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I. INTRODUCTION

Most research on changes in wages and earnings over the life cycle has concentrated on the period during which human capital is rising: the years of full-time schooling, entry into the labor force, and prime working years. The typical model, as developed in Mincer (1974), Brown (1976), and Hanushek and Quigley (1978), tries to explain the individual's decision to accumulate skills, and the effect this has on his wage rate or earnings. Depreciation rates, if they are mentioned at all, are often not estimated. The main exceptions to this have been studies by Johnson (1970), Johnson and Hebein (1974), Heckman (1975) (1976), Haley (1976), and Rosen (1976). They have estimated gross depreciation rates as part of ambitious and only partially successful attempts to model the individual's optimal accumulation problem, first examined by Ben Porath (1967).

While these authors have estimated gross depreciation rates ranging from 0.6 to 13.3 percent per year, they have not provided estimates of net depreciation for the years immediately prior to retirement. Moreover, except for Heckman (1975), their estimates have been based on annual earnings or income rather than hourly wage data. Blinder and Weiss (1976) and Heckman (1976) have hypothesized, and Parsons (1977) has found that annual hours fall with age for men prior to retirement. Therefore, earnings will fall faster than wage rates and depreciation rates based on earnings will overstate the decline in earnings capacity. Thus, because gross rather than net depreciation rates are estimated, and because earnings rather than wage data is generally used, little is known about the actual pattern of wage rates towards the end of the life cycle.

To measure wage decline among older men, this paper will estimate
net depreciation rates of earnings capacity from pooled time series-cross
section wage data. It will also examine whether or not the depreciation
rate is a function of the level of education, and whether increasingly
poor health contributes to the effect of aging on wage rates. By using
panel data, it will also be possible to see if the general rise in wage
levels outweighs the relative decline that occurs with age, or whether
real wage rates actually fall with age. Section II develops a simple
model for estimating the rate of depreciation of earnings capacity and
the age at which net depreciation begins, Section III presents estimators
for pooled data when the number of observations is not identical for all
individuals in the sample, and Section IV discusses the results.

II.   ESTIMATING NET DEPRECIATION

Earnings capacity may decline with age for several reasons.
If acquired human capital depreciates over time because of technical ob-
solescence, then the stock will tend to fall unless continued gross in-
vestment exceeds this gross depreciation. Ben Porath (1967) has shown
that as the individual approaches retirement and the repayment period for
new investment shrinks, gross investment will decrease. At some age prior
to retirement, gross depreciation will exceed gross investment, and earn-
ings capacity will decline.

The standard human capital model, in which earnings depend on ex-
perience rather than on age, is based on the assumption that work life is
constant across education groups. Mincer (1974, p. 8) presents estimates
of retirement age that rise with the level of schooling. However, in the
data set used here, the median age of retirement (the youngest age at

which half the sample was no longer in the labor force) was 65 regardless of race or education. Therefore, the net depreciation rate and the year in which it is first positive depend on age rather than experience.

A second source of depreciation in earnings capacity is the decline in physical and mental abilities that may or may not occur with age. Clark, Kreps, and Spengler (1978) cite several studies which argue that the correlation between age and productivity is weak and should not be used to justify mandatory retirement or similar age-based policies. Nevertheless, on average, declining wage rates among older men may reflect declining strength or acuity. Kalish (1975), Troll (1975), and Fogarty (1975) cite evidence on the decline after age 50 of a large variety of abilities, including hearing, lung capacity, manual dexterity, short term memory, peripheral vision, and basal metabolism. While it is not clear how all of these affect productivity on the job, they may be correlated with abilities which do.[1]

