DOCUMENT RESUME

ED 264 290 TM 850 780

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TITLE	Comparing the Long and Short Forms of the Student Version of the Jenkins Activity Survey.
PUB DATE	[85]
NOTE	29p.
PUB TYPE	Reports - Research/Technical (143)
EDRS PRICE	MF01/PC02 Plus Postage.
DESCRIPTORS	Behavior Rating Scales; College Students; *Comparative Testing; *Factor Analysis; *Factor Structure; Goodness of Fit; Higher Education; Mathematical Models; Measurement Techniques; *Personality Measures; *Test Theory
IDENTIFIERS	*Jenkins Activity Survey (Student Version); LISREL Computer Program; Person Product Moment Correlation; Type A Behavior; Type B Behavior

ABSTRACT

This paper reports on a short version of the Student Jenkins Activity Survey (JAS), a multiple choice questionnaire that measures Type A "coronary-prone" behavior in assessing subjects' A/B types. The primary objective was to determine if the short and long forms of the student JAS represent similar measurement instruments. A secondary objective was to determine if the short and/or the long JAS results in comparable factor patterns and covariance structures for As and Bs. The short Student JAS gives factors that are more independent than those of the long JAS, and provides a more comparable structure for Type As and Bs. The present results suggest that researchers should still use the short form since its structure most closely resembles that hypothesized by the underlying theory. An important finding is that neither the Glass orthogonal nor the oblique two-factor model fit the data well in an absolute sense: less than half of the common variance was explained by any model, on either form. Examination of the residual correlation matrices indicated that two classes of large residuals (self versus perceived others' ratings and actions versus feelings), which were largely responsible for the significantly poor fits, were observed on every LISREL computer run. (PN)

ED264290

Comparing the Long and Short Forms of the Student Version of the Jenkins Activity Survey

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Running Head: Short versus long JAS

Research examining Type A "coronary-prone" behavior has proliferated over the past decade (Dembroski, Weiss, Shields, Haynes, & Feinleib, 1978; Friedman & Rosenman, 1974; Glass, 1977; Matthews, 1982; Price, 1982). Although no prospective research with students has yet been reported, research with adults suggests that the standardized stress interview method of assessing level of Type A behavior provides a superior (i.e., more valid), qualitatively different measure than questionnaire assessments (see, e.g., Byrne, Rosenman, Schiller, & Chesney, 1985; Matthews, Krantz, Dembroski, & MacDougall, 1982; Mueser & Yarnold, 1985; Yarnold, Mueser, & Grimm, 1985). However, due to the relative cost-efficiency of the questionnaire over the interview, many of the studies in this field have employed a self-report questionnaire measure of Type A in assessing subjects' A/B Types, and most of these have used the Student version of the Jenkins Activity Survey (JAS). A growing trend, however, is toward the use of multimeasure, multiattribute assessment of Type A behavior, and the therefore necessary development of shorter, faster, more precise measure procedures (e.g., Perlman, Hartman, & Lenahan, 1984; Tasto, Chesney, & Chadwick, 1978). Accordingly, this paper reports on a short version of the frequently used student JAS.

The Student JAS consists of 44 multiple choice questions of which 21 are actually scored using a unit-weighting procedure (Glass, 1977; Yarnold & Mueser, 1985). Data from our laboratory suggests that randomly sampled undergraduates require a mean of 10.2 minutes (sd = 2.77, N = 52) to complete the 44-item (long) version, as compared with a mean of 4.3 minutes (sd = 1.25) to complete the 21-item (short) version, which contains only the scored items [$t(51) = 14.49$, $p < .0001$]. Thus, administering the short version of the JAS instead of the long version should save an average of at least 5.9 minutes.

For researchers who mass-screen and/or use college undergraduates participating in exchange for "subject pool" credit, this time savings may be of considerable practical importance. Indeed, reducing the time required to assess individual's A/B Types could help simplify such logistical operations as scheduling subjects, experimenters, and facilities.[1]

Before researchers may use the short JAS, however, it must be determined whether the short and long versions measure the same construct. That is, one must address the question of whether independent, random samples of students who complete versions of the 21 and 44-item Student JAS have score distributions and factor structures that are comparable. No previous research has yet addressed this issue. Accordingly, the present study was designed to evaluate the relative fit of different measurement models for the JAS and to test hypotheses about equality of factor structures and score distributions between the short and long form.

