REGIONAL DIFFERENCES IN THE STRUCTURE OF EARNINGS

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THERE have been a number of attempts to estimate the relationship between schooling and earnings of individuals, but the most common feature of these studies has been severe data limitation which, in turn, has dictated how the analysis could proceed. Perhaps the most serious restriction imposed by the data has been the assumption that earnings relationships are the same across the nation or, at least, across very sizable aggregations of states. This paper examines the viability of such assumptions by looking at differences in earnings functions among smaller, more homogeneous labor markets. This research reveals large differences across labor markets in the returns to human capital and indicates that much of the observed difference in regional and racial earnings results from structural differences in earnings functions.

I Models and Data

This study, as with its predecessors, relies upon a very simple model of individual earnings

$$\text{Earnings} = f\ (\text{education, ability, experience}). \tag{1}$$

Nevertheless, adequate data for analyzing even such simple earnings functions have not been plentiful. Representative samples, such as those from the United States censuses (e.g., Becker (1964) and Hanoch (1967)), do not adequately measure important quality differences in individual ability or schooling. But, samples which provide more accurate measures of embodied human capital tend to be nonrepresentative of the whole population (e.g., Hansen, Weisbrod and Scanlon (1970) and Griliches and Mason (1972)). This study falls into the second class — one with a relatively rich description of characteristics of individuals in a specific subset of the entire labor force.

From a survey of all enlisted men departing from the U.S. Army during the fiscal year (FY) 1969, a sample of slightly over 180,000 individuals was formed by selecting individuals who (1) had been in the Army two years or less,[1] (2) had completed the survey with respect to income, and (3) had been working full time.[2] The survey provided data on marital status, employment, occupation, and wages, ten months after separation from the Army. These data were merged with service information about education, Armed Forces Qualification Test (AFQT) score, age, race, military occupation, and home of record of the individual.[3]

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[1] Limiting the sample to a uniform length of service minimizes problems arising from variations in length and depth of military training (about which there is no information) and makes the sample more representative of young males in the labor market.

[2] From the sample of those meeting the first criteria and connected with the labor force, 92 per cent were working full time, 2 per cent were working part time, and 6 per cent were unemployed but looking for work. Separate analyses of unemployed individuals indicate that the probability of being unemployed is independent of the measured human capital attributes used in the earnings functions here. Further, the relationship between mean regional earnings (for the urban regions described below) and the unemployment rate in the regional samples was statistically insignificant at the 10 per cent level (and even slightly negative in value).

[3] Sampling distortions could be introduced by either systematic nonresponse to the survey or by selection biases in the military draft. (For the entire sample, the survey response rate was 73.9 per cent.) Because of these factors, the variance of schooling and ability scores is less in the sample than in the population as a whole, but the mean values of schooling completed (12 years), AFQT percentile (53.5) and earnings ($7,031 annually) do not appear unreasonable for this age group. Further, both ends of the educational and ability distributions appear adequately represented, insuring a good data base for investigating the underlying earnings structure as long as any sampling biases are not correlated with important but unmeasured earnings characteristics. A detailed discussion of sample and population characteristics can be found in Hanushek (1973).

The actual model estimated was

$$\log Y = \beta_1 + \beta_2 S + \beta_3 AFQT + \beta_4 E + \epsilon \quad (2)$$

where Y is weekly earnings, S is years of schooling, $AFQT$ is the percentile score of the individual, and E is experience.[4]

The relative richness of the sample does allow for testing several interesting extensions of the standard human capital hypotheses. It is likely that many sampled individuals are in some training status, and, as has been developed previously (Mincer, 1962), people undergoing training by the firm would be expected to receive lower wages. The extent of this can be tested, albeit crudely, since the survey recorded whether or not individuals were undergoing training at the time of survey. Information indicating the military occupation of the individual (roughly a one-digit occupational breakdown) allows testing for differential transferability of various military skills.[5] Further, data on the civilian occupation of the individual at the one-digit level permit analysis of occupational differences which exist over and above human capital or training differences. Such differences could exist through differences in the monopoly position of either labor suppliers or demanders.

