 2020; Spada & Caselli, 2017 [Study 1]) and Turkish (Nazligül, & Süsen, 2021) gamers. However, an alternative, three-factor model, which further divided N-MOG into two factors (i.e., uncontrollability [N-MOG1] and dangers [N-MOG2]), was proposed based on a superior model fit than the two-factor model, via confirmatory factor analysis (CFA), among Italian (Spada & Caselli, 2017 [Study 2]), Iranian (Akbari et al., 2021), and U.S. (Gandolfi et al., 2021) gamers. Given the variation in the latent structure of MOGS across cultural groups, our first objective of the present study was to confirm the latent structure of MOGS among Chinese gamers.

Despite its short history, MOGS, as a relatively new assessment tool of specific metacognitions, has displayed generally good reliability (i.e., *α* = .79 to .94) and satisfactory convergent validity with psychological distress and IGD in Italian, Turkish, Iranian, and USA gamer samples (Akbari et al., 2021; Caselli et al., 2021; Gandolfi et al., 2021; Marino et al., 2020; Nazligül, & Süsen, 2021; Spada & Caselli, 2017). As no empirical study has yet evaluated the psychometric properties of MOGS among Chinese people, we intended to begin by assessing its reliability (i.e., internal consistency and one-month test-retest reliability) and convergent validity as the second objective of the present study. As for indicators of convergent validity, we selected generic metacognitions, stress, and IGD tendency and expected to find a positive association of MOGS with these indicators similarly as those positive associations were consistently reported in extant studies (Akbari et al., 2021; Gandolfi et al., 2021; Nazligül, & Süsen, 2021; Marino et al., 2020). Compared to its generic counterpart, we would expect to find metacognitions about online gaming to be a stronger indicator of its corresponding maladaptive behavioral pattern (i.e., IGD), which would attest to the advantages of using specific measures for specific behaviorsas reported in previous studies (e.g., Chen et al., 2021a). In addition to the psychometric validation of the 12-item MOGS, the third objective of the present study was to devise a short-form MOGS as a more flexible alternative for researchers and practitioners to assess metacognitions about online gaming in mass-scale screening and/or intervention programs.

**2. Method**

*2.1 Participants and procedure*

We conducted an anonymous survey among university students in China from June to July 2021, with the inclusion criteria being of Chinese ethnicity, aged 18 years or above, and with lifetime online gaming experience (i.e., gameplay via the Internet or local area networks, including single-player or multi-player online games using mobile phones, computers, tablets, game consoles, and other devices). Every eligible participant was briefed on the study purpose and participation rights before survey administration. A total of 1340 consenting participants (59.3% female) aged from 18 to 26 years (*M* = 19.84, *SD* = 1.33), in which 91.9% (*n* = 1232) were past-year gamers, completed the survey voluntarily. To assess the test-retest reliability of MOGS, we collected a one-month follow-up sample of 262 matched participants (75.6% female, age range: 18-23 years, *M* = 20.31, *SD* = 1.24). Ethics approval was obtained from the affiliated department at the University of the corresponding author.

*2.2 Measures*

*2.2.1 Metacognitions about Online Gaming Scale (MOGS)*

The 12-item MOGS, developed by Spada and Caselli (2017), assesses positive and negative metacognitions about online gaming on a 4-point Likert scale (from 1 = *do not agree to* 4 = *agree very much*). A higher sum score indicates higher levels of positive or negative metacognitions. Following the standard translation and back-translation procedure (Brislin, 1970), we created the Chinese version of MOGS and further made minor revisions based on the pilot test results collected from 30 university students with lifetime gaming experience (e.g., we specified the term “my functioning” as “my work/study”). The psychometric properties of this Chinese version will be further explored among Chinese gamers in the following sections.

*2.2.2 Generic metacognitions*

The 30-item Chinese version of Wells et al.’s Metacognitions Questionnaire (MCQ-30; Zhang et al., 2020) was used to assess generic metacognitions from five dimensions: (a) positive beliefs about worry (e.g., “Worrying helps me cope”), (b) negative beliefs about worry (e.g., “When I start worrying, I cannot stop”), (c) cognitive confidence (e.g., “I have a poor memory”), (d) need for control (e.g., “It is bad to think certain thoughts”), (e) cognitive self-consciousness (e.g., “I think a lot about my thoughts”). Each subscale has six items and adopts a 4-point Likert scale (from 1= *do not agree* to 4= *agree very much*; *α* = .82 to .85 in present study). A higher sum score for each subscale represents higher levels of the corresponding dimension of generic metacognitions.

