# Examining the Dimensionality of the Hong Psychological Reactance Scale

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**Abstract**

The Hong Psychological Reactance Scale (HPRS) purports to measure trait reactance (Hong & Page, 1989). Trait reactance refers to a propensity to experience psychological reactance, a motivational state that is aroused when a behavioral freedom is eliminated or threatened with elimination. To date, five studies have examined the psychometric properties of the HPRS, but reached different conclusions regarding its factor structure. The current study further investigated the factor structure and reliability of the scale, using confirmatory factor analysis to test four competing models. Four a priori models were tested: a one-factor model, four-factor model, higher-order model, and bifactor model. A modified bifactor model in which a general reactance factor explained common variance among all of the items and specific factors explained additional variance, among sets of items after controlling for reactance was championed. As such, it is recommended that researchers interested in relating general trait reactance to external criteria employ structural equation modeling (SEM). Implications for estimating reliability and scoring of the HPRS are discussed.

**Examining the Dimensionality of the Hong Psychological Reactance Scale**

Numerous researchers have argued that assessing noncognitive variables (e.g., attitudes, personality) is important in educational and workplace settings, as they influence people's experiences and outcomes (e.g., Kyllonen; Kyllonen, Walters, & Kaufman, 2005). For example, Kyllonen et al. asserted that the measurement of noncognitive constructs could contribute to three areas of higher education, specifically: (1) admissions, (2) programming (e.g., guidance, policies, and interventions), and (3) outcome evaluation. One such non-cognitive variable that could be useful to universities and organizations is psychological reactance.

**Defining Psychological Reactance**

*Psychological reactance* is a motivational state that functions to restore behavioral freedoms that are taken away or threatened with elimination (Brehm, 1966; Brehm & Brehm, 1981). For example, consider an employee, Jen, who received a request by her supervisor to work late one evening. Jen had already put in a full day, had plans with her friends that night, and consequently, felt angry and frustrated at the thought of having to cancel her night with friends. She felt a strong urge to meet with her friends anyway, so she decided not to work late. Jen's arousal in motivation aimed at restoring her ability to spend time with friends is an example of psychological reactance.

Psychological reactance theory (PRT) originated from the tradition of social psychology; thus, researchers initially conceptualized reactance to be situational. As such, researchers primarily focused on aspects of the situation that gave rise to reactance (Brehm, 1966; Brehm & Brehm, 1981). However, PRT has evolved to include the conceptualization of reactance as a trait representing one's propensity for reactance, independent of the situation (e.g., Brehm & Brehm; Dowd, Milne, & Wise, 1991). Said another way, individuals can vary in how "reactant" they are *in general*. From this perspective, trait reactance has been related to numerous individual difference variables such as trait anger (Hong & Giannakopoulos, 1993; Hong & Faedda, 1996), depression (Hong & Giannakopoulos; Hong & Faedda), aggression (Dowd & Wallbrown, 1993), happiness (Joubert, 1990), conformity (Goldsmith, Clark, & Lafferty, 1992), and several measures of personality (Buboltz, Woller, & Pepper, 1999; Dowd & Wallbrown; Dowd, Wallbrown, Sanders & Yesenosky, 1994).

To accurately assess the correlates of psychological reactance, a measure with sound psychometric properties is required. As such, researchers have created three measures of trait reactance: the Questionnaire for the Measurement of Psychological Reactance (QMPR; Merz, 1983 as cited in Tucker & Byers, 1987), the Therapeutic Reactance Scale (TRS; Dowd et al., 1991), and the Hong Psychological Reactance Scale (HPRS; Hong & Page, 1989; see Appendix). The QMPR was evaluated in several studies, all of which concluded that the scale was unstable and should not be used (Donnell, Thomas, & Buboltz, 2001; Hong & Ostini, 1989; Tucker & Byers). Consequently, the TRS and HPRS were both developed in response to the unmet need for a quality measure of trait reactance, but the TRS was developed for use with people in therapy and has only been examined in two studies (i.e., Dowd et al., 1991; Buboltz, Thomas, & Donnell, 2002). In contrast, the HPRS was developed for use with the general population, its psychometric properties have received considerable study, and it has been employed in several substantive studies (e.g., Dillard & Shen, 2005; Hellman & McMillin, 1995; Hong, 1990; Hong & Giannakopoulos, 1993; Hong, Giannakopoulos, Laing, & Williams, 2001; Hong & Langovski, 1994; Joubert, 1990; Joubert, 1992). For these reasons, the HPRS is the focus of the current study (see Appendix). To date, five studies have examined the dimensionality and reliability of the HPRS. Unfortunately, the researchers arrived at very different conclusions regarding the factor structure of the items and the scoring of the measure, which complicates the inferences made from these scores in the substantive studies.

**The Hong Psychological Reactance Scale: Empirical Studies of its Psychometric Properties**

**Hong and Page (1989).** In the first psychometric study of the scale, Hong and Page (1989) administered the HPRS to 257 university students in Australia. They used a principle components analysis (PCA) to study the dimensionality of the scale. It can be inferred that their intention was to identify interpretable latent constructs that explain relationships among the items. Unfortunately, the use of PCA for this purpose is inappropriate (Benson & Nasser, 1998; Preacher & MacCallum, 2003). Exploratory factor analysis (EFA), unlike PCA, recognizes and accounts for measurement error, which results in less biased parameter estimates than produced by PCA. Thus, it would have been more appropriate to use this technique, rather than PCA. Nonetheless, Hong and Page employed PCA with both orthogonal and oblique rotation. The researchers justified interpreting the orthogonal rotation, because the components produced by both rotations were "virtually identical" and the correlations between the components were "minimal" (Hong & Page, p. 1324). However, using orthogonal rotation forces the components to be uncorrelated, which is seldom appropriate (Preacher & MacCallum, 2003). Rather, they should have used oblique rotation because it allows the components to correlate. Unfortunately, Hong and Page never reported the correlations between the components, much less the solution using oblique rotation. Four components were labeled but not clearly defined: *Freedom from Choice*, *Conformity Reactance*, *Behavioral Freedom,* and *Reactance to Advice and Recommendations* (see Table 1). Hong and Page only reported internal consistency estimates for the total score ($\alpha$ = .77, $\omega$ = .82). Unfortunately, they did not report the reliabilities of the subscales defined by the four components, even though they found four “orthogonal” components of reactance.