To focus on the net depreciation of earning capacity as a function of age rather than on its accumulation as a function of experience, we specify a model in which an individual has a certain level of human capital (K_0) which begins to depreciate after a certain age (t_0). For $t > t_0$

$$(1) \quad K_t = K_{t-1}(1 - \delta) \, .$$

Until that age, the stock is fixed, but after that it depreciates at a constant rate (δ) each year until retirement. By recursion

$$(2) \quad K_t = K_0 \qquad\qquad\qquad t \leq t_0$$

$$= K_0(1 - \delta)^{(t - t_0)} \qquad t > t_0$$

The observed wage (y_t) is equal to the capital stock times the rental rate on human capital (R) times other measured $(e^{\beta X})$ and

unmeasured (e^u) factors reflecting differences among individuals and their labor markets. The rental rate is the same for all individuals in all time periods, the other characteristics are not affected by the wage rate, and the error term is normally distributed.

$$(3) \quad y_t = R \, K_o (1 - \delta)^{(t - t_o)} e^{\beta X + u}$$

If we assume that the initial capital stock is some function of the individual's schooling (S), for example

$$(4) \quad K_o = e^{\alpha S}$$

then we can substitute (4) into (3) and take logs to obtain

$$(5) \quad \log y_t = \log R + \beta X + \alpha S + (t - t_o) \log (1 - \delta) + u.$$

Since $\delta \cong \log(1 - \delta)$ when δ is small, the depreciation rate can be estimated from the age coefficient of an OLS regression of (5), with aging beginning at t_o.

This simple model is suitable for analyzing the wages of age groups with flat or declining profiles, though obviously it is inappropriate for groups whose human capital and wages are growing. It suffers from the same identification problems that most accumulation models have, namely, that gross investment and depreciation are not separately identified. Those models can measure only net investment, while the present model can measure only net, not gross, depreciation. However, for most purposes, questions about the decline of wages and earnings of older men concern net depreciation rather than the gross depreciation that would occur in the absence of any investments.

The sample used for estimation includes observations on men 45 to 64 years old. It is hypothesized that earnings capacity begins to depreciate at some age t_o, but there is no reason to believe that t_o equals 45. To estimate t_o, depreciation was constrained to 0 for $t \leq t_o$ by setting t equal to t_o for $t \leq t_o$. This new age variable was included in the wage regression (5). The intercept and age terms were thus

$$
(6) \quad
\begin{array}{ll}
\delta_o - \delta t & t > t_o \\
\delta_o - \delta t_o \equiv \gamma & t \leq t_o .
\end{array}
$$

Since (6) is constant for $t \leq t_o$, $\delta_o = \gamma + \delta t_o$. Substituting this into the first part of (6) shows that this specification is equivalent to $\gamma - \delta(t - t_o)$, which is the desired form for estimating (5). The wage regression was then run for different values of t_o. The value which maximizes R^2 is the maximum likelihood estimator of t_o, given the standard normality assumptions concerning the error term.

As a check on the specification in (6), an alternative functional form which did not constrain the age-wage profile to be flat for $t \leq t_o$ was used. See Gujerati (1978). The intercept and age terms were

$$
(7) \quad
\begin{array}{ll}
\gamma_o + \gamma_1 t + \gamma_2(t - t_o) = \gamma_o - \delta t - \gamma_2 t_o & t > t_o \\
\gamma_o + \gamma_1 t & t \leq t_o
\end{array}
$$

where $-\delta = \gamma_1 + \gamma_2$.

Mincer (1974, p. 22) has suggested that depreciation rates may vary with age, education, and other factors. Indeed, it has become common practice to estimate human capital models separately by education category.[2] Depreciation rates can only vary by education if human capital

is not homogeneous. If workers differ only in the level and not the composition of their capital stock, rates in change of this stock, and in wage rates, will not vary. Even in Mincer or Ben Porath models which permit accumulation of human capital and in which interest rates and parameters of human capital production functions might vary by education category, depreciation rates will not vary as long as there is only one type of human capital. Thus testing hypotheses that depreciation rates differ implicitly assumes that different types of human capital exist, and that the composition of human capital is systematically related to the level of education.