*2.2.3 IGD tendency and probable IGD*

The 27-item Chinese Internet Gaming Disorder Checklist (C-IGDC; Chen et al., 2020) was used to evaluate participants’ past-year IGD tendency on a 3-point Likert scale (0 = *never*, 1 = *sometimes*, and 2 = *often*; *α* = .95 in present study). Only 1232 participants with past-year gaming experience were prompted to these questions. A sample item is, “How often did you feel like losing everything while not being able to play Internet games (in the past 12 months)?” A total C-IGDC score was computed to indicate participants’ IGD tendency, with a higher summated score representing a higher tendency for IGD. In line with Chen et al.’s (2020) guidelines, we adopted the optimal screening cutoff score of ≥ 20 and identified 191 (15.5%[[1]](#footnote-1) among the past-year gamers) probable IGD gamers (1 = *probable IGD*, 0 = *non-IGD*) based on their C-IGDC sum score.

*2.2.4 Stress*

The 7-item stress subscale from the Chinese version of the Depression Anxiety Stress Scale (DASS-21; Wang et al., 2016) was used (*α* = .90 in present study). Participants rated how often they experienced stress (e.g., “I found it hard to wind down”) during the last week on a 4-point response scale (from 0 = *Did not apply to me at all* to 3 = *Applied to me very much, or most of the time*). A higher summated score indicates a higher level of stress.

*2.2.5 Demographic Information*

Sex and age were included in the questionnaire.

*2.3 Data analysis*

Only past-year gamers answered IGD tendency items; other than that, no missing data was found in the data set of the current study. In the first stage of data analysis, we intended to establish two samples for structure validity testing with the split-half method, in which the overall sample was randomly divided into Subsample A (*n* = 670) for the exploratory factor analysis (EFA) and Subsample B (*n* = 670) for CFA. In Subsample A, the appropriateness of EFA for our data was first assessed by Kaiser-Meyer-Olkin (KMO) Measure of Sampling Adequacy (KMO > .80 as meritorious; Kaiser & Rice, 1974) and Bartlett's test of sphericity (*p* < .05; Bartlett, 1950). Subsequently, we used the parallel analysis (O’Connor, 2000) to decide the number of factors to be extracted, and then performed EFA with the principal axis factoring method for extraction and the varimax method for rotation in SPSS 26.0 (IBM Corp, 2019).

In Subsample B, the sampling adequacy for CFA was determined by the N : *q* = 20:1 rule (Jackson, 2003), which suggested a desirable sample size of 540. Given that our current sample size of 670 is larger than this threshold, we proceeded to conduct CFA in Mplus 7.3 (Muthén & Muthén, 2012) with robust Maximum Likelihood (MLR) estimation to examine the conceptualized latent structure of MOGS indicated by the EFA results. The most parsimonious, one-factor model of MOGS was tested as a baseline model for comparison. The goodness of model fit was evaluated by χ2 test (*p* > .05), comparative fit index (CFI ≥ .90), root mean square error of approximation (RMSEA ≤ .08 and a nonsignificant p-close value), and standardized root mean square residual (SRMR ≤ .08; Bentler, 1990; Browne & Cudeck, 1993; Kline, 2016). A smaller Akaike information criteria (AIC) value was preferred for model comparisons (Kline, 2016). The standardized factor loading estimates were expected to be .05 or higher (Hair, 2014). On the factor level, we further examined the average variance extracted (AVE ≥ .50) and the composite reliability (ρc ≥ .70) as CFA-derived reliability measures (Hair et al., 2014). Additional measurement invariance was explored with multi-group CFA to assess the invariance of MOGS across both sexes on configural, metric, and scalar levels. A noninvariance was indicated with ΔCFI ≥ .01 and ΔRMSEA ≥ .015 (F. F. Chen, 2007).

