TEMPORAL CHANGE IN OCCUPATIONAL MOBILITY: EVIDENCE FOR MEN IN THE UNITED STATES*

ROBERT M. HAUSER, JOHN N. KOFFEL, HARRY P. TRAVIS AND PETER J. DICKINSON

The University of Wisconsin – Madison

American Sociological Review 1975, Vol. 40 (June): 279-297

In a reanalysis of trend data on occupational mobility among men in the United States, we define ways in which movement among occupation groups across generations might be constant even when the occupational structure is changing. Our analysis suggests that no change has taken place in occupational mobility (as specified by our definition). Rather, the changing occupational distribution is the major factor affecting patterns of intergenerational occupational mobility. This implies the possibility of constructing occupational mobility tables for times when the age-specific occupation distribution is known, but no mobility survey has been carried out. Moreover, rather than treating the underlying process of mobility as a variable in comparative research and the variation in the distribution of occupations as a disturbance, it may be more fruitful to treat transformations of the occupational structure as problematic in comparative mobility research.

Trends in intergenerational occupational mobility or stratification have long been subjects of interest and controversy among sociologists. There are numerous reasons for an interest in mobility trends: a concern with the prospects for social equality or equality of opportunity, efforts to understand the transformation of the labor force in economic development, attempts to analyze the rise and fall of groups competing for power. Research and writing about mobility trends have often suffered from conceptual or analytic failures to separate the several aspects of stratification and mobility. Our purpose is to provide empirical estimates of mobility trends among U.S. men which avoid this confusion.

Variations in professional opinion about occupational mobility trends in the United States have probably been greater than any well-documented trends (Duncan, 1968: 675-80). Indeed, students of social mobility have reached no greater consensus on the matter than has the society they have sought to enlighten. Relying on the few available data, or in some cases on no data, mobility analysts have concluded that American society is becoming more rigid, that it is not becoming more rigid, that there has been no change in rates of mobility, or that we are moving toward a situation of full equality of opportunity. That observers have reached disparate conclusions from the same statistics is a problem in the sociology of knowledge (see Koffel, 1974). In other cases it may be possible to trace differences in conclusions about mobility trends to differences among data and statistical measures applied to them. We shall explore the latter possibility.

Relatively few facts are available about

*Prepared for meetings of the Research Committee on Social Stratification, International Sociological Association in Warsaw, Poland, March 1974. Computations were carried out in the Center for Demography and Ecology, University of Wisconsin – Madison, which has core support from the National Institute for Child Health and Human Development (Grant No. HD05876). This research was supported in part by the Division of Social Systems and Human Resources, RANN-NSF (Grant No. GI-31604X), by the American College Testing Program, by the College of Agriculture and Life Sciences, University of Wisconsin – Madison, and by funds granted to the Institute for Research on Poverty at the University of Wisconsin by the Office of Economic Opportunity pursuant to the provisions of the Economic Opportunity Act of 1964. We thank O. D. Duncan, P. Duncan-Jones, A. S. Goldberger, L. A. Goodman, K. Hope, R. M. Stoltzenberg and H. White for their comments on earlier drafts of this paper. The opinions expressed are those of the authors.

trends in intergenerational occupational mobility among U.S. men. Opportunities to enter high status occupations appear to have improved in successive cohorts of U.S. men for at least the last 40 years, irrespective of those men's occupational origins (Duncan, 1965; Hauser and Featherman, 1973; 1974a; 1974b). There is less evidence about changes in the ease of movement among occupational positions from one generation to the next, but a serious and comprehensive effort to assess trend from available data has suggested that the dependence of son's on father's occupation has been remarkably stable for more than half a century (Duncan, 1968; but see Duncan, 1966; Jackson and Crockett, 1964; Blau and Duncan, 1967; Tully et al., 1970). For related evidence of occupational mobility trends among U.S. women, see Featherman and Hauser (1974) and Hauser, Featherman and Hogan (1974).

In this paper we re-examine trend data on occupational mobility among men in the United States. Our analysis of temporal change follows the traditional interest of social mobility analysts in separating parameters of the occupational structure from the process of occupational mobility. This long-standing concern is well expressed by Raymond Boudon (1973) in his exhaustive review of mobility measures: "... a good mobility index should make a distinction between the amount of mobility generated by the changes in the social structure and the amount of mobility generated by other factors. Indeed, the former should be eliminated." Andrea Tyree (1973) has ably documented the parallel arguments by which several authors mistakenly hit upon the ratios of actual frequencies in a mobility table to those expected under independence as indices of "social distance mobility." These ratios are defective because the index for each cell in a mobility table varies inversely with the marginal proportion in its row and column and because the set of such ratios in a table determines both the row and column marginal distributions up to a constant of proportionality (Duncan, 1966; also see Goodman, 1969b). Thus, social distance mobility ratios are intimately related to the marginal distributions of the mobility table from whose influence they were supposedly freed.

Applying the work of Leo Goodman (1969a; 1969b; 1970; 1972a; 1972b) and Otis Dudley Duncan (1966) we shall define particular ways in which the pattern of movement among occupation groups across generations might be constant even when the occupational structure is changing. (Also see Haberman, 1974a:215-27.) We then reanalyze most (but not all) of the existing trend data on father to son occupational mobility in the United States, and our analysis suggests that no change has taken place in occupational mobility (as specified by our definition). That is, there is minimal evidence of change in the process of occupational mobility beyond that induced by the changing occupational structure and the succession of cohorts. This suggests a possible redirection of comparative studies of occupational mobility. It is no longer possible to assume that the underlying process of mobility is problematic in comparative analysis, while variations in the occupation distribution are a nuisance factor or disturbance. Rather, attention should be directed both to the causes of shifts in the occupation distribution and to their consequences in rates and patterns of mobility. We will take up the latter issue in a sequel to the present article (Hauser et al., 1975).

A Multiplicative Model

Suppose we observe tables of son's occupation by father's occupation at several different times. Denote the categories of father's occupation by $P(i=1,\ldots,I)$, those of son's occupation by $S(j=1,\ldots,J)$, and those of time by $T(k=1,\ldots,K)$. We may specify the observed frequencies, f_{ijk}, in the three-way classification of P by S by T by the multiplicative identity,

$$f_{ijk} = \eta\, \tau_i^P \tau_j^S \tau_k^T \tau_{ij}^{PS} \tau_{ik}^{PT} \tau_{jk}^{ST} \tau_{ijk}^{PST} \cdot (1)$$

Here, η is the geometric mean of the f_{ijk}, and the τ-parameters pertain to the probability that an observation appears in the subscripted cell of the superscripted univariate or joint distribution (Goodman, 1972). Thus, the parameters τ_i^P, τ_j^S, τ_k^T pertain to the respective probabilities that an observation appears in the i^{th}, j^{th}, or k^{th} cell of the marginal distributions of father's occupation, son's occupation, or time, relative to the grand mean. The parameters τ^{PS}_{ij}, τ^{PT}_{ik}, and τ^{ST}_{jk} pertain to the respective chances that an observation appears in the ijth, ikth, or jkth cell of the marginal classifications of father's occupation by son's occupation, father's occupation by time, or son's occupation by time, relative to the probabilities given by the products of lower-order parameters. Finally, the parameters τ^{PST}_{ijk} pertain to the probability that an observation appears in the ijkth cell of the complete three-way classification, relative to the probability given by products of lower-order parameters. (For an introductory discussion of multiplicative models for contingency tables, see Ku and Kullback, 1974.)

