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Cross-Cultural Validation of the Malevolent Creativity Behavior Scale in 7 Countries

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ABSTRACT

This study examined the psychometric properties and cross-cultural validity of the Malevolent Creativity Behavior Scale (MCBS). A total of 2937 participants from Brazil, China, Ghana, Kenya, Peru, the United Kingdom, and the United States completed the 13-item MCBS. Confirmatory factor analyses compared multiple factor structures (unidimensional, three-factor, hierarchical, and bifactor), and measurement invariance was tested both across countries and by sex. The original three-factor solution demonstrated a generally acceptable fit. The measurement invariance findings indicated that the MCBS retains stable thresholds and factor loadings across groups, supporting the meaningfulness of comparisons. No significant item bias emerged by sex. However, most MCBS items do not reference novelty, a defining feature of creativity, posing the concern that the MCBS focuses more on malevolent ideation or antagonistic behaviors rather than creative malevolent processes. Overall, the results underscore the MCBS as a reliable tool for measuring harmful and creative behaviors in diverse cultural and demographic contexts. These findings contribute to the growing understanding of how malevolent creativity manifests and can be measured worldwide.

Creativity, a behavioral characteristic, is widely valued for its ability to drive progress, solve problems in innovative ways, find alternatives when something does not work, invent practical solutions, and transform ideas into real improvements (Runco and Jaeger 2012). However, this same potential can also manifest itself in aversive ways, which has been referred to as malevolent creativity, understood as the deliberate application of novel ideas for the purpose of causing harm to others (Cropley et al. 2008). In adulthood, this type of creativity is reflected in covert defamation strategies to damage someone's reputation (Kapoor and Khan 2016), maneuvers to hinder a peer's work (Harris and Reiter-Palmon 2015), deceptions for personal gain, or the strategic use of information to discredit or pressure others in social and professional settings (James et al. 1999).

The original definition proposed by Cropley et al. (2008) emphasizes intentionality as the central feature of malevolent creativity. Nevertheless, subsequent studies have introduced variations, in some cases broadening the phenomenon to include behaviors aimed at personal gain without necessarily involving deliberate harm to others (e.g., Hao et al. 2016). Such formulations, while relevant for understanding manipulative or deceptive behaviors, do not always strictly align with the original conceptualization. In the present study, we adopt Cropley et al.'s (2008) and Perchtold-Stefan et al.'s (2023) definition as our main framework. This formulation emphasizes both the originality of ideas and the deliberate motivation to inflict harm, unifying perspectives in the literature that have sometimes described the construct in slightly different ways.

Building on this conceptual foundation, malevolent creativity has been documented as a concrete and observable phenomenon in different areas, where it is expressed in subtle or obvious ways for harmful purposes toward others (Cropley et al. 2008). Over the years, there has been discussion about whether its appearance is always intentional or whether, in some cases, it arises as a not entirely premeditated consequence of human actions (James et al. 2010). In the work context, for example, it can manifest itself as strategic sabotage when an individual manipulates information or discredits a colleague to obtain personal advantages (Harris and Reiter-Palmon 2015). In the digital environment, it is observed in the planned dissemination of disinformation to influence public opinion or alter electoral processes (de Saint-Laurent et al. 2020; Nguyen et al. 2025). In the interpersonal sphere, one of its most studied expressions is gaslighting, a type of psychological manipulation in which a person intentionally distorts another's perception of reality to exert control or personal benefit (Sweet 2019).

This ability to devise novel ways to deceive, manipulate, or harm underlines the phenomenon's complexity. Studies such as Hao et al. (2016) have identified similar manipulative behaviors in everyday life, while Harris and Reiter-Palmon (2015) highlight that premeditation and aggression are key elements in generating creative ideas for destructive purposes. In any context where it manifests itself, its effects are consistent: it not only harms the direct victims but also undermines trust in the environment, generates conflicts, and affects collective and organizational well-being (Cropley 2010; Fousiani et al. 2025). One of the contexts where it also manifests itself is social media and the internet, linked to cyberviolence and cyberbullying, which has even allowed the development of the concept in this field called cyber malevolent creativity (Liu et al. 2024).

Malevolent creativity (Cropley et al. 2008) has a significant impact that transcends everyday examples and extends to various areas of human life. In the field of education, for example, this phenomenon can manifest itself in behaviors such as plagiarism or sabotage among students to gain academic advantages, which not only affects individual learning but also compromises the integrity of the educational system (Kapoor 2019). In the realm of politics, malevolent creativity is evident in the development of disinformation campaigns capable of influencing elections or polarizing societies, as documented in the use of social media to spread anti-immigrant narratives (de Saint-Laurent et al. 2020).

Malevolent creativity, as a psychological variable, is defined as the ability to generate original and effective ideas with the deliberate intention of causing physical, psychological, or social harm (Perchtold-Stefan et al. 2023b). Rather than being an impulsive or unreflective reaction, malevolent creativity involves a deliberate process. It requires cognitive flexibility, understood as the ability to generate multiple solutions or approaches, combined with malicious motivation, the purposeful intent to harm others (Perchtold-Stefan, Fink, et al. 2021; Lee and Dow 2011). This differentiates it from reactive aggression, an instinctive behavior triggered by an immediate stimulus without prior elaboration (Geen 2001). Malevolent creativity, on the other hand, requires premeditation, intentional design of a harmful strategy, and a novelty component, which makes it unpredictable and sophisticated (Gino and Ariely 2012).

Malevolent creativity can be understood through three distinct manifestations: “harming others”, “lying”, and “playing tricks” (Hao et al. 2016). “Harming others” involves causing physical, emotional, or social harm through creative planning that maximizes the negative impact on the victim. “Lying” refers to the construction of original and clever falsehoods for the purposes of deception or manipulation and demonstrates an ability to craft novel narratives with harmful intentions. “Playing tricks” encompasses playful but harmful strategies, such as pranks or deceptions designed to humiliate or destabilize, where creativity is employed for the purposes of malicious entertainment or covert revenge (Hao et al. 2016). The MCBS was developed to measure self-reported behaviors associated with malevolent creativity (Hao et al. 2016). However, because these reports also show how people imagine and admit to engaging in such acts, the instrument indirectly explores the underlying psychological traits.

