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Age, Education and Occupational Earnings Inequality

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Age, Education, and Occupational Earnings Inequality

The presence of a wide dispersion of earnings within narrowly defined Census occupations raises two important questions about the determinants of earnings behavior. First, what factors are operative that cause individuals in the same occupation to be rewarded differently? Second, do there exist systematic differences in these determinants between occupations or groups of occupations? A response to the second will have to await empirical investigation. In response to the first, six explanations may be offered as hypotheses:

(1) Differences in tasks. A given Census occupation may represent a variety of tasks and functions. A "secretary", for example, may engage in typing, shorthand and dictation work, reception work, administrative chores, budgeting, or some combination of these. A "lawyer" may work on corporate tax returns, wills, estates, contracts or litigation. Tasks within an occupation usually vary in difficulty, and the more difficult are usually accorded higher compensation because of the additional training required or because of the limited supply of qualified personnel. Differences in schooling or ability, insofar as they are related to the tasks performed, may thus lead to differences in pay within an occupation.

(2) Different levels of efficiency. Workers in the same occupation may perform the same task at different speeds or levels of efficiency: Typists, programmers and many kinds of operatives differ in the speed and accuracy of their work. Differences in experience and ability, insofar as they reflect the efficiency and reliability of

the work performed, may result in differences in earnings within an occupation.

(3) Institutional factors. Institutions will often set up differential pay scales for a given occupation on the basis of education and experience. Public school systems, branches of the civil service, and large corporations have set pay scales based on education or years of service for many lines of work. The formulas used may reflect worker productivity or, as Ruggles (1970) argues, they may reflect cultural and institutional systems of remuneration quite contrary to actual productivity. For whichever reason, earnings may vary directly with age or education within an occupation.

(4) Time worked. Differences in hours worked per week and employment stability over the year will lead to differences in annual earnings.

(5) The demand for labor. Labor market conditions will often vary from one locality to another, depending on the availability of labor and the industrial mix of the regions. Differences in the demand and supply conditions between areas will cause differences in compensation for similar work.

(6) Discrimination. The differential compensation of workers with similar ability, qualifications, and work experience on the basis of their sex or race will lead to a dispersion of earnings within an occupation.

In this paper, we will investigate the effect of these factors on occupational earnings inequality across all occupations in our sample and across occupations in five major Census subgroups. Age and schooling will receive primary attention in our work and it will be shown that they are important determinants of earnings inequality among professional and clerical occupations but not among skilled, semi-skilled or unskilled occupations. Ability is also hypothesized as an important factor, but no measure of ability is provided in our sample. Differences in time worked and labor demand conditions, as measured by industrial and urban-rural mix, will also be analyzed, and their effect on earnings inequality is strong in most of the occupational subsamples. Differences in the race and sex composition of occupations do not appear to be significant factors in occupational earnings inequality, and the explanation offered is that discrimination takes the form of occupational segregation rather than differences in pay for similar work. In the conclusion a sketch of a "structural" theory of income distribution is proposed to account for our results.

I. Empirical Results and Interpretation

To analyze the determinants of occupational earnings inequality, we randomly drew a sample of approximately 200 workers for each of 291 occupations in 1960 and 439 in 1970 from the 1960 and 1970 Census 1/100 Public Use Samples.[1] We then computed the mean and standard deviation of earnings, age, schooling and hours worked, an industrial dispersion coefficient, and the percent urban, white male, white female, black male, and black female for each occupation. For the purpose of regression analysis, the observational unit was the occupation, and regressions

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[1] In the 1960 Census there were 295 occupations. Four, however, had fewer than 50 occupational members and were omitted. Many occupations in the two years had less than 200 members but more than 50 and were kept in the sample. The total sample size was 41,349 in 1960 and 63,661 in 1970.

were performed across all occupations and across occupations in each of five major sub-groups -- professional, technical, and managerial (professional); clerical and sales (clerical); skilled and craft (skilled); semi-skilled and operative (operative); and service and unskilled (service).[2]

A. The Level of Occupational Earnings Inequality

We chose the coefficient of variation of earnings, defined as the ratio of the standard deviation of earnings to mean earnings ($SD(E)/\overline{E}$), as our measure of earnings inequality, since mean earnings varied considerably between occupations and for a given occupation between 1960 and 1970. The coefficient of variation for all those earning wage and salary income in the economy was .953 in 1960 and .952 in 1970 (See Wolff (1975)).[3] The average (unweighted) occupational coefficient of variation across all occupations was .683 in 1960 and .707 in 1970, lower than the overall level of inequality in the two years (Table 1). The slight rise in the average level of occupational inequality, despite the finer occupational classification scheme in 1970, suggests that the presence of occupational

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[2]The Public Use Sample Occupational Codes for the five major sub-groups are as follows:

	1960	1970
Professional	000-290	001-246, 801-806
Clerical	301-395	260-396
Skilled	401-555	401-586
Operative	601-775	601-726
Service	801-985	740-796, 821-986

[3]Earnings reported in the 1960 and 1970 Censuses refer to years 1959 and 1969, respectively. All other variables, unless otherwise indicated, are as of the time the Census questionnaire was answered.

