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Dispositional Resistance to Change: Measurement Equivalence and the Link to Personal Values Across 17 Nations

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The concept of dispositional resistance to change has been introduced in a series of exploratory and confirmatory analyses through which the validity of the Resistance to Change (RTC) Scale has been established (S. Oreg, 2003). However, the vast majority of participants with whom the scale was validated were from the United States. The purpose of the present work was to examine the meaningfulness of the construct and the validity of the scale across nations. Measurement equivalence analyses of data from 17 countries, representing 13 languages and 4 continents, confirmed the cross-national validity of the scale. Equivalent patterns of relationships between personal values and RTC across samples extend the nomological net of the construct and provide further evidence that dispositional resistance to change holds equivalent meanings across nations.

Keywords: resistance to change, personal values, measurement equivalence, scale validation

Change is ubiquitous. It exists in practically all aspects of life and affects virtually every individual worldwide. People, however, differ in how they respond to change. Whereas some gladly embrace the notion of change and actively seek it out, others tend to avoid it when possible and to resist it otherwise. The concept of dispositional resistance to change, which embodies these individual differences, has been established in an article previously published in the *Journal of Applied Psychology* (Oreg, 2003). The trait of dispositional resistance to change and its measurement scale (henceforth called the *RTC Scale*) have been developed through a series of studies in which the scale's structural, construct, concurrent, and predictive validities were demonstrated. As Oreg's (2003) studies show, the trait is related to, yet distinct from, other personality traits such as intolerance for ambiguity (Budner, 1962), sensation seeking (Zuckerman, 1994), dogmatism (Rokeach, 1960), risk aversion (Slovic, 1972), and openness to experience (Digman, 1990). Overall, those who are dispositionally resistant to change are less likely to voluntarily initiate changes and more likely to form negative attitudes toward the changes they encounter (e.g., Oreg, 2006).

The trait comprises four dimensions: *Routine seeking* involves the extent to which one enjoys and seeks out stable and routine environments; *emotional reaction* reflects the extent to which individuals feel stressed and uncomfortable in response to imposed change; *short-term focus* involves the degree to which individuals are preoccupied with the short-term inconveniences versus the potential long-term benefits of the change; finally, *cognitive rigidity* represents a form of stubbornness and an unwillingness to consider alternative ideas and perspectives. Although different dimensions become salient in different contexts, the composite RTC score has been shown to predict individuals' reactions to change in a variety of contexts under both voluntary and imposed conditions (Oreg, 2003, Studies 5–7; Oreg, 2006).

When considering evidence for the RTC Scale's validity, however, it is worth noting that six of the seven samples in Oreg's (2003) studies involved participants from the United States alone. The seventh sample was nationally mixed, with a majority of U.S. participants. This raises the question of whether the trait also represents meaningful individual differences in countries that embody other national cultures. If so, to what extent is that meaning equivalent across these countries? To what extent does the trait's structure replicate? From a practical perspective, we ask whether the RTC Scale can be used as a measure of dispositional resistance to change across national cultures. In the present work we aim at answering these questions by testing the measurement equivalence of the scale and by extending its nomological net across 17 countries.

Although we aim to examine measurement equivalence across cultures, in practice our comparisons are among national samples. Clearly, cultural boundaries do not necessarily coincide with the geographical boundaries of nations, and each nation is likely to comprise several subcultures (Schwartz, 1999). Nevertheless, nations typically include a dominant system, including the dominant official language, educational, and political systems, as well as shared media, symbols, and markets (Liu, Borg, & Spector, 2004). Thus, each nation can be said to include a dominant culture, with core attributes that are shared among its subcultures. Accordingly, and in line with most cross-cultural studies, our comparison of cultures is based on the comparison of national samples under the premise that there is greater similarity within countries than there is between them.

Establishing the Meaningfulness of a Construct Across Cultures

The establishment of new constructs and measures is typically done in the context of a specific culture. However, cross-cultural research suggests that something is shared in the way people from a given culture think, feel, and behave and that this something varies across cultures (e.g., Hofstede, 1980). This shared narrative evolves from the shared history and language and is ultimately reflected in beliefs, values, and other psychological attributes that are similar among individuals within a given culture, and it is likely to be different among individuals from other cultures (M. W. L. Cheung, Leung, & Au, 2006). Accordingly, it is now well accepted that before a psychological construct can be informatively used across cultures and languages, one should verify that its meaning is invariant, or at least similar, across these cultures (e.g., M. W. L. Cheung et al., 2006; Church & Lonner, 1998; Ghorpade, Hattrup, & Lackritz, 1999; Liu et al., 2004). From a methodological perspective, evidence for measurement equivalence is required before an instrument can be validly used across cultures (e.g., Liu et al., 2004; Vandenberg & Lance, 2000). Such evidence would indicate that constructs are comparable and that the instrument taps the same psychological meanings (Ghorpade et al., 1999). In turn, this would suggest that models with constructs developed in a given culture can be validly applied in other cultures.

