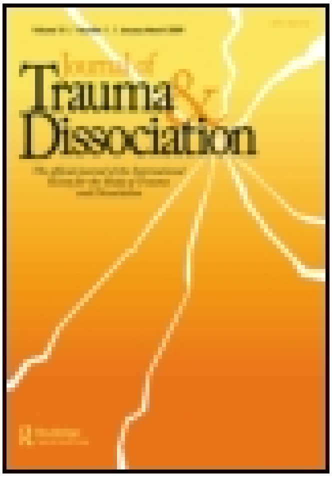

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On the Dimensionalities of the Dissociative Experiences Scale (DES) and the Dissociation Questionnaire (DIS-Q)

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ABSTRACT. The dimensionalities of the Dissociative Experiences Scale (DES) and the Dissociation Questionnaire (DIS-Q) have been explored by factor analysis by a number of investigators. The general claim is that both scales are multidimensional. However, we argue that, in both cases, this apparent multidimensionality is a form of range restriction that arises because some symptoms are more common than other symptoms. As a result, the resulting subscales derived from the multifactor solutions are so highly intercorrelated as to be incapable of meaningful discriminant validity. The current findings support a more parsimonious single factor solution for both scales.

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Bernstein and Putnam's Dissociative Experiences Scale (DES; E. Bernstein & Putnam, 1986; Carlson, Putnam, Ross, et al., 1993) and Vanderlinden and colleagues' (1993) Dissociation Questionnaire (DIS-Q; Vanderlinden, van Dyck, Vandereycken, & Verkes, 1993) are both used in the evaluation of dissociative identity disorder (DID). In both cases, arguments have been made that they are multidimensional (Ross, Joshi, & Currie, 1991; Ross, Ellason, & Anderson, 1995). This paper examines the dimensionality of these two scales in the light of previous work showing that much (Bernstein & Eveland, 1982; Bernstein & Garbin, 1985; Bernstein, & Keith, 1991; Bissonnette & Bernstein, 1990; Bernstein, Teng, & Garbin, 1986) but not all (Bernstein & Gesn, 1997; also see Bernstein, Garbin, & Teng, 1988; Nunnally & Bernstein, 1994) of the evidence for the multidimensionality of various psychological scales is an artifact of differences in item distributions, in the present case simply meaning that some symptoms are more common than others.

STUDY 1: THE DISSOCIATIVE EXPERIENCES SCALE

Ross et al. (1991, 1995) examined the factor structure of the 28-item DES in two studies that respectively employed samples derived from the general population (GP, $N = 1055$) and patients diagnosed with dissociative identity disorder (DID, $N = 274$). Both samples provided a three-factor solution. Consistent with the scale's overall discriminant validity, there was also a substantial mean difference between the two samples; Ross et al. (1995) reported that the mean $\pm$ SD DES scores was 44.6 ± 19.2 for individuals with DID vs. 10.8 ± 20.2 for the general population. Scales were derived from a varimax rotation of the principal components using a cutoff of .45 to define a salient variable. Setting aside slight differences in the results of the two analyses, the scales were labeled (a) absorption-imaginative involvement (items 2, 14, 15, 17, 18, 19, 20, 21, 22, 23, 24, and 25), (b) activities of dissociated states (items 3, 4, 5, 6, 8, 10, 25, and 26), and (c) depersonalization-derealization (items 1, 7, 11, 12, 13, 16, 27, 28). Ross et al.'s (1991, 1995) conclusions differed from those of Fischer and Elnitsky (1990) who argued that the DES was adequately represented by a single factor.

We suggest that Ross et al.'s (1995) critique of the scale's original unidimensionality is in error because they failed to consider issues that are important when *categorical* (*discrete*) data such as items, as opposed to *continuous* variables, are analyzed. Although the definition of "continuous" is perhaps somewhat arbitrary in empirical research, Nunnally and Bernstein's (1994) definition will be used. They define a variable as continuous if it takes on 11 or more levels with measurable probability. Items on the DES approach this criterion but, as we hope to show below, they behave like items on other tests in being discrete.

Ross et al. (1995) argued that Fischer and Elnitsky's college student population was less adequate than their own stratified general population cluster. However, a college student population, being more homogeneous, should yield lower rather than higher inter-item correlations than a more diverse sample. These lower correlations, in turn, should cause more rather than fewer factors to be extracted since, by definition, less variance is concentrated in the early components. Ross et al. (1995) also criticized Fischer and Elnitsky's use of Cattell's (1966) scree criterion to define the number of factors because of its subjectivity (Kim & Mueller, 1978). However, part of Cattell's intent in developing the scree criterion was to minimize artifacts present when items, as opposed to whole tests, are factored. For reviews of the number-of-factors issue see Gorsuch (1973) and Nunnally and Bernstein (1994). Indeed, it is safest to say that there is vigorous dispute and therefore no "litmus test" to define the number of factors unequivocally, despite the passion with which proponents of the various criteria define their position and the wide use of such indices as the Kaiser-Guttman rule (Guttman, 1954; Kaiser, 1960, 1970) to retain as many factors as there are eigenvalues exceeding 1.0.

Ross et al. (1995) used a component solution followed by an orthogonal rotation, whereas Fischer and Elnitsky (1990) used a common factor solution followed by an oblique rotation. These possible differences can be ignored in examining dimensionality. Although Fisher and Elnitsky defined their factor loadings from a common factor, they, like most, used a preliminary component solution to determine the number of factors. Finally, rotation is irrelevant since it does not apply to a single factor solution.

Nunnally and Bernstein (1994), among others, noted that items are considerably less reliable than whole scales, by definition, so that the average correlation among a series of items will be lower than the average correlation of scales that are an aggregate of these items. In turn, this means that the factor structure of items will be diffused (spread among more than one factor) relative to the factor structure of scales (also see Gorsuch, 1983; Bernstein & Eveland, 1982; Bernstein & Garbin, 1985; Bernstein & Keith, 1991; Bernstein & Teng, 1986, 1989; Bernstein et al., 1994, 1995; Bissonnette & Bernstein, 1990). Moreover, assuming that the data contain a reasonably strong first (general) factor, range restriction will lead to higher correlations among items whose distributions are more similar to one another. This produces factors that are defined by the similarity of these item distributions.

