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VALIDATION OF POPULIST ATTITUDES SCALES IN SLOVAKIA

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ABSTRACT

Objectives. Research on populism has progressed over the last decade and several scales have been proposed to measure populist attitudes. None of these has been validated in the context of Slovakia, where populists are a long-term part of both coalition and opposition. This study aimed to verify the psychometric properties of four populist attitudes scales that are frequently used and verified in international research, on a Slovak sample.
Participants and setting. Data of 832 respondents collected using an online panel in November 2021 were analyzed using the R software. The research sample was representative according to the distribution of gender, age, education, and regions in Slovakia.
Statistical analysis. The original scales were tested using an exploratory dataset (N = 416). The modified scales were verified using a confirmatory dataset (N = 416).
*Results.*The results showed that the original scales did not fit the data. However, after several modifications, the two scales were validated on the Slovak sample. The scales were invariant across the gender, age, and educational groups.
Limits. A possible shortcoming of the validated populist attitudes scales is the instability in predicting electoral behavior, which is discussed in connection with the results of other studies, and the Slovak political and cultural context.

key words:
populist attitudes scale,
validation,
Slovakia

kľúčové slová:
škály populistických postojov,
validácia,
Slovensko

INTRODUCTION

The rise of populist political parties since the 1990s has stimulated research into this phenomenon. The definition of populism with only the necessary and sufficient conditions, is: "Populism is a thin centered ideology that considers society divided into two homogeneous and antagonistic groups – the pure people and the corrupt elite, and which argues that politics should be an expression of general will (volonté générale) of the people" (Mudde, 2007, p. 23). People are perceived as sovereign, homogeneous, clean, and virtuous, and the elite (most often politicians or government establishments) threaten the purity and unity of sovereign people. The distinction between "good" and "evil" is an essential distinguishing and defining feature of populism. The tension between the people and the elite is Manichaean, which means that it has moral quality (Hawkins, 2009; Mudde, 2004). Populism can have a harmful side, i.e., the well-known demagogic populism, but also a more friendly one in the sense of sympathetically listening to the masses and fulfilling their legitimate needs.

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Empirical studies aimed at identifying and measuring populist attitudes have expanded only in the last decade (for a comprehensive overview, see Piterová et al., 2021). The genesis of the development brought several scales, which led to an interest in their comparison and identification of those with the best psychometric properties. Based on an extensive comparative study in eight countries with at least 250 respondents from each country, it was shown that the following three best met the required criteria (Castanho Silva et al., 2020): 1. Akkerman et al. (2014), 2. Schulz et al. (2018), and 3. Castanho Silva et al. (2018), each of them representing a different approach to populism.

Based on the American research of Hawkins et al. (2012) and Akkerman et al. (2014) captured three main features of populism through their scale: popular sovereignty, anti-elitism (the distinction between people and the elite), and the Manichaean distinction between good (people) and evil (elites). The approach presented by these authors (confirmed through empirical research on Dutch respondents) reflects a one-dimensional understanding of populism. Their populist attitudes scale consists of six items (three items represent popular sovereignty, two represent anti-elitism, and one represents the Manichaean outlook). The authors also confirmed a significant positive correlation between the range of populist attitudes surveyed and the intention to elect (Dutch) populist political parties, regardless of their right- or left-wing orientation (Akkerman et al., 2014). The same results were supported in an international context (Castanho Silva et al., 2020).

A few years later, Van Hauwaert et al. (2020) added two more items to the original items of Akkerman et al. (2014) to capture how individuals perceive the continuing negative impact of elites' behavior and interests in their lives, and perceptions of (political) elites as negative and self-serving authorities. According to the authors, these items exceeded those of the original anti-elitist scale by examining an individual's perception of the negative antagonistic relationship between people and elites. Based on a large international sample (from 9 European countries, including Poland as the closest country to Slovakia), the authors proved that this eight-items extended scale is not only the most informative and has the most coverage (compared to the original six-items scale of Akkerman et al., 2014 and/or reduced three-items version of Van Hauwaert et al., 2020), but also has good dicrimination, predictive power, and validity.

A milestone in the development of measurement tools was overcoming the operationalization of populism as a one-dimensional construct, which dominated the research. Schulz et al. (2018) operationalized populism in three subdimensions: 1. anti-elitism when elites are perceived as corrupt and deceptive by people; 2. belief in popular sovereignty, which leaves power to the people; and 3. understanding people as homogeneous and virtuous (Wirth et al., 2016, in Schulz et al., 2018). According to Schulz et al. (2018), the idea that people are a homogeneous group that is wise and honest is necessary for an adequate definition of populism. The Manichaean perspective of populism, through the depiction of an entity of people against an entity of politicians or government, has been captured across all dimensions (Schulz et al., 2018). According to the authors, populist attitudes are formed (by the stated dimensions) as a second-order latent construct. The populist scale consists of 12 items (four items per each dimension) and has been empirically verified on Swiss respondents (Schulz et al., 2018). It also adequately predicts the choice of populist political parties in an international context, regardless of their right- or left-wing orientation (Castanho Silva et al., 2020).

Based on the same definition of populism as the authors before (Mudde, 2004), Castanho Silva et al. (2018) emphasized three slightly different elements. The basic components of populist attitudes are: people centrism (which combines the two dimensions of Schulz et al. (2018) – popular sovereignty and homogeneity of people), as well as the anti-elitism and Manichaean outlook on politics. Through the last dimension, the authors also reemphasized the importance of this aspect of populist attitudes ("politics as a moral struggle" was part of the anti-establishment items, e.g., the populist attitudes scale of Akkerman et al., 2014; Hawkins et al., 2012; Schulz et al., 2018). The authors intended to measure the Manichean element in its "pure" form, meaning the perception of politics as a moral struggle without connecting with the people or the elite. Castanho Silva et al. (2018) also suggested that the populist attitudes scale should include various components (dimensions) separately, which has rarely been considered in the past, with the exception of Schulz et al. (2018), Stanley (2011), or Oliver and Rahn (2016). Assuming that populism has different components, it is natural that not all items would behave as if they were measuring one construct (Castanho Silva et al., 2018). This scale was (as was Van Hauwaert et al., 2020) developed by testing the items on international samples, not only from European countries, but also on respondents from the USA, Mexico and South America. However, European countries did not include any countries from Central or Eastern Europe, only its western and southern parts.

The validity of the populist atttitudes scales was verified in relation to several variables. At the beginning of the era of populism measurement development, pluralism and elitism were negatively related to populism and used in the validation process (Akkerman et al., 2014; Hawkins et al., 2012; Schulz et al., 2018). While pluralism emphasizes the diversity of opinions (Akkerman et al., 2014; Hawkins et al., 2012; Schulz et al., 2018), populism demands moral clarity and fulfilling the people's will. Holding pluralistic attitudes is perceived as suspicious or dangerous. Elitism (Akkerman et al., 2014; Hawkins et al., 2012; Schulz et al., 2018) represents a mirror image of populism. The elites are considered pure and virtuous and the people corrupt. In practice, however, their relationship is sometimes ambiguous because the elitism and charismatic leaders of populist movements express a similar hierarchical concept of leadership (Akkerman et al., 2014).

Additionally, there are two widely held correlates of populist attitudes – political trust (Castanho Silva et al., 2020; Erisen et al., 2021; Rooduijn et al., 2016) and conspiratorial mentality (Castanho Silva et al., 2017). Negative correlations with political trust reflect the known and empirically supported relationship between political discontent and populist voting (Rooduijn et al., 2016). Positive correlations with conspiracy thinking (Castanho Silva et al., 2020; Erisen et al., 2021; Oliver & Rahn, 2016) or belief in conspiracies is related to negative perceptions of elites and helps people fulfill their need for control and certainty. According to the GLOBSEC report (Hajdu & Klingová, 2020), the Slovak Republic is among the European countries with the highest incidence of people who have been shown to be susceptible to disinformation and conspiracies.

There is another construct associated with the aforementioned variables: mistrust of experts. This indicates skepticism about science and expert opinion. According to Oliver and Rahn (2016), anti-expertise was specifically related to anti-elitism sentiments. Mistrust of experts is operationalised through several questions where the common sense knowledge of ordinary folk is played up as superior to experts' ideas, which is one way populists praise common people (Oliver & Rahn, 2016).

There also exists a widespread belief that those who vote for populists are those with a lower socioeconomic status and/or lower education, who (therefore) feel threatened by others (Kriesi et al., 2006). Such people can also identify with the category of "ordinary citizens" or "the common man", and therefore feel attracted to the message that they are victims of the behavior of the evil elite (Kriesi et al., 2006). However, feelings of relative deprivation (Elchardus & Spruyt, 2016) emphasize that it is not the low economic position of the individual itself, but rather the subjective interpretation of a certain vulnerability that has a direct effect on populism. The results of conducted research support the existence of a negative relationship between populist attitudes and education (Balta et al., 2021; Boscán et al., 2018; Elchardus & Spruyt, 2016). Moreover, Elchardus and Spruyt (2016) found that men were more prone to populism, whereas age had no effect on populist attitudes (p. 124). Sociodemographic characteristics such as gender, age, and education were also strong differentiation factors in the 2020 Slovak general elections (see Gyárfašová & Slosiarik, 2020). Evidently, populist preferences have different demographic correlates depending on the country and its thick ideological attachment (Rooduijn, 2018).

