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The Zimbardo time perspective inventory: Time for a new strategy, not more new shortened versions

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Abstract

Researchers continue to attempt to resolve the psychometric problems associated with the five-factor Zimbardo Time Perspective Inventory through the development of shortened forms of the scale. These atheoretical efforts have been data driven and have resulted in scales whose reliability and validity have not been subsequently supported. The purpose of this paper was to explore the factorial validity and reliability of new short scales on samples independent from

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which they were developed. We used data from five different samples in four different countries (Australia, Britain, Slovenia, and the United States) to examine the psychometric validity and reliability of three recently developed scales, the ZTPI-20, ZTPI-17, and ZTPI-15. Results regarding validity were equivocal for all scales and reliability coefficients were suboptimal in all samples. We conclude by stressing the necessity for a theoretically driven approach to enhancing the psychometric assessment of time perspective rather than simply sacrificing reliability or discriminant validity for improved model fit in a shorter scale.

Keywords

Confirmatory factor analysis, exploratory structural equation modelling, Zimbardo Time Perspective Inventory, cross-cultural, internal consistency

Introduction

The Zimbardo Time Perspective Inventory (ZTPI; Zimbardo and Boyd, 1999) was developed to serve as an integrated measure of the cognitive, affective, and behavioural dimensions of temporal psychology and, in part, to overcome the conceptual and measurement issues that pertained in the temporal psychology field at the time. Despite these noble aspirations, evidence has emerged suggesting that issues remain regarding the reliability and factorial validity of the ZTPI (e.g. Crockett et al., 2009; McKay et al., 2015; Perry et al., 2015; Sircova et al., 2014). Of particular concern are the claims that the ZTPI consists of some items that do not even assess time perspective (Crockett et al., 2009; Shipp et al., 2009). For example, the ZTPI item, 'It upsets me to be late for appointments' is proposed to assess future time perspective; however, it might as easily serve as an assessment of conscientiousness.

The psychometric shortcomings of the original 56 items have precipitated the development of multiple short forms of the scale; specifically, there are six recently published shortened versions of which we are aware.We have contended (McKay et al., 2015; Perry et al., 2015) that three of these attempts to strengthen the reliability and validity of the ZTPI solely through the elimination of items, and the production of shortened versions of the scale (e.g. Laghi et al., 2013; Sircova et al., 2014; Zhang et al., 2013) is atheoretical and, therefore, does not promote the development of research in temporal psychology. In other words, they are the product of the elimination of items that load poorly. Instead, we argue that theory-driven attempts to enhance the psychometric validity of the ZTPI are the way forward. For example, Worrell et al. (in press) reported acceptable cross-cultural indices for a shortened version of the ZTPI which retained only items with a specific temporal content (e.g. '*past*', '*tomorrow*', '*future*', etc.). Such an approach contrasts with those which simply retain high loading items in one sample, whose loadings are often not replicated in the next sample.

The three most recently published shortened versions of the ZTPI are the focus of this study. All of these versions retained the original five-factor structure of the ZTPI, though included different combinations and numbers of the original 56 items.The first is a 20-item Hebrew version (ZTPI-20), with four original ZTPI items in each factor (Orkibi, 2015); the second, a 15-item version developed for use in Czech and Slovak adults (ZTPI-15), with three items in each factor (Košt'ál et al., 2016); while the third is a 17-item Hungarian version (ZTPI-17; Orosz et al., 2017). These 'new scales' each resulted from either data-driven approaches (e.g. using the highest loading items or permitting correlated error terms to improve fit), or through the deliberate choice of items that loaded well in other versions of the ZTPI.

We contend that such atheoretical approaches for the development of short versions of the ZTPI limit the generalizability of the resulting scales. Moreover, as concluded in relation to previous examinations of similarly derived short versions of the ZTPI (McKay et al., 2015; Perry et al., 2015), we are of the opinion that shortened versions of the ZTPI are only useful to the development of study of time perspective if two important criteria are fulfilled. Specifically: (1) they must yield scores that demonstrate good fit indices across multiple samples and cultures that support the five-factor structure and (2) they must demonstrate good internal consistency. Thus, the purpose of the present study was to examine the factorial validity and reliability of these latest shortened versions of the ZTPI in multiple samples from different cultures to determine if they meet these criteria.