These spurious factors were first observed in abilities tests with dichotomous items and were denoted "difficulty factors" (Gorsuch, 1983; Carroll, 1945). This is perhaps unfortunate as they can also be observed in personality assessment and obtained with multi-category as well as dichotomously scored items (sometimes more easily, depending upon the criterion used to define number of factors, see Bernstein & Teng, 1989). A third consideration raised by Bernstein and Teng (1989), among others, is that items keyed in the positive direction typically correlate more poorly with items keyed in the negative direction than with other items keyed in the positive direction, and vice versa, because of differences in method rather than content variance (Campbell & Fiske, 1959). This third consideration is not relevant here because all DES items are keyed in the same direction (the higher the number assigned the activity, the more strongly the item is indicative of DID). However, the first two points are, as we hope to demonstrate, major considerations.

Methods for Study 1

Study 1 used methods previously employed by the first author to investigate the dimensionality of the DES. The following results were obtained from the general population database previously reported in Ross et al. (1991) and the DID database reported in Ross et al. (1995).

Results of Study 1

Properties of Ross et al.'s Proposed Solution

The proposed solution is the assignment of items to scales presented in Ross et al. (1995). Note that even though the factors obtained in their analyses are orthogonal by definition, the scores that they produce by averaging the items can be quite highly correlated when, as is almost invariably the case, items are either included or excluded from scales rather than given their factor score weights. Specifically, the correlations among the three proposed scales ranged from .69 to .81 within the general population database and from .72 to .83 within the dissociative identity disorder (DID) database. Even though the individual scale values of coefficient a are quite reasonable, ranging from .80 to .83 within the general population database and from .84 to .90 within the DID database, disattenuating the scale correlations increases these correlations among the scales to .93-1.00 within the general population and .83-.97 within DID. The three scales are therefore nearly totally redundant when scored conventionally. The coefficient a values for the full scale are higher yet at .92 and .95, respectively. This latter difference also suggests there is slightly more true score variance among the DID patients.

Factor Structure

Figure 1 is the scree plot for the two samples. The DID data provide four eigenvalues greater than 1.0 and the general population data provide five. However, a scree criterion is consistent with the adequacy of a single factor in both cases. Note that the first component only accounted for 41% and 35% of the variance in the DID and Normal samples respectively, and that the second component accounted for 8% and 7%, respectively. However, we suggest that the relatively small amount of variance accounted for by the first component is typical of item-level analysis because of their unreliability and consequently poor intercorrelation, and that it is also a reflection of the heterogeneity of the item distributions.

This paper is concerned with the extent to which the three-factor solution previously obtained is an artifact of the similarities and dissimilarities of item distributions. The first author's previous research, described above suggests one approach. Means, standard deviations, and skewnesses were obtained for each item over participants within the two databases, and items were then ranked with regard to each of these three statistics for each database. For example, item 2 (missing amnesia symptoms for parts of a conversation) had the second highest mean within the general population database (24.3) and the third highest mean within the DID database (60.9). It was thus commonly endorsed in both groups. Conversely, item 4 (measuring amnesia symptoms for changes in clothing) had the second lowest mean within the general population database (= 1.9) and the lowest mean within the DID database (= 20.3). It was thus relatively unlikely to be endorsed by either group.

Table 1 contains all six item rankings (three statistics × two databases). Similar data appear in Table 2 of Ross et al. (1991) and Table 1 of Ross et al. (1995). The two sets of item means and item skewnesses were at least modestly correlated between the two samples ($r = .69$ and $.74$, respectively) but the two sets of standard deviations were unrelated ($r = .00$). Moreover, the item means were highly correlated with the item skewnesses ($r = -.99$ and $-.86$ for the DID and general population databases, respectively). However, item means and item skewnesses were poorly correlated with item standard deviations within the DID database ($r = .15$ and $-.17$) but were highly correlated within the general population database ($r = .94$ and $-.93$). The main points are the following: (a) items which are highly endorsed in one group tend to be highly endorsed in the other, and (b) such items tend to have negative skewness and, among normals, high variability over participants.

Table 1 indicates that factors defined on the basis of the varimax rotated component analysis are also well defined on the basis of their means and, therefore, skewnesses in both samples. Although Ross et al. (1995) labeled factor I "Absorption-imaginative involvement," it could just as well be called "commonly

FIGURE 1. Scree Plots for DES Items Derived from the Dissociative Identity Disorder (DID) and General Population (GP) Databases.

