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THE ITEMETRIC PROPERTIES OF THE EYSENCK PERSONALITY QUESTIONNAIRE: A REPLY TO HELMES

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Summary—Helmes (1980) presented results demonstrating that the purported factors of the Eysenck Personality Questionnaire (EPQ) did not emerge in a sample of 191 undergraduates. However, this conclusion was reached using incorrect factor analytic methodology. Helmes not only used an idiosyncratic version of Eysenck's original methodology but also failed to take into account recent evidence on factor analytic procedures. To illustrate the remarkable clarity of the EPQ factors, the results from three samples of data are presented here. The sample data are diverse, taken from three different populations: 1198 British adults, 406 British students and 116 Thai students. The consistency of various psychometric parameters adds further evidence justifying the existence of these factors.

INTRODUCTION

Helmes (1980) presented results from an exhaustive item and multiple-order factor analysis of the Eysenck Personality Questionnaire (EPQ), the EPQ being a relatively new personality inventory introduced by Eysenck and Eysenck (1975) as a measure of Extraversion (E), Neuroticism (N), Psychoticism (P) and Social Desirability (the 'Lie' scale (L)). Helmes' item analysis indicated that theP scale was of dubious validity, due to a common occurrence of extreme item-response distributions (item splits), and its correlation with-Desirability as measured by Jackson's Desirability scale from the Personality Research Form (Jackson, 1974). Furthermore, the mean item-total correlations for each scale were very low, for instance a value of −0.22 for the L scale and 0.08 for the P scale. The item factor analysis at the first, second and third order failed to yield the four factors expected, many items failing to even load an interpretable factor. Helmes concluded that... "it is arguable whether the EPQ originally had the structure claimed for it..." (Helmes, 1980, p. 54). This conclusion is also reached indirectly by Loo (1979) in his psychometric investigation of the EPQ.

Both Loo and Helmes reference their factor analyses against Howarth's (1976) investigation of the EPI, with neither author reaching a satisfactory structural agreement between either themselves or with Howarth's results. Finally, Jackson and Paunonen (1980) concluding their review of the EPI and EPQ state... "the data bespeak the difficulty of attempting to relate EPI and EPQ test responses to a general underlying Introversion–Extraversion construct..." (p. 542).

Problems with the study

The statements above combine to form a depressing outlook for the continued use of the EPQ in its present form. Helmes principal component (PC) factored the sample data using two procedures, the first method involved extracting the first four PCs and rotating them to approximate simple structure using Varimax and Promax algorithms. Additionally, a Procrustes rotation to the EPQ scoring key was implemented. The second method was purportedly based upon that used by Eysenck and Eysenck (1969). That is, retaining all components with eigenvalues greater than 1.00 from an initial PC analysis, rotating these using Promax and subsequently factoring the interfactor correlations to obtain second-order factors. The items are then projected onto these unrotated factors, another Promax rotation is carried out, and thus the procedure repeats itself to the third order and above. The second method as described by Helmes is simply not that as used by the Eysencks. While the use of the KG criterion is maintained, it would appear that Helmes has used the Cattell–White (C–W; Cattell and White, 1962) rather than the Hendrickson–White (H–W; Hendrickson and White 1966) higher-order item projection algorithms. The C–W algorithm projects items onto rotated higher-order factors, thus for any high-order factor rotation there are progressively fewer variables left to define hyperplanes. The rotated solutions will thus tend to be extremely sensitive to any slight shifting of factor axis positions. The advantage of the H–W algorithm is that at each higher-order stage of analysis, the items are projected onto the factors prior to rotation. Thus at each stage, the number of possible hyperplane variables is at a maximum.

