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Addressing a need for valid measures of trait reactance in adolescents: A further test of the Hong Reactance Scale

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Abstract

Research is scant concerning the developmental aspects of trait reactance. If measures are not validated for use across different age groups, it is difficult for researchers to investigate the specificities of reactance across the lifespan. So far, the factor structure and psychometric properties of the Hong Psychological Reactance Scale (HPRS) have not been tested in adolescents. In Study 1, using data from 1,301 Portuguese adolescents (*M* = 14.8 years), we conducted confirmatory factor analysis to test a series of competing factor models. *Post hoc* modifications resulted in a bifactor model with acceptable fit. Bifactor statistical indices showed that HPRS scores are unidimensional. Path analysis via SEM indicated HPRS scores were strongly related to scores from another measure of trait reactance. Study 2, using an independent sample of 327 Portuguese adolescents (*M* = 14.2 years), supported modelling the HPRS with a bifactor model. Finally, our results indicated HPRS scores were negatively correlated with indicators of emotional and cognitive wellbeing, supporting a conceptualization of reactance as patterns of negative cognitions and negative affect. Overall, this study indicates the HPRS is an appropriate measure for assessing trait reactance in adolescents.

Keywords: trait reactance; adolescents; psychometrics; bifactor model; Hong Psychological Reactance Scale

Addressing a need for valid measures of trait reactance in adolescents: A further test of the Hong Reactance Scale

According to the theory of psychological reactance (J. W. Brehm, 1966; S. S. Brehm & Brehm, 1981) most individuals believe that they are able to conduct a variety of behaviours at will. When this perceived ability is threatened, or removed entirely, individuals experience an aversive motivational state with affective and cognitive components (*state reactance*) that drives behavioural and cognitive efforts towards reinstating their sense of freedom. Such efforts can include direct engagement with the restricted behaviour, denying the threat, or exercising the freedom in an alternative way (S. S. Brehm & Brehm, 1981; Dillard & Shen, 2005; Rains, 2013; Rains & Turner, 2007; Smith, 1977; Wicklund, 1974).

The theory of psychological reactance also acknowledges that individual differences play a role in determining tendency to react (S. S. Brehm & Brehm, 1981). *Trait reactance*, operationalized as the propensity to experience the affective and cognitive components of state reactance when freedom is threatened, has been of particular interest in clinical psychology where the susceptibility of a client to react has direct implications for the behavioural tasks used (Jahn & Lichstein, 1980) and the effectiveness of interventions (Weeks & L'Abate, 1982). Given its practical relevance, it is important that measures of trait reactance assess this construct in their target populations in a reliable and valid manner. The present study aimed to validate a well-studied measure of trait reactance, the Hong Psychological Reactance Scale (Hong & Page, 1989), for use with a sample not yet tested in prior research: adolescents.

Psychometric properties of the Hong Psychological Reactance Scale

The Hong Psychological Reactance Scale (HPRS), an adaptation of Merz's (Merz, 1983) original 18-item Psychological Reactance Questionnaire, has had its factorial structure assessed on numerous occasions in the almost 30 years since in publication by Hong and Page (Hong & Page, 1989). Studies have supported both multidimensional (four and two factors) and unidimensional factor solutions (see Table 1), although as will become apparent, it appears that evidence is converging on the latter.

Four-factor models. A first series of articles by Hong and her colleagues have championed two differing four-factor structures of the HPRS (Hong, 1992; Hong & Faedda, 1996; Hong & Page, 1989). The first structure, proposed by Hong and Page (Hong & Page, 1989) and Hong (Hong, 1992), included factors labelled as Freedom from Choice, Conformity Reactance, Behavioural Freedom, and Reactance to Advice and Recommendations. As is evident in Table 1, there were differences in item loadings even across these two studies. Hong and Faedda (Hong & Faedda, 1996) re-examined the HPRS, both the full version and a reduced 11-item version, in a larger sample of undergraduates. These principle component analyses revealed a different four-factor structure, comprising Reactance to Compliance, Resisting Influence from Others, Reactance to Advice and Recommendations, and Emotional Response to Restricted Choice factors. It is important to note that the analyses conducted by Hong et al. all used exploratory analyses with an orthogonal extraction method, in other words, assuming that the HPRS scale is multidimensional. This approach is curious considering that these authors then provided measures of reliability based on total scores, rather than for separate subscales, in essence assuming unidimensionality.

In a subsequent study using confirmatory factor analysis (Thomas, Donnell, & Buboltz, 2001), the 11- and 14-item factor models offered by Hong and Faedda (Hong & Faedda, 1996) were shown to have adequate fit in a sample of American undergraduates, but only when factors were allowed to correlate. The factor structures offered by Hong and Page (Hong & Page, 1989) and Hong (Hong, 1992) did not fit the data adequately. Furthermore, while the correlations between factors were indicative of a higher-order relationship, models including higher-order factors (which are not reported in the article) were said not to support this hypothesis. This conclusion is questioned by a later study by Shen and Dillard (Shen & Dillard, 2005) who conducted confirmatory factor analyses with a maximum likelihood method to assess the 11-item factor structure offered by Hong and Faedda (Hong & Faedda, 1996). These authors assessed a four-factor model and a higher-order model, and ultimately concluded that the higher-order model was an acceptable description of the HPRS items. While this hinted that the HPRS might be a unidimensional scale, none of the aforementioned studies using confirmatory methods directly tested a one-factor model.

Unidimensional models. Other studies have since gone further to support the proposal that the HPRS is a unidimensional scale. Jonason and Knowles (Jonason & Knowles, 2006) aggregated data from past studies (Hong, 1992; Hong & Faedda, 1996; Hong & Ostini, 1989; Thomas et al., 2001; Tucker & Byers, 1987) and concluded that the HPRS is a unidimensional rather than multidimensional scale. However, a later study by Jonason (Jonason, 2007) revealed that while the unidimensional version of the HPRS had generally adequate internal consistency (with alphas between .74 and .67) it failed to show previously demonstrated relationships with sex, age, and race, thus questioning its usefulness. The authors argued that this outcome was likely a result of the scale containing poor items. Jonason, Bryan and Herrera (Jonason, Bryan, & Herrera, 2010) attempted to remove the "bad items" from the model, and presented a 10-item reduced version of the scale.

A two-factor model. Hong and Faedda (Hong & Faedda, 1996), and Hong, Giannakopoulos, Laing and Williams (Hong, Giannakopoulos, Laing, & Williams, 1994) explicitly called for research investigating the effect of culture on psychological reactance, but to date just one study has examined the properties of a non-English version of the HPRS in a culture other than the U.S.A or Australia (De las Cuevas, Peñate, Betancort, & de Rivera, 2014). This article assessed the factorial structure of the Spanish version of the HPRS (Pérez García, 1993). An exploratory analysis revealed two dimensions, labelled as Affective and Cognitive reactance, although the relatively high correlation between these dimensions (r = .58) was indicative of a higher-order one-factor solution. They then used CFA to test four-factor (14 and 11 item versions), two-factor, and one-factor models. Ultimately, the two-factor model with affective and cognitive dimensions had the best fit.

Bifactor models. A final, more recent, series of investigations of the HPRS have been conducted by Yost and her colleagues (Brown, Finney, & France, 2011; Yost & Finney, 2018). While Brown et al. (Brown et al., 2011) demonstrated that a four-factor model fit well to their data, they added to our understanding of the HPRS by testing a bifactor model. In this bifactor model the 14 items load on a general reactance factor, but specific factors are included to account for shared variance between items due to similar wording/context. The good fit of data to this model demonstrated that trait reactance should be modelled as a unidimensional construct after partialing out the nuisance effects accounted for by four specific factors. By calculating bifactor statistical indices (McDonald, 1999), (Yost & Finney, 2018) further demonstrated that the 14-item HPRS was sufficiently unidimensional to use and interpret a total score of reactance.

