Decomposing Global Self-Esteem

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ABSTRACT We argue in this paper for distinguishing two dimensions of global self-esteem, self-competence and self-liking. Studies 1 and 2 identify a corresponding pair of factors in Rosenberg's (1965) Self-Esteem Scale. Studies 3 and 4 examine the predictive value of the two-dimensional approach to self-esteem as reflected in the unique associations of self-competence and self-liking with negative life events and word recognition.

Decomposing Global Self-Esteem

Self-esteem has emerged as a central construct in psychological theory. The wide variety of definitions, models, and measures, however, reflects a lack of consensus on how self-esteem should be conceived. This problem has been lamented by reviewers of the literature over the years (e.g., Blascovich & Tomaka, 1991; Crandall, 1973; Shavelson, Hubner, & Stanton, 1976; Wells & Marwell, 1976;

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Wylie, 1974). In this paper, we highlight a distinction that appears to be one source of the inconsistency. First, we provide a theoretical basis for distinguishing two aspects of self-esteem. We then show how these aspects are represented in the most frequently used measure of the construct. Finally, we examine their construct validity in relation to negative life events and word recognition.

Self-Competence and Self-Liking

As an object of value, the self can be understood according to the axiological distinction between instrumental and intrinsic value (Dewey, 1939). Instrumental value refers to the utility of an object, or what it can do. Intrinsic value refers to qualities of an object that are considered good in themselves, without causal extension. Applied to persons, the distinction is reflected in personal competence on the one hand and characterological worth on the other. That is, an individual takes on value both by merit of what she can do and what she is. Informally, this is expressed as the difference between "respect" and "liking." The former is manifest as observable abilities, skills, and talents; the latter as moral character, attractiveness, and other aspects of social worth. The distinction, however, does not imply mutual exclusiveness, for abilities are often viewed as virtues, and virtues are often used to great effect. Despite this overlap, the distinction is worth maintaining for understanding the twofold genesis and expression of self-esteem.

The claim for two basic sources or forms of self-esteem, one based on ability and the other on worth or "goodness," is not without precedent (Brissett, 1972; Brown, 1998; Diggory, 1966; Franks & Marolla, 1976; Gecas, 1971; Silverberg, 1952; White, 1963). Tafarodi and Swann (1995) recently formalized the distinction in proposing self-competence and self-liking as interdependent but separate dimensions of global self-esteem. Self-competence is defined as the valuative experience of oneself as a causal agent, an intentional being with efficacy and power. Self-liking, on the other hand, is defined as the valuative experience of oneself as a social object, a good or bad person according to internalized criteria for worth. Accordingly, global self-esteem is conceived as consisting of two distinct dimensions of value. The conceptual separation reflects our existential duality as both autonomous agents and social beings (Bakan, 1966; Guisinger & Blatt, 1994). We suggest here that the frustrating diversity of definitions, measures, and theories of self-esteem is partly due to lack of formal recognition of this duality. Theorists have often emphasized one half or the other without appreciating their integral complementarity, or have blurred the distinction by adopting a unidimensional view that straddles both (see Wells & Marwell, 1976, for a review).

The most influential account of self-esteem was offered by Morris Rosenberg. During his career, Rosenberg consistently argued for a simple, unitary conception of self-esteem as "the feeling that one is 'good enough'" (1965, p. 31). The conceptual simplicity was achieved by subordinating self-competence to self-liking, which was understood as global self-esteem (Rosenberg, 1979). In claiming that self-competence "may contribute" to self-esteem, Rosenberg saw the former as a source of a higher-order valuation rather than as a constitutive dimension. This is a significant conceptual commitment. It implies that measurement of global self-esteem is tantamount to measurement of self-liking, broadly conceived as "goodness." In contrast, our view that self-esteem consists of self-liking and self-competence in the same way that length and width define a rectangle questions the experiential reality of any higher-order valuation more generalized than either of the two dimensions. Rather, self-esteem may be nothing more or less than self-competence and self-liking, just as the size of the rectangle is nothing more nor less than the composite of its length and width.

Deciding between these two conceptual positions is partly a pragmatic matter. A researcher's interest may not extend beyond self-liking. If so, Rosenberg's unitary construct may serve well enough. Elsewhere, adoption of a truncated view of self-esteem may be self-limiting. This is especially true in contexts where self-liking and self-competence hold divergent unique relations to variables of interest. Here, a unidimensional approach to theory and measurement would be inappropriate. The problem is compounded when measures designed to measure self-esteem as self-liking, or, alternatively, as some value superordinate to self-liking and self-competence, end up inadvertently tapping both dimensions without distinguishing them. Such is the case, we suggest, with Rosenberg's own well-known measure.

Rosenberg's (1965) Self-Esteem Scale

Rosenberg's version of unidimensionality is tacitly supported by a majority of today's social and personality psychologists, due mainly to the continuing popularity and widespread influence of his Self-Esteem Scale (SES). The SES, a simple, ten-item self-report instrument (see Appendix) with compelling face validity, has been the measure of choice in one-quarter of self-esteem studies published since its creation (Blascovich & Tomaka, 1991). Reflecting the conceptual commitment of its author, the scale is assumed to be unidimensional. Structural analysis, however, has revealed a somewhat different picture.

Exploratory factor analyses (EFA) of the items have revealed two distinguishable factors. Often, positively-worded items have loaded higher on one factor and negatively-worded items on the other. This pattern has led some to dismiss the apparent two-dimensional structure of the scale as method artifact due to response set (Carmines & Zeller, 1974; Hensley & Roberts, 1976; Zeller & Carmines, 1980). Others have interpreted the split as reflecting a substantive distinction between positive and negative self-esteem (Barber, 1990; Kaplan & Pokorny, 1969; Kohn & Schooler, 1969; Openshaw, Thomas, & Rollins, 1981; Shahani, Dipboye, & Phillips, 1990; Owens, 1993). Still others have remained noncommital on the meaning and importance of the distinction (Dobson, Goudy, Keith, & Powers, 1979).

Confirmatory factor analyses (CFA) have produced similarly mixed conclusions. Shelvin, Bunting, and Lewis (1995) found a one-factor model to provide acceptable statistical fit, but failed to assess the relative fit of a two-factor model. Goldsmith (1986) found better fit for a two-factor than one-factor model, and marshalled additional data to conclude that the split reflects a method artifact in some populations but a substantive distinction in others. Marsh (1996) examined responses to seven of the ten SES items, assessing the comparative fit of one-factor and two-factor measurement models. The best fit to the data was achieved by a one-factor model with correlated measurement errors among the three negative items and between two of the four positive items. Marsh took these findings to "support the existence of a single latent construct underlying responses to the SES items" and "undermine claims that two factors are needed" (p. 817). His conclusion, however, overstates the extent to which the results resolve the question of substantive vs. artifactual dimensionality. Any one-factor measurement model with correlated errors reflects multidimensionality and can therefore be alternatively specified using multiple latent factors (see Kline, 1998). Thus, the differential fit of a valence model specified using correlated errors rather than explicit factors is not instructive for deciding if the factors represent more than a method effect. A second limitation is that Marsh examined only two multidimensional measurement models, excluding any informed by the self-liking/ self-competence distinction. Finally, the study's reliance on only seven of ten SES items may misrepresent the dimensionality of the full scale.

Item response theory (IRT) offers an alternative means of examining scale characteristics. Rather than providing a framework for testing unidimensionality, however, IRT analysis assumes it (Hambleton, Swaminathan, & Rogers, 1991). As such, this form of analysis does not provide a formal means of resolving the issue of the SES's dimensionality. Even so, the response functions generated in IRT do allow metric consistency across items to be assessed and are therefore relevant to determining scale structure. The item discrimination parameter, *a*, estimated by IRT models is comparable in meaning to the item-factor loadings estimated in CFA. Both gauge the strength of association between response to the item and the latent trait it represents (Reise, Widaman, & Pugh, 1993). Estimates of *a* should be high across a set of items that purport to measure the same latent trait. Consistent with this, Gray-Little, Williams, and Hancock (1997) found all ten SES items to have adequate values of *a* when fit to a graded response model. The authors interpreted this finding as supporting Rosenberg's claim that the SES items define a single underlying construct, global self-esteem. The main limitation of their data, however, is that a set of items representing two highly correlated self-esteem dimensions would produce comparable results, with the single latent trait in the model representing the covariance of the two dimensions. This possibility is given credence by the authors' own principal components analysis of the items. The second factor extracted had an eigenvalue of 1.25, exceeding Kaiser's (1960) criterion for further consideration. Despite this result, the authors opted for a single-factor solution, partly because of a large disparity between first and second factor eigenvalues. A large primary and small secondary factor, however, is precisely what would emerge through orthogonal extraction in the case of two highly correlated but substantively distinct dimensions. Subsequent oblique rotation may have helped reveal a more balanced structure, but the authors chose to discount rather than pursue a two-dimensional solution.