Table 1

Unweighted Mean Values Across Occupations of Regression Variables

Variables	1960 All	Professional	Clerical	Skilled	Operative	Service
$SD(E)/\overline{E}$	.683	.715	.769	.495	.606	.880
$\overline{E}$	4428	6148	3570	4727	3513	2110
$\overline{H}$	1824	1961	1710	1874	1783	1604
SD(H)	639	646	646	527	625	785
$\overline{A}$	39.2	39.4	36.8	41.7	36.9	40.1
SD(A)	12.8	12.3	13.4	12.8	11.4	15.1
$\overline{S}$	11.6	14.6	11.9	9.9	9.8	9.4
SD(S)	2.37	2.17	2.17	2.52	2.42	2.71
$(\overline{S \cdot A})$	453	574	439	411	355	377
C	616	782	623	585	471	469
% SMSA	55.0	58.6	58.6	57.1	50.5	46.6
% White male	69.7	74.5	51.3	87.7	72.5	47.2
% White female	20.8	21.8	42.7	5.0	17.2	25.8
% Black male	6.7	2.6	4.3	6.9	8.0	16.1
% Black female	2.8	1.2	1.7	0	2.3	10.9
Number of Occupations	291	97	37	59	53	45
Fraction of Labor Force	1.00	.213	.254	.120	.195	.218

Key: 1) Bar ($^{-}$) = sample mean; 2) SD = sample standard deviation; 3) E = wage and salary earnings; 4) $SD(E)/\overline{E}$ = coefficient of variation; 5) H = total hours worked per year (average weeks per year times average hours per week); 6) A = age in years; 7) S = schooling in years; 8) C = "industrial dispersion coefficient" of occupational employment, defined as

$$C_i = \sum_{j=1}^{16} f_{ij}(\overline{E}_{ij} - \overline{E}_i)^2$$

where f_{ij} is the fraction of occupation i employed in industrial group j for each of the 16 major Census industrial groupings, $\overline{E}_{ij}$ is the mean earnings of occupation i in industrial group j, and $\overline{E}_i$ is the mean earnings of occupation i; 9) % SMSA = percent of occupation employed in an SMSA (Standard Metropolitan Statistical Area); 10) Whites = non-Spanish whites, Chinese, Japanese; 11) Blacks = Blacks, Spanish whites, and others.

Data Sources: Stratified samples drawn from the 1960 and 1970 Census 1/100 Public Use Samples.

Table 1 (continued)

Unweighted Mean Values Across Occupations of Regression Variables

Variables	1970					
	All	Professional	Clerical	Skilled	Operative	Service
$SD(E)/\overline{E}$	.707	.696	.794	.560	.647	.934
$\overline{E}$	6816	9597	5403	6802	5287	3259
$\overline{H}$	1821	1913	1694	1921	1859	1538
SD(H)	635	643	656	558	603	739
$\overline{A}$	38.4	39.4	37.1	37.8	39.1	37.4
SD(A)	13.6	12.8	14.4	12.7	14.2	15.3
$\overline{S}$	12.3	14.9	12.3	10.9	10.2	10.5
SD(S)	2.25	2.11	1.99	2.22	2.49	2.64
$(\overline{S \cdot A})$	471	584	458	410	397	387
C	1060	1356	1125	988	785	697
% SMSA	50.0	56.9	57.1	54.5	44.1	26.1
% White male	63.6	69.4	42.0	84.1	58.7	45.0
% White female	25.2	25.1	49.0	6.6	24.7	31.4
% Black male	7.0	3.1	4.1	8.4	11.7	12.0
% Black female	4.1	2.3	4.9	1.9	5.0	11.6
Number of Occupations	439	151	61	96	66	65
Fraction of Labor Force	1.00	.212	.259	.117	.170	.242

Key: 1) Bar $(\overline{\ \ })$ = sample mean; 2) SD = sample standard deviation; 3) E = wage and salary earnings; 4) $SD(E)/\overline{E}$ = coefficient of variation; 5) H = total hours worked per year (average weeks per year times average hours per week); 6) A = age in years; 7) S = schooling in years; 8) C = "industrial dispersion coefficient" of occupational employment, defined as

$$C_i = \sum_{j=1}^{16} f_{ij}(\overline{E}_{ij} - \overline{E}_i)^2$$

where f_{ij} is the fraction of occupation i employed in industrial group j for each of the 16 major Census industrial groupings, $\overline{E}_{ij}$ is the mean earnings of occupation i in industrial group j, and $\overline{E}_i$ is the mean earnings of occupation i; 9) % SMSA = percent of occupation employed in an SMSA (Standard Metropolitan Statistical Area); 10) Whites = non-Spanish whites, Chinese, Japanese; 11) Blacks = Blacks, Spanish whites, and others.

Data Sources: Stratified samples drawn from the 1960 and 1970 Census 1/100 Public Use Samples.

inequality is not a classification artifact but a substantive phenomenon. On the other hand, when the coefficient of variation was regressed on the occupational size (in millions of workers) across all occupations, occupational size had a positive and significant coefficient.[4] This suggests that occupations with larger populations are characterized by more tasks and functions and that a finer occupational breakdown may reduce the measured level of inequality.

The mean coefficient of variation varied considerably between occupational groups. In 1960 it was highest among service occupations, followed by clerical, professional, operative and skilled occupations. Between 1960 and 1970 the average degree of inequality rose in all groups, except professionals. However, the rank order among the five groups was identical in 1970 to that of 1960, suggesting the same set of factors at work in the two years within each of the five groups.

