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SUBSTITUTION OVER TIME: ANOTHER
LOOK AT LIFE CYCLE LABOR SUPPLY

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ABSTRACT

Most studies of the intertemporal substitution of work use life cycle data and, from those studies, many have concluded that intertemporal labor substitution is unimportant for macroeconomics. This paper takes another look at life cycle data and argues that a consideration of measurement errors, taxes, on-the-job training, older workers, hours reporting bias, and all of the “margins” composing aggregate labor supply over the life cycle suggests that substitution over time may be very important for macro fluctuations. The life cycle data used includes fairly standard male cross-section and panel data samples as well as a sample of women experiencing the termination of AFDC benefits as their youngest child turns 18 years old.

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I. Introduction

Today macroeconomics relies almost exclusively on intertemporal models, models which are also quite common in other areas of economic inquiry. One important question that must be answered in the building of a dynamic model is “What are the possibilities for substitution over time?” This paper is about one potential margin of substitution, the substitution of work over time.

An analysis of the substitution of work over time dates at least back to Hicks [1939, chapters XVI and XVII] and since Lucas and Rapping [1969] this margin of substitution has been included in a substantial number of macro models. The presence of this margin affects predictions for labor market and output fluctuations in response to aggregate productivity shocks (Hall [1980], Hansen [1985]), to changes in the timing of taxes (Judd [1987]), to changes in the timing of government spending (Aiyagari et al [1992], Barro [1981, 1987], Baxter and King [1993]), to monetary shocks (Lucas [1972]), and many other questions of substantial interest. But since a series of studies by Altonji [1986], MaCurdy [1981], Abowd and Card [1989], and analyses of tax experiments (some of which are surveyed by Hausman [1985] and Pencavel [1986]), it has been widely believed that the substitution of work over time is a quantitatively unimportant margin. This belief and its origins can be seen in Pencavel's [1986] and Card's [1994] surveys, the comments by Ashenfelter [1984], Mankiw [1989], and Plosser [1989], and in the calibration of macro models by Auerbach and Kotlikoff [1987, p. 50], Judd [1987], and many others exploring the consequences of some interesting and important policy changes.

This paper revisits the “life cycle” evidence - evidence on low frequency changes in wages and time worked with age. My goal here is *not* to reconcile the variety of life cycle and non-life-cycle studies that are available (I do attempt this in another paper, Mulligan [1995]). Nor is my goal to argue that the substitution of work over time is the single missing piece for every macroeconomic puzzle. Instead, I revisit *life cycle* evidence to show that it is consistent with quantitatively important substitution of labor over time. Although I do not think that life cycle evidence is necessarily the best evidence for getting at this issue, it is true that the widely cited PSID and other micro data studies are basically life cycle studies. I then show that the usual cohort data might be reasonably modified in several ways: including the aggregation of micro measures of labor supply, accounting for time spent training on the job, using

different techniques for measuring work time, measuring marginal tax rates, and including older workers.

Although the usual cohort data is consistent with a lot of substitution over time, there are some problems with its construction. There is no obvious way to infer the wages of those who are not working at various points in the life cycle. Nor is it particularly plausible that tastes, health, and other relevant variables are independent of age or even uncorrelated with wages. Nor is it particularly obvious exactly what procedure should be used to obtain the “right” estimate from the usual cohort samples. These problems are enough to eagerly search elsewhere for anticipated wage changes and associated changes in work. My search led me to another life cycle event that has not yet been studied: the termination of AFDC at the 18th birthday of a family's youngest child.² Section IV explains how this study can alleviate some of the apparent problems with a synthetic cohort sample, that it is consistent with substantial substitution over time, and - perhaps surprisingly - that estimates are similar to the synthetic cohort estimates.

II. A Life Cycle Model of Labor Supply

II.A. An Overview of the Intertemporal Substitution Hypothesis

The intertemporal substitution hypothesis (ISH) can be simply stated: workers intertemporally reallocate their work in response to changes in the relevant relative price. A statistical model that captures this idea is:

$$E_t \left[\ln \frac{n_{t+1}^i}{n_t^i} \right] = -\sigma r_t + \sigma E_t \left[\ln \frac{w_{t+1}^i}{w_t^i} \right] + \varepsilon_t^i \quad (1)$$

where n_t^i is the labor supply of worker i at date t , w_t^i is worker i 's date t after-tax market value of time, and r_t is the real rate of return to savings between periods t and $t+1$. ε_t^i includes preference parameters and, in some models, a precautionary motive for working. $\sigma > 0$ is the intertemporal elasticity of substitution (IES). In many environments, labor supply decisions are made sequentially in time rather than at the “beginning of time” so the law of demand more appropriately applies to “plans” for labor supply as a function of anticipated wages. E_t denotes date t expectations of future variables and is therefore included in equation (1).

Notice that the model does not necessarily predict that workers will work the hardest when the

²I have also been led to look at the anticipated wage changes associated with non-life cycle events including seasons, wars, and agricultural shocks (Mulligan [1995]).

conventionally measured "real wage" is the highest. There are two relevant forces at work. First, the relative price of leisure in any two periods depends not only on the ratio of the wages but also on an interest rate. Second, through the taste term ε_t , the model allows for impatience, prudence or other reasons that workers may have different preferences for current and future leisure. Consider, for example, a perfectly flat time profile for wages. A worker will choose more leisure in the later periods if the interest rate is high enough or more leisure in the earlier periods if he is impatient enough. Without data on the interest rate and preferences, it is necessary to estimate the trend component of the intertemporal labor allocation together with the responsiveness of work effort to incentives.

Beneath any notational complexities, the econometrics of estimating the elasticity σ is also quite straight-forward: identify situations with different anticipated rates of wage growth and measure the associated differences in anticipated labor supply growth. The implementation of this can be quite challenging (eg., how is "anticipated wage growth" measured?), but the thought experiment is simple enough. I (and many others in the empirical literature) do not intend to say or do anything more complicated than this, but a more complicated model is needed to be clear about the measurement of the variables of interest and interpretation of the parameter σ . My life cycle model includes several complications necessary for understanding the data: (a) uncertainty about future wages, (b) various measurement errors in hours and wages, (c) discrete choices about labor force "participation" at various points in time, (d) time aggregation in the measurement of labor supply, and (e) potentially nonlinear labor income tax schedules.

II.B. Indivisible Labor

A person's lifetime includes many potential "work sessions," equally spaced in time. For our purposes, it may be useful to think of a potential work session as a day or week, although a month is in some ways a more convenient definition for the AFDC application in section IV.C. If work is to occur during a potential session, it must occur for exactly \bar{n} units of time. This "indivisibility" of labor might be interpreted as the optimal bunching of labor in continuous time in the presence of fixed "commuting" costs and high frequency fatigue effects (see Mulligan [1998]), in which case \bar{n} is a function of the magnitude of the fixed cost and of the form of the fatigue effects, but not the other variables in the model such as wages, taxes, or tastes for leisure. w denotes the average product of labor for the session, so that $w\bar{n}$ is the total amount produced by a worker who chooses to work the session. w and \bar{n} may vary over time and across states of nature.

II.C. Human Capital Accumulation

Two things can happen during a work session: goods production or the production of one's own human capital. When goods production occurs, the worker is paid $w\bar{n}$. When human capital accumulation occurs, the worker may or may not be paid depending on whether the firm is financing the training or not. If the firm does not finance the training, the worker does not have any earnings for that session but still values the training for the future earnings it produces. In this case, I let $w\bar{n}$ denote that valuation.³

In Section III, I use two measures of time spent "producing human capital." The first is time spent searching for a job. The second is time at work spent learning new things.

II.D. Time Aggregation

It will be assumed that the econometrician observes only time aggregated measures of the relevant variables, so it is convenient to use two indices t and k to identify a potential work session. t indexes the time aggregated periods or "time intervals" and varies from 0 to T during a consumer's lifetime. k indexes the potential "sessions" within any particular time interval and varies from 1 to K for each t . Thus there is a total of $K(T+1)$ potential sessions in a consumer's lifetime.

II.E. Tax and Benefit System

During any potential session, a consumer receives government transfer payments in the amount $b_{t,k}$ (the consumer is a net taxpayer in the case $b < 0$). Benefits are determined according to the formula:

$$b_{t,k} = \max\left\{\underline{b}_{t,k}, \bar{b}_{t,k} - R_{t,k} \max[w_{t,k}n_{t,k}I_{t,k} - d_{t,k}, 0]\right\} \quad (2)$$

where $w_{t,k}$ is the average product of time during the k th session of time interval t (time which may be spent either in goods or human capital production), $n_{t,k}$ is time worked, $I_{t,k}$ is an indicator variable for goods production or firm-financed human capital production, $I_{t,k}w_{t,k}n_{t,k}$ is labor earnings, $d_{t,k}$ are deductions from earnings and "earnings disregards", $R_{t,k}$ is the "benefit reduction rate" or "marginal labor income tax rate," $\bar{b}_{t,k}$ is the maximum benefit available (and in some applications may be a function of family composition, asset holdings, and other variables), and $\underline{b}_{t,k}$ is the minimum benefit available. Benefits are reduced $R_{t,k}$ dollars for every dollar of net earnings, where net earnings are computed as gross earnings net of deductions.

³Human capital accumulation can be valued more than $w\bar{n}$ without changing the interpretation of my results, although see my discussion of time aggregation.

Notice that earnings and deductions are not aggregated across sessions for the purposes of computing taxes and benefits. I note in the text where this assumption may be unrealistic and of some consequence for the results.

II.F. Uncertainty, Budget Constraints, and Utility Functions

I index a realization of a consumer's life history by ω . For each possible realization $\omega \in \Omega$, consumer choices of stochastic processes for consumption $c_{t,k}(\omega)$, work $n_{t,k}(\omega)$, and tax deductions $d_{t,k}(\omega)$ must satisfy a present value budget constraint:

$$\sum_{t=0}^T \sum_{k=1}^K e^{-\rho(tK+k)} Q_{t,k}(\omega) [c_{t,k}(\omega) - w_{t,k}(\omega)n_{t,k}(\omega) - b_{t,k}(\omega) + f_{t,k}(d_{t,k}(\omega), \omega) - H_{t,k}(\omega)w_{t,k}(\omega)\bar{n}_{t,k}(\omega)] = A_0 \quad (3)$$

$$\ln Q_{t,k}(\omega) \equiv - \sum_{s=0}^{t-1} \sum_{l=1}^K [r_{s,l}(\omega) - \rho] - \sum_{l=1}^k [r_{t,l}(\omega) - \rho]$$

where $r_{t,k}$ is the ex poste real return on a one period bond purchased at date $t,k-1$. $H_{t,k}$ is an indicator variable for "purchases" of self-financed human capital. In addition to (3), we have the constraints (2) and $n_{t,k}(\omega) \in \{0, \bar{n}_{t,k}(\omega)\}$ for all $t=0, \dots, T, k=1, \dots, K$, and $\omega \in \Omega$.

Deductions are costly. Deductions in the amount d cost $f_{t,k}(d, \omega)$. For each t,k,ω , the deduction cost function is nondecreasing and nonconcave in the amount deducted. There is no cost if no deductions are taken.

The revelation of information over time is modeled with the filtered probability space $(\Omega, \mathcal{F}, F, \pi)$. Each state of nature ω has unconditional probability $\pi(\omega)$. The filtration $F = \{\mathcal{F}_{0,1}, \mathcal{F}_{0,2}, \dots, \mathcal{F}_{T,K}\}$ on Ω is assumed to be increasing and the stochastic processes $w_{t,k}, \bar{n}_{t,k}, \gamma_{t,k}, r_{t,k}, \bar{b}_{t,k}, R_{t,k}$ and $\underline{b}_{t,k}$ are adapted to it.⁴

Consumption and work are assumed to evolve as if workers choose functions $c_{t,k}(\omega)$, $n_{t,k}(\omega)$, and $d_{t,k}(\omega)$ adapted to the filtration F with the objective of maximizing the expected value of an intertemporally and intratemporally separable utility function (4) subject to the constraints (2), (3) and $n_{t,k}(\omega) \in \{0, \bar{n}_{t,k}(\omega)\}$:

$$\sum_{\omega \in \Omega} \pi(\omega) \sum_{t=0}^T \sum_{k=1}^K e^{-\rho(tK+k)} [u(c_{t,k}(\omega)) - \gamma_{t,k}(\omega)v(n_{t,k}(\omega))] \quad (4)$$

$$\rho > 0, \quad v(0) \equiv 0, \quad v_{t,k}(\omega) \equiv v(\bar{n}_{t,k}(\omega)) > 0$$

Notice that I do not treat $H_{t,k}$ and $I_{t,k}$ as choice variables and do not say exactly how sessions devoted to

⁴The random function $f_{t,k}$ should be measurable with respect to $\mathcal{F}_{t,k}$.

human capital production translate into higher future wages, but the first order conditions of the problem described above are among those of the larger problem that include optimal human capital accumulation over the life cycle.⁵

II.G. When to Work

When describing the decision to work, it is useful to define $\tau_{t,k}(\omega)$ as the implicit tax rate on work in session k of time interval t :

$$\tau_{t,k}(\omega) \equiv \min \left\{ \frac{b_{t,k}(\omega) - \underline{b}_{t,k}(\omega)}{w_{t,k}(\omega) \bar{n}_{t,k}(\omega)}, R_{t,k}(\omega) - \frac{R_{t,k}(\omega) d_{t,k}^*(\omega) - f_{t,k}(d_{t,k}^*(\omega), \omega)}{w_{t,k}(\omega) \bar{n}_{t,k}(\omega)} \right\} \leq R_{t,k}(\omega) \quad (5)$$

$$d_{t,k}^*(\omega) \equiv \underset{d \in [0, w_{t,k}(\omega) \bar{n}_{t,k}(\omega)]}{\operatorname{argmin}} \left\{ f_{t,k}(d, \omega) - R_{t,k}(\omega) d \right\}$$

If $\bar{b}_{t,k}$ is large enough (in what follows, I suppress the ω index), the implicit tax rate is related to the “benefit reduction rate” $R_{t,k}$. However, when positive deductions are optimal and the deduction cost function is strictly concave, the indivisibility of labor means that the implicit tax rate is strictly less than the benefit reduction rate. If $\bar{b}_{t,k}$ is small enough, $\tau_{t,k}$ is unrelated to $R_{t,k}$.