At present, it is not clear whether the concept of dispositional resistance to change shares its meaning across cultures and whether the RTC Scale taps this shared meaning. Resistance to change may appear to be a universal phenomenon, in particular through the effects of technology and globalization and with changes increasing in frequency and magnitude. Yet equivalence in the construct's meaning across cultures should not be taken for granted. Existing typologies of culture suggest that societies differ in their general inclinations toward change (e.g., Hofstede, 2001). In particular, among Hofstede's (2001) five dimensions, *uncertainty avoidance* appears to distinguish between societies on aspects that are related to the notion of change. Uncertainty avoidance involves the extent to which a society aims to reduce uncertainty as a means of coping with anxiety. This dimension reflects the extent to which a culture emphasizes structure and stability (Hofstede, 2001). Thus, individuals' interpretations of change and their inherent reactions to it may vary across cultures. A test of measurement equivalence is required before we can determine whether individual differences in the orientation toward change are comparable across cultures.

Evidence for construct validity across cultures is typically established by showing that relationships among scale items are equivalent across cultures. Another approach on which construct validation is frequently based, yet which is less typically used in cross-cultural research, involves the establishment of the construct's nomological net. Evidence for the equivalence of a construct's meaning can be provided by demonstrating the existence of the same nomological net across cultures. Both approaches are used in the present study. Specifically, we begin by using multiple-group confirmatory factor analysis (CFA), in which the homogeneity of correlation matrices across nations is tested (M. W. L. Cheung et al., 2006). We then extend the nomological net of dispositional resistance to change and further establish the scale's construct validity by comparing relationships between RTC and a set of well established, cross-culturally validated dispositions. For this, we chose to focus on personal values.

Resistance to Change and Personal Values

As an anchor for comparison, we sought an individual differences framework that has already been validated across cultures and that could be theoretically linked with dispositional resistance to change. Schwartz's (1992) theory of personal values constitutes one such framework. Values are often defined as transituational goals that vary in their importance and serve as guiding principles in people's lives (Kluckhohn, 1951; Rokeach, 1973; Schwartz, 1992). Contrary to traits (e.g., dispositional resistance to change), values are typically conceptualized (e.g., Rokeach, 1973; Schwartz, 1992) as being subjectively rank ordered, with their combination forming a system of value priorities or hierarchies.

Schwartz's (1992, 2005) research identifies 10 values that can each be categorized into one of four broad value dimensions. These dimensions are represented in two primary contrasts. The first involves *self-enhancement* versus *self-transcendence* values and describes the tension between an individual's emphasis on his or her own success and dominance versus an emphasis on the welfare of others. The second contrast is of central relevance to the present work and involves the tension between *openness to change* values and *conservation* values. Openness values represent an emphasis on the proactive and voluntary search for stimulation, novelty, and change and on free and autonomous thinking and behavior. Conversely, conservation values prescribe the status quo, the preservation of security and social order, and submissive self-restriction. Accordingly, dispositional resistance to change is expected to yield negative correlations with openness to change values and positive correlations with conservation values.

Considering that the structure of Schwartz's value system has been validated in more than 200 samples from more then 70 cultural groups (e.g., Schwartz, 1992, 2005; Schwartz & Sagiv, 1995) and that value dimensions have been found to hold equivalent meanings across cultures, we used value dimensions in the present study as an anchor around which we aimed to establish cross-cultural equivalence in the meaning of dispositional resistance to change. Assuming that dispositional resistance, as measured with the RTC Scale, shares its meaning across cultures, values of openness to change should show a negative correlation and values of conservation a positive correlation with RTC scores across all of the countries sampled.

Method

Participants

As a prerequisite to testing measurement invariance across samples, it is necessary to ensure sample comparability. A common and suitable procedure for this is to use samples that are matched on the basis of a predetermined set of characteristics (Steenkamp & Baumgartner, 1998). Although such matching is not without its faults (we elaborate on this in the Discussion section), it is nonetheless essential when random samples are not available. Therefore, samples in the present study consisted of undergraduates, thus matching on the basis of level of education and age. A total of 4,201 undergraduate students from 17 countries participated in the study for course credit or as part of the course requirements. Table 1 provides descriptive statistics of the samples' characteristics, including the geographical location, sample size, language in which surveys were administered, religion, percentage of female participants, and mean age.

The countries sampled in the present study are Australia, China, Croatia, the Czech Republic, Germany, Greece, Israel, Japan, Lithuania, Mexico, the Netherlands, Norway, Slovakia, Spain, Turkey, the United Kingdom, and the United States. The average sample size was 241, ranging from 171 to 386. Except for China and Slovakia, for which sample sizes were 194 and 171, respectively, all samples included more than 200 respondents.

Measures and Procedure

Following recommended procedures (e.g., Schaffer & Riordan, 2003), we translated RTC and values scales to the native language of each participating country through a translation and back-translation process by two individuals who were fluent in both English and the country's native language. Any differences found between the original and back-translated versions were discussed until agreement was reached concerning the most appropriate translation.

