TWO COUNTRIES, ONE HAPPINESS? THE INTERDEPENDENT HAPPINESS SCALE IN ITALY AND POLAND

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Although Interdependent Happiness Scale (IHS) has been used in different socio-cultural contexts, no study has tested the measurement invariance across countries. Because previous research did not agree on the dimensionality of the scale, we primarily investigated the factorial structure of the IHS, comparing the one-factor model to three-factor, second-order, and bifactor models. We also evaluated the applicability of the IHS to Italian (*N* = 290) and Polish samples (students: *N* = 253; community: *N* = 241) using a multigroup confirmatory factor analysis. We investigated also whether the scale measures an aspect of happiness different from satisfaction with life. The bifactor model fitted the sample data better than alternative models in all samples. The general factor accounted for over 90% of the reliable variance in the total IHS score for all samples. Our study supported the configural and metric invariance of the bifactor model in the Italian and Polish versions of the IHS, ensuring the comparability of correlations involving the IHS scores in the two different cultures and samples. We showed that IHS is not redundant with the SWLS, although the two concepts are deeply intertwined. Our findings provided support for a conceptualization of interdependent happiness as a hierarchically organized construct.

Keywords: Interdependent happiness; Satisfaction with life; Bifactor modeling; CFA; Construct validity.

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Happiness is vital to both individual and social life. Happiness is one of the key components in the mutual construction of culture/society and the human mind (Hitokoto & Uchida, 2018). Happiness experiences can be both common and specific across countries, depending on cultural and individual factors. For example, some research has found cultural similarities in the factors influencing happiness levels, such as economic and material wealth (Diener, 2000). However, while these factors can be fundamental for society, securing a certain level of collective well-being, a good amount of happiness results from the actualization of psychological meaning in life, influenced by sociocultural context. For example, Diener, Suh, Smith, and Shao (1995) found a positive correlation between national individualism — the national characteristic of valuing people's freedom — and average subjective well-being (SWB). Cross-cultural researchers frame these findings as an opportunity to examine the meaning of happiness in various cultures (Markus & Kitayama, 1991) and to seek true cultural diversity in happiness.

In happiness research, the individual level has been extensively investigated, while the social aspects have long been overlooked. According to relationship research across cultures (Antonucci, Ajrouch, & Birditt, 2013; Delle Fave et al., 2016; Lansford, Antonucci, Akiyama, & Takahashi, 2005), recent studies have proposed culturally sensitive models of well-being (Uchida & Kitayama 2009; Uchida, Norasakkunkit, & Kitayama, 2004; Uchida & Ogihara 2012). People in different cultures may pursue happiness in different ways (Krys, Capaldi et al., 2019). The most popular scale of subjective well-being — Satisfaction with Life Scale (SWLS; Diener, Emmons, Larsen, & Griffin, 1985) — used in thousands of studies worldwide seems to track the individualistic-themed type of well-being (Krys, Zelenski et al., 2019). To bridge the cultural gap in happiness research, Hitokoto and Uchida (2015) developed the Interdependent Happiness Scale (IHS) to measure individual perceptions of the interpersonally harmonized, quiescent, and ordinary nuances of happiness, which are the shared meaning of happiness in an interdependent cultural context. The IHS is one of the leading propositions to complement SWLS. The IHS was prepared in the Japanese cultural context and has been validated in several other collectivistic and individualistic societies. Therefore, in social indicator research, it seems essential to develop and validate complementary scales that will track more collectivism-themed types of well-being.

The construct of interdependent happiness was proposed to echo the differences between independent and interdependent self-construals in well-being studies (Markus & Kitayama, 1991). Hitokoto and Uchida theorized that life satisfaction might be the type of happiness pursued in individualistic cultures and by individuals with a predominantly independent construal of self. In contrast, interdependent happiness is supposed to be pursued in collectivistic cultures (particularly in Confucian Asia) and individuals who hold a predominantly interdependent construal of self. Surprisingly, previous empirical studies on the association between self-construals and interdependent happiness documented that the latter was predicted by independent self-construals better than by interdependent self-construals (Krys, Zelenski et al., 2019; Park, Norasakkunkit, & Kashima, 2017; Pilarska, 2014; Yamaguchi & Kim, 2015). However, most previous studies relied on the self-construals scale proposed by Singelis (1994), which has problematic psychometric qualities (Grace & Kramer, 2003). Recent research has also cast doubts over the assumption that individuals from different cultures have reliable differences in self-construal of Markus and Kitayama's theory (Matsumoto, 1999). Some authors, thinking that Markus and Kitayama used dichotomous categories, assumed, on the contrary, that individuals may be characterized by specific combinations of individualistic and collectivistic co-existing attitudes (Green, Deschamps, & Paez, 2005).

According to the most recent revision of the self-construal scales (Vignoles et al. 2016), selves may be construed in terms of defining the self (i.e., difference vs. similarity), experiencing the self (i.e., self-containment vs. connection to others), making decisions (i.e., self-direction vs. receptiveness to influence), looking after oneself (i.e., self-reliance vs. dependence on others), moving between contexts (i.e., consistency vs. variability), communicating with others (i.e., self-expression vs. harmony), or dealing with conflicting interests (i.e., self-interest vs. commitment to others). This way, we expect to provide a fine-grained picture of the associations between interdependent happiness and self-construals.

Different collectivistic cultures came up with different ideas about the sources of an individual's worth. Leung and Cohen (2011) framed these differences into two different cultural logics: "honor" versus "face" cultures. The former are collectivistic cultures based on reputation (Mosquera, Manstead, & Fischer, 2002). A sense of honor depends on the ability to defend one's reputation — honor can be lost, but can also be regained (Osterman & Brown, 2011). The logic of the second type of collectivistic culture — face cultures — is based on "what other people see." In face cultures, self-worth is based on others' assessments of whether they fulfill stable social role obligations (Kim & Cohen, 2010). Leung and Cohen (2011) also proposed a third type of cultural logic — cultures of "dignity" — that they theorized to characterize individualistic cultures. In these cultures, each person's worth is permanent and essential and does not rely on others' admiration.

Uchida and Ogihara (2012) theorized that the construal of happiness is conceptualized more interdependently among members of collectivistic cultures than among members of individualistic cultures. People in individualist societies are inclined to endorse an achievement-oriented meaning of happiness, while those in collectivist ones are more inclined to espouse a relationship-oriented construal of happiness. These cultural differences in the meaning of happiness could explain why individuals in interdependent cultures scored lower than those in individualist contexts in Western-derived measures of subjective well-being, like the SWLS (Diener et al. 1985). This cultural-psychological theory on the shared meaning of happiness in a cultural context has started to become an essential milestone for studying happiness (Hitokoto & Uchida, 2018).