Method

Participants

Data from five independent samples were analysed. Participants in the British Adolescent sample were 913 pupils (aged 12–16; 49.8% male) from 10 High schools in Northern Ireland. A total of 943 questionnaires were completed with 913 included in analyses. Thirty were excluded as a result of having been partially completed or spoiled.

Participants in the United States study were 816 academically talented adolescents (aged 11–18; 46.6% male) attending a summer programme at a research university in a Western state. Students were accepted into the summer programme using several criteria, including school achievement, teacher recommendations, and an academic product. Participants were predominantly in the seventh to 11th grades.

Participants in the Australian sample were a general population sample of 667 (aged 17–70, $M = 29.45$; 67.8% female) recruited online through social media (e.g. Facebook, forums), via email snowballing, and through posts placed on university learning management systems (e.g. Moodle, Blackboard).

Participants in the British Undergraduate sample were 455 university undergraduates (aged 18–25; 49.7% male) recruited from a university in the North West of England through opportunistic and snowball sampling.

Participants in the Slovenian sample were 425 adolescents and young adults (aged 15–29, 70.4% female), who completed an online questionnaire sent to them via email or social media (e.g. Facebook). The scale was adapted to the Slovenian language using the back-translation technique (Geisinger, 2003).

Measures

The ZTPI (Zimbardo and Boyd, 1999) is a 56-item scale measuring time perspective in five factors: Past Negative (PN), Past Positive (PP), Present Fatalistic (PF), Present Hedonistic (PH), and Future (F). PN reflects a negative or aversive view of the past (e.g. '*I think about the bad things that have happened to me in the past*') and PP reflects a warm, sentimental attitude towards the past (e.g. '*Happy memories of good times spring readily to mind*'). Scores on PH reflect an orientation towards present pleasure with little concern for future consequences (e.g. '*Taking risks keeps my life from becoming boring*') whereas PF describes a helpless and hopeless attitude towards the future and their life (e.g. '*My life is controlled by forces I cannot influence*'), and F indicates behaviour dominated by striving for future goals and rewards (e.g. '*When I want to achieve something, I set goals and consider specific means for reaching those goals*'). Responses were on a 5-point Likert-type scale from 1 (*very unlike me*) to 5 (*very like me*).

The ZTPI-20 is comprised of 20 items, with four assigned to each of the five factors: PN, PP, PH, PF, and F.Orkibi (2015) reported the following model fit indices for the 20 items: comparative fit index (CFI) = .895, Tucker–Lewis index (TLI) = .912, root mean square error of approximation (RMSEA) = .054. Moreover, internal reliability estimates (Cronbach's α) for the five factors were as follows: PP = .69, PN = .80, PH = .73, PF = .65, F = .70.

The ZTPI-short (referred to here as the ZTPI-15; Košt'ál et al., 2016) is comprised of 15 ZTPI items, with three items in each of the five factors.

These authors reported the following model fit indices for their 15-item version: CFI = .944, TLI = .921, RMSEA = .047.

The ZTPI-17 (Orosz et al., 2017) is comprised of 17 items with four PN and F items, and three items in the remaining three factors. The authors reported the following model fit indices for their 15-item version: CFI = .953, TLI = .941, RMSEA = .040.

For all three previous studies and the current one, participants respond to questions using the ZTPI's original 5-point Likert scale (1 = *very uncharacteristic of me*; 5 = *very characteristic of me*). Furthermore, it should be noted that participants in the current study and ZTPI-20 and ZTPI-17 studies completed the 56-item ZTPI, while the authors of the ZTPI-15 study actually tested an 18-item version of the scale by adding three researcher-derived items to create distinct future negative and positive dimensions. For a full list of items in each version of the scale tested in the present study, please see Appendix 1.

Statistical analyses

Model fit was assessed using confirmatory factor analyses (CFAs) and exploratory structural equation modelling (ESEM) in Mplus7 (Muthén and Muthén, 1998–2012) and the MLM estimator. The MLM maximum likelihood parameter estimates with standard errors and a mean-adjusted chi-square (χ^2) test statistic that is robust to non-normality. The MLM χ^2 test statistic is also referred to as the Satorra-Bentler χ^2. While in CFA items are restricted to load on individual factors, constraining coefficients on all other factors to zero, ESEM allows for non-significant cross-loading of items. This approach enables freely estimated cross-loadings, has less restrictive assumptions than CFA, and potentially provides more valid estimates (Marsh et al., 2012).