After confirming the factor structure and items of the Chinese version of MOGS, we performed all the remaining statistical tests in SPSS 26.0 to evaluate its other psychometric properties in the overall sample. We started with assessing the internal consistency of MOGS by Cronbach’s α (> .70 as acceptable and > .80 as good; Cortina, 1993) and its one-month test-retest reliability by intraclass correlation coefficient (ICC > 0.50 as acceptable; Koo & Li, 2016). The convergent validity was tested by bivariate correlations of MOGS with generic metacognitions, stress, and IGD tendency (*r* = .10, .30, and .50 corresponding to small, medium, and large effect sizes; Cohen, 1988), as well as *t*-tests for detecting hypothesized, significant differences of MOGS scores between probable IGD group and non-IGD group. Additionally, we followed the procedure recommended by Stanton et al. (2002) and created a short-form MOGS.

**3. Results**

*3.1 Factor analysis of the full-scale MOGS*

In Subsample A, the appropriateness of EFA for the 12-item MOGS with our data was established by a KMO of .93 and a significant Bartlett's test of sphericity (*p* < .001). Subsequently, we extracted two factors from the 12-item MOGS with EFA as indicated by the parallel analysis. As shown in Table 1 (12-item version), 11 of the 12 items were dominantly loaded on one factor with a standardized factor loading greater than .40. However, Item 7 displayed a double-loading (.52 and .60) on both factors, and hence was excluded for lacking distinctiveness of factor loading with one specific factor. After the removal of Item 7, we re-performed the parallel analysis and EFA on the remaining 11 items (i.e., revised MOGS) based on a KMO of .92 and a significant Bartlett's test of sphericity (*p* < .001). This revised 11-item MOGS demonstrated adequate factor loading and distinctiveness of factor loading, with six items dominantly loaded on one factor (.57 to .86) and five items on the other (.68 to .85; see Table 1, 11-item version). We followed the original factor naming of MOGS for the two extracted factors because all these 11 items were grouped in the same way as Spada and Caselli (2017) initially proposed during the development of MOGS.

Based on the EFA results, we proceeded to test the fitness between the 11-item, two-factor revised MOGS and our data in Subsample B. The data fitted the two-factor model fairly (χ2(43) = 344.10, *p* < .001, CFI = .90, RMSEA = .10, 90% CI [.09, .11], p-close < .001, SRMR= .06, AIC = 11367.08) and outperformed the baseline one-factor model (χ2(44) = 1206.23, *p* < .001, CFI = .62, RMSEA = .20, 90% CI [.19, .21], p-close < .001, SRMR = .15, AIC = 12784.27). As the two-factor model appeared to be the superior model, we further included three pairs of within-factor residual covariances (Item 8 with 9, Item 5 with 6, and Item 3 with 4) and enhanced the model fit to a satisfactory level (χ2(40) = 130.70, *p* < .001, CFI = .97, RMSEA = .06, 90% CI [.05, .07], p-close = .11, SRMR= .05, AIC = 11043.48). The standardized factor loadings of the 11 items in this two-factor model ranged from .65 to .90, exceeding the minimally acceptable threshold of .50. This 11-item, two-factor revised MOGS, as shown in Figure 1, is the final latent structure for the Chinese version of MOGS, which also demonstrated satisfactory CFA-based reliability (AVE: F1 = .56, F2 = .70; ρc:F1 = .88, F2 = .92).

Furthermore, we conducted multi-group CFA to assess the invariance of MOGS across both sexes on configural, metric, and scalar levels. Table 2 demonstrated a good model fit of the 11-item Chinese version of MOGS on the configural, metric, and scalar levels (CFI = .962 to .966, RMSEA = .059 to .063, SRMR = .50 to .54). Further comparisons among the three levels suggested the measurement invariance was held at the scalar level between both sex groups (metric versus configural: ΔCFI = .001, ΔRMSEA = .004; scalar versus metric: ΔCFI = .005, ΔRMSEA = .000).

*3.2 Internal consistency, test-retest reliability, and validity*

The Chinese version of MOGS displayed high internal consistency (*α* = .90 for N-MOG and .92 for P-MOG) and acceptable one-month test-retest reliability (ICC = .60 for N-MOG and .64 for P-MOG). As shown in Table 3, both N-MOG and P-MOG showed a good convergent validity in terms of significant, positive correlations with the five generic metacognitive factors (*r* (1338) = .36 to .48, *p* < .001), stress (*r* (1338) = .46 and .29, respectively, *p* < .001), and IGD tendency (*r* (1338) = .55 and .47, respectively, *p* < .001). It is also worth noting that both MOGS factors displayed stronger correlations with IGD tendency (*r* (1338) = .47 to .55, *p* < .001) than the five generic metacognitive factors[[2]](#footnote-2) (*r* (1338) = .31 to .37, *p* < .001). Moreover, there was an expected and significant, positive correlation between N-MOG and P-MOG (*r* (1338)= .53, *p* < .001).