We are not especially interested in the complete or saturated model given by equation 1, but rather with the possibility that a more parsimonious model will accurately reproduce the observed cross-classification. According to our first alternative,

$$F_{ijk} = \eta\, \tau^{P}_{i}\, \tau^{S}_{j}\, \tau^{T}_{k}\, \tau^{PT}_{ik}\, \tau^{ST}_{jk}, \qquad (2)$$

where F_{ijk} is the expected frequency in the ijkth cell. The hypothesis expressed by the model is that the distributions of father's and son's occupations are determined exogenously at each time, and father's and son's occupations are conditionally independent. That is, equation 2 gives the same model as equation 1, except $\tau^{PS}_{ij} = \tau^{PST}_{ijk} = 1$ for all i, j and k. Using Leo Goodman's computer program, ECTA, or other available programs, it is a routine matter to obtain maximum-likelihood estimates of the F_{ijk} and to run a likelihood-ratio (LR) χ^2 test of the departure of the F_{ijk} from the f_{ijk}. A large value of χ^2_{LR} indicates that net of change in occupational structure there is association between father's occupation and son's occupation, or that association varies over time, or both. (For a lucid account of statistical inference in log-linear models, see Davis, 1974.) We do not believe the model of equation 2, since it postulates the statistical independence of father's and son's occupation at each time. (The χ^2_{LR} for this test may be written as the sum over time periods of the χ^2_{LR} for the test of independence between father's and son's occupations at each time.) Rather, it represents a set of baseline conditions of temporal change in the marginal distributions of father's and son's occupation against which we may assess and compare the associations between father's and son's occupations net of shifting occupation distributions.

We can also supplement the model of equation 2 to reflect temporally constant associations between son's and father's occupations:

$$F_{ijk} = \eta\, \tau^{P}_{i}\, \tau^{S}_{j}\, \tau^{T}_{k}\, \tau^{PS}_{ij}\, \tau^{PT}_{ik}\, \tau^{ST}_{jk}. \qquad (3)$$

This differs from the model of equation 1 in that $\tau^{PST}_{ijk} = 1$ for all i, j and k, and it differs from the model of equation 2, because the τ^{PS}_{ij} are not constrained to equal one. Equation 3 corresponds to a substantively interesting null hypothesis: there are changes over time in the marginal distributions of father's and son's occupations, but the latter associations are invariant with respect to time. Again, we may obtain maximum-likelihood estimates of the frequencies under the model of equation 3 and run a χ^2_{LR} test of goodness of fit. Here, a large value of χ^2_{LR} indicates the association between father's occupation and son's occupation is not constant across time (after controlling change over time in the occupational structure). Moreover, because the model in equation 3 includes all of the parameters of the model in equation 2, plus the parameters of the father-son association, the difference between the χ^2_{LR} for the model of equation 2 and that of equation 3 gives a test of the null hypothesis that there is no temporally constant association between father's occupation and son's occupation (after controlling change over time in the occupational structure).

While our analyses are not limited to the models specified in equations 2 and 3, all of our analysis does follow the same pattern. That is, in looking at a mobility table (or some aspect of it) we take change in the distributions of father's occupation and of son's occupation as a baseline condition. Then we measure the extent of temporally constant and of temporally variable association between father's occupation and son's occupation.

An illustration may clarify the implications of the model in equation 3. Suppose at time k the observed frequencies in the sub-table classifying a pair of categories of father's occupation by a pair of categories of son's occupation are as follows:

Father's occupation	Son's occupation 1	2
1	f_{11k}	f_{12k}
2	f_{21k}	f_{22k}.

The model of equation 3 specifies that the odds-ratio

$$\frac{(f_{11k}/f_{12k})}{(f_{21k}/f_{22k})}=\frac{(f_{11k}/f_{21k})}{(f_{12k}/f_{22k})}=\frac{f_{11k}f_{22k}}{f_{12k}f_{21k}} \quad (4)$$

is the same for all k within the limits of sampling error. For example, expression 4 might describe the chance that the son of a white-collar worker became a white-collar worker rather than a blue-collar worker relative to the chance that the son of a blue-collar worker became a white-collar worker rather than a blue-collar worker. The model of equation 3 says that odds-ratio does not vary across time. Moreover, the model says this invariance over time in the odds-ratios holds for any pair of categories of father's occupation and for any pair of categories of son's occupation, which need not be the same pair of father's and son's occupations. In general, the models investigated here may be described in terms of odds-ratios or certain functions of them, which may or may not vary across time. An implication of the model with temporally constant odds-ratios is that if we observe a table of father's occupation by son's occupation at any one time, and if we know the marginal distributions of son's occupation and of father's occupation at a second time, we can estimate the frequencies in the cross-classification of father's occupation by son's occupation at the second time by a simple iterative procedure (for example, see Deming, 1943; Mosteller, 1968).

In the following section we apply the models given by equations 2 and 3 and other instructive models in a reanalysis of available data on trends in occupational mobility among U.S. men. We shall see that models which are similar to equation 3 fit the data rather well. It should be noted that our choice of years, ages or time periods forms a part of the null hypothesis. That is, aggregation in the temporal dimension of our three-way classification may affect the association between father's occupation and son's occupation within time periods and its variation over time periods. However, our major empirical result does appear to stand up under alternative representations of the temporal dimension.

1962 OCG: Mobility to First Job

Figure 1 gives the names of 12 major occupation groups which combine detailed titles in the 1960 classification system of the U.S. Bureau of the Census. The groups are arrayed in descending order of the socioeconomic status of constituent occupations on the Duncan scale (1961), but much of our analysis requires no assumptions about hierarchy in the occupational structure. Using the classification scheme of Figure 1, we have tabulated son's first full-time civilian occupation after leaving school for the last time by father's occupation when the son was about 16 years old in each of nine 5-year cohorts covered in the March 1962 Current Population Survey supplement, "Occupational Changes in a Generation" (OCG; see Blau and Duncan, 1967). Our analyses pertain to 17,200 male OCG respondents who reported father's occupation and own first occupation. Since father's occupation refers to the sixteenth birthday of the son, while the son's first occupation refers to a fixed point in the son's career (even though son's age at first job varies), comparisons of the nine mobility tables for men of different ages in 1962 represent true intercohort comparisons, flawed only by the possibilities of differential mortality and recall.

We have proportionately adjusted values of χ^2_{LR} for the OCG data downward by a factor of 0.62 to reflect the efficiency of the Current Population Survey sample design relative to simple random sampling. This design factor is based on published standard errors of percentages (Blau and Duncan, 1967:477), but the

Figure 1. Twelve Category Occupation Distribution, U. S. Data from "Occupational Changes in a Generation"

Occupation Group	Duncan SEI
Professional, technical and kindred workers, self-employed	84
Professional, technical and kindred workers, salaried and others	73
Managers, officials and proprietors, exc. farm, salaried	68
Sales workers	49
Managers, officials and proprietors, exc. farm, self-employed	47
Clerical and kindred workers	45
Craftsmen, foremen and kindred workers	31
Operatives and kindred workers	18
Service workers, including private household	17
Farmers and farm managers	14
Farm laborers and foremen	9
Laborers, except farm and mine	7

efficiency of the CPS design may be greater for more complex statistics like those used here (Kish and Frankel, 1974). Thus, our selection of a design factor may incorrectly reduce the likelihood of our rejecting the null hypothesis when it is false. This potential bias may be offset because we have not made an adjustment for simultaneous inference. The effect of this adjustment on our analysis could be large. For example, Table 1 reports 16 tests of significance (specified without looking at the data). Thus, if we wished to maintain an overall significance level of $p = 0.05$ in the set of analyses reported in Table 1, we should reach a nominal probability level of $0.05/16 = 0.003$ in any one test before rejecting the null hypothesis in that test (Goodman, 1969a: 8-10). Moreover, even after the application of our design factor, the OCG sample is so large that we run a substantial risk of rejecting null hypotheses in favor of trivial alternatives at conventional levels of statistical significance.