B. Age[5]

On *a priori* grounds we would expect the impact of age on earnings to vary considerably from occupation to occupation. For those whose

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[4]The regression results were as follows (t-ratios in parentheses):

	1960	1970
Constant	.664	.691
	(33.8)	(44.8)
Occupational Size	.064	.065
	(2.6)	(2.3)

[5]Mincer (1974) shows quite convincingly that individual experience is a better predictor of an individual's earnings than his age. In our sample, where the unit of observation is the occupation, there is little quantitative difference between using age and experience. (The correlation between mean age and mean experience, estimated as age less years of schooling less five, across occupations was .93 in both 1960 and 1970.) Moreover, the use of the standard estimate for experience is known to be poor for females and minorities, who are included in our sample, and the use of experience rather than age raises the question of whether total work experience or experience in a particular occupation is more relevant for occupational earnings.

experience leads to on-the-job training and either increased efficiency at a given task (typing, for example) or the assumption of more difficult tasks (engineering, for example), we would anticipate a rise in the pay rate with age. Among other occupational groups like civil servants, school teachers and unionized crafts, where wage scales are apportioned to age for institutional reasons, we would anticipate a similar cross-sectional profile. Among semi-skilled and unskilled occupations, like taxi drivers and maids, where experience leads to little increase in productivity, we would expect little bearing of age on earnings. Previous cross-sectional evidence tends to confirm this pattern (Wolff (1975), pp. 21-22 and 60-61). Within most professional and clerical occupations earnings rise with age until about age 60 and then level off. Within most skilled, semi-skilled and unskilled occupations, earnings rise until about age 30, level off, and then decline. For the labor force as a whole, the hourly wage rate rises steeply with age until about age 27, then rises less steeply, and finally levels off at about age 40. These results conform with common observation. For many professional groups, like programmers, air pilots, and college professors, administrative positions, particularly in public administration, and clerical jobs, like secretaries, salaries will start low, increase rapidly in the early years, then rise less rapidly, and, in many instances, level off as institutionally-imposed ceilings are reached. For occupational groups, like garage attendants, bartenders, and janitors, earnings will show little systematic relation to age.

Moreover, the previous results suggest that the age-earnings cross-sectional profile is approximately logarithmic in shape for the pooled labor force and for professional and clerical occupations and flat for skilled, semi-skilled and service groups. In the former case, an increase in mean age, with the age distribution around the mean constant, will result in higher mean earnings, a lower variance of earnings, and thus a lower coefficient of variation.[6] In the latter case, a higher mean age will have no effect on the coefficient of variation. The regression results in Table 2 confirm this hypothesis.

The impact of the standard deviation of age on earnings inequality is more difficult to predict a priori, since it depends on how the shape of the age distribution shifts. If an increased standard deviation of age reflects a greater concentration of younger workers, then earnings inequality would rise, since mean earnings would fall but the variance of earnings rise. Conversely, if it reflects a greater concentration of older workers, inequality would fall, since mean earnings would increase but the variance decline. In 1960

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[6]Suppose in a given occupation, earnings $E_i = k(\ln A_i) + c$ for an individual i of age A_i. Mean earnings $\overline{E}$ and the variance of earnings V(E) are thus given by

$$\overline{E} = \frac{k}{N} \Sigma \ln A_i + c \quad \text{and} \quad V(E) = \frac{1}{N^2} \sum_i \sum_j (E_i - E_j)^2$$

where N is the occupational size. Suppose everyone in the occupation ages by Δt. Then

$$E_i' = E_i + \Delta E_i \approx E_i + k \frac{\Delta t}{A_i}$$

Mean earnings will therefore rise. Moreover,

$$E_i' - E_j' = E_i - E_j + k \, \Delta t \left(\frac{1}{A_i} - \frac{1}{A_j}\right)$$

If $E_i > E_j$, then $|E_i' - E_j'| < |E_i - E_j|$.

If $E_i < E_j$, then $|E_i' - E_j'| < |E_i - E_j|$.

Therefore, the variance of earnings will fall.

Table 2

Results of Regressing the Coefficient of Variation on Indicated Variables Across Occupations (t-ratios in parentheses)

Independent Variables	1960 All	Professional	Clerical	Skilled	Operative	Service
Constant	2.01* (3.8)	2.78** (2.5)	11.57 (6.9)	-.85 (0.7)	1.66 (0.9)	-4.35** (2.0)
$\overline{E}$	.0000089 (.82)	.000018 (1.3)	-.000048 (1.1)	-.000034** (2.0)	.000055 (1.6)	-.000069 (1.4)
$\overline{H}$	-.00040* (8.3)	-.00015 (1.7)	-.00073* (5.1)	-.00017* (3.2)	-.00033 (1.9)	-.00029** (2.3)
SD(H)	.00089* (10.4)	.0013* (9.2)	-.00028 (1.0)	.00049* (3.6)	.00068* (3.1)	.00088* (3.6)
$\overline{A}$	-.041* (3.5)	-.053** (2.1)	-.266* (5.2)	.026 (0.9)	-.015 (0.3)	.073 (1.6)
SD(A)	.0016 (.24)	-.036** (2.4)	-.022 (1.1)	.0033 (0.3)	.0057 (0.3)	.025 (1.7)
$\overline{S}$	-.116* (2.6)	-.180** (2.2)	-.822* (5.3)	.152 (1.3)	-.097 (0.6)	.450** (2.4)
SD(S)	.076** (2.5)	.067 (1.4)	.225** (2.3)	.034 (0.7)	-.036 (0.4)	.180 (1.6)
$(\overline{S \cdot A})$	.0032* (2.9)	.0039** (2.0)	.023* (5.0)	-.0029 (1.0)	.0017 (0.4)	-.0081 (1.7)
C	.000075* (3.3)	.000061** (2.2)	.00015 (1.9)	.000082** (2.1)	.00039* (4.6)	.00013** (2.0)
$(\%SMSA - .5)^2$	.012* (3.7)	.015* (2.8)	.013 (0.7)	.0034 (0.9)	.0084 (1.3)	.029* (4.1)
R^2	.606	.644	.912	.750	.670	.835

