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TRAINING, TENURE, AND PRODUCTIVITY

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ABSTRACT

There is substantial evidence from the literature on individual wage determination that length of service to the firm is an important determinant of earnings and thus of labor productivity, holding constant employee attributes such as age, sex, and education. Earnings growth associated with increased tenure is usually interpreted as a reflection of firm-specific on-the-job training (OJT). In this paper a model of producer technology consistent with the hypothesis of firm-specific OJT is formulated and estimated. Empirical implementation of the model on data for U.S. manufacturing provides the basis for estimation of the marginal productivity of workers classified by length of service to the firm, i.e., of the tenure-productivity profile. The parameter estimates also enable us to determine the effect of recent changes in the tenure distribution (due to changes in labor turnover behavior) on manufacturing productivity performance.

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Recent studies in the literature on individual wage determination reveal that tenure is an important determinant of earnings, holding constant the employee's total amount of work experience, education, and personal characteristics such as sex and race.[1] The significance of tenure in earnings equations is generally interpreted as providing support for the hypothesis that a substantial fraction of the skills acquired by workers via on-the-job training (OJT) are firm-specific. It is true that the high partial correlation between earnings and tenure may be partly spurious--an artifact of unobserved worker heterogeneity.[2] But the strong association between earnings and tenure is not destoyed by attempts to control for heterogeneity. For example, wage regressions estimated by Mincer and Jovanovic[3] on two different samples--the National Longitudinal Survey sample of young men, and the Michigan Income Dynamics sample of men of all ages--in which an attempt is made to control for heterogeneity, suggest that <u>more than half</u> of skills acquired on the job are firm-specific. According to their parameter estimates, a typical worker would have increased his current hourly earnings more by moving to his current job a year earlier (holding total time in the labor force constant) than he would have by entering the labor force a year earlier (holding time in the current job constant).

The role of specific OJT as a determinant of worker productivity is recognized to be of even greater importance when one considers the fact that the tenure coefficient in wage equations captures merely the worker's private return to investment in firm-specific skills, not the social return. According to the theory of OJT, the costs and returns of specific investment are shared (in theoretically indeterminate proportions) by worker and firm. This implies that the rate of growth of wages paid to the worker as he accumulates skills will be lower than the rate of growth of his marginal productivity (MP), which reflects the combined return to employer and worker. By the nature of the case, the wage equation framework is incapable of providing an estimate of the social return to specific training; what it can, in principle, provide--assuming that the heterogeneity problem can be solved, arguably a heroic assumption--is a lower-bound estimate of the growth of MP attributable to training.

Although there is ample evidence that firm-specific training contributes significantly to the productivity of labor resources, the specific training hypothesis has not been integrated or reflected in empirical analyses of production behavior. Failure to account for specific training in the analysis of production activity is striking because OJT is, by definition, integral with the process of production. (Of course, it is the very integration of the two activities that is the source of many of the measurement problems associated with training; the process of human capital formation is "submerged" within production.) Failure to acknowledge the presence of training in this context is unfortunate in several respects. First, information about the level and structure of training costs and returns, which cannot be obtained by any other approach, may be lost. Second, the representation of technology will, in general, be misspecified, possibly resulting in distorted estimates of parameters characterizing the structure of production. The distortion arises from an attempt to isolate for purposes of analysis two activities (training and production) which economic agents have found appropriate or efficient to combine. Finally, the direct analysis of production behavior offers an alternative to wage equation estimation for the investigation of training issues.

The objective of this research is to formulate and estimate a model of producer technology consistent with the hypothesis of firm-specific OJT.

The model is implemented empirically using data for the U.S. manufacturing sector. Firm-specific training is often thought to play a particularly important role in the market for manufacturing production workers[4], who comprise roughly seventy percent of manufacturing employment. Estimation of the model enables us to perform tests of the training hypothesis and to obtain estimates of parameters which identify the tenure-MP profile. Moreover, the implications of specific training for productivity trends in manufacturing may be assessed.

This paper is organized as follows. In Section I, the general form of a model of technology consistent with the specific training hypothesis is developed. This model is compared and contrasted to two other general classes of production models which have been specified by previous investigators. In Section II, a specific functional form and estimation procedure are selected, estimation results are presented and hypothesis tests performed. Problems of statistical inference associated with the empirical analysis of firm-specific training are discussed in Section III.

I

The essence of the OJT hypothesis is that firms utilize current employed resources to augment the future MP of employees. Consequently, the appropriate representation of the technology of an enterprise which engages in OJT is a multi-product model of production. Such enterprises produce intangible investment in human capital as well as ordinary output, using a stock of human capital (the sum of depreciated past investments) and other resources. The production possibility frontier (PPF) characterizing the technology may be written in general, implicit form as

$$F(Q, IHC, K, SHC, A) = 0 \qquad (1)$$

where Q = quantity of ordinary output
IHC = quantity of investment in human capital

K = quantity of physical capital input
SHC = quantity (stock) of human capital
A = index of technology

Assume for simplicity that all training is firm-specific. Evolution of the stock of human capital is determined by past investments in OJT according to the "perpetual inventory" equation

$$SHC(t) = \int_0^t IHC(x) \exp[-k(t-x)]\, dx \qquad (2)$$

where k is the (presumed constant) depreciation rate of human capital. This depreciation rate reflects the separations behavior of trained employees as well as the effects of skill atrophy.

This model of production is formally identical to that formulated by Christensen, Jorgenson, and Lau[5] to describe the PPF of the entire economy. They represent the aggregate production frontier by

$$F(C,I,K,L,A) = 0 \qquad (3)$$

where
C = quantity of consumption goods output
I = quantity of investment goods output
K = quantity of capital input
L = quantity of labor input
A = index of technology

K and I are related by an accumulation equation similar to (2). Just as physical capital is a produced means of production at the level of the total economy, so are specific skills produced means of production at the enterprise level. In both cases, past decisions to allocate resources to the production of one output in favor of the other condition the current availability of inputs. Also, less-than-full utilization of resources weakens the tradeoff between (human or physical) capital formation and current production.

Conventional specifications of technology abstract from firms' production and utilization of specific human capital. The PPF is written as

$$F(Q,K,L,A) = 0 \qquad (4)$$

where L = total manhours employed

It is instructive to compare the index of total factor productivity (TFP) consistent with the "misspecified" technology (4) with the index consistent with the "true" technology (1). This comparison enables us to determine the direction and magnitude of biases in TFP measurements based on the maintained hypothesis that (4) represents the structure of production.

The rate of growth of TFP is defined as follows:

$$(\dot{P}/P) = \sum_i w_i \ (\dot{Y}_i/Y_i) - \sum_j v_j \ (\dot{X}_j/X_j) \tag{5}$$

where
P = total factor productivity
Y_i = quantity of i-th output $i = 1,\ldots,N$
w_i = share of i-th ouput in value of total output
X_j = quantity of j-th input $j = 1,\ldots,M$
v_j = share of j-th input in value of total input

and dotted variables indicate differentiation with respect to time. As Jorgenson and Griliches[6] have shown, if the production function is characterized by constant returns to scale, and if the necessary conditions for producer equilibrium--all marginal rates of transformation between pairs of inputs and outputs are equal to the corresponding (shadow) price ratios--are satisfied, this definition of TFP measures the "shift" in the production frontier. The definition of TFP applied to (1) yields

$$(\dot{P}/P) = w_Q(\dot{Q}/Q) + w_{IHC}(\dot{IHC}/IHC) - v_K(\dot{K}/K) - v_{SHC}(\dot{SHC}/SHC) \tag{6}$$

