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Assessing the Cross-Cultural Validity of the Succession, Identity and Consumption (SIC) Scale Across Four French-Speaking Countries

RESEARCH ARTICLE

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ABSTRACT

French-speaking countries are aging fast, forcing them to accommodate their older population and, most likely, generating ageism. The present study aims to investigate this issue by examining the cross-cultural validity of a scale assessing prescriptive ageism: the Succession, Identity, Consumption scale. In four French-speaking countries (Canada, France, Belgium and Switzerland), Confirmatory Factor Analysis (CFA) results reveal the suitability of both a three-factor 15-item model and a higher-order factor model. Multiple-group CFA revealed measurement invariance of both models across countries. MIMIC (Multiple Indicators MultIple Causes) modeling showed that the Canadians obtained the highest ageism scores, followed by the French, Belgians, and Swiss. Second, using dynamic indicators of socioeconomic parameters, we observed that rapid aging populations and additional pressure on the workforce could be viewed as key factors for understanding global ageism as well as succession and consumption-based prescriptions.

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prescriptive ageism; SIC scale; multiple-group confirmatory factor analyses; MIMIC modeling; relative speed of aging; senior dependency ratio increase

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The population of older adults is presently estimated at approximately 900 million in the world. This number will likely rise to 2 billion by 2050, moving from 12% to 22% of the global population (World Health Organization, 2017). Western countries are among those whose populations are aging rapidly, due to factors such as increased life expectancy and decreasing birth rates. In Canada for example, the number of 60-year-old individuals has outnumbered those aged 20 since 2016 (Statistics Canada, 2016). Such demographic shifts are likely to affect intergenerational relationships, notably by intensifying intergenerational tensions stemming from the perception that an aging population puts pressure on limited resources. In turn, such intergenerational tensions may pave the way for ageism (North & Fiske, 2012). Aiming at shedding further light on this relationship, previous research has identified three domains of prescription (Succession, Identity, and Consumption), all relating to these intergenerational or resource-based tensions, and comprising distinct factors of ageism in the Succession Identity Consumption (SIC) scale (North & Fiske, 2013a).

The merits and deficiencies of the SIC scale have been exposed elsewhere (e.g., Hancock & Talley, 2020). Our point is not to elaborate upon it now, except to note that (1) the SIC scale has been developed based solely on American English-speaking samples, in spite of the prevalence of ageism worldwide (e.g., Rychtaříková, 2019, for Europe), and (2) the potential determinants of prescriptive ageism as measured by the SIC scale have not been examined yet. The present study directly addresses both issues, by (1) assessing the generalizability of the SIC scale across four French-speaking countries (Canada, France, Belgium, and Switzerland), and (2) using criteria such as demographic and economic dynamic indicators to better understand the sociostructural roots of prescriptive ageism in these countries.

DEFINING AND MAPPING AGEISM

Ageism is defined as a psychosocial mechanism generated by conscious or unconscious perception of the intrinsic qualities of an individual (or group) in relation to his or her age (e.g., Boudjemadi & Gana, 2009). The underlying process is implicit and/or explicit and is expressed individually or collectively through discriminatory behavior, prejudices, and stereotypes that may be positive but are most often negative. Stereotypes are defined as cognitive structures containing shared knowledge, beliefs, and expectancies about social groups (Hamilton & Sherman, 1994). The Stereotype Embodiment Theory (Levy, 2009) proposes that stereotypes are internalized across the lifespan and are present into every age category, even among preschool children (Flamion et al., 2020). Once internalized, these stereotypes exert their influence in multiple ways (e.g., Levy et al., 2009). In other words, ageism is a particularly pervasive, detrimental, insidious, and deleterious phenomenon leading to many negative outcomes in various contexts (Nelson, 2016; see also WHO, 2021): ageism can lead to mistreatment, neglect, and abuse of older adults (McDonald, 2017), impact the mental and physical health of older adults as well as their overall quality of life in many ways (Chang et al., 2020; Mikton et al., 2021; Officer et al., 2020), or impair older adults' cognitive and functional performance and shorten their lifspan (Ayalon et al., 2019).

Stereotypes encompass two components: a descriptive one reflecting how groups are typically perceived, and a prescriptive one that reflects how groups 'should be' and how they 'should not be'. Each component has distinct features. First, prescriptive stereotypes tend to differ across groups while descriptive stereotypes are widely supported (Burgess & Borgida, 1999; North & Fiske, 2012). Second, activating prescriptive stereotypes has long-term negative consequences (e.g., persistent negative judgment about a target), which is not the case when descriptive stereotypes are activated (Gill, 2004). Third, prescriptive stereotypes play an important role in maintaining the social status quo (while descriptive ones just describe social groups), allowing higher societal status groups to protect and preserve their privileges over lower societal status groups (Delacollette et al., 2010; North & Fiske, 2013b). Controlling prescriptions about dominated groups directly benefits dominant groups, to such an extent that members of dominated groups violating prescriptions face backlash while those adhering to them earn reward (Pratto et al., 1997; Rudman et al., 2012).

DOMAINS OF RESOURCE-BASED AGEIST PRESCRIPTIONS

Age groups (young people, middle-aged adults, older people) have common interests in as much as they are interdependent in a variety of domains such as family life or work. For example, some working adults may have to provide for their children or parents, while others are financially dependent on their parents. At a societal level, promoting young people employment may reduce their dependency on their parents and enhance their productivity, contributing as such to supplying retirement pensions (Dykstra & Hagenstad, 2016; Williamson et al., 2003). Meanwhile, age groups do not hold equal status and take turn at different levels of privilege, reaching different levels of societal resources (North, 2019). In modern societies, young people often have fewer resources. As they grow up, middle-aged people acquire more resources (e.g., prestige, assets, leadership positions, or influence), before surrendering them when they reach an older age. In other words, middle-aged individuals are perceived as benefiting from a higher status while younger and older people have a lower status (North & Fiske, 2012, 2013b). This imbalance in access to resources creates intergenerational tensions of which older adults are the main victims, at least in Western societies in which young people are generally perceived as an investment or a potentiality, while elders represent a debt or a burden (Bergman, 2017). A rapidly aging population seems to intensify this perception that older people are undermining the economy and monopolizing resources. Moreover, young people seem to believe that older people have already been in a dominant position in the past and that they must step aside and make way for the next generation (North, 2019; North & Fiske, 2012).

These findings provided North and Fiske (2012) with a starting point for offering a theoretical framework on prescriptive ageism in which three general types of tensions were identified. First, the Succession domain refers to tensions over enviable resources, status, societal positions, or influence. Here, ageism is based on the view that older adults have had too much influence and have accumulated wealth, hence, they should step aside and pass on these resources to younger generations. Second, the Consumption domain relates to tensions over the passive depletion of shared societal resources. In other words, older adults should avoid overusing public resources (such as government healthcare resources) or family resources (energy, ownership) at the expense of younger individuals. Third, the Identity domain refers to symbolic tensions over symbolic or figurative resources. Here, ageism is based on the view that older adults should not intrude into younger adults' culture, context, and territory. Over the last decade, findings of empirical studies have supported this resource-based ageist prescription framework. For example, studies that used vignettes describing adherence to or violation of prescriptive expectations in each domain (Succession, Identity, and Consumption) by younger, older, or middle-aged people found that (1) older targets triggered the most positive reactions when adhering to prescriptive expectations but also the most negative reactions when violating them; (2) younger participants emerged as the primary perpetrators of age-based prescriptions; that is, they were the most polarized toward older targets, rewarding them the most for prescription adherence and punishing them the most for violation (as participants grew older, they were more forgiving of older targets' violation of prescriptive expectations) (North & Fiske, 2013b). In a series of more recent experiments, North and Fiske (2016) investigated the impact of resource scarcity on intergenerational exclusion of older workers. In the first three studies, the authors recruited younger participants and manipulated the perceived availability of resources (Succession, Identity, and Consumption) between generations. Results showed that resource scarcity generates a backlash effect, assessed here as negative views and avoidance of older workers who violate prescriptions compared with those adhering to them. A fourth study assigned participants to the role of a manager allocating funding between younger, middle-aged, and older (equally qualified) workers. Results showed that participants, in particular the younger ones, were more likely to deprive the older workers of resources at the benefit of others.