The first order condition determining where $n_{t,k}$ equals 0 or $\bar{n}_{t,k}$ is:

$$\gamma_{t,k} \nu_{t,k} \begin{matrix} > \\ < \end{matrix} Q_{t,k} w_{t,k} \bar{n}_{t,k} (1 - \tau_{t,k}) E_{t,k} \lambda_{t,k} \quad (6)$$

where $E_{t,k}$ denotes date t,k expectations and $\lambda_{t,k}$ is the lifetime “marginal” utility of wealth.⁶ (6) simply says that people work during those periods when the benefit exceeds the cost. The benefit of working is discounted after-tax earnings (or the value of self-financed human capital production) for session k of interval t $Q_{t,k} w_{t,k} \bar{n}_{t,k} (1 - \tau_{t,k})$ times the expected “marginal” utility of wealth $E_{t,k} \lambda_{t,k}$. The cost is the disutility of work $\gamma_{t,k} \nu_{t,k}$. Comparative static changes in $Q_{t,k} w_{t,k} (1 - \tau_{t,k})$ can be interpreted as generating a “substitution

⁵See Ghez and Becker [1975]. The larger problem with $H_{t,k}$ and $I_{t,k}$ as choice variables has the constraint $H_{t,k} \bar{n}_{t,k} = I_{t,k} n_{t,k}$ for all t,k,ω .

⁶Because of its discreteness, the labor supply decision at any date and state has a discrete effect on expected remaining lifetime discounted utility of consumption and $\lambda_{t,k}$ is the size of this effect per dollar of date t,k after-tax earnings $w_{t,k} \bar{n}_{t,k} (1 - \tau_{t,k})$. $\lambda_{t,k}$ approaches “marginal” utility as $w_{t,k} \bar{n}_{t,k} (1 - \tau_{t,k})$ goes to zero. Mulligan [1998] also shows that $\lambda_{t,k}$ is literally a marginal utility when particular lotteries and “taste insurance” contracts are introduced as choice variables.

effect” while changes in $E_{t,k}\lambda_{t,k}$ can be interpreted as generating a “wealth effect.”

II.H. Consumption Insurance?

The problem described above presumes that insurance against surprises to wages, taxes, and interest rates is unavailable.⁷ It is straight-forward to allow for perfect consumption insurance by collapsing the series of budget constraints (3) into a single budget constraint that equates the expected present value of expenditures to the expected present value of resources. In this case, the lifetime “marginal” utility of wealth $\lambda_{t,k}$ is no longer a random variable with respect to $\mathcal{F}_{t,k}$ ($E_{t,k}\lambda_{t,k} = \lambda_{t,k}$). Otherwise the condition for working session k of interval t is the same as (6).

II.I. Information Lost from Time Aggregation

Given time aggregated lifetime data on all of the exogenous variables of the model, we cannot make sharp predictions about time aggregated labor supply without answering three questions.

- (1) How is information about future wages and interest rates revealed during a time interval?
- (2) What is a worker's market value of time at each potential session, including those sessions he did not work?
- (3) What is a worker's nonmarket value of time at each potential session?

I make three assumptions about this:

- (A1) No information is revealed within time intervals: $\mathcal{F}_{t,k} = \mathcal{F}_{t,l}$ all $k = 1 \dots K$
- (A2) The market value of time and the indivisibility of labor does not vary within time intervals: $Q_{t,k}w_{t,k}\bar{n}_{t,k}(1-\tau_{t,k}) = Q_{t,l}w_{t,l}\bar{n}_{t,l}(1-\tau_{t,l})$ and $\bar{n}_{t,k} = \bar{n}_{t,l}$ all $k = 1 \dots K$
- (A3) $\gamma_{t,k} = \gamma g_{t,k}$ where $\sigma \ln(g_{t,k})$ $k = 1 \dots K$ are drawn according to the distribution function G with unbounded support. $\sigma > 0$ is a constant. $E[\sigma \ln(g_{t,k})] = 0$ and $E[(\sigma \ln(g_{t,k}))^2] = 1$. Without loss of generality, the K draws are assumed to be independent.

Assumptions (A1) and (A2), although often only implicit, are extremely common in the labor supply literature. (A3) produces “increasing marginal disutility of work” within a time interval, an assumption which appears in one form or another in the literature. To see mechanically how (A3) produces increasing marginal disutility, notice that, given (A1) and (A2), workers that work at all during a time interval will

⁷Although one of Mulligan's [1998] interpretations of equation (6) presumes that “tastes insurance” or some other mechanism is available to compensate those who are unlucky enough to especially dislike work when wages are high.

work during the low disutility sessions. Since sessions differ in their marginal disutility, workers require a higher wage to work a larger fraction of the time interval.⁸ I use (A1)-(A3) to derive some empirical specifications, but comment further on their relevance as I discuss the empirical results.

In accordance with assumption (A2), I suppress the k subscript except where necessary for clarity.

II.J. Consumer Heterogeneity and the Individual Labor Supply Equation

Consumers may differ in terms of the realizations of the exogenous variables $w_{t,k}$ and $\gamma_{t,k}$. They are the same in terms of the stochastic processes generating these realizations, the utility functions $u(c)$ and $v(n)$, the parameter σ , and the way in which information arrives (including the function G defined above).

Consider a group of consumers who are identical in terms of the two random variables $\gamma_t v_t$ and $Q_t w_t \bar{n}_t (1 - \tau_t) E_t[\lambda_t]$. We compute the group average N_t of each individual's time worked during the time interval t , including those who did not work at all during the interval:

$$G^{-1}\left(\frac{N_t}{K \bar{n}_t}\right) = \sigma \ln w_t (1 - \tau_t) + \sigma \ln \bar{n}_t + \sigma \ln \frac{Q_t E_t[\lambda_t]}{\gamma_t v_t} \quad (7)$$

This is a nonlinear version of MaCurdy's [1981] " λ -constant labor supply function," which may be surprising given MaCurdy's apparently different description of the labor supply problem. In fact, the only difference is in the measurement of labor supply.

For a group of consumers who, for some integer $s > 0$, have the four random variables $\gamma_t v_t$, $Q_t w_t \bar{n}_t (1 - \tau_t) E_t[\lambda_t]$, $E_t[\gamma_{t+s} v_{t+s}]$, $E_t[Q_{t+s} w_{t+s} \bar{n}_{t+s} (1 - \tau_{t+s}) \lambda_{t+s}]$ in common, we have a version of equation (1):

$$\begin{aligned} E_t \left[G^{-1}\left(\frac{N_{t+s}}{K \bar{n}_{t+s}}\right) - G^{-1}\left(\frac{N_t}{K \bar{n}_t}\right) \right] &= \sigma K \sum_{s'=1}^s (E_t r_{t+s'} - \rho) + \sigma E_t \ln \frac{w_{t+s} (1 - \tau_{t+s})}{w_t (1 - \tau_t)} + \sigma E_t \ln \frac{\bar{n}_{t+s} \gamma_{t+s} v_{t+s}}{\bar{n}_t \gamma_t v_t} \\ &+ \sigma [E_t \ln E_{t+s} \lambda_{t+s} - \ln E_t E_{t+s} \lambda_{t+s}] \end{aligned} \quad (8)$$

The only difference between an individual's labor supply plans and those of the group is the "sampling" error due to the fact that a single individual only samples K times from the distribution G during a time interval. The probability that a particular individual does not work at all during a particular time

⁸Lucas (1970, p. 25) has exactly the specification (A3), except that he has a continuum of potential work sessions per time interval (as compared to my K).

interval is:

$$\left[1 - G \left(\sigma \ln \frac{w_t R_t Q_t (1 - \tau_t) E_t \lambda_t}{y_t v_t} \right) \right]^K$$

Thus extended intervals of participation and nonparticipation during a person's lifetime are evidence of either (a) large and persistent changes in tastes or the value of time or (b) a large "willingness" to substitute over time σ .

The responsiveness of labor supply to wages depends on the level of labor supply in the model (8). Typically there will be little response of N when N is near 0 or $K\bar{n}$. This is even true in the special case that $g_{t,k}^0$ is distributed uniformly on $[0,e]$ and the equation (1) obtains - a special case which gets a lot of attention in the literature - once the possibilities of corner solutions for N_t and N_{t+s} are taken into account.⁹ The bulk of my analysis follows Altonji [1986], Ghez and Becker [1975], MaCurdy [1981], and others by focusing on this special case and ignoring possible corner solutions. My section IV returns to the more general model (8) and includes some analysis of how the responsiveness of labor supply might vary with its level.

The derivation of equation (8) reveals several points that are quite relevant for the empirical applications in this paper:

- (i) the market value of time w_t for time interval t can, for those who have some earnings during the interval, be measured as $y/[N(1-h_t)]$, where h_t is the fraction of sessions worked that were devoted to self-financed human capital production, and y_t is total pre-tax labor earnings for the time interval.
- (ii) σ measures the responsiveness of *anticipated* labor supply changes to anticipated wage changes, not *ex poste* labor changes to *ex poste* wage changes.
- (iii) the implicit tax on work has the "benefit reduction rate" or "marginal labor income tax rate" $R_{t,k}$ as its upper bound
- (iv) how time aggregation is related to the measurement of labor supply, wages, tax rates, and other variables
- (v) how measured "employment" and "hours" are related
- (vi) those who are not working during any particular work session are not a random sample of the population, but those for whom work yields a greater disutility

⁹See Smith [1977, p. 249] for a discussion and Rogerson and Rupert [1991] for estimates of a labor supply model with corners at "year-round" work.

- (vii) the effect of consumption insurance on comovements of wages and labor supply

II.K. Macro "Experiments" to be Calibrated from Life Cycle Data

Before reviewing empirical studies and proposing ways to improve them, something must be said about why life cycle substitution is of interest for macroeconomics. One item of substantial interest is the response of *aggregate* labor supply to aggregate temporary "shocks" to the market value of time. Candidate shocks to the market value of time include productivity shocks such as those in Kydland and Prescott [1982], monetary shocks such as those in Lucas [1972], temporary government spending shocks such as those modeled by Hall [1980], Barro [1981], Baxter and King [1993], an income tax cut that is "phased in" over time, or a tax or subsidy on savings. Computing the response to these shocks probably requires a general equilibrium model of which my model (8) would be one piece, but the parameter σ in my model is the most crucial ingredient. When we, for example, compare two equilibria (which differ, say, according to the processes generating policy or productivity shocks) which have different anticipated rates of growth of the after-tax discounted market value of time, σ will - up to an aggregation bias term - measure the cross-equilibria difference in anticipated rates of aggregate labor supply growth. This paper uses various life cycle data to estimate σ . Whether the world actually exhibits the temporary wage fluctuations predicted by these models is an interesting empirical question but one beyond the scope of this paper.

Although very high frequency nonseparabilities can be used to motivate my indivisible labor model (see Mulligan [1998]), labor supply should be separable over time at low frequencies in order for the degree of intertemporal substitution found in life cycle data to be the same as that applicable to higher frequency temporary wage movements. If, for example, nonmarket capital were accumulated while a person was not working and that capital increased the marginal utility of leisure, then people would be more willing to substitute time over long periods than over short periods. Or, if as modeled by Kydland and Prescott [1982], extended nonmarket time lowers the marginal utility of leisure, then people would be more willing to substitute time over short periods than over long periods.

The model also has strong predictions for consumption which have been the subject of extensive testing in the literature (eg., Friedman [1957], Hall [1978], Hall [1988], Shea [1995]). One area of testing relates to the responsiveness of consumption growth to interest rates and is intimately related to my analysis of labor supply growth. However, variations in ex ante real interest rates that are uncorrelated with tastes and other relevant variables are even tougher to find than are variations in anticipated wage growth. Furthermore, my model is perfectly consistent with lots of substitution of work over time but little

substitution of consumption (just set σ large and make u'' highly negative).

A second area of consumption testing relates to the permanent income hypothesis' predictions for the effect of income shocks on consumption. But, because the consumption side of the model could easily be modified to include intertemporal consumption nonseparabilities (eg., Becker and Murphy [1989], Becker and Mulligan [1997]) or even consumption-leisure nonseparabilities (Heckman [1974], Ghez and Becker [1975]), a variety of observed responses of consumption to income shocks are consistent with the hypothesis that labor supply is correlated with anticipated wage growth.

If workers in my data were literally unable to transfer resources across periods by borrowing, saving, or consuming assets, then the static model of labor supply would apply period by period (whatever a "period" means). The sign and magnitude of the synthetic cohort correlation between work and wages would depend on offsetting "income" and "substitution" effects and, assuming the income effect is positive, the regression (10) would underestimate σ in my synthetic cohort data.¹⁰ As discussed below, σ would be overestimated in my AFDC samples because the income and substitution effects are in the same direction.