Measurement Issues

In the original study, Hitokoto and Uchida (2015) administered the IHS to Japanese and American student samples and community samples collected in the United States, Germany, Japan, and Korea. The study aimed to compare how the scale score would predict subjective well-being across countries and samples. The IHS score tended to be more strongly associated with overall well-being than self-esteem, especially in Korea, where the impact of self-esteem was negligible. When Hitokoto and Uchida (2015) addressed the psychometric properties of the IHS, a single factor solution was found to apply to both American and Japanese student samples, as indicated by the acceptable goodness-of-fit indices; other factorial models were not tested. Using a Filipino high school student sample, Datu, King, and Valdez (2016) demonstrated the reliability and discriminant validity of the IHS against the sense of relatedness to others — whether one feels accepted or special when one is with one's parents, teachers, and peers. The same study provided evidence that a hierarchical model of interdependent happiness was more applicable to different Filipino samples, with three first-order factors (i.e., relationship-oriented happiness or harmony, quiescent happiness, and ordinary happiness) and a second-order general factor. Unfortunately, this model was fit-equivalent to a standard three-factor model with correlated factors, thus precluding the possibility of studying interdependent happiness at different levels of analysis as a general construct with specific subordinate facets. Furthermore, the presence of three different factors in the study by Datu et al. (2016) seems inconsistent with the confirmatory factor analysis of Hitokoto and Uchida (2015).

Another measurement issue concerns the invariance of IHS across different cultures. Cross-cultural research cannot take for granted that a scale will elicit equivalent responses because of differences in context, culture, and language. Thus, the measurement model's equivalence must be investigated to determine whether different populations' responses differ by more than chance. Because the interdependent happiness construct has been proposed in the context of differences in Eastern-Western conceptualization of happiness (e.g., Uchida & Kitayama 2009; Uchida & Ogihara 2012; Uchida et al., 2004), to make cross-cultural comparisons valid, a scale devised to assess this construct should ensure equivalence of translated versions. Multigroup confirmatory factor analysis (MGCFA) is the elective method for determining the equivalence of test scores obtained from different groups of respondents in different countries (Byrne, 2016). In particular, valid inferences at the construct level are legitimate depending on the degree of measurement invariance of the observed scores. First, *configural* invariance ensures that the same construct is measured in different cultures by the same set of items, although the measurement unit and the average level of the items can be different. Second, *metric* invariance secures that each item's measurement units are the same in different cultures, although each item's average level can be different. Third, *scalar* invariance guarantees that the items have the same average level. When all conditions are met, the average level of the construct can be compared between groups. When only metric invariance is supported, one can compare between groups the correlations obtained in each group. Any quantitative comparison between countries is precluded if metric invariance is not supported. To the best of our knowledge, the measurement invariance of the IHS has not been addressed yet. Invariance studies are not available, although the instrument has been applied in different cultures and languages.

A third measurement issue concerns the IHS's construct validity: whether the scale measures an aspect of happiness different from subjective well-being. To be considered distinct, two constructs must satisfy two requirements (Singh, 1991): they must be theoretically and empirically different. While the first requirement has been addressed elsewhere (Hitokoto, & Uchida, 2018; Uchida & Oishi, 2016), the second one still is open. Except for Datu et al. (2016), the few studies that administered the IHS and the SWLS reported large to very large correlations (i.e., r = .68-.74), not too much, however, as expected for two sister-concepts of happiness (Krys, Zelenski et al., 2019; Arimitsu, Hitokoto, Kind, & Hofmann, 2019). A regression study also showed that the IHS predicted SWLS with a large effect size (i.e., β = .59), controlling several other variables, including demographics and personality (Hitokoto & Tanaka-Matsumi, 2014). Unfortunately, this study did not report the zero-order correlation useful for comparison with other researches. All these findings did not consider the measurement error in both variables, a factor that could have deflated the correlation at the observed variable level (see Hunter & Schmidt, 2004). Without appropriate corrections for unreliability, the observed correlations between scales putatively measuring different constructs may be small when the construct-level correlations are very large (Le, Schmidt, Harter, & Lauver, 2010). In sum, although there are sound theoretical arguments for separating interdependent happiness from life satisfaction, it is unclear whether the two constructs might be redundant (Shaffer, DeGeest, & Li, 2016).

The Present Study

Hitokoto and Uchida (2015) showed that IHS captures interdependent happiness in cultures that focus on harmonious interpersonal relations. Even with the potential advantages of using IHS to compare collectivist and individualistic cultures, there is only one empirical investigation (Datu et al. 2016) that examined the IHS applicability. Hence, our study addressed this important empirical gap by assessing the psychometric validity and the IHS factorial structure among students in Italy and Poland and community participants in Poland. The two countries are considered balanced in individualism and collectivism, reporting a score very close to the middle of the ranking on Hofstede's revised individualism-collectivism index (e.g., Minkov et al., 2017). Accordingly, we expect cultural differences on the IHS to be of lesser importance in measuring the construct.

Before testing invariance, we aimed to establish the best factorial structure to represent IHS in the two countries. We examined which first-order, second-order, or bifactor models best fit the IHS scores. According to Datu et al. (2016), the first-order factors and the higher-order factors had the same fit to the data, so testing a bifactor model could shed light on the hierarchical structure of the IHS. According to

Bonifay, Lane, and Reise (2017), a bifactor model provides invaluable information to evaluate the extent to which a psychometric scale yields a univocal total score and whether it provides a reliable assessment of theoretically distinct constructs after accounting for variance in the general factor. To our knowledge, this issue has not been addressed in the extant literature for the IHS. We expected the bifactor model to be the best fitting one for both countries because it could best represent the hierarchical structure of the IHS than a second-order model, whose fit is equivalent to a standard three-factor model.