The validity of populist attitudes scales is often verified through the prediction of the intention to vote for populist parties or their identification with them (Akkerman et al., 2014; Castanho Silva et al., 2020; Stanley, 2011; Van Hauwaert et al., 2020). In this context, however, it should be noted that despite their good psychometric properties, these scales do not predict election intentions without problems. Their effectiveness has been proven to differ in international comparisons and often has low predictive power (Castanho Silva et al., 2020), which raises the question of the contextual specificity of populism. However, although some authors suggest potential differences due to different cultural settings (Van Hauwaert et al., 2018), their specifications have not been mentioned (Castanho Silva et al., 2020; Van Hauwaert et al., 2018). Research also suggests a problem when populist parties are part of the government (Stanley, 2011) because it is difficult for a respondent tending to populist parties to accept items that government elites are against the interests of the people, as they are currently the government elites themselves. Even in an extensive international comparison, the scales with the best psychometric properties failed to predict support for populists in countries with populists in the government (Castanho Silva et al., 2020).

This aspect may be relevant to the Slovak context, as after the 2020 general election, two populist parties, one purely populist and the other with populist rhetoric, switched places between the government and the opposition. Since the fall of communism, populist parties have succeeded in winning the support of the people. While the division between right and left is not transparent in Slovakia, populism is present across the political spectrum, both in the coalition and the opposition. However, the extent and form of populism among the parties has changed over time. As Školkay et al. (2021, p.7) stated in their report, populism has become less nationalistic and more socioeconomic over the 2000–2020 period. When the authors compared support for populist parties over this period of time, purely populist parties maintained the same support, but support for parties with populist rhetoric, authentic and anti-system ideology doubled (p.11). After the 2020 general election, the populist left-leaning party SMER-SD was replaced by a populist right, anti-corruption movement OĽaNO, that during its pre-election campaign highlighted the misconduct of the ruling party, exposing corruption, clientelism and other scandals. Additionally, crises such as the two years of the COVID-19 pandemic, restrictions and inadequate support or compensation from the state have also changed populism in Slovakia. As can be seen in the Eurobarometer, among 27 European countries, Slovak people are the least satisfied with the functioning of democracy in their own country (67% are dissatisfied) (Eurobarometer, 2022). This frustration is picked up by politicians who further polarize society. This "transformed" populism can be characterized by creating new and artificial categories with the aim of finding “new enemies” that can represent “evil groups” (Vašečka, 2022).

Another potential issue is the low voter turnout and volatility of electoral behavior in Slovakia. Voter turnout it the 2020 general election was 65.8%, which was a slight increase compared to previous years (e.g., 2016 – 59.8%, 2012 – 59.1%, 2010 – 58.8%)[1]. Such volatility was seen in the support for new and/or non-systemic parties, while traditional political parties were expelled from parliament (Gyárfášová et al., 2017; Gyárfášová & Slosiarik, 2020). High volatility is also typical in other Central and Eastern European countries (Gyárfašová et al., 2017). Taken together, this cultural setting can foment a phenomenon when people express their worries through political means other than the established parties (Kriesi et al., 2006). In other words, an inclination toward populism (and its correlates) can be the way to cope with specific situations. However, populism is compared to a chameleon: it takes on the color of the environment in which it occurs (Taggart, 2000). This situation raises the question of how to measure populist attitudes (and with what approach) in such an environment, the findings of which can enrich the knowledge regarding such measures of populism.

This study aimed to verify the psychometric properties of three populist attitudes scales (Akkerman et al., 2014, including the items added by Van Hauwaert et al., 2020; Schulz et al. 2018; Castanho Silva et al., 2018) that achieved the best results in an international comparative study conducted by Castanho Silva et al. (2020). The validation process included construct validity by conducting confirmatory factor analysis (CFA), and convergent and discriminant validity by correlation analysis of populist attitudes scales and selected variables. Based on the results of the aforementioned studies, we expected positive correlations between similar subscales of distinct populist attitudes scales as well as between populist attitudes and conspiracy thinking, relative deprivation, mistrust of experts, and negative correlations with elitist attitudes, pluralist attitudes, and political trust. The predictive validity was tested by the ability of populist attitudes to predict past and planned voting for both left- and right-wing populist parties. Finally, measurement invariance and differences between age, gender, and educational groups in the level of populist attitudes were calculated. We hypothesized that men, older and less educated people would hold stronger populist attitudes (Balta et al., 2021; Boscán et al., 2018; Elchardus & Spruyt, 2016).

METHODS

Participants and procedure

The consent of the ethics committee of the Centre with the research project was obtained. Data were collected between 5th and 15th November 2021 using an online panel of respondents from the 2Muse Agency. The sample of respondents was selected based on quota sampling, but limited resources constrained the choice of the sample size. Quotas for gender, age, education, and region of residence were calculated based on data from the Statistical Office of the Slovak Republic (Sčítanie.sk, 2021). The research agency’s panelists who met the conditions of sampling were approached to participate in the research, based on their prior consent to participate in further research. Data collection continued until the established quotas were met. The research agency provides for the respondents a small incentive for completing the questionnaire. Informed

[1] https://volby.sme.sk/parlamentne-volby/2020/vysledky
https://volby.statistics.sk/nrsr/nrsr2016/sk/
https://volby.sme.sk/parlamentne-volby/2012/vysledky
https://volby.sme.sk/parlamentne-volby/2010/vysledky

consent was obtained from all the participants. The participants were assured that their data would remain anonymous and confidential. From the initial sample (N = 902), records from 23 participants who answered all questions in less than 2 minutes were removed. Four items (Q3_2, Q3_5, Q3_8, Q3_32) had negative-wording, therefore, the scale was reversed. Data were checked for careless responses by calculating the Mahalanobis distance and longstrings using the "careless" package (Yentes & Wilhelm, 2021) in R software (R Core Team, 2021; RStudio team, 2019). 34 participants with a longstring and Mahalanobis distance > 3 SD were excluded. Thirteen people were excluded as careless, based on additional inspection of longstrings and Mahalanobis distance values, and response time. This yielded a total of 832 respondents who were representative of the Slovak population in terms of gender (48.8 % men), with small deviations in age composition (aged 18 to 75 years, with an average value of 43.79 (SD = 15.01), level of education (primary 10.9 %, secondary 70.4 %, tertiary 18.6 %), and in terms of affiliation to eight regions. The composition of the sample is presented in Table 1.

Table 1 Research sample composition and population statistical data

		N	%	Statistical Office of the SR (%)[1]
GENDER	Men	406	48.8	48.91
	Women	426	51.2	51.09
AGE	18-25	115	13.8	11.8
	26-35	174	20.9	19.48
	36-45	162	19.5	22.49
	46-55	167	20.1	18.39
	56-65	139	16.7	16.8
	66-75	75	9	11.63
EDUCATION	Primary	91	10.9	16.97
	Lower Secondary	259	31.1	19.22
	Upper Secondary	327	39.3	29.57
	Tertiary	155	18.6	18.38
AREA	Big city	225	27	53.19
	Suburb of big city	38	4.6	
	Small city	256	30.8	
	Village, farm	313	37.6	46.81
REGION	Bratislava	97	11.7	13.2
	Trnava	84	10.1	10.38
	Trenčín	96	11.5	10.59
	Nitra	106	12.7	12.44
	Žilina	114	13.7	12.69
	Banska Bystrica	102	12.3	11.48
	Prešov	117	14.1	14.85
	Košice	116	13.9	14.35

Note. N total = 832

Measures

The Slovak versions of each scale were provided using a forward translation procedure, three translations by two independent experts and one layperson were compared, discussed, and synthesized into the final version. In this section, each scale is briefly described, but the full Slovak and English versions are provided in *Populist attitudes scales* in the Supplemental materials[2].

Populists attitudes scales

The scale of Castanho Silva et al. (2018) consists of nine items divided into three dimensions: *anti-elitism* (3 items, e.g., "The government is pretty much run by a few big interests looking out for themselves"), *people's centrism* (3 items, e.g., "Politicians should always listen closely to the problems of the people"), and *Manichean outlook* (3 items, e.g., "You can tell if a person is good or bad if you know their politics"). Three items were negatively worded.

The scale of Schulz et al. (2018) consists of twelve items divided into three dimensions: *anti-elitism* (4 items, e.g., "The differences between ordinary people and the ruling elite are much greater than the differences between ordinary people"), *sovereignty of people* (4 items, e.g., "The people should have the final say on the most important political issues by voting on them directly in referendums"), and *homogeneity of people* (4 items, e.g., "Ordinary people all pull together").

The scale of Akkerman et al. (2014) is a unidimensional scale that consists of six items. Items capture three aspects: *anti-elitism* (2 items, e.g., "The political differences between the elite and the people are larger than the differences among the people"), *popular sovereignty* (3 items, e.g., "The politicians in the Slovak parliament need to follow the will of the people"), and *Manichean outlook* (1 item, "What people call ʿcompromise' in politics is really just selling out on one's principles").

The scale of Van Hauwaert et al. (2020) is a unidimensional scale consisting of eight items. Six items were from Akkerman et al.'s (2014) scale, and two items representing anti-elitism were added by the authors. However, in this study, four items were substituted with items from Schulz et al.'s (2018) scale, because they were almost identical (e.g., Schulz et al. (2018) – The differences between ordinary people and the ruling elite are much greater than the differences between ordinary people; Van Hauwaert et al. (2020) – The political differences between the elite and the people are larger than the differences among the people).

The scale of Akkerman et al. (2014), Schulz et al. (2018), and Van Hauwaert et al. (2020) originally used a 5-point Likert scale, while Castanho Silva et al. (2018) used a 7-point Likert scale. Instead of shortening, we expanded the range of scale, which should not affect the results. Participants indicated their agreement on a scale, ranging from 1 (strongly disagree) to 7 (strongly agree). The order of items was randomized within parts with the same answer options.