III. Synthetic Cohort Samples

III.A. Cohort Cross-section and Individual Panel Approaches Compared

There are two approaches that have commonly been used to estimate σ (with pretty similar results) and each has its own relative advantages. The first focuses on equation (1) and uses individual panel data. Anticipated labor supply growth is measured as actual annual hours growth for those working at t and $t+1$ (as reported in the survey) and anticipated wage growth is estimated in a first stage regression of actual average hourly earnings growth on a variety of variables presumed to be in the date t information set. However, it is important to note that, among those variables that have been used to predict wage growth, functions of age are the best predictors. To see this, consider Altonji's [1986] prediction of average hourly earning growth in a sample of 10,036 continuously married prime aged man-years from the 1968-81 waves of the PSID. With two socioeconomic indicators for parents, years of schooling of father and mother, age, a schooling quadratic, age interacted with a schooling quadratic, and year dummies as explanatory variables, his prediction equation has an R^2 of 0.0054 and a standard error of 0.254. In other words, the standard deviation of his predicted wage growth is 0.0187, which can be compared with the standard deviation of average hourly earnings growth of 0.037 in my CPS synthetic cohort data for cohorts aged 25-

¹⁰An exhaustive list of tests of "liquidity constraints" is beyond the scope of this paper, but I point out that an inability to transfer resources across periods would also imply extraordinary rates of return to schooling and OJT, a prediction which is at odds with the empirical findings of Mincer [1974] and others.

60 and 0.087 for cohorts aged 25-79.¹¹ Hence, among the variables that have been used in the literature to predict wage growth, age dummies have the vast majority of the explanatory power. This, of course, does not rule out the possibility that someday someone will use the PSID or other individual panel data to identify quantitatively important anticipated wage changes that are not associated with age.

The second approach studies cross-sectional cohort ("synthetic cohort") specifications motivated by equation (7). First, we make the decomposition (9) and assume that the cross-sectional covariance is zero:

$$\sigma \ln \frac{Q_t^i(E_t \lambda_t^i)}{\gamma_t^i v_t^i} = \alpha_t a_t^i + \varepsilon_t^i, \quad \text{cov}_t(\varepsilon_t^i, w_t^i(1 - \tau_t^i)) = 0 \quad (9)$$

where i indexes individuals and a denotes a person's age. Second, assume that $g_{t,k}^\sigma$ is distributed uniformly on $[0, e]$ (or, equivalently, integrate equation (1) over time) and obtain (10) by averaging (7) across consumers of the same age:

$$\ln N_t^a = \sigma \ln \tilde{w}_t^a + (1 + \sigma) \ln \bar{n}_t + \ln K + \alpha_t a_t + \varepsilon_t^a + A_t^a \quad (10)$$

where N_t^a is the cohort *arithmetic* average annual hours worked, \tilde{w}_t^a is the cohort *geometric* average after-tax market value of time, and A_t^a is a within-cohort "aggregation" bias depending on the second and higher moments of the within-cohort distributions of ε_t and $\ln \tilde{w}_t$.¹² A very similar expression can be derived relating arithmetic averages of work hours to arithmetic averages of after-tax market values of time, but I use the geometric average \tilde{w}_t^a because it can be decomposed in a straight-forward way into geometric

¹¹It is interesting to note that, if the intertemporal substitution elasticity of "hours" were as large as one and the intertemporal model fit Altonji's data perfectly, the standard deviation of log hours growth would be 0.0187. This is the difference between no wage growth and the growth of annual hours from 2000 to 2038.

¹² $A_t^a = \ln \left[1 + \sum_{i=2}^{\infty} \frac{\mu_i^a(\varepsilon) + \sigma^i \mu_i^a(\ln w)}{i!} \right]$ where $\mu_i^a(\varepsilon)$ is the i th moment of the within cohort distribution of ε , and $\mu_i^a(\ln \tilde{w})$ is the i th moment of the within cohort distribution of $\ln \tilde{w}$.

averages of tax factors, measured pretax wages, and proportional measurement errors.¹³

Treating \bar{n} and K as constants in the cross-section and assuming that the aggregation bias is uncorrelated with \bar{w}_t^a across cohorts, equation (10) is then estimated by least squares in the cross-section of workers aggregated by cohort. Notice that age is included in the regression because - as mentioned in the above discussion of equation (1) - the ISH has nothing to say about the rate of growth of labor supply that would occur in the absence of growth of the after-tax value of time. Age is also included in the regression (10) because cohorts may differ in their time t tastes or their expected lifetime marginal utility of wealth $E_t \lambda_t$.

The slope parameter σ in equation (10) is a well-defined "structural parameter" in my life cycle model. Indeed, contrary to the claims of Smith [1977, p. 249], σ is a parameter of an individual's preferences - it describes the amount of intertemporal heterogeneity in his marginal disutility of work (see also Lucas [1970]). σ also dictates the response of *aggregate* labor supply - including all its components (the fraction working sometime during the year, weeks worked conditional on working, weekly hours, etc.) - to temporary wage fluctuations, as well as the welfare implications of those fluctuations. Most important, the value of σ is the major determinant of the response of macro variables to temporary monetary, fiscal, technological, and other "shocks."

Assumption (9) is crucial, so it deserves some interpretation. It says that those "tastes" for work and interest rates which cannot be explained by a linear term in age are uncorrelated with after-tax market values of time. Part of this is effectively a recursivity assumption - that, given wage growth, workers of all ages trade off working this period or next in the same way. But (9) is also an assumption about "cohort effects."¹⁴ Assumption (9) is violated, for example, when health affects tastes for work and health deteriorates at an increasing rate with age (ie, recursivity is violated) or the detrended expected lifetime marginal utility of wealth $E_t \lambda_t^a$ varies with date of birth in a way that is correlated with detrended \bar{w}_t^a (ie, "cohort effects" do not follow a trend). A weaker assumption than (9) can be used when proxies for tastes or interest rates are included in the regression (10). The presence of cohort effects can also be tested by obtaining cohort cross-section estimates at different points in time. Mulligan [1995] does so and finds similar regression estimates of σ for 1979, 1980, and 1985 (which are in turn similar to the 1976 estimates

¹³Similar regression estimates of σ are obtained when arithmetic averages are used. Even if the aggregation bias term were correlated with \bar{w}_t^a , an estimate of σ inclusive that correlation may be the more relevant for macroeconomic forecasting because macro data is, by definition, aggregated.

¹⁴Depending on the interpretation of λ_t , (9) may also limit the length of a time period or the degree to which λ_t diminishes with lifetime wealth. See Mulligan [1998].

reported here).

The cross-section approach has three important advantages. First, cross-section studies offer a number of interesting and highly relevant variables - such as measures of health, training, and alternative measures of time use - that are unavailable in panel studies. Second, cross-section samples are typically much larger than panel samples so older workers may be more reliably studied and a much richer relationship between age and wages can be accurately estimated in cross-sections. Third, some of the disadvantages of cross-section data can be overcome by complementing the cross-sections with some information from panel studies.

Another difference between individual panel and cross-sectional cohort analyses is that the latter is subject to a "composition bias" due to the death of cohort members and, because in practice the market value of time can only be observed for those who work sometime during the year, the variation across age groups in the fraction and types of individuals who are employed during the year. However, I argue in subsection III.F that cross-section cohort data can be supplemented with individual panel data to correct for the composition bias.

Both individual panel and cross-sectional cohort studies must make inferences about the market value of time when a man is not working. When someone is not working at instant t , it is typically assumed that the market value of time can be inferred from earnings during some other nearby period or from the earnings of a similar person who is working, or some combination of these. But it might be the case that a person's market value of time when not working is low when compared to the times he does work or when compared with the market value of time of apparently similar people who are working. I see no solution to this form of "selection bias" for the individual panel and cohort studies in the literature or for my own cross-section cohort study, but point out that another advantage of my panel study of welfare mothers in section IV is that we can be confident that changes in the tax rules dominate any unobserved changes in the market value of time.

I begin by revisiting the cross-section cohort ("synthetic cohort") samples, sticking with the same basic specification (10) but emphasizing the measurement of the key variables of interest. First, I include the "employment" margin in my measures of "labor supply" as suggested by my model. Second, I include income and social security taxes in the calculations of the value of time. Third, I assume that average hourly earnings misestimates the value of time in a way that is related to the amount of time training on the job. Fourth, I consider the possibility that hours worked as reported by employees to standard demographic surveys are systematically biased. Because much of the data needed to address these issues is only available in micro cross-sections, I am necessarily constrained to construct cross-section cohort samples

instead of individual panel samples. But, fifth, I do use panel data to supplement the cross-section data and address a measurement problem that is peculiar to the latter - "composition bias". I return also to an individual panel sample later in the paper.

Formally, I decompose cohort a 's geometric average after-tax market value of time \tilde{w}_t^a into five components:

$$\tilde{w}_t^a = \hat{w}_t^a (1 - \tau_t^a) \frac{1}{1 - h_t^a} v_t^a B_t^a$$

$$v^a \equiv \frac{\hat{N}_t^a}{N_t^a} \quad (11)$$

- (i) pretax CPS average hourly earnings, \hat{w}_t^a
- (ii) tax factor, $(1 - \tau_t^a)$
- (iii) correction for those reported hours that are self-financed OJT, $1/(1 - h_t^a)$
- (iv) correction for CPS hours reporting error, v_t^a
- (v) composition bias, B_t^a

Below I show how \hat{w}_t^a , $(1 - \tau_t^a)$, and B_t^a are computed as cohort geometric averages from micro-level data, and h_t^a is computed as an arithmetic average from micro-level data. N_t^a and \hat{N}_t^a are computed as cohort-arithmetic averages from separate micro data sets and then their ratio is used to compute v_t^a .

III.B. Total Labor Supply in a Cross-section of Cohorts

Figure 1 displays hours worked for male cohorts aged 20-80 in 1976.¹⁵ The solid line displays CPS average annual hours for all men in each of 61 age brackets. Because annual hours worked are zero for some men, this solid line lies below the dashed line which displays average annual hours for only those men who worked positive hours in 1976.

¹⁵The cross-section is the March 1977 CPS (ICPSR study #7784). Hours are measured as "usual weekly hours last year" times "weeks worked last year." Average annual earnings is measured as "earnings last year" divided by annual hours.

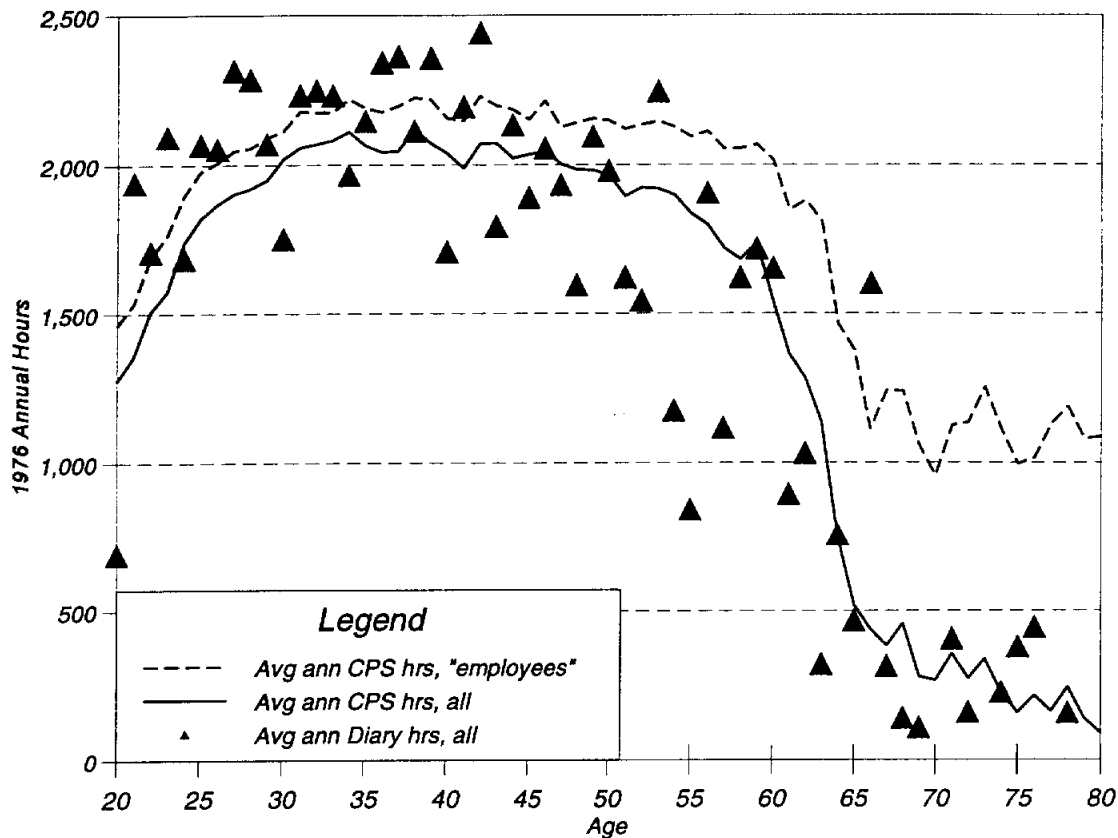


Figure 1 Male Hours Worked by Age Group

The CPS cohort geometric mean of average hourly earnings (which are measured only for those working positive hours in 1976) are displayed as a solid line in Figure 2. Judging from the cross-sectional data displayed in that figure, wages almost double between ages 20 and 40. CPS annual hours worked by “employees” shown in Figure 1 increase, but more modestly. The increase in average hours by all between ages 20 and 40 is more substantial, almost doubling. The data in Figures 1 and 2 suggest that both hours and wages fall dramatically between ages 55 and 80.