The empirical measurement of this construct has been validated in various contexts in Europe, South America, and Asia (Farias et al. 2025; Fousiani et al. 2025; Osman 2024). Some studies have revealed the connections between the Malevolent Creativity Behavior Scale and related psychological variables, including conscientiousness (Lee and Dow 2011), less premeditation (Harris and Reiter-Palmon 2015), high need for cognition (Baas et al. 2019), low emotional intelligence (Harris et al. 2013), antagonist personality (Perchtold-Stefan, Rominger, et al. 2021), anger (Cheng et al. 2021; Jiaq et al. 2024; Li et al. 2022), social exclusion (Perchtold-Stefan et al. 2024), the Dark Triad (Jia et al. 2020) and the Dark Tetrad (Dow 2023), approach motivation (Hao et al. 2020), perceived unfairness, moral disengagement, aggression (Zhang et al. 2024), the Light Triad (Malik et al. 2020), schizotypal personality traits and cognitive disorganization (Perchtold-Stefan, Rominger, et al. 2021). Other life experiences like abusive supervision (Malik et al. 2020), childhood trauma (Li et al. 2022), and early life adversity (Geng et al. 2024) can also influence the malevolent creativity.

Hao et al.'s (2016) Malevolent Creativity Behavior Scale (MCBS) stands out as the most widely used instrument to assess the frequency of malevolent creative acts in daily life. Initially, an initial set of 20 items was generated that reflected behaviors such as cheating, revenge, and rumors. These items were evaluated by four judges, who identified seven redundant items, reducing the scale to 13 final items applied to 908 Chinese university students. The psychometric validation process was divided into two stages. First, an exploratory factor analysis (EFA) was performed with a subsample of 454 participants, where three factors were identified: “harming others” (6 items), “lying” (4 items), and “playing tricks” (3 items), explaining 55.88% of the total variance, and factor loadings higher than 0.40. Second, a confirmatory factor analysis (CFA) in the other subsample of 454 participants confirmed this three-factor structure, with fit indices meeting psychometric criteria and interfactor correlations ranging from 0.69 to 0.85. Internal reliability was satisfactory for the total scale (Cronbach's $\alpha = 0.80$) and the subscales ("harming others", $\alpha = 0.80$; "lying", $\alpha = 0.76$; "playing tricks", $\alpha = 0.61$), and correlational validity indicating positive associations with aggression, openness, and extraversion.

This scale has been subject to subsequent adaptations and validations that provide data on its psychometric performance in other settings. In Sudan, Al-Mahdawi et al. (2022) translated the MCBS into Arabic and administered it to 1619 university students, using EFA to verify its structure. They first considered the item-test correlation of the instrument and decided to eliminate items 4 and 5, and considered only 11 items for the respective analyses. The results confirmed all three factors, albeit with lower explained variance (47%), reliability values were higher than 0.70 on the overall scale and at the dimension level, except for "harming others" ($\alpha = 0.55$). Additionally, they found that women scored slightly higher on all three dimensions; however, in this study, factorial invariance analysis was not performed based on men and women. The low consistency in "harming others" suggests that the items may not adequately capture the construct in this context, possibly due to cultural norms that inhibit reporting aggressive behaviors, while the lack of CFA and a detailed cultural adaptation process compromises the validity of this version compared to the Chinese original.

In Turkey, Akben et al. (2024) adapted the MCBS in a sample of 555 university students. They used CFA to validate the 13-item trifactor structure, obtaining acceptable fit indices, and assessed criterion validity with significant correlations with the three traits of the Dark Triad, unstable emotionality, and conscientiousness. Adequate reliability values were presented in the general MCBS ($\alpha = 0.84$), including the subscales "harming others" (0.76) and "lying" (0.81), unlike the "playing tricks" dimension ($\alpha = 0.51$). A test–retest with 49 participants after 15 days showed significant correlations for the total scale ($r = 0.62$) and the subscales ("harming others", $r = 0.57$; "lying", $r = 0.55$; "playing tricks", r = 0.67).

In Iran, Nahavandi et al. (2025) adapted the MCBS in a sample of 393 students, selected by cluster random sampling. They applied CFA to validate the 13-item three-factor structure with adequate fit indices and factor loadings between 0.42 and 0.87, and moderate correlations between dimensions of 0.44 and 0.52, which allows for confirming the internal structure according to psychometric criteria. In addition, adequate reliability values of $\alpha = 0.83$ were reported for the total scale, with $\alpha = 0.79$ for "harming others", $\alpha = 0.82$ for "lying", and $\alpha = 0.68$ for "playing tricks". Significant correlations were reported with creative thinking ($r = 0.24$) and aggression (r = 0.50).

Several studies have reported differences in malevolent creativity according to various sociodemographic aspects. In a study in China, Nie (2024) assessed malevolent creativity in 310 university students and found a significant correlation between sex and malevolent creativity, with males scoring higher than females. In Sudan, Osman (2024) considered a sample of 579 students from a public university. Significantly higher scores of malevolent creativity were reported in sociodemographic aspects toward the younger age group and lower academic level, and significant differences were only presented in the dimension of "harming others" with respect to sex (males). Another study in this context (Sudan) by Al-Mahdawi et al. (2022), on the contrary, reports significant differences in MCBS scores among the group of females. While other research on American participants indicates that differences were not significant regarding age and sex (Ceballos and Watt 2023), a lack of confirmation of differences has been reported even in scores on malevolent creativity tasks (Dumas and Strickland 2018; Perchtold-Stefan et al. 2023a).

Despite its promise as a psychometric instrument, the MCBS faces challenges that question its robustness outside the original Chinese context. Variability in reliability, especially on the "harming others" and "playing tricks" dimension, and inconsistencies in reported sex differences in females in Sudan and males in China suggest potential measurement biases or unaddressed cultural influences. Furthermore, the "playing tricks" dimension may overlap with mischievous or playful behaviors rather than intentional harm, and responses to the scale may reflect a propensity toward malevolent ideation rather than enacted malevolent creative acts (Waldie et al. 2021). Such critiques highlight the need to clarify whether the MCBS actually measures malevolent creativity, or merely the frequency of malevolent (but not creative) ideation. Moreover, the wording of the MCBS items provides fewer cues about creativity, explicitly present in only two items (i.e., "How often do you have ideas about new ways to punish people" and "How often do you engage in an original form of sabotage"), thereby posing doubts about whether it captures malevolent creativity or the frequency of harmful thoughts. Replication of the three-factor structure across countries is encouraging, but the absence of factor equivalence analyses across sexes and cultures, along with a lack of transparency in adaptations, compromises its utility for valid comparisons.

Despite reporting few validations of the instrument, the MBCS has been used in various contexts and languages other than English or Chinese, which have only reported reliability and factor-loading findings at the item level. Such as the study by Khorakian et al. (2020) on Muslim university students, while several studies applied to adults in Europe only reported internal consistency findings (Klomp 2024; Perchtold-Stefan, Fink, et al. 2021; Silva et al. 2025). Other research has not reported any reliability or internal validity values for the MCBS (Szabó et al. 2022; Ceballos and Watt 2023). Furthermore, while the construct of malevolent creativity is defined as the production of novel ideas intended to cause harm (Cropley et al. 2008; Perchtold-Stefan et al. 2023a), studies have broadened or blurred this definition, including manipulative or mischievous behaviors that do not strictly align with intentional harm (e.g., Hao et al. 2016). This variability raises conceptual questions about whether the MCBS captures malevolent creativity as initially defined, or somewhat related tendencies, such as everyday deviance or malevolent ideation.