MEASURING RESOURCE-BASED AGEIST PRESCRIPTIONS

Most of the instruments assessing ageist stereotypes focus on content-based, descriptive stereotypes (for a review of available ageism scales, see Ayalon et al., 2019). For instance, the Fraboni Scale of Ageism (FSA; Fraboni et al., 1990; Rupp et al., 2005) measures the cognitive and affective components of ageism. These descriptive measures mainly assess beliefs about what older people are, but not the expectations about what they should be. Hence, descriptive scales have largely overlooked the role of more controlling, prescriptive beliefs in ageism.

To the best of our knowledge, only one scale directly assesses ageism from the prescriptive standpoint: The aforementioned Succession, Identity, and Consumption Scale (SIC; North & Fiske, 2013a). Theoretically speaking, the original American SIC items and subscales were developed with reference to the three key domains in which older people are expected to relinquish the shared and limited social resources and privileges for the benefit of middle-agers and younger people, each likely to lead to intergenerational tensions if not met (North & Fiske, 2012). Methodologically speaking, North and Fiske (2013a) developed and validated the SIC measure using both exploratory and confirmatory factor analyses. In addition, they showed demographic differences in ageism in terms of respondents' gender, age, and ethnicity. However, as recently pointed out by Hancock and Talley (2020), this initial validation did not include any measurement invariance analyses for the SIC scale, thus casting doubt on the appropriateness of observed demographic differences. To address this shortcoming, Hancock and Talley (2020) tested and showed measurement invariance of the SIC scale across gender and ethnic groups with a sample of US students. Using both theoretical and statistical criteria, they also suggested some modifications, recommending moving one item from the consumption subscale to succession subscale ('Older people shouldn't be so miserly with their money if younger relatives need it') and deleting one other item ('Older people probably shouldn't use Facebook').

Thus far, the SIC scale benefited from advanced statistical techniques (e.g., exploratory and confirmatory factor analyses, measurement invariance analyses for gender and ethnicity) during its development and validation. However, all analyses have been conducted only in the US context with English-speaking populations; yet issues concerning intergenerational tensions are particularly salient in Western societies as a whole, including French-speaking countries, raising the need for a similar instrument in other languages than English.

AGEISM IN WESTERN SOCIETIES AND FRENCH-SPEAKING COUNTRIES

Ageism is prevalent in Western societies (Fiske, 2017; Rychtaříková, 2019), as well-documented by findings from various ageism surveys conducted in Canada and Europe. In Canada, 63% of respondents aged 66 and older indicated that they had been treated unfairly or differently because of their age (Revera Inc., 2012). Half of the sample identified ageism as the most tolerated social prejudice, as compared to gender or race-based discrimination. From 2009 to 2018, the rate of police-reported violence by family members against older adults increased by 11%, while spousal violence and violence against children remained stable or even decreased during the same period (National Institute on Ageing, 2020; Statistics Canada, 2019). In Europe, findings from the European Social Survey (ESS, 2020) revealed that 44% of participants believed ageism was a serious problem for society; 39% of participants reported having experienced disrespect, and 29% thought they had been treated badly because of their age. Noteworthily, the number of people who declared having experienced unfair treatment because of their age was higher (35%) than those declaring having experienced unfair treatment based on their gender (25%) or ethnicity (17%).

Focusing on French-speaking countries and French provinces, similar findings arise. Findings from the ESS showed that 54% of Belgian, 48% of French, and 42% of Swiss respondents reported having been a victim of ageism (Abrams & Swift, 2012). In these countries, ageism is more experienced than sexism and racism (Abrams et al., 2011; Swift et al., 2018). In Belgium, Schroyen et al. (2015) found that older people are more likely to be excluded from clinical trials and undertreated while undergoing oncology treatment. A series of experiments conducted in Switzerland with human resources professionals and university students indicated that older candidates were considered as warmer but less competent than younger candidates by employers. Consequently, the former were less often chosen for the job even when the position required the person to be warm (Krings et al., 2011*;* cited by Froidevaux & Maggiori, 2020). Finally, experiments with French university students also showed that older people faced denial of their full humanness, while younger people did not (Boudjemadi et al., 2017).

Despite its prevalence, ageism is still under-researched compared with other intergroup constructs such as racism or sexism (North & Fiske, 2012), especially in European French-speaking countries where the concept of ageism is not yet well known. Consequently, there are very few instruments assessing descriptive stereotypes and no instrument assessing prescriptive stereotypes in French (Boudjemadi & Gana, 2009); besides, there is no empirical research about the determinants of prescriptive ageism in these countries.

DETERMINANTS OF AGEISM

Various contextual factors like legislation, demography, and socioeconomic conditions have been examined for illuminating our understanding of ageism (e.g., North & Fiske, 2012) as well as ageism differences across countries (Abrams & Swift, 2012; Swift et al., 2018). For example, in their systematic review, Marques et al. (2020) identified two determinants as robustly associated with ageism at a societal level: the availability of resources in society and the percentage of older people in the country. Interestingly, however, a cross-cultural study involving 26 countries from Eastern and Western societies did not find any relation between ageism and demographic trends like the average percentage of older people (Löckenhoff et al., 2009). Such demographic criteria are static indicators that might fail to capture the rapidly changing dynamics underpinning intergenerational tensions, North and Fiske (2015) observed, leading them to favor more dynamic indicators of sociostructural parameters like the relative speed of aging (RSA) and the senior dependency ratio increase (SDRI). RSA represents the increase of the relative proportion of those over 65 years old in the country. It can be measured as the average annual percentage growth of over 65 relative to the total population. SDRI is a measure of the pressure on the productive population (the labor force group) and reflects the number of individuals (65 and over) who are likely to be economically dependent from those typically in the labor force (aged 15 to 64). A low ratio indicates there are likely enough people working to support the dependent population (e.g., better pensions and better health care for citizens) while a high ratio implies more financial stress on working people and possible political instability (Simon et al., 2012). SDRI can be measured as the average annual increase in number of people aged 65 and above compared with people in the labor force. In their cross-cultural meta-analysis of modern attitudes toward older adults, North and Fiske (2015) showed that in comparison with Western societies, greater rates of ageism have been found in Asian countries and were associated with a greater acceleration of aging and unprecedented age burden forms for Asian families and societies.

This suggests that rapidly aging populations and additional pressure on the workforce create a societal strain, forcing people to adapt to the context in an abrupt way and ultimately enhancing older people derogation (North & Fiske, 2015). As noted above, Western societies are aging fast, faster than they have ever been, forcing them to accommodate their older population and, most likely, generating ageism. Whether and – if any – to what extent dynamic indicators of sociostructural parameters are associated with prescriptive ageism (as measured by the SIC scale) in these countries remains an open question that needs an answer.

OBJECTIVES AND HYPOTHESES OF THE PRESENT STUDY

Our first objective was to examine the generalizability of the SIC scale across four French-speaking countries, namely: Canada, France, Belgium, and Switzerland. Our approach involved testing the factorial validity, construct validity, and reliability of the SIC scale for each country; as well as testing the multigroup invariance of the SIC scale across countries. More precisely, measurement invariance – which ensures the suitability of the scale for measuring the same construct in the same way in different countries – was first tested as a prerequisite for the interpretation of population heterogeneity, that is, between-group differences in latent factor means (Sass, 2011). In other words, measurement invariance implies that 'constructs are fundamentally the same in each sociocultural group, and thus comparable. Under this condition, hypotheses about the nature of sociocultural differences and similarities can be confidently and meaningfully tested' (Little, 1997: 53).

Our second objective was to explore the relationships between prescriptive ageism and dynamic indicators of sociostructural parameters for the purpose of illuminating ageism differences and similarities across the four countries. Based on the World Bank World Development Indicators (2020), we first computed the relative speed of aging (RSA) and the senior dependency ratio increase (SDRI) for France, Canada, Switzerland, and Belgium for one generation (from 1981 to 2018). RSA (measured as the average annual increase in percentage of people aged 65 and above) was 0.21 for Canada, 0.17 for France, 0.13 for Switzerland, and 0.12 for Belgium. SDRI (measured as the average annual increase in number of people aged 65 and above compared with people in the labor force aged 15 to 64) was 0.31 for Canada, 0.29 for France, 0.20 for Belgium, and 0.19 for Switzerland. Second, we compared both of them to a global score of ageism as well as to each domain of prescription (the Succession, the Identity, and the Consumption) as measured by the SIC scale.