Using the 1960 Census 1/1000 Public Use Micro Sample (PUMS), Ghez and Becker [1975] study the hourly earnings of *employed* white men aged 22 to 65. They compute the average annual hours and average hourly earnings by age for the men in their sample. In a regression of each age group’s log average hours on its age and its log average pretax average hourly earnings, they obtain a wage coefficient of 0.39.¹⁶

¹⁶They smooth their data by taking three-year moving averages. Their more widely cited estimate of 0.45 is obtained when income and family size regressors are included. See their Table 3.5.

When the log of "leisure" (5096 - annual hours worked) is used as the dependent variable, they find a coefficient of 0.25. Because they study samples of employed men, measure annual hours and earnings, and isolate the wage variation that occurs with age, the Ghez-Becker specification is essentially the same as those of Altonji [1986] and MaCurdy [1981] (subject to the caveats mentioned in section III.A. above).



Figure 2 Male Pretax Market Value of Time by Age Group

My cohort analysis begins with 1976 annual hours and earnings as reported in the 1977 CPS. In a regression of each age group's log average hours of "employees" on its age and its average log pretax average hourly earnings of "employees", I report a wage coefficient of 0.37 in the first cell of Table 1 for a sample of men aged 25-55. Like Ghez and Becker's, my estimate of 0.37 is from a sample of only those

employed sometime during the year.¹⁷

However, hours of “employees” are difficult to interpret in my life cycle model and may be of limited interest for making predictions for aggregate labor supply N_t^a . For example, hours of “employees” are constant over time and equal to \bar{n} in the special case of $K=1$ while aggregate labor supply might be quite sensitive to temporary wage changes. When aggregate labor supply is measured as in the model - annual hours averaged across all men including those who did not work during the year - I obtain a σ estimate of 0.57.

Heckman [1993, pp. 116, 119] suggests that changes in labor supply over the life cycle consist of changes at both the “intensive” and “extensive” margins. This is not true in my model in which *all* changes are at the extensive margin. However, my model, Heckman's discussion, and Coleman's [1984] and Alogoskoufos' [1987a,b] studies of business cycle fluctuations agree that annual hours worked by those working positive hours is not the same as aggregate labor supply and that the difference between the two - the fraction of workers who work positive hours during a year - may also respond to wage fluctuations. And the larger elasticity estimates in my Table 1 for “all” as compared to “employees only” confirm Heckman's [1993] life cycle conjecture.¹⁸

CPS estimates of σ are sensitive to the number of older age groups included. Table 1 reports estimates about twice as large for samples aged 24-64.¹⁹ After introducing data on taxes and health, I return in subsection III.H to some of the differences between young and older workers.

¹⁷Our samples sizes are quite similar: they have 33,591 men and I have 34,654 men. I include all men regardless of race and do not smooth my cohort data by computing moving averages. Estimates of σ obtained with 3 year moving averages (not reported in this paper) are typically 20% larger than the corresponding estimates reported here. Ghez and Becker also include workers aged 56-65 but, since retirement ages have fallen over time and the importance of social security has grown over time (Costa 1998, chapter 2), my aged 25-55 sample is probably the better comparison.

¹⁸Given the maximum feasible work during a time interval $(K\bar{n})$, the parameter K dictates in my model how aggregate labor supply responses are partitioned between “employment” and “employee hours.” For $K=1$, “employment” is the entire response. For, say, $K=2$, “employment” is half of the response at the margin for a group with an 89% employment rate.

¹⁹Depending on the specification, including the 4 age groups 20-23 or excluding some of the younger age groups affect estimates of σ although not systematically in direction or another. These results are available from the author.

Table 1: Estimates of σ , with OJT and Tax Factors

		IES Estimates			
OJT Source	Tax factor	ages 25-55		ages 24-64	
		"employees" only	All men	"employees" only	All men
none	none	0.370	0.569	0.650	1.406
	"Marginal" A (MA)	0.388	0.588	0.665	1.462
	"Marginal" B (MB)	0.406	0.613	0.693	1.525
	"Average" A (AA)	0.366	0.555	0.661	1.407
	"Average" B (AB)	0.376	0.571	0.673	1.440
SRC	none	0.397	0.614	0.710	1.551
	"Marginal" A (MA)	0.424	0.646	0.786	1.695
	"Marginal" B (MB)	0.442	0.671	0.818	1.778
	"Average" A (AA)	0.425	0.635	0.764	1.590
	"Average" B (AB)	0.435	0.651	0.783	1.641
Heckman et al (1998)	none	0.463	0.761	1.184	2.770
	"Marginal" A (MA)	0.500	0.784	1.348	3.163
	"Marginal" B (MB)	0.487	0.755	1.363	3.223
	"Average" A (AA)	0.512	0.767	1.265	2.829
	"Average" B (AB)	0.501	0.756	1.335	3.029

Notes: (1) IES estimates are coefficients from regressions of log age group hours on age and age group workers' average log after-tax average hourly earnings. Earnings and hours (inclusive of OJT) from the 1977 CPS.
(2) "Marginal" tax factor A is one minus the 1976 Federal Individual Income Tax (IIT), Social Security Old Age, Survivors, Disability, and Hospital Insurance (OASDI & HI) tax on an additional dollar of gross earnings minus - for those aged 62-71 and receiving Social Security - the implicit marginal tax of social security benefits.
(3) "Marginal" tax factor B is "Marginal" tax factor A plus the accumulation of Social Security Wealth (SSW) associated with an additional dollar of gross earnings.
(4) "Average" tax factor A is the increment to 1976 IIT, OASDI & HI, and the implicit tax on social security benefits that would result from not working at all during 1976, as a fraction of 1976 gross earnings, plus one.
(5) "Average" tax factor B is an age weighted average of "Average" tax factor A and "Marginal" tax factor B.
(6) The log of each tax factor is averaged across workers within each cohort to obtain a cohort tax factor. See Appendix I for more details of the computation of the four tax factors.
(7) "employees" are those men reporting positive work hours and average hourly earnings between \$1 and \$100 for calendar year 1976.
(8) SRC OJT is cohort average self-financed on-the-job training, computing from the 1976 Time Use Study as indicated in the text.
(9) Heckman et al OJT is on-the-job training inferred from earnings panel data for young men. See Heckman et al [1998].

III.C. Taxes and Incentives to Work over the Life Cycle

One difference between CPS average hourly earnings and the after-tax market value of time $w_t(1-\tau_t)$

is the labor income tax factor ($1 - \tau_l$). This factor can vary with age because of the progressivity of the federal individual income tax system and its dependence on marital status, the regressivity of social security payroll taxes, and the effect of age on the rate of accumulation of social security wealth. Four components of federal tax and benefit rules are included in my calculations of the tax factor:

- (1) Individual Income Taxes
- (2) Social Security OAS, DI, and HI payroll taxes (employee component only)²⁰
- (3) Implicit taxation by social security benefit formulas from “earnings limits” (applies only to men aged 62-71 who receive a Social Security benefit) and inadequate “delayed retirement credits” (applies only to men age 65-71).
- (4) The accumulation of social security wealth (applies to men under age 72)

The fourth component is the least straight-forward both because rules are complicated and cohort specific and because computation of the tax factor requires information about workers' expectations of future benefit formulas. I therefore report calculations with and without this fourth component.

The accounting period for Individual Income and Social Security taxes is a calendar year. If a potential work session were to coincide with the tax accounting period (as they do in my model), then incorporating taxes into the estimation of σ is fairly straight-forward (see equation (5)), although the CPS does not provide much information about the deductions that families might be taking from their individual income tax or the costs they bear in the acquisition of those deductions. I therefore assume that all families take the standard deduction and, according to equation (5), compare the taxes paid with taxes that would be paid if the man's earnings were set to zero. This produces an “average tax rate” measure of the tax factor.

Another implication of a long work session is that the decision to work before age 62 typically has a negligible effect on the accumulation of social security wealth. 1976 earnings affects Social Security benefit formulas by affecting the lifetime average of one's top index earnings years (AIME). If someone plans to work most of his prime age years, then not working a long session in a particular year can only affect AIME by dropping that year from the calculation and adding another year from the calculation.²¹ I therefore exclude SSW from any calculations of the “average” tax rate.

²⁰It is assumed that the earnings data measured by the CPS is net of employer Social Security contributions.

²¹In the extreme case that a worker's earnings during the years he works grows at the same rate of national index, this substitution of one year for another has zero effect on AIME and therefore zero effect on SSW.

If the tax accounting period includes multiple potential work sessions, incorporating taxes into the estimation of σ is significantly more complicated. In one extreme (and counterfactual) case, however, the appropriate tax factor is one minus the marginal tax rate on an additional dollar of earnings.²² Shorter work sessions also imply that a work decision might have an important effect on social security wealth, so I include one computation of the marginal tax rate that includes the effect of work on SSW. Three of the four measures of the tax factor used in my analysis are:

- (MA) "Marginal" tax rate, PIA fixed (IIT, OASDI & HI payroll, phaseout of OA benefits)
- (MB) "Marginal" tax rate with accumulation of SSW (IIT, OASDI & HI payroll, phaseout of OA benefits, and accumulation of SSW)
- (AA) "Average" tax rate (IIT, OASDI & HI payroll, phaseout of OA benefits, and effect of retirement decision on OA benefits)

My model supposes that K and \bar{n} are the same for all age groups, but another interesting model might allow \bar{n} to increase with age (while holding constant $K\bar{n}$). The growth of \bar{n} with age would, for example, explain why the "employment rate" appears to be (at least in CPS data) the relatively more important margin for the old than for the young, even before the "old" reach age 62. I therefore consider a fourth measure of the tax rate that averages the tax rates (MB) and (AA) with the weights depending on age:

- (AB) age-weighted average of (MB) and (AA)

The four tax rates for each cohort (one minus the within-cohort geometric average of their corresponding tax factor) are displayed in Figure 3. With the exception of men aged 65-71, average rates (AA) are lower than marginal rates (MA and MB). But what is more relevant for estimation of σ is the change in the tax rate with age and we see that average rates (AA) rise more rapidly for prime aged men, a difference which can be attributed to the fact that marginal rates are well above average rates for young men and taxable income is rising with age and, for tax measure (MB), the increased rate of accumulation of SSW with age.²³

Rates of change of the four tax series differ most at ages 62-65. Here average tax rates remain relatively low because most men have not yet retired by age 61 and retirement at age 62 is not particularly

²²This special case requires that the marginal tax rate is a continuous function of earnings (otherwise one has to account for the "kinky" behavior described by Hausman [1985]) and that each tax accounting period includes very many potential work sessions.

²³See Appendix I for some details of my calculations and Feldstein and Samwick [1992] for a study of the accumulation of SSW.

encourage or discouraged by social security benefit formulas. For ages 65-71, however, the social security disincentive for work in the following year is pretty large - over 90% of a year's benefits are lost by delaying retirement one year (Appendix I). Thus Figure 3 displays low cohort-average average tax rates except for ages 65-71. Marginal tax rates increase a bit at age 62 because some of the men who do begin to take social security still earn above the earnings limit and are subject to an additional 50% marginal tax. As men age, more are in this situation so cohort-average marginal tax rates continue to rise in Figure 3.