This study aims, first, to describe the cross-cultural validation of the Malevolent Creativity Behavior Scale (MCBS).

Second, it aims to examine its underlying factor structure, test its equivalence across sex and culture, and investigate its psychometric properties in a multinational sample from Brazil, China, Ghana, Kenya, Peru, the United Kingdom, and the United States. The goal is to establish a robust measure that facilitates the assessment of malevolent creativity in diverse cultural and demographic contexts. A cross-cultural validation is especially important because cultural norms, social desirability, and local interpretations of harmful intent highly influence malevolent behaviors. For instance, some cultures may discourage the endorsement of items related to direct aggression, while others may accept certain forms of manipulation or trickery. In previous adaptations, some dimensions of the MCBS have shown lower reliability, raising concerns that the scale may operate differently across various contexts. Therefore, establishing measurement equivalence is necessary to ensure that differences between countries reflect variations in malevolent creativity rather than cultural response biases or translation issues.

1 | Materials and Methods

1.1 | Participants

We selected seven countries (Brazil, China, Ghana, Kenya, Peru, the UK, and the USA) to cover a broad range of world regions, cultural traditions, and socio-economic backgrounds, thus maximizing the detection of cross-cultural differences. Beyond practical considerations (i.e., the availability of collaborators and samples), these countries differ in their cultural value profiles, particularly on dimensions such as individualism–collectivism, power distance, and uncertainty avoidance, which are theorized to influence creativity (Erez and Nouri 2010).

The study had a total of 2937 participants from 7 countries, ranging in age from 18 to 77 years (M = 28.18; SD = 9.49; 55.84% female). The countries surveyed were represented by 376 from Brazil (M = 32.1; SD = 12.87; 68.08% female), 592 from China (M = 23.9; SD = 6.85; 61.31% female), 623 from Ghana (M = 24.1; SD = 3.61; 43.01% female), 310 from Kenya (M = 20.1; SD = 2.05; 60.32% female), 398 from Peru (M = 22.7; SD = 4.48; 65.07% female), 316 from the UK (M = 41.1; SD = 13.85; 50.31% female), and 322 from the USA (M = 41.4; SD = 11.95; 45.96% female).

The sampling was non-probabilistic by convenience because eligibility criteria were considered as: (1) being of legal age, (2) minimum educational level of completed secondary school, (3) residing in one of the countries that participated in the study, (4) responding completely to the measurement instruments, and (5) providing informed consent. The number of participants was determined using an a priori sample size calculator (Soper 2025) in which the number of observed ($n = 13$) and latent variables of the model ($n = 3$) was considered, together with the effect size ($\lambda = 0.3$), the statistical power ($1-\beta = 0.95$) and the desired probability ($\alpha = 0.05$). With these values, we were able to obtain a recommended number of 184 participants for each national sample.

1.2 | Instrument

Malevolent Creativity Behavior Scale (MCBS, Hao et al. 2016). This instrument is composed of 13 items, which evaluate different daily behaviors of malevolent creativity, composed of 3 dimensions named harming others (6 items), lying (4 items), and playing tricks (3 items). This instrument has a Likert-type scale ranging from 1 (Never) to 5 (Usually). Most countries applied the original English version except for two countries, where translations were made to the Brazilian (Portuguese; Farias et al. 2025) and Peruvian (Spanish) contexts. The initial translation was reviewed in collaboration with two other researchers: disagreements and differences were discussed as a group to refine the scale in both applied languages (i.e., Spanish and Portuguese) for its application to the Latin American context. The process required a linguistic equivalence based on back-translation (Cruchinho et al. 2024; Lira and Caballero 2020) to assess the adequacy of the original scale and to obtain the versions used in the present study for survey collection.

1.3 | Data Analysis

We conducted a confirmatory factor analysis (CFA) to examine various factor structures for each country. The models tested were: (a) a three-factor model (replicating the structure reported by Hao et al. 2016), (b) a unidimensional model, (c) a hierarchical model with 1 s-order factor and three specific factors, (d) a bifactor model with three orthogonal specific factors (standard bifactor model), and (e) a bifactor model with three correlated specific factors (bifactor S-1 model). We included the models b, d, and e to evaluate the plausibility of a unidimensional latent structure for the MCBS, since the three-factor solution, despite showing good fit, revealed highly correlated factors, suggesting considerable shared variance. We tested one additional bifactor (f) and one additional bifactor S-1 (g) model based on the results observed in previous models. Figure 1 presents a graphical representation of all tested models.

All CFA models were estimated using the weighted least squares mean and variance adjusted (WLSMV) estimator. The model fit was assessed using the following indices: values of 0.95 or higher on the Comparative Fit Index (CFI) and Tucker-Lewis Index (TLI) indicate good fit, while values above 0.90 are considered adequate. For the Root Mean Square Error of Approximation (RMSEA), values below 0.05 indicate good fit, and values below 0.10 are considered acceptable (Hu and Bentler 1999; MacCallum et al. 1996; McDonald and Ho 2002). In addition to these indices, we examined model nesting using the likelihood ratio (LR) test.

To investigate the equivalence of the factor structures across countries, we assessed the measurement invariance across countries for the best fitting and best interpretable solution using a multigroup confirmatory factor analysis (MGCFA). We also tested fit indices and measurement invariance with MGCFA of the best-fitting structure across sexes and across age groups. The age groups were divided as follows: 18–25 years ($N = 1729$; 58.86%), 26–39 years ($N = 771$; 26.25%), and 40 years