Following North and Fiske's (2015) reasoning, we hypothesized that the worse a country situation in terms of both RSA and SDRI, the more it would endorse prescriptive ageist stereotypes; concretely and based on the country-level indicators, our hypothesis (#1) was Canada > France > Belgium ≈ Switzerland for global ageism. Zooming in on each dynamic indicator, we expected that RSA would match the succession domain in particular, because a higher speed of aging is likely to exacerbate tensions over status and societal positions that (a larger number of) older adults hold and refuse to pass on. Hence, our hypothesis (#2a) was Canada > France > Belgium ≈ Switzerland for the succession domain. We also expected that SDRI would especially fit the consumption domain, because a higher senior dependency is likely to strengthen tensions over shared societal resources that (a larger part of) non-productive population is using in full sight of the productive population. In concrete terms, our hypothesis (#2b) was Canada ≈ France > Belgium ≈ Switzerland for the consumption domain. No specific expectation was made about the identity domain mainly because no sociostructural parameter was adjusted to it.

RESEARCH DESIGN AND METHOD

PARTICIPANTS AND PROCEDURE

Participants were undergraduate students enrolled in introductory psychology courses and recruited on a voluntary basis. The sample consisted of 1,827 French-speaking students from Canada ($n = 433$), France ($n = 575$), Belgium ($n = 434$), and Switzerland ($n = 385$). There was a high degree of similarity between the four subsamples regarding their gender ratio and age. Canadian respondents predominantly identified as female (78%) and had a mean age of 19.66 years ($SD = 2.82$, range = 17–38). Of the French sample, 73% identified as female; the mean age was 20.17 years ($SD = 2.19$, range = 16–33). The Belgian sample was 84% female with a mean age of 25.31 years ($SD = 2.90$, range = 22–46). Of the Swiss sample, 72% identified as female and the mean age was 22.56 years ($SD = 3.14$, range = 17–47).[1]

Participants completed a newly translated French version of the SIC scale (North & Fiske, 2013a) in electronic format. The questionnaire was completely anonymous, and no participant was identifiable. Apart from gender and age, no personal data was requested. At the time of data collection, the majority of European universities involved did not require ethics committee in social science. However, we followed the recommendations of such a committee regarding the non-retention of IP addresses in the data files.

MEASURE: THE FRENCH SIC SCALE

Team members proceeded with the French translation of the scale following best-practice guidelines in cross-cultural methodology, that is, independent/blind/back translation, educated translation, small-scale pretests (ITC Guidelines for Translating and Adapting Tests, ITC, 2017). The translation process was led by a bilingual translator and sent to another bilingual translator for back-translation. Following discussions and experts' consultation, we excluded three problematic items (one per dimension) from the original questionnaire, resulting in a 17-item self-reported measure of ageism. These specific items were: 'AARP (American Association of Retired Persons) wastes charity money' (C7 in the original study) because there is no equivalent association in France, Belgium, Canada, or Switzerland; 'Most older workers don't know when it's time to make way for the younger generation' (S4) and 'Older adults typically shouldn't go to places where younger people hang out' (I1) were both identified as highly redundant with other items in the original study (see North & Fiske, 2013a, for details). Additionally, based on Hancock and Talley's (2020) study, we substituted 'social network' for 'Facebook' in the item 'Older people probably shouldn't use Facebook' (I4).[2] We tested the clarity of the remaining 17 items among a sample of 20 French citizens (50% female; mean age = 32 years, *SD* = 15.07, range = 18–65). Clarity was measured on a 6-point scale ranging from 1 (not clear at all) to 6 (very clear). All items were rated as rather clear (*M* = 5.35; *SD* = 0.35; range = 4.20–5.75).

As in the original study, responses were measured on a 6-point scale ranging from 1 (strongly disagree) to 6 (strongly agree). Higher scores indicated stronger endorsement of prescriptive ageist stereotypes.

DATA ANALYSIS

All analyses were performed using M*plus* (Muthén & Muthén, 1998–2012). Inspection of the data from each country revealed few missing data (from 0.05% to 1% of the item scores), resulting in a very good minimum coverage of .926 or more for any item. The 'MISSING=ALL' option in M*plus* was used for handling them. Test for multivariate skewness and kurtosis revealed departure from normality. For each country, both multivariate skewness and multivariate kurtosis tests of fit were significant (at $p < .01$). Precisely, between one (Canada) and four (Switzerland) items exhibited a skewness value and/or a kurtosis value of 2.0 or more. As recommended in such cases (e.g., Brown, 2006), we relied on a robust Maximum Likelihood estimator for all analyses (abbreviated MLR in M*plus*), which provides test statistics that are robust to non-normal and missing data.

To assess the factorial structure and model fit to the data for each country, we first examined single-group baseline models using Confirmatory Factor Analyses (CFA). For each CFA, ageism was partitioned into three latent subscales which were defined by their respective indicators as originally indicated by North and Fiske (2013a) (minus the three excluded items), and correlations between latent factors were permitted. To assess the reliability of each factor, we used Rho (ρ) coefficient,[3] for which values over .50–.60 are considered as acceptable in the literature (e.g., Raines-Eudy, 2000). In addition, just as North and Fiske (2013a), we compared the three-factor SIC model with two competing configurations: (1) a one-factor model specifying paths from a single factor of Ageism to each of the retained items, and (2) a higher-order factor model in which a broader dimension of Ageism was specified to account for the pattern of relationships among the three first-order SIC factors.[4]

To test for multigroup invariance across countries, we used a combined strategy that implemented both the Multiple-Groups CFA (MGCFA) approach and the MIMIC modeling (Jöreskog & Goldberger, 1975), also referred to as CFA with covariates. As a first step, MGCFA was used to test for *measurement invariance*; that is, whether there was equivalence of all measurement parameters of the model across countries. Measurement invariance was tested with a least restricted solution – that is, equal factor structure across groups (or *configural* invariance) – first evaluated. Nested models were then used to test subsequent models that maintained previous equality constraints and incrementally added more restrictive constraints across groups in the following order: equal factor loadings (or *weak factorial* invariance), and equal indicator intercepts (or *strong factorial* invariance).[5] If the imposition of additional constraints results in a non-significant degradation in the fit of the model, then there is evidence for invariance (Brown, 2006). As a second step, MIMIC modeling was used to test for *population heterogeneity*; that is, whether the latent factor means differ from one country to another. Although MGCFA also allows testing for group concordance of factor means, we opted for MIMIC modeling in the light of three important advantages: Unlike MGCFA – which entails the simultaneous analysis of two or more measurement models (one per group), MIMIC methodology involves a single measurement model. As a consequence, MIMIC modeling is more parsimonious, more powerful (smaller sample sizes are required), and, most importantly, easier to implement especially when the number of groups exceeds two (see Brown, 2006, for details). Regarding the MIMIC procedure, we first checked that the model was reasonable and good fitting in the full sample, collapsing across groups (i.e., the country covariate was not included in this step); next, we added the country covariate (representing group membership) to the model to examine its direct effects on the latent factors. A significant direct effect of the covariate on one or more latent factors would indicate population heterogeneity; that is, a difference in the factor(s) means at different levels of the covariate, between one country and another (Brown, 2006).[6] Effect size of mean differences between countries was measured using Cohen's *d* (*d* = .20, .50, and .80 for small, medium, and large effects, respectively; Cohen, 1992).

To assess the overall fit of the models, we reporte the MLR chi-square statistic (MLR χ^2) of each model (a non-significant value indicates a well-fitting model).

However, because this statistic is sensitive to sample size (Brown, 2006), we also report alternative fit indices (and cut-off values) recommended by Hu and Bentler (1999): Comparative Fit Index (CFI), Tucker-Lewis fit Index (TLI) (acceptable to good fit when both values are close to .90 and .95, respectively), Standardized Root Mean Square Residual (SRMR; excellent fit when value is .05 or lower), and Root Mean Square Error of Approximation (RMSEA; excellent fit when value is .05 or lower) along with its 90% confidence interval (the lower and upper bounds of the RMSEA's 90% CI should be lower than .05 and .10, respectively; Chen et al., 2008). In addition, to evaluate the difference in model fit, we used a scaled difference in MLR χ^2 test because the difference between two MLR χ^2 values for nested models is not distributed as χ^2 (Muthén & Muthén, 1998–2012). If the difference test is non-significant, the fit of the two models is considered as equivalent. To test for multigroup invariance, we used fit criteria proposed by Cheung and Rensvold (2002). Evidence for configural invariance is found when RMSEA for the configural model is below .093; for the subsequent models, evidence for invariance is based on a decrease in CFI of less than .010.