A need for validation in adolescent samples

A limitation of the studies outlined above is that all used adult samples. There are theoretical reasons to anticipate mean differences in levels of reactance between adolescents and adults. Adolescence is a critical period of cognitive, emotional and social transformation (Milyavskaya et al., 2009; Steinberg, 2001) during which individuals develop a sense of identity and autonomy (Erikson, 1968a, 1968b). The pursuit of independence and individuality across this period means that adolescents are particularly sensitive to the rules, regulations, and increased responsibilities that can be perceived as a hindrance to their ability to establish a sense of self-determination. Restrictions, like those given by over-controlling parents, can be seen to block the basic psychological needs required for an adaptive social development (Deci & Ryan, 2008), particularly the needs for autonomy and competence (Grolnick, Deci, & Ryan, 1997). Adolescence is also a period of important life transitions, such as the movement from middle to high school, and high school to higher education. These educational transitions alone provide the conditions for high levels of anxiety and uncertainty, such as the trend for a deterioration in grades (Barone, Aguirre-Deandreis, & Trickett, 1991; Isakson & Jarvis, 1999) and increased feelings of loneliness (Benner & Graham, 2015), which might also thwart the fulfilment of autonomy and competence needs, and lead to reactance and non-compliance. Although research on age and reactance is scarce, studies have shown younger participants have higher reactance than older participants (Hong et al., 1994), although this relationship may not be linear, with older adults (55+) showing increased reactance (Woller, Buboltz, & Loveland, 2007).

Despite the indications that adolescents are more reactive than adults, there is little theoretical reason or evidence to suggest that the factorial structure or psychometric properties of the HPRS would be different for adolescents compared to adults. Indeed, many of the studies examining the factorial structure of the HPRS used undergraduate student samples (early twenties) who were, therefore, only a few years beyond adolescence. This may, for some, be sufficient to make an assessment of the HPRS in adolescents redundant, particularly when one considers the large number of studies that have tested the HPRS in adults. We argue that this would be mistaken. Our reasoning for this is that current research has made surprisingly little consideration of the developmental aspects of reactance (but see Dowd, Pepper, & Seibel, 2001).

Research in developmental psychology has shown that some constructs show continuity over time (Caspi & Roberts, 2001) and others, such as personality dimensions, complexify over the lifespan. Adolescence is a critical period for this progressive differentiation and complexification (Sebastian, Burnett, & Blakemore, 2008). This is evident in both emotional and cognitive dimensions, especially for higher order socio-cognitive processes such as self-concept, which are central aspects of personal agency mechanisms and the self-regulation of dispositional tendencies. This is confirmed by longitudinal and meta-analytic studies using different personality models such as the Five-Factor Model (with the dimension of conscientiousness registering a significant change across different development stages) (Roberts, Walton, & Viechtbauer, 2006) or the psychobiological model of personality (with dimensions of self-directedness, cooperation and self-transcendence registering different values across different ages) (Josefsson et al., 2013; Moreira et al., 2015). Considering this evidence, we consider it warranted for research to investigate whether the same might be true for reactance. However, without a validated tool for measuring reactance in both adult and adolescent samples it is difficult to fully understand the specificities of human reactance at different psychological stages of development.

Such potential additions to current knowledge are likely to have important practical implications. Adolescent compliance with contextual, institutional, and relational rules and expectations is associated with the presentation of adaptive behaviors. These in turn have implications for adolescents' subjective experiences, behavioural outcomes, and developmental trajectories. Practitioners working with adolescents in different contexts (from schools to therapeutic settings) have a need to identify individual clients at risk of reactance and non-compliance, including failing to comply with suggestions and to adhere to expectations (e.g. following rules, attend appointments), and those at risk of dropping out of therapy, treatment or school (e.g. Madsen, McQuaid, & Craighead, 2009).

Objectives

The overall aim of the present investigation was to further assess the psychometric properties of the HPRS. Specifically, we aimed to add to the current psychometric literature surrounding this measure by testing its factorial structure and psychometric properties in adolescents. In Study 1, our aim was to test a number of competing models using factor analysis, making *post hoc* modifications when necessary, and to assess other psychometric properties including scale unidimensionality, measurement invariance, and convergent validity. The aim of Study 2 was to replicate the factorial structure identified in Study 1 in an independent sample of adolescents, and to further test the nomological network of trait reactance.

Study 1

Method

Participants and procedure

Schools and students were recruited using a convenience sampling strategy. We directly contacted schools from the north of Portugal because this is the most densely populated region and where the majority of schools are located (although note that the north and south of Portugal are culturally similar). The principle aim of Study 1 was to validate the HPRS in a previously unvalidated age group: the transitional period between childhood and adulthood historically considered to span between the ages of 12 and 18. We therefore purposefully wanted to recruit participants representing a wide age-range to capture the broad characteristics and variation within this developmental period. Consequently, for the four schools that agreed to participate (all of which were secondary schools), teachers administered consent forms to all students. Data were then collected (paper format questionnaires in the context of a supervised class) from students who consented to participate, and whom obtained parental consent if younger than 18 years. Each school accounted for between 6.4% and 50.1% of the total sample. Students did not receive any reward for participating in the study. We obtained ethical approval from the directive board of the Psychology for Positive Development Research Center (CIPD), Portugal, prior to conducting the study.

A total of 1,301 adolescents completed the study measures. The students were distributed approximately evenly across the 9th (27.3%), 10th (27.0%), 11th (25.5%) and 12th (20.2%) grades. The mean age of participants was 14.8 years (*SD* = 1.2, *Range* = 12 – 18). Of the total sample, 46.7% were male, and 52.7% were female. Almost all students were Portuguese (98.8%). Overall, as is typical in Portugal, the parents of the students had a low level of education; 61.5% of mothers and 68.9% of fathers had achieved a high school education or lower.

Measures

Hong Psychological Reactance Scale. Participants completed the 14-item Portuguese-language version of the HPRS (Hong & Page, 1989). This version was adapted following a well-known translation-backtranslation procedure (van der Vijver & Leung, 1997) in which the original English-language items of the HPRS were translated into Portuguese by the authors (who are Portuguese), and then backtranslated into English by two independent bilingual researchers. The original, Portuguese, and backtranslated versions of the HPRS were then examined by the authors to ensure accuracy and cultural sensitivity.

The HPRS items are scored on a 5-point Likert scale ranging from 1 (*completely disagree*) to 5 (*completely agree*). Prior studies have described these items in terms of emotional responses to restricted choice (e.g. “The thought of being dependent on others aggravates me”), reactance to compliance (e.g. “Regulations trigger a sense of resistance in me”), resisting influence from others (e.g. “I resist the attempts of others to influence me”), and reactance to advice (e.g. “I consider advice from others to be an intrusion”). For the present sample the HPRS had acceptable reliability (α = .75).

The Therapeutic Reactance Scale. We also administered a Portuguese translation of the 28-item Therapeutic Reactance Scale (TSR; Dowd, Milne, & Wise, 1991) translated into Portuguese using the translation-backtranslation procedure described above. Dowd et al. (Dowd et al., 1991) identified a two-factor structure of the TRS via factor analysis, including Behavioral Reactance (17 items, e.g. “I consider myself more competitive than cooperative”) and Verbal Reactance (11 items, e.g. “I enjoy debates with other people”). All items are scored on a four-point Likert scale from 1 (*completely disagree*) to 4 (*completely agree*). Consistent with the scoring of the original TRS (Dowd et al., 1991), we calculated a total reactance score by summing the items scores. For the study sample, the TRS was found to have good reliability (α = .74).

Data Analysis

All data analyses were conducted using R (R Core Team, 2019). Prior to our main analyses we assessed the nature of missing data using the *MissMech* package (Jamshidian, Jalal, & Jansen, 2014). This analysis indicated that missing data were missing completely at random (MCAR).