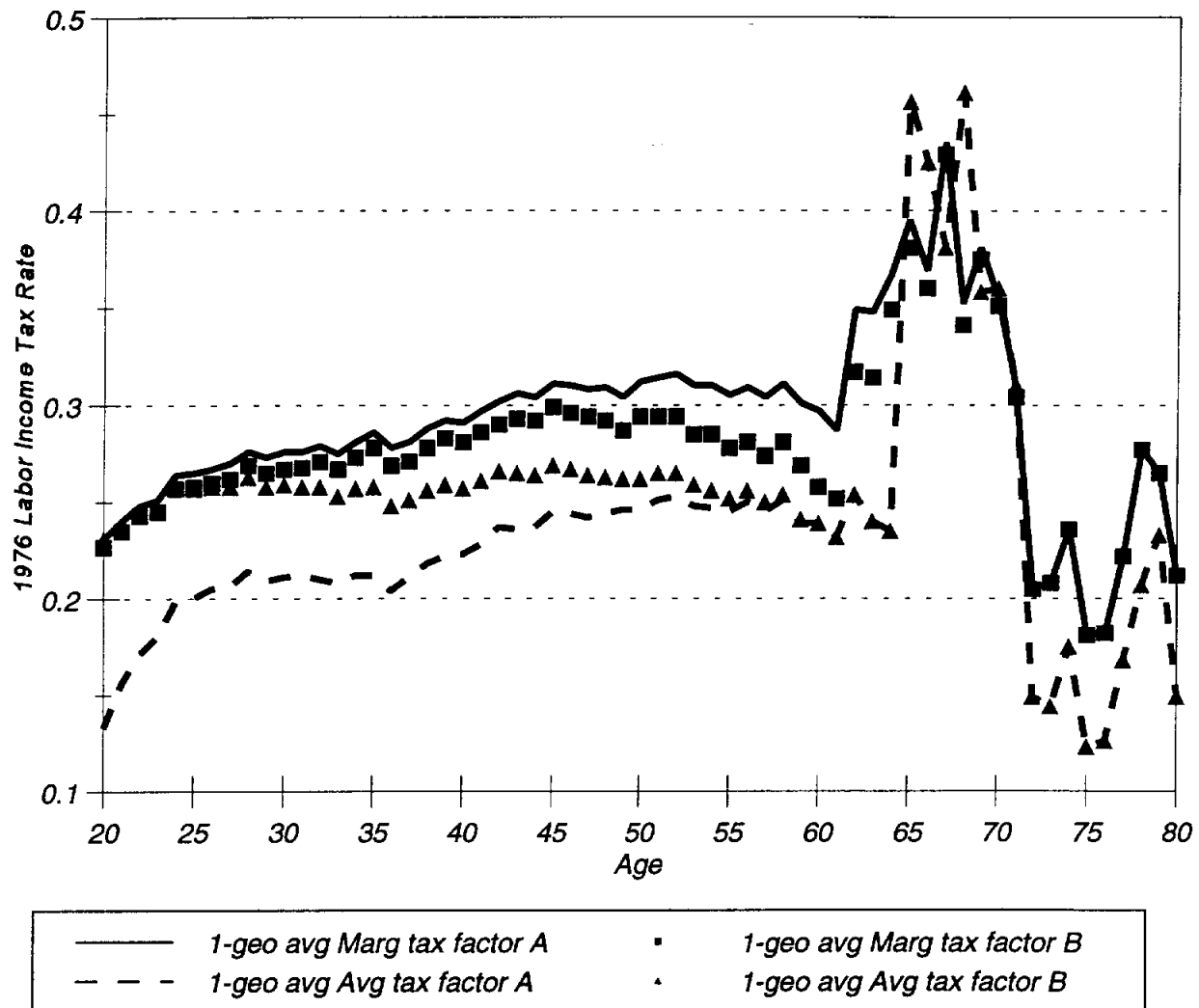


Figure 3 Labor Income Tax Rates by Age Group

It is important to note that none of the series in Figure 3 fully capture the work disincentives of Social Security. One consequence of the social security benefit formulas is that some men aged 62-71 switch to part time jobs to keep their earnings at or below the earnings limit. The part time jobs have lower

hourly pretax wages so a man's acceptance of a part time job - which shows up in Figure 2 as a lower pretax wage - is itself a tax even if he does not pay a dime of the 50% tax implicit in the social security benefit formulas. Thus there are a variety of ways one might partition the after-tax market value of time \bar{w}_t^a into a pre-tax component and a tax factor; my Figures 2 and 3 are only one such way.

Estimates of σ derived from the various tax factors are shown in Tables 1 and 3. The sign and magnitude of the effect of including the tax factor depends on the method of its calculation. To the extent that detrended tastes for work do vary with age (even in a way that is uncorrelated with pre-tax wages), the tax accounting period differs from the period of labor indivisibility, and the tax system is progressive, $(1 - \tau_t)$ is negatively correlated with ε_t and estimates of σ derived from the various tax factors are biased downwards.

III.D. On-the-job training over the life cycle

My empirical analysis so far treats a worker's productivity growth as exogenous. It is quite plausible that the increase in productivity of young men is due to the accumulation of skills over time and that the decrease (or slower growth) in old age is due to depreciation and low levels of skill accumulation. If time is an important component of skill accumulation and skill accumulation occurs on the job, then the correlation of skill accumulation with age leads to a problem with the interpretation of life cycle wage estimates such as those as the solid line in Figure 2. The problem is that worker productivity is estimated as the ratio of earnings to hours spent at work but some of the hours at work are not spent producing and are not compensated by the employer. Younger men are presumably engaged in the most skill accumulation so that their productivity is underestimated the most. As men age, the underestimation is mitigated as skill accumulation time falls. For this reason, the growth of the market value of time of young men is overstated.²⁴

Little data is available on time spent training on the job, and even less on the question of who finances that training. A 1976 study of time use by the Survey Research Center (SRC) asked study

²⁴Jacob Mincer has made this point in several studies (including Mincer [1974, 1997]), distinguishing between average hourly earnings ("wages") and the market value of time ("capacity wages") and deriving implications for the life cycle pattern of hours and earnings. Ghez and Becker [1975, pp. 94, 100] and Heckman [1975, p. 228] also make this point in their studies of labor supply, but Heckman treats training time as unobservable while Ghez and Becker correct for the bias only by including age squared in equation (10).

participants two relevant questions:²⁵

- (i) "Do you feel you are learning skills on your job that could lead to a better job or to a promotion?" (IF YES) "Sometimes people learn these skills as part of their regular work, while others use time at work to learn skills that are *not* part of their regular job. About how many hours *per week* do you usually spend learning new things as *part of* your regular work?"
- (ii) "And how many hours *per week* do you spend learning new things that are *not* part of your regular work?" (Stafford and Duncan [1985], p. 284).

Under the assumption that 100% of the hours reported as a response to (ii) and 50% of the hours reported as a response to (i) are not compensated by employers, I display as a solid line in Figure 4 the total hours of worker-financed training time for each of 61 male age groups.

Rather than attempting to measure time spent training on the job, Heckman et al [1998] infer OJT time from the earnings growth they observe in a panel study of young men. Figure 4 displays their estimates as a dashed line. Because their method of measuring training time is so different from the SRC survey method, it is a nice complement to the SRC measure although it should be noted that the Heckman et al measure is derived from a model with exogenous labor supply.²⁶

The 1976 SRC OJT data is first smoothed by regressing OJT on a quadratic in experience and then the two OJT estimates are used to adjust the CPS estimates of the after-tax market value of time according to equation (11) and displayed in Figure 4. Both measures result in a substantially reduced rate of growth in the after-tax market value of time while the Heckman et al data actually imply a decline in that value of time between the ages of 20 and 23. Not surprisingly, larger estimates of σ are reported in Table 1 when OJT time is used to adjust CPS average hourly earnings. The Heckman et al OJT numbers typically make the biggest difference.²⁷

²⁵The 2406 respondents are the same 2406 respondents who filled out the time diaries that are the object of my study in Subsection III.E below, although the two OJT questions were part of the questionnaire and not part of the diary section of the study (my sample sizes are also smaller because I use only men and require that the necessary variables have valid codes).

²⁶Heckman [1975, p. 255] reports estimates of training time derived from a model with endogenous labor supply that are very similar to the Heckman-Lochner numbers. Heckman [1975, p. 251] also reports another set of estimates which are very much like the SRC numbers before age 30 and like the Heckman-Lochner estimates after age 30.

²⁷Notice that log CPS hours are still used as the independent variable, as suggested by my model where OJT yields disutility just like other work.

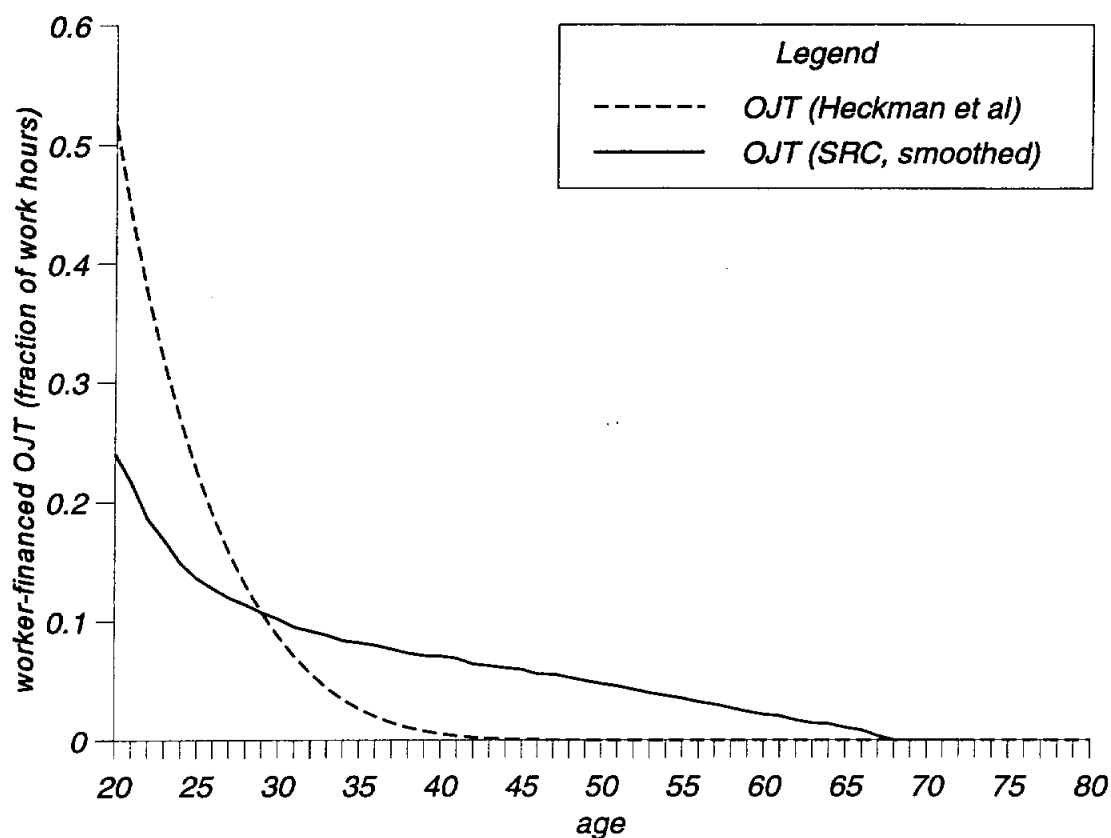


Figure 4 Estimates of the Life Cycle Pattern of Worker-Financed On-the-Job Training

My measurement of OJT implicitly assumes the Mincer/Ben Porath model of training where human capital is accumulated through training time that cannot be used to produce current output. But an upward adjustment of the value of time that declines with age and hence flattens the life cycle value of time profile can be derived from a number of other accumulation models including those with “learning by doing,” “signaling,” or firm-financing of some general training (see Rosen [1972] or Gibbons and Murphy [1992]).

III.E. Hours Reporting Bias

It has been argued by Juster and Stafford [1991, p. 496], Stafford and Duncan [1985], and others that life cycle studies of labor supply are sensitive to the method of measuring labor supply. In particular, they claim that larger life cycle changes in hours are found in time diary data than in CPS-type survey data.

There are a variety of reasons one expects the time diary estimates to be more accurate than

measures based on CPS-type surveys. First, CPS respondents merely answer the questions "how many weeks did you work last year" and "In the weeks you worked, how many hours did you usually work?"²⁸ whereas time diary respondents are obligated to account for all of their time in a particular day or days. The stereotypical response of "40 hours" might be expected for the CPS respondent, but such a response by a time diary respondent would create inconsistencies in his diary unless he actually worked 40 hours (although Pencavel [1986, p. 14] suggests that common reports of 40 are real and the result of legal restrictions). Suppose, for example, that a time diary respondent works more than forty hours in a week. A response of "40 hours" would leave a hole in his schedule which he would have to fill by fabricating a story for what he was doing when he was actually working. Casual "40 hours" responses by CPS respondents would generate the appearance that hours do not fluctuate over the life cycle. A second reason to prefer the time diary is that an attempt is made to measure travel time, coffee breaks, lunches at work, and other activities done "at work" while not actually working. A third advantage of the diary data is that it measures time devoted to finding a job. According to my model, time spent searching for a job is work even though it is not compensated. A fourth reason that CPS annual hours are misreported is that they include time spent on paid vacation or sick leave. Fifth, there is some evidence that - aside from the distinction between weeks worked and weeks paid - workers make systematic errors in their responses to retrospective questions about weeks worked (Horvath [1982]).

I suggest four separate corrections for hours reporting bias. The first and preferred method discards CPS hours all together and computes age group hours from a study of time use. Figure 1 compares the reports of 1977 CPS respondents with results from a 1975-76 SRC time diary study by age group. 624 male respondents completed time diaries for three or four days between October 1975 and September 1976. Total minutes of "normal work," "work on a second job," and "unemployment activities" were summed over the diary days, weighted in such a way as to represent a seven day "synthetic"

²⁸CPS respondents are also asked about hours worked in the week prior to the interview, but there is no corresponding earnings question (for example, there are many CPS men who did not work in the week prior to the interview but had substantial earnings in the prior calendar year). One might use log CPS cohort hours last week as the LHS variable in equation (10) and CPS average log hourly earnings last year as the RHS variable. Slightly higher IES estimates are found, but closer to those reported in the first two columns in Table 1 than those in the last two columns (results available upon request, see also Mulligan [1995]).

week.^{29,30} The sum does not include minutes spent on coffee breaks at work, eating lunch at work, or commuting. There is more idiosyncratic variability in the diary data which is certainly due to (a) the smaller sample sizes used to compute age group means and (b) the greater micro-level variation in diary hours. But it seems clear that, as compared to the CPS survey, somewhat more hours are measured for younger men by the diaries and substantially fewer hours for men nearing retirement age (1976 hours for men aged 50-64 are 1448 in the diary and 1685 in the CPS). This second discrepancy with the CPS is consistent with Ruhm's [1990] study of "bridge jobs" (switches away from career occupation or industry) and, for those aged 62-64, his study of "partial retirement" (periods of employment separated by spells of retirement).