The definition of TFP corresponding to the null hypothesis of no specific training (i.e., equation (4)) is

$$(\dot{P}'/P') = (\dot{Q}/Q) - v_K(\dot{K}/K) - v_L(\dot{L}/L) \tag{7}$$

The bias in conventional productivity accounting is equal to the difference between (7) and (6):

$$(\dot{P}'/P') - (\dot{P}/P) = w_{IHC}[(\dot{Q}/Q) - (\dot{IHC}/IHC)] - v_L[(\dot{L}/L) - (\dot{SHC}/SHC)] \tag{8}$$

We have used the condition that $v_L = v_{SHC}$, i.e. that the share of labor in total input cost equals the share of human capital. It is convenient to rewrite (8) as

$$(\dot{P}'/P') - (\dot{P}/P) = v_L([d \ln(SHC/L)]/dT) - w_{IHC}([d \ln(IHC/Q)]/dT) \quad (9)$$

Equation (9) reveals that the conventional measure of the rate of growth of TFP overstates the true measure when the growth rate of the stock of human capital exceeds the growth rate of manhours, i.e., when the stock of human capital per manhour, or "labor quality", is increasing. The overstatement is higher the larger is labor's share in total input cost. The conventional measure understates the true measure when the growth rate of human capital investment exceeds the growth rate of output. The understatement is greater the higher is the share of human capital investment in the total (shadow) value of the firm's production-cum-training activity. The analysis which follows suggests that $[(\dot{Q}/Q) - (\dot{IHC}/IHC)]$ and $[(\dot{L}/L) - (\dot{SHC}/SHC)]$ will generally have opposite signs at any given moment, implying that the two sources of bias will be reinforcing rather than offsetting. The hypothesis that they have opposite signs is consistent with the notion of an "equilibrium" skill distribution, which the firm attempts to maintain by accelerating training investment when labor quality declines and vice versa.

Although the multi-product formulation (1) is useful in enabling us to assess potential biases attending conventional productivity measures, it cannot provide a basis for empirical research, since investment in firm-specific training, and hence the stock of human capital, are not directly observable. Development of a model of training-cum-production capable of empirical implementation requires us to make an assumption about the <u>determinants</u> of training activity. A hypothesis about the determinants of the firm's "demand" for training may enable us to identify the path of training investment and the stock of human capital. It is therefore postulated that the typical firm has a "standard training program," which consists of a sequence of investments in an employee undertaken at specified points in his career within the enterprise. In other words, the firm is <u>committed</u>

to provide certain training investments as the worker accumulates years of service. The notion of a standard training program abstracts from variation in investment activity which might arise due to such factors as differences in learning ability among new entrants and the stage of the business cycle. The implications of individual variation in the intensity of investment will be considered in Section III.

Under the standard training program hypothesis, both the current level of an employee's MP and his rate of investment in OJT are indexed by his length of service (LOS) to the enterprise, or tenure. Similarly, the distribution of employees by length of service determines both the marginal productivity of total labor input and the rate of aggregate training investment; i.e., it determines both the stock and the flow of specific human capital. If the training program is characterized by a monotone decreasing (with tenure) rate of investment in the typical employee--a policy consistent with standard models of optimal human capital accumulation--and there is no depreciation of skills, individual MP-tenure profiles will be monotone increasing and concave. As the distribution of employees shifts in the direction of lower tenure, the stock of specific human capital per worker falls, and the flow of training investment per worker rises.

The production possibility frontier of an enterprise which has a standard training program may be approximated by

$$F(Q,K,L_1,L_2,\ldots,L_N,A) = 0 \qquad (9)$$

where L_i = employment in the i-th length-of-service category $i = 1,2,\ldots,N$

In principle, this formulation captures all relevant information concerning the firm's past and current training investments. Estimates of MP by tenure group may be interpreted as ordinates of the tenure-MP profile, and thus enable us to determine the position and slope of the profile. The height of the profile at any given length of service indicates the contribution to the stock of human capital by a worker with that length of service, and the slope of the profile indicates the rate of investment in the worker.

We therefore propose to estimate a model of producer technology in which labor is classified by length of service to the establishment. Before turning to the discussion of estimation procedure and results, it is desirable to contrast this model with other models of disaggregated labor in production. Previous investigators have estimated models of production--either cost or production functions, or marginal productivity (first-order) conditions derived from them--in which labor is classified by one or more attributes postulated to determine marginal productivity. The most common disaggregation criteria are age, occupation, and educational attainment of workers. Although these models bear an obvious formal resemblance to a model of the general form (9), these models are based on assumptions about the operation of the labor market which are inappropriate to the analysis of firm-specific training. These differences with respect to assumptions about labor market structure are reflected in different research objectives.

Previous analyses of labor disaggregated by characteristics such as age, education, and occupation conventionally assume that the market for each type of labor is competitive, i.e., the firm can hire all it wants of each type of labor at an exogenously determined, constant wage rate. Also, the firm is postulated to be in competitive equilibrium with respect to each type of labor at every moment, equating the MP of each type to its wage. The primary objective of these studies is to obtain estimates of demand elasticities for and substitution elasticities among the different categories of labor. Knowledge of these elasticities would enable the analyst to assess the effects on relative employment of policy- or otherwise-induced changes in the relative prices of different types of labor and, in some cases, capital.

The assumptions of the competitive labor market paradigm are appropriate to the analysis of the firm's demand for labor only if the skills required for production are completely transferable across firms, i.e., only if all training is general. An alternative paradigm, that of the internal labor market, is needed to account for the behavior of employment and wages classified by skill when skill is a product of firm-specific training and experience. The constraints on the wages and employment of the various skill groups are fundamentally different when skills are generated by the firm's own training activity than they are when skills may be obtained on the external market. In the latter case, the firm is presumably free of employment (quantity) constraints, being able to hire arbitrary amounts of each type of labor in each period, but subject to N price constraints, required to pay each type of labor its respective market-determined wage. Where necessary skills are developed by firm-specific training--and assuming that the quantity of training embodied in a worker is determined by tenure--the firm operates under (N - 1) employment constraints: the maximum currently available supply of workers with t years' experience and training is limited by the number of entrants t years ago. Only the level of employment of new workers is unconstrained in each period. It is true that the firm can increase its supply of skilled workers in the short run by accelerating the training and promotion of workers lower down in the skill hierarchy, but it is often assumed[7] that training is subject to increasing instantaneous marginal costs, which means that it is more expensive to upgrade workers rapidly than slowly. This suggests that the need for workers with different amounts of firm-specific training imparts an intertemporal interdependence to hiring and employment decisions, in contrast to the myopic (one-period) character of decisions to employ externally available labor.

There is also a sharp distinction between the postulated behavior of wage rates classified by skill when skills are firm-specific and relative wage behavior when skills are perfectly general. Because specific training drives a wedge between a worker's MP and his opportunity wage, producers are not constrained by the external market to pay a worker a wage equal to his MP at any particular point in his career. Competition among employers for prospective trainees would function to constrain the expected present value of wages over the worker's career to equal the expected present value of his MP. Rather than operating under a regime of N wage constraints--one for each job classification in the skill hierarchy--the employer operates under two broad constraints, one on the level and one on the slope of the wage profile. The height of the profile must be high enough to attract an adequate quantity and quality of new entrants; and the rate of growth of wages must be sufficient to encourage the worker to accept training and promotions. Changes in external labor market conditions may force employers to adjust the general level of wages, but not the relative wage structure. It is often argued in the personnel management/industrial relations literature that employers seek to maintain stable relative wages within the internal labor market in order to promote a sense of fairness among employees. There is evidence that relative (occupational) wage rates among manufacturing production jobs are extremely stable.