A five-factor model for the ZTPI was assessed. The indices used to test model fit were χ^2, CFI, TLI, RMSEA, and the standardized root mean square residual (SRMR). Although Hu and Bentler's (1999) cut-offs for 'good' fit (i.e. > .95 for CFI and TLI, < .06 for RMSEA, and < .08 for SRMR) and 'acceptable' fit (i.e. > .90 for CFI and TLI, < .08 for RMSEA, and < .08 for SRMR) are typically cited, Marsh et al. (2004) suggested that strict adherence to these cut-off values is likely to lead to erroneous results, as factor loadings in social sciences are typically lower (see, e.g. Heene et al., 2011).

Internal consistency was assessed in a variety of ways. As has been noted by many (e.g. Cortina, 1993), Cronbach's alpha is largely a function of the number of items in a scale. Schmitt (1996) stressed the point that alpha alone is not a clear representation of reliability, and that researchers should present a variety of estimates, one of which should be the inter-correlations of items. Consequently, we also calculated mean inter-item correlations (MICs), omega point estimates and confidence intervals, and average variance extracted (AVE). As omega has fewer assumptions than alpha, problems associated with inflation of internal consistency are less likely (Dunn et al., 2013).

Results

Results of CFA analyses for the full ZTPI, the ZTPI-20, ZTPI-17, and ZTPI-15 are reported in Table 1. For all scales, the absolute fit indices (SRMR and RMSEA) were within or around the acceptable .08 and .05 values, respectively. However, in the case of all scales the chi-square values were significant, and the relative indices (CFI and TLI) fell short of the optimal .95 values and, in many cases, acceptable fit (i.e. .90).

Results of the ESEM analyses for all of the short scales are displayed in Table 2. In each case there was a marked improvement for all scales in terms of the absolute indices, and of note, the chi-square values for four out of the five samples were non-significant for the ZTPI-15 and two samples achieved this in the ZTPI-17. Such a dramatic improvement in model fit is the result of the estimation of non-significant cross-loadings. Across all samples, only one item (PP20) loaded substantively (> .30) onto another factor and this was only evident in the Australian sample.

To examine the extent to which the measurement model was consistent across samples, we performed multigroup CFAs on each measure, using sample as the grouping variable. Measurement invariance was examined on a series of increasingly constrained models. First, configural invariance was assessed by replicating the CFA-ICM (independent cluster model) across all samples. Second, factors were constrained to test metric invariance. Third, we examined scalar invariance by constraining factors and item intercepts. Finally, residual variance was tested by factors, item intercepts, and factor means. Model invariance is supported by little or no change in model fit on the increasingly constrained models. Cheung and Rensvold (2002) suggested $\Delta CFI \leq .01$, although Meade et al. (2008) suggested a stricter criterion of $\Delta CFI \leq .002$ to support invariance. The results of invariance testing are presented in Table 3. Generally, all versions displayed similar findings in that configural and metric invariance were largely supported but scalar invariance was not, indicating that there are differences in the way the items are interpreted in the different samples.

Cronbach's alpha reliability estimates, MIC values, omega point estimates and confidence intervals, AVE, and factor correlations are displayed in Supplementary Material Table 1 for the ZTPI-20, in Table 2 for the

Table 1. Confirmatory factor analysis model fits for shortened versions of the ZTPI.

	$\chi^2_{s\text{–}b}$	*df*	CFI	TLI	SRMR	RMSEA (90% CI)
CFA ZTPI						
American sample	4484.84*	1474	.653	.634	.081	.050 (.048, .052)
Australian sample	4297.11*	1474	.612	.591	.083	.058 (.055, .059)
British adolescent	4517.21*	1474	.713	.703	.073	.048 (.046, .049)
British university	3629.35*	1474	.664	.652	.081	.057 (.054, .059)
Slovenian sample	3789.61*	1474	.651	.632	.102	.060 (.058, .063)
CFA ZTPI-20						
American sample	408.50*	160	.898	.879	.054	.044 (.038, .049)
Australian sample	543.90*	160	.785	.745	.073	.064 (.058, .070)
British adolescent	443.19*	160	.911	.894	.051	.044 (.039, .049)
British university	352.71*	160	.889	.868	.059	.051 (.044, .059)
Slovenian sample	407.68*	160	.842	.812	.070	.060 (.053, .068)
CFA ZTPI-17						
American sample	333.53*	109	.891	.864	.059	.050 (.044, .056)
Australian sample[a]	293.56*	109	.895	.869	.051	.051 (.044, .058)
British adolescent	324.16*	109	.917	.897	.050	.046 (.041, .052)
British university[a]	268.46*	109	.886	.858	.059	.057 (.048, .065)
Slovenian sample	324.82*	109	.858	.823	.079	.068 (.060, .077)
CFA ZTPI-short						
American sample	190.88*	80	.925	.901	.042	.041 (.034, .049)
Australian sample	236.43*	80	.847	.799	.059	.058 (.049, .066)
British adolescent	207.34*	80	.919	.894	.045	.042 (.035, .049)
British university	221.69*	80	.865	.822	.057	.062 (.053, .072)
Slovenian sample	208.82*	80	.879	.841	.054	.062 (.051, .072)

