How Not to Measure Intergenerational Occupational Persistence[1]

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This article is an invited critique of Rytina's proposal that canonically weighted scales for detailed occupations be used to assess intergenerational occupational persistence. These canonical scales yield higher intergenerational correlations than the Duncan SEI primarily because they capitalize on chance fluctuations in the data used to construct them. Thus, there is no reliable basis for revising previous findings about the relative importance of ascription and achievement in the stratification process. Canonical scaling (and related log-linear and log-multiplicative models) are no more than a useful adjunct to models that locate the sources of intergenerational occupational persistence in the measurable characteristics of occupations and persons.

In "Scaling the Intergenerational Continuity of Occupation," Steve Rytina (1992, in this vol.) attempts to improve the measurement of occupational social standing and to revise prior findings about the persistence of occupational standing across generations by specifying the scale of occupational standing with a canonical weighting of detailed occupational categories. As he notes, Klatzky and Hodge (1971) carried out a canonical analysis of Blau and Duncan's (1967) 17 × 17 classification of

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occupational mobility among U.S. men in 1962;[2] they noted the similarity between socioeconomic characteristics of occupations and their estimates of canonical scores for the 17 occupational categories. Rytina confirms this key finding of Klatzky and Hodge by analyzing mobility among 308 detailed occupational categories in the cumulative General Social Survey (GSS), 1972–86. He also proposes a new scale of occupational status, the symmetric scaling of intergenerational continuity (SSIC), which is based on constrained canonical weights for detailed occupations.

Rytina's analysis yields three striking empirical findings. First, the observed correlation between offspring's and father's status increases from .256 using the NORC (Siegel 1971) prestige scores and .338 using the SEI (Duncan 1961) scale to .450 with the author's preferred SSIC scale, in which the canonical weights (scale values of occupations) are constrained to be equal in the generations of parents and offspring. Rytina says that the SSIC corrects the SEI and, we assume, prestige scores, by improving the correspondence of occupational ranks with "the details of ascent and descent." Thus, he finds that "continuity of occupation" is understated when traditional scales are used.

Second, aside from the larger estimate of the intergenerational correlation, Rytina finds that educational attainment is much less important in mediating the persistence of occupational social standing from one generation to the next. This reduction occurs through an increasing dependence of schooling on father's occupational standing and a roughly compensating decrease in the influence of schooling on offspring's occupational standing, along with the larger overall correlation between father's and offspring's occupational standing. That is, the indirect effect of father's occupational standing is roughly constant, but the direct effect increases. This, he says, contradicts widely accepted findings by showing that ascription is more important than achievement in explaining intergenerational occupational persistence.

Third, contrary to Hout's (1988) analyses of (nearly) the same data, Rytina finds that the dependence of offspring's occupational social standing on parental occupational standing is maintained among college graduates. Hout (1988, p. 1391) had argued that, for college graduates, "current occupational status is independent of origin status. This finding provides a new answer to the old question about education's overcoming disadvantaged origins. A college degree can do it." Rytina finds that origins have almost no effect on destinations for college graduates when

[2] Rytina does not mention two applications of canonical analysis to the estimation of trend in intergenerational occupational mobility (Hauser and Featherman 1977, pp. 169–87; Featherman and Hauser 1978, pp. 63–138).

occupations are indexed by the SEI, but that there is a strong effect if occupations are indexed instead by the SSIC. The correlation between father's occupational standing and offspring's occupational standing among college graduates is only .0687 when the SEI is used, but it is .2844 when the SSIC is used.

Finally, and somewhat independently of his empirical findings, Rytina argues for the theoretical superiority of his canonical scale in assessments of the intergenerational "continuity of occupation." He argues that the SSIC is ideal for assessing whether ascription or achievement is more important.

In our discussion, we use "occupational standing" as a generic term covering prestige scales, socioeconomic indexes, and Rytina's canonical scores for occupations. We agree with Jencks's (1990) emphasis on the substantive difference between occupational prestige and occupational socioeconomic status and, thus, on the different interpretations of their intergenerational correlations. Occupational prestige scores measure prestige, and their validity does not stand or fall on the size of intergenerational prestige correlations. However, in this discussion, we ignore differences between prestige scales and socioeconomic indexes or canonical scores and focus on those between socioeconomic indexes and canonical scores. Also, we have restricted our discussion to intergenerational correlations of occupational characteristics and ignored correlations of other variables. We think it entirely possible that intergenerational correlations of other social and economic characteristics or combinations of such characteristics, including occupations at different stages of the life course, may be higher or lower than the correlations of characteristics of occupations considered by Rytina. Finally, other measures of the influence of family origins on social and economic success, like the similarity of siblings, may suggest greater or lesser levels of intergenerational persistence. In short, we do not want our discussion of intergenerational occupational correlations to be read as more general remarks about the process of social stratification.

WHAT'S WRONG WITH THE SSIC?

Our critique distinguishes Rytina's findings from his conclusions. In our opinion, the findings are correct and the conclusions are all wrong. Rytina obtains large intergenerational correlations in the 1972–86 GSS because the canonical scores depend on a very large number of parameters. Analytically, this alone accounts for almost all of the difference between Rytina's intergenerational correlations and those based on the SEI. We also used Rytina's scores to cross-validate Rytina's empirical findings in

one or more of three bodies of data, the 1987–90 GSS (Davis and Smith 1990), the 1962 Occupational Changes in a Generation Survey (OCG), and the 1973 OCG (Featherman and Hauser 1983). The cross-validations do not support Rytina's findings: (1) intergenerational correlations of occupational standing are about the same or smaller when Rytina's scales are used as when the Duncan SEI is used; (2) the direct intergenerational effect of father's occupational standing—the effect that is not mediated by educational attainment—is smaller than the share estimated by Rytina in the 1972–86 GSS; (3) within levels of educational attainment, intergenerational correlations of occupational standing are about the same when the SSIC is used as when the Duncan SEI is used, that is, they are negligible. Since we cannot confirm Rytina's findings in other samples, we cannot agree with his conclusion that the empirical evidence supports a preference for the SSIC over the Duncan SEI.