Complicating interpretation of the SES's two factors is their high intercorrelation. Proponents of unidimensionality have taken the commonality to suggest the dominance of a general self-esteem factor and the relative unimportance of more refined distinctions between scale items. Such reasoning, though tempting, oversteps caution. All

SES item intercorrelations can be assumed to be inflated by shared method variance, exaggerating the commonality between factors and thereby obscuring simple structure. More importantly, though, high correlation between two factors is not itself sufficient justification for abandoning their conceptual separation. Rather, discriminant validity, especially evidence for incremental prediction, should be examined in such cases (Macmann & Barnett, 1994). If each factor offers predictive utility beyond the other, then their unique or independent variances should not be collapsed together. Furthermore, if divergent patterns of associations with theoretically linked variables can be built around the unique variances, then the factors represent a meaningful distinction, irrespective of the magnitude of their intercorrelation (Cronbach & Meehl, 1955).

No published research specifically addressing the incremental validity of the two SES factors could be found at the time of writing. Examination of other aspects of discriminant validity has been rare and evidence for associative divergence has been mixed (Carmines & Zeller, 1974; Goldsmith, 1986; Hagborg, 1996; Owens, 1993, 1994). One problem common to these studies is their restriction to a valence-based factor structure in examining discriminant validity. The separation of self-esteem into positive and negative attitudinal dimensions, though plausible (see Cacioppo & Bernston, 1994), is neither intuitively nor theoretically compelling. For example, the argument that "I am a good person" and "I am not a good person" are statements representing different dimensions is somewhat perplexing. This limitation lends credence to the claim that the distinction between positive and negative self-esteem is more artifactual than substantive (Marsh, 1996; Zeller & Carmines, 1980).

In revisiting the dimensionality of the SES, we propose that there are indeed two substantive dimensions underlying the scale, but they represent a semantic distinction that ultimately reduces to self-liking and self-competence rather than positive and negative self-esteem.

The SES as Assessment and Acceptance

The SES appears to split equally into two types of items: those that invite assessment of qualities and those that imply self-acceptance. Thus, the statements "I feel I have a number of good qualities," "I am able to do things as well as most other people," "I feel I do not have much to be proud of," "I feel I'm a person of worth, at least on an equal plane with others," and "All in all, I am inclined to feel that I am a failure" orient the respondent toward objective self-assessment of personal qualities, especially abilities ("do things," "failure"). The feeling of worth is linked to objective comparison ("on an equal plane with others"); the feeling of pride is presented as having grounds, which, in western culture, reduces mainly to one's past accomplishments; and the feeling of failure is prefaced as implicitly cumulative or summative ("all in all"), suggesting a generalization of what one has and has not achieved. The "good qualities" to be numbered are left vague, but justification is implicit in the tallying.

In contrast, "On the whole, I am satisfied with myself," "At times I think I am no good at all," "I wish I could have more respect for myself," "I take a positive attitude toward myself," and "I certainly feel useless at times" are more subjective in orientation, highlighting the extent to which one is happy with and accepting of oneself, justifiably or not. This is clearest for three of the items ("satisfied," "positive attitude," "no good"). Wishing for more self-respect implies recognition that one ought to feel better about (like?) oneself as one is, rather than a desire to become more competent and thereby self-respectable. Finally, though ostensibly referring to competence, "useless" is often used informally as a synonym for "unworthy," again implying acceptance of who and what one is.

This speculative distinction between *assessment* and *acceptance* describes the unique semantic differentiation of the SES items. On closer inspection, however, the ad hoc dichotomy is largely redundant with self-competence and self-liking and may be less fundamental for our understanding of global self-esteem. Specifically, the objective, ability-oriented thrust of the assessment items clearly reflects self-competence, whereas the more subjective, worth-oriented thrust of the acceptance items clearly reflects self-liking. To explore this redundancy, Tafarodi and Swann (1995) formed two distinct item parcels from the SES by combining three assessment and three acceptance items. Using CFA, the assessment parcel was modeled as an indicator of self-competence and the acceptance parcel as an indicator of self-liking. Results revealed that the SES indicators loaded on self-competence and self-liking about as well as did validated indicators of the dimensions, suggesting that the dimensionality of the SES may be reducible to self-competence and self-liking. This conclusion, however, is undermined by the limitations of the study, most notably the exclusion of four SES items and reliance on item parcels rather than individual items. Moreover, no compelling support for the theoretical basis and functional significance of the competence-liking distinction was offered.

In the studies that follow, we sought to better gauge the significance of self-competence and self-liking for understanding the SES and, more importantly, global self-esteem in general. We first examined the validity of the assessment-acceptance differentiation of the SES items (Study 1). Next, we examined the validity of interpreting these scale-specific factors as self-competence and self-liking (Study 2). Finally, we moved beyond the SES to assess the discriminant validity of the two-dimensional conception of self-esteem in predictive contexts. Specifically, we examined the unique associations of self-competence and self-liking with negative life events (Study 3) and word recognition (Study 4).

STUDY 1

Overview

Using CFA, three SES measurement models, representing unidimensional, positive-negative and assessment-acceptance interpretations of the items, were compared in their fit with the responses of a large sample of university students. Of the three models, assessment-acceptance provided the best fit. Combining all three models into a five-factor structure provided even better fit and clarified the full dimensionality of the scale.

METHOD

Participants

Participants were 836 students (418 men and 418 women)[1] enrolled in an introductory psychology course at the University of Toronto. The modal age was 19.

1. The strict gender equivalence was achieved through random exclusion of women's data at the time of analysis. The parity permitted balanced comparison of models across gender, to maximize the generalizability of findings. Throughout this paper, all reported results pertaining to comparative model fit and the significance of parameter estimates held equally for men and women, as confirmed through separate testing. For economy, quantitative results are presented only for the combined sample.

Materials and Procedure

The ten SES items were completed in standard administration order (Rosenberg, 1979; see Appendix) as part of a mass testing session conducted at the beginning of the fall term. Designed as a 6-point Guttman scale, the SES has been most often used in research as a simple summated scale, with 4- or 5-point Likert-type ratings. Consistent with this, responses were made on a 5-point scale, anchored with *strongly disagree* and *strongly agree*.

RESULTS

As a precondition for collapsing the sample across gender, the discrete response distributions for men and women were compared using the χ^2 statistic. Negative items were reverse-scored prior to analysis so that higher values uniformly represented higher self-esteem. Results revealed that the male and female distributions were not significantly different ($\alpha = .05$) for any of the ten items, reflecting gender equivalence in item response. Consistent with previous findings, the response distributions were clearly nonnormal. For six of the items, the modal response was 5, the right endpoint of the scale. A majority of the item distributions had skewness or kurtosis values > 1. Consistent with these characteristics, the Shapiro-Wilk test of normality (Shapiro & Wilk, 1965) was significant at $p < .0001$ for all items.

CFA of the SES

With only five response values, the SES items are best viewed as ordered categorical indicators of underlying continuous dimensions. One approach to dealing with such items in CFA requires estimation of asymptotic covariances from polychoric correlations (Lee, Poon, & Bentler, 1990; Jöreskog, 1994; Muthén, 1984). Reliable estimation in this context, however, requires very large samples. Moreover, it may be unnecessary in the present case. Bollen and Barb (1981) found that the simple correlation of five-category variables is fairly accurate, reproducing about 90% of the true (continuous) correlation. They concluded that the five-category case may define a "threshold," in that using more than five categories affords only modest gains toward reproducing the underlying correlation. Consistent with this, Tepper and Hoyle (1996) found that CFA model estimation using the asymptotic rather than standard covariance matrix produced only negligible differences in the five-category case. Given these findings, we deemed it justified to analyze the categorical SES items as if they were continuous. The clear nonnormality of the items, however, precluded the straightforward use of maximum likelihood (ML) or normal theory generalized least squares methods (Potthast, 1993; West, Finch, & Curran, 1995). We therefore opted for ML estimation with Satorra-Bentler (1994) "robust" statistics, which are adjusted in proportion to sample non-normality. The Satorra-Bentler statistics appear to perform nearly as well in moderate and large samples as do newer test statistics based on either ML or Browne's (1984) asymptotically distribution free (ADF) estimator (Bentler & Yuan, 1999; Yuan & Bentler, 1998, 1999). Model specification and testing was conducted using EQS 5.7b.

