LONG TERM LINEAR TRENDS IN STATUS ATTAINMENT IN ITALY

Harry B. G. Ganzeboom

Donald J. Treiman

Last Revised: March 31, 1993

Please cite this paper as:

Ganzeboom, Harry B.G. & Donald J. Treiman (1993). Long Term Linear Trens in Status Attainment in Italy. Nijmegen: Department of Sociology Nijmegen University [Working Paper].

An Italian translation of this paper has been published in a volume in memory of Prof. Paolo Ammassari:

Ganzeboom, Harry B.G.; Treiman, Donald J. (1996). "Tendenze lineari di lungo periodo nel conseguimento di status in Italia [Long Term Linear Trends in Status Attainment in Italy]". Pp. 187-219 in Carlo Mongardini (Ed.), "Teoria sociologica e stratificazione sociale [Sociological Theory and Social Stratification]". Roma: Nuova Italia Scientifica.

ABSTRACT

In this article we analyze six sample surveys that contain information on intergenerational mobility in Italy and estimate linear trends in the determinants of educational and occupational attainment for birth cohorts and labor market entry cohorts that together cover most of the 20th century (1904-1985). By pooling data from surveys that were carried out over a period of 24 years and applying linearly restricted regression equations, we obtain a statistically powerful model of Italy's stratification regime in this period. We find that the effect of social origins (father's occupational status) on educational attainment has significantly <u>in</u>creased for women, but not for men. With respect to occupational attainment, the trend for Italy is in line with that found in other countries: education has become more important and parental status (again measured by father's occupational status) has become less important. When these trends are studied for the non-farm population, the increasing importance of education continues to hold but the declining importance of father's occupation does not.

INTRODUCTION

Paolo Ammassari's promise as a student of social stratification, alas, remained unfulfilled at the time of his untimely death. He was very active in the 1970's, playing a prominent role in the reactivation of the ISA Research Committee on Social Stratification and Mobility after a period of dormancy in the 1960s; hosting one of the first meetings in the modern era of the Committee, the 1973 meeting in Rome (Featherman, Hauser, and Sewell, 1974); and organizing a major empirical study of social mobility in Italy, a national sample survey of about 3,500 persons. Unfortunately, very little on mobility has appeared from this survey. We know only of a single unpublished paper presented at the Dublin meeting of the ISA Research Committee on Social Stratification (Ammassari, 1977); however, the mobility tables he presented at the Dublin meeting have been used in several cross-temporal and international comparisons (e.g., Heath, 1983; Ganzeboom, Luijkx, and Treiman, 1989).[1]2 After a number of years in which he was distracted from the study of social mobility by other obligations, Prof. Ammassari had planned to return to this topic, and even after the onset of his final illness had hoped to organize a volume of papers on methods and techniques for studying social stratification and social mobility.[2] It is to be hoped that the data from the Ammassari survey can be deposited in a data archive and in that way others can carry on Prof. Ammassari's work.

For students of social stratification, Italy represents an extremely interesting case, since it is simultaneously a major industrial country, with a labor force similar to those of the countries of North West Europe,[3] and, until recently, a quite traditional country with respect to its educational system, which was one of the most unequal in Europe, with a fraction of the population as highly educated as anywhere in the world and another fraction obtaining less education than in virtually any Western European country.[4] In this contribution we address the question of how distinctive Italy is with respect to the process of status attainment by studying long term trends in the determinants of educational attainment and occupational status attainment, using data from six surveys conducted over a period of 24 years.

DATA

Our data[5] derive from six general population sample surveys, which all contain information on education, occupation, and father's occupation. Oddly, to our best knowledge four of these six surveys have never before been used by Italian scholars (or others) to study the stratification regime of Italy.[6] Of the two that have been used, the best known among Italian researchers is the 1985 "Study of Social Mobility and Education in Italy" conducted by a team of five: Barbagli, Capecchi, Cobalti, de Lillo, and Schizzerotto [ITA85].[7] The data collected by Lopreato in 1963/64 [ITA63] are also well known, since Lopreato published an intergenerational occupational mobility table in the American Journal of Sociology (Lopreato, 1965; see also Lopreato and Hazelrigg, 1972:376-390), which refuted the claim of Lipset and Bendix (1959) that Italy was more rigid than other industrialized countries (however see our results below). Lopreato's data have not been included in the collections of any of the major data archives; however, they have been preserved by the Data Program and Library Service of the University of Wisconsin-Madison, from which we obtained them[8]. To the best of our knowledge, the two election surveys of Barnes (1968) [ITA68] and Barnes and Sani (1972) [ITA72] have never before been used in stratification analysis, although they have been accessible through the major U.S. data archive, the Inter-university Consortium for Political and Social Research [ICPSR], for a long time. The fifth data set [ITA75p] was collected in 1974 by Sartori and Marradi (in association with Sani) for Barnes and Kaase's (1979) Political Action project. The sixth data set [ITA87] was collected by Calvi and Anselmi as the Italian contribution to the 1987 module ('Social Inequality') of the International Social Survey program [ISSP].

All six data sets utilize a nearly uniform coding system for educational attainment, which we have converted into a metric of years of education:[9]

Educational level	Years of education

None, illiterate	0
Scuola elementare non conclusa	4
Scuola elementare con licenza	5
Scuola media inferiore non conclusa	7
Scuola media inferiore con licensa	8
Scuola media superiore non conclusa	11
Scuola media superiore con dipploma	13
Universita, ma non laurea	15
Laurea	18
Post Laurea	20

By contrast, the six data sets differ considerably with respect to the way occupations are classified. Three files [ITA63, ITA68, ITA72] utilize various coarse codes, which differ substantially across studies. [ITA85] uses a fairly detailed (97 category) code and [ITA75p] uses a highly detailed (three-digit) code for both father and respondent. Finally, [ITA87] uses _different_ coarse codes for fathers and respondents. In order to convert these diverse codes into a standard metric, we matched each category in each classification to a category of the 1968 International Standard Classification of Occupations [ISCO] (International Labor Office, 1969; Treiman, 1977) and then converted it to the International Socio-Economic Index of occupational status [ISEI] (Ganzeboom, De Graaf, and Treiman, 1992), which ranges from 10 (for "Farm Laborers") to 90 (for "Judges").

From each file we selected all men and women who were between 25 and 64 years of age at the time of the survey and who had valid responses to all variables in the analysis. The age specification was imposed to avoid selection bias due to the omission of those still in school or those already retired. In three of the six data sets respondent's age is recorded in broad, 10-year intervals. Since this is inconvenient in analysis that uses cohorts and cohort characteristics, we added a uniformly distributed random jitter to the age variables in these data sets in order to distribute cases evenly throughout each interval.

Descriptive statistics for each of the six files are shown in Table 1.