Method

Subjects

A total of 1,248 university undergraduates (approximately equal numbers of men and women) participated in the study in exchange for class credit. Data were collected during mass-testings of between two- and four-hundred students, in which either the long (44-item) or short (21-item) JAS was included as one of a battery of questionnaires being distributed. Half of the subjects, randomly selected, completed the long version (N=624), while the other half completed the short version (N=624). JAS scores were later used to identify groups of Type As and Type Bs. Subjects with scores above the median (7 on both forms) were classified as Type A and subjects with scores below the median were classified as Type B (see Glass, 1977, for a discussion regarding the "median-split" method of assignment to A/B categories). For the short form of the JAS, there were 236 Type As and 234 Type Bs; for the long form, 253 As and 225 Bs.

Procedure

Subjects' responses to the 21 A/B items were coded according to Glass (1977), using all items from the short form and only the 21 scored items from the long form. To scale variables in a comparable metric across groups, all items were first standardized separately for As and Bs within both the long and short forms before being intercorrelated (see Cunningham, 1978; Sorbom & Joreskog, 1976). As input data for factor analyses, we built separate Pearson product-moment intercorrelation matrices of these items for the short-form sample and the long-form sample.[2] We also built separate intercorrelation matrices for Type As and Type Bs for both the short and long form.

Analysis Strategy for Structural Comparisons

Model testing. We performed separate confirmatory factor analyses on each group's data using LISREL IV (Joreskog & Sorbom, 1978) to test three alternative measurement models. The first model had two orthogonal factors -- "hard-driving" (Factor H) and "speed and impatience" (Factor S) -- that emerged in Glass' (1977) original work on the student version of the JAS. These factors replicated those found earlier for adults, the difference being that the job-involvement factor of the adult JAS was missing -- by design -- on the student JAS (Glass, 1977; Yarnold & Mueser, 1985). To estimate this model using LISREL, factor loadings from Glass' (1977) results that were below .35 were fixed at zero, and factor loadings above .35 were designated as free parameters to be calculated by the program. This two-factor model was not intended to explain as much variance in JAS scores as possible, but rather was designed as a measurement model to represent Glass' (1977) independent dimensions of coronary-prone behavior. To correspond with Glass' (1977) original orthogonal analysis, this model constrained the correlation between Factors H and S to be zero. Table 1 presents the JAS items and LISREL factor loadings for this orthogonal two-factor model.

Insert Table 1 About Here

The second measurement model that we tested was an alternative, oblique version of the orthogonal two-factor model. The only difference between the first and second models was that the former forced the correlation between factors to be zero, whereas the latter designated this factor intercorrelation as a free parameter to be calculated by the LISREL program. This allowed us to test empirically the hypothesis that Factors H and S are uncorrelated, by directly comparing the fit of the orthogonal and oblique models.

The third measurement model assumed that the only source of variation in each variable was sampling error and represented a baseline against which to compare the fit of the other models (see Alwin & Jackson, 1979; McGaw & Joreskog, 1971). To estimate this zero common-factor model, we gave each variable its own factor, specified an identity matrix of factor loadings, fixed all correlations among factors to zero, and designated as a free parameter the variance of unique error for each variable. As with all other LISREL models, the unique error in each variable was specified to be independent of the unique error in other variables. We used this zero common-factor model to compute a measure of relative fit -- the Tucker-Lewis coefficient (Tucker & Lewis, 1973) -- for each of the two-factor models. This coefficient reflects the improvement in variance explained by the particular model over a model that assumes there is no common variance (see Bryant & Veroff, 1982, 1984, for the formula and procedural details in computing this coefficient).

Hypothesis testing. Each LISREL analysis produced a maximum-likelihood chi-square and degrees of freedom that we used to test three sets of hypotheses about factor structure. The first set of hypotheses concerned whether the oblique two-factor model provided a significant improvement in fit over the orthogonal two-factor model. We tested improvement in fit by examining the statistical significance of the difference between the chi-square values obtained using each two-factor model (see Bentler & Bonett, 1980). We did this separately for the pooled sample, for Type As, and for Type Bs, within both the short and long forms of the JAS.

The second set of hypotheses concerned whether the short and long forms of the JAS have a comparable factor structure. To evaluate the equality of factor loadings between the short and long forms, we used LISREL IV (Joreskog & Sorbom, 1978) to perform simultaneous confirmatory factor analysis. We did this separately for the data of the pooled sample, of Type As, and of Type Bs, using both the orthogonal and oblique two-factor models.

The third set of hypotheses concerned whether Type As and Type Bs have a comparable factor structure. To evaluate the equality of factor loadings between A/B groups, we again used simultaneous confirmatory factor analysis. We did this separately for short and long forms using both the orthogonal and oblique models.