Confirmatory factor analysis. We conducted CFAs using the *lavaan* package (Rosseel, 2012) to test a series of competing factor models. Path diagrams for these models are presented in supplementary materials. Models 1 to 5 were a one-factor model, a correlated four-factor model, a higher-order model with four first-order factors, the two-factor model championed by De las Cuevas et al. (De las Cuevas et al., 2014), and the bifactor model incorporating the *Anger*, *Rules and Regulations*, *Advice*, and *Independence* specific factors proposed by Brown et al. (Brown et al., 2011) (see Figure S1 for models). The Advice specific factor was modelled as a correlated error instead of a specific factor because a factor with only two items would not be identified. Because none of these models showed adequate fit, *post hoc* structural modifications were made to form two incomplete bifactor models (models 6 and 7; see Figure S2 for proposed models). The four first-order factors included in models 2 and 3 were those identified by Hong and Faedda (Hong & Faedda, 1996). These were Emotional Response toward Restricted Choice (items 4, 6, 7 and 8), Reactance to Compliance (items 1, 2, 3 and 14), Resisting Influence from Others (items 10, 11, 12 and 13) and Reactance to Advice and Recommendation (items 5 and 9). Items 5 and 9 were constrained to have equal loadings on their factor, as advised by Kenny, Kashy, and Bolger (Kenny, Kashy, & Bolger, 1998), as otherwise the model would not be identified. As is recommended with bifactor models, the specific factors were constrained to be orthogonal with the general factor, and with one another (Chen, West, & Sousa, 2006).

To account for the fact that the endogenous variables were categorical, we used a robust weighted least squares (WLSMV) estimator, which is specifically designed for ordinal data and less biased than robust maximum likelihood (Li, 2016). Items were declared as being ordinal using the *ordered* argument of the *cfa* function (*lavaan* package). Because missing data were MCAR, we used a full-information technique (Asparouhov & Muthén, 2010), meaning that models were estimated from the entire data set. The metrics of each latent factor for all models were defined through constraining one item loading to 1. We report standardized estimates despite using unstandardized parameters in the models. To assess the goodness of fit for these models we used a number of indicators and heuristics: 1) the Chi-square test (χ^2) and χ^2/df ratios, which are recommended to be ≤ 5 (Schumaker & Lomax, 2010); 2) the Comparative Fit Index (CFI), and 3) the Tucker Lewis Index (TLI), which are both recommended to be ≥ .95 (Cangur & Ercan, 2015; Hu & Bentler, 1999); 4) the Root-Mean Square Error Approximation (RMSEA), for which values < .08 indicate acceptable fit (Browne & Cudeck, 1992); and 5) the Standardized Root Mean Square Residual Index (SRMR), for which values < .05 indicate of good fit (Hu & Bentler, 1999).

Assessment of unidimensionality. Considering the work of Reise and colleagues (Reise, 2012; Reise, Moore, & Haviland, 2010) we made our first assessment of scale unidimensionality by comparing the standardized factor loadings for the general and specific factors of the championed bifactor model. Stronger loadings on the general reactance factor compared to the specific factors are an indication of score unidimensionality. We also calculated Omega Hierarchical (ω_H) (McDonald, 1999; Zinbarg, Revelle, & Yovel, 2007), which determines the proportion of variance accounted for by the general factor. To interpret scores as a measure of a single construct, ω_H should be around .75, and no less than .50 (Reise, Scheines, Widaman, & Haviland, 2013). We then compared ω_H with Omega (ω), which estimates the proportion of variance explained by both general and specific factors (McDonald, 1999). This difference illustrates how much a total HPRS score represents the construct of reactance despite multidimensionality; low scores indicate that systematic variance is mostly a result of the general factor. Finally, Reise (Reise, 2012) proposed that if a bifactor model is truly unidimensional then similar general factor loadings will be identified if the analysis is conducted with subsets of items. We therefore reran the bifactor model with four different subsets of nine items.

Measurement invariance. To assess measurement invariance across gender we used a multi-group CFA approach (*measurementInvarianceCat* function of the *semTools* package) (62) that tests nested models differing in the number of applied restrictions. Increasing restrictions allows the testing of configural, metric, and scalar invariance (Meredith & Teresi, 2006). Invariance was determined based on the change in CFI and RMSEA between models (ΔCFI $\leq$ -.010 and ΔRMSEA $\geq$ .015 indicative of non-invariance) (Chen, 2007).

Construct validity. If the HPRS has construct validity participant scores for this scale should be positively related to scores from other validated instruments measuring reactance. We therefore used Structural Equation Modelling (*sem* function of *lavaan* package) to test the association between HPRS and TRS scores.

Results

Item statistics and inter-item polychoric correlations for Study 1 are presented in Supplementary Table 1 (Table S1).

Tests of factorial structure via CFA. Table 2 presents the fit indices for each of the eight models tested. The one-factor model (Model 1), the correlated four-factor model (Model 2), the higher-order model (Model 3), and the two-factor correlated model (Model 4) did not show acceptable fit. Model 2 was inadmissible because of issues with discriminant validity, specifically $r > 1.0$ between the Reactance to Compliance and Reactance to Advice and Recommendation factors. Model 3 was also inadmissible because of a negative error variance for the Reactance to Compliance latent factor. This indicates that data are being over-factored (Chen et al., 2006) and the Reactance to Compliance factor is not represented in the data.

The bifactor model proposed by Brown et al. (Brown et al., 2011) did not converge (see Model 5) and an inspection of the output revealed large negative error terms for two items. Our solution was to test an incomplete bifactor model (Model 6) that excluded the Rules and Regulations and Independence specific factors (Chen et al., 2006). This model did converge, but did not have acceptable fit. Finally, in accordance with Brown et al. (Brown et al., 2011), Model 7 incorporated an Opposite specific factor. We also decided that the Anger specific factor would include two further items; items 12 ("It makes me angry when another person is held up as a role model for me to follow") and 14 ("It disappoints me to see others submitting to society's standards and rules"). Values for CFI (.97), RMSEA (.06), TLI (.95), SRMR (.05), for this model were indicative of acceptable fit. The χ^2/df ratio (5.46), however, fell just short of the threshold for acceptable model fit.

Assessment of unidimensionality. The factor loadings and error terms for Model 7 are summarised in Table 3. These data reveal that four of the nine items loading on specific factors had larger standardized loadings on the general factor. This is inconsistent with other similar comparisons (Brown et al., 2011; Yost & Finney,

2018) that showed almost all factor loadings were stronger for the general reactance factor. However, values for Omega and Omega Hierarchical were calculated as, $\omega = .82$ and $\omega_H = .63$, meaning 77% (.63/.82 = .77) of the reliable variance in HPRS scores was due to the general factor. Calculations of the systematic variance associated with the Anger and Opposite specific factors (ω_{HS}) indicated that 43% of the remaining modelled total score variance was due to variance in the Anger specific factor, and 54% was due to variance in the Opposite specific factor. Finally, as a test of parameter invariance, Model 7 was retested with four different subsets of items. Standardized factor loadings summarized in Supplementary Table 2 (Table S3). Consistent with what would be expected with a unidimensional scale (Reise, 2012), factor loadings were broadly similar across the subsets of items.

Measurement invariance. Model 7 had generally acceptable fit to the data for both male (CFI = .95, TLI = .93, RMSEA = .06, SRMR = .06) and female (CFI = .97, TLI = .96, RMSEA = .05, SRMR = .06) students. Based on ΔCFI = .006, ΔRMSEA = .008, and $\Delta\chi^2 = 21.54$, $p = .37$, our results suggest that metric invariance across gender was established. Scalar variance, however, was not achieved, as the value for ΔCFI (.016) was above the .010 threshold, and $\Delta\chi^2 = 106.63$, $p < .001$. ΔRMSEA was acceptable at .001.

Construct validity. We used SEM to test the association between mean HPRS scores and a latent reactance factor with two TRS dimensions (path diagram shown in Figure 1). The hypothesized model had acceptable fit based on CFI (.96) and SRMR (.08), but did not meet the thresholds for RMSEA (.16) or TLI (.87). The model showed that trait reactance as measured by the HPRS was strongly related to trait reactance measured by the TRS (β = .89, SE = .01).

Study 2

The purpose of Study 1 was to test the psychometric properties of the Portuguese version of the HPRS in sample of adolescents. In summary, the results converged with extant research with adults in their indication that the HPRS can be considered a unidimensional measure of reactance. Based on these findings, the purpose of Study 2 was twofold; (a) to replicate the modified incomplete bifactor solution (Model 7) in an independent sample of Portuguese adolescents; and (b) to further build on our understanding of the scales nomological network by examining the relationship between reactance and multiple external variables for which theory and prior research allow for the creation of *a priori* hypotheses.