Measurement Models

Three competing models were tested, as shown in Figure 1. For the Unidimensional model, all ten items were specified as indicators of a single factor. For the Positive-Negative model, the five negatively-

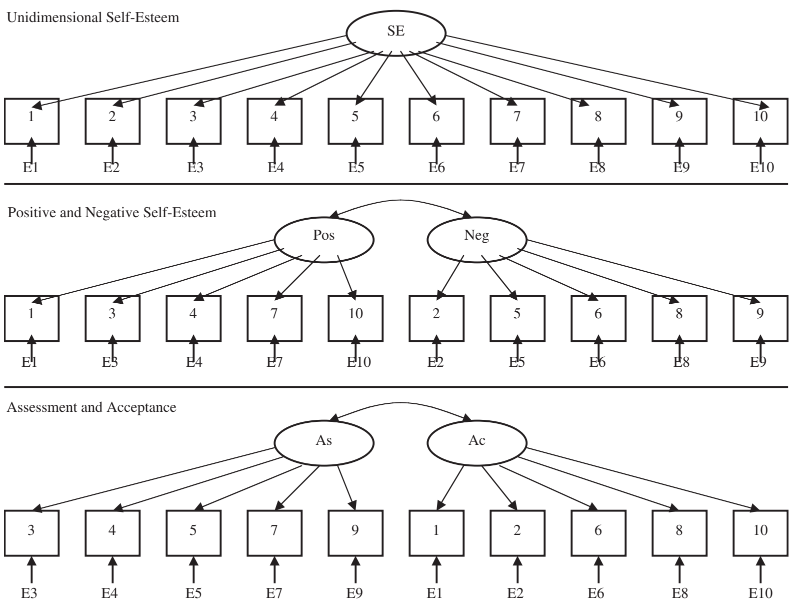

Figure 1
SES measurement models. Items are identified by their ordinal position (1–10) in the administration version of the scale (see Appendix). E = error/uniqueness.

worded items were specified as indicators of one factor and the five positively-worded items as indicators of a second factor. For the Assessment-Acceptance model, the five assessment items were specified as indicators of one factor and the five acceptance items as indicators of a second factor. For all three models, error covariances were constrained to zero. The factor correlation was freely estimated for both two-factor models. The model fit indices were: the χ^2 test statistic, Satorra-Bentler's scaled (nonnormality-adjusted) χ^2 (S-Bχ^2), the robust Comparative Fit Index (RCFI; Bentler, 1995), the Root Mean Squared Error Approximation (RMSEA; Steiger, 1990), and the consistent version of Akaike's (1987) information criterion (CAIC; Bozdogan, 1987). Lower values indicate better fit for all indices other than RCFI, where higher values indicate better fit.

Consistent across models, all standardized factor loadings were > .50 and highly significant (p < .0001). The latent factor correlation was .83 for Positive-Negative and .80 for Assessment-Acceptance. Goodness-of-fit results appear in Table 1. As the Unidimensional model is essentially a singly constrained version of the other two, its fit was

Table 1
Goodness-of-Fit for SES Measurement Models in Studies 1 and 2

Model	*df*	χ^2	S-Bχ^2	RCFI	RMSEA	CAIC
Study 1						
Null	45	3763	–	–	–	–
Unidimensional Self-Esteem	35	621	448	.82	.14 (.13–.15)	350
Positive and Negative Self-Esteem	34	497	390	.84	.13 (.12–.14)	234
Assessment and Acceptance	34	425	309	.88	.12 (.11–.13)	162
Combined	15	19	16	.99	.02 (.00–.04)	−97
Study 2						
Null	45	7105	–	–	–	–
Unidimensional Self-Esteem	35	864	661	.87	.12 (.11–.13)	570
Positive and Negative Self-Esteem	34	693	543	.89	.11 (.10–.12)	407
Assessment and Acceptance	34	618	474	.91	.10 (.09–.11)	333
Combined – Fixed Loadings	45	67	56	.99	.02 (.01–.03)	−311

Note. S-Bχ^2 = Satorra-Bentler scaled χ^2; RCFI = robust Comparative Fit Index; RMSEA = Root Mean Square Error Approximation (90% confidence interval appears in parentheses); CAIC = consistent version of Akaike's Information Criterion.

statistically compared against them using χ^2 difference tests. The Unidimensional model provided significantly worse fit than did both Positive-Negative, S-B$\chi^2_{\text{diff}}(1) = 58$, $p < .0001$, and Assessment-Acceptance, S-B$\chi^2_{\text{diff}}(1) = 139$, $p < .0001$. These comparisons are tantamount to testing the null hypothesis that the factor intercorrelation is equal to one. Thus, they also support the discriminant validity of the highly correlated factors in both two-dimensional models. Because these two models are not hierarchically related, it was not possible to formally test the difference in their fit. Nonetheless, the superiority of Assessment-Acceptance over Positive-Negative was consistently apparent across fit indices.

The fit of the Assessment-Acceptance model, although better than its competitors, was itself inadequate, failing to account for at least 90% of the covariation among the observed variables. Moreover, the superiority of the Positive-Negative over the Unidimensional model was consistent with Marsh's (1996) and Zeller and Carmines' (1980) proposal that the SES is characterized by correlated errors among items of the same valence. We therefore decided to effectively combine all three models[2] to determine whether significant item variance was accounted by the assessment-acceptance and valence distinctions, respectively, beyond variance common to all ten items. In this combined five-factor model, each item was modeled as loading on three factors: a common factor, a positive (for positively-worded items) or negative (for negatively-worded items) factor, and an assessment (for assessment items) or acceptance (for acceptance items) factor. The factors were specified as uncorrelated.

The combined model fit very well (see Table 1) and was not rejected, S-B$\chi^2(15) = 16$, $p = .36$. All 30 factor loadings were positive, as expected. The common factor loadings were consistently significant and 7/10 positive/negative loadings and 7/10 assessment/acceptance loadings were significant (see Table 2). Same-item factor loadings were statistically compared using univariate Lagrange multiplier tests. Results revealed that the common factor loading differed significantly

2. Because the Positive-Negative and Assessment-Acceptance item groupings are partially overlapping (more positive than negative Assessment items and more negative than positive Acceptance items), combining the Unidimensional model with each of the two-dimensional models alone would have obscured the unique contributions of each of the latter to item variance.

Table 2
Standardized Factor Loadings for SES Combined Model in Study 1

SES Item	Item R^2
1 = .63*Cf + .28*Po + .24*Ac	.53
2 = .49*Cf + .57*Ne + .30*Ac	.66
3 = .59*Cf + .42*Po + .53*As	.80
4 = .53*Cf + .36*Po + .35*As	.53
5 = .71*Cf + .10 Ne + .14 As	.54
6 = .52*Cf + .56*Ne + .21*Ac	.63
7 = .63*Cf + .45*Po + .16 As	.63
8 = .53*Cf + .08 Ne + .42*Ac	.47
9 = .74*Cf + .05 Ne + .10 As	.55
10 = .69*Cf + .21* Po + .45*Ac	.72

Note. Cf = Common Factor; Po = Positive; Ne = Negative; Ac = Acceptance; As = Assessment. Factor loadings marked with an asterisk are positive at $p < .05$.

from the positive/negative loading for 5/10 items (1, 5, 8, 9, 10). For all five, the common factor loading was higher than the positive/negative loading in the unconstrained solution. The common factor loading also differed significantly from the assessment/acceptance loading for 6/10 items (1, 5, 6, 7, 9, 10). For all six, the common factor loading was higher than the assessment/acceptance loading in the unconstrained solution. Finally, the positive/negative and assessment/acceptance loadings were compared, revealing significant differences for items 1, 2, 10. The positive/negative loading was higher than the assessment/acceptance loading for the first two but lower for the last item in the unconstrained solution.

The combined model accounted for 60% of item variance on average. Of this amount, 61% was attributable to the common factor, 22% to the positive and negative factors, and 17% to the assessment and acceptance factors.

DISCUSSION

The CFA results provide some support for our claim that the SES is more than a unidimensional scale. The single-factor measurement model did not fit as well as either of the two-factor models.