Analysis Strategy for Distributional Comparisons

The long and short forms of the JAS each resulted in distributions of 624 replications of 21 items. Since both distributions represented random samplings of undergraduate psychology students, the means and variances of total and Factor H and S JAS scores between distributions should be comparable, as should means and covariances across all 21 JAS items. The alternative hypotheses that differences in means and covariances between forms existed were evaluated using multivariate analyses of variance and chi-square tests, respectively (Green, 1978).

Results

Significance Levels

Because the present study involved multiple statistical comparisons, we decided to make our alpha level more stringent to avoid capitalizing on Type 1 errors (Cook & Campbell, 1979). Specifically, we divided our desired alpha (i.e., .05) by the number of statistical comparisons that we planned in order to obtain a new alpha adjusted for the error-rate per experiment (see Ryan, 1959). We planned 28 statistical comparisons of structural hypotheses (3 model contrasts × 3 datasets × 2 forms, plus comparison of the short and long forms within 3 datasets × 2 models, plus comparison of As and Bs for 2 forms × 2 models = 28). For each subject we also created two profiles: one profile consisted of the total JAS score and Factor H (hard-driving/competitive) and Factor S (speed/impatience) scores (the "3-scale" profile); and the other profile consisted of all 21 scored JAS items (the "21-item" profile). We planned a 2(A/B Type) × 2(short/long form) multivariate analysis of variance for each profile (2 additional statistical comparisons). Finally, we planned 8 statistical comparisons of covariance matrices (4 comparisons for each profile: one between forms pooling over A/B Type, one between A/B Types pooling over form, and one between forms for each A/B Type separately). Thus, we planned a total of 38 statistical comparisons. To achieve an actual alpha level of .05, we thus used $\underline{p} < .002$ as a criterion for establishing statistical significance with the present data (i.e., $.05/38 < .002$).[3]

Model Testing

Table 2 displays the chi-square statistics and measures of relative fit for the three LISREL models for both the short and long form of the JAS.

Insert Table 2 About Here

Focusing first on the pooled sample, it is evident that none of the models fits the data well in an absolute sense. The ratio of chi-square to degrees of freedom, which approaches zero as the fit of the given model improves (Joreskog, 1971), revealed that all three models provide a relativaly poor fit for both the pooled short form and pooled long form samples. Tucker-Lewis coefficients also indicated that the orthogonal and oblique models each account for less than half of the common variance on the short and long form. Neither of these models, however, was designed to maximize the amount of explained variance. Indeed, the purpose of the present study was not to find the absolute best-fitting model for the JAS, but rather was to compare the relative fit of alternative measurement models. Clearly, further research is necessary to develop a more precise measurement model for the JAS.

Turning to the samples of As and Bs, we see that the ratio of chi-square to degrees of freedom for each LISREL model is smaller for these groups than for the pooled sample with both the short and long form. This is not surprising -- since the maximum-likelihood chi-square is a function of sample size (Bentler & Bonett, 1980; Joreskog & Sorbom, 1978), the smaller samples of As and Bs would be expected to reduce the ratio of chi-square to degrees of freedom. Tucker-Lewis coefficients again indicated that the orthogonal and oblique models each account for less than half of the common variance on the short and long forms. For this reason, the reduction in the ratio of chi-square to degrees of freedom associated with the A/B samples appears to be largely an artifact of reduced sample size.

Direct contrasts of chi-square values and degrees of freedom confirmed that each of the two-factor models represents a significant improvement in fit over the zero common-factor model (see Table 3). This was true for both the short and long forms as well as for the pooled, Type A, and Type B samples (all p's $< .0001$). The orthogonal and oblique models thus fit the data better than the zero common-factor model, although they each account for less than half of the explainable variance in JAS scores.

Insert Table 3 About Here

Hypothesis Testing

Orthogonality of the two-factor model. One purpose of the present study was to determine whether the oblique version of the two-factor model fit the data better than the orthogonal version. Table 3 presents the comparative statistics contrasting these two models. Adjusting alpha for the per experiment error rate (Ryan, 1959), the oblique model provided a significant improvement in fit over the orthogonal model using the long form of the JAS, but not using the short form. This effect held only for the pooled sample and not for As or Bs examined separately -- a result consistent with the fact that smaller sample sizes reduce statistical power and limit our ability to detect true structural differences (Bentler & Bonett, 1980). Thus, letting Factors H and S correlate improved the fit of the two-factor model significantly when using the long form of the JAS. LISREL estimates of the oblique model using the pooled samples indicated that the two factors correlated .17 for the long form and .09 for the short form. Evidently, Factors H and S are more independent when using the short form of the JAS.