However, we have not analyzed these files separately, but rather have pooled all six samples into a single file, which is used for all of our analysis; we do, however, carry out separate analyses of men and women. Inspecting the statistics in Table 1, it is reassuring to note that most of the differences across studies are either quite small or are fairly orderly, following clear time trends. Thus, we can have reasonable confidence in the quality of the data.

METHODS AND MODELS

In order to evaluate historical trends in status attainment, we estimate an elementary status attainment model involving only three status variables: father's occupation and respondent's education and occupation. We first assess trends in the determinants of educational attainment and then assess trends in the determinants of occupational status.

Educational attainment

For educational attainment, the kernel of this model is a simple regression equation of the form:

$$\hat{EDU} = b_0 + b_1*FIS + b_2*BYR \tag{1}$$

where EDU = years of school completed; FIS = father's occupational status (measured by the ISEI); and BYR = year of birth.

Historical trends are estimated by adding a multiplicative interaction between father's occupational status and year of birth:

$$\hat{EDU} = b_0 + b_1*FIS + b_2*BYR + b_3*FIS*BYR \tag{2}$$

In this type of model, the main terms for the two independent variables cannot be evaluated independently since the effect of each depends on the level of the other. Precisely, b_1 refers to the effect of father's occupation when BYR=0 whereas b_2 refers to the effect of year of birth on level of educational attainment when FIS=0. But these statements are not very meaningful, since both points are outside the ranges of these variables.

In addition to problems of interpretability, interaction models such as (2) suffer from two other difficulties. First, there is often multicollinearity (extremely high correlation) between the main terms and the interaction terms, which may lead to numerical problems in estimation. Second, since FIS and BYR are correlated, in Model (2) the effect of father's occupational status is confounded by the effect of year of birth.

In order to avoid these complications, it is convenient to center one independent variable within the other. This is accomplished by subtracting the within-cohort means:

$$FIS' = FIS - MFIS_c \tag{3}$$

where $MFIS_c$ = the mean FIS score for those in cohort c. We used 5-year birth cohorts to calculate the means.[10] This transformation solves all three problems discussed above. First, centering father's occupational status within cohorts removes the correlation between the two main effects, and its potentially confounding consequences. Second, the main effect for birth year now refers to those who have the average father's ISEI score (FIS) of those within the same five-year birth cohort. Third, centering mitigates multicollinearity between main terms and interaction terms.

Note that transformation (3) is <u>non-linear</u>, since we subtract within-cohort means, not the overall mean. The effect of this transformation is that the model becomes similar (but not identical) to a procedure in which the bivariate regression of years of education on father's occupation is estimated cohort-by-cohort. The difference is that our specification constrains cohort differences in the effect of father's occupational status on educational attainment to be linear, a constraint that would be relaxed in a cohort-by-cohort analysis. In our specification, we retain the advantage of a one-parameter test of the hypothesis that the relationship between father's occupation and educational attainment has changed over time, although at the cost of the simplifying assumption that any trend is linear.

In order to make the model still more interpretable we take two additional steps, both of which are <u>linear</u> transformations. First, we subtract a constant from the year of birth, in order to move the main effect coefficient for father's occupational status (FIS) to within the range of the data. In other contexts, it would be appropriate to subtract the sample mean of year of birth. In that case both main effects would refer to the mean of the data, which would have the advantage of rendering them identical in (1) and (2). However, in the data that we analyze here, the mean year of birth differs between men and women. In this situation it is more convenient to choose a fixed point in time. We choose (19)30 for the analysis of educational attainment, which is close to the average year of birth:

$$BYR' = BYR - 30 \tag{4}$$

Second, the multicollinearity between main terms and interaction terms is dependent upon the differences in range of the constituent components. To minimize the risk of multicollinearity problems, we shrink the standard deviations of BYR' and FIS' by dividing them by 10, on order to bring them into approximately the same range as the standard deviation for education.

$$BYR'' = BYR'/10 \quad \text{and} \quad FIS''=FIS'/10 \tag{5}$$

Note that transformations (5) are linear transformations that affect the estimated parameters only in case of numerical problems. They have the additional advantage of making the parameter estimates somewhat more readable.

If we integrate the several steps into one model, we have:

$$\hat{EDU} = b_0 + b_1{*}FIS'' + b_2{*}BYR'' + b_3{*}(FIS''){*}(BYR'')$$

or

$$\hat{EDU} = b_0$$
$$+ b_1*((FIS-MFIS_c)/10)$$
$$+ b_2*((BYR-30)/10)$$
$$+ b_3*((FIS-MFIS_c)/10)*((BYR-30)/10) \quad (6)$$

Note that b_0 gives the expected education for a person born in 1930 whose father's occupational status (FIS) is at the average for the 1928-1932 birth cohort. For practical purposes such a respondent can be regarded as the 'average respondent' in our data. b_1 gives the effect of father's occupation for those born in 1930, where the effect is measured as years of education per 10 points of FIS. b_2 gives the (linear) effect of educational expansion over time--precisely, the expected difference in average years of schooling for persons born 10 years apart whose fathers have the mean occupational status of their cohort. b_3 gives the (linear) change in the effect of father's occupational status on educational attainment over time (or, alternatively, the temporal trend in educational attainment for those at varying levels of father's occupational status). This is the coefficient of greatest interest with respect to the analysis of trends.

Occupational Attainment

Our model of occupational attainment is specified following a similar logic. However, the kernel of the model now has three independent variables:

$$\hat{ISE} = b_0 + b_1*FIS + b_2*EDU + b_3*EXP \quad (7)$$

where ISE is the respondent's occupational status, measured by the ISEI; EXP stands for years of experience; and the two remaining variables are defined as above. Since we have no direct measure of labor market experience, this term is approximated (in the conventional way) as:

$$EXP = AGE - EDU - 6 \quad (8)$$

that is, the years elapsed since the respondent left school (on the assumption that he or she started school at age six and continued uninterrupted). This measure, which we hereafter refer to as "seniority," approximates labor force experience to the extent that the respondent worked continuously since leaving school. The approximation is therefore much better for men than for women, who tend in Italy, as elsewhere, to interrupt their careers for childbearing and rearing. Moreover, in Italy many women leave the labor force permanently after marriage (Durand, 1975:190-191), so that women with high seniority are a highly selected set of all women. Therefore, the coefficients for seniority should be interpreted with caution, especially for women. However, they will effectively control for career effects in occupational attainment of whatever nature; the inclusion of an experience term makes it possible to interpret the coefficients of the other two terms (father's occupational status and respondent's education) as the effect of these endowments at the beginning of the respondent's career, that is, at the point of school leaving and labor market entry.

In order to capture career dynamics more fully, we add interactions between experience and the two other main terms. Our model then becomes:

$$\hat{ISE} = b_0 + b_1*EXP + b_2*FIS + b_3*EDU + b_4*FIS*EXP + b_5*EDU*EXP \quad (9)$$

The two interaction terms model changes in the effect of father's occupational status and respondent's education over the respondent's career. We expect the effect of father's occupational status to decline over the life-cycle; for education, the expectation is less obvious, since well educated people tend to experience more career advancement than poorly educated people (Sørensen and Tuma, 1981:84).