The estimation of static, technical elasticities of substitution and demand--the principal objective of previous analyses of disaggregated labor in production--is not the goal of the present research, nor is the economic significance of such parameters clear in a labor market characterized by intertemporal constraints on the supply of various skills and relative wages which do not reflect relative marginal productivity and are not exogenous to employers.

II

In this section we formulate and estimate an econometric model of producer technology--a cost function--in which labor is classified by length of service. The estimation procedure to be implemented requires that we adopt the maintained hypothesis of weak separability between labor and non-labor inputs. Separability constrains the marginal rate of substitution between any two types of labor to be independent of the level of nonlabor inputs, and the elasticity of substitution between any nonlabor input and labor to be the same for all types of labor. It is generally not desirable to impose separability restrictions a priori. When the labor separability hypothesis has been subjected to test in previous studies of disaggregated labor in production, it has usually been rejected. The "stylized fact" is that more highly skilled labor and capital are complementary inputs, and both are substitutes for unskilled labor in production.[8] Evidently, the potential bias of parameter estimates arising from inappropriate imposition of the separability restriction is attenuated if the technology is specified to have the translog form. According to Chinloy[9], the translog function has the property of "approximate consistency in aggregation. This implies that little error arises from a two-stage construction of value-added by forming subaggregates of labor and nonlabor inputs, and subsequently aggregating the two. This vitiates in large part any error from incorrect aggregation if separability does not obtain."

Under the maintained hypothesis of separability, there exists a consistent aggregate index of labor input. It is convenient to begin specification of the production model with the labor input index. It is assumed that this index, denoted L^*, is a translog index of its arguments:

$$\ln L^* = \sum_{i=1}^{N} B_i \ln L_i + (1/2)\sum_{i}^{N}\sum_{j}^{N} w_{ij} \ln L_i \ln L_j \qquad (10)$$

The translog may be viewed as a local second-order approximation to any arbitrary functional form.

Symmetry and homogeneity constraints are imposed on (10). Symmetry implies that $w_{ij} = w_{ji}$, for $i,j = 1,\ldots,N$. Parameter restrictions implied by linear homogeneity, which ensures that an x percent change in the labor services of each group yields an x percent change in total labor input, are

$$\sum_i^N B_i = 1; \sum_j^N w_{ij} = 0$$

Imposing symmetry and homogeneity constraints by direct substitution into (10) yields

$$\ln L^* = \sum_{i=1}^{N-1} B_i \ln(L_i/L_N) + \ln L_N +$$

$$(1/2) \sum_{i<j}^N \sum^N w_{ij}[2 \ln L_i \ln L_j - (\ln L_i)^2 - (\ln L_j)^2] \qquad (11)$$

Following Chinloy[10], we define an index of labor quality, Z, consistent with the quantity index of labor input. Labor quality is defined as labor input per manhour:

$$Z = L^* / L$$

where $L = \sum_i^N L_i$

Consequently

$$\ln Z = \ln L^* - \ln L$$

Substituting the expression (11) for $\ln L^*$, we obtain

$$\ln Z = \sum_i^{N-1} B_i \ln(L_i/L_N) + \ln(L_N/L) +$$

$$(1/2)\sum_{i<j}^N \sum^N w_{ij}[2 \ln L_i \ln L_j - (\ln L_i)^2 - (\ln L_j)^2] \qquad (12)$$

The overall structure of production is represented by a translog cost function, which relates the minimum cost, C, of producing a given level of output, Q, to a vector of input prices, P, and an index of technology, T:

$$\ln C = a_0 + a_Q \ln Q + \sum_h^M a_h \ln P_h +$$

$$(1/2)\sum_h^M \sum_k^M g_{hk} \ln P_h \ln P_k + a_T T + \sum_h^M g_{hT} T \ln P_h \qquad (13)$$

Allowing for nonzero g_{hT} permits us to test for nonneutral technical change. Differentiating the cost function with respect to $\ln P_h$, and imposing the symmetry condition $g_{hk} = g_{kh}$ $(h,k = 1,\ldots,M)$ we obtain a system of M equations:

$$(\partial \ln C)/(\partial \ln P_h) = a_h + \sum_{k}^{M} g_{hk} \ln P_k + a_{hT} \qquad h = 1,\ldots,M \qquad (14)$$

By Shephard's lemma (which applies to arbitrary cost functions),

$$(\partial C)/(\partial P_h) = X_h$$

where X_h denotes quantity of input h. Hence

$$(\partial \ln C)/(\partial \ln P_h) = (\partial C/\partial P_h)(P_h/C) = X_h\ (P_h/C) = s_h \qquad (15)$$

where s_h is the share of input h in total cost. Under the maintained hypothesis of cnstant returns to scale, the following restrictions on the cost function parameters obtain:

$$\sum_{h}^{M} a_h = 1; \sum_{h}^{M} g_{hk} = 0$$

These restrictions enable us to eliminate (for example) the M^{th} equation and g_{hM} from the remaining (M - 1) equations in (14). The resulting system of equations may then be written

$$s_h = a_h + \sum_{h}^{M-1} g_{hk} \ln(P_h/P_M) + g_{hT}\ T \qquad h = 1,\ldots,M-1 \qquad (16)$$

Estimation of the parameters of the cost function by (16) requires data on the price of each factor of production. To maintain a consistent accounting framework, the (implicit) price of each factor is defined as the ratio of total expenditure on the factor, C_h, to factor quantity:

$$P_h = C_h/X_h \qquad h = 1,\ldots,M \qquad (17)$$

Thus

$$\ln P_h = \ln C_h - \ln X_h \qquad (18)$$

The price of labor corresponding to the quantity index of labor input is

$$\ln P_L^* = \ln C_L - \ln L^* \qquad (19)$$

This is related to the conventional definition of the price of labor, labor cost per hour worked (P_L), by

$$\ln P_L^* = \ln P_L - \ln Z \qquad (20)$$

Clearly, our measure of labor quality determines how given total expenditure on labor input is divided into price and quantity components. By evaluating the time derivative of (20), it is evident that the change in the price of a unit of labor equals the change in the price per manhour minus the change in labor quality (input per manhour). If labor quality increases (decreases) over time, cost per unit of input has been falling (rising) relative to cost per manhour.

The appropriate measure of the price of labor to be included in the cost function and derived factor share equations is P_L^*, since this indexes the cost to the firm of employing a unit of labor of constant quality. Moreover, P_L^* is more likely than P_L to be exogenous to producer behavior, since the latter is affected by changes in employment mix as well as changes in group-specific wage rates.

For purposes of empirical analysis, the cost function was specified to be of the value-added or net output variety: output is defined as value added, and only primary factor (capital and labor) prices are included in the cost function. It is well known that a value-added specification is based on the maintained hypothesis of strong separability between primary and intermediate (energy and materials) inputs. The separability hypothesis has been subjected to statistical tests and decisively rejected in recent econometric work. Unfortunately, the data required to generalize the model to include intermediate inputs were not available for this investigation.

In a two-factor setting, the system of input share equations (16) reduces to a single equation:

$$s_L = a_L + g_{LL} \ln(P_L^*/ P_K) + g_{LT} T \qquad (21)$$

Substituting for $\ln P_L^* = \ln P_L - \ln Z$ and using the expression (12) for $\ln Z$, we obtain the estimating equation

$$s_L = a_L + g_{LL}[[\ln(P_L/ P_K) - \ln(L_N/L)] - \sum_i^{N-1} B_i \ln(L_i/L_N) -$$

$$(1/2) \sum_{i<j}^{N}\sum^{N} w_{ij}(2 \ln L_i \ln L_j - (\ln L_i)^2 - (\ln L_j)^2)] \qquad (22)$$

Ordinary least squares estimation of (22) yields simultaneous estimates of the parameters of the labor input (or labor quality) index and cost function parameters. The equation is exactly identified: the coefficient on the difference $[\ln(P_L/ P_K) - \ln(L_N/ L)]$ equals g_{LL}, the coefficient on $\ln(L_i/ L)$

equals $g_{LL}B_i$, and so forth.