Table 1 summarizes our analyses of the classification of father's occupation by son's first occupation by age in the 1962 OCG data. There is a line in the table for each model or comparison among models. The left-hand column gives a verbal statement of the null hypothesis, and the columns to the right report a goodness-of-fit test and other descriptive results. In reading this and subsequent tables, it will be helpful to recall that χ^2 with ν degrees of freedom has expectation ν and variance 2ν and $\chi^2(\nu)$ is distributed normally when ν is large. It is also helpful to look at the ratio, $\chi^2(\nu)/\nu$, in assessing the size of a component of association relative to others based on the same number of observations (Goodman, 1972a:1058; Haberman, 1974b: 592-3).

Panel A of Table 1 reports the analysis of change over time in the occupational distributions of fathers and sons. Clearly, those distributions have changed over the half-century represented in the OCG data. Against the null hypothesis of temporally constant occupation distributions, we obtain χ^2_{LR} of 432.22 in the case of father's occupation and 429.95 in the case of son's first occupation. The marginal tables of occupation by time each have 88 degrees of freedom, so the changes over time in the occupation distributions are statistically significant. (We thank Keith Hope for recommending these two calculations. See Hope, 1974, for an elegant analysis of mobility trends in British society.)

In line B1 we estimate the baseline model of equation 2. This model makes the temporal

Table 1. Analyses of Temporal Homogeneity in Mobility from Father's Occupation to Son's First Full Time Occupation: U. S. Men Aged 20 - 64 in 1962 by 5-Year Cohorts (N = 17,200)

Null Hypothesis	χ^2_{LR}	df	p	Δ	χ^2_H/χ^2_T
A. Marginal trends					
1. Constant distribution of father's occupation	432.22	88	.000	--	--
2. Constant distribution of son's first occupation	429.95	88	.000	--	--
B. Full matrix (12 x 12)					
1. Margins vary with time (age)	5344.57	1089	.000	26.31	100.0%
2. Margins vary with time (age), constant father-son interactions	854.19	968	>0.5	8.56	16.0
3. No father-son interaction (B2 vs. B1)	4490.38	121	.000	17.75	84.0
C. Main diagonal blocked					
1. Margins vary with time (age)	3512.87	981	.000	18.94	65.7
2. Margins vary with time (age, constant father-son interactions	739.99	872	>0.5	7.03	13.8
3. No father-son interaction (C2 vs. C1)	2772.88	109	.000	11.91	51.9
D. Five broad groups blocked					
1. Margins vary with time (age)	1213.48	801	.000	8.71	22.7
2. Margins vary with time (age), constant father-son interactions	593.93	712	>0.5	5.56	11.1
3. No father-son interaction (D2 vs. D1)	619.55	89	.000	3.15	11.6
E. Selected tests for trend					
1. Constant occupational inheritance (C2 vs. B2)	114.20	96	>0.1	1.53	2.2
2. Constant interactions within major groups off the main diagonal (C2 vs. D2)	146.06	160	>0.5	1.47	2.7
3. Constant mobility	18.12	8	.022	0.24	0.3
4. Constant upward and downward mobility	25.54	16	.064	0.28	0.5
5. Constant occupation-specific inheritance	96.09	88	.271	1.29	1.8

changes in the occupation distributions of fathers and sons a part of the null hypothesis, but it also attempts to account for all of the frequencies in the three-way classification of father's occupation by son's first occupation by age. Against the baseline model, we obtain χ^2_{LR} of 5344.57 with 1089 degrees of freedom, which is obviously statistically significant at any of the conventional levels. We reject the model with temporally variable marginals but no association between father's occupation and son's first occupation.

The column labeled Δ gives the index of dissimilarity between the distributions of observed frequencies and the maximum-likelihood estimates. (See Taeuber and Taeuber, 1965:195-245, for a discussion of Δ and related measures, and Goodman, 1965, for a related application.) In contrasts between models and in tests for trend, the value of Δ measures the improvement in fit between expected and observed distributions. Values of Δ may legitimately be compared within Table 1 and within later tables, but not between tables, because the index is sensitive to differing levels of aggregation. The model of line B1 misclassifies 26.31 percent of the distribution. This is the percentage of cases which would be shifted to another cell of the classification by a proper specification of the relationship between father's occupation and son's first occupation and changes in it over time.

Finally, the entry of 100.0% in the column labeled χ^2_H/χ^2_T indicates that we take $\chi^2_{LR} =$ 5344.57 to represent all of the variation or association in the data which we might wish to explain by subsequent complications of our baseline model. That is, we shall treat the baseline χ^2_{LR} just like the total sum of squares in a conventional analysis of variance. This analogy is developed extensively by Goodman (1970; 1971; 1972a). However, in the present context, explaining all of the variance does not imply accounting for inter-unit variability, but fitting a joint frequency distribution. Thus, the present effort has more in common with the testing of overidentifying restrictions in a path model than with increasing the size of a multiple correlation.

When we estimate the model of equation 3, we obtain the results shown in line B2 of Table 1. Under the hypothesis of temporally constant father's occupation-son's first occupation interactions, we obtain $\chi^2_{LR} = 854.19$ with df = 968. (By "interactions" in this context, we refer to the multiplicative parameters of association between father's occupation and son's first occupation.) Our test statistic is well below its expected value, so clearly we do not reject the null hypothesis. The model of temporally constant father-son interactions accounts for all but 16 percent of the variation in the baseline model, and it misclassifies only 8.56 percent of the distribution.

It is instructive to compare lines A1 and A2 of Table 1 with line B2, for these, respectively, represent change over time in father's occupation, in son's occupation and in their association. The χ^2_{LR} in line B2 is almost as large as the sum of the corresponding entries in lines A1 and A2 but this is misleading because the χ^2_{LR} in line B2 uses many more degrees of freedom. Note the χ^2_{LR} in line A1 and in line A2 is each about 5 times larger than its expectation, while the χ^2_{LR} in line B2 is less than its expected value. (Adjusting components of χ^2_{LR} for their differing degrees of freedom is similar to adjusting R^2, the coefficient of determination, for loss of degrees of freedom.) Thus, we conclude that changes over time in the distributions of father's occupation and of son's first occupation are far more important than changes over time in the association between father's occupation and son's first occupation.

In line B3 we report the differences between corresponding entries in lines B1 and B2. This gives us a test of the hypothesis of no father's occupation-son's first occupation interactions, given son's birth cohort. As might be expected, we obtain a large and statistically significant χ^2_{LR} which accounts for 84 percent of the variation in the baseline model and correctly classifies an additional 17.75 percent of the distribution.

While we have not detected any changes in father-son occupation interactions in the analysis of panel B, we are not yet content to conclude that no change has occurred. The failure to reject null in a global test does not imply that a more narrowly specified hypothesis will not be rejected. For this reason, we have specified and tested a series of hypotheses about changes in the relationship between father's and son's occupation which involve occupational mobility and/or occupational inheritance or disinheritance. It should be kept in mind that "inheritance" does not here refer to job inheritance in the strict sense, but only to the possibility that a son may enter the same major occupation group as his father (Goodman, 1969a:14).