Futhermore, of the two-factor alternatives, the assessment-acceptance model fit the best, suggesting that our proposed semantic distinction captures an important aspect of the scale's dimensionality. On the other hand, the combined, five-factor model, revealed that a common factor accounted for the lion's share of reliable variance across items, with positive/negative and assessment/acceptance contributing only modest increments beyond that. Moreover, the increment offered by assessment/acceptance independent of positive/negative was not significant for 3 of 10 items. The dominance of the common factor is why many researchers have been unmoved by competing claims of a more complex dimensionality. We certainly agree that the strong expression of a common factor offers some justification for claiming that the items are more alike than different. We are not convinced, however, that the common factor must be interpreted as "general" or "global" self-esteem, as some have been quick to argue. Such an interpretation may well be correct, but other possibilities are equally plausible. First, the common factor may represent a method artifact, the "halo effect" being the most obvious candidate (Murphy, Jako, & Anhalt, 1993). The halo effect would manifest here as an invalid transfer or generalization of positive (or negative) judgment across substantively distinct SES items. If strong enough, such an effect would grossly inflate the inter-item correlations and express itself as a dominant common factor. Other, contextual factors that vary across individuals and have generalized effects on responding might further inflate the correlations, worsening the problem. Second, the common factor may represent nothing more than the considerable causal interdependence, over time, of the two generalized dimensions that correspond to assessment and acceptance. We will have more to say on this later.

In sum, the assessment-acceptance distinction helped account for the full dimensionality of the SES. But how useful are these dimensions for understanding self-esteem in general? Deriving ad hoc distinctions from the structural peculiarities of a single measure may be counterproductive if it adds to the proliferation of highly redundant trait constructs. When such distinctions mirror established ones that are grounded in theory, it is appropriate to examine the possibility of parsimonious absorption. As described earlier, the close parallel between assessment-acceptance and competence-liking suggests that the former might be practically reduced to the latter. This was explored next.

STUDY 2

Overview

To provide replication, the three competing SES measurement models were again compared using CFA. Study 1 parameter estimates for the combined, five-factor measurement model was then cross-validated in the new sample. All models were also compared on their fit with the SLCS, a self-report measure of self-competence and self-liking. Finally, a combined SES-SLCS model was tested to gauge the redundancy of the assessment-acceptance and competence-liking distinctions.

METHOD

Participants

Participants were 1648 students (824 men and 824 women) enrolled in introductory psychology courses at the University of Toronto. The modal age was 19.

Material and Procedure

The SES was administered as before, but together with the 20-item Self-Competence/Self-Liking Scale (SLCS; Tafarodi & Swann, 1995; see Appendix). As the two measures use the same response scale, they were administered in combined form with SLCS items following SES items. The SLCS divides into two 10-item subscales, one designed to measure self-competence and the other self-liking. Both subscales have an equal number of positively and negatively worded items. In validating the measure, Tafarodi and Swann (1995) found coefficient alphas of .89 and .92 and uncorrected test-retest (3-week interval) reliabilities of .80 and .78 for self-liking and self-competence, respectively, with the subscales correlated at .69.

RESULTS

SES

The three competing SES models—Unidimensional, Positive-Negative, and Assessment-Acceptance—were specified as before. Results revealed that, across models, all standardized factor loadings were $> .50$ and highly significant ($p < .0001$). The factor correlation was .87 for Positive-Negative and .85 for Assessment-Acceptance. Goodness-of-fit results appear in Table 1. As before, the Unidimensional model provided significantly worse fit than did both Positive-Negative, S-B$\chi^2_{\text{diff}}(1) = 118$, $p < .0001$, and Assessment-Acceptance, S-B$\chi^2_{\text{diff}}(1) = 187$, $p < .0001$, with Assessment-Acceptance superior to Positive-Negative on all indices.

The combined, five-factor model was then tested. To provide a strict cross-validation, we fixed the 30 factor loadings to the values found in Study 1. This highly constrained model fit very well (see Table 1) and was not rejected, S-Bχ^2 (45) = 56, $p = .13$. Morever, simultaneous Lagrange multiplier testing of the fixed loadings revealed that only one (the loading of item 7 on the assessment factor) would provide a significant increment in fit ($p = .007$) if freely estimated. This limited difference, coupled with the non-rejection of the constrained model, amounts to a high degree of cross-validity for the five-factor measurement model (see MacCallum, Roznowski, & Necowitz, 1992).

SLCS

As a precondition for examining the redundancy of the SES and SLCS, the a priori two-dimensional measurement structure of the SLCS was tested using CFA. This structure has been confirmed elsewhere (e.g., Tafarodi & Swann, 1995), but never against competing models at item level, as is most appropriate. Only stringent CFA validation would warrant using the measure as a reference for reinterpreting the SES. The distributional properties of the SLCS items are highly similar to those of the SES. Thus, ML estimation with Satorra-Bentler statistics were again used for model testing. Negative items were reverse-scored prior to analysis.

Three competing measurement models were tested (see Figure 2). The Unidimensional and Positive-Negative models mirror those tested for the SES, with the latter distinguishing positively from negatively worded items. The Competence-Liking model reflects the a priori target loadings of the items as indicators of self-competence and self-liking. For all three models, error covariances were constrained to zero. Factor correlations were freely estimated.

Across models, all standardized factor loadings were > .50 and highly significant ($p < .0001$). The latent factor correlation was .91 for Positive-Negative and .80 for Competence-Liking. Goodness-of-fit results appear in Table 3. Chi-square difference tests confirmed that the Unidimensional model provided significantly worse fit than did both of the two-dimensional models. Of the latter two, the a priori

Competence-Liking model was clearly superior, accounting for 92% of the covariation among the observed variables.

Although the fit of the Competence-Liking model was arguably adequate, we chose to examine the unique significance of the positive-negative and competence-liking distinctions beyond variance common to all 20 items, just as we did for the SES in Study 1. We again specified a combined, five-factor model (common, positive, negative, competence, and liking), with each item loading on three factors.

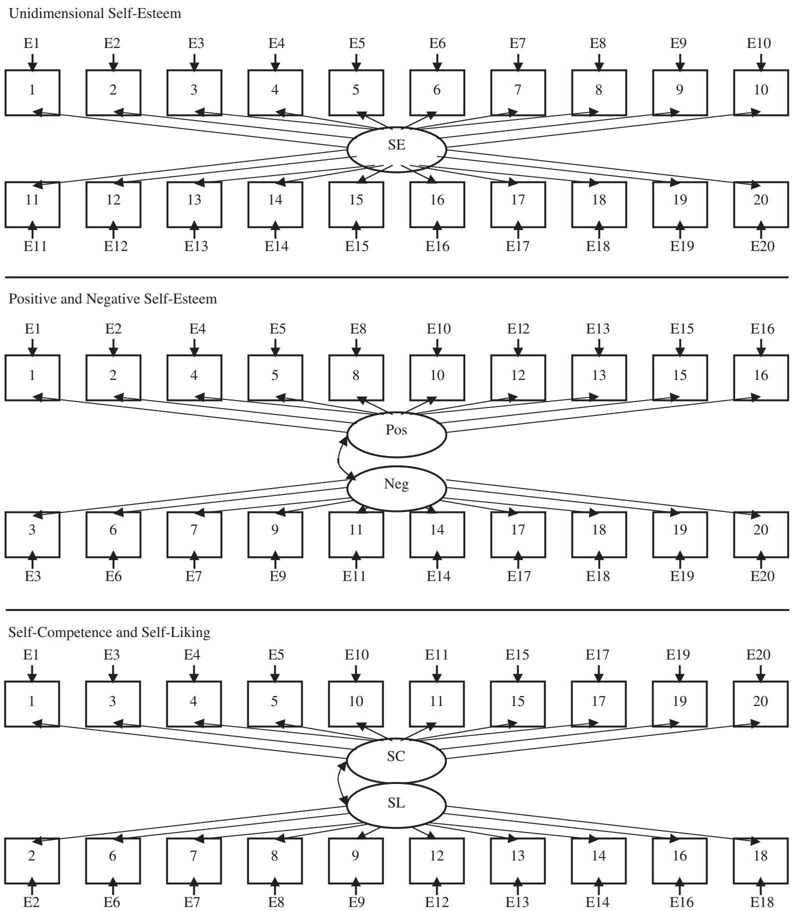

Figure 2
SLCS measurement models. Items are identified by their ordinal position (1–20) in the administration version of the scale (see Appendix). E = error/uniqueness.