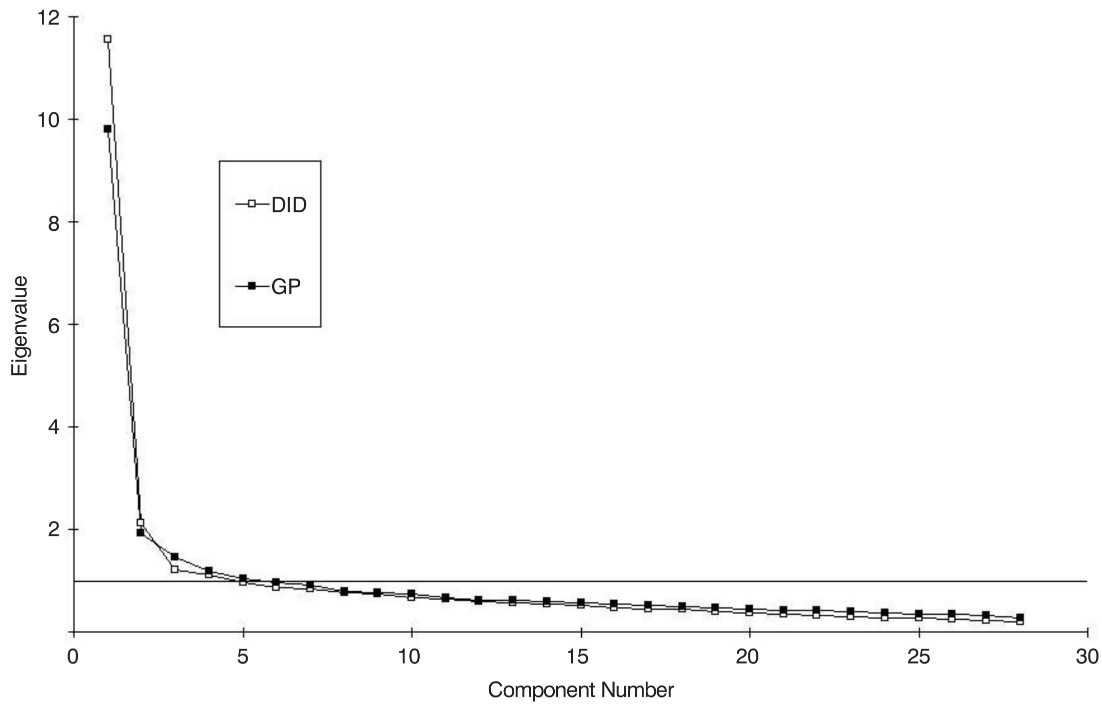

TABLE 1. Rank Orderings of DES Items Over Participants on the Basis of Their Means, Standard Deviations (SD), and Skewnesses (SK), Separately Within Databases.

			Database					
			DID			GP		
Item	Symptom	Factor	Mean	SD	SK	Mean	SD	SK
2	Amnesia/attention loss for conversations	I	3	28	26	2	6	27
14	Past memory/event revivification	I	8	12	21	6	4	23
15	Uncertainty about the reality of an event	I	9	16	20	10	9	17
17	Absorption of attention to movies, TV, etc.	I	10	6	19	5	2	24
18	Reality confusion for fantasy or daydream	I	14	14	10	13	13	16
19	Ability to ignore pain	I	7	11	23	1	1	28
20	Trance episodes/staring into space	I	4	24	25	7	8	21
21	Vocal self-dialogue while alone	I	1	4	16	8	7	22
22	Situational identity/behavior alteration	I	1	22	28	12	12	18
23	Puzzling alteration of abilities	I	6	23	22	3	3	26
24	Uncertainty of task completion vs. intention	I	5	17	24	4	5	25
25	Amnesia for one's own behavior	I	12	18	18	9	11	20
3	Discovery of amnesic travel/arrival	II	25	25	4	26	26	3
4	Amnesia for a clothing/apparel change	II	28	27	1	27	27	1
5	Discovery of amnesic item acquisition	II	26	21	2	24	21	4
6	Unexplained friendliness from strangers	II	23	15	6	11	10	19
8	Inability to recognize friends/family	II	27	26	3	21	20	8
10	False accusation by others of lying	II	22	3	7	17	19	12
25	Amnesia for one's own behavior	II	12	18	18	9	11	20
26	Failure to recognize one's own art/writing	II	21	2	8	18	17	11
1	Amnesia while driving/traveling	III	16	19	14	14	16	15
11	Failure to recognize self in the mirror	III	24	9	5	28	28	2
12	Derealization for one's surroundings	III	20	10	9	22	22	6
13	Depersonalization of one's own body	III	13	13	17	25	25	5
16	Unfamiliarity for a familiar location	III	19	20	15	16	15	14
27	Auditory voices inside one's head	III	2	1	27	19	18	9
28	Unclear/foggy visual sense of the world	III	15	7	11	23	24	7
7	Out-of-body experience	III	17	5	12	20	23	10
9	Amnesia for important life events	-	18	8	13	15	14	13

Note: Item 25 is assigned to both factors I and II, and item 9 belongs to no factor.

endorsed dissociative symptoms." None of these 12 item means had a ranking less than 14 among the DID and 13 among the general population. Likewise, factor II or "activities of dissociative states" could be called "infrequently endorsed dissociative symptoms." Only one item (#25) failed to rank below position 21 among the DID patients though two others (#6 and #10) had relatively large means in the general population. Factor III or "depersonalization/ derealization" can be called "dissociative symptoms endorsed at an intermediate level" since item means tend to fall between those of the other two factors. In addition, even though the response scale nominally allowed a range of 100 points, the data fell in relatively few categories, especially for items on putative factor II.

This relationship between item statistics and factor structure may be pursued as in our previous work using multiple group factor analysis, which defines factors on the basis of sums of items (Bernstein et al., 1988; Gorsuch, 1983; Nunnally & Bernstein, 1994) and thus examines properties of conventionally scored scales. This is a method of confirmatory factor analysis developed apparently independently by Holzinger (1944) and Thurstone (1945). In particular, we obtained the fit as the percentage of variance explained for the following models within each database: (a) the first three principal component (PCs) using the factor-score weights obtained from PC analyses of each database, (b) Ross et al.'s (1995) proposed item assignments, (c) three factors defined by the rankings based upon the item means so that the nine items with the highest means comprised factor I, the nine items with intermediate means comprised factor II, and the ten items with the lowest means comprised factor III, (d) three factors analogously defined by rankings based upon item standard deviations, (e) three factors analogously defined by rankings based upon item skewnesses, and (f) "pseudofactors" in which items were assigned arbitrarily to factors (items 1, 4, 7, 10, 13, 16, 19, 22, 25, and 28 comprised one factor, items 2, 5, 8, 11, 14, 17, 20, 23, and 26 comprised a second factor, and items 3, 6, 9, 12, 15, 18, 21, 24, and 27 comprised a third factor). Models using item statistics from one sample and the other sample's correlation matrix were also evaluated, but these results did not add to an understanding of the three factors. The method is thus confirmatory, since it tests proposed structures, but not in the sense of structural equation modeling.

Models fit better in the DID database than the GP database because of the former's marginally greater reliability and, therefore, true variance. Most critically, the proposed factors, which are basically PC weights rounded to 0 or 1, fit only .3% and .1% better than weights defined on the basis of item means in the DID and GP samples, respectively. Indeed, the model based upon item means only fits

TABLE 2. Fits as Percentage of Variance Explained for Alternative DES Item Factor Models Data in Dissociative Identity Disorder (DID) and General Population (GP) Samples.