Measures for validity assessment

Conspiracy thinking was measured using the Conspiracy Mentality Questionnaire of Bruder et al. (2013). It consists of five items, e.g., "I think that many very important things happen in the world, which the public is never informed about". Respondents answered the extent to which they think or do not think the following statements are true, on an 11-point scale, ranging from 0 (0% = certainly not) to 10 (100% = certain).

[2] Supplemental materials are located at https://osf.io/gd2qv/

McDonald's omega (ω) of the scale is .85 (exploratory dataset), .87 (confirmatory dataset).

Political trust was measured using six items from the European Social Survey (2018), "How much do you personally trust the following?" the Slovak parliament, the European parliament, the Slovak national government, Slovak politicians, Slovak political parties, the Slovak legal system" on an 11-point scale, ranging from 0 (complete distrust) to 10 (complete trust). McDonald's omega (ω) of the scale is .92 (exploratory dataset), .93 (confirmatory dataset).

Pluralist attitudes were measured using the scale provided by Schulz et al. (2018), which consisted of two items adopted from previous studies (Akkerman et al., 2014; Hawkins & Rovira Kaltwasser, 2014) and two items added by Schulz et al., 2018), e.g., "Democracy is about achieving compromise among differing viewpoints". Respondents answered the extent to which they agreed with individual statements on a 7-point scale, ranging from 1 (strongly disagree) to 7 (strongly agree). McDonald's omega (ω) of the scale is .74 (exploratory dataset), .73 (confirmatory dataset).

Elitist attitudes were measured using the scale provided by Schulz et al. (2018), which consisted of two items (considering politicians and successful businessmen as elite) from Hawkins et al. (2012), and one item (considering independent experts as elite) added by Akkerman et al. (2014), e.g., "Our country would be governed better if important decisions were left up to successful business people" Respondents answered the extent to which they agreed with individual statements on a 7-point scale, ranging from 1 (strongly disagree) to 7 (strongly agree). McDonald's omega (ω) of the scale is .47 (exploratory dataset), .44 (confirmatory dataset). Items are used in correlation analysis separately, not as a scale.

Mistrust of experts was assessed using the scale developed by Oliver and Rahn (2018), which consists of three items, e.g., "I'd rather put my trust in the wisdom of ordinary people than the opinions of experts and intellectuals". Respondents answered the extent to which they agreed with individual statements on a 7-point scale, ranging from 1 (strongly disagree) to 7 (strongly agree). One item was negatively worded. McDonald's omega (ω) of the scale is .7 (exploratory dataset), .63 (confirmatory dataset).

Relative deprivation was measured using seven items from the scale of Elchardus and Spruyt (2016), e.g., "It is always other people who can profit from all kinds of advantages offered in this society". Respondents answered the extent to which they agreed with individual statements on a 5-point scale, ranging from 1 (strongly disagree) to 5 (strongly agree). McDonald's omega (ω) of the scale is .81 (exploratory dataset), .84 (confirmatory dataset).

Electoral behavior was measured by two items: "Who did you vote for in the last general elections in 2020?" and "Who would you vote for in the next general elections?" A list of Slovak political parties was provided to the respondents.

Data analysis

The dataset was randomly split into exploratory (N = 416) and confirmatory (N = 416) datasets. The research sample composition for each dataset is available in Table S1 in the Supplemental materials.

All statistical procedures were conducted using the R software (R Core Team, 2021; RStudio team, 2019). Raw dataset, R code, and codebook are available upon request from the corresponding author. Descriptive statistics and correlation matrices were examined. None of the analyzed observations were treated as outliers, since all the observed data points could theoretically represent the distribution of the respective constructs in the population.

To determine construct validity, we applied confirmatory factor analysis (CFA) to the exploratory dataset separately for each scale. The WLSMV estimator was used in the "lavaan" package (Rosseel, 2012). The χ^2 test of model fit, comparative fit index (CFI > .95 or acceptable > .90), Tucker–Lewis index (TLI > .95 or acceptable > .90), root mean square error of approximation (RMSEA < .06 or acceptable < .08), and standardized root mean squared residual (SRMR < .06 or acceptable < .08) were used to evaluate the parameters. The cut-off criteria for the fit indices were set in line with those of Hu and Bentler (2009), Bentler (1990), and Fabrigar et al. (1999). Modification indices were examined to improve the model fit. The final models were tested using a confirmatory dataset. The differences in the results between the explanatory and confirmatory datasets are an indicator of the instability of the results and indicate the need to verify the findings in future research. Further validation was conducted on the confirmatory dataset, but the results for the exploratory dataset are available in the Supplemental materials.

The reliability of scales was determined using the McDonald's omega (ω). Descriptive statistics of the original and modified populist scales/subscales and selected correlating variables can be found in Table S2 in the Supplemental materials.

In further analyses, we used continuous quantifiers of concept structures that differed according to the degree to which they held populist attitudes. We tested one unidimensional (Van Hauwaert et al., 2020) and one three-dimensional modified scale (Schulz et al., 2018); therefore, we used two aggregation methods. A simple mean score was calculated for a unidimensional scale. For the multidimensional scale, we used a simple mean score, as well as the Goertz approach (2006, 2020, as cited in Wuttke, 2020), which understands populism as a non-compensatory concept (i.e., a low score in one dimension cannot be substituted by a high score in another). Mean values across dimensions were calculated. To score high on populist attitudes, respondents were required to score high on all the dimensions. Therefore, the minimum value across the three dimensions of the multidimensional scale represents the intersection of all dimensions, that is, the level of populist attitudes. This method is epistemologically and methodologically preferred to operationalize populist attitudes, however, at first sight it disregards information on all but the lowest attribute (Wuttke, 2020, p. 361).

The convergent and discriminant validity of populist attitudes measured by the modified scales of the Schulz et al. (2018) and Van Hauwaert et al. (2020) were evaluated using nonparametric Spearman's correlation coefficients on the confirmatory dataset. The correlation matrix was visualized using the "corrplot" package (Wei & Simko, 2021).

The predictive validity of the populist scales was estimated by the effect of populist attitudes (average and minimum score on the modified scale of Schulz et al. (2018); average score on the modified scale of Van Hauwaert et al. (2020)) on voting behavior in general elections in 2020, as well as voting intention in the case of future elections. The coding of Slovak political parties as populist or non-populist was carried out along the schema provided by PopuList (Rooduijn et al., 2019), TIMBRO Authoritarian Index (TAPI) (2019), and the Populism and Political Parties Expert Survey (POPPA) (Meijers et al., 2020). There are inconsistencies for some political parties. For example, TAPI did not include the clearly populist movement OĽaNO and also SMER-SD party but did include the far-right-wing party ĽSNS. According to PopuList, OĽaNO and SMER-SD are populist, but ĽSNS is considered right-wing, Eurosceptic, but not populist. In contrast, ĽSNS received the highest ranking on populism in the POPPA survey (9.27 out of 10), while SMER-SD had a relatively low score (under 4 out of 10). In this paper, we have followed the report of Školkay et al. (2021) and we have coded political parties as populist if they are either purely populist, with populist ideology (OĽaNO, We are family, Homeland), with populist rhetoric and authentic ideology (SMER-SD, ĽSNS), and/or are considered as borderline cases (SNS). For more details about coding populist parties, see S9. *Coding of populist parties in Slovakia* in the Supplemental materials. A series of binomial logistic regression models were tested on a confirmatory dataset using the glm function from the "stats" package (R Core Team, 2021).

Finally, configural, metric, scalar, and strict invariance across gender, age and educational groups were tested for modified scales of Schulz et al. (2018) and Van Hauwaert et al. (2020) using exploratory and confirmatory dataset (for the results, see *Measurement invariance* in the Supplemental materials). For measurement invariance, package "lavaan" (Rosseel, 2012) and "semTools" (Jorgensen et al., 2021) were used. Differences between groups (gender, age, and education) were determined by t-test and ANOVA with post-hoc pairwise t-tests and Tukey's HSD correction using the "rstatix" package (Kassambara, 2021). Confidence intervals for each pairwise comparison were obtained. Cohen's d was estimated as an indicator of the effect sizes of the differences between groups.

RESULTS

Construct validity

The original scale of Castanho Silva et al. (2018) consists of nine items divided into three dimensions: anti-elitism (Q3_1 – Q3_3), centrism (Q3_4 – Q3_6), and Manichean outlook (Q3_7 – Q3_9). For wording of the items named Q3_1 – Q3_9 in our dataset, see the *Populist Attitudes Scales* in the Supplemental materials. The CFA of the original version showed that three items had loading less than .5 (Q3_5, Q3_7, Q3_9), the latter had loading only .26. Q3_8 and Q3_9 loaded on other factors than indicated. SRMR was < .08, χ^2 was significant, but it is sensitive when the sample size is over 100 respondents. CFI and TLI were > 0.9, and RMSEA < .08, but SRMR > .08. After excluding item Q3_9, the fit index values improved slightly, while Q3_2, Q3_5, and Q3_8 co-variated, and Q3_7 still had low factor loading (0.322). Without Q3_7 and Q3_9, the Manichean outlook contained only one item. The factor was excluded, which significantly improved all fit indices, including χ^2/df (see Table 2). The original version of the Castanho Silva et al. (2018) scale was not internally coherent in our dataset, but a model with two latent factors, anti-elitism and centrism (without item Q3_5), fit the data well in both the datasets. It is recommended to inspect the items, translations, etc., and to test the entire scale in future research. This scale was excluded from further analyses in this paper. McDonald's omega coefficients for the Manichean Outlook were $\omega = .46$ (exploratory dataset), $\omega = .51$ (confirmatory dataset). Reliability coefficients for anti-elitism were $\omega = .79$ (exploratory dataset), $\omega = .76$ (confirmatory dataset), for modified subscale of centrism $\omega = .72$ (exploratory dataset), $\omega = .82$ (confirmatory dataset).