Equation (22) was estimated using quarterly observations on the durable and nondurable sectors of U.S. manufacturing for the period 196101 to 197604. Because no regular time-series data on employment classified by job tenure are available, it was necessary to construct estimates of these series using "new hires" data from the Bureau of Labor Statistics' labor turnover survey, in conjunction with job tenure data collected periodically in the Current Population Survey. These sources and the assumptions underlying the construction of the tenure-group series are documented in the Appendix. In a nutshell, a perpetual inventory algorithm was developed, and the estimate of the number of employees with n periods' tenure in period t is proportional to the number of new hires in period t-n. The available data sources dictated, to a certain extent, the specific partitioning of employees into tenure groups. After experimenting with several alternative partitionings, disaggregation into three mutually exclusive and exhaustive categories--employees with 0-6 months',7-24 months', and 25 or more months' tenure in the current establishment--was adopted. Further disaggregation tended to generate unstable estimates of the parameters of the labor input index, evidently as a consequence of multicollinearity among the constructed tenure-group series. Together, the three groups account for roughly a third of employment, on the average.

Three variants of equation (22) were estimated for each of the two sectors. All equations were estimated using the Cochrane-Orcutt adjustment for serial correlation of residuals. The estimation results are presented in Tables 1-D (durables) and 1-N (nondurables). Model III represents the unrestricted form of the labor share equation, corresponding to the translog index of labor input (linear homogeneity and symmetry imposed). Model II corresponds to a Cobb-Douglas labor input index, in which all second-order terms (w_{12}, w_{13}, w_{23}) are constrained to equal zero. Model I corresponds to

TABLE 1-D

ESTIMATES OF VARIANTS OF LABOR SHARE EQUATION

DURABLE MANUFACTURING

Parameter*	Model I	Model II	Model III
a_L	.8217	.8250	.8251
	(121.5)	(303.5)	(298.3)
G_{LT}	-.5969E-03	-.5820E-03	-.5515E-03
	(3.7)	(8.3)	(7.7)
G_{LL}	.1104	.1094	.1082
	(29.3)	(37.1)	(33.7)
B_i		.0369	-.0174
		(2.0)	(0.4)
B_2		.1218	.0514
		(4.7)	(0.9)
W_{12}			-.1033
			(1.5)
W_{13}			.1986
			(2.0)
W_{23}			.2395
			(1.3)
rho	.8274	.5769	.5408
	(11.7)	(5.6)	(5.1)
D-W	1.97	2.16	2.22
R^2	.9905	.9932	.9936
SSR	.7422E-03	.5302E-03	.4952E-03

* Absolute t-values in parentheses
All equations estimated using Cochrane-Orcutt adjustment for first-order serial correlation.

TABLE 1-N
ESTIMATES OF VARIANTS OF LABOR SHARE EQUATION
NONDURABLE MANUFACTURING

Parameter*	Model I	Model II	Model III
a_L	.7988	.8023	.8138
	(99.6)	(160.4)	(164.4)
G_{LT}	-.1424E-02	-.1597E-02	-.1614E-02
	(7.8)	(14.2)	(15.1)
G_{LL}	.1394	.1508	.1523
	(16.5)	(26.6)	(26.5)
B_1		.9321E-02	.0264
		(0.6)	(0.9)
B_2		.1089	.1475
		(3.8)	(3.5)
W_{12}			.0494
			(0.7)
W_{13}			-.0854
			(0.9)
W_{23}			-.2429
			(1.3)
rho	.8320 (11.9)	.8134 (11.1)	.7954 (10.4)
D-W	1.72	1.74	1.70
R^2	.9631	.9850	.9855
SSR	.1063E-02	.4327E-03	.4179E-03

* Absolute t-values in parentheses.
All equations estimated using Cochrane-Orcutt adjustment for first-order serial correlation.

the null hypothesis of no specific OJT, since the price of labor is defined as labor cost per manhour; this hypothesis implies the following restriction on the parameters of the Cobb-Douglas index:

$$B_1 \ln(L_1/L_3) + B_2 \ln(L_2/L_3) + \ln(L_3/L) = 0$$

To test the hypotheses that (1) the labor index is Cobb-Douglas, and (2) labor input is equivalent to total manhours, we performed F-tests on the three specifications. This consists of calculating the change in the sum of squared residuals from imposing restrictions; dividing this change by the sum of squared residuals of the unrestricted equation; and dividing numerator and denominator of this ratio by the appropriate number of degrees of freedom (number of parameter restrictions and residual degrees of freedom, respectively). The resulting test statistic is distributed, asymptotically, as $F(v_1,v_2)$, where v_1 is the numerator degrees of freedom and v_2 is the denominator degrees of freedom. Performing the test for Models II and III indicates that we cannot reject the hypothesis of a Cobb-Douglas index, in favor of the more general translog specification, for either sector. $F(3,55) = 1.29$ for durables and 0.28 for nondurables; the critical value of $F(3,55)$ at the .95 level of significance is 2.78. Imposing the Cobb-Douglas restrictions results in a negligible decrease in the fraction of variance explained by the independent variables.

The hypothesis that labor input is equivalent to total manhours is decisively rejected for both sectors. The test statistic $F(2,58)$ equals 7.87 for durables and 28.6 for nondurables, compared to a critical value of 5.03 at the .99 significance level. Besides providing superior overall goodness of fit, Model II yields estimates of individual parameters considerably lower in variance than does Model I. Except in the case of the capital-labor substitution parameter ($g_{LL} = -g_{LK}$) for nondurables, the differences between point estimates of the parameters common to Models

I and II were not statistically significant. This suggests that failure to include information about the tenure distribution in the specification of technology does not result in seriously distorted estimates of parameters characterizing the technology.

Before examining the labor index parameters and implied relative marginal productivity estimates in detail, we consider the behavior of the cost function implicit in our estimates of equation (22). In order to qualify as a well-behaved neoclassical cost function, a cost function must have the properties of monotonicity and convexity. The translog function does not satisfy these conditions globally; we therefore check to establish that they are satisfied for the sample space on which the model is estimated. A sufficient condition for monotonicity to obtain is that the fitted cost shares for all factors be positive in each period. This condition is satisfied for all variants of the model for both sectors (fitted values are not reported here). Convexity of the cost function is guaranteed if the own-price elasticity of demand for each factor is negative. Own-demand elasticities and Allen partial elasticities of substitution are related to estimated model parameters as follows:

$$ED_i = (g_{ii}/s_i) + s_i - 1 \qquad i = L,K$$

$$ES_{ij} = (g_{ij}/(s_i \cdot s_j)) + 1 \qquad i,j = L,K;\ i \neq j$$

$$ES_{ii} = (g_{ii} + s_i^{\,2} - s_i)/s_i^{\,2} \qquad i = L,K$$

where s_i = fitted value of share of factor i in total cost

ED_i = own-price elasticity of demand for factor i

ES_{ij} = Allen elasticity of substitution between factors i and j

Estimates of ES_{ij} and ED_i vary, with s_i and s_j, over the sample period; convexity was satisfied in each period. Estimates of the substitution and demand elasticities implied by Models I and II, for the first, middle, and last quarter of the period, are presented in Table 2. Although the estimates are well-behaved, the substitution and demand elasticities are quite small,