There are obvious practical limits to the value of the SSIC in research. One of the advantages of other widely used scales is that the relevance and validity of the criteria on which they are based, for example, popular ratings of social standing or occupational education and income, are widely acknowledged. We doubt that there will be similar demand for an occupational scale that maximizes the intergenerational correlation between occupations. This is important because there is scientific value in using the same occupational scale in different fields and disciplines, for example, in studies of political behavior, educational performance, and mortality, as well as in studies of social mobility.

Also, despite Rytina's efforts to enlarge the number of occupation lines in his analysis, the SSIC ignores about one-quarter of the census occupation titles and about one-tenth of the work force. We do not believe that it would be acceptable to the research community to censor a substantial number of occupations or workers in this way. For example, Duncan's (1961, pp. 110–13) SEI was developed and accepted in large part because, for the first time, it provided an index of the status of *all* U.S. occupations; it took another decade to develop a prestige scale for all occupations (Siegel 1971). It would be possible to cover all occupations with an index constructed in the same fashion as the SSIC by adopting an intermediate-level aggregation of occupation lines rather than using all detailed lines. However, this could only be carried out with a corresponding loss of occupational detail.

We have also assessed the conceptual basis for Rytina's claim that the SSIC (or a similar scale) provides a superior framework for the measurement of intergenerational occupational persistence, and we find none. While we think there is a place in mobility research for canonical correlation, association models, and related criterion scoring methods, we do not believe there is any reason to replace scales based on measured occu-

pational characteristics with the SSIC (or a similar scale).[3] Even if the empirical evidence showed that correlations based on the SSIC (or a similar scale) were consistently larger than those based on conventional scales, we would not accept the SSIC as a new basis for mobility studies. On the contrary, we think that there should be more basic research on measurable dimensions of occupations that persist across generations.

CORRECTING THE CANONICAL CORRELATIONS

Rytina acknowledges that his analysis is based on a single sample and that replication in other samples is desirable. He also acknowledges the possibility that the scale values he obtained may be affected by overfitting chance fluctuations in the sample data. Unfortunately, rather than carrying out a true cross-validation, he relies upon various tests internal to the 1972–86 GSS sample, for example, a comparison of the intergenerational SSIC correlation between populous and sparse occupational lines.

The key problem is that Rytina ignores the sampling variability of his intergenerational correlations.[4] When the SSIC is estimated, the intergenerational correlations increase (relative to conventional scales or indexes) for two reasons: first, the SSIC may reflect dimensions of occupations that persist, but are not correlated with the conventional measures, and, second, the SSIC capitalizes on chance associations between occupations in the original sample.[5] In effect, when he estimates the canonical correlation between father's occupation and offspring's occupation, Rytina is computing a multiple correlation based on 614 independent variables. When he constrains the scales to be the same for offspring as for fathers, he is still estimating an equation with 307 independent variables. Even in a sample of nearly 8,000 cases, the effect of this profligate use of

[3] However, association models, which have been extended to multidimensional classifications, provide a superior methodological basis for empirically based rankings of occupational categories, comparisons of those rankings across populations, and comparisons of the openness of mobility regimes. Goodman (1984) provides the key exposition of these models and methods, including a comparative analysis of association models and canonical correlation models.

[4] In the one place (1992, table 6) where Rytina reports standard errors of correlations based on the SSIC, they are apparently based on the naive and erroneous assumption that the SSIC scores were not estimated in the same data used to estimate the correlations.

[5] There is, or was, a similar problem in the initial construction of the Duncan SEI, which is a weighted average of aggregate measures of the educational attainment and income of men in the 1950 census. This gave rise to criticisms that the correlations between the SEI and education or income were built in, which is true, and that they were artifacts, which is not true (Blau and Duncan 1967, pp. 117–28). In any event, all of the analyses reported herein are far removed from the data that were used to construct the SEI.

degrees of freedom cannot be ignored; the canonical correlation is much larger in the sample than we would expect it to be in the population.

Rytina estimates a canonical correlation of 0.489 between father's occupation and offspring's occupation in 7,965 cases from the 1972–86 GSS. Using one of Lawley's (1959) formulas, we estimate a corrected correlation of 0.411 when the sample canonical correlation is 0.489 in a simple random sample of 7,965. But the GSS sample is actually much less efficient than a simple random sample; it is approximately two-thirds as efficient (Davis and Smith 1991). Thus, the problem of statistical unreliability is greater than would be suggested by the size of the GSS sample.[6] If we apply Lawley's formula again, assuming a simple random sample two-thirds the size of the GSS, we estimate the corrected correlation, 0.337. This is virtually the same as the observed intergenerational SEI correlation, 0.338.[7]

Thus, the "dramatic" difference between intergenerational SEI correlations and intergenerational canonical correlations is just about what we should expect if the true intergenerational correlation were that observed for the SEI, given only the size and design of the GSS sample and the number of parameters to be estimated. Rytina's findings of greater ascription are essentially an artifact of his failure to take account of the sample-population distinction.

INTERGENERATIONAL CORRELATIONS OF OCCUPATIONAL STANDING

In our cross-validations, we compare the behavior from sample to sample of intergenerational correlations based on Rytina's canonical scales with that of correlations based on the NORC prestige scale and the Duncan SEI. The first column of table 1 reproduces the correlations from which Rytina concludes that the NORC prestige scale and the SEI are less highly correlated across generations than the rough and ready (hereafter referred to as R&R), the SSIC, or the canonical scales in which father's

[6] This problem also may have a significant effect on Rytina's identification of nonlinearities in correlations based on the SEI. He identifies 47 occupation lines for which nonlinearities are nominally significant with $P < .05$, but if the GSS is two-thirds as efficient as a simple random sample of the same size, the true probability level exceeds .10. We think this is a loose criterion for the identification of outliers.