Previous studies have already validated the HPRS in relation to personality variables (Yost & Finney, 2018), perceived threat to freedom (Shen & Dillard, 2005), attitudes (De las Cuevas et al., 2014; Shen & Dillard, 2005), satisfaction with life (Hong & Faedda, 1996; Jonason & Knowles, 2006), locus of control (De las Cuevas et al., 2014; Jonason & Knowles, 2006), depression (Jonason & Knowles, 2006), tendencies to abdicate authority to others and to aggress (Jonason et al., 2010), and self-efficacy (De las Cuevas et al., 2014). To further expand the nomological network of psychological reactance we were interested in examining the relations between psychological reactance and the affect and cognitions related to wellbeing. Our justification for doing so was that since Brehm's (J. W. Brehm, 1966) writings, evidence has mostly converged on a conceptualization of reactance as an intertwined interaction between negative cognition and negative affect (Dillard & Shen, 2005; Kim, Levine, & Allen, 2013; Quick, 2012; Quick & Kim, 2009; Quick & Stephenson, 2007, 2008; Rains, 2013). Prior studies have shown that affective wellbeing, which has been referred to as *happiness* (Cloninger & Zohar, 2011), is negatively related to psychological reactance (Jourbert, 1990), and that negative affect, in this case anger, is positively related to reactance (Hong & Faedda, 1996). Based on these findings, we hypothesized that reactance would have a negative association with affective wellbeing (happiness). Hong and Faedda (Hong & Faedda, 1996) have also shown significant negative correlations between life satisfaction, which is considered a crucial cognitive element of wellbeing (Diener, 1994), and the HPRS. Not all studies, however, have replicated the association between life satisfaction and reactance (Jonason & Knowles, 2006), and consequently we did not hypothesize the relation between cognitive wellbeing (which has been described as *wellness*) (Cloninger & Zohar, 2011) and reactance, instead merely exploring the association. No studies have assessed the relation of trait reactance with wellbeing as a higher-order construct, incorporating both affective (happiness) and cognitive (wellness) components, and so we also explored this association.

Method

Participants and procedure

Schools and students were recruited in the same manner as they were in Study 1. In total, 327 students from eight schools in the north of Portugal (98.1% Portuguese nationals) chose to participate in this study (each school accounted for between 3.4% and 35.2% of the total sample). Unlike for Study 1, we were able to recruit students from middle schools as well as high schools. Consequently, students were from the 5th to 12th grades (5th = 28 students, 6th = 60 students; 7th = 34 students; 8th = 15 students; 9th = 65 students; 10th = 99 students, and 12th = 26 students) and were thus aged between 10 and 17 years old, with a mean of 14.24 (*SD* = 2.19). Overall, 151 participants were male and 176 were female. As in Study 1, all students were offered informed consent forms to be signed by legal guardians, and those who returned these forms were able to complete paper versions of all measures while being supervised by a member of the research team in a classroom context.

Measures

Participants completed the Portuguese language versions of the HPRS (Hong & Page, 1989) and TRS (Dowd et al., 1991), the details of which are presented in Study 1. The HPRS (α = .78) and TRS (α = .71) had acceptable reliability when applied to the present sample.

Wellbeing.

Life Satisfaction. Life satisfaction was assessed using the Brief Multidimensional Students' Life Satisfaction Scale (BMSLSS; original version by Seligson, Huebner, & Valois, 2003), which we had translated into Portuguese using a translation-backtranslation procedure. This scale comprises six items that address degree of satisfaction across six domains including family, friends, school experience, self, environment, and life in general (e.g. "My family life is…", "My satisfaction with where I live is…"). Responses are given on a seven point Likert-type scale from 0 (*terrible*) to 6 (*fantastic*). A total life satisfaction score is determined by summing the six items. Within the sample of the present study, the BMSLSS had good reliability (α = .82).

Satisfaction with Social Support. Students' satisfaction with their social support was measured using the Portuguese version of the Brief Satisfaction with Social Support Scale for children and adolescents (Gaspar, Ribeiro, Matos, Leal, & Ferreira, 2009). This instrument comprises 12 items (e.g. "I am satisfied with the amount of friend I have", and "I'm not with friends as much as I'd like to be") scored on a five point Likert-type scale from 1 (*totally agree*) to 5 (*totally disagree*). Seven items were reverse coded so that higher scores across all items reflected higher satisfaction. For the present sample, this measure had good reliability (α = .83).

Quality of Life. We used a Portuguese translation of KIDSCREEN-10 (Erhart et al., 2009; Matos, Gaspar, & Simões, 2012) to measure mental-health and wellbeing. The ten items of this instrument (e.g. "Do you feel fit and well?"), scored on a Likert-type scale from 1 (*not at all*) to 5 (*extremely*), measure the affective, cognitive, and psycho-social aspects of mental health. Two items were recoded so that higher scores reflected better quality of life. This short instrument had good reliability in the study sample (α = .82).

Affect. We used a Portuguese adaptation of the Positive and Negative Affect Scale (PANAS) (Watson, Clark, & Tellegen, 1988) to assess the emotional component of wellbeing. This scale consists of 12 positive (e.g. "Enthusiastic", and "Content") and 15 negative (e.g. "Sad", and "Nervous") adjectives for describing feelings and emotions, which participants indicate the extent to which they feel on a 5-point Likert-type scale from 1 (*very slightly or not at all*) to 5 (*extremely*). For the present study, the positive (α = .93) and negative (α = .94) subscales of PANAS had excellent reliability.

Composite indicators of affective and cognitive wellbeing. In accordance with prior studies interested in the higher-order dimensions of wellbeing (Cloninger & Zohar, 2011; Josefsson et al., 2011; Moreira et al., 2015), we estimated composite indicators corresponding to wellness (cognitive wellbeing) and happiness (affective wellbeing). The composite wellness index was the mean average score for satisfaction with social support, satisfaction with life, and quality of life scales. The composite affective wellbeing (happiness) index was computed by subtracting the mean score of negative adjectives from the mean score of positive adjectives using the PANAS. Negative scores indicate a predominantly negative emotional experience and positive scores indicate a predominantly positive emotional experience.

Data Analysis

Data were screened as they were in Study 1, and there was no evidence that missing data were not MCAR. In an attempt to replicate the bifactor solution championed in Study 1, the data were subjected to CFA using Model 7. In all respects, this analysis was identical to that run in Study 1, and we interpreted goodness-of-fit indices with the same heuristics. Next, we assessed the construct validity of the HPRS by calculating Pearson's correlation coefficients for the relationship between total HPRS scores, TRS scores, and multiple indicators of wellbeing. Bootstrapping was used to compute 95% confidence intervals for these coefficients.

Results

Inter-item polychoric correlations and item statistics for Study 2 are shown in Supplementary Table 3 (Table S3).

Test of factorial structure via CFA. The modified incomplete bifactor model championed in Study 1 was found to have acceptable fit based on CFI (.96), χ^2/df (2.18), and RMSEA (.07, 95% CIs [.05, .08]). The values for TLI (.94) and SRMR (.07) fell just short of the thresholds of acceptable fit. Factor loadings for this model are summarized in Table 3.

Construct validity. Table 4 summarizes the correlations between the HPRS, the TRS, and multiple indicators and composite scores related to wellbeing. Total HPRS scores had moderate to large positive correlations with the behavioral (r = .50) and verbal (r = .36) reactance subscales of the TRS, and a large positive correlation with the total TRS score (r = .50). Consistent with our expectations, reactance had a positive correlation with negative affect (r = .21), and negative correlation with positive affect (r = -.16). The result was a negative correlation with student composite affective wellbeing, or happiness (r = -.22). Reactance also had negative correlations with the individual indicators of cognitive wellbeing (r = -.19 to - .22) as well as with the composite indicator of wellness (r =-.26). Finally, reactance had a moderate negative correlation with the overall composite indicator of wellbeing (r = -.27).

