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WHY DO WAGE PROFILES SLOPE UPWARDS? TESTS OF THE GENERAL HUMAN CAPITAL MODEL

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ABSTRACT

This paper tests some empirical implications of the general human capital model's explanation of rising wage profiles. At the individual level, the model implies that there will be a negative relationship between the initial wage level and wage growth of young, inexperienced workers. At the market level, the model implies that the present value of the wage profile of an investor equals that of an otherwise identical non-investor, or that the ratio of the present values equals one. We test both of these hypotheses.

Evidence on the wage level-wage growth tradeoff points to a negative relationship between initial wage levels and wage growth, even after correcting for negative biases that may have influenced existing estimates of this relationship. Evidence on present values of wage profiles suggests that the ratio of the present value of rising wage profiles to flat wage profiles is quite close to one. Alternative estimates of this ratio are tightly clustered around one, and more often than not are insignificantly different from one. Overall, then, the evidence is largely consistent with the general human capital model.

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I. Introduction

One of the most robust findings in labor economics is a positive return to time in the labor market in standard log wage equations, throughout much of the working life. But economists disagree on the source of this relationship. The dominant explanation is probably the general human capital model, in which the stock of human capital rises with experience as people invest in general human capital (not limited to a specific firm), resulting initially in lower real wages and subsequently in higher real wages (Ben-Porath, 1967; Mincer, 1974; Becker, 1975). A second explanation that has attracted considerable interest is that long-term incentive-compatible contracts must pay workers less than their marginal product when young, and more when old; thus, even if productivity is constant over a worker's life, wages will rise (Lazear, 1979 and 1981).[1]

Much existing research testing the competing models considers whether wage increases are closely positively correlated with productivity increases (e.g., Medoff and Abraham, 1981; Brown, 1989; Kotlikoff and Gokhale, 1992; Hellerstein and Neumark, forthcoming).[2] In this paper, in contrast to attempting to measure workers' productivity (or increments thereto), we test theoretical implications of the general human capital for wage profiles, tests that do not require estimates or inferences regarding productivity.

We test two implications of the human capital model of general investment. At the individual level, this model implies that there will be a negative relationship between the initial wage level and wage growth of inexperienced workers. This implication has been

[1] A third, more recent explanation, is that workers prefer rising wage profiles, as a form of forced saving (Frank and Hutchens, forthcoming; Loewenstein and Sicherman, 1991).

[2] Implications of specific human capital investment are ambiguous. Becker (1975) argues that such investment implies that wages grow slower than marginal product, but Carmichael (1983) and Blinder (1981) develop specific human capital models in which wages grow faster than marginal product. Some implications of the forced-saving model are similar to those of Lazear's model, since workers are paid less than their marginal products when young, and more when old.

considered in the literature (Hause, 1977 and 1980; Chamberlain, 1978; Lillard and Weiss, 1979; Kearl, 1988). The existing estimates, however, may be affected by a spurious correlation stemming from regression to the mean, which leads to a bias towards a negative relationship between wage levels and wage growth. In contrast, we attempt to test this implication correcting for this negative bias. At the market level, the model implies that the ratio of the present value of the wage profile of an investor to that of an otherwise identical non-investor equals one.[3] We also test this implication, which does not appear to have been considered in the literature.[4]

To summarize the results, the evidence from the wage level-wage growth test points to a negative relationship between initial wage levels and wage growth, even after correcting for negative biases that may have influenced existing estimates of this relationship. The evidence from the present value test suggests that the ratio of the present value of rising wage profiles to flat wage profiles is quite close to one. Alternative estimates of this ratio are tightly clustered around one, and more often than not are insignificantly different from one. Overall, then, the evidence is largely consistent with the general human capital model.

II. Testing the General Human Capital Model

As Weiss (1986) emphasizes, the human capital model of general on-the-job investment generates hypotheses at the individual and the market level. At the individual level, the model implies that there will be a negative correlation between the initial wage level and wage growth of young workers, controlling for other factors that affect wages. This arises because individuals face a tradeoff between lower current wages (entailing investment)

[3]Throughout, we abstract from endogenous labor supply variation, and hence refer to wage rather than earnings profiles.

[4]A partial exception is Hause (1977), who notes that equality of present values implies lower variability across individuals of lifetime earnings than of annual earnings, and finds evidence of this. Regression to the mean could partly drive this result as well.

and higher future wages. At the market level, the model has implications for the equilibrium wage structure. In particular, "at the start of working life the present value of the constant earnings stream...must equal the present value of the observed earnings profile..." (Mincer, 1974, p. 18). In this paper, we test both of these implications of the human capital model. The first test has received the most attention in the literature. In our view, though, the equilibrium implication regarding present values of wage profiles is at least as important, if not more so, because it appears to specify a much sharper restriction on the data.[5]

Testing for a Negative Relationship between Wage Levels and Wage Changes

The first test is based on estimates of the relationship between levels and changes of log wages (or residuals from the corresponding regressions). This component of our research parallels some earlier research, although much of this earlier research utilizes unusual data sources. Hause (1977) provides estimates of partial correlations among earnings at different ages using data from Sweden and the U.S. However, his data sources suffer from small sample sizes and unsatisfactory earnings measures: the U.S. data are based on recalled full-time earnings covering a fifteen-year period, while the Swedish data are based on total taxable income. Reflecting these problems, perhaps, Hause's estimated signs of the partial correlations vary widely, ranging from .75 to -.6, depending on the precise sample used and the years or ages for which the correlations are computed. Using the Swedish data, Hause (1980) embeds a similar test in estimation of the covariance structure of earnings, and finds a

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[5]For example, nothing in the Lazear or forced-saving models rules out a negative correlation between initial wage levels and wage growth within a single job. And job shopping (see Topel and Ward, 1992) may generate a negative correlation as individuals move through jobs, if those with the lowest wages seek out and move to higher-wage jobs. These alternative models do not make any predictions regarding present values of wage profiles over the life cycle, because they apply only to wage growth within a job. But heuristically, if all jobs lasted a lifetime, the models would appear to have different implications. The Lazear model implies that steeper wage profiles entail higher wages, either because costly bonding requires firms to pay a worker in excess of their marginal product (Akerlof and Katz, 1989), or because the incentive contract boosts productivity (Lazear and Moore, 1984). The forced-saving model would have the opposite implication, since it asserts that workers sacrifice present value to obtain rising wage profiles.

significant negative correlation (-.49) between earnings and earnings growth. Kearl (1988) studies individual-level data on all income (rather than earnings) from 19th century Utah; estimating a similar covariance structure, he finds strong negative correlations (ranging from -.2 to -.6). Lillard and Weiss (1979) study scientists in the U.S., and fail to find a negative relationship between earnings levels and growth rates, which they suggest may be because there are strong ability effects on both the level and growth of earnings for scientists.[6] Chamberlain (1978), using the early years of the National Longitudinal Survey of Young Men, finds strong evidence of a negative correlation (-.76) between the intercepts of wage level and wage growth equations.