Participants filled out the 17-item RTC Scale, Schwartz's 40-item Portrait Value Questionnaire (PVQ; Schwartz et al., 2001), and a demographics questionnaire. Items on the RTC Scale consist of statements concerning one's typical orientation toward and reaction to change (see the first column of Table 2). Response

Table 1
Descriptive Statistics on Samples' Demographics

Country	Town	*N*	Language	Religion (majority)	% female	Age (in years) *M*	Age (in years) *SD*
Australia	Burwood and St. Lucia	251	English	30% atheist	67	21.09	3.61
China	Beijing	194	Chinese	—	56	20.72	1.09
Croatia	Zagreb	246	Croatian	81% Roman Catholic	83	21.43	1.79
Czech Republic	Brno	224	Czech	50% Roman Catholic	78	22.49	2.10
Germany	Braunschweig	206	German	51% Protestant	49	23.03	4.35
Greece	Athens	386	Greek	87% Greek Orthodox	60	20.97	2.31
Israel	Haifa	241	Hebrew	83% Jewish	82	24.35	3.21
Japan	Tsukuba	337	Japanese	—	23	19.71	1.62
Lithuania	Vilnius	212	Lithuanian	96% Catholic	77	20.31	1.67
Mexico	Mexico City, Tampico, and Merida	265	Spanish	82% Catholic	51	20.62	2.19
Netherlands	Tilburg	205	Dutch	—	80	20.22	3.45
Norway	Bergen	266	Norwegian	67% Christian	74	23.24	4.40
Slovakia	Bratislava	171	Slovakian	50% Catholic	54	21.40	1.10
Spain	Salamanca	288	Spanish	—	59	21.90	1.55
Turkey	Istanbul	241	Turkish	98% Muslim	39	21.04	1.52
United Kingdom	Durham	204	English	95% Christian	45	19.22	1.83
United States	Auburn, AL	264	English	49% Christian	50	21.19	2.38
Total or *M*		4201			60.41	21.35	2.37

Note. In some countries, it was deemed inappropriate to collect data on respondents' religion in the context of this study. Dashes represent the missing information for these countries.

options range from 1 (*strongly disagree*) to 6 (*strongly agree*). The reliability (Cronbach's alpha) of the scale in each country is presented in the seventh column of Table 3.

The PVQ Scale includes short verbal portraits of hypothetical individuals. Each portrait describes a person's goals or aspirations that point implicitly to the importance of a value. The verbal portraits describe each person in terms of what is important to him or her. For example, the item "Thinking up new ideas and being creative is important to him. He likes to do things in his own original way"[1] describes a person who values openness to change. For each portrait, participants respond to the question "How much like you is this person?" Responses range from 1 (*not like me at all*) to 6 (*very much like me*). Values are inferred from their self-reported similarity to the individuals described in the various items. In line with Schwartz's (1992) prescriptions, we centered respondents' value scores to control for individual differences in how people distribute importance ratings across value items. We subtracted the mean response to all value items from the mean response to items within each dimension. For example, an individual's centered score on conservation is obtained by subtracting that individual's mean response to all 40 PVQ items from his or her mean response to the conservation items.

The PVQ has been used in several studies across numerous countries and has been shown to be a valid measure of values (e.g., Capanna, Vecchione, & Schwartz, 2005; Koivula & Verkasalo, 2006; Schwartz et al., 2001). Coefficient alphas of the four value dimensions in each of the samples are presented in Columns 3–6 of Table 3. As can be seen, except for the Openness subscale in one sample and the Conservation subscale in two samples, the coefficient alphas of the value scales achieved acceptable levels. The mean coefficient alphas across samples were .77, .75, .80, and .83 for the Openness, Conservation, Self-Transcendence, and Self-Enhancement subscales, respectively.

Analyses

To test for the scale's measurement equivalence across samples, we used a multigroup CFA procedure (using AMOS Version 7.0; Arbuckle, 2006). Although many approaches have been proposed for this procedure, tests of measurement equivalence in cross-cultural research typically follow a three-step series of nested constraints that are placed on parameters across samples (e.g., Grouzet et al., 2005; Meredith, 1993; Steenkamp & Baumgartner, 1998). The first two steps include tests of *configural* and *metric invariance* and establish that a construct holds the same psychological meaning across samples. The third step, testing for *scalar invariance,* aims at verifying that sample means, which should reflect some theoretically established higher level construct, can be meaningfully compared. Because in the present study our focus is on establishing the measurement equivalence of an individual-level personality construct, and because a culture-level resistance construct is yet to be conceptualized, we restrict our analyses to the first two steps. We elaborate on this issue in the Discussion section.

Configural invariance constitutes the most basic test of measurement equivalence and involves an examination of the configuration of relationships between items and latent variables across samples. Each of the scale items is required to show the same pattern of zero and nonzero loadings on the latent factors in each of the samples. In our case, this step would involve a test of the extent to which the same four-factor RTC Scale structure is justified in all samples.

Replicating a construct's structure by demonstrating configural invariance, however, provides only preliminary evidence that the

[1] The PVQ Scale has different forms for male and female respondents.

Table 2
Resistance to Change Scale Items and Confirmatory Factor Analysis Factor-Loading Ranges Across the 17 Samples