Klevmarken and Quigley (1978, p. 56) to the contrary, as long as depreciation from technical obsolescence occurs at a constant rate, human capital stocks of different vintages will not depreciate at different rates. If the current stock (K_t) is the sum of depreciated past investments (Q_i),

$$(8) \quad K_t = \sum_{j=1}^{t} \delta^{t-j} Q_j = \delta K_{t-1} + Q_t \ .$$

Thus regardless of the person's age, experience, or the time pattern of his human capital investments, gross depreciation of the previous period's stock will be the same. Furthermore, since human capital is accumulated during schooling as well as on the job, there is no justification for having gross depreciation depend on experience rather than on age.

Without specifying a precise functional relationship for heterogeneous human capital, the hypothesis that an individual's depreciation rate depends on his level of schooling can be examined by including a schooling-age inter-action term in the wage equation.[3] If $\delta_i = \delta + \emptyset \, S_i$, then

$$(9) \quad \log y_t = \log R + \beta X + \alpha S - (t - t_o)\delta - \phi(t - t_o)S + u \ .$$

If the estimate of ϕ is significantly different from zero, then the hypothesis that human capital is homogeneous must be rejected.

III.   ERROR STRUCTURE

The error structure assumed for the model is the same as that used by other recent papers using pooled data, including Hanushek and Quigley (1978) and Weiss and Lillard (1978). The error term of the ith individual in the jth year has three components.

$$(10) \quad u_{ij} = D_j + v_i + w_{ij} \ .$$

One component (D_j) varies from year to year, but is identical for all observations in any year. This reflects changes in the wage rate that result from changes in the general level of productivity due to improvements in technology, increases in the capital stock, or fluctuations in business conditions. A more elaborate model might distinguish between shifts in market conditions for workers of different ages, schooling, or locations, but here it is assumed that D_j is the same for all workers. These macro factors may or may not offset decreases in an individual's relative (cross section) wage as he ages. Because there are only six time periods and because normality is not a reasonable assumption for this component, the D_j were assumed to be fixed, not random, and they were estimated with dummy variables.

The second component of the error term (v_i) is the same in each time period but varies across individuals. It reflects systematic

differences between workers in ability, motivation, under or overreporting of wage rates, or other factors which persist over time and are not captured by measured characteristics. These include cohort or vintage effects which might result from changes in the quality of school, in labor market demand, or in interruptions of work experience due to extended military service during wartime or prolonged unemployment. Since cohort effects need not be a linear function of time, the hypothesis that wages differ among otherwise similar men born in different years can be tested by seeing if the mean of this individual component of the error term varies by year of birth. Of course the mean for all men in a sample must be zero, but not for men of any one cohort.

The final component of the error term (w_{ij}) is uncorrelated over time and across individuals, and reflects purely random disturbances for each worker in each period. A substantial portion of this component is probably due to measurement error, especially when the dependent hourly wage variable is calculated by dividing annual earnings by annual hours worked.[4] Since it will never be possible to separate measurement error from actual changes in an individual's wage, interpretations of the variance of w or its ratio to the variance of v should be done with caution. Large w may simply indicate a large measurement error rather than a high degree of wage instability over time.

Two alternative assumptions were made concerning the components of the error term. The first was that the individual specific (v_i) and observation specific (w_{ij}) components were both normally distributed random variables with zero means and variances σ_v^2 and σ_w^2 respectively. In this case,

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$$E(u_{ij})^2 = \sigma_v^2 + \sigma_w^2$$

$$(11) \quad E(u_{ij}u_{ik}) = \sigma_v^2 \qquad\qquad\qquad j \neq k$$

$$E(u_{ij}u_{hk}) = 0 . \qquad\qquad\qquad i \neq h$$

Wallace and Hussain (1969) have derived estimators for the variance

of the error components when the D_j are also normal variables and there are

T observations on N individuals. When the D_j are not random and can be

estimated with dummy variables, these estimators reduce to

$$\hat{\sigma}_w^2 = \frac{1}{NT-N-T-K-1} \sum_{i=1}^{N} \sum_{j=1}^{N} (\hat{u}_{ij} - \frac{1}{T} \sum_{j=1}^{T} \hat{u}_{ij})^2$$

$$(12)$$

$$\hat{\sigma}_v^2 = \frac{1}{T^2(N-T-K-1)} \sum_{i=1}^{N} (\sum_{j=1}^{T} \hat{u}_{ij})^2 - \frac{\hat{\sigma}_w^2}{T}$$

where the $\hat{u}_{ij}$ are the residuals from a least squares regression of all NxT

observations and K is the number of independent variables not including the

time dummies.