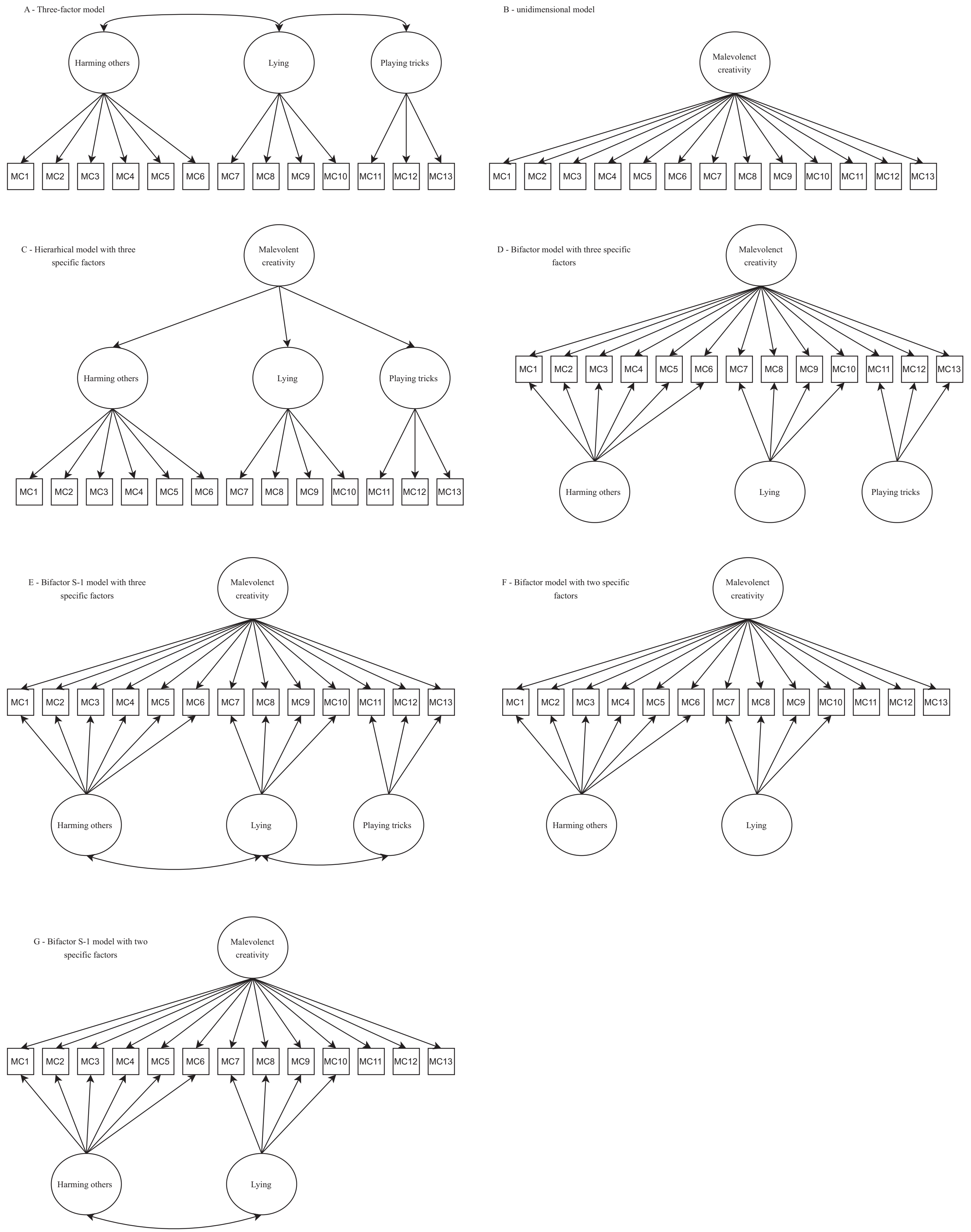

FIGURE 1 | Graphical representation of the factor models tested.

TABLE 1 | Fit indices for the confirmatory factor models by country.

Country	χ^2	df	CFI	TLI	RMSEA [95% CI]
Three-factor model					
Brazil	277.02	62	0.968	0.960	0.096 [0.085–0.108]
China	399.60	62	0.964	0.955	0.096 [0.087–0.105]
Ghana	249.40	62	0.887	0.857	0.070 [0.061–0.079]
Kenya	172.47	62	0.947	0.933	0.076 [0.063–0.090]
Peru	186.86	62	0.980	0.975	0.071 [0.060–0.083]
UK	158.52	62	0.971	0.964	0.070 [0.057–0.084]
USA	111.68	62	0.998	0.998	0.050 [0.035–0.065]
Total sample	1201.18	62	0.982	0.977	0.079 [0.075–0.083]
Unidimensional model					
Brazil	662.59	65	0.912	0.894	0.157 [0.146–0.168]
China	496.42	65	0.954	0.945	0.106 [0.097–0.115]
Ghana	369.61	65	0.816	0.779	0.087 [0.078–0.096]
Kenya	388.69	65	0.844	0.812	0.127 [0.115–0.139]
Peru	711.04	65	0.897	0.877	0.158 [0.148–0.169]
UK	273.29	65	0.938	0.926	0.101 [0.089–0.113]
USA	162.48	65	0.997	0.996	0.068 [0.055–0.082]
Total sample	3004.72	65	0.953	0.944	0.131 [0.126–0.136]
Hierarchical with three specific factors					
Brazil	277.02	62	0.968	0.960	0.960 [0.061–0.108]
China	Convergence problems				
Ghana	249.40	62	0.887	0.857	0.070 [0.061–0.079]
Kenya	172.47	62	0.947	0.933	0.076 [0.063–0.090]
Peru	186.86	62	0.980	0.975	0.071 [0.060–0.083]
UK	158.52	62	0.971	0.964	0.070 [0.057–0.084]
USA	111.68	62	0.998	0.998	0.050 [0.035–0.065]
Total sample	1201.18	62	0.982	0.977	0.079 [0.075–0.083]

(Continues)

TABLE 1 | (Continued)

Note: Factor correlations on the three-factor solution. $F1_{Brazil} \sim F2_{Brazil} = 0.686$; $F1_{Brazil} \sim F3_{Brazil} = 0.842$; $F2_{Brazil} \sim F3_{Brazil} = 0.718$. $F1_{China} \sim F2_{China} = 0.945$; $F1_{China} \sim F3_{China} = 0.848$; $F2_{China} \sim F3_{China} = 0.953$. $F1_{Ghana} \sim F2_{Ghana} = 0.581$; $F1_{Ghana} \sim F3_{Ghana} = 0.795$; $F2_{Ghana} \sim F3_{Ghana} = 0.594$. $F1_{Kenia} \sim F2_{Kenia} = 0.595$; $F1_{Kenia} \sim F3_{Kenia} = 0.764$; $F2_{Kenia} \sim F3_{Kenia} = 0.690$. $F1_{Peru} \sim F2_{Peru} = 0.728$; $F1_{Peru} \sim F3_{Peru} = 0.721$; $F2_{Peru} \sim F3_{Peru} = 0.737$. $F1_{UK} \sim F2_{UK} = 0.783$; $F1_{UK} \sim F3_{UK} = 0.912$; $F2_{UK} \sim F3_{UK} = 0.727$. $F1_{USA} \sim F2_{USA} = 0.956$; $F1_{USA} \sim F3_{USA} = 0.992$; $F2_{USA} \sim F3_{USA} = 0.963$; $F1_{Total\ sample} \sim F2_{Total\ sample} = 0.826$; $F1_{Total\ sample} \sim F3_{Total\ sample} = 0.862$; $F2_{Total\ sample} \sim F3_{Total\ sample} = 0.866$. Factor loadings of the first-order factors on the second-order factor: $F1_{Brazil} = 0.898$; $F2_{Brazil} = 0.765$; $F3_{Brazil} = 0.938$. $F1_{Ghana} = 0.881$; $F2_{Ghana} = 0.659$; $F3_{Ghana} = 0.902$. $F1_{Kenia} = 0.811$; $F2_{Kenia} = 0.733$; $F3_{Kenia} = 0.942$. $F1_{Peru} = 0.844$; $F2_{Peru} = 0.863$; $F3_{Peru} = 0.854$. $F1_{UK} = 0.991$; $F2_{UK} = 0.790$; $F3_{UK} = 0.920$. $F1_{USA} = 0.992$; $F2_{USA} = 0.963$; $F3_{USA} = 1.000$; $F1_{Total\ sample} = 0.907$; $F2_{Total\ sample} = 0.911$; $F3_{Total\ sample} = 0.950$.