[7] We are grateful to Steve Rytina for correcting an error in our application of a related formula in Lawley (1959). Incidentally, it is unnecessary to correct SEI correlations for loss of degrees of freedom, as we have corrected Rytina's canonical correlations, because the SEI uses only one degree of freedom.

TABLE 1

INTERGENERATIONAL CORRELATIONS OF OCCUPATIONAL STANDING

	General Social Survey					Occupational Changes in a Generation, 1973			
				20 Outliers Deleted				20 Outliers Deleted	
	308 Occ, 1972–86	306 Occ, 1972–86	306 Occ, 1987–90	1972–86	1987–90	306 Occ, Current	306 Occ, First	Current	First
Prestige	.256	.256	.261	.217	.227	.263	.294	.218	.266
SEI	.338	.348	.301	.319	.274	.375	.423	.343	.365
R&R	.427	.430	.267	.390	.249	. . .	. . .	. . .	. . .
SSIC	.450	.453	.267	.412	.261	.376	.474	.349	.369
Canonical	.489	.493	.220	.454	.208	.320	.386	.294	.298
Ratios of correlations:									
R&R/SEI	1.26	1.24	.89	1.22	.91	. . .	. . .	. . .	. . .
SSIC/SEI	1.33	1.30	.89	1.29	.95	1.00	1.12	1.02	1.01
Canonical/SEI	1.45	1.42	.73	1.42	.76	.85	.91	.86	.82
N	7,965	7,777	2,602	7,355	2,455	23,016	21,202	20,559	17,694

NOTE.—All findings except those in the first column or those including NORC prestige scores exclude military occupations; in the 1972–86 GSS, there are 7,777 cases after fathers or offspring with military occupations have been excluded. Occ = occupations.

and offspring's occupations are permitted to have distinct scores.[8] In the second column of table 1, we show similar correlations, based upon a slightly different use of the 1972–86 GSS sample. First, we have deleted the two military lines. There is no reason to believe that the SEI scores for those lines accurately represent current members of the armed forces; the data used by Duncan (1961, pp. 263–75) pertained to former members of the armed forces who were unemployed at the time of the 1950 census. Also, there are no directly measured NORC prestige scores for members of the armed forces, and it is conventional to limit analyses of social mobility to the civilian noninstitutional population. Second, we have corrected the NORC prestige score for carpenter's helpers (750) from "0," as listed incorrectly in the source used by Rytina, to the correct value of "23." These changes barely affect the correlations.

In the last three rows of table 1, we report the ratios of correlations based on Rytina's three scales to the correlations based on the SEI. We believe these ratios provide an appropriate basis for comparing the performance of the four scales. We see no reason to dwell on the lower intergenerational correlations of prestige scores (Jencks 1990). The comparison with correlations based on the SEI takes account of the variation from sample to sample in the intergenerational correlation of occupational socioeconomic status. Both in the first and second columns of table 1, each of Rytina's scales has a substantially higher intergenerational correlation than the SEI, and the differential increases from the scale that is partly dependent on the SEI—the R&R score—to the unconstrained canonical scoring. Correlations based on Rytina's scales are 25%–45% larger than those based on the SEI.

In the third column of table 1, we report intergenerational correlations in an independent sample, the 1987–90 GSS. The population definition and codings of all variables are identical to those in the 1972–86 GSS; only the years and the random sampling outcomes differ. The findings are very different in this sample. While the prestige correlations are virtually identical, the other three correlations are smaller. The SEI correlation declines by 13.5%, from .348 to .301, suggesting that there has been a decline in occupational status persistence in the United States. However, the R&R correlation declines by 37.9%, from .430 to .267, and the SSIC correlation declines by 41.1%, from .453 to .267, so it is smaller than the SEI correlation and about as large as the NORC prestige correlation. The correlation of canonical scores declines even more, by 55.4%, from .493 to .220. Rytina's three scales yield correlations that are, respectively,

[8] Rytina did not give us his R&R scale values, and we reconstructed them from the description in his text. We have reproduced the findings in his table 1 using our version of the R&R scale, and we are satisfied that it is the same scale he constructed.

89%, 89%, and 73% as large as those based on the SEI in the 1987–90 GSS.

Notably, the correlation based on the R&R scale shrinks almost as much as that based on the SSIC. This is significant because the R&R scale is the result of a single iteration of one method for estimating a symmetrically scaled canonical correlation.[9] It is not overfitting in the later stages of estimation that accounts for sample-to-sample shrinkage of the intergenerational correlations, but a failure to replicate the most obvious nonlinearities in the intergenerational relationship of SEI scores. That is, Rytina claims "superior validity" for the SSIC because intergenerational correlations increase when he smooths out nonlinearities in the relationship between father's and offspring's SEI; the claim fails because the nonlinearities are not replicable. When we use Rytina's scales in the 1987–90 GSS, there is no longer any indication that they yield larger intergenerational correlations than the SEI.[10]

To be sure, one could obtain higher intergenerational correlations in the 1987–90 GSS sample by estimating another set of constrained and unconstrained canonical scores, but these correlations would be subject to exactly the same problem of overfitting as those in the 1972–86 sample. Furthermore, just as the SSIC correlations shrink in the 1987–90 GSS sample, we expect that correlations based on canonical scores in the 1987–90 GSS will shrink in other samples. The problem is that we want a measure of occupational standing that can be used to assess differences in intergenerational correlations from one population to another, but the SSIC cannot help us to do this as long as one of the samples of interest is the one from which it was estimated. And outside that sample, the SSIC loses its seeming advantage.[11]

We agree with Rytina that statistical outliers—extreme values in the distribution of the SSIC—may be responsible for the unreliability of correlations based upon it. For this reason, we repeated our analysis of the 1972–86 and the 1987–90 GSS samples, after we eliminated 20 occupation lines falling more than 2.5 standard deviations (SDs) from the mean of the SSIC. As shown in column 4 of table 1, these exclusions have little effect on the findings in the 1972–86 data. However, the 1987–90 GSS data again show no change in the NORC prestige correla-

[9] See Rytina (in this vol.), n. 14.