Discussion

The present article presents a series of psychometric assessments using two independent samples of adolescents. The overall finding of the investigation was that the HPRS is an adequate measure of trait reactance in adolescents. Following similar assessment steps to those used by Brown et al. (2011), an evaluation of several competing models using CFA did not support modelling the HPRS via a unidimensional model, a four-factor model, a higher-order four-factor model, or a two-factor model. The bifactor model championed by Brown et al. (2011) did not converge, and an inspection of the output for this analysis revealed a number of items with negative disturbance, which can be interpreted as meaning that the model was misspecified, most likely because of over-factoring (Chen et al., 2006). We consequently proceeded to make *post hoc* model modifications. An incomplete bifactor model with one specific factor – two were removed to account for nuisance variance – was not suitable for modelling the HPRS. Finally, a structure including an expanded Anger specific factor and additional Opposite specific factor was found to be acceptable for modelling the HPRS in adolescents. The principle finding of this investigation, therefore, was that none of the competing models from prior studies were suitable for modelling the HPRS in adolescents. Nonetheless, our championed bifactor model, despite being modified, converged with past findings that indicate the HPRS is best modelled as a general reactance factor with specific factors that account for shared variance between sets of items.

In addition to examining the fit of competing models, Study 1 used model-based psychometric indexes (ω_H) to directly test the unidimensionality of HPRS scores. One initial finding that differed from past studies was that around half of the items loaded more strongly on the general factor than their specific factor. Most past studies have shown a clear weighting toward loadings on the general factor (Yost & Finney, 2018). This finding thus makes it less clear whether HPRS scores are truly unidimensional. However, after calculating ω and $\omega_{H,}$ it was clear from our results that the HPRS scores remain unidimensional when applied to adolescents.

Although model-based psychometric indexes have been used to assess the unidimensionality of HPRSs scores in other studies (Yost & Finney, 2018), Study 1 adds to current literature by testing the item parameter invariance of the bifactor model. These analyses, using subsets of items, further confirmed that the HPRS is unidimensional, a finding consistent with the assertions made by Yost and colleagues (2018) that trait reactance is a broad unidimensional construct that becomes multidimensional when operationalized in the HPRS. It therefore seems that multiple factors emerged in past research because the HPRS items, worded purposefully to be heterogeneous, converged into artefact factors based on similar wording and/or content.

The SEM model used to test concurrent validity in Study 1 indicates that the HPRS has validity as a measure of trait reactance in adolescents. Trait reactance as measured by the HPRS had a strong positive association with trait reactance measured using the TRS (Dowd et al., 1991). This demonstrates that these scales measure similar constructs, and thus serves to validate the HPRS as a measure of reactance. In Study 2, we calculated correlations between total HPRS score and indicators of wellbeing. One of our *a priori* predictions about these correlations was that HPRS would share a negative correlation with affective wellbeing. Our analyses supported this hypothesis by revealing a positive correlation with negative affect, a negative correlation with positive affect, and significant negative correlation with a composite indicator of affective wellbeing. Although past studies have shown only tentative links between reactance and cognitions relevant to wellbeing (Hong & Faedda, 1996), leading us to simply explore the association rather than form solid hypotheses, our results indicated a clear link between negative cognitions and reactance. These analyses therefore support the conceptualization of reactance in terms of negative affect and negative cognition (Dillard & Shen, 2005; Kim et al., 2013; Quick, 2012; Quick & Kim, 2009; Quick & Stephenson, 2007, 2008; Rains, 2013), and indicate the HPRs has convergent validity in adolescents.

Our study also adds to the current understanding of the HPRS by demonstrating its validity in a sample from a country and culture (Portugal) different from those used in the majority of past research (U.S. and Australia). The results imply that it is acceptable to model the HPRS in terms of general and specific factors and to calculate total reactance scores with individuals from a collectivist society. This finding aligns with one other study testing perceived threats to freedom in a sample from South Korea (Quick & Kim, 2009), which also concluded the construct of trait reactance was applicable to collectivist cultures.

When interpreting the results of the present study it is important to consider that the sample differed from those most frequently used in past studies in two ways; 1) the sample comprised adolescents rather than adults; and 2) these adolescents were from a collectivist rather than individualist society. This point is relevant because it precludes any clear inference about why the factorial structure championed for Portuguese adolescents differs to that observed for adult samples from different cultures. Future studies testing the factorial structure in Portuguese adults would be an important step to clarifying this issue. Nonetheless, from the present study it remains valid to conclude the version of the HPRS tested is adequate for use with Portuguese adolescents, and to infer that the HPRS is best modelled as a general reactance factor with specific factors.

Implications

These findings complement past research by validating the HPRS for use with adolescents. For researchers, this validation serves as a preliminary indication that this measure of trait reactance may be suitable for investigating the development of trait reactance across the life span and in important contexts. Understanding how adolescents' subjective experiences of school are expressions of reactance, for example, is a major research challenge for education (Moreira & Garcia, 2019).

At a practical level, having a quick-to-administer and validated tool for assessing trait reactance in adolescents will help practitioners working in different contexts (from schools to therapeutic settings) identify individuals at risk of non-compliance (e.g. school disengagement/dropout, failure to follow rules, non-adherence with therapy/treatment). As well as being useful for identifying non-compliant individuals, understanding an adolescent's trait reactance may be useful for practitioners to consider when performing interventions because reactance is theoretically expected to influence behavioral compliance (Jahn & Lichstein, 1980). Indeed, reactance has been shown to be associated with noncompliance with medical recommendations (Fogarty & Youngs, 2000). Moreover, communication research has indicated that trait reactance has a moderating effect on the influence of different types of language/message on compliance (Dillard & Shen, 2005).

Limitations and future research

One limitation specific to Study 2 is that the sample size was relatively small (N = 327). Inadequate sample sizes can lead to improper solutions in confirmatory factor analyses, with the implication that our championed model may have inaccurate parameter estimates and fit statistics. However, although rules of thumb for sample size should be considered as nothing more than general guidelines – the actual required size is dependent on multiple sample and model characteristics - it is worth noting that our sample exceeds, or is close to, many of such guidelines e.g. $N \geq 200$ (Boomsma & Hoogland, 2001), N/number of variables ≥ 10 (Nunnally, 1967), and N/number of free parameters = 5-10 (Bentler & Chou, 1987). The second analysis of Study 2 was a series of Pearson correlations. While the sample size may give cause to question the results of the CFA, power calculations indicate that our study had sufficient statistical power (level of significance, α, set a .05 and β set at .20 resulting in a power of 80%) to detect correlations of a medium effect size ($r = |.30|$) (Cohen, 1988).

To strengthen the evidence that the Portuguese HPRS is a high-quality measure in adolescents, future studies may also wish to consider testing whether reactance is predictive of outcome variables relevant to this age group, such as school related processes and outcomes (e.g. academic performance, engagement with school and school absenteeism), pro-social / disruptive behaviours and emotional functioning. While we are aware of no studies that have explicitly tested these relations, we expect, for example, that some students will perceive the compulsory attendance of school and imposed workloads as a threat to personal freedom, and thus may make behavioral and cognitive efforts to reinstate their freedom. Demonstrating that total reactance, as measured by the Portuguese HPRS, is predictive of academic related outcomes will, therefore, not only offer evidence of predictive validity, but also have important implications for identifying students at risk of poor academic outcomes and informing interventions.

Conclusions

In summary, this research provides a synthesis of past research into the HPRS and highlights that evidence is converging on a single hypothesis – trait reactance is a unidimensional construct, but with some variation in HPRS scores accounted for by item wording/content. Most prior studies, however, used adults from just two individualist countries. By assessing competing factor models, we have shown a version of the HPRS adheres to this pattern of results when tested in adolescents from a different culture. Our results also demonstrated that high trait reactance is linked to decreased affective and cognitive wellbeing, supporting a conceptualization of reactance in terms of negative affect and negative cognition.