Abbreviations: CFI, Comparative Fit Index; CI, confidence interval; df, degrees of freedom; RMSEA, Root Mean Square Error of Approximation; TLI, Tucker-Lewis Index.

or older ($N = 437$; 14.79%). We initially divided the 40+ group into two categories: 40–59 and 60 or older. However, the sample size for the 60+ group ($N = 95$; 3.23%) was insufficient for MGCFA (as at least 200 participants are typically required; Koh and Zumbo 2008). Therefore, we collapsed the groups into a single 40+ category to ensure an adequate sample size.

In line with the recommendations of Wu and Estabrook (2016) for categorical data, we tested the following models for each factor structure: (a) the configural model, which assumes the number and composition of factors are equivalent across groups; (b) the proposition 4 model, which constrains the thresholds to be equal across groups; and (c) the proposition 7 model, which further constrains both thresholds and factor loadings to be equal across groups. All models were estimated using delta parameterization and the WLSMV estimator. A factor structure is considered invariant when changes in fit indices between progressively restricted models do not worsen the fit indices with values that meet the following criteria: $\Delta\chi^2/\Delta df < 5$, $\Delta RMSEA < 0.015$, $\Delta CFI < 0.010$, and $\Delta TLI < 0.010$ (Chen 2007; Cheung and Rensvold 2002).

We performed all the analyses in R software (version 4.4.1). Confirmatory factor analyses and LR tests were conducted with the lavaan package (Rosseel 2012), while multigroup confirmatory factor analyses were carried out using both lavaan (Rosseel 2012) and semTools (Jorgensen et al. 2025). Nesting relationships were also checked using the semTools package (Jorgensen et al. 2025).

2 | Results

We initially tested a three-factor confirmatory model, replicating the structure reported by Hao et al. (2016). While the fit indices were generally acceptable to excellent across all countries, except for Ghana, where the CFI and TLI fell below the expected thresholds, the three factors displayed high correlations (see Table 1 note). In some countries, such as China and the USA, these correlations approached 0.90, indicating a substantial overlap between the factors and suggesting the plausibility of a unidimensional latent construct. Given these results, we also tested unidimensional models, as well as hierarchical models (in which the three first-order factors load onto a single second-order factor), and bifactor, and bifactor S-1 models. Table 1 provides the fit indices for all tested models across countries and for the total sample, except for the bifactor and bifactor S-1 models, which are presented in Table 2. The factor loadings of all the tested factor models are presented in the Supporting Information.

The fit indices of the unidimensional models were lower compared to the three-factor model. However, it is important to note that models with more parameters typically demonstrate a better fit. The unidimensional model exhibited at least one good or excellent fit index for Brazil, China, the UK, and Peru, somewhat supporting the unidimensionality of the MCBS for these countries. Given that the unidimensional model is nested within the three-factor model (as verified using the net() function from the semmTools package), we conducted likelihood ratio tests. Across all countries, the three-factor model provided a significantly better fit to the data than the unidimensional model (Brazil: $\Delta\chi^2(3) = 187.34$, $p < 0.001$; China: $\Delta\chi^2(3) = 92.00$, $p < 0.001$; Ghana: $\Delta\chi^2(3) = 81.56$, $p < 0.001$; Kenya: $\Delta\chi^2(3) = 115.09$, $p < 0.001$; Peru: $\Delta\chi^2(3) = 185.85$, $p < 0.001$; UK: $\Delta\chi^2(3) = 73.97$, $p < 0.001$; USA: $\Delta\chi^2(3) = 42.36$, $p < 0.001$; All countries combined: $\Delta\chi^2(3) = 813.03$, $p < 0.001$).

We also tested a second-order (hierarchical) model, in which the three first-order factors were specified as indicators of a higher-order latent construct. Because the correlated and hierarchical models are statistically equivalent,[1] as both imply the same covariance structure among the observed items, their global fit indices were expected to be identical. Although the models cannot be formally compared using likelihood-ratio tests, we examined the hierarchical specification to assess whether malevolent creativity could also be represented by a general factor, thereby justifying the use of a total score, an interpretation not afforded by the correlated model. For most countries, the results were identical to the three-factor solution. However, in the Chinese sample, the hierarchical model produced a Heywood case (a standardized second-order loading greater than 1 and a negative residual variance), regardless of whether unit loading or unit variance identification was used. Constraining the residual variance to be greater than zero resolved the issue, but this represents a technical rather than empirical solution, as it imposes an assumption not clearly supported by the data. Given these limitations, we retained the three correlated first-order factors as the more faithful representation of the data.

Regarding the bifactor and bifactor S-1 models, we initially tested models with three specific factors and one general factor across all countries. For the bifactor model, most countries showed convergence issues, often resulting in negative residual variances. The only countries for which the model converged were China and the USA. For the bifactor S-1 model, we observed similar convergence issues for some countries. Notably, in both cases (bifactor and bifactor S-1), the negative residual variances were consistently associated with items 11 or 12, both part of the third factor (comprised of items 11, 12, and 13). Consequently, we tested both a bifactor model and a bifactor S-1 model with two specific factors (excluding factor 3) and one general factor that included all items. Overall, the fit indices for these models were excellent across all countries, somewhat supporting the plausibility of a unidimensional latent construct of malevolent creativity behavior, besides specific factors. However, although the fit improved, the interpretation of the specific factors was compromised in both the bifactor and the bifactor S-1 solutions, indicating that after controlling for a general factor, where all items had adequate loadings, there was insufficient variance to maintain the specific factors with their original composition. All the bifactor and bifactor S-1 models are nested within the three-factor model, and likelihood-ratio tests would therefore be possible. However, given the convergence problems observed in several of the bifactor and bifactor S-1 models with three specific factors, and the reduced interpretability of the specific factors in the bifactor and bifactor S-1 models with only two specific factors, we concluded that the three-factor model is the most interpretable and retained it for further analyses.