Table 2 shows that the elasticity of hours (workers & nonworkers) with respect to the CPS hourly wage is about 1.4 for the 25-55 year old sample and larger for samples that include older men. Elasticities of 2 are typical when the CPS hourly wage is corrected using the SRC OJT hours (see Table 3), and even larger when Heckman et al's numbers are used (not reported).³¹

Like those based on CPS hours, estimates of σ based on diary hours are sensitive to the number of older age groups included and this sensitivity can be seen in Tables 2 and 3. The diary estimates are even more sensitive to the truncation of older workers in the 50-64 age range because this is the largest discrepancy between CPS and diary hours.

If it is true that the life cycle hours profile is misestimated with CPS data, then average hourly earnings are also likely to be misestimated. Ideally, one would like to discard the CPS hours data from the computation of average hourly earnings and replace it with diary data but, due to the relatively small sample size, Figure 1 shows that the age-grouped diary data is fairly idiosyncratic. This means that average

²⁹My sample is larger than that of Stafford and Duncan [1985] because, apparently, restrict attention to people aged 64 or less with "regular work schedules", exclude "supplemental respondents," and use only one of the four waves of the 1975-6 time study.

³⁰I assume that each respondent's three or four diary days are "randomly" chosen from the year (actually they're chosen from October, November, February, May, June, and September) so that the SRC's calculation of minutes per synthetic week is a calculation of minutes per "representative week". I multiply minutes per synthetic week by 52/60 to get annual hours. Mechanically, the "scaling up" is done first by the SRC who computes minutes per representative week and then I multiply by 52 and divide by 60 to get hours per year.

³¹As is obvious from Figure 1, there is substantially more sampling error when Time Diary cohorts are used to estimate cohort hours. Not surprisingly, OLS standard errors for σ are larger, typically between 0.2 and 0.6 as compared to a typical range of 0.02 to 0.25 when CPS hours are used. Of course, serial correlation and a number of "first stage estimation errors" mean that OLS standard errors overstate the precision of my estimates.

hourly earnings computed with diary data would be idiosyncratic and *in a way that is correlated with idiosyncratic errors in measured hours*. One approach is to instrument for $\log(\text{CPS earnings}/\text{diary hours})$ with $\log(\text{CPS earnings}/\text{CPS hours})$ in the regression of diary hours on average hourly earnings. We see in the first two rows of Table 2 that doing so substantially increases point estimates of σ . However, because of the poor fit of the first stage regression, the magnitude of the increase is quite sensitive to the tax factor, OJT factor, and sample used.

Table 2: Estimates of σ Corrected For Hours Reporting Errors				
		IES Estimates		
Age Group	Method	only hours corrected	hours and wages corrected	Addendum: no correction
25-55	Compute age group hours from Time Diary	1.357	6.417	0.569
24-64		1.991	4.797	1.406
25-55	Compression Model, $\theta = 0.7$	0.729	0.815	0.569
24-64		1.685	2.088	1.406
25-55	Liars Model, $\theta = 0.3$	0.909	1.354	0.569
24-64		1.824	4.354	1.406
25-55	Drop CPS men reporting exactly 40 hrs/wk	1.000	0.809	0.569
24-64		2.261	1.942	1.406

Notes: (1) IES estimates are coefficients from regressions of log age group hours on age and age group workers' average log market value of time. "All" sample (workers and nonworkers) is used.
(2) No tax or OJT factors are used.
(3) When diary hours are used to compute the wage (2nd column, first two rows), log CPS wage instruments for the "corrected" wage.
(4) "No correction" from Tables 1

Diary age group sample sizes are fairly small, with as few as 5 men and 18 days sampled in any single group aged 24-55.³² The typical age group samples 12 men and 47 days. Hence, my second and third corrections for hours reporting bias are of some interest because they do not rely on grouping the diary study by age. Both methods assume a model of hours reporting bias, calibrate the model by comparing histograms of reported hours for 365 diary men and 23899 CPS men aged 25-55 reporting positive hours, and infer the true hours elasticity from estimated CPS hours elasticities. The first model is a "compression model," assuming that reported hours \hat{n} is closer to some stereotypical number \tilde{n} than is

³²Among groups aged 56-69, only the aged 63 group samples fewer days, 8.

true hours n for those whose true hours are positive: $\ln \hat{n} - \ln \tilde{n} = \theta (\ln n - \ln \tilde{n})$. The second liars model assumes that a fraction θ those working positive hours report the stereotypical number \tilde{n} while the rest report the truth. When conservatively calibrated to the cross-section of 365 diary men, both models imply substantially larger elasticities (see the middle four rows of Table 2). Details of these calculations can be found in Appendix II.

A critical review of the literature on time measurement is beyond the scope of this paper (see Juster [1986] for a favorable evaluation of diary studies and Juster and Stafford [1991] for a literature review), but there may be other reasons to distrust time diaries, so my fourth method is based on the CPS data only. It discards those men reporting *exactly* 40 weekly hours (roughly half of the sample) before computing cohort averages and estimating cross-cohort regressions. The method implicitly assumes that all reporting 40 are “liars” and that the propensity to lie is unrelated to determinants of labor supply other than the wage. Estimates of 0.8 and 1.0 are found with this method for the 25-55 age group. Correcting wages in addition to hours increases estimates in the first six rows of the table, but this is not true with the fourth method.

III.F. Older Workers I: Sources of Wage Variation

It appears from the pre-tax wage series in Figure 2 and is obvious from the tax measures shown in Figure 3 that there is more age-related variation in the after-tax market value of time when older workers are included in the sample. All else the same, more variation in the after-tax market value of time might be expected to minimize underestimation of σ due to measurement and other errors. Because labor supply may not be as linear as suggested by my equation (10), it may also be desirable - from the point of view of making predictions for aggregate labor supply - to have an older sample with mean hours worked closer to the population mean.³³

However, including older workers may increase the difficulty of inferring an age group's average market value of time because relatively few work sometime during the year (although see my discussion in the next subsection). Panel data complements the cross-section data in my subsection III.G to make some progress on this problem.

Another problem is that older workers are more likely to hold part-time jobs for “noneconomic reasons” and for this reason part-time jobs may pay less than full-time jobs (author's calculations from the March 1977 CPS). If the old age movement to part-time jobs is in response to the declining labor

³³ Average annual hours worked for those aged 24 and over in the 1975-76 Diary Study (including women) is 1165, which can be compared to 2047 for men aged 25-55 and 1677 for men aged 24-79.

productivity growth that occurs with age, my measures of wages overstate the declining rate of growth and may understate σ . However, an old age movement to part-time employment may be in response to an exogenous change in preferences for work so that the decline in wages is a response to a decline in labor supply (not the other way around!) and σ is overestimated. One adjustment for this is to estimate the age group average pre-tax market value of time from a sample of full-time workers only. Doing so produces estimates of σ which are very similar to those displayed in Tables 1-3, suggesting that these two biases are nearly offsetting.³⁴ It should be noted, however, that even the average hourly earnings of full time workers in their 60's are as low or lower than those of full time workers aged 30.

III.G. "Composition Bias" over the Life Cycle

Because "employment rates" vary with age, the average hourly earnings of a cohort's workers is an average over a sample that varies with age. In contrast, the implications of the theory have been derived for a sample which (ignoring mortality) is constant over time - the entire cohort. To the extent that the market value of time differs for employees and nonemployees, the average log hourly earnings of cohort a 's workers X_t^{*a} is a biased estimated of the average log market value of time of the entire cohort X_t^a . I refer to this bias as a "composition bias" and denote it B_t^a :

$$\text{composition bias}_t^a = \ln B_t^a = (1 - \Pi_t^a) (X_t^{*a} - \tilde{X}_t^a)$$

The composition bias corrections are done separately for "young" and "old" men (age \leq 55 and age $>$ 55). For young men, PSID waves 1976-78 (referring to hours and earnings for 1975-77) are used to correct for the composition bias arising when the average hourly earnings of workers in 1976 are used to estimate the value of time of all cohort members. Both PSID-SRC and SEO samples are used but, by weighting each man with his 1976 family weight, SEO members are effectively down-weighted for the purpose of computing various sample average log wages.

At the annual level of time aggregation, transitions to and from employment are often associated with changes in household status. For example, a young man is likely to form his own household when he

³⁴Full-time wage results - including those that make OJT, tax, composition bias, and hours reporting bias corrections and include workers as old as 79 - are available upon request. Of course, these calculations do not rule out the possibility that the decline in the taste for work mainly reduces wages of full-time workers rather than increasing the propensity to take low wage part-time jobs. It is also possible that full-time wage estimates understate σ because the movement to part-time employment is part of the implicit social security tax.

goes from a full year of no employment to a year of some employment. I therefore utilize the individual files from the PSID (in addition to the family files), because the individual files include information about hours and incomes of men who are not heads of households. The individual files report an individual's "taxable income" and indicate whether that income includes any nonlabor income, but not how much. This is the only measure of labor income available for PSID men who are not heads of households but, not surprisingly, taxable income includes "asset income" only for 2 percent of these men.³⁵ I therefore exclude those nonheads with asset income from my calculations of average log hourly earnings.

In order to estimate the average value of time of nonworkers, some assumptions are required. I assume that, conditional on age, the annual growth rate of a person's market value of time is uncorrelated with changes in his employment status. This assumption is true under the null hypothesis of a zero labor supply elasticity and can be true with a nonzero elasticity, but is false under other conditions. If those who change employment status have a different growth rate of their market value of time, then my estimates of the "composition bias" are subject to a "selection bias." There is no magic correction for selection bias, so I only note that it may be a problem and also point out that nearly every other study of labor supply - including those individual panel studies of the "hours" margin in isolation from the "employment" margin (see my discussion of assumption (A2) on page 8) - has this kind of selection bias. The selection bias arises in so many studies of labor supply because there is no such person who "works continuously" - even the hardest working man has some time when he is not working (eg., evenings, weekends, sick days, holidays, vacations) and the econometrician is typically unable to directly measure the market value of time during those nonwork periods.

III.G.1. Young men

Over a three year period, enough young men (95%) are employed at some time that we can obtain a reasonable estimate of the market value of time of those not employed in any particular year by looking at the earnings of those men in adjacent years. First, I compute the $(t-1), t$ and $t, (t+1)$ average log hourly earnings growth for those employed at least 200 hours in each of the two relevant years. For those not employed at t but employed in an adjacent year, I estimate their market value of time at t by adding (or subtracting) the relevant growth rate to (or from) their average log hourly earnings when they were working

³⁵Based on comparisons of individual taxable income, individual transfer income, and household head labor income for men who are heads of households and report no asset income, it is clear that "taxable income" excludes "transfer income" even though some forms of transfer income are taxable by the IRS.

at $(t-1)$ or $(t+1)$.³⁶

Since employment rates are much greater than 50% for all young age groups, one might expect that the relative market value of time of those not employed is decreasing in the employment rate because those not employed are a “more select” sample in the high employment age groups. In fact, the data - together with the estimation method outlined above - support this.

It is too much to use the relatively few available PSID observations to estimate the market value of time of those not employed as a *function* of age, so I compute instead the mean percentage difference between a nonworker's market value of time and the average value of time of workers his age - a 20% (in log points) difference in my sample.³⁷ Bils' (1985) obtains a similar estimate of 20% for young men from the NLSY. Assuming that the 20% gap is independent of age, the percentage composition bias of my CPS value of time estimates is:

$$\ln B_t^a = (1 - \Pi_t^a) 0.20 \quad (12)$$

where Π_t^a is the fraction of cohort a employed at some time during year t .

If instead relative market value of time of those not employed is decreasing with age (as the PSID data suggests), my calculation (12) overstates the correlation between composition bias and employment rates and I overstate the growth of the market value of time with age among young men. This error tends to reduce estimates of the labor supply elasticity.

III.G.2. Old men

The annual employment status is much more persistent for old men because it is so typical for a man to “retire” and never return to employment again. One therefore might compute the composition bias for old men by accumulating the composition bias. However, this accumulation requires many years of a panel data set large enough to estimate age-specific employment rates and retirement hazards for each cohort. Assuming that $B_{t-1}^a = B_t^{a-1}$, the composition bias can be computed according to:

$$B_t^a = (X_t^{*a} - g_t^a) - X_{t-1}^{*a} + B_t^{a-1} \quad , \quad age > 55$$

³⁶For those working in both adjacent years, I average the results of the two methods.

³⁷The median difference is 11%.

The first term is the (log of) $t-1$ hourly earnings averaged across those working at t . The second term is estimated (log of) date $t-1$ hourly earnings for those working at $t-1$. The third term is an estimate of the lagged bias, which differs from the actual lagged bias to the extent that retirement rates were different for the a and $(a-1)$ cohorts. Hence, the first two terms are the “increment to the composition bias.”

Another relevant complication for old men is that some change from full-time to part-time work and the part-time work tends to have lower hourly earnings. As discussed above, an important part of the difference between part-time and full-time wages by the elderly is the form in which they pay the implicit social security tax so I do not want to count the retirement of a full-time worker as the exit of a relatively high wage person unless his wage is higher than that of other high wage workers. Yet another complication is that workers may receive a bonus, accumulated sick pay, or other extra earnings upon retirement that cannot be attributed to the work they did during the year prior to retirement. This would also make it appear that high wage men are more likely to retire. Using the 1980 5% Census PUMS, I therefore compute the increment to the composition bias by imputing wages for sample men aged 55 and over according to the median average hourly earnings for full-time male workers aged 50-54 in the same schooling and two digit occupation category and then, cohort-by-cohort, comparing the log imputed wage for those retiring and those remaining in the labor force.³⁸ Those who continue to work typically earn 8% per hour more than those retiring in cohorts aged 55-59, 4% more in cohorts aged 60-69, and 6% more in cohorts aged 70-79. In other words, low wage men are more likely to retire, especially before age 60. This finding is consistent with the fact that social security rules encourage retirement most for low earnings people and with the patterns of retirement by occupation and schooling documented by Costa [1998] and others.