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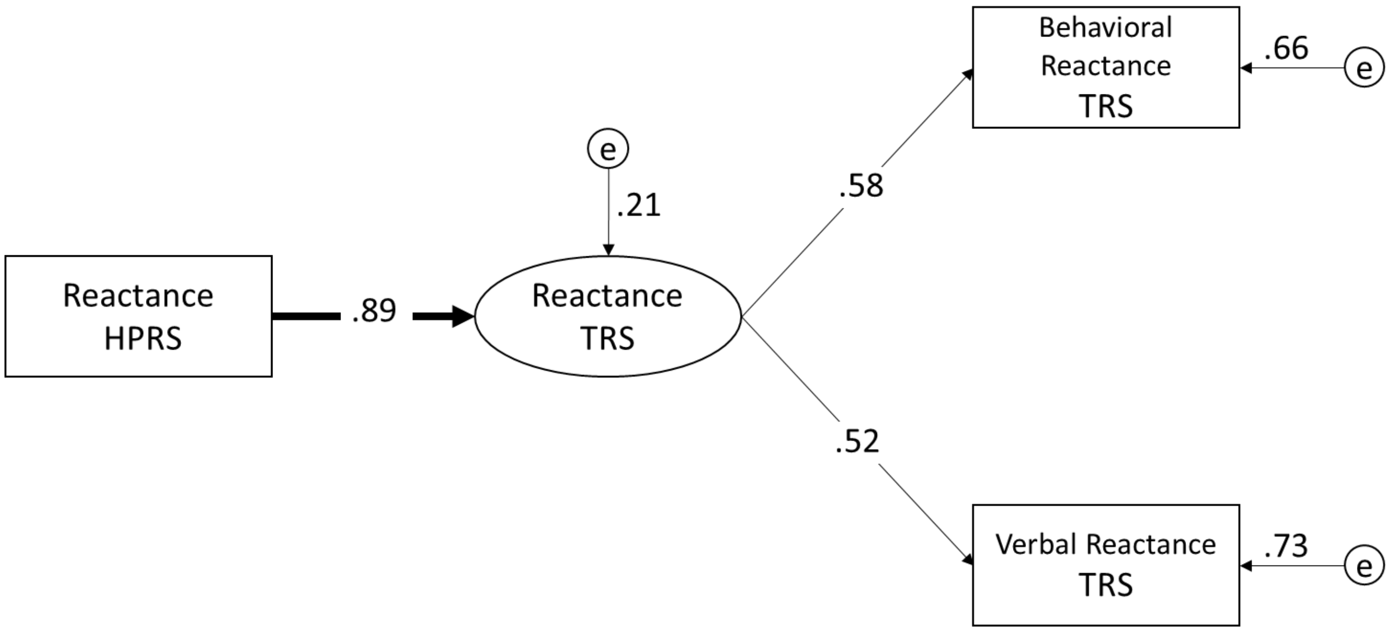

Figure 1. Outcome of SEM analysis. Boldface arrow indicates structural component of model. Values correspond to standardized coefficients. e = error. HPRS = Hong Psychological Reactance Scale. TRS = Therapeutic Reactance Scale.

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Table 1.

Study characteristics and extracted factor structures of past research investigating the psychometric properties of the HPRS.

Study	Model	Items	Sample Characteristics	Extracted factors and corresponding items			
Tucker & Byer (1987)	4-factor	18	College students	10, 13, 15, 12, 18, 7, 14, 9	6, 11, 4, 3, 2, 9	17, 18, 8, 2	1, 4, 2
Hong & Ostini (1989)	2-factor	18	Australian college students	4, 8, 9, 11, 12, 14, 16, 17, 18		5, 7, 10	
Hong & Page (1989)	4-factor	14	Australian undergraduates	4, 6, 8, 10	1, 2, 3	11, 12, 13, 14	5, 7, 9
Hong (1992)	4-factor	14	Australian adults	4, 6, 7, 10	1, 2, 3	12, 13, 14	5, 9, 10, 11
Hong & Faedda (1996)	4-factor	14	Australian undergraduates	4, 6, 7, 8	1, 2, 3, 14	10, 11, 12, 13	5, 9
	4-factor	11	Non-student adults	6, 7, 8	1, 2, 3	11, 12, 13	5, 9
Thomas et al. (2001)	4-factor	11	American undergraduates	6, 7, 8	1, 2, 3	11, 12, 13	5, 9
Shen & Dillard (2005)	4 first-order factors, 1 second- order factor	14	American undergraduates	6, 7, 8	1, 2, 3	11, 12, 13	5, 9
Jonason & Knowles (2006)	1-factor	18	NA	1, 2, 4, 6, 8, 9, 11, 12, 13, 14, 15, 16, 17, 18			
Jonason et al. (2010)	1-factor	18	American undergraduates	1, 2, 3, 6, 7, 10, 12, 13, 14, 15			
Brown et al. (2011)	4-factor	14	American undergraduates	4, 6, 7, 8	1, 2, 3, 14	10, 11, 12, 13	5, 9
	Bifactor	14		GF = 1, 2, 3, 4, 5, 6, 7, 8, 9, 10, 11, 12, 13, 14 SF_1 = 1, 2, 3, 14; SF_2 = 3, 9, 13; SF_3 = 4, 6, 7, 8			
De las Cuevas et al. (2014)	2 factor model	14	Spanish psychiatric outpatients 18+ years	4, 6, 7, 8, 12, 14		1, 2, 3, 5, 9, 10, 11, 13	
Yost & Finney (2018)	Bifactor model	14	American undergraduates	GF = 1, 2, 3, 4, 5, 6, 7, 8, 9, 10, 11, 12, 13, 14 SF_1 = 4, 6, 7, 8; SF_2 = 1, 2, 3, 14; SF_3 = 3, 9, 10; SF_4 = 5, 9			

1. Regulations trigger a sense of resistance in me
2. I find contradicting others stimulating
3. When something is prohibited, I usually think “that’s exactly what I am going to do”
4. The thought of being dependent on others aggravates me
5. I consider advice from others to be an intrusion
6. I become frustrated when I am unable to make free and independent decisions
7. It irritates me when someone points out things which are obvious to me
8. I become angry when my freedom of choice is restricted
9. Advice and recommendations induce me to do just the opposite
10. I am content only when I am acting of my own free will
11. I resist the attempts of others to influence me
12. It makes me angry when another person is held up as a model for me to follow
13. When someone forces me to do something, I feel like doing the opposite
14. It disappoints me to see others submitting to society’s standards and rules
15. When someone forces me to do something I say to myself: Now that’s exactly what I don’t want to do*
16. It pleases me to see how others submit to social norms and constraints*
17. Strong praise makes me sceptical*
18. I react negatively when someone tries to tell me what I should or should not do*

Note. *Items included in Merz's original trait reactance questionnaire, and also reincorporated by Jonason and colleagues, but not in the 14-item Hong Pyschological Reactance Scale; GF = General Factor; SF = Specific Factor; NA = Not Applicable

Table 2.

Model fit indices (*N* = 1,301).

Model	χ^2	*df*	χ^2/df	CFI	TLI	SRMR	RMSEA
1. One factor model	1328	77	17.25	.85	.83	.09	.11, CI [.11, .12]
2. Four factor correlated model*	1045	72	14.51	.89	.86	.09	.10, CI [.10, .11]
3. Higher-order model**	1199	74	16.20	.87	.84	.09	.11, CI [.10, .11]
4. Two factor correlated model	719	72	14.51	.89	.86	.09	.10, CI [.10, .11]
5. Bifactor Model: Anger + rules + independence	Model did not converge						
6. Incomplete Bifactor Model: Anger	966	72	13.42	.90	.87	.08	.10, CI [.09, .10]
7. Modified Incomplete Bifactor Model: Anger + opposite	366	67	5.46	.97	.95	.05	.06, CI [.05, .07]

Note. * = Model not admissible because covariance matrix of latent variables is not positive definite.
** = Model not admissible because of negative error terms for one factor. This is indicative of over-fitting a first-order factor that is not represented in the data. It is for this reason that *rules* and *independence* were not included in model 5. For all models, see Supplementary Figures 1 and 2 (Figures S1 and S2) for path diagrams.

Table 3.

Standardized (and fully standardized) factor loadings and unstandardized error terms for the Hong Psychological Reactance Scale items based on the bi-factor model (Model 7) (N = 1,301).