Finally, a central part of this analysis concerns the homogeneity of labor markets throughout the country. For data reasons past analyses have made the very strong homogeneity assumption that returns to human capital are equal over very large regions. The common treatment of geographic location within the country has been to use large region intercept dummy variables (Griliches and Mason, 1972) or to stratify on very large regions (Hanoch, 1967). These crude techniques always display large, significant differences in earnings by regions and provide little reason to believe that there has been sufficient accounting for differences in labor markets across the country. There is also direct evidence that there are significant differences, at least in rural areas, by states (Welch, 1965).

With the knowledge of the current residence of each individual in the sample, individuals can be divided into fairly precise regions or labor markets. For blacks and whites separately, the criteria for forming urban regions were as follows: (1) except in the South, all Standard Metropolitan Statistical Areas (SMSA's) with over 200,000 people in 1960 and with 75 or more sample observations were considered separate regions; (2) in the South, all SMSA's with over 75 sample observations were considered regions. Remaining areas (not regions by (1) and (2) were grouped by states into 24 "rural" regions, provided there were 75 or more sample observations. This stratification provided 24 rural white, 126 urban white, 16 rural black and 27 urban black samples from which separate regression models are estimated.[6]

[4] Earnings defined as both hourly and weekly wages and in both linear and logarithmic form were considered with equation (2) providing the best estimates.

One concern in using weekly earnings instead of hourly earnings is that aggregate economic conditions could contaminate the model. The observations were recorded over a twelve-month period of changing economic conditions. The unadjusted national unemployment rate changed from a low of 2.9 per cent to a high of 4.7 per cent. In order to allow for this possible factor, models were estimated which included, as one of the independent variables, the unadjusted national unemployment rate during the month in which the individual answered the survey. This was invariably insignificant according to traditional statistical tests. Other analyses of these data indicate that whether or not a person is employed depends upon the aggregate conditions. These models indicate no sensitivity of earnings to aggregate conditions, given that the individual is employed. An alternative mode of analysis would be the development of structural models for both hourly earnings and hours worked; this seemed much more difficult and also would require more data — particularly in the hours-worked model.

Actual labor force experience of individuals is unknown. Here the assumption was made that all time outside of school and the Army was spent in the labor force. The models were actually estimated using age instead of E, and the coefficients were transformed to equation (2) from using the relationship $E = \text{Age}-S-8$. This assumption is partially relaxed below.

[5] The data file does not contain information on length of training within each occupational field. An assumption must be made that individuals within the same one-digit field receive equal amounts of training in order to use this information. This is probably not a bad assumption for first-term draftees but becomes increasingly tenuous as individuals are in the military for longer periods of time.

[6] Before applying the sample size criterion, there was the possibility of 165 regions for both whites and blacks. Initial analyses of the data indicated that a fairly large number of individuals is needed within a sample in order to obtain good parameters estimates; that is, small regions may contain too little variation in the observed characteristics of individuals. A rather arbitrary sample size cutoff of seventy-five observations was placed upon the individual regional samples, resulting in the elimination of two per cent of the individuals in the sample. The effect of the sample size limitation was more pronounced in the case of blacks. For that group, the loss rate was nineteen per cent,

II Aggregate Characteristics

For descriptive purposes, the set of regional analyses is divided by urban and rural location and black and white individuals. An overall summary of the performance of the models is displayed in table 1. The data set for this analysis contained, relative to most past samples, a much richer view of the individuals in the sample; but even so, the results are far from overwhelming. The simple human capital models — both here and elsewhere — fail to explain as much as half of the variation in incomes.

In table 1, the variation in the log of individual weekly earnings is partitioned into the proportion of individual variation between regions (column 1), or the variation in regional means, and the proportion within regions (column 2). The proportion of the within region variance which is explained by the set of regional earnings regression models is displayed in column 3 of the table.[7] Finally, the sum of column 3 times 2 and column 1 yields a crude estimate of the aggregate explained variance since column 1 is the amount "explained" by stratifying the sample into regions, column 3 is the amount explained in each of the regions, and column 2 is amount of within region variance which could be explained by the regression models.

as compared with one-half per cent for whites. Further, data problems caused the elimination of San Francisco and Philadelphia from the black analysis.

[7] This is calculated as the total explained sum of squares over the total sum of squares for all of the regions. Thus, it is a weighted average of the R^2's in the regional models.