When comparing the probable IGD group with the non-IGD one, we further found that probable IGD gamers scored significantly higher in both N-MOG and P-MOG than their non-IGD counterparts (N-MOG: *M*IGD = 12.31, *M*non-IGD = 8.15, *t* (1190) = -15.12, *p* < .001; P-MOG: *M*IGD = 10.67, *M*non-IGD = 8.40, *t* (169.18) = -8.11, *p* < .001), lending extra support to the good convergent validity of MOGS as well as its discriminant power to differentiate the probable IGD and non-IGD gamers.

*3.3 The short-form MOGS*

According to the guidelines recommended by Stanton et al. (2002), we mainly considered the equivalence between the short form and its full-scale form in terms of latent structure, reliability, validity, and coverage. Given a minimum of three items is required for constructing each latent factor (Hair, 2014), we created two possible short forms (i.e., 8-item and 6-item) for comparison by selecting items bearing the highest factor loading from each factor (e.g., the top three items with the highest factor loadings from the two factors for the 6-item version; for a discussion on this method, see Marsh et al., 2005). When fitting the conceptualized two-factor model with the data, we found the 6-item version (Items 2, 3, 4, 9, 11, and 12) manifested excellent goodness of model fit (χ2(8) = 17.64, *p* = .02, CFI = .99, RMSEA = .04, 90% CI [.02, .07], p-close = .64, SRMR= .02, AIC = 6307.71; see Figure 2 for the standardized factor loadings), which significantly outperformed the 8-item versions (χ2(19) = 79.77, *p* < .001, CFI = .97, RMSEA = .07, 90% CI [.05, .09], p-close = .02, SRMR= .05, AIC = 8170.77) and its alternative one-factor baseline model (χ2(9) = 602.95, *p* < .001, CFI = .55, RMSEA = .31, 90% CI [.29, .34], p-close < .001, SRMR= .18, AIC = 7307.12).

This 6-item, short-form MOGS also displayed good CFA-based reliability (AVE: F1 = .74, F2 = .74; ρc: F1 = .89, F2 = .89) and supported measurement invariance at the scalar level for both sex groups (metric versus configural: ΔCFI = .002, ΔRMSEA = .006; scalar versus metric: ΔCFI = .002, ΔRMSEA = .006; see Table 2). Subsequent analyses also revealed this short form had equivalent reliability (*α* = .89 and .89 and ICC = .52 and .61), convergent validity (*r* (1338) = .27 to .47, *p* < .001; see Table 3; also significant differences in MOGS scores for probable IGD and non-IGD cases, *p* < .001) and coverage (i.e., endorsed by an invited panel of experts in the related fields) as its full-scale version.

**4. Discussion**

To address the missing link of lacking a proper assessment tool to assess metacognitions about online gaming among Chinese gamers, the present study made the first attempt at validating MOGS with a sample of Chinese lifetime gamers recruited from universities in China. With respect to our first objective, regarding the factor structure of MOGS in Chinese gamers, results provided additional evidence for rejecting the one-factor model that displayed an unacceptable model fit, indicating the multi-dimensionality of MOGS, in line with all previous validation studies on MOGS (Akbari et al., 2021; Gandolfi et al., 2021; Nazligül, & Süsen, 2021; Marino et al., 2020; Spada & Caselli, 2017). As Item 7 lacked the distinctiveness of factor loading with one specific factor, we removed it from the Chinese version of MOGS, and then identified the two-factor model of MOGS as that was initially proposed by Spada and Caselli (2017) with the remaining 11 items. This 11-item, two-factor Chinese version of MOGS displayed not only satisfactory structure validity in both EFA and CFA but also scalar measurement invariance between both sex groups. Nevertheless, we would still like to call for more validations with data collected from multiple sources and longitudinal waves to confirm the robustness of this 11-item, two-factor model structure in Chinese populations.