TABLE 2
ESTIMATES OF SUBSTITUTION AND PRICE ELASTICITIES IMPLIED BY
LABOR SHARE EQUATION PARAMETER ESTIMATES

	ES_{KL}	ES_{KK}	ES_{LL}	ED_L	ED_K
Durables					
Model I					
196101	.0476	-.3083	-.0074	-.0064	-.0413
196804	.3096	-1.239	-.0773	-.0619	-.2477
197604	.1548	-.8474	-.0282	-.0239	-.1309
Model II					
196101	.0982	-.5974	-.0162	-.0139	-.0844
196804	.3030	-1.251	-.0734	-.0591	-.2439
197604	.1772	-.9452	-.0332	-.0280	-.1492
Nondurables					
Model I					
196101	.1736	-.6343	-.0475	-.0373	-.1363
196804	.2613	-.7736	-.0883	-.0660	-.1953
197604	.2921	-.7910	-.1079	-.0788	-.2133
Model II					
196101	.1264	-.4436	-.0360	-.0280	-.0984
196804	.1840	-.5681	-.0596	-.0450	-.1390
197604	.2451	-.6434	-.0934	-.0676	-.1775

ES_{ij} denotes Allen elasticity of substitution between i and j, $i,j = K,L$

ED_j denotes own-price price elasticity of demand.

relative to those obtained by other reserchers. For example, the "consensus" point estimate of the elasticity of demand for labor is of the order -0.3. I have not been able to determine why the estimated elasticities are substantially smaller than one would expect.

We turn now to the focal point of the empirical work, the analysis of estimates of the parameters of the index of labor quality. The analysis has two objectives. The first is to obtain estimates of the relative marginal productivity of the three tenure groups of employees, and to perform significance tests on the estimated productivity differentials. The second is to examine the behavior of the labor quality index over time, and to consider the implications of the latter for trends in aggregate labor productivity.

Since labor input is postulated to be weakly separable from nonlabor inputs, comparisons of the marginal productivity of the three tenure groups may be made entirely in terms of the labor index parameters. The elasticity of output with respect to the quantity of the i-th group's labor services is

$$(\partial \ln Q)/(\partial \ln L_i) = ((\partial \ln Q)/(\partial \ln L^*))((\partial \ln L^*)/(\partial \ln L_i)) =$$
$$((\partial \ln Q)/(\partial \ln L^*))(B_i + \sum_j w_{ij} \ln L_j)$$

The output elasticity measures the percentage change in output attributable to a one percent change in L_i. The marginal product of L_i, i.e., the absolute change in output resulting from a unit (one manhour) change in L_i, is obtained by multiplying the elasticity by the average product, Q/L_i:

$$MP_i = ((\partial \ln Q)/(\partial \ln L^*))(Q/L_i)(B_i + \sum_j w_{ij} \ln L_j)$$

Hence the ratio of the marginal productivity of group i to that of group k is

$$RMP_{ik} = MP_i/MP_k = [(B_i + \sum_j w_{ij} \ln L_j)/(B_k + \sum_j w_{kj} \ln L_j)](L_k/L_i)$$

This reduces to $(B_i/B_k)(L_k/L_i)$ in the Cobb-Douglas case. RMP_{13} and RMP_{23}--the ratios of the marginal productivities of each of the two "junior" groups to that of the "senior" employees--were computed for each quarter for the Cobb-Douglas variant of the labor index (Model II). For the nondurables sector, the inequalities $0 < RMP_{13} < RMP_{23} < 1$ were satisfied in every quarter. These inequalities also held in all quarters for durable goods industries, with the exception of three quarters in which RMP_{23} slightly exceeded unity. In order to obtain summary measures of the relative marginal productivity of the three tenure groups, sample means of RMP_{13} and RMP_{23} were calculated. These are presented in the top panel of Table 3. These figures may be interpreted to signify that, for example, workers with 0-6 months' tenure in durable goods industries are 24.0 percent as productive, on the average, as workers with over two years' experience in their current establishment. Although the magnitudes of the marginal productivity ratios constitute presumptive evidence of the unequal contributions (per manhour) of the various tenure groups, formal tests of the statistical significance of the productivity differentials were conducted. The (absolute) difference in marginal productivity between two tenure groups is proportional to

$$DMP_{ij} = (B_j/L_j) - (B_i/L_i)$$

Since this is a linear combination of normally distributed variables with known variances and covariance, calculation of the t-ratio of DMP_{ij} is straightforward. T-ratios were calculated for the two pairs of "adjacent" tenure groups, i.e., groups 1 and 2, and groups 2 and 3. For the durables sector, DMP_{12} was significantly greater than zero at the 97.5 percent level ($t(64) = 2.00$) in 41 quarters (out of 64); DMP_{23} was significantly greater than zero in 33 quarters. For nondurables, DMP_{12} was always sigfificantly greater than zero, and DMP_{23} was in 43 quarters. Sample average t-values are displayed in the bottom panel of Table 3; all pass the significance test easily.

Because workers with different amounts of tenure exhibit differences in marginal productivity, variation in the tenure distribution of employees induces variation in the average quality of utilized labor services. The logarithm of the index of average labor quality corresponding to the Cobb-

TABLE 3

SUMMARY RELATIVE MARGINAL PRODUCTIVITY STATISTICS

	Durables	Nondurables
marginal productivity ratios:		
RMP_{13}	.240	.054
RMP_{23}	.654	.537
t-statistics on marginal productivity differences:		
DMP_{12}	2.17	4.39
DMP_{23}	2.43	2.74

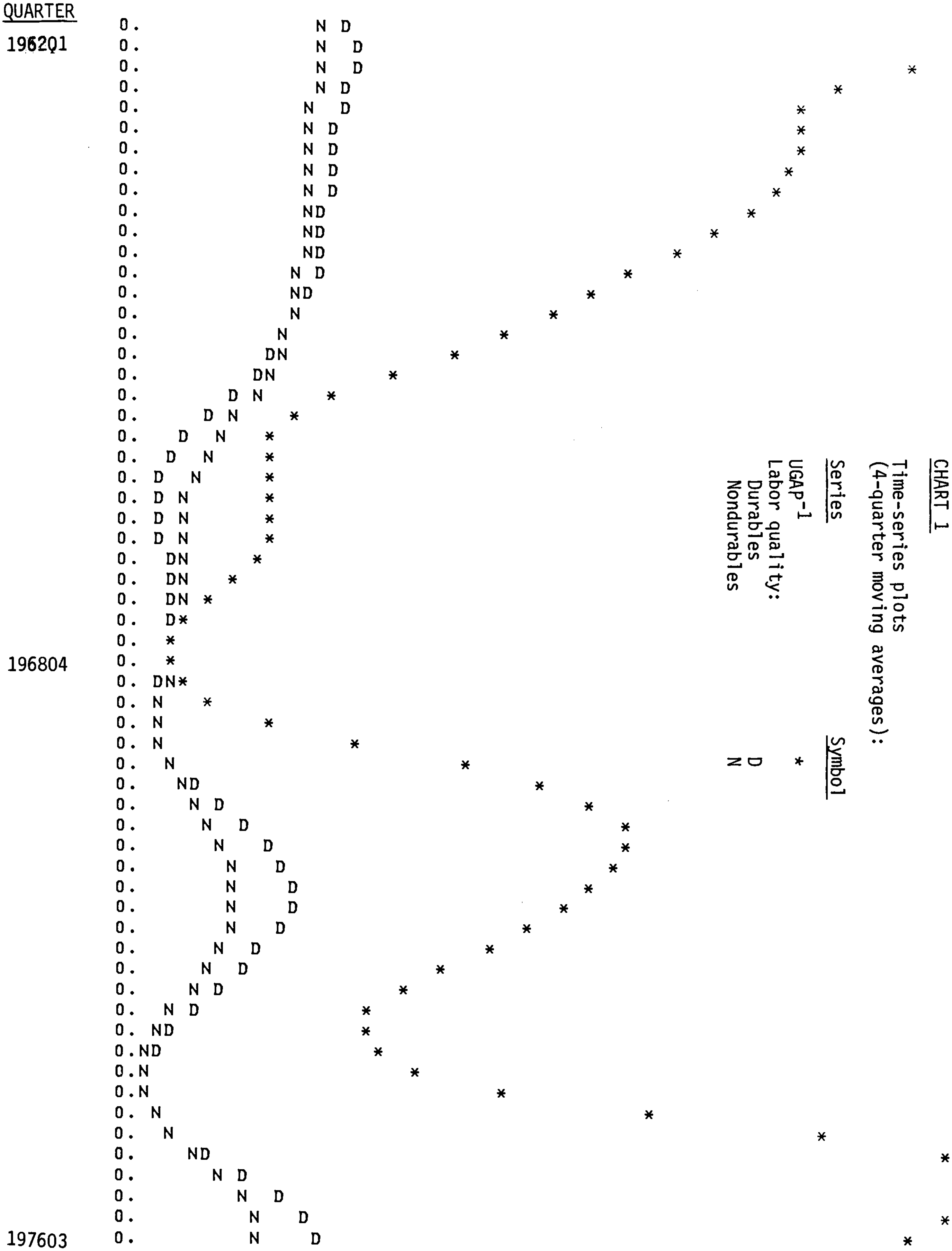
note: See text for definitions of variables. Figures reported represent sample means.