In panel C of Table 1 we test a series of models following the same logic as in panel B, but we block the entries along the main much less than its expectation, and we are unable to reject the model of quasi-homogeneity, i.e., temporally constant father-son interactions off the main diagonal. Line C3 contrasts the models of lines C1 and C2, and from it we conclude that there is a significant set of father-son interactions off the main diagonal. Thus, while the cross-classification is quasi-homogeneous in respect to patterns of occupational mobility, it does not fit the model of quasi-perfect mobility or quasi-independence (White, 1963; Goodman, 1965; 1968; 1969a), at least in respect to the distinction between occupational inheritance and occupational mobility as defined by the 12 major occupational groups. Indeed, the temporally constant, off-diagonal interactions account for more of the variation from the baseline model than do the observed frequencies in the main diagonals. (Compare line C3 with the difference between lines B1 and C1.) diagonal of each father-son mobility table. That is, we constrain the entries involving occupational inheritance, forcing each frequency on the main diagonal to take on its observed value. In this way, the null hypothesis presumes the observed pattern of occupational inheritance in each time period. Thus, entries in panel C of Table 1 differ from corresponding entries in panel B by excluding the effects of departures from the null hypothesis where father and son are in the same major occupation group. By comparing entries in panels B and C, we will be able to isolate the effects of occupational mobility and changes therein from the effects of differential occupational inheritance and changes therein (Goodman, 1969a:29-39).

In line C1, we fit the model of temporally-variable margins with no father-son interactions off the main diagonal. Clearly, this model can be rejected, although the main diagonals do account for about 35 percent of the association in the tables (compare lines B1 and C1). In line C2, we fit the model with observed frequencies in the main diagonals and a temporally constant set of father-son interactions off the diagonal. Goodman (1969a:29-30) uses the term "quasi-homogeneity" to refer to models in which two (or more) cross-classifications are specified to be alike in some interactions but not in others. In the model of line C2, patterns of occupational mobility, but not of occupational inheritance, are assumed stable across time. Here, χ^2_{LR} is much less than its expectation, and we are unable to reject the model of quasi-homogeneity, i.e., temporally constant father-son interactions off the main diagonal. Line C3 contrasts the models of lines C1 and C2, and from it we conclude that there is a significant set of father-son interactions off the main diagonal. Thus, while the cross-classification is quasi-homogeneous in respect to patterns of occupational mobility, it does not fit the model of quasi-perfect mobility or quasi-independence (White, 1963; Goodman, 1965; 1968; 1969a), at least in respect to the distinction between occupational inheritance and occupational mobility as defined by the 12 major occupational groups. Indeed, the temporally constant, off-diagonal interactions account for more of the variation from the baseline model than do the observed frequencies in the main diagonals. (Compare line C3 with the difference between lines B1 and C1.) However, we shall take up this hypothesis again.

In line E1 of Table 1, we contrast the models of line B2 and line C2. Since the former model posits a temporally constant pattern of association between father's and son's occupations, and the latter differs only in permitting occupational inheritance to fluctuate with time, the comparison in line E1 tells whether the pattern of occupational inheritance has changed over time. Changes in occupational inheritance account for only 2.2 percent of the association in the table, and they reclassify only 1.5 percent of the distribution. Effects this small might have occurred by chance more than 10 percent of the time. Thus, we may specify our earlier finding of no significant changes in father-son interactions to say there have been no changes of association in the aggregate of cells involving occupational inheritance or in the aggregate of cells involving occupational mobility.

In the analyses of panel D of Table 1, we again block the main diagonal entries of each father-son mobility table, and additionally we block cells off the main diagonal of the tables which pertain to movement within five very broad occupation groups. The broad occupation groups are upper white-collar (professional, technical and kindred workers; managers, officials and proprietors, except farm), lower white-collar (sales workers; clerical and kindred workers), upper blue-collar (craftsmen, foremen and kindred workers; operatives and kindred workers), lower blue-collars (service workers, including private household; laborers, except farm and mine), and farm (farmers and farm managers; farm laborers and foremen). As before, when we block cells within the five broad groups (including the main diagonal), we are constraining the model to reproduce exactly the observed frequencies in those cells in each time period. Thus, the null hypotheses in panel D presume the observed patterns of association within the five broad groups and changes in them over time. Looking at panel D, we can assess the amount of association between father's occupation and son's occupation outside the five broad groups; the null hypothesis here (line D3) is quasi-perfect mobility, but in a smaller subset of cells than in the analysis of panel C. Also, we can assess changes over time in association outside the five broad groups, testing the null hypothesis of quasi-homogeneity (line D2), but in a smaller subset of cells than in the analyses of panel C. Finally, by comparing corresponding entries in panels D and C, we can assess the amount and changes in the amount of association between father's occupation and son's first occupation which involves sons in the same broad group, but not the same major group as their fathers.

As shown in line D1, even when we block all of the cells within these very broad occupation groups, there remains a statistically significant and substantial amount of association in the tables, about a quarter of the association in the complete tables. In line D2, we estimate a model which contains temporally constant father-son occupation interactions outside the five broad occupation groups, and χ^2_{LR} is less than its expected value. We are unable to reject this model of quasi-homogeneity, which misclassifies only 5.56 percent of the distribution and accounts for almost 89 percent of the association in the baseline model. As reported in line D3, the temporally constant interactions outside the five broad groups account for a statistically significant 11.6 percent of the association in the baseline model, thus permitting us to reject this weaker statement of the hypothesis of quasi-perfect mobility. Finally, in line E2, we contrast the models of lines C2 and D2, which differ only in that the latter permits temporal variation in the association between father's and son's occupation in cells which are off the main diagonal but within the five broad groups defined above. The χ^2_{LR} in line E2 is well below its expected value, so we are unable to reject the hypothesis of temporal homogeneity in the association between father's occupation and son's occupation off the main diagonal, but within the five broad occupation groups. Thus, our analysis continues to support the hypothesis of temporal homogeneity in occupational mobility between generations, and at the same time, it fails to support the hypothesis of quasi-perfect mobility outside the five broad groups.

Panel E of Table 1 reports other tests for trend in mobility which are related to, but distinct from the earlier ones. Line E3 tests the hypothesis that the pattern of mobility from father's occupation to son's first occupation is the same in every cohort, except the rate of mobility, i.e., the propensity to move versus stay, has changed over time. This model differs from the model of line E1 in introducing time-varying parameters for occupational mobility *per se*, rather than the several occupation-specific parameters of the model in line E1. Our mobility parameter is a special case of the "triangles" parameters described by Goodman (1972b), but our model under the null hypothesis is not a triangles model, but the model of equation 3. If the alternative hypothesis were true, the model of equation 3 would have to be supplemented by time-varying mobility parameters. The addition of the time-varying parameters is nominally significant at the 0.022 level. The changing mobility parameters account for only 0.3 percent of the association in the baseline model, and they allocate only 0.24 percent of the distribution to a different cell in the classification. We regard this effect as substantively trivial and possibly random for the reasons given earlier.

Under the alternative hypothesis just discussed, the mobility parameter only makes the distinction between movers and stayers, thus neglecting the possibility that propensities to move upward and downward may change over time in different directions. That is, offsetting changes in upward and downward mobility rates may yield no change in overall mobility rates. In line E4 of Table 1 we test the hypothesis that rates of upward or downward mobility have changed (as well as rates of stability). To define "upward" or "downward" mobility (a distinction not used in any earlier model), we ordered the major occupation groups by Duncan scores (see Figure 1), and we introduced separate time-varying parameters for the triangular aggregates of cells pertaining to upward and to downward mobility. However, the model under the null hypothesis is not a triangles model, but that of equation 3. In line E4, the χ^2_{LR} for changes in upward and downward mobility is not even significant at a nominal 0.05 level. Moreover, there is no significant difference between the changes over time in upward and in downward mobility. Changes over time in rates of upward relative to downward mobility account for a χ^2_{LR} of only 7.42 with 8 degrees of freedom, and they account for only 0.2 percent of the association in the baseline model (compare lines E3 and E4 of Table 1).

Finally, in line E5 of Table 1 we report a test of changes across time in occupation-specific inheritance. That is, we ask whether there are any changes over time in propensities to inherit specific occupations, aside from possible trends in propensity toward mobility *per se* (tested in line E3). The reader will note this test contrasts the model of line E1 with that of line E3. That is, changing occupational inheritance permits changes both in general and occupation-specific propensities to move. Again, as shown in line E5, we fail to reject the null hypothesis of temporal homogeneity in occupation-specific inheritance. By an extension of the methods used here it would be possible to assess changes over time in occupational inheritance in each major occupation group, but we have not done so.