The original scale of Schulz et al. (2018) consists of twelve items divided into three dimensions: anti-elitism (Q3_10 – Q3_13), sovereignty (Q3_14 – Q3_17) and homogeneity of people (Q3_18 – Q3_21). For wording of the items named Q3_10 – Q3_21 in our dataset, see the *Populist Attitudes Scales* in the Supplemental materials. The results for the original scale showed high factor loadings (all greater than 0.61). The model had CFI and TLI above the recommended minimums of .95, RMSEA just slightly below .08 at .074, and SRMR < .05, however, χ^2/df was slightly above

3, which is not acceptable with $p < 0.05$. According to the modification indices, one item (Q3_16) had cross-loadings with other factors. The item was removed, which improved the values of fit indices including χ^2/df, which was 2.2 (see Table 2). To improve the model, Q3_12 was removed because it loaded on other factors, and was correlated with items from other factors. After removing the item, the model fit increased, χ^2/df = 1.7, but $p = .028$, CFI and TLI > .99, RMSEA and SRMR < .05. Q3_19, which loaded on other factors, was removed. The modified model (Q3_12, Q3_16, Q3_19 removed) had a good fit for all indices, as well as insignificant χ^2 ($p = .076$). The same model tested on the confirmatory dataset had RMSEA < .08 and SRMR < .05, TLI and CFI > .95, and χ^2/df was lower than the acceptable value of 3 ($p < .001$) (see Table 2). The original scale performed well as a first-order model, however, the authors operationalized populism as a second-order factor. In our dataset, the sovereignty factor caused problems in converging the model; thus, further inspection of the items within the factor in the second-order factor model is required. The model was further tested by using a first-order factor model. Reliability coefficients for modified subscales of anti-elitism were $\omega = .8$ (exploratory dataset), $\omega = .83$ (confirmatory dataset); for the sovereignty $\omega = .85$ (exploratory dataset), $\omega = .87$ (confirmatory dataset); and for the homogeneity of people $\omega = .82$ (exploratory dataset), $\omega = .8$ (confirmatory dataset).

The original unidimensional scale of Akkerman et al. (2014) consists of six items (anti-elitism: Q3_11, Q3_13; popular sovereignty: Q3_16, Q3_17, Q3_22; Manichean outlook: Q3_23). The original scale had CFI, TLI > .95, SRMR < .08, but also significant χ^2 and RMSEA > .08. After removing two items (Q3_16 and Q3_11, in that order), fit indices improved. A modified scale with four items representing the three aspects of populist attitudes was confirmed on an exploratory dataset. The fit indices indicated a good internal consistency. However, the model had a positive χ^2 and RMSEA > .08 on a confirmatory dataset, therefore, the scale was not fully confirmed and thus was not part of further validation (see Table 2).

The original unidimensional scale of Van Hauwaert et al. (2020) consists of eight items (six items from Akkerman scale, and two items added by the authors: Q3_24 and Q3_25 representing anti-elitism). For wording of the items in our dataset, see the *Populist Attitudes Scales* in the Supplemental materials. The original scale had an acceptable fit for the TLI (> .90) and SRMR (< .08) indices, but not for the other indices. However, the factor loadings for all the items were greater than .649. Items suggested in the modification indices as problematic (Q3_16, Q3_22, Q3_25) were gradually removed, which increased all fit indices to desirable levels (TLI, CFI > .99, RMSEA, SRMR < .06). The modified model was not fully confirmed on a confirmatory dataset, but the removal of Q3_11 improved all the fit indices. The four-items version (anti-elitism: Q3_13, Q3_24; sovereignty: Q3_17, Manichean outlook: Q3_23) fits the data well (see Table 2). Reliability coefficients for modified scale were $\omega = .82$ (exploratory dataset), $\omega = .84$ (confirmatory dataset).

The results of the CFA for original and modified scales are summarized in Table 2.

Convergent and discriminant validity

A correlation matrix of the scales/subscales on the exploratory dataset is provided in Figure S1 in the Supplemental materials. The results for the confirmatory dataset are depicted in Figure 1. The convergent validity of the modified scale by Van Hauwaert et al. (2020) was supported by the assumed high positive correlation with the anti-elitism (.78) and sovereignty subscales (.77) of the modified scale of Schulz et al. (2018), and with conspiracy thinking (.55). The discriminant validity of the scale was proven by the anticipated negative moderate correlation with political trust (-.43).

Table 2 Confirmatory Factor Analysis Models

Populist attitudes scales		Dataset	Fit statistics							Factor loading			
			χ^2	df	p	CFI	TLI	RMSEA	SRMR	Min	Avg	Max	Lowest
Castanho Silva et al. (2018)[a]	Original	E	103.648	14	<.001	.933	.919	.124	.075	.261	.6	.819	Q3_9
	Modified[c]	E	6.302	4	.178	.998	.998	.037	.017	.58	.75	.84	Q3_2
		C	1.364	3	.714	1	1	<.001	.008	.49	.6	.909	Q3_2
Schulz et al. (2018)[a]	Original	E	88.444	27	<.001	.972	.984	.074	.047	.61	.76	.879	Q3_12
	Modified	E	23.423	15	.076	.0995	.996	.037	.029	.647	.78	.897	Q3_11
		C	43.488	15	<.001	.985	.99	.068	.033	.68	.79	.869	Q3_21
Akkerman et al. (2014)[b]	Original	E	67.76	8	<.001	.95	.96	.134	.056	.531	.69	.777	Q3_11
	Modified	E	3.006	2	.222	.998	.998	.035	.014	.693	.72	.774	Q3_22
	Modified	C	13.289	2	.001	.962	.904	.116	.028	.602	.697	.768	Q3_22
Van Hauwaert et al. (2020)[b]	Original	E	175.308	14	<.001	.875	.929	.167	.071	.589	.69	.794	Q3_11
	Modified	E	7.803	5	.167	.996	.997	.037	.019	.62	.71	.807	Q3_11
		C	22.384	5	<.001	.984	.987	.092	.023	.713	.75	.848	Q3_11
	Modified $_{\text{-Q3\_11)}}$	C	3.777	2	.151	.998	.997	.046	.012	.685	.76	.866	Q3_24

Notes. E = Exploratory dataset, N_{exp} = 416; C = Confirmatory dataset, N_{conf} = 416; [a] Three-dimensional scale, [b] Unidimensional scale, [c] Two-dimensional scale; X^2 = chi-square, df = degrees of freedom, p = statistical significance, CFI = comparative fit index, TLI = Tucker–Lewis index, RMSEA = root mean square error of approximation, SRMR = standardized root mean squared residual, MIN = minimum AVG = average, MAX = maximum.

Figure 1 also indicates moderate correlation with relative deprivation (.45), which suggests that those who feel deprived in society score higher on the populist attitudes scale. The scale had a moderate positive correlation with pluralist attitudes (.29) and the homogeneity of people (.27) and a small correlation with mistrust of experts (.19). Confirmed moderate to high associations with the subscales of populist attitudes measured by the modified scale of Schulz et al. (2018), conspiracy thinking, relative deprivation and political distrust indicate convergent and discriminant validity of the scale. The anticipated negative relations of populist attitudes with elitist and pluralist attitudes were not supported, which will be the subject of the discussion.

In the case of the modified scale of Schulz et al. (2018), average and minimum scores for the aggregate version were calculated. The average score in the scale of Schulz et al. (2018) was mostly driven by the sovereignty subscale, whereas the minimum score was mostly driven by the homogeneity subscale. The average score also showed higher correlations with the other constructs than the minimum score (see Figure 1). Anti-elitism was highly correlated with the sovereignty of people (.61), conspiracy thinking (.50) and moderately with relative deprivation (.37) and pluralist attitudes (.34). Sovereignty was positively correlated with conspiracy thinking (.52), relative deprivation (.42), pluralist attitudes (.24), and mistrust of experts (.20) and was negatively correlated with political trust (-.32). Homogeneity had a moderate positive correlation with sovereignty of people (.37), relative deprivation (.37), and conspiracy thinking (.33). The convergent validity of the modified scale proposed by Schulz et al. (2018) was supported by the anticipated high positive correlations with conspiracy thinking (.58) and relative deprivation (.51). Discriminant validity was proven by anticipated negative moderate correlation with political trust (-.34). The anticipated negative associations of populist attitudes and pluralist and elitist attitudes were not supported and are discussed later.

Predictive validity

The results of the binary logistic regression suggest that the populist attitudes measured by the modified scale of Schulz et al. (2018) (minimum score across dimensions) allow us to predict planned voting for populist political parties. The average score across dimensions yielded the best results, allowing for the prediction of past and future voting for populist parties. The modified scale of Van Hauwaert et al. (2020) was able to predict past but not future voting for populist parties (see Table 3).

Table 3 Binomial Logistic Regression (Can a vote (previous or future) for populist parties be predicted based on the populist attitude?)