**Hong (1992).** Given the need to study the functioning of the measure in a different population, Hong (1992) assessed the structure of the HPRS using 462 non-student participants from the general public in Australia. Unfortunately, Hong replicated the analytic procedure used by Hong and Page (1992; PCA with varimax rotation). Again, four components emerged and the same labels used in the previous study were applied. Interestingly, three items “loaded” on different components than in Hong and Page’s study (see Table 1). Specifically, item 10 (“I am content only when I am acting of my own free will”) moved from the *Freedom of Choice* to the *Reactance to Advice and Recommendations* component, item 11 (“I resist the attempts of others to influence me”) switched from the *Behavioral Freedom* to *Reactance to Advice and Recommendations* component, and item 7 (“It irritates me when someone points out things which are obvious to me”) shifted from the *Reactance to Advice and Recommendations* to the *Freedom of Choice* component, which indicates a lack of stability in structure across studies. Hong did acknowledge these differences, but curiously argued that the factor structures across the two studies were nearly the same. Again, Hong reported reliability estimates associated with the total score ($\alpha$ = .81, split-half = .76), but not for the subscale scores.

**Hong and Faedda (1996).** To re-examine the structure of the HPRS, Hong and Faedda (1996) administered the scale to a much larger sample than had been used in the past (1,425 university students as well as 1,660 non-student participants from the general public in Australia). The responses from the total sample ($N$ = 3,085) were analyzed via PCA and EFA with both orthogonal and oblique rotation. Hong and Faedda deemed the results from the PCA with varimax rotation as most appropriate because of the “meaningful and interpretable factor structure" (p. 176) as well as the ability to compare the current results to the previous two studies.

Four components emerged once again but some items loaded on other components and the researchers altered the component names to reflect these changes: *Reactance to Compliance*, *Resisting Influence from Others, Reactance Toward Advice and Recommendations*, and *Emotional Response Toward Restricted Choice* (see Table 1). For the first time, the researchers defined the components. *Reactance to Compliance* represents the tendency to experience reactance when expected to abide by the rules or desires of other people. This component was represented by four items, three of which previously represented *Conformity Reactance* in the two previous studies by Hong and her colleagues. *Resisting Influence from Others* represents the tendency to experience reactance when it appears as though others are trying to control one's behavior. This component was represented by four items that previously represented *Behavioral Freedom* and *Reactance Toward Advice and Recommendations* in previous studies. *Reactance Toward Advice and Recommendations* represents the tendency to experience reactance in response to advice and suggestions offered by other people concerning what one should do. The same two items consistently represented this component across studies. Interestingly, one component retained the same items, yet changed names. *Emotional Response Toward Restricted Choice* was represented by the four items that previously represented *freedom of choice* (Hong, 1992). *Emotional Response Toward Restricted Choice* refers the tendency to experience reactance when one is unable to make choices without other people interfering. It is unclear why Hong and Faedda changed the name of the *Freedom of Choice* component to *Emotional Response Toward Restricted Choice* given that they are represented by the same items. This change in structure and component labeling further complicates our understanding of the reactance construct and interpretation of reactance scores in substantive studies.

In addition, Hong and Faedda (1996) concluded that three items had non-salient loadings, complicating the interpretation of the factor structure: items 4 ("The thought of being dependent on others aggravates me"), 10 ("I am content only when I am acting of my own free will"), and 14 ("It disappoints me submitting to standards and rules"). These items were removed resulting in an 11-item version of the HPRS. Using the same sample as the initial analysis, the researchers conducted a PCA with varimax rotation to examine the factor structure of the remaining 11 items. Not surprisingly, a four-component structure emerged with no cross loadings, and paralleled the four components that emerged when a PCA was conducted using responses to all 14 items. The correlations among the four subscales ranged from .21 to .44. Hong and Faedda recommended using this 11-item version of the HPRS but did not clearly state how to score the measure. Furthermore, reliability was only estimated for the total score ($\alpha$ = .80 and .77 for the 14- and 11-item versions, respectively), which suggests that one "reactance" score should be computed. This is surprising given that they found empirical support for a multi-dimensional solution, suggesting that four subscale scores be computed instead.

In addition to their use of questionable data analytic techniques (i.e., PCA with varimax rotation), a further limitation of all three studies conducted by Hong and her colleagues is that reliability estimates for the total score were reported instead of the more justifiable computation of reliability estimates for each of the subscale scores. This is a problem because estimating Cronbach's coefficient alpha associated with the total score implies that a unidimensional solution underlies the responses (McDonald, 1999). However, the distinctiveness of the four components does not support the computation of a total score. If four factors underlie the reactance responses, these four scores, not a total score, should be used to profile a person's psychological reactance. Thus, the properties of these four scores (e.g., reliability, validity) should be reported. Fortunately, Thomas, Donnell, and Buboltz (2001) improved upon these limitations in their study of the HPRS.

**Thomas et al. (2001).** Thomas et al. (2001) were the first to explicitly note the inconsistencies found in the factor structure of the HPRS (e.g., items switching factors and factors switching names) in addition to Hong and colleague's seemingly inappropriate choice of varimax rotation. For instance, in response to Hong (1992), Thomas et al. stated, "even though some minor fluctuation in pattern coefficients may be expected to occur by chance, the number and magnitude of the differences between these two studies merit some concern" (p. 4). Thus, despite the consistency of uncovering four components, how to calculate reactance scores was still unclear because four different structures were produced. To address this issue, Thomas et al. used confirmatory factor analysis (CFA) to further test the competing models introduced by Hong and her colleagues (see Table 1). Both the 11- and 14-item versions of the scale were assessed using two independent American college student samples (*N* = 539 and 905, respectively). This study was important because in addition to using a more appropriate analytic technique (i.e., CFA), it was the first time the scale was studied using an American sample. Thomas et al. tested each of Hong's four models (see Table 1). Each model was evaluated twice in order to assess if the factors were correlated: once forcing the factors to be orthogonal and once allowing them to correlate. Adequate fit for the 14- and 11-item versions was only found when allowing the four factors from the Hong and Faedda model to correlate; no other models had adequate empirical fit. Factor correlations ranged from .29 and .72 for the 11-item scale, and .30 to .81 for the 14-item scale.