It may be useful to summarize our analyses of mobility from father's occupation to son's first occupation by cohort in the 1962 OCG data. First, we have found substantial change in the occupational structure over time, as evidenced by change in the distributions of father's occupation and of son's occupation. Second, we have found strong patterns of association between father's occupation and son's occupation. Considering only the temporally constant association between father's occupation and son's first occupation, the ratios of χ^2_{LR} to its degrees of freedom are 143.13 for the aggregate of cells on the main diagonal, 107.67 for the aggregate of cells off the main diagonal but within the five broad occupation groups, and 6.96 for the aggregate of cells outside the five broad occupation groups. While much of the association in the tables is attributable to occupational inheritance or to movement between closely related occupation groups, there appears to be association throughout most, if not all of the mobility table. Thus, the OCG data on mobility to first occupations do not appear to conform to the pattern of quasi-perfect mobility which Goodman (1965; 1969a; 1969b) has observed in British and Danish mobility tables. This may partly be a consequence of the larger number of occupation groups employed in the present analysis (Haberman, 1974a:216), or perhaps we have not located the subset of cells in which quasi-perfect mobility holds.

Third, we have found remarkable homogeneity in the patterns of association between

Figure 2. Ten Category Occupation Distribution, Indianapolis Data from Rogoff's Recent Trends in Occupational Mobility

Professional
Semi-professional
Proprietors, managers, officials
Clerks and salesmen
Skilled
Semi-skilled
Protective service
Farming
Personal service
Unskilled

father's occupation and son's first occupation when changes in the occupational structure have been controlled. There appear to be negligible differences among cohorts in the propensity to move versus stay; in the propensities to move up, move down, or stay; in the propensity to move up relative to the propensity to move down; in the propensity to inherit one's father's occupation; in patterns of movement among similar occupation groups; and in patterns of movement among disssimilar occupation groups.

Blau and Duncan (1967:107-11) have preceded us in attempting to measure mobility trends by comparing tables of mobility from father's occupation to son's first occupation among men of different ages in the 1962 OCG survey. They compared 4 ten-year age cohorts using a three-category occupation classification (manual, nonmanual, farm), a ten-category classification and scores of detailed titles on Duncan's scale (1961). From standard mobility indexes and contingency measures in the three by three and ten by ten classifications, Blau and Duncan (1967:109) concluded, "the extent of association between origin and destination has been less recently than at an earlier date." However, the product-moment correlations between status scores of fathers and sons are virtually the same in each cohort. We shall see that at least the first of these results may be consistent with the hypothesis of temporal invariance which we have advanced (Hauser et al., 1975).

Five National Surveys: 1947 to 1972

Much of the recent analysis of trends in occupational mobility in the U.S. rests on comparisons among four national surveys in which both father's occupation and son's current occupation were ascertained. These are a 1947 survey by the National Opinion Research Center (NORC), 1952 to 1957 surveys by the Survey Research Center (SRC) at the University of Michigan and the 1962 OCG survey carried out by the U.S. Bureau of the Census. The first three surveys were compared by Jackson and Crockett (1964), and the last was added to the series by Blau and Duncan (1967:97-105). We have added a fifth observation to the series of mobility tables for all U.S. men. This is a table for about 500 adult men in 1972 obtained from an NORC (1972) survey.

Jackson and Crockett emphasized the lack of evidence of increasing rigidity in the stratification system and suggested there may have been some lessening of the dependence of son's on father's occupation. They conclude (Jackson and Crockett, 1964:15), "The data suggest, however, that the rate of occupational mobility in the United States has increased somewhat since the end of World War II. At the least, we found scant evidence that the system of occupational inheritance is growing more rigid." Blau and Duncan (1967:105) reached a similar conclusion.

There are serious questions about the comparability of data across years since there was some variation in the items used to ascertain occupation, and the five surveys were carried out by three different agencies (Jackson and Crockett, 1964:11; Blau and Duncan, 1967: 98-103). In particular, there are peculiarities in the marginal distributions of father's and of son's occupations in the 1947 NORC data. These were recognized by Jackson and Crockett and by Blau and Duncan, but still leave their conclusions in doubt (Koffel, 1974). Unfortunately, the four mobility tables were rendered comparable only in respect to a division among manual, nonmanual and farm occupations.

Our analyses of the three by three tables from five national surveys are summarized in Table 2. We shall not recapitulate the logic of our analysis, which follows that of Table 1. In panel A we report the likelihood-ratio χ^2 for the variation across surveys of father's occupation and of son's current occupation. Each of these is large, relative to its degrees of freedom, but the variation in the occupation distributions across surveys may reflect methodological factors as well as change over time. (We caution the reader against comparison of χ^2_{LR} and our other measures across tables, because of their differing levels of aggregation and numbers of observations.) In the baseline model of line B1, the margins are free to vary across time, but there is no association between father's occupation and son's occupation. The baseline model yields a large and statistically significant χ^2_{LR}, and it misclassifies 17.74 percent of the distribution. As shown in line B2, when we introduce a set of time-constant parameters for the interactions between father's and son's occupations, we account for 99.4 percent of the association in

Table 2. Analyses of Temporal Homogeneity in Mobility from Father's Occupation to Son's Current Occupation: U. S. Men in 1947 (NORC), 1952 (SRC), 1957 (SRC), 1962 (OCG), and 1972 (NORC) by Three Major Occupation Groups

Null Hypothesis	χ^2_{LR}	df	p	Δ	χ^2_H/χ^2_T
A. Marginal trends					
1. Constant distribution of father's occupation	118.40	8	.000	--	--
2. Constant distribution of son's occupation	140.05	8	.000	--	--
B. Full matrix (3 x 3)					
1. Margins vary with time	3268.18	20	.000	17.74	100.0%
2. Margins vary with time, constant father-son interactions	19.52	16	.242	1.02	0.6
3. No father-son interaction	3248.66	4	.000	16.72	99.4
C. Main diagonal blocked					
1. Margins vary with time	30.87	5	.000	0.61	0.9
2. Margins vary with time, constant father-son interactions	1.00	4	>0.5	0.06	0.0
3. No father-son interaction	29.87	1	.000	0.55	0.9
D. Selected tests for trend					
1. Constant occupational inheritance (C2 vs. B2)	18.52	12	.100	0.96	0.6
2. Constant mobility	2.95	4	>0.5	0.16	0.1
3. Constant upward and downward mobility	11.10	8	>0.1	0.64	0.3
4. Constant occupation-specific inheritance	15.57	8	.050	0.80	0.5

Note: N's are NORC-1947, 1153; SRC-1952, 747; SRC-1957, 1023; OCG-1962, 17,615; NORC-1972, 483.

the baseline model. The observed results might easily have occurred by chance if the model of time-constant interactions were correct, and the model misclassifies only one percent of the cases. As shown in line B3, the time-constant interactions are large and statistically significant.

In panel C of Table 2 we report analyses in which the three main diagonal cells of each mobility table have been blocked to permit occupational inheritance to vary freely over time. (Recall that occupational inheritance is defined by our occupation categories, which are very broad.) When the margins are fixed and the main diagonals are blocked, there remains a very small but statistically significant association in the tables (line C1). Slightly more than one percent of the distribution is misclassified, and less than one percent of the association in the baseline model remains. Thus, the five-survey data come very close to fitting a model of quasi-perfect mobility, when they are analyzed in a three by three mobility table. As shown in lines C2 and C3 of the table, virtually all of the remaining association may be explained by temporally constant interaction between father's occupation and son's occupation off the main diagonal. When that association is entered into the model, a virtually perfect fit of the data is obtained.