In the present case, not all individuals have complete observations

in each period. One common solution is to drop all individuals with missing

values, but if retirement or other withdrawal from the labor market depends

on the market wage, Heckman (1974) and Gronau (1974) have shown that omitting

these men might result in biased estimates of the parameters of the wage

equation. Instead, new expressions for $\hat{\sigma}_w^2$ and $\hat{\sigma}_v^2$ were derived to take into

account the missing observations. The only difference from (10) is the

substitution of T_i, the number of observations for individual i, for T, the

number of periods in the sample.

$$\hat{\sigma}^2_w = \frac{1}{(N-1)(T^*-1)-K} \sum_{i=1}^{N} \sum_{j=1}^{T_i} (\hat{u}_{ij} - \frac{1}{T_i} \sum_{j=1}^{T_i} \hat{u}_{ij})^2$$

(13)

$$\hat{\sigma}^2_v = \frac{1}{N-T^*-K-1} \sum_{i=1}^{N} (\frac{1}{T_i} \sum_{j=1}^{T_i} \hat{u}_{ij})^2 - \frac{\sigma^2_w}{T^*}$$

where $T^* = \sum_{i=1}^{N} T_i/N$, the average number of observations per individual.

These estimates were then used to define

$$(14) \quad \lambda = 1 - \sqrt{\frac{\hat{\sigma}^2_w}{\hat{\sigma}^2_w + T^* \hat{\sigma}^2_v}}$$

Hausman (1978) has shown that regressing y^* on x^*, where $y^*_{ij} = y_{ij} - \lambda \sum_{j=1}^{T_i} y_{ij}/T_i$

and $x^*_{ij} = x_{ij} - \lambda \sum_{j=1}^{T_i} x_{ij}/T_i$, is equivalent to the generalized least squares

estimator of the regression parameters. Note that if there are no persistent

differences between individuals ($\sigma^2_v = 0$) then $\lambda = 0$ and least squares regression

on the unadjusted variables is efficient.

The second assumption made about the error term was that the individual

component (v_i) was fixed rather than random. Because of the large number of

individuals, estimating this component by dummy variables was not practical.

Instead, $\bar{y}_{ij} = y_{ij} - \sum_{j=1}^{T_i} y_{ij}/T_i$ was regressed on $\bar{x}_{ij} = x_{ij} - \sum_{j=1}^{T_i} x_{ij}/T_i$.

This is equivalent to setting $\lambda = 1$ and running y^* on x^*.

IV.   DATA AND RESULTS

The model described above was estimated using pooled data from the National Longitudinal Survey of Older Men (NLS) for six periods between 1966 and 1975. The dependent variable was the log of the hourly wage, deflated by the CPI. Reported hourly wage rates (calculated by the NLS for men who were paid by the week, month, or year by dividing the rate of pay by usual hours worked per week times the number of weeks in the reported pay period) were available for 1966, 1967, 1969, and 1971. In the last two surveys, only questions on annual wage and salary income for the previous year, annual weeks worked, and usual weekly hours were asked, so the calculated hourly wage in the last two periods is actually for 1972 and 1974.

Because this calculated wage is less reliable than the measures available for the earlier years, observations were excluded from the sample for the last two periods if the respondent received business, professional, or farm income or if his calculated wage increased by more than 100 percent or decreased by more than 50 percent.[5] Observations were also excluded for years in which the individual was over 65. The reduced sample used for estimation contained 11691 observations on 2603 whites and 4806 observations on 1091 blacks.

In addition to age and schooling, the other independent variables in the wage regression (the X's in equation 5) included dummy variables for living in labor markets with more than one million workers (Big) or less than 50,000 workers (Small), and for living in the South. Because of price variations and possible differences in labor market demand, a positive sign is anticipated for Big, and negative signs for Small and South. A dummy variable also identified men with health conditions that limited the type of work they could do (H). Because F tests revealed that wage functions differed significantly by race, all regressions were run separately for whites and blacks.