In addition to testing the four-factor models, the researchers also fit a second-order model to evaluate the appropriateness of creating a total reactance score. However, the second-order model did not adequately fit the data. Consequently, it would be most appropriate to score the HPRS as four subscales, not a total score. Unfortunately, Thomas et al. found that the reliabilities for the HPRS subscales ranged from .48 to .64 for the 11-item scale and .48 to .63 for the 14-item scale. These estimates are concerning but not surprising considering the low number of items representing each factor. Because a total score would not be appropriate to calculate (items are clearly multidimensional) and the subscales had low reliabilities, Thomas et al. cautioned against using the scale.

**Shen and Dillard (2005).** Shen and Dillard (2005) conducted the most recent study of the structure of the HPRS. They administered the HPRS to three different samples of American undergraduate students. The first two samples ($N$ = 188 and 200) completed the 14-item version of the HPRS, and the third sample ($N$= 233) completed the 11-item version. Shen and Dillard only reported results for the 11-item version, even though the additional three items may have improved reliability. Nonetheless, they fit a correlated four-factor model to the 11-item version of the scale and found that it fit adequately across the three samples. Again, factor correlations ranged from .45 to .76, which further supports the appropriateness of oblique rotation over orthogonal rotation.

In addition to fitting a four-factor model, Shen and Dillard (2005) also fit a higher-order model and reported adequate fit. The reliabilities for the total score across the three samples were .75, .80, and .79, which the authors deemed adequate. Thus, Shen and Dillard recommended calculating a total score. Being able to calculate the total score, in place of four subscale scores, would avoid problems with the low reliabilities associated with the subscale scores (which ranged from .45 to .71 in their study).

**The Need for Further Study of the Hong Psychological Reactance Scale**

There are several competing and conflicting recommendations regarding the scoring and use of the HPRS. Thomas et al. (2001) found that the 11-item and 14-item four-factor structure interpreted by Hong and Faedda (1996) best fit the data, but only when the factors were allowed to correlate, thus supporting the use of four subscale scores. However, the reliabilities of the subscales were too low, and Thomas et al. concluded that the scale should not be used. In contrast, Shen and Dillard (2005) were able to fit a higher-order model to the data, enabling practitioners to justify the computation and use of a reliable total reactance score.

**Purpose of the Current Study and Hypothesized Models**

Given the conflicting findings and recommendations concerning the HPRS, additional study of the factor structure is needed. Moreover, we feel that one model, which has not been tested, is very plausible: the bifactor model.

The bifactor model is similar to the second-order model in that both model an overall factor. In a second-order model, this overall factor (i.e., higher-order factor) accounts for the relationship among the lower-order factors. In a bi-factor model, however, this overall factor accounts for relationships among individual items (akin to a single-factor model) and is termed a *general factor*. In addition to this general factor, there are also *specific factors* that account for unique variance among the items *over and above* the general factor; as such, the specific factors are uncorrelated with the general factor (Chen, West, & Sousa, 2006). Bifactor models can be useful when it is hypothesized that (1) there is a general factor that can account for the shared variance among a set of items, (2) there are sets of items that share variance beyond what can be explained by the general factor (i.e., specific factors), and (3) there may be external variables that are differentially related to the general factor and specific factors (Chen et al., 2006).

In reviewing how the HPRS was developed, trait reactance was never theorized to be multidimensional. In fact, it is more plausible that trait reactance is unidimensional and the factors that have emerged are simply due to the specific nature of their context or wording. For example, two questions ask about reactance in a specific situation that involves receiving advice: “I consider advice from others to be an intrusion” and “Advice and recommendations usually induce me to do just the opposite.” In addition, some items ask participants about emotions they experience (e.g., “The thought of being dependent on others aggravates me” and “I become angry when my freedom of choice is restricted”). Although several items may share a similar context or wording, it seems as though each question was written to measure a unidimensional conceptualization of trait reactance; whether the dimensions found by researchers are theoretically meaningful dimensions of trait reactance still needs to be established. Moreover, substantive researchers have consistently used a total score regardless of the empirical factor structure that has emerged (Dillard & Shen, 2005; Hellman & McMillin, 1995; Hong, 1990; Hong & Giannakopoulos, 1993; Hong et al., 2001; Hong & Langovski, 1994; Joubert, 1990; Joubert, 1992). Consequently, researchers appear to have interpreted the HPRS as measuring general trait reactance, which further supports testing a bifactor model of trait reactance.

The purpose of the current study is to test four models representing the factor structure of the HPRS: a single-factor model, four-factor model, second-order model, and bifactor model (see Figure 1). Although a single-factor model was never tested, much less supported in previous psychometric study of the scale, Hong and her colleagues reported reliability estimates for the total score, and most substantive studies employing the HPRS have used a total score (Dillard &

Shen, 2005; Hellman & McMillin, 1995; Hong, 1990; Hong, 1992; Hong & Faedda, 1996; Hong & Giannakopoulos, 1993; Hong et al., 2001; Hong & Langovski, 1994; Hong & Page, 1989; Joubert, 1990; Joubert, 1992), implying unidimensionality. In addition to a single-factor model, we will also test the four-factor model proposed by Hong and Faedda and supported by Thomas et al. (2001), as well as the second-order model championed by Shen and Dillard (2005). We believe, however, that a bifactor model will best represent the relationships among the items. That is, we believe that all 14 items of the HPRS represent a general factor (reactance), yet sets of items share additional variance over and above the general factor of reactance due to item wording and context. These sets of items align with the four "factors" found by Hong and Faedda (1996) and later supported by Thomas et al. (2001) in that they explain variance among the same groups of items. However, a critical point to be made is that these specific factors are *not* the same as the first-order factors found in earlier studies. This is because these specific factors represent common variance after controlling for reactance, whereas the first order factors represent dimensions of reactance. In other words, specific factors do not describe dimensions of trait reactance; they are orthogonal to trait reactance.