Previous research on the overlap of IHS and SWLS has investigated the construct overlap using only both measures' total scores. In the present study, we aimed to address this issue by parsing general and specific variance components in the IHS using a bifactor model. Foreshadowing the method section, it is worth noting that a bifactor model is a latent structure in which a general factor common to all items represents the variance in the target domain of interest (i.e., the interdependent happiness). In contrast, group factors orthogonal to the general factor reflect the common variance in specific groups of items (i.e., interpersonally harmonized, quiescent, and ordinary nuances of happiness). This feature allows to separate general and specific variance components and assess the IHS criterion validity. We expect the general IHS factor to be more strongly associated with the SWLS than other specific IHS factors. For comparison, we also used a domain-specific definition of satisfaction with life to see whether the potential redundancy could be resolved using top-down and bottom-up life satisfaction measures (Bechtel, 2007; Headey, 2014).

Because IHS is expected to be associated with contextual and social variables linked to selfhoods conceptualization and identity construction, like self-construals (Vignoles at al., 2016) and cultural logics (Leung and Cohen, 2011), we aim to study their associations with the IHS factors in the two countries involved in the present study. In doing so, this paper extends the current discussion on well-being metrics by providing an empirical argument that IHS can be used as a valid well-being scale in various cultural contexts.

METHOD

Participants and Procedures

Italian data: 290 University students from different Universities located in Rome aged from 18 to 50 (*M* = 25.13, *SD* = 4.43; 53.4% females) participated in the study. The ethical review board for psychological research of Sapienza University of Rome approved the study (Prot. 1362).

Polish data: the Polish sample was composed by two subsamples: a) students, *N* = 253, *M* = 22.50, *SD* = 2.60; 79.8% females; b) community: *N* = 241; *M* = 43.05, *SD* = 14.97; 56.4% females. The Polish part of the study was approved by the Polish Academy of Sciences (prot. #7/11/2017).

Participants were asked to fill in an on-line battery of self-report measures. The declared completion time of the on-line survey was around 45 minutes. Before collecting the data and obtaining on-line informed consent, all participants were informed about the study's general aims and characteristics.

Measures

Socio-demographic variables. We collected participants' age, gender, and level of urbanization of where they live (from 1 = *residing in a rural zone or village* to 9 = *residing in a metropolitan area*). Furthermore, we asked participants to report personal and family monthly income. Because Italy and Poland have different average national incomes, we transformed incomes into categories based on a tertiary split. Low-, medium-, and high-income participants were those below the 33rd percentile, between the 34th and 66th percentile, and above the 67th percentile of each country's distribution, respectively.

Interdependent Happiness Scale (IHS; Hitokoto & Uchida, 2015). It comprises nine items to assess happiness in terms of interdependent existence and interpersonal harmony. Respondents rated their answers on a 9-point Likert scale (from 1 = *not describing me at all* to 9 = *it fully describes me*). The scale measures three different factors, each comprised of three items. Relationship oriented happiness or harmony refers to a harmonious state with a balance between oneself and significant others. Quiescence is a conception of wellness considered to be a state of low activation to preserve social norms. Ordinariness refers to the quality of not being different or special or unexpected in any way. In the current study, the total score's internal consistency was .87, .90, and .92 for Italian students, Polish students, and the Polish community, respectively.

Satisfaction with Life Scale (SWLS; Diener et al., 1985). It is comprised of five items to assess individual life satisfaction. Respondents rated their answers on a 9-point Likert scale (from 1 = *not describing me at all*, to 9 = *it fully describes me*). Higher scores correspond to higher life satisfaction levels. In the current study, the total score's internal consistency was .86, .84, and .88 for Italian students, Polish students, and Polish community, respectively.

The 48-item Self-Construal Scale (SCS; Owe et al., 2013; Vignoles, et al., 2016). It is a scale aimed to measure how much participants perceive themselves as independent versus interdependent from their context. The scale includes eight dimensions: difference versus similarity, self-containment versus connection to others, self-direction versus receptiveness to influence, self-reliance versus dependence on others, self-expression versus harmony, self-interest versus commitment to others, consistency versus variability, and decontextualized versus contextualized self. Respondents rated their answers on a 9-point Likert scale (from 1 = *not describing me at all*, to 9 = *it fully describes me*). In this study, we used the dimension oriented to the self so that higher scores correspond to higher independent self-construal. In the current study, the internal consistencies of the eight dimensions were the following. Difference: .78, .70, and .60; self-containment: .39, .36, .15; self-direction: .74, .73, .64; self-reliance: .78, .72, .62; self-expression: .73, .72, .45; self-interest: 71., .73, .48; consistency: .82, .80, .70; de-contextualized self: .71, .71, .55, for Italian students, Polish students, and Polish community, respectively.

Honor, Dignity, and Face Scales (HDF; Leung, and Cohen, 2011). We measured honor, dignity, and face cultural logics based on 18-items originally developed by Severance et al. (2013). Respondents rated their answers on a 9-point Likert Scale (from 1 = *not describing me at all*, to 9 = *it fully describes me*). In the current study, the internal consistency coefficients of honor were .68, .83, and .74 for Italian students, Polish students, and Polish community, respectively. For dignity, the internal consistencies were .79, .87, and .86. For the face subscale, the coefficients were .64, .82, and .84.

Domain-specific life satisfaction. These data were collected using nine items measuring satisfaction with: health, relationship with family, relationship with friends, relationship with other people, own skills, personal growth, life achievements, individual and family's financial situations. Respondents rated their answers on a 9-point Likert Scale (from 1 = *not describing me at all*, to 9 = *it fully describes me*). Higher scores correspond to higher life domains satisfaction levels. We conducted an exploratory factor analysis on our data, which showed that the items were organized into two dimensions: nonmaterialistic and materialistic satisfaction. The internal consistency of the nonmaterialistic satisfaction was .87, .90, and .89, and that of the materialistic satisfaction was .80, .78, .79 for Italian students, Polish students, and Polish community, respectively.

Statistical Analyses

We tested the following models on both Italian and Polish samples using the "lavaan" (Rosseel, 2012) package for R: a one-factor model, a three-factor model with uncorrelated and correlated factors (i.e., relationship oriented or harmony, quiescence, and ordinariness), a second-order factor model, a bifactor model with a common factor and three-group factors. In the one-factor model, each item reflected the variance of a single latent factor common to all IHS items. In the three-factor model, the IHS domain was divided into separate but related subcomponents. As such, the three-factor model does not require assumptions regarding a common latent construct directly or indirectly impacting item responses. A second-order factor model explicitly imposes a hierarchical structure on the data, in which first-order factors are not correlated, but they load on a general factor. Because the second-order and the three-factor model are equivalent fit models, previous research (Datu et al., 2016) could not compare the two models to establish whether the IHS construct has a hierarchically multifaceted structure. In the present study, we fitted a bifactor model with a general factor common to all items and group factors common to the three specific groups of items.