Although the fit indices were broadly comparable across countries for the retained three-factor solution, we further examined their equivalence by performing an MGCFA. Table 3 presents the fit indices for the baseline, proposition 4 (thresholds), and proposition 7 (thresholds + loadings) models. It is important to note that Ghana and the UK were not included in the measurement invariance results presented in this table due to discrepancies in the number of categories for all items. Specifically, Ghana did not endorse the 1 category in the scale for any item, while the UK did not endorse the 5th category for item 12 and showed limited endorsements of category 5 in other items. A table presenting the percentage of respondents in each country who endorsed each response category for each item is provided in the Supporting Information. As an additional analysis, we performed MGCFA including both countries after harmonizing the response categories across all samples. Specifically, we merged categories 1 and 2 to address differences in the Ghanaian data, and we merged categories 4 and 5 for item 12 to address differences in the UK data. We report these results in the Supporting Information as supplementary evidence of measurement invariance between these countries and the others. However, we did not consider them part of the main analysis, since the lack of endorsement of the affected categories already indicates that the measure functions differently in these countries. The MGCFA results demonstrated full measurement invariance of the MCBS three-factor structure across countries. A latent mean comparison after confirming proposition 7 invariance revealed that Peru and Brazil presented significantly lower scores on Factor 1 compared with Kenya (the reference group; $z = -2.21$, $p = 0.027$; $z = -3.08$, $p = 0.002$, respectively), whereas the USA and China did not ($z = 1.76$, $p = 0.078$; $z = -1.71$, $p = 0.087$, respectively). All countries showed significantly lower scores on Factor 2 compared with Kenya, with the lowest means observed in China

TABLE 2 | Fit indices for the bifactor confirmatory factor models by country.

	χ^2	df	CFI	TLI	RMSEA [95% CI]	ECV	ωh	PUC	H
Bifactor with three specific factors									
Brazil					Convergence problems				
China	347.63	52	0.969	0.953	0.098 [0.088–0.108]	0.88	0.92	0.69	0.95
Ghana					Convergence problems				
Kenia					Convergence problems				
Peru					Convergence problems				
UK					Convergence problems				
USA	71.35	52	0.999	0.999	0.034 [0.008–0.052]	0.95	0.96	0.69	0.98
Total sample	1025.46	52	0.985	0.977	0.080 [0.076–0.084]	0.83	0.94	0.69	0.94
Bifactor S-1 with three specific factors									
Brazil					Convergence problems				
China	132.19	49	0.991	0.986	0.054 [0.043–0.065]	0.77	0.87	0.69	0.94
Ghana	98.59	49	0.970	0.952	0.040 [0.029–0.052]	0.58	0.68	0.69	0.80
Kenia	112.96	49	0.969	0.951	0.065 [0.049–0.081]	0.62	0.75	0.69	0.87
Peru	89.47	49	0.994	0.990	0.046 [0.030–0.060]	0.63	0.79	0.69	0.93
UK					Convergence problems				
USA					Convergence problems				
Total sample	516.404	49	0.993	0.988	0.057 [0.053–0.062]	0.61	0.77	0.69	0.92
Bifactor with two specific factors									
Brazil	174.41	55	0.982	0.975	0.076 [0.064–0.089]	0.76	0.83	0.73	0.94
China					Convergence problems				
Ghana	122.86	55	0.959	0.942	0.045 [0.034–0.055]	0.64	0.71	0.73	0.80
Kenia	128.72	55	0.964	0.950	0.066 [0.051–0.081]	0.68	0.76	0.73	0.86
Peru	211.34	55	0.975	0.965	0.085 [0.073–0.097]	0.76	0.82	0.73	0.94
UK	124.45	55	0.979	0.971	0.063 [0.049–0.078]	0.79	0.82	0.73	0.93
USA	75.70	55	0.999	0.999	0.034 [0.010–0.052]	0.96	0.96	0.73	0.98
Total sample	1218.03	55	0.982	0.974	0.085 [0.081–0.089]	0.86	0.88	0.73	0.94
Bifactor S-1 with two specific factors									
Brazil	181.69	54	0.981	0.973	0.079 [0.067–0.092]	0.78	0.84	0.73	0.94
China	274.63	54	0.977	0.966	0.083 [0.074–0.093]	0.84	0.87	0.73	0.95
Ghana	126.98	54	0.956	0.936	0.047 [0.036–0.057]	0.64	0.70	0.73	0.80
Kenia	130.28	54	0.963	0.947	0.068 [0.053–0.083]	0.67	0.75	0.73	0.86
Peru	131.12	54	0.988	0.982	0.060 [0.047–0.073]	0.64	0.75	0.73	0.95
UK	129.05	54	0.978	0.968	0.066 [0.052–0.081]	0.79	0.81	0.73	0.93
USA	75.54	54	0.999	0.999	0.035 [0.012–0.053]	0.96	0.96	0.73	0.98
Total sample	1146.91	54	0.983	0.975	0.083 [0.079–0.087]	0.81	0.86	0.73	0.94

Note: Factor correlations on the three-factor solution.
Abbreviations: ωh, hierarchical omega; CFI, Comparative Fit Index; CI, confidence interval; df, degrees of freedom; ECV, explained common variance; H, construct replicability; PUC, percentage of uncontaminated correlations; RMSEA, Root Mean Square Error of Approximation; TLI, Tucker–Lewis Index.

TABLE 3 | Fit indices of the multigroup confirmatory factor analysis across countries for the three-factor model.

	χ^2	df	$\Delta\chi^2/\Delta$df	CFI (Δ)	TLI (Δ)	RMSEA (Δ)
Baseline	1147.41	310	—	0.985	0.982	0.082
Prop 4	1423.31	414	2.65	0.982 (0.003)	0.984 (0.002)	0.078 (0.004)
Prop 7	1556.79	454	3.36	0.981 (0.001)	0.984 (0.000)	0.078 (0.004)

Abbreviations: CFI, Comparative Fit Index; df, degrees of freedom; Prop 4, proposition 4 (thresholds); Prop 7, proposition 7 (thresholds + loadings); RMSEA, Root Mean Square Error of Approximation; TLI, Tucker-Lewis Index.

($z = -12.47$, $p < 0.001$), followed by Peru ($z = -8.32$, $p < 0.001$), the USA ($z = -7.67$, $p < 0.001$), and Brazil ($z = -7.66$, $p < 0.001$). For Factor 3, all countries also presented lower scores compared with Kenya, from the lowest to the highest: China ($z = -13.61$, $p < 0.001$), Brazil ($z = -9.55$, $p < 0.001$), Peru ($z = -7.03$, $p < 0.001$), and the USA ($z = -3.92$, $p < 0.001$).