What we do differently is to consider the problem of negative bias induced by a regression to the mean type of problem in these micro-level estimates.[7] To illustrate this problem, suppose we have one observation on a level (defined for period t) and a change (defined from period t to t+1) of the log wage for each individual. If there is any measurement error in wages, or more generally if there are any influences on wages unrelated to human capital investment on the job which are not perfectly correlated over time, then a spurious negative correlation is induced between the wage level and the wage change. Denote by w_t either the log wage level, or the residual from the regression of log

[6]This paper was not directed towards testing the human capital model. In fact, the authors find a negative correlation (-.14) in raw wages, when they do not condition on variables such as experience. To the extent that experience effects reflect human capital investment, the raw correlation may provide a better test.

[7]In addition, we use more complete data to estimate the wage equations. We use micro-level panel data, as did Chamberlain, but construct measures of actual experience while he used only potential experience. Also, of the papers cited, only Chamberlain used a proxy for unobserved ability, specifically an IQ score collected from high schools. Unobserved ability is potentially important since higher-ability individuals may invest more (and indeed appear to do so in Chamberlain's data), which imparts an upward bias to the correlation between wage levels and wage growth (see also Hause, 1972). We use the National Longitudinal Survey Youth cohort (NLSY) which has arguably superior proxies for ability.

wages on control variables that may shift wages independently of human capital investment.[8] The human capital model implies that w_t is lower the larger the amount of current period investment, denoted k_t, but w_t is higher the greater the amount of accumulated human capital, denoted H_t, at the beginning of the period. Assume that measured w_t is determined by these two effects, plus an error term ϵ_t,

$$(1) \qquad w_t = -k_t + \beta H_t + \epsilon_t = w_t^* + \epsilon_t \ ,$$

where β is the return on past investment, ϵ is assumed for the moment to be serially uncorrelated, with mean zero in every period, and ϵ_t is uncorrelated with w_t^*, which denotes the unobserved component of the wage that is the return on human capital minus net investment. (We also assume no depreciation.) We can use (1) to obtain an equation for the change in log wages

$$(2) \qquad \begin{aligned} \Delta w_t &= -k_{t+1} + k_t + \beta H_{t+1} - \beta H_t + \epsilon_{t+1} - \epsilon_t \ , \\ &= -k_{t+1} + k_t + \beta k_t + \epsilon_{t+1} - \epsilon_t \\ &= \Delta w_t^* + \epsilon_{t+1} - \epsilon_t \ . \end{aligned}$$

We also assume that human capital investment declines over time. In particular, we assume that for all t, $k_{t+1} = \theta k_t$, with $0 < \theta < 1$. We can then rewrite equation (2) as

$$(3) \qquad \Delta w_t = (1 - \theta + \beta)k_t + \epsilon_{t+1} - \epsilon_t \ ,$$

where $(1 - \theta + \beta) > 0$. Equations (1) and (3) reveal the essence of the human capital model's prediction; investment in period t (k_t) lowers period t wages but increases wage growth from period t to t+1.

We are interested in the sign of the coefficient from a regression of Δw_t on w_t. In

[8]Such variables might include schooling, ability, union membership, marital status, etc. We look at results for both raw wages and wage residuals. We construct the residuals from regressions of wage levels and changes on the levels and changes, respectively, of these standard control variables. However, we exclude experience terms from the regressions that create the residuals because, according to the human capital model, these reflect investment effects.

data on inexperienced workers, we can assume that the correlation between the terms involving k_t in equations (1) and (3) swamps the correlation between k_t in equation (3) and βH_t in equation (1), because for such workers the return on cumulative past investments is small. (More precisely, if workers are observed before the overtaking age, k_t exceeds βH_t.) Thus, as long as workers are observed before the overtaking age, the human capital model predicts that in the regression (omitting the constant)

$$\Delta w_t^* = \gamma w_t^* + \eta_t \ , \tag{4}$$

we should find $\gamma < 0$.

The problem, however, is that as equations (1) and (2) make clear, both the dependent and independent variables are measured with error. As in the usual case, measurement error in w_t biases the estimate of γ toward zero. But unlike the usual case, measurement error in the dependent variable also biases the estimate of γ because this measurement error is correlated with w_t. To see this, note that the observed model corresponding to equation (4) is

$$\Delta w_t = \gamma w_t - \gamma \epsilon_t + \epsilon_{t+1} - \epsilon_t + \eta_t \ . \tag{5}$$

In addition to the usual measurement error bias associated with the term $\gamma\epsilon_t$, ϵ_t enters the equation a second time, because it affects the measurement of Δw_t. Thus, the plim of the OLS estimate of γ is

$$\gamma - (\gamma + 1)\cdot \mathrm{Var}(\epsilon)/\mathrm{Var}(w) \ , \tag{6}$$

so that the estimate is biased downward (rather than toward zero), as long as the absolute value of γ is less than one.[9]

[9]This is not the precise approach taken in the existing literature. Chamberlain (1978), Hause (1980), and Kearl (1988) estimate covariance structures for earnings (or for earnings or wage residuals) that remove individual fixed effects and estimate the remaining correlation between individual-specific intercepts and slopes. In this set-up, as well, uncorrelated measurement error in wages introduces the same potential for a spurious negative correlation. If an individual's wage is low in one period, relative to his average, then his wage must be high in other periods, relative to his average, and hence his wage growth must be high.

However, we can obtain a consistent estimate of γ by instrumenting for w_t in equation (5). In particular, we instrument with the lagged value w_{t-1}. Given the assumption that ϵ is serially uncorrelated, w_{t-1} is uncorrelated with both ϵ_t and ϵ_{t+1}, and it is correlated with w_t. The variation in w_t that should be explained by the instrument is that related to the extent of human capital investment. The notion behind using w_{t-1} is that a young person investing relatively heavily in human capital will have low w_t as well as low w_{t-1}. However, if ϵ is serially correlated, then w_{t-1} is correlated with ϵ_t, and w_{t-1} is not a valid instrument. Consequently, we also report results using w_{t-2} as an instrument instead of w_{t-1}. For these estimates, any bias from non-independent errors should be much smaller.[10]

Comparing Present Values of Rising and Flat Wage Profiles

Present value calculations rely on assumptions regarding the structure of the wage profile for two reasons. First, it is not enough to estimate a wage regression including the standard controls as well as experience effects, and then to compute the present values of earnings profiles first including the experience effects and then zeroing them out. Rather, we need to be able to estimate the reduction in initial wages for investors implied by the human

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[10]To consider some specific examples, if ϵ is AR(1) with autocorrelation parameter ρ, then the plim of the IV estimate of γ using w_{t-1} is $\gamma - \rho(\gamma - \rho + 1)\cdot Var(\epsilon)/Cov(w,w_{-1})$. If $\rho > 0$, $\gamma < 0$, and $\rho + |\gamma| < 1$, the bias is negative, so a negative estimate of γ could plausibly still reflect bias. However, if w_{t-2} is used as an instrument instead of w_{t-1}, the bias should be much smaller. In this case the inconsistency in the IV estimate of γ is $\rho^2(\gamma - \rho + 1)\cdot Var(\epsilon)/Cov(w,w_{-2})$, which equals the bias when w_{t-1} is used as an instrument, multiplied by the ratio $\rho/\{Cov(w,w_{-2})/Cov(w,w_{-1})\}$. This latter ratio should be considerably less than one because ρ, which measures the persistence in the unexplained component of the wage, should be less than $\{Cov(w,w_{-2})/Cov(w,w_{-1})\}$, which measures the persistence in the overall wage. If ϵ is MA(1), then w_{t-2} is uncorrelated with ϵ_t, and is a valid instrument. Therefore, if the two IV estimates using w_{t-1} and w_{t-2} are very close, there cannot be much bias from ϵ following either an MA(1) or AR(1) process.