Because the data violated multivariate normality assumptions (Mardia's normalized coefficient = 29.57 for Italy, and 16.96 and 46.57 for Polish students and community, respectively), we used the robust maximum likelihood (ML) method to estimate model parameters. Robust ML provides unbiased parameter estimates, corrects standard errors for non-normal data, and adjusts many of the model's fit indices (Satorra & Bentler, 2001). We assessed the model's fit using the Satorra–Bentler scaled χ^2 (SBχ^2) and other indices to have a more comprehensive evaluation (Kline, 2011). The comparative fit index (CFI) and Tucker-Lewis index (TLI) compare the factor model's fit to that of a baseline model in which all observed variables are expected to be uncorrelated. CFI and TLI > .90 indicate acceptable fit, while values > .95 indicate a good fit. The standardized root mean square residual (SRMR) is a residual-based statistic for which values less than .10 and .05 are considered acceptable and good fit, respectively. The root mean square error of approximation (RMSEA) measures the difference between the model-implied covariance matrix and the population one to control sampling variability. An RMSEA of .05 or less indicates a close fit, and values up to.08 represent a reasonable error of approximation. Because all models tested were nested (i.e., all parameters of a more restricted model are included in a less restrictive one), a scaled chi-square difference test was used to assess whether a more restrictive model produced a better fit than a less restricted one (Satorra & Bentler, 2001).

We tested the invariance across subgroups using a MGCFA (Cheung & Rensvold 2002; Vandenberg & Lance, 2000). We followed a sequential constraint approach (Dimitrov, 2010). First, we tested whether the factor loading pattern was the same across groups (i.e., configural invariance). Second, we constrained the factor loadings to be equal across groups and compared the fit of this model (i.e., metric invariance) to the configural invariance model. Metric invariance is a prerequisite for testing the invariance of the intercepts of the underlying items setting them equal across groups (i.e., scalar invariance). Although the SBχ^2 difference test has been used, Cheung and Rensvold (2002) suggested that it is more reasonable to base invariance decisions on a difference in CFI. ΔCFI $\leq$ .010 supports the invariance hypothesis.

Using the bifactor model parameters, we computed the hierarchical factor scores using the Anderson-Rubin method, which ensures the estimated factors' orthogonality and produces scores with a mean of 0, and a standard deviation of 1 (DiStefano, Zhu, & Mindrila, 2009). Because the general and the group factors are orthogonal in a bifactor model, the factor scores account for separate quotas of the variance in IHS ratings. We assessed the IHS correlations with SWLS and domain-specific life satisfaction scores, self-construal dimensions, and cultural logics, using both summated ratings and hierarchical factor scores in subsequent validation analyses. Correlations with positive coefficients approaching or exceeding .10, .30, or .50 were considered small, moderate, or large in magnitude (Cohen, 1988), respectively.

RESULTS

Descriptive Statistics

The IHS score was moderately high in the Italian sample (M = 5.64, SD = 1.32). Polish data showed lower levels: student subsample (M = 5.03, SD = 1.42), community subsample (M = 4.80, SD = 1.37). The Italian and Polish student samples were different (F = 26.92, p = .000). The Italian student sample was different from the Polish community subsample (F = 26.92, p = .000). No significant differences were found between Polish student and community participants.

Table 1 reports means and standard deviations for IHS broken down by country, gender (male/female), urbanization level (rural/urban), personal income (low-medium-high), family income (low-medium-high). The analysis of Italian data showed significant IHS differences by gender (F = 4.14, p = .043), showing that males had higher levels of IHS compared to females. No significant differences in IHS emerged from urbanization, personal, and family monthly income. The analysis of Polish students' data showed no differences in IHS by gender, personal and family monthly income, but we found significant differences as we compared IHS scores by urbanization (F = 5.52, p = .020). People living in urban areas showed higher IHS levels than those living in smaller rural areas. The analysis of Polish community data showed no gender, family monthly income, and urbanization differences. Personal monthly income was associated with IHS scores (F = 3.92, p = .021): as personal monthly income grew, IHS scores were higher

TABLE 1
Means and standard deviations for IHS broken down by country, gender (male/female), urbanization level (rural/urban), personal income (low-medium-high), family income (low-medium-high)

	Italy (Students)			Poland (Students)			Poland (Community)		
	N	*M* (*SD*)	*F*	*N*	*M* (*SD*)	*F*	*N*	*M* (*SD*)	*F*
Gender			4.14*			1.73			1.48
Females	155	5.49 (1.34)		202	4.96 (1.40)		136	4.90 (1.44)	
Males	135	5.80 (1.27)		49	5.28 (1.42)		105	4.69 (1.26)	
Urbanization			0.96			5.52*			2.59
Rural	146	5.56 (1.23)		126	4.82 (1.46)		126	4.67 (1.49)	
Urban	144	5.71 (1.41)		127	5.23 (1.34)		127	4.94 (1.23)	
Personal Income			1.47			0.36			3.92*
Low	96	5.48 (1.44)		84	4.92 (1.44)		80	4.68 (1.51)	
Medium	100	5.63 (1.31)		84	5.07 (1.33)		81	4.60 (1.33)	
High	94	5.81 (1.19)		85	5.08 (1.48)		80	5.15 (1.20)	
Family Income			0.03			1.39			2.17
Low	96	5.61 (1.45)		84	5.00 (1.44)		80	4.55 (1.52)	
Medium	100	5.65 (1.38)		84	4.85 (1.39)		81	4.91 (1.35)	
High	94	5.66 (1.11)		85	5.08 (1.48)		80	4.97 (1.20)	

Note. * p < .05.