	Sample	
Model	DID	GP
Principal components	53.1	47.1
Proposed	51.0	43.7
Means	50.7	43.6
Variances	47.3	43.6
Skewness	50.5	43.1
Pseudofactor	45.7	39.8

2.4% and 3.5% poorer than the maximum imposed by the first three PCs in these samples. Some of this difference is due to the fact that a PC solution capitalizes upon chance. In order to estimate this capitalization, the two main samples were randomly split in half. The PC factor score weights from each half sample were then applied to the other half sample. The difference in fit between this solution and that generated by the sample's own PCs estimates the shrinkage within a half-sample. These values were 1.6% and 1.4% for the DID half-samples and 2.1 and 2.4% for the GP half-samples. Assume further that shrinkage would diminish as a square root factor for the whole sample. If so, roughly half of the disparity between the fit of the item means and the PCs can be attributed to random error. This means that differences in means explains all but 1% of the variance among correlations. Consequently, we suggest that no model of item organization can really produce separable factors, at least in a three-factor model.

Discussion of Study 1

We thus suggest that Ross et al.'s (1995) proposed three-factor structure is an artifact of differences in item statistics so that there is no need for a multifactor solution. To be sure, their factor analysis obtained three factors "by the book," but we suggest that is due to the differences in item distributions that are normally present (and desirable, see Nunnally & Bernstein, 1994) when items are factored. It should be quite clear from the tables that their factors simply describe differences in perceived frequency of occurrence. These frequency differences can be very important to an understanding of DID, but they do not imply that the DES is multidimensional. In sum, these results support Fischer and Elnitsky (1990).

Three additional studies using either common factor or component analysis have been reported. Gleaves and Eberenz (1995) evaluated 125 females with bulimia or anorexia nervosa ($n = 125$); Dunn, Ryan, and Paolo (1994) evaluated 493 male substance abusers, and Ray, June, Turaj, and Lundy (1992) evaluated 260 college students. Although multiple group analyses can evaluate their proposed structures, it does not seem useful because of the extent to which overall differences in correlation magnitude and, therefore, PC structure have been shown simply to reflect differences in item means.

Dubester and Braun (1995) did a thorough psychometric analysis of the DES. They replicated the more positive features of the scale, including its high internal consistency but also noted a negative skew, which was true here though incidental to the goals of the study. They properly avoided a factor analysis because their sample size was small, but they did assume the Ross et al. (1995) three-factor model. As evidence for the construct validity of the subscales, they noted a significant ($p < .001$) difference among the three means in an analysis of variance. We suggest that this finding, which was obviously true here, does not imply three separable dimensions. Indeed, it precisely suggests what is necessary to our conclusion–differences in reported symptom frequency produce differences in correlation magnitude. In other words, their results are quite consistent with our view that the subscales actually represent points along one dimension rather than separable dimensions. One point to keep in mind, though, is that it is possible that an apparent disparity in correlation magnitude might be found in predicting criteria which are skewed–Ross et al.'s "Absorption-imaginative involvement," as it consists of commonly endorsed items, could well correlate more highly with criteria with high base rate and vice versa for their "activities of dissociative state" as regards criteria with low base rate. However, the scale as a whole would also accomplish this end.

Frankel (1990), among others, implies that the concept of dissociation is multidimensional on theoretical grounds by noting: "Whereas suggestibility might play an important role over the whole spectrum of hypnotic experience and behavior, it is possibly more relevant to the less demanding aspects. . . . The deep trance with spontaneous amnesia . . . might be something else." However, the use the term "multidimensional" is really more consistent with the notion of a qualitative "taxon" (Meehl, 1995; Meehl & Golden, 1982; Meehl & Yonce, 1994) that needs not be multidimensional in the mathematical sense. Waller (Waller, Putnam, & Carlson, 1996; also see Waller & Ross, 1997) has argued that normal variation in dissociative tendencies is continuous, but pathological dissociative experiences represent a qualitative taxon. The present factor analytic model assumes within-group differences are quantitative and does not bear upon Waller's concern with between-group differences. Nonetheless, we conclude that the DES is a unidimensional instrument within the context of a factor analytic model and thus simpler to work with than the subscales previously suggested.

STUDY 2: THE DISSOCIATION QUESTIONNAIRE

Vanderlinden (Vanderlinden, van Dyck, Vandereycken, & Vertommen, 1991; Vanderlinden, van Dyck, Vandereycken, Vertommen, & Verkes, 1993) developed the Dissociation Questionnaire (DIS-Q) to assess trauma-based symptoms connected with both adult and childhood events. The DIS-Q consists of 63 self-descriptive items that are responded to on a five-point scale so that one item reads "It occurs that I want to do several things at the same time." The response format is 1 = not at all, 2 = a little bit, 3 = moderately, 4 = quite a bit, and 5 = extremely. Vanderlinden, van der Hart, and Varga (1996) have summarized evidence for the scale's validity.

This study is concerned with the dimensionality of the DIS-Q as it is relatively new and potentially important. Using criteria that are conventional for factoring scales, Vanderlinden, van Dyck, Vandereycken, and Vertommen (1991) and Vanderlinden, van Dyck, Vandereycken, Vertommen, and Verkes (1993) found evidence for a four-factor structure: (a) confusion and fragmentation (25 items), (b) loss of control (18 items), (c) amnesia symptoms (14 items), and (d) absorption (6 items). Evidence that these factors may simply reflect differences in item statistics may be found in Table 1 of Vanderlinden et al. (1993): three of the six correlations among the four subscales exceed .66. Part of the reason that these correlations were not higher was that one of the subscales (presumably Absorption, since it consists of but six items) only had an internal consistency reliability of .42 (though, on the other hand, the authors reported the most reliable scale to have an internal consistency reliability of .99). Disattenuation of these correlations would most likely lead to average scale intercorrelations in excess of .8, which would preclude any of the subscales correlating differentially with a criterion, one of the major reasons subscales are retained over the full scale. As a result, this study parallels study 1 in investigating the bases for the apparent multidimensionality.