Table 3
Goodness-of-Fit for SLCS Measurement Models in Study 2

Model	*df*	χ^2	S-Bχ^2	RCFI	RMSEA	CAIC
Null	190	19999	–	–	–	–
Unidimensional Self-Esteem	170	3268	2309	.84	.11 (.10–.11)	1839
Positive and Negative Self-Esteem	169	2874	2051	.86	.10 (.10–.10)	1453
Assessment and Acceptance	169	1715	1220	.92	.08 (.07–.08)	294
Combined	130	402	293	.99	.04 (.03–.04)	−691

Note. S-Bχ^2 = Satorra-Bentler scaled χ^2; RCFI = robust Comparative Fit Index; RMSEA = Root Mean Square Error Approximation (90% confidence interval appears in parentheses); CAIC = consistent version of Akaike's Information Criterion.

Although this model was rejected, S-B$\chi^2(130) = 293$, $p < .0001$, the fit was very good (see Table 3). Absolute rejection of well-specified but complex models is typical when sample size is large, owing to the high power sensitivity of the χ^2 test. All significant factor loadings were positive, as expected. The common factor loadings were consistently significant. 18/20 positive/negative loadings and 19/20 competence/liking loadings were significant (see Table 4). Same-item factor loadings were statistically compared using univariate Lagrange multiplier tests. Results revealed that the common factor loading differed significantly from the positive/negative loading for 12/20 items (2, 3, 4, 6, 8, 10, 11, 12, 14, 16, 17, 18). For all twelve, the common factor loading was higher than the positive/negative loading in the unconstrained solution. The common factor loading also differed significantly from the competence/liking loading for 11/20 items (1, 4, 6, 7, 8, 9, 10, 11, 12, 13, 18). For all but one of the eleven, the common factor loading was higher than the competence/liking loading in the unconstrained solution. Finally, the positive/negative and competence/liking loadings were compared, revealing significant differences for 11/20 items. The positive/negative loading was higher than the competence/liking loading for three of the eleven (12, 16, 20) but lower for the others (3, 4, 6, 7, 10, 11, 14, 17) in the unconstrained solution.

The combined model accounted for 60% of item variance on average. Of this amount, 68% was attributable to the common factor,

Table 4
Standardized Factor Loadings for SLCS Combined Model in Study 2

SLCS Item	Item R^2
1 = .46*Cf + .37*Po + .48*Co	.58
2 = .77*Cf + .22*Po + .10*Li	.64
3 = .59*Cf + .05 Ne + .46*Co	.56
4 = .62*Cf + .16*Po + .23*Co	.47
5 = .54*Cf + .40*Po + .44*Co	.64
6 = .67*Cf + .07*Ne + .21*Li	.49
7 = .62*Cf + .07*Ne + .55*Li	.69
8 = .55*Cf + .25*Po + .24*Li	.43
9 = .65*Cf + .08*Ne + .34*Li	.54
10 = .63*Cf + .26*Po + .41*Co	.63
11 = .66*Cf − .05 Ne + .35*Co	.56
12 = .79*Cf + .22*Po + .15*Li	.69
13 = .82*Cf + .24*Po + .01 Li	.73
14 = .65*Cf + .06*Ne + .39*Li	.58
15 = .50*Cf + .42*Po + .44*Co	.62
16 = .83*Cf + .22*Po + .07*Li	.75
17 = .55*Cf + .14*Ne + .43*Co	.50
18 = .75*Cf + .07*Ne + .34*Li	.69
19 = .41*Cf + .44*Ne + .34*Co	.48
20 = .48*Cf + .61*Ne + .25*Co	.66

Note. Cf = Common Factor; Po = Positive; Ne = Negative; Co = Competence; Li = Liking. Factor loadings marked with as asterisk are positive at $p < .05$.

12% to the positive and negative factors, and 20% to the competence and liking factors.

Combined Measurement Model

Absorption of the ad hoc assessment-acceptance distinction within the more established competence-liking distinction, if justified, would simplify matters by wedding the multidimensionality of the SES to a pair of constructs with known nomological relations.

The overlap of the SES and SLCS is reflected in the similarity of items, some of which are nearly identical ("I wish I could have more respect for myself" [SES] vs. "I do not have enough respect for myself" [SLCS], and "I feel I do not have much to be proud of" [SES] vs. "I do not have much to be proud of" [SLCS]). Reduction of one measure to the other presumes a high proportion of shared variability. To gauge this commonality, SES total score was simultaneously regressed on the self-competence and self-liking subscale scores and the interaction of these predictors. Gender and the interactions of gender with the three previous predictors were included in the initial model. As neither gender nor any of the interactions were uniquely associated with SES score (smallest $p = .23$), these predictors were eliminated from the model. R^2 for the reduced model, with only self-liking and self-competence as predictors, was quite high at .83. Moreover, both dimensions of the SLCS were uniquely and strongly associated with SES score: $\beta = .36$, $t\,(1645) = 24.24$, $p < .0001$ for self-competence, and $\beta = .61$, $t\,(1645) = 40.88$, $p < .0001$ for self-liking.

Reduction also requires evidence that the SES assessment and acceptance items are congeneric indicators of competence and liking (see Nunnally & Bernstein, 1994). To look at this, we specified an expanded, seven-factor (common, positive, negative, assessment, acceptance, competence, liking) combined model that included both SES and SLCS items. Each item was defined as loading on three factors: common, positive/negative, and assessment/acceptance (for SES items) or competence/liking (for SLCS items). All factor correlations were implicitly fixed to zero except for assessment-competence and acceptance-liking. If the two pairs of parallel constructs are as highly redundant as expected, their freely estimated correlations should be very strong, despite the high degree of residualization on the common, positive, and negative factors. The model fit well (see Table 5), although it was again strictly rejected, S-B$\chi^2(343) = 1011$, $p < .0001$. More importantly, the assessment-competence and acceptance-liking correlations were .80 and .92, respectively. Correlations this high in a context where shared item variance attributable to common and valence factors has already been accounted for is consistent with the claim that the assessment-acceptance and competence-liking are largely redundant. Even so, the redundancy was not complete. Constraining the factors correlations to unity produced a significant decrement in fit, S-B$\chi^2_{\text{diff}}(2) = 62$, $p < .0001$, with Lagrange multiplier tests revealing that both constraints contributed significantly to the decrement. For practical purposes, however, the high redundancy justifies subsuming the assessment and acceptance factors of the SES to the more general self-esteem dimensions of competence and liking.

Table 5
Goodness-Of-Fit for Combined SES-SLCS Measurement Models in Study 2

Model	*df*	χ^2	S-Bχ^2	RCFI	RMSEA	CAIC
Null	435	32308	–	–	–	–
7-Factor w/ Correlated Parallel Factors:						
Freely Estimated	343	1347	1011	.97	.04 (.04–.05)	−1537
Constrained to Unity	345	1430	1073	.97	.04 (.04–.05)	−1470

Note. S-Bχ^2 = Satorra-Bentler scaled χ^2; RCFI = robust Comparative Fit Index; RMSEA = Root Mean Square Error Approximation (90% confidence interval appears in parentheses); CAIC = consistent version of Akaike's Information Criterion.

Finally, we examined the τ-equivalence (equal true-score variability) of the two measures on the critical factors, as reflected in the estimates from the constrained seven-factor model. The average standardized loadings for SES vs. SLCS indicators of acceptance-qua-liking were comparable (.24 vs. .23, respectively). The average standardized loading for SES indicators of assessment-qua-competence, however, was somewhat lower than that of SLCS indicators (.32 vs. .39, respectively), implying somewhat better per-item measurement of competence by the SLCS.

DISCUSSION

The successful cross-validation of the CFA results of Study 1 reinforces our multidimensional account of the SES. Clearly, there is a dominant common factor that binds all ten items. Beyond this common factor, however, valence and assessment-acceptance distinctions are needed to fully characterize the dimensionality of the scale. Furthermore, the assessment and acceptance factors appear to be highly redundant with self-competence and self-liking, justifying the conceptual absorption of the former within the latter. The SES items, however, are slightly weaker indicators of competence than are their SLCS counterparts.

The dominance of the common factor is reflected in the high cross-correlations of competence and liking items. Formal tests of discriminant validity aside, the considerable overlap cannot be ignored. Specifically, conceptual separation must be shown to correspond to predictive efficiency (Sechrest, 1963). If the competence-liking distinction is to be of any real value in research on self-esteem, the reliable variance unique to each dimension must be shown to relate to psychological phenomena that the other does not. The divergent relations should be consistent with the distinct hypothetical origins and ramifications of each dimension (Cronbach & Meehl, 1955). If this cannot be shown, then the discriminant validity confirmed through CFA is of trifling practical significance. The final two studies were conducted to provide such evidence. Study 3 examines the impact of distinct types of negative life events on each dimension of self-esteem over a 4-week period. Study 4 examines perceptual selectivity in the recognition of positive and negative words reflecting the two dimensions.