The discussion so far has suggested that wage rates may be stable or may increase between age 45 and the age at which depreciation begins (t_o); that the depreciation rate may or may not be a linear function of the individual's level of schooling; and that the individual component of the error term may be nonrandom and fixed over time or may be a normally distributed random variable. Ideally, t_o should be estimated and the hypotheses of flat age-wage hills and schooling-depreciation interactions should be tested using the two stage generalized least squares techniques appropriate for the two assumptions about the error term. However, to save on computer time, OLS was used to estimate t_o and to test the two hypotheses, and GLS was run only after t_o and the correct functional form had been determined.

The first step taken to estimate depreciation rates of earnings capacity was thus to run ordinary least squares regressions to find the values of t_o which maximized R^2, under the assumptions that the age-wage hill was flat prior to t_o and that schooling had no effect on depreciation. This specification produced the coefficients presented in columns 1 and 4 of Table 1. For whites, wage depreciation (δ) at the rate of 0.62 percent per year begins at age 51. For blacks, depreciation of 1.43 percent per year begins at age 53. As a check on this technique, t_o was also estimated by including a dummy variable for each year of age in the wage regressions to find the age of peak wages. Holding other factors constant, peak wages occurred at age 48 for whites and age 50 for blacks. Using the same specification as before, the estimates of δ for these two ages were 0.53 percent for whites and 1.10 percent for blacks.

To test the hypothesis of a flat age-wage hill prior to t_o, wage regressions were re-estimated to allow the age coefficients to differ for ages less than and greater than t_o. Using the specification of equation (7) and

TABLE 1

Wage Regressions

	Whites ($t_o = 51$)			Blacks ($t_o = 53$)		
	OLS	Random Effect	Fixed Effect	OLS	Random Effect	Fixed Effect
	(1)	(2)	(3)	(4)	(5)	(6)
age($-\delta$)	-.0062 (.0012)	-.0057 (.0016)	-.0048 (.0019)	-.0143 (.0021)	-.0125 (.0025)	-.0099 (.0030)
Education	.0586 (.001)	.060 (.002)	-	.023 (.002)	.025 (.003)	-
Health	-.137 (.010)	-.031 (.007)	-.017 (.006)	-.071 (.014)	-.017 (.011)	-.003 (.010)
South	-.065 (.010)	-.082 (.017)	-.106 (.027)	-.205 (.016)	-.228 (.028)	-.313 (.066)
Big	.062 (.011)	.067 (.018)	.027 (.026)	.053 (.018)	.027 (.032)	-.256 (.090)
Small	-.187 (.009)	-.113 (.014)	-.023 (.017)	-.359 (.014)	-.328 (.024)	-.169 (.051)
D67	.031 (.012)	.031 (.006)	.031 (.005)	.043 (.017)	.041 (.010)	.041 (.009)
D69	.081 (.012)	.090 (.007)	.088 (.006)	.126 (.018)	.119 (.011)	.117 (.010)
D71	.119 (.013)	.127 (.008)	.124 (.008)	.184 (.019)	.176 (.013)	.172 (.012)
D72	.175 (.015)	.175 (.009)	.170 (.010)	.228 (.021)	.202 (.015)	.194 (.014)
D74	.192 (.017)	.197 (.012)	.190 (.013)	.280 (.024)	.252 (.018)	.240 (.019)
Constant	.950	.192	.132	1.536	.445	.148
R^2	.272	.131	.081	.367	.193	.094

setting t_0 equal to 51 whites and 53 for blacks yielded estimates of δ_1 (the age coefficient for $t \leq t_0$) for both races that were small and insignificant and estimates of δ that were virtually identical to those using the functional form of equation (6).