Methods for Study 2

Following approval of the human subjects review committee at the University of North Texas and informed, signed consent from each participant, data were collected from introductory psychology student volunteers. Demographic data were also recorded. The DIS-Q items were responded to and scored in the manner described by Vanderlinden, et al. (1993).

Results of Study 2

Demographic Data

The participants were 405 introductory undergraduate psychology students. Demographic data were available on gender ($n = 401$), age ($n = 386$), and marital status ($n = 404$). The sample was predominantly female (66.7%) and ranged in age from 18 to 52 ($m = 21.6$, $s = 5.0$). Most were single (89.1%). Approximately the same number was married (7.9%) as was cohabiting (7.1%), and 1.2% was divorced.

Table 3 contains the mean responses and the rank ordering of that mean out of the 63 items as a function of the factor to which they were assigned by Vanderlinden et al. (1993). These factors will be referred to by number rather than by the name (Vanderlinden et al., 1993), because interpretation of these factors is what is at issue. There are large mean differences among these groups of items. Factors II and IV items were most highly endorsed, both having means of 2.2 over participants and respective standard deviations of .3 and .2. The corresponding mean ranks were 17.8 and 16.3 with standard deviations of 10.2 and 9.6. In contrast, the means for factors I and III were 1.7 and 1.6 ($s = .4$ in both cases); the corresponding data based upon rank orderings were 38.6 and 43.2 ($s =$

TABLE 3. Means of DIS-Q Items and Their Rankings as a Function of Putative Factor Assignment.

Factor I			
Item	Symptom	Mean	Rank
2	Feeling of unreality	1.62	38
3	Loss of contact with own body	1.48	46
7	Feeling of being someone else	1.28	55
9	Possession by a strange internal power during fatigue	1.46	48
10	Finding self in unwanted situations	2.36	7
11	Great detachment between self and own activities or thoughts	1.94	26
12	Identity confusion	2.22	16
16	Discontrol of behavior	1.56	42
17	Feelings of confusion	2.23	15
22	Ego dystonic communication to others	1.67	37
27	Inability to recognize self in mirror	1.23	59
28	Experience of bodily alteration	1.24	58
29	Perception of unreality for surroundings	1.40	50
30	Feeling estranged from ones own body	1.17	63
34	Feeling of internal foreign control over decisions	1.35	54
36	Concern about how to prevent ones own actions	1.92	28
39	Finding a familiar place to feel strange or new	1.38	51
40	Inability to understand or explain own actions	1.76	34
41	Ego dystonic actions	1.75	35
50	Desire/wish for more self-control	2.33	10
57	Experience of one's own mind as split	1.48	46
59	Sense of having two or more internal persons	1.24	57
61	Internal voices	1.48	45
62	Discrepancy between ones private self-concept and public self	2.62	4
63	Foggy/distant perception of world around oneself	1.57	41

Factor II			
Item	Symptom	Mean	Rank
1	Dream-like feeling	2.19	19
4	Unconscious awareness while gorging on food	1.69	36
5	Amnesia for travel	2.07	24
6	Unexplained laughing or crying	1.83	31
8	Amnesia for part or all of a conversation	2.77	3
14	Eating urges while not being hungry	2.36	8
15	Undesired anger	2.20	17
23	Unexplainable mood changes	2.35	9
24	Unconscious behavior	2.27	11
38	Confusion of whether an event was real or dreamt	2.03	25
43	Absorption into fantasy or daydream	1.84	30
44	Aimless staring without thought	2.20	18
46	Difficulty resisting bad habits	2.44	6
48	Eating without conscious attention	2.18	20
49	Day-dreaming	2.78	1
54	Loss of sense of time	2.15	22
60	Unconscious activity	2.15	23

Factor III			
Item	Symptom	Mean	Rank
13	Puzzling discovery of new belongings	1.20	61
18	Amnesia for previous day	1.61	39
19	Unfamiliarity with friends or family	1.25	56
20	Situationally experienced split personality	1.41	49
21	Amnesia for important life events	1.36	53
25	Immediate amnesia for what one has been told	1.89	29
26	Sudden memory black-outs during activities	1.22	60
31	Hyperattentional focus for TV/movies	1.83	32
32	Amnesia for large blocks of time	1.49	44
35	Disremembered actions	1.55	43
37	Sudden discovery of self in unexplained locations	1.18	62
45	Absence of thoughts	1.93	27
47	Misplaced belongings	2.78	2
55	Amnesia for the actuality vs. mere planning of activities	1.58	40
58	Handwriting or drawings not remembered	1.38	51

Factor IV			
Item	Symptom	Mean	Rank
33	Revivification of past events	2.23	14
42	Close observation of one's own activities	2.26	12
51	Hyperawareness of each step one takes	2.16	21
52	Unexplained shift in skill or ability	2.60	5
53	Hyperawareness of each bite while eating	1.83	33
56	Internal arguments involving a simultaneous desire to do separate activities	2.24	13

17.3 and 16.1). Thus, factors II and IV can be thought of as more commonly endorsed items, and factors I and III can be thought of as less commonly endorsed items. The means are very similar numerically to Vanderlinden et al.'s despite the fact their Flemish participants responded to a Dutch version and our participants answered an English translation. Their corresponding item means for factors I–IV were 1.4, 1.7, 1.4, and 1.9. Their sample was slightly more variable than ours as their scale standard deviations ranged from .4 to .6. A similar tendency for items to fall in a limited number of categories seen in Study I was observed here.

Table 4 contains the internal consistency reliabilities of the four scales (coefficients alpha) along the diagonal, the observed correlations among the scales above the diagonal, and the disattenuated correlations among the scales below the diagonal using conventional unit weighting. Note that even though the factors are orthogonal in the original sample and would be approximately so here using fractional factor score weights, conventional unit scoring provides scales that are too highly correlated to relate differentially to criteria of practical interest. The alpha for the full scale was .97. These values are almost identical to those reported by Vanderlinden et al; (1993) as the largest discrepancy, including the full scale

TABLE 4. Factor Intercorrelations for DIS-Q Items.