CFA Model Fit and Comparison Results

To find a baseline model for MGCFA, we compared different classes of models within each sample and subsample. For Italian students (Panel a, Table 2), the one-factor model (Model 1) yielded acceptable fit indices. The three-factor model with no second-order factor (Model 2) was a good fit. With an overarching factor affecting first-order factors, the second-order model, relationship-oriented happiness, quiescent happiness, and ordinary happiness (Model 3), was equivalent in fit to Model 2. Models 2 and 3 had better fit indices than Model 1, $\Delta SB\chi^2(3) = 12.80$, $p < .001$. The bifactor model, with a single latent factor common to all items and three separated group factors (Model 4), was an excellent fit. It showed a significant improvement in fit indices compared to Model 2 and 3, $\Delta SB\chi^2(6) = 57.36$, $p < .001$. For Polish students (Panel b, Table 2), Model 1 yielded acceptable fit indices. Models 2 and 3 were a good fit. Model 2 and 3 fit indices were significantly better than the ones of Model 1, $\Delta SB\chi^2(3) = 20.60$, $p < .001$. Model 4 was an excellent fit. Model 4 showed a significant improvement in terms of goodness-of-fit indices compared to Model 2 and 3, $\Delta SB\chi^2(6) = 48.58$, $p < .001$. Last, for Polish community participants (Panel c, Table 2), Model 1 was acceptable. Models 2 and 3 were a good fit and improved over Model 1, $\Delta SB\chi^2(3) = 12.35$, $p < .001$. Model 4 was a better fit compared to Models 2 and 3, $\Delta SB\chi^2(6) = 24.90$, $p < .001$. In sum, because the bifactor model outperformed all alternative models, it was used as the configural equality model in MGCFA.

TABLE 2
Fit indices of the hypothesized and alternative models of the Interdependent Happiness Scale

	χ^2	*df*	*p*	CFI	TLI	GFI	IFI	RMSEA [90% CI]	SRMR
Italy students (Panel a)									
Model 1 (one factor model)	83.78	27	.000	.92	.89	.90	.90	.09 [.07, .10]	.06
Models 2 and 3 (three factor model and hierarchical model)	68.07	24	.000	.94	.91	.92	.93	.08 [.06, .10]	.05
Model 4 (bifactor model)	24.91	18	.127	.99	.98	.97	.98	.04 [.00, .06]	.03
Poland students (Panel b)									
Model 1 (one factor model)	99.32	27	.000	.91	.88	.89	.91	.12 [.10, .15]	.06
Models 2 and 3 (three-factor model and hierarchical model)	75.33	24	.000	.94	.91	.91	.94	.09 [.07, .11]	.05
Model 4 (bifactor model)	31.62	18	.024	.98	.97	.96	.98	.05 [.03, .08]	.03
Poland community (Panel c)									
Model 1 (one factor model)	87.08	27	.000	.90	.87	.85	.89	.10 [.09, .12]	.06
Models 2 and 3 (three-factor model and hierarchical model)	73.99	24	.000	.92	.88	.87	.91	.13 [.10, .17]	.06
Model 4 (bifactor model)	47.67	18	.000	.96	.91	.92	.91	.08 [.06, .10]	.05

Note. CFI = comparative fit index; TLI = Tucker-Lewis index; GFI = goodness of fit index; IFI = incremental fit index; RMSEA = root mean square error of approximation; SRMR = standardized root mean square residual.

Multigroup Analyses (MGCFA)

The bifactor model's configuration equality produced fit indices exceeding the recommended minimum, suggesting the number of factors and the factor-loadings were the same across the three samples.

The model constraining the factor loadings to be equal across groups for items on the general factor (FL1 in Table 3) exceeded the recommended standard for acceptable fit. FL1 was not different from the configuration equality model, $\Delta SB\chi^2(18) = 12.91$, $p = .80$. Next, we constrained the factor loadings to be equal for items on the group factors (Model FL2). The analysis yielded an overall good fit. Although FL2 was a significant loss of fit compared to FL1, $\Delta SB\chi^2(18) = 33.03$, $p = .02$, the ΔCFI = .006 was lesser than .01 indicating that the models were not substantially different (Cheung & Rensvold, 2002). The metric invariance of the bifactor model was supported.

TABLE 3
Fit indices for multigroup analysis of the Interdependent Happiness

	χ^2	*df*	*p*	CFI	TLI	GFI	IFI	RMSEA [90 % CI]	SRMR
Configuration equality model	109.07	54	.000	.97	.97	.90	.90	.08 [.06, .10]	.06
Factor loading invariance Model 1 (FL1)	121.83	72	.000	.97	.98	.92	.93	.07 [.05, .09]	.05
Factor loading invariance Model 2 (FL2)	154.15	90	.000	.97	.97	.97	.98	.06 [.05, .08]	.03
Scalar invariance model	355.26	111	.000	.90	.91	.89	.91	.11 [.10, .12]	.06
Partial scalar invariance Model 1 (PS1)	185.38	102	.000	.97	.96	.91	.94	.07 [.05, .08]	.05
Partial scalar invariance Model 2 (PS2)	296.31	102	.000	.91	.92	.96	.98	.10 [.09, .12]	.03

Note. CFI = comparative fit index; TLI = Tucker-Lewis index; GFI = goodness of fit index; IFI = incremental fit index; RMSEA = root mean square error of approximation; SRMR = standardized root mean square residual.

The scalar invariance model yielded an acceptable fit. Again, the $\Delta SB\chi^2(21) = 260.04$, $p = .000$, showed a significant loss of fit for the scalar invariance model relative to FL2 (Table 3). The ΔCFI = .068 confirmed that the two models did not fit the data equally, and the full scalar invariance was rejected. Indeed, we tested the hypothesis that only two groups were different, namely whether the scalar invariance held between Polish students and community subsamples, not imposing equality of intercepts on the Italian sample (partial scalar model, PS1 in Table 3). The fit indices also exceeded the recommended minimum. Although PS1 was a significant loss of fit compared to FL2, $\Delta SB\chi^2(12) = 34.90$, $p = .000$, the ΔCFI = .007 was lesser than .01, indicating that the two models were not substantially different. The scalar invariance of the bifactor model between Polish students and community participants was supported. A subsequent analysis investigated the latent mean differences between the Polish student and community groups (given that they showed scalar invariance). The only significant latent mean difference was for the quiescence factor ($p = .033$). Having fixed to 0 the latent mean of the community group, the value estimate for the student group was +0.25 (Wald = 2.138). Last, we tested the scalar invariance between Italian and Polish students, not imposing equality constraints on the intercepts of the Polish community subsample (partial scalar model, PS2 in Table 3). The fit indices exceeded the recommended minimum. Although PS2 was a significant loss of fit compared to FL2, $\Delta SB\chi^2(12) = 186.84$, $p = .000$, the ΔCFI = .051 was higher than .01, indicating that the models were substantially different. The scalar invariance between countries was rejected.