As indicated by column 4, the regional models explain between an eighth and a quarter of the variance in incomes in each aggregation.[8] This actually sets an outer bound on our knowledge of earnings relationships because about one half of the explanatory power lies in the regional division of the sample. The causes of such regional variations in mean income are not well understood — although a later section of this paper does delve into part of the explanation.

III Characteristics of the Earnings Functions

Was it necessary to go to the fineness of regional definition used here, or were more aggregate regions of the type used in the past satisfactory? This question was looked at carefully during the course of analysis, and it is clear that this detail is warranted. The appropriate tests for equality of coefficients were applied to aggregate regions, and homogeneity within broad regions was consistently rejected at the one per cent level even when each of the

[8] Models from this sample of new entrants into the job market would be expected to possess lower explanatory power than ones from other conceivable samples taken later in individuals' work profiles. Since the survey information applies to a time ten months after separation from the Army, the earnings figure almost certainly contains a sizable transitory component. This would have the effect of increasing the unexplained error in the models.

The ten-month period does seem long enough to minimize the high transitory component of earnings which would exist if the sample included many people in temporary jobs while waiting entrance into school. After ten months, one would not expect many still waiting to enter school. (Full-time students have been excluded from the sample so there is no contamination from temporary jobs held by students.)

TABLE 1. — ANALYSIS OF VARIANCE OF INDIVIDUAL EARNINGS [a]

			Proportion of Variance			
Grouping	No. of Regions	No. of Individuals	Between 1	Within 2	Explained 3	$4=1+3\cdot 2$
All	193	180,330	.158	.842	.091	.235
Urban	153	72,882	.135	.865	.118	.237
Rural	40	107,448	.126	.874	.079	.195
White	150	168,069	.067	.933	.092	.152
Urban	126	65,599	.083	.917	.122	.195
Rural	24	102,470	.046	.954	.080	.122
Black	43	12,261	.195	.805	.068	.250
Urban	27	7,283	.129	.871	.074	.193
Rural	16	4,978	.158	.842	.060	.209

[a] Individual regional models for the log of income from which this table is derived are displayed in Hanushek (1973).

micro regions of this study was allowed to have its own intercept.[9]

According to traditional significance tests, there are mixed levels of significance for individual parameter estimates. (See appendix table A–3.) The schooling coefficients are consistently well estimated with only 17 out of 193 coefficients having a t-statistic less than 1.67. For whites, the estimated experience coefficients are also quite precise while experience estimates for blacks and the ability coefficients are less precise. A large part of the imprecision that does arise in the estimates appears to arise from lack of enough variation in the inputs; larger samples — ones with more observed independent variation in the exogenous variables — consistently have more precise estimates. The trade-off of imprecision for the larger numbers of regions seemed warranted.[10]

The difference in earnings structure by labor market implies that it is not possible to present a single point estimate of the marginal importance of additional inputs. It is really necessary to think in terms of a distribution of returns dependent upon geographical location. In order to place the models in perspective, however, mean parameters are presented for varying aggregations of the models (urban/rural, black/white).

Schooling: The majority of attention in earnings analysis has naturally gone to the return to formal schooling. The distribution of the individual schooling parameter estimates and coefficient means based on the sample distribution of individuals are displayed in table 2. In the semi-log models estimated, the schooling coefficient (times 100) can be interpreted as the percentage increase in earnings associated with one additional year of schooling.

The mean schooling coefficient of 0.049 implies that an additional year of schooling is worth 345 dollars per year to the average individual in the sample. The corresponding figures for whites and blacks are 348 and 291 dollars, respectively. (The advisability of additional schooling for any individual depends upon this return, plus his discount rate and the precise shape of his earnings profile. This calculation is not considered here.)

While we would like to have "ability-free" estimates of the returns to schooling, these estimates cannot be interpreted strictly in this manner. If more able or more motivated individuals tend to continue longer in schooling and these abilities or motivations lead to increased earnings, estimates of earnings as a function of only schooling and experience would overstate the return to schooling. But, the *AFQT* percentile, included in the models to

[9] The covariance, or Chow, tests used are described in Fisher (1970). The country was divided into seven aggregate regions. Separate covariance tests for urban and rural, white and black were performed for each of the seven regions described in section IV. As an example of the F values, allowing for different regional intercepts, the core South had the lowest F-statistic for the seven rural white tests with $F(30, 13050) = 2.05$. (If constrained to the same intercept, the same test yields $F(36, 13056) = 14.61$.)