* Significant at the one percent level

** Significant at the five percent level

(Refer to Table 1 for symbol definitions.)

Table 2 (continued)

Results of Regressing the Coefficient of Variation on Indicated Variables Across Occupations (t-ratios in parentheses)

Independent Variables	1970 All	Professional	Clerical	Skilled	Operative	Service
Constant	2.30* (4.7)	3.87* (3.7)	8.27* (4.7)	1.41 (1.3)	2.24 (0.5)	1.17 (0.9)
$\overline{E}$	-.000016 (0.2)	.000017** (2.4)	.000033 (1.1)	-.000011 (1.0)	.000012 (0.3)	-.000104* (4.6)
$\overline{H}$	-.00040* (9.2)	-.00025* (3.7)	-.00058* (3.1)	-.00031* (4.9)	-.0010* (3.7)	-.00003 (0.3)
SD(H)	.00040* (5.4)	.00097* (9.4)	.00092* (3.7)	.00044* (4.2)	-.00046 (1.5)	-.00003 (0.2)
$\overline{A}$	-.045* (3.9)	-.078* (3.2)	-.198* (3.5)	-.032 (1.3)	-.010 (0.0)	-.003 (0.1)
SD(A)	.018* (3.3)	.021** (2.0)	.028 (1.7)	.015* (2.7)	.020 (0.6)	-.011 (1.1)
$\overline{S}$	-.124* (3.1)	-.264* (3.7)	-.657* (4.0)	-.057 (.64)	.037 (0.1)	.009 (1.1)
SD(S)	.089* (3.5)	.037 (1.1)	.104 (1.1)	.084** (2.2)	.109 (0.8)	.129** (2.0)
$(\overline{S \cdot A})$	.0035* (3.4)	.0053* (3.0)	.016* (3.4)	.0024 (1.1)	-.00031 (0.0)	-.00010 (0.0)
C	.000040* (3.7)	.000025** (2.0)	.000042 (1.7)	.000027 (1.4)	.000087 (1.3)	-.000043 (1.0)
$(\%\ SMSA - .5)^2$	.016* (6.2)	.022* (4.3)	.044* (4.9)	.0058** (2.2)	.0176 (1.9)	.0101 (1.8)
R^2	.534	.619	.826	.773	.502	.674

* Significant at the one percent level

** Significant at the five percent level the only significant coefficient of the standard deviation of age at the five percent significance level was in the professional group, but it was negative. In 1970 the coefficients in the pooled, professional, and skilled samples were positive and significant at, at least, the five percent level and the coefficient in the clerical subsample was positive and significant at the ten percent level. This difference in results between 1960 and 1970 can be attributed to the change in the age composition of the work force during this period. Between 1960 and 1970, the median age of those employed fell from 41 to 39, the percentage of workers 21 or under increased from 16 to 20, and the percentage of workers 26 or under from 25 to 30 (see Wolff (1975), p. 27). In the professional and clerical groups the mean age was stable but the standard deviation of age increased and in the skilled group mean age fell sharply while the standard deviation of age remained constant, suggesting that in each of these groups there was a higher concentration of new entrants in 1970 than in 1960 (Table 1). Given the change in the age composition of the work force, the stronger positive effect of the standard deviation of age on earnings inequality in 1970 than in 1960 would serve to confirm the hypothesized cross-sectional age profiles for the occupational groups.

C. Schooling

Previous work suggests a wide variation in education-earnings profiles among occupations (Wolff (1977)). For professions like lawyers, doctors, and university teachers, where an advanced degree is required, there is almost no variance in schooling and thus little relation of earnings to education within the occupation. Among other professions like school teachers, nurses and engineers, schooling generally varies from two to eight years of college, and earnings generally rise with schooling in this range. Among many clerical occupations, education varies from two years of high school to four years of college and earnings tend to rise with education. Among skilled occupations there is a wide variance of educational attainment, and earnings tend to rise with schooling through the high school years and then level off. For operatives and service workers, for whom relatively little training is required, there is a wide range of schooling levels but little relation of earnings to schooling. Additional evidence suggests that among many professional, most clerical, and some skilled occupations, earnings rise with schooling after some minimal level of education, though less slowly as advanced education becomes increasingly less relevant to the actual tasks performed, and usually up to a ceiling (Wolff (1975), pp. 53-56). Thus, we would expect logarithmic education-earnings cross-sectional profiles within the professional, clerical and perhaps skilled occupational groups, and flat profiles within the operative and service groups.