With respect to the second objective, we confirmed the good reliability and validity of the revised 11-item MOGS among Chinese gamers. The satisfactory internal consistency of MOGS (i.e., *α* = .90 and .92) in the current study was consistent with previous studies (e.g., Akbari et al., 2021; Caselli et al., 2021), whereas our finding regarding its acceptable one-month test-retest reliability (i.e., .60 and .64) was comparable to those found in Turkish gamers (i.e., .58 and .72; Nazligül, & Süsen, 2021). The CFA-based reliability measures, AVE (i.e., .56 and .70) and composite reliability (i.e., .88 and .92), also supported the satisfactory reliability of the Chinese version of MOGS on the latent factor level. For convergent validity, the two types of specific metacognitions about online gaming (i.e., N-MOG and P-MOG), measured by MOGS, demonstrated significant, positive correlations with the five dimensions of generic metacognitions (*r* (1338) = .36 to .48, *p* < .001); these medium-sized correlation coefficients, taken together with the additional EFA results on MOGS and the generic metacognition items, indicated that specific metacognitions were sufficiently distinct from generic metacognitions albeit possessing a potentially common root. In line with previous studies on generic metacognitions (e.g., Akbari et al., 2021; Capobianco et al., 2018; Spada et al., 2008; Zhang et al., 2020), these two MOGS constructs also demonstrated significant, positive correlations with both stress and IGD tendency among Chinese gamers. In specific, IGD tendency displayed a generally greater effect size (i.e., from marginally large to large) in correlation with N-MOG and P-MOG than those with generic metacognitions (i.e., medium); this difference in effect size implied that specific metacognitions about online gaming appear to be a better indicator of IGD tendency than the generic metacognitions, underscoring the good convergent validity of the Chinese version of MOGS. It is not surprising that MOGS is a superior indicator of IGD tendency than the generic metacognitions because past literature has consistently revealed that specific and targeted assessment tools measures have better predictive power than the general ones (e.g., Chen et al., 2021a; Rydell et al., 2007). Furthermore, our finding on the significant differences in both N-MOG and P-MOG between probable IGD and non-IGD gamers suggested a promising potential of including intervention strategies, such as metacognitive therapy (Caselli et al., 2018), targeting on maladaptive metacognitions about online gaming in various levels of IGD presentations.

With respect to the third objective, we created a 6-item, short-form MOGS to provide researchers and practitioners with more versatility to measure metacognitions about online gaming for individualized needs. This short-form MOGS exhibited an excellent model fit with satisfactorily high factor loadings and scalar measurement invariance for both sex groups; it also performed in a similar way as its full-scale version, as evidenced by equivalent and adequate reliability (i.e., internal consistency, test-retest reliability, AVE, and composite reliability), convergent validity, coverage, and discriminant power to differentiate probable IGD and non-IGD gamers. Therefore, this short form could be considered as an alternative for measuring metacognitions about online gaming to reduce the burden on both researchers with limited resources (Richins, 2014) and respondents (Wiriyakijja et al., 2020).

In addition, we observed male gamers tended to report a higher level of IGD tendency than their female counterparts (*r* (1338) = .26, *p* < .001), which is consistent with recent IGD studies in China (e.g., Chen et al., 2021b; Long et al., 2018) and mainstream studies worldwide (Mihara & Higuchi, 2017; Stevens et al., 2021). A similar, elevated level in N-MOG and P-MOG, of both full-scale and short-form versions, was identified among male gamers than female ones (*r* (1338) = .19 to .26, *p* < .001). This pattern indicates that male gamers tended to have significantly more negative and positive metacognitions about online gaming than females, which is presumably associated with their considerably more gameplay involvement and hence greater risk for IGD. Such sex difference in N-MOG and P-MOG has not been fully studied in previous studies, partly due to the limited proportion of female gamers recruited (e.g., 5.3-25%; Akbari et al., 2021; Caselli et al., 2021; Gandolfi et al., 2021; Spada & Caselli, 2017); thus, we called for future studies to further explore such sex difference in metacognitions about online gaming and its potential application for IGD assessment and intervention.

The current study is also limited in three aspects. First, data was only collected on a sample of Chinese university students, in part limiting the generalizability of the results, future studies on a general population or other age groups (e.g., adolescents or older adults) are needed. Second, although the 6-item, short-form MOGS exhibited excellent goodness of model fit and good reliability and validity, this version still needs further validation in other samples, and hence we call for more replication studies to evaluate the robustness of its psychometric properties (e.g., predictive validity) across samples. Finally, the cross-sectional design adopted by the present study cannot track the changes in metacognitions about online gaming with stress and IGD tendency over time to evaluate the predictive validity in a natural setting, which could be addressed by a longitudinal design in future studies.