[10] As an example of the effect of sample size, if the additional 16 black regions with between 75 and 125 observations were eliminated, the number of coefficients with t-statistics less than 1.0 would go from 2 to 1 for schooling, from 24 to 14 for *AFQT* score, and from 16 to 7 for age. Similarly, the t-statistics between 1.0 and 1.67 would go from 7 to 5, 11 to 8, and 9 to 5 for the respective variables. While increasing the size cutoff eliminates only 11 per cent of the number of black observations, it eliminates 36 per cent of the black regions.

TABLE 2. — FREQUENCY DISTRIBUTION AND MEANS OF SCHOOLING COEFFICIENTS

Grouping	Value <.034	.035 – .049	.050 – .064	.065 – .079	>.080	Mean [a]
All	35	54	61	28	15	.049
Urban	28	42	43	25	15	.052
Rural	7	12	18	3	0	.046
White	27	39	48	22	14	.049
Urban	24	31	36	21	14	.053
Rural	3	8	12	1	0	.046
Black	8	15	13	6	1	.047
Urban	4	11	7	4	1	.050
Rural	4	4	6	2	0	.043

[a] Mean is weighted by number of observations in each region.

compensate for these factors, surely includes some achievement that resulted from schooling and does not accurately portray motivational factors.

More precisely, decomposing $AFQT_i$ into a school component and a nonschool component and, as supported by Coleman et al. (1966) and Hanushek (1972), letting the relationship between achievement and schooling depend upon the quality of the school attended by the individual, we can then represent $AFQT$ as

$$AFQT_i = a_0 + a_1 AFQT^*_i + \delta_j S_i + \epsilon_i \qquad (3)$$

where $AFQT^*_i$ is the nonschool component, δ_j is the quality coefficient for the j^{th} school and ϵ_i is a stochastic component.[11] If the $AFQT$ coefficient in the earnings model is β, $\beta\delta_j$ should be added to the estimated returns to schooling but is instead attributed to the $AFQT$ variable. However, assuming that length of schooling is independent of school quality, models without $AFQT$ attribute $(\beta a_1 r + \beta\delta)$ to schooling where r is the correlation between $AFQT_i$ and S_i and δ is mean school quality. (Variation due to differences in regional school quality is ignored when $AFQT$ is excluded from the model.) Rough bounds on the returns to schooling are found by estimating models with and without $AFQT$ as displayed in table 3.[12]

TABLE 3. — MEAN SCHOOLING COEFFICIENTS WITH AND WITHOUT *AFQT* VARIABLE

Grouping	With *AFQT* (1)	Without *AFQT* (2)	(1)/(2)[a]
Total	.049	.057	.86
Urban	.052	.059	.88
Rural	.046	.055	.83
White	.049	.057	.86
Urban	.053	.060	.89
Rural	.046	.056	.82
Black	.047	.049	.96
Urban	.050	.052	.97
Rural	.043	.045	.96

[a] In the age adjusted form of the model (instead of experience adjusted), the ratio of coefficients with and without *AFQT* is 0.71, 0.77, and 0.67 for Total, Urban and Rural, respectively.

The consistency of the estimated relationships at the extremes of the educational distribution was tested through the introduction of intercept dummy variables for individuals with a college education or more and individuals with less than a high school education. Neither of these variables proved to be significantly different from zero.

Ability: Ability and achievement differences of individuals were measured by Armed Forces Qualification Test (*AFQT*) percentile scores. The coefficient estimates for this variable are small and consistently less precise than those for the other variable in the model. This could arise from a number of sources. First, cognitive ability and achievement could have little or no impact on the earnings.[13] Alternatively this could be a very poor measure of the ability quantity which is important; the test could be unreliable (i.e., a large sampling error of the test) or the test could be invalid (i.e., it doesn't measure what it is intended to measure).[14] Finally, this test could be a reasonable measure of productivity, but employers could lag in adjusting wages to these differences. For example, employers could hire on the basis of known characteristics (schooling and experience) and promote on the basis of ability. With the short work history in this sample, the ability effect may not be adequately observed. Within these data, it is not possible to distinguish adequately among the competing explanations.