We then assess historical trends in these effects by comparing labor market entry cohorts. The pertinent variable is constructed as:

$$ENT = BYR + 6 + EDU \quad (10)$$

Experience and labor market entry year are thus reparameterizations of age, year of survey, and years of education. Expressed in the vernacular of age-period-cohort analysis (Mason, et al., 1973), we assume that there are no period (i.e., survey) effects, but that occupational attainment can be understood completely in terms of cohort and age (life cycle) effects.[11]

To evaluate historical trends in occupational attainment, we add to our equation a term for year of labor force entry (ENT) and the corresponding interaction terms:

$$\begin{aligned} I\hat{S}E = {} & b_0 + b_1{*}EXP + b_2{*}FIS + b_3{*}EDU \\ & + b_4{*}FIS{*}EXP + b_5{*}EDU{*}EXP \\ & + b_6{*}ENT \\ & + b_7{*}FIS{*}ENT + b_8{*}EDU{*}ENT \end{aligned} \tag{11}$$

As before, a model thus specified yields coefficients that are pertinent outside the range of data points, and has the additional disadvantage of being vulnerable to numerical estimation problems. Therefore, as in our model of educational attainment, we center father's occupation and respondent's education within five-year labor market entry cohorts. We also center year of labor force entry around (19)40, following the same logic as before. However, we do not center EXP since doing so would render the associated coefficient difficult to interpret: by controlling seniority, we evaluate the effect of father's occupation and respondent's occupation at the beginning of the career. Again, the variances of all variables are converted to the same order of magnitude by dividing FIS, ENT, and EXP by 10. Since the dependent variable is measured in the same metric as FIS, it is convenient to divided ISE also by 10. Our final model for occupational attainment thus becomes:

$$\begin{aligned} I\hat{S}E/10 = {} & b_0 \\ & + b_1{*}(EXP/10) \\ & + b_2{*}((FIS-MFIS_c)/10) \\ & + b_3{*}(EDU-MEDU_c) \end{aligned}$$

$$+ b_4*((FIS-MFIS_c)/10)*(EXP)/10)$$

$$+ b_5*(EDU-MEDU_c)*(EXP/10)$$

$$+ b_6*((ENT-40)/10)$$

$$+ b_7*((FIS-MFIS_c)/10)*((ENT-40)/10)$$

$$+ b_8*(EDU-MEDU_c)*((ENT-40)/10)$$

or

$$\hat{ISE}' = b_0 + b_1*EXP' + b_2*FIS'' + b_3*EDU'' + b_4*FIS''*EXP' + b_5*EDU''*EXP' + b_6*ENT' + b_7*FIS''*ENT' + b_8*EDU''*ENT' \qquad (12)$$

where EXP' = EXP/10; FIS" = (FIS-$MFIS_c$)/10; EDU" = EDU-$MEDU_c$; and ENT' = (ENT-40)/10). In Eq. (12), b_0 (*10) gives the expected ISEI score at the point of labor force entry for a person entering the labor force in 1940 whose father's occupational status (FIS) and own education (EDU) are at the average for the 1938-1942 entry cohort. b_1 gives the effect of seniority (EXP) among those with the average education and average father's ISEI score of the labor force entry cohort. b_2 and b_3 give the effects of father's occupation and years of education on occupational status at the beginning of the career for those entering the labor force in 1940. Note that because of the rescaling of FIS, EXP, and ISE, but not of EDU, b_1 and b_2 can be interpreted directly, as the effect of one year difference in seniority or one point difference in father's occupational status or a on the respondent's ISEI score, but b_3 must be multiplied by 10 to put it in the same metric as the original variables. b_6 gives the (linear) effect of the upward shift of the occupational distribution over time--precisely, the expected difference in the average ISEI score for persons born in successive years whose fathers' have the mean occupational status of their cohort and who themselves have the mean years of schooling of their cohort, but no seniority. The interaction term b_4 gives the (linear) change in the effect of father's occupational status on occupational attainment resulting from each additional 10 years of seniority (or, alternatively, the change in the effect of seniority on occupational status for those who vary by 10 points in father's occupational status). b_5

gives the (linear) change in the effect of educational attainment on occupational attainment for each additional year of seniority (or, alternatively, the change in the effect of seniority on occupational status resulting from each additional year of schooling). Similarly, b_7 gives the difference in the effect of father's occupational status for labor force entry cohorts 10 years apart and b_8 gives the change in the effect of educational attainment on occupational attainment for successive entry cohorts; alternatively, these coefficients give the temporal trend in occupational attainment for those at varying level of education and father's occupational status. b_7 and b_8 are the coefficients of greatest interest with respect to the analysis of trends.

RESULTS

Tables 2 and 3 display the results of our regression analyses, separately for men and women. In Table 2 we see the results for models of educational attainment, and in Table 3 the results for models of occupational attainment. Appendix A contains the descriptive statistics used to estimate these models.

Educational attainment

Model (a) of Table 2 shows estimates for Eq. (1), in which the effect of father's occupation on educational attainment is not allowed to vary over time, while Model (b) shows the coefficients for Eq. (6) as presented above. In the present case, our main interest is in Model (b), the point of presenting Model (a) being simply to show that when independent variables are centered, the coefficients change hardly at all when interaction terms are introduced. Turning to Model (b), note first the coefficient of the intercept (b_0), which indicates that men born in 1930 whose fathers' had the average occupational status of their cohorts would be expected to have about 7 1/2 years of schooling on average, while women would be expected to have just under 6 1/2 years of schooling.

b_1 gives the effect of father's occupational status for those born in 1930. Recall that because of the way FIS was scaled, the effect is measured as expected years of education per 10 points of FIS. For the 1930 cohort, the effect of father's occupation was very strong: each 10 points of FIS was worth about 1.3 years of schooling for men and about 1.1 year of schooling for women. Since the total range in ISEI is 80 points (from 10 for farm laborers to 90 for judges), the difference in expected years of education between men born in 1930 into the lowest and the highest rungs in the Italian stratification system is something over 10 years of schooling (precisely, 10.26 = 1.282*80/10); for women, the difference is also large, about eight years (precisely, 8.70 = 1.088*80/10). Our impression is that the effect of father's occupational status on educational attainment was much stronger for the 1930 birth cohort in Italy than it was for the corresponding cohort in most other industrialized countries. Of course, confirmation of this conjecture must await further analysis.

Next consider b_2, the coefficient for birth year, which gives the (linear) effect of educational expansion over time, for persons whose fathers' had the mean occupational status of their cohort. This coefficient, which is 1.0 for men and upto 1.1 for women gives the total educational expansion of the Italian population. Since year of birth is measured by decade, the coefficient implies that mean occupation has grown by 5 to 6 years over the 50 years we study. Since father's occupation is centered within cohorts, this number represents the total effect and is not confounded with expansions of the occupational structure of fathers.