RESULTS

BASELINE MODEL FOR EACH COUNTRY

The first set of analyses evaluated baseline models separately in each country. In every country, the CFA revealed issues with two items, both from the Consumption subscale. The first problematic item was 'Older adults shouldn't be so miserly with their money if younger relatives need it' (C5). In each sample, this item obtained the lowest – and less than acceptable – loading on its intended factor (recommended cut-off > .40; Velicer et al., 1982). The second problematic item was 'Older adults don't really need to get the best seats on buses and trains' (C6). This item showed (1) very low loadings on its intended factor (< .25) in the Swiss and Belgian samples and (2) consistent correlations with several items from different subscales in the French and Canadian samples. In addition, preliminary descriptive analyses revealed it was the most skewed item (value > 2) and associated with the highest kurtosis value (> 5) regardless of the country. We therefore decided to exclude both items from the model and to test for a three-factor SIC model consisting of 15 items (**Table 1**). **Figure 1** presents this model.[7]

For each country, results supported a three-factor model of prescriptive ageism. As shown in **Table 2**, fit indices indicated acceptable to good model fit across countries. All standardized factor loadings were higher than .40 (except for two and three loadings in the Canadian and Belgian samples, respectively) and significant at $p < .01$. They ranged from .48 to .74 in the French sample, from .34 to .65 in the Canadian sample, from .40 to .75 in the Swiss sample, and from .34 to .72 in the Belgian sample. Cross-scale correlations showed that the SIC subscales were positively as well as moderately to highly correlated

ITEMS	SUBSCALES
La société progresserait probablement plus rapidement sans les personnes âgées qui s'opposent au changement.	Succession (S1)
Les personnes âgées ont des pouvoirs politiques trop importants en comparaison à la jeune génération.	Succession (S2)
La plupart des personnes âgées ne savent pas quand laisser la place aux jeunes.	Succession (S3)
Les personnes âgées sont souvent trop bornées pour réaliser qu'elles n'arrivent plus à faire les choses comme avant.	Succession (S4)
Les jeunes sont généralement plus productifs au travail que les personnes âgées.	Succession (S5)
Les promotions accordées aux travailleurs âgés ne devraient pas être basées sur leur expérience mais plutôt sur leur productivité.	Succession (S6)
Il est injuste que les personnes âgées puissent prendre part à des votes relatifs à des sujets qui concerneront bien plus les jeunes.	Succession (S7)
Les personnes âgées n'ont rien à faire dans les lieux fréquentés par les jeunes.	Identité (I1)
En général, les personnes âgées ne devraient pas aller en boîte de nuit.	Identité (I2)
Les personnes âgées n'ont rien à faire sur les réseaux sociaux.	Identité (I3)
Les personnes âgées ne devraient même pas essayer d'avoir l'air cool.	Identité (I4)
Les médecins passent trop de temps à soigner les personnes âgées malades.	Consommation (C1)
Les personnes âgées coûtent trop cher au système de santé.	Consommation (C2)
Les personnes âgées représentent souvent un poids trop important pour leurs familles.	Consommation (C3)
Passé un certain âge, le meilleur bénéfice que les personnes âgées puissent apporter à la société est de léguer leurs biens.	Consommation (C4)

Table 1 French version of the SIC scale.

Note: Responses are measured on a 6-point scale ranging from 1 (strongly disagree) to 6 (strongly agree).

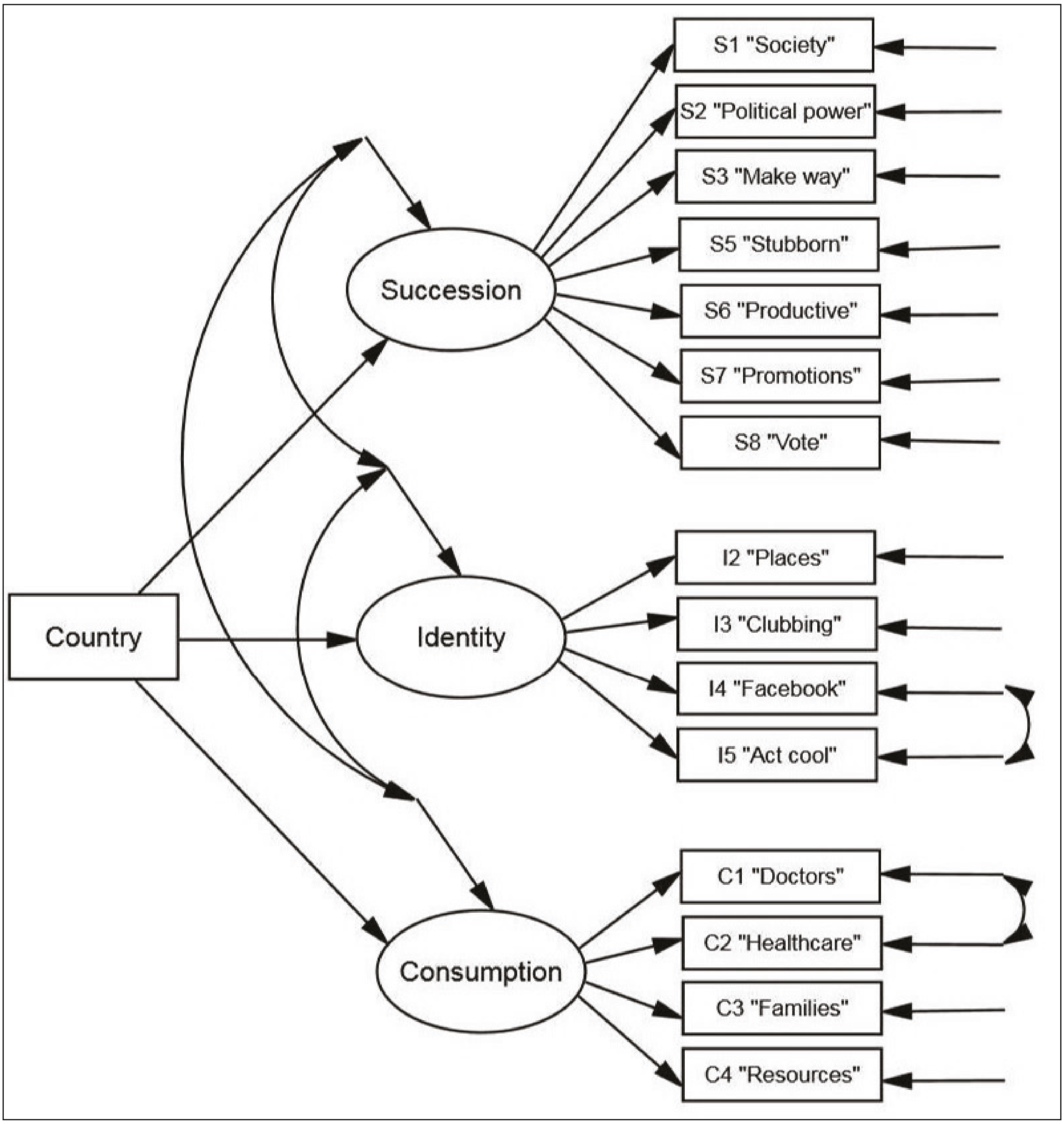

Figure 1 Three-factor model for the 15-item SIC (see North & Fiske, 2013a, Table 1, for the full wording of each original item; and Table 1 of the present article for the French translation).

Note: Inclusion of the country covariate is referred to as the MIMIC version of the model (paths from the covariate to latent factors represent the direct effects of the covariate on these factors; residual variances of the three factors are specified to be correlated because Succession, Identity, and Consumption are not completely orthogonal, and this overlap cannot be fully accounted for by the country covariate).

to each other. They ranged from .62 to .76 for the French sample, from .77 to .84 for the Canadian sample, from .46 to .83 for the Swiss sample, and from .55 to .92 for the Belgian sample. In any country, the lowest association was between Identity and Consumption and the highest association was between Succession and Consumption.[8] Scale reliability as estimated by the Rho coefficient was good in the French sample (ρ = .74 for Succession, .76 for Identity, .73 for Consumption), and acceptable to good in the Canadian (ρ = .69 for Succession, .64 for Identity, .62 for Consumption), Swiss (ρ = .73 for Succession, .70 for Identity, .61 for Consumption), and Belgian samples (ρ = .65 for Succession, .72 for Identity, .55 for Consumption).