Structural equivalence of the short and long forms. Another purpose of the present study was to determine whether the short and long forms of the JAS have a comparable factor structure. We tested the equality of factor loadings between forms by contrasting the chi-square values obtained from two types of simultaneous factor analyses. The first type specified that the short and long forms had identical factor loadings, while the second specified that factor loadings be computed separately for the short and long forms. We performed both types of simultaneous analyses separately for the pooled, Type A, and Type B samples using both the orthogonal and oblique models. The difference between the chi-square values (with their accompanying difference in degrees of freedom) obtained from these two types of analyses was used to test the hypothesis that the particular model provided an equivalent fit for both forms (see Alwin & Jackson, 1979, 1980; Bryant & Veroff, 1982, 1984; Joreskog, 1971).

Using the adjusted alpha-level with the pooled sample, the orthogonal model fit both forms equally well ($\chi^2 = 25.767$, df = 9, $p < .01$), whereas the oblique model produced a significant overall difference in factor loadings between forms ($\chi^2 = 28.128$, df = 9, $p < .001$). With the sample of Type As, both forms had a comparable factor structure using either the orthogonal ($\chi^2 = 15.730$, df = 9, $p < .10$) or the oblique model ($\chi^2 = 15.642$, df = 9, $p < .10$).

With the sample of Type Bs, there were significant structural differences between forms using both the orthogonal ($\chi^2 = 52.210$, df = 9, $p < .0001$) and oblique models ($\chi^2 = 55.226$, df = 9, $p < .0001$). Considered together, these results suggest that the orthogonal model provides a more equivalent factor structure between forms than does the oblique model, although neither model yields an equivalent structure across forms for the sample of Type Bs.

Structural equivalence for As and Bs. We also wished to determine whether the factor structures of the short and long forms were equivalent for As and Bs. We tested the equality of factor loadings between groups of As and Bs by again performing two types of simultaneous factor analyses. The first type specified that the two A/B groups had identical factor loadings, while the second specified that factor loadings be computed separately for As and Bs. Both types of analyses were performed separately for the short and long forms using both two-factor models.

Using the adjusted alpha-level with the short form, Type As and Type Bs had an equivalent factor structure using the orthogonal model ($\chi^2 = 27.537$, df = 9, $p < .01$), but showed a significant overall difference in factor loadings using the oblique model ($\chi^2 = 30.422$, df = 9, $p < .001$). With the long form, As and Bs had a different structure regardless of whether the orthogonal ($\chi^2 = 32.087$, df = 9, $p < .001$) or oblique model ($\chi^2 = 31.636$, df = 9, $p < .001$) was used. Taken as a whole, these findings suggest that the orthogonal model provides a more equivalent structure for Type As and Type Bs than does the oblique model, but only with the short form of the JAS.

Comparing Distributions of Total and Factor H and S Scores

Means and standard deviations of the total, hard-driving (Factor H) and speed-impatience (Factor S) JAS scores are given by A/B Type and form in Table 4.

Insert Table 4 About Here

In order to determine whether means on these scores varied as a main or interactive effect of A/B Type or form, a 2(A/B) × 2(short/long) multivariate analysis of variance (MANOVA) was performed. The results indicated a highly statistically significant main effect of A/B Type, $F(3,942) = 874.48$, $p < .0001$. Subsequent contrasts (analyses of variance; ANOVAs) revealed that Type As scored significantly higher total ($F = 2{,}615.48$; df = 1,945; $p < .0001$), Factor H ($F = 1{,}041.96$; df = 1,945; $p < .0001$) and Factor S ($F = 230.97$; df = 1,945; $p < .0001$) scores than Type Bs. There was no evidence supporting a multivariate main effect of form ($F = 0.48$; df = 3,943; $p < .71$), or an interaction between A/B Type and form ($F = 0.08$; df = 3,943; $p < .97$), however. Thus, As from the long- and short-form samples had comparable total and Factor H and S JAS scores, as did Bs. Further, scores of As on all three measures were significantly higher than those of Bs.

A chi-square test of the homogeneity of the covariance matrices of the short- and long-forms of the JAS (pooled over A/B Type) was not statistically significant, $\chi^2 = 3.56$, df = 6, $p < .74$. Thus, the short and long JAS resulted in homogeneous score distributions (i.e., for total and Factor H and S JAS scores).