Helmes' use of the KG bound also does not match the methodology attributed to Eysenck. All first-order factors with eigenvalues greater than 1.00 are not extracted and rotated by Eysenck, rather he accepts the arbitrary figure of 20 as being the cutoff (Eysenck and Eysenck, 1969, p. 197; 1976, p. 54). Thus having established that Helmes' second method of analysis bears little relation to Eysenck's methodology, it is pertinent to query why such importance is attached to this particular sequence of operations. Evidence has accumulated over the years indicating the the Eysenck methodology is not the best procedure for factoring. For example, the arguments of Hakstian and Muller (1973) concerning the use of the KG bound in PCs. They point out that Guttman's statements on the rank of correlation matrices do not apply to PC analysis. Thus accepting the validity of Kaiser's (1960, 1965) arguments concerning the use of factor alpha's, a very different approach to factor extraction is required. Based upon extensive empirical analyses on 26 sets of data using three tests of factor extraction, Barrett and Kline (1982b) found little to recommend any one test of factor extraction. As both Veldman (1974), and Kaiser (1963) note, overextraction and subsequent rotation of factors leads to disintegration of factor structure. The support for this statement derives from both orthogonal (Varimax) and oblique rotated sets of data. Barrett and Kline (1981b) also provide evidence for this effect in a scale factoring of Cattell's 16 PF. Factoring correlation matrices composed from increasing subject sample sizes yielded idiosyncratic rotated solutions when the quantity of factors rotated was too large. Evidence presented by Cattell and Vogelmann (1977) clearly demonstrates that the simplistic use of the KG bound leads to overestimation of factors from large data matrices. Thus both Loo and Helmes have almost certainly overextracted at the first order, leading to a disintegrated first-order rotated factor pattern.

With regard to the rotation methodology, Hakstian and Abell (1974) have presented evidence suggesting that the Direct Oblimin (Jennrich and Sampson, 1966) algorithm is marginally superior to Promax in attaining simple structure as defined by the maximum Hyperplane Count (HC). By increasing the range of permissible δ values and adjusting the step sizes between these values, the efficiency of the Direct Oblimin algorithm with regard to maximizing HCs is enhanced. This is not to say that the resultant factor patterns given by Promax and Direct Oblimin will differ significantly. The aim is simply to maximize the HC and thus attain a better approximation to simple structure.

Finally, it is helpful to assess the validity coefficients of each factor as defined by Cattell and Tsujioka (1964). The factor validity coefficient of a scale or set of items may be defined as the ratio of mean validity (mean item-factor correlation) to mean homogeneity (mean interitem correlation). The optimal homogeneity is a value which is low relative to the correlation of the items with the factor. The factor validity coefficient is in fact a multiple correlation coefficient, thus it can be viewed as indirectly assessing the similarity of a set of items that load upon a factor. If the items are no more than reworded counterparts of one or two basic items, then their homogeneity is likely to be high with the result that the mean item-factor correlation will reach a ceiling value. The factor validity will then be low. Of course, in the converse limit, the factor validity will also be low. In this way, the coefficient can be used to assess the validity of any set of items that the investigator intends to call a factor scale or at least an item group that measure a common component.

Some new data

Having, therefore, demonstrated that the factoring methodology of both Loo and Helmes is of a poor standard with regard to recent evidence, no attempt has been made to replicate it in the studies reported below. This is not simply an attempt to confound the factor structure of the EPQ by introducing novel or different analysis methods, rather the best techniques available (as suggested by empirical evidence) have been implemented. Velicer (1976b, 1977) especially has shown that very little difference exists between various solutions given by different eigenvalue decomposition routines. The major deviations between Helmes' methodology and that reported below are in the factor extraction, rotation and higher-order factoring stages. The studies, from which the data below has been extracted, are reported in greater detail in Barrett and Kline (1980a,b) and Kline *et al.* (1981).

