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Adaptation and Shortening of the Right-Wing Authoritarianism (RWA) Scale for the Ukrainian Context: Evidence from a Bias-Corrected Confirmatory Factor Analysis of Longitudinal Data

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The Right-Wing Authoritarianism (RWA) scale remains a cornerstone in the study of authoritarian attitudes, yet its adaptation and validation outside Western contexts is limited. This study develops a Ukrainian adaptation of the RWA scale and evaluates its psychometric properties using longitudinal data ($N = 99$). Two bias-correction techniques – acquiescent response style (ARS) and common method bias (CMB) correction – were validated, employed, and compared. While response bias appeared negligible at the aggregate level, its distorting effect on the factor structure was substantial, leading previous studies to misinterpret RWA's dimensional composition. Bias detection through split-half correlation density analysis confirmed systematic method effects, while bias-corrected data revealed a markedly different factor structure, free from confounding artifacts. Findings indicate that applying bias-correction techniques is essential when analyzing RWA's factor structure. Two shortened versions of the scale – a unidimensional and a three-dimensional – were developed for the Ukrainian context, both demonstrating strong reliability and longitudinal measurement invariance.

Key words: Right-Wing Authoritarianism scale (RWA), Ukraine, bias correction, common method bias (CMB), acquiescent response style (ARS), longitudinal measurement invariance.

Цимбал Тарас, Сазонова Валерія. Адаптація та скорочення шкали правої авторитарності (RWA) для українського контексту: результати конфірматорного факторного аналізу лонгітюдних даних, скоригованих на похибку вимірювання. Шкала правої авторитарності (Right-Wing Authoritarianism, RWA) залишається ключовим інструментом у дослідженні авторитарних установок, проте її адаптація та валідизація поза межами західного контексту є обмеженою. У цьому дослідженні розроблено українську адаптацію шкали RWA та оцінено її психометричні властивості на основі лонгітюдних даних ($N = 99$). Було валідизовано, застосовано та порівняно дві процедури корекції похибки вимірювання – корекцію на схильність погоджуватися (ARS) і корекцію на загальну методологічну похибку (CMB). Хоча на агрегованому рівні відповідні похибки виявилися незначними, їх вплив на факторну структуру був суттєвим, що призвело до хибної інтерпретації вимірів RWA в попередніх дослідженнях. Аналіз щільності розподілу кореляцій випадкових поділів шкали підтвердив систематичні методологічні ефекти, а очищені від похибок дані продемонстрували значно відмінну факторну структуру, вільну від спотворювальних артефактів. Отримані результати вказують на необхідність застосування процедур корекції похибок під час аналізу факторної структури RWA. Для українського контексту розроблено дві скорочені версії шкали – одновимірну та тривимірну; обидві продемонстрували високу надійність і лонгітюдну інваріантність вимірювання.

Ключові слова: шкала правої авторитарності (RWA), Україна, корекція похибки вимірювання, загальна методологічна похибка (CMB), схильність погоджуватися (ARS), лонгітюдна інваріантність вимірювання.

INTRODUCTION

For decades, the Right-Wing Authoritarianism (RWA) scale has remained the industry standard in authoritarianism research. Initially developed by Altemeyer in the early 1980s to replace the deeply problematic yet once-dominant F-scale (Altemeyer, 1981), RWA was designed to overcome its predecessor's flaws and has since become the most widely used measurement tool among authoritarianism researchers worldwide. Continued maintenance of the scale, provided both by its author (Altemeyer, 1996, 1998, 2006, 2022) and other researchers (Bizumic & Duckitt, 2018; Duckitt et al., 2010; Funke, 2005) has sustained its relevance by addressing ongoing challenges and meeting the demand for an ever-shorter yet reliable measurement tool, replicable across diverse national contexts.

RWA's enduring success arises from several advantages over the F-scale and its derivatives, including a balanced composition, a simpler dimensional structure, brevity, relevant stimuli, and strong reliability and validity (Sazonova & Tsymbal, 2024, p. 22). So endowed with virtues, the scale proliferated far beyond its Canadian cradle, reaching English-speaking countries and much beyond. Its reported national adaptations include Argentina (Etchezahar et al., 2011), Brazil (Vilanova et al., 2023), Chile (Cárdenas & Parra, 2010), Colombia (García-Sánchez et al., 2022), El Salvador (Orellana, 2018), and Peru (Rottenbacher de Rojas, 2012) in Latin America; Czechia (Chylíkova & Buchtík, 2016), France (Dru, 2003), Germany (Schneider & Lederer, 1995), Greece (Sochos, 2021), Hungary (Zsolt, 2004), Italy (Rattazzi et al., 2007), Poland (Radkiewicz, 2011), Spain (Garzón, 1992), and Sweden (Zakrisson, 2005) in Europe; Israel (Rubinstein, 1996) and Turkey (Güldü, 2011) in the Middle East; South Africa (Edwards & Leger, 1995; Gray & Durrheim, 2006) in Africa; Indonesia (Ji, 2007), Japan (Takano et al., 2021), and Taiwan (Huang, 2007) in Asia.

Reports of RWA usage extend to numerous other countries, although without detailed accounts of its adaptation or psychometric performance. These include Belgium (Duriez & Van Hiel, 2002), China (Hsu & Wang, 2024), Croatia (Šram, 2020), Czechia (Dunbar & Simonova, 2003), Denmark (Bartusevičius et al., 2020), Egypt and Morocco (Lemieux et al., 2017), Georgia (Despotashvili, 2016), India (Felix & Chaube, 2021), Iraq (Jabr, 2021), Malaysia (Rashid, 2021), Norway (Halkjelsvik & Rise, 2014), the Netherlands (Onraet et al., 2021), Pakistan (Siraaj et al., 2022), Serbia (Nikolov, 2024), and the former Soviet Union (McFarland et al., 1992). Additionally, several cross-national studies have included samples from Bangladesh (Peterson et al., 2011), Ethiopia (D'Urso et al., 2024), Kuwait, Saudi Arabia, Yemen, Oman, Bahrain, UAE, Syria, and Sudan (Albaghli & Carlucci, 2021), Bulgaria (Kemmelmeier et al., 2003), and Ukraine (Van Hiel & Kossowska, 2007). RWA's global reach cements its status as a tool of choice for emerging authoritarianism researchers.

Its global success notwithstanding, scholars have questioned RWA's theoretical validity, particularly its conflation with conservatism. Ray (1985, p. 272) famously critiqued RWA as "just another conservatism scale," arguing that it measures political ideology rather than authoritarianism per se. Subsequent studies have reinforced this critique, demonstrating that right-wing ideology explains 60 % of RWA variance, with the remaining 40 % due to non-ideological sources (Radkiewicz, 2011). Additionally, its validation largely relied on other ideological scales, raising concerns of tautological measurement (Oesterreich, 2005). Empirical findings also indicate that RWA exhibits weak or non-existent correlations with non-attitudinal authoritarianism measures. Ray (1985, p. 272) reported an RWA correlation $r = -.024$ with his Directiveness scale, while Sazonova & Tsymbal (2024, pp. 25–26) obtained $r = .14$ ($p > .05$) and $r = .21$ ($p < .05$) correlations with Oesterreich's authoritarianism scale in two respective waves of a longitudinal study in Ukraine.

Despite these theoretical reservations, RWA remains one of the most widely used tools in political psychology due to its predictive power. It has a proven record of predicting prejudice and discrimination (Dunbar & Simonova, 2003; Hunsberger, 1996; Laythe et al., 2001; Nesdale et al., 2012; Putranto et al., 2021; Whitley & Lee, 2000) and has been linked to support for violence (Benjamin, 2006), capital punishment (McKee & Feather, 2008), torture (Benjamin, Jr, 2016), and war (Felix & Chaube, 2021; Hsu & Wang, 2024; Terrizzi & Drews, 2005; Wollast et al., 2024). RWA has also been associated with conspiracy thinking and rejection of science (Grzesiak-Feldman & Irzycka, 2009; Kerr & Wilson, 2021), as well as political orientation (Ballout et al., 2023; Güldü, 2020; Kehn et al., 2023; Vasilopoulos & Lachat, 2018), anti-egalitarianism (Cotterill et al., 2014), and affective polarization (Renström et al., 2022). Individuals with higher RWA scores are more prone to stigmatizing people with mental health issues (DeLuca et al., 2018; Szabó et al., 2024), supporting suspension of civil liberties (Cohrs et al., 2005; Duckitt, 1993; Manson, 2020), reacting negatively to asylum seekers (Onraet et al., 2021), and denying environmental concerns (Lalot et al., 2022).

While RWA primarily captures political attitudes, some studies have noted its connections to personality traits (Choma et al., 2019; Ekehammar et al., 2004; Hotchin & West, 2018; Hou et al., 2024; Zebarjadi et al., 2023), and found its genetic underpinnings (Kandler et al., 2016).

With the immense literature devoted to RWA and the plethora of research opportunities it offers for collaboration and intellectual exchange across disciplines and borders, it is unfounded to expect its demise, despite its theoretical inconsistency. A cautious and measured application of the scale, fully recognizing its scope as a measure of conservative-traditional worldview rather than authoritarianism per se, can help tip the balance its misnomer engenders and the benefits it offers to international scholarship in favor of the latter.

In the last few decades, the resurgence of populist and extremist parties, alongside democratic backsliding, has renewed interest in authoritarianism research (Tsymbal & Sazonova, 2023, pp. 11–12). Authoritarianism theory has long linked perceived threat to authoritarian responses, predating the RWA scale (Feldman & Stenner, 1997, pp. 741–742; Sales, 1973). Recent studies confirm this, distinguishing dormant authoritarian predispositions from threat-activated manifestations (Feldman & Stenner, 1997; Hastings & Shaffer, 2005; Shaffer & Hastings, 2007). Since RWA measures political attitudes rather than fixed traits, it should fluctuate with perceived threat.

Russia’s 2022 invasion of Ukraine presents an anomaly in this respect. Despite intensified threats, support for democracy in Ukraine rose after 2014 and surged post-2022 (Alexseev & Dembitskyi, 2024). The lack of RWA-based studies in Ukraine limits theoretical insights into authoritarian traits and threat responses. While cross-national studies included Ukrainian samples (Van Hiel & Kossowska, 2007; Vargas-Salfate et al., 2020), no detailed RWA adaptation exists, hindering analysis of Ukraine’s resistance to authoritarianism amid global trend toward it and contrary to theoretical predictions.

Objective of the Study. This article develops a Ukrainian adaptation of the RWA scale and evaluates its psychometric properties through a longitudinal study. It contributes to methodological debates by addressing measurement biases, examining dimensionality, and comparing bias-correction techniques. In addition, it proposes shortened versions of the scale, enhancing its applicability for future research.

1. RIGHT-WING AUTHORITARIANISM (RWA) SCALE

1.1. Composition and Versions of RWA Scale

Altemeyer conceptualized right-wing authoritarianism (RWA) as a unitary construct encompassing authoritarian submission (deference to authority), authoritarian aggression (hostility toward dissenters), and conventionalism (adherence to societal norms). The original 24-item RWA scale (Altemeyer, 1981, pp. 148, 305–306) was designed to capture this interplay, later expanded to 30 items (Altemeyer, 1996, pp. 14–15). The most widely used version, referred to hereafter as RWA20, was introduced in 2006 with 20 retained items and two practice questions (Altemeyer, 2006, pp. 11–12). While a 10-item version was published in 2022, RWA20 remains dominant in research (Chylíkova & Buchtík, 2016; DeLuca et al., 2018; Güldü, 2011; Peterson et al., 2011; Siraaj et al., 2022; Vilela et al., 2016; Wedell & Bravo, 2022).

Several abridged versions of the RWA scale have also gained traction. Some national adaptations replicated Zakrisson’s 15-item Swedish version (Ekehammar et al., 2004; Etchezahar et al., 2011; García-Sánchez et al., 2022; Tan et al., 2016; Zakrisson, 2005), while others followed Rattazzi et al.’s 14-item Italian version (Kehn et al., 2023; Rashid, 2021; Rattazzi et al., 2007). Another significant variant, Funke’s RWA^{3}D scale, was explicitly designed to separate the three subdomains and has been employed in studies beyond its German origins (D’Urso et al., 2024; Funke, 2005; Guidetti et al., 2021; Jackson & Gaertner, 2010; Kandler et al., 2016). Additionally, researchers have created ad hoc adaptations, some reducing the scale to as few as three items (Duckitt & Sibley, 2010; Lilly et al., 2024; Satherley et al., 2021; Vargas-Salfate et al., 2020). The proliferation of modified versions reflects the scale's flexibility and ability to maintain conceptual and empirical coherence across studies (Altemeyer, 1996, p. 52).

Most RWA items contain multiple semantic elements referring to several subdomains in order to reinforce its cohesion. This design, however, has led to a significant degree of confusion, when the need to classify scale items by subdomains emerged (Funke, 2005, pp. 197–198; Mavor et al., 2009). Three primary approaches have been advanced: a priori classification based on conceptual definitions (García-Sánchez et al., 2022; Smith & Winter, 2002), rater-based classification (Funke, 2005), and post-hoc classification through factor analysis (Cárdenas & Parra, 2010; Etchezahar, 2012; Mavor et al., 2010; Rattazzi et al., 2007; Takano et al., 2021).

Table 1 summarizes classification attempts, exposing inconsistencies in subdomain assignments. The multi-barreled nature of the items is evident in cross-loading items, which are attributed to multiple subdomains within a study, and swing items, which shift their subdomain alignment across studies. Some items, such as the cross-loading item 3, have been assigned to both aggression and submission in multiple studies (Etchezahar, 2012; Rattazzi et al., 2007; Takano et al., 2021). Others, like swing item 15, have been classified under submission in some studies (Cárdenas & Parra, 2010; García-Sánchez et al., 2022) but as conventionalism in others (Takano et al., 2021). A particularly extreme case, swing item 9, has been assigned to all three subdomains across different studies.

In response to these inconsistencies, some researchers have excluded cross-loading items (Mavor, 2012; Rattazzi et al., 2007), reworded ambiguous items (Funke, 2005), or created separate scales for each subdomain (Duckitt et al., 2010). However, most studies fail to systematically analyze item semantics before applying factor analysis, leaving ample room for confirmation bias in the post-hoc interpretation of factor analysis results.

To address these issues, we conducted an a priori semantic analysis of RWA20 items, independent of factor analysis. Prior studies defined aggression, submission, and conventionalism narrowly, emphasizing high-intensity expressions while neglecting weaker manifestations or their conspicuous absence. This asymmetry, particularly for aggression, wrongly confined it to pro-trait items, overlooking its expected presence in con-trait items, thus leading to its statistical overlap with the direction of wording (Duckitt et al., 2010, p. 689; Funke, 2005, p. 202; Mavor et al., 2010, p. 28).

We propose a bipolar conceptualization of RWA subdomains, accounting for both authoritarian tendencies and their opposites. Aggression includes both punitive inclinations and their manifest absence in con-trait items. Conventionalism reflects both norm adherence and openness to change, while submission encompasses deference to authority and active skepticism toward it. This framework offers a more balanced approach to understanding RWA's subdomains, avoiding structural imbalances caused by earlier one-sided interpretations.

1.2. Semantic Breakdown of RWA20 Items

Each RWA item consists of a stimulus and a disposition toward it, with some items containing multiple stimuli and dispositions linked to different subdomains. In pro-trait item 3, "mighty leader" represents submission, "destroy" signals aggression, and "radical new ways" belong to conventionalism. Con-trait item 9 mirrors this structure: "free thinkers" reflect anti-submission, "needs" denotes anti-aggression (embracing dissent), and "traditional ways" represent conventionalism, but "defy" reverses the stance into anti-conventionalism.

Aggression in pro-trait items appears in verbs denoting suppression (e.g., "destroy," "silence"), whereas con-trait items express anti-aggression via positive wording (e.g., "are good and virtuous," "nothing wrong"). The symmetry between these expressions implies aggression's continuity across the pro/con divide.

Conventionalism is marked by references to tradition, moral authority, and societal norms, while submission involves figures of authority and their challengers. Aggression is embedded in disposition, whereas conventionalism and submission reside in stimuli, making aggression present in all items referencing challengers to authority and norms. Table 1 presents our semantic classification of RWA20 items (column 3).