A chi-square test of the homogeneity of the covariance matrices of Type As and Type Bs (pooled over form) was statistically significant, however, $\chi^2 =$

184.16, df = 6, $p < .0001$. Thus it appears that the covariance structures between total and Factor H and S JAS scores are not comparable for As and Bs. Examination of the covariance matrices, however, indicated that whereas the relative magnitudes and signs of the variance and covariance terms of As and Bs were in the same pattern and ordering, every term for the As had a greater absolute value than the corresponding term for Bs. This may reflect the effect of greater range restriction in the data of Type Bs, because the maximum possible (and of course, observed) range of scores on each of the three scales is greater for As than for Bs. The effect of range restriction (limiting the number of levels variables may assume) is often to constrain the maximum obtainable correlation (and therefore covariance) between variables to being less than 1.0 in absolute value (Bishop, Fienberg, & Holland, 1975). Thus, the significantly different covariance matrices of As and Bs may be artifactual -- reflecting range restriction and corresponding constrained absolute correlations -- rather than underlying differences in the nature of the interrelations among items.

Finally, the covariance matrices (for the three JAS scales) of the long and short JAS were contrasted separately for As and Bs. The results indicated that total and Factor H and S JAS scores of Type As on the long- and short-form were homogeneous, $\chi^2 = 4.52$, df = 6, $p < .61$, as were the covariance matrices of Type Bs on the long- and short-form, $\chi^2 = 7.32$, df = 6, $p < .30$. Thus, when As and Bs are examined separately, the 3-scale covariance matrices generated by the long- and short-form are homogeneous.

Comparing Distributions of 21 JAS Items

The analyses presented above were also performed on profiles consisting of all 21 scored JAS items, rather than on 3-scale profiles.[4] The MANOVA revealed a highly statistically significant A/B Type main effect, $F(21,924) = 80.84$, $p < .0001$. ANOVAs revealed that As scored significantly higher on all of the individual JAS items (all F's > 13.13, df's = 1,944, p's < .0003) except one -- maintaining a fixed work schedule over holidays ($F = 1.34$, df = 1,944, $p < .25$). There was no evidence supporting a multivariate main effect of form, $F(21,924) = 1.28$, $p < .18$, or for an interaction between A/B Type and form, $F(21,924) = 1.18$, $p < .27$, however. Thus, it appears that As on the long- and short-forms had comparable means on the JAS items; the same was true for Type Bs; and As scored significantly higher than Bs on the items.

A chi-square test suggested that the 21-item covariance matrices of the long and short JAS (pooled over A/B Type) were homogeneous, $\chi^2 = 270.23$, df = 231, $p < .04$. Thus, the long and short JAS resulted in homogeneous distributions of 21 JAS items.

The 21-item covariance matrices of As and Bs (pooled over form) were not homogeneous, however, $\chi^2 = 749.25$, df = 231, $p < .0001$. Examination of the matrices, as in the previous analysis, revealed that virtually all the terms in the Type A matrix were greater in absolute value than corresponding terms for Type Bs, but reflected the same pattern of interrelationships. Range restriction may thus be invoked again to explain this significant difference.

Finally, As had homogeneous 21-item covariance matrices across form, $\chi^2 = 275.1$, df = 231, $p < .03$, as did Type Bs, $\chi^2 = 254.2$, df = 231, $p < .15$. Thus, when As and Bs are examined separately, the 21-item covariance matrices generated by the long and short JAS are homogeneous.

Discussion

The primary objective of the present study was to determine whether the short- and long-forms of the Student JAS represent similar measurement instruments. Analyses of the factor structure each form generated, pooling over A/B Type, revealed that Glass' (1977) orthogonal two-factor model fit both forms equally well. When Type As were considered separately, the long- and short-forms resulted in comparable factor structures whether the orthogonal or oblique two-factor models were employed. For Type Bs the long- and short-forms resulted in different factor structures regardless of which two-factor model was chosen. However, analyses of 3-scale and 21-item mean profiles and covariance structures revealed that score distributions generated by the short and long JAS are comparable. Since the factor patterns and actual score distributions obtained using the long and short JAS are largely comparable, researchers may use the short-form without significantly affecting subjects' responses to the items, or their measured location along the A/B continuum.

A secondary objective was to determine whether the short and/or the long JAS results in comparable factor patterns and covariance structures for As and Bs. Factor analyses revealed that As and Bs had a similar structure on the short-form (using the orthogonal but not the oblique two-factor model), whereas the factor structure of As and Bs on the long JAS were statistically different regardless of which two-factor model was employed. The 3-scale and 21-item covariance structures of As and Bs were significantly different, but that may reflect range restriction in the Type B data. Nevertheless, covariance matrices of As were similar across form, as were the corresponding matrices of Type Bs. Thus we conclude that the short-form of the Student JAS results in structurally similar but less covariant scores/items for Bs relative to As, whereas the long-form results in significantly different factor patterns and covariances.