STUDY 3

Overview

College students completed the SLCS on two occasions four weeks apart. They also provided a retrospective record of negative life events on the second occasion. Time 1 → Time 2 change in self-competence and self-liking was examined as a function of intervening life events. Consistent with theory, we predicted that each dimension of self-esteem would be affected by a specific type of adversity.

METHOD

Participants

Participants were 244 students (174 women and 70 men) enrolled in introductory psychology courses at the University of Wales, Cardiff, and the University of Toronto. The modal age was 19.

Materials and Procedure

Participants completed several paper-and-pencil measures on two occasions, 4 weeks apart, including two that are relevant here: the SLCS and a modified form of the Life Events Record (LER; Tafarodi & Walters, 1999).

The SLCS was the first measure completed at both Time 1 and Time 2. We chose it over the SES as a measure of self-competence and self-liking because of its greater reliability (10- vs. 5-item scales) and the superiority of its competence items.

The LER is a retrospective measure of life events. Respondents recall any personally significant events that occurred during a specified time period. The abbreviated form used here asks only for negative life events experienced during the four weeks between testing sessions. Each event is briefly described in writing and the subjective intensity of its negative impact is rated on a 9-point scale anchored by *mild* (1) and *very strong* (9). Space is provided for up to 10 events. Frequency of negative events, optionally weighted by intensity ratings, is calculated. Given its reliance on deliberate recall, the LER is best used for relatively short retrospective periods. In contrast to standard life event inventories (checklists), its open-ended format provides a personalized record of what the *respondent* experienced as significant, irrespective of how notable these events would have been for others.

The LER was completed only at Time 2, following the SLCS and separated from it by several unrelated measures. The instructions asked participants to report any personally significant negative events that had occurred in the four weeks since Time 1.

Although characterized by a fair degree of stability over time, global self-esteem has been shown to decrease in response to negative life events (Joiner, Katz, & Lew, 1999). Given the distinct hypothetical origins of self-competence and self-liking, described at the outset, each dimension of self-esteem should be especially sensitive to those events that are thematically matched with it in self-valuative relevance. That is, self-competence should be more responsive to achievement-related events than to social events involving negative evaluation by others or other events. Self-liking, in contrast, should be more responsive to social events involving negative evaluation than to achievement-related or other events. These predictions were tested using path analysis.

RESULTS

Two female participants were eliminated as clear multivariate outliers on the basis of discontinuously high values of robust Mahalanobis D^2 for the variables analyzed below (see Khattree & Naik, 1995). This left a final sample size of 242. As all results reported below were parallel for men and women, and for students from the two universities, gender and nationality will not be discussed further.

Responses on the LER were categorically differentiated to create domain-specific negative life event scores. Specifically, a pair of judges blind to the purpose of the study independently classified all events reported by participants using a three-category scheme.

Events of primarily interpersonal significance involving negative evaluation (e.g., criticism from parents, romantic rejection) were placed in the *Social* category.[3] Events of primary significance for ability or competence (e.g., failing an exam, arriving late for a job interview) were placed in the *Achievement* category. Remaining events (e.g., suffering the flu, witnessing a horrible accident) were placed in the *Other* category. Cohen's (1960) κ across judges was .90, a high level of chance-corrected agreement. Disagreements were resolved through discussion.

To avoid any distortion of results due to potential confounding of self-esteem with intensity ratings, only the *unweighted* or simple event frequencies were used in hypothesis-testing.[4] For each participant, separate frequencies were calculated for Social (M = .46), Achievement (M = .47), and Other events (M = .92).

Simple (manifest variable) path analysis was conducted to test the predictions. Both self-competence and self-liking at Time 2 were modeled as endogenous variables predicted by self-competence and self-liking at Time 1 and the three frequencies of negative life events. This is tantamount to predicting change in self-esteem from reported events (see Cronbach & Furby, 1970). The Time 1 SC-SL covariance was included, as were all covariances among life event categories. Finally, the Time 2 SC-SL error covariance was included to accommodate shared residual variance due to concurrent measurement. The resulting model was estimated solely for the purpose of testing critical path coefficients. The standardized coefficients appear in Figure 3, where error variances are omitted for economy.

As expected, negative Social events experienced in the 4-week period uniquely predicted loss of SL, $p < .0001$. The unique contributions of negative Achievement and Other events, in contrast, were not significant, $p = .31$ and $p = .09$, respectively. Furthermore, Lagrange multiplier tests of imposed equality constraints confirmed

3. Obviously, not all instances of negative evaluation by others have primary significance for self-liking. Most of the "social" events reported, however, involved hostility, conflict, disapproval, or rejection that did not directly reflect on the respondent's competence.

4. Analyses using intensity-weighted frequencies produced results parallel to those reported.

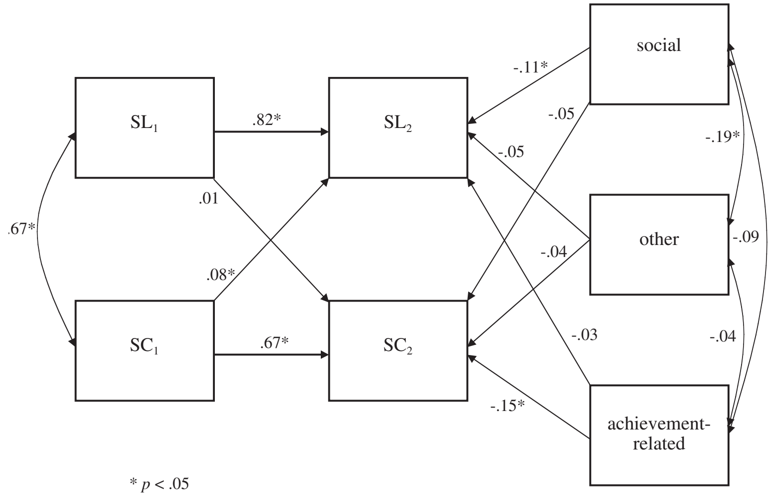

Figure 3
Path diagram for 4-week change in self-esteem as a function of negative life events. All path coefficients are standardized. Error variances are omitted.

that the unique association of Social events was more negative than those of Achievement and Other events, p = .02 and p = .03, respectively.

Also as expected, negative Achievement events uniquely predicted loss of SC, p = .001. The unique contributions of negative Social and Other events, in contrast, were not significant, p = .29 and p = .45, respectively. Furthermore, Lagrange multiplier tests confirmed that the unique association of Achievement events was more negative than those of Social and Other events, p = .02 and p = .03, respectively.

Two additional Lagrange multiplier tests confirmed that Social events were more strongly associated with loss of SL than loss of SC, p = .04, and Achievement events were more strongly associated with loss of SC than loss of SL, p = .002.

DISCUSSION

Self-competence, defined as the valuative derivative of personal agency, is assumed to correspond to one's history of success and failure at achieving goals. Self-liking, defined as the valuative representation of oneself as a moral-aesthetic social object, is assumed to derive from appraisals of worth conveyed by others or reflexively generated by the "generalized other" of self-judgment. Confirming this theoretical casting, negative achievement-related events (failure and frustration) were found to diminish self-competence, whereas negativity from others (rejection, disapproval, and interpersonal conflict) diminished self-liking. Thus, each dimension appeared to be sensitive to the specific type of adversity that matched its hypothetical antecedents. Investigation into the impact of positive events, not examined here, would be a useful extension of this line of research.

Although the results support the construct validity of the two dimensions of self-esteem, caution must be taken not to over-interpret the associations found. First, the critical partial relations were of modest magnitude. Second, the LER's respondent-centered approach is limited by its reliance on deliberate memory retrieval. Ideally, retrieval would not be systematically biased. Low self-esteem, however, is associated with negative memory bias (e.g., Story, 1998; Tafarodi, 1998). This raises the possibility that participants who experienced even modest loss of self-liking or self-competence over a 4-week period recalled more negative events than they would have in happier times. In this case, the critical associations of life events with change in self-esteem would be inflated. In the extreme, these associations might be due entirely to recall bias, effectively reversing the critical causal arrows in Figure 3. The specific pattern of associations, however, suggests that any retrieval bias of this sort would have to be quite selective, with loss of self-competence uniquely promoting recall of negative achievement-related but not social events and loss of self-liking uniquely promoting recall of negative social but not achievement-related events. Such specificity of bias would, ironically, support the construct validity of the two dimensions nearly as much as would self-esteem's being contingent on life events as assumed.