Although the wage gap between workers and recent retirees is largest for cohorts aged 55-60, the increment to the composition bias is larger for cohorts aged 60 and over because the retirement hazard is so much larger. Figure 2 displays as crosses the age-wage profile corrected for OJT and composition biases. We see that correcting for composition bias lowers the estimate cohort market value of time (compare the crosses with the squares), especially for young and old men. Table 3 reports corresponding estimates of σ and, when compared with Tables 1 and 2, shows that correcting for composition bias in the aged 24-64 sample typically lowers estimates of σ although the effect is quite small relative to the other adjustments I've made. Although not reported in the tables, the same is true for my estimates for the aged

³⁸The 1980 PUMS is used to have enough observations in each age-labor force status cell and in each schooling-occupation cell.

25-55 and aged 24-79 samples.

Table 3: Estimates of σ including Older Men, with Tax, OJT, Health, and Composition Bias Corrections							
IES Estimates							
Age Group	Tax factor	1977 CPS			1975-6 Time Diary Hours, CPS Average Hourly Earnings		
		health regressors?			health regressors?		
		no	yes	yes, arith avg	no	yes	w/o pers care
24-64	none	1.488	1.238	0.984	2.088	1.167	1.897
	"Marginal" A (MA)	1.538	1.270	1.147	2.132	1.262	1.898
	"Marginal" B (MB)	1.604	1.317	1.225	2.181	1.251	1.888
	"Average" A (AA)	1.509	1.135	0.944	2.128	1.058	1.749
	"Average" B (AB)	1.542	1.191	1.052	2.173	1.117	1.829
24-79	none	2.953	2.799	1.027	2.923	1.139	1.766
	"Marginal" A (MA)	2.672	2.685	1.168	2.438	1.186	1.695
	"Marginal" B (MB)	2.668	2.597	1.363	2.413	1.110	1.583
	"Average" A (AA)	2.193	1.991	1.366	1.916	0.780	1.167
	"Average" B (AB)	2.198	1.982	1.425	1.924	0.777	1.159
Notes: (1) IES estimates are coefficients from regressions of log age group hours on age and age group workers' average log after-tax market value of time. "All" sample (workers and nonworkers) is used. (2) See Notes to Table 1 and Appendix I for explanation of the tax factors (3) OJT is cohort average self-financed on-the-job training, computing from the 1976 Time Use Study as indicated in the text. (4) Health regressors are age group averages of hours spent on personal care and four indicators of physical condition. Personal care hours are omitted from the column labeled "w/o pers care". (5) Column labeled "yes, arith avg" uses arithmetic average CPS wage rather than geometric average (6) See text for details of computation of composition bias correction of wages.							

III.H. Older Workers II: Health as an Independent Determinant of Labor Supply

Aging is associated with changes in physical capabilities, especially at older ages. Some of these changes affect labor productivity and, presuming my various measures of the market value of time are related to labor productivity, are useful sources of life cycle wage variation. But aging may also affect the marginal disutility of work. Without a proxy for the marginal disutility of work, my basic labor supply specifications attribute all of the age-related labor supply changes (apart from a trend) to age-related wage changes.

I introduce several proxies for the marginal disutility of work in the specifications that include older

men. One proxy is annual hours spent on “personal care”: dressing, bathing, toilet, trips to the doctor, helping another adult with personal care, sleeping, and napping. There are two reasons why one might expect this to be a good proxy for the marginal disutility of work. First, a sick and frail person who must spend extra time on these activities effectively has a shorter day to divide between work and “leisure”. Second, one might expect age-related changes in personal care time to be correlated with other factors shifting the marginal disutility of work.

Men spend a substantial amount of time on personal care: 8 minutes per week on medical appointments, 268 minutes per week washing and dressing, 59 minutes per week on medical care for oneself or another adult, 2 minutes per week on other personal care, and 3455 minutes per week sleeping or napping.³⁹ The sum of these personal care minutes are correlated with age, although the only subcategories with some positive correlation are medical appointments, night sleep, and resting or napping (medical care for other adults is slightly negatively correlated with age).

I use four measures of physical capabilities from Wave I of the Health and Retirement Survey (HRS) and the survey of Assets and Health Dynamics of the Oldest Old (AHEAD) to proxy for health status, although it is unfortunate that these health surveys occurred 15 years after my CPS and diary surveys were conducted.⁴⁰ Physical incapacities are likely to be associated with pain experienced during work activities and would thus be a reason why an older person might not work even if he were as productive as when young. Physical incapacity measures are also expected to be correlated with other determinants of the marginal disutility of work.

The first measure of physical (in)capabilities is the fraction of cohort affirmative answers to three questions:⁴¹

“Is it very difficult or impossible for you to...

- (i) ...walk several blocks?
- (ii) ...climb a flight of stairs without resting?
- (iii) ...lift or carry weights over 10 pounds, like a heavy bag of groceries?

My second measure is the fraction of the cohort often troubled with pain that makes normal work

³⁹These are averages for men aged 24-79.

⁴⁰Neither HRS nor AHEAD survey many men aged 50 or less, so I assign all cohorts aged 24-50 the values of the health variables for the 51-52 year-old cohort.

⁴¹Those respondents who say they cannot answer the question because they do not do the activity are assumed to be unable to do it.

difficult.⁴² The third measure is the fraction of the cohort that had been an overnight patient in a hospital during the year prior to the interview.⁴³ The fourth is Body Mass Index (weight in kilograms)/(height in meters, squared)), which has been shown to be closely related to mortality, health, and labor force participation [see Costa 1998, chapter 4, for a review of the relevant literature]. Age group average BMI and age group average squared BMI are included in the cross-age group regressions.

It is perhaps unsurprising that the inclusion of health measures does not substantially affect elasticity estimates, especially in the 24-64 age group. After all, health deteriorated relatively rapidly with age in the early part of the 20th century while gainful employment rates did not (Costa 1998, chapters 2 and 4). For example, Costa (1998, Fig 2.3) reports gainful employment rates of over 80% for white men aged 65 and over in 1880, 1900, 1910, and 1920 (as compared with 20 or 25% for the 1970's). At the same time she reports sizable majorities of veterans aged 65 and over in 1910 suffering from chronic musculoskeletal (67.7%), chronic digestive (84.0%), and chronic circulatory (90.1%) conditions and compares them to much lower rates for WWII veterans aged 65 and over in 1983 (47.2% with chronic musculoskeletal, 48.9% with chronic digestive, and 39.9% with chronic circulatory conditions). Costa also displays data from 1930 and 1992 showing a much steeper 1930 age gradient for risks of heart disease, arteriosclerosis, hypertension, and other chronic conditions. With an age-health gradient that is so small by historical standards, health probably should not explain much of the modern age-employment pattern. Furthermore, my measures of physical incapacity are not that high (eg., only 13% of those aged 65 report that pain prevents them from working) and follow a pretty linear trend with age.

The only case in which the health variables make a substantial difference for the 24-64 group is when cohort hours are measured with the time diary. But this may be spurious because diaries constrain that all uses of time - including the LHS variable (work hours) and one of the RHS variables (personal care hours) - sum to 24 hours per day. So the final column of Table 3 reports wage elasticities from regressions including measures of physical incapacity but not personal care hours.

While modern changes in health with age may not produce dramatic changes in the marginal disutility of work, this is not to say that health does not explain a lot of life cycle labor supply through its

⁴²AHEAD asks a slightly different question - whether pain "kept you from doing things you wanted to do" during the last 12 months. For the age groups that appear in both HRS and AHEAD, affirmative response rates to the AHEAD question are twice as high so I cut AHEAD responses in half to make comparable with AHEAD.

⁴³The HRS also includes measures of cognitive ability such as memory skills. I exclude these measures under the assumption that they are related to labor productivity but not the marginal disutility of work.

effect on the wage. Pre-tax wages grow much less rapidly and even fall with age which partly reflects some declining physical capacity with age and partly reflects human capital investment decisions made in response to that declining physical capacity. Bartel and Taubman's [1979] study of four health conditions finds that, among those working during the year, adverse health reduces pre-tax wages twice as much as it reduces hours and presumably some of this reduction in hours is a response to the wage. The U.S. House Ways and Means Committee [1996, Section 1, p. 5], Diamond and Mirlees [1978], and others suggest that even the implicit tax disincentives of the social security system are a response to the age-health relationship.

III.I. Within-cohort Aggregation Bias

I estimate equation (10) by regressing log cohort average hours on age and average log after-tax market value of time. A within cohort aggregation bias A_t^a , a function of the second and higher moments of the within-cohort distribution of ε_t and $\ln \tilde{w}_t$, is an omitted variable in this regression. One check for the presence of an omitted variable bias is to use log average after-tax market value of time as a independent variable. Doing so has little effect on estimates of σ (results available upon request) with the exception of the aged 24-79 CPS sample and, to a much lesser degree, the aged 24-64 CPS sample (see the third column of Table 3). An aggregation bias term is still an omitted variable, but may be weakly or even negatively correlated with A_t^a so the similarity of results with average log wage and with log average wage suggests that aggregation bias is not serious.

Nor is it clear that purging aggregation bias from estimates of σ is especially interesting for macroeconomic forecasting. Shocks affecting the market value of time may also affect the distribution of ε_t and $\ln \tilde{w}_t$ and do so in a way that is correlated with \tilde{w}_t in much the same way it is in life cycle data.

III.I. Overview of Results from Synthetic Cohort Samples

Life cycle data are consistent with a substantial willingness to substitute leisure over time. Hours worked by a cohort grow almost twice as fast relative to trend as that cohort's average after-tax market value of time. An important - but not the only - component of the cohort's labor supply response is due to changes in its annual "employment rate" while many previous studies of similar life cycle data focus only on "employee hours." Moreover, time diary data suggest that the life cycle changes in aggregate work hours are understated by CPS-type surveys. Training data suggest that life cycle changes in the market value of time are overstated by CPS average hourly earnings. Another departure from previous studies is my consideration of behavior after age 60 or 65. Tables 1-3 review the six modifications I have made to life cycle-based calculations of the IES. Using a sample of prime aged working CPS males, I arrive at an

estimate of 0.37 that is similar to Ghez and Becker's. Expanding the sample to include all prime aged men increases the estimate to 0.57. Modifying this estimate by introducing older workers delivers estimates of 1.5 or higher. Or, instead, an elasticity of 1.36 is the result when aggregate labor supply is considered with time diary data. If the time diaries are correct, young men slightly underreport and retirement aged men substantially over report annual hours in CPS-type surveys. Larger elasticity estimates are found even with CPS data when either a correction for self-financed on-the-job training or a correction for life cycle changes in the marginal tax rate is made. Slightly smaller estimates are obtained after a correction for "composition bias" and an introduction of health regressors. The result of making all seven modifications - considering aggregate cohort labor supply, adding older workers, correcting for life cycle changes in the marginal tax rate and training time, using diary data, and correcting for composition bias produces elasticity estimates of 1 or 2. Even if one is dubious about some of my departures from previous studies of life cycle data, it seems difficult to argue that the life cycle data offer a powerful rejection of the ISH.

IV. A Life Event: the Termination of AFDC Benefits

IV.A. Overview and Comparison with Synthetic Cohort Samples

Clearly the value of time grows at different rates at different points of the life cycle. It is also clear that labor supply grows at different rates over the life cycle. But the discussion and analysis above shows that it is difficult to quantitatively determine how anticipated wage growth changes with age and to what degree labor supply responds to these changes as well as changes in the willingness to work. As an attempt to get better measures of anticipated changes in the value of time and corresponding changes in labor supply - changes which cannot be attributed to changes in the willingness to work or other unobservables - I look at a special life cycle event, the termination of AFDC benefits at the 18th birthday of the youngest child.

IV.B. Termination of AFDC as a Life Cycle Event

For three reasons, the termination of AFDC benefits on the 18th birthday of the youngest child is an interesting life cycle event. First, it is *fully anticipated*. Children can only get older or die, either of which eventually terminates AFDC benefits. Furthermore, since policy for the period 1970-1995 had been rather stable over time regarding the age at which benefits are terminated,⁴⁴ it is only reasonable for most

⁴⁴It can be argued that the 1996 "welfare reform" introduced substantial uncertainty about a particular family's future welfare eligibility. During the prior 25 years, the most important legislative changes were in 1981, changes which affected to some degree the magnitude of the AFDC tax on work (House Ways and Means Committee, pp. 517-18).

of this period for families with youngest child aged 17 to anticipate termination of benefits. Since benefits are reduced to zero on the 18th birthday of the youngest child, the magnitude of the benefit reduction that a family with youngest child aged 17 might expect is easy for them to calculate: a reduction from current benefit levels to zero.