In summary, the short Student JAS gives factors that are more independent than those of the long JAS, and provides a more comparable structure for Type As and Type Bs. In addition, the short-form requires significantly less time (sixty percent) to complete. In situations where a time savings of approximately six minutes is of little practical significance, however, the present results suggest that researchers should still use the short-form since its structure most closely resembles that hypothesized by the underlying theory.

An important finding is that neither the Glass (1977) orthogonal nor the oblique two-factor model fit the data well in an absolute sense: less than half of the common variance was explained by any model, on either form. Examination of the residual correlation matrices indicated that two classes of large residuals -- which were largely responsible for the significantly poor fits -- were observed on <u>every</u> LISREL run. Specifically, inter-correlations between variables which measured (a) self versus (perceived) other ratings (e.g., "how would <u>you</u> rate yourself...," versus "how would <u>your spouse</u> (or best friend) rate you..."), and (b) actions versus feelings (e.g., "how often <u>do you</u>...," versus "how often <u>do you feel like</u>...") were significantly misrepresented in the present two-factor model. Of interest, preliminary data from our laboratory suggests that individuals' best friends are relatively poor judges of what their friend's actual Type A (JAS total score) scores are (N = 150, r = .30). This suggests that the first class of residuals (self versus perceived-others' ratings) is an important method factor to be included in future optimizing factor analyses, and may represent a potentially fruitful avenue to pursue toward altering the Type A's beliefs regarding the relationship between their Type A hyper-responding and others' evaluations of them (see e.g., Price, 1982).

Future research should also examine the impact of subject variables such as sex and ethnicity upon the factor pattern and score/item distributions of the JAS and other measures of Type A behavior (see e.g., Perlman et al., 1984). Such efforts should lead toward a final integration of measurement models and prospective validation for coronary-artery and heart disease.

Footnotes

Appreciation is extended to the test-scoring service of the University of Illinois at Chicago for assistance in data management, and to Janet Goranson for preparing the manuscript. Portions of the paper were presented at the 1985 Annual Meeting of the American Psychological Association (Los Angeles).
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[1] Approximately 95% of the subjects completing either form should require 1.96 standardized minutes or less to finish. Based upon our samples, this translates into 15.6 minutes for the long form, and 5.4 minutes for the short form. Thus, the relatively efficient short form appears to be a potential candidate for inclusion in large-scale mass screenings, in which time is scarce and availability limited. At the University of Illinois at Chicago, for example, the psychology department mass-screens approximately 4,000 undergraduates per academic year. Participating researchers are limited to ten minutes with each subject, calculated as the mean completion latency plus one standard deviation (i.e., the long JAS could not be included in this mass-testing, whereas the short JAS could). Similarly, we have been granted up to fifty minutes to interview/test each of 6,000 inmates incarcerated in a large metropolitan jail, and obviously wish to minimize A/B assessment time. Finally, such time-efficient procedures are more appropriate for researchers working with geriatric or clinical medical subject populations.

[2]Since the student JAS contains some items with two response alternatives, Kendall's tau might be a more appropriate measure of association (see, e.g., Veroff, Feld, & Gurin, 1962). However, when using Kendall's tau we obtained results parallel to those we report — which we present in order to increase comparability to related studies (which typically report parametric correlational procedures).

[3]These procedures were intended to make the paper technically accurate. For example, establishing separate alphas for hypotheses on the basis of whether they were directional (e.g., As would receive higher means on all 21 JAS items) or non-directional (e.g., the factor structure of As and Bs on the short form would be different) would represent an even more technically accurate procedure. As will be seen, however, the effects which did emerge were robust, such that choice of liberal or conservative decision criteria made little difference on the conclusions of the analyses.

[4]A Table providing means and standard deviations for each of the 21 JAS items, by A/B type and form, is available upon request.