The regression coefficients of mean schooling ($\overline{S}$) confirm this hypothesis at the five percent significance level for professionals in 1960 and at the one percent level for professionals in 1970 and clerical workers in both years. Moreover, the coefficient of mean schooling is negative and significant at the five percent level in 1960 and at the one percent level in 1970 for the pooled sample. The other coefficients of mean schooling are insignificant, except for service workers in 1960, where it is positive. The effect of the standard deviation of schooling on earnings inequality is more difficult to predict, since, as in the case of the standard deviation of age, it depends on the other moments of the education distribution. The coefficient of the standard deviation of schooling was positive and significant at, at least, the five percent significance level for the pooled sample in both years, clerical workers in 1960, and skilled and service workers in 1970.

Between 1960 and 1970 the educational composition of the work force changed substantially. The percentage with 8 or less years of schooling fell from 33 to 22; the percentage with a high school degree or more rose from 51 to 55; and those who attended college from 17 to 23 (Wolff (1975), pp. 21-22). However, the distribution of the labor force among the five occupational groups was almost constant in this period, except for a shift between operatives and service workers (Table 1). As a result, the more educated "filtered down" the occupational ladder. The mean education among professionals increased slightly between 1960 and 1970, that of clerical workers and operatives somewhat more, and that of skilled and service workers by about a full year. To determine whether the effect of schooling on mean hourly earnings had altered over this period, we tried two additional regression forms across the pooled sample and the subsamples:

$$\text{(i)} \quad \frac{\bar{E}}{\bar{H}} = a_0 + a_1\bar{S} + a_2\bar{A} + a_3\bar{H} + u$$

$$\text{(ii)} \quad \frac{\bar{E}}{\bar{H}} = b_0 + b_1 \ln(\bar{S}) + b_2 \ln(\bar{A}) + b_3\bar{H} + v$$

The coefficient estimates of mean schooling and the logarithm of mean schooling were significant at, at least, the five percent level for the pooled, professional and clerical samples in 1960, insignificant for the skilled, operative, and service groups in 1960, and significant for all groups in 1970 (Table 3). These results suggest that as job opportunities closed out for the more highly educated new entrants in the professional and clerical field, they filtered down to the higher-paying occupations in the skilled, operative and service groups. This would explain the significant correlation of mean schooling and mean hourly earnings in these groups in 1970 and its absence in 1960.[7] However, this is a relation across occupations. Within skilled, operative, and service occupations there is as little relation between education and earnings in 1970 as there was in 1960.

D. The Interaction of Age and Schooling

For professional and clerical workers, we would expect a positive relation between earnings inequality and the age-education interactive variable, whereas for the other groups no relation. In fields like engineering, programming, administration and secretarial work, entry salaries tend to be fixed primarily by schooling level. The variance in earnings will thus be low for a young cohort. As the cohort ages, the importance of formal schooling for job performance lessens and that of experience increases, and earnings spread out as factors like

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[7]The 1970 coefficient estimates of mean schooling and the logarithm of mean schooling in these three groups are approximately half of what they are in the other groups, suggesting a much weaker relation between schooling and earnings in these groups as opposed to the professional and clerical groups.

Table 3

Results of Regressing Hourly Earnings on Schooling, Controlled for Age and Hours Worked*
(Coefficients a_1 and b_1 and their t-ratios shown)

	All	Professional	Clerical	Skilled	Operative	Service
1960						
Form (i)	.211	.227	.269	.123	.110	.044
	(13.2)	(5.8)	(2.7)	(1.6)	(1.3)	(0.7)
Form (ii)	2.53	3.10	3.04	1.49	0.95	0.37
	(12.9)	(5.8)	(2.5)	(1.9)	(1.1)	(0.6)
1970						
Form (i)	.448	.472	.469	.255	.262	.234
	(24.4)	(9.1)	(4.4)	(3.0)	(2.8)	(3.1)
Form (ii)	5.62	6.77	5.68	2.82	2.71	2.43
	(23.4)	(9.0)	(4.1)	(3.1)	(2.9)	(3.1)

* See text for equations (i) and (ii).

ability, drive, and the occurrence of opportunities assert themselves. The positive impact of the interaction variable on earnings inequality is verified for the professional, clerical and pooled samples in Table 2.

E. Other Factors

1. Time Worked

With a fixed hourly wage rate and a fixed distribution of hours worked around the mean, an increase in mean hours worked will cause an increase in mean earnings and a decline in the coefficient of variation. With the hourly wage rate and mean hours worked fixed, an increase in the standard deviation of hours worked will result in increased earnings inequality. This is confirmed for the pooled sample and most of the occupational subsamples (Table 2).[8]

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[8] The variable used was total hours worked per year (H), which was computed as the product of weeks worked in the year preceding the Census year and the hours worked in a prespecified week in the Census year. Only those with a positive number of weeks and hours worked were included in calculating the mean and standard deviation of hours worked.

Results from regression forms (i) and (ii) (see Section C) indicated that the hourly wage rate was positively and significantly related to hours worked in the pooled and all subsamples, except for professionals in both years and clerical workers in 1960. This may reflect overtime payments, the lower compensation of part-time and marginal employees and the like. This finding is, in fact, what Barzel (1973) predicts. Mincer's (1974, p. 94) results also suggest a positive correlation between weekly earnings and weeks worked, after controlling for schooling and experience. Mincer (1975) later argues that firms which train employees are more reluctant to lay them off and employees are more reluctant to switch jobs because their marginal product would be lower elsewhere. This, plus more efficient search procedures, would account for the smaller unemployment of those with the higher wage rates. This issue deserves fuller exploration at a later time. For our present purposes, the hourly wage rate gradient with respect to hours worked does not appear sufficiently great to offset the postulated effect of the mean and standard deviation of total hours on earnings inequality.