Douglas input index is

$$\ln Z = (B_1 \ln(L_1/L_3)) + (B_2 \ln(L_2/L_3)) + \ln(L_3/L)$$

The labor quality index was evaluated, using point estimates of B_1 and B_2, for the sample interval 196101 to 197604. A time-series plot of a four-quarter moving average of the index (normalized to unity in 196804) is shown in Chart 1. An index of aggregate labor market tightness, Wachter's UGAP measure, is also indicated. The chart shows that the durables and nondurables quality indices exhibit marked and similar cyclical variation, although the durables series is characterized by somewhat wider fluctuations. As expected, labor quality moves countercyclically, rising as firms curtail hiring in response to weakening sales, declining as firms dilute their experienced workforces with new entrants during recoveries. Short-term changes in labor input per manhour may be quite pronounced: labor quality changed by as much as 12.9 percent in durables, and 11.4 percent in durables, within six quarters.

The secular behavior of labor quality has direct implications for long-term trends in total factor productivity (see eq. (9),p. 6). Labor quality tended to decline in both sectors over the sample period: the average quarterly decline was -.048 percent in durables and -.107 percent in nondurables. These figures correspond to cumulative reductions in quality of 3.3 percent and 8.3 percent, respectively, over 16 years. To determine whether the secular decline in quality was statistically significant, the logarithm of labor quality was regressed on a time trend and a constant. The t-value of the trend coefficient was 1.77 for durables and 5.46 for nondurables, indicating significance at the .95 level (t(64) = 1.67). Because both beginning and ending quarters of the sample interval were periods at or near cyclical troughs in macroeconomic activity, the average rates of change do not appear to be seriously distorted by sample period definition; if anything, the greater severity of the mid-Seventies recession might produce an


QUARTER
196201
196804
197603
CHART 1
Time-series plots
(4-quarter moving averages):
Series
Symbol
$UGAP^{-1}$
*
Labor quality:
Durables
D
Nondurables
N

underestimate of the quality decline.

III

The estimates of the parameters of the index of labor input enabled us to measure the MP of each of the various tenure groups. In this section we discuss some problems of attempting to make inferences about the shape of a "representative individual's" MP-profile from these estimates of MP by tenure group. In essence, all of these problems of inference are caused by the possibility of unobserved worker heterogeneity, combined with a process of selection (by firms, and/or self-selection by workers) which determines which workers will accumulate tenure. It is likely that our inability to control for various unobserved characteristics results in an overestimate of the slope of the "representative individual's " MP-profile by the derived relative marginal productivity estimates. We shall consider the effect of two "types" of heterogeneity: heterogeneity with respect to human capital investments undertaken prior to employment in the current firm, and heterogeneity with respect to the intensity of investment within the current firm.

Due to data limitations, labor input is classified by a single characteristic -- length of service to the establishment -- in the model of production underlying the empirical investigation. The human capital theory of marginal productivity determination implies that tenure is not the only attribute which determines the user value of labor services. In particular, education and general training acquired in previous employment are usually postulated to augment the MP of workers. In principle, if there are n characteristics which determine the productivity of labor resources, a complete n-way classification is required for the consistent measurement of labor input. If the classification scheme is of less than order n, and some of the attributes by which labor is <u>not</u> classified are correlated with those by which it is classified, estimated differences in MP between cells in the classification will reflect differences in attributes not "controlled for". There is an omitted classification criterion, analagous to an omitted variable in regression analysis.

Although it is not possible to cross-classify labor input by tenure, previous experience, and education in our data set, extraneous information concerning the correlation of tenure with education and age (from which previous experience may be inferred) may enable us to evaluate the influence of these omitted criteria on the estimated productivity differentials. Such information is available in the form of two-way classifications of employees by (1) tenure and education, and (2) tenure and age, derived from the Current Population Survey. Estimates of average educational attainment and average age of employees in four lenght-of-service categories were calculated from these crosstabs and are presented in Table 4. The first column of the table indicates that there is no clear association between tenure and educational attainment. This finding appears to warrant rejection of the hypothesis that part of the estimated returns (in the form of higher productivity) to increased tenure should be attributed to differences in education across tenure groups.

In contrast to education, age exhibits a strong positive association with tenure. Moreover, the differences in average age between tenure groups always exceed the differences in tenure (measured as differences between midpoints of the class intervals). The implication is that workers with more tenure in their current firm generally also have more previous work experience. Rough estimates of differences in average age, tenure, and previous work experience between "adjacent" tenure groups are as follows:

group	age difference	tenure difference	previous exper. difference
7-12 m./0-6 m.	2.1 yrs.	0.5 yrs.	1.6 yrs.
13-24 m./7-12 m.	1.7	0.8	0.9
over 25 m./13-24 m.	10.6	5.3	5.3

It is likely, then, that workers with greater length of service had higher endowments of general human capital, and higher MP, at the time they entered the firm than workers with low tenure. Viewed from a different perspective, workers with greater previous experience are more likely to "survive" to advanced service than those with less prior experience. These differences

TABLE 4

AVERAGE YEARS OF SCHOOL COMPLETED AND AVERAGE AGE
BY LENGTH OF SERVICE TO CURRENT EMPLOYER

Length of service	Average years of school completed[1]	Average age[2]
0 - 6 months	12.5 yrs.	29.6 yrs.
7 - 12 months	12.4	31.7
13 - 24 months	12.9	33.4
Over 25 months	12.3	44.0

Notes: 1. Education data refer to males, 25 years old and over, employed in January, 1978.

2. Age data refer to all workers (male and female), 16 years old and over employed in January, 1973.

Source: Crosstabulations from Current Population Surveys, January 1973 and 1978. All workers assumed to be at the midpoint of their respective education or age class interval.

in MP at time of entry would, in themselves, tend to cause the estimated productivity differentials to overstate the returns to specific training. However, this overstatement would be attenuated if older, more experienced workers tend to invest a smaller fraction of their time in specific OJT -- behavior consistent with optimal life-cycle human capital accumulation. The justification for this statement will be provided by the following discussion, in which the implications of heterogeneity with respect to the intensity of specific investment are analyzed.

