Intertemporal Labor Supply and Long-Term Employment Contracts

By JOHN M. ABOWD AND DAVID CARD*

We compare a contracting model and a labor supply model. One test is whether earnings changes are more variable than hours changes, as predicted by the labor supply model, or less variable, as predicted by the contracting model. We apply this test to two longitudinal surveys and find that earnings are somewhat more variable than hours for men who never change employers. The estimates suggest that changes in earnings and hours not associated with measurement error occur at fixed wage rates.

Despite rapid progress over the last decade in modeling employment contracts and recent evidence on the importance of long-term jobs in the economy, microeconomic studies of labor supply continue to interpret individual hours and earnings data in terms of an auction model of the labor market.[1] Traditional labor supply models assume that earnings represent the product of desired hours and market wage rates. Contracting models, on the other hand, interpret earnings as optimal consumption for the payment period, including savings and insurance payments from firms to workers.[2] If savings and insurance are important components of earnings, then average hourly earnings provide, at best, noisy information on underlying productivity. Contract models, therefore, offer a simple explanation for the weak link between wage rates and hours that has confounded empirical studies of intertemporal labor supply.[3]

In this paper we compare the implications of a life cycle labor supply model and an intertemporal contracting model for changes in individual earnings and hours over time. Specifically, we compare a dynamic labor supply model in which individuals have access to complete capital markets to a symmetric information contracting model in which employees receive complete insurance from their employers. The critical distinction between these models is whether earnings represent optimal consumption or the product of wage rates and hours of work. We develop a simple test between the two models based on the relative variability of earnings and hours with respect to changes in productivity. If earnings represent the product of wages and hours, then changes in productivity generate bigger changes in earnings than hours. If earnings represent consumption, on the other hand, then changes in productivity generate smaller changes in earnings than hours, provided that leisure is a normal good.

This simple test is complicated by changes in earnings and hours that may occur with movements across jobs. Although the labor

*Department of Economics, Princeton University, Princeton, NJ 08544. We acknowledge the helpful comments of Gary Chamberlain, Henry Farber, Zvi Griliches, and John Kennan, on earlier drafts. We are also grateful to the Department of Economics at the Massachusetts Institute of Technology for their hospitality while this paper was completed. This paper was originally circulated under the title "Intertemporal Labor Supply in the Presence of Long Term Contracts." Our research was partially supported by the University of Chicago Graduate School of Business.

[1] The recent literature on labor supply is surveyed by Mark Killingsworth (1983) and John Pencavel (1987). The long duration of jobs is emphasized by Robert Hall (1982). Estimates of completed job durations for adult males are presented by Katharine Abraham and Henry Farber (1985).

[2] Oliver Hart (1983) and Sherwin Rosen (1985) provide useful surveys of the literature on contracting models.

[3] This point is emphasized by Rosen (1985). James Brown (1982) estimates aggregate employment equations that include savings and insurance components in earnings.

supply model makes no distinction between changes within and across jobs, the contracting model is employer-specific. We therefore propose the following test of the two alternative models: compare the relative contribution of productivity shocks to changes in earnings and changes in hours for workers who are observed on the same job over time. If, as the intertemporal contracting model suggests, these workers are fully insured, then the contribution of productivity shocks to changes in earnings should be smaller than the contribution of productivity shocks to changes in hours. If the labor supply model is correct, on the other hand, then the contribution of these shocks to changes in earnings should be greater than the contribution to changes in hours.

Our empirical analysis is conducted with data from the *Panel Study of Income Dynamics* (*PSID*) and the *National Longitudinal Survey of Men 45–59* (*NLS*). In order to identify workers who are potentially covered by long-term contracts, we distinguish between individuals who report the same employer during the sample period, and individuals who change employers at least once. We find that earnings and hours changes are substantially less variable for individuals who do not change employers. For both groups of workers, the contribution of productivity shocks to earnings is *greater* than the contribution to hours, although we cannot reject the hypothesis that earnings and hours move proportionately with changes in productivity. This finding casts doubt on the usefulness of either consumption-smoothing contract models or dynamic labor supply models. In fact, a simple interpretation of the data is that earnings and hours vary at fixed hourly wage rates.

Since the focus of this paper is on the contrast between labor supply and contracting models of earnings and hours variation, we concentrate on relatively simple specifications of the components of earnings and hours. In our other work (1986), we investigate more general factor-analytic models of the covariance structure of earnings and hours changes. Although these models provide a somewhat better description of earnings and hours changes, we find that for adult male workers, most systematic hours variation occurs at fixed wage rates. The conclusions in this paper, therefore, are not affected by extensions to the simple components of variance models of earnings and hours presented here.

Section I presents a simple theoretical analysis of intertemporal contracting and intertemporal labor supply models. For both models we derive the implications of changes in productivity for relative changes in earnings and hours. These implications provide the basis for our empirical test between the models.

In Section II, we show how to estimate the theoretical models using the variances, autocovariances, and cross-covariances of earnings and hours changes from individual longitudinal data. A two-factor variance components model provides a convenient framework for distinguishing changes in productivity from other sources of earnings and hours variation, including changes in tastes and measurement error.

Section III summarizes the data from both surveys and presents estimates of the structural parameter that distinguishes the contracting and labor supply models. The covariance structure of earnings and hours changes is remarkably similar in the two surveys. Our main empirical finding is that productivity variation affects earnings at least as much as hours. This is true for individuals who have the same employer in all years and for those who change employers. The data therefore provide some evidence against a contracting interpretation. They also suggest, however, that productivity-related changes in earnings and hours occur at fixed wage rates.

I. Earnings and Hours under Long-Term Contracting Models and Life Cycle Labor Supply Models

Here we present a simple dynamic model of earnings and hours determination under long-term employment contracts.[4] We also

[4] The modern analysis of implicit contracts begins with Walter Oi (1962) and Rosen (1968). The macroeco-

present a model of earnings and hours determination under a standard life cycle labor supply framework.[5] We make identical assumptions about preferences and individual productivity in the two models. For the contracting model, we assume that employers have access to complete capital and insurance markets. For the labor supply model, we assume that individuals have direct access to these markets. Our models, therefore, contrast a widely used version of the intertemporal labor supply model with a class of testable contracting models.

Individual productivity is modeled as a random variable drawn from a sequence of distributions that are common knowledge for both workers and firms. Productivity is the only source of uncertainty in the model. Apart from firm-specific training and recruiting costs, individuals are equally productive at all firms. Long-term attachments between workers and firms arise from two sources: first, the desire to avoid recurrent training costs, which occurs in either the contracting or labor supply model; and, second, the desire to smooth consumption vis-à-vis productivity, which is associated with long-term attachments in the contracting model.

Preferences for consumption and leisure within periods are modeled as a general function of consumption, hours of work, and age. Preferences are assumed to be additively separable over time and across states of productivity. The worker's intertemporal objective is to maximize the expected discounted value of lifetime utility. In the contracting model, the expectation is taken over the distribution of individual productivities. In the labor supply model, the expectation is taken over the distribution of market wages, which is assumed to be identical to the distribution of individual productivities.

Let θ_t represent the productivity of a given individual in period t.[6] Assume that θ_t is distributed on the interval (θ_l, θ_u) according to a known distribution function $F_t(\theta_t)$. Let $u(c_t(\theta_t), h_t(\theta_t), t)$ represent a concave von Neumann-Morgenstern utility function over consumption (c) and hours of work (h) in period t. Let the utility discount rate be ρ. The worker's objective is to maximize expected utility denoted by

$$(1)\quad \sum_{t=0}^{T}\left(\frac{1}{1+\rho}\right)^{t} \times \int_{\theta_l}^{\theta_u} u(c_t(\theta_t), h_t(\theta_t), t)\, dF_t(\theta_t),$$

where T represents a fixed planning horizon.

Consider the long-term contracting model first. Firms offer contracts consisting of contingent labor demand functions $h_t(\theta_t)$ and contingent earnings functions $g_t(\theta_t)$ for $t = 1,\dots,T$. Since workers have no access to capital markets, $g_t(\theta_t) = c_t(\theta_t)$ for all t. If productivity is θ_t in period t, the firm's revenues are $\theta_t h_t(\theta_t)$ and its costs are $g_t(\theta_t)$. We assume that θ_t is observable and, therefore, contracts are fully enforceable. We also assume that firms are risk neutral and can borrow and lend at the constant real interest rate r.[7] Competition among firms for the services of a worker with the sequence of productivity distributions $\{F_t(\theta_t)\}$ implies that contracts offered to that worker have expected present value equal to the training

nomic implications of employment contracts were emphasized by Martin Baily (1974), Donald Gordon (1974), and Costas Azariadis (1975).