Because ϵ is interpreted as influences on the wage other than investment, it might be presumed to be positively serially correlated. Note, however, that variables associated with permanent differences in wage levels (which may cover most of the variables included in standard wage regressions) drop out of $\epsilon_{t+1} - \epsilon_t$, and hence may generate little bias. Rather, it is the transitory, non-human capital influences on wages that are of central interest, and these may be largely uncorrelated over time. To garner some evidence on the non-independence of ϵ, we estimated ρ from the residuals of wage regressions including individual fixed effects, and experience terms. Whereas the estimate of ρ from the residuals of the regression excluding fixed effects was .48, the estimate using the fixed-effects regression was -.06. While suggesting that ϵ is serially independent, this is not definitive because ϵ is meant to be the residual net of all human capital investment, some of which may be unobserved and unrelated to the observables.

capital model. Second, the estimated parameters of the log wage equation may provide information on the appropriate discount rate to use in the present value calculations. Thus, in contrast with the wage level-wage growth test, the present value test must be more closely integrated with structural models of human capital investment.

The most familiar form of the human capital model's implication that in equilibrium the present values of alternative wage profiles for an individual are equal is embodied in Mincer's (1974) model in which individuals invest in schooling only. This model is useful for clarifying some issues that arise in interpreting standard wage equations in the context of the human capital model. Given the assumption that individuals work for the same number of years after completing schooling, and imposing the equality of present values of the earnings streams resulting from alternative schooling choices, the wage equation

(7) $$w_t = w_0 + r_s S + \epsilon_t$$

results, where w_0 is the log wage of the individual with no schooling, S is years of schooling, and r_s is the rate of return to schooling, interpreted roughly as an average rate of return across individuals. If this framework describes the determination of wages, then (perhaps controlling for ability differences that influence w_0), the equilibrium implication can be tested by comparing the present value of the wage profile at one level of schooling with that of the wage profile at another level of schooling, using an appropriate discount rate r.[11]

Rosen (1977) has pointed out some difficulties related to using equation (7) to test the human capital model. If individuals have varying rates of return to schooling, then the Mincer model implies corner solutions, depending on whether an individual's rate of return to schooling is greater than or less than r. Rosen showed that corner solutions can be

[11] Of course, the wage profile would have to be "dated" to begin at the completion of schooling, and to continue for a fixed number of years.

eliminated by introducing individual variation in ability and in costs of financing education, which leads to a determinate equilibrium of schooling-wage combinations. However, the slope of the log wage-schooling relationship does not identify the return to schooling, since it confounds the role of education in producing human capital with covariation between wages and schooling stemming from variation in ability and access to finance.

However, Willis (1986) develops an alternative interpretation of Mincer's model in which, as a special case, equation (7) describes a determinate equilibrium, and the coefficient on S is the rate of return to schooling. Willis considers a model with heterogeneous, rather than homogeneous human capital.[12] He shows that in such a model, as long as there is equality of opportunity (*i.e.*, all individuals face the same interest rate r), and equality of "relative ability" (*i.e.*, ability has a single factor that shifts productivity equally in all pursuits), then equation (7) can be derived as a market equilibrium, with r as the return to schooling, and the present value of lifetime earnings, discounted at the rate r, equated across the alternative human capital investment decisions facing an individual.[13] Thus, under Willis's reinterpretation of equation (7), comparison of the present values of alternative wage profiles is a more compelling test of the human capital model.

While these issues have been addressed with respect to schooling decisions, they also pertain to on-the-job investments, which must be brought into the model for the purposes of this paper. The optimal path of post-schooling human capital accumulation depends on the technology for producing human capital. Mincer's approach was to assume that the solution

[12]Heterogeneous human capital implies that workers embodying different human capital investments are only imperfect substitutes in production, rather than representing different "efficiency units" of a single type of labor.

[13]The restriction regarding equal relative ability can be relaxed so that it holds for only some workers, while for other workers, ability is completely specialized to specific schooling levels (or occupations), as long as the former type of workers are the marginal workers that determine equilibrium earnings differentials (Willis, p. 570).

to the optimal investment problem is given by

(8) $$k_t = k_0 - (k_0/T) \cdot P_t \ ,$$

where k_t is the ratio of time spent investing in period t, T is the period over which investments are made, the 0 subscript denotes the initial period, and P_t is potential experience in period t. In this case, denoting by r_p the rate of return to post-schooling investments, Mincer derived the log wage equation for period t as

(9) $$w_t = r_s S + \ln\{1 - k_0 + (k_0/T) \cdot P_t\} + r_p k_0 \cdot P_t + (-r_p k_0/2T) \cdot P_t^2 \ .$$

The terms involving k_0, r_p, and T reflect the accumulated returns to past investments, and the negative impact on earnings of investment in period t (when w_t is observed).

There is, in fact, little *a priori* reason to believe that equation (8) is the optimal investment profile. We thus also consider an alternative specification of the investment profile that was proposed by Mincer. In addition, we consider some specifications of the wage equation augmented to account for differences in labor force attachment. These specifications are listed in Table 1, and are discussed in Appendix B.

According to Willis (1986), in a model with heterogeneous human capital, under the same assumptions described above regarding equality of opportunity and ability, along with the assumption that opportunities for post-schooling investment are independent of schooling, the wage profiles that Mincer derived--incorporating returns to on-the-job investment--can be derived as equilibrium wage profiles with the coefficient of schooling still measuring the discount rate.[14] Arguably, r_p should also equal this discount rate because, in equilibrium, individuals invest up to the point where the marginal gain from investment equals the

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[14]This also requires that the optimal investment profile be the same for each "type" of human capital in the heterogeneous human capital model. This can hold if it is schooling and not on-the-job investment that differentiates human capital. If it does not, though, then we can only estimate some average experience-earnings profile, and a structural interpretation of the parameters is less compelling.

marginal cost (as, for example, in the model in Ben-Porath, 1970).