Table 1

Decomposition of RWA20 Items into Subdomains: a Priori (etic), CFA-Based (emic), and Literature Classifications

Item Codes	**Semantic breakdown of items into subdomain indicators**	**A Priori Sub-Domain Assignment**	**Post-hoc CFA-Based Classification**	**(Mavor, 2012)**	**(Smith & Winter, 2002)**	**(Rattazzi et al., 2007)**	**(Funke, 2005)**	**(García-Sánchez et al., 2022)**	**(Etchezahar, 2012)**	**(Cárdenas & Parra, 2010)**	**(Takano et al., 2021)**
rwa03	Our country desperately needs *a mighty leader* **[Submission]** who will do what has to be done to *destroy* **[Aggression]** the *radical new ways and sinfulness* **[Conventionalism]** that are ruining us	ACS	F1	A		AS		A	AS	A	AS
rwa04	*Gays and lesbians* **[anti-Conventionalism]** are just as *healthy and moral* **[anti-Aggression]** as anybody else	**AC**		**C**	**C**	**C**			**C**		**C**
rwa05	It is always better to trust the judgment of *the proper authorities in government* **[Submission]** *and religion* **[Conventionalism]** *than to listen* **[Aggression]** to the noisy rabble-rousers in our society who are trying to create doubt in people's minds	ACS	F1	AS	S				AS		AS
rwa06	*Atheists and others who have rebelled against the established religions* **[anti-Conventionalism]** are no doubt every bit *as good and virtuous* **[anti-Aggression]** as those who attend church regularly	**AC**		**C**		**C**			**C**		**C**
rwa07	The only way our country can get through the crisis ahead is to get back to *our traditional values* **[Conventionalism]**, put some *tough leaders in power* **[Submission]**, and *silence* **[Aggression]** the troublemakers spreading bad ideas	ACS	F1	A		AS			AS		AS
rwa08	There is *absolutely nothing wrong* **[anti-Aggression]** with *nudist camps* **[anti-Conventionalism]**	**AC**		**C**		**C**			**C**		**C**
rwa09	Our country *needs* [anti-Aggression] *free thinkers* **[anti-Submission]** who have the courage to defy *traditional ways* **[anti-Conventionalism]**, even if this upsets many people	**ACS**	**F2∩ F3**	**SC**				**S**		**A**	**C**
rwa10	Our country will be destroyed someday if we do not *smash* **[Aggression]** the perversions eating away at *our moral fiber and traditional beliefs* **[Conventionalism]**	AC	F2	A							AS
rwa11	Everyone *should have their own* **[anti-Aggression]** *lifestyle, religious beliefs, and sexual preferences* **[anti-Conventionalism]**, even if it makes them different from everyone else	**AC**	**F2**	**C**	**C**	**C**			**C**		**C**

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rwa12	*The "old-fashioned ways" and the "old-fashioned values"* **[Conventionalism]** still show the best way to live	C	F1∩F2					C		C	
rwa13	You have to *admire* **[anti-Aggression]** those who challenged *the law* **[anti-Submission]** and the *majority's view by protesting for women's abortion rights, for animal rights, or to abolish school prayer* **[anti-Conventionalism]**	**ACS**	**F2∩F3**								
rwa14	What our country really needs is a *strong, determined leader* **[Submission]** who will *crush evil* **[Aggression]**, and *take us back to our true path* **[Conventionalism]**	ACS		A	A	AS	A		AS		AS
rwa15	Some of the *best* **[anti-Aggression]** people in our country are those who are *challenging our government* **[anti-Submission]**, *criticizing religion, and ignoring the "normal way things are supposed to be done."* **[anti-Conventionalism]**	**ACS**	**F3**	**SC**				**S**		**S**	**C**
rwa16	*God's laws about abortion, pornography and marriage must be strictly followed* **[Conventionalism]** before it is too late, and *those who break* **[Submission]** them *must be strongly punished* **[Aggression]**	ACS	F1							C	
rwa17	There are many *radical, immoral people* **[Conventionalism]** in our country today, who are trying to ruin it for their own godless purposes, whom the *authorities should put out of action* **[Aggression]**	AC		A	A			A		C	AS
rwa18	*A "woman's place" should be wherever she wants to be* **[anti-Aggression]**. *The days when women are submissive to their husbands* **[anti-Submission]** *and social conventions* **[anti-Conventionalism]** belong strictly in the past	**ACS**	**F1**	**SC**			**S**				**C**
rwa19	Our country will be great if we *honor the ways of our forefathers* **[Conventionalism]**, *do what the authorities tell us to do* **[Submission]**, and *get rid of* **[Aggression]** the "rotten apples" who are ruining everything	ACS	F3	A							AS
rwa20	There is *no "ONE right way" to live life* **[anti-Conventionalism]**; *everybody has to create their own way* **[anti-Aggression]**	**AC**		**C**		**C**			**C**		**C**
rwa21	Homosexuals and feminists should be *praised* **[anti-Aggression]** for being brave enough to *defy* **[anti-Submission]** *"traditional family values"* **[anti-Conventionalism]**	**ACS**	**F1∩F2**	**C**		**C**			**C**		**C**
rwa22	This country would work a lot better if certain groups of *troublemakers* **[Submission]** would just *shut up* **[Aggression]** and accept their group's *traditional place* **[Conventionalism]** in society	ACS									

Note: Codes (column 1) and subdomain markers (columns 3-12) for con-trait items are presented in **bold**. The classes from column 4 were interpreted as follows: F1 – aggression, F2 – conventionalism, F3 – submission (see section 3.7).

The semantic breakdown reveals that twelve items incorporate all three subdomains (ACS), seven exhibit aggression-conventionalism nexus (AC), and only one represents pure conventionalism (C). Aggression appears in 19 items, conventionalism in all 20, and submission in 12, indicating that conventionalism is the most consistently represented subdomain.

Interpretation of scale items can be approached from two perspectives: an etic perspective, which represents an external, researcher-driven categorization of item content, and an emic perspective, which reflects how respondents internally perceive, process, and respond to these items. While the etic perspective defines the full range of possible semantic elements embedded in the scale, the emic perspective determines which of these elements participants recognize and incorporate into their responses (Iliescu et al., 2024, p. 98; Pike, 1967, p. 38).

Our semantic breakdown outlines the etic meanings embedded in the scale but does not dictate how respondents interpret items. Emic meanings emerge from subjective perceptions, reflected in the factor structure of the data. While some semantic components influence responses (dominant), others may remain latent (dormant), depending on social context and cognition. Identifying which components respondents engage with and what meaning factors assume in empirical models requires post hoc analysis, which follows in later sections.

Several key expectations emerge from the semantic decomposition:

1. Factor structure should not follow the pro/con divide. Since all three subdomains are evenly distributed across both pro-trait and con-trait items, a factor structure that collapses into two factors mirroring this divide would strongly suggest a measurement artifact.

2. Factor instability across time and samples is expected. Given the multi-barreled design of items, their alignment with factors is likely to vary depending on emic conditions such as social context and individual cognitive focus. Consequently, longitudinal and cross-sample inconsistencies in factor structure are anticipated, with swing items emerging as a natural outcome.

3. Cross-loading items provide interpretive value. Items that relate to multiple subdomains are expected and should not be dismissed. Their semantic composition can offer insights into the latent variables they bridge, helping to clarify the conceptual structure of subdomains within the scale.

4. A three-factor solution is the most theoretically coherent outcome. The presence of three distinct subdomains suggests that the scale is most likely to resolve into a three-factor structure. However, this solution may be attenuated by emic reception.

5. Aggression and conventionalism are the most interwoven subdomains due to their frequent co-presence in item semantics, while submission is the most structurally distinct. The latter is thus the most likely candidate to form a statistically distinct factor.

2. METHODOLOGY

2.1. Sample

The study included 99 participants (76 women, 23 men; M_{age} = 21.4, SD_{age} = 2.4, range = 17–28), comprising 55 BA students, 19 MA students, and 25 recent graduates (within the past five years) from the Faculty of Sociology, Taras Shevchenko National University of Kyiv. By the second survey wave, all participants were one year older, with the distribution shifting to 41 BA, 25 MA, and 33 graduates.

2.2. Materials

2.2.1. Right-Wing Authoritarianism Scale

This study employed Altemeyer’s 20-item Right-Wing Authoritarianism Scale (RWA20) (Altemeyer, 2006). As no formal Ukrainian adaptation had been reported, we conducted a new translation. Two translators independently produced Ukrainian versions, which were reconciled in a meeting with a third expert judge. The finalized translation was pretested with five respondents, whose feedback informed minor revisions.

The translation strategy prioritized fidelity to Altemeyer’s original wording while upholding linguistic naturalness in Ukrainian. Some cultural and contextual adjustments were necessary:

- *Religion* (items 5, 15) was replaced with *church* to emphasize its institutional role and association with religious authorities, rather than implying a pluralistic or individual choice.
- *Nudist camps* (item 8) were changed to *nudist beaches*, a more recognizable concept in Ukraine.
- *School prayers* (item 13), an unfamiliar practice in Ukraine, were replaced with *religious instruction at schools*, referencing the elective *Foundations of Christian Ethics* course, which was approved for school curricula in 2005 and later became a subject of societal debate in 2019–2022.

- *Country* (item 3) was substituted with *state* to clarify that the *mighty leader* refers to a formal political leader rather than an informal figure.

Apart from these modifications, all original item references were retained. Participants rated their agreement with each statement on a 9-point Likert-type response scale (*totally disagree* to *totally agree*), presented as empty checkboxes without numerical labels. The final RWA score was calculated as the unweighted mean of all included items, with con-trait items reverse-coded. The full Ukrainian translation and item statistics are provided in Appendix A.

2.2.2. Oesterreich's Authoritarianism Scale

Oesterreich's Authoritarianism Scale (OAS) measures authoritarianism through 23 non-attitudinal items (Oesterreich, 1996, 2005). A validated Ukrainian translation was used (Sazonova, 2018, pp. 113–114). Each item contrasts two statements (e.g., "I always do things in the same way" vs. "I like to give new things a try"), rated on a five-point scale. This format minimizes acquiescence bias by avoiding explicit disagreement. Response options were displayed as unnumbered checkboxes to minimize directional influence.

Originally, the scale had 9 pro-trait and 14 con-trait items. To enhance comparability with RWA20, five con-trait items with the lowest item-total correlations were removed, yielding an 18-item balanced version (OAS18). Final scores were computed as unweighted means of pro-trait items and reverse-coded con-trait items.

Unlike RWA, which starts with a pro-trait item, OAS begins with a con-trait item, potentially shifting response patterns (Weijters et al., 2013). Prior research found OAS weakly correlated with attitudinal authoritarianism, including right-wing extremism (Oesterreich, 2005, p. 292), the F-scale, and VSA (Sazonova & Tsymbal, 2024, pp. 25–26). In this study, OAS18 primarily serves as a methodological control for bias in RWA20 responses.

2.3. Procedure

2.3.1. Data Collection

Participants were invited via email to complete an online questionnaire hosted on the *LimeSurvey* platform. They were informed that the study aimed to examine the social and political attitudes of Ukrainian youth and that participation was voluntary, responses were anonymous, and all data would be treated confidentially.

The first wave of data collection took place in June 2023, followed by the second wave in June 2024. In the first wave, 399 invitations were sent with two follow-up reminders, resulting in 172 initiated responses, of which 133 were completed. Due to panel attrition, the second wave yielded 103 completed responses.

2.3.2. Response Screening

Responses were screened based on inter-wave correlations of RWA20 scores for each participant. The lowest 1 % were excluded, removing three responses.

After adjusting for acquiescent response style, responses with the most extreme 1 % of absolute inter-wave differences were also removed. The final sample included 99 responses.

2.4. Data Analysis

Due to the small, homogeneous sample, a number of sample size-adjusted methodological choices were made. Bias correction has led to redundancy between the similarly worded items 3 and 14, requiring application of a shrinkage estimator (Ledoit & Wolf, 2004; Opgen-Rhein & Strimmer, 2007; Schäfer & Strimmer, 2005) to preserve the full item set and resolve matrix singularity. Shrinkage intensity parameters for CMB-corrected data were estimated with the *cov.shrink()* function from the *corpcor* package in R (Schafer et al., 2022): $\lambda_{var1} = .3918$, $\lambda_{var2} = .3388$ for variance shrinkage and $\lambda_{cor1} = .1661$, $\lambda_{cor2} = .1724$ for correlation shrinkage. MLR estimator was applied to uncorrected data to address small sample and model complexity (Shi et al., 2021).

Model fit was evaluated following SEM guidelines (Kline, 2023). Models were compared using the corrected Akaike Information Criterion (AICc) (Burnham & Anderson, 2004; Hurvich & Tsai, 1989, p. 270) and the sample-size adjusted Bayesian Information Criterion (SABIC) (referred to as *IB* in Sclove, 1987, p. 336; and *BIC** in Yang, 2006) instead of AIC/BIC. Consistent with Hu & Bentler (1999, pp. 27–28), the combination of CFI ≥ .95 and SRMR ≤ .09 was prioritized given our study conditions. Longitudinal measurement invariance (MI) was evaluated, following standard MI criteria (Chen, 2007, p. 501). All CFA models were estimated with R package *lavaan* (Rosseel et al., 2024).

3. RESULTS

3.1. Overview of Analysis

Our analysis followed a structured sequence: (1) exploratory factor analysis (EFA) on the raw data to assess factor structure and identify bias effects, (2) statistical correction to isolate the content-based factor structure, (3) EFA followed by confirmatory factor analysis (CFA) to validate the dimensionality of bias-corrected data, (4) testing the same models with an additional method factor on the raw data, (5) constructing an abridged version of RWA20 for better interpretability of factors and (6) assessing the longitudinal MI of the abridged version.

Table 2

Scale Reliability Statistics for RWA20 Based on Raw, ARS-Corrected, and CMB-Corrected Data

	RAW DATA		ARS-CORRECTED DATA		CMB-CORRECTED DATA	
	Wave 1	Wave 2	Wave 1	Wave 2	Wave 1	Wave 2
Cronbach's α (standardized)	0,87	0,88	0,89	0,90	0,89	0,89
Guttman's λ_6	0,91	0,92	0,90	0,92	0,90	0,92
McDonald's ω	0,87	0,88	0,89	0,90	0,89	0,90
Mean interitem correlation	0,25	0,27	0,28	0,31	0,28	0,30
Share of CMB variance in the scale score	6 %	9 %	-	-	-	-
Test-retest reliability (ICC)	$0{,}69^{***}$		$0{,}69^{***}$		$0{,}72^{***}$	

*** p < .001

Our findings indicate high reliability of RWA20 in Ukraine and position it within the reported international range. The observed α-reliabilities (.87 and .88) and mean inter-item correlations (.25 and .27) (Table 2) closely correspond to those reported by Altemeyer in the UK, Australia, Italy, Spain, Germany, the United States, and Canada, where α ranged from .82 to .90, while inter-item correlations fell between .18 and .32 (Altemeyer, 2022). Similar levels were found in Turkey (α = .85; Güldü, 2011) and Pakistan (.90; Siraaj et al., 2022), whereas lower values were reported in Czechia (.78; Chylíkova & Buchtík, 2016, p. 14) and Brazil (.69; Vilela et al., 2016). Across alternative RWA versions, reliability estimates range between .68 and .94 in the US (Wilson & Sibley, 2013, p. 279), Sweden (Zakrisson, 2005, p. 867), Italy (Rattazzi et al., 2007, p. 1129), Poland and Ukraine (Van Hiel & Kossowska, 2007, p. 20).

The scale demonstrated moderate test-retest reliability (*ICC* = .69, $p < .001$), assessed using a two-way random-effects model with absolute agreement (single measures). This falls within Cicchetti's (1994, p. 286) "good" reliability range and Koo & Li's (2016, p. 158) "moderate" category. Few studies report test-retest reliability of RWA: Rubinstein documented .83 in Israel and Palestine, Altemeyer obtained .85 (Rubinstein, 1996, p. 22), Sibley et al. (2007, p. 363) indicated .81 (*N*=165), Asbrock et al., (2010, p. 332) (*N* = 127) and Sibley & Duckitt (2013, p. 456) (*N* = 147) reported .79, all over significantly shorter intervals of six months or less, except for Sibley & Duckitt, where two waves of study were 13 months apart. Given the one-year interval between the waves, our reliability level is sufficient to establish the scale as a reliable measure. Importantly, this year was formative for many participants, shaped by both academic exposure to sociopolitical topics and escalating existential threats due to the ongoing Russian war on Ukraine, both of which are known to influence RWA (Dunwoody & McFarland, 2018; Nikolov, 2024). Despite these influences, our test-retest reliability is close to values obtained by Liu et al. (2008, p. 120) from a partly student sample in Taiwan for an 8-item version of RWA: .68 (*N* = 88) and .71 (*N* = 73).

Item-level test-retest reliabilities ranged from .30 to .66, except for item 6 (*ICC* = .18; see Appendix A), likely due to participants' growing familiarity with atheism through their studies. However, removing any item did not improve α-reliability. Item 8 (*nudist beaches*) stood out as the clearest outlier in both waves, making it the strongest candidate for elimination.