In line with the original study, we also tested a competing one-factor model that consisted of a single dimension of Ageism comprising all 15 items. Across the four samples, the one-factor model provided a poorer fit to the data with values for CFI, TLI, SRMR and RMSEA along with its 90% confidence interval mostly outside the recommended cut-offs. In each country, comparison of the one-factor model with the three-factor model yielded a significant scaled difference in MLR χ^2, indicating a better model fit for the three-factor solution compared to the one-factor solution. Fit statistics for these models are also shown in ***Table 2***.

Finally, the pattern of correlations between subscales provided empirical support to proceed with a higher-order factor analysis. In this analysis, a single second-order factor indicating a general tendency for Ageism was specified to account for the correlations among the SIC subscales.[9] Across countries, as expected, the fit of the higher-order factor model was strictly equivalent to the fit of the three-factor model.[10] Besides, what speaks to the viability of a second-order factor of Ageism (alongside the Succession, Identity, and Consumption subscales), was its excellent reliability (ranging from .84 in the Swiss sample to .92 in the Canadian sample) and the large size of the higher-order factor loadings (ranging from .52 to .98 depending on subscales and countries).

MULTIGROUP INVARIANCE ACROSS COUNTRIES

Having ensured the three-factor structure of the scale in all samples, we carried on to conduct MGCFA and test for *measurement invariance* which, if confirmed, would allow to validly compare the latent means of the Succession, Identity, and Consumption subscales as well as the broader dimension of Ageism across countries. ***Table 3*** (MGCFA part) shows the goodness-of-fit statistics and invariance tests for the three- and higher-order factor models.

For both models, results provided evidence for configural invariance (good model fit indices and RMSEA value less than .093 cut-off), weak factorial invariance (good model fit indices and decrement in CFI less than .010), and *partial* strong factorial invariance. Indeed, the *full* strong factorial invariance failed to be established (poor fit indices and decrement in CFI higher than .010). However, relaxing the equality constraints for one intercept in the French group, two intercepts in the Canadian group, four intercepts in the Swiss group, and two intercepts in the Belgian group (at the three-factor model level), along with for one intercept in each country (at the higher-order factor model level) yielded a non-significant degradation in the fit compared to the weak factorial invariance model (good fit indices, decrement in CFI less than .010). It should be noted here that partial (strong factorial) invariance neither is unusual with such complex models (involving several latent factors) nor prevents invariance evaluation to continue. Byrne, Shavelson, and Muthén (1989) argued that the meaningfulness of latent mean comparisons is ensured as long as at least two invariant parameters per construct remain.

Having established (partial) strong invariance across countries, we then used the MIMIC methodology to test for *population heterogeneity*; that is, whether the latent factor means differ from one country to another one. ***Table 3*** (MIMIC part) shows the goodness-of-fit statistics for the three- and higher-order factor models (first excluding and then including the country covariate in models). Both models provided a good fit to the data. All

MODEL	MLRχ^2(*df*)	CFI	TLI	SRMR	RMSEA (90% CI)	ΔMLRχ^2(Δ*DF*)
France						102.03(3)***
Three-factor model, $\theta_{I4-I5}, \theta_{C1-C2}$	180.58*** (85)	.940	.926	.040	.044 (.035–.053)	
One-factor model, $\theta_{I4-I5}, \theta_{C1-C2}$	345.01*** (88)	.839	.808	.061	.071 (.063–.079)	
Canada						21.93(3)***
Three-factor model, $\theta_{I4-I5}, \theta_{C1-C2}$	140.27*** (85)	.938	.923	.043	.039 (.027–.050)	
One-factor model, $\theta_{I4-I5}, \theta_{C1-C2}$	165.52*** (88)	.913	.896	.047	.045 (.034–.056)	
Switzerland						60.63(3)***
Three-factor model, $\theta_{I4-I5}, \theta_{C1-C2}$	140.92*** (85)	.937	.923	.049	.041 (.029–.053)	
One-factor model, $\theta_{I4-I5}, \theta_{C1-C2}$	248.34*** (88)	.820	.786	.065	.069 (.059–.079)	
Belgium						109.09(3)***
Three-factor model, $\theta_{I4-I5}, \theta_{C1-C2}$	159.94*** (85)	.913	.892	.047	.045 (.034–.056)	
One-factor model, $\theta_{I4-I5}, \theta_{C1-C2}$	263.93*** (88)	.795	.755	.062	.068 (.059–.077)	

Table 2 Fit indices for the CFA of various SIC models across countries.

Note: θ_{I4-I5}, θ_{C1-C2} = residuals of items I4 and I5, and residuals of items C1 and C2, respectively, are allowed to covary. Fit indices of the higher-factor model were not shown because they are strictly equivalent to those of the three-factor model. MLR χ^2 = Robust Chi-square; *df* = degree of freedom; CFI = Comparative Fit Index; TLI = Tucker-Lewis Index; SRMR = Standardized Root Mean Square Residual; RMSEA = Root Mean Square Error of Approximation; CI = Confidence Interval for RMSEA; ΔMLRχ^2 = scaled difference between the three-factor and the one-factor model MLR χ^2 values for each country. *** $p < .001$.

MODEL	MLRχ^2(*df*)	CFI	TLI	SRMR	RMSEA (90%CI)	$CFI_{decrease}$
MGCFA (measurement invariance)						
Three-factor model						
Configural invariance	621.16*** (340)	.934	.919	.044	.043 (.037–.048)	—
Weak factorial invariance	696.85*** (376)	.925	.916	.052	.043 (.038–.048)	.009
Strong factorial invariance	965.81*** (412)	.871	.868	.060	.054 (.050–.059)	.054
Partial strong factorial invariance, $\tau_{I4(F)}$, $\tau_{S3(C)}$, $\tau_{S5(C)}$, $\tau_{S1(S)}$, $\tau_{S2(S)}$, $\tau_{I5(S)}$, $\tau_{C2(S)}$, $\tau_{S1(B)}$, $\tau_{I3(B)}$ free	759.37*** (403)	.917	.913	.055	.044 (.039–.049)	.008
Higher-order factor model						
Strong factorial invariance	1021.17*** (418)	.859	.858	.064	.056 (.052–.061)	.066
Partial strong factorial invariance, $\tau_{I4(F)}$, $\tau_{S3(C)}$, $\tau_{S5(C)}$, $\tau_{S1(S)}$, $\tau_{S2(S)}$, $\tau_{I5(S)}$, $\tau_{C2(S)}$, $\tau_{S1(B)}$, $\tau_{I3(B)}$ free + $\tau_{Cons(F)}$, $\tau_{Succ(C)}$, $\tau_{Ident(S)}$, $\tau_{Succ(B)}$ free	760.44*** (405)	.917	.914	.055	.044 (.039–.049)	.008
MIMIC models (population heterogeneity)						
Three-factor model						
Country covariate not included (full sample)	336.30***(85)	.941	.927	.034	.040 (.036–.045)	—
Country covariate added, $\beta_{S1(F-B)}$,						
$\beta_{S3(F-C)}$, $\beta_{S5(F-C)}$, $\beta_{C2(F-S)}$ estimated	446.62***(117)	.933	.914	.032	.039 (.035–.043)	—
Higher-order model						
Country covariate added, $\beta_{S1(F-B)}$,						
$\beta_{S3(F-C)}$, $\beta_{S5(F-C)}$, $\beta_{C2(F-S)}$ estimated	485.19*** (123)	.926	.910	.035	.040 (.036–.044)	—

Table 3 Fit indices and invariance across countries for MGCFA and MIMIC models of the three-factor SIC model and the higher-order factor ageism model.