2. Alternative Employment Opportunities

To control for the industrial mix of an occupation, we introduced the following "industrial dispersion coefficient" C:

$$C_i = \sum_{j=1}^{16} f_{ij} (\overline{E}_{ij} - \overline{E}_i)^2$$

for occupation i across industries j, where $\overline{E}_{ij}$ is the mean earnings of occupation i in industry j, $\overline{E}_i$ the mean earnings of occupation i, and f_{ij} is the percentage of occupation i employed in industry j.[9] This index is higher the larger the dispersion of occupational membership across industries and the larger the difference in mean earnings across industries. It roughly measures the availability of alternative employment opportunities for an occupation in different industries and tends to reflect differences in tasks between industries. (For example, a bank economist might forecast interest rates, while a government economist might design national account estimating procedures.) We would expect the dispersion index to have a positive impact on earnings inequality. In 1960 this relation is confirmed in five of the six cases at, at least, the five percent level and in the sixth case at the ten percent level. In 1960, it is confirmed at, at least, the five percent level in two cases and at the ten percent level in one.

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[9] The 16 industrial groups used in the index are as follows: 1) Agriculture, Forestry, and Fisheries; 2) Mining; 3) Construction; 4) Manufacturing (durables); 5) Manufacturing (nondurables); 6) Transportation; 7) Communication; 8) Utilities and Sanitary Services; 9) Wholesale Trade; 10) Retail Trade; 11) Finance, Insurance, and Real Estate; 12) Business and Repair Services; 13) Personal Services; 14) Entertainment and Recreation Services, 15) Professional and Related Services; and 16) Public Administration.

To control for the geographical spread of an occupation's employment, we introduced into the regression the square of the difference between the percentage of an occupation employed in an SMSA and 50 percent.[10] Previous work had indicated that mean annual earnings were almost uniformly higher for workers employed in SMSA's than those employed outside them in a given occupation (Wolff (1975), p. 50). Moreover, in running the following regression:

$$\text{(iii)} \qquad \frac{\overline{E}}{\overline{H}} = c_0 + c_1\overline{S} + c_2\overline{A} + c_3\overline{H} + c_4(\%\ \text{SMSA}) + w$$

we obtained a positive and significant coefficient at, at least, the five percent level, in the pooled samples and all subsamples in both years, except for clerical workers in both years and operatives in 1960. The previous work had also suggested that the variance of earnings was approximately the same for SMSA workers and non-SMSA workers in most occupations (Wolff (1975), p. 50). The overall variance of earnings is therefore maximized when there is an even split in the occupational work force between SMSA and non-SMSA employment. Mean earnings, on the other hand, rise as the percentage employed in SMSA's increases. Thus, the coefficient of variation either declines as the percentage employed in an SMSA increases or rises as the SMSA percentage increases from zero, peaks at or before fifty percent, and then declines. The percent SMSA less fifty percent should therefore have a negative coefficient. The results for 1960 show significant and positive coefficients for the pooled, professional, and service

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[10] SMSA stands for Standard Metropolitan Statistical Area which includes not only the central city area but also the "ring."

samples at the one percent level, and those for 1970 show three coefficients positive and significant at the one percent level, one at the five percent level, and the other two at the ten percent level.[11]

One possible explanation for this somewhat paradoxical result is that the urban-rural mix of an occupation is a relatively invariant characteristic of that occupation. Except for services, the percentage employed in SMSA's remained relatively stable in the occupational subsamples between 1960 and 1970 (Table 1). Previous results had showed this even more strongly (Wolff (1975), p. 50). Thus, what our results indicate is not that shifts in the urban-rural mix lead to greater or less inequality within an occupation but rather that between occupations more mixed occupations have less inequality than less mixed ones. This latter possibility might be due to greater competitive forces in the mixed occupations and a greater tendency to wage rate equalization for similarly qualified workers. This hypothesis warrants further investigation in the future.

Another interesting feature of the results is that, except for services, all the coefficients of the deviation form of the percent SMSA variable were more significant in 1970 than in 1960.[12] On the other hand,

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[11]We also tried the percent employed in an SMSA instead of the deviation form of the percent employed in the regression with far fewer significant coefficients. Farbman (1973), whose unit of observation was the county and whose sample consisted of southern states, found relatively few significant coefficients for percent rural.

[12]The rise in the minimum wage during the sixties and the extension of its coverage to many kinds of service workers may have narrowed the earnings differential between urban and rural areas for service workers and accounted for the lower t-ratio of the variable for service workers in 1970.

all the regression coefficients of the industrial dispersion coefficient were less significant in 1970 than in 1960, except for the pooled sample. This suggests a trend towards reducing occupational earnings differences across industries while increasing them between rural and urban areas. This too might warrant further analysis in the future.