Thus, in our empirical work we estimate wage equations such as equation (9), substituting r for r_s and r_p, and using the estimate of r as the discount rate in the present value calculations.[15] Of course, Willis' reinterpretation of Mincer's model makes it clear that asking whether the present values of observed, estimated wage profiles are equal to the present values of estimates of what these wage profiles would be in the absence of post-schooling investments tests only a restricted version of the human capital model. Nonetheless, it is an important version because, based on Willis' work, it must be the one that researchers have in mind when they interpret the standard log wage equation as a human capital earnings function.

III. The Data

We rely primarily on data through 1987 from the National Longitudinal Survey Youth cohort (NLSY), whose respondents were aged 14-21 at its inception in 1979. This sample offers some advantages for this research. For the test based on the wage level-wage growth correlation, we want a sample with a large number of observations before the overtaking age. Also, as discussed earlier, unobserved ability may be positively associated with both the intercept and the slope of the wage equation, hence biasing the results against finding a tradeoff between wage levels and wage growth of young workers. The NLSY has test scores from the Armed Services Vocational Aptitude Battery Test which can be incorporated into the model to control for this possible source of bias. Research with these data suggests that they are not contaminated with measurement error with respect to estimating their effects on

[15]A more subtle problem with leaving r_s and r_p unconstrained, and treating them as average rates of return but not estimates of the discount rate, is that we then have a multiplicity of possible discount rates. But the parameters of the optimal investment profile (such as equation (8)) depend in part on this discount rate. Thus, it would be inconsistent to vary the discount rate while assuming that these parameters remain the same.

wages (Blackburn and Neumark, 1993).[16]

We restrict the analysis to men to reduce problems from selection into employment, and because we focus on the implications of the human capital model for continuous workers.[17] We impose some standard sample restrictions for estimating wage equations, and focus solely on white men. In addition, we impose some non-standard restrictions to obtain a sample of individuals who are "typical" in terms of the human capital model--going to school continuously and then working more or less continuously. These restrictions lead to a final sample of 7,480 observations on 1,437 individuals. Details regarding the sample restrictions and their effects on the potential sample size are given in Appendix A. Finally, in many of the analyses we restrict the sample further, because we require data to define changes in variables, or lagged values of variables. The effects of these restrictions on the sample are also explained in Appendix A.

To explore the robustness of some of our results, we also use the January 1987 CPS Occupational Mobility and Job Tenure supplement, imposing sample restrictions similar to those used for the NLSY data. With this data set we can obtain a random sample of the same age group included in the NLSY as of 1987 (ages 22-29), as well as a random sample of workers of all ages. This data set includes information on the amount of time that respondents have been engaged in their current kind of work, a variable that provides some basis for comparison with results from the NLSY using actual experience. With both data

[16]A disadvantage of the NLSY is that the sample period covered is one of sharp changes in the wage structure, in particular in the estimated returns to schooling and experience (see, *e.g.*, Murphy and Welch, 1991). The increased return to experience over this period may make it difficult to identify changes in returns to experience for individuals as they gain experience (*i.e.*, the structure of the wage profile). Thus, it would be worthwhile to reexamine evidence on these tests using data from a period with a more stable wage structure. We did, however, verify that our results were unchanged if we dropped data from 1979-1982, when, for these data, the returns to schooling rose most rapidly (Blackburn and Neumark, 1993).

[17]For the implications of the human capital model for discontinuous workers, see Polachek (1975).

sets, of course, results using potential experience can be compared.

Table 2 reports descriptive statistics and OLS estimates of regressions for the level and change in log wages, for the NLSY sample. Most of the estimates of the level equations correspond to estimates from other samples. One exception is that the peak experience effect occurs at 9.4 years, which is low compared to estimates from samples covering a broader age range.[18] The wage changes with experience are positive for about the first 13 years of experience, and then begin to decline, roughly paralleling the results for levels.

IV. Empirical Results

Evidence on the Wage Level-Wage Growth Tradeoff

Table 3 presents estimates of the regressions of log wage changes on log wage levels, for both raw data and residuals. Looking first at the raw data, column (1) of Panel A suggests a strong negative relationship between wage levels and changes, with an estimated coefficient of -.259.[19] However, as explained above, this negative relationship may be partly spurious. This is confirmed by the estimates in column (2), in which the lagged wage is used as an instrument for the contemporaneous wage. The estimated coefficient (-.058) is still significant and negative, but is considerably closer to zero. The same result holds in column

[18]Appendix Table A2 provides estimates of comparable specifications for the 1987 observations from the NLSY, and the January 1987 CPS supplement. For the latter data set, time doing the current kind of work is used as a proxy for actual experience. For the 1987 NLSY data, the peak of the profile is at 12.9 years of experience, close to that in Table 1. For the CPS data for the same age group, the peak is 8.5 years, even lower. In contrast, for all ages from the CPS data, the peak is 25.7 years. Thus, the strong quadratic in the returns to experience, and the low implied experience level at which wages peak, appears to be largely attributable to the use of data on young workers. The table also reports comparisons for specifications using potential experience. In this case, there is evidence that the experience level at which wages peak is lower in the NLSY (17.0 *vs.* 26.1 years in the CPS for 22-29 year-olds). But the quadratic coefficient is imprecisely estimated for young workers in both samples.

[19]The regression estimate of γ was not corrected for observations that are not contiguous, which may bias this estimate towards zero. But when we reestimated γ using only the contiguous observations, the results were virtually unchanged.

(3), where we instrument instead with the wage lagged twice.[21] The similarity of the estimated coefficients in columns (2) and (3) suggests that serial correlation in the error term does not bias the instrumental variable results.

The lower panel of the table reports the same set of results using residuals from regressions of log wage levels and changes on the levels and changes, respectively, of the control variables (other than experience) used in Table 2. The results are qualitatively and quantitatively very similar.

The estimates in Table 3 make two points. Looking simply at wage levels in one period, and wage changes from that period to the next period, results in negative bias in the estimated relationship between wage levels and wage growth, suggesting that previous estimates have overstated the wage level-wage growth tradeoff. On the other hand, estimates that should reduce or eliminate this negative bias do not overturn the result that the relationship is negative, as predicted by the general human capital model.

We can delve further into the implications of the general human capital model for wage level-wage growth relationships like those reported in Table 3. According to the model, this relationship should be most strongly negative for the least experienced workers, rise to zero at the overtaking age, and subsequently be positive. In Table 4, we explore these implications by looking at the earliest and latest available observations on each individual, imposing a maximum of three years of potential experience on the early observations, and a minimum of six years of potential experience on the late observations. Given the similarity

[21] An alternative to instrumenting with w_{t-1}, for example, is simply to substitute w_{t-1} for w_t. As equation (5) shows, this removes most of the bias from regression to the mean, because only the component of the error $-\gamma\epsilon_{t-1}$ is correlated with the regressor (in contrast to $-(\gamma+1)\epsilon_t$ when w_t is used as a regressor). Thus, if γ is relatively small, there should be little bias from using w_{t-1}. In fact, the estimate of γ obtained from this method should be biased towards zero, relative to the estimate using w_{t-1} as an IV for w_t. For all of the specifications reported in the paper, this was the case. For example, for the estimates corresponding to Table 3, but using w_{t-1} or w_{t-2} as regressors, instead of instruments, the estimates of γ ranged from -.029 to -.041, with standard errors of .013 to .014.

of the results in Table 3 using w_{t-1} and w_{t-2} as instruments, in this table we use only the former (to obtain larger samples).