For this reason, the specific factors will not share the same name as the first-order factors that previously represented the common variance among the same sets of items. That is, in addition to measuring reactance, the *Emotional Response Toward Restricted Choice* items also seem to represent anger and include words such as "aggravates," "frustrated," "irritates," and "angry." As such, the specific factor representing these items will be labeled *Anger*. The specific factor for the *Reactance To Compliance* items will be called *Rules* because of the similar wording and context having to do with rules and regulations. Likewise, the additional variance shared between the two *Reactance Toward Advice and Recommendations* items will be called

*Advice* because both items refer to attitudes toward receiving advice and recommendations. Finally, it is possible that the *Resisting Influence from Others* items share additional variance, beyond reactance, because each deals with a feeling of being independent. Thus, this specific factor will be called *Independence.*

## Methods

### Participants and Procedures

The HPRS was administered to first-year students at a Mid-Atlantic, mid-sized, public university as part of a required university-wide assessment day. Students were randomly assigned to various batteries of tests (some of which included the HPRS), which were administered by trained proctors. Of the 1,286 participants completing the HPRS, 61.6% were female, 79.2% were White non-Hispanic, 4.6% were Asian, 3.4% were African American, 2.2% were Hispanic, and 8.6% did not specify their ethnicity, and the average age was 18.42 (*SD* = 0.39). These demographics are representative of the population at this university.

### Instrument

Hong and Page (1989) developed the HPRS scale by refining an English version of Merz's Psychological Reactance Scale (originally in German). Specifically, they refined several of Merz's items and created new ones, resulting a pool of 60 items. These items were subsequently evaluated according to relevance, clarity, and semantics, with 15 items selected as the most representative of trait reactance. Hong and Page then presented the 15 items to nine behavioral scientists for further evaluation. One item was removed due to redundancy, resulting in the 14-item Hong Psychological Reactance scale (see Appendix). Participants indicated the extent to which they endorsed each statement on a five-point scale (1 = strongly disagree to 5 = strongly agree).

# Results

## Descriptive Statistics

Before conducting the CFA analyses, data were screened for outliers and non-normality. Using Mahalanobis distances and an examination of response patterns, four multivariate outliers were removed due to response sets, resulting in an effective sample size of 1,282 participants. The properties of the data indicated both univariate and multivariate normality, and as a result, maximum likelihood estimation was employed for the CFA analyses (Finney & DiStefano, 2006; West, Finch & Curran, 1995). Correlations, means, standard deviations, skew, and kurtosis for the 14 items are reported in Table 2. Item correlations ranged from .08 to .49; thus, there were no issues of multicollinearity. There was also no indication of a ceiling or floor effect. That is, item means tended to fall toward the middle of the response scale and standard deviations were close to 1.

## Confirmatory Factor Analysis

**Criteria for Assessing Model-Data Fit.** Both absolute and incremental fit indices were utilized to assess global model-data fit. Absolute indices simply consider how well the model accounts for observed covariances in the data (Hu & Bentler, 1995), whereas incremental fit indices consider the improvement in fit of the hypothesized model over a null baseline model. Based upon work examining the sensitivity of various fit indices (Hu & Bentler, 1998, 1999), the comparative fit index (CFI), the root mean square error of approximation (RMSEA), and the standardized root mean square residual (SRMR) were used to assess model fit.

In addition to assessing global fit, local areas of misfit were also examined in order to diagnose specific areas of misfit, and thus help uncover the structure of the items. This was done by examining the standardized covariance residuals, which indicate how well the model reproduces the bivariate relationships among each individual pair of items. Large positive values (reported on a *z*-score metric) indicate that the relationship between the pair of items is not well-reproduced by the model (i.e., underestimated), implying that the items share variability after controlling for the latent constructs being modeled (construct irrelevant variance).

**Testing Hypothesized Models.** CFA analyses were conducted with LISREL 8.72 (Jöreskog & Sörbom, 1993), and the variance/covariance matrix of the scores was used as input. Results (see Table 3) indicated that two of the three indices associated with the one-factor model did not meet recommended standards (i.e., CFI ≥ .95, RMSEA ≤ .06, SRMR ≤ .08; Hu & Bentler, 1999). Moreover, ten standardized covariance residuals were above a value of 5.0. Thus, the one-factor model did not represent the data adequately.

The four-factor model fit significantly better than the one-factor model ($\Delta\chi^2(6)= 486.51$, $p < .001$). The lack of relative fit of the one-factor model compared to the four-factor model was not surprising given the range of correlations among the four factors (.48 to .81); the scores were clearly not unidimensional in nature. Although the various fit indices associated with the four-factor model indicated superior fit compared to the one-factor model, the CFI was less than the recommended .95 (CFI = .92) and four standardized covariance residuals were above a value of 5.0. Thus, we did not find support for the correlated four-factor Hong and Faedda (1996) model.

Given that the higher-order model is more parsimonious than the four-factor model, we knew the higher-order model would result in less than adequate fit. Interestingly, the higher-order model did not converge to an admissible solution. When examining the results, the disturbance term for the RIO first-order factor was negative (Heywood case). This gave us insight into possible model misspecification. As clearly explained by Chen, West, and Sousa (2005), if a first-order factor, such as *Resisting Influence from Others*, reflects only the general second-order factor (reactance in this case), this can be manifested in a low or nonsignificant disturbance term. That is, the data are being over-factored by forcing a first-order factor that is not represented by the data (*Resisting Influence from Others* in this case). Over-factoring can result in inadmissible solutions (e.g., Heywood cases), in addition to low or nonsignificant parameter estimates (Rindskopf, 1984).

Given the misfit of the four-factor model and the inadmissible solution due to over-factoring for the higher-order model, at this point, we knew the HPRS could not be scored or interpreted as previous studies suggested. That is, neither a unidimensional nor higher-order solution was supported, suggesting that the summation of items to create a total score is not supported by the empirical structure underlying the scores. Moreover, the four-factor model specified by Hong and Faedda (1996) did not adequately align with the relationships between the items. However, a bifactor model was still a plausible representation of the inter-item relationships.