3. Expected Earnings

To determine whether a trade-off existed between expected earnings and its risk, as measured by the coefficient of variation, we included mean earnings ($\bar{E}$) in the regression. The only case of a significant and positive coefficient was for professionals in 1970, suggesting that this group has a choice between a low risk, low return job, like teaching, and a high risk, high return job, like administration.[13]

4. Discrimination

Having controlled for time worked, age, schooling, industrial dispersion, and geographical mix, we wanted to determine whether there existed any systematic tendency to underpay females and minorities with respect to white males within a given occupation. In the case of minorities, which constitute a small percentage of most occupations, the effect should show up as a positive relation between the coefficient of variation and the percentage of minorities employed in the occupation. In the case of females, where they constitute approximately one-third or more of an occupation, this should show up as a negative relation between earnings inequality and the square of the

[13] Both skilled workers in 1960 and service workers in 1970 had significant negative coefficients. This suggests that the high wage earners in these groups also enjoy the steadier employment. In the case of skilled workers, this may be due to the heavy presence of unions.

difference between percent female and fifty percent.

Different forms were tried for the pooled sample and each of the subsamples in the two years. The results varied very little between forms, and Table 4 shows the results of adding the deviation form of percent females and the percent black and Hispanic to the variables in Table 2.[14] The change in the R-squared from adding these two variables was relatively small in all cases except for operatives in 1960 and service workers. The only occurrences of significant coefficients in the hypothesized directions were for the percent black and Hispanic in the operative group in 1960 at the one percent level and for the deviation form of percent females in 1970 in the pooled sample at the five percent level and for the service group at the one percent level. The coefficient estimates of the other variables changed very little from adding these two variables.[15] By and large no systematic pattern of differential compensation for white males, females, and minorities within occupation could be inferred from our regression results.

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[14] Other variables tried were the percent female, the percent white female, the deviation form of the percent white female, the deviation form of percent black and Hispanic, the percent black and Hispanic males, the percent black and Hispanic females, and the deviation forms of the latter two.

[15] The major exceptions were the following: 1) the standard deviation of schooling became insignificant for clerical workers in 1960; 2) mean earnings became insignficant for skilled workers in 1960; 3) mean age, the standard deviation of age and the age-education interactive variable became significant, at the five percent level, for service workers in 1960; and 4) the standard deviation of schooling became insignificant for service workers in 1970.

Table 4

The Change in the R-Squared Statistic Resulting from the Inclusion of Percent Black and Hispanic and the Square of the Difference Between Percent Female and Fifty Percent with the Variables in Table 2

	1960	1970
All	.009	.030
Professionals	.010	.012
Clerical	.014	.001
Skilled	.031	.020
Operative	.083	.011
Service	.030	.078

These results conform to those reported in an earlier study (Wolff (1976)). In this study the work force was divided into 32 occupational groups, and the overall differential in earnings between whites, and blacks and Hispanics, and males and females was decomposed into two effects: one from differences in their occupational distribution and the other from differences in mean earnings within occupation. In the case of blacks and Hispanics, "occupational segregation" was the more important factor of the two, while in the case of females both factors were of similar import.[16] However, no adjustments were made for hours worked, age, education, or differences in labor demand conditions, as in the present study. Moreover, here we use a much

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[16] Even with a division of the labor force into five occupational groups, we can see sizeable differences in the occupational distribution of white males, white females, black and Hispanic males, and black and Hispanic females (Table 1).

finer occupation breakdown. Therefore, in our sample occupational segregation should show up as a much stronger determinant of differences in earnings between white males and other groups, and differences in pay for the same work a much weaker force, as the results in Table 4 tend to support.

This argument would also help explain the findings of Chiswick (1974) of a flatter age-earnings profile for non-white males than white males, of Hanushek (1973) that the mean regression coefficient of experience for blacks was about half that of whites, and of Hanushek (1973) and of others of a smaller rate of return to schooling for blacks than whites. These results would occur if discrimination forced blacks into those occupations where schooling and experience were unrelated to earnings, which are precisely the low-paid occupations. Chiswick argued that the flatter age-earnings profile of non-whites was probably due to their smaller post-school investment. Yet their smaller "investment" may be due to their placement in those occupations where training is irrelevant. Welch (1973) found that the returns to schooling for those entering the work force in 1960 were greater for blacks (compared to other blacks) than for whites (compared to other whites), while for those entering the work force in the 1930's and 1940's the returns were less for blacks (and in some cases negative). This would be the case if the more educated blacks in the younger cohorts now have access to the upper part of the occupational ladder while the less educated are forced into their "traditional" slots at the bottom of the occupational ladder. This hypothesis might warrant further investigation in the future.

II. Conclusion and Comparison with Other Studies

One way of summarizing our findings is to look at the percent of the inter-occupation variance in earnings inequality explained by the factors we have used in our model (Table 2). The coefficients of determination (R^2) vary between .50 and .91. The explained variance is highest for the clerical group in both years, followed by skilled and service workers, and then by professionals and operatives. If the residual is primarily due to the structured and institutional characteristics of the occupation, then these characteristics seem strongest for professionals and operatives and weakest for clerical workers. Between 1960 and 1970, the R^2 declined in each group, except skilled workers, suggesting that institutional forces became more important during the decade.