Reliability Analyses

A psychometric question in the light of MGCFA analyses was whether the harmony with others, quiescence, and ordinariness could give specific information to researchers using subscale scores, in addition to that provided by using the IHS total score. Using the standardized factor loadings from the metric invariance model (Table 4), we calculated the proportion of variance explained by the general factor (ECV) in the three groups. The ECV was .80 in all groups, supporting the use of a total score instead of separate subscale scores (Rodriguez, Reise, & Haviland, 2016).

TABLE 4
Standardized solution for the bifactor model of the IHS

	Item	Factor	GEN loading	HAR loading	QUI loading	ORD loading	Unique variance
Italy (Students)	IH1	HAR	.77	.12			.40
	IH2	HAR	.61	.47			.41
	IH3	HAR	.58	.67			.21
	IH4	QUI	.61		.35		.51
	IH5	QUI	.62		.06		.62
	IH6	QUI	.55		.14		.67
	IH7	ORD	.82			.10	.31
	IH8	ORD	.75			.43	.26
	IH9	ORD	.80			.22	.31
Poland (Students)	IH1	HAR	.78	.12			.38
	IH2	HAR	.56	.44			.49
	IH3	HAR	.63	.72			.09
	IH4	QUI	.54		.32		.61
	IH5	QUI	.67		.06		.55
	IH6	QUI	.65		.17		.55
	IH7	ORD	.82			.10	.32
	IH8	ORD	.73			.42	.29
	IH9	ORD	.78			.22	.34
Poland (Community)	IH1	HAR	.80	.13			.34
	IH2	HAR	.63	.48			.37
	IH3	HAR	.64	.73			.06
	IH4	QUI	.61		.36		.49
	IH5	QUI	.71		.06		.49
	IH6	QUI	.69		.17		.50
	IH7	ORD	.86			.10	.26
	IH8	ORD	.76			.44	.23
	IH9	ORD	.82			.23	.28

Note. IH = interdependent happiness; GEN = general factor; HAR = harmony group factor; QUI = quiescence group factor; ORD = ordinariness group factor.

Furthermore, a fine-grained analysis based on variants of the reliability coefficient ω compared the relative utility of total and subscale scores. The proportion of reliable variance in the total score accounted for by the general IHS factor (ω hierarchical = .86, .86, and .88 for Italian students, Polish students, and Polish community participants, respectively) was very similar in size to the total proportion of reliable variance (ω = .92, .92, and .94, respectively). It follows that the general factor accounted for about 93%, 93%, and 94% of the reliable variance in the total IHS score for Italian students, Polish students, and Polish community participants, respectively. Subsequently, we assessed the reliability coefficients in the three groups separately for harmony (ω = .84, .85, and .89 for Italian students and the two polish samples, respectively), quiescence (ω = .66, .69, and .75, respectively), and ordinariness (ω = .88, .86, and .90, respectively). Partialling out the general factor variance from the subscale scores, the ω hierarchical coefficients fell for harmony (ω hierarchical = .25, .25, and .26 for Italian students and the two Polish samples, respectively), becoming negligible for quiescence (ω = .06, .05, and .06, respectively), and ordinariness (ω = .08, .08, and .08, respectively). Overall, the IHS subscales reliably measured common variance in interdependent happiness, but only harmony preserved a small proportion of specific information not encompassed in the IHS total score.

Validity Analyses

IHS and Life Satisfaction

Metric invariance is enough to compare the correlations across groups. Accordingly, we assessed whether the IHS measured the same happiness as SWLS and the domain-specific life-satisfaction in the three samples. The correlations were calculated using both the summated and hierarchical factors scores. When using summated scores (Table 5a), all the IHS scores were correlated positively with the SWLS and the domain-specific life-satisfactions, and the effect sizes were medium-large. Using hierarchical factor scores (Table 5b), only the general IHS factor was correlated positively with life satisfaction, with a large effect size in all the different sub-samples.

TABLE 5
Zero-order correlations between IHS total score and IHS sub-scales scores with SWLS and SWL-domains sub-scales using IHS summated ratings and factor scores

IHS factor		a) Summated ratings			b) Factor scores		
Country		SWLS	Mater.	Nonmater.	SWLS	Mater.	Non-mater.
Italy (Students)	General	.76**	.46**	.61**[b]	75**	.48**	.59**[b]
	Harmony	.67**	.34**	.60**[b]			.12*[a]
	Quiescence	.53**[a]	.57**	.55**[b]			
	Ordinariness	.52**	.51**	.49**	.20**	.15**	
Poland (Students)	General	.79**	.53**	.67**[c]	.76**	.54**	.62**[c]
	Harmony	.72**	.43**	.66**[c]	.21**		.30**
	Quiescence	.39**[c]	.51**	.50**			.12*
	Ordinariness	.43**	.58**	.50**	.14*		
Poland (Community)	General	.78**	.47**	.49**	.80**	.44**	.42**
	Harmony	.69**	.40**	.38**			
	Quiescence	.54**	.51**	.44**		.15*	.17**
	Ordinariness	.53**	.50**	.42**			

Note. IHS = Interdependent Happiness Scale; SWLS = Satisfaction with Life Scale; SWL = Satisfaction with Life; Mater. = materialistic subscale of SWL-domains; Nonmater. = nonmaterialistic subscale of SWL-domains. [a] pair of correlations among Italy/Poland students which are significantly different; [b] pair of correlations among Italy/Poland community which are significantly different; [c] pair of correlations among Poland students and Poland community which are significantly different.
** $p < .01$. * $p < .05$. Not significant correlations are not reported.

Regarding the life-satisfaction domains, the effect sizes were medium-large, especially in the two student samples and for the nonmaterialistic domain. For Polish communities, the general IHS factor was almost equally associated with both domains of life satisfaction.

In most cases, the associations of IHS subscales with life satisfaction constructs did not differ from those involving the total score (Table 5a). The coefficients fell when using the factor scores (Table 5b), and minor differences also emerged across samples. For instance, harmony was only weakly associated with the nonmaterialistic domain in Italy. In Poland, harmony was not associated with any life satisfaction constructs in the community sample, while it was positively related to the SWLS and the nonmaterialistic domain for students. The quiescence factor score had no significant correlations in Italy, while it was weakly associated with the nonmaterialistic domain in both Polish samples and the materialistic domain in the community sample. Last, ordinariness was not associated with the life satisfaction domains in the two Polish samples, while in Italy and Polish students, it was related to the SWLS. In sum, the summated ratings were almost equally associated with all life satisfaction constructs in all samples, while the factor scores showed divergent patterns of correlation with the life satisfaction measures and by country or sample.