Despite these limitations, this study is the first to validate MOGS among Chinese gamers. Our findings supported the two-factor structure for the revised 11-item MOGS and confirmed its satisfactory reliability and validity for measuring Chinese gamers’ metacognitions about online gaming. To further promote the use of MOGS for various settings and purposes, we also created a 6-item, short-form MOGS and found it to display equivalent psychometric soundness as its full-scale version. We envisioned this Chinese set of MOGS, as a gaming-specific metacognitions measurement tool, would lay the foundation for studying metacognitions (e.g., explore the role of repetitive negative thinking in the association between gaming-specific metacognitions and IGD) among Chinese gamers and inspire more metacognition-based applications for screening high-risk gamers and as well as provide more cost-effective interventions to safeguard a healthy development of the largest gamer population in the world.

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**Table 1**

Item characteristics and exploratory factor analyses of MOGS items for the 12-item and 11-item versions based on Subsample A (*n* = 670).

|  |  |  |  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- |
|  | Skew-ness | Kur-tosis | **12-item version** | | | | **11-item version** | | | |
| F1a | F2a | Commu-  nality | Corrected item-total correlationb | F1a | F2a | Commu-  nality | Corrected item-total correlationb |
| **Factor 1: negative metacognition about online gaming (N-MOG)** |  |  |  |  |  |  |  |  |  |  |
| 1. I continue to play despite I think it would be better to stop | 0.85 | 0.31 | **.57** | .30 | .41 | 0.61 | **.57** | .30 | .41 | 0.61 |
| 2. I have no control over how much time I play | 1.29 | 1.02 | **.79** | .27 | .69 | 0.80 | **.79** | .27 | .70 | 0.80 |
| 3. Once I start online gaming, I cannot stop | 1.80 | 2.46 | **.86** | .19 | .77 | 0.81 | **.86** | .19 | .77 | 0.81 |
| 4. Online gaming makes me lose control | 1.87 | 2.97 | **.83** | .20 | .74 | 0.79 | **.83** | .20 | .74 | 0.79 |
| 5. Thoughts about online gaming interfere with my functioning | 1.24 | 0.86 | **.73** | .29 | .61 | 0.75 | **.73** | .28 | .61 | 0.75 |
| 6. Thoughts about online gaming are becoming an obsession | 1.73 | 2.73 | **.75** | .29 | .64 | 0.75 | **.75** | .28 | .64 | 0.75 |
| **Factor 2: positive metacognition about online gaming (P-MOG)** |  |  |  |  |  |  |  |  |  |  |
| 7. Online gaming makes my worries more bearable | 1.12 | 0.74 | **.52** | **.60** | .63 | 0.70 | -- | -- | -- | -- |
| 8. Online gaming reduces my negative feelings | 0.63 | -0.21 | .23 | **.80** | .69 | 0.79 | .23 | **.78** | .67 | 0.78 |
| 9. Online gaming helps me to control my negative thoughts | 0.70 | -0.22 | .28 | **.83** | .76 | 0.83 | .29 | **.83** | .76 | 0.83 |
| 10. Online gaming stops me from worrying | 1.05 | 0.30 | .38 | **.69** | .61 | 0.74 | .38 | **.68** | .61 | 0.73 |
| 11. Online gaming reduces my anxious feelings | 0.70 | -0.23 | .20 | **.84** | .75 | 0.81 | .21 | **.85** | .76 | 0.83 |
| 12. Online gaming distracts my mind from problems | 0.66 | -0.08 | .23 | **.79** | .67 | 0.77 | .23 | **.79** | .67 | 0.77 |

*Note:* a Standardized factor loadings were reported. b Corrected item-total correlation was tested on the factor basis and within each factor. The bold values indicate item loading.