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Table 1

Variables and LISREL Factor Loadings Using the Orthogonal Two-Factor Model for the Short and Long Forms of the JAS

Variable	Factor H (Hard-Driving) Short Form	Factor H (Hard-Driving) Long Form	Factor S (Speed and Impatience) Short Form	Factor S (Speed and Impatience) Long Form
1. Everyday life challenging	.0	.0	.0	.0
2. Acts immediately when stressed	.0	.0	.0	.0
3. Normally eats rapidly	.0	.0	.65	.76
4. Told of rapid eating	.0	.0	.84*	.84*
5. Feels like hurrying others' speech	.0	.0	.09	.27
6. Hurrys others' speech	.0	.0	.16	.19
7. Punctual for appointments	.0	.0	.0	.0
8. Others consider hard-driving/ competitive	.0	.0	.0	.0
9. Self considers hard-driving/ competitive	.78	.85	.0	.0
10. Spouse rates as hard-driving/ competitive	.75*	.75*	.0	.0
11. Spouse rates as active	.0	.0	.0	.0
12. Friends rate as high energy	.20	.16	.0	.0
13. Fiery temper when younger	.0	.0	.0	.0
14. Deadlines normal in courses	.0	.0	.0	.0
15. Sets deadlines for self	.32	.19	.0	.0
16. Works at >2 projects simultaneously	.0	.0	.0	.0
17. Maintains work schedule over holidays	-.11	.08	.0	.0
18. Brings work home at night	.0	.0	.0	.0
19. Other's look to for leadership	.0	.0	.0	.0
20. Much more responsible	.38	.24	.0	.0
21. Approaches life seriously	.35	.36	.0	.0

Goodness-of-fit $\chi^2 = 25.767$, df = 9, $p < .01$.

These factor loadings were obtained from LISREL analyses performed simultaneously on short and long form data. The loadings for each form have been rescaled by the pooled item standard deviations and factor variances to permit direct between-form comparisons of the magnitude of the loadings within each factor.

An * indicates a constrained value. These loadings were fixed at unstandardized values of 1.0 both for the separate analyses of each group and for the simultaneous analyses (see Alwin & Jackson, 1979, 1980; Sorbom, 1974; Sorbom & Joreskog, 1976). All loadings of .0 were also constrained. These items have been scored in a consistent manner -- such that agreement with an item is scored as a Type A response. On the actual JAS, item numbers 7, 11, 12, 15, 16, 18 and 19 are presented with a reverse polarity relative to the other items.

Table 2

Chi-Square Statistics and Measures of Relative Fit for the Three LISREL Models Using the Short and Long Forms of the JAS

Data Set	Model	Form	N	χ^2	df	χ^2/df[a]	TLC[b]
Pooled Sample	zero common-factor	Short	624	2052.214	210	9.772	----
		Long	624	2092.502	210	9.964	----
	orthogonal two-factor	Short	624	1233.701	199	6.200	.407
		Long	624	1391.656	199	6.993	.331
	oblique two-factor	Short	624	1230.507	198	6.215	.406
		Long	624	1379.034	198	6.965	.335
Type As	zero common-factor	Short	236	742.914	210	3.538	----
		Long	253	853.605	210	4.065	----
	orthogonal two-factor	Short	236	499.908	199	2.512	.404
		Long	253	636.612	199	3.199	.282
	oblique two-factor	Short	236	494.406	198	2.497	.410
		Long	253	634.564	198	3.205	.280
Type Bs	zero common-factor	Short	234	615.286	210	2.930	----
		Long	225	715.322	210	3.406	----
	orthogonal two-factor	Short	234	439.044	199	2.206	.375
		Long	225	557.536	199	2.802	.251
	oblique two-factor	Short	234	432.512	198	2.184	.386
		Long	225	557.413	198	2.815	.246

[a] As this ratio decreases and approaches zero, the fit of the given model improves (see Joreskog, 1971).

[b] This coefficient reflects the amount of variance explained by the given model relative to the amount of total variance. As the coefficient increases and approaches 1.0, the fit of the model improves (see Tucker & Lewis, 1973).

Table 3

Comparative Statistics From Contrasts of LISREL Models Using the Short and Long Forms of the JAS

Data Set	Models Contrasted	Form	$\Delta\chi^2$	Δdf	p
Pooled Sample	zero common-factor vs. orthogonal two-factor	Short	818.513	11	< .0001*
		Long	700.846	11	< .0001*
	zero common-factor vs. oblique two-factor	Short	821.707	12	< .0001*
		Long	713.468	12	< .0001*
	orthogonal two-factor vs. oblique two-factor	Short	3.194	1	< .05
		Long	12.622	1	< .001*
Type As	zero common-factor vs. orthogonal two-factor	Short	243.005	11	< .0001*
		Long	216.993	11	< .0001*
	zero common-factor vs. oblique two-factor	Short	248.508	10	< .0001*
		Long	219.040	10	< .0001*
	orthogonal two-factor vs. oblique two-factor	Short	5.502	1	< .01
		Long	2.047	1	< .10
Type Bs	zero common-factor vs. orthogonal two-factor	Short	176.241	11	< .0001*
		Long	157.768	11	< .0001*
	zero common-factor vs. oblique two-factor	Short	182.774	10	< .0001*
		Long	157.909	10	< .0001*
	orthogonal two-factor vs. oblique two-factor	Short	6.533	1	< .01
		Long	0.123	1	< n.s.