To test the hypothesis that the depreciation rate is independent of the level of schooling, wage regressions were also run using the specification of equation (8). They included a schooling-age interaction term for observations with $t > 51$ for whites and $t > 53$ for blacks. The coefficients on this variable were small and insignificant for both blacks and whites, and the remaining coefficients, including those on age and education, were almost identical to the OLS coefficients in Table 1. These results do not allow rejection of the null hypotheses that there is little or no increase in wage rates between age 45 and the early 50's, and that the decline in wages after 51 or 53 does not differ by level of schooling. Therefore, the simplest specification, that of equations (5) and (6), was used for GLS estimation.

The two sets of GLS coefficients, assuming either a random or fixed individual component, are presented in Table 1. For both whites and blacks, the estimate of the depreciation rate is highest and of the standard error lowest using OLS, while the lowest coefficient and highest standard error estimates come from the fixed effects model. The random effects results lie in the middle. However, the differences are not very large. For both races, the depreciation rate from the random effects model is about 10 percent lower than the OLS estimate, and the fixed effects estimate is about 25 percent lower than the OLS estimate. The confidence intervals for the three estimators all include the other two coefficient estimates.

The OLS, random effects, and fixed effects estimates of the coefficients on the calendar year dummy variables were even closer than the depreciation

estimates, especially for whites. They all indicate that the general level
of wages for older men rose monotonically between 1966 and 1974. These in-
creases were large enough to offset the relative declines experienced due
to aging. For instance, the wage rates of black men over 53 would have de-
clined 11.4 percent during the eight years from 1966 to 1974 (8x.0143) be-
cause of depreciation, but the 28 percent rise in the general level far out-
weighed this relative decline. Similarly for whites over 51, the 19 percent
general rise far outweighed the 5 percent relative decline. Thus, at least
during the period 1966-74, the average older man did not suffer declining
real wage rates over time, even though wage rates did decline in the cross
section.

A further interesting point is that the general wage level rose faster
for blacks than for whites. The year increase was greater for blacks for
every period but 1971-72, and the total rise for 1966-74 was significantly
larger, i.e., there was no overlap of five percent confidence intervals of
the black and white coefficients for 1974. Traditionally, the black-white
wage or earnings ratio has been positively correlated with the level of aggre-
gate demand. However, this improvement in the relative position of blacks
was not due to changes in business conditions, since the unemployment rate
went from 3.8 percent in 1966 and 1967 to 5.9 percent in 1971 to 516 per-
cent in 1972 and 1974. Instead, this evidence supports findings by Freeman
(1973), Smith and Welch (1977), and Long (1977) that decreased labor market
discrimination has benefitted older black workers, and indicates that these
gains were not eroded by the economic downturn from 1970 to 1974.

Coefficients on education and South, on the other hand, suggest that
the effects of past discrimination persist. For whites, the education coef-
ficient was .060, similar to the levels found for data including men of all
ages. For blacks, the much lower education coefficient of .025 may reflect

discrimination in the provision of schooling as well as greater labor market discrimination against relatively well educated blacks. Living in the South results in lower wages for both whites and blacks, but the South-nonSouth ratio is three times larger than whites, probably because past or present discrimination has been stronger in the South.

The coefficients on the three remaining variables, dummies for poor health, and large and small labor markets, all have the expected signs, but their size and significance vary according to estimation technique. Although all three procedures indicate that poor health reduces wage rates for white workers, only the OLS health coefficient is significant for blacks. To see how much of the effect of age on wage rates is the result of declining health, the health variable was regressed on age. The age coefficients in these regressions were not significant. Furthermore, when health was omitted from the wage regressions, the age coefficients barely changed. Thus the decline in wage rates that occurs with advancing age is not the result of increasingly poor health.