Consistent with student samples globally, RWA scores clustered in the lower half of the 9-point scale (M_1 = 2,62, SD_1 = 0,98; M_2 = 2,68, SD_2 = 1,04). RWA20 and OAS18 showed no significant correlation in wave 1 (r_1 = .06, p_1 = .57) and only a marginal correlation in wave 2 (r_2 = .20, p_2 = .05), reinforcing the argument that RWA weakly captures pure authoritarianism.

Table 3

Factor Loadings for One-, Two-, and Three-Factor EFA Solutions Based on Raw 1st-Wave Data

Item Code	1-Factor	2-Factor		3-Factor		
	F1	F1	F2	F1	F2	F3
rwa21	**0,610**	**0,690**		**0,612**		**0,207**
rwa04	**0,437**	**0,652**		**0,646**		
rwa18	**0,485**	**0,643**		**0,595**		
rwa13	**0,550**	**0,564**		**0,452**		**0,349**
rwa20	**0,399**	**0,559**		**0,615**		
rwa11	**0,253**	**0,509**	**-0,233**	**0,551**		
rwa16	0,639	0,482	0,248	0,537	0,317	
rwa12	0,562	0,422	0,219	0,455	0,260	
rwa22	0,558	0,414	0,222	0,325		0,307
rwa06	**0,483**	**0,409**		**0,348**		**0,200**
rwa09	**0,363**	**0,302**		**0,218**		**0,269**
rwa14	0,439		0,759		0,763	
rwa03	0,437		0,726		0,705	
rwa05	0,326	-0,222	0,644		0,616	
rwa07	0,552		0,578		0,612	
rwa10	0,685	0,266	0,539	0,342	0,594	
rwa17	0,453		0,470		0,337	0,252
rwa19	0,318		0,423		0,279	0,269
rwa15	**0,293**		**0,231**			**0,806**
rwa08	**0,191**		**0,182**			**0,295**
Variance per factor	0,22	0,16	0,15	0,15	0,13	0,07
Variance per model	0,22	0.31		0,35		

Table 4

Factor Loadings for One-, Two-, and Three-Factor EFA Solutions Based on 1st-Wave CMB-Corrected Data

Item code	1-factor	2-factor		3-factor		
	F1	F1	F2	F1	F2	F3
rwa10	0,693	0,717		0,317	0,464	
rwa16	0,647	0,673		0,576	0,230	
rwa20	**0,482**	**0,642**			**0,665**	
rwa04	**0,577**	**0,592**		**0,536**		
rwa14	0,555	0,583			0,697	
rwa12	0,565	0,555		0,456	0,205	
rwa07	0,552	0,542		0,414	0,214	
rwa11	**0,348**	**0,512**			**0,580**	
rwa21	**0,713**	**0,505**	**0,285**	**0,260**	**0,297**	**0,284**
rwa18	**0,606**	**0,495**		**0,655**		
rwa03	0,516	0,471		0,335	0,207	
rwa06	**0,496**	**0,309**	**0,251**		**0,258**	**0,280**
rwa05	0,292	0,201		0,612	-0,282	
rwa15	**0,269**	**-0,403**	**0,887**		**-0,290**	**0,775**
rwa13	**0,607**	**0,311**	**0,397**		**0,326**	**0,481**
rwa22	0,562	0,267	0,395	0,419		0,333
rwa19	0,278		0,344			0,348
rwa17	0,424		0,341			0,352
rwa08	**0,175**		**0,318**			**0,309**
rwa09	**0,369**		**0,283**		**0,247**	**0,358**
Variance per factor	0,26	0,21	0,09	0,12	0,11	0,08
Variance per model	0,26	0,30		0,31		

Table 5

Factor Loadings for One-, Two-, and Three-Factor EFA Solutions Based on Raw 2nd-Wave Data

Item Code	1-Factor	2-Factor		3-Factor		
	F1	F1	F2	F1	F2	F3
rwa13	**0,521**	**0,725**		**0,681**		
rwa20	**0,428**	**0,675**		**0,640**		
rwa18	**0,407**	**0,639**		**0,607**		
rwa11	**0,498**	**0,634**		**0,611**		**0,222**
rwa04	**0,632**	**0,570**		**0,557**		**0,394**
rwa21	**0,576**	**0,536**		**0,500**		
rwa15	**0,316**	**0,422**		**0,395**		
rwa16	0,598	0,416	0,263	0,381	0,426	
rwa12	0,549	0,359	0,264	0,329	0,254	
rwa06	**0,415**	**0,279**		**0,251**	**0,405**	
rwa10	0,628		0,740		0,503	0,381
rwa14	0,397	-0,235	0,715	-0,231	0,470	0,347
rwa03	0,462		0,677		0,589	
rwa07	0,583		0,660		0,628	
rwa22	0,495		0,453		0,282	0,289
rwa17	0,260		0,409		0,331	
rwa19	0,395		0,385		0,532	
rwa05	0,400		0,370		0,402	
rwa09	**0,524**	**0,250**	**0,346**	**0,245**		**0,569**
rwa08	**0,194**		**0,243**			**0,280**
Variance per factor	0,23	0,16	0,15	0,14	0,12	0,06
Variance per model	0,23	0.31		0,32		

Table 6

Factor Loadings for One-, Two-, and Three-Factor EFA Solutions Based on 2nd-Wave CMB-Corrected Data

Item Code	1-Factor	2-Factor		3-Factor		
	F1	F1	F2	F1	F2	F3
rwa18	**0,568**	**0,677**		**0,603**		
rwa07	0,584	0,671		0,566	0,264	-0,226
rwa20	**0,601**	**0,635**		**0,549**		
rwa16	0,598	0,594		0,506		
rwa06	**0,412**	**0,520**		**0,467**		
rwa05	0,371	0,474		0,445		
rwa19	0,358	0,470		0,483		0,201
rwa03	0,474	0,395		0,410		0,310
rwa21	**0,618**	**0,363**	**0,311**	**0,275**	**0,459**	
rwa12	0,539	0,347	0,240	0,266	0,384	
rwa22	0,467	0,259	0,252		0,462	
rwa17	0,207	0,127		0,122		
rwa09	**0,503**		**0,705**		**0,578**	**0,224**
rwa04	**0,673**		**0,616**		**0,713**	
rwa11	**0,616**		**0,600**		**0,537**	
rwa13	**0,687**	**0,249**	**0,510**	**0,236**	**0,290**	**0,399**
rwa10	0,694	0,289	0,474		0,605	
rwa14	0,385		0,467			0,486
rwa15	**0,395**		**0,378**			**0,586**
rwa08	**0,158**		**0,314**		**0,237**	
Variance per factor	0,27	0,15	0,12	0,12	0,12	0,05
Variance per model	0,27	0,28		0,29		

Note: The EFA solutions were obtained using promax oblique rotation and principal factor solution as the factoring method from shrinkage-optimized covariance matrix. Items are ordered by decreasing loadings on the first factor in the two-factor solution and, subsequently, on the second factor to illustrate the confounding effect of the pro-trait/con-trait differential. Loadings below |.200| were suppressed, except when representing the highest loading for a given item (e.g., item 8 in tables 3–6, and item 17 in table 6).

3.2. Exploratory Factor Analysis of Raw Data

Altemeyer's factor analyses of RWA produced one- or two-factor models, with the latter dismissed as a statistical artifact caused by wording direction (Altemeyer, 1981, p. 188, 1996, p. 54). Subsequent studies have yielded diverse solutions, supporting one-factor (Duckitt, 1993; Imhoff & Brussino, 2013), two-factor (Chylíkova & Buchtík, 2016; Etchezahar, 2012), and three-factor models (Cárdenas & Parra, 2010; Mavor et al., 2010; Orellana, 2018; Passini, 2008; Zakrisson, 2005). Others have identified method factors alongside the substantive ones (Rattazzi et al., 2007; Takano et al., 2021), while the ACT scale, an RWA derivative, has been successfully tested for four dimensions in Brazil (Vilanova et al., 2020). While cultural variability may account for the diversity of outcomes, measurement bias has not been systematically tested to determine whether it also plays a role in these inconsistencies.

In our data, five and six eigenvalues exceeded 1 in the 1st and 2nd waves, respectively, with scree plots indicating an elbow after two components. This suggested multiple plausible models: (1) a content factor (authoritarianism) with a method factor, (2) a content factor with overshadowed subdomains, or (3) two asymmetrically expressed subdomain factors. To evaluate these possibilities, one-, two-, three-, and four-factor EFAs were conducted. The four-factor model was discarded, as the fourth factor isolated only correlated errors between similarly worded items (tables 3 and 5).

EFA results showed strong wording-based confounding effects, in line with prior studies (e.g., Mavor et al., 2010, p. 30). In the two-factor models, all uniquely loading items on the first factor were con-trait, while most of the uniquely loading items on the second factor were pro-trait (Tables 3 and 5). This pattern contradicts our theoretical expectation 1 that factor structure should not mirror the pro/con divide. Furthermore, the equal and stable variance split (16 % and 15 %) between the two factors contradicts the expected hierarchy between authoritarianism and its subdomains: if one factor represents authoritarianism and the other its strongest subdomain, the former should explain more variance. If both factors are subdomains, with one merging aggression and submission and the other reflecting conventionalism (e.g., Etchezahar, 2012; Rattazzi et al., 2007; Takano et al., 2021), the aggression-submission factor should not load uniquely on pro-trait items, since submission typically spans across pro/con items (Mavor et al., 2010, p. 29; Rattazzi et al., 2007, p. 1230). Furthermore, the absorption of submission by aggression contradicts our theoretical expectation 5.

Another interpretation of a two-factor model as representing pro- and anti-authoritarian attitudes (Güldü, 2011, pp. 40–41) is also problematic. A bipolar construct should yield strongly negative correlations between its opposite halves, yet such factors correlated weakly (r = -.13, p = .01 in Chylíkova & Buchtík, 2016, p. 19), suggesting that the correlation reflects association between response biases inherent in pro-trait and con-trait items rather than substantive variance, which inevitably cancels out when con-trait items are not reverse-coded. Emergence of models with a positive and a negative factors is an expected outcome in balanced scales where indicators that operationalize a bipolar construct are confounded by a response style factor (Billiet & McClendon, 2000, pp. 609–610; Cambré et al., 2002).

Table 7

Confirmatory Factor Analysis of 1-, 2-, and 3-Factor EFA-Derived Models on Raw Data

Model	χ^2	p-value	df	RMSEA	SRMR	CFI	TLI	AICc	SABIC
1st wave:									
1-factor	403,661	0,000	170	0,118	0,114	0,643	0,601	7598,48	7519,41
2-factor	270,309	0,000	165	0,080	0,089	0,839	0,815	**7496,69**	7393,25
3-factor	217,668	0,001	157	**0,062**	**0,070**	**0,907**	**0,888**	7509,13	**7352,10**
2nd wave:									
1-factor	480,877	0,000	170	0,136	0,114	0,591	0,543	7678,35	7599,28
2-factor	354,266	0,000	166	0,107	0,088	0,752	0,716	**7576,52**	7478,42
3-factor	330,351	0,000	159	**0,104**	**0,085**	**0,774**	**0,730**	7606,13	**7464,56**

Note: MLR estimator was used. Two- and three-factor models included a second-order general factor. All unique loadings and cross-loadings above .200 from Tables 3 and 5 were included. Exact fit is indicated by p-value > .05, good fit is indicated by RMSEA < .06, SRMR < .08, TLI and CFI ≥ .95; acceptable fit is indicated by RMSEA < .08, SRMR < .09, TLI and CFI ≥ .90. The best-performing index values for each wave are highlighted in **bold**.

Three-factor models appeared superficially interpretable, with one factor reflecting conventionalism (mostly con-trait), another aggression (mostly pro-trait), and the third submission (mixed) (e.g., Mavor et al., 2010, p. 31). However, nearly perfect pro/con separation in the first two factors contradicts the semantic entanglement of aggression and conventionalism across items. Additionally, the third factor was unstable, with only one uniquely loading item across both waves, which had already been flagged for elimination due to poor performance (item 8).

CFA of these EFA-derived models found that none achieved an exact or good fit (table 7). The findings indicate that substantive factors alone cannot explain the observed structure, thus pointing to perceptible presence of method bias.

3.3. Detection and Partialling Out of Method Bias

Method bias refers to a variety of influences unrelated to substantive constructs that systematically affects responses and stem from data collection techniques. A major source is satisficing, where respondents provide responses without fully engaging with cognitively demanding questions (Krosnick, 1991). It occurs due to question difficulty, low motivation, or low ability (Podsakoff et al., 2012, pp. 559–561). As satisficing emerges from respondent-questionnaire interaction, its extent depends on measurement characteristics (Podsakoff et al., 2012, pp. 546–547).

Criticism of method bias contamination in fully unbalanced scales (where all items are worded in the same direction), contributed to the decline of F-scale and the rise of RWA, which introduced a balanced mix of items to mitigate common method bias (CMB), particularly acquiescent response style (ARS). The assumption was that ARS in pro-trait and con-trait items would counterbalance each other, yielding a purer measure of the substantive factor. However, while balanced scales reduce some measurement artifacts, they are not entirely immune (Podsakoff et al., 2012, p. 552). Unlike unbalanced scales, though, they contain built-in mechanisms to detect and control bias (Ray, 1979, 1983).

Weijters et al. (2013) identified three components of CMB in balanced scales: ARS, careless responding, and confirmation bias. ARS reflects a preference for one side of the response scale, careless responding ignores con-trait items, while confirmation bias occurs when the direction of the first item wording determines direction of ARS in subsequent responses.

Billiet & McClendon (2000, pp. 611, 621) established criteria for confirming that a variable measures ARS:

A. It must act as a common factor in semantically balanced items,
B. Appear in multiple constructs,
C. Remain stable over time,
D. Correlate with the sum of agreements across all items, and
E. Correlate negatively with education and positively with old age.

Chylíkova & Buchtík (2016), building on Cambré et al. (2002), added:

F. ARS should have smaller factor loadings than the content factor.

We add two further criteria, capitalizing, in particular, on our use of longitudinal data:

G. ARS should correlate oppositely with pro-trait and con-trait items.
H. ARS should be less stable over time than the content factor, as it is more circumstantially influenced.

These criteria will guide our validation of ARS and CMB measures. Evidence suggests that CMB accounts for up to 32 % of scale variance (Podsakoff et al., 2012, p. 543), yet its impact on RWA remains largely unexamined, apart from Mavor et al.'s (2010, p. 32) attempt to extract ARS, which found its effect negligible.

A useful way to detect the presence of method bias in balanced psychometric scales is split-half correlation distribution analysis (SHCD). It involves repeatedly dividing the full set of items into two equal half-scales, computing correlations between summated scores of each subset and its complementary half, and analyzing the distribution of these correlations across a large number of randomized permutations.

Under minimal method bias, the correlations should form a symmetric, approximately normal distribution, as each randomly drawn half-scale consistently measures the same construct. However, systematic method effects distort this pattern, creating skewed, bimodal, or kurtotic distributions. Comparing the observed correlation distribution with a bias-corrected version allows for empirical assessment of bias magnitude and impact.

Fig. 1 (top-left and top-right) presents the SHCDs for wave 1 and wave 2 raw data, respectively, based on 10,000 sampled splits out of the 184,756 possible permutations. In both waves, the purely pro-trait vs. con-trait split-half correlation (r_1 = .49, r_2 = .53) is an extreme outlier, falling beyond three standard deviations from the mean (M_1 = .77, M_2 = .79). The modal correlations (Mo_1 = .7959, Mo_2 = .8112) are overwhelmingly produced by balanced (5-pro, 5-con) or near-balanced (6-4, 4-6) splits. Despite accounting for only 82% of all possible splits, balanced permutations almost exclusively dominate the peak of the distribution, while moderately (3-7, 7-3) or highly unbalanced (2-8, 1-9, 0-10) splits are underrepresented.

The low correlation between purely pro/con half-scales and the overrepresentation of balanced splits point to a systematic method effect related to item wording. The empirical correlation density curve exceeds the reference normal curve around r = .60 (or r = .65 in the 2nd wave) on the x-axis but drops below expected values near r = .88, meaning fewer strong correlations and an excess of weaker correlations. As a result, the mean is pulled leftward relative to the mode, leaving balanced splits artificially overrepresented at the peak of the distribution. The recurrence of this pattern across waves confirms its stability, indicating that method bias suppresses correlations for certain unbalanced half-scales, thereby distorting the observed factor structure.