Finally, b_3 gives the (linear) change in the effect of father's occupational status on educational attainment over time. As it turns out, there is <u>no</u> temporal trend for men, so that the very strong effect of father's occupational status on educational attainment remains the same for the most recent cohorts. For women the results are even more striking, because the effect of father's occupation on educational attainment has <u>increased</u> significantly over time. For the most recent cohort of women, those born around 1960, the coefficient for father's occupational status is 1.342 (= 1.012 + .110*30/10), which implies an expected difference in the educational attainment of women from the lowest and highest status origins of 10.7 years (= 1.342*80/10), or about the same as what it is for men. Projecting these coefficients backwards to the earliest cohort, women born around 1905, we see that for such women the difference in expected educational attainment for those from the lowest and highest status origins was then a little more than three years (precisely, 3.2 = [(1.012 + .110*25/10)*80/10]). So, over the period for which we have data, educational inequality between women from the lowest and highest status origins has <u>increased</u> by 7 1/2 years (precisely, 7.5 = 10.7 - 3.2).

In sum, as we noted above, Italy has a highly inegalitarian educational system; the degree of inequality with respect to years of schooling has been increasing substantially over time, but even more substantially for women than for men, as we see from the standard deviations in Table 1; and therefore (Treiman and Yip, 1989) inequality with respect to educational opportunity (as indicated by the effect of father's occupation on educational attainment) is very high in Italy and shows no evidence of abating; indeed, for women it has increased over the course of the 20th century.

Occupational status

Table 3 reports our analysis of occupational status attainment. Four models are estimated, separately for men and women, building up to the full model of Eq. (12)--which is shown here as Model (d) in each set. Since this analysis includes only those with valid occupation codes, the number of cases is reduced relative to Table 2, particularly for women.[12]

Model (a) refers again to the kernel of the stratification process and reports the effects of father's occupational status (relative to other members of the same labor force entry cohort), entry cohort, and seniority, averaged over the total data set. The main effect of father's occupation, b_2, is about .45 for men and is slightly smaller, about .40, for women. Since the standard deviations of father's and respondent's occupational status are similar, these coefficients can be interpreted as standardized regression coefficients, net of the confounding effects of entry year and seniority. They imply an average level of intergenerational occupational mobility during the middle years of the century that is neither exceptionally high nor exceptionally low compared to other Western European countries at this time. The large size of the coefficient associated with labor force entry year, b_6, indicates very rapid upward shifting of the occupational structure, such that a man would be expected to obtain an occupation about one-third of an ISEI point higher if he entered the labor force one year later (precisely, .364 ISEI points) and a woman would be expected to gain about half an ISEI point per year (precisely, .487 ISEI points). Thus, for example, if two men entered the labor force in 1930 and 1960, respectively, and they had similar seniority and were from similar social origins (measured by their father's (relative) ISEI), the man who entered the labor force in 1960 would be expected to hold a job with an ISEI score about 11 points higher than the man who entered the labor force in 1930 (precisely, 10.92 = .364*3*10). Note that 11 points is not trivial; it is larger than the difference between a university professor and a primary school teacher, or the difference between a carpenter and an unskilled construction laborer. For women, the difference is even larger, as we have noted: a woman entering the labor force in 1960 would be expected to hold a job about 15 points higher than a woman with similar seniority and social origins who entered the labor force in 1930 (precisely, 14.61 = .487*3*10). These results are, of course, just what we would expect from the sharp upward shift in the average occupational status of respondents, especially women, over time. The third variable in Model (a), seniority, has no effect on occupational status attainment for either men or women. As we shall see below, this reflects the fact that we have not yet taken account of the effect of education on occupational attainment.

Model (b) simply allows the effect of father's occupational status to vary with year of labor force entry and with seniority. None of the interaction terms is significant, for either men or women, and this model appears to be of little interest.

Model (c) is, however, informative. It differs from Model (a) by including educational attainment (measured relative to the average for the labor force entry cohort) as a determinant of occupational status. Its inclusion drastically changes all of the other coefficients in the model, which serves to remind us how much information is lost in analyses that are restricted to the explication of intergenerational occupational mobility tables: the variance explained by the models is increased by nearly half; the effect of father's occupational status is reduced by nearly half; and the effect of both entry year and seniority substantially increase. The increase in the effect of seniority is straight forward: within cohorts, seniority and education are negatively correlated, so that when education is omitted, the effect of seniority is suppressed. The increase in the effect of entry year follows from the negative correlation between entry year and seniority in our data: recent entry cohorts cannot have been in the labor force for a long time. The negative correlation between entry year and seniority can be vividly seen in Figs. 1-4, in which there is no data in the upper left and lower right hand cells (see also Appendix A, Table 2) .

As expected, educational attainment has a substantial effect on occupational status: for men, each additional year of schooling (relative to the mean for the cohort) results in an expected increase of nearly two ISEI points (precisely, 1.90; recall that the coefficients must be multiplied by 10 to bring them into the original ISEI metric), net of father's occupational status, seniority, and entry year; for women, the expected increase is something over two ISEI points (precisely, 2.34). However, since the model (c) results are averaged over entry year cohorts, they cannot tell us about the question of greatest interest: how does the effect of social origins and educational attainment on occupational status change over time in Italy? To answer this question, we must turn to Model (d).

Model (d) is analogous to Model (b) in Table 2, but is more complicated, since it include three basic determinants of occupational attainment: father's occupation, education and experience. We start with the intercept, b_0, which indicates the expected occupational status (ISEI score) for persons who entered the labor force in 1940 and have the average education and social origins (father's occupational status) of the 1938-1942 labor force entry cohort and no labor force experience.[13] Note that the expected ISEI score is a bit higher for men than for women: 34.8 compared to 30.9, which is consistent with the result for educational attainment in Table 2.

The coefficients b_2 and b_3 give the net effects of father's occupation and years of education (relative to the cohort averages) on occupational status <u>at the beginning of the career</u> for those entering the labor force in 1940. Interestingly, educational attainment appears to be unrelated to occupational attainment at the beginning of the career for women: the coefficient is -.06 (= -.006*10), compared to .61 for men. However, the size of these coefficients is misleading since they refer to points for which there are virtually no cases in our sample. Recall that we have restricted our analysis to those age 25-64; that the highest level of education in the data, "<u>post laurea</u>," is specified as the equivalent of 20 years of education; and that the labor force entry year is defined as "age - years of education - 6". Thus, only the handful of 25 or 26 year olds in our sample with post graduate ("<u>post laurea</u>") education would have zero seniority.