[10] We have also cross-validated Rytina's scales separately for men and women in the 1972–86 and in the 1987–90 GSS, and we have looked at the behavior of Rytina's scales in relation to that of more recent versions of the Duncan SEI (Stevens and Featherman 1981). All of these analyses yield essentially the same findings as those reported herein.

[11] The proper way to deal with this problem would be to test for changes between populations in canonical scores and in correlations based upon them.

tion, a small decline in the SEI correlation, and very large declines in the correlations based on the SSIC, R&R, and canonical scores. Evidently, the problem of sample-to-sample shrinkage in correlations based on Rytina's scales cannot be attributed to this small set of extreme values on the SSIC.

In the last four columns of table 1, we report cross-validations of intergenerational correlations of Rytina's SSIC and canonical scales in the 1973 Occupational Changes in a Generation Survey (Featherman and Hauser 1978). The data pertain to intergenerational correlations of father's occupation with son's current occupation and with son's first occupation among men 25–64 years old who were in the experienced civilian labor force in March 1973. Again, the analysis is restricted to the 306 civilian occupation lines that were included in Rytina's analysis. Here, the findings are more favorable to the SSIC than in the 1987–90 GSS sample. The intergenerational correlations of SEI and SSIC scores are almost the same, except the SSIC correlation is slightly larger (0.474 vs. 0.423) in the case of father's occupation and son's first occupation; this differential disappears when the 20 outliers are removed. The intergenerational correlations of the canonical scales are consistently lower than those of either the SSIC or the SEI; we think this is clear evidence of shrinkage due to overfitting in the 1972–86 GSS data. As is well known, the intergenerational correlations of the NORC prestige scores are lower than those of the SEI and, consequently, also lower than those of Rytina's scales. Finally, deletion of the 20 outliers lowers all of the correlations slightly, but—excepting the negligible difference between SEI and SSIC correlations—the general pattern of findings remains the same. From this second cross-validation, we again find no reason to conclude that correlations based on the SSIC will be larger than those based on the SEI, outside of the sample in which the SSIC was estimated.

EDUCATIONAL ATTAINMENT AS AN INTERVENING VARIABLE

Rytina argues that his scales yield larger estimates of the direct effect of father's occupational standing on offspring's occupational standing using a simple three-variable causal model: father's occupational standing affects educational attainment, and both of those variables affect offspring's occupational standing. As shown in the first two columns of table 2, Rytina's scales lead to estimates of the direct intergenerational effect that are 1.9–2.7 times larger than estimates based on the SEI. Table 3 contains evidence that his scales also lead to much larger estimates of the share of the intergenerational effect that is direct: 53.4% for R&R, 58.8% for SSIC, and 67.3% for the canonical scores, compared to only 36.3% for the SEI.

TABLE 2

DIRECT EFFECTS OF FATHER'S OCCUPATIONAL STANDING ON OFFSPRING'S OCCUPATIONAL STANDING

	General Social Survey					Occupational Changes in a Generation, 1973			
				20 Outliers Deleted				20 Outliers Deleted	
	308 Occ, 1972–86	306 Occ, 1972–86	306 Occ, 1987–90	1972–86	1987–90	306 Occ, Current	306 Occ, First	Current	First
Prestige	.080	.080	.091	.068	.084	.097	.118	.094	.121
SEI	.123	.124	.110	.113	.104	.149	.181	.140	.149
R&R	.228	.231	.092	.207	.101	. . .	. . .	. . .	. . .
SSIC	.265	.269	.097	.236	.115	.174	.265	.169	.187
Canonical	.329	.333	.080	.304	.093	.146	.195	.144	.148
Ratios of effects:									
R&R/SEI	1.86	1.87	.84	1.83	.97	. . .	. . .	. . .	. . .
SSIC/SEI	2.16	2.17	.88	2.09	1.10	1.17	1.46	1.21	1.25
Canonical/SEI	2.68	2.69	.72	2.69	.89	.98	1.07	1.03	1.00
N	7,965	7,777	2,602	7,355	2,455	23,016	21,202	20,559	17,694

NOTE.—All findings except those in the first column or those including NORC prestige scores exclude military occupations; in the 1972–86 GSS, there are 7,777 cases after fathers or offspring with military occupations have been excluded.

TABLE 3

PERCENTAGE OF THE CORRELATION BETWEEN FATHER'S AND OFFSPRING'S OCCUPATIONAL STANDING THAT IS NOT MEDIATED BY EDUCATIONAL ATTAINMENT

	General Social Survey					Occupational Changes in a Generation, 1973			
				20 Outliers Deleted				20 Outliers Deleted	
	308 Occ, 1972–86	306 Occ, 1972–86	306 Occ, 1987–90	1972–86	1987–90	306 Occ, Current	306 Occ, First	Current	First
Prestige	31.4	31.3	34.8	31.2	36.9	37.1	40.1	42.9	45.4
SEI	36.3	35.6	36.5	35.4	37.9	39.7	42.8	40.7	40.8
R&R	53.4	53.7	34.6	53.1	40.7	. . .	. . .	. . .	. . .
SSIC	58.8	59.3	36.4	57.4	44.0	46.2	55.9	48.3	50.7
Canonical	67.3	67.7	36.3	67.0	44.6	45.7	50.4	48.9	49.8
Ratios of percentages:									
R&R/SEI	1.47	1.51	.95	1.50	1.07	. . .	. . .	. . .	. . .
SSIC/SEI	1.62	1.67	1.00	1.62	1.16	1.16	1.31	1.19	1.24
Canonical/SEI	1.85	1.90	1.00	1.89	1.18	1.15	1.18	1.20	1.22
N	7,965	7,777	2,602	7,355	2,455	23,016	21,202	20,559	17,694

NOTE.—All findings except those in the first column or those including NORC prestige scores exclude military occupations; in the 1972–86 GSS there are 7,777 cases after fathers or offspring with military occupations have been excluded.