We also investigated the factor structure and the measurement invariance of the MCBS across sexes and age groups. To do so, we first tested the three-factor model for males and females and for each age group. In all cases, the fit indices were adequate, as presented in Table 4. The factor loadings are presented in the Supporting Information. To investigate the equivalence of item parameterization across sexes and across age groups, we conducted an MGCFA for each case. The results are presented in Table 5. They indicate that the three-factor structure of the MCBS is fully invariant across both sexes and age groups. A latent mean comparison after confirming proposition 7 invariance revealed that males had significantly higher scores than females on Factor 1 ($z = 6.02$, $p < 0.001$), Factor 2 ($z = 6.96$, $p < 0.001$), and Factor 3 ($z = 9.15$, $p < 0.001$). Regarding age groups, the latent mean comparison revealed that both the middle adulthood and older group, as well as the early adulthood group, presented lower scores than the emerging adulthood group across all three factors. In all cases, the middle adulthood plus older group exhibited the lowest scores. The differences were as follows: Factor 1—middle adulthood plus older ($z = -11.99$, $p < 0.001$), early adulthood ($z = -6.03$, $p < 0.001$); Factor 2—middle adulthood plus older ($z = -12.85$, $p < 0.001$), early adulthood ($z = -7.35$, $p < 0.001$); Factor 3—middle adulthood plus older ($z = -10.33$, $p < 0.001$), early adulthood ($z = -3.36$, $p = 0.001$).

TABLE 4 | Fit indices for the three-factor confirmatory factor models by sex and age groups.

	χ^2	df	CFI	TLI	RMSEA [95% CI]
Sex					
Males	568.14	62	0.982	0.977	0.079 [0.073–0.085]
Females	710.37	62	0.980	0.930	0.080 [0.075–0.085]
Age groups					
18–25	783.693	62	0.971	0.963	0.082 [0.077–0.087]
26–39	356.801	62	0.989	0.987	0.079 [0.071–0.087]
40+	224.057	62	0.987	0.983	0.078 [0.067–0.089]

Note: $F1_{Males}$~$F2_{Males} = 0.841$; $F1_{Males}$~$F3_{Males} = 0.850$; $F2_{Males}$~$F3_{Males} = 0.863$. $F1_{Females}$~$F2_{Females} = 0.807$; $F1_{Females}$~$F3_{Females} = 0.869$; $F2_{Females}$~$F3_{Females} = 0.863$. $F1_{18\text{-}25}$~$F2_{18\text{-}25} = 0.760$; $F1_{18\text{-}25}$~$F3_{18\text{-}25} = 0.795$; $F2_{18\text{-}25}$~$F3_{18\text{-}25} = 0.833$. $F1_{26\text{-}39}$~$F2_{26\text{-}39} = 0.875$; $F1_{26\text{-}39}$~$F3_{26\text{-}39} = 0.931$; $F2_{26\text{-}39}$~$F3_{26\text{-}39} = 0.905$. $F1_{40+}$~$F2_{40+} = 0.832$; $F1_{40+}$~$F3_{40+} = 0.900$; $F2_{40+}$~$F3_{40+} = 0.835$.
Abbreviations: CFI, Comparative Fit Index; CI, confidence interval. Factor correlations on the three-factor solution; df, degrees of freedom; RMSEA, Root Mean Square Error of Approximation; TLI, Tucker-Lewis Index.

3 | Discussion

The current findings elucidate the structural properties, cross-cultural applicability, and sex invariance of the Malevolent Creativity Behavior Scale (MCBS) initially introduced by Hao et al. (2016). Across various countries (i.e., Brazil, China, Ghana, Kenya, Peru, the UK, and the USA) and including both male and female samples, and participants from different age groups, our analyses reveal a range of potential factor solutions, with results generally favoring a three-factor structure of malevolent creativity.

A key point of the investigation involved comparing the established three-factor model (Hao et al. 2016) with alternative solutions, including unidimensional, hierarchical, and bifactor models. Our replication of the three-factor model exhibited acceptable to excellent fit indices across most countries, aligning with previous findings (Farias et al. 2025; Meshkova et al. 2018). However, the high inter-factor correlations observed, often exceeding 0.80, suggest that these distinct factors share a significant portion of their variance, which could align with the previous suggestion of the scale lacking a broad measurement of malevolent creativity (Waldie et al. 2021). Both hierarchical and bifactor models further supported a central malevolent creativity factor, with the specific dimensions adding limited additional explanatory power once the general factor was accounted for. Although the bifactor solution often improved certain fit indices, convergence issues in some countries (possibly due to sample size and response distribution constraints), and the lack of interpretability of the remaining specific factors, highlight the complexity of this modeling approach. Overall, finding a better fit for the three-factor structure has important implications. Conceptually, it indicates that malevolent creativity may function differently than a single construct, contrary to Cropley et al.'s (2008) original definition that highlights intentional novelty used to cause harm. On the other hand, the high correlations among the three subscales seen in our study and previous research (e.g., Hao et al. 2016; Waldie et al. 2021) may suggest that harming others, lying, and playing tricks are better

TABLE 5 | Fit indices of the multigroup confirmatory factor analysis across sexes and age groups.

	χ^2	df	$\Delta\chi^2/\Delta df$	CFI (Δ)	TLI (Δ)	RMSEA (Δ)
MGCFA across sexes						
Baseline	1280.96	124	—	0.981	0.976	0.080
Prop 4	1335.21	150	2.08	0.980 (0.001)	0.980 (0.004)	0.073 (0.007)
Prop 7	1218.05	160	11.72	0.983 (0.003)	0.983 (0.003)	0.064 (0.009)
MGCFA across age groups						
Baseline	1357.67	186	—	0.981	0.977	0.080
Prop 4	1442.62	238	1.63	0.981 (0.000)	0.981 (0.004)	0.072 (0.008)
Prop 7	1429.23	258	0.67	0.981 (0.000)	0.983 (0.002)	0.068 (0.004)

Abbreviations: CFI, Comparative Fit Index; df, degrees of freedom; Prop 4, proposition 4 (thresholds); Prop 7, proposition 7 (thresholds + loadings); RMSEA, Root Mean Square Error of Approximation; TLI, Tucker-Lewis Index.

understood as different outward expressions of a broader malevolent creativity tendency.