A limitation common to the preceding studies is their exclusive reliance on self-report measurement. Insofar as self-competence and self-liking are represented as distinct valuative concepts in the nexus of the self-concept, they should also be associated with involuntary, non-communicative patterns of behavior. Such implicit expression of self-esteem should reveal the same associative divergence as seen in the case of negative life events. Evidence of this sort would circumvent the hazards of self-report, including self-presentation and other forms of deliberate and non-deliberate distortion. Support for construct validity using implicit measurement would strengthen the case for two-dimensional self-esteem. We designed the final study with this in mind.

The personal relevance of valuative information should vary as a function of self-competence and self-liking. Those low in self-competence tend to be preoccupied with their perceived inability and lack of success. Therefore, they should be quick to discern information suggestive of failure or inefficacy (Bargh & Pratto, 1986; Higgins & King, 1981; Sanbonmatsu & Fazio, 1991; Sedikides & Skowronski, 1990). This translates into the prediction that those low in self-competence should be better than those high in self-competence at recognizing content related to weak agency. Such a difference represents a passive form of perceptual selectivity. The opposite prediction, however, cannot be made for content related to strong agency. Those with negative self-views often hold stringent self-ideals and experience intense dissatisfaction when falling short of these ideals (Higgins, Klein, & Strauman, 1987; Kuiper, Olinger, & MacDonald, 1988). Preoccupation with one's failings entails preoccupation with what one has failed to achieve, embody, or otherwise live up to. This suggests that conceptual nodes representing imperatives of success, achievement, and the realization of goals are at least as strongly associated in memory with the representation of low self-competence as with high self-competence, implying comparable accessibility (Teasdale, Taylor, Cooper, Hayhurst, & Paykel, 1995; see also Segal, Gemar, Truchon, Guirguis, & Horowitz, 1995). As such, there is little reason to expect superior recognition of information related to strong agency by those high in self-competence. Such information is equally relevant for those low and high on the dimension.

A parallel argument applies to self-liking. Those low on this dimension tend to be preoccupied with concerns about their social worth. Dominant themes include guilt over perceived transgressions, concerns about physical appearance, dissatisfaction with social identity, and fears of rejection or disapproval by others. For the reasons outlined above, the heightened personal relevance of information suggestive of "badness" or unworthiness should render it especially recognizable by those who lack self-liking. This translates into the prediction that those low in self-liking should be better than those high in self-liking at recognizing content related to low social worth. As before, however, content related to high social worth is expected to be as relevant to those low as those high in self-liking, implying similar recognition of it. The predicted valence asymmetry is consistent with research on social perception showing that negative instances of morally relevant behavior are perceived as more diagnostic of personality than are positive instances (Skowronski & Carlston, 1987). More broadly, negative information appears to receive greater attention and weight (Fiske, 1980; Peeters, 1971; Yzerbyt & Leyens, 1991).

To test these predictions, the ability to recognize common words reflecting agency and social worth was examined as a function of self-competence and self-liking.

STUDY 4

Overview

College students viewed a series of gradually unmasked trait words, including words representing low and high agency and social worth. The number of successive presentations required to correctly recognize each word was recorded. Students' self-competence and self-liking scores were used to predict ease of recognition within each semantic category. Perceptual selectivity uniquely attributable to each dimension of self-esteem was tested for fit with the personal relevance hypothesis.

METHOD

Participants

Participants were 126 introductory psychology students (87 women and 39 men) at the University of Toronto. Their modal age was 19.

Materials and Procedure

Subjects are tested individually on a computer using an adaptation of the gradual unmasking procedure developed by Macrae, Stangor, and Milne (1994). After a short practice phase, 48 personality trait words were presented in a randomized series. Each word was repeatedly presented at the center of the monitor screen. Each presentation lasted 200 ms with an interstimulus interval of 2 s. On initial presentation, the word was densely masked with dots such that it could not be identified. With each successive presentation, approximately 11% of the initial mask was diffusely removed, such that the word was completely unmasked by the 10th presentation. The participant was required to hit the space bar as soon as the word was recognized. At that point, the repeated presentations ended and the participant was required to write down the word on a form. The participant then pressed the return key to begin presentation of the next word.

The fewer successive presentations required to correctly recognize a word, holding general ability constant, the more accessible the word can be assumed to be in memory. The personal relevance hypothesis predicts a specific pattern of perceptual selectivity reflecting differential accessibility. Namely, accessibility of negative but not positive agency-related words is predicted to vary as a function of self-competence. Similarly, accessibility of negative but not positive worth-related words is predicted to vary as a function of self-liking. To test these predictions, five categories of words were used: High Competence (C+; e.g., competent, capable, effective), Low Competence (C−; e.g., weak, failure, defeated), High Social Worth (W+; e.g., attractive, worthy, likable), Low Social Worth (W−; e.g., inferior, despised, rejected), and Neutral (N; e.g., subtle, serious, talkative). The 16 N words were selected from the neutral range (neither positive nor negative in perceived meaning) in Anderson's (1968) normed list of trait adjectives and were indicative of neither competence nor social worth. These words were included to permit estimation of general recognition ability, an important individual difference to control for in refined tests of perceptual selectivity. The remaining four categories were represented by eight words each. These words had been confirmed through preliminary research to be highly indicative of their semantic category, as reflected in college students' judgments of their applicability to the experience of low and high self-competence and self-liking. There was no need to match categories on normative recognizability (word frequency, word length, etc.), as the predictions were associative rather than cross-categorical. To reduce structural similarity of words across valence categories, negated versions of the positive words were avoided in representing the negative categories (e.g., "incompetent" was not used with "competent," nor "unlikable" with "likable").

Finally, participants completed the SLCS at the end of the experimental session to provide measures of self-competence and self-liking in the nonselect sample.

RESULTS

The sample was screened for univariate and multivariate outliers. None were found. Gender did not qualify any of the results reported below and therefore will not be discussed further. Analysis of the written responses revealed that the words were correctly recognized 98.76% of the time. For each participant, the average number of presentations required to correctly recognize words was calculated within each semantic category. The sample means were 5.55 for C+, 5.60 for C−, 5.57 for W+, 5.92 for W−, and 5.86 for N. N word recognition was used solely as a covariate to control for general ability. Simultaneous multiple regression was used to examine perceptual selectivity uniquely attributable to each dimension of self-esteem. Four parallel models were tested, corresponding to the four semantic categories of interest. For all regression models tested, the SC × SL interaction and all quadratic terms were initially included as predictors. As none of these terms were close to significance, they were dropped from the models to preserve degrees of freedom and focus testing (Darlington, 1990). Results for the reduced models, with only SC, SL, and recognition of N words (covariate) as predictors, are reported below.

For C+ words, neither SC, $\beta = .00$, $t(122) = .06$, $p = .95$, nor SL, $\beta = -.06$, $t(122) = -1.20$, $p = .23$ emerged as significant predictors. For C− words, however, SC, $\beta = .16$, $t(122) = 2.86$, $p = .005$, but not SL, $\beta = -.01$, $t(122) = -.13$, $p = .90$, was uniquely associated. Additional testing revealed that the unique association of SC was significantly greater than that of SL, $p = .04$. The results can also be expressed in incremental terms, using sequential tests. Entering SC into a model that already included SL and N word recognition predicted an additional 2% ($p = .005$) of the variance in C− word recognition. In contrast, entering SL into a model that already included SC and N word recognition predicted only .009% ($p = .90$) more variance. The form of partial association reveals that those lower in SC (holding SL constant) were quicker to recognize C− words.

For W+ words, neither SC, $\beta = .06$, $t(122) = .93$, $p = .35$, nor SL, $\beta = .06$, $t(122) = .98$, $p = .33$, emerged as significant predictors. For W− words, however, SL, $\beta = .16$, $t(122) = 2.94$, $p = .004$, but not SC, $\beta = -.05$, $t(122) = -.93$, $p = .36$, was uniquely associated. Further testing revealed that the unique association of SL was significantly greater than that of SC, $p = .01$. Incrementally, entering SL into a model that already included SC and N word recognition predicted an additional 2% ($p = .004$) of the variance in W-word recognition. In contrast, entering SC into a model that already included SL and N word recognition predicted only .15% ($p = .36$) more variance. The form of partial association reveals that those lower in SL (holding SC constant) were quicker to recognize W− words.

Regression coefficients were also compared across models. Multivariate tests using Wilks's Λ confirmed that the SC coefficient obtained for C− words was greater than those obtained for C+ words, $p = .005$, W+ words, $p = .04$, and W− words, $p = .0003$. Similarly, the SL coefficient obtained for W− words was greater than those obtained for W+ words, albeit marginally so at $p =. 06$, C+ words, $p = .0005$, and C− words, $p = .004$.