When the wage change is regressed on the contemporaneous wage, as reported in Panel A, the estimated coefficient is strongly negative for both the early (-.279) and late (-.274) observations. The latter negative estimate, and the near equality of these estimated coefficients, both contradict the general human capital model. However, upon instrumenting for w_t with w_{t-1}, only the estimated coefficient for the early observations is negative (-.099) and significant, while that for the late observations is positive and insignificant. The results in Panel B, using the residuals, are qualitatively similar. Thus, the regression to the mean problem inherent in the estimates in columns (1) and (3) apparently obscures evidence in favor of the general human capital model.[22]

Evidence on Present Value Comparisons of Wage Profiles

Next, we use estimates of alternative wage profiles to compute the ratio of the present value of the wage profile of an investor to that of a non-investor. The wage equation estimates for the alternative specifications of the wage equation are reported in Table 5. The first four columns report estimates of quadratic earnings functions, based on a linearly declining investment profile. Columns (1) and (2) report estimates of the original Mincer formulation, first excluding and then including the test scores. In both cases, the estimates of the investment period (T) appear rather low, at 4.4-4.5 years. The estimates of k_0 indicate that initial wages are reduced by about 30-40 percent by human capital investment.[23] Columns (5) and (6) instead report estimates of Mincer's Gompertz earnings functions, based

[22] Also, the fact that the estimated coefficient in column (2) is more strongly negative than that in column (2) of Table 3 is consistent with the fact that in Table 4, the estimate is based only on the least experienced workers, while in Table 4 it is based on all workers.

[23] Note that the positive estimates of k_0 imply that there is a tradeoff between initial wage levels and wage growth, consistent with the results in Tables 3 and 4 that examine this tradeoff in a less restrictive manner.

on an exponentially declining investment profile. These estimates suggest a longer investment period. For example, $\beta = .17$ implies that it takes 17.6 years for k_t to decline to five percent of its initial value.

Our principal concern is with the present value calculations, reported in the last three rows of the table. We report estimates for three alternative levels of final potential experience (30, 40, and 50 years). In columns (1) and (2), the ratios of the present value of the wage profile of an investor to that of a non-investor are significantly greater than one, although the estimated ratios are all quite close to one, ranging from 1.02 to 1.09. In addition, while the standard errors (constructed from first-order approximations to the nonlinear function of the parameters being tested) are quite small for these specifications, there are other sources of uncertainty that are not captured in the statistical formulas, such as mortality, productivity growth in different jobs, *etc.*, so that the precision of these estimates should probably be regarded as overstated.[24] Similar calculations are reported in columns (5) and (6) for the Gompertz earnings functions. For this specification, the estimated ratio of the present value of the wage profile of an investor to that of a non-investor is very close to one under any of the assumptions regarding years of work. The estimates range from 1.00 to 1.02, and are never significantly different from one.

For both earnings functions the estimate of k_0 rises slightly upon including the test scores, suggesting that omitted ability obscures part of the reduction in initial wages of investors, because they have on average higher ability. Given this result, we attempt to go

[24]The precision of these estimates may seem surprising. However, only the parameters of the experience profile (*i.e.*, r, k_0, and T) are required to estimate the ratio of the present values of the earnings profiles. Let $\ln(w_t) = X\beta$ be the profile of non-investors, and $\ln(w_t) = X\beta + Z_t\gamma$ be the profile of investors. Then letting r denote the interest rate, and taking sums from $t=1$ to T, this ratio is:

$$\{\Sigma \exp(X\beta+Z_t\gamma)/(1+r)^t\}/\{\Sigma \exp(X\beta)/(1+r)^t\} = \{\exp(X\beta)\cdot\Sigma \exp(Z_t\gamma)/(1+r)^t\}/\{\exp(X\beta)\cdot\Sigma 1/(1+r)^t\} = \{\Sigma \exp(Z_t\gamma)/(1+r)^t\}/\{\Sigma 1/(1+r)^t\} \ .$$

one step further than including the test scores, including instead individual fixed effects. For these estimates, second-order Taylor-series approximations to the nonlinear wage equations are used, in order to obtain models that are linear in the variables. However, because schooling is fixed for each individual, the coefficient of schooling cannot be identified. Thus, we assume that the specifications including the test scores adequately remove any omitted ability bias in the schooling coefficient, and, in columns (3) and (7), impose the estimate of the return to schooling from the previous column. As expected given more complete controlling for unobservable ability, for both the quadratic and Gompertz earnings functions, the estimates indicate higher initial investment (a larger estimate of k_0), and a longer investment period (a larger estimate of T in column (3), and a smaller estimate of ß in column (7)), although, admittedly, an estimate of k_0 of .95 for the Gompertz earnings function is implausible. For the quadratic earnings function, in column (3), the effect of removing the fixed effects on the estimated present value ratio is rather small, although the relative present value of the wage profile of investors rises slightly. For the Gompertz earnings function, the effect is somewhat more severe. In particular, for an assumed work life of 30 years, the ratio of present values is significantly below one.[25]

As a final exercise with the NLSY data, we turn to estimates of the specifications incorporating data on actual experience, using the specifications described in Table 1 and Appendix B.[26] For the quadratic earnings function, the estimates indicate an investment

[25]The differences relative to columns (2) and (6) stem principally from the fixed-effects estimation, not the use of second-order linear approximations. When we reestimated the specification in column (2) using the second-order approximation, the estimates of r, k_0, and T were similar to those in column (2), and the estimated ratios of present values were identical. When we reestimated the specification in column (6) using the second-order approximation, the estimate of k_0 rose to .56 while the estimate of ß fell to .07. The estimated ratios of present values were .96, 1.00, and 1.02, for an assumed work life of 30, 40, and 50 years, respectively.

[26]For these specifications we revert to including the test scores, and excluding the fixed effects. The estimates with fixed effects were quite imprecise, and the estimates of k_0 exceeded one for both earnings functions. The problem is likely that λ, the ratio of actual to potential experience, varies little for individuals (it varies somewhat

period of 13.02 years, which may be more plausible than the estimate of 4.43 years from the parallel specification using only potential experience. The estimates for the Gompertz specification are little changed relative to the estimates in column (6). For both earnings functions, the estimated ratios of present values of the wage profiles are close to one, and not significantly different from one, whatever the assumed years of work.