[5]Studies of life cycle labor supply and consumption originate in Franco Modigliani and Richard Brumberg's (1954) analysis of the divergence between planned consumption and earnings over the life cycle and in Milton Friedman's (1957) study of the consumption function. Robert Lucas and Leonard Rapping (1969) use a two-period model in their influential study of intertemporal substitution and labor supply. Multiperiod labor supply is considered by James Heckman (1974, 1976), Gilbert Ghez and Gary Becker (1975), and many subsequent authors, in particular, Thomas MaCurdy (1981).

[6]For notational simplicity, we suppress the dependence of θ_t on the individual. Randomness of θ_t is over *ex ante* identical individuals.

[7]Implicitly we are assuming that productivity risks are fully diversifiable. See Rosen (1985, pp. 1153–54) for a discussion of aggregate vs. idiosyncratic productivity risks and the implications of nondiversifiability.

costs, R, for that worker:

$$(2)\quad \sum_{t=0}^{T}\left(\frac{1}{1+r}\right)^{t} \times \int_{\theta_l}^{\theta_u}[\theta_t h_t(\theta_t) - g_t(\theta_t)]\, dF_t(\theta_t) = R.$$

Pointwise optimization of the Lagrangian expression for the maximization of (1), subject to (2), leads to the first-order conditions:

$$(3a)\quad \left(\frac{1+r}{1+\rho}\right)^{t} u_c(c_t(\theta_t), h_t(\theta_t), t) - \lambda = 0,$$

$$(3b)\quad \left(\frac{1+r}{1+\rho}\right)^{t} u_h(c_t(\theta_t), h_t(\theta_t), t) + \lambda\theta_t = 0,$$

where u_c and u_h represent the partial derivatives of $u(\cdot,\cdot,\cdot)$ with respect to c and h, and λ represents the multiplier associated with the constraint (2). Equations (3a) and (3b) have the familiar implications that the marginal utility of consumption follows a deterministic trend, while the marginal rate of substitution between consumption and leisure equals θ_t for each realization of productivity.

Differentiation of the first-order conditions (3a) and (3b) and some rearrangement yields

$$(4)\quad \frac{\partial \log h_t}{\partial \log \theta_t} - \frac{\partial \log h_t}{\partial \log \nu_t} = \frac{c_t}{\theta_t h_t}\frac{\partial \log c_t}{\partial \log \theta_t},$$

where $\nu_t \equiv \lambda((1+\rho)/(1+r))^t$. To understand the implications of equation (4), consider the log-linear approximation to the solution of equations (3a) and (3b):

$$(5a)\quad \log c_t = \phi \log \theta_t - \alpha \log \nu_t + a_t,$$

$$(5b)\quad \log h_t = \eta \log \theta_t + \delta \log \nu_t + b_t,$$

where a_t and b_t are time-varying terms in the log-linear approximation representing shifts in tastes for consumption and leisure. The parameter ϕ represents the substitution elasticity between consumption and leisure holding constant the marginal utility of wealth: the sign of ϕ depends on the sign of u_{ch}. If the permanent income hypothesis is correct, for example, then $\phi = 0$ and consumption is independent of productivity. The parameter η represents the elasticity of substitution of labor supply over time and across states of θ; therefore, $\eta \geq 0$. The parameter $-\alpha$ represents the elasticity of consumption demand with respect to the marginal utility of wealth; if consumption is a normal good, then $\alpha \geq 0$.[8] Finally, the parameter δ represents the elasticity of labor supply with respect to the marginal utility of wealth; if leisure is a normal good, then $\delta \geq 0$. Since $E[c_t] \cong E[\theta_t h_t]$ by constraint (2),[9] the restriction (4) implies (to a first-order approximation) $\eta - \delta \cong \phi$, or

$$(6)\qquad \mu \equiv \phi/\eta \cong 1 - (\delta/\eta).$$

The parameter μ represents the relative sensitivity of consumption and hours choices to changes in productivity. Even in the absence of direct information on productivity, μ is identifiable from information on the relative variability of earnings and hours. If $\mu \geq 1$, then $\delta \leq 0$; that is, if consumption is more variable than hours with respect to changes in productivity, then leisure is an inferior good. If $\delta > 0$ is treated as a maintained hypothesis, then the intertemporal contracting model presents one testable implication: namely, that changes in productivity influence earnings *less* than hours, on average.

Now consider the intertemporal labor supply model. We assume that workers have access to risk-neutral insurance and capital markets so that the life cycle budget con-

[8]A concave utility function implies that λ is a decreasing function of wealth, and therefore that the sign of the derivative of the demand for consumption goods with respect to λ is the same as the sign of the derivative of demand for consumption goods with respect to income.

[9]This holds for small training costs, R.

straint can be replaced by its expectation:[10]

$$(7)\quad \sum_{t=0}^{T}\left(\frac{1}{1+r}\right)^{t} \times \int_{\theta_l}^{\theta_u}[\theta_t h_t(\theta_t)-c_t(\theta_t)]\,dF_t(\theta_t)=0.$$

The first-order conditions for the maximization of (1) subject to the constraint (7) are identical to (3a) and (3b). Labor earnings, however, are now described by $g_t(\theta_t)=\theta_t h_t(\theta_t)$. The log-linear form of the solution for g_t and h_t becomes

$$(8a)\quad \log g_t=(1+\eta)\log\theta_t+\delta\log\nu_t+b_t;$$

$$(8b)\quad \log h_t=\eta\log\theta_t+\delta\log\nu_t+b_t,$$

where, as before, b_t represents a time-varying component of tastes for leisure. Under the labor supply interpretation, the variability of earnings relative to hours with respect to changes in productivity is given by

$$(9)\quad \mu\equiv(1+\eta)/\eta.$$

Since earnings represent the product of wages and hours in the labor supply model, earnings must respond *more* than hours to changes in productivity.

Our analysis of the contracting model shows that the elasticity of earnings with respect to productivity is less than the elasticity of hours with respect to productivity if leisure is a normal good. Our analysis of the intertemporal labor supply model shows that the relation between these elasticities is reversed under identical assumptions.[11] In the next section we develop a statistical model for estimating the critical parameter μ, the ratio of the two elasticities.

[10] Robert Topel and Finis Welch (1986) refer to this model as a contracting model with self-insurance.

[11] If the worker does not have access to complete insurance markets, the implications of the intertemporal labor supply model are unchanged, provided that the marginal utility of income is held constant. Since changes in productivity may be correlated with changes in the marginal utility of wealth, estimation of μ requires additional structure. (See our other paper.)

II. Econometric Models for the Covariance Structure of Earnings and Hours Changes

Our empirical strategy is to fit the model of earnings and hours implied by the contracting model (equations (5a) and (5b)) and the labor supply model (equations (8a) and (8b)) to estimated covariance matrices of earnings and hours changes from longitudinal survey data. We develop a two-factor interpretation of earnings and hours changes that allows us to distinguish between productivity components of earnings and hours, on one hand, and components of variance associated with preference variation and survey measurement error, on the other. The two-factor structure permits direct estimation of μ, the relative variability of earnings and hours with respect to productivity changes, as well as an overall measure of the goodness of fit of this simple class of models to the covariance structure of earnings and hours changes.

For the contracting model, the first step is to express equations (5a) and (5b) in first-difference form, taking account of individual-specific components. Since earnings are identical to consumption in this model, we substitute $\log g_t$ for $\log c_t$. Let $\Delta\log g_{it}$ and $\Delta\log h_{it}$ represent the observed changes in the logarithms of real annual earnings and annual hours for individual i between periods $t-1$ and t, respectively. Append a survey measurement error u_{it}^* to the expression for $\log g_{it}$ and a survey measurement error v_{it}^* to the expression for $\log h_{it}$. Then, equations (5) imply

$$(10a)\quad \Delta\log g_{it}=\phi\Delta\log\theta_{it} -\alpha(\rho-r)+\Delta a_{it}+\Delta u_{it}^*,$$

$$(10b)\quad \Delta\log h_{it}=\eta\Delta\log\theta_{it} +\delta(\rho-r)+\Delta b_{it}+\Delta v_{it}^*.$$

These equations express observed changes in earnings and hours in terms of changes in productivity, changes in tastes for consumption and leisure, and changes in measurement error. Since employees can perfectly insure individual productivity variation, the marginal utility of wealth follows a deterministic trend and contributes only the constant $(\rho - r)$ to earnings and hours changes.