$\chi^2_{s\text{–}b}$: Satorra–Bentler adjusted chi-square; *df*: degrees of freedom; CFI: comparative fit index; RMSEA: root mean square error of approximation; SRMR: standardized root measure square residual; TLI: Tucker–Lewis index; ZTPI: Zimbardo Time Perspective Inventory.
[a]Non-positive definite residual covariance matrix (both related to item 20).
*Statistically significant at $p < .001$.

ZTPI-17, and in Table 3 for the ZTPI-15. Omega point estimates and confidence intervals were calculated using the MBESS package (Kelley and Lai, 2012) in R (R Development Core Team, 2012) with 1000 bootstrapped samples. Results for the ZTPI-20 show that only seven out of the 25 alpha and omega values reached the .70 threshold for acceptable reliability, and five of these were for the PN factor (> .70 in all five samples). The same pattern was evident for omega point estimates. Only four MICs

Table 2. ESEM model fits for shortened versions of the ZTPI.

	$\chi^2_{s–b}$	*df*	CFI	TLI	SRMR	RMSEA (90% CI)
ESEM ZTPI						
American sample	3100.73*	1270	.823	.790	.042	.041 (.040, .044)
Australian sample	2843.40*	1270	.811	.773	.041	.047 (.044, .048)
British adolescent	3068.98*	1270	.853	.824	.041	.040 (.038, .041)
British university	2832.84*	1270	.782	.732	.042	.052 (.049, .055)
Slovenian sample	2678.21*	1270	.822	.782	.041	.050 (.048, .054)
ESEM ZTPI-20						
American sample	205.40*	100	.964	.931	.022	.036 (.029, .043)
Australian sample	246.74*	100	.929	.865	.033	.050 (.042, .058)
British adolescent	183.93*	100	.976	.955	.021	.030 (.023, .037)
British university	207.62*	100	.946	.898	.029	.049 (.039, .058)
Slovenian sample	181.72*	100	.956	.917	.027	.044 (.034, .054)
ESEM ZTPI-17						
American sample	127.66*	61	.973	.940	.020	.037 (.028, .046)
Australian sample	101.79*	61	.980	.956	.019	.032 (.021, .043)
British adolescent	76.45	61	.995	.988	.014	.017 (.000, .027)
British university	109.98*	61	.968	.930	.024	.042 (.029, .054)
Slovenian sample	100.03	61	.979	.953	.021	.039 (.024, .052)
ESEM ZTPI-Short						
American sample	68.81	40	.984	.957	.016	.030 (.017, .041)
Australian sample	93.23*	40	.958	.890	.025	.048 (.035, .060)
British adolescent	59.57	40	.989	.972	.015	.023 (.009, .035)
British university	68.15	40	.976	.937	.020	.039 (.022, .055)
Slovenian sample	69.45	40	.977	.939	.022	.042 (.024, .058)

CFI: comparative fit index; *df*: degrees of freedom; RMSEA: root mean square error of approximation; SRMR: standardized root measure square residual; TLI: Tucker–Lewis index; $\chi^2_{s–b}$: Satorra–Bentler adjusted chi-square; ZTPI: Zimbardo Time Perspective Inventory.
*Statistically significant at $p < .001$.

were greater than .40, all of which were from the PN factor. The PN factor also consistently produced AVE > .40, which was only found in one other instance. Alpha values were all greater than .70 on the PN factor for the ZTPI-17, although only three other values of the remaining 20 reached this level. In total, eight omega point estimates reached a .70 level, again mainly deriving from the PN factor. Reliabilities for the ZTPI-15 show that only three alpha values reached the .70 threshold, and all were for the PN factor. A further two omega point estimates reached .70. However, given that

Table 3. Measurement invariance across samples for shortened versions of the ZTPI.