Note: Fit indices for configural invariance, weak factorial invariance, and MIMIC model without the country covariate were not shown for the higher-order model because they are strictly equivalent to those of the three-factor model. $\tau_{I4(F)}$, $\tau_{S3(C)}$, $\tau_{S5(C)}$, $\tau_{S1(S)}$, $\tau_{S2(S)}$, $\tau_{I5(S)}$, $\tau_{C2(S)}$, $\tau_{S1(B)}$, $\tau_{I3(B)}$ free = equality constraints of indicator intercepts were relaxed for item I4 for France, items S3 and S5 for Canada, items S1, S2, I5 and C2 for Switzerland, items S1 and I3 for Belgium; $\tau_{Cons(F)}$, $\tau_{Succ(C)}$, $\tau_{Ident(S)}$, $\tau_{Succ(B)}$ free = equality constraints of construct intercepts were relaxed for Consumption for France, Succession for Canada, Identity for Switzerland, and Succession for Belgium; $\beta_{S1(F-B)}$, $\beta_{S3(F-C)}$, $\beta_{S5(F-C)}$, $\beta_{C2(F-S)}$ estimated = direct effects of the Country covariate on indicators were estimated for item S1 (between France and Belgium), items S3 and S5 (between France and Canada), item C2 (between France and Switzerland). MLR χ^2 = Robust Chi-square; *df* = degree of freedom; CFI = Comparative Fit Index; TLI = Tucker-Lewis Index; SRMR = Standardized Root Mean Square Residual; RMSEA = Root Mean Square Error of Approximation; CI = Confidence Interval for RMSEA; $CFI_{decrease} = CFI_{unconstrained\ model} - CFI_{constrained\ model}$ (if $CFI_{decrease}$ > cut-off = .010, then the invariance hypothesis is rejected; otherwise, there is evidence for invariance). *** $p < .001$.

parameter estimates were reasonable and statistically significant (e.g., range of standardized factor loadings = .42 to .72; cross-scale correlations between .63 and .82). Inclusion of the country covariate somewhat reduced the fit of both models to the data (fit indices not shown), without altering other parameter estimates. As recommended by Brown (2006) in such a case, we inspected modification indices and identified salient direct effects of the country covariate on four indicator intercepts (among the nine indicator intercepts for which equality constraint was relaxed during the test of strong factorial invariance) that could be the most responsible for the decrease in model fit (the S1, S3, S5, and C2 indicators). Indeed, after allowing these four direct effects to be freely estimated, all fit indices became good for both models.[11] Of greater significance for our purpose, there is evidence of population heterogeneity; that is, the factor means are not invariant at different levels of the country covariate but instead behave differently across countries. ***Table 4*** shows the unstandardized and standardized estimates as well as the size of the direct effects of the country membership on the latent factors of Succession, Identity, Consumption, and Ageism.

CROSS-CULTURAL DIFFERENCES IN PRESCRIPTIVE AGEISM AND DYNAMIC INDICATORS OF SOCIOSTRUCTURAL PARAMETERS

Inspection of ***Table 4*** shows that all direct effects of the country covariate were significant for the Ageism dimension and a large majority of them was significant for the Succession, Identity, and Consumption subscales; that is, latent means significantly differed from one country to another one.

For Ageism (hypothesis #1), Canadian respondents reported higher means than the French, Belgian, and

COUNTRY COMPARISONS	UNSTANDARDIZED ESTIMATES (STANDARD ERRORS)				*Z*-SCORE				EFFECT SIZE (COHEN'S *d*)			
	S	I	C	AGEISM	S	I	C	AGEISM	S	I	C	AGEISM
France (ref) *vs*												
Canada	.199** (.062)	.150* (.065)	.021 (.048)	.093* (.041)	3.22	2.31	0.45	2.28	.26	.19	.04	.18
Switzerland	–.253** (.060)	–.418** (.060)	–.287** (.047)	–.258** (.043)	–4.22	–6.97	–6.10	–5.97	.33	.51	.51	.51
Belgium	–.240** (.056)	–.053 (.063)	–.217** (.045)	–.167** (.039)	–4.29	–0.84	–4.78	–4.33	.32	.06	.39	.33
Canada (ref) *vs*												
Switzerland	–.452** (.066)	–.568** (.062)	–.308** (.050)	–.352** (.046)	–6.84	–9.22	–6.11	–7.59	.58	.69	.55	.69
Belgium	–.440** (.062)	–.203** (.063)	–.239** (.049)	–.261** (.042)	–7.10	–3.24	–4.90	–6.18	.56	.25	.43	.51
Switz. (ref) *vs* Belgium	.012 (.060)	.365** (.059)	.069 (.040)	.091* (.037)	0.21	6.20	1.75	2.50	.02	.44	.12	.18

Table 4 Results of MIMIC models for the Succession (S), Identity (I) and Consumption (C) subscales and the Ageism construct.

Note. A positive sign of unstandardized estimates means that the parameter estimate of the reference group (ref) is lower than the parameter estimate of the comparison group; a negative sign means the opposite. *z*-score = unstandardized estimate/standard error (if *z*-score > |1.96| or |2.58|, then the direct effect is statistically significant at * $p < .05$ or ** $p < .01$, respectively); Effect size (Cohen's *d*) = standardized estimate from a standardized solution (Std in M*plus*) where only the latent variables are standardized, not the categorical predictor (the Country covariate; $d = .20$, .50, and .80 for small, medium, and large effects, respectively).

Swiss. French respondents had higher means than the Belgian and Swiss. Finally, Belgian respondents scored higher than the Swiss. Remembering that the relative speed of aging (RSA) and the senior dependency ratio increase (SDRI) are 0.21 and 0.31 for Canada, 0.17 and 0.29 for France, 0.12 and 0.20 for Belgium, 0.13 and 0.19 for Switzerland, this shows that ageism differences closely follow the country rankings regarding RSA in combination with SDRI.

For Succession (hypothesis #2a), we observed a similar pattern of mean differences across countries, except for Belgian respondents who did not score higher than the Swiss. Such a pattern precisely follows RSA hierarchy between countries; that is, Canada (0.21) > France (0.17) > Belgium (0.12) ≈ Switzerland (0.13). For Consumption (hypothesis #2b), we found no differences between the Canadians and French as well as no differences between the Belgians and Swiss; we also found that the former group scored consistently higher than the latter. This set of equality-inequality mirrors the SDRI ordering between countries; that is; Canada (0.31) ≈ France (0.29) > Belgium (0.20) ≈ Switzerland (0.19). For Identity, Canadian respondents exhibited higher scores than the three remaining countries; the French and Belgians had equivalent scores, and the Swiss scored at the lowest.

DISCUSSION

Ageism is a psychosocial mechanism that is particularly deleterious to older people. It is associated with ageist prescriptions, which convey intergenerational tensions based on access to and sharing of limited resources (North & Fiske, 2012). Given its negative impacts on the older population combined with its prevalence worldwide, developing international research about prescriptive ageism is becoming crucial to better understand and, ultimately, reduce it. So far, the vast majority of studies on prescriptive ageism have been conducted in the US, with English-speaking populations (e.g., North & Fiske, 2013b, 2015). This is true for both research about the determinants of prescriptive ageism (e.g., North & Fiske, 2015) and research devoted to the measurement of prescriptive ageism, from which only one scale to date has been developed, namely the Succession, Identity, Consumption scale (SIC; North & Fiske, 2013a). Since other, non-English-speaking countries are no exception regarding ageism, there is a strong need for similar research elsewhere and in other languages.

Accordingly, our first goal was to investigate the relevance of the three-factor SIC model of prescriptive ageism outside the USA, focusing on four distinct French-speaking countries (Canada, France, Belgium, and Switzerland). In each country tested, findings from CFAs provided evidence of factorial validity (good fit indices, magnitude of primary loadings) as well as valuable clues for construct validity (medium to high-sized correlations between subscales, presence of Ageism as a broader dimension) of the three-factor solution of the French SIC scale. Furthermore, in all countries, the total SIC scale exhibited good reliability, as did each subscale. This set of results provides support for the relevance of distinguishing three prescriptive beliefs as specific components of a global construct of ageism, referring to tensions over succession resources, consumption resources, and identity, respectively. As a whole, the present findings match well those obtained amongst American English-speaking samples (North & Fiske, 2013a). Two differences, however, should be noted.