Factor and item	Loading *M*	Loading *SD*
Routine seeking		
1. I generally consider changes to be a negative thing.	0.54	0.14
2. I'll take a routine day over a day full of unexpected events any time.	0.64	0.10
3. I like to do the same old things rather than try new and different ones.	0.70	0.08
4. Whenever my life forms a stable routine, I look for ways to change it.[a]	0.44	0.11
5. I'd rather be bored than surprised.	0.50	0.09
Emotional reaction		
6. If I were to be informed that there's going to be a significant change regarding the way things are done at school, I would probably feel stressed.[b]	0.64	0.08
7. When I am informed of a change of plans, I tense up a bit.	0.72	0.08
8. When things don't go according to plans, it stresses me out.	0.64	0.08
9. If one of my professors changed the grading criteria, it would probably make me feel uncomfortable even if I thought I'd do just as well without having to do any extra work.[b]	0.54	0.10
Short-term focus		
10. Changing plans seems like a real hassle to me.	0.62	0.11
11. Often, I feel a bit uncomfortable even about changes that may potentially improve my life.	0.72	0.09
12. When someone pressures me to change something, I tend to resist it even if I think the change may ultimately benefit me.	0.49	0.10
13. I sometimes find myself avoiding changes that I know will be good for me.	0.50	0.09
Cognitive rigidity		
14. I often change my mind.[a]	0.48	0.17
15. I don't change my mind easily.	0.63	0.11
16. Once I've come to a conclusion, I'm not likely to change my mind.	0.68	0.08
17. My views are very consistent over time.	0.64	0.13

[a] This item is reverse coded. [b]When used in a job setting, these items are rephrased to fit the organizational context.

construct shares its meaning across samples. A much stronger case is made if item loadings are of the same magnitude across samples (Meredith, 1993; Vandenberg & Lance, 2000). This indicates that members of the different samples calibrate the measure and thus interpret the construct in the same way. This form of invariance is called *metric invariance* and involves a model identical to that tested for configural invariance with the added constraint of factor loadings across samples.

In line with Coovert and Craiger's (2000) recommendations, we included the two indexes considered most important for determining model fit: the root-mean-square error of approximation (RMSEA) and the comparative fit index (CFI). We also looked at the goodness-of-fit index (GFI), which is commonly considered in CFAs. CFI and GFI values range from 0 to 1.00, where values greater than .95 indicate good fit and values greater than .90 are considered satisfactory (Hoyle, 1995). For RMSEA, values of .05 or less indicates a close fit and values of up to .08 represent reasonable errors of approximation (Browne & Cudeck, 1993).

As evidence for metric invariance, beyond having a good fit, the fit of the *metric* model should not be significantly worse than that of the *configural* model. Although traditionally only the chi-square difference test has been used, it is well acknowledged that a statistically significant chi-square is often obtained even when there are only minor differences between groups' factor patterns (Vandenberg & Lance, 2000). Thus, as in the case of establishing model fit, differences between models should be established through the use of fit indices beyond the chi-square (Bollen, 1989; Vandenberg & Lance, 2000). In particular, use of the differences between indexes such as the RMSEAs and CFIs of both models has been recommended (G. W. Cheung & Rensvold, 2002). For the ΔCFI, an absolute value of .01 or smaller indicates that the invariance hypothesis should not be rejected. Values over .02 indicate a lack of invariance, and values between .01 and .02 suggest that some differences may exist between models (G. W. Cheung & Rensvold, 2002). No critical values have been indicated in the literature for the RMSEA.

It should be noted, however, that full measurement invariance is quite rare, with some researchers arguing that it is particularly unlikely when testing forms of invariance beyond configural invariance (Horn, 1991; Steenkamp & Baumgartner, 1998). Therefore, for many constructs, and in particular when testing invariance across a large variety of samples, it may be that only *partial measurement invariance* exists (Byrne, Shavelson, & Muthen, 1989; Steenkamp & Baumgartner, 1998). A test of partial metric invariance would require relaxing some of the item loading constraints. Although all of the items would still be required to load on the same factors in each of the samples, the requirement that the loadings be of the same magnitude across samples may be dropped for some of the items. Whenever possible, the choice of constraints to be relaxed should be based on substantive rather than on sample-specific empirical data.

After testing measurement invariance, we compared the pattern of relationships between dispositional resistance and personal values across the 17 samples as a second source of evidence for the construct's shared meaning across cultures.

Results

As can be seen in Column 7 of Table 3, all RTC coefficient alphas achieved a satisfactory level of .70 or above. The mean alpha was .80, with coefficients ranging from .72 to .85. As a first step in establishing measurement equivalence, a separate CFA was run for each sample. In each of the 17 CFAs, all of the items significantly loaded ($p < .05$) on their expected factor (see Figure 1 and Table 2). Furthermore, as can be seen in Table 3, the four-factor RTC Scale presented satisfactory fit across all countries, with the exception of the GFI in Slovakia, which was just below .90. RMSEAs ranged from .033 to .065, CFIs ranged from .90 to .97, and GFIs ranged from .89 to .94. The mean RMSEA, CFI, and GFI values across the 17 samples were .050, .93, and .92, respectively.

Table 3
Coefficient Alphas, Descriptive Statistics, and Fit Indexes for the 17 Samples