The overall pattern of results is consistent with the prediction that SC and SL are uniquely associated with the ability to recognize negative but not positive words specifically relevant to each dimension.

DISCUSSION

Self-competence and self-liking are conceived as complementary representations of value in the structure of the self-concept. They are assumed to occupy the same global tier in the semantic network. The results of this study reveal a symmetrical pattern of semantic specificity in the relation of self-esteem with perceptual selectivity. Supporting the contention that information reflecting weak agency is of greater personal relevance to those who believe they lack competence, such individuals were quickest to recognize words representing this specific form of valuative deficit. Similarly, those with doubts about their social worth were quickest to recognize words representing this alternate form of deficit, again consistent with the claim that such information is especially relevant for them. The symmetry of this pattern fits with the casting of self-competence and self-liking as closely related but distinguishable dimensions of self-esteem, each with a predictable degree of uniqueness in its cognitive and motivational expression.

GENERAL DISCUSSION

We have argued in this paper for differentiating two correlated aspects of self-esteem. That human beings experience themselves as both willful agents and social objects gives rise to two forms of personal valuation: self-competence and self-liking. Measurement of self-esteem has often merged the two, as illustrated by the multidimensionality of Rosenberg's (1965) SES, which appears to include both assessment and acceptance factors. These scale-specific factors are highly redundant with self-competence and self-liking, and therefore can be interpreted as expressions of the more generalized dimensions.

A Higher Love?

The unique predictive efficiency of self-competence and self-liking was confirmed in Studies 3 and 4. This evidence, however, does not diminish the fact that the two dimensions are highly correlated as measured. Their considerable overlap is clearly reflected in the dominance of the "common factor" found in Studies 1 and 2. Does this commonality represent a third, higher-order dimension of global self-esteem? Perhaps. Alternatively, the high correlation may be due to self-report method factors, as suggested earlier. Furthermore, even if the correlation of self-competence and self-liking is substantive rather than artifactual, it is not clear that an additional construct is needed to account for it. Positing a superordinate construct to explain a high correlation is justified when: 1) the correlation cannot be more parsimoniously explained by unidirectional or bidirectional effects between the two subordinate constructs; and 2) the superordinate construct holds clear surplus meaning beyond the subordinate constructs. Neither justification applies well here, as the analogy of height and weight helps illustrate.

The intercorrelation of height and weight is at least as high as that between the two dimensions of self-esteem. It would seem odd to suggest that this is so because height and weight are both influenced by a third, "size" variable. Rather, the association is more simply explained by the fact that taller people tend to have larger bodies than shorter people and therefore weigh more. Hence, the additional variable appears to be unnecessary. This does not, of course, preclude some composite of height and weight from proving useful beyond the separate constructs in the context of prediction. Such surplus meaning, however, is captured in the height $\times$ weight interaction. There is no theoretical justification for interpreting the interaction as anything more than just that—an interaction between two highest-order constructs.

Arguably, the case of self-competence and self-liking is similar. Their high correlation is consistent with the reciprocal determination that is hypothesized to bind them through development (see Tafarodi

& Swann, 2001). These paths of influence are enough to explain any overlap of self-competence and self-liking beyond artifactual covariance. As in the case of height and weight, an additional concept may be unnecessary.

Neither is it clear that a third, more global dimension of self-esteem provides surplus meaning. Just as height and weight jointly constitute what we think of as body "size," self-competence and self-liking can be thought of as jointly constituting "general" self-esteem. In both cases, the higher-order concept holds no independent referential meaning; it serves only to represent a correlated composite of dimensions. Such umbrella concepts are useful as expedients in discourse but not as self-standing theoretical constructs. On the other hand, future research may reveal that superordinate, general self-esteem does in fact represent an aspect of feeling or belief that is not captured by self-competence, self-liking, and their interaction. Until then, however, the more parsimonious interpretation is justified.

Self-Competence and Self-Efficacy

Bandura (1989, 1992) has defined *self-efficacy* as "people's beliefs about their capabilities to exercise control over events that control their lives" (1989, p. 1175). In its most generalized form, this refers to the overall assurance or faith that individuals have in their ability to achieve their goals (Sherer et al., 1982; Tipton & Worthington, 1984; Woodruff & Cashman, 1993). The correspondence of this trait-like conception of general self-efficacy with self-competence invites discussion of just how the two constructs are related.

Bandura (1990) has argued that self-efficacy is separate from self-esteem. This is clearly the case for task-specific self-efficacy. The conceptual separation also applies to general self-efficacy, but its representation in experience at this higher level may be more difficult to discern. We should recognize that human development is characterized as much by the need to know "who we are" as "what we can do." Accordingly, one's personal history of success and failure inevitably gives rise to a generalized attitude toward the self as agent. The more successful one has been in fulfilling the myriad conscious intentions that constitute a history of action, the stronger one feels. As an aspect of personal identity, this strength is experienced as positive value, irrespective of any secondary, moral meaning that attaches to it. This is so because the value of successful action is twofold, one aspect being the primitive and immediate pleasure of "effectance" (White, 1963), and the other the moral interpretation of the success. The two aspects need not be consistent. All of this suggests that general self-efficacy, defined as a *global expectancy*, and self-competence, defined as a global dimension of *self-value*, are but two consequences of the same cumulative process. Namely, self-competence is the valuative imprint of general self-efficacy on identity. Consistent with this unifying interpretation, examinations of the discriminant validity of global self-esteem and general self-efficacy have failed to clearly distinguish the two constructs (Bernard, Hutchison, Lavin, & Pennington, 1996; Stanley & Murphy, 1997).

Conclusion

We suggest that the diffraction of global self-esteem into self-competence and self-liking helps explain the conceptual differences that continue to impede research in this area. The wide range in viewpoints noted by critics of the literature is due in part to the tendency of some to focus on one or the other dimension. For example, William James's (1890/1950) oft-cited definition of self-esteem as the ratio of successes to pretensions captures the essence of self-competence but not self-liking. In contrast, Carl Rogers' (1961) equation of self-esteem with authenticity and self-acceptance is more relevant to self-liking than to self-competence. More recently, Baumeister (1997), while accepting that there are "two main sources of self-esteem" (p. 688), focuses on self-liking in discussing the determination of self-esteem (e.g., Baumeister, Dori, & Hastings, 1998). The same emphasis appears in the work of Leary and his colleagues on self-esteem as a "sociometer" (e.g., Leary & Downs, 1995). Finally, Seligman's (1995) extolling of optimism over self-esteem stems mainly from his doubts about the adaptive value of unwarranted self-liking, not self-competence. Similarly focused treatments appear throughout the literature, and have been the source of much dispute over the years. A balanced understanding of the nature and importance of both dimensions would go far in reconciling seemingly discrepant perspectives and preventing future inconsistencies. The result promises to be a clearer account of what is arguably our most important attitude.

Appendix

SES Items (Rosenberg, 1965, 1979).

1. On the whole, I am satisfied with myself. (Ac)
2. At times I think I am no good at all. (Ac)
3. I feel that I have a number of good qualities. (As)
4. I am able to do things as well as most other people. (As)
5. I feel I do not have much to be proud of. (As)
6. I certainly feel useless at times. (Ac)
7. I feel that I'm a person of worth, at least on an equal plane with others. (As)
8. I wish I could have more respect for myself. (Ac)
9. All in all, I am inclined to feel that I am a failure. (As)
10. I take a positive attitude toward myself. (Ac)

SLCS Items (Tafarodi & Swann, 1995).

1. Owing to my capabilities, I have much potential. (C)
2. I feel comfortable about myself. (L)
3. I don't succeed at much. (C)
4. I have done well in life so far. (C)
5. I perform very well at a number of things. (C)
6. It is often unpleasant for me to think about myself. (L)
7. I tend to devalue myself. (L)
8. I focus on my strengths. (L)
9. I feel worthless at times. (L)
10. I am a capable person. (C)
11. I do not have much to be proud of. (C)
12. I'm secure in my sense of self-worth. (L)
13. I like myself. (L)
14. I do not have enough respect for myself. (L)
15. I am talented. (C)
16. I feel good about who I am. (L)
17. I am not very competent. (C)
18. I have a negative attitude toward myself. (L)
19. I deal poorly with challenges. (C)
20. I perform inadequately in many important situations. (C)

Note. Ac= Acceptance; As= Assessment; C= Self-Competence; L= Self-Liking.

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