METHOD

Subjects

Three sets of subject data were used for the analysis. A small sample of Thai University students ($N = 116$) was administered a back translated version of the EPQ. A larger sample of 406 British University students also provided test responses. Finally, a Gallup quota sample of $N = 1198$ British adults commissioned by Eysenck was used. The details of this sample are provided in Eysenck (1979). The raw data from this sample was kindly loaned to the authors by Eysenck, thus enabling a full psychometric analysis to be undertaken. All samples were of mixed sex. For the purposes of the analyses below, the samples were left unseparated, however, the British student and British adult samples were split in the previous analyses. Further details about the samples (and analyses) can be obtained from the relevant papers referenced above.

Analyses

The item analyses were computed using point-biserial correlation coefficients. As these are mathematically equivalent to Pearson product–moment correlations (Nunnally, 1978), mean item-total correlations were computed using Fisher's Z transformation. The results of the factor analyses are taken from 1st order PC solutions where, contrary to the Scree test (Cattell, 1966) and the MAP test (Velicer, 1976a), only the first four factors were extracted and rotated to simple structure. Rotation was varied from near orthogonality to near maximum obliquity using Direct Oblimin rotation, varying the parameter δ from -40 to 0.6 in steps of 0.1. With the convergence criterion set at 0.00001, the 'best' solution was determined by the maximum HC (± 0.1 bound) with the minimum overall sum of squared loadings within an HC fixing the 'best' plateau HC solution.

By disregarding the results from factor extraction tests and simply using the first four PCs, Eysenck's claim that his factors can be found at any order is highlighted. In Barrett and Kline (1980b) the factors E, N, P and L emerged at the second order, the first-order quantity of factors being defined by the Scree and Velicer MAP test. The loading pattern of these second-order factors is in almost perfect equivalence with the solutions computed below. A slight increase in eigenvalue size is the only noticeable effect in extracting four factors at the first order. This is, of course, to be expected due to the compounding of error variance in factoring at the second and third order. The Thai sample was not taken to higher order analysis; this was unnecessary given the results from the two other samples.

Finally, alpha reliability and factor validity coefficients were computed for the Gallup sample data. As this is the largest and most representative sample reported here, these coefficients are likely to be near their optimal values.

RESULTS

Table 1 presents a summary of the results from the item analyses. Contained within this table are the mean full-scale item-total correlations. Additionally, the number of other scale items significantly ($P < 0.01$) correlated with a particular scale total are listed. The number in brackets below the mean scale item-total correlation is the mean non-scale item-total coefficient.

As can be seen from this table, compared with Helmes' values there is a great improvement in size of correlation on the P and L scales especially. Table 2 presents a breakdown of the frequency of extreme item-split P values for each complete item scale.

The values overall are slightly better than those reported by Helmes.

Table 3 provides information as to the percentage of variance extracted in each of the three sample factorings. Two other special factor solutions are presented here demonstrating the amount of variance that can be accounted for by these four factors. The solutions are taken from a Radial parcel (Cattell, 1974) factoring of the EPQ, with parcels composed of two and four items. (Barret and Kline, 1981a). Parcelling of items reduces individual item measurement error and thus provides a more clear representation of the factor structure.

Table 4 gives the number and percentage of items with their highest loading on the appropriate keyed scale.

In this table the Promax and Varimax rotated solutions reported by Helmes can be compared with the Direct Oblimin solutions.

Table 1. Mean item-total correlation for keyed scale...and significant item breakdown

Scale	Helmes	Thai students		British students		British adults	
			$P < 0.01$		$P < 0.01$		$P < 0.01$
P (25)	0.08	0.2956 (0.0037)	17P 1L 1E 2N	0.3621 (−0.0577)	25P 10L 5E 2N	0.4026 (−0.0066)	25P 12L 8E 16N
E (21)	0.32	0.4356 (0.0000)	19E	0.5336 (−0.0577)	21E 3L 5N 5P	0.5210 (−0.0890)	21E 8L 15N 6P
N (23)	0.49	0.4177	21N 1L 4P	0.4833	23N 1L 4E 7P	0.5170	23N 3L 15E 9P
L (21)	−0.22	0.3981 (−0.2065)	18L 1E 2N 4P	0.4032 (−0.1868)	21L 1E 1N 11P	0.4456 (−0.0934)	21L 12E 13N 6P

Figures in parentheses = mean correlation of significant non-scale items.