Populist attitudes	B (SE)	OR[CI]	B (SE)	OR[CI]
Minimum score on Schulz et al. (2018) modified	.1 (.07)	1.10 [.95, 1.28]	.23** (.07)	1.26[1.08, 1.47]
Average score on Schulz et al. (2018) modified	.28** (.10)	1.33[1.09, 1.64]	.21* (.10)	1.24[1.01, 1.54]
Average score on Van Hauwaert et al. (2020) modified	.25** (.09)	1.29[1.08, 1.55]	.07 (.09)	1.08[.90, .29]

Notes. Confirmatory dataset (N = 416); B = Estimate; SE = Standard Error; CI = Confidence Interval; OR = Odds Ratio

Different results were estimated on an exploratory dataset, where populist attitudes measured by average and minimum scores in the modified scales of Schulz et al. (2018) and Van Hauwaert et al. (2020) succeeded in predicting the previous vote for

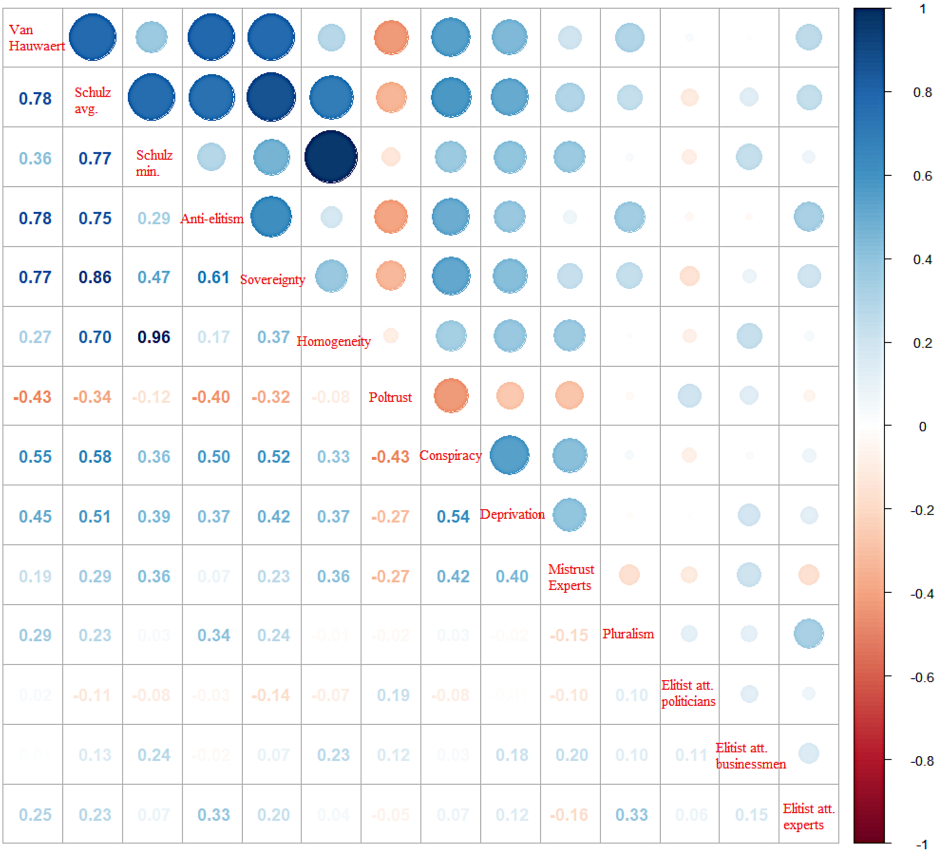

Figure 1 Correlation matrix of populist attitudes scales and selected covariates

Notes. Confirmatory dataset (N = 416), Van Hauwaert = average score on modified unidimensional scale of Van Hauwaert et al. (2020); Schulz avg. = average score on three dimensions of modified scale of Schulz et al. (2016); Schulz min. = the lowest score on three dimensions of modified scale of Schulz et al. (2016); Anti-elitism, Sovereignty, Homogeneity = Average score on individual dimension from Schulz et al. (2018) modified scale; Poltrust = political trust scale; Conspiracy = = conspiracy thinking scale; Deprivation = relative deprivation scale; Mistrust experts = Mistrust of experts scale; Pluralism = pluralist attitudes scale; Elitist att. politicians, businessmen, experts - individual items of Elitist attitudes scale

populist parties, but just a minimum score on Schulz's scale also predicted potential future votes for populist parties (see Results of binary logistic regressions on an Exploratory dataset that are provided in Table S3 in the Supplemental materials). Increasing populist attitudes were in both cases associated with an increased likelihood of previous (or future) voting for populist political parties.

Intergroup differences – sociodemographic variables

Gender

Two sample t-tests proved that men and women did not differ in the level of populist attitudes measured by the minimum score on Schulz et al.'s (2018) modified scale (t(412.81) = -.738, p = .461), average score on Schulz et al.'s (2018) modified scale (t(413.34) = -.407, p = .685), or average score on Van Hauwaert et al.'s (2020) modified scale (t(413.96) = -.017, p = .987). For descriptive statistics, see Table 4. Our assumption about men having a higher level of populist attitudes has not been supported.

Table 4 Descriptive statistics for populist attitudes across gender subgroups

Populist attitudes	Group	N	Mean	SD
Schulz et al. (2018) modified scale[1]	Men	210	3.76	1.30
	Women	206	3.86	1.35
Schulz et al. (2018) modified scale[2]	Men	210	5.13	.95
	Women	206	5.13	.99
Van Hauwaert et al. (2020) modified scale[2]	Men	210	5.72	1.09
	Women	206	5.72	1.08

Notes. Confirmatory dataset (N = 416), [1] the lowest score on three dimensions, [2] average score; 7-point response scale

Age categories

The ANOVA results showed no significant differences between age groups in the level of populist attitudes measured by the minimum score on Schulz et al.'s (2018)

Table 5 Descriptive statistics for populist attitudes across age groups

Populist attitudes	Group	N	Mean	SD
Schulz et al. (2018) modified scale[1]	18-25	54	3.60	1.17
	26-35	84	3.61	1.17
	36-45	84	3.87	1.20
	46-55	74	3.86	1.54
	56-65	77	4.06	1.47
	66-75	43	3.82	1.30
Schulz et al. (2018) modified scale[2]	18-25	54	4.78	.95
	26-35	84	4.88	.91
	36-45	84	5.12	.93
	46-55	74	5.28	1.11
	56-65	77	5.34	.89
	66-75	43	5.42	.85
Van Hauwaert et al. (2020) modified scale[2]	18-25	54	5.34	1.14
	26-35	84	5.46	1.17
	36-45	84	5.71	.98
	46-55	74	5.83	1.16
	56-65	77	6	.94
	66-75	43	6.02	.94

Notes. Confirmatory dataset (N = 416), [1] the lowest score on three dimensions, [2] average score; 7-point response scale modified scale (F(5) = 1.29, p = .267), but there were moderate significant differences between the two youngest and two oldest age groups (d = .51 – .70) in the case of an average score on Schulz et al.'s (2018) modified scale (F(5) = 4.512, p = .0005). Our assumption that older respondents have stronger populist attitudes compared to younger participants has been supported. For descriptive statistics, see Table 5. Confidence intervals for each pairwise comparison are shown in Figure 2.

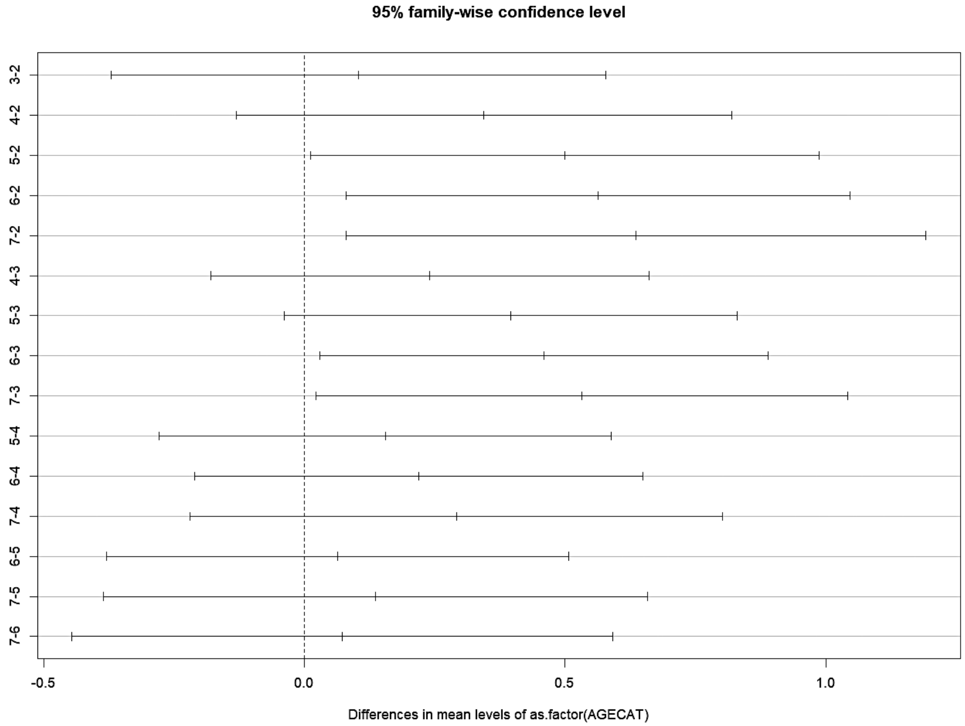

Figure 2 Confidence intervals for each pairwise comparison (Average score on the modified scale of Schulz et al., 2018)

Notes. Confirmatory dataset (N = 416), Age groups: 2 (18–25), 3 (26–35), 4 (36–45), 5 (46–55), 6 (56–65), 7 (66–75)

In the case of the modified scale of Van Hauwaert et al. (2020), the ANOVA results showed significant differences among the age groups (F(1) = 4.258, p = < .001). Pairwise comparisons using the t-test with Tukey's HSD correction showed moderate differences between the youngest and the two oldest age groups, and between age groups 26–35 and 56–65 (d = .51 – .65). As expected, older respondents had stronger populist attitudes compared to younger participants. For descriptive statistics, see Table 5. Confidence intervals for each pairwise comparison are shown in Figure 3.