The effect of education increases substantially with seniority, as indicated by the positive values for both men and women of, b_5, the coefficient for the interaction term EXP'*EDU". For those in the 1940 labor force entry cohort with five years of seniority, the effects on ISEI of each year of education are, respectively for men and women, .76 (=10*(.061+.031*.5)) and .24 (=10*(-.006+.059*.5); for those with 15 years of seniority, they are 1.08 and .83; for those with 25 years of seniority, they are 1.39 and 1.42; and so on. Thus, by the height of the careers of those entering the labor force in 1940, each year of education (relative to the cohort average) was worth nearly a point and a half on the ISEI scale, net of social origins (see Figs. 1 and 2 for a graphical representation of the same point; in this figure the second row, which includes entry cohorts from 1925 through 1944, contains the entry cohort of 1940). While this may seem relatively unimpressive, it yields nearly a 10 point ISEI difference between male high school graduates and those with post-graduate degrees (precisely, 9.73 = 1.39*(20-13)) and an equal difference for women.

Still, the effect of education on occupational status attainment is relatively modest for the 1940 labor force entry cohort. Also, the effect of father's occupational status is relatively strong for this entry cohort, and for this effect there is little change over the course of the career. At the beginning of the career, each point increase in father's occupational status (relative to the cohort average) is worth something over a quarter of an ISEI point for men, net of education (precisely, .285), and something over a third of an ISEI point for women (precisely, .347). Compared to similar figures for other Western European countries, these effects are relatively strong. Moreover, they do not change much with seniority, since the coefficient b_4, for the interaction term EXP'*FIS', is significant neither for men nor for women.

Our real interest is not in the 1940 entry cohort, which is simply a convenient reference category, but in how occupational opportunities have changed over time. We start by noting that for both men and women there has been a very substantial upgrading of occupational opportunities: net of social origins, educational attainment, and seniority, each later year of labor force entry (b_6) results in about a half point increase in ISEI for men (precisely, .459) and nearly two-thirds of an ISEI point for women (precisely, .645). We have noted this point before, in our discussion of Model (a), but the coefficients are substantially larger in Model (d).

But the question of central interest is not simply whether people have better chances than their fathers but whether the basis of occupational achievement has changed. There is some evidence that the process of status attainment in industrialized countries is relatively more achievement-based and less ascription-based than in non-industrialized countries (Treiman and Yip, 1989); and there is some evidence that societies throughout the world have become more open over the course of the 20th century with respect to the intergenerational transmission of status (Ganzeboom, Luijkx, and Treiman, 1989). But there is yet relatively little evidence regarding shifts in the relative importance of achievement and ascription over time (but see De Graaf and Luijkx, 1992, for the Netherlands; De Graaf and Huinink, 1992, for Germany; and Featherman and Hauser, 1978:Ch. 5, for the U.S.). Here we are able to provide such evidence for Italy.

Inspecting b_7 and b_8, the interaction terms involving year of labor force entry, it is clear that in the course of the 20th century the process of occupational attainment in Italy has become increasingly more dependent on achievement (educational attainment) and increasingly less dependent on social origins (father's occupational status). Consider first the role of education, expressed by coefficient b_8. The returns to an additional year of education at the time of labor force entry are expect to differ by .041 of an ISEI point for men who enter the labor force ten years apart, and by .062 of an ISEI point for women. Thus, over the course of the 75 years for which we have data on labor force entry (from 1910 to 1985), the number of ISEI points gain expected per year of schooling (relative to the cohort average) increased by about three for men (precisely, 3.08 = .041*7.5*10) and by well over four for women (precisely, 4.65 = .062*7.5*10), which are very substantial increases.

As we have already noted, one difficulty in evaluating Model (d) is that seemingly reasonable projections often carry us beyond the range of our data. For example, were we to project the effect of education on occupational status at the beginning of the career to those who began their careers early in the century, we would show a negative effect for women, clearly a nonsensical result, and a misleading one as well since we have no data in this range. As a solution to this problem, we have graphed the net relationship between education and occupational status implied by Model (d) separately for 16 categories created by subdividing our sample into four labor force entry cohorts (ranging from entry years 1910-1924 to entry years 1965-1985) cross-tabulated by four seniority categories (those with less than 15 years since labor force entry, those with 15-27 years, those with 28-41 years, and those with 42 or more years). The graphs for males are shown as Fig. 1 and for females as Fig. 2. Figs. 3 and 4 (to be discussed below) show corresponding graphs displaying the net relationships between father's occupational status and respondent's occupational status implied by Model (d). In each of the four figures, each cell is identified by the midpoint of the range of labor force entry dates and years of seniority included in the cell. In each of the figures, cells were left blank if there were not sufficient cases to produce reliable estimates, and the sizes of the dots are proportional to the size of the sample with specified combinations of characteristics. Thus, it is immediately evident from inspection of the figures to which combinations of entry year and seniority the bulk of our data obtain.

Both of the figures relating educational attainment to occupational status show the same story--which is, of course, just the story implied by the coefficients of Model (d). If we read down each column, we see that the slopes relating education to occupational status systematically increase, which indicates that the effect of education increases over the course of the 20th century for successive labor force entry cohorts. If we read across each row, we see that the slopes also systematically increase, which indicates that the effect of education increases with seniority. In both cases, the increases are steeper for women than for men, which reflects the larger size of the interaction terms involving seniority and labor force entry year for women than for men.

In the same way, we can visually examine the impact of entry year and seniority on the (net) relationship between father's occupational status and respondent's occupational status. Before turning to the graphs, however, we note that the coefficient b_7, for the interaction between labor force entry year and father's occupational status, is significantly negative--which is just what we would expect under the hypothesis that over the course of the century the Italian stratification system has moved away from an ascriptive basis of occupational allocation. Evaluating the impact of this change over the same 75 year period as before, we note that the impact of fathers' occupational status on the occupational status of their sons decreased by about a third of an ISEI point (precisely, -.383 = -.051*75/10) per point of fathers ISEI and the impact on the occupational status of daughters decreased by about half an ISEI point (precisely -.548 = -.073*75/10). Again, these are large decreases. Consider, for example, two fathers, one a secondary school teacher (ISEI = 63) and the other a bookkeeper (ISEI = 45). At the beginning of the period under study (the early years of the century), we would expect the difference in the status of their sons, net of education, to be nearly seven points--about half a standard deviation--larger than at the end of the period we are studying (precisely, -6.89 = (-.383)(63-45)); and for the daughters of two such men the expected difference would be even larger, nearly ten points (precisely, -9.86 = (-.548)(63-45)).

This pattern is very clearly revealed in Figs. 3 and 4. Reading down each column, we see that the slopes relating father's occupational status to occupational status systematically increase, which is exactly as expected from what we just discussed. However, reading across each row, we see that there is virtually no change in the slopes, which is of course implied by the very small size of b_4, the coefficient for the interaction term EXP'*FIS", for both men and women.