Table 2. Frequency of extreme P-values

Scale	Range	Helmes	Thai students	British students	British adults
P	<0.2 or >0.8	22 (88%)	11 (44%)	16 (64%)	22 (88%)
	<0.1 or >0.9	13 (52%)	4 (16%)	9 (36%)	9 (36%)
E	<0.2 or >0.8	6 (29%)	3 (14%)	3 (14%)	2 (10%)
	<0.1 or >0.9	1 (5%)	0 (0%)	0 (0%)	0 (0%)
N	<0.2 or >0.8	5 (22%)	5 (22%)	1 (4%)	1 (4%)
	<0.1 or >0.9	0 (0%0	3 (13%)	0 (0%)	0 (0%)
L	<0.2 or >0.8	9 (43%)	4 (19%)	10 (48%)	3 (14%)
	<0.1 or >0.9	1 (5%)	1 (5%)	3 (14%)	0 (0%)

Table 3. Percentage of variance accounted for by the factor solutions

Description	Helmes	Thai students	British students	British adults
4 Primary item factors	23.5	22.2	23.1	24.8
4 Secondary item factors	—	—	22.8	24.1
Size = 2 Radial Parcels 4 Factors	—	—	35.6	37.7
Size = 4 Radial Parcels 4 Factors	—	—	52.5	54.8

Table 4. No. and % of items with their highest loading on keyed scale

Scale	No. of items	Helmes Promax/Varimax	Thai students	British students	British adults
P	25	14 (56%)	10 (40%)	21 (84%)	25 (100%)
E	21	16 (76%)/17 (81%) 17 (81%)	19 (90%)	20 (95%)	21 (100%)
N	23	22 (96%)	20 (87%)	23 (100%)	23 (100%)
L	21	18 (86%)	18 (86%)	19 (90%)	21 (100%)

Table 5. Mean absolute factor loading of scale items

Scale	Helmes (Procrustes)	Thai students	British students	British adults
P	0.255	0.167	0.333	0.390
E	0.455	0.401	0.516	0.511
N	0.472	0.388	0.462	0.506
L	0.377	0.374	0.374	0.429

Table 6. Factor validities and coefficient alphas

Factor	Gallup sample Factor validity	Gallup sample coefficient alpha	Helmes sample coefficient alpha
E	0.99	0.86	0.84
N	0.99	0.87	0.86
L	0.98	0.80	0.75
P	0.97	0.77	0.59

Table 5 presents the mean absolute factor loadings of the full-item scale. All items in the British student and adult samples loaded in the same direction. For the Thai sample item E36(-0.115) and items P9, 33, 57, 61, 79 and 87 failed to do so. (All these values had loadings less than 0.2). Thus the P scale value is very slightly inflated. This is unimportant as the factor could not be identified at all clearly in the Thai data set.

The mean loading of the non-scale items was not calculated. Given the overall $\sim 50\%$ hyperplane count on each factor from data set, the clarity of the scale loadings (>0.3), and the dubious nature of the stability of such loadings, it is apparent how little these values contribute to any understanding of the factor scales.

Table 6 presents the alpha and factor validity coefficients for the first four primary factors identified as E, N, L and P within the Gallup sample data. The alpha coefficients for the total sample reported by Helmes are also included for comparison in this table.

DISCUSSION

Before entering upon the discussion of the results it is necessary to point out certain deficiencies in the Thai data, the most obvious being the low sample size. However, the data set was included here for this very reason. Both Helmes and Loo report results based upon fairly low subject numbers. Thus if the Thai data could yield results comparable to their data, a check on the replicability of the factors in populations other than British can be made. As it was, the Thai data surpassed the results of both Helmes and Loo. However, the P factor did not appear with any strength in the Thai data and in fact it was concluded that this scale was inappropriate for measurement in the Thai culture. However, the remaining three scales did appear strongly. Thus small sample size and cultural specificity can be ruled out as reasons for the nonappearance of E, N and L. The Canadian culture is more similar than the Thai to British culture.