	I	II	III	IV
I	.95	.81	.84	.58
II	.88	.91	.78	.56
III	.92	.88	.87	.49
IV	.72	.71	.62	.70

Coefficient alpha internal consistency reliabilities are along the diagonal, observed correlations are above the diagonal; and disattenuated correlations are below the diagonal. Coefficient alpha for the full scale is .97.

was .03. However, our factor correlations were slightly larger than theirs, particularly those involving factor IV.

Figure 2 is a scree plot of the principal components of the item intercorrelations. Although a total of 13 components met the Kaiser-Guttman ($\lambda > 1$) criterion, a scree criterion is consistent with a single factor solution. Part of the problem in this case is that a component needs only account for 1/63 or 1.6% of the variance to meet the Kaiser-Guttman criterion. Vanderlinden et al. (1993) were cognizant of this issue, as they also required a factor to account for 5% of the variance in component space. Again the first principal component only accounted for 35% of the variance (the second accounted for 5%), which we suggest is a reflection of the unreliability of items.

Multiple group factoring was conducted as in Study 1. Models describing four factors based upon the proposed assignment, the first principal components, item means, item standard deviations, item skewnesses and arbitrary assignment (pseudofactors) were successively fit to the data. Table 5 describes the fits in terms of the percentages of variance accounted for. As can be seen, the proposed model fit only 1.5% better than the model based upon item means.

Discussion of Study 2

The differences in item means among the four factors are not as extreme as they were in the DES, where there was almost no overlap among the item means comprising the respective DES factors. However, they are sufficiently great here so that it appears reasonable for us to argue that both scales are really unifactor measuring instruments on groups of parsimony. At the same time, the psychometric properties of the DIS-Q as a whole are strong. Save for shortening the DIS-Q to 20 or so items and comparing it to similar tests like the DES, little more seems needed for its further exploration. Because of the nonclinical nature of the present population, we are hesitant to do the requisite item analysis with the data presented in this paper.

FIGURE 2. Scree Plot for DIS-Q Items.

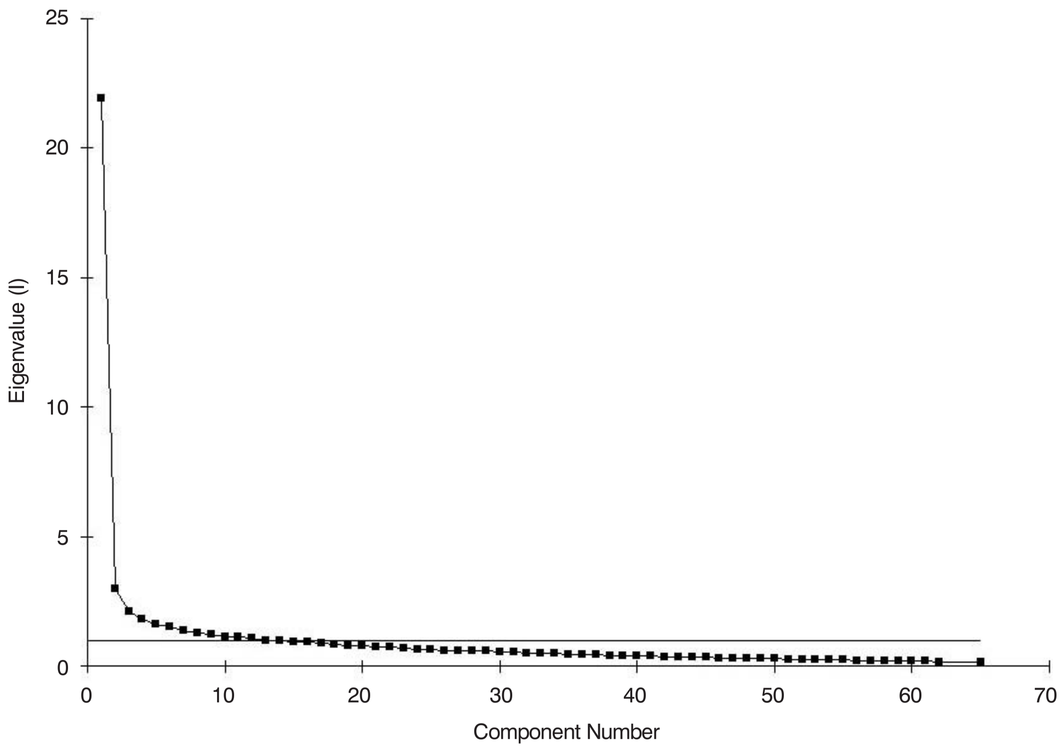

TABLE 5. Fits as Percentage of Variance Explained (% Var) for Alternative DIS-Q Item Factor Models.

Model	% Var
Principal components	45.7
Proposed	43.8
Mean	42.3
Variance	42.5
Skewness	42.0
Pseudofactor	41.6

GENERAL DISCUSSION

Any factor analysis assumes that scores can locate individuals anywhere within a space or continuum. The present issue is the number of directions along which the individuals can vary using the DES or DIS-Q items and, therefore, whether or not the sets of items need to be separated into subscales. We suggest that the need previously expressed for more than one direction (dimension) and thus subscales is an artifact of treating items when they are categorical as if they were continuous scales. Assuming that dissociation is in fact a continuum, the data also clearly indicate that the DID patients and the general population occupy sharply different regions of this space. The standardized mean differences between the two groups on their composite DES scores is perhaps as large as any that are found between a normal reference group and a target group anywhere in the psychiatric literature. However, our results do not deal with whether or not a dimensional representation of dissociation is appropriate in either case. We merely state that *if* a dimensional representation is assumed, one dimension is sufficient to account for individual differences generated by the two scales. We cannot rule out the possibility that additional items might establish a multidimensional structure. However, both scales are already reasonably long and sample content broadly, so this may be unlikely.