Fitting the bi-factor model resulted in a CFI that was slightly less than the recommended .95 (CFI = .93) and two standardized covariance residuals that were above a value of 5.0. As expected based on the higher-order results, none of the *Resisting Influence from Others* items had significant factor loadings on the *Independence* factor. This is an interesting finding, in that it implies that the *Independence* factor does not exist as a domain specific factor after controlling for variance due to the general reactance factor (Chen et al., 2006). In other words, these items share no common variance over and above what is shared with the general reactance factor. Thus, although the originally specified bifactor model resulted in the most adequate fit of the completing models, the model was not satisfactory.

**Testing Ancillary Models.** The original bi-factor model was modified to create an incomplete bifactor model that removed the *Independence* factor (Figure 2; Chen et al., 2006). Although the fit of the incomplete bi-factor model appeared promising, it was not adequate. An examination of the standardized covariance residuals indicated that items 3 and 9 (residual = 7.56) and items 3 and 13 (residual = 5.37) shared variance after controlling for the general reactance factor and the domain specific factors. Of note, all three of these items focused on doing the opposite ("When something is prohibited, I usually think, 'That's exactly what I am going to do';" "Advice and recommendations usually induce me to do just the opposite;" "When someone forces me to do something, I feel like doing the opposite"). It appeared that there was shared variance due to item wording or the extreme nature of these items. Thus, we modeled these relationships via an *Opposite* factor (see Figure 2). The incomplete bifactor model with the *Opposite* factor fit the data well and significantly better than the incomplete bifactor model ($\Delta\chi^2(3) = 100.94$, $p < .001$).

**Parameter Estimates and Reliability.** Given the adequate fit of the modified incomplete bi-factor model to the 14-items, the unstandardized coefficients, standardized coefficients, and error terms were examined (see Table 4). All unstandardized paths were significant ($p < .05$), as was the error covariance representing the domain specific *Advice* items. The majority of the items had higher standardized coefficients associated with the general reactance factor than the specific factor for which the item served as an indictor. We believe this supports the existence of an overall general reactance factor. However, for all but two items (items 3 and 6), less than 50% of the item's variance was explained by the model, suggesting random variation (i.e., unreliability) or unique systematic variance associated with something other than the factors being modeled (e.g., variance associated with another construct, wording issues, or method variance; DeShon, 1998; Jöreskog, 1993). In other words, although the incomplete bifactor model fit responses from the 14-item HPRS, some of the items had a large amount of unexplained variance.

By championing the incomplete bifactor model, we support the modeling of a general trait reactance factor, partialling out the effects of the specific factors. Unfortunately, there is not a straightforward way to calculate an observed total score to represent this general trait reactance factor from the bifactor model. This is because the total score as a measure of general reactance is contaminated with the effects of the specific factors. In contrast, the general reactance factor is "pure" because the effects of other systematic sources of variance are explained by the specific factors. As such, the reliability of this latent factor is of greatest interest. The reliability of a latent factor (i.e., construct reliability) can be estimated via Coefficient H, which represents the amount of variance among the items explained by the latent factor (Hancock & Mueller, 2001). In the current study, Coefficient H is .849, which is considered adequate.

That said, if one were to insist on using a total HPRS score, it would be important to report how well that total score represents the purified general reactance factor. Traditionally, coefficient alpha has been used as a reliability index of composite scores. However, because the items of the HPRS are "factorially complex," it does not make sense to estimate reliability as a function of all systematic variance, especially because not all of that variance is due to reactance. As such, alpha tends to overestimate reliability in such cases because it includes systematic variance from all the factors in the model instead of only the factor of interest. It is recommended that only the variance due to the factor of interest be included as systematic variance, as given by $\omega_H$ (McDonald, 1999; Zinbarg, Revelle, Yovel, 2007; Zinbarg, Revelle, Yovel, & Li, 2005). Specifically, $\omega_H$ is equivalent to the squared correlation between the total score and the factor score and represents how well the total score reflects the latent factor score. For the current study, $\omega_H$ is .78.

**Discussion**

The goal of the current study was to address the need for additional examination of the HPRS factor structure given the conflicting conclusions regarding its scoring and use (e.g., Shen & Dillard, 2005; Thomas et al., 2001). As such, we tested four hypothesized models as well as two ancillary models. The current research contributes to the study of the HPRS in five ways: (1) a one-factor model was tested and rejected, which had not been done in prior research, (2) the factor models proposed in previous studies were rigorously tested and compared against one-another, none of them fitting adequately, (3) a bifactor model was proposed and tested, a model that had not been considered until now, (4) the *Independence* factor (representing the RIO items) was not found to be distinct from general trait reactance in the bifactor model, suggesting that this is not a unique factor, and (5) a new, *Opposite* factor was identified and found to improve the fit of the model to the data.

The current study went beyond testing the models supported in prior research by proposing a bifactor model to represent the relationships among the items. The bifactor model fit well globally, however, one specific factor (*Independence*) did not explain any additional variance beyond general trait reactance, indicating that this factor fails to exist after controlling for general trait reactance; the common variance shared by these items is almost entirely explained by the general reactance factor. This finding from the bifactor model also explains the negative disturbance term associated with the *Resisting Influence from Others* first-order factor in the higher-order model. The *Resisting Influence from Others* factor correlated highly with the second-order factor due to a lack of specific systematic variance shared across these items (both the correlation with the second-order factor and the disturbance term were inadmissible values). This aligns with the specification of the higher-order versus the bifactor model because the disturbance term of the *Resisting Influence from Others* first-order factor in the higher-order model is similar to the *Independence* specific factor in the bifactor model. More specifically, both represent variance in the same set of items unrelated to the general reactance factor. Although the higher-order and bifactor results lead to the same conclusion regarding the *Resisting Influence from Others* and *Independence* factors, in our opinion, the bifactor model more clearly indicated that these items shared no common variance over and above general trait reactance. That is, the non-salient factor loadings on the *Independence* factor of the bifactor model were much easier to interpret than the inadmissible solution produced by the higher-order model. This highlights the utility of the bifactor model for assessing if the data are being over-factored (i.e., if certain factors are not needed to explain inter-item relationships; Chen et al., 2006).