	Study 1				Study 2			
Item	λ_{GEN}	λ_{ANG}	λ_{OPP}	Error Variance	λ_{GEN}	λ_{ANG}	λ_{OPP}	Error Variance
(1) Regulations trigger a sense of resistance in me	1.00 (.42)			.83	1.00 (.17)			.97
(2) I find contradicting others stimulating	1.33 (.56)			.69	1.64 (.29)			.92
(3) When something is prohibited, I usually think "that's exactly what I am going to do"	1.15 (.48)		1.00 (.52)	.50	2.28 (.40)		1.00 (.48)	.61
(4) The thought of being dependent on others aggravates me	0.81 (.34)	1.00 (.47)		.66	1.41 (.25)	1.00 (.29)		.86
(5) I consider advice from others to be an intrusion	0.87 (.36)			.87	2.24 (.39)			.85
(6) I become frustrated when I am unable to make free and independent decisions	1.20 (.50)	1.01 (.48)		.52	2.95 (.51)	1.52 (.44)		.54
(7) It irritates me when someone points out things which are obvious to me	1.09 (.46)	0.47 (.22)		.74	2.66 (.46)	1.17 (.34)		.67
(8) I become angry when my freedom of choice is restricted	0.91 (.38)	1.47 (.70)		.37	3.08 (.54)	2.25 (.65)		.29
(9) Advice and recommendations induce me to do just the opposite	1.11 (.47)		1.39 (.72)	.27	2.99 (.52)		1.07 (.51)	.47
(10) I am content only when I am acting of my own free will	1.21 (.50)			.75	3.43 (.60)			.64
(11) I resist the attempts of others to influence me	0.30 (.13)			.98	2.64 (.46)			.79
(12) It makes me angry when another person is held up as a model for me to follow	0.90 (.38)	0.74 (.35)		.74	2.99 (.52)	0.99 (.29)		.65
(13) When someone forces me to do something, I feel like doing the opposite	1.31 (.55)		0.40 (.21)	.66	3.21 (.56)		0.54 (.26)	.62

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(14) It disappoints me to see others submitting to society's standards and rules	0.34 (.14)	0.97 (.46)		.77	2.18 (.38)	0.28 (.08)		.85

Note. λ_{GEN} = factor loadings on the general factor; λ_{ANG}= factor loadings on the anger specific factor; λ_{ANG}= factor loadings on the opposite specific factor

Table 4.

Pearson correlation coefficients with bootstrapped 95% confidence intervals between HPRS, TRS and indicators of wellbeing.

	Correlations with the HPRS			
	n	*r*	*p*	*95% CI*
TRS				
Behavioral Reactance	280	.50	<.001	[.40, .58]
Verbal Reactance	285	.36	<.001	[.25, .45]
Total Reactance	263	.50	<.001	[.41, .59]
Wellbeing				
Quality of Life	297	-.20	<.001	[-.30, -.09]
Life Satisfaction	302	-.22	<.001	[-.32, -.11]
Satisfaction with Social Support	290	-.19	.002	[-.29, -.07]
Composite Cognitive Wellbeing (Wellness) Index	247	-.26	<.001	[-.36, -.14]
Positive Affect	293	-.16	.007	[-.27, -.04]
Negative Affect	294	.21	<.001	[.10, .32]
Composite Affective Wellbeing (Happiness) Index	279	-.22	<.001	[-.33, -.11]
Composite Wellbeing Index	247	-.27	<.001	[-.38, -.15]

Note. $^{a}p < .10$, $^{*}p < .05$, $^{**}p < .01$, $^{***}p < .001$; HPRS = Hong Psychological Reactance Scale; TRS = Therapeutic Reactance Scale

Table S1.

Polychoric inter-item correlations, response frequencies, and descriptive statistics for the HPRS items in Study 1 (*N* = 1,301).

Item Number	1	2	3	4	5	6	7	8	9	10	11	12	13	14
1	1.00													
2	0.32	1.00												
3	0.18	0.28	1.00											
4	0.17	0.22	0.15	1.00										
5	0.11	0.21	0.30	0.05	1.00									
6	0.23	0.25	0.24	0.42	0.16	1.00								
7	0.15	0.24	0.20	0.19	0.25	0.30	1.00							
8	0.21	0.31	0.09	0.45	0.01	0.55	0.34	1.00						
9	0.20	0.28	0.59	0.13	0.40	0.22	0.26	0.10	1.00					
10	0.15	0.18	0.28	0.15	0.23	0.29	0.21	0.19	0.30	1.00				
11	-0.03	0.04	0.00	0.09	0.04	0.08	0.10	0.07	-0.01	0.11	1.00			
12	0.13	0.23	0.11	0.29	0.04	0.29	0.32	0.37	0.08	0.19	0.19	1.00		
13	0.21	0.20	0.37	0.24	0.19	0.29	0.23	0.24	0.39	0.26	-0.01	0.32	1.00	
14	0.08	0.13	0.02	0.29	-.07	0.25	0.19	0.35	-0.01	0.07	0.17	0.28	0.14	1.00
Frequencies														
1 (Completely Disagree)	132	297	431	114	253	36	65	61	331	87	80	68	84	93
2	265	470	475	213	565	132	188	141	614	328	109	102	248	156
3	650	407	318	353	468	294	374	367	370	497	344	270	480	552
4	406	289	216	453	196	655	560	633	167	391	590	460	444	435
5 (Completely Agree)	68	72	94	400	44	417	348	427	53	236	397	635	278	300
Missing	21	7	8	9	16	8	7	13	7	3	22	7	8	6
Mean	2.98	2.60	2.40	3.58	2.49	3.87	3.63	3.81	2.34	3.21	3.78	4.01	3.40	3.49
SD	0.96	1.12	1.18	1.19	0.97	0.95	1.07	1.06	1.02	1.11	1.05	1.09	1.10	1.08

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Table S2.

Invariance analysis of general factor loadings for the HPRS

Item	Specific Factor	λ_{GEN}	λ_{GEN}	λ_{GEN}	λ_{GEN}	λ_{GEN}
1		.42		.38	.46	
2		.56				.54
3	OPP	.48	.56	.52	.43	.46
4	ANG	.34	.25	.33	.31	
5		.36	.50	.40		
6	ANG	.50		.50	.50	.53
7	ANG	.46	.41	.44	.43	.44
8	ANG	.38	.24		.34	.42
9	OPP	.47	.68		.45	
10		.50	.49			.49
11		.13		.09		.16
12	ANG	.38			.26	
13	OPP	.55	.57	.57		
14	ANG	.14	.06	.13	.08	.19

Note. λ_{GEN} = fully standardized factor loadings on the general factor (last four columns refer to fully standardized factor loadings on the general factor in four different subsets of nine items): ANG = Anger specific factor; OPP = Opposite specific factor

Table S3.

Polychoric inter-item correlations and descriptive statistics for Study 2 based on the HPRS (*N* = 327).

Item Number	1	2	3	4	5	6	7	8	9	10	11	12	13	14
1	1.00													
2	.23	1.00												
3	.11	.23	1.00											
4	.05	-.03	.04	1.00										
5	.07	.22	.29	.22	1.00									
6	-.02	.15	.22	.20	.15	1.00								
7	.17	.14	.12	.22	.18	.40	1.00							
8	.09	.04	.18	.31	.06	.55	.46	1.00						
9	.07	.19	.44	.05	.37	.20	.26	.23	1.00					
10	.11	.18	.26	.05	.19	.38	.21	.34	.38	1.00				
11	-.05	-.04	.01	.25	.13	.22	.26	.36	.16	.25	1.00			
12	.05	-.10	16	.26	.13	.38	.31	.46	.17	.22	.37	1.00		
13	.08	.13	.34	.13	.24	.28	.26	.29	.41	.32	.18	.37	1.00	
14	.12	.18	.13	.18	.19	.22	.22	.23	.19	.16	.17	.23	.19	1.00
Frequencies														
1 (Completely Disagree)	19	34	78	28	40	13	13	7	47	9	15	8	21	16
2	52	85	96	41	106	23	27	22	110	49	28	27	47	24
3	150	130	89	81	122	74	95	69	110	113	102	74	99	144
4	80	65	49	98	38	145	129	147	43	103	113	100	103	88
5 (Completely Agree)	24	13	15	77	20	71	63	79	14	53	65	116	54	53
Missing	2	0	0	2	1	1	0	3	3	0	4	2	3	2
Mean	3.12	2.81	2.47	3.48	2.67	3.73	3.62	3.83	2.59	3.43	3.57	3.89	3.38	3.42
SD	.96	1.00	1.14	1.22	1.04	1.01	1.01	0.95	1.03	1.02	1.05	1.06	1.12	1.01