The specification of our model of training-cum-production was based on the hypothesis of a standard training program, i.e. a sequence of training investments which is the same for all entrants. Under this hypothesis, the estimates of MP by tenure group reflect the slope of the (uniform across workers) tenure-MP profile. Suppose we relax the assumption of uniform investment profiles, and assume that workers are free to choose among a continuum of investment intensities, i.e. they may choose the fraction of time at work to be devoted to investment. Now if the probability that a worker would attain a given length of service were independent of his intensity of investment, we could still interpret our estimates of MP by tenure group as indicating the "average" of the slopes of the various MP-tenure profiles. But the assumption of independence has little appeal. On the contrary: ceteris paribus, individuals with greater expected completed length of service (CLS) -- for whatever reason -- should devote more time to firm-specific investment. In a cross-section of individuals, the people with higher tenure, or uncompleted length of service, will tend to have higher CLS. Thus, workers with higher tenure will have invested more intensively at every point in their careers than individuals with lower tenure. In the presence of heterogeneity with respect to investment intensity, labor input should be classified by intensity of specific investment as well as by its duration (tenure) in the specification of labor input. Since intensity of investment is an omitted classification criterion positively correlated with tenure, the estimated returns to increased tenure capture the returns to higher intensity as well. Similarly, omission of the relevant but unobservable variable "intensity" in wage equations leads to upwardly biased tenure coefficients.

Although the intensity on an individuals' investment is unobservable, we have postulated that this intensity is determined by his expected CLS, which, in turn, is inversely related to his probability of separation. Thus, if one could identify an attribute or set of attributes of workers which indicated or determined expected CLS (or the propensity to separate), this could be used to "control" for heterogeneity in investment behavior.

Two recent studies have adopted the strategy of including variables postulated to reflect separations propensity in individual wage regressions in order to purge the tenure coefficients of the effects of heterogeneous investment. Both studies utilize longitudinal data and implicitly assume the determinant of investment intensity to be stable over time, so that past observations on the behavior of the individual can be used to control for heterogeneity. Mincer and Jovanovic[11] include a "previous mobility" variable, in addition to education and linear and quadratic terms in both total work experience and tenure, in log-wage regressions for NLS young men, NLS mature men, and MID men of all ages. The effect on the tenure coefficients of introducing the previous mobility variable into the equation varied form sample to sample. In the young men's sample, it had no effect: "Prior mobility is not related to current wages and does not affect the tenure coefficients... Apparently differences in early mobility of young men are not indicative of future differences in specific capital investments nor do thay capture differences in wage levels which are positively related to the length of current tenure."[12] Introduction of the prior mobility variable into the regression for mature men "cuts the linear (tenure) term in half and reduces its signigicance," suggesting that "repeated mobility at an advanced stage of the life-cycle is an indicator of persistant turnover, denoting little investment in human capital."[13] As one might expect, in the wage equation for MID men of all ages "the inclusion of prior mobility variables reduces the tenure slope by close to 20 percent." This implies that "heterogeneity biases the tenure - wage slope coefficient upward by about 25 percent."[14]

Bartel and Borjas[15] also estimate earnings functions on the NLS mature men's data set, but they attempt to control for heterogeneity with respect to specific investment behavior in a different way. Instead of introducing an additional right-hand-side variable to control for heterogeneity, they specify the dependent variable to be (Y_t-Y_o), where Y_t is current earnings and Y_o is imputed earnings in the first year of the life cycle. (The initial-earnings imputation is made using data on initial occupation.) By analyzing wage growth rather than wage levels they claim to "net out individual differences that are unobserved but affect the individual's earnings throughout the life cycle."[16] Although substituting lifetime wage growth for wage level in an earnings equation probably does control for some aspects of unobserved heterogeneity, it is not at all obvious that it controls for characteristics which determine specific investment behavior; controlling for previous mobility appears to be a superior procedure for accomplishing this. Bartel and Borjas find that tenure is a highly significant determinant of earnings growth of mature men; that is, it is a highly significant determinant of their current earnings, holding initial earnings constant. Unfortunately, they don't report results for a comparable wage level model, so it is not possible to assess the effect on the tenure coefficient of controlling for initial earnings.

The attempts by Mincer/Jovanovic and Bartel/Borjas to control for unobserved variation in the intensity of specific investment are implicitly based on the hypothesis that the differences across individuals in characteristics which determine investment behavior are permanent differences. To the extent that investment heterogeneity is an artifact of permanent (or relatively stable) differences in individual characteristics (e.g., tastes), inclusion of additional information in wage equations may help to attenuate bias in tenure coefficients. But recent theoretical work by Jovanovic[17] suggests that an individuals' desired intensity of investment may also be determined by sheer chance, i.e. by the quality of his "match" with his current employer. This result emerges from a model of permanent job separations in which the intensity of on-the-job specific training and on-the-job search are endogenous. Jovanovic assumes that there exists a nondegenerate distribution of a worker's productivity across different employers, the nondegeneracy of this distribution being due to the assumption that the quality of the worker-firm match differs across prospective matches. In his model, the quality of the match determines the worker's productivity upon entering the firm. Once employed, the worker devotes a fraction of his time, I, to specific OJT and a fraction of his time, S, to on-the-job search. These fractions are endogenous: the model determines the evolution of I(T) and S(T), where T indexes tenure on the current job. The following equilibrium conditions are yielded by solution of the model: (1) "Separation probabilities regarded as a function of job tenure are uniformly lower for those who are well matched, for two reasons: workers that are well matched spend less time searching for alternative work, and when they do receive alternative offers, they are less likely to accept them."[18] (2) Well matched individuals spend higher fractions of time investing (at any tenure), because (by (1)) they have a lower probability of future separation.

This theory has disturbing implications for our ability to control for heterogeneity of investment behavior in wage equations or other applications. For the quality of the worker-firm match is not merely unobservable; it is also likely to be uncorrelated with <u>any</u> observable individual attribute.

APPENDIX

Construction of Length-of-service Distributions of Employees from Labor Turnover Data

Empirical implementation of the model developed in this paper requires time-series estimates of the length-of-service distribution of employment. Regular time-series data on the tenure distribution are not available. In the absence of such data, it is necessary to construct these series using the perpetual inventory technique, by which means we attempt to make inferences about the current tenure distribution on the basis of the past history of the gross flow of workers into employment.

The data base for the perpetual inventory algorithm includes monthly data on the number of "new hires" in an industry, derived from the Bureau of Labor Statistics' Labor Turnover survey. New hires are, for the most part, "temporary or permanent additions to the employment roll of persons who have never before been employed by the establishment." These employment accessions are distinguished in the turnover survey from "rehires," i.e., additions to the roll of former employees recalled by the establishment.

The tenure distribution of employees in an industry in period t is related to the past hiring history of the industry by the following set of identities:

$$EMP(n,t) = NH(t-n) * s(n,t) \qquad n = 1,\ldots,w$$

where $EMP(n,t)$ = number of employees with n months' tenure in period t

$NH(t-n)$ = number of new hires in period t-n

$s(n,t)$ = probability that a worker hired in t-n will remain employed ("survive") until t

w = maximum observed length of service

It is evident from this set of equations that given the (known) path of new hires, the problem of estimating tenure-distributions reduces to the problem of estimating the "survival probabilities" s(n,t). Estimation of survival probabilities is feasible because, although regular series on the tenure distribution do not exist, direct observations of the distribution are available for five dates during the period 1963-1978. Under certain assumptions concerning the survival probabilities and the observed distributions, these data, in conjunction with the new hires data, enable us to construct high-frequency series on employment classified by length of service.