SHCD analysis clearly confirms that correlational structure of raw data is contaminated with bias associated with the direction of wording, necessitating its removal and further analysis of bias-corrected dataset.

3.3.1. Validating Acquiescent Response Style (ARS) Measure

By design, any balanced scale can be repurposed to measure acquiescence (Ray, 1983, p. 85). Since ARS reflects a systematic tendency to agree or disagree regardless of content, a balanced scale's equal number of pro-trait and non-reverse-coded con-trait items cancels out substantive content. What remains is the tendency to over-agree or over-disagree, which is the operationalization of ARS (Hofstee et al., 1998, pp. 899–900).

Mavor et al. (2010) applied this ARS extraction technique to RWA using the social dominance orientation (SDO) scale, another balanced scale, to compute an ARS score. However, no significant acquiescence effects emerged in regression coefficients or EFA factor loadings, which remained nearly identical to uncorrected data (largest difference: .007), making the bias-extraction procedure seemingly unnecessary.

By deriving ARS score from another scale, Mavor et al. implicitly treated acquiescence as a stable personality trait across scales – a frequently challenged notion. Ray (1983) has demonstrated early on that ARS includes both a stable component and a scale-specific effect. To assess this, we calculated ARS separately on RWA20 and OAS18 to compare their effectiveness in decontaminating RWA20 from acquiescence.

To compute the ARS score for each scale, all items were z-normalized, and the mean was taken across unreversed con-trait and regular pro-trait items. Each pro-trait and reverse-coded con-trait item was then regressed on the ARS score, and residuals were retained as ARS-corrected item scores. The final ARS-corrected scale score was calculated as the mean of all ARS-corrected items.

In wave 1, the ARS score from RWA20 significantly correlated ($p \leq .05$) with 14 of 20 items, while in wave 2, it correlated with 15 items, meeting criterion *A* of Billiet & McClendon's test. Although its correlations with ARS from OAS18 were weak (r_1 = -.22, r_2 = -.19), they were significant or marginally significant (p_1 = .026, p_2 = .054), aligning with criterion *B*. This pattern suggests ARS variance contains more salient scale-specific and less pronounced person-specific components, supporting Ray's (1983) argument on the presence of both. The negative correlation between ARS scores from RWA20 and OAS18 is consistent with Weijters et al.' (2013) definition of confirmation bias (since the two scales start with oppositely worded items). The ARS score was stable across waves (r = .59, $p \leq .001$), satisfying criterion *C*.

ARS also correlated strongly with the sum-of-agreements variable (r_1 = .72, r_2 = .77, $p < .001$), meeting criterion *D*. A significant education effect on ARS (criterion *E*) was found, with lower ARS in participants with at least six years of university ($n = 33$) compared to those with 1–2 years ($n = 18$). Welch t-tests confirmed this pattern in both waves ($p = .004$, $p = .038$), indicating that university education reduces ARS over time.

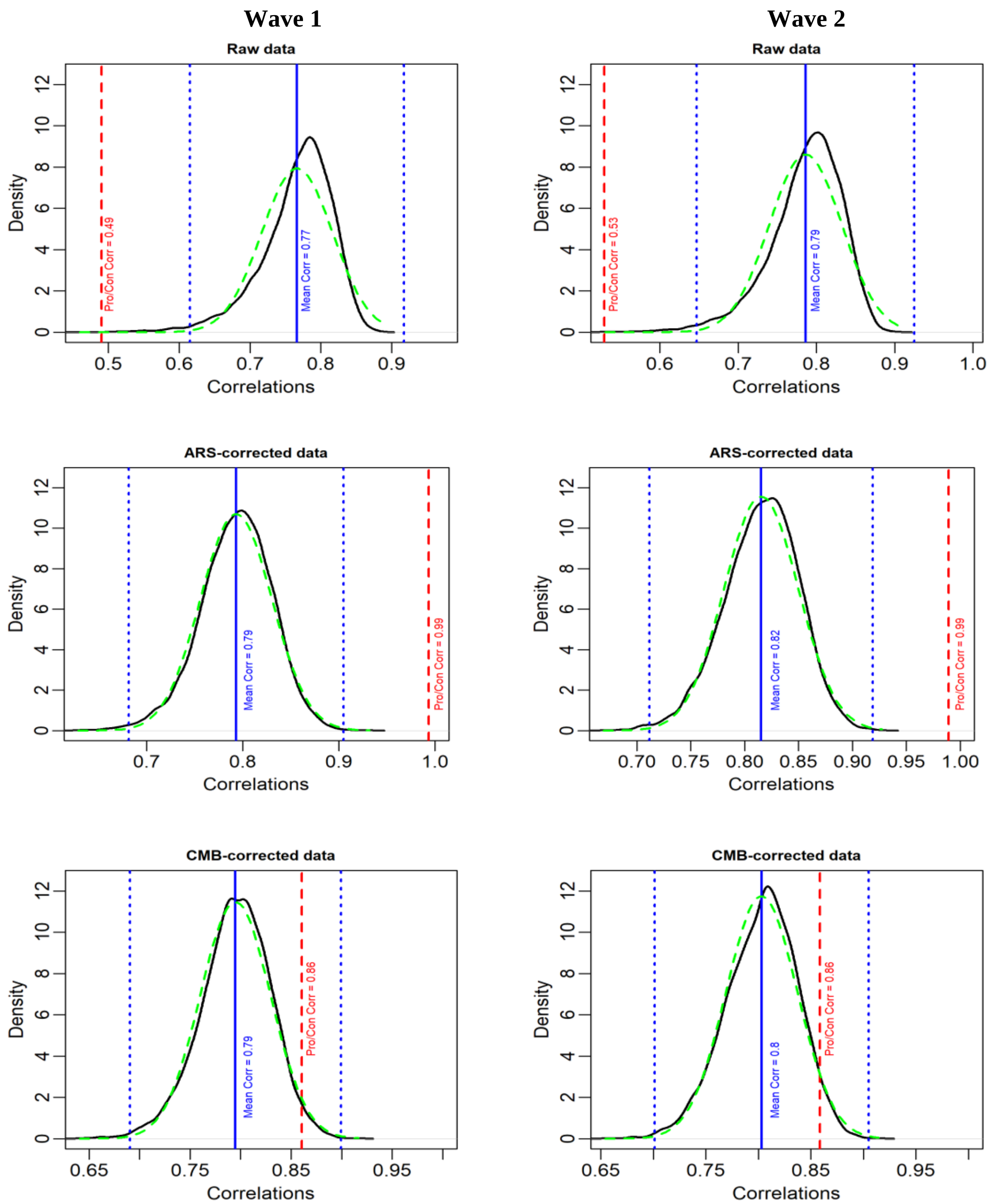

Fig. 1. *Distribution of Correlations Between 10,000 Randomly Selected Permutations of Half-Scales (Subsets of Items, Each Containing Half of the Total) and Their Complementary Scales. Distributions are Presented for First-Wave Raw (Top Left), ARS-Corrected (Middle Left), and CMB-Corrected (Bottom Left) Data, as Well as Second-Wave Raw (Top Right), ARS-Corrected (Middle Right), and CMB-Corrected (Bottom Right) Data (N = 99). Dotted Blue Lines Indicate Three Standard Deviations from the Mean, While the Dashed Red Line Represents the Correlation Between Pro-Trait and Con-Trait Half-Scales. A Reference Normal Curve is Plotted as a Dashed Green Line.*

Note the normalizing effect of the bias-correction procedures on the distributions. In the first wave, skewness decreased from -1,03 (raw data) to -0,33 (both ARS- and CMB-corrected data), and excess kurtosis declined from 1,82 to 0,42 (ARS-corrected) and 0,38 (CMB-corrected). In the second wave, skewness dropped from -0,93 to -0,30 (both ARS- and CMB-corrected data), and excess kurtosis diminished from 1,51 to 0,15 (ARS-corrected) and 0,07 (CMB-corrected data). Despite the ARS correction shifting the pro-trait and con-trait half-scale correlation from the left-side to the right-side tail of the distribution, it remained an extreme outlier, exceeding three standard deviations. Only the CMB correction successfully resolved this issue, bringing the pro/con correlation within acceptable bounds.

ARS score's mean correlation with items in both waves (r_1 = .30, r_2 = .29) was much weaker than that of the ARS-corrected scale score (r_1 = .53, r_2 = .55), meeting the criterion *F*. It also correlated positively with all pro-trait items and negatively with all con-trait items (criterion *G*). Finally, it showed lower test-retest stability (criterion *H*) than the scale score: ICC for the scale score was .68, while for ARS, it was lower at .59, although their 95% confidence intervals overlapped (CI_1: .57 – .78, CI_2: .44 – .71). Therefore, the ARS score, obtained from RWA20, is a fully validated measure of acquiescence for this scale.

The ARS score, derived from OAS18, has failed the test, however. In wave 1, it correlated significantly with only two items (item 3: r = -.25, p = .01; item 14: r = -.28, p = .00), and in wave 2, with just six items (3, 5, 6, 7, 9, 22, r = -.25 to .27, p ≤ .05), failing the baseline criterion *A*. Its weak correlation with the sum-of-agreements variable (r_1 = -.23, p_1 = .02; r_2 = .16, p_2 = .10) further suggests that it does not reflect ARS in RWA20, casting doubt on the cross-scale validity of acquiescence measures.

To confirm its invalidity, we applied the same bias-correction procedure using OAS18-derived ARS score and conducted EFA on the bias-corrected RWA data. Factor loadings in the 3-factor solution remained nearly unchanged from raw data (largest difference: .031 in wave 1, .077 in wave 2), consistent with Mavor et al. Thus, ARS extracted from another scale does not reliably measure ARS in RWA20, undermining Mavor et al.' (2010) claim that ARS has a trivial effect and underscoring the need for reexamining ARS presence in RWA.

3.3.2. Validating Common Method Bias (CMB) Measure

CMB encompasses all systematic variance unrelated to the substantive construct, including acquiescent response style (ARS). To assess CMB in our data, we applied the unmeasured latent method factor technique (Podsakoff et al., 2012, p. 553), using a bifactor CFA model where each item loaded on both a substantive construct and a method factor (e.g., Billiet & McClendon, 2000; Savalei & Falk, 2014; Weijters et al., 2013). Unlike constrained models, our approach allowed method factor loadings to vary freely, acknowledging previously established variability in ARS correlations with items (r_1: .08 – .58, r_2: .05 – .48). The method and substantive factors were set to orthogonality based on weak, non-significant correlations between ARS and ARS-corrected scale scores (r_1 = -.12, p_1 = .24; r_2 = .02, p_2 = .87). To ensure model identification, both general factors' variances were fixed at 1 (Fig. 2).

Expectedly, con-trait items loaded negatively or near-zero on CMB factor, while pro-trait items exhibited stronger positive loadings, supporting prior findings on their higher contamination with method variance. The extracted method factor met validation criteria: it significantly correlated with 12 items in the 1st wave and 15 items in the 2nd, thus meeting criterion *A*; was weakly but consistently linked to ARS score in OAS18 (r_1 = -.25, p_1 = .01; r_2 = -.23, p_2 = .02, criterion *B*), and remained stable across waves (r = .67, p = .00, criterion *C*). It correlated strongly with sum-of-agreements (r_1 = .66, p_1 =.00; r_2 = .65, p_2 =.00, criterion *D*).

CMB scores were negatively associated with education (criterion *E*), with lower scores in participants with more years of university education. Welch *t*-tests confirmed this effect in both waves (p < .001, p = .027), indicating that higher education reduces CMB. Consistent with criterion *F*, CMB factor loadings were lower than those of the content factor (Fig. 2). All negative loadings were confined to con-trait items (criterion *G*). CMB also exhibited lower stability than the content factor (criterion *H*): $ICC_{content}$ = .72, ICC_{CMB} = .67, both p < .001 (95 % *CI*: .61–.81 and .55–.77, respectively).

A strong correlation between ARS and CMB scores (r_1 = .88, p_1 < .00; r_2 = .86, p_2 < .00) suggests they capture the same construct, validating both. However, while ARS captures variance tied to acquiescence and confirmation bias, CMB encompasses broader common variance, including careless responding and other sources of bias. This suggests that CMB extracts a larger share of variance from items, making it a more comprehensive measure of method bias.

3.3.3. Selection of Bias-Correcting Technique

Both bias measures accounted for substantial shares of variance in certain items (Table 8), sometimes surpassing content variance, particularly in items 3 and 14 across both waves, where bias explained 21–57 % of total variance. These two similarly worded items, along with item 7, refer to strong leadership, making them cognitively demanding for young respondents navigating the political complexities of a society torn between the fresh anti-authoritarian legacy of the Revolution of Dignity (2014) and the need to rally around a leader to fight off the Russian invasion since 2022. Conflicting attitudes – aversion toward authoritarian Fig.s and appreciation for President Zelenskyi – likely intensified response difficulties, leading to satisficing behavior.

Table 8

Decomposition of Item Variance Into Content, Method Bias, and Residual Components

Items	**Shares of Content, Acquiescent Response Style (ARS), and Residuals (in %) by Survey Waves**						**Share of Content, Common Method Bias (CMB), and Residuals (in %) by Survey Waves**					
	Content-related variance		**ARS-Related Variance**		**Residual Variance**		**Content-Related Variance**		**CMB-Related Variance**		**Residual Variance**	
	Wave 1	**Wave 2**	**Wave 1**	**Wave 2**	**Wave 1**	**Wave 2**	**Wave 1**	**Wave 2**	**Wave 1**	**Wave 2**	**Wave 1**	**Wave 2**
rwa03	***0,22***	*0,30*	***0,32***	*0,23*	*0,46*	*0,47*	***0,16***	***0,18***	***0,49***	***0,39***	*0,35*	*0,43*
rwa04r	0,31	0,53	0,10	0,09	0,59	0,38	0,36	0,55	0,14	0	0,50	0,45
rwa05	***0,11***	*0,23*	***0,36***	*0,07*	*0,53*	*0,70*	***0,09***	*0,19*	***0,36***	*0,08*	*0,56*	*0,73*
rwa06r	0,34	0,22	0,02	0	0,64	0,78	0,33	0,21	0	0,01	0,66	0,78
rwa07	*0,31*	*0,41*	*0,24*	*0,15*	*0,46*	*0,44*	*0,29*	*0,29*	*0,21*	*0,29*	*0,50*	*0,42*
rwa08r	*0,12*	*0,09*	*0,01*	*0*	***0,86***	***0,91***	*0,09*	*0,05*	*0,01*	*0,04*	***0,90***	***0,90***
rwa09r	0,26	0,41	0,04	0,01	0,70	0,58	0,23	0,35	0	0,04	0,77	0,61
rwa10	*0,48*	*0,48*	*0,16*	*0,15*	*0,36*	*0,37*	*0,47*	***0,31***	*0,15*	***0,41***	*0,39*	*0,28*
rwa11r	*0,13*	*0,36*	*0,12*	*0,17*	*0,75*	*0,47*	*0,16*	*0,43*	*0,12*	*0,06*	*0,73*	*0,51*
rwa12	0,34	0,36	0,03	0,02	0,63	0,62	0,39	0,38	0	0,01	0,61	0,61
rwa13r	0,48	0,42	0,08	0,19	0,44	0,39	0,48	0,55	0,03	0,13	0,49	0,32
rwa14	***0,23***	*0,25*	***0,34***	*0,21*	*0,43*	*0,54*	***0,16***	***0,12***	***0,57***	***0,45***	*0,27*	*0,42*
rwa15r	0,19	0,21	0,01	0,15	0,80	0,64	0,17	0,25	0,01	0,05	0,83	0,70
rwa16	0,41	0,42	0,04	0,01	0,56	0,57	0,46	0,44	0	0,01	0,54	0,55
rwa17	*0,29*	***0,12***	*0,10*	***0,17***	*0,61*	*0,71*	*0,27*	***0,09***	*0,12*	***0,14***	*0,62*	*0,78*
rwa18r	0,31	0,23	0,05	0,14	0,64	0,63	0,38	0,33	0,11	0,13	0,52	0,54
rwa19	*0,15*	*0,23*	*0,14*	*0,13*	*0,72*	*0,65*	*0,14*	*0,20*	*0,13*	*0,08*	*0,73*	*0,72*
rwa20r	0,26	0,25	0,07	0,12	0,66	0,62	0,28	0,38	0,06	0,15	0,66	0,47
rwa21r	0,51	0,45	0,06	0,04	0,43	0,51	0,56	0,49	0,07	0,01	0,37	0,50
rwa22	0,37	0,31	0,01	0,10	0,62	0,60	0,41	0,25	0	0,12	0,59	0,63

Note: Items labeled with "r" indicate con-trait (reverse-scored) items. To facilitate comparison, columns referring to wave 1 are shaded. Items identified as worst-performing – defined as those where method variance exceeds content variance or residual variance exceeds 85 % – are marked in bold.