Education

The ANOVA results showed significant differences across educational groups in the level of populist attitudes (the lowest score on the three dimensions), measured using Schulz et al.'s (2018) modified scale (F(3) = 9.043, p < .001). Pairwise comparisons using the t-test with Tukey's HSD correction showed moderate differences between

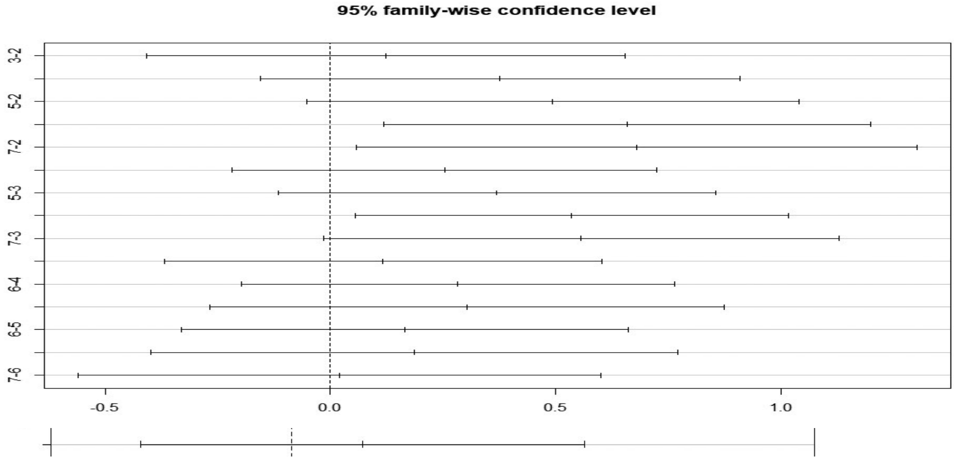

Figure 3 Confidence intervals for each pairwise comparison (Average score on the modified scale of Van Hauwaert et al., 2020*)*

Notes. Confirmatory dataset (N = 416), Age groups: 2 (18–25), 3 (26–35), 4 (36–45), 5 (46–55), 6 (56–65), 7 (66–75)

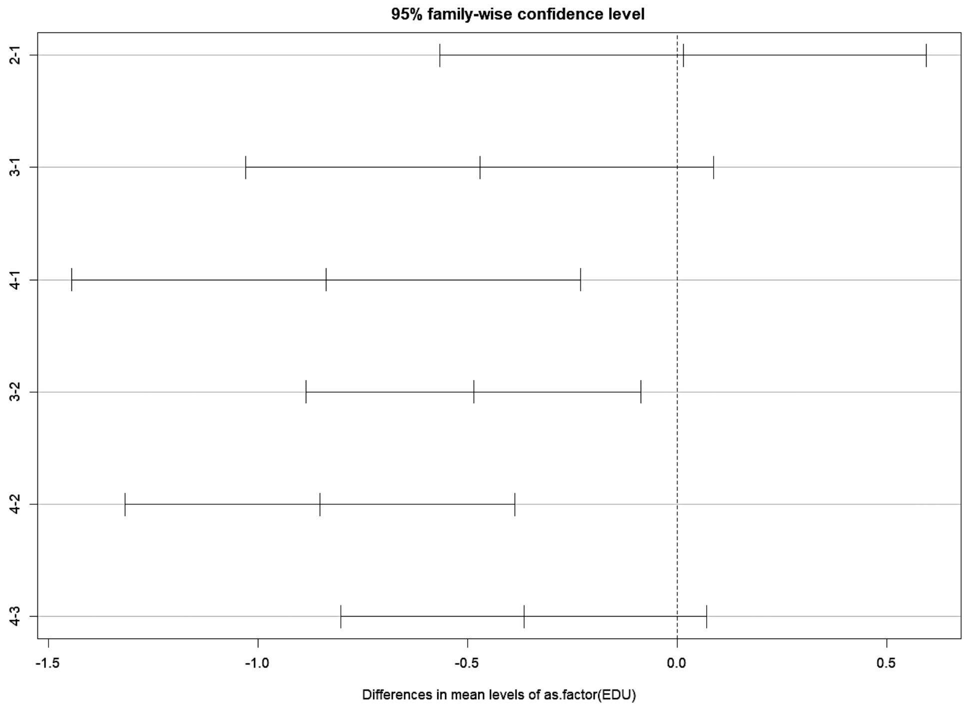

Figure 4 Confidence intervals for each pairwise comparison (Minimum score on the modified scale of Schulz et al., 2018)

Notes. Confirmatory dataset (N = 416), 1 = primary education, 2 = lower secondary education, 3 = upper secondary education, 4 = tertiary education

Table 6 Descriptive statistics for populist attitudes across educational groups

Populist attitudes	Group	N	Mean	SD
Schulz et al. (2018) modified scale[1]	Primary	45	4.17	1.47
	Lower Secondary	119	4.18	1.21
	Upper Secondary	163	3.70	1.38
	Tertiary	89	3.33	1.08
Schulz et al. (2018) modified scale[2]	Primary	45	5.09	1.28
	Lower Secondary	119	5.39	.94
	Upper Secondary	163	5.14	.89
	Tertiary	89	4.77	.87
Van Hauwaert et al. (2020) modified scale[2]	Primary	45	5.6	1.28
	Lower Secondary	119	5.91	1.0
	Upper Secondary	163	5.76	1.02
	Tertiary	89	5.44	1.06

people with tertiary education and those with primary education ($p = .0023$, $d = .65$) or lower secondary education ($p < .0001$, $d = .74$); and small differences between people with lower and upper secondary education ($p = .009$, $d = .37$). As expected, respondents with lower education had stronger populist attitudes compared to more educated participants. For the descriptive statistics, see Table 6. For confidence intervals for each pairwise comparison, see Figure 4.

Differences among education groups were proven in the case of the average score on three dimensions of the modified scale of Schulz et al. (2018) ($F(3) = 7.418$, $p <.001$). Pairwise comparison using t-test with Tukey's HSD correction proved small to moderate ($d = .43 – .69$) differences between people with tertiary education and both lower and upper secondary education. Our assumption that respondents with lower education have stronger populist attitudes compared to more educated participants has been supported. For the descriptive statistics, see Table 6. For confidence intervals for each pairwise comparison, see Figure 5.

The ANOVA results showed statistically significant differences in populist attitudes measured by the modified scale of Van Hauwaert et al. (2020) ($F(3) = 3.529$, $p = .015$). Pairwise comparisons using t-test with Tukey's HSD correction proved small differences between people with lower secondary education compared to tertiary education ($p = .01$, $d = .44$). For the descriptive statistics, see Table 6. For confidence intervals for each pairwise comparison, see Figure 6.

DISCUSSION

This study aimed to test selected populist attitudes scales (Akkerman et al., 2014; Castanho Silva et al., 2018; Schulz et al., 2018; Van Hauwaert et al., 2020) on a Slovak sample and validate the scale/s for future research in Slovakia. This study is worthwhile because it analyzes populist attitudes in a setting where populist parties have been both in power and in the opposition for more than a decade, unlike the Western European context, where populist forces have been for years mostly in opposition.

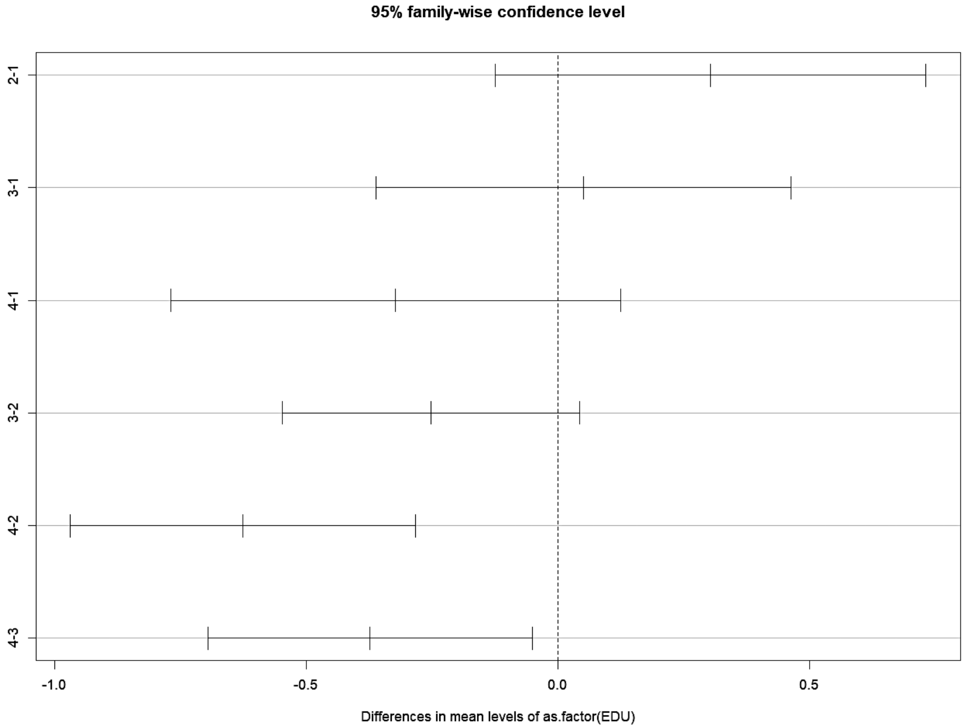

Figure 5 Confidence intervals for each pairwise comparison (Average score on the modified scale of Schulz et al., 2018)