We excluded five problematic items from the original scale based on content-related (3 items) as well as psychometric considerations (2 items), resulting in a 15-item version of the SIC scale. Specifically, we excluded the item, 'Older adults don't really need to get the best seats on buses and trains' from the Consumption subscale. In our view, this item did not perform as well as other items forming the underlying dimension of Consumption. Specifically, compared to others (e.g., 'Older people are too big a burden on the healthcare system'), it appeared to be less closely associated with the core content of the consumption construct (which focuses on the 'apparent exploitativeness of shared resources' or 'using up everything before others get there'; see North & Fiske, 2013a). Further support for this exclusion came from the findings of the baseline model analyses for each country (i.e., very low loadings on its intended factor, consistent correlations with items from different subscales). Such an approach is echoing some previous research that suggested similar revisions before using the SIC scale (e.g., Hancock & Talley, 2020, who also identified as problematic the second item we discarded, 'Older adults shouldn't be so miserly with their money if younger relatives need it'). Knowing whether this reduced version is as efficient – or even more – as the original one (in terms of predictive validity for example, in the USA and elsewhere) is an intriguing question, though beyond the scope of the present study. Another difference with the original scale concerns the hierarchical factor model of ageism we highlighted.[12] In our view, this model fits well with previous research on ageism, which showed some patterns of high correlations among all subscales or used a global score of ageism in predictive validity analyses for example (e.g., Fraboni et al., 1990; North & Fiske, 2013a). As such, establishing a hierarchical factor model of ageism represents a clear contribution of the present study to the research on ageism and ageism-related prejudices.

Our second goal was to examine the potential determinants of prescriptive ageism in Canada, France, Belgium, and Switzerland, by using two dynamic indicators of sociostructural parameters related to high-speed aging and additional pressure on the productive population, respectively. We focused on these indicators because previous research provided valuable insights about their role in ageism and ageism-related differences across countries (North & Fiske, 2015). Thus, following North and Fiske's (2015) reasoning, we expected variations in Ageism (hypothesis #1), Succession (hypothesis #2a), and Consumption (hypothesis #2b) between the four countries to follow the level of disparity in terms of relative speed of aging (RSA), senior dependency ratio increase (SDRI), or both.

Before testing such hypotheses, we ran an innovative two-step approach that combined Multiple-group CFA (MGCFA) and MIMIC modeling, thereby ensuring valid cross-group comparisons (Sass, 2011). MGCFA demonstrated (partial) strong factorial measurement invariance of the SIC scale across countries, extending previous research that demonstrated SIC invariance for race and gender in the US (Hancock & Talley, 2020). For our French adaptation of the SIC scale, this is a good point because it demonstrates its overall ability to measure the same underlying construct (Ageism) and sub-constructs (Succession, Identity and Consumption) in all countries under study. In turn, whether or not there would be a country effect on ageism could be confidently tested. Results from MIMIC modelling revealed such an effect for the global score of ageism and for most of the subscores as well.

Specifically, Canadians exhibited higher levels of ageism than the three other countries, French scored higher than Belgians and Swiss, and Swiss scored the lowest. Overall, this distinctive pattern of ageism endorsement is very close to the country ranking regarding RSA combined with SDRI, thus supporting mostly our first hypothesis.[13] Regarding Succession, Canadians scored consistently higher than French, Belgians, and Swiss, French exhibited higher scores than Belgians and Swiss, but Belgians did not score higher than Swiss. Regarding Consumption, Canadians and French did not differ, Belgians and Swiss did not either, but the former group scored consistently higher than the latter one. The first pattern of differences is strictly echoing RSA hierarchy between countries, while the second one is exactly following SDRI ordering between countries, thus providing strong support for our hypotheses (2a and 2b).

Thus, in accordance with North & Fiske's (2015) line of reasoning, it seems that such dynamic indicators adequately capture the strain from rises in population aging. In this respect, the first rank of Canada both for global ageism and for each prescriptive domain in the present study can be viewed as the most vivid evidence: for almost 40 years, Canada's population has been growing old faster, and its senior dependency ratio is increasing at a stronger pace than that of European countries (France-Belgium-Switzerland). As a whole, our results suggest that countries facing the greatest challenges of aging population for one generation are more ageist than countries having experienced more graduate aging and less abrupt accommodation to the aged (North & Fiske, 2015). Accordingly, it could be argued that ageism is partly due to the level of societal strain commonly associated with population aging and the related view that older adults represent a burden to modern society (Nelson, 2005; North & Fiske, 2015). This is in line with the idea that population aging is likely to produce some depreciation of older adults and societal conflict based on intergenerational tensions over resources allocation (North & Fiske, 2012; The International Longevity Center, 2006).

LIMITATIONS AND FUTURE DIRECTIONS

Despite its strength, variety of samples, and consistent results, the present study suffers from some limitations that must be acknowledged. First, only French-speaking and Western citizen ratings were used in this study.

To allow for a widespread use of the SIC scale, future research should aim to gather data in countries with a diversity of other cultures and languages (e.g., Asian and African cultures). Also, largely due to recruitment among psychology students, a large proportion of the samples was female. In light of evidence that men tend to exhibit higher levels of ageism than women (e.g., North & Fiske, 2013a), a more balanced sampling should be used to allow for the generalizability of our findings.

Second, DIF analyses using Wood's (2009) approach provided some evidence that a few items were associated with differential item functioning across countries. This non-equivalence was less than 10% of the response scale for any item in any two-group comparisons. Although this country-based item bias was minor (as well as beyond the scope of investigation of the present study), understanding why some items behave differently as indicators of Succession, Identity, or Consumption in France and other countries represents both an interesting and relevant focus for future research. Additionally, the convergent validity of the SIC scale with other measures of ageism (e.g., the FSA-14; Boudjemadi & Gana, 2009) was not tested. As a result, establishing the empirical similarity of our scale with other scales of ageism was not possible. This point serves as the next logical step in the psychometric evaluation of our scale (as suggested by Boateng et al., 2018, for example), for the purpose of determining whether or not the SIC scale is strongly related to tried-and-tested measures of ageism.

Third, we did not make any assumption about the factors that could explain variation in Identity ratings. As a reminder, Identity refers to intergenerational tensions over more symbolic resources than Succession or Consumption do; it is about 'unwanted intrusions into young ingroup territory', noting that 'those who are considered *old* often depends on context in this case' (North & Fiske, 2013a: 707). Thus, as a symbolic and inherently cultural dimension, Identity is less likely to generate cross-cultural differences in ageist prescriptions determined by sociostructural parameters. Rather, such differences might be rooted in specific threats activated by older people. This is in line with the Intergroup Threat Theory (Stephan et al., 2016), which posits that symbolic (and realistic) threats determined by in-group identification or inter-group conflict impact intergroup relations. Future studies that apply this kind of theoretical framework within the scope of ageism and its determinants are therefore warranted.

Fourth, despite relying on two dynamic indicators of sociostructural parameters to enlighten ageism in various countries, this study by itself cannot provide a full exploration of how economic and demographic conditions affect age-based stereotypes around the world. Future research will need to address this issue more directly, for example by examining very different countries from the socioeconomic standpoint. A longitudinal study would also be informative, making it possible to focus on the evolution of ageist prescriptions over time, thus allowing to assess the link between some (targeted) dynamic indicators of sociostructural parameters and ageism scores over time, beyond a single comparison of patterns.

CONCLUSION

This study provides support for the cross-cultural validity of the SIC scale. Results also revealed across-country differences in prescriptive ageism that largely reflect demographic as well as socioeconomic disparities between the four countries considered. As such, the SIC scale has several attractive features that make it a suitable instrument for assessing prescriptive ageism and its roots. Considering its substantial social consequences, it appears important to counter ageism. Practical suggestions can notably be drawn from the PEACE model (Lytle & Levy, 2017), a theoretical framework focusing on two interrelated factors that have the potential to reduce ageism: education about aging and positive contact experiences with older adults (see also Bousfield & Hutchison, 2010; Marshall, 2015). In this sense, promoting knowledge about aging and positive intergenerational contact could help reducing ageism, and as such are beneficial to anyone (scholar, decision-maker, etc.) interested in countering ageism and improving older adults' well-being and quality of life.

NOTES

1 Gender distribution for the French, Canadian, and Swiss samples is homogeneous (Pearson Chi-square (2) = 5.15, p = .08). For the Belgian sample, yet, the ratio of women to men (about 5:1) is higher than for the three other samples (about 3:1). Given men score consistently higher than women on ageism (e.g., North & Fiske, 2013a; see also Kornadt et al., 2013, for more specific gender effects), it could be argued that such an unbalanced distribution artificially decreased the mean score of ageism in our Belgian sample. To test this eventuality, we compared the observed mean of the total sample of Belgians on ageism (N = 434; m = 2.28; SD = 0.60) with two other observed means from two randomly truncated samples excluding 100 women (the first 100 women and the last 100 women, respectively), resulting in the same ratio as other countries. Both means remained almost unaltered (N = 334; m = 2.26, SD = 0.60, and N = 334; m = 2.31, SD = 0.61), and did not differ from the observed mean of the total sample (p = .65 and p = .50, respectively). This provides support for considering our Belgian sample just like others.