Scale	$\chi^2_{s\text{–}b}$	*df*	$\Delta\chi^2_{s\text{–}b}$	Δ*df*	CFI	ΔCFI	TLI	SRMR	RMSEA (90% CI)
ZTPI-20									
Configural invariance	1915.23*	800	–	–	.898	–	.879	.056	.046 (.044, .049)
Metric invariance	2084.79*	860	168.77	60	.888	.010	.877	.060	.047 (.044, .049)
Scalar invariance	3463.86*	920	1378.21	60	.768	.120	.761	.074	.065 (.063, .067)
Residual invariance	4218.03*	940	754.17	20	.701	.067	.698	.099	.073 (.071, .075)
ZTPI-17									
Configural invariance[a]	1544.98*	545	–	–	.893	–	.866	.058	.053 (.050, .056)
Metric invariance	1751.50*	593	206.52	48	.876	.017	.857	.064	.055 (.052, .058)
Scalar invariance	2483.62*	641	732.12	48	.802	.070	.790	.073	.066 (.064, .069)
Residual invariance	3258.48*	661	774.86	20	.721	.081	.713	.102	.078 (.075, .080)
ZTPI-15									
Configural invariance	986.31*	400	–	–	.910	–	.882	.048	.047 (.044, .051)
Metric invariance	1099.91*	440	113.60	40	.899	.011	.879	.052	.048 (.044, .052)
Scalar invariance	1997.92*	480	898.01	40	.767	.132	.745	.070	.070 (.066, .073)
Residual invariance	2765.10*	500	767.18	20	.653	.114	.635	.111	.083 (.080, .086)

CFI: comparative fit index; *df*: degrees of freedom; RMSEA: root mean square error of approximation; SRMR: standardized root measure square residual; TLI: Tucker–Lewis index; $\chi^2_{s\text{–}b}$: Satorra–Bentler adjusted chi-square; ZTPI: Zimbardo Time Perspective Inventory.
[a]Non-positive definite residual covariance matrix (related to item 20).
*Statistically significant at $p < .001$.

Clark and Watson (1995) make a convincing argument for why estimates of .60 are acceptable for research purposes, especially when applied to broad constructs such as time perspective, we decided to additionally compute MICs and AVE for all factors.

MICs were largely between .20 and .40, which is in line with the observed internal consistency estimates for alpha and omega. AVE was calculated by averaging the squared standardized parameter estimates for each factor. Fornell and Larcker (1981) recommend a value of at least .50 to represent a meaningful amount of explained variance, though this is subject to scale length. AVE $\geq$ .50 was achieved in only three of the 25 values calculated in the ZTPI-20 (all PN factors), only once in the ZTPI-17, and not at all in the ZTPI-15. Of greater concern were the particularly low AVE values however. The ZTPI-20 yielded nine values < .30 from the CFA loadings. The ZTPI-17 presented five factors with AVE < .30 and the ZTPI-short had four values below < .30. Noticeable is that the PP factor was the least reliable in terms of AVE on the ZTPI-20 but performed markedly better in the ZTPI-15 and in between on the ZTPI-17. The opposite was true for the PH factor. Only seven scales from the 25 reached a MIC of .40 or greater, only one of which was > .50. AVE was typically greater than .40 for the PN and PH factors but lower for the PP, PF, and F factors.

Discussion

The present results clearly demonstrate that, while all of the ZTPI short versions examined in the present study presented some evidence of factorial validity to a greater or lesser extent, all of the scales presented significant reliability concerns. Further, measurement invariance is consistently unsatisfied, which is problematic for a concept that is used to identify cultural differences. Each has achieved improved model fit on the original by simplifying the model. However, simply eliminating items based on individual samples serves to sacrifice internal consistency in favour of model fit. In essence, this is a short cut. The real solution to overcoming psychometric problems should be to develop theoretically based new items or to determine a new theoretical approach to assessing the construct.