One argument that might be raised is that there are indeed multiple true dimensions of dissociation, but that the items reflecting these are endorsed with different frequencies. This hypothesis states that there is a dimension of absorption, which happens to be assessed in the DES with commonly endorsed items, and there is a dimension of activities of dissociative states, which happens to be assessed with infrequently endorsed items. This further implies that one can generate infrequently endorsed absorption items and frequently endorsed activities items. Though it cannot be categorically ruled out, the following arguments can be made against it. First, these putative factors were not proposed–they emerged from an analysis which we argue leads to the generation of spurious factors.

Thus, the authors did not set out to construct the resulting subscales. Second, the usual goal of scientific parsimony is that one should prefer one explanation to two. The argument based upon differential endorsement is readily demonstrable both here and in the other context "difficulty" factors have been noted. Third, even if suitable items could be constructed, the increased similarities of the subscale distributions would increase the already high intercorrelations.

Gleaves, Eberenz, Warner, and Fine (1995) provided results that apparently contradict our single factor hypothesis. They performed a principal component analysis on a composite of three groups of participants: 170 college undergraduates, 15 individuals diagnosed with DID (the authors used the older term "multiple personality" in their article), and 15 individuals with eating disorders. Although details of the component analysis were not presented, they accepted a four-component solution. One component, which they denoted "amnesia," was similar in content to the present factor II ("activities of dissociative state" or "infrequently endorsed symptoms" in previous discussion). A second, which they denoted "depersonalization and derealization," was similar to the present factor III ("depersonalization/derealization" or "symptoms endorsed at an intermediate level"), and a third, which they denoted "common dissociative experiences" was similar to the present factor I ("Absorption-imaginative involvement" or "commonly endorsed symptoms"). Only two items defined the fourth component, which they labeled "absorption." The authors performed a discriminant analysis with the four components as predictors. The amnesia and depersonalization/derealization components both had large weights in discriminating DID patients from normals. Further analysis revealed that all 170 college students and 13 of 15 DID patients were correctly classified with this equation.

Since discriminant function weights control for one another, both amnesia and depersonalization apparently contributed to the discrimination of DID patients and normals, which apparently contradicts our view that the factors are too highly correlated to provide meaningful discrimination.

Our endorsement probability differences imply that the highest scoring DID patients endorse both amnesiac/dissociative and depersonalization/derealization symptoms, but lower scoring DID patients only endorse depersonalization/derealization items (in contrast to both DID subgroups, college students would uniformly tend to denying both item types, perhaps endorsing some of the absorption-imaginative involvement items). We thus suggest that amnesiac/dissociative items thus detect one DID subgroup and the *difference* between these items and depersonalization/derealization items detects the other subgroup. However, we suggest that these two sets of items simply fall at different regions of the same dimension rather than along separable dimensions. In fact, good test construction dictates a spread of endorsement probability (e.g., Nunnally & Bernstein, 1994, Chap. 8) to improve discrimination along the length of a continuum. What would falsify this argument is for one subgroup to endorse one type of item and the other subgroup to endorse the other. Unfortunately, even though

Gleaves et al (1995) was otherwise well reported, these subscale (component) score data were not presented. Given our findings that the two types of items are reported with highly disproportionate frequency, this latter outcome appears unlikely.

To illustrate further, assume a group of nondrinkers, moderate drinkers, and heavy drinkers answer two items truthfully: "I have at least one drink a week" and "I have at least one drink a day." Nondrinkers would not endorse either item, moderate drinkers would endorse the first but not the second, and heavy drinkers would endorse both items. Discriminant analysis would indicate that the two items each contribute uniquely to discrimination since their each helps identify a subgroup, as above. However, they clearly measure two points along the same dimension and not separate ones. Similarly, depression and schizophrenia are separable and not simply different regions along a single dimension even though they have different base rates because one can find depressives who are not schizophrenic and schizophrenics who are not depressed.

In sum, Tables 1 and 3 show that the items forming the previously defined "dimensions" of the DES and DIS-Q can be characterized in terms of their univariate distributions. Using one confirmatory factor analytic approach, multiple group analysis, we found that these differences in distribution explained the apparent multidimensionality of the two scales. On grounds of parsimony, we therefore conclude that the scales, and therefore the trait, are in fact unidimensional. Although these distributional differences are a fact, we acknowledge that, other methods, particularly those derived from structural equation modeling may suggest that there is multidimensionality, above and beyond these distributional differences. However, one must be very careful and show that the result is not an artifact of differences in item distributions.

REFERENCES

Bernstein, E. M., & Putnam, F. W. (1986). Development, reliability, and validity of a dissociation scale. *Journal of Nervous and Mental Diseases*, *174*, 727-735.

Bernstein I. H., & Eveland D. C. (1982). State vs. trait anxiety: A case study in confirmatory factor analysis. *Personality and Individual Differences*, *3*, 361-372.

Bernstein, I. H., & Garbin, C. P. (1985). A comparison of alternative proposed subscale structures for MMPI scale 2. *Multivariate Behavior Research*, *20*, 223-225.

Bernstein, I. H., & Gesn, P. R. (1997). On the dimensionality of the Buss/Perry Questionnaire, *Behaviour Research and Therapy*, *35*, 563-568.

Bernstein I. H., & Keith J. B. (1991). Reexamination of Eisen, Zellman, and McAlister's Health Belief Model Questionnaire. *Health Education Quarterly*, *18*, 207-220.

Bernstein, I. H., & Teng, G. (1989). Factoring items and factoring scales are different: Spurious evidence for multidimensionality due to item categorization. *Psychological Bulletin*, *105*, 467-477.

Bernstein, I. H., Teng, G., & Garbin, C. P. (1986). A confirmatory factoring of the self-consciousness scale. *Multivariate Behavioral Research*, *21*, 459-475.

Bissonnette, V., & Bernstein I. H. (1990). Artifacts can replicate: A reply to Pliavin and Charng. *Personality and Social Psychology Bulletin*, *16*, 554-561.

Campbell, D. T., & Fiske, D. W. (1959). Convergent and discriminant validation by the multitrait-multimethod matrix. *Psychological Bulletin*, *56*, 81-85.