Line D1 of Table 2 contrasts the models of lines C2 and B2, which differ only with respect to the possibility of changes in occupational inheritance over time. The χ^2_{LR} for this contrast is not significant at even the 0.05 level, so we fail to reject the hypothesis of no change in occupational inheritance. Changes over time in father-son interactions on the main diagonal of the five-survey tables account for less than one percent of the association in the baseline model and reclassify less than one percent of the distribution. We next consider the possibility that there is a changing propensity to move versus stay, but there is no support for this in the data (line D2). There is no support for the hypothesis that propensities toward upward mobility, downward mobility and stability have changed across surveys (line D3), nor is there any support for the hypothesis that propensities toward upward relative to downward mobility have changed across surveys (compare lines D2 and D3). Finally, a test for temporal changes in occupation-specific inheritance reaches a nominal 0.05 level of significance, but such changes account for a negligible share of the association in the five-survey mobility tables. Thus, our analysis of the 25-year time series of mobility tables for U.S. men supports the analyses of other researchers insofar as they have emphasized a lack of change in the association between father's occupation and son's occupation.

Indianapolis: 1910 and 1940

The original and classic study of mobility trends in the United States was that of Natalie Rogoff (1953a; 1953b), who analyzed occupation reports of men about themselves and their fathers obtained from marriage license applications in Indianapolis in years centered around 1910 and 1940. About 10,000 reports were obtained in each period. Rogoff's analyses of these data were based primarily on comparisons of the fatally flawed "social distance mobility ratio" (Duncan, 1966; Blau and Duncan, 1967; Tyree, 1973). Depending on where one looks in Rogoff's writings, her conclusions are discrepant with regard to changes in the dependence of son's on father's occupation. In her monograph, Rogoff (1953a:106) writes: "... the processes by which men selected and were selected for occupations were more closely related to social origins in 1940 than they had been in 1910." However, in a summary of her work (Rogoff, 1953b:451), she concludes, "... no great change has taken place in recent times in the extent to which men may move from the occupational origins represented by their fathers' positions. The channels to social mobility afforded by the contemporary occupational structure are about as easily traversed now as they were at the beginning of the century."

Otis Dudley Duncan's (1966) careful reanalysis of Rogoff's data was primarily methodological in intent. With the possible exception of the analyses most closely resembling our own, his results suggested that the 1910 and 1940 mobility tables were very similar with regard to the dependence of son's on father's occupations. For example, after running regression analyses of the 1910 and 1940 data from Rogoff's detailed tables using several alternative scales of occupational status, Duncan (1966:69-70) writes, "... the socioeconomic status of occupations held by white Indianapolis men marrying in 1940 was no more closely related to the socioeconomic status of their occupational origins than had been the case for white men marrying in 1910."

Another of Duncan's reanalyses of Rogoff's data was a major stimulus to our thinking about change in mobility tables, though with somewhat different empirical results. After verifying that the 1910 and 1940 ouflow matrices produce different destination vectors when multiplied by the same origin vector, and that the two matrices have correspondingly different fixed point distributions, Duncan (1966:73) asks,

> Is it possible that differences in the 1910 and 1940 mobility patterns are due solely to shifts in the distribution of job opportunities open to young men? . . . Can we, in other words, contrive a comparison between the two mobility tables putting the change in occupation structure in the role of an exogenously determined factor, which then induces a change in mobility patterns? The starting point of the comparison is to test the null hypothesis that all changes in the mobility table are due to proportional adjustments occasioned by changes, 1910 to 1940, in the two marginal distributions—the distribution of sons by their father's occupations, and the distribution of sons by their own occupations.

In order to test this hypothesis, Duncan used a least-square procedure suggested by Deming (1943) to estimate the frequencies in the 1940 table by proportional adjustment of rows and columns of the 1910 table. This null hypothesis is exactly the same as that of our equation 3, namely that the father's and son's occupation distributions vary over time, and there is a temporally constant set of father-son interactions. However, Duncan's statistical analysis differs from our own, in that Duncan fitted the 1940 margins to the 1910 table, while we have obtained maximum-likelihood estimates of the frequencies using the data for both periods.

Figure 2 gives the titles of the ten major occupation groups into which the Indianapolis data were aggregated, listed in order of socioeconomic standing on the Duncan scale. Our analyses of the Indianapolis data (Rogoff, 1953a:44-5) are summarized in Table 3. Again, the logic of our analysis is the same as in Table 1, but the reader should not compare components of variation or indexes of dissimilarity across tables. Between 1910 and 1940, there were large shifts in the distributions of father's occupation and of son's occupation in Indianapolis (lines A1 and A2). Obviously, we reject the baseline model of changes in the occupational structure, but no association between father's occupation and son's occupation (line B1), which misclassifies 17.68 percent of the distribution. As shown on lines B2 and B3, when we also specify a temporally constant set of father-son interactions, we correctly classify an additional 15.04 percent of the distribution, leaving only 2.64 percent wrongly classified. The model of temporally constant interactions accounts for 97.5 percent of the association in the baseline model. As in the OCG data on mobility to first occupations, changes over time in the marginal distributions of the mobility table for Indianapolis are much larger relative to their degrees of freedom than are changes over time in the association between father's occupation and son's occupation.

Since the Indianapolis data are not from a probability sample, it is not clear how seriously we should take the probability levels associated with our goodness-of-fit tests. At conventional levels, we would reject the hypothesis of temporally constant interactions, for $p = 0.003$ in line B2. However, given the very large number of observations and the very small amount of association attributable to changes in the father-son interactions, our analyses of the Indianapolis data lead us to the same conclusion as our analyses of the OCG data on mobility to first job and the five national surveys. There has been little if any change in the association between father's occupation and son's occupation. We have also analyzed the five by four tables of mobility for men in Indianapolis with urban occupations in 1910, 1940 and 1967 with similar results (Tully et al., 1970;192). There are nominally significant differences in father-son interactions between years, but these account for only 1.7 percent of the association in the baseline model. We have given less attention to the 1967 Indianapolis data than to the 1910 and 1940 data because of the different methods used in the 1967 survey.

In panel C of Table 3, we report analyses in which the main diagonals of the 1910 and 1940 tables have been blocked, permitting occupational inheritance to vary over time. When the margins are fixed and the main diagonal is blocked, there is still substantial

Table 3. Analyses of Temporal Homogeneity in Mobility from Father's Occupation to Son's Current Occupation: Men Married in Indianapolis in 1910 and 1940 by Ten Major Occupation Groups (N = 20,146)

Null Hypothesis	χ^2_{LR}	df	p	Δ	χ^2_H/χ^2_T
A. Marginal trends					
1. Constant distribution of father's occupation	686.06	9	.000	--	--
2. Constant distribution of son's occupation	848.55	9	.000	--	--
B. Full matrix (10 x 10)					
1. Margins vary with time	4794.62	162	.000	17.68	100.0
2. Margins vary with time, constant father-son interactions	121.73	81	.003	2.64	2.5
3. No father-son interaction	4672.89	81	.000	15.04	97.5
C. Main diagonal blocked					
1. Margins vary with time	1201.19	142	.000	7.56	25.1
2. Margins vary with time, constant father-son interactions	96.71	71	.046	1.93	2.0
3. No father-son interaction	1104.48	71	.000	5.63	23.0
D. Selected tests for trend					
1. Constant occupational inheritance (C2 vs. B2)	25.02	10	.008	0.71	0.5
2. Constant mobility	4.98	1	.013	0.23	0.1
3. Constant upward and downward mobility	5.13	2	.081	0.26	0.1
4. Constant occupation-specific inheritance	20.03	9	.019	0.48	0.4

association between father's and son's occupations (line C1), but three-quarters of the association in the baseline model for Indianapolis data may be attributed to occupational inheritance (compare lines B1 and C1). When we estimate frequencies based on temporally constant father-son interactions off the main diagonal (line C2), we account for all but two percent of the association in the baseline model, and we misclassify less than two percent of the distribution. The temporally homogeneous father-son interactions off the main diagonal are overwhelmingly significant (line C3).