In our opinion, Rytina invites the reader to assess the role of education as an intervening variable in the stratification process by reference *only* to the effects of father's occupational status in the 1972–86 GSS, and he emphasizes the revisionist implications of his findings in those data. But his findings are not a great deal different from those of Blau and Duncan (1967, pp. 169–70) that are based on the Duncan SEI. They estimated a correlation of .405 between father's SEI and son's current SEI and a correlation of .417 between father's SEI and the SEI of the son's first job. These two estimates—in an earlier sample, to be sure—are not much smaller than those estimated by Rytina using the R&R and SSIC in the 1972–86 GSS. Using the original Blau-Duncan data, and applying Rytina's three-variable model, 44.0% of the intergenerational effect on son's current SEI is direct and 53.8% of the intergenerational effect on the SEI of son's first occupation is direct. Both of these are substantially larger than Rytina's estimate of 36.3% for SEI of current occupation in the 1972–86 GSS data, and the latter, at least, is comparable to Rytina's estimates based on the R&R and SSIC.

Where, then, is the disagreement between Rytina's finding and the conventional view? The problem is that Rytina looks only at the effect of father's occupational standing on offspring's occupational standing and not, more broadly, at the effect of social background on occupational standing. Even by the standards of a generation ago, Rytina's three-variable model cannot be regarded as a plausible or complete representation of the stratification process. For example, the original Blau-Duncan model included father's education as well as father's occupational standing among the exogenous variables. In the original Blau-Duncan model, 55.2% of the effect of father's occupational SEI on son's current SEI is direct, and 64.6% of the effect of father's occupational SEI on the SEI of son's first occupation is direct.[12] However, in that same model, *none* of the effect of father's education on the SEI of current or of first occupation is direct; all of the effect of paternal education on son's occupational attainment is mediated by schooling. Subsequent models of status attainment typically include other social background variables, like number of siblings, farm origin, southern birth, race, and intact family. The general finding that educational attainment is a key intervening variable in the stratification process is not based merely upon relations between the occupational ranks of parents and offspring, but upon a more extensive set of findings about the effects of several social background characteristics. For example, in Featherman and Hauser's (1978, p. 274) regressions of the SEI of first job on social and economic background characteristics among birth cohorts reaching age 16 between 1925 and 1965, the direct

[12] These are percentages of reduced-form coefficients, not of correlations.

effect of father's occupational SEI averages 42% of its total effect, while the direct effect of number of siblings averages 34.5% of its total effect; the direct effect of father's education averages 3.5% of its total effect; and the direct effect of broken family averages 5.6% of its total effect.[13] That is, there is no substantial basis for the claim that education mediates more than about half of the effect of father's occupational standing on offspring's occupational standing, but there is a good deal of evidence of a stronger mediating role in the case of other background variables.

Despite the weak rationale for Rytina's analysis of direct and indirect effects, we have cross-validated it in the 1987–90 GSS and in the 1973 OCG surveys. The third column of table 2 gives our estimates of the direct intergenerational effects of father's occupational standing that are obtained when Rytina's three-variable model is estimated in the 1987–90 GSS. Again, Rytina's major finding disappears in the 1987–90 data. The direct effects of the R&R, SSIC, and canonical scales are somewhat smaller than those of the SEI and similar to that of the prestige scale. The direct effect of the SEI is almost 90% as large in 1987–90 as in 1972–86, but the direct effect of the SSIC is only 36% as large in 1987–90 as in 1972–86. As shown in table 3, in the 1987–90 GSS, the direct effects of father's occupational standing account for about one-third of its correlation with offspring's standing, regardless of the scale that is used. The fourth and fifth columns of tables 2 and 3 repeat the comparison between the 1972–86 and 1987–90 GSS surveys, but with 20 outlying occupations removed; again, Rytina's findings almost disappear in the 1987–90 GSS. In the last four columns of tables 2 and 3, we report cross-validations in analyses of mobility to current occupations and to first occupations in the 1973 OCG survey. Here, the SSIC, but not the canonical scale, yields larger direct effects than the SEI. We do not know why the SSIC has larger direct effects in the OCG data; it may be the closer temporal proximity of the 1973 OCG to the period covered by the 1972–86 GSS.[14] Still, taking into account the findings in both the GSS and the 1973 OCG, we are not convinced that analyses based on the SSIC will yield evidence of greater direct ascription than those based on the SEI.

ASCRIPTION WITHIN LEVELS OF SCHOOLING

In addition to estimating the direct effect of father's occupational standing in the general population, Rytina looks at the effect within six specific levels of completed schooling: 0–7 years, 8 years, 9–11 years, 12 years,

[13] See n. 11 above.

[14] See the discussion of table 5 below.

13–15 years, and 16 or more years. His analysis addresses and claims to disconfirm Hout's (1988) finding that the effect of father's occupational status on offspring's status is negligible among college graduates.[15] Our slightly modified version of Rytina's analysis, shown in the first of each pair of columns in table 4, reproduces and extends his findings. In the 1972–86 GSS, within each level of completed schooling, intergenerational correlations of occupational standing based on the SSIC and canonical scales greatly exceed those based on the SEI or prestige scales. The disparity in findings is indeed "dramatic" among college graduates, the locus of Hout's analysis. The intergenerational correlation of the SEI is only .072, while those of the SSIC and canonical scales are .288 and .359.