Cross-cultural analyses provided evidence for the MCBS's structural robustness. Measurement invariance testing indicated that the MCBS three-factor structure and factor loadings are stable across the sampled countries. This level of invariance supports meaningful comparisons of malevolent creativity across different cultural groups. However, some challenges arise regarding response category distributions in specific countries, such as Ghana and the UK, where certain response categories were either not endorsed or used sparingly. These patterns highlight potential cultural differences in item interpretation or respondents' willingness to select extreme categories. While these issues did not undermine the applicability of the MCBS, they emphasize the need for careful consideration of cultural context and response style. According to Erez and Nouri's (2010) model, creativity is influenced by cultural values that determine whether novelty is aimed at individual or collective goals, and whether deviation is tolerated. Countries in our sample vary significantly on cultural dimensions such as individualism–collectivism and power distance (Hofstede 2001). For example, the UK and the USA are characterized by high individualism, where independent expression may support the endorsement of malevolent strategies, such as lying or playing tricks. In contrast, Ghana, Kenya, and Peru tend to be more collectivist, with social harmony norms that may discourage behaviors aimed at causing harm. These differences may help explain why specific response categories were rarely endorsed in Ghana and the UK, highlighting that cultural values influence not only behavior but also the self-reporting of malevolent creativity.

The results also extend to comparisons by sex and age groups. The tested factor solution demonstrated invariance across male and female respondents and age groups, indicating that the MCBS functions equivalently in all groups, as found previously (Kapoor and Kaufman 2022; Perchtold-Stefan et al. 2023b). After establishing measurement invariance across males and females, we observed that males scored significantly higher on the three malevolent creativity factors, consistent with previous research indicating that males often report higher levels of malevolent creativity, partly due to gender socialization processes that promote assertiveness, competitiveness, and risk-taking (Perchtold-Stefan et al. 2023b). Females, on the other hand, may be socialized to prioritize prosocial behaviors and relational harmony, which could reduce their willingness to endorse items involving direct harm, even if indirect manipulation is present. In contrast, age showed a consistent decline, with emerging adults scoring the highest, early adults scoring lower, and middle to older adults scoring the lowest on all factors. Developmentally, this pattern aligns with increased self-regulation, decreased sensation-seeking, and occupational costs as age advances (Maples-Keller et al. 2016). Overall, the pattern suggests a general tendency toward malevolent creativity that is more prevalent in men and peaks in younger adulthood, then declines with maturation and the shifting of social roles.

Although the present study supports the structural validity and cross-group equivalence of the MCBS, its content does not fully capture the full range of malevolent creativity behaviors. The 13 items mainly measure interpersonal deception and low-to-moderate severity antagonism. Key areas are underrepresented or missing, such as property damage (e.g., creative sabotage of materials or digital assets), harm to animals, economic or organizational harms (e.g., clever embezzlement or procurement fraud), technologically mediated harms (e.g., doxxing, deepfakes, social engineering exploits), and collective or institutional targets (e.g., coordinated misinformation, infrastructure disruption). Thus, our psychometric results cannot be interpreted independently of the scale's conceptual ambiguity. As noted, only two items of the MCBS directly reference creative or original actions. The other items describe harmful or deceptive behaviors without specifying novelty, a core criterion of creativity (Runco and Jaeger 2012). This means that, while we present that the MCBS functions consistently across cultures and demographic groups, what is being measured may be malevolent ideation or antagonistic tendencies rather than malevolent creativity.

Despite these promising results, the study's limitations deserve attention. First, the convergence difficulties encountered with some bifactor models, especially in countries with smaller sample sizes, underscore the relationship between sample heterogeneity and model complexity. Bifactor models with multiple specific factors require large, diverse data sets to estimate additional parameters reliably. Future investigations could implement larger samples in countries where fit indices indicated convergence issues or adopt streamlined modeling approaches that explicitly target potential sources of misfit. Second, certain items may need more nuanced cultural adaptation. For instance, if a specific behavior measured by the MCBS does not align well with local norms in each country, it may be necessary to adjust item wording or provide clearer definitions. Such adaptations should be conducted carefully and tested thoroughly to maintain conceptual consistency with the original scale. Third, our multigroup analyses by age required combining the two oldest categories because the 60+ subgroup ($N=95$) was too small for MGCFA (recommended ≥ 200; Koh and Zumbo 2008). This choice may hide differences in later adulthood and limit conclusions about age-specific noninvariance. Future research should oversample older adults and test alternative groupings and continuous-age methods to better assess age-related differences.

Looking ahead, future research could further refine the MCBS by exploring simplified models or alternative item groupings, particularly in contexts where convergence issues persist. Developing and testing shorter forms of the MCBS might also be beneficial. To strengthen the external validity of the MCBS, future research could include a broader range of countries and cultures, particularly from regions that have been underrepresented in previous studies. This would examine whether malevolent creativity patterns are consistent across more diverse cultures, such as emerging economies or indigenous cultures. Additionally, research on how malevolent creativity relates to other psychosocial variables, such as aggressiveness, morality, and leadership styles, could provide further insights into the underlying mechanisms of malevolent creativity and how it manifests across different groups and contexts. In addition, we can examine the differences in these relationships based on participants' country and sex. Several studies have indicated that certain variables contribute to malevolent creativity, such as moral disengagement (Shi et al. 2023), aggression (Zhou et al. 2024), the Dark Triad (Gao et al. 2022), unfairness (Zhang et al. 2024), early life adversity (Ceballos and Watt 2023; Geng et al. 2024; Jia et al. 2020), and moral reasoning (Zhao et al. 2022). It has been suggested, however, that empathy, conscientiousness, social support, prosocial motivation (Wu et al. 2024), resilience (Wang et al. 2022), and emotional intelligence (Harris et al. 2013) may mitigate this type of creativity.

4 | Conclusion

In this study, we present a comprehensive cross-cultural examination of the Malevolent Creativity Behavior Scale, assessing its structure, measurement invariance, and demographic differences across seven countries and age groups. Our findings provide evidence that the three-factor solution proposed initially by Hao et al. (2016) is stable across cultural contexts, sexes, and age groups. The MCBS demonstrated acceptable fit indexes in most countries, supporting its use for cross-national comparisons. At the same time, our results highlight an important conceptual limitation, that is, the MCBS may not completely capture malevolent creativity. It also measures malevolent ideation or antagonistic tendencies. This distinction is important for interpreting differences and cultural patterns, as future findings possibly reflect variations in willingness to report antagonistic behavior rather than differences in creative problem-solving conducted with harmful intent. Nonetheless, our study advances the understanding of the MCBS, points out conceptual gaps that require refinement, and offers directions for future work.

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Conflicts of Interest

The authors declare no conflicts of interest.

Data Availability Statement

The data that support the findings of this study are available from the corresponding author upon reasonable request.

Endnotes

[1] Although both models imply the same covariance structure among the observed items, they differ in parameterization and interpretation: in the correlated first-order model, factor covariances are estimated directly, whereas in the hierarchical model they are reproduced through a higher-order factor.

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Supporting Information

Additional supporting information can be found online in the Supporting Information section. **Data S1.**