When we contrast the models with and without changes over time in occupational inheritance (line D1), we find the temporal shifts to be extremely small, though they would be statistically significant at conventional probability levels. Changes over time in occupational inheritance account for only 0.5 percent of the association in the baseline model, and they reclassify only 0.71 percent of the distribution. The test for changes in the propensity to move between 1910 and 1940 is not even nominally significant at the 0.01 level (line D2), and it accounts for only a tenth of a percent of the variation in the baseline model. Even less substantial results are obtained when we test for possible changes both in upward and downward mobility (line D3), and there is virtually no change over time in the propensity toward upward relative to downward mobility (compare lines D2 and D3). Finally, there is little evidence of change in occupation-specific inheritance in Indianapolis (line D4). Considering the strict comparability of measurement in the Indianapolis data, the thirty-year separation between the two measurements and the very large samples in the two years, the stability of occupational mobility patterns is remarkable.

It is instructive to compare our results with those in Duncan's test of proportional change in the Indianapolis data (Duncan, 1966:74-7). Taking the 1910 matrix as the standard, Duncan's model misclassifies 7 percent of the distribution under the model of proportional change, while we misclassify 2.64 percent of the distribution under the same model. Duncan reports a decomposition of the sum of squares of ratios of 1940 to 1910 frequencies whose interpretation parallels that of the components of association in our baseline model. In Duncan's analysis, 76 percent of the sum of squares is attributable to proportional change in the row or column marginals. There are no corresponding figures in our decomposition since our baseline model incorporates change in the marginal distributions; however, we did note the substantial shifts in the occupational structure from 1910 to 1940. Duncan reports that 7 percent of the sum of squares is "due to proportional change in both distributions, reflecting the initial correlation between son's and father's occcupation," and 17 percent is "due to nonproportional change, or interaction" (1966:74). In our analysis, 97.5 percent of the association in the baseline model is due to temporally constant patterns of association between father's and son's occupations. In our view, this gives a stronger basis to Duncan's conclusion (1966:96) that

> some considerable modifications of the mobility *pattern* (emphasis in the original) . . . occurred in consequence of the change in 'structure' represented by alterations of the frequency distributions of origin and destination classes.

Other Data on Mobility Trends

The analyses reported above do not exhaust the possibilities for measuring trend in occupational mobility among U.S. men. For example, the Johns Hopkins University sample of men aged 30 to 39 in 1968 might be compared with the cohort of that age in the 1962 OCG survey (Coleman et al., 1972). Data from the Six-City Survey of Labor Mobility might be used to contrast mobility patterns in 1940 and 1950, and it may be possible to construct a time-series of data from the Detroit Area Study (Duncan, 1968:697-703). Comparisons might be made between certain male OCG respondents and those in the two male panels of Herbert Parnes' National Longitudinal Surveys (Parnes, Miljus and Spitz, 1970; Parnes, Fleisher, Miljus and Spitz, 1970). These examples probably do not exhaust the resources presently available, and another important test of our hypothesis will become possible when the 1973 OCG data from the national survey arrive (Featherman and Hauser, 1975).

While the data listed above are available in published or machine-readable form, we have not exploited any of them in the present analysis. Rather, we made two other comparisons, neither of which pertains strictly to temporal change, but both of which bear on the extension of our finding of invariant mobility patterns. First, we compared mobility of (OCG) men in the U.S. in 1962 to the mobility of Wisconsin men in 1973, using data from a statewide survey commissioned by Featherman and Hauser (1975). Second, we compared the mobility of U.S. men in 1962 from father's occupation to son's current occupation, using the same nine five-year cohorts as in our earlier analysis of mobility to first full-time occupations. While both of these comparisons are sound with regard to survey method, the first confounds time and place, and the second confounds time (cohort) with chronological age. We carried out these two comparisons using the same methods as in the analysis of Table 1, with the same results. That is, once controlling for change in the occupational structure, we found no change in occupational mobility.

Discussion

The foregoing analyses apply an appropriate solution to the standard problem in comparative mobility analysis–how to separate the effects of changes of occupational structure, given by shifting marginal distributions, from those of the process of mobility, given by associations between occupational origins and destinations, for the purpose of measuring change in the latter. Analytically, the solution is to adopt the multiplicative or log-linear specification of the frequencies, which posits separate parameters for the marginal distributions and changes therein and for the underlying associations and changes therein. In several large bodies of data on U.S. men, we have observed that the pattern of association between father's occupation and son's occupation is largely invariant with respect to time.

It is always a logically difficult matter to maintain the null hypothesis, but we think we have offered sufficient evidence of temporal invariance to challenge the ingenuity of other researchers who would offer a more plausible alternative hypothesis and supporting data. One simple model of change we have not treated explicitly is linear trend; but in scanning reams of computer output, we have seen little evidence of order in deviations from the model of temporal homogeneity. We did locate one trend across OCG cohorts to increase the ratio of actual to expected chances of mobility from father's occupation to son's first full-time occupation. This trend accounts for about half the variation in mobility propensity across nine OCG cohorts, but the total of that variation represents less than a third of one percent of the association in the baseline model (see line E3 of Table 1). Other models may capture systematic patterns of change in mobility with relatively few parameters, but the explanatory potential of such models is necessarily limited.

We have used our imagination in specifying models of change in mobility, but beyond noting that the OCG data do not conform to the hypothesis of quasi-perfect mobility, the present analysis says little about the pattern of occupational mobility at any one time. That is an important, but distinct, problem. For alternative models of the mobility table, see Goodman (1972b) and Haberman (1974a).

Among serious students of mobility in the United States there has been agreement of late that trends, if any, have been slight. Thus, one might ask what motivates so long and tedious an effort to sustain the hypothesis of no trend. We think the present essay is amply justified by the variety of hypotheses about trend which we have been able to test. However, we believe it equally important that we have eliminated trends in the occupational structure from our measurements of trend in mobility.

Sociologists have long recognized that the occupational structure and changes in it affect mobility patterns. This recognition has generally taken the form of injunctions to control variation in the occupational structure before venturing comparisons between mobility tables and of efforts to construct mobility indexes which would make this possible. However, once trends in the occupational structure are controlled, there are no trends in the occupational mobility of U.S. men. This suggests an inversion of the traditional problem of comparative mobility analysis.

Rather than treating shifts in the occupational structure as a nuisance factor, to be set aside before undertaking comparative mobility analysis, we suggest that shifts in the occupational structure may be both the driving force and the problematic issue in comparative mobility studies. We ought to be asking what changes in observed mobility chances may be induced by transformations of the occupational structure, such as occur in the processes of urbanization and industrialization. That is, even when the relative chances of men (as defined by our model) do not change, it is possible for other interesting properties of the mobility tables to vary systematically with changes in the marginal distributions (Duncan, 1966:76-7). Moreover, if changing occupational mixes affect mobility patterns, students of mobility will want to take more than a casual interest in the sources of occupational transformations. We believe our analysis adds force to Wilbert E. Moore's (1966:196) observation:

> the analytical separation of mobility accounted for by changes in the distribution of occupation within the socioeconomic structure from that accounted for by change in the distribution of opportunity or accessibility is a perfectly legitimate procedure. But there is no reason to say that only the second datum is interesting.