The second of each pair of columns in table 4 reports corresponding correlations estimated in the 1987—90 GSS. In every case the decline in correlations based on Rytina's scales is as dramatic as the previously noted disparity in findings within the 1972–86 GSS. For example, among college graduates, the intergenerational correlation of the SSIC falls from .288 to .097, while that of the SEI falls from .072 to .050. We think that both correlations in the 1987–90 data are negligible by any reasonable standard. Overall, while there is variation in the correlations from one level of schooling to the next, we see no evidence that any of those correlations are more than modest in size or that those based on Rytina's canonical scales are larger than those based on the SEI or NORC prestige scales.[16] As in the case of his first two findings, Rytina's third major empirical finding fails the test of cross-validation. There is no reason to conclude from Rytina's analysis that SEI correlations underestimate intergenerational status ascription.

THE USE AND ABUSE OF CANONICAL CORRELATIONS

Apart from the empirical issue of cross-validation, there is an important question about the meaning of the SSIC and the intention of Rytina's proposal. Would use of the SSIC be justified if there were no evident

[15] The SEs that Rytina reports in his table 6 for the SSIC are incorrect. He apparently has tried to use a standard formula for the standard error of a zero-order correlation, but this standard formula does not take into account the degrees of freedom lost in estimation of the SSIC. That is, Rytina underestimates the sampling variability of correlations based on the SSIC in the 1972–86 GSS data. Thus, Rytina's reports of t-statistics and probability values associated with the SSIC are also incorrect. Also, Rytina reports 1,847 as the sample count in two education subgroups, leading to an inconsistency between the counts in the table and the total number of observations in the sample.

[16] We have repeated these analyses after removing 20 extreme observations on the SSIC, and the findings are the same.

TABLE 4

INTERGENERATIONAL CORRELATIONS OF OCCUPATIONAL STANDING BY LEVEL OF COMPLETED SCHOOLING

	0–7 Years		8 Years		9–11 Years		12 Years		13–15 Years		16 or More Years	
	1972–86	1987–90	1972–86	1987–90	1972–86	1987–90	1972–86	1987–90	1972–86	1987–90	1972–86	1987–90
Prestige	.111	−.006	.186	.007	.119	.160	.113	.123	.062	.091	.042	.071
SEI	.137	.272	.082	−.031	.135	.157	.178	.159	.145	.116	.072	.050
SSIC	.331	.114	.279	.058	.276	.134	.299	.138	.310	.076	.288	.097
Canonical	.402	−.017	.343	.036	.302	.123	.377	.119	.356	.065	.359	.081
N	341	61	331	46	1,059	251	2,696	826	1,567	688	1,783	730

problems of cross-validation? Answering this question requires an understanding of what the SSIC is supposed to measure and whether this is in fact the same thing that is supposed to be measured by the occupational scales used in models of status attainment.

Rytina says he intends to improve the measurement of the "intergenerational continuity of occupation." What is "occupation" (in the singular) and in what sense is it continuous across generations? An occupation is a socially recognized collection of similar jobs, where similarity is usually defined with respect to activities and duties performed, goods or services produced, forms of employment tenure, or—in broader classifications—educational requirements and economic rewards. As such, occupation is not a quantitative variable, and we can only assign numeric values to occupations in relation to one or more occupational characteristics.

In previous mobility studies continuity has been studied in two different respects. First, whether the child follows the same occupation as the father is treated as a question of continuity of occupations (plural). This is an essential definition of continuity in discrete models and corresponds to a concern with the proportion of cases that fall on the main diagonal of a mobility table. Unfortunately, the proportion of cases that is continuous in this sense depends on the level of detail of the occupational categorization and on the distribution of persons across occupations in each generation, as well as on specific mechanisms of occupational persistence or mobility.

A second mode of studying continuity has been achieved by refining the question and concentrating on only one or a few measurable properties of all occupations. With measures of such properties as prestige, education, or income it is possible to ask how much the offspring's occupation is like or unlike the parent's, even when the two are not in the same category. The degree of likeness or difference between the generations, in such an approach, has a specific metric corresponding to the units of the measured property. This refinement of the problem of continuity of occupations to become the problem of the continuity of properties associated with occupations is the essential shift in perspective in linear status attainment models, as compared with traditional models for mobility tables.[17]

Of course, the earlier focus on the continuity between generations in occupational categories as such did not imply a disregard for the properties of the occupations in the categories. Instead, it was implicit that

[17] However, recent developments of log-linear and log-multiplicative models have made it possible to combine the two perspectives (Goodman 1979; Hout 1984*a;* Hout 1984*b;* Hout 1988).

occupational categories had stable and relatively homogeneous properties, and this stability and homogeneity made it worth inquiring into continuity across generations.

Rytina's SSIC, however, deliberately disregards the existential, measurable aspects of life experienced in occupations. The SSIC places occupations close to one another purely on the basis of relative rates of intergenerational movement. To Rytina, an occupation that characteristically has better prospects for the next generation, say, in terms of SEI, should be ranked higher than one that has the same SEI but poorer next-generation prospects. Though SSIC manifestly does not take the status of occupations into consideration at all, Rytina justifies its use in part on the grounds that it corrects the SEI by repositioning occupations in closer correspondence with "the details of ascent and descent."

The idea that SEI values of occupations should be corrected by adjusting them on the basis of average intergenerational moves from and to higher and lower occupations is misguided. Though the precise meaning of SEI may be unclear, an undoubted advantage of the measure, which it shares with prestige, is that it is determined independently of the intergenerational movements that it is intended to gauge and predict. Socioeconomic status and prestige are properties of occupations, not properties of the relative standings between occupations of one generation and those of its offspring or parents. If two occupations have the same SEI but the first shows higher average SEI for its offspring, it may seem reasonable to raise its ranking to reflect the good fortune of its children. But this good fortune is not enjoyed by the parents and should not be imputed to them. On the other hand, if two occupations share the same SEI but the first has higher average levels of SEI among its parents, it does not seem reasonable to raise the first occupation's ranking, since there is no obvious reason why downwardly mobile offspring should be judged better off than their stable or upwardly mobile peers. Yet Rytina's procedure does give downwardly mobile occupations a higher ranking, and, furthermore, it adjusts an occupation's ranking by the same amount whether it is downward mobility from the preceding generation or upward mobility to the next that is involved.