Finally, we explored the robustness of the results for the original quadratic and Gompertz earnings functions, using our 1987 CPS data set. These estimates are reported in Table 6, for 22-29 year-olds (the age range of the NLSY sample in 1987), and for workers of all ages. The estimates are quite similar in either case, although for the 22-29 year-olds, the estimate of T is quite imprecise. Most importantly, for both age ranges and both earnings functions, we again obtain estimated ratios of present values of wage profiles that are close to one, ranging from .99 to 1.04, most of which are not significantly different from one.

We interpret the evidence from the present value comparisons as providing striking consistency with the prediction of the general human capital model that, in equilibrium, the present value of the wage profile an individual faces if he invests in human capital equals the present value of the wage profile if he does not invest. For a number of specifications, samples, and estimators our estimates of these ratios are tightly clustered around one, although sometimes the estimates are significantly different from one. However, an estimate of 1.02, even with a standard error less than .01, does not strike us as providing any substantive rejection of the general human capital model.

V. Conclusions

The human capital model of general investment implies that, at the individual level, inexperienced workers give up present wages in return for wage growth. At the level of the

because it is updated each year).

market, equilibrium implies that the ratio of the present value of the wage profile of an investor to that of an otherwise identical non-investor equals one. In this paper we test both of these implications. In general, we view the evidence as consistent with the general human capital model.

The evidence on the wage level-wage growth tradeoff points to a negative relationship between initial wage levels and wage growth. However, failure to account for the problem of regression to the mean leads to severe negative bias in estimates of this relationship, and, more importantly, results in a negative estimate even for workers who are likely at or beyond the overtaking age. In contrast, after correcting for regression to the mean, the estimated relationship is positive for more experienced workers, and negative only for inexperienced workers, as predicted by the general human capital model.

The evidence from the present value estimates is that the ratio of the present value of the wage profile of an investor to that of an otherwise identical non-investor is generally near one, depending on the exact earnings function specification, sample, and estimator. While the estimates are sometimes significantly different from one, they are tightly clustered around one, and rarely differ by more than .05. Thus, in our view, the present value tests also corroborate the predictions of the human capital model.

We hasten to emphasize that, to the best of our knowledge, our tests of the human capital model are new. While other researchers have estimated wage level-wage growth correlations, none have addressed the regression to the mean problem. And we have come across no papers that implement the present value tests. Thus, we expect further refinements and replications of these and related tests to yield more definitive answers.

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Appendix A: Sample Construction

A number of restrictions are imposed on the NLSY sample in order to focus on men who follow the "typical" pattern of school, work, and investment in the original human capital models. First, we ensure that they are enrolled in school relatively continuously from age six. In particular, the sample is restricted to individuals whose age at departure from school (i.e., the last time they are observed in school in the sample) was within two years of what this age would have been had the individual begun grade 1 at age 6, remained continuously enrolled in school, and completed one grade each year; only post-schooling observations are used. Second, we focus on individuals who work relatively continuously. Specifically, measuring potential experience from the actual date of departure from schooling, we restrict the sample to individuals whose actual experience (based on cumulative weeks worked since leaving school) is more than one-half of their potential experience. We also impose some relatively standard restrictions on the sample, including omitting individuals who spent any time in the military, and observations for which respondents report a wage less than one-half of the minimum wage, report their industry as agriculture, forestry, or fisheries, or in which they report they are self-employed. We also require a minimum of two observations per person, for some of the analysis. Finally, some of the analyses require multiple observations per person, or entail other restrictions, hence leading to smaller sample sizes. The effects of the sample selection rules and the data requirements on the sample size and number of individuals available are documented in Appendix Table A1.

Appendix B: Alternative Investment Profiles and Wage Equations

An alternative investment profile that was originally considered by Mincer, but has been little used in the subsequent literature, is one in which the investment ratio k_t declines exponentially, or

(B1) $$k_t = k_0 \exp(-\beta P_t) \ .$$

This leads to the wage equation

(B2) $$w_t = r_p k_0/\beta - (r_p k_0/\beta) \cdot \exp(-\beta P_t) + \ln(1 - k_0 \exp(-\beta P_t)) \ .$$

We also consider specifications that bring data on actual experience into the estimation. We follow Mincer (1978) in assuming that full-time earnings capacity in period t (E_t) is given by

(B3) $$E_t = E_{t-1} \cdot (1 + r_p \lambda k_t) \ ,$$

where k_t is given by equation (7). As Mincer (1978) points out, this specification captures the notion that "opportunities and incentives for market-oriented (job) investments should increase with the amount of time devoted to the labor market" (Mincer 1978, p. 3), by entailing a positive relationship between dollar investment costs and hours of work. λ represents the fraction of potential experience an individual actually works, assumed constant and estimated by the ratio of actual experience to potential experience at a point in time. This investment profile yields the wage equation

(B4) $$w_t = (r_p k_0) \cdot \lambda P_t - (r_p k_0/2T) \cdot (\lambda P_t^2) + \ln\{1 - k_0 + (k_0/T) \cdot P_t\} \ .$$ [1]

Letting $k_t = k_0 \exp(-\beta \lambda P_t)$, the exponentially-declining investment ratio can also be

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[1] This alternative wage equation was proposed by Mincer (1978) in response to an alternative equation proposed by Hanushek and Quigley (1978). The exchange arose because in the linear approximations to equations (9) and (B2) used in the literature, r_s and r_p are not separately identified along with either k_0 and T, or k_0 and β, whereas bringing actual experience into the wage equation (as in equations (B4) and (B5) permits all of the parameters to be separately identified. For reasons explained in the text, we do not exploit these augmented specifications for this purpose, but rather only as a coherent means of incorporating data on actual experience.

We did, however, explore the consequences of letting r_s and r_p be unequal. As a general matter, this led to rather nonsensical estimates, with r_p in the range of .4-.5. Such results have to be viewed as casting doubt either on these specifications, on the human capital model generally, or on the specific implication that r_s should equal r_p, because in equilibrium both of these are driven to equal the discount rate.

used in the context of this extension of the model, yielding the wage equation

(B5) $$w_t = \lambda r_p k_0/\beta - (\lambda r_p k_0/\beta)\cdot \exp(-\beta P_t) + \ln\{1 - k_0 \exp(-\beta P_t)\}$$.[2]

[2]This can be derived by substituting equation (B1) into equation (B3), repeated substituting for lagged E, substituting the integral from 0 to t for the summation, and integrating.