Country	*N*	Reliabilities (α) Openness	Conservation	Self-Transcendence	Self-Enhancement	RTC	RTC *M*	RTC *SD*	$\chi^2(107)$	RMSEA	CFI	GFI
Australia	251	.78	.73	.80	.85	.82	3.09	0.57	172.56	.050	.93	.93
China	194	.82	.75	.81	.85	.85	3.14	0.62	170.07	.055	.94	.91
Croatia	246	.82	.84	.83	.84	.84	3.01	0.61	159.88	.045	.97	.93
Czech Republic	224	.85	.83	.87	.86	.84	3.13	0.56	184.24	.057	.92	.91
Germany	206	.79	.65	.75	.88	.77	3.12	0.48	131.36	.033	.97	.93
Greece	386	.71	.58	.75	.87	.72	3.03	0.50	227.29	.054	.93	.94
Israel	241	.80	.81	.80	.84	.85	3.15	0.59	193.42	.058	.93	.92
Japan	337	.81	.72	.80	.78	.75	3.22	0.52	199.46	.051	.91	.93
Lithuania	212	.76	.81	.82	.87	.77	2.86	0.51	171.39	.053	.92	.91
Mexico	265	.71	.76	.80	.78	.79	2.79	0.58	216.74	.062	.92	.90
Netherlands	205	.78	.74	.83	.83	.85	3.17	0.52	177.59	.058	.94	.91
Norway	266	.79	.76	.73	.86	.84	2.91	0.56	218.21	.063	.92	.91
Slovakia	171	.77	.75	.79	.84	.79	3.27	0.51	184.28	.065	.90	.89
Spain	288	.79	.76	.84	.82	.81	3.01	0.58	165.97	.044	.95	.94
Turkey	241	.74	.77	.80	.83	.77	3.03	0.54	188.86	.056	.90	.91
United Kingdom	204	.77	.77	.83	.84	.78	3.02	0.51	190.22	.062	.90	.90
United States	264	.64	.73	.71	.72	.83	3.05	0.54	160.90	.044	.95	.94
M	247.12	.77	.75	.80	.83	.80	3.06	0.55	183.08	.050	.93	.92

RTC = Resistance to Change Scale; RMSEA = root-mean-square error of approximation; CFI = comparative fit index; GFI = goodness of fit index

In line with Oreg's (2003) findings, there were significant correlations among RTC subscales (Table 4, top section), with the highest correlation being between the Emotional Reaction and Short-Term Focus subscales and the lowest involving the Cognitive Rigidity subscale. The high correlation between the Emotional Reaction and Short-Term Focus subscales has been previously explained on the basis of both dimensions being affective in nature. Following Oreg (2003), we compared the four-factor model with a three-factor model, whereby items of the Emotional Reaction and Short-Term Focus subscales were all set to load on a single affective factor (see Figure 2). The three-factor model presented poorer fit on all three fit indexes in all samples, with the exception of Lithuania, in which the two models presented virtually equal fit. With this exception of Lithuania, the chi-square tests comparing the three- and four-factor models in each of the samples indicated that the fit of the four-factor model was significantly ($p < .01$) better.

Considering the particularly low correlations of cognitive rigidity with the remaining subscales in some of the samples, we wanted to test the extent to which each of the subscales was associated with the overarching resistance construct. We therefore tested a second-order latent factor model, in which the RTC subscale factors loaded on a latent factor representing the overall resistance to change disposition (see Figure 3). In all samples, model fit was satisfactory and was very similar to the fit of the first-order factor model that was initially tested. In addition, the Routine Seeking, Emotional Reaction, and Short-Term Focus subscales yielded significant loadings on the second-order factor in all samples. The Cognitive Rigidity subscale yielded significant loadings in 14 of the 17 samples. The loading of this subscale was not significant in Greece, Slovakia, and the United Kingdom. We elaborate on this finding in the Discussion section.

Measurement Equivalence

After testing model fit in each country separately, we proceeded with the intercorrelated four-factor model (Figure 1) to test the configural and metric invariance of the scale across the 17 samples. In the configural invariance model, all items yielded a significant loading on their corresponding factors in all 17 countries. Furthermore, the model presented a satisfactory fit (RMSEA = .013, CFI = .928, GFI = .919), indicating that the pattern of item loadings is consistent across samples. Similar results were obtained for the metric invariance model. All items loaded on their corresponding factors and model fit was satisfactory (RMSEA = .014, CFI = .915, GFI = .909), suggesting that the magnitude of item loadings was consistent across samples.

Stronger evidence for metric invariance is established by comparing the extent to which the fit of the metric model is poorer than that of the configural model. We therefore ran the chi-square difference test and looked at the differences in RMSEAs, CFIs, and GFIs across models. The chi-square difference test was significant, $\Delta\chi^2(195) = 446$, $p < .01$. However, as noted above, given that the chi-square difference test suffers from the same problems as the chi-square test for determining model fit, this should not be considered evidence for the lack of invariance. Calculated from the fit indexes of the two models (configural and metric), the differences in the fit indexes were .001, .013, and .010 for the ΔRMSEA, ΔCFI, and ΔGFI, respectively. Although the ΔRMSEA and ΔGFI meet the .01 threshold and indicate a negligible difference between the fit of the two models, the ΔCFI suggests that some of the constraints on item loadings may not be justified. Therefore, even though the fully constrained metric model yielded satisfactory fit, we wanted to gain additional insights as to why its fit was somewhat poorer than the fit of the configural invariance model. We therefore considered the possi-