2 Authors explained that 'many older adults are active on Facebook and younger participants may not [see] use of Facebook as violating prescriptive norms for older adults' (2020: 9). Thus, they recommended the removal of this item. However, this is not necessarily true for all social networks, and we hence decided to reformulate, not delete, the item. In the same vein, at that time, we decided against moving the item 'Older people shouldn't be so miserly with their money if younger relatives need it' to the Succession dimension for preliminary analyses because we had no preconceived idea about how the item would behave in non-US samples.

3 Rho is a Structural Equation Modeling estimate of reliability which is defined as the ratio of a scale's estimated true score variance to the scale's estimated total score variance (Yang & Green, 2011). This measure is identical to McDonald's Omega

(ω) coefficient (McDonald, 1999) which represents a good alternative to the more established alpha coefficient (Stone et al., 2013; see Revelle & Condon, 2019; Revelle & Zinbarg, 2009).

4 There are two main reasons why a higher-order factor model was tested in the present study: first and theoretically speaking, Succession, Identity and Consumption were designed as the three key domains of one single broader construct, viz. the prescriptive ageism; second and empirically speaking, North and Fiske (2013a) observed moderate to high correlations among the SIC subscales (from .46 to .66 depending on studies and scales).

5 In the present study, we decided not to assess equal item residual variances (*strict factorial* invariance) for two imbricated reasons: first, in the recommended sequence of MGCFA invariance evaluation, testing this equality is referred to as 'optional' (Brown, 2006: 270); second, when there are many groups involved in the comparison, it may be unreasonable to assume invariance of residuals across groups (Little, 1997).

6 In the present study, the Country covariate has four levels (France, Canada, Switzerland, and Belgium). Consequently, three dummy codes have been created in which France was treated as the reference group (0 = France).

7 Guided by two high modification indices (found in each sample), the residuals of items 'Doctors spend too much time treating sickly older people' and 'Older people are too big a burden on the healthcare system' were allowed to covary, as well as the residuals of items 'Older people probably shouldn't use social networks' and 'Older people shouldn't even try to act cool'. In both cases, semantic closeness between the two items of the pair justifies letting their residuals covary.

8 Considering the correlations between factors, the reader might wonder why we did not include ESEM (Exploratory Structural Equation Modeling; Asparouhov & Muthén, 2009) and/or BI-ESEM (Bifactor-ESEM; e.g., Morin et al., 2016, 2020) models to investigate the cross-loadings. Comparing the findings from both analysis (ESEM and BI-ESEM) in our reference (French) sample with that of the CFA (no residual correlations were permitted for any model), we found the fit for the CFA solution did not differ substantially from the corresponding ESEM model ($MLR\chi^2_{CFA} < MLR\chi^2_{ESEM}$ and $RMSEA_{CFA} < RMSEA_{ESEM}$, but CFI_{CFA} and $TLI_{CFA} < CFI_{ESEM}$ and TLI_{ESEM}), while the BI-ESEM model showed the best fit indices. However, as compared to the CFA solution, both the ESEM and BI-ESEM models displayed much poorer parameter estimates as illustrated by a majority of item-indicators with lower primary loadings on their expected factors (e.g., on the Succession factor, the mean of the target factor loadings was .54 for CFA versus .44 for ESEM versus .33 for BI-ESEM). Moreover, (1) nontarget cross-loadings were virtually non-existent (e.g., zero or one depending on factors for ESEM); (2) factor correlations for ESEM (S-C = .60, S-I = .59, I-C = .60; *m* = .60) were not substantially lower than CFA (S-C = .74, S-I = .58, I-C = .59; *m* = .64); (3) only five out of 15 item-indicators loaded significantly ($p < .05$) on the general factor defined in BI-ESEM; (4) the specific factor related to the Consumption subscale was very underdetermined (only one significant loading at $p < .05$, but still low = .22) in the BI-ESEM model. These findings provide support for our CFA strategy at the expense of the ESEM or BI-ESEM solutions.

9 An alternative approach for representing a general construct comprised of several related domain-specific factors is the (confirmatory form of) bifactor model (Chen et al., 2006). However, as it often does, the bifactor model for the SIC scale failed to converge for the most part (see Gana & Broc, 2019, for details).

10 'As expected', since our first-order model has three factors, then the higher-order solution is just-identified, producing the same goodness of fit as the first-order model. Even so, we were interested in testing such a solution in order to examine the magnitude (and statistical significance) of the higher-order factor loadings as well as the size of the reliability of the higher-order factor (Brown, 2006).

11 More specifically, there was a significant direct effect of the covariate on each indicator intercept (holding latent factors constant). This could be evidence of Differential Item Functioning (DIF); that is, the SIC seems to work differently in distinct countries for these items. This encouraged one reviewer to suggest that we use Wood's (2009) approach to test systematically for DIF. In short, this is a quick strategy for empirically identifying DIF-free items and no DIF-free items in step 1, then testing the direct effect of any covariate on each no DIF-free or studied item in step 2 (using DIF-free items as anchors). This stepwise analysis relies on (a) likelihood ratio tests (used in both steps) that compare nested two-group models to evaluate whether the items, examined one at a time, function differently across groups (controlling for group differences in the latent factors) and (b) the Benjamini-Hochberg (BH; Benjamini & Hochberg, 1995) procedure that enables to control the false positive rate in multiple comparisons (used between step 1 and step 2). Broadly, findings from DIF testing showed that no DIF-free items (after adjustment for false discovery and after step 2) were primarily those associated with highest modification indices in MIMIC modeling, and that DIF effects were less than 10% of the response scale. Specifically, we found significant and small to medium direct effects of country on the S1 ($z = -5.40$, $p < .001$, $d = .41$), S3 ($z = -6.25$, $p < .001$, $d = .53$), S5, ($z = -5.46$, $p < .001$, $d = .49$) and C2 ($z = 4.77$, $p < .001$, $d = .36$) indicators. Controlling for country differences in latent factors, the Belgian sample has lower scores on S1 than the French (by .41 units, or nearly half a point on the 1–6 scale), the Canadians have lower scores on S3 and S5 than the French (by .53 and .49 units, respectively, or about half a point on the 1–6 scale), and the Swiss have higher scores on C2 than the French (by .36 units, or about a third of one point on the 1–6 scale). DIF testing involving other two-group comparisons (Canada *vs* Switzerland, Canada *vs* Belgium, and Switzerland *vs* Belgium) as well as DIF testing resulting in less than small to medium effects are available by request from the co-first author (BC).

12 In their study, North and Fiske also evaluated such a model but decided to report only the first-order model for 'simplicity's sake' (2013a: 708).

13 'Mostly' because we did not expect any difference for ageism between Belgium and Switzerland in view of their strong similarity in terms of both dynamic indicators. In practice, we found such a difference probably because ageism not only consisted of Succession and Consumption (for which there was indeed no difference between Belgium and Switzerland), but also Identity (a domain which no sociostructural parameter was connected with, then no hypothesis was made, but for which a Belgium-Switzerland difference has been found). Further research about Identity-based prescriptions is warranted to clarify this difference (as well as other results such as the lack of difference between France and Belgium on the Identity subscale).

ADDITONAL FILE

The additonal file for this article can be found as follows:

- **Supplementary file.** SIC scale data. DOI: *https://doi.org/10.5334/irsp.544.s1*

COMPETING INTERESTS

The authors have no competing interests to declare.

AUTHOR CONTRIBUTIONS

The first two authors named are the lead and corresponding authors. We describe contributions to the paper using the CRediT taxonomy. Conceptualization: V.B.; Methodology: V.B., and B.C.; Software: B.C.; Validation: V.B. and B.C.; Formal analysis: B.C.; Investigation: V.B., S.A., C.I.P., M.L., F.L., and W.S.; Resources: V.B.; Data Curation: B.C.; Writing - Original Draft: V.B. and B.C.; Writing - Review & Editing: V.B., B.C., S.A., K.G., M.L., and F.L.; Visualization: V.B., B.C. and K.G; Supervision: V.B.; Project administration: V.B.

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