The available distributions are based on responses of employed persons to the question, "When did [respondent] start working at his present job or business?," a supplementary question included in the Current Population Surveys (CPS) of January 1963, 1966, 1968, 1973, and 1978. For wage and salary workers (which include the vast majority of manufacturing workers), "a job is defined as a continuous period of employment [as defined below] with a single employer, even though a person may have worked at several occupations while working for that employer." A period of employment is considered "continuous" if there have been "no interruptions except for vacations, temporary illness, strikes, short-term layoffs (less than 30 days), and similar temporary factors." A representative tenure distribution (that of males employed in durable and nondurable manufacturing at the time of the 1978 survey) is shown in Table A-1. As indicated, the class intervals of the length-of-service distribution are 0-6 months, 7-12 months, 13-24 months, and so forth. The percentages of all employees (male and female) in each of the first three class intervals, at each survey date, are presented for the two sectors in Table A-2. Also shown are the average tenure distributions, in which each cell is computed as the simple average of the corresponding cells from the five survey tabulations.

Two important assumptions were made to permit construction of a time-

TABLE A-1

PERCENT DISTRIBUTION OF MALE EMPLOYEES
BY PERIOD WHEN CURRENT JOB STARTED:
JANUARY, 1978
DURABLE AND NONDURABLE MANUFACTURING

Period when current job started:	Durables	Nondurables
July 77 - Jan 78	14.2	14.3
Jan - June 77	8.2	7.5
Jan - Dec 76	9.6	9.8
Jan - Dec 75	5.6	6.8
Jan 73 - Dec 74	11.8	12.1
Jan 68 - Dec 72	17.0	18.6
Jan 63 - Dec 67	11.8	9.7
Jan 58 - Dec 62	6.9	6.5
Jan 53 - Dec 57	5.8	5.4
Jan 48 - Dec 52	5.2	4.9
Jan 43 - Dec 47	2.3	2.8
Prior to Jan 43	1.7	1.6

Source: Unpublished BLS crosstabulation based on January, 1978 Current Population Survey

TABLE A-2

PERCENT OF EMPLOYEES IN THREE LENGTH-OF-SERVICE CATEGORIES, SELECTED DATES 1963-1978, AND ESTIMATED AVERAGE PROBABILITIES OF REMAINING EMPLOYED

	0-6 MONTHS	7-12 MONTHS	13-24 MONTHS
DURABLES			
Percent of employees in length-of-service category:			
1963	11.9	7.0	7.8
1966	14.5	7.7	8.6
1968	14.8	7.5	11.3
1973	14.8	6.6	9.4
1978	15.2	8.4	10.5
average 1963-1978	14.2	7.4	9.5
Estimated average probability of remaining employed for specified length of time:			
	.811	.445	.292
NONDURABLES			
Percent of employees in length-of-service category:			
1963	11.7	7.2	8.3
1966	14.4	8.1	8.2
1968	16.0	7.9	10.0
1973	15.4	6.3	10.2
1978	15.4	8.4	10.9
average 1963-1978	14.6	7.6	9.5
Estimated average probability of remaining employed for specified length of time:			
	.679	.374	.233

series of tenure distributions. First, it was assumed that the survival probabilities were independent of calendar time, i.e. that $s(n,t) = s(n)$ for all n. This assumption is analagous to the common assumption of constant depreciation rates in the construction of capital stock series. The second assumption was that the average of the published CPS distributions represents the average tenure distribution for the entire period spanned by the job tenure surveys (January 1963 - January 1978). Because the surveys were conducted at almost regular intervals throughout the period, and general economic conditions varied considerably from one survey date to the next, the average of these distributions should be a reasonable approximation to the average length-of-service distribution for the entire period.

Given these two assumptions, the estimation of survival probabilities proceeds as follows. The number of employees with n months' tenure in period t is

$$EMP(n,t) = s(n) * NH(t-n)$$

and their share in total employment is

$$SHR(n,t) = (s(n) * NH(t-n)) / EMP(t)$$

where $$EMP(t) = \sum_{n} EMP(n,t)$$

The expected value of $SHR(n,t)$ over the 1963-1978 period is assumed to be equal to the corresponding average share in the CPS tabulations, $CPS(n)$:

$$E[SHR(n,t)] = E[(s(n) * NH(t-n)) / EMP(t)] = CPS(n)$$

where E[ ] represents the mathematical expectation operator. Factoring out $s(n)$, which is assumed constant, and rearranging terms,

$$s(n) = CPS(n) / E[NH(t-n) / EMP(t)]$$

This was the equation used to calculate the survival probabilities. In a sense, this procedure consists in benchmarking the new hires data to the infrequently observed tenure distributions. It is appealing because it eliminates systematic biases in the turnover data which are known to exist, and it abstracts from differences in employment levels reported in the household and establishment (turnover) surveys. Survival probabilities were calculated separately for the durables and nondurables sectors, for the first three tenure categories distinguished in the published tabulations: 0-6 months, 7-12 months, and 13-24 months. These estimates are shown in Table A-2.

The estimated survival rate for a particular group is, of course, an average of the survival rates of its members. For example, the survival rate of the 0-6 month group is an average of the rates of workers with 0 months', 1 month's, 2 months', etc., tenure. In order to obtain more precise estimates of the tenure distribution, the survival rate <u>within</u> the 0-6 month group was specified to decline at a constant monthly rate (corresponding to a constant monthly conditional probability of separation), such that the average of the monthly rates equalled the estimate for the entire group . The "monthly" survival rates were obtained by numerically solving the following expression for q:

$$(1/7) * \sum_{m=0}^{6} q^m = s(0\text{-}6)$$

where s(0-6) is the average survival rate of workers hired 0-6 months ago, and q^m is the survival rate of employees hired "exactly" m months ago, m = 0,1,...,6. As for the 7-12 and 13-24 month groups, the average survival rate for each group was simply imputed to all persons in the group, i.e., no allowance was made for within-group variation in survival rates.

Time-series estimates of the number of employees in a given length-of-service category were obtained by multiplying lagged new hires by the appropriate survival rate; this was performed for the three groups specified. The number of employees with greater than 24 months' tenure was computed as a residual, by subtracting the sum of the estimates of employment in the 0-6, 7-12, 13-24 month categories from the total employment figure.

<u>FOOTNOTES</u>

1. See, for example, Mincer, J., and B. Jovanovic, "Labor mobility and wages," NBER Conference Paper No. 35, June, 1980, and Duncan, G., and S. Hoffman, "On-the-job training and earnings differences by race and sex," Rev. Econ. Stat., November, 1979.

2. The heterogeneity problem is discussed in detail in Section III of this paper.

3. Op. cit.

4. See, for example, Doeringer, P., and M. Piore, Internal labor markets and Manpower Analysis, Lexington, Mass.: Heath, 1971.

5. Christensen, L., D. Jorgenson, and L. Lau, "Transcendental logarithmic production frontiers," Rev. Econ. Stat., February, 1973.

6. Jorgenson, D., and Z. Griliches, "The explanation of productivity change," Rev. Econ. Stud., July, 1967.

7. See, for example, Rosen, S., "Learning and experience in the labor market," Jnl. Hum. Res., VII:3, and Ehrenberg, R., Fringe benefits and overtime behavior, Lexington, Mass.: Heath

8. See Hamermesh, D., "Econometric studies of labor-labor substitution and their implications for policy," Jnl. Hum. Res., XIV:4.

9. Chinloy, P., "Sources of quality change in labor input," Amer. Econ. Rev., March, 1980, p. 118.

10. Op. cit.

11. Op. cit.

12. Op. cit., p. 29.

13. Ibid.

14. Op. cit., p. 32.

15. Bartel, A., and G. Borjas, "Wage growth and job turnover: an empirical analysis," NBER Conference Paper No. 36, June, 1980.

16. Op. cit., p. 23

17. Jovanovic, B., "Firm-specific capital and turnover," Jnl. Polit. Econ., Vol. 87, No. 6, 1979.

18. Op. cit., p. 1255.