Notes. Confirmatory dataset (N = 416), 1 = primary education, 2 = lower secondary education, 3 = upper secondary education, 4 = tertiary education

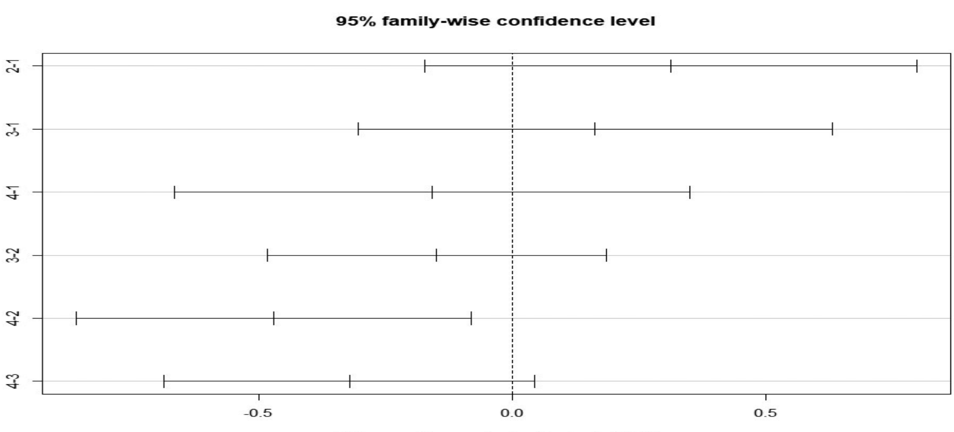

Figure 6 Confidence intervals for each pairwise comparison (Average score on the modified scale of Van Hauwaert et al., 2020)

Notes. Confirmatory dataset (N= 416), 1 = primary education, 2 = lower secondary education, 3 = upper secondary education, 4 = tertiary education

Construct validity

The original versions of all selected populist attitude scales did not fit the data well, while some items had small factor loadings or cross-loadings, thus, according to the results of CFA, they didn't prove to be internally coherent. However, reliability coefficients measured by McDonalds ω were at least acceptable for most subscales. Two populist attitudes scales were excluded from further validation:

- The modified scale (without 2 items with the lowest factor loadings – Q3_16 and Q3_11) of Akkerman et al. (2014) with four items was not fully confirmed in the confirmatory dataset, therefore, it was not included in the validation process.
- The populist attitudes scale of Castanho Silva et al. (2018), with adequate psychometric properties among Slovak respondents, consists of five rather than nine items: three items (with one negative worded) representing anti-elitism (the only subscale that works adequately) and two items representing people centrism. The Manichean outlook subscale had low reliability and was excluded in the CFA. The exclusion of Q3_5 from the subscale of people centrism indicates that spending time with people as only one of the ways of knowing the will of people (but not very common in Slovakia), does not seem to be so important because it is sufficient for politicians to listen to the will of the people and follow it. Due to a two-factor structure that failed to fully capture populist attitudes, the scale was excluded from further validation. However, these items should be inspected and tested in future studies.

After several modifications to the models (excluding three or four problematic items), two populist attitude scales (Schulz et al., 2018; Van Hauwaert et al., 2020) presented a good model fit and high factor loadings, indicating high internal coherence. All the subscales also had high reliability and so the scales were considered appropriate for further validation.

The populist attitudes scale of Schulz et al. (2018)

The final version of Schulz et al.' s (2018) scale, with adequate psychometric properties among Slovak respondents, consisted of three items per subscale: anti-elitism, popular sovereignty and homogeneity. However, the items of anti-elitism were operationalized differently than in the case of Castanho Silva et al. (2018), where more emphasis was placed on the negative practices and activities of government politicians. Schulz et al.'s (2018) anti-elitism items describe the processes through which (rather passive) politicians in general (not just the government) are involved. However, both ways of operationalising the items of anti-elitism work adequately in the Slovak context. Popular sovereignty is also unlike Castanho Silva et al.'s (2018) people centrism (where only simple listening to and following the will of the people is required) but is perceived through a more intense participation of people in the government (through the expression of opinions and referendums). This fact may not be popular in every country and does not have to be perceived as something positive, but rather somewhat annoying as probably is the case in Slovakia. This may be the reason for the problem of the subscale of popular sovereignty when testing populism as a second-order factor, and this way of measuring popular sovereignty is probably less appropriate for the Slovak context. Items with cross-loadings (Q3_12, Q3_16, Q3_19, one from each dimension) were removed from the model (some reasons are mentioned later in a contextual explanation).

The populist attitudes scale of Van Hauwaert et al. (2020)

The final scale of Van Hauwaert et al. (2020), with adequate psychometric properties among Slovak respondents, was a unidimensional scale consisting of four instead of eight items representing all three aspects: anti-elitism, centrism, and the Manichean outlook. Anti-elitism was measured by two items; the first (Q3_24) examined the perception of political elites as negative and self-serving authorities. It was the one of two items added by Van Hauwaert et al. (2020) that was useful and without which Akkerman et al.'s (2014) version does not work very well. The second (Q3_13) examines perceptions of the distance between people and elites. This item has a long history and is effective in many populist attitudes scales (Akkerman et al., 2014; Hawkins et al., 2012; Schulz et al., 2018). Popular sovereignty was measured only by the belief that politicians should follow the people's will (Q3_17). This item was also effective in many populist attitudes scales (Akkerman et al., 2014; Hawkins et al., 2012; Schulz et al., 2018). The Manichean outlook was measured by the item (Q3_23) of Akkerman et al. (2014), originally suggested by Hawkins et al. (2012) based on the concept of the stealth democracy (Hibbing & Theiss-Morse, 2002 in Hawkins et al., 2012). The Manichaean outlook of politics differs from those of Castanho Silva et al. (2018), which was embedded in how populism / populists framed elites and people and was rather a part of anti-system (anti-elitism) item(s). Moreover, the components can overlap, because they are viewed as unidimensional. This Manichean worldview, as a part of anti-system (anti-elitism) item(s), seems more suitable in the Slovak context. However, more items are required for a valid conclusion.

The exclusion of item Q3_16 and Q3_22 (popular sovereignty) may be due to aspects specific to the country or period of data collection, which is discussed below in the contextual explanation. Q3_25 (anti-elitism) examined the perceived continuous negative impact of the behavior and interest of elites on respondents' lives. As one of two items added by Van Hauwaert et al. (2020) to Akkerman et al.'s (2014) populist attitudes scale, was not beneficial for the Slovak environment. Finally, Q3_11 (also anti-elitism) was excluded based on the confirmatory dataset, but was not problematic for the exploratory dataset. Schulz et al. (2018) retained this item in the model, therefore, the five-items version should be verified in future studies.

Convergent and discriminant validity

The modified scales of Schulz et al. (2018) and Van Hauwaert et al. (2020) scored similarly in the external validity tests. The scales were moderately correlated with political trust (negatively) and conspiracy thinking (positively), which was empirically supported by Erisen et al. (2021) or Castanho Silva et al. (2020). The expected negative correlations with pluralist and elitist attitudes were not observed in this study, however, Schulz et al. (2018) found small negative correlations. Akkerman et al. (2014) suggested that populist, elitist, and pluralist scales are not mutually exclusive (p. 1336). A positive correlation between populist and pluralist attitudes was also obtained in Akkerman et al. (2014), which may indicate that at least some people holding populist attitudes accept different views, supporting the need to compromise or prefer politicians to ordinary people. The correlation matrix showed a positive relationship between populist attitudes and relative deprivation, indicating that populist attitudes rise with increasing feelings of exclusion and deprivation. According to Elchardus and Spruyt (2016), feelings of relative deprivation are related not only to populism, but also to economic vulnerability and a declinist view of how a society is evolving. Because socioeconomic characteristics can affect the extent to which individuals are included in the political system, populism is particularly appealing to the marginalized or excluded (Mudde & Rovira Kaltwasser, 2017). This also corresponds to the main idea of populism, which is to return the „average" citizen to political interest and return the voice to people who do not feel represented by the political elite.

Predictive validity

The ability of populist attitude scales to predict voting behaviour was not consistent between datasets. However, for both dataset, the average score on the modified scale of Schulz et al. (2018) and Van Hauwaert et al. (2020) allowed a prediction of vote for populist parties in the last general elections. The intention to vote for populists can only be consistently predicted by using the minimum score on the scale of Schulz et al. (2018). The distinctions in other predictions raise the question regarding the accuracy of predictions and calls for future verifications.

Similar results were provided by Jungkunz et al. (2021), who pointed out the failure of populist attitudes scales to explain voting for populist parties in countries where populist leaders are in power. Stanley (2011) had a similar consideration when his battery of populism-specific questions failed to predict electoral behavior of populist political parties among Slovak respondents more than a decade ago. This is the case of Slovakia, which should focus future research attention on the "anti-elite" feeling, which is dominantly directed towards political parties and government within the existing populist attitudes scales. As soon as populists are in power, they may shift their anti-elite rhetoric onto other groups, such as academics, journalists, judges, minorities, which will also change the follower's perception (Jungkunz et al., 2021, p.16). Another relevant aspect is that many people do not participate in elections in Slovakia[3], and many voters who do, decide immediately before voting, based on their current emotions/sympathies, the election campaigns, or immediate circumstances, which is seen in the high volatility of electoral behavior.