In the sequel to this paper (Hauser et al., 1975), we shall attack the first of the problems just posed. Given a fixed pattern of mobility from father's occupation to son's current occupation, we shall ask what the implications are of changes in the marginal distributions of father's and of son's occupations. In discussing his analysis of proportional change in the Indianapolis data, Duncan wrote (1966:75), "Had only the 1910 table been available, plus the 1940 marginals, the assumption of proportional changes would have been a reasonable basis for estimating the 1940 table." Applying a well-known technique of proportional adjustment (Mosteller, 1968), we have taken tables of U.S. men's mobility from father's occupation to current occupation for certain cohorts in 1962, and we have projected them backward and forward in time to fit earlier and later occupation distributions, while preserving the underlying patterns of association in the mobility tables. Several properties of the resulting hypothetical mobility tables vary systematically across years both within and between cohorts. Changes in the occupational structure and the succession of cohorts have led to increased rates of mobility and of upward mobility, even though the underlying process of mobility has not changed over time.

REFERENCES

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

Boudon, Raymond
1973 Mathematical Structures of Social Mobility. San Francisco: Jossey-Bass.

Coleman, James S., Zahava D. Blum, Aage B. Sørensen and Peter H. Rossi
1972 "White and black careers during the first decade of labor force experience. Part I: occupational status." Social Science Research 1:243-70.

Davis, James A.
1974 "Hierarchical models for significance tests in multivariate contingency tables: an exegesis of Goodman's recent papers." Pp. 189-231 in Herbert L. Costner (ed.), Sociological Methodology, 1973-1974. San Francisco: Jossey-Bass.

Deming, W. Edwards
1943 Statistical Adjustment of Data. New York: Wiley.

Duncan, Otis Dudley
1961 "A socioeconomic index for all occupations." Pp. 109-38 in A. J. Reiss, P. K. Hatt and O. D. Duncan, Occupations and Social Status. New York: Free Press.
1965 "The trend of occupational mobility in the United States." American Sociological Review 30:491-8.
1966 "Methodological issues in the analysis of social mobility." Pp. 51-97 in Neil J. Smelser and Seymour Martin Lipset (eds.), Social Structure and Mobility in Economic Development. Chicago: Aldine.
1968 "Social stratification and mobility: problems in the measurement of trend." Pp. 675-719 in Eleanor Bernert Sheldon and Wilbert E. Moore (eds.), Indicators of Social Change. New York: Russell Sage.

Featherman, David L. and Robert M. Hauser
1974 "Trends in occupational mobility by race and sex in the United States, 1962-1972." Working Paper 74-25, Center for Demography and Ecology, The University of Wisconsin–Madison.
1975a "Design for a replicate study of social mobility in the United States." Pp. 219-51 in Kenneth Land and Seymour Spilerman (eds.), Social Indicator Models. New York: Russell Sage.

Goodman, Leo A.
1965 "On the statistical analysis of mobility tables." American Journal of Sociology 70:564-85.
1968 "The analysis of cross-classified data: independence, quasi-independence, and interactions in contingency tables with or without missing entries." Journal of the American Statistical Association 63:1091-131.
1969a "How to ransack social mobility tables and other kinds of cross-classification tables." American Journal of Sociology 75:1-39.
1969b "On the measurement of social mobility: an index of status persistence." American Sociological Review 34:831-50.
1970 "The multivariate analysis of qualitative data: interactions among multiple classifi-

cations." Journal of the American Statistical Association 65:226-56.
1971 "The analysis of multidimensional contingency tables: stepwise procedures and direct estimation methods for building models for multiple classifications." Technometrics 13:33-61.
1972a "A general model for the analysis of surveys." American Journal of Sociology 77:1035-86.
1972b "Some multiplicative models for the analysis of cross-classified data." Pp. 649-96 in Proceedings of the Sixth Berkeley Symposium on Mathematical Statistics and Probability. Berkeley: University of California Press.

Haberman, Shelby J.
1974a The Analysis of Frequency Data. Chicago: University of Chicago Press.
1974b "Log-linear models for frequency tables with ordered classifications." Biometrics 30:589-600.

Hauser, Robert M. and David L. Featherman
1973 "Trends in the occupational mobility of U.S. men, 1962-1972." American Sociological Review 38:302-10.
1974a "Socioeconomic achievements of U.S. men, 1962-1972." Science 185:325-31.
1974b "White-nonwhite differentials in occupational mobility among men in the United States, 1962-1972." Demography 11: 247-65.

Hauser, Robert M., David L. Featherman and Dennis P. Hogan
1974 "Sex, age and race in the structure of occupational mobility in the United States, 1962." Presented at the Eighth World Congress of Sociology, Toronto (August).

Hauser, Robert M., John N. Koffel, Harry P. Travis and Peter J. Dickinson
1975 "Structural changes in occupational mobility among men in the United States." To appear in the American Sociological Review.

Hope, Keith
1972 "Quantifying constraints on social mobility: the latent hierarchies of a contingency table." Pp. 121-90 in Keith Hope (ed.), The Analysis of Social Mobility: Methods and Approaches. Oxford Studies in Social Mobility, Working Papers 1. Oxford: Clarendon Press.
1974 "Trends in the openness of British society in the present century." Presented at the Mathematical Social Sciences Board Conference on Measurement and Models in Social Stratification. Toronto (August).

Jackson, Elton F. and Harry J. Crockett, Jr.
1964 "Occupational mobility in the United States: a point estimate and trend comparison." American Sociological Review 29:5-15.

Kish, Leslie and Martin R. Frankel
1974 "Inferences from complex samples." Journal of the Royal Statistical Society, Series B. 1-22.

Koffel, John N.
1974 Subjectivity in Mobility Trend Literature: Sociology of Knowledge and Statistical Approaches. Unpublished M. A. thesis. University of Wisconsin–Madison.

Ku, Harry H. and Solomon Kullback
1974 "Loglinear models in contingency table analysis." The American Statistician 28:115-22.

Moore, Wilbert E.
1966 "Changes in occupational structures." Pp. 194-212 in Neil J. Smelser and Seymour Martin Lipset (eds.), Social Structure and Mobility in Economic Development. Chicago: Aldine.

Mosteller, Frederick
1968 "Association and estimation in contingency tables." Journal of the American Statistical Association 63:1-28.

National Opinion Research Center
1972 National Data Program for the Social Sciences, Codebook for the Spring 1972 General Social Survey. University of Chicago (July).

Parnes, Herbert S., Belton M. Fleisher, Robert C. Miljus and Ruth S. Spitz
1970 The Pre-Retirement Years: A Longitudinal Study of the Labor Market Experience of Men, Volume 1 (1966 survey). Manpower Administration, U.S. Department of Labor, Manpower Research Monograph No. 15.

Parnes, Herbert S., Robert C. Miljus and Ruth S. Spitz
1970 Career Thresholds: A Longitudinal Study of the Educational and Labor Market Experience of Male Youth. Volume 1 (1966 survey), Manpower Administration, U.S. Department of Labor, Manpower Research Monograph No. 16.

Rogoff, Natalie
1953a Recent Trends in Occupational Mobility. New York: Free Press.
1953b "Recent trends in urban occupational mobility." Pp. 442-54 in Reinhard Bendix and Seymour Martin Lipset (eds.), Class, Status, and Power: A Reader in Social Stratification. New York: Free Press.

Taeuber, Karl E. and Alma F. Taeuber
1965 Negroes in Cities: Residential Segregation and Neighborhood Change. Chicago: Aldine.

Tully, Judy Corder, Elton F. Jackson and Richard F. Curtis
1970 "Trends in occupational mobility in Indianapolis." Social Forces 49:186-200.

Tyree, Andrea
1973 "Mobility ratios and association in mobility tables." Population Studies 27:577-88.

White, Harrison C.
1963 "Cause and effect in social mobility tables." Behavioral Science 7:14-27.