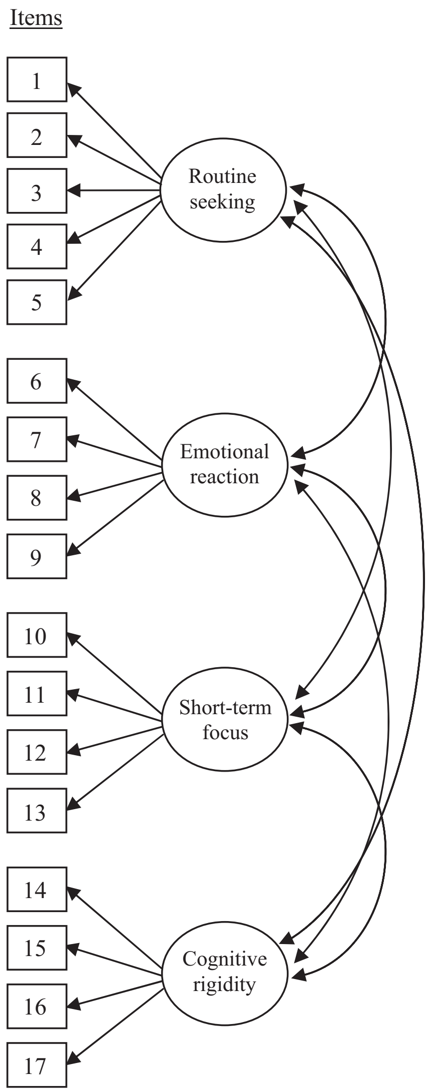

Figure 1. Four-factor model with intercorrelated factors.

bility of partial metric invariance by relaxing some of the loading constraints.

Although previous research with the RTC Scale provides little basis for determining which of the RTC items would hold a more consistent meaning across cultures, differences in the responses of participants from different cultures might be expected for the scale's two negatively worded items (i.e., Items 4 and 14). Previous research on the use of personality scales across cultures indicates that negatively worded items often yield different responses and have a differential effect across cultures (e.g., Lai & Yue, 2000; Schmitt & Allik, 2005). Thus, despite the fact that Items 4 and 14 loaded significantly on their expected factors in all 17 samples, the magnitude of their loadings may not be invariant across samples. We therefore relaxed the loading constraints for these two items and retested the metric invariance. As expected, model fit improved: RMSEA = 0.13, CFI=.919, and GFI=.911. The differences in fit indexes between the configural model and this metric model were now .001, .009, and .008 for the RMSEA,

Table 4
Mean Correlations Between Resistance to Change Scale (RTC) Subscales and Portrait Value Questionnaire (PVQ) Dimensions Across Samples

Subscale	Routine Seeking	Emotional Reaction	Short-Term Focus	Cognitive Rigidity
RTC subscales				
1. Routine Seeking	—			
2. Emotional Reaction	.49	—		
3. Short-Term Focus	.61	.77	—	
4. Cognitive Rigidity	.23	.16	.21	—
PVQ subscales				
5. Openness	−.50	−.30	−.32	−.14
6. Conservation	.46	.26	.28	.17
7. Self-Transcendence	−.05	−.05	−.03	−.06
8. Self-Enhancement	−.02	.03	−.01	.00

Note. Correlations (estimated) among RTC subscales were derived within the CFAs. The correlations between RTC and PVQ subscales were calculated from the means of the RTC and PVQ subscale scores.

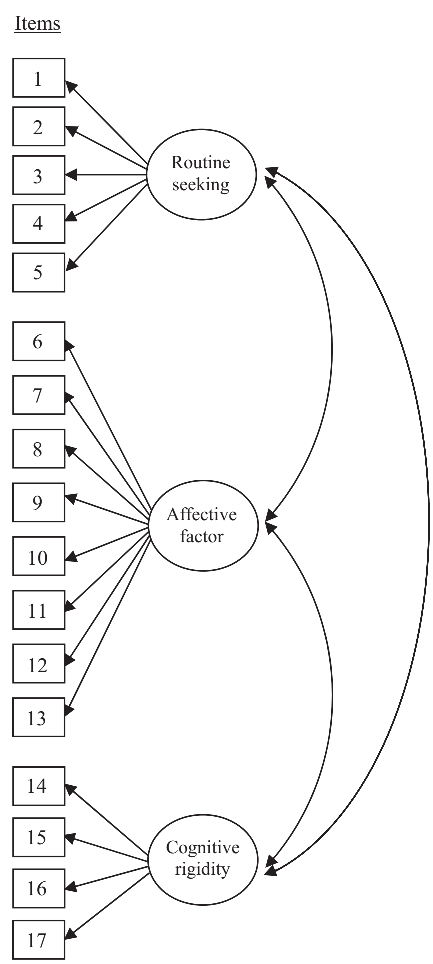

Figure 2. Three-factor model, with the emotional reaction and short-term focus factors merged into one affective factor.

CFI, and GFI, respectively, thus meeting G. W. Cheung and Rensvold's (2002) criteria for invariance.[2]

Resistance and Values

To test for the equivalence in the meanings of the dispositional resistance construct in yet another way, we tested the correlations between RTC Scale scores and the value dimension scores in each of the samples (see Table 5). As expected, in all countries, the RTC Scale yielded significant negative correlations with Openness values (ranging from $-.27$ to $-.57$, $p < .01$, in all cases) and positive correlations with Conservation values (ranging from .23 to .58, $p < .01$). The relationships between RTC Scale scores and the other two value dimensions were substantially weaker and mostly nonsignificant. This pattern of relationships was replicated for each of the four RTC Scale dimensions, with the strongest relationships being found with the Routine Seeking subscale and the weakest with the Cognitive Rigidity subscale (see the bottom section of Table 4). This is consistent with the fact that the Cognitive Rigidity dimension exhibits the lowest correlation with the RTC Scale composite score, both in the present study as well as in Oreg's (2003) studies.