Apart from these heavily biased items, overall scale performance remained strong, with few items showing method-related variance exceeding 10 % across both waves. The impact of bias removal was minimal at the aggregate level: correlations between raw and bias-corrected scale scores were high (CMB-corrected: $r_1 = .96$, $r_2 = .95$; ARS-corrected: $r_1 = .98$, $r_2 = .99$, both $p < .001$). ARS explained only 3 % and 1 % of total scale score variance in waves 1 and 2, while CMB accounted for 6 % and 9 %. This suggests that, for practical purposes where RWA20 is treated as a unidimensional construct, simple summation of raw item scores (with con-trait items reverse-coded) remains a fully viable and justified approach.

However, structural relationships within RWA20 were more sensitive to method bias. SHDCs (fig. 1) show that both techniques improved data structure, as seen in the normalization of the empirical curve. Yet, ARS correction artificially inflated the pro/con correlation, thus introducing its own artifact. In contrast, CMB correction preserved scale structure while normalizing skewness and excess kurtosis, thereby producing a more reliable dataset.

Removal of CMB improved test-retest reliability at the aggregate level ($ICC = .72$, $p < .001$; table 2), though individual item reliability decreased slightly (mean $ICC_{raw} = .47$, mean $ICC_{corrected} = .44$), as expected given that CMB inflated item correlations across waves. Internal consistency improved within waves, with Cronbach's α and mean inter-item correlation increasing in CMB-corrected data (table 2). Although both data-correction techniques were validated, partialling out CMB was preferred as a more comprehensive and less intrusive method.

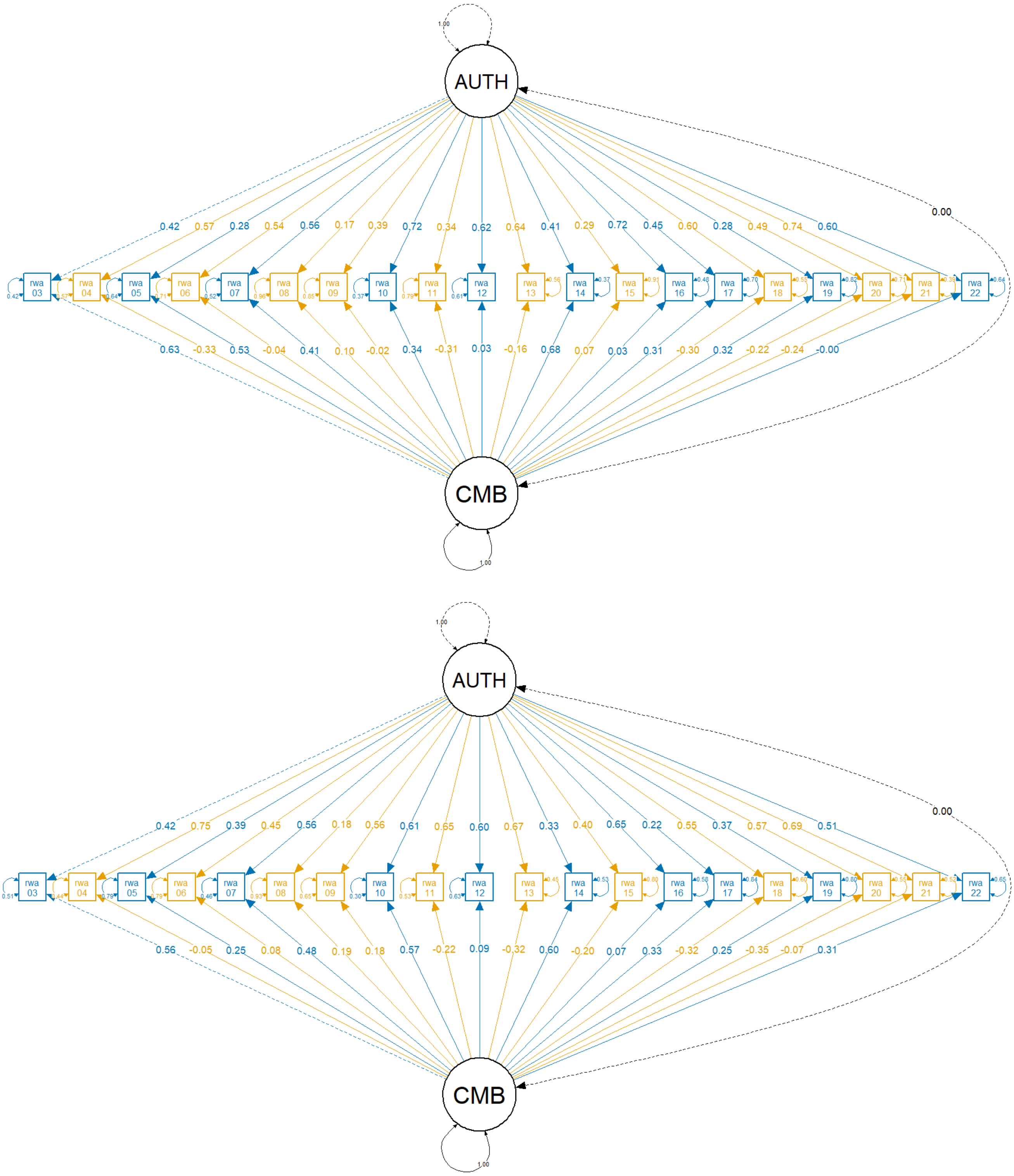

Fig. 2. *Bifactor CFA Model with a Single Content Factor (AUTH) and a Method Factor (CMB) for Wave 1 (top) and Wave 2 (bottom), Used to Extract CMB Scores. Pro-Trait Items and Their Loadings are Shown in Blue, Con-Trait Items in Orange*

3.4. EFA and CFA of CMB-Corrected Data

Despite barely detectable aggregate-level effects, CMB correction has markedly altered the correlational structure. Two- and three-factor EFA solutions no longer reflected item-wording effects, revealing a different factor structure (tables 4 and 6). All multidimensional solutions exhibited unstable factor composition, with items shifting across factors between waves.

All three solutions underwent CFA (table 9), which strongly supported the scale's three-dimensional structure. The model with three first-order factors and a general second-order factor achieved exact fit in both waves (p > .05), with all fit indices meeting good-fit criteria, except for CFI_2 = .945 and TLI_2 = .934, which remained within the acceptable range. Given its statistical fit, theoretical justification, and parsimony, the three-factor hierarchical structure emerges as the most empirically justified and interpretable representation of CMB-corrected RWA20.

Table 9

Confirmatory Factor Analysis of 1-, 2-, and 3-Factor EFA-Derived Models on CMB-Corrected Data

Model	χ^2	p-value	df	RMSEA	SRMR	CFI	TLI	AICc	SABIC
1st wave:									
1-factor	209,648	0,021	170	0,049	0,071	0,908	0,897	**7389,95**	7310,88
2-factor	187,256	0,113	165	0,037	0,066	0,948	0,940	7399,12	7295,68
3-factor	164,576	0,303	156	**0,024**	**0,062**	**0,980**	**0,976**	7451,33	**7285,93**
2nd wave:									
1-factor	237,114	0,001	170	0,063	0,074	0,859	0,843	**7530,65**	7451,58
2-factor	206,850	0,013	164	0,051	0,067	0,910	0,896	7538,99	7429,94
3-factor	185,417	0,074	159	**0,041**	**0,064**	**0,945**	**0,934**	7557,25	**7415,69**

Note: ML estimator was applied to shrinkage-optimized covariance matrix. Two- and three-factor models included a second-order general factor. All unique loadings and cross-loadings above .200 from Tables 4 and 6 were retained. The best-performing index values for each wave are highlighted in **bold**.

3.5. Reintegrating Method Effects: Testing Bifactor Content-Method Models on Raw Data

To assess whether models derived from CMB-corrected data generalize to raw data, the bias-correction process was reversed by integrating the CMB extraction model (fig. 2) with the wave-specific content structures (tables 4, 6, and 9), resulting in bifactor models that included a method factor and a content factor. In these models, the method factor loaded on all items, while content factor structure varied: in one-factor model, it loaded directly on items (fig. 2); in two- and three-factor models, it loaded onto subdomains, which then loaded onto items. Content and method factors were constrained to be orthogonal.

The 3-factor + CMB model demonstrated the best fit in wave 1, outperforming simpler models (table 10). In wave 2, no model fully met Hu & Bentler's (1999) strict criteria, but the 3-factor + CMB model again exhibited the strongest fit despite a suboptimal RMSEA. Given our study conditions (N = 99, df > 140), RMSEA up to .10 and CFI/TLI above .80–.85 can still indicate an acceptable approximation of the data structure (Kenny et al., 2015; Niemand & Mai, 2018; Shi et al., 2021).

Table 10

Fit of Models Derived from CMB-Corrected Data to Raw Data with a Method Factor

Model	χ^2	df	p-value	RMSEA	SRMR	CFI	TLI	AICc	SABIC
1	2	3	4	5	6	7	8	9	10
1st wave:									
1-factor + CMB	244,406	151	0,000	0,079	**0,070**	0,857	0,820	**7602,21**	7387,46
2-factor + CMB	223,289	147	0,000	0,072	0,083	0,883	0,849	7637,96	7372,09
2-factor + CMB + correlated errors	195,419	145	0,003	0,059	0,081	0,923	0,899	7643,69	7347,10
3-factor + CMB	188,785	140	0,004	**0,059**	0,071	**0,925**	**0,899**	7742,05	**7347,65**
3-factor + CMB + correlated errors	160,117	138	0,096	0,040	0,068	0,966	0,953	7766,69	7321,86
2nd wave:									
1-factor + CMB	324,966	151	0,000	0,108	0,074	0,771	0,712	**7685,42**	7470,68
2-factor + CMB	284,028	146	0,000	0,098	**0,070**	0,818	0,763	7717,66	7436,92

The End of the Table 10

1	2	3	4	5	6	7	8	9	10
2-factor + CMB + correlated errors	261,102	144	0,000	0,091	0,067	0,846	0,797	7730,40	7416,87
3-factor + CMB	265,633	145	0,000	**0,092**	0,079	**0,841**	**0,792**	7716,56	**7419,97**
3-factor + CMB + correlated errors	246,437	143	0,000	0,085	0,077	0,864	0,819	7735,30	7403,64

Note: MLR estimator was used. Two- and three-factor models included a second-order general factor. All unique loadings and cross-loadings above .200 from Tables 4 and 6 were retained. Correlated errors included similarly worded items: rwa11~~rwa20 and rwa03~~rwa14 in the 1st wave, rwa12~~rwa16, rwa03~~rwa14 in the 2nd wave. The best-performing index values for each wave are highlighted in **bold**, excluding models with correlated errors.

Adding correlated errors improved fit in both waves, making the 3-factor + CMB + correlated errors model the best-fitting solution in wave 1, where it met all goodness-of-fit thresholds. In wave 2, while no model fully met cutoff criteria, this model again demonstrated the strongest fit across most indices, achieving marginally acceptable fit by lenient criteria.

Overall, the findings support a three-factor structure as the most empirically justified representation of the scale across waves, with wave 1 providing clearer support and wave 2 remaining somewhat less conclusive due to fit constraints. Importantly, all bifactor models incorporating a method factor (table 10) fit raw data much better than their simple counterparts (table 7) by all indices, confirming method bias's role in obscuring factor structure, despite its unnoticeable presence at the aggregate level.

In both waves, 3-factor models (fig. 3) explained 30–70 % of item variance, except for item 8 in both waves and item 17 in wave 2, which poorly fit the scale. To improve interpretability in the face of model instability, the models required simplification by removing redundant, outlying, and unstable items.

3.6. Abridged Three-Dimensional Version of RWA (RWA13)

The abridgement aimed to create a shorter version of RWA20 that preserves its three-factor structure while also holding longitudinal MI.

The procedure eliminated items falling out of scale (items 8 and 17), as well as swing items 6 and 22, which had the lowest test-retest reliability (Appendix A). Items loading solely on the same single factor in both waves were retained as core indicators (items 3, 5, 11, 15, and 18) (Fig. 3). Redundant items with similar wording were removed (item 14 as duplicating item 3, item 20 as duplicating item 11). Item 12, cross-loading on two factors in both waves, was retained with both loadings. Items 7, 9, 10, 13, 16, 19, and 21 were kept, retaining only the loadings that remained stable across waves, while item 4 was excluded as a swing item.

Table 11

Scale Reliability Statistics for RWA13 Based on Raw, ARS-Corrected, and CMB-Corrected Data

	RAW DATA		ARS-CORRECTED DATA		CMB-CORRECTED DATA	
	Wave 1	**Wave 2**	**Wave 1**	**Wave 2**	**Wave 1**	**Wave 2**
Cronbach's α (standardized)	0,82	0,85	0,84	0,88	0,84	0,87
Guttman's λ_6	0,86	0,89	0,87	0,90	0,87	0,91
McDonald's ω (1-factor model)	0,82	0,85	0,85	0,88	0,85	0,88
Mean interitem correlation	0,25	0,30	0,29	0,35	0,29	0,35
Share of CMB variance in the scale score	6 %	5 %	-	-	-	-
Correlation with RWA20 score	$0,98^{***}$	$0,98^{***}$	$0,98^{***}$	$0,98^{***}$	$0,97^{***}$	$0,98^{***}$
Test-retest reliability (ICC)	$0,72^{***}$		$0,71^{***}$		$0,74^{***}$	

*** p < .001

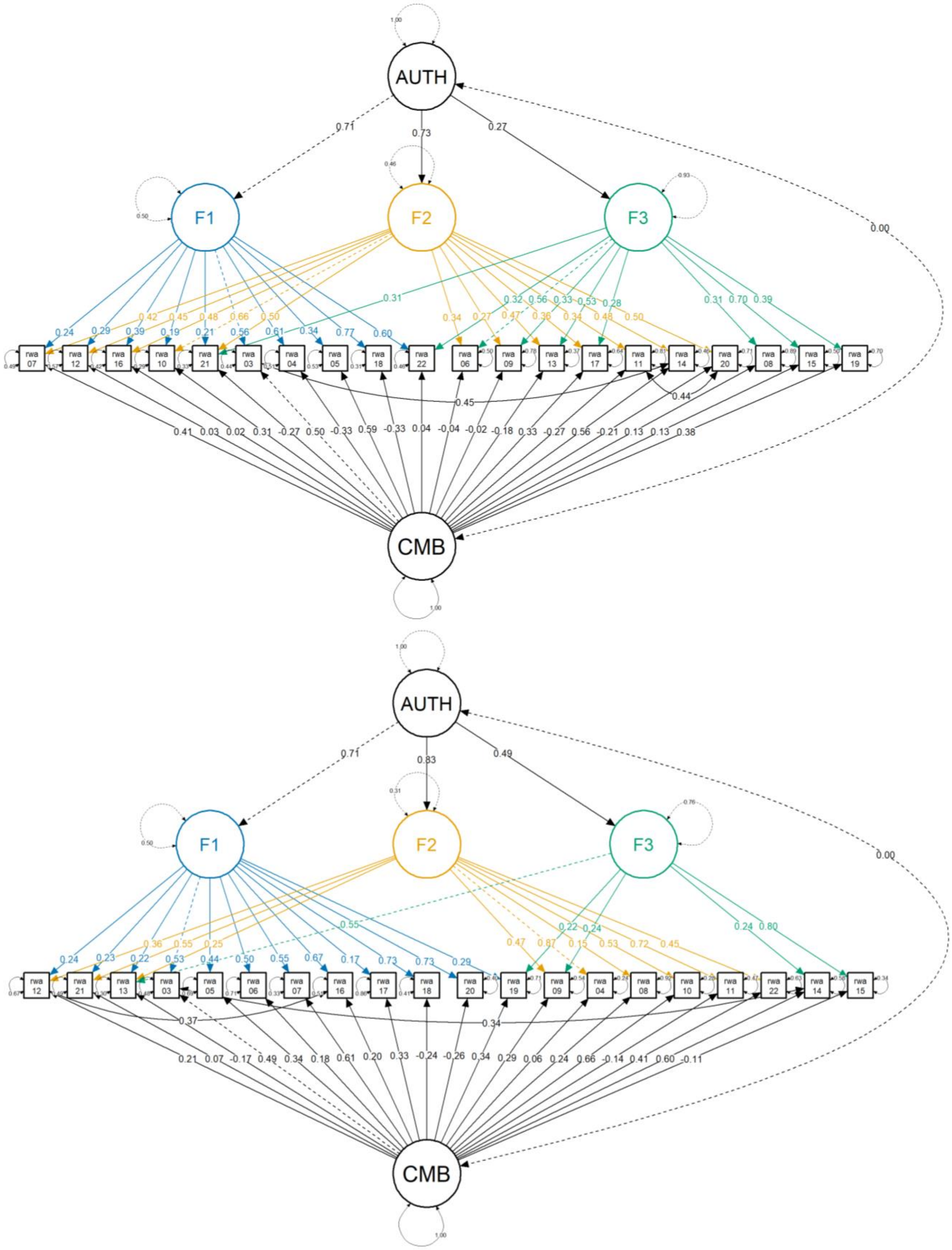

Fig. 3. *Best-Fitting Bifactor Models for Raw RWA20 Data*
Bifactor three-factor model with a method factor (CMB) and a second-order general content factor (AUTH) for wave 1 (top) and wave 2 (bottom). Factor loadings for the three subdomains (F1, F2, and F3) are color-coded for clarity

The resulting abridged scale consisted of thirteen items (hence RWA13), most of which have been used in other abridged versions, attesting to their cross-cultural relevance (see last column in Appendix B). Mean test-retest reliability of items improved (ICC_{raw} = .50, $ICC_{corrected}$ = .47), compared to ICC_{raw} = .47 and $ICC_{corrected}$ = .44 in RWA20. At the aggregate level, RWA13 showed good reliability (Table 11), with test-retest reliability and mean interitem correlation improving, though expected reductions in α, λ_6, and ω-reliabilities occurred due to fewer items (table 2). Despite this, reliability remained high, confirming RWA13's robust psychometric properties.