The distribution of estimated coefficients (displayed in appendix table A–1) indicates less variance within groupings in these coefficients than in the schooling coefficients and thus less worry about aggregations. Further, the black coefficients are consistently about one half the magnitude of the white coefficients.

These estimates, which are very similar to the basic estimates of Griliches and Mason (1972), appear quite small.[15] They indicate that a decile change in position in the test score

[11] For a description of the *AFQT* tests, see Karpinos (1966). In this development, we make a rather strong assumption that there is no interaction between these components.

[12] Note from table 3 that in the larger rural regions, where school quality is probably more heterogeneous, the reduction in the schooling coefficient from including *AFQT* is larger than in the more homogeneous urban regions.

[13] This hypothesis is developed in Gintis (1971).

[14] This hypothesis and an attempt to deal with it are contained in Griliches and Mason (1972).

[15] This is considerably different from the findings of Hansen, Weisbrod and Scanlon (1970) for low achievers. They found that including an achievement measure reduced the schooling coefficient to insignificance. The Griliches and Mason instrumental variable estimates are also larger.

leads to only a one per cent change in white earnings or a one-half per cent change in black earnings. This implies that one to two years of additional schooling is equivalent to moving the entire range of the ability scale in terms of the change in earnings.

Tests were also made for the consistency of the ability relationship across the entire range. Both a series of dummy variables for different percentile intervals and intercept dummy variables for the top and bottom ranges in conjunction with the continuous *AFQT* variable were insignificant.

Experience: The measure of experience of the individual has a consistently strong effect on earnings, especially earnings of whites. Interestingly, the mean white experience coefficient (.029) is over double that for blacks (.013) (see appendix table A–2). Part of this could arise from age being an imperfect measure of work experience. (Note that this variable is not experience on the current job but lifetime experience levels.) Since the unemployment rate for black teenagers is considerably higher than that for white teenagers (historically almost double), the same chronological age for a white and a black is not associated with the same average work experience level. However, using the alternative experience normalization:

$$E = P_i\,(AGE - S - 8) \tag{4}$$

where P_i is the teenage employment rate for blacks or whites and assuming an employment rate of 0.85 for whites and 0.70 for blacks (which roughly corresponds to conditions in 1964–1966) yield an adjusted mean experience coefficient for whites of 0.034 compared to 0.015 for blacks. This factor, therefore, does not seem to be the explanation of the differences in experience rewards.

Occupation: From analyzing intercept dummy variables for one-digit civilian occupations, it appears that being in agricultural or structural occupations significantly affects earnings. Within 23 of the 24 rural white regions and 9 of the 16 rural black regions (comprising 99.1 and 52.8 per cent of the total individuals), a significantly negative relationship between earnings and agricultural occupations was estimated.[16] The mean coefficients across all regions were −.216 for whites and −.145 for blacks and −.275 in black regions where it was statistically significant. This probably reflects measurement problems (when there is income in kind) as well as a depressed wage market.

In 56 urban white regions and 9 urban black regions (comprising 76.1 and 53.5 per cent of the total individuals), significant positive effects were estimated for the structural trades. Within these regions, the mean coefficients for whites and blacks were 0.085 and 0.110, probably reflecting labor supply restrictions by the building trade unions.

Formal Training: While virtually everybody in the sample is probably undergoing some sort of on-the-job training, being identified as in a formal training program exhibited a significant effect on earnings in 22 of the 126 urban white regions and 2 of the 27 urban black regions. Within these regions the average estimated effect of being trained was −.092 and −.086 for whites and blacks, respectively. This is, however, difficult to interpret without more detailed information about the terms of training.

Other Factors: As mentioned earlier, several other hypotheses about earnings functions were tested but proved to have an insignificant impact on earnings. Surprisingly, the military occupation of the individual — measured at the one-digit level — never displayed any independent impact on post-service earnings, although this probably reflects the low level of training which is provided inductees (as opposed to enlistees) into the Army.[17] The marital status and family size of the individual were also considered but rejected as a significant factor in determining earnings. Finally, although ten to twenty per cent of the individuals in each region entered the military service from a different region, no systematic differences between migrants and nonmigrants were

[16] The term rural, it should be remembered, has a special meaning in the context of the regions for this analysis. Rural refers to all land area left after removing the included SMSA's. In all but the South, this rural includes anybody not in an SMSA of 200,000 people or more. The percentage of rural individuals in agriculture is fairly low — 3.4 per cent for whites, 1.6 per cent for blacks.