Carlson, E. B., Putnam, F. W., Ross, C. A., Torem, M., Coons, P., Dill, D. L., Loewenstein, R. J., & Braun, B. G. (1993). Validity of the Dissociative Experiences Scale in screening for multiple personality disorder: A multicenter study. *American Journal of Psychiatry*, *150*, 1030-1036.

Carroll, J. B. (1945). The effect of task difficulty and chance success on correlations between items or between tests. *Psychometrika*, *10*, 1-20.

Cattell, R. B. (1966). The scree test for the number of factors. *Multivariate Behavioral Research*, *1*, 256-276.

Dubester, K. A., & Braun, B. G. (1995). Psychometric properties of the dissociative experiences scale. *Journal of Nervous and Mental Diseases*, *183*, 231-235.

Dunn, G. E., Ryan, J. J., & Paolo, A. M. (1994). A principal component analysis of the Dissociative Experiences Scale in a new substance abuse population. *Journal of Clinical Psychology*, *50*, 936-940.

Fischer, D. G., & Elnitsky, S. (1990). A factor analytic study of two scales measuring dissociation. *American Journal of Clinical Hypnosis*, *30*, 201-207.

Frankel, F. H. (1990). Hypnosis and dissociation. *The American Journal of Psychology*, *147*, 823-829.

Garbin, C. P., Robertson, E., & Bernstein, I. H. (1986). An analysis of the internal structure of Levenson and Gottman's Social Competency scale and suggestions for its use. *Journal of Consulting and Clinical Psychology*, *54*, 406-407.

Gleaves, D H., & Eberenz, K. P. (1995). Assessing dissociative symptoms in eating disordered patients: Construct validation of two self-report measures. *International Journal of Eating Disorders*, *18*, 99-102.

Gleaves, D. H., Eberenz, K. P., Warner, M. S., & Fine, C. G. (1995). Measuring clinical and non-clinical dissociation: A comparison of the DESW and QED. *Dissociation*, *8*, 24-31.

Gorsuch, R. L. (1983) *Factor analysis* (2nd Ed.). Hillsdale, NJ: LEA

Guttman, L. (1954). Some necessary conditions for common factor analysis. *Psychometrika*, *19*, 149-161.

Holzinger, K. J. (1944). A simple method of factor analysis. *Psychometrika*, *9*, 257-262.

Kaiser, H. J. (1970). A second generation little jiffy. *Psychometrika*, *35*, 401-417,

Kaiser, H. J. (1960). The application of electronic computers to factor analysis. *Educational and Psychological Measurement*, *20*, 141-151.

Kim, J. O., & Mueller, C. W. (1978). *Factor analysis: Statistical methods and practical issues*. Beverly Hills CA: Sage Publications.

Meehl, P. E., & Golden, R. R. (1982). Taxometric methods. In P. Kendall & J. Butcher (Eds.), *Handbook of research methods in clinical psychology* (pp. 127-181). New York: Wiley.

Meehl, P. E., & Yonce, L. J. (1994). Taxometric analysis: I. Detecting taxonicity with two quantitative indicators using means above and means below a sliding cut (MAMBAC procedure). *Psychological Reports*, *74*, 1059-1274.

Meehl, P. E. (1995). Bootstraps taxometrics: Solving the classification problem in psychopathology. *American Psychologist*, *50*, 266-275.

Nunnally, J. C., & Bernstein, I. H. (1994). *Psychometric Theory* (3rd Ed.). New York: McGraw-Hill.

Ray, W. J., June, K., Turaj, K., & Lundy, R. (1992). Dissociative experiences in a college age population: A factor analytic study of two dissociation scales. *Personality & Individual Differences*, *13*, 417-424.

Ross, C. A., Ellason, J. W., & Anderson, G. (1995). A factor analysis of the dissociative experiences scale (DES) in dissociative identify disorder. *Dissociation*, *8*, 229-235.

Ross, C. A., Joshi, S., & Currie, R. (1991). Dissociative experiences in the general population: A factor analysis. *Hospital and Community Psychiatry*, *42*, 297-301.

Thurstone, L. L. (1945). A multiple group method of factoring the correlation matrix. *Psychometrika*, *10*, 73-78.

Vanderlinden, J., van der Hart, O., & Varga, K. (1996). European studies on dissociation. In L. Michelson & W. Ray (Eds.) *Handbook on dissociation: Theoretical, empirical and clinical perspectives.* New York: Plenum Press. (pp. 25-49).

Vanderlinden, J., van Dyck, R., Vandereycken, W., & Vertommen, H. (1991). Dissociative experiences in the general population in the Netherlands and Belgium. *Dissociation*, *4*, 180-184.

Vanderlinden, J., van Dyck, R., Vandereycken, W., & Vertommen, H. (1993). Dissociation and traumatic experiences in the general population of the Netherlands. *Hospital and Community Psychiatry*, *44*, 786-788.

Vanderlinden, J., van Dyck, R., Vandereycken, W., & Vertommen, H., Verkes, R. J. (1993). The Dissociation Questionnaire (DIS-Q): Development and Characteristics of a new self-report questionnaire. *Clinical Psychology & Psychotherapy*, *1*, 21-27.

Vanderlinden, J., Vandereycken, W., van Dyck, R., & Vertommen, H. (1993). Dissociative experiences and trauma in eating disorders. *International Journal of Eating Disorders*, *13*, 187-193.

Vanderlinden, J., Varga, K., Peuskens, J., & Pieters, G. (1995). Dissociative symptoms in a population sample of Hungary. *Dissociation*, *8*, 205-208.

Waller, N. G., Putnam, F. W., & Carlson, E. B. (1996). Types of dissociation and dissociative types: A taxometric analysis of dissociative experiences. *Psychological Methods*, *1*, 300-321.

Waller, N. G., & Ross, C. A. (1997). The prevalence and biometric structure of pathological dissociation in the general population: taxometric and behavior genetic findings. *Journal of Abnormal Psychology*, *106*(4), 499-510.

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