The "human capital" factors, schooling and age, were strongly significant for the professional and clerical groups, marginally significant for skilled labor, and almost entirely insignificant for operatives and service workers. This suggests that human capital factors may be relevant to perhaps half the labor force. The fact that the human capital factors are highly significant for the pooled samples must be construed as an artifact of aggregation, where the significant relations between earnings, schooling and age in the professional and clerical groups dominate the random relations in the other groups. Moreover, the fact that Mincer (1974) and Chiswick (1974) obtain significant coefficients for schooling and experience across the entire labor force is likewise due to the aggregation of relevant with non-relevant groups.[17]

An alternative explanation for our results to one that might be inferred from a human capital model is that the distribution of earnings slots is fairly well fixed within occupation and that the occupational distribution (and by implication the industrial composition) is the primary determinant of the overall distribution of earnings. This interpretation is consistent with Soltow's (1960) early findings that historically the change in income inequality in the United States was due more to the redistribution of workers over occupations than their distribution over schooling level or age. It is also consistent with Hanushek's (1973) finding that differences in industrial and occupational structure among regions account for over 80 percent of the variance in mean earnings across regions and to Osberg's (1975) finding that once controlling for differences in structure between counties, personal characteristics like age and education are unimportant in explaining differences in mean earnings and earnings inequality across counties.

The function of schooling would then be to "rank" individuals within a relatively well defined set of occupational slots. This is consistent with our finding that the secular change in the schooling distribution between 1960 and 1970 had relatively little impact on occupational earnings inequality. What appeared to happen instead was that the better educated new entrants in the labor force "filtered down" the occupational ladder as professional and clerical positions

[17] Actually, their samples are restricted to males between 25 and 65, to white males between 25 and 65 in some cases, and white non-farm adult males in other cases.

closed up to the higher paying occupations within the skilled, operative, and service groups. This would explain the significant relation between mean earnings and mean education across occupations in these latter three groups in 1970 and their insignificant relation in 1960. However, the more highly educated new entrants in these occupations apparently did not receive greater wages than their less educated co-workers, since the increased mean schooling in these groups still had an insignificant impact on earnings inequality. It appears instead that the educational requirement for a given job in these occupational groups was inflated during the 1960's to meet the composition of the new labor force, but the tasks remained relatively unchanged and pay remained commensurate with the task.[18] Thus, it does not appear that, given a structure of occupational tasks, education is intrinsically productivity-augmenting. Rather, it appears to serve as a ranking device for prospective employees throughout much of the occupational ladder.

The effect of experience on earnings would also vary among occupations, depending on the characteristics of the occupation. For professional, clerical, and some skilled occupations Mincer's (1962, 1974, 1975) argument that differences in post-school investment in training leads to differences in earnings is consistent with our findings. Moreover, special studies of professional groups by Link (1973) for chemical engineers, Katz (1973) for university faculty,

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[18]From our personal experience we know that in a public utility company in New Jersey an appliance repairman did not require a high school diploma in 1965 but he did in 1970, though the job changed very little over this period.

Johnson and Stafford (1974) for academic economists, and Klevmarken and Quigley (1976) for Swedish engineers all show a positive relation between earnings and experience. However, for operative, service and certain skilled occupations age (and by implication experience) seems to have little bearing on earnings.

Our finding that earnings inequality decreases with mean age within an occupation is not inconsistent with Chiswick's (1974) finding that the variance of age is positively related to earnings inequality across states, to Mincer's that the relative dispersion in gross earnings rises as a schooling cohort ages until peak earnings are reached, or to a previous finding that the Gini coefficient fell for age cohorts between 18 and 30 and rose thereafter (Wolff (1975)). These findings can be reconciled if the occupational composition shifts as a cohort ages and, in particular, if the more able, more ambitious or better schooled move up the job ladder to the more technical or administrative positions while the less able, less ambitious, or less schooled remain "stuck" in their entry-level occupations. This is consistent with the findings of both Hause (1972) and Taubman and Wales (1973) of increasing impact of ability on earnings with age. This interpretation is similar to one offered by Becker (1975, p. 217) to explain the earlier peak of earnings with age for less skilled than more skilled workers. This interpretation is also consistent with Alexander's (1974) finding that firm experience is a more important determinant of earnings than age for low and medium income workers, but a less important one for high income workers. This would be the case if the high income workers can move up the occupational ladder faster by switching firms, whereas the low and medium income workers, as Alexander argues, start at the bottom of the ladder when they switch firms.

The attenuation of the impact of schooling on earnings with age, observed by us, Chiswick (1974), and Mincer (1974), can also be explained by this model. Suppose that entry into occupations in the top part of the ladder is set by schooling level and that in the bottom part is relatively free and that entry-level wages are relatively fixed in an occupation. As the cohort ages, promotions and raises will be based on performance, and the more able, better educated and more ambitious will obtain better positions and receive higher earnings. (See Taubman and Wales (1973) for a similar argument.) However, since post-school training and ability play a more important role in the later working life, the effect of schooling on earnings will decay over time. This model can help explain Mincer's (1974) findings that earnings inequality is less for the highly educated than the less educated early in working life, but the order is reversed later in working life, results which Mincer says "are in no obvious way related to secular trends in human capital, such as the upward trend in schooling" (pp. 108-9). At the entry level stages, the highly educated are more concentrated in the occupational ladder than the less educated and thus have a smaller relative dispersion of earnings. As the cohort ages, the less educated remain in their initial occupations and, relatively speaking, their earnings tend to converge. Among the more educated there will be greater opportunity for some to move up the occupational ladder and the earnings within this group will tend to diverge with age.

The alternative argument we have presented is, at this stage, simply a sketch of a model in which the function of schooling and experience can be understood in the context of an exogenously determined occupational structure. We hope this work will lead to further development of the argument by ourselves and other researchers.

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