[17] Since everybody served the same time in the Army, it is not possible to estimate the importance of Army experience relative to civilian experience. The estimates of Griliches and Mason (1972) imply that a year in the Army is worth virtually nothing in terms of civilian earnings.

found. This last finding supports the contention that the observed differences in returns to human capital are "pure" regional effects rather than further manifestations of the embodied human capital in each individual.

IV Regional Variations in Earnings

Almost 16 per cent of the observed variance in earnings results from differences in the mean earnings among regions. Are these mean earnings differentials simply a reflection of input differentials, or is the structure of earnings (the various model coefficients) the dominant factor?

The answer to this question comes from some manipulation of the expression for the variance of the means between regions. For the j^{th} regions, the mean earnings (E^j) equals the vector of mean inputs — S, $AFQT$, etc. — (X^j), times the vector of estimated coefficients (b^j), as in equation (5).

$$E^j = X^j b^j. \tag{5}$$

Let M be the vector of national mean levels for the inputs into (5),[18]

$$E^{*j} = Mb^j \tag{6}$$

so that E^{*j} in equation (6) would be the predicted mean earnings in region j with the national mean level of inputs. Then, letting NE equal the national mean earnings levels, n_j equal the number of observations in region j, and T equal the total number of observations, variance in the regional means equals

$$\frac{\Sigma n_j (E^j - NE)^2}{T} = \frac{\Sigma n_j [(E^j - E^{*j}) + (E^{*j} - NE)]^2}{T} \tag{7}$$

$$= \frac{\Sigma n_j (E^j - E^{*j})^2}{T} + \frac{\Sigma n_j (E^{*j} - NE)^2}{T} + \frac{2\Sigma n_j (E^j - E^{*j})\,(E^{*j} - NE)}{T}. \tag{8}$$

On the right-hand side in equation (8), the first term is the variance due to input differences, holding structure constant.[19] The second term is the difference in earnings due to differences in earnings structure by region with the input levels held constant.[20] The final term is an interaction component reflecting whether individuals with above average input levels tend to locate in regions that pay above average (+) or vice versa (−). Dividing through equation (8) by the total variance in mean earnings yields a proportion due to mean input differences (levels of education, etc.), a proportion due to structural differences in the earnings relationships (values of b^j), and a proportion due to the interaction of inputs and earnings structures.

The implications of this decomposition of variance, shown in table 4, are clear. The first

TABLE 4. — DECOMPOSITION OF MEAN EARNINGS VARIATION

Grouping	Variation Between Regions	Decomposition: Input Differences	Structural Differences	Interaction
Total	.158	.028	.994	−.022
Urban	.136	.055	.806	.139
Rural	.126	.018	1.042	−.059
White	.067	.034	.837	.129
Urban	.083	.040	.839	.121
Rural	.046	.018	.811	.171
Black	.195	.013	1.016	−.029
Urban	.129	.018	.979	.039
Rural	.158	.006	.956	.038

column is the proportion of the total variation in incomes which is explained by differences in the mean earnings levels among regions, while the remaining columns distribute this variance among different sources. In no case do input differences account for more than six per cent of the total variance in mean regional earnings. On the other hand, structural differences in the earnings relationships among regions account for over 80 per cent of the variance in mean earnings.[21]

This suggests quite strongly that more effort should be devoted to analyzing the structure of

[18] Both M and NE below refer to national means within the group being considered; for example, within black rural regions.

[19] The term $(E^j - E^{*j})$ can be rewritten as $(X^j - M)b^j$.

[20] The term $(E^{*j} - NE)$ can be rewritten as $M(b^j - b)$ where b is a vector of "national" coefficients such that $NE = Mb$.

[21] To be precise, this analysis applies to the variance in the log of earnings. For the total rural sample and for the total black sample, the variation that would result from structural difference alone is greater than the total variance in mean earnings, because there is a negative interaction term (high earners located in low paying regions) which suppresses the variance from what would be observed if individuals were located randomly.