In the labor supply model, the equation for the change in hours is identical to (10b). The equation for the change in earnings, however, is

$$(11)\quad \Delta \log g_{it} = (1+\eta)\Delta \log \theta_{it} + \delta(\rho - r) + \Delta b_{it} + \Delta u^*_{it}.$$

Equations (10a) and (11) are very similar. The statistically identifiable difference between the contracting and labor supply models arises from the different coefficients on the change in individual productivity. To clarify this point, we complete the models by specifying the covariance structures of individual productivity, preference variation, and measurement error.

We adopt a linear specification for individual productivity consisting of a permanent individual effect (θ_i), an aggregate time effect (d_t), a quadratic labor force experience effect, and a purely stochastic component (z_{it}):

$$\log \theta_{it} = \theta_i + d_t + \zeta_\theta x_{it} + \tfrac{1}{2}\xi_\theta x^2_{it} + z_{it},$$

where x_{it} represents the labor force experience of individual i at the beginning of year t. Since labor force experience increases by one unit each year, the change in the logarithm of individual productivity is

$$(12)\quad \Delta \log \theta_{it} = \kappa_{\theta t} + \xi_\theta x_{i0} + \Delta z_{it},$$

where x_{i0} represents the labor force experience of individual i at the beginning of the survey and $\kappa_{\theta t}$ is a time effect that incorporates the change in the aggregate productivity shock as well as the change in average labor force experience.[12]

In a similar fashion, we assume that the preference variations (a_{it} and b_{it}) contain permanent individual effects, aggregate time effects, quadratic experience effects, and stationary, serially uncorrelated random components:

$$a_{it} = a_i + a_t + \zeta_a x_{it} + \tfrac{1}{2}\xi_a x^2_{it} + \varepsilon_{ait},$$
$$b_{it} = b_i + b_t + \zeta_b x_{it} + \tfrac{1}{2}\xi_b x^2_{it} + \varepsilon_{bit}.$$

These specifications permit tastes for consumption and leisure to exhibit homogeneous curvature over the life cycle. The vector of transitory deviations from the life cycle profile of preferences $(\varepsilon_{ait}, \varepsilon_{bit})$ is assumed to be independent and identically distributed for all i and t with an unrestricted contemporaneous covariance matrix. The first differences of the preference variations can be written as

$$(13a)\quad \Delta a_{it} = \kappa_{at} + \xi_a x_{i0} + \Delta \varepsilon_{ait},$$

$$(13b)\quad \Delta b_{it} = \kappa_{bt} + \xi_b x_{i0} + \Delta \varepsilon_{bit},$$

where κ_{at} and κ_{bt} are composite time effects that incorporate changes in a_t and b_t as well as changes in average labor force experience.[13]

Finally, we assume that the vector of survey measurement errors (u^*_{it}, v^*_{it}) contains permanent and purely transitory errors:

$$u^*_{it} = u^*_i + \varepsilon_{uit}, \quad v^*_{it} = v^*_i + \varepsilon_{vit}.$$

The permanent measurement error components, represented by u^*_i and v^*_i, model systematic deviations of the survey instrument from the theoretically appropriate concepts. We assume that the vector of transitory errors $(\varepsilon_{uit}, \varepsilon_{vit})$ is independent and identically distributed with an unrestricted contemporaneous covariance matrix. The first differences of the measurement errors can be written as

$$(14a)\quad \Delta u^*_{it} = \Delta \varepsilon_{uit},$$

$$(14b)\quad \Delta v^*_{it} = \Delta \varepsilon_{vit}.$$

[12] The effect $\kappa_{\theta t} = \Delta d_t + \zeta_\theta - \frac{1}{2}\xi_\theta + \xi_\theta t$, since $x_{it} - x_{it-1} = 1$ and $x^2_{it} - x^2_{it-1} = 2x_{i0} + 2t - 1$, where $x_{i0} =$ labor force experience at the beginning of the first survey period.

[13] The term $\kappa_{at} = \Delta a_t + \zeta_a + \xi_a t - \frac{1}{2}\xi_a$ and similarly for κ_{bt}.

Equations (14a) and (14b) indicate that only the transitory measurement errors contribute to the covariance structure of earnings and hours changes. Permanent response biases are eliminated by differencing.

Under the assumptions we have made, preference variation and survey measurement errors are statistically indistinguishable, since the first differences of both components represent first differences of uncorrelated vectors. For notational simplicity, we combine the transitory preference variation components, $\Delta\varepsilon_{ait}$ and $\Delta\varepsilon_{bit}$, with the transitory survey measurement error components, $\Delta\varepsilon_{uit}$ and $\Delta\varepsilon_{vit}$, to form a single vector of variance components $(\Delta u_{it}, \Delta v_{it})$. In the labor contract model, the preference variation and measurement error components of variance in earnings and hours changes are given by

$$(15a) \qquad \Delta u_{it} = \Delta\varepsilon_{ait} + \Delta\varepsilon_{uit},$$

$$(15b) \qquad \Delta v_{it} = \Delta\varepsilon_{bit} + \Delta\varepsilon_{vit}.$$

In the labor supply model, on the other hand, the preference variation and measurement error components of variance in earnings and hours are given by

$$(15a') \qquad \Delta u_{it} = \Delta\varepsilon_{bit} + \Delta\varepsilon_{uit},$$

$$(15b') \qquad \Delta v_{it} = \Delta\varepsilon_{bit} + \Delta\varepsilon_{vit}.$$

In either case, the vector $(\Delta u_{it}, \Delta v_{it})$ is independently and identically distributed across individuals with an arbitrary contemporaneous covariance matrix and a known autocovariance structure. Specifically, the vector $(\Delta u_{it}, \Delta v_{it})$ is a bivariate first-order moving average process with first-order autocorrelations equal to $-\frac{1}{2}$.[14] This simple autocorrelation structure reflects the following observation: if y_t is serially uncorrelated with variance σ^2, then the variance of Δy_t is $2\sigma^2$, the covariance of Δy_t with Δy_{t-1} is $-\sigma^2$, and the covariance between Δy_t and Δy_s is zero for $|t-s|>1$.

Combining equations (12)–(15), the equations for the changes in log earnings and log hours in the labor contracting model can be simplified to

$$(16a) \qquad \Delta \log g_{it} = \kappa_{gt} + \xi_g x_{i0} + \phi\Delta z_{it} + \Delta u_{it}$$

$$(16b) \qquad \Delta \log h_{it} = \kappa_{ht} + \xi_h x_{i0} + \eta\Delta z_{it} + \Delta v_{it},$$

where κ_{gt} and κ_{ht} combine the aggregate time effects of equations (12) and (13); ξ_g and ξ_h combine the linear labor force experience effects of equations (12) and (13); Δu_{it} and Δv_{it} represent the combined preference variation and survey measurement errors from equation (15); and Δz_{it} represents the individual productivity variation from equation (12). For the labor supply model, the hours equations is the same as (16b). The equation for earnings, on the other hand, becomes

$$(17) \qquad \Delta \log g_{it} = \kappa'_{gt} + \xi'_g x_{i0} + (1+\eta)\Delta z_{it} + \Delta u_{it},$$

where κ'_{gt} combines the aggregate time effects of equations (12) and (13), and ξ'_g combines the linear labor force experience effects of changes in productivity and preferences. In general, the year effects κ_{gt} and κ'_{gt} and the experience slopes ξ_g and ξ'_g are different in (16a) and (17).

Neither the labor supply model nor the labor contracting model, however, imposes any restrictions on the year effects or experience slopes of equations (16) and (17). Under our assumptions, individual productivity changes and preference variations contribute three unrestricted time effects ($\kappa_{\theta t}$, κ_{at}, and κ_{bt}) to the changes in log earnings and log hours in the labor contract model, or two unrestricted time effects ($\kappa_{\theta t}$ and κ_{bt}) in the labor supply model. The cross-sectional means of $\Delta \log g_{it}$ and $\Delta \log h_{it}$ in period t (controlling for experience) are sufficient to estimate only two linear combinations of these effects (κ_{gt} and κ_{ht}). Similarly, there are three unrestricted labor force experience effects (ξ_θ, ξ_a, and ξ_b) in the labor contract model, or two unrestricted experience effects (ξ_θ and ξ_b) in the labor supply model. Again,

[14] This model also implies that the ratio of either first-order cross covariance ($\mathrm{Cov}(\Delta \log g_{it}, \Delta \log h_{it+1})$, or $\mathrm{Cov}(\Delta \log g_{it}, \Delta \log h_{it-1})$) to the zero-order covariance ($\mathrm{Cov}(\Delta \log g_{it}, \Delta \log h_{it})$) is $-\frac{1}{2}$.

however, we can only identify two experience slopes (ξ_g and ξ_h).[15] Therefore, the coefficients of the multivariate regression of individual i's changes in log earnings and log hours on time effects and initial labor force experience are unrestricted by either model.[16]

Equations (16) and (17) do, however, provide a simple two-factor model for the residuals from the regression of changes in individual earnings and hours on time effects and labor force experience. According to these equations, unpredicted changes in earnings and hours contain a time-stationary preference and measurement error component, with a known autocorrelation structure, and a productivity component, with an arbitrary autocorrelation structure.