To summarize: it is clear that, taking the country as a whole, the Italian status system has shifted substantially away from ascription and toward achievement over the course of the 20th century.

Non-farm labor force

There is one consideration that may give us pause, however. We know that over the period we are studying Italy shifted from a predominantly agricultural economy to a predominantly urban industrial economy.[14] This suggests the possibility that our results simply reflect differences in the stratification systems of agricultural and non-agricultural societies. To investigate this possibility, we replicate our occupational analysis for the non-farm population only. The results are shown in the second panel of Table 3.

As it turns out, it does indeed appear to be the case that the reduction in the effect of father's occupation on respondent's occupation is primarily due to the shift of the labor force out of agriculture. Consider Model (d). Although virtually all the remaining coefficients remain relatively unaffected when the analysis is restricted to those not engaged in agriculture, the two non-trivial coefficients involving father's occupation, b_2 and b_6, are substantially reduced in size, and b_7, the coefficient expressing trends in the relationship between father's occupation and respondent's occupation, becomes non-significant. Although this coefficient remains negative for both men and women, it becomes quite small when the analysis is restricted to the non-farm population. Thus, in Italy, as elsewhere, it appears that the primary determinant of the shift toward greater societal openness is the movement of the population off the land.

On the other hand, the increase in the effect of educational attainment on occupational status is not affected at all by the restriction of the analysis to those not engaged in agriculture; in fact, the non-farm coefficients are slightly larger than those for the population as a whole. Hence, our initial conclusion, that the importance of achievement is increasing <u>relative to ascription</u> continues to hold up under the restricted analysis. So, in still another way, we come back to the same conclusion: over the course of the 20th century the process of status attainment has become more open in Italy.

NOTES

[1]The same survey did yield a new occupational prestige scale for Italy (Ammassari, 1978).

[2]Personal communication to Treiman, 13 December 1990.

[3]In 1971, 16 per cent of the Italian labor force was engaged in agriculture, about the same percentage as France although somewhat higher than in Germany and the United Kingdom, and the distribution of the non-agricultural labor force was quite similar to that of the other three countries (International Labor Office, 1977:Table 2B).

[4]Schooling was not compulsory in Italy until the 1950's (Haycraft, 1985:226). Barbagli (1973:16, quoted in Acquaviva and Santuccio, 1976:169), notes that "In 1950 about 30 per cent of the Italian population had either never learned or else had forgotten how to read and write. Italy had the highest illiteracy rate in Europe. At the same time, however, our country had the highest proportion of university students in its population when compared with Switzerland, Sweden, Germany, France, Great Britain, Holland and Belgium." Even in more recent years the variance in years of school completed has remained large in Italy relative to other Western European countries (Ganzeboom and Treiman, 1991).

[5] The data used in this article are all part of International Stratification and Mobility File (Ganzeboom & Treiman, 1992) that contains standardized extract files for surveys from around the world with social mobility information. Many colleagues and data archives have kindly helped us in collecting and processing this information. Special thanks are due to Elisabeth Stephenson, the project archivist.

[6]Indeed, the four are not even mentioned by Cobalti (1992) in his comprehensive review of Italian stratification research.

[7]We thank Antonio Schizzerotto and Luisa Saviori, of the University of Trento, for making these data available to us.

[8] Laura Guy (DPLS) was helpful in retrieving them.

[9]We thank Angela Prando and Cinzia Meraviglia for their advice on this matter.

[10]These cohorts are centered on decades and half-decades. For example, the (19)30 birth cohort includes persons born between 1928 and 1932, inclusive. Our oldest cohort is centered at (19)05 and our youngest at (19)60. This is the only point in our analysis where we treat cohorts in a discrete mode. In the remainder of the analysis we measure single years of birth (or single years of labor force entry date); the numerical value of our variable refers to the last two digits of the year, e.g., 30 for those born in 1930, and so on.

[11]Period effects in this instance could be either effects associated with specific historical events or effects that arise from differences in quality or measurement among the six surveys. The way we specify and interpret our model effectively assumes that such effects do not exist. We think this a reasonable assumption. Given the dates of the surveys, there is no reason to suspect historical effects; and, as we have noted, the coefficients in Table 1 give us confidence that there are no anomalous surveys among the six.

[12] Comparing the number of women available for the education and occupational attainment analyses, respectively (shown in the two panels of Table 1), is somewhat disturbing. In the [ITA75p] and [ITA85] files, more than 60 per cent of the women in the education analysis are also included in the occupation analysis, compared to less than 40 per cent in the other three files (recall that [ITA63] did not include women at all). This difference reflects the inclusion of information on the most recent job for women not currently working in [ITA75p] and [ITA85].

[13]The averages are computed separately for men and women.

[14]The proportion of the economically active population engaged in agriculture declined from 57 per cent in 1901 to 18 per cent in 1971 (Acquaviva and Santuccio, 1976:48) and has undoubtedly declined still further since then.

ACKNOWLEDGEMENT

This paper was prepared, in part, while the first author was a Visiting Scholar in the Department of Sociology, UCLA, and, in part, while the second author was a Fellow at the Center for Advanced Study in the Behavioral Sciences. His tenure at the Center was supported, in part, by National Science Foundation grant SES-9022192.

REFERENCES

Articles and books

Acquaviva, Sabino, and Mario Santuccio. 1976. Social Structure in Italy: Crisis of a System. Translated from the Italian by Colin Hamer. Boulder: Westview Press.

Ammassari, Paolo. 1977. "Occupational Mobility Trends in Italy." Paper presented at the meeting of the Research Committee on Social Stratification and Mobility, International Sociological Association, Dublin, April 5-7.

Ammassari, Paolo. 1978. La Soddisfazione Professionale, il Prestigio e le Tradizioniu Familiari. Rome: Instituto di Statistica Economica dell' Universita' de Roma.

Barbagli, Marzio. 1973. "Scuola e Mercato del Lavoro." Cited in Acquaviva and Santuccio, 1976.

Barnes, Samuel H., Max Kaase, et al. 1979. Political Action: Mass Participation in Five Western Democracies. Beverly Hills, CA: Sage.

Cobalti, Antonio. 1992. "Italian Mobility in an International Perspective: a Comprehensive View." Paper presented at the meeting of the Research Committee on Social Stratification and Mobility, International Sociological Association, Trento, May 14-16.

De Graaf, Paul M., and Johannes J. Huinink. 1992. "Trends in Measured and Unmeasured Effects of Family Background on Educational Attainment and Occupational Status in the Federal Republic of Germany." Social Science

Research 21:84-112.

De Graaf, Paul M., and Ruud Luijkx. 1992. "Van 'Ascription' naar 'Achievement'? Trends in Statusverwerving in Nederland tussen 1930 en 1980." Mens en Maatschappij 67:412-33.

Durand, John D. 1975. The Labor Force in Economic Development: A Comparison of International Census Data, 1946-1966. Princeton: Princeton University Press.