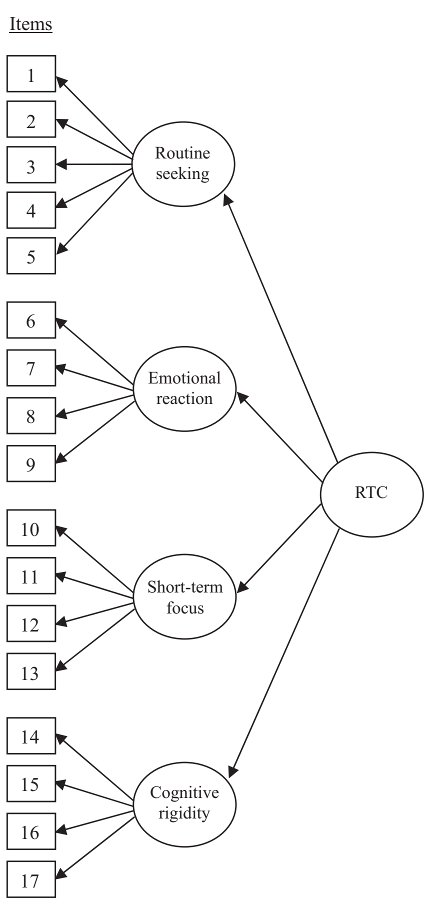

Figure 3. Four-factor model with a second-order resistance to change (RTC) factor.

Table 5
Pearson Correlations Between Resistance to Change (RTC) and Portrait Value Questionnaire Dimensions for the 17 Samples

Country	RTC & Openness	RTC & Conservation	RTC & Self-Transcendence	RTC & Self-Enhancement
Australia	$-.44^{**}$	$.32^{**}$	.06	$-.07$
China	$-.54^{**}$	$.55^{**}$	$-.05$	$-.02$
Croatia	$-.56^{**}$	$.58^{**}$	$-.19^{*}$	$-.12$
Czech Republic	$-.57^{**}$	$.54^{**}$	$-.03$	$-.10$
Germany	$-.50^{**}$	$.54^{**}$	$-.04$	.03
Greece	$-.44^{**}$	$.45^{**}$	$-.13^{**}$	$-.05$
Israel	$-.57^{**}$	$.51^{**}$	$-.03$	.00
Japan	$-.45^{**}$	$.42^{**}$	$-.06$	.01
Lithuania	$-.44^{**}$	$.48^{**}$	$-.13$	$-.01$
Mexico	$-.30^{**}$	$.34^{**}$	$-.20^{**}$	.08
Netherlands	$-.56^{**}$	$.45^{**}$	.06	.10
Norway	$-.54^{**}$	$.44^{**}$	$-.10$	.08
Slovakia	$-.34^{**}$	$.33^{**}$	$-.13$	.03
Spain	$-.50^{**}$	$.43^{**}$	$-.06$	.03
Turkey	$-.31^{**}$	$.23^{**}$	$-.12$	.06
United Kingdom	$-.47^{**}$	$.44^{**}$	$-.13$	.06
United States	$-.27^{**}$	$.28^{**}$	$-.17^{*}$	.06
M	$-.43$	.44	$-.08$	.00

* $p < .05$. ** $p < .01$ (two-tailed significance tests).

Discussion

Does dispositional resistance to change take on equivalent meanings across cultures? We addressed this question using data from 17 countries on four continents with 13 different languages. With some exceptions, the cross-national validity of the construct was established through a replication of the scale's structure and with evidence for partial measurement equivalence. The scale yielded the same structure across all samples and, with the exception of the two negatively worded items, all of the items had invariant loadings across countries. Furthermore, expected relationships between the scale and Schwartz's (1992) value dimensions further expand the construct's nomological net and help establish the scale's convergent and discriminant validities. These results suggest that the construct of dispositional resistance to change carries equivalent meanings across nations and that its measurement scale can be reliably and validly used in the countries sampled for this study. Caution should be taken, however, when interpreting individuals' responses to the two negatively worded

[2] Certainly, relaxing any loading constraints would be expected to improve model fit to some degree. Yet our choice to relax constraints for the two reverse-coded items was based on the strong recommendation (e.g., Byrne, Shavelson, & Muthen, 1989; Steenkamp & Baumgartner, 1998) to relax constraints only for items for which there is a priori reason to suspect lack of invariance rather than relaxing them arbitrarily or on the basis of sample-specific results. To examine the extent to which the relaxation of loadings for the reverse-coded items is substantive, we tested whether this relaxation improved fit beyond the relaxation of any two items. We therefore ran three additional tests of metric invariance, each time relaxing loadings of a different pair of items, determined randomly. In two of the three tests, the fit of the new metric model somewhat improved over the fully constrained metric model, but in none of the cases did the ΔCFI reach below the .01 threshold. This supports the indication that negatively worded items may be particularly prone to different interpretations across cultures (e.g., Lai & Yue, 2000; Schmitt & Allik, 2005).

items. The negative wording appears to yield different responses in different cultures.