The relative contribution of productivity changes to earnings and hours changes depends on the parameters ϕ and η. In the absence of direct information on the variance of individual productivity shocks, these parameters are not separately identifiable from the covariance structure of earnings and hours changes. The critical parameter μ, which represents the ratio of ϕ to η, is identifiable from the *relative* covariances of earnings and hours changes, however. In particular, μ is identifiable if changes in earnings and hours exhibit second-order or higher autocorrelation, since the preference variation and measurement error components are assumed to contribute only first-order autocorrelation. Alternatively, μ is identifiable if the first-order autocorrelations of earnings and hours changes are not identically equal to $-\frac{1}{2}$, since the preference variation and measurement error components are assumed to have first-order autocorrelations equal to $-\frac{1}{2}$. Lastly, μ is identifiable if the autocovariances and cross covariances of earnings and hours changes are not time stationary, since the preference variation and measurement error components are assumed to be stationary.

While μ is potentially identifiable from the covariance structure of earnings and hours changes, the separate variance contributions of changes in productivity, preference shifts, and measurement errors are not identifiable. All three components contribute to the variances and first-order autocovariances of changes in earnings and hours. It is impossible to determine their separate contributions, however, without further assumptions on either the serial correlation properties of the productivity shocks, or the correlations of the measurement errors and preference variations.

Table 1 displays the theoretical formulas for the autocovariances and cross covariances of earnings and hours changes implied by equations (16) and (17). The variables $\Delta \log \tilde{g}_{it}$ and $\Delta \log \tilde{h}_{it}$ are defined as the deviations of $\Delta \log g_{it}$ and $\Delta \log h_{it}$, respectively, from their conditional (regression-adjusted) means given t and x_{i0}. We refer to these variables as experience-adjusted changes in log earnings and log hours. The formulas are written in terms of the parameter μ so that they apply to either the contracting or labor supply model.[17] Table 1 shows how the covariance structure of earnings and hours changes depends on the covariance structure of individual productivity changes (Δz_{it}), the covariance structure of

[15] If one assumes that life cycle tastes for leisure are linear (rather than quadratic) in labor market experience, then the life cycle labor supply model implies that the ratio of the labor force experience coefficient of earnings changes (ξ_g) to the labor force experience coefficient of hours (ξ_h) equals $(1+\eta)/\eta$. Provided that tastes for leisure are linear in experience, then, the intertemporal substitution elasticity may be estimated directly from the covariances of earnings and hours changes with experience. This technique is equivalent to the instrumental variables estimation schemes used by MaCurdy and by Joseph Altonji (1986). For the samples considered in this paper, the *PSID* yields an estimate of η equal to 1.52 (with a standard error of .44) and the *NLS* yields an estimate of η equal to -1.62 (with a standard error of .61) when this technique is used.

[16] The experience slopes of earnings and hours changes are actually restricted to be constant over time. Both the contracting model and the labor supply model imply this restriction, given our model for individual productivity and preference variations. The experience slopes from the *PSID* are consistent with this restriction ($\chi^2 = 25.80$ with 18 degrees of freedom, probability value $=.104$). The experience slopes from the *NLS* are not consistent with this restriction ($\chi^2 = 34.66$ with 8 degrees of freedom, probability value $\cong 0$).

[17] Expressing the covariances as functions of μ requires the definition of $\Delta \tilde{z}_{it} \equiv \eta \Delta z_{it}$.

TABLE 1—IMPLIED COVARIANCES OF EXPERIENCE-ADJUSTED CHANGES IN LOG EARNINGS AND LOG HOURS: CONTRACTING AND LABOR SUPPLY MODELS

Covariance Element	Implied Formulae
1) $\text{Var}(\Delta \log \tilde{g}_{it})$	$\mu^2 \text{Var}(\Delta \tilde{z}_{it}) + 2\sigma_u^2$
2) $\text{Var}(\Delta \log \tilde{h}_{it})$	$\text{Var}(\Delta \tilde{z}_{it}) + 2\sigma_v^2$
3) $\text{Cov}(\Delta \log \tilde{g}_{it}, \Delta \log \tilde{h}_{it})$	$\mu \text{Var}(\Delta \tilde{z}_{it}) + 2\rho_{uv}\sigma_u\sigma_v$
4) $\text{Cov}(\Delta \log \tilde{g}_{it}, \Delta \log \tilde{g}_{it-1})$	$\mu^2 \text{Cov}(\Delta \tilde{z}_{it}, \Delta \tilde{z}_{it-1}) - \sigma_u^2$
5) $\text{Cov}(\Delta \log \tilde{h}_{it}, \Delta \log \tilde{h}_{it-1})$	$\text{Cov}(\Delta \tilde{z}_{it}, \Delta \tilde{z}_{it-1}) - \sigma_v^2$
6) $\text{Cov}(\Delta \log \tilde{g}_{it}, \Delta \log \tilde{h}_{it-1})$, $\text{Cov}(\Delta \log \tilde{h}_{it}, \Delta \log \tilde{g}_{it-1})$	$\mu \text{Cov}(\Delta \tilde{z}_{it}, \Delta \tilde{z}_{it-1}) - \rho_{uv}\sigma_u\sigma_v$
7) $\text{Cov}(\Delta \log \tilde{g}_{it}, \Delta \log \tilde{g}_{it-2})$	$\mu^2 \text{Cov}(\Delta \tilde{z}_{it}, \Delta \tilde{z}_{it-2})$
8) $\text{Cov}(\Delta \log \tilde{h}_{it}, \Delta \log \tilde{h}_{it-2})$	$\text{Cov}(\Delta \tilde{z}_{it}, \Delta \tilde{z}_{it-2})$
9) $\text{Cov}(\Delta \log \tilde{g}_{it}, \Delta \log \tilde{h}_{it-2})$, $\text{Cov}(\Delta \log \tilde{h}_{it}, \Delta \log \tilde{g}_{it-2})$	$\mu \text{Cov}(\Delta \tilde{z}_{it}, \Delta \tilde{z}_{it-2})$

Note: $\Delta \log \tilde{g}_{it} \equiv \Delta \log g_{it} - \kappa_{gt} - \xi_g x_{i0}$, $\Delta \log \tilde{h}_{it} \equiv \Delta \log h_{it} - \kappa_{ht} - \xi_h x_{i0}$, $\Delta \tilde{z}_{it} \equiv \eta \Delta z_{it}$, $\rho_{uv} \equiv$ Correlation $(\Delta u_{it}, \Delta v_{it})$, $\sigma_u^2 \equiv \text{Var}(\Delta u_{it})$, $\sigma_v^2 \equiv \text{Var}(\Delta v_{it})$,

the preference variation and measurement error process (Δu_{it} and Δv_{it}), and μ. The formulas in Table 1 form the basis for our empirical test of the contracting model vs. the labor supply model.

III. A Test of the Contracting Model vs. the Labor Supply Model, using Longitudinal Data on Adult Males

The longitudinal earnings and hours data used in this paper are drawn from the *Panel Study of Income Dynamics* and the *National Longitudinal Survey of Men 45–59*. These surveys use substantially different methods for determining annual earnings and annual hours worked. The *PSID* collects information on various components of labor earnings while the *NLS* collects data on wage and salary income in a single question. The *PSID* likewise collects information on hours for both primary and secondary jobs. The *NLS*, on the other hand, collects annual hours information for the main job only. Our Data Appendix describes the variables we actually used and the survey questions from which these variables were derived.[18]

[18]See Survey Research Center (1981) and Center for Human Resources Research (1977, 1980) for documentation of the survey variables and procedures.

From the *PSID*, we selected 1448 male household heads whose records indicate nonzero earnings and hours in each year from 1969 to 1979 (the third through thirteenth waves of the survey). We included only those male household heads who were between the ages of 21 and 64 in all eleven sample years. The "one-employer" subsample was defined on the basis of answers to the questions about present employment status and reason for changing employment status. If an individual was currently employed, or temporarily laid off and reported having his present job for at least one year (including promotions), then the individual was considered to have the same employer as in the previous year. An individual with the same employer as in the previous year for all years from 1970 to 1979 was included in the one-employer subsample. There were 618 individuals who satisfied this condition. The remaining 830 individuals were included in the "multiple-employers" subsample. Every member of the multiple-employers subsample experienced at least one change of employer during the period from 1969 to 1979. Table 2 presents means and standard deviations of the changes in log real annual earnings, and log annual hours, as well as basic demographic variables for the *PSID* sample and subsamples.