TABLE 5. — WEIGHTED MEAN STRUCTURAL CHARACTERISTICS FOR MACRO GEOGRAPHIC REGIONS

Group and Parameter	Region						
	West	Central	New South	Core South	Appal.	Northeast	Gr. Lakes
Urban White							
Earnings [a]	$139	$138	$124	$121	$126	$132	$146
Education	.051	.053	.057	.076	.058	.057	.051
Experience	.034	.026	.030	.044	.032	.029	.034
Rural White							
Earnings [a]	$130	$122	$117	$112	$114	$128	$136
Education	.031	.051	.057	.052	.056	.044	.008
Experience	.023	.029	.027	.026	.029	.017	.003
Urban Black							
Earnings [a]	$132	$131	$112	$106	$115	$116	$137
Education	.039	.060	.047	.051	.056	.052	.044
Experience	.014	.007	.006	.008	.019	.014	.013
Rural Black							
Earnings [a]	..	..	$101	$95	$102	$124	$134
Education	..	..	.049	.043	.053	.038	.039
Experience	..	..	.022	.011	.014	.022	.012

[a] Geometric mean of weekly earnings within the region.

labor markets than looking at the distributions of individuals and their characteristics in analyzing regional income patterns. Studies which account for variations in regional incomes by variations in aggregate education and experience levels overlook more basic, structural differences in the labor markets within each of the regions.

Aggregate differences in the structural estimates can be seen better by grouping the estimated functions into seven macro geographic regions.[22] Table 5 displays mean weekly earnings and mean coefficients for education and age in each of the seven regions by race and place of residence. (The distribution of ability coefficients does not show much regional variance and has, thus, been omitted from the table. Remember, however, that there are significant racial disparities in them.) The earnings functions are slightly steeper in terms of schooling in the urban areas of a given region than in the rural remainder regions, and the returns to education in the three southeastern regions tend to be higher than elsewhere (even though there are large schooling quality differences between the South and elsewhere). Also, in 10 out of 12 comparisons, the marginal returns for an extra year of education are higher for whites than for blacks. While it is tempting to explain the regional differences in schooling returns by differential demands for skilled labor (as between urban and rural) and differential supplies of educated labor (north and west versus south), the complexity of such explanations requires considerably more analysis than is feasible here.

The pattern of experience parameters is not as consistent as the education parameters. The returns in experience tend to be slightly higher for urban areas than for rural areas and higher for whites than blacks. There are, however, several exceptions to these observations.

V Black-White Differentials

The overall picture of black-white earnings differentials from this sample does not look as bleak as that from national averages. In the sample, black earnings are 87.2 per cent of white earnings while the figures for urban and rural are 87.3 per cent and 82.0 per cent, respectively. Nationally in 1970, median income for black males age 25 and over was 61 per cent of the corresponding white income. Part of the relative improvement shown in the sample

[22] *WEST*: Wash., Ore., Calif., Nev., Ariz., Utah, Idaho, Mont., Wyo., Colo., N. Mex.; *CENTRAL*: N. Dak., S. Dak., Neb., Kan., Mo., Iowa, Minn.; *NEW SOUTH*: Texas, Okla., Fla., *CORE SOUTH*: Ark., La., Miss., Ala., Ga., S.C., N.C.; *APPALACHIA*: Tenn., Ky., Va., W. Va., Md., Del., D.C.; *NORTHEAST*: Pa., N.J., N.Y., Conn., R.I., Mass., Vt., N.H., Maine; *GREAT LAKES*: Wis., Ill., Ind., Mich., Ohio.

is due to a slightly more favorable geographic distribution and the restricted age range of sample blacks along with consideration of only employed individuals. Nevertheless, it is interesting to look at the differentials which do exist within the sample and attempt to identify the causes of these differentials.

Since both earnings structures and input levels differ between blacks and whites, separate estimates can be made of the effects of these differences. Within the 27 urban regions and 16 rural regions in which both black and white earnings models were estimated, two predictions were made: (1) mean black earnings from the black earnings models but using the white input mean characteristics for each region; and (2) mean black earnings from the white earnings models using mean black input characteristics.[23]

These predictions for black earnings along with the actual mean earnings for blacks and whites are displayed in table 6. For rural areas, if blacks had the same characteristics as whites in each region, the disparity in earnings would remain the same; however, if they could receive wages according to the white earnings structure (without changing any input characteristics), 56 per cent of the racial gap would be eliminated. In the urban areas, the predicted black earnings using white input levels reduces the earnings disparity from 0.87 to 0.89. However, receiving the same reimbursement for their input characteristics as whites increases the earnings ratio to 0.96, or 69 per cent of the earnings differential.