3.6.1. Modelling Factor Structure of RWA13

One-, two-, and three-factor models of RWA13 were first tested on CMB-corrected data, all achieving exact fit, but only one- and three-factor models crossed fit thresholds in the 1st wave. In the 2nd wave, however, exact fit and all other indices identified the three-factor model as the best representation of the data (table 12).

Table 12

Fit of RWA13 Models to CMB-Corrected Data

Model	χ^2	p-value	df	RMSEA	SRMR	CFI	TLI	AICc	SABIC
1st wave:									
1-factor	68,298	0,366	65	**0,023**	**0,064**	**0,986**	**0,984**	**4802,87**	**4768,74**
2-factor	66,103	0,274	60	0,032	0,118	0,975	0,967	4820,79	4773,73
3-factor	67,652	0,290	62	0,030	0,088	0,976	0,970	4813,95	4772,40
2nd wave:									
1-factor	95,223	0,009	65	0,069	0,072	0,899	0,879	**4802,87**	4858,13
2-factor	89,899	0,007	60	0,071	0,103	0,900	0,870	4907,05	4860,00
3-factor	78,382	0,078	62	**0,052**	**0,082**	**0,945**	**0,931**	4887,14	**4845,60**

Note: ML estimator was applied to shrinkage-optimized covariance matrix. The models were derived from associations of RWA13 items with factors in EFA solutions for CMB-corrected RWA20 (Tables 4 and 6). The best-performing index values for each wave are highlighted in **bold**.

The models were then converted to a bifactor design, with the CMB factor loading on all items. Model identification was achieved by adding the CMB score from RWA20, which was set to correlate with the CMB factor to ensure that CMB captured method variance (Fig. 4). The one-factor model performed well in the 1st wave but lacked stability, failing in the 2nd wave. In contrast, the three-factor model was the most consistent, meeting good or acceptable fit criteria across both waves, except for TLI_2 = .869, which, while suboptimal, was the best among all models and remained within a reasonable range. Moreover, adding two correlated errors improved the three-factor model to good or acceptable fit across all indices, a result not achieved in simpler models (table 13).

Table 13

Fit of 1-, 2-, and 3-Factor Models to RWA13 Raw Data with a Method Factor

Model	χ^2	p-value	df	RMSEA	SRMR	CFI	TLI	AICc	SABIC
1st wave:									
1-factor + CMB	96,733	0,007	65	**0,070**	**0,064**	0,945	0,923	**5032,01**	4952,94
2-factor + CMB	80,339	0,034	59	0,060	0,091	**0,963**	**0,943**	5054,21	**4945,16**
3-factor + CMB	95,226	0,004	62	0,074	0,075	0,942	0,916	5048,75	4955,74
3-factor + CMB + correlated errors	87,524	0,012	60	0,068	0,072	0,952	0,928	5054,36	4950,91
2nd wave:									
1-factor + CMB	141,581	0,000	65	0,109	**0,073**	0,888	0,843	5053,00	4973,93
2-factor + CMB	125,912	0,000	59	0,107	0,082	0,902	0,849	5075,93	4966,88
3-factor + CMB	122,739	0,000	62	**0,099**	0,084	**0,911**	**0,869**	5052,41	4959,40
3-factor + CMB + correlated errors	101,251	0,001	60	0,083	0,082	0,940	0,908	**5044,23**	**4940,79**

Note: MLR estimator was applied. Two- and three-factor models included a second-order general factor. Correlated errors, where indicated, were allowed between items rwa11~~rwa20 and rwa03~~rwa14 in both waves. The method factor (CMB) was specified to correlate with the CMB score for identification purposes (as shown in fig. 4). The best-performing index values for each wave are highlighted in **bold**, excluding models with correlated errors.

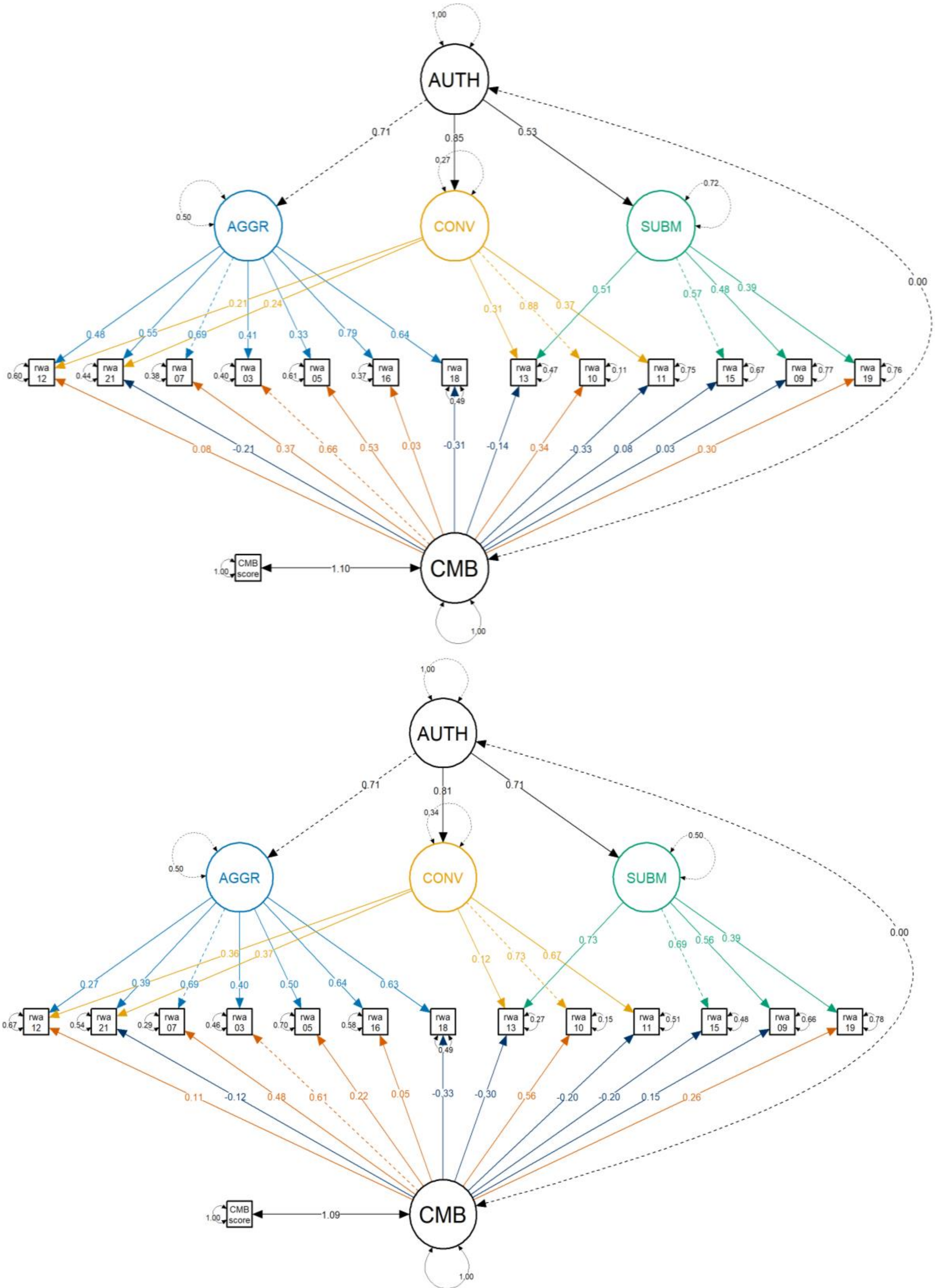

Fig. 4. *Best-Fitting Bifactor Model for Raw RWA13 Data: 3-Factor Hierarchical Model for Wave 1 (top) and Wave 2 (Bottom). Factor loadings for the three subdomains (AGGR, CONV, and SUBM) are color-coded for clarity. The CMB score, derived from the full RWA20 scale, was introduced to aid model identification and ensure that method variance is captured by the CMB factor. The correlation between CMB score and CMB exceeding 1 reflects a scaling artifact rather than a substantive issue.*

3.6.2. Longitudinal MI of RWA13 Bifactor Models

Longitudinal measurement invariance (LMI) was tested for all three bifactor models (without correlated errors) at configural, metric, and scalar levels to identify the most stable structure. In metric and scalar tests, method factor loadings were unconstrained, as the focus was on evaluating content factor invariance.

Table 14

LMI of Bifactor Models with a Common Method Factor for RWA13: Absolute Fit Indices

Model	Configural CFI	Metric CFI	Scalar CFI	Configural RMSEA	Metric RMSEA	Scalar RMSEA	Configural SRMR	Metric SRMR	Scalar SRMR
3-factor	**0,945**	**0,949**	**0,949**	**0,076**	**0,070**	**0,067**	0,073	0,080	0,081
2-factor	0,930	0,935	0,935	0,087	0,078	0,075	0,081	0,093	0,095
1-factor	**0,945**	0,916	0,918	0,092	0,087	0,082	**0,064**	**0,077**	**0,078**
Acceptable fit thresholds	≥ 0,950				≤ 0,08			≤ 0,09	

Note: In metric and scalar invariance tests, loadings on the method factor were unconstrained. The best-performing fit index values for each wave are highlighted in **bold**.

Table 15

LMI of Bifactor Models with a Common Method Factor for RWA13: Comparative Fit Indices

Model	Configural AICc	Metric AICc	Scalar AICc	Configural SABIC	Metric SABIC	Scalar SABIC
3-factor	**10353,85**	**10207,12**	**10143,80**	**10012,55**	**9991,60**	**9981,70**
2-factor	10396,98	10216,40	10148,02	10034,27	**9991,60**	9996,11
1-factor	**10292,44**	**10197,39**	**10124,76**	10034,27	1002,54	10008,49

Note: In metric and scalar invariance tests, loadings on the method factor were unconstrained. The best-performing fit index values for each wave are highlighted in **bold**.

LMI analysis confirms that the 3-factor bifactor model best represents the data across waves (tables 14 and 15), consistently showing the strongest absolute fit, with highest CFI/TLI values and lowest RMSEA at all MI levels. The 1- and 2-factor models failed to meet the CFI ≥ .950 benchmark, with higher RMSEA values, indicating poorer fit. Although the 1-factor model had the lowest SRMR, all models remained within acceptable limits, making this criterion less decisive. Therefore, the 3-factor bifactor model (fig. 4) is the most empirically supported and theoretically coherent solution for assessing the construct longitudinally.

Table 16

Fit Indices for Configural, Metric, and Scalar LMI of the 3-Factor Hierarchical Model of RWA13 with a Common Method Factor

Model	N	Number of Parameters	χ^2 (df)	p-Value	CFI	TLI	RMSEA [90 % CI]	SRMR	AICc	SABIC
Configural	99	118	188,775 (120)	0,000	0,945	0,917	0,076 [0,055, 0,096]	0,073	10353,85	10012,55
Metric (partial)	99	103	199,622 (135)	0,000	0,949	0,931	0,070 [0,048, 0,089]	0,080	10207,12	9991,60
Scalar (partial)	99	94	208,811 (144)	0,000	0,949	0,935	0,067 [0,046, 0,087]	0,081	10143,80	9981,70
Acceptable fit thresholds				> 0,050	≥ 0,950	≥ 0,950	≤ 0,080	≤ 0,090		

Table 17

Changes in Fit Indices Across Configural, Metric, and Scalar LMI Levels for the 3-Factor Hierarchical Model of RWA13 with a Common Method Factor

Model	χ^2 Difference Statistics		ΔCFI	ΔTLI	ΔRMSEA	ΔSRMR	ΔAICc	ΔSABIC	Meets Criteria
	$\Delta\chi^2$ (Δdf)	P-Value							
Configural	—	—	—	—	—	—	—	—	
Metric (partial)	10,8480 (15)	0,76	0,003	0,014	-0,007	0,007	-146,73	-20,95	Yes
Scalar (partial)	9,1889 (9)	0,42	0,000	0,004	-0,002	0,001	-63,32	-9,9	Yes
Threshold values		> 0,050	≤ 0,010	≤ -0,010	≤ 0,015	≤ 0,03 (metric) ≤ 0,01 (scalar)			

Note: Threshold values are based on Chen (2007, p. 501). Partial invariance was established at metric and scalar levels by freeing method factor loadings to allow for cross-wave differences in response styles.

Tables 16 and 17 confirm that the 3-factor model not only outperforms alternatives but also maintains measurement equivalence across waves, indicating that imposition of stricter invariance constraints improved or maintained model fit.

3.7. Interpretation of RWA13 Factors

All RWA13 items can be grouped into five parcels based on their factor associations (table 18): F1 (items, loading on factor 1), F1∩F2 (cross-loading on factors 1 and 2), F2 (loading on factor 2), F2∩F3 (cross-loading on factors 2 and 3), and F3 (loading on factor 3).

Table 18

Association of RWA13 Items with Parcels and Their Etic Semantic Composition

F1	F1∩F2	F2	F2∩F3	F3
rwa03 (ACS)	rwa12 (C)	rwa10 (AC)	**rwa13 (ACS)**	**rwa09 (ACS)**
rwa05 (ACS)	**rwa21 (ACS)**	**rwa11 (AC)**		**rwa15 (ACS)**
rwa07 (ACS)				rwa19 (ACS)
rwa16 (ACS)				
rwa18 (ACS)				

Note: Con-trait items are marked in bold.

Juxtaposing items with their etic semantic structures reveals a clear pattern: F1 and F3 parcels contain ACS items, whereas F2 consists solely of AC items. Cross-loading parcels also follow a pattern: F2∩F3 includes an ACS item, while F1∩F2 mixes C and ACS. By cross-referencing the etic semantic composition of the parcels with the statistical procedure that generated them, the factors' emic identities can be deduced (fig. 5).

Parcels F3 and F2∩F3 uniquely share only the S component of their ACS composition, since the AC component of F2∩F3 is also present in F2. This suggests that F2∩F3 is linked to F2 through shared AC variance, whereas its connection to F3 stems from their common S variance. Thus, in F3, S emerges as the dominant (emically pronounced) component, distinguishing it from F2, which contains only AC.

Statistically, this means that participants responding to F2∩F3 items varied in their emphasis on A, C, or S components. Subsamples focusing on AC contributed to F2∩F3's shared variance with F2, while those emphasizing S account for its shared variance with F3. Although some AC variance may still be present between F3 and F2∩F3, its impact appears marginal, as it is largely absorbed by the F2∩F3–F2 nexus (fig. 5).

This evidence supports the identification of the F3 parcel – and, by extension, the F3 factor – as the submission subdomain in RWA13 (fig. 4).