From the *National Longitudinal Survey of Men 45–59*, we selected 1309 men whose

TABLE 2—SAMPLE AND SUBSAMPLE CHARACTERISTICS FOR THE *PSID* AND THE *NLS* OF OLDER MEN: MEANS AND STANDARD DEVIATIONS FOR THE INDICATED YEARS

	PSID				*NLS*[a]		
Year	All	One Employer	Multiple Employers	Year	All	One Employer	Multiple Employers
Change in Log Real Earnings[b]							
1969–70	2.5	1.8	3.0	1966–67	4.4	2.6	6.7
	(40.)	(24.)	(49.)		(31.)	(20.)	(41.)
1970–71	3.0	2.5	3.4	1967–69	4.6	4.0	5.3
	(40.)	(24.)	(48.)		(31.)	(23.)	(39.)
1971–72	6.9	5.8	7.7	1969–71	2.5	3.2	1.6
	(41.)	(30.)	(48.)		(31.)	(23.)	(38.)
1972–73	4.7	3.5	5.6	1971–73	−1.2	1.2	−4.4
	(37.)	(29.)	(41.)		(38.)	(29.)	(47.)
1973–74	−5.5	−3.8	−6.8	1973–75	−16.4	−6.9	−28.7
	(36.)	(26.)	(42.)		(54.)	(32.)	(72.)
1974–75	−4.2	−2.8	−5.3				
	(43.)	(28.)	(51.)				
1975–76	4.1	2.7	5.2				
	(47.)	(30.)	(57.)				
1976–77	2.5	2.8	2.3				
	(44.)	(25.)	(54.)				
1977–78	0.2	0.0	0.4				
	(44.)	(27.)	(53.)				
1978–79	−5.5	−4.4	−6.4				
	(42.)	(28.)	(51.)				
Change in Log Annual Hours[b]							
1969–70	−0.8	−1.8	−0.1	1966–67	−0.1	−0.8	0.8
	(35.)	(21.)	(42.)		(29.)	(19.)	(37.)
1970–71	−0.3	−0.6	−0.0	1967–69	−0.3	−0.5	−0.2
	(34.)	(18.)	(42.)		(27.)	(21.)	(33.)
1971–72	2.0	−0.1	3.6	1969–71	−0.9	−0.3	−1.6
	(34.)	(19.)	(42.)		(26.)	(19.)	(33.)
1972–73	1.9	2.1	1.7	1971–73	−1.2	0.9	−3.9
	(28.)	(19.)	(34.)		(29.)	(18.)	(38.)
1973–74	−4.1	−2.5	−5.3	1973–75	−10.9	−3.0	−21.1
	(27.)	(18.)	(32.)		(46.)	(22.)	(64.)
1974–75	−2.4	−1.1	−3.4				
	(33.)	(20.)	(41.)				
1975–76	0.6	−0.2	1.2				
	(38.)	(22.)	(47.)				
1976–77	0.3	0.6	0.1				
	(38.)	(20.)	(47.)				
1977–78	0.5	−0.9	1.6				
	(37.)	(21.)	(45.)				
1978–79	−4.2	−1.3	−6.4				
	(37.)	(24.)	(44.)				
Demographic Characteristics							
Age	35.8	38.2	34.1		49.1	48.9	49.5
	(9.)	(8.)	(9.)		(3.)	(3.)	(3.)
Potential Experience	18.9	21.4	17.1		34.4	33.8	35.2
	(11.)	(10.)	(11.)		(6.)	(5.)	(5.)
Percent Nonwhite	27.3	27.7	27.0		29.6	30.8	28.2
Sample Size	1448[c]	618	830		1309[d]	735	574

Note: Standard deviations are shown in parentheses.

[a] Statistics from the *NLS* are *not* at annual rates.

[b] Means and standard deviations times 100.

[c] Eight outliers with average hourly earnings greater than $100/hour (1967 dollars) have been deleted.

[d] Nine outliers with absolute changes in log earnings or log hours in excess of 3.5 have been deleted.

records indicate nonzero earnings and hours for each of the years 1966, 1967, 1969, 1971, 1973, and 1975.[19] We included only those males who were between the ages of 45 and 64 in all six sample years. The one-employer subsample was defined on the basis of the number of years the individual had worked for his current employer in 1971, and whether or not the individual worked for a different employer in 1973 or 1975. An individual who had worked for his current employer at least five years in 1971 and who did not change employers in either 1973 or 1975 was included in the one-employer subsample. There were 735 individuals who satisfied this condition. The remaining 574 individuals were included in the multiple-employers subsample. Means and standard deviations for the *NLS* sample are also presented in Table 2. For the interpretation of the *NLS* data, it is important to note that later waves of the survey were administered biennially. The changes in earnings and hours from 1969 to 1975 for the *NLS* therefore refer to changes in annual totals differenced over two-year intervals. These changes are *not* reported at annual rates in Table 2.

Table 2 shows that average age and potential labor force experience (age minus years of education minus five) differ substantially between the *PSID* and *NLS* samples because of design differences in the underlying surveys. The *NLS* workers are an average of 13.3 years older and 15.5 years more experienced than the *PSID* workers. We use experience-adjusted changes in earnings and hours in all subsequent calculations in this paper to correct for any systematic differences between the *PSID* and *NLS* samples arising from this difference in labor force experience. Both surveys also oversampled nonwhites. The percentage of nonwhites is similar in our two samples, however, and we make no further adjustments to account for the small difference in racial composition between them.[20] There are no important differences in age, labor force experience, or percentage nonwhite between the one-employer and multiple-employers subsamples of either survey.

Table 2 reveals three striking features of the individual earnings and hours changes from our two samples. First, the overall pattern of changes in earnings and hours is similar in both surveys. This conclusion applies when comparing all individuals, individuals with one employer, and individuals with multiple employers. The older sample (*NLS*) experienced slightly larger changes in earnings and hours during the 1973 to 1975 period than the younger sample (*PSID*). Individuals in the multiple-employer subsample of each survey also experienced substantially larger earnings and hours changes during this period than those in the one-employer subsamples. Second, there is significant nonstationarity in the cross-sectional variation of changes in earnings and hours. In the *PSID* sample, earnings and hours changes are most variable in the 1975–76 period and least variable in the 1972–73 period. In the *NLS* sample, on the other hand, these changes are most variable in the 1973–75 period and least variable in the 1969–71 period. Nonstationarity is equally apparent in the one-employer and multiple-employers subsamples of both surveys. Finally, earnings and hours changes are much less variable for the one-employer subsamples of both surveys.[21]

[19] These six survey years were the only waves in which comparable earnings and hours data were collected.

[20] Our *PSID* sample includes the *Survey of Economic Opportunity* subsample, which oversampled low-income households. Our *NLS* sample includes the black enumeration districts, which also oversampled low-income households.

[21] One might ask if variability of earnings and hours changes for the multiple-employers subsamples is similar to the one-employer subsamples if we consider only those years that do not involve employer changes. The answer is yes. In the *PSID* sample, in which we have annual data sampled at an annual rate, individuals experience substantial variability in earnings and hours changes during the three-year period surrounding the change in employer. For the *NLS* sample, in which we cannot perform such a detailed year-to-year analysis because we have annual data sampled at a biennial rate, it is still true that most of the difference in variation between the multiple-employers and one-employer subsamples occurs because of the variability contributed by the period in which the employer change actually occurred. Put differently, most of the added variability in earnings and hours changes for the multiple-employers subsample occurs around the time of employer changes.