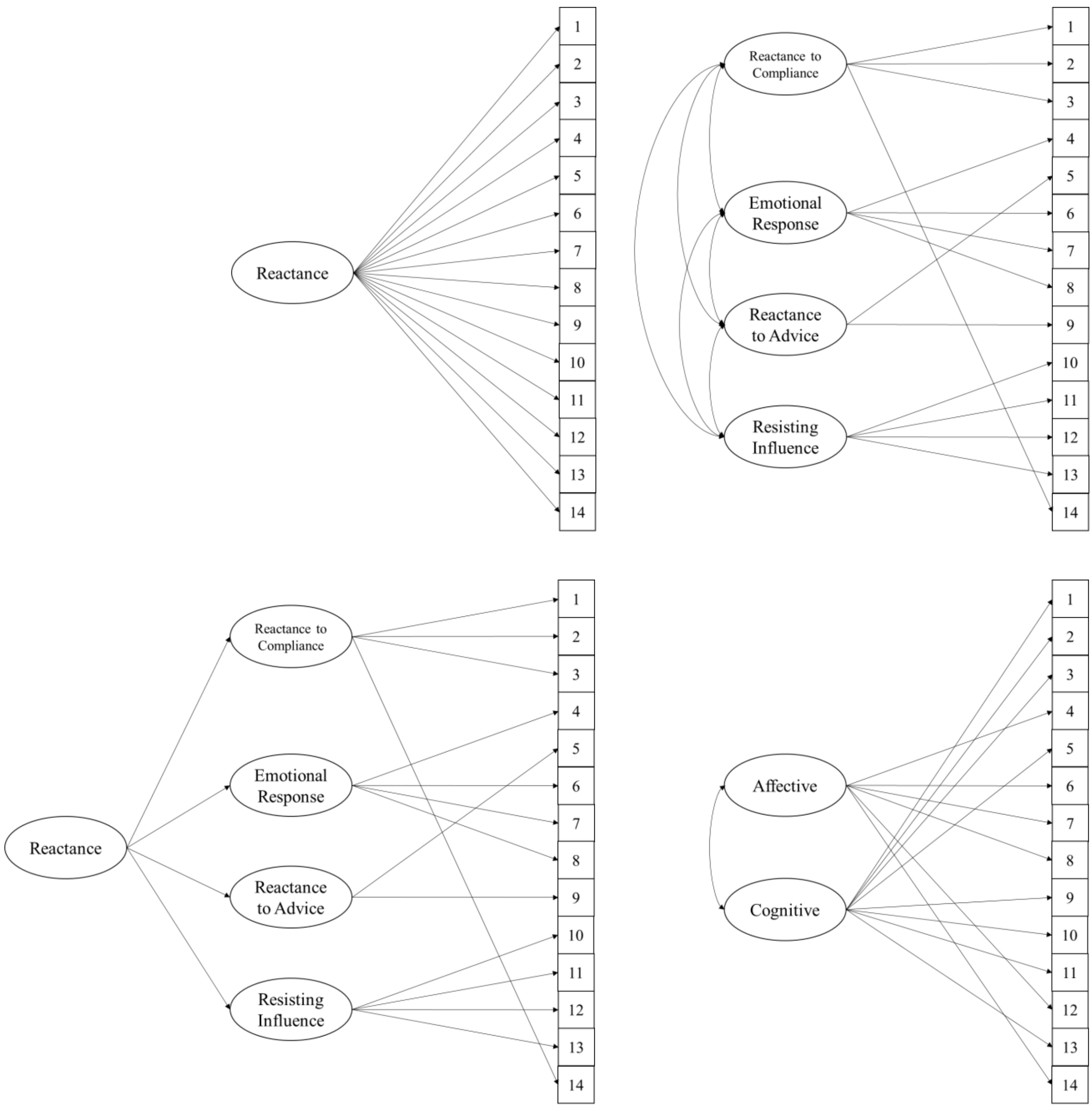

Figure S1. Models 1 – 4. Model 1 = one factor model. Model 2 = four factor correlated model. Model 3 = higher order model. Model 4 = two factor correlated model.

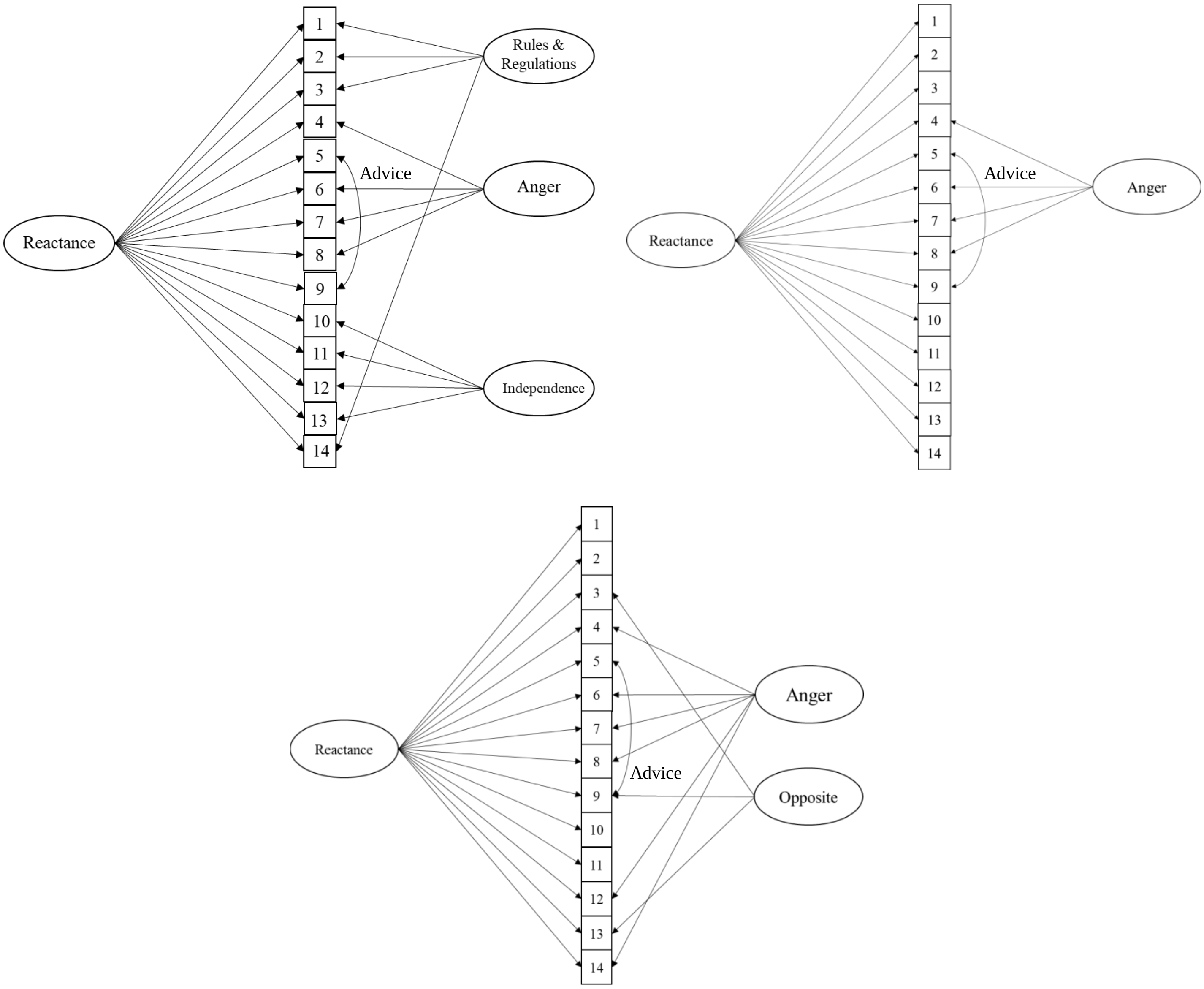

Figure S2. Bifactor models (models 5 – 8). Model 5 (top-left model) is the same as the Bifactor model tested by Brown, Finney, & France (2011). Models 6 (top right), and 7 (lower) are ancillary models. Model 7 had the best fit to our data. Note the specific factor *Advice* was modelled as a correlated error between items 5 and 9 and not a specific factor as otherwise the model would not be identified.