In addition, a new *Opposite* factor was specified represented the relationship among items that referenced a desire to do the opposite. However, whether this specific factor represents a method effect (i.e., wording or item extremeness) or a substantive construct still needs to be determined via relationships between this factor and external criteria. An examination of the distributions suggests that two out of the three items (3 and 9) appear to be slightly positively skewed with more people responding that they disagree with these items, which may indicate that these items share variance because they are more extreme representations of reactance. In addition to the *Opposite* factor, the other three specific factors (*Anger, Rules and Regulations,* and *Advice*) also need to be related to external variables to evaluate whether they represent method effects due to similar wording or context, or whether they represent substantively meaningful constructs.

**How does one score the HPRS?**

Although the current study established a model that appears to adequately represent the inter-item relationships, scoring the HPRS in a way that aligns with the bifactor model becomes a challenge. One might rationalize that calculating a total observed score is justifiable given that the bifactor model includes an overall, general trait reactance factor. However, computing a total observed score implies that the items are unidimensional, which has not been empirically supported. That is, calculating a total score confounds the variance associated with the general and specific factors. On the other hand, we did not find support for the computation of four subscale scores of reactance. Unlike Thomas et al. (2001) who supported the four-factor model and suggested that the HPRS be scored by reporting subscale values instead of a total score, we rejected the correlated four-factor model. If one wishes to relate general trait reactance to external criteria, we recommend employing structural equation modeling (SEM) to facilitate modeling a “purified” general reactance factor. The use of SEM also would allow the estimation of relationships between the specific factors and external criteria providing insight into what these specific factors represent.

As previously mentioned, theoretically-related external variables should be examined in relation to both general trait reactance as well as specific factors in future research to begin to give them meaning (Benson, 1998; Chen et al., 2006). If the general and specific factors related differentially to various constructs, this would further emphasize that the computation and use of a total reactance score would result in biased relationships with external criteria. For example, say the *Anger* specific factor had a strong relationship with trait anger. If one were to create a total observed reactance score, the total score would contain this variance due to trait anger and thus, the correlation between total observed reactance and trait anger would be inflated compared to the relationship between trait reactance and general trait reactance factor from the bi-factor model. Consequently, the relationships between the HPRS total score and external criteria are difficult to interpret because these total scores consist of systematic variance representing trait reactance as well as variance due to factors *unrelated* to trait reactance. Thus, a strong program of research investigating the relationship between the general trait factor and criteria is needed.

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Table 1

*Factor Structure of Hong Psychological Reactance Scale Across Studies (Adapted from Thomas et al., 2001)*

| Model | Factor 1 items | Factor 2 items | Factor 3 items | Factor 4 items |
|---|---|---|---|---|
| Model 1: Hong & Page (1989) | 4, 6, 8, 10<br>Freedom of Choice | 1, 2, 3<br>Conformity Reactance | 11, 12, 13, 14<br>Behavioral Freedom | 5, 7, 9<br>Reactance to Advice & Recommendations |
| Model 2: Hong (1992) | 4, 6, 7, 8<br>Freedom of Choice | 1, 2, 3<br>Conformity Reactance | 5, 9, 10, 11<br>Reactance to Advice & Recommendations | 12, 13, 14<br>Behavioral Freedom |
| Model 3: Hong & Faedda (14 items; 1996) | 4, 6, 7, 8<br>Emotional Response toward Restricted Choice | 1, 2, 3, 14<br>Resistance to Compliance | 10, 11, 12, 13<br>Resisting Influence from Others | 5, 9<br>Reactance toward Advice & Recommendations |
| Model 4: Hong & Faedda (11 items; 1996) | 6, 7, 8<br>Emotional Response toward Restricted Choice | 1, 2, 3<br>Resistance to Compliance | 11, 12, 13<br>Resisting Influence from Others | 5, 9<br>Reactance toward Advice & Recommendations |

Table 2

*Item Correlations and Descriptive Statistics (N = 1,282)*

| Item | 1 | 2 | 3 | 4 | 5 | 6 | 7 | 8 | 9 | 10 | 11 | 12 | 13 | 14 |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| 1 | 1.00 | | | | | | | | | | | | | |
| 2 | 0.38 | 1.00 | | | | | | | | | | | | |
| 3 | 0.39 | 0.40 | 1.00 | | | | | | | | | | | |
| 4 | 0.24 | 0.26 | 0.20 | 1.00 | | | | | | | | | | |
| 5 | 0.17 | 0.28 | 0.27 | 0.21 | 1.00 | | | | | | | | | |
| 6 | 0.26 | 0.28 | 0.19 | 0.46 | 0.23 | 1.00 | | | | | | | | |
| 7 | 0.24 | 0.32 | 0.15 | 0.32 | 0.31 | 0.40 | 1.00 | | | | | | | |
| 8 | 0.28 | 0.26 | 0.20 | 0.30 | 0.15 | 0.49 | 0.38 | 1.00 | | | | | | |
| 9 | 0.26 | 0.30 | 0.44 | 0.18 | 0.48 | 0.21 | 0.30 | 0.21 | 1.00 | | | | | |
| 10 | 0.27 | 0.30 | 0.31 | 0.25 | 0.34 | 0.39 | 0.30 | 0.40 | 0.42 | 1.00 | | | | |
| 11 | 0.10 | 0.16 | 0.08 | 0.20 | 0.19 | 0.26 | 0.17 | 0.19 | 0.22 | 0.30 | 1.00 | | | |
| 12 | 0.20 | 0.29 | 0.26 | 0.22 | 0.33 | 0.30 | 0.34 | 0.31 | 0.32 | 0.40 | 0.29 | 1.00 | | |
| 13 | 0.28 | 0.35 | 0.42 | 0.27 | 0.29 | 0.32 | 0.32 | 0.33 | 0.42 | 0.40 | 0.23 | 0.40 | 1.00 | |
| 14 | 0.29 | 0.32 | 0.42 | 0.19 | 0.38 | 0.24 | 0.21 | 0.28 | 0.39 | 0.39 | 0.21 | 0.37 | 0.37 | 1.00 |
| Mean | 2.96 | 2.62 | 2.07 | 3.39 | 2.05 | 3.37 | 3.44 | 3.53 | 2.01 | 2.75 | 2.90 | 2.61 | 2.69 | 2.34 |
| *SD* | 1.06 | 1.12 | 1.00 | 1.09 | 0.93 | 1.09 | 1.12 | 1.02 | 0.90 | 1.03 | 1.04 | 1.04 | 1.09 | 1.03 |
| Skew | -0.13 | 0.19 | 0.79 | -0.27 | 0.79 | -0.31 | -0.44 | -0.48 | 0.81 | 0.16 | 0.09 | 0.35 | 0.23 | 0.54 |
| Kurt | -0.44 | -0.79 | 0.13 | -0.67 | 0.37 | -0.63 | -0.53 | -0.20 | 0.68 | -0.42 | -0.50 | -0.35 | -0.64 | -0.12 |