**Table 2**

Measurement invariance of the full-scale MOGS between both sex groups based on Subsample B (*n* = 670).

|  |  |  |  |  |  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- |
| Model | | Model Fit | | | | | | Model Comparison | | | | |
| χ2 | *df* | *p* | CFI | RMSEA [90% CI] | SRMR | Δχ2\* | Δ*df* | *p* | ΔCFI | ΔRMSEA |
| **Full-scale MOGS** | |  |  |  |  |  |  |  |  |  |  |  |
|  | Configural | 185.49 | 80 | <.001 | .966 | .063 [.051, .075] | .050 |  |  |  |  |  |
|  | Metric | 191.21 | 89 | <.001 | .967 | .059 [.047, .070] | .053 | 4.84 | 9 | .85 | .001 | .004 |
|  | Scalar | 214.02 | 98 | <.001 | .962 | .059 [.049, .070] | .054 | 24.05 | 9 | .004 | .005 | .000 |
| **Short-form MOGS** | |  |  |  |  |  |  |  |  |  |  |  |
|  | Configural | 30.73 | 16 | .02 | .989 | .052 [.023, .080] | .023 |  |  |  |  |  |
|  | Metric | 37.02 | 20 | .01 | .987 | .050 [.023, .075] | .036 | 6.42 | 4 | 0.17 | .002 | .002 |
|  | Scalar | 48.80 | 24 | .002 | .981 | .056 [.033, .078] | .033 | 13.13 | 4 | 0.01 | .006 | .006 |
|  |  |  |  |  |  |  |  |  |  |  |  |  |

*Note*. \* Chi-square difference testing was carried out using scaling correction factor for robust Maximum Likelihood (MLR) estimation.

**Table 3**

Descriptive statistics and bivariate correlations key constructs (*N* = 1340).

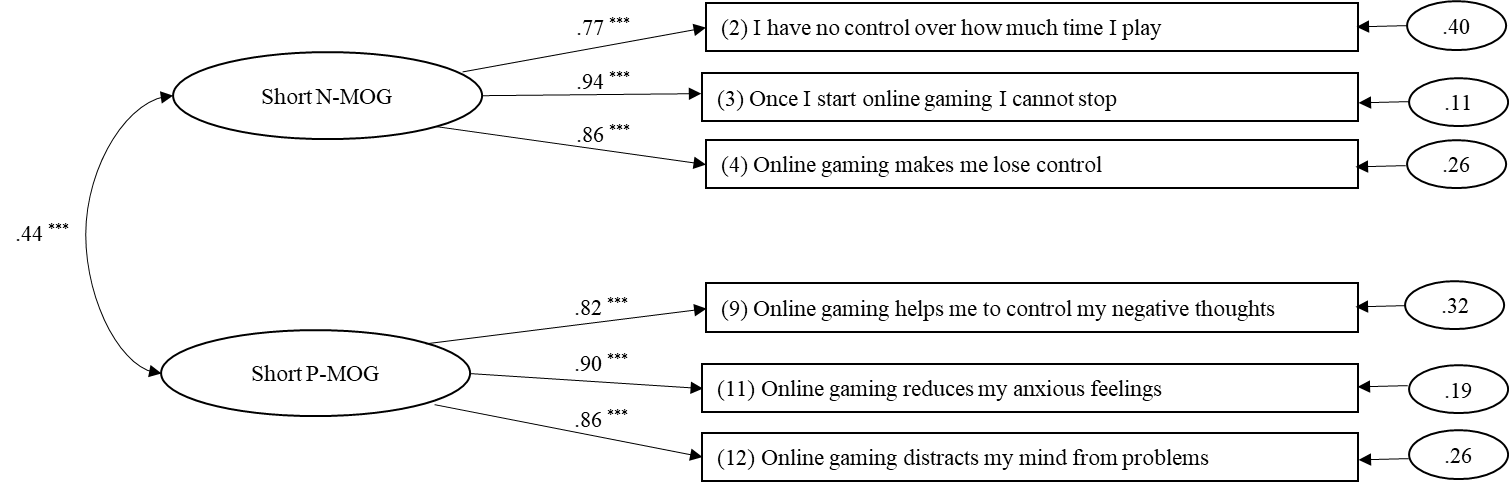
|  |  |  |  |  |  |  |  |  |  |  |  |  |  |  |
| --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- | --- |
|  | *M* | *SD* | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 | 9 | 10 | 11 | 12 |
| 1. N-MOG | 8.79 | 3.38 | (.90) |  |  |  |  |  |  |  |  |  |  |  |
| 2. P-MOG | 8.70 | 3.35 | .53\*\*\* | (.92) |  |  |  |  |  |  |  |  |  |  |
| 3. Short N-MOG | 4.21 | 1.86 | .94\*\*\* | .47\*\*\* | (.89) |  |  |  |  |  |  |  |  |  |
| 4. Short P-MOG | 5.33 | 2.11 | .49\*\*\* | .98\*\*\* | .43\*\*\* | (.90) |  |  |  |  |  |  |  |  |
| 5. POSa | 11.10 | 3.30 | .46\*\*\* | .39\*\*\* | .44\*\*\* | .36\*\*\* | (.85) |  |  |  |  |  |  |  |
| 6. NEGa | 11.90 | 3.42 | .41\*\*\* | .40\*\*\* | .36\*\*\* | .38\*\*\* | .76\*\*\* | (.84) |  |  |  |  |  |  |
| 7. CCa | 11.16 | 3.22 | .47\*\*\* | .36\*\*\* | .44\*\*\* | .34\*\*\* | .84\*\*\* | .75\*\*\* | (.85) |  |  |  |  |  |
| 8. NCa | 11.47 | 3.41 | .48\*\*\* | .40\*\*\* | .46\*\*\* | .38\*\*\* | .81\*\*\* | .81\*\*\* | .83\*\*\* | (.82) |  |  |  |  |
| 9. CSCa | 12.34 | 3.15 | .37\*\*\* | .36\*\*\* | .33\*\*\* | .35\*\*\* | .73\*\*\* | .81\*\*\* | .74\*\*\* | .78\*\*\* | (.84) |  |  |  |
| 10. Stress | 4.96 | 4.56 | .46\*\*\* | .29\*\*\* | .45\*\*\* | .27\*\*\* | .50\*\*\* | .47\*\*\* | .51\*\*\* | .52\*\*\* | .41\*\*\* | (.90) |  |  |
| 11. IGD tendencyb | 9.57 | 8.85 | .55\*\*\* | .47\*\*\* | .47\*\*\* | .46\*\*\* | .35\*\*\* | .36\*\*\* | .36\*\*\* | .37\*\*\* | .31\*\*\* | .38\*\*\* | (.95) |  |
| 12. Sexc | -- | -- | .26\*\* | .19\*\*\* | .23\*\*\* | .19\*\*\* | .18\*\*\* | .09\*\*\* | .18\*\*\* | .14\*\*\* | .07\* | .26\*\*\* | .31\*\*\* | (--) |
| 13. Age | 19.84 | 1.33 | .01 | -.01 | .00 | .00 | -.07\*\* | -.04 | -.04 | -.06\* | -.05 | -.07\* | -.00 | -.03 |