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

Featherman, David L., Robert M. Hauser, and William H. Sewell. 1974. "Toward Comparable Data on Inequality and Stratification: Perspectives on the Second Generation of National Mobility Studies." Current Sociology 2:383-97.

Ganzeboom, Harry B. G., and Donald J. Treiman. 1991. "Educational Expansion and Educational Achievement in Comparative Perspective." Paper presented at the meeting of ISA Research Committee 28, Social Stratification, Ohio State University, Columbus, Ohio, August 20-21.

Ganzeboom, Harry B. G., Paul M. de Graaf, and Donald J. Treiman. 1992. "A Standard International Socio-Economic Index of Occupational Status." Social Science Research 21:1-56.

Ganzeboom, Harry B. G., Ruud Luijkx, and Donald J. Treiman. 1989. "Intergenerational Class Mobility in Comparative Perspective." Research in Social Stratification and Mobility 8:3-84.

Haycraft, John. 1985. Italian Labyrinth: Italy in the 1980s. London: Secker & Warburg.

Heath, Anthony. 1983. La Mobilitá Sociale. Bologna: Il Mulino.

International Labor Office. 1969. International Standard Classification of Occupations. Revised edition, 1968. Geneva: International Labor Office.

International Labor Office. 1977. Yearbook of Labour Statistics 1977. Geneva: International Labor office.

Lipset, Seymour Martin, and Reinhard Bendix. 1959. Social Mobility in Industrial Society. Berkeley: University of California Press.

Lopreato, Joseph. 1965. "Social Mobility in Italy." American Journal of Sociology 71:311-314.

Lopreato, Joseph, and Lawrence E. Hazelrigg. 1972. Class, Conflict, and Mobility. San Francisco: Chandler.

Mason, Karen Oppenheim, William M. Mason, H. H. Winsborough, and W. Kenneth Poole. 1973. "Some Methodological issues in Cohort Analysis of Archival Data." American Sociological Review 38:242-58.

Sørensen, Aage, and Nancy Brandon Tuma. 1981. "Labor Market Structures and Job Mobility." Research in Social Stratification and Mobility 1:67-94.

Treiman, Donald J. 1977. Occupational Prestige in Comparative Perspective. New York: Academic Press.

Treiman, Donald J., and Kam-Bor Yip. 1989. "Educational and Occupational Attainment in 21 Countries." Pp. 373-394 in Cross-National Research in Sociology, edited by Melvin L. Kohn. Newbury Park: Sage.

Data files

Barbagli, Marzio, Vittorio Capecchi, Antonio Cobalti, Antonio de Lillo, and Antonio Schizzerotto. Italian Social Mobility Survey, 1985. [MRDF] Trento, Italy: Department of Sociology, University of Trento [distributor].

Barnes, Samuel H. Italian Mass Election Survey, 1968. [MRDF] ICPSR ed. Rome, Italy: CISER [producer]. Ann Arbor, MI: Inter-university Consortium for Political and Social Research [distributor] [ICPSR:7953].

Barnes, Samuel H., and Giacomo Sani. Italian Mass Election Survey, 1972. [MRDF] ICPSR ed. Ann Arbor, MI: Inter-university Consortium for Political and Social Research [distributor] [ICPSR:7954].

Barnes, Samuel H., Max Kaase, et al.: Political Action: An Eight Nation Study, 1973-1976. [MRDF] ICPSR ed. Ann Arbor, MI: Inter-university Consortium for Political and Social Research [distributor] [ICPSR:7777].

Ganzeboom, Harry B. G., and Donald J. Treiman (1992). International Stratification and Mobility File. [MRDF] Amsterdam/Nijmegen: Steinmetz Archives/International Archives of Comparative Social Research.

ISSP [International Social Science Program]. International Social Survey Program: Social Inequality, 1987. [MRDF] Ann Arbor, MI: Inter-university

Consortium for Political and Social Research [distributor] [ICPSR:9383].

Lopreato, Joseph. Class Consciousness in Italy, 1963-1964. [MRDF] Madison, WI: Data and Program Library Service. University of Wisconsin [distributor] [DPLS:SB-507-001-1-1].

Figures 1 - 4 go here.

Table 1.--Descriptive Statistics for the Surveys Used in the Analysis.

			Survey					
			ITA63	ITA68	1TA72	ITA75	ITA85	ITA87
Education Analysis								
Number of cases		Men	1255	919	654	558	1747	286
		Women	-	922	690	582	1812	305
Years of schooling (EDU)	Mean	Men	6.51	6.56	7.18	8.36	8.53	9.76
		Women	-	5.79	6.05	7.05	7.71	8.40
	S.D.	Men	4.05	3.48	3.63	4.22	4.36	4.58
		Women	-	2.85	2.67	3.87	4.24	4.42
Father's occ. status (FIS)	Mean	Men	34.7	30.0	31.3	33.8	36.6	31.6
		Women	-	29.1	30.0	34.0	36.8	31.7
	S.D.	Men	14.7	15.2	15.0	11.8	15.0	16.0
		Women	-	15.0	15.1	13.1	15.3	18.2
Birth year [2 digits] (BYR)	Range		04-42	04-43	08-47	11-50	21-60	23-61
Occupation Analysis								
Number of cases		Men	1210	884	553	546	1666	212
		Women	-	352	214	363	1191	92
Occupational status (ISE)	Mean	Men	37.7	38.4	40.7	40.1	45.4	47.7
		Women	-	36.3	40.5	41.4	46.3	49.9
	S.D.	Men	14.7	14.3	13.9	14.0	15.9	17.4
		Women	-	15.2	16.2	16.3	14.9	19.0
Seniority (EXP)	Mean	Men	31.8	31.3	28.9	29.5	29.0	21.4
		Women	-	30.5	27.3	28.3	28.7	20.8
	S.D.	Men	11.8	11.9	11.5	13.4	13.3	10.6
		Women	-	12.6	11.4	13.0	13.9	12.9
Father's occ. status (FIS)	Mean	Men	34.4	29.7	31.5	33.6	36.4	32.6
		Women	-	30.0	30.6	34.2	37.2	36.3
	S.D.	Men	14.5	15.0	15.1	11.7	14.7	15.7
		Women	-	15.4	15.5	12.7	15.0	20.2
Year's of schooling (EDU)	Mean	Men	6.43	6.47	7.34	8.22	8.44	10.36
		Women	-	6.30	6.92	7.69	8.24	10.88
	S.D.	Men	3.99	3.40	3.68	4.16	4.31	4.53
		Women	-	3.42	3.38	4.29	4.39	4.71
Labor force entry year (ENT)	Range		11-66	10-64	18-70	18-74	27-84	33-85

For definitions of variables see text. For descriptions of data files, see text and References. Source: Harry B.G. Ganzeboom & Donald J. Treiman (1992) "International Stratification and Mobility File." [MRDF] Amsterdam/Nijmegen: Steinmetz Archives/International Archives of Comparative Social Research.