IHS and Cultural Logics

When using summated scores, all the IHS ratings were correlated positively with honor, dignity, and face in all samples (Table 6a). However, while there was a tendency for dignity to be associated the most with the IHS total for the Italian students, both honor and dignity were the most important cultural ideals about interdependent happiness for the Polish communities. The Polish student sample showed less differentiated associations. Similar findings were found using the general factor score (Table 6b).

TABLE 6
Zero-order correlations between IHS total score and IHS subscales scores with honor, dignity and face using IHS summated ratings and factor scores

		a) Summated ratings			b) Factor scores		
Country	IHS factor	Honor	Dignity	Face	Honor	Dignity	Face
Italy (Students)	General	.25**[a, b]	.41**	.33**	.21**[b]	.37**	.31**
	Harmony	.31**[b]	.38**	.26**	.16**		
	Quiescence	.18**[a, b]	.36**[b]	.26**[b]			
	Ordinariness	.19**[b]	.35**[b]	.22**	.13*	.18**	
Poland (Students)	General	.38**[c]	.37**	.30**	.32**[c]	.34**	.28**
	Harmony	.37**	.37**	.25**	.20**	.16**	
	Quiescence	.35**[c]	.27**[c]	.31**	.16*		.30**
	Ordinariness	.29**[c]	.24**[c]	.21**[c]	.17**		
Poland (Community)	General	.54**	.48**	.33**	.53**	.45**	.31**
	Harmony	.48**	.46**	.29**		.17**	
	Quiescence	.57**	.56**	.40**		.14*	.22**
	Ordinariness	.58**	.53**	.38**			

Note. IHS = Interdependent Happiness Scale. Not significant correlations are not reported; [a] pair of correlations among Italy/Poland students which are significantly different; [b] pair of correlations among Italy/Poland community which are significantly different; [c] pair of correlations among Poland students and Poland community which are significantly different.
** $p < .01$. * $p < .05$.

Using the subscales, the correlation analysis was difficult to interpret in terms of specific cultural associations (Table 6a). The only systematic difference was the larger average effect size for Polish community participants than both student samples. The analysis of factor scores showed that harmony was associated only with honor in Italian students, with dignity and honor in Polish students, with dignity only in Polish community participants. The quiescence factor was not associated with any of the cultural logics in Italy, while it was significant with face in both Polish samples. The ordinariness factor was relevant to dignity and honor for Italians and to honor for Polish students. The correlations involving group factors were small, suggesting that the general factor only was associated with cultural logic.

IHS and Self-Construals

Concerning the relation between IHS and self-construal facets, the coefficients were lower than those assessed with cultural logics and life satisfaction constructs. Self-definition was not linked to IHS in Italy and was weakly correlated with IHS in Poland, in both subsamples (see Table 7). Decision-making oriented to the self, self-consistency, and self-expression were weakly correlated with IHS in all subsamples. Self-reliance was weakly correlated only with quiescence in all subsamples and only when considering factor scores. Self-interest was weakly correlated with ordinariness in all subsample and negatively correlated with harmony in the Polish community subsample, but only when considering factor scores. The major difference between scoring methods appeared in the case of the decontextualized self subscale: in general, the strongest association was the one with IHS general factor, but the coefficients decreased remarkably when considering the factor scores. Unexpectedly, the self-containment score was negatively associated with IHS in all subsamples. Foreshadowing the discussion, this might be due to the very low reliability of this scale in all three samples. In general, the self-construal factors resulted in smaller correlations with IHS in Italy and larger in Poland.

TABLE 7
Zero-order correlations between IHS total score and IHS subscales scores with self-construal subscales using IHS summated ratings and factor scores

		a) Summated ratings							
Country		Difference	Contain.	Direction	Reliance	Expression	Interest	Consist.	Decontext.
Italy	General		–.15**a,b	.12*		.13*		.30**	.59**b
	Harmony		–.24**b	.14*		.18**		.31**	.12*a
	Quiescence					.16**		.25**	
	Ordinariness							.17**	
Poland Students	General	.21**	–.36**	.25**		.25**		.22**	.62**
	Harmony	.26**	–.34**	.23**		.29**		.26**	.30**
	Quiescence	.21**	–.40**	.14*		.19**		.18**	.12*
	Ordinariness	.23**	–.34**	.16**		.22**		.22**	
Poland Community	General		–.45**					.28**	.42**
	Harmony	.17**	–.47**			.15*	-.15*	.23**	
	Quiescence		–.45**					.15*	.17**
	Ordinariness		–.39**						

(Table 7 continues)

Table 7 (continued)

		b) Factor scores							
Country		Difference	Contain.	Direction	Reliance	Expression	Interest	Consist.	Decontext.
Italy	General		–.13*[a, b]					.28**	.17**
	Harmony		–.18**	.14*		.13*			
	Quiescence			.15**	.22**				
	Ordinariness						.22**		
Poland Students	General	.14*	–.34**	.17**		.19**		.19**	.17**
	Harmony	.23**	–.20**	.18**		.21**		.20**	
	Quiescence		–.22**	.22**	.26**				
	Ordinariness	.15*		.23**		.15*	.14*		.15*
Poland Community	General		–.42**					.26**	.25**
	Harmony	.19**	–.24**	.14*		.17**	-.16*		.15*
	Quiescence		–.25**	.19**	.21**	.23**		.26**	.13*
	Ordinariness		.20**			.19**	.16*		

Note. IHS = Interdependent Happiness Scale; Contain. = self-containment; Consist. = self-consistency; Decontext. = decontextualized-self. [a] pair of correlations among Italy/Poland students which are significantly different; [b] pair of correlations among Italy/Poland community which are significantly different; [c] pair of correlations among Poland students and Poland community which are significantly different. Not significant correlations are not reported.
** $p < .01$. * $p < .05$.

DISCUSSION

The IHS was developed to capture the interdependent and interpersonal nature of happiness (Hitokoto & Uchida, 2015). Although this scale has been used in diverse socio-cultural and linguistic contexts (Datu et al., 2016; Krys, Capaldi et al., 2019; Krys, Zelenski et al., 2019), no study has considered if the items of the IHS have similar meaning among different cultures. The measurement invariance is essential for the comparison and interpretation of psychological constructs in different groups regarding socio-cultural and linguistic profiles. The current study was designed to address this issue and other measurement concerns regarding the IHS. Because previous research did not agree on the scale's dimensionality, we primarily investigated the factorial structure of the IHS, hypothesizing that a hierarchical model might reconcile previous inconsistent findings. Given the inherently cultural nature of interdependent happiness, we also evaluated the IHS applicability to Italian and Polish samples using a multigroup analysis. The last — and perhaps most important — issue regards whether the scale measures an aspect of happiness different from satisfaction with life.