TABLE 3—STATIONARY CROSS-COVARIANCE STRUCTURE FOR *PSID* AND *NLS* SAMPLES AND SUBSAMPLES[a]

Sample Cross-Covariance	*PSID* All	*PSID* One Employer	*PSID* Multiple Employers	*NLS*[b] All	*NLS*[b] One Employer	*NLS*[b] Multiple Employers
Earnings Autocovariances						
1) $\mathrm{Var}[\Delta \log \tilde{g}_t]$	.172	.074	.245	.158	.074	.259
	(.011)	(.006)	(.017)	(.011)	(.010)	(.020)
2) $\mathrm{Cov}[\Delta \log \tilde{g}_t, \Delta \log \tilde{g}_{t-1}]$	−.060	−.031	−.081	−.043	−.028	−.069
	(.006)	(.004)	(.009)	(.006)	(.006)	(.011)
3) $\mathrm{Cov}[\Delta \log \tilde{g}_t, \Delta \log \tilde{g}_{t-2}]$	−.007	−.002	−.010	−.001	.002	−.004
	(.003)	(.002)	(.004)	(.003)	(.002)	(.008)
Hours Autocovariances						
4) $\mathrm{Var}[\Delta \log \tilde{h}_t]$	.117	.040	.174	.108	.039	.191
	(.007)	(.003)	(.012)	(.010)	(.005)	(.020)
5) $\mathrm{Cov}[\Delta \log \tilde{h}_t, \Delta \log \tilde{h}_{t-1}]$	−.035	−.016	−.050	−.038	−.018	−.066
	(.003)	(.002)	(.006)	(.006)	(.003)	(.014)
6) $\mathrm{Cov}[\Delta \log \tilde{h}_t, \Delta \log \tilde{h}_{t-2}]$	−.011	−.000	−.019	.008	.001	.017
	(.002)	(.001)	(.003)	(.005)	(.003)	(.011)
Earnings/Hours Cross-Covariances						
7) $\mathrm{Cov}[\Delta \log \tilde{g}_t, \Delta \log \tilde{h}_{t+2}]$	−.006	−.001	−.010	.001	.000	.002
	(.002)	(.001)	(.004)	(.004)	(.001)	(.009)
8) $\mathrm{Cov}[\Delta \log \tilde{g}_t, \Delta \log \tilde{h}_{t+1}]$	−.023	−.005	−.037	−.015	−.002	−.033
	(.004)	(.001)	(.007)	(.004)	(.001)	(.010)
9) $\mathrm{Cov}[\Delta \log \tilde{g}_t, \Delta \log \tilde{h}_t]$	.073	.011	.119	.063	.008	.126
	(.007)	(.001)	(.012)	(.007)	(.002)	(.014)
10) $\mathrm{Cov}[\Delta \log \tilde{g}_t, \Delta \log \tilde{h}_{t-1}]$	−.020	−.003	−.033	−.010	−.002	−.022
	(.004)	(.001)	(.007)	(.004)	(.001)	(.009)
11) $\mathrm{Cov}[\Delta \log \tilde{g}_t, \Delta \log \tilde{h}_{t-2}]$	−.002	.001	−.004	.007	.000	.015
	(.003)	(.001)	(.005)	(.005)	(.002)	(.010)
12) Goodness of Fit for Nonstationary $MA(2)$[c]	137.19	168.09	153.54	15.15	11.84	16.09
	(.053)	(.000)	(.006)	(.233)	(.459)	(.187)
13) Goodness of Fit for Stationary $MA(2)$[d]	180.74	168.18	175.03	79.11	33.21	72.33
	(.000)	(.000)	(.000)	(.000)	(.043)	(.000)

[a] Covariance matrix and standard errors based on equally weighted minimum distance estimates of cross-covariances. The standard errors are shown in parentheses.

[b] *NLS* estimates based on the four changes with gaps of two years: 1967–69, 1969–71, 1971–73, 1973–75.

[c] χ^2 statistic for nonstationary bivariate $MA(2)$ vs. arbitrary process for the bivariate cross-covariance function. The statistic has 112 degrees of freedom for the *PSID* sample and 12 degrees of freedom for the *NLS* sample. Probability values are shown in parentheses.

[d] χ^2 statistic for stationary bivariate $MA(2)$ vs. nonstationary bivariate $MA(2)$ process for cross-covariance function. The statistic has 87 degrees of freedom for the *PSID* sample and 21 degrees of freedom for the *NLS* sample. Probability values are shown in parentheses.

Our theoretical analysis of the contracting and labor supply models focuses on their implications for the autocovariances and cross-covariances of earnings and hours changes. In particular, the identifiability of the parameter that distinguishes the two models depends critically on observable characteristics of the covariance matrix of earnings and hours changes. Therefore, we study this matrix in detail. Table 3 presents the average cross-covariances of the *PSID* and *NLS* samples.[22] The similarity between

There is no substantial difference between the one-employer and multiple-employer subsamples when data from employer changes are excluded. See Altonji and Christina Paxson (1985) for a detailed comparison of hours variability between job changers and stayers in the *PSID*.

[22] The complete covariance matrix of earnings and hours changes for the *PSID* contains 210 unique elements; the complete covariance matrix for the *NLS*

the two samples is also evident in their covariance structure. Both samples and all the subsamples exhibit strong positive correlations between contemporaneous changes in earnings and hours, and strong negative autocorrelation in earnings and hours changes. This similarity is even more remarkable since the *PSID* data represent year-to-year changes, while the *NLS* data represent changes in annual data over two-year intervals.

In comparison to the first-order autocovariances of earnings and hours changes, the second-order autocovariances are relatively small, although nonzero in the *PSID* at least. The higher-order autocovariances in both data sets (not reported in Table 3) are generally small and mixed in sign. Row 12 of Table 3 contains the statistics for a test that the third- and higher-order autocovariances of earnings and hours changes are jointly equal to zero. This hypothesis is not rejected for any of the *NLS* samples or for the complete *PSID* sample. These samples are therefore consistent with a (nonstationary) bivariate second-order moving average ($MA(2)$) model of earnings and hours changes.[23] On the basis of the test statistics in row 12 there is some evidence of third- and higher-order serial covariation in the two subsamples of the *PSID*. These covariances are of trivial magnitude, however, and we choose to assume that they are zero in the interest of parameteric simplicity.[24]

For both complete samples and for all the subsamples except the one-employer subsample of the *NLS*, there is also strong evidence of nonstationarity in the covariances of earnings and hours changes. The goodness of fit statistics for a stationary model of the cross-covariances of earnings and hours (up to second order) are recorded in the last row of Table 3. Judging by these statistics, at least one of the variance components generating the changes in earnings and hours in the *PSID* and *NLS* surveys is nonstationary.

Table 3 also shows that the first-order autocorrelations of earnings and hours changes are negative and smaller than one-half in absolute value for both samples and all the subsamples.[25] Similarly, the ratios of the first-order cross-covariances of earnings and hours changes to their corresponding zero-order covariances are all negative and smaller than one-half in absolute value.[26] In the framework of our two-factor model, the fact that these autocorrelations are smaller than one-half in absolute value is evidence of a productivity component in earnings and hours. A pure measurement error model of the data implies that these autocorrelations are all exactly equal to $-\frac{1}{2}$.

To summarize the evidence in Table 3, the covariance structure of changes in earnings and hours is consistent with a second-order bivariate moving average model. Third- and higher-order autocovariances and cross-covariances are approximately zero in both the *PSID* and *NLS* surveys. In addition,

contains 55 unique elements. Estimates of these matrices (with standard errors) for the complete samples are contained in our other paper. The covariances reported in Table 3 represent simple averages of the covariances for each year of the *PSID* or *NLS* survey.

[23] By a nonstationary second-order moving average representation, we mean that $\mathrm{Cov}(\Delta \log g_{it}, \Delta \log g_{it-j}) = 0$, $\mathrm{Cov}(\Delta \log h_{it}, \Delta \log h_{it-j}) = 0$, $\mathrm{Cov}(\Delta \log g_{it}, \Delta \log h_{it+j}) = 0$, and $\mathrm{Cov}(\Delta \log g_{it}, \Delta \log h_{it-j}) = 0$, for all $j \geq 3$; and all other variances and covariances are unrestricted.

[24] One explanation for the higher-order serial correlation of the *PSID* data as compared to the *NLS* data is the fact that the *NLS* data are sampled biennially. If an $MA(2)$ model is appropriate for year-to-year (*PSID*) changes, for example, then biennial (*NLS*) changes follow an $MA(1)$ model.

[25] In the *PSID* sample, the first-order autocorrelations of earnings changes are −.35 (overall), −.42 (one-employer), and −.33 (multiple-employers). In the *NLS* sample, the first-order autocorrelations of earnings changes are −.27 (overall), −.39 (one-employer), and −.27 (multiple-employers). Similarly, the *PSID* first-order autocorrelations of hours changes are −.30 (overall), −.40 (one-employer), and −.29 (multiple-employers). The *NLS* first-order autocorrelations of hours changes are −.35 (overall), −.46 (one-employer), and −.35 (multiple-employers).

[26] In the *PSID* sample the ratios $\mathrm{Cov}(\Delta \log g_{it}, \Delta \log h_{it+1})/\mathrm{Cov}(\Delta \log g_{it}, \Delta \log h_{it})$ and $\mathrm{Cov}(\Delta \log g_{it}, \Delta \log h_{it-1})/\mathrm{Cov}(\Delta \log g_{it}, \Delta \log h_{it})$ are −.32 and −.27 (overall), −.46 and −.27 (one-employer), and −.31 and −.28 (multiple-employers), respectively. In the *NLS* sample these ratios are −.24 and −.16 (overall), −.25 and −.25 (one-employer), and −.26 and −.18 (multiple-employers), respectively.

both samples and all the subsamples exhibit (*i*) second-order serial correlation, (*ii*) covariance nonstationarity, and (*iii*) first-order autocorrelations of earnings and hours changes less than one-half in absolute value. Since any one of these three conditions is sufficient to identify the relative contribution of productivity changes to earnings as compared to hours in our two-factor variance components model, the parameter μ is empirically identified.