However, it should be recognized that when abbreviating scales, alpha will necessarily decrease (Streiner, 2003), but this does not necessarily have to signify a decrease in reliability. In fact, low internal consistency might be seen as a positive when it comes to short scales because the presence of heterogeneous items may maximize the area of the domain covered while allowing for an increase in the efficiency of data collection. The average inter-item correlations in the present study suggest that this might be what is happening in the present data, as the correlations are not so high as to imply redundancy and not so low as to imply that the items are measuring different constructs. Additionally, the relatively good omegas, considering the short length of the scales (> .50 in most cases), suggest that each domain taps into a relatively reliable (or at least defined) general factor – these estimates are higher than the omegas obtained for the so-called General Factor of Personality in most cases.

The ZTPI versions in the present study are the fourth, fifth, and sixth such scales that we have assessed using some or all of these datasets in recent times (McKay et al., 2015; Perry et al., 2015). However, in each case we have been unable to replicate the results reported by the scale developers.The adequate to poor CFI indices reported herein are slightly higher overall than those previously reported (McKay et al., 2015; Perry et al., 2015) for the scales proposed by Zhang et al. (2013; $CFI = .81 < CFI < .90$), Laghi et al. (2013; $CFI = .74 < CFI < .86$), and Sircova et al. (2014; $CFI = .72 < CFI < .78$).Similarly, the RMSEA values in the present study were marginally better than those reported by Zhang et al. (2013; $RMSEA = .05 < CFI < .07$), Laghi et al. (2013; $RMSEA = .06 < CFI < .09$), and Sircova et al. (2014;

RMSEA = .05 < CFI < .06). Nevertheless, as well as unacceptable factorial validity, reliability estimates in the present study were unsatisfactory. Specifically, although the alpha reliability estimates reported by Košt'ál et al. (2016) and Orkibi (2015) were all above or around 0.70, these values were not replicated in the present samples. The alpha values for PP, PH, and F were particularly problematic in all samples and for all versions of the ZTPI. The inability to replicate the findings of the developing authors does not bode well for the conceptual development of the construct.

An examination of the items used in all six short forms of the ZTPI reveals that 45 of the 56 items have been used, and that only four items (#2, #10, #20, and #50) appear in all versions of the measure. One possible future direction is to apply a more theoretical approach to amending the ZTPI, perhaps by attempting to eliminate items that do not measure time perspective but do measure related constructs. Researchers could also consider different measures altogether, such as those that focus on more specific temporal domains. It may be that the time perspective construct is too broad for a single instrument and it would be better assessed with temporal domain-specific measures that assess distinct dimensions, for example, the Adolescent Time Inventory (Mello and Worrell, 2015), the Consideration of Future Consequences Scale (Strathman et al., 1994), or the Temporal Focus Scale (Shipp et al., 2009).

The temporal psychology literature is increasingly becoming fractured, and the multiplicity of 'versions' of scales is not helping advance our conceptualization or application of the construct. To be fully meaningful, the assessment of this construct needs to be able to transcend cultural variance. The results of the present study should be interpreted in the context of problems assessing 'short forms' of scales administered in their original longer format (Knowles and Condon, 2000), in particular the fact that responses to items on scales

> often involves more than responding to the semantic content of the item. Respondents interpret the items within a context. As the context for an item changes, even as its position in the test changes, the meaning of the item may shift. (p. 250)

However, in conclusion, the shortened versions of the ZTPI examined here did not achieve the psychometric criteria that would justify their intended purpose. We contend that the advancement of the international study of time perspective, and the resolution of previously identified measurement problems (e.g. Worrell and Mello, 2007), is not served by the atheoretical development of a multiplicity of short forms of the ZTPI.

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Supplementary material

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Appendix 1

	Past positive	Past negative	Present hedonistic	Present fatalistic	Future
ZTPI-20	Items 2, 20, 29, 49	Items 16, 22, 34, 50	Items 1, 31, 42, 55	Items 14, 37, 38, 39,	Items 10, 13, 45, 51
ZTPI-15	Items 2, 7, 20	Items 4, 50, 54	Items 26, 42, 46	Items 14, 38, 39	Items 10, 40, 45
ZTPI-17	Items 15, 20, 29	Items 22, 25, 34, 50	Items 31, 42, 46	Items 37, 38, 39	Items 13, 21, 40, 45

ZTPI: Zimbardo Time Perspective Inventory.