An interesting and unexpected finding was the nonsignificant loading of the Cognitive Rigidity subscale on the second-order resistance construct in 3 of the 17 samples. When considering these three samples—Greece, Slovakia, and the United Kingdom—we cannot identify a common denominator that could explain why these particular three should yield lower loadings of cognitive rigidity. We should note, however, that the overall divergence of cognitive rigidity from the remaining three facets is consistent with Oreg's (2003) findings.

From a conceptual perspective, we can offer at least one explanation for the inconsistent patterns that are sometimes observed with cognitive rigidity. Routine seeking and, even more so, emotional reaction and short-term focus reflect a form of insecurity. Those who enjoy routines, who react emotionally to changes, and who tend to focus on the short-term hassles that change creates resist change because it elicits discomfort and stress. In this respect, the subscales should be negatively associated with traits such as emotional stability and self-confidence. However, much like dogmatism (Rokeach, 1960), cognitive rigidity involves strong personal convictions and a form of stubbornness that are typically associated with higher levels of self-confidence. In line with such a rationale, the former three subscales have been shown to correlate positively with neuroticism and negatively with self-esteem and self-efficacy, whereas cognitive rigidity has shown the reverse pattern of relationships (Oreg, 2003, Study 3).

Nevertheless, the relationship patterns between cognitive rigidity and several other external variables, both in this study (with values) and previously (Oreg, 2003, Study 3, with risk aversion, sensation seeking, and dogmatism), correspond with the patterns found for the other subscales. Furthermore, the significant relationships between cognitive rigidity and actual resistance reactions (e.g., Naus, van Iterson, & Roe, 2007; Oreg, 2003, Studies 5 and 6) suggest that it taps a unique yet meaningful portion of variance in individuals' reactions to change. We therefore suggest that when considering using the RTC Scale to predict reactions to specific changes, the Cognitive Rigidity subscale should be maintained as part of the general measure.

Individual reactions to change have long been explained and predicted by the nature of the change as well as by the context in which change occurs (e.g., Armenakis, Harris, & Mossholder, 1993; Jones, Jimmieson, & Griffiths, 2005; van Dam, Oreg, & Schyns, 2008). Dispositional resistance to change predisposes some people to show an adverse reaction to a change even if the change is docile and its context is relatively welcoming. Such people find comfort in routines, are less flexible cognitively, and find it more difficult to set aside the short-term inconveniences of change. Not only do they react more negatively than others to harmful changes, but they also resist changes that may turn out to be beneficial. Knowledge of who these people are is important for organizational change management and for career counseling. Evidence from this article suggests that information gathered with the RTC Scale can be validly used to try to identify these individuals, at least within the 17 nations we sampled.

Two limitations of the present study are worth noting. First, because we could not collect data from random samples, we used a matched-sample design to increase sample comparability. Although such comparability is essential for cross-cultural validation studies, the price, in this case, was that our samples were not representative of their national cultures. Moreover, matching on a given variable may inadvertently result in unmatching on other variables (Meehl, 1970), thus raising further doubts about the generalizability of one's findings. However, some evidence for the external validity of our findings lies in the fact that Oreg's (2003) findings with U.S. undergraduates were equivalent to those found for U.S. employees. Furthermore, in Schwartz's (1992, 2005) studies of values, comparisons of undergraduate samples yielded patterns that were equivalent to those found through comparisons of representative national samples. Combined with the consistent patterns of relationships found in the present study between values and dispositional resistance to change, this somewhat attenuates the concern for nongeneralizability. Nevertheless, a replication of our findings with representative samples would be necessary before the measurement equivalence of the RTC Scale could be more definitively concluded.

Related to the first limitation, the cultures sampled for our study do not represent all existing national cultures around the globe. Although our results indicate measurement equivalence across the 17 countries sampled, it is yet to be determined whether such equivalence exists across additional countries. In particular, none of our samples were from Africa or South America. Additional data from countries on these continents would be necessary to more confidently argue for a universally shared meaning of the dispositional resistance to change construct.

Our findings suggest that dispositional resistance to change shares its meaning, as an individual-level construct, across cultures. Thus, the RTC Scale can be used to compare individuals within a given culture across a large variety of cultures. This is distinct from being able to compare cultures in their aggregate level of dispositional resistance. Measurement equivalence at the individual level is a necessary yet insufficient criterion for making comparisons between units at a higher level (van de Vijver & Poortinga, 2002). This is because differences between individuals across cultures are not necessarily equivalent in meaning to differences between cultures (M. W. L. Cheung et al., 2006). The extent to which resistance to change can be viewed as a culture-level construct requires a separate validation process, which would explore, both theoretically and empirically, the meaningfulness of resistance to change as a cultural dimension. Such a validation process should include an explicit discussion of what it means for a culture to be change resistant, as is done in discussions of other culture-level constructs such as uncertainty avoidance (Hofstede, 1980). This discussion may or may not be similar to the discussion of individual differences in resistance. From an empirical perspective, data would need to be collected from a large number of cultures to permit the necessary statistical analyses. Furthermore, correlations between culture-level resistance and other culture-level variables would have to be tested to establish construct validity.

In the meantime, researchers from different cultures can join discussions on dispositional resistance with evidence that the constructs' meaning extends the boundaries of a given culture. Correspondingly, the evidence provided in the present study serves to further validate the RTC Scale and to endorse its use in empirical studies across a variety of cultures.

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