* Statistically significant at beyond the .05-level, when adjusted for the overall per comparison error-rate (see Ryan, 1959).

Table 4

Descriptive Statistics for Total and Factor H and S JAS Scores, by A/B Type and Form

A/B Type	JAS Score	Short Form M	Short Form sd	Long Form M	Long Form sd
Type A	Total	10.30[a]	2.09	10.43[b]	2.22
	Factor H	3.18	1.36	3.26	1.34
	Factor S	1.62	1.08	1.61	1.17
Type B	Total	4.37[c]	1.44	4.45[d]	1.36
	Factor H	0.83	0.85	0.87	0.93
	Factor S	0.70	0.77	0.65	0.80

[a]N = 236.
[b]N = 253.
[c]N = 234.
[d]N = 225.

Means and Standard Deviations of 21 JAS Items, by A/B Type and Form[a]

	Short JAS[b]				Long JAS[c]			
	Type A[d]		Type B[e]		Type A[f]		Type B[g]	
JAS Item[h]	$\bar{X}$	sd	$\bar{X}$	sd	$\bar{X}$	sd	$\bar{X}$	sd
1. Everyday life challenging	1.14	0.71	1.66	0.75	1.14	0.72	1.51	0.79
2. Acts immediately when stressed	0.58	0.49	0.64	0.49	0.54	0.51	0.71	0.45
3. Normally eats rapidly	1.26	0.94	1.71	0.92	1.13	0.94	1.67	0.85
4. Told of rapid eating	1.12	0.78	1.58	0.58	1.10	0.81	1.44	0.65
5. Feels like hurrying others' speech	0.60	0.59	0.92	0.52	0.61	0.60	0.96	0.51
6. Hurrys others' speech	0.89	0.66	1.24	0.54	1.00	0.64	1.25	0.57
7. Punctual for appointments	0.81	0.75	0.70	0.67	0.93	0.75	0.72	0.66
8. Others consider hard-driving/ competitive	1.05	0.80	1.94	0.51	1.11	0.82	1.92	0.51
9. Self considers hard-driving/ competitive	0.99	0.78	1.96	0.56	0.98	0.80	1.90	0.60
10. Spouse rates as hard-driving/ competitive	1.04	0.85	1.86	0.58	1.07	0.84	1.91	0.59
11. Spouse rates as active	1.30	0.57	0.88	0.46	1.32	0.56	0.93	0.46
12. Friends rate as high energy	2.43	0.72	1.97	0.66	2.50	0.64	2.08	0.67
13. Fiery temper when younger	0.97	0.77	1.37	0.80	0.99	0.75	1.36	0.76
14. Deadlines normal in courses	0.92	0.67	1.24	0.64	0.92	0.64	1.22	0.53
15. Sets deadlines for self	1.56	0.59	1.09	0.55	1.52	0.57	1.13	0.55
16. Works at >2 projects simultaneously	1.18	0.73	0.88	0.61	1.16	0.73	0.91	0.64
17. Maintains work schedule over holidays	1.20	0.63	1.25	0.49	1.24	0.60	1.27	0.51
18. Brings work home at night	1.87	0.40	1.77	0.51	1.94	0.29	1.80	0.49
19. Other's look to for leadership	1.40	0.60	0.96	0.54	1.38	0.62	0.91	0.56
20. Much more responsible	0.73	0.69	1.13	0.65	0.60	0.66	1.16	0.63
21. Approaches life seriously	0.82	0.77	1.30	0.62	0.79	0.76	1.22	0.62

[a]Table to accompany, "Comparing the Long and Short Forms of the Student Version of the Jenkins Activity Survey," by P. Yarnold, F. Bryant and L. Grimm.

[b]Consists of only the 21 scored JAS items.

[c]Consists of 21 scored and 23 "filler" (unscored) items.

[d]N = 236.

[e]N = 234.

[f]N = 253.

[g]N = 225.

[h]For these items, a <u>lower</u> number is scored as <u>more</u> Type A, except for items 7, 11, 12, 15, 16, 18 and 19, for which the opposite is true (i.e., higher numbers indicate more Type A). For every item, Type A's scored significantly higher than type B's, in the expected direction (the effect was only marginally statistically significant for item 17).