The estimates of the variances of the components of the error term, presented in Table 2, are also of some interest. For both blacks and whites, over two thirds of the total unexplained variance is due to the individual term, and less than one third from the observation specific component. This result, similar to the findings of Hanushek and Quigley (1978) and Hausman (1978), indicates that most of the variation in measured wage rates that is not explained by observable characteristics comes from persistant differences among individuals. In other words, less than one third of the unexplained variation in one year's wages will even out over time. Moreover, casual inspection of the raw data suggests that a substantial portion of the observation specific error term is simple measurement error, especially when wage rates are calculated from annual earnings and hours. If this is the

TABLE 2

Variance Estimates

	Whites	Blacks
s_v^2	.148	.113
s_w^2	.041	.053
s^2	.189	.166
Number of Observations	11661	4806
Number of Individuals	2603	1096
$\dfrac{s_v^2}{s_v^2 + s_w^2}$	.783	.680

source of the error, then the percentage of true variation in cross section wage rates that is permanent will be even higher.

Finally, several recent papers have suggested that cohort effects might be important in explaining wage rates. To test this hypothesis, dummy variables for year of birth were added to the wage regressions.[6] An F test indicated that the hypothesis that year of birth has no effect on wage rates, i.e., that all the additional coefficients were zero, should be rejected for both races. However, there was no obvious pattern to the coefficients. For instance, the largest and smallest coefficients for whites were for 1907 and 1917, while the next largest and smallest were for 1920 and 1919. When the year of birth variables were added to the wage regressions, the calendar year coefficients remained unchanged, but the estimates of the depreciation rate rose from 0.62 percent to 0.86 percent for whites and fell from 1.43 percent to 1.16 percent for blacks. When the dummy variables were replaced by a linear year of birth variable, the new coefficient was small and insignificant for both races.

V.   CONCLUSION

Results of regressions on a longitudinal sample of men aged 45 to 64 indicate that wage rates begin to decline around age 50, at the rate of 0.6 percent per year for whites and 1.3 percent per year for blacks. The estimates of the depreciation rate were relatively robust for OLS, random effects, and fixed effects estimators.

Although relative wage rates declined with age, these declines were outweighed by the general increase in wage levels, so that on average, the real wages of men approaching retirement did not decline. The general wage level increased significantly faster for blacks than for whites, probably as

a result of decreases in labor market discrimination during 1966-74. On the other hand, lower returns to education for blacks and higher South-nonSouth wage differentials probably reflect past or present discrimination in schooling and in Southern labor markets.

Analysis of individual specific and observation specific components of the error term indicate that over two thirds of the unexplained variance in measured wage rates persists over time. Only one third of the measured, and probably less of the actural, variance, is due to temporary fluctuations. Cohort effects had a significant effect on wage rates. However, these effects were not a linear function of year of birth.

Footnotes

[1] A third source of depreciation may be atrophy through nonuse. While not relevant to older male workers, this source of depreciation has been examined by Mincer and Polacheck (1973) in their study of the earnings of women with discontinuous labor force experience.

[2] See for example Heckman (1976), Haley (1976), Rosen (1976), Hanushek and Quigley (1978), Johnson (1970), and Johnson and Hebein (1974).

[3] The most attractive specification is to have two kinds of earnings capacity which are multiplicative and depreciate at different rates, beginning at age t_o. Then earnings capacity (E) at age t is

$$E_t = K_o P_o \qquad\qquad t \leq t_o$$

$$E_t = K_o(1-\delta)^{t-t_o} P_o(1-\pi)^{t-t_o}$$

If, as before

$$y = R E e^{\beta x + u} \quad .$$

Unfortunately, δ and π are not separately identified under their specification, since

$$\log y = \log R + \log K_o + \log P_o + (t-t_o) [\log(1-\delta) + \log (1-\pi)] + \beta x + u .$$

More complicated multiplicative functions of K and P, such as Cobb-Douglas, also do not yield identification of δ and π .

[4] See below for further discussion of this problem.

[5] Such large changes were very uncommon for 1966-71, but all too frequent for the 1973 and 1975 samples. They were also associated with large changes in the number of hours or weeks worked. This suggests that these large changes were primarily the result of measurement error rather than changes in the actual wage received.

[6] Weiss and Lillard (1978) have analyzed a model in which cohort effects are a linear function of year of birth. When the effects of calendar time and age are also linear, the parameters of their model cannot be identified from panel data. In the present specification, the relation between the wage rate and all three of these effects is nonlinear, so all are identified.

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