Table 2.--Linear Trends in Educational Attainment in Italy for Cohorts Born Between 1904 and 1961, Persons Age 25-64. Metric Regression Coefficients.

	Men		Women	
	(a)	(b)	(a)	(b)
b_0: Intercept	7.51	7.51	6.38	6.38
b_1: Father's ISEI (FIS")	1.281	1.282	1.083	1.012
b_2: Year of Birth (BYR")	1.010	1.010	1.180	1.190
b_3: FIS"*BYR"		**-.010** (0.0)		**.110** (5.3)
Adj. R^2	.332	.331	.358	.362

For definitions of variables, see text.

Table 3.--Linear Trends in Occupational Attainment in Italy for Cohorts Born Between 1904 and 1961, Persons Age 25-64. Metric Regression Coefficients.

	Men				Women			
	(a)	(b)	(c)	(d)	(a)	(b)	(c)	(d)
A. Respondent's Occupational Status								
b_0: Intercept	3.95	3.96	3.47	3.48	3.82	3.84	2.97	3.09
b_1: Seniority (EXP')	.008	.006	.155	.155	-.004	-.011	.228	.199
b_2: Father's ISEI (FIS")	.452	.436	.254	.285	.403	.299	.164	.347
b_3: Years of schooling (EDU")			.190	.061			.234	-.006
b_4: EXP'*FIS"		.011		.003		.038		-.033
		(0.6)		(0.1)		(1.2)		(1.0)
b_5: EXP'*EDU"				.031				.059
				(4.9)				(4.4)
b_6: Lab. Frc. Entry Yr. (ENT')	.364	.363	.456	.459	.487	.483	.666	.645
b_7: ENT'*FIS"		**-.021**		**-.051**		**.004**		**-.073**
		(1.5)		(3.7)		(0.1)		(2.8)
b_8: ENT'*EDU"				**.041**				**.062**
				(7.1)				(4.9)
Adj. R^2	.322	.323	.460	.469	.354	.355	.520	.528
B. Respondent's Occupational Status (Non-farm Labor Force)								
b_0: Intercept	4.37	3.38	3.86	3.88	4.27	4.27	3.37	3.52
b_1: Seniority (EXP')	-.011	-.013	.146	.145	-.006	-.008	.231	.197
b_2: Father's ISEI (FIS")	.354	.301	.165	.184	.310	-.010	.077	.109
b_3: Years of schooling (EDU")			.184	.045			.227	-.031
b_4: EXP'*FIS"		.017		.002		.084		-.003
		(0.9)		(0.0)		(2.6)		(0.2)
b_5: EXP'*EDU"				.039				.061
				(5.3)				(4.6)
b_6: Lab. Frc. Entry Yr. (ENT')	.250	.248	.347	.349	.353	.353	.542	.517
b_7: ENT'*FIS'		**.011**		**-.023**		**.072**		**-.018**
		(0.5)		(1.7)		(2.6)		(0.7)
b_8: ENT'*EDU'				**.046**				**.069**
				(8.1)				(5.7)
Adj. R^2	.226		.226	.403	.413	.249	.250	.487

For definitions of variables, see text.

APPENDIX A: DESCRIPTIVE STATISTICS

Table 1.--Means, Standard Deviations, and Correlations for the Education Analysis (males above the diagonal
[N=5419]; females below [N=4331]).

	EDU	FIS"	BYR'	FIS"*BYR'	Mean	Std. Dev.
EDU: Years of school		.452	.359	.054	7.613	4.190
FIS": Father's ISEI	.434		.000	.119	.000	1.473
BYR': Year of birth	.417	.004		.008	1.045	14.832
FIS"*BYR'	.239	.407	.006		.002	22.308
Mean	6.991	.000	5.122	.077		
Std. Dev.	3.824	1.517	13.425	22.771		

Table 2.--Means, Standard Deviations, and Correlations for the Occupation Analysis (males above the diagonal [N=5071]; females below [N=2212]).

	ISE'	EDU"	FIS"	EXP'	EYR'	EDU"*EXP'	FIS"*EXP'	EDU"*EYR'	FIS"*EYR'	Mean	S.D.
ISE': Occ. status		.516	.420	-.338	.383	.473	.396	.300	.121	4.140	1.539
EDU": Years of school	.526		.448	-.112	.005	.918	.382	.544	.197	.000	3.412
FIS": Father's ISEI	.365	.453		-.048	.001	.419	.913	.204	.372	.000	1.425
EXP': Seniority	-.416	-.086	.014		-.836	-.089	-.030	-.064	-.022	2.979	1.259
EYR': Entry year	.473	.009	.014	-.892		-.010	-.003	.003	.004	.458	1.641
EDU"*EXP'	.481	.901	.417	-.062	-.003		.423	.244	.072	-.482	9.468
FIS"*EXP'	.341	.388	.895	.031	-.004	.431		.068	.026	-.086	4.334
EDU"*EYR'	.424	.780	.346	-.066	.013	.471	.204		.400	.028	5.765
FIS"*EYR'	.214	.332	.661	-.005	-.003	.210	.293	.426		.003	2.498
Mean	4.346	.000	.000	2.848	1.110	-.368	.026	.045	.009		
Std. dev.	1.598	3.216	1.439	1.339	1.537	8.722	4.197	6.408	2.992		

Table 3.--Means, Standard Deviations, and Correlations for the Non-farm Occupation Analysis (males above the diagonal [N=4273]; females below [N=1900]).

	ISE	EDU"	FIS"	EXP"	EYR'	EDU"*EXP'	FIS"*EXP'	EDU"*EYR'	FIS"*EYR'	Mean	S.D.
ISE': Occ. status		.537	.373	-.268	.296	.489	.339	.368	.168	4.504	1.387
EDU": Years of school	.559		.433	-.121	.004	.918	.369	.591	.212	.000	3.548
FIS": Father's ISEI	.327	.443		-.055	.001	.404	.909	.222	.437	.000	1.454
EXP': Seniority	-.338	-.092	-.001		-.826	-.102	-.039	-.070	-.025	2.842	1.239
EYR': Entry year	.379	.010	.004	-.889		-.006	-.001	.003	.003	.667	1.600
EDU"*EXP'	.504	.902	.414	-.064	-.001		.410	.304	.093	-.533	9.596
FIS"*EXP'	.294	.383	.891	.014	-.002	.434		.089	.091	-.100	4.310
EDU"*EYR'	.493	.814	.355	-.076	.013	.525	.218		.391	.025	6.051
FIS"*EYR'	.241	.336	.707	-.008	-.002	.223	.345	.419		.003	2.599
Mean	4.717	.000	.000	2.684	1.313	-.397	-.001	.049	.008		
Std. dev.	1.398	3.312	1.469	1.298	1.481	8.654	4.142	6.739	3.152		