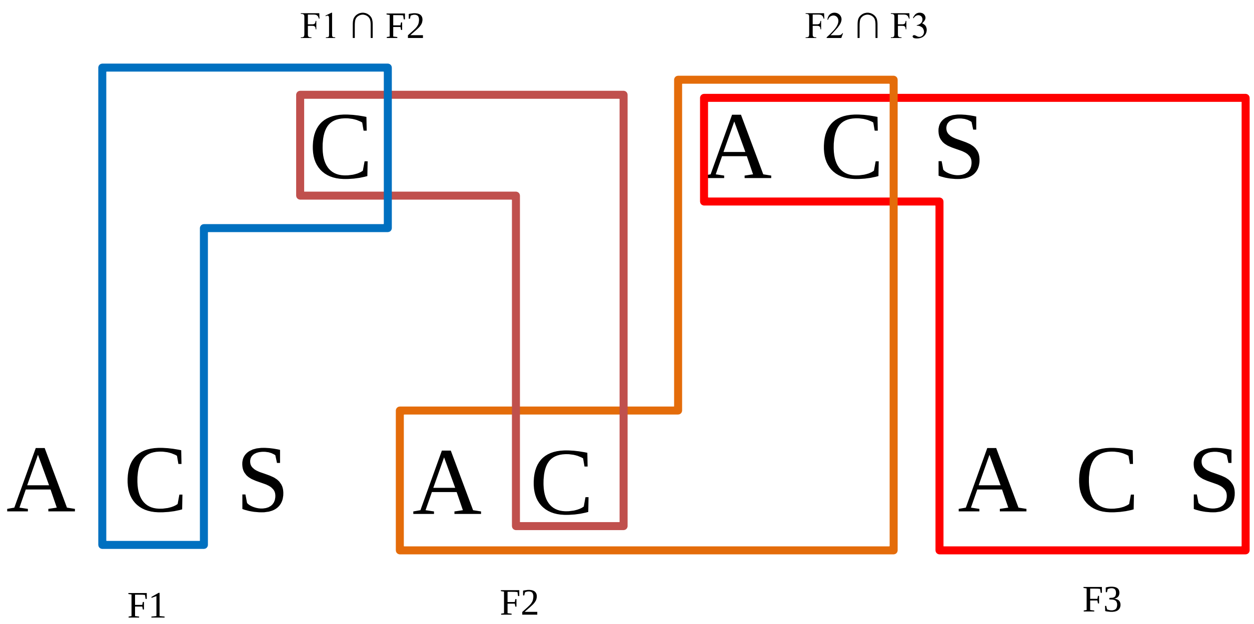

Fig. 5. *Semantic Intersections and Differences of RWA Parcels*
A = Aggression, C = Conventionalism, S = Submission.

Etically, parcel F1 contains all three components (ACS) but shares only C with the F1∩F2 parcel. For simplicity, we reduced F1∩F2 to its C component as the only common element across its two items (table 18). The uniqueness of F1 is therefore tied to A and S components. However, since F1 lacks cross-loadings with F3, which is dominated by S, this suggests that S is dormant in F1, leaving A as its primary unconsumed variance, which distinguishes it from other parcels.

This conclusion is reinforced by factor score correlations: F1 showed a stronger correlation with F2 in both waves (r_1 = .79, r_2 = .77, p < .001) than with F3 (r_1 = .59, r_2 = .66, p < .001), indicating closer statistical and semantic alignment with F2. The evidence highlights the A component as the defining feature of F1, identifying it with the aggression subdomain (fig. 4).

Unlike F1 and F3, the F2 parcel intersects with two others: F1∩F2 and F2∩F3. It connects to other factors solely through the C-containing intersectional parcels, indicating that F2 is predominantly a C parcel. Given that C is the only semantic component not yet assigned to a factor, F2 is best identified with the conventionalism subdomain.

This classification of RWA13 items and factor interpretations largely align with prior literature, reinforcing the validity of our approach (table 1). While identifying three extracted factors with RWA20's subdomains is a theoretical success, it does not imply purity in factor scores. The pervasive etic co-presence of A and C across items limits complete statistical separation of subdomains.

Impurity of subdomain scores is evident in cross-wave correlations: aggression (wave 1) correlates more strongly with conventionalism (wave 2) than with itself, and submission (wave 2) correlates more with aggression (wave 1) than with itself. While remaining correlations follow expected patterns and confidence intervals are not exceeded, these findings caution against treating factors in RWA13, and especially in RWA20, as pure measures of the subdomains.

Nonetheless, our analysis supports a three-factor structure with a second-order general factor and a method factor, both empirically and theoretically. To enhance subdomain validity, introducing pure S and AS items in place of some ACS bundles could improve differentiation between A and C, refining subdomain measurement in RWA.

3.7. Abridged Unidimensional Version of RWA (RWA6)

A six-item abridged RWA20 version (RWA6) was developed for practical applications, maintaining a balanced mix of con-trait and pro-trait items (Appendix C). Items with the lowest loadings were sequentially eliminated from the bifactor unidimensional CFA model until achieving good model fit and a minimum .400 loading on the content factor across all items.

The final RWA6 items demonstrated test-retest reliabilities from .47 to .61 (raw data), item-total and item-rest correlations above .50, and superior aggregate test-retest reliability compared to RWA20 (table 2) and slightly better than RWA13 (table 11). Reliability indicators (α, λ_6, ω) ranged between .81 and .86, slightly lower than RWA20 but comparable to RWA13, while mean interitem correlation exceeded both. High correlations between RWA6 and RWA20 (table 19) confirm RWA6 as a valid, concise measure of the aggregate RWA construct for sociological surveys.

Table 19

Scale Reliability Statistics for RWA6 Based on Raw, ARS-Corrected, and CMB-Corrected Data

	RAW DATA		ARS-CORRECTED DATA		CMB-CORRECTED DATA	
	Wave 1	Wave 2	Wave 1	Wave 2	Wave 1	Wave 2
Cronbach's α (standardized)	0,83	0,82	0,85	0,85	0,85	0,86
Guttman's λ_6	0,82	0,81	0,84	0,84	0,84	0,85
McDonald's ω	0,83	0,82	0,86	0,85	0,85	0,86
Mean interitem correlation	0,44	0,43	0,50	0,49	0,48	0,50
Share of CMB variance in the scale score	0 %	0 %	-	-	-	-
Correlation with RWA20 score	$0{,}86^{***}$	$0{,}90^{***}$	$0{,}89^{***}$	$0{,}92^{***}$	$0{,}93^{***}$	$0{,}93^{***}$
Correlation with RWA13 score	$0{,}86^{***}$	$0{,}91^{***}$	$0{,}90^{***}$	$0{,}93^{***}$	$0{,}92^{***}$	$0{,}93^{***}$
Test-retest reliability (ICC)	$0{,}73^{***}$		$0{,}72^{***}$		$0{,}74^{***}$	

*** $p < .001$

Due to the elimination of bias-contaminated items, RWA6 is virtually free from CMB. Regression-based estimation showed 0 % variance attributable to CMB, compared to 5–9 % in RWA20 and RWA13. Consequently, a one-factor model fits both waves acceptably ($SRMR_1$ = .06, CFI_1 = .91; $SRMR_2$ = .07, CFI_2 = .90), demonstrating configural ($SRMR_c$ = .06, CFI_c = .90), metric ($SRMR_m$ = .07, CFI_m = .92), and scalar ($SRMR_s$ = .07, CFI_s = .92) LMI. $\Delta SRMR_m$ = .007, $\Delta SRMR_s$ = .003 and ΔCFI_m = .017, ΔCFI_s = .000 confirm both metric and scalar invariance.

4. DISCUSSION

Few longitudinal studies on RWA have been published (Asbrock et al., 2010; Liu et al., 2008; Sibley et al., 2007; Sibley & Duckitt, 2013), with none focusing on the scale's validation or measurement invariance over time. This study provides the first psychometric validation and adaptation of the RWA scale for Ukraine, leveraging its longitudinal design to assess reliability, dimensionality, and measurement artifacts. Results confirm the scale's internal consistency over time but expose systematic response biases, necessitating correction to uncover a clearer latent structure.

The procedures developed in this study contribute to the existing literature in several ways. By distinguishing between etic and emic perspectives on the RWA scale, we introduce a structured method for interpreting empirically derived factor structures through comparative analysis of the scale's semantic composition (etic) alongside empirical patterns of overlap and differentiation between item parcels (emic). It offers a more systematic alternative to subjective face-valid classification, prone to confirmation bias. Given the multi-barreled nature of RWA20 items, nearly any item can be interpreted as reflecting any subdomain, enabling unfalsifiable classifications in the absence of a standardized method. By limiting researcher discretion, our approach enhances reproducibility and comparability, even with recognition of cross-cultural variations in RWA's emic structure.

A crucial step in the semantic breakdown of RWA items was redefining aggression as a bipolar construct. Treating it as unipolar hindered meaningful factor interpretations by restricting it to pro-trait items, which resulted in illogical outcomes – such as near-zero correlations between factors composed solely of pro- and con-trait item clusters (Chylíkova & Buchtík, 2016; Güldü, 2011).

Across all levels of analysis, the three-dimensional RWA model provided the best fit, when controlling for method artifacts. The abridged scale (RWA13), composed of emically stable items, demonstrated longitudinal MI at configural, metric, and scalar levels. The subdomains of aggression, conventionalism, and submission were successfully mapped to the factors extracted from bias-corrected data, by systematically matching them to the etic semantic structure of RWA.

While our results are consistent with recent studies supporting a three-dimensional RWA structure, our analysis questions the validity of models derived without accounting for method bias. Both ARS and broader CMB, though minor nuisances in unidimensional balanced scales, markedly distort complex factor structures by inflating correlations among same-direction-of-wording items while deflating those between opposite-worded items. This pattern is evident in many reported three-factor models, where inter-factor correlations were clearly deflated, ranging from .13 to .58 (Cárdenas & Parra, 2010, p. 71; Mavor et al., 2010, p. 30; Orellana, 2018, p. 21; Passini, 2008, p. 55; Zakrisson, 2005, p. 868).

Our raw data factor solutions exhibited similar distortions, with correlations ranging from .39 to .55 (two-factor) and .24 to .48 (three-factor) across waves. However, introducing a CMB factor to bifactor models raised these values to .56 – .73 (two-factor) and .33 – .71 (three-factor). The optimized RWA13, free of malfunctioning items, further increased inter-factor correlations to .58 – .80 across waves – closer to the expected magnitude for a construct defined by subdomain covariation. SHDCs (Fig. 1) further support that CMB, particularly ARS, distorts factor structures, emphasizing the necessity of bias correction in RWA modeling.

Our findings suggest that due to the tight semantic entanglement between aggression and conventionalism in RWA20, these two subdomains are most prone to collapsing into a single factor – aggressive conventionalism – than either aggression-submission or conventionalism-submission combinations are to form. While some studies have reported an aggression-submission merger (e.g., Etchezahar, 2012), our analysis shows that aggression and conventionalism consistently correlate more strongly with each other than with submission in bias-corrected datasets. The apparent merging of aggression and submission, along with submission's general instability, stems from method artifacts. In three-factor models that do not account for method bias, aggression and conventionalism are typically defined by same-wording items, leaving submission – a mix of pro- and con-trait items – vulnerable to dissolution into the factor dominated by pro-trait items (aggression) due to stronger contamination of pro-trait items with ARS. Indeed, in our raw data, the third factor (submission) correlated more strongly with the predominantly pro-trait factor (aggression), noticeably exceeding all other inter-factor correlations – an artifact that disappears in bias-corrected data.

The SHCD analysis procedure we proposed can be used to detect method artifacts in data with a general content factor and to evaluate bias-correction techniques. Deviations from normality in SHCD plots (Fig. 1) were all attributable to acquiescence effects, which were successfully eliminated through CMB correction. Pro/con subscale correlations that deviate by more than three standard deviations should be considered a clear red flag for method bias.

This study employed two methods to partial out method bias: ARS and CMB corrections, both effectively removing bias and revealing unbiased latent structures. While Mavor et al. (2010) rejected ARS correction, we found that ARS carries strong scale-specific effects and its measure should be derived from the same scale it corrects, not across scales. ARS correction removed bias but artificially inflated pro/con correlations, whereas CMB correction normalized distributions without introducing artifacts of its own. Strong correlations between these bias measures, despite their different derivations, validated both and quantified method bias, which accounted for up to 57 % of variance in some items. We also extended Billiet & McClendon' (2000) ARS validation criteria and found support for Weijters et al.' (2013) operationalization of confirmation bias as an effect of the first item's directionality in a set.

Method bias undermines data-driven strategies by distorting EFA results, leaving researchers without reliable tools to empirically discern factor structure. Two main responses have emerged: ignoring ARS bias, while taking EFA results at face value, or abandoning EFA in favor of theory-driven models refined through CFA iterations. While the former strategy has shown surprising robustness (Savalei & Falk, 2014), and controlling for acquiescence appeared unnecessary (Billiet & McClendon, 2000, p. 626), our findings suggest that in multifactor models, acquiescence can significantly obscure factor structure. Bias-correction techniques allow EFA to be retained even in the presence of bias. Rather than following the standard raw data → EFA → CFA sequence, data correction can be integrated into the process to address the issue: raw data → bias-corrected data → EFA → CFA → bias measure integration → bifactor CFA on raw data.

CONCLUSIONS

This study provides a psychometric assessment of RWA in Ukraine, confirming its empirical robustness and exposing susceptibility to measurement distortions. By applying novel methods to detect and control response biases, we achieved a more accurate representation of RWA's latent structure. Despite some methodological drawbacks, the scale remains a widely used tool for measuring right-wing authoritarian attitudes, motivating us to develop three-dimensional (RWA13) and unidimensional (RWA6) abridged versions, adapted for the Ukrainian context, for further research (Appendices B and C). An important caveat remains that the scale primarily captures a conservative-traditional worldview rather than pure authoritarianism.

LIMITATIONS

This study relied on a student sample, which may limit its generalizability to broader populations. Among the 90 RWA studies cited here 23,8 % of 120,012 participants were students; excluding two large New Zealand samples, this percentage rises to 52 % of 53,864 participants. To enhance generalizability, future research should test these findings on more diverse samples. Additionally, the small sample size ($N = 99$) influenced our methodological choices.

REPLICABILITY

All data and code are publicly available in an open-access GitHub repository, including the dataset with relevant variables and an R script detailing data processing, bias correction, factor analysis, and model comparisons. Researchers can use these materials for replication or further analysis.

Repository: https://github.com/tarasts/RWA20_Analysis_Ukraine

For inquiries, contact the corresponding author.