Table 1 describes the expected values of the variances, autocovariances, and cross-covariances of experience-adjusted earnings and hours changes in terms of the autocovariance structure of individual productivity and the covariance structure of preference variation and measurement error. Estimation of μ and tests of the goodness of fit of the statistical model described in Table 1 require that we parameterize the autocovariance structure of individual productivity changes. We use two different parameterizations. In the first case, we assume that Δz_{it} is a stationary second-order moving average.[27] In the second case, we assume that Δz_{it} is a nonstationary second-order moving average.[28] If individual productivity is stationary, the bivariate process for earnings and hours changes described in Table 1 is stationary. While we have strong evidence against a stationary covariance structure, the advantage of a stationary model is that the sufficient statistics for estimation of the structural parameters are just the average variances and covariances reported in Table 3. If individual productivity is a nonstationary second-order moving average, on the other hand, the sufficient statistics for estimation of the structural parameters are all the elements of the complete covariance matrix of earnings and hours changes up to second order. In both cases, we use a method of moments estimator based on minimizing the distance between the sample covariance matrix and the theoretical covariance matrix implied by Table 1 to estimate μ and the goodness of fit of the structural models.[29]

Table 4 reports the goodness of fit statistics and the associated estimates of μ for the various samples and subsamples of the *PSID* and *NLS* data.[30] Panel A contains the estimates for a stationary parameterization of individual productivity changes. The goodness of fit statistics are large, even in comparison to the goodness of fit statistics for an unrestricted stationary covariance model (reported in row 13, Table 3). The estimates of μ are all bigger than one, and are actually larger for the one-employer subsamples than for the multiple-employer subsamples or the overall samples. The estimates of μ for the one-employer samples are relatively imprecise, however, and one is within two standard errors of both estimates.

Panel B of Table 4 contains the estimates of μ and the goodness of fit statistics for a nonstationary parameterization of individual productivity changes. This model fits the data better in all cases, although the estimates of μ are not much affected. The nonstationary model actually provides an acceptable fit to the one-employer subsample of the *NLS*. For the other subsamples and the two complete samples, the two-factor model of the covariance structure of earnings and hours changes is rejected.

The point estimates of μ from the one-employer subsamples provide evidence against the contracting model of earnings and hours changes, and in favor of the intertemporal labor supply model. The associated estimates of the intertemporal substitution elasticity

[27] In the *PSID*, this results in the addition of three parameters for the productivity process. In the *NLS*, because of the irregular timing of the survey, this results in the addition of six parameters.

[28] In the *PSID*, this results in 27 parameters for the productivity process. In the *NLS*, this results in 12 parameters for the productivity process.

[29] See Gary Chamberlain (1984) for a discussion of the statistical theory of these estimators, and the comparison between these estimators and the maximum likelihood estimators. Our goodness of fit measures are derived in Whitney Newey (1985).

[30] Estimation of μ requires one arbitrary normalization of the variance parameters in Table 1. We set the correlation of Δu_{it} and Δv_{it} to 0. All statistics reported in Table 4, including the estimate of μ, are invariant to the choice of normalization.

TABLE 4—ESTIMATED RELATIVE CONTRIBUTION OF PRODUCTIVITY TO EARNINGS AND HOURS FOR *PSID* AND *NLS* SAMPLES AND SUBSAMPLES USING STATIONARY AND NONSTATIONARY SPECIFICATIONS

	PSID			*NLS*[b]		
Definition	All	One Employer	Multiple Employers	All	One Employer	Multiple Employers
A. Stationary Model						
1. Relative Contribution of Productivity to Variance of Change in Log Earnings (μ)	1.05 (.078)	1.54 (.439)	1.02 (.079)	1.56 (.174)	4.39 (2.19)	1.39 (.154)
2. Elasticity of Intertemporal Labor Supply (η)	19.96 (30.9)	1.85 (1.50)	51.81 (213.)	1.77 (.563)	.29 (.190)	2.53 (.980)
3. Goodness of Fit for Structural Model[a]	335.90 (.000)	229.64 (.000)	304.32 (.000)	145.54 (.000)	50.23 (.036)	144.72 (.000)
B. Nonstationary Model						
4. Relative Contribution of Productivity to Variance of Change in Log Earnings (μ)	1.14 (.091)	2.46 (.781)	1.16 (.089)	1.23 (.100)	4.10 (2.15)	1.14 (.080)
5. Elasticity of Intertemporal Labor Supply (η)	7.27 (4.81)	.68 (.365)	6.40 (3.64)	4.34 (1.89)	.32 (.220)	7.14 (4.08)
6. Goodness of Fit for Structural Model[b]	261.15 (.000)	142.13 (.000)	223.01 (.000)	127.02 (.000)	33.84 (.206)	97.80 (.000)

Note: The standard errors are shown in parentheses.

[a]χ^2 statistic for stationary structural model vs. nonstationary bivariate *MA*(2) model. The statistic has 92 degrees of freedom for the *PSID* sample and 34 degrees of freedom for the *NLS* sample. Probability values are shown in parentheses.

[b]χ^2 statistic for nonstationary model vs. nonstationary bivariate *MA*(2) model. The statistic has 68 degrees of freedom for the *PSID* sample and 28 degrees of freedom for the *NLS* sample. Probability values are shown in parentheses.

(η) are recorded in rows 2 and 5 of Table 4.[31] In the *PSID* one-employer subsample, the estimates of η from the stationary and nonstationary models are 1.84 and .68, respectively. These estimates are larger than the instrumental variables estimates reported by Joseph Altonji (1986) and Thomas MaCurdy (1981) for *PSID* males, although they are based on a very different methodology. In the *NLS* one-employer sample, the estimates of η are .29 for the stationary model and .32 for the nonstationary model. These estimates are comparable to other estimates based on individual longitudinal data.

While the results from the one-employer subsamples are relatively favorable to the labor supply interpretation of earnings and hours changes, the results from the multiple-employer subsamples and the overall samples reveal a major difficulty with this interpretation. In the contracting model, changes in employer represent changes between contracts. The contract model therefore offers a simple explanation for the greater variability of earnings and hours for those who change employers than those who do not. The labor supply model, on the other hand, predicts the same structure of earnings and hours changes within and across jobs. The labor supply model by itself does not explain the higher variation in earnings and hours changes for those who change jobs. The labor supply model also predicts the same relative effect of productivity changes on earnings and hours for job changers and stayers. The point estimates of μ for the multiple-employer subsamples, however, are very different from the estimates based on the one-employer subsamples. In both the *PSID* and *NLS* multiple-employer subsamples, μ is precisely estimated and close to, but greater than, one. The implied estimates of the inter-

[31]The estimates of η are obtained from the formula $\eta = 1/(\mu - 1)$. If μ is near one, η will be imprecisely estimated and the point estimate of η will fluctuate substantially with relatively small changes in the point estimate of μ.

temporal substitution elasticity are large and imprecise.

The estimates of μ for individuals who change employers suggest that productivity changes affect earnings and hours proportionately. In other words, for these individuals, hours vary at fixed wage rates. One potential explanation for this finding in the framework of a labor supply model is that individuals cannot fully insure productivity risks. In this case, our estimation strategy confounds changes in productivity and changes in the marginal utility of wealth. Since changes in the marginal utility of wealth influence earnings and hours proportionately in the labor supply model, μ is biased towards one if the component of variance that we attribute to productivity changes includes changes in the marginal utility of wealth. In our other work, however, we find that estimates of μ are unaffected by controlling for changes in the marginal utility of wealth. The evidence that changes in earnings and hours occur at constant wage rates is inconsistent with either labor supply models or the contracting models considered in this paper. Fixed wage contract models have been considered by Abowd and Orley Ashenfelter (1981), and applied in the macroeconomics literature by Stanley Fischer (1977) and John Taylor (1980), among others. Our results for the job changers suggest that these models may be useful in the empirical analysis of individual data as well.

Finally, Table 4 also reports parameter estimates for the complete *PSID* and *NLS* samples. It is clear from these estimates that the characteristics of the multiple-employer subsamples carry over to the complete samples. In the complete samples, changes in productivity have slightly larger effects on earnings than hours, although we cannot easily reject the hypothesis that productivity-induced changes in hours occur at fixed wage rates (i.e., $\mu = 1$).