Akkerman et al. (2014) confirmed relations of populist attitudes and intention to vote for populist political parties among Dutch respondents in the original study. Van Hauwaert et al. (2020) also affirmed the predictive power of the modified Akkerman et al. (2014) scale in nine European countries but without specific information about the political situation in those countries or any comparison of results. In addition, on a cross-national comparison of several scales by Castanho Silva et al. (2020), the scales of Akkerman et al. (2014) and Schulz et al. (2018) had high external validity, allowing the prediction of populist party identification in selected countries (Italy, Spain, France, but not in Mexico, the United Kingdom, or Greece, the only country with populists in government during the data collection). The predictive power of the populist scale yields mixed results, so scales with good psychometric characteristics predicted populist support in some countries but not in others. This problem persists also in countries with populists in government.

On the other hand, inconsistency between measured attitudes and overt behavior is a well known in psychology research (Wicker, 1969). Later, it was revealed that different contextual factors (situational, personal, etc.) may play a role in stronger or weaker attitude-behavior (in)consistency (Glasman & Albarracín, 2006; Haddock & Maio, 2004).

Measurement invariance and group differences

Configural, metric, (partial) scalar, and strict invariance were established for both scales in case of gender and education. In the case of age groups, strict invariance was not established for the scale of Van Hauwaert et al. (2020) on the exploratory dataset; however it is not considered a limitation for calculating group differences because the residuals are not part of the latent factor (Vandenberg & Lance, 2000). Differences between younger and older respondents and also between more and less educated

[3] https://volby.sme.sk/parlamentne-volby/2020/vysledky

respondents were proved for at least one of the tested scales. Higher populist attitudes were supported among older and less educated people. Other studies also support the existence of a negative relationship between education and populist attitudes (Balta et al., 2021; Boscán et al., 2018). Elchardus and Spruyt (2016) found that age has no effect on populist attitudes but men and less educated people were more prone to populism (p. 124). In our study, men and women scored the same in both scales of populist attitudes. As populist preferences can have different demographic correlates depending on the country (Rooduijn, 2018), it seems that it is higher age and lower education in Slovakia, which is in line with the vulnerability of these specific groups in the Slovak cultural context.

Contextual explanations

There are several possible reasons why some items or subscales (e.g., low reliability of the Manichean outlook and elitist attitudes; low score in homogeneity of people; troubles with sovereignty in a second-order model) have proven problematic.

First, the specific political context of Slovakia should be considered. The Manichaean outlook on politics means a moral struggle, where one side is obviously good and the other bad. The political scene in Slovakia is characterised by a higher number of heterogeneous political parties. Populist parties can be found both in opposition and in government, so populism is a common part of the political culture in Slovakia. However, specific cultural context and history could also play a role. Many people are aware that the actions of political parties can be judged as both good and bad. People may get lost in the evaluation of the moral quality of such a political scene, or their evaluations may arise as a secondary aspect of probably more relevant ingroup and outgroup processes. As it was said in the introduction, populist politicians in Slovakia are creating new categories, artificial „evil groups" that are still changing (Vašečka, 2022). However, how to operationalise the Manichaean outlook as a component of populism with such chameleon characteristics requires further research. Furthermore, some authors do not even consider the Manichaean outlook as a separate component of populist attitudes, but is captured across all dimensions, e.g., Schulz et al. (2018); Elchardus and Spruyt (2016); Hobolt et al. (2016). This approach seems more appropriate for the Slovak context.

Second, it is possible that responses have been influenced by the immediate experience of people through the period of the COVID-19 pandemic. Data collection during the third wave of the pandemic may increase the awareness of the great heterogeneity of Slovak people, its diversification or even irrationality, and their own helplessness. It may explain the exclusion of items (Q3_12, Q3_16 and Q3_19 in scale of Schulz et al. (2018)) in the Slovak context and also the lower average level of perceived homogeneity of people. Also, the pandemic time may be understood as a period when we do not want to hand power over to common people, which may explain also the exclusion of Q3_16 and Q3_22 in the scales of Schulz et al. (2018) and Van Hauwaert et al. (2020) applied in the Slovak context.

Limitations

The present study also has some limitations. Data were collected through a research agency using an online panel of respondents who are willing to participate. The nature of the online population carries a risk of "non-naive participants". The drawbacks of two instruments (Schulz et al., 2018; Van Hauwaert et al., 2020) is the absence of negatively-worded items that could lead to acquiescence bias. Furthermore, data collection was conducted during the third wave of the COVID-19 pandemic, thus some anti-elitism and elitism items might cause troubles for some respondents who may hold populist attitudes but also prefer the country to be run by politicians over ordinary people. Also the polarization of society and differences in attitudes towards vaccinations or government measures were present both in public and in the government, which may have an impact on responses on the perceived homogeneity of people, considered to be one of the essential dimensions of populist attitudes.

Practical recommendations

Out of all the scales included in the study, two of them, one proposed by Schulz et al. (2018) and the second by Van Hauwaert et al. (2020), were in modified versions validated for measuring populist attitudes in Slovakia. Both had their strengths and weaknesses on our dataset.

It can be stated that the modified scale of Van Hauwaert et al. (2020) seems complex in the Slovak context because it includes items representing different components of populism on a single scale, i.e, anti-elitism, which also includes Manichean outlook, and popular sovereignty. This scale was based on the first and most intensively used populist attitudes scale by Akkerman et al. (2014) (Van Hauwaert et al., 2020). Although the scale was able to predict voting for populist parties in past elections, it could not robustly predict voting for populist parties in future elections based on our data. Electoral behaviour in Slovakia is, however, characterized by high volatility (Gyárfášová et al., 2017; Gyárfášová & Slosiarik, 2020) and may be influenced by many factors. The populist attitudes measured by the modified version of this scale were uniformly understood by gender, age and educational groups, so the scale is appropriate for measuring differences between these groups, and the results may be interpreted in the same way.

Schulz et al. (2018) enriched the scale of populist attitudes through the dimension of homogeneity of people but did not include a Manichean outlook as a separate component. The weakness but also the strength of this scale is precisely this dimension, which may not reliably capture populist attitudes in times of crises and increasing polarization of society, when even ordinary people do not perceive themselves as united or homogeneous. And since populism thrives in times of crises, when society becomes even more fragmented, it may be difficult to capture the unity of ordinary people. The solution could be to measure the unity of the people, not on the basis of common values and interests, but on the basis of a common goal, which is to return power to the people. Such an item would thus combine two other dimensions – the homogeneity of the people and their sovereignty into a single item, which has been already proposed as a centrism scale by Castanho Silva et al. (2018). The relevance of this approach is supported by the fact that even multidimensional approaches have abbreviated versions of populist scales (with one item from each dimension) and are ultimately understood as a unidimensional construct (Castanho Silva et al., 2018; Van Hauwaert et al., 2020). On the other hand, this subscale may more accurately differentiate people with populist attitudes from people who "just" reject the elites and call for a "more democratic" governance of the country. The scale is invariant across gender, age and education groups so this latent construct can be measured and interpreted in the same way across groups. Like the previous scale, this scale also had problems predicting voting behavior. The differences in predictions depend on whether we operationalize populist attitudes as an average on the three subscales or as a minimum score following Goertz's approach (2006, 2020, as cited in Wuttke, 2020). The average score (saturated mainly by people's sovereignty) allowed past behavior to be stably predicted, but it was the minimum score (saturated by homogeneity of people) that stably predicted the future choice of populist parties on both datasets, confirming the need for further investigation of populist attitudes in conjunction with voting behavior.

CONCLUSION

Despite the mentioned limitations, the results of the present study provide support for good psychometric properties of the Slovak version of two scales for measuring populist attitudes – the modified scale of Schulz et al. (2018) and the scale proposed by Van Hauwaert et al. (2020), thus creating room for its further use in research of populist attitudes in Slovakia.

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SÚHRN

Validácia škál populistických postojov na Slovensku

Ciele. Výskum populizmu posledné desaťročie napreduje, čoho znakom je navrhnutie niekoľkých škál na meranie populistických postojov. Žiadna z nich nebola overená v podmienkach Slovenska, kde sú populisti dlhodobo súčasťou koalície aj opozície. Cieľom tejto štúdie bolo overiť psychometrické vlastnosti štyroch škál populistických postojov, ktoré sa často používajú a overujú v medzinárodných výskumoch, na slovenskej vzorke.

Participanti a postup výskumu. Dáta 832 respondentov, získané pomocou online panelu v novembri 2021, boli analyzované pomocou softvéru R. Výskumná vzorka bola reprezentatívna z hľadiska distribúcie pohlavia, veku, úrovne vzdelania a príslušnosti k regiónom Slovenska.

Štatistická analýza. Pôvodné škály boli testované pomocou exploračného súboru dát (N = 416). Modifikované škály boli overené pomocou konfirmačného súboru dát (N = 416).

Výsledky. Výsledky ukázali, že pôvodné škály nezodpovedajú dátam. Po niekoľkých úpravách však boli dve škály validované na slovenskej vzorke. Škály boli invariantné pre pohlavie, vek a dosiahnuté vzdelanie.

Limity. Možným nedostatkom validovaných škál populistických postojov je nestabilita pri predikcii volebného správania, ktorá je diskutovaná v kontexte výsledkov iných štúdií a tiež v politickom a kultúrnom kontexte Slovenska.