We believe, in contrast, that an average decline or rise in SEI between generations is a social fact, which, if true, should be revealed by the measures used in mobility studies, rather than intentionally obscured by them. The study of social mobility should, to the greatest degree possible, be a study of the measurable properties of individuals, including the properties of their occupations, and the means by which individuals come to have those properties. We should study status, prestige, income, wealth, health, and well-being, but not a measure that places occupations

close to one another solely by relative rates of interchange. Our empirical and theoretical interest lies in the degree to which relative rates of interchange do or do not correspond to stability in measurable properties of occupational roles.

This said, it is still true that canonical correlation and similar procedures have their place. The canonical correlation provides a useful reference point for judging the degree to which a single scale or set of scales based on measured properties of occupations can account for the observed pattern of intergenerational movement. This is the use that Klatzky and Hodge (1971) made of the method in their exemplary analysis of Blau and Duncan's (1967) 17-category mobility table. Similarly, Hout's (1984*a*) status, autonomy, and training (SAT) model of the mobility table might be compared with log-linear or log-multiplicative models that generate intrinsic scalings of occupational categories. It is quite a long step from such uses to substituting canonical scores for measured properties of occupations.

A second, proper use of canonical scaling and related procedures is to determine the number of linear dimensions required to account for the association in a mobility table. For example, Klatzky and Hodge (1971) identified two distinct dimensions of occupational mobility in Blau and Duncan's (1967) 17-category mobility table, and similar tests for dimensionality can be constructed within the framework of log-linear and log-multiplicative models.

One seeming advantage of scales based on log-linear and log-multiplicative models is that they are marginally invariant, while canonical scales and correlations are not. However, it is possible to sidestep this problem. Hauser and Featherman (1977) and Featherman and Hauser (1978) estimated canonical correlations in a series of mobility tables that were standardized to fit a common set of relative mobility chances and thus demonstrated the effects of trends in occupational distributions on intergenerational occupational correlations. These trends (toward lower intergenerational correlations) closely followed the trends in correlations in the observed tables.

However, because of the problem of shrinkage, great caution should be exercised in any canonical analysis of mobility classifications, especially if the samples are small and the number of occupations is large. Even at the high levels of aggregation common in most studies of mobility tables, canonical scores exhibit shrinkage when they are applied to independent samples. Table 5 shows this effect in the two GSS samples and in Blau and Duncan's 1962 OCG data (1967, p. 496, table J2.1), using the standard 17-category classification. The rows of table 5 identify scales for the categories according to their sources, while the columns identify the

TABLE 5

FATHER-OFFSPRING CORRELATIONS OF 17-CATEGORY SEI, CANONICAL SCORES, AND AVERAGED CANONICAL SCORES IN THREE U.S. SAMPLES

	Sample, Correlation, and Shrinkage		
	OCG 1962	GSS 1972–86	GSS 1987–90
SEI	.403	.338	.305
Scores based on 1962 OCG:			
EIG1	.449	.364	.301
Shrinkage	1.000	.967	.886
Average EIG1	.445	.368	.297
Shrinkage	1.000	.986	.882
Scores based on 1972 to 1986 GSS:			
EIG1	.427	.383	.296
Shrinkage	.935	1.000	.856
Average EIG1	.437	.374	.296
Shrinkage	.980	1.000	.877
Scores based on 1987 to 1990 GSS:			
EIG1	.371	.324	.337
Shrinkage	.833	.868	1.000
Average EIG1	.368	.330	.326
Shrinkage	.854	.913	1.000

NOTE.—Rows labeled EIG1 give the father-offspring sample correlations of EIG*f* and EIG*o* from the specified sources. Rows labeled "Average EIG1" replace EIG*f* and EIG*o* with their arithmetic averages. SEI is the average Duncan score within each of the 17 categories. Shrinkage is the ratio of the correlation of a canonical score to the correlation of the SEI in a given sample and year relative to the corresponding ratio in the sample and year in which the canonical scores were estimated.

samples to which the scales are separately applied. The entries in the table are the correlations between fathers' and offsprings' occupations obtained with each combination of a scale and sample.

Row 1 of table 5 gives the intergenerational correlation of SEI at the 17-category level of aggregation. The temporal decline in this correlation again suggests a decrease in the intergenerational association of occupational standing over time. Row 2, EIG1, shows the correlations observed in the samples using canonical scores derived from the 1962 OCG data. The value .449 in column 1 of row 2 is the canonical correlation in the 1962 data, while the values .364 and .301 in columns 2 and 3 are intergenerational correlations in the GSS samples and are based on the canonical scores estimated in the 1962 data. Row 4, average EIG1, gives the same information using a simple arithmetic average of the fathers' and sons' canonical scores in the 1962 data. Because the SEI correlation declines over time, we have normed our comparisons of canonical correlations with SEI correlations on the ratios of canonical correlations to those of the SEI. Thus, in rows 3 and 5, both labeled "shrinkage," we

show the ratio of the canonical correlation to the SEI correlation in each sample to the corresponding ratio in the (1962 OCG) sample in which the canonical scores were estimated. In each of the GSS samples, the intergenerational correlation of canonical scores shrinks relative to the correlation based on the SEI. Table 5 also reports similar analyses of canonical scores estimated in the 1972–86 GSS and in the 1987–90 GSS. In each case, there is shrinkage in the correlation of canonical scores estimated in another sample, and the shrinkage is larger than we would expect from sampling variability alone.[18] Even at the high level of aggregation employed by Klatzky and Hodge (1971), canonical weighting cannot be relied upon to yield scores permitting valid, direct comparisons of intergenerational occupational persistence across populations. Even a modest expansion of occupational detail will yield substantial differences between sample values of the canonical correlation and estimates adjusted for degrees of freedom. For example, if the true correlation lies between 0.4 and 0.5, and we use a 100-category occupational classification, Lawley's correction leads us to expect sample estimates from 6% to 10% higher than the true intergenerational correlation in a sample like the 1972–86 GSS.