Table 3

*Fit Statistics for the Alternative Reactance Models with 14 items*

| Model | $\chi^2$ | df | CFI | RMSEA | SRMR |
|---|---|---|---|---|---|
| *Hypothesized Models* | | | | | |
| One-factor Model | 930.83 | 77 | .82 | .09 | .063 |
| Four-factor Model | 444.32 | 71 | .92 | .06 | .044 |
| Higher-order | Didn't converge | | | | |
| Bi-factor Model | 401.20 | 64 | .93 | .06 | .041 |
| *Ancillary Models* | | | | | |
| Incomplete Bi-factor Model | 421.38 | 68 | .93 | .06 | .042 |
| Incomplete Bi-factor with "Opposite" specific factor | 320.44 | 65 | .95 | .06 | .037 |

*Note*. CFI = comparative fit index; RMSEA = root mean square error of approximation; SRMR = standardized root mean square residual. *N* = 1282. **p* < .001.

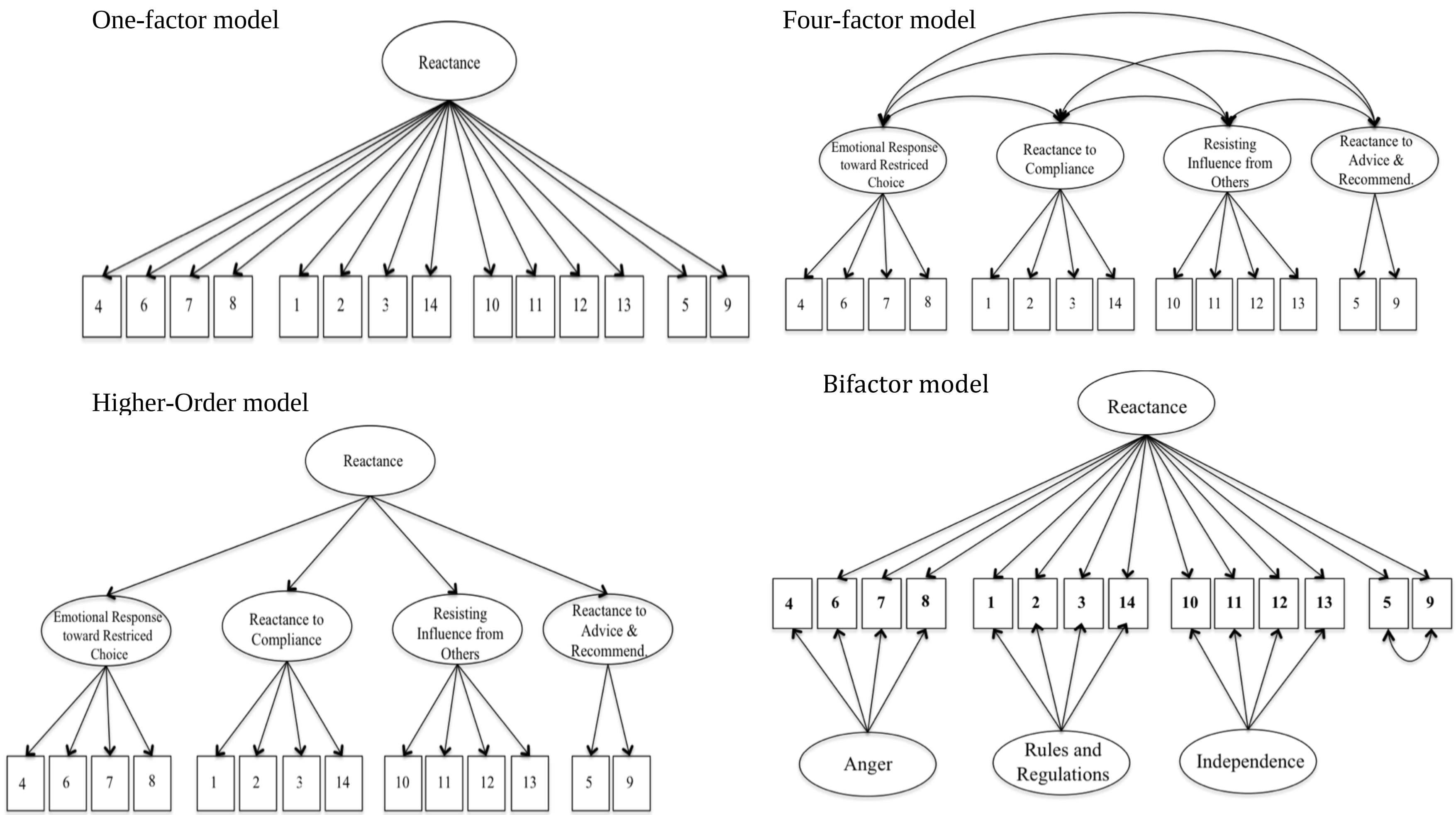
Table 4

*Completely Standardized Pattern Coefficients and Error Terms from Modified Incomplete Bi-factor Model*

| Items | General Factor | Rules and Regulations | Anger | Opposite | Error Terms ($1-R^2$) |
|---|---|---|---|---|---|
| 1 | .42 | .36 | | | .70 |
| 2 | .50 | .33 | | | .64 |
| 3 | .44 | .57 | | .58 | .14 |
| 14 | .58 | .22 | | | .62 |
| 4 | .40 | | .39 | | .69 |
| 6 | .51 | | .65 | | .32 |
| 7 | .51 | | .24 | | .68 |
| 8 | .51 | | .35 | | .62 |
| 9 | .57 | | | .28 | .59 |
| 13 | .63 | | | .23 | .55 |
| 5 | .51 | | | | .74 |
| 10 | .67 | | | | .55 |
| 11 | .40 | | | | .84 |
| 12 | .61 | | | | .63 |

*Note*. The *Advice* domain specific factor (representing variance among the *Reactance to Advice and Recommendations* items) could not be modeled given only two indicators (items 5 and 9) and no correlations among factors. Thus, the standardized correlated error term represented the relationship between these two items after controlling for the general reactance factor and the opposites factor ($r$ = .18).