In nearly all cases, the mean item-total correlation for each scale was far superior to Helmes' results. Noticeably the values for the P scale are much higher. However, this particular scale contains many items with relatively severe item splits outside the range 0.2 and 0.8. This is perhaps not surprising considering the samples tested. The scale was primarily designed for clinical samples although as can be seen from the factor analyses it does have some variance in normal samples. The characteristics of the E, N and L scales are of course superior to those of the P. The Gallup sample of adults providing a particularly low number of poor item splits. Noting the consistency across the three data samples of the mean item-total coefficients for the L scale (Table 1), Helmes' value of -0.22 is indicative of a gross effect running through his raw data on this scale. The existence of this negative correlation should not be taken lightly, unfortunately its causation is inexplicable in view of the evidence presented here. Each sample of data presented here was collected by a different team of investigators in different geographic locations, thus no procedural test bias is likely to be evident in these samples.

With regard to the pervasiveness of the factors P, E and N, Eysenck (1978) has referred to the factors as 'superfactors', that is, factors which are found to exist at higher orders. Both conceptual and methodological reasons are cited by Eysenck for the importance of such factors over and above the primaries usually extracted in factor analytic studies. Thus in a cojoint scale factoring of the Howarth Personality Questionnaire and Additional Personality Factor Inventory 2 (Howarth, 1970, 1971, 1972) and the EPQ, P, E, N and L were located separately and clearly in the resultant factor pattern (Barrett and Kline, 1980a).

Finally, a small pilot study undertaken by the authors examined the comparability of the EPI and EPQ E, N and L scales within a sample of 55 individuals. The sample was composed primarily of university undergraduates in addition to a small number of adults. Pearson correlations were computed for each scale yielding the values 0.8262 for E, 0.9080 for N and 0.6850 for L. All correlations were significant at $P < 0.01$. Bearing in mind the exclusion of the Impulsivity items from the EPQ E scale (Eysenck and Eysenck, 1977, 1978) it would appear that the Eysencks' assumption of scale comparability for E and N is upheld by these results. However, the L scale comparability must be of doubtful validity.

Because of Helmes' incorrect use of Eysenck's methodology in addition to the somewhat poor psychometric standard of these techniques, very little of Helmes' methodology could be replicated here. However, this is by far the least important point being made in this paper. In all the tables presented, clear trends in the data demonstrate a consistency of psychometric parameters. As the samples increase in size and become more representative of their populations, so do the summary statistics increase. The use of the Thai sample data simply illustrates that under the worst conditions for factoring (low sample size and a different culture) the E, N and L factors appear quite clearly. The radial parcel factor results also demonstrate just how much variance can be accounted for by the four questionnaire factors; a virtual doubling of the item factor variance percentages. Overall the data presented here suggest that the EPQ items do in fact load upon four factors in a consistent and coherent fashion. The data presented by both Loo and Helmes do not show such consistency between samples. This, as noted in the Introduction, is due to the overextraction of factors at the first order. The reason for the failure of Helmes to find E, N, P and L at the first order probably resides in the raw data. This method of simply extracting and rotating the first four factors is the one used in the above analyses. Once again, the consistency of the results argue against any claims of spurious or chance factor structures.

According to a comprehensive range of psychometric criteria, the factor structure of the EPQ is replicable. As to whether or not the EPQ was factorially devised is another issue beyond the scope of this paper. However, this paper does demonstrate that the EPQ scales are certainly factors as defined by principal components analysis. As Velicer (1976b, 1977) has shown that little difference exists between various factor analysis solutions, in addition to Harman's (1976) assertions of the redundancy of the diagonal in any large data matrix, it is expected that the results here would be generalizable to any other method of factoring.

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