Table 1
Alternative Wage Equation Parameterizations

	Investment Profile	Definitions	Wage Equation	Author
I.	$k_t = k_0 - (k_0/T)\cdot P_t$	In text	$w_t = r_s S + \ln\{1 - k_0 + (k_0/T)\cdot P_t\} + r_p k_0 \cdot P_t + (-r_p k_0/2T)\cdot P_t^2$	Mincer (1974)
II.	$k_t = k_0 \exp(-\beta P_t)$	In text	$w_t = r_p k_0/\beta - (r_p k_0/\beta)\cdot \exp(-\beta P_t) + \ln(1 - k_0 \exp(-\beta P_t))$	Mincer (1974)
III.	$k_t = k_0 - (k_0/T)\cdot P_t$	λ = actual/potential experience	$w_t = (r_p k_0)\cdot \lambda P_t - (r_p k_0/2T)\cdot(\lambda P_t^2) + \ln\{1 - k_0 + (k_0/T)\cdot P_t\}$	Mincer (1978)
IV.	$k_t = k_0 \exp(-\beta P_t)$	λ = actual/potential experience	$w_t = \lambda r_p k_0/\beta - (\lambda r_p k_0/\beta)\cdot \exp(-\beta P_t) + \ln\{1 - k_0 \exp(-\beta P_t)\}$	

Table 2
Descriptive Statistics and Wage Level and Wage Change Regressions, NLSY White Male Non-Military Sample, OLS Estimates, 1979-1987,
Dependent Variable for Regressions: Logarithm or Change in Logarithm of Hourly Wage[1]

	Means (Std. Dev.) (1)	Level Regression (2)		Change Regression (3)
Years of schooling	12.46	.07		--
	(1.83)	(.003)		
Experience	3.78	.13	Δ Experience	.22
	(2.17)	(.01)		(.02)
Experience^2 × 10^{-2}	18.98	-.71	Δ Experience^2 × 10^{-2}	-.80
	(20.07)	(.09)		(.10)
Married, spouse present	.35	.07	Δ Married, spouse present	-.02
	(.48)	(.01)		(.02)
Divorced, widowed, or separated	.04	.00	Δ Divorced, widowed, or separated	-.06
	(.20)	(.02)		(.03)
Union member	.21	.23	Δ Union member	.11
	(.41)	(.01)		(.01)
SMSA	.71	.12	Δ SMSA	.03
	(.46)	(.01)		(.02)
South	.30	.02	Δ South	-.00
	(.46)	(.01)		(.03)
Academic test	.29	-.04		--
	(.72)	(.01)		
Technical test	.44	.06		--
	(.68)	(.01)		
Computational test	.28	.05		--
	(.85)	(.01)		
$\bar{R}^2$	--	.32		.04

1. The constructed sample has 7,480 cross-sectional observations on 1,437 individuals. In this table, the last observation on each individual has to be dropped, to define the change, resulting in 6,043 observations on each individual. Standard errors of coefficient estimates (not corrected for dependence among observations on the same individual) are reported in parentheses. The changes in column (3) are defined from period t (the period of the level) to the next available observation; changes in continuous variables were computed as the actual change divided by the number of years elapsed between observations. The experience measure is accumulated weeks worked since leaving school, divided by 52. Years of schooling and test scores are constant in the sample, and therefore drop out of the change specifications. All specifications include dummy variables for the year from which the observation was drawn. Coefficients for these dummies, and the intercept, are not reported. The test scores were "pre-processed" by regressing them on age dummies for all observations (on white males) for whom test scores were available. The residuals from these regressions were then used. The "academic test" is the average of the residuals for the tests of arithmetics, mathematics, word knowledge, paragraph comprehension, and general science. The "technical test" is the average of the residuals for the test of auto and shop knowledge, electronics, and mechanical knowledge.

Table 3

Regressions of Changes in Log Wages on Levels of Log Wages, NLSY Sample[1]

A. Regressions Using Changes and Levels of Log Wages, Dependent Variable: Change in Log Wages from t to t+1

	(1)	(2)	(3)
Log wage at t	-.259 (.012)	-.058 (.018)	-.049 (.021)
$\tilde{R}^2$	.125	--	--
Instrument for log wage	None	Log $wage_{t-1}$	Log $wage_{t-2}$

B. Regressions Using Residuals for Changes and Levels of Log Wages, Dependent Variable: Residual Change in Log Wages from t to t+1[2]

	(1)	(2)	(3)
Log wage residual at t	-.318 (.013)	-.065 (.022)	-.062 (.027)
$\tilde{R}^2$	.153	--	--
Instrument for log wage	None	Log $wage_{t-1}$	Log $wage_{t-2}$

1. The sample is restricted to individuals with four or more observations, to define the changes and the lags. There are 5,455 observations on 1,037 individuals satisfying this restriction. The correlations are based on all but the first two observations for each of these individuals, or 3,381 observations. The regressions in Panel A also include dummy variables for the year from which the observation was drawn.

2. The regressions from which the residuals are computed are the same as in Table 2, except that the experience variables are excluded, and are estimated using all (5,455) observations on individuals with four or more observations.

Table 4
Regressions of Changes in Log Wages on Levels of Log Wages, Earliest and Latest Available Observations, NLSY Sample[1]

A. Regressions Using Changes and Levels of Log Wages, Dependent Variable: Change in Log Wages from t to t+1

	Earliest Observation on Each Person, Potential Experience ≤ 3 (N=663, Mean Potential Experience=2.19)		Latest Observation on Each Person, Potential Experience ≥ 6 (N=688, Mean Potential Experience=8.23)	
	(1)	(2)	(3)	(4)
Log wage at t	-.279 (.028)	-.099 (.046)	-.274 (.028)	.008 (.045)
$\bar{R}^2$	.123	--	.121	--
Instrument for log wage	None	Log $wage_{t-1}$	None	Log $wage_{t-1}$

B. Regressions Using Residuals for Changes and Levels of Log Wages, Dependent Variable: Residual Change in Log Wages from t to t+1[2]

	(1)	(2)	(3)	(4)
Log wage residual at t	-.353 (.030)	-.150 (.055)	-.329 (.030)	.031 (.056)
$\bar{R}^2$	.175	--	.152	--
Instrument for log wage	None	Log $wage_{t-1}$	None	Log $wage_{t-1}$

1. The sample is restricted to individuals with three or more observations, to define the changes and the lags. The regressions from which the residuals are computed are the same as in Table 2, except that the experience variables are excluded, and are estimated using all (5,831) observations on the 1,225 individuals with three or more observations.

Table 5
Non-Linear Least Squares Estimates of Structural Parameters of General Human Capital Investment Profiles, and Ratios of Present Values of Wage Profiles of Investors Relative to Non-Investors, NLSY Sample[1]

	Quadratic Earnings Functions				Gompertz Earnings Functions			
	I	I	I	III	II	II	II	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Parameter estimates:								
r	.09 (.003)	.08 (.003)	.08	.08 (.003)	.09 (.003)	.08 (.003)	.08	.08 (.003)
k_o	.29 (.02)	.30 (.02)	.65 (.27)	.43 (.03)	.43 (.01)	.45 (.01)	.95 (.29)	.48 (.01)
T	4.48 (.66)	4.43 (.69)	9.27 (3.83)	13.02 (3.18)	--	--	--	--
ß	--	--	--	--	.17 (.02)	.16 (.02)	.05 (.01)	.12 (.02)
Test scores included	No	Yes	No	Yes	No	Yes	No	Yes
Individual fixed effects included	No	No	Yes	No	No	No	Yes	No
$\bar{R}^2$	.32	.33	--	.34	.32	.33	--	.34
PDV ratios, investors/non-investors:								
Max. potential exper.=30	1.08 (.01)	1.09 (.01)	1.10 (.004)	1.02 (.03)	1.01 (.02)	1.00 (.02)	.84 (.06)	.96 (.03)
Max. potential exper.=40	1.04 (.003)	1.05 (.004)	1.07 (.02)	1.03 (.02)	1.02 (.02)	1.01 (.02)	.95 (.17)	.98 (.03)
Max. potential exper.=50	1.02 (.002)	1.02 (.002)	1.05 (.02)	1.02 (.01)	1.02 (.02)	1.02 (.04)	1.02 (.23)	.98 (.03)