Regarding the first issue, we showed that a bifactor model was almost a perfect fit in all the subsamples examined. While the original study (Hitokoto & Uchida, 2015) supported the unidimensionality of the IHS, subsequent research cast doubt on whether multiple factors could be the best solution (Datu et al., 2016). The same study alluded to a possible hierarchical arrangement of IHS with three first-order factors and a second-order factor on the top. However, because of model equivalence, the study could not compare the second-order hierarchical model and the standard three-factor model with correlated factors. We overcame the equivalent-fit problem using a bifactor model that captured the variance common to all IHS items and parsed it from that associated with harmony, quiescence, and ordinariness. Previous research pointed out the superior fit of a bifactor model over other classes of models (Bonifay et al., 2016); however, it also recommended caution in interpreting the bifactor model, especially when the general factor has a weak loading pattern (e.g., Heubeck & Wilkinson, 2019). While we agree with this view, the strong loading pattern for the general factor emerging from three samples and two countries supports the bifactor model's interpretation as the best way to represent the IHS's internal structure.

Establishing measurement invariance is crucial for cross-cultural research (Milfont & Fischer, 2010). No previous study has addressed these issues between countries, but only by gender-invariance in a single country (Datu et al., 2016). Our study supported the bifactor model's configural and metric invariance in the Italian and Polish language versions of the IHS. This finding ensures the comparison of correlations involving the IHS and other instruments in the two different cultures and samples. Unfortunately, the scalar invariance, a prerequisite for comparing the latent means obtained in the two countries, was not achieved. Notwithstanding this, the IHS scores were scalar invariant between students and communities within the Polish cultural context. Polish students and community participants did not differ on the general IHS factor. Nor the groups differed on happiness based on the quality of social interaction (harmony) or the extent to which people have a similar level of accomplishment with the people around them (ordinariness). A rather slightly higher level of happiness derived from being able to meet normative expectations in the Polish sociocultural context characterized the youngest group. No previous study has compared student samples to community samples using the IHS or a bifactor model. However, one must interpret this result with caution, considering that the amount of specific variance associated with quiescence happiness after controlling for the general factor was negligible.

Established that the bifactor model was metric-invariant across all samples used, we examined the relative amount of information conveyed in the total and subscale scores (Rodriguez et al., 2016). Our study showed that the use of the total score is more supported than using subscale scores. The general factor accounted for over 90% of the total IHS score's reliable variance for all samples. This finding is consistent with the intents of Hitokoto & Uchida (2015) to develop a common metric for interdependent happiness. The harmony facet was the only subscale score to preserve a small, but not negligible, proportion of specific interdependent happiness, deriving from the perception that one's reference group, or society, is overall happy. Noteworthy, Hitokoto & Uchida (2015) themselves defined interdependent happiness "as basically relationship-oriented, and as a state of harmony with a certain balance being achieved between the self and significant others" (p. 214). This concept has a psychometric match in the reliable variance of the harmony scale in Italy and Poland.

Overall, the analyses carried out so far support the view that the IHS is based on a single well-being construct. Therefore, it was logical to ask what construct it is, and its relationships with life satisfaction, the most widely used well-being measure in cross-cultural research. The high correlation between IHS and SWLS found in previous studies (Arimitsu et al., 2019; Hitokoto & Tanaka-Matsumi, 2014; Krys, Zelenskiet al., 2019) suggested that interdependent happiness and life satisfaction might be empirically overlapping, or, at best, difficult to separate, even if there are sound theoretical arguments to do so. Our study showed that the general IHS factor was correlated in the range .75-.80 with the SWLS, either using the total score or the hierarchical factor scores that reflect all IHS items' common core. Van Mierlo, Vermunt, and Rutte (2009) showed that a correlation between .70 and .85 points to substantial overlap between constructs, but not complete redundancy, a conclusion that might apply to IHS and SWLS, too. After all, interdependent happiness and satisfaction with life are sister constructs, and therefore, it still is entirely possible that higher IHS scores might be associated with higher SWLS ones, especially if one considers the global factor underlying interdependent happiness. Using the subscales led us to a more optimistic conclusion, with correlations of harmony, quiescence, and ordinariness, with SWLS in the range .39-.72, still large but indicating considerable distinctness (Van Mierlo et al., 2009). Using the hierarchical factor scores for harmony, quiescence, and ordinariness, there is no overlap with SWLS.

This study is not free from limitations. In Italy, participants consisted only of undergraduate students and were not representative of the entire population; in Poland, participants were both students and community participants. Also, the three groups' different sociodemographic characteristics and how they were enrolled did not warrant generalizations. These differences might, perhaps, account for the lack of scalar invariance between countries. Future research, with more extensive and stratified randomly selected samples, could be useful to follow up the investigation of the measurement invariance of the IHS and to assess its validity with other measures of well-being, quality of life, and social indicators.

Relatedly, the IHS was devised to be sensitive to differences in independent and interdependent self-construals. However, our study involved only two countries in the middle of the individualism-collectivism (I-C) continuum. This might have prevented us from highlighting country-specific association with the self-construals and cultural logics. Future studies should involve countries that are considered more separated on the I-C continuum.

Last, due to the lack of scalar invariance across countries, IHS's average level differences do not necessarily imply unequal levels of interdependent happiness. There might be a cultural bias in how the different groups responded to the items. For instance, Italian and Polish participants might have a different tendency to social desirability or acquiescence, response style factors which are well-known to affect item responses and indeed item intercepts in CFA (Steinmetz, 2013). Unfortunately, we did not assess these tendencies in the present study; therefore, this issue could be addressed and eventually resolved in IHS studies that will include control scales.

Notwithstanding these limitations, our findings provided support for a conceptualization of interdependent happiness as a hierarchically organized construct; established the metric invariance of the scale between countries so that the correlations of the scale with external criteria can be generalized reliably across countries; showed that the IHS is not redundant with the SWLS, although the two concepts are deeply intertwined. The search for an alternative measure to describe a happy society using the IHS has led to a promising research line focusing on the social aspects of happiness. It is essential to further validate the IHS in different cultures and contexts to explore cross-culturally the social conceptualization of happiness.

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