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Appendix A

Ukrainian Translation and Item Statistics of Right-Wing Authoritarianism Scale (RWA20) Calculated for Raw and CMB-Corrected Data

Item Code	Item Content	RAW DATA					CMB-CORRECTED DATA					Ukrainian translation of item content
		Standardized Item-Total Correlation		Item-Rest Correlation		Test-Retest Reliability (ICC)	Standardized Item-Total Correlation		Item-Rest Correlation		Test-Retest Reliability (ICC)	
		Wave 1	Wave 2	Wave 1	Wave 2		Wave 1	Wave 2	Wave 1	Wave 2		
1	**2**	**3**	**4**	**5**	**6**	**7**	**8**	**9**	**10**	**11**	**12**	**13**
rwa01	*The established authorities generally turn out to be right about things, while the radicals and protestors are usually just “loud mouths” showing off their ignorance*	*non-scored practice item*										*Органи влади зазвичай праві у важливих питаннях, а радикали та протестувальники – просто «крикуни», які демонструють власну некомпетентність*
rwa02	*Women should have to promise to obey their husbands when they get married*	*non-scored practice item*										*Вийшовши заміж, жінка зобов’язана слухатися свого чоловіка*
rwa03	Our country desperately needs a mighty leader who will do what has to be done to destroy the radical new ways and sinfulness that are ruining us	0,53	0,56	0,48	0,52	0,53***	0,57	0,55	0,50	0,48	0,30**	Нашій державі життєво необхідний сильний лідер, який викорінить радикальні нововведення та аморальність, що руйнують нашу країну
rwa04	**Gays and lesbians are just as healthy and moral as anybody else**	0,51	0,69	0,39	0,63	0,48***	0,65	0,73	0,58	0,69	0,45***	**Геї та лесбійки – такі самі здорові й моральні люди, як усі інші**
rwa05	It is always better to trust the judgment of the proper authorities in government and religion than to listen to the noisy rabble-rousers in our society who are trying to create doubt in people’s minds	0,42	0,50	0,36	0,44	0,54***	0,38	0,46	0,30	0,38	0,47***	Завжди краще довіряти представникам влади та церкви, ніж слухати галасливих провокаторів, які баламутять людям голови
rwa06	**Atheists and others who have rebelled against the established religions are no doubt every bit as good and virtuous as those who attend church regularly**	0,56	0,50	0,48	0,42	0,18*	0,57	0,49	0,52	0,41	0,20*	**Атеїсти та інші, хто виступає проти релігії, – не менш добрі та порядні люди, ніж ті, хто регулярно відвідує церкву**
rwa07	The only way our country can get through the crisis ahead is to get back to our traditional values, put some tough leaders in power, and silence the troublemakers spreading bad ideas	0,62	0,67	0,58	0,63	0,33***	0,61	0,66	0,54	0,59	0,40***	Єдиний шлях, яким наша країна може подолати майбутню кризу, – повернутися до традиційних цінностей, привести жорстких лідерів до влади та закрити роти баламутам, які поширюють згубні ідеї

The Continuation of the Appendix A

1	2	3	4	5	6	7	8	9	10	11	12	13
rwa08	**There is absolutely nothing wrong with nudist camps**	0,31	0,29	0,23	0,21	0,44***	0,29	0,24	0,21	0,15	0,44***	**Немає нічого поганого в нудистських пляжах**
rwa09	**Our country needs free thinkers who have the courage to defy traditional ways, even if this upsets many people**	0,46	0,61	0,38	0,56	0,33***	0,47	0,58	0,40	0,54	0,33***	**Нашій країні потрібні вільнодумці, які наважуються виступати проти традиційних цінностей, навіть якщо їхні дії обурюють багатьох людей**
rwa10	Our country will be destroyed someday if we do not smash the perversions eating away at our moral fiber and traditional beliefs	0,74	0,71	0,71	0,68	0,55***	0,74	0,74	0,69	0,68	0,60***	Наша країна рано чи пізно буде знищена, якщо ми не викорінимо всі ті збочення, які роз'їдають нашу мораль і традиційні цінності
rwa11	**Everyone should have their own lifestyle, religious beliefs, and sexual preferences, even if it makes them different from everyone else**	0,34	0,58	0,22	0,50	0,56***	0,44	0,69	0,36	0,64	0,51***	**Кожен має жити по-своєму, дотримуючись власних релігійних переконань та сексуальних орієнтацій, навіть якщо вони не такі, як у всіх інших**
rwa12	The "old-fashioned ways" and the "old-fashioned values" still show the best way to live	0,63	0,62	0,55	0,55	0,61***	0,63	0,61	0,56	0,55	0,62***	«Старомодні» погляди та цінності залишаються найкращим орієнтиром у житті
rwa13	**You have to admire those who challenged the law and the majority's view by protesting for women's abortion rights, for animal rights, or to abolish school prayer**	0,63	0,62	0,55	0,52	0,47***	0,69	0,78	0,65	0,75	0,42***	**Нам слід захоплюватися людьми, які кидають виклик закону та поглядам більшості, виборюючи право жінок на аборти, захищаючи права тварин та протестуючи проти релігійного виховання в школах**
rwa14	What our country really needs is a strong, determined leader who will crush evil, and take us back to our true path	0,53	0,50	0,49	0,46	0,66***	0,61	0,48	0,55	0,40	0,47***	Нам потрібен сильний та рішучий лідер, який очистить країну від зла й поверне нас на шлях істинний
rwa15	**Some of the best people in our country are those who are challenging our government, criticizing religion, and ignoring the "normal way things are supposed to be done."**	0,40	0,41	0,32	0,32	0,40***	0,38	0,50	0,31	0,44	0,39***	**Найкращі люди нашої країни – це ті, хто кидає виклик владі, критикує церкву та не погоджується жити за усталеними порядками**
rwa16	God's laws about abortion, pornography and marriage must be strictly followed before it is too late, and those who break them must be strongly punished	0,69	0,68	0,62	0,61	0,54***	0,70	0,68	0,64	0,62	0,55***	Поки не пізно, нам слід схаменутися й почати неухильно дотримуватися Божих заповідей щодо абортів, порнографії та шлюбу, а їх порушників – суворо карати

The End of the Appendix A

1	2	3	4	5	6	7	8	9	10	11	12	13
rwa17	There are many radical, immoral people in our country today, who are trying to ruin it for their own godless purposes, whom the authorities should put out of action	0,56	0,35	0,51	0,29	0,32***	0,53	0,29	0,46	0,21	0,32***	У нашій країні багато радикальних й аморальних людей, які намагаються зруйнувати її заради власних нечестивих намірів і яких влада повинна зупинити
rwa18	**A "woman's place" should be wherever she wants to be. The days when women are submissive to their husbands and social conventions belong strictly in the past**	0,55	0,49	0,46	0,39	0,44***	0,67	0,65	0,61	0,58	0,46***	**«Місце» жінки там, де вона сама того хоче. Часи, коли жінки підкорювалися своїм чоловікам та суспільним нормам, безповоротно минули**
rwa19	Our country will be great if we honor the ways of our forefathers, do what the authorities tell us to do, and get rid of the "rotten apples" who are ruining everything	0,43	0,50	0,36	0,44	0,58***	0,39	0,46	0,31	0,38	0,55***	Наша країна стане сильнішою, якщо ми будемо шанувати звичаї предків, виконувати вказівки влади та позбудемося «паршивих овець», які все псують
rwa20	**There is no "ONE right way" to live life; everybody has to create their own way**	0,49	0,51	0,39	0,40	0,56***	0,57	0,68	0,50	0,62	0,49***	**Не існує «єдино правильного» шляху в житті. Кожен має жити по-своєму**
rwa21	**Homosexuals and feminists should be praised for being brave enough to defy "traditional family values"**	0,67	0,64	0,58	0,57	0,59***	0,76	0,68	0,73	0,64	0,57***	**Гомосексуали та феміністки заслуговують на повагу за те, що не бояться кинути виклик «традиційним сімейним цінностям»**
rwa22	This country would work a lot better if certain groups of troublemakers would just shut up and accept their group's traditional place in society	0,65	0,57	0,58	0,53	0,30***	0,66	0,54	0,60	0,47	0,28**	У нашій країні все було б значно краще, якби активісти з «невдоволених» груп закрили роти й прийняли традиційне місце своєї групи в суспільстві

* $p < 0{,}05$ ** $p < 0{,}01$ *** $p < 0{,}001$

Note: Con-trait (reverse-scored) items are presented in bold. Item numbering, embedded in the item codes, follows the same order as in Altemeyer (2006, pp. 11–12), including the two practice items at the beginning. Color shading is applied to facilitate pairwise comparisons of the same indicators calculated for raw and CMB-corrected data. All indicators are based on a dataset of 99 complete cases.

The item-rest correlation assesses the relationship between each item and the total score of the remaining items and corresponds to the *r.drop* column in the output of the alpha() function from the *psych* package in R (Revelle, 2024). Test-retest reliability is measured using the Intraclass Correlation Coefficient (ICC) based on a two-way random-effects model with absolute agreement and single measures, as implemented in the *irr* package in R (Gamer et al., 2022).

The procedure employed to correct for common method bias (CMB) in the data and produce bias-corrected dataset is described in detail within the article.

Appendix B

Item Statistics of 13-Item Abridged Ukrainian Version of Right-Wing Authoritarianism Scale (RWA13) Calculated for Raw and CMB-Corrected Data

Item Code	Item Content	RAW DATA					CMB-CORRECTED DATA					Usage in Other Abridged Versions of RWA Scale
		Standardized Item-Total Correlation		Item-Rest Correlation		Test-Retest Reliability (ICC)	Standardized Item-Total Correlation		Item-Rest Correlation		Test-Retest Reliability (ICC)	
		Wave 1	Wave 2	Wave 1	Wave 2		Wave 1	Wave 2	Wave 1	Wave 2		
1	2	3	4	5	6	7	8	9	10	11	12	13
rwa03	Our country desperately needs a mighty leader who will do what has to be done to destroy the radical new ways and sinfulness that are ruining us	0,54	0,55	0,44	0,47	0,53***	0,58	0,58	0,47	0,47	0,30**	Altemeyer-10, Zakrisson-15
rwa05	It is always better to trust the judgment of the proper authorities in government and religion than to listen to the noisy rabble-rousers in our society who are trying to create doubt in people’s minds	0,44	0,54	0,35	0,46	0,54***	0,42	0,52	0,29	0,42	0,47***	
rwa07	The only way our country can get through the crisis ahead is to get back to our traditional values, put some tough leaders in power, and silence the troublemakers spreading bad ideas	0,66	0,65	0,58	0,58	0,33***	0,65	0,66	0,55	0,58	0,40***	Mavor-14, Duckitt-20, Altemeyer-10
rwa09	**Our country needs free thinkers who have the courage to defy traditional ways, even if this upsets many people**	0,46	0,61	0,33	0,53	0,33***	0,46	0,59	0,35	0,51	0,33***	Mavor-14, Zakrisson-15, Duckitt-20
rwa10	Our country will be destroyed someday if we do not smash the perversions eating away at our moral fiber and traditional beliefs	0,77	0,69	0,70	0,63	0,55***	0,76	0,76	0,68	0,68	0,60***	Duckitt-20
rwa11	**Everyone should have their own lifestyle, religious beliefs, and sexual preferences, even if it makes them different from everyone else**	0,32	0,56	0,17	0,45	0,56***	0,42	0,65	0,31	0,56	0,51***	Altemeyer-10, Mavor-14, Rattazzi-14
rwa12	The “old-fashioned ways” and the “old-fashioned values” still show the best way to live	0,65	0,64	0,53	0,54	0,61***	0,65	0,62	0,55	0,54	0,62***	Zakrisson-15, Altemeyer-10
rwa13	**You have to admire those who challenged the law and the majority’s view by protesting for women’s abortion rights, for animal rights, or to abolish school prayer**	0,63	0,66	0,53	0,55	0,47***	0,69	0,8	0,63	0,75	0,42***	Altemeyer-10
rwa15	**Some of the best people in our country are those who are challenging our government, criticizing religion, and ignoring the “normal way things are supposed to be done.”**	0,41	0,48	0,30	0,36	0,40***	0,4	0,54	0,29	0,46	0,39***	Mavor-14, Zakrisson-15
rwa16	God’s laws about abortion, pornography and marriage must be strictly followed before it is too late, and those who break them must be strongly punished	0,71	0,68	0,62	0,59	0,54***	0,71	0,67	0,63	0,59	0,55***	Zakrisson-15, Altemeyer-10

The End of the Appendix B

1	2	3	4	5	6	7	8	9	10	11	12	13
rwa18	**A "woman's place" should be wherever she wants to be. The days when women are submissive to their husbands and social conventions belong strictly in the past**	0,55	0,53	0,43	0,40	0,44***	0,68	0,66	0,59	0,57	0,46***	Duckitt-20, Funke-12
rwa19	Our country will be great if we honor the ways of our forefathers, do what the authorities tell us to do, and get rid of the "rotten apples" who are ruining everything	0,45	0,52	0,35	0,43	0,58***	0,41	0,49	0,3	0,39	0,55***	Zakrisson-15, Duckitt-20
rwa21	**Homosexuals and feminists should be praised for being brave enough to defy "traditional family values"**	0,68	0,66	0,57	0,57	0,59***	0,78	0,69	0,73	0,62	0,57***	Rattazzi-14, Duckitt-20, Altemeyer-10

* $p < 0{,}05$ ** $p < 0{,}01$ *** $p < 0{,}001$

Note: Con-trait (reverse-scored) items are presented in bold, and item numbering follows the sequence in *Altemeyer* (2006, pp. 11–12). Color shading is applied to facilitate pairwise comparisons between raw and CMB-corrected data. All indicators are based on a dataset of 99 complete cases.

The item-rest correlation assesses the relationship between each item and the total score of the remaining items, corresponding to the *r.drop* column in the output of the alpha() function from the *psych* package in *R* (Revelle, 2024). Test-retest reliability is measured using the Intraclass Correlation Coefficient (ICC), calculated based on a two-way random-effects model with absolute agreement and single measures, as implemented in the *irr* package in *R* (Gamer et al., 2022).

CMB-corrected item scores were computed relative to the full scale (RWA20) and remain unchanged from Appendix A, following the procedure detailed in the article. In contrast, item-total and item-rest correlations are calculated against the abridged scale (RWA13), which explains their differences from the same indicators in Appendix A, where they were computed using the full scale. However, test-retest reliability for each item remains unchanged compared to the full RWA20 (Appendix A).

The last column provides references to other abridged versions of the RWA scale that included these items: *Altemeyer-10* (Altemeyer's 10-item short form, Altemeyer, 2022), *Duckitt-20* (Duckitt & Fisher's 20-item version, Duckitt & Fisher, 2003), *Funke-12* (Funke's 12-item German adaptation, Funke, 2005), *Mavor-14* (Mavor's 14-item abridgment, Mavor, 2012), *Rattazzi-14* (Rattazzi et al.'s 14-item Italian version, Rattazzi et al., 2007), and *Zakrisson-15* (Zakrisson's 15-item Swedish version, Zakrisson, 2005).

Appendix C

Item Statistics of 6-Item Abridged Ukrainian Version of Right-Wing Authoritarianism Scale (RWA6) Calculated for Raw and CMB-Corrected Data

Item Code	Item Content	RAW DATA					CMB-CORRECTED DATA					Usage in Other Abridged Version of RWA Scale
		Standardized Item-Total Correlation		Item-Rest Correlation		Test-Retest Reliability (ICC)	Standardized Item-Total Correlation		Item-Rest Correlation		Test-Retest Reliability (ICC)	
		Wave 1	Wave 2	Wave 1	Wave 2		Wave 1	Wave 2	Wave 1	Wave 2		
1	2	3	4	5	6	7	8	9	10	11	12	13
rwa04	**Gays and lesbians are just as healthy and moral as anybody else**	0,66	0,78	0,49	0,68	0,48***	0,69	0,79	0,54	0,69	0,45***	Mavor-14, Rattazzi-14, Duckitt-20, Smith-9, Altemeyer-10
rwa10	Our country will be destroyed someday if we do not smash the perversions eating away at our moral fiber and traditional beliefs	0,73	0,67	0,58	0,50	0,55***	0,83	0,83	0,74	0,74	0,60***	Duckitt-20
rwa12	The "old-fashioned ways" and the "old-fashioned values" still show the best way to live	0,69	0,73	0,53	0,58	0,61***	0,69	0,72	0,54	0,57	0,62***	Zakrisson-15, Altemeyer-10

The End of the Appendix C

1	2	3	4	5	6	7	8	9	10	11	12	13
rwa13	**You have to admire those who challenged the law and the majority's view by protesting for women's abortion rights, for animal rights, or to abolish school prayer**	0,70	0,68	0,56	0,51	0,47***	0,70	0,76	0,56	0,63	0,42***	Altemeyer-10
rwa16	God's laws about abortion, pornography and marriage must be strictly followed before it is too late, and those who break them must be strongly punished	0,79	0,73	0,67	0,59	0,54***	0,79	0,72	0,68	0,58	0,55***	Zakrisson-15, Altemeyer-10
rwa21	**Homosexuals and feminists should be praised for being brave enough to defy "traditional family values"**	0,81	0,76	0,70	0,64	0,59***	0,81	0,77	0,71	0,66	0,57***	Rattazzi-14, Duckitt-20, Altemeyer-10

* *p < 0,05* ** *p < 0,01* *** *p < 0,001*

Note: Con-trait (reverse-scored) items are presented in bold, and item numbering follows the sequence in *Altemeyer* (2006, pp. 11–12). Color shading is applied to facilitate pairwise comparisons between raw and CMB-corrected data. All indicators are based on a dataset of 99 complete cases.

The item-rest correlation assesses the relationship between each item and the total score of the remaining items, corresponding to the *r.drop* column in the output of the alpha() function from the *psych* package in *R* (Revelle, 2024). Test-retest reliability is measured using the Intraclass Correlation Coefficient (ICC), calculated based on a two-way random-effects model with absolute agreement and single measures, as implemented in the *irr* package in *R* (Gamer et al., 2022).

CMB-corrected item scores were computed relative to the full scale (RWA20) and remain unchanged from Appendix A, following the procedure detailed in the article. In contrast, item-total and item-rest correlations are calculated against the abridged scale (RWA6), which explains their differences from the same indicators in Appendix A, where they were computed using the full scale. However, test-retest reliability for each item remains unchanged compared to the full RWA20 (Appendix A).

The last column provides references to other abridged versions of the RWA scale that included these items: *Altemeyer-10* (Altemeyer's 10-item short form, Altemeyer, 2022), *Duckitt-20* (Duckitt & Fisher's 20-item version, Duckitt & Fisher, 2003), *Funke-12* (Funke's 12-item German adaptation, Funke, 2005), *Mavor-14* (Mavor's 14-item abridgment, Mavor, 2012), *Rattazzi-14* (Rattazzi et al.'s 14-item Italian version, Rattazzi et al., 2007), and *Zakrisson-15* (Zakrisson's 15-item Swedish version, Zakrisson, 2005).