If canonical scores or similar scores are used as baseline associations in tabular models seeking to describe nonlinear patterns of mobility, the shrinkage of association when using a canonical score in a new sample may be inconsequential, since new scores will be extracted in each sample, and exact score values in any single sample will be of secondary interest. However, Rytina's proposal that constrained canonical scores be taken as measures of social standing strongly suggests that they should be used in more than one sample. Only by using the scores in more than a single sample can questions of cross-sample or cross-time variability in the intergenerational returns to social standing be addressed. Reestimating an SSIC in every new sample would confound changes in canonical scores, changes in occupational distributions, and sampling variability.[19]

However, in the present instance, no matter how the data are scaled, there does appear to be a decline in the intergenerational correlation across the three surveys. The decline appears in the SEI correlations, in the unconstrained canonical correlations (.449, .383, and .337), and in the correlations of average canonical scores (.445, .374, .326). Based upon the work of Featherman and Hauser (1978), we suspect that changing

[18] Lawley's (1959) formula implies that the sample canonical correlation should be about 98.4% of the sample correlation in the 1972–86 GSS when the analysis is based on 17 occupational categories.

[19] Again, there are appropriate methods to address these problems in the case of log-linear and log-multiplicative models for ordered categories (Goodman 1984).

occupational distributions, as well as changes in the association between father's and offspring's occupation (Hout 1988) are sources of these trends.

CONCLUSION

In the context of linear models of the stratification process, Rytina's proposed use of canonical scores must be judged wanting, in comparison with the well-established use of measured properties of occupations, such as NORC prestige scores or the Duncan SEI. Although we have used the SEI as a standard in our evaluation of Rytina's proposals, we do not want to advocate its use too strongly.[20] The most that can be said for it is that the SEI does appear to represent several of the sources of intergenerational occupational persistence and that future uses of it will be comparable to past uses. At the same time, we believe that the study of social stratification would be well served by the development of models that disaggregate occupational education from occupational income and that add other measurable occupational characteristics to models of intergenerational occupational stratification.

In models of mobility tables, canonical and related scaling models can be justified as an economical means of extracting linear associations from the table to evaluate nonlinear, categorical effects, to assess the dimensionality of association, and—with better statistical methods than Rytina has used—to provide a baseline against which substantive models of occupational mobility may be assessed. The ultimate goal in the analysis of mobility tables should also be to model relationships in measurable properties of occupations and persons. Only such models can give substance to our measurements of the intergenerational persistence of occupational roles.

REFERENCES

Blau, Peter M., and Otis Dudley Duncan. 1967. *The American Occupational Structure*. New York: Wiley.

Davis, James A., and Tom W. Smith. 1990. *General Social Surveys, 1972 to 1990* (MRDF). NORC ed. Chicago: NORC. Distributed by the Roper Center for Public Opinion Research, Storrs, Conn.

———. 1991. *The NORC General Social Survey: A User's Guide*. Chicago: NORC.

Duncan, Otis Dudley. 1961. "A Socioeconomic Index for All Occupations." Pp. 109–38 in *Occupations and Social Status*, edited by Albert J. Reiss, Jr. New York: Free Press.

Featherman, David L., and Robert M. Hauser. 1978. *Opportunity and Change*. New York: Academic.

[20] Hodge (1981) offers an insightful critique of the Duncan SEI.

———. 1983. *Occupational Changes in a Generation, 1973* (MRDF). DPLS ed. Washington, D.C.: U.S. Bureau of the Census. Distributed by Data and Program Library Service, University of Wisconsin, Madison.

Goodman, Leo A. 1979. "Simple Models for the Analysis of Association in Cross-Classifications Having Ordered Categories." *Journal of the American Statistical Association* 74 (September): 537–52.

———. 1984. *The Analysis of Cross-Classified Data Having Ordered Categories*. Cambridge, Mass.: Harvard University Press.

Hauser, Robert M., and David L. Featherman. 1977. *The Process of Stratification: Trends and Analyses*. New York: Academic.

Hodge, Robert W. 1981. "The Measurement of Occupational Status." *Social Science Research* 10 (December): 396–415.

Hout, Michael. 1984*a*. "Status, Autonomy, and Training in Occupational Mobility." *American Journal of Sociology* 89 (May): 1379–1409.

———. 1984*b*. "Occupational Mobility of Black Men in the United States, 1962–1973." *American Sociological Review* 49 (June): 308–22.

———. 1988. "More Universalism, Less Structural Mobility: The American Occupational Structure in the 1980s." *American Journal of Sociology* 93, no. 6 (May): 1358–1400.

Jencks, Christopher. 1990. "What Is the True Rate of Social Mobility?" Pp. 103–30 in *Social Mobility and Social Structure,* edited by Ronald L. Breiger. Cambridge: Cambridge University Press.

Klatzky, Sheila R., and Robert W. Hodge. 1971. "A Canonical Correlation Analysis of Occupational Mobility." *Journal of the American Statistical Association* 66, no. 333 (March): 16–22.

Lawley, D. N. 1959. "Tests of Significance in Canonical Analysis." *Biometrika* 46:59–66.

Siegel, Paul M. 1971. "Prestige in the American Occupational Structure." Doctoral dissertation. University of Chicago, Department of Sociology.

Stevens, Gillian, and David L. Featherman. 1981. "A Revised Socioeconomic Index of Occupational Status." *Social Science Research* 10 (December): 364–95.