IV. Conclusion

Our goal in this paper was to develop an empirical strategy for testing between contracting and labor supply models. Such a test must rely on the fundamental distinction between these models: in contracting models, earnings represent optimal consumption, whereas in labor supply models, earnings represent the product of wage rates and hours of work. We derive a testable contrast between the two models based on the relative variability of changes in earnings and changes in hours. If the contracting model is correct, earnings are less variable than hours with respect to changes in productivity. If the labor supply model is correct, the reverse is true.

In order to apply the test, we specify a complete model of earnings and hours variation, including productivity components and components due to changes in tastes and measurement errors. This statistical model is itself testable, providing a check on the ability of either theory to explain the covariance properties of earnings and hours changes in longitudinal data.

We apply the model to longitudinal data from the *PSID* and *NLS* surveys. Generally speaking, the data are inconsistent with the simple covariance structure implied by either the labor supply or contracting model. Contrary to the implications of the contracting model, the contribution of productivity shocks to earnings is at least as large as the contribution to hours. This is true for individuals with the same employer over the entire period of the *PSID* and *NLS* surveys and more generally. From the point of view of the labor supply model, however, the implied intertemporal substitution elasticities are large and imprecise. A simpler interpretation of the data is that productivity-related changes in hours occur at fixed wage rates. We conclude that the specification and testing of fixed wage models for individual earnings and hours data should be a high priority for future research.

DATA APPENDIX

For the *Panel Study of Income Dynamics* we used an extract from the thirteen-year merged individual tape distributed through the Inter-University Consortium for Political and Social Research and documented by the Survey Research Center of the Institute for Social Research (1981 and previous years). Our sample consisted of all males on the thirteen-year merged individual tape with complete age and schooling data who were continuously heads of household from wave III to wave

XIII of the survey and who reported nonzero annual labor earnings and annual hours in each of the 11 waves. We included individuals from both the *Survey of Economic Opportunity* subsample and the Survey Research Center national probability subsample.

The following is a description of the *PSID* variables used. Numbers like Vxxxx refer to the variable numbers in the Survey Research Center codebooks for the thirteen-year merged individual tape. Survey questions are referenced by the question number and the exact question from the questionnaire.

ANNUAL EARNINGS: The variables used were: V1196, V1897, V2498, V3051, V3463, V3863, V5031, V5627, V6174, V6767, and V7413. These correspond to Survey Research Center's computed values for the head of household's total labor income in the calendar year before the survey. Annual earnings are computed from questions that changed somewhat from year to year. For 1979 (wave XIII) earnings are based on the sum of the answers to the following survey questions:

(K8) How much did you (HEAD) receive from wage and salaries in 1979, that is, before anything was deducted for taxes and other things?
(K9) In addition to this did you (HEAD) have any income from bonuses, overtime, or commissions?
(K10) How much?
(K11) Did you (HEAD) receive any other income in 1979 from professional practice or trade?
(K12a) How much from professional practice?
(K12b) How much from farming or market-gardening?
(K12c) How much from roomers or boarders?

Farmers and others with business income also answer a battery of questions on net farm income and total business income. Only the labor part of farm, business, and roomer income is added to variable V7413. The determination of the labor part of these variables is part of the coding process at the Survey Research Center.

ANNUAL HOURS: The variables used were V1138, V1839, V2439, V3027, V3423, V3823, V4332, V5232, V5731, V6336, and V6934. These correspond to Survey Research Center's computed value for the head of household's annual hours worked in the calendar year before the survey. The actual survey questions on which this variable was based changed from year to year. For 1979 (wave XIII) the questions were:

(C26) How many weeks did you actually work on your main job in 1979?
(C27) And, on average, how many hours a week did you work on your main job in 1979?
(C28) Did you work any overtime which isn't included in that?
(C29) How many hours did that overtime amount to in 1979?
(C42) How many weeks did you work on your extra jobs in 1979?
(C43) On the average, how many hours a week did you work on your extra jobs?

CHANGE OF EMPLOYER: Individuals were considered to have changed employer between the past calendar year and the survey year if they either (a) reported being unemployed at the time of the survey (approximately March of the survey years), or (b) reported being employed or temporarily laid off at the time of the survey but reported a change in jobs not associated with a promotion. Current employment status was measured from variables V1278, V1983, V2581, V3114, V3528, V3967, V4458, V5373, V5872, V6492, and V7095. This variable is the filter question that determines whether the head of household is asked the battery of questions about current employment (if he is employed or temporarily laid off) or unemployment (if he is unemployed). For individuals who were employed or temporarily laid off at the time of the survey, a change of employer was measured by variables V1282, V1988, V2586, V3119, V3534, V3986, V4490, V5399, V5890, V6501, and V7104. This variable is based on the answer to the question about what happened to the head of household's previous job. In wave XIII this question was: "(C12) What happened to the job you had before—did the company go out to business, were you laid off, promoted, were you not working, or what?". Promotions were not counted as changes of employer. All other answers were counted as a change of employer.

SEX: The variables used were V7492, V7509, V7526, V7547, V7561, V7576, V7601, V7653, V7687, and V7714. Only males were included.

HEAD OF HOUSEHOLD: The variables used were V7490, V7507, V7524, V7545, V7559, V7574, V7599, V7624, V7651, V7685, and V7712. Only heads of household for all eleven waves from wave III to wave XIII were included.

SCHOOLING: The variables used were V0313, V0794, V1485, V2197, V2838, V3241, V3663, V4198, V5074, V5647, V6194, V6787, and V7433. Maximum completed schooling was determined by examining all schooling variables. This was the schooling level used.

AGE: The variables used were V7460, V7476, and V7491. Age was determined for the first year an individual entered the sample, then adjusted to age in 1969.

For the *National Longitudinal Survey of Men 45–59, 1966 to 1975*, we used an extract from the public use data file release 75A distributed through the Inter-University Consortium for Political and Social Research and documented by the Center for Human Resource Research (1977, 1980). We used data from survey years 1966, 1967, 1969, 1971, 1973, and 1975. Annual data from the survey year 1966 refer to calendar year 1965. Subsequent annual data refer to the twelve months preceding the actual interview—approximately June 1966 to May 1967 for the 1967 survey and approximately July of the previous year to June of the survey year for the subsequent surveys. Our sample consisted of all males who had valid age and schooling data and reported nonzero annual earnings and annual hours for the years we studied.

The following is a description of the *NLS* variables used. Numbers like Vxxxx refer to the Center for Human Resource Research codebook variables numbers (not the reference numbers) for the release 75A public use tape. (Some variables are assigned two consecutive variable numbers.) Survey questions are referenced by the question number but only the facimile question in the public use codebooks is reproduced.

ANNUAL EARNINGS: The variables used were V0263-4, V0784-5, V1280-1, V3166-7, V2528-9, V2685-6. For the first two survey years, these variables represent the answer to the question: "(63A) What was your income from wages and salary in 1965?" (Example from 1966 survey). In the subsequent years these variables represent the answer to the question: "(16) What was your income from wages and salary in the past year?" (Example from 1975 survey).

ANNUAL HOURS: The variables used were hours per week: V0082, V0660, V1128, V1581, V2520, V2675, and weeks per year: V0589, V1022, V1168, (V2421 with V2461), V2519, V2674. There is substantial survey-to-survey variation in the questions used to measure these hours concepts. In 1966 the hours per week question was: "(11B) What were the usual number of hours per week worked in 1965?". In 1967 the question was: "(7B) What is the number of hours worked at your current or last job?". In 1971 the question referred to the current job only. In 1973 and 1975 the question was "(12B) What is the number of hours per week usually worked during the weeks worked in the past year?" (Example from 1975 survey). The weeks worked per year variable is a Center for Human Resource Research recode of the raw data for the survey years 1966 to 1971. For the 1971 survey we recoded the weeks worked variable, which refers to weeks worked since the last interview, into weeks worked in the last year by dividing the number of weeks worked since the last interview (V2421) by the number of weeks since the last interview (V2461) and multiplying by 52. In 1973 and 1975 the variable refers to the question: "(12A) What is the number of weeks worked in the past year?" (Example from 1975 survey).

EMPLOYER CHANGES: The variables used were V2406, V2548, and V2708. The first of these is reported "tenure at current job" in the 1971 survey, which is a recode of the question: "(6H) What is the year you started working at your current job?". The last two variables are a recode of the answer to the question: "(Check Item C) Is the date you started working at your current job September 1, 1971 or later?" (Example from 1973 survey). Individuals with 1971 tenure greater than five years and no reported change of employer in the 1973 and 1975 surveys were treated as having the same employer for all years.

AGE: The variable used was V0024, age in 1966.

EDUCATION: The variable used was V0611, highest grade completed.

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