1. There are 7480 observations. Asymptotic standard errors of coefficient estimates and approximate standard errors of PDV ratios are reported in parentheses. The standard errors of the PDV ratios were estimated using a first-order linear approximation of the function used in the calculation, and computing its standard error using the variance-covariance matrix of the coefficient estimates. Dummy variables for marital status, union membership, SMSA, South, and for the year from which observations were drawn were included. The PDV ratios in columns (4) and (8) are estimated using the sample mean of λ = .84. Roman numerals I-IV refer to the specifications listed in Table 1. For the fixed-effects estimates in columns (3) and (7), second-order Taylor series approximations to the earnings function were used to obtain models that were linear in the variables. In these columns the estimate of r from the preceding column was imposed, since the sample was constructed so that schooling is fixed for each individual.

Table 6
Non-Linear Least Squares Estimates of Structural Parameters of General Human Capital Investment Profiles, and Ratios of Present Values of Wage Profiles of Investors Relative to Non-Investors, White Male January 1987 CPS Sample[1]

	Quadratic Earnings Functions (I)		Gompertz Earnings Functions (II)	
	All ages (1)	22-29 (2)	All ages (3)	22-29 (4)
Parameter estimates:				
r	.08 (.002)	.10 (.01)	.08 (.002)	.10 (.01)
k_0	.27 (.01)	.40 (.04)	.40 (.02)	.43 (.03)
T	19.23 (.58)	19.35 (11.32)	--	--
β	--	--	.14 (.01)	.06 (.03)
$\bar{R}^2$	.40	.22	.41	.22
PDV ratios, investors/non-investors:				
Max. potential exper.=30	1.00 (.001)	1.02 (.03)	1.00 (.01)	.99 (.18)
Max. potential exper.=40	1.02 (.001)	1.03 (.01)	1.01 (.01)	1.02 (.23)
Max. potential exper.=50	1.02 (.001)	1.04 (.01)	1.02 (.01)	1.03 (.19)

1. There are 6,221 observations for the CPS all ages sample, and 1,405 observations for the sample aged 22-29. Asymptotic standard errors of coefficient estimates and approximate standard errors of PDV ratios are reported in parentheses. The standard errors of the PDV ratios were estimated by using a first-order linear approximation of the function used in the calculation, and computing its standard error using the variance-covariance matrix of the coefficient estimates. Dummy variables for marital status, union membership, SMSA, and South were included. Roman numerals I and II refer to the specifications listed in Table 1.

Table A1
Sample Construction and Samples Available for Each Analysis, NLSY Data Set

	Observations	Individuals
Sample construction (successive restrictions):		
Full NLSY sample, 1979-1987	114174	12686
Males	57627	6403
Whites (non-black, non-hispanic)	34110	3790
Non-military sample	28521	3169
Individuals with no military service	28134	3126
Interviews	26607	3126
Employed survey week with reported wage	17165	3016
Wage > one-half of minimum wage	16837	2998
Not self-employed	15929	2970
Not in agriculture, forestry, or fisheries	15365	2941
Non-enrolled	11634	2725
Post-schooling	10666	2533
Continuous (and non-missing) schooling	7712	1801
Final constructed sample, non-missing data, actual experience ≥ one-half of potential experience, minimum of two observations per person	7480	1437
Table 1: wage level and wage growth regressions (requires dropping last observation on each person)	6043	1437
Table 2: regressions of wage growth on wage level (wages and residuals), instrumenting with wage lagged once and wage lagged twice (requires minimum of four observations per person, and dropping first two observations on each person)	3381	1037
Wage level and wage growth regressions to estimate residuals	5455	1037
Table 3: regressions of wage growth on wage level, earliest observation with potential experience ≤ 3, instrumenting with wage lagged once	663	663
Table 3: regressions of wage growth on wage level, latest observation with potential experience ≥ 6, instrumenting with wage lagged once	688	688
Wage level and wage growth regressions to estimate residuals	5831	1225
Table 4: full sample, minimum of two observations per person	7480	1437

Appendix Table A2
Wage Level and Wage Change Regressions,
White Males, January 1987 CPS and NLSY 1987 Samples,
Dependent Variable: Logarithm or Change in Logarithm of Hourly Wage[1]

	NLSY 1987		CPS 1987, 22-29		CPS 1987, All Ages	
	(1)	(2)	(3)	(4)	(5)	(6)
Years of schooling	.09 (.01)	.09 (.01)	.06 (.005)	.10 (.01)	.07 (.002)	.08 (.002)
Experience	.11 (.03)	--	.09 (.01)	--	.04 (.002)	--
Experience^2 × 10^{-2}	-.42 (.18)	--	-.51 (.10)	--	-.07 (.005)	--
Potential experience	--	.07 (.02)	--	.07 (.01)	--	.04 (.002)
Potential experience^2 × 10^{-2}	--	-.21 (.15)	--	-.14 (.08)	--	-.06 (.03)
Married, spouse present	.10 (.03)	.12 (.03)	.12 (.02)	.08 (.02)	.29 (.01)	.19 (.02)
Divorced, widowed, or separated	-.01 (.04)	.01 (.05)	.11 (.05)	.08 (.05)	.24 (.02)	.13 (.02)
Union member	.17 (.03)	.17 (.03)	.22 (.03)	.21 (.03)	.15 (.01)	.17 (.01)
SMSA	.13 (.03)	.13 (.03)	.14 (.02)	.14 (.02)	.15 (.01)	.15 (.01)
South	-.01 (.03)	-.01 (.03)	-.03 (.02)	-.04 (.02)	-.04 (.01)	-.04 (.01)
$\bar{R}^2$	.22	.19	.22	.22	.41	.40
Peak of experience profile (years)[2]	12.9	17.0	8.5	26.1	25.7	30.6

1. There are 1,405 observations for the CPS sample of 22-29 year olds, 6,221 observations for the CPS all ages sample, and 1,213 observations for the NLSY sample. Standard errors of coefficient estimates are reported in parentheses. The intercept is not reported. The sample restrictions and variable definitions applied to the NLSY sample are the same as those in Table 2. In the CPS sample, the Occupational Mobility and Job Tenure Supplement has a question on the length of time doing the current kind of work; this is used as an estimate of actual experience for this sample.
2. For all years of the NLSY data (7,480 observations) the peaks for columns (1) and (2) are 10.8 and 11.6 years.