*Note.* N-MOG = Negative metacognitions about online gaming; P-MOG = Positive metacognitions about online gaming; POS= Positive beliefs about worry; NEG = Negative beliefs about worry; CC = Cognitive confidence; NC= Need for control; CSC= Cognitive self-consciousness; IGD = Internet Gaming Disorder. a the five generic metacognitions from the Metacognitions Questionnaire (MCQ-30). Cronbach’s α was reported in parentheses along the diagonal line. b only 1232 past-year gamers responded to items of this construct. c Female = 0, Male = 1. \*\*\**p*< .001; \*\* *p*< .01; \* *p*<.05.

图形用户界面, 文本

描述已自动生成

**Figure 1**. The latent two-factor structure of the 11-item, full-scale Chinese version of MOGS based on Subsample B (*n* = 670). Standardized factor loadings were reported. N-MOG = negative metacognition about online gaming, P-MOG = positive metacognition about online gaming.



**Figure 2**. The latent two-factor structure of the 6-item, short-form Chinese version of MOGS based on Subsample B (*n* = 670). Standardized factor loadings were reported. N-MOG = negative metacognition about online gaming, P-MOG = positive metacognition about online gaming.

1. This proportion of probable IGD gamers falls within the range of problematic online gamer (POG) estimation (i.e., 3.5-17.0%) revealed in a recent systematic review of POG studies in China (Long et al., 2018). [↑](#footnote-ref-1)
2. Additional comparisons between MOGS and the generic metacognition were conducted by EFA, in which all the 11 MOGS items and 30 items generic metacognition were put together for factor extraction with Subsample A’s data. Following the extracted factor number suggested by the parallel analysis, the EFA successfully extracted two MOGS factors, N-MOG and P-MOG, as expected, in addition to generic metacognition factors, indicating the differences among MOGS and the generic metacognition items on the latent factor level. [↑](#footnote-ref-2)