The picture is clear. The largest cause of differences in earnings between blacks and whites is a difference in the rates of reimbursement for skills and abilities (as reflected by education, *AFQT*, age, training status and occupation). Although blacks have lower schooling levels, lower *AFQT* levels and lower levels of participation in the high paying construction industries, these factors do not account for much of the difference in earnings.

Table 6. — Actual and Predicted Mean Earnings

	Actual		Predicted Black Earnings			
	White (1)	Black (2)	Black Struct. White Means [a] (3)	White Struct. Black Means [b] (4)	(3)/(1)	(4)/(1)
Rural	$126	$103	$103	$116	.82	.92
Urban	138	120	123	133	.89	.96

[a] Predicted geometric mean of weekly earnings using the black earnings function for each region and the mean values of white inputs in that region.
[b] Predicted geometric mean of weekly earnings using the white earnings function for each region and the mean values of white inputs in that region.

VI Conclusions

This analysis indicates that the value of education or other inputs cannot be described by a single statistic but instead appear to be a function of the geographical area in which the individual lives. Considering major metropolitan areas as separate labor markets, one finds significant variation in the returns to human capital across labor markets. This implies that past analyses of the returns to schooling, ability and experience will be very dependent upon the geographic distribution of the individuals in the sample and, thus, upon the specific aggregation of relationships for different labor markets.

The effect of structural differences in the earnings functions by labor markets is dramatic: within the sample over 80 per cent of the differences in mean earnings among labor markets is attributable to differences in earnings structure as opposed to differences in input means among regions, and the choice of region is equivalent in many cases to the marginal earnings of several years of schooling.

At the same time, differences in earnings by race appear to arise fundamentally from differ-

[23] The estimates here all relate to geometric mean weekly earnings. They are weighted by the sample distribution of the black population. Looking at the reverse situation of the decrease in white earnings associated with black mean inputs and black earnings structure but weighting by the white population distribution makes only very slight changes in the predictions.

ences in the earnings functions for blacks and whites. Within the sample, virtually none of the racial difference in earnings is accounted for by differences in schooling, ability or experience levels. In terms of individual coefficients, the schooling estimates for blacks are slightly less than those for whites, while the estimates of *AFQT* and experience effects are dramatically less for blacks.

Because of the nonrepresentativeness of the sample, these results require further confirmation. However, they suggest strongly that more attention be given to disaggregated structural models which incorporate differences in individual labor markets.

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APPENDIX

TABLE A-1. — FREQUENCY DISTRIBUTION AND MEANS OF ABILITY COEFFICIENTS

Grouping	Parameter Value <0	0 – .001	.001 – .002	.002 – .003	>.003	Means
Total	27	64	72	24	6	.0012
Urban	26	55	52	14	6	.0009
Rural	1	9	20	10	0	.0013
White	17	47	59	22	5	.0012
Urban	17	45	45	14	5	.0010
Rural	0	2	14	8	0	.0013
Black	10	17	13	2	1	.0005
Urban	9	10	7	0	1	.0005
Rural	1	7	6	2	0	.0005

TABLE A-2. — FREQUENCY DISTRIBUTION AND MEANS OF EXPERIENCE COEFFICIENTS

Grouping	Value <.020	.021 – .040	>.040	Mean
Total	66	88	39	.028
Urban	52	64	37	.028
Rural	14	24	2	.027
White	32	79	39	.029
Urban	29	60	37	.030
Rural	3	19	2	.028
Black	34	9	0	.013
Urban	23	4	0	.013
Rural	11	5	0	.014

TABLE A-3. — DISTRIBUTION OF *t*-STATISTICS

Coefficient	$t<1.0$	$1.0<t<1.67$	$t\geq 1.67$
White Models			
Schooling	4	13	133
AFQT	51	30	69
Experience	8	7	135
Black Models			
Schooling	2	7	34
AFQT	24	11	8
Experience	16	9	18