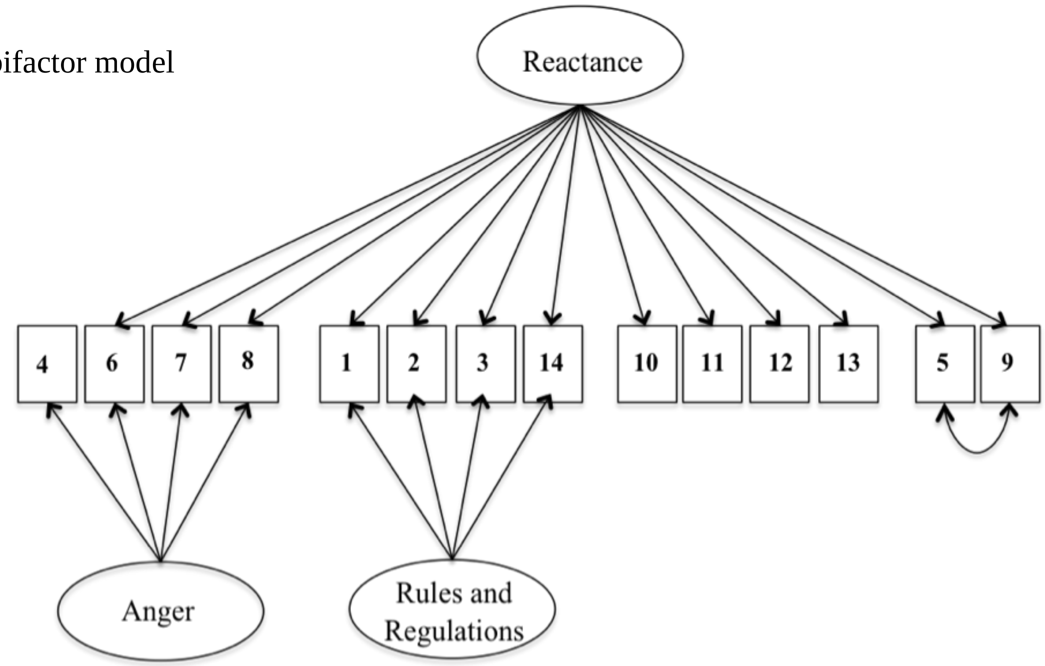
*Figure 1.* Tested Models



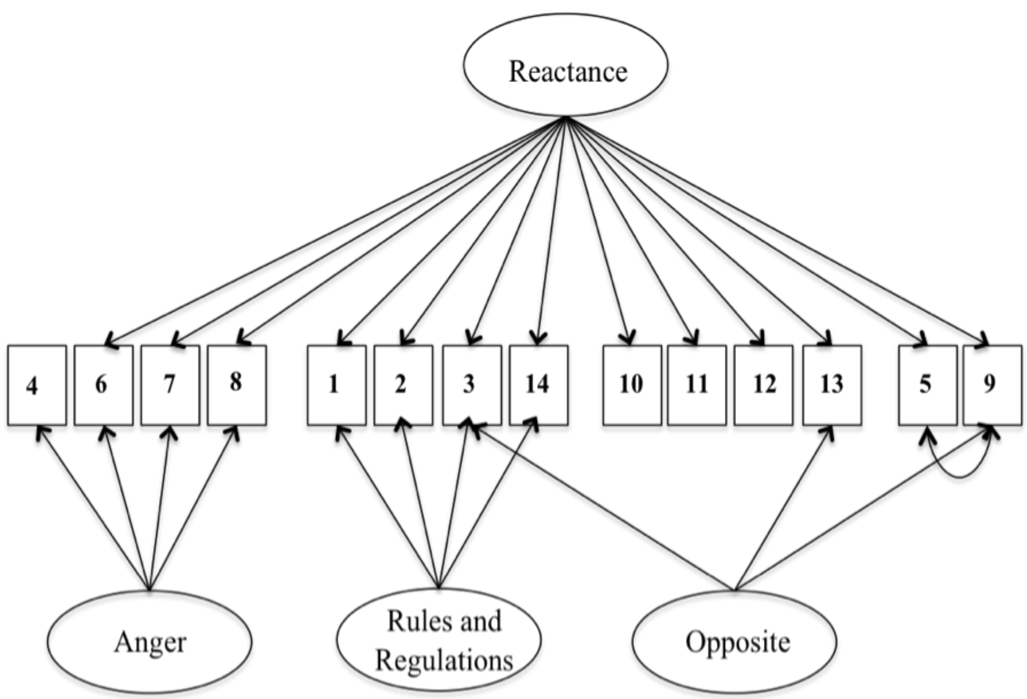

*Note:* In the bifactor model, a correlated error, not a specific factor, was modeled for *Reactance Toward Advice and Recommendations* items because a factor with only two indicators (items 5 and 9) would not be identified in this model. Fortunately, the correlated error represents the shared variance in items 5 and 9 in the same manner as modeling a specific factor for these items.

*Figure 2.* Modified Models

Incomplete bifactor model




Incomplete bifactor model with “opposite” specific factor



## Appendix

Hong's Psychological Reactance Scale

The following statements concern your general attitudes. Read each statement and please indicate how much you agree or disagree with each statement. If you *strongly agree* mark a 5. If you *strongly disagree,* mark a 1. If the statement is more or less true of you, find the number between 5 and 1 that best describes you. Realize that students do not feel the same nor are they expected to feel the same. Simply answer how **you** feel. **There are no right or wrong answers. Just answer as accurately as possible.**

| Strongly Disagree | Disagree | Neither Agree nor Disagree | Agree | Strongly Agree |
|---|---|---|---|---|
| 1 | 2 | 3 | 4 | 5 |

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 usually 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 role 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 standards and rules.