Second, because AFDC benefits act as an implicit tax on the earnings of AFDC household heads, *the event produces a quantitatively important change in the value of time*. To a good approximation, AFDC and food stamp benefits b_t for an *eligible* family at date t can be computed according to equation (2) with $\underline{b}_t = 0$, and b_t a function of calendar time, family size, state of residence, and other demographic variables. In the language of the House Ways and Means Committee, d_t are “expenses reasonably attributable to the earning of income” (including child care costs, transportation costs, and payroll taxes) and “earnings disregards” and other deductions.⁴⁵ R_t is the “benefit reduction rate,” approximately equal to 0.5 for the years 1970-95. The benefit formula (2) thereby acts as a significant tax on work for pay, and represents a change in the value of time that is an order of magnitude larger than those teased out of synthetic cohort samples.⁴⁶

If a “work session” were very short relative to the AFDC accounting period (one month) and the constraint $d_t \leq w p_t$ does not bind, then labor supply might be viewed as perfectly divisible and the tax rate on an additional hour of work for otherwise eligible families is $R_t(t)$ (see Barro and Sahasakul [1983] for a proof). However, if participation in the labor force during an AFDC accounting period requires a discrete increase in hours, then the implicit tax rate $\tau_{t,k}$ on participation is less than $R_{t,k}$ and - to the extent that the length of a work session is similar to the AFDC accounting period - can be computed according to (5) for otherwise eligible families.⁴⁷

While the change in the market value of time is large and anticipated, one disadvantage of a study of AFDC is that the magnitude of the change is difficult to compute exactly. An exact calculation requires information about “earnings disregards”, the costs of obtaining those disregards, the degree to which work is “indivisible,” tastes, the pre-tax market value of time, and other determinants of family eligibility.

⁴⁵Ways and Means Committee 1996, Section 8, page 390. See also for a discussion of the rules concerning deductions and how they changed over time.

⁴⁶Computing the implicit tax on work may be a difficult task (for an economist or for a welfare participant!) but, since families in my sample have youngest child aged 17, it is reasonable to expect that their experience with the welfare system would cause them to act as if they have pretty accurate estimates of the tax.

⁴⁷This formula ignores the effect of AFDC labor supply on asset holdings which, in turn, affects b . Mulligan [1997a] shows that the effect is small.

However, equation (5) shows that the benefit reduction rate is an upper bound on this implicit tax rate. Zero is a lower bound and applies to families whose family structure, asset holdings, or other characteristic makes them ineligible. The implicit tax rate is close to zero for a family with a relatively high pre-tax market value of time.

Many studies of labor supply face the problem of estimating a person's market value of time when he or she is not working. A third advantage of studying the termination of AFDC is that we can plausibly argue that *the change in the value of time associated with the termination of benefits dominates other changes* over the relatively short section of the life cycle being studied. I argue that, while government policy produces an important change in the value of time, the 18th birthday of the youngest child is not associated with large and rapid changes in "tastes," health, productivity, and other variables. However, it turns out that one disadvantage of this episode is changes over time in the implicit tax rate cannot be calculated precisely.

Another disadvantage of my AFDC study is that I do not have OJT or diary hours measures. Indeed, hours reporting bias may be especially systematic before the termination of AFDC since AFDC beneficiaries have an incentive to underreport their earnings (and, hence, their hours) to government agencies - although they may not have the same incentive to misreport to the PSID. If such PSID misreporting does occur, then I am likely to underestimate hours before age 18 and overestimate σ .

IV.C. Data Description and Tabulation

Using the 1970-1990 waves of the PSID and including both the PSID-SRC and SEO samples, I extract all households with youngest child aged 17 and a wife or female head present. I add to these records information on the employment and family situation as well as some consumption expenditures of the wife or female household head two years earlier (when the child was 15) and two years later (when the child was 19). This main sample can be divided into two:

<u>AFDC sample</u>	AFDC income > 0 in calendar year of the interview when the youngest child was aged 17
<u>nonAFDC sample</u>	AFDC income = 0 in calendar year of the interview when the youngest child was aged 17

If the 18th birthday of the youngest child is not associated with large and rapid changes in "tastes," health, productivity, and other variables, then we expect little change in the employment of nonAFDC women whose youngest child turns 18. Column (1) of Table 4 verifies this conjecture, showing a slight decrease in the fraction of women employed sometime during the year as their youngest child ages from

15 to 19. Perhaps this decrease is to be expected given that the women in the sample are typically in their 40's and 50's and labor force participation rates of women are declining slightly in this range (Sweet [1973, Table 1-4] and author's own calculations using the 1981-91 CPS outgoing rotation groups). Conditional on working, annual hours worked for nonAFDC increase slightly.

Table 4: Women's Market Work as a function of Age of Youngest Child					
	(1)	(2)	(3)	(4)	(5)
fract. hours>0 at age 15	0.69	0.30	0.28	0.32	0.35
fract. hours>0 at age 19	0.67	0.33	0.34	0.38	0.43
annual hours, 15	1051	348	314	382	387
annual hours, 19	1090	422	406	486	560
% change annual hours	0.04	0.19	0.26	0.24	0.37
annual hrs of workers, 15	1533	1146	1135	1192	1112
annual hrs of workers, 19	1630	1283	1200	1288	1289
% change ann hrs of wrkrs	0.06	0.11	0.06	0.08	0.15
ann hrs of continuous, 15	1594	1264	1243	1363	1363
ann hrs of continuous, 19	1679	1391	1374	1517	1517
% change ann hrs of cont	0.05	0.10	0.10	0.11	0.11
Sample:	nonAFDC	AFDC	AFDC	AFDC	AFDC
Subsample:	all	all	female heads	fem hd, afdc19=0	fem hd, afdc19=0, hdage<62
Sample Size:	1622	79	65	53	46
Notes: (1) Annual measures of mother's market work according to the age of the youngest child. (2) Percentage changes are computed as log differences. (3) "hrs of continuous" is hours of mothers employed sometime during both of the years when the youngest was aged 15 and 19. (4) Source: Author's calculation using the PSID.					

Column (2) of Table 4 reports the fraction of women in the AFDC sample who were employed when the youngest child was 15 and the fraction of the same women who were employed four years later when the youngest child was 19. We see that the fraction working positive hours increases from 0.30 to

0.33. Considering that many of the families were not eligible for AFDC when the child was aged 15, the percentage change in the annual hours worked by this group is a substantial 0.19 log points. If we exclude the 14 male-headed AFDC households, the fraction increases from 0.28 to 0.34 and annual hours increase by 26 percent.

A few of the 65 age 17 female-headed AFDC households - 12 to be exact - still receive AFDC payments after the youngest child turns 18. For all of these 12, a younger child entered the household and may be the reason for continued AFDC eligibility. If we exclude those households who do not appear to have eligibility terminated, the fraction working increases from 0.32 to 0.38 and annual hours increase by 24 percent.

Seven of the female household heads are old enough to be eligible for social security, although it is not clear that all of them would have had long enough work histories to be eligible. For those who do collect social security after their youngest child turns 18, the change in the incentives to work is different from those who do not. One way to separate the effect of social security is to delete all households with heads aged 62 or more when the child was 19. For those in this sample who have their AFDC eligibility terminated, annual hours increase 37%.

“Annual hours” reported in the third and fourth rows is an average across the same sample of women at two points in time, with women who do not work at all (who are 60-70% of the AFDC sample) counted as zeros. What about hours among those who do some work during the year? This is equal to annual hours of all divided by the fraction with positive hours and is reported in the sixth and seventh rows of Table 4. “Hours of workers” increase substantially in the AFDC sample, about as much as the fractions working some hours. Notice from the first two rows of the table that “hours of workers” is an average over two different samples at the two points in time because the fractions of those working positive hours is different. Another interesting statistic is hours worked by those working in both years (“annual hours of continuous workers”), a statistic which is necessarily computed over the same sample at both points in time. We see even large increases here, and it can be inferred from the table that those “entering the labor force” between ages 15 and 19 (ie, those who have zero hours at 15 and positive hours at 19) work fewer hours than those who were already in the labor force and remain there.

Notice that, after the termination of benefits, annual hours of continuous workers are quite similar to those of women in the nonAFDC sample. However, annual hours and the fractions with positive annual hours in the AFDC and nonAFDC groups are still fairly different. Thus, although the termination of benefits has a substantial affect on work, it is clear that AFDC eligibility cannot explain all of the difference between the employment of the AFDC and nonAFDC samples.

IV.D. Tax Rate and Elasticity Estimates

What do we learn about willingness to substitute work over time from this life cycle event? An estimated intertemporal elasticity of substitution of work is one summary of the data which is comparable to other intertemporal studies wage and work changes. This first requires an estimate of the change in the sample average implicit tax rate on work. I consider three estimates of the sample average τ before age 18: 0.2, 0.3, and 0.4.

At the micro level, the implicit tax rate τ is zero for families who are ineligible regardless of mother's earnings and can be bounded above by the benefit reduction rate (including the Food Stamp program, 0.47 before 1981 and somewhat larger after) for otherwise eligible families. Even conditional on non-earnings determinants of eligibility, it is difficult to compute the implicit tax rate because the tax on earnings is nonlinear, is determined by both AFDC and Food Stamp rules, depends on earnings-related deductions and how they interact with other determinants of eligibility, depends on the costs of obtaining various deductions and "earnings disregards," depends on the pre-tax market value of time and degree to which work is "indivisible," and changes with calendar time. Fraker, Moffitt and Wolf [1985] compute average tax rates of 0.25 in 1971, 0.32 in 1979 and 0.70 in 1982 and even these are too high because they ignore the costs of obtaining various deductions (shown as f in equation (5)). Thus I view 0.5 as a conservative upper bound on the implicit tax rate τ for families satisfying eligibility requirements other than mother's earnings.

In order to compute a sample average τ , we must compute the fraction of families satisfying eligibility requirements other than mother's earnings. We know that a family in my AFDC sample is eligible (or at least appears eligible to the welfare agency) when the youngest child is aged 17 because they receive AFDC income in that year. We also know that the 53 families in columns (4) and (5) of Table 4 are ineligible when the child is aged 19. What we do not know is exactly how many families are eligible when the child is aged 15. However, studies of AFDC mobility surveyed by the House Ways and Means Committee [1996, pp. 500-10] suggest that an important fraction - perhaps as much as one half⁴⁸ - of those receiving AFDC at child age 17 would not be eligible at child age 15 because of changes in eligibility other than high earnings at that age. If so, then the sample average age 15 implicit tax rate could easily be as small as 0.2. Thus I view 0.4 as a conservative upper bound on the sample average τ .

Second, as shown in my model, we need to say something about the shape of the distribution of unobserved determinants of participation in order to say something about how a, say, 50% wage change

⁴⁸80% of the sample of 46 families in column (5) receive AFDC sometime during the calendar year the youngest child was 15.

would affect participation rates in a group with a participation rate of 25% vs. a group with a rate of 75%. We make two assumptions in this regard - normal and uniform distributions. Then the estimated elasticity of participation is the percentage change in $G^{-1}(I_t)$ (where I_t is participation at child's age t) divided by the percentage change in the value of time, ($\ln(0.6)$, $\ln(0.7)$, or $\ln(0.8)$). Table 5 reports elasticity estimates for a variety of work and tax rate estimates. An elasticity of one is fairly typical.

Table 5: Intertemporal Elasticity Estimates					
	(1)	(2)	(3)	(4)	(5)
fract of full-year work at age 15	0.17	0.17	0.19	0.19	0.19
fract of full-year work at age 19	0.20	0.20	0.27	0.27	0.27
implicit tax rate	0.2	0.4	0.2	0.3	0.4
ΔG^{-1} , std lognormal	0.13	0.13	0.28	0.28	0.28
σ	0.60	0.26	1.25	0.78	0.55
ΔG^{-1} , std uniform	0.19	0.19	0.37	0.37	0.37
σ	0.86	0.38	1.66	1.04	0.73
Subsample:	all female heads, welf19 =0, hdage<62				
Notes: (1) "Fraction of full year work" is AFDC sample average annual hours divided by 2080. (2) See text for computation of the implicit tax rate.					

It is well known that AFDC participation involves movement into and out of the labor force (see the studies surveyed by the U.S. House Ways and Means Committee [1996]) so this fact alone strongly suggests that at least some people are willing to substitute work over time. It may even be the case that people select into the AFDC program on the basis of their willingness to substitute over time σ as well as income, tastes, and other variables. If so, then the elasticity estimates reported in Table 5 are not indicative of a typical person's willingness to substitute leisure over time, with the deviation determined by the importance of σ relative to income, tastes and other variables.

AFDC is most relevant for women workers and women may have a different willingness to substitute leisure over time (Mulligan [1995], for example, estimates larger σ 's using synthetic cohorts of women). One potential reason for a difference between women and men is the women are often the "secondary" and male the "primary" household worker in two worker households. The bearing of children

may be another source of differences. But both of these differences are not especially relevant for our AFDC sample because 65 of the 79 AFDC households are headed by women and the youngest child for these women is 17.

Even if the AFDC sample were special with respect to willingness to substitute over time, this select sample may be especially interesting for business cycle and other applications where the workers whose labor supply is changing over time may also be a select sample with respect to σ .

IV.E. Sample Selection Bias

AFDC eligibility depends on employment status, so my AFDC sample is a select sample in two ways. First, as can be seen from the age 19 employment variables for the AFDC and nonAFDC samples, the AFDC sample selects adults whose lifetime propensity to work is relatively low. Second, to the extent that the propensity to work varies over time in ways unrelated to the AFDC benefit formulas, my AFDC sample selects parents who happened to have a low propensity to work in the year their youngest child was 17 years old. For this reason alone, employment and earnings in the AFDC sample would tend to be higher during years the child was not 17.

However, the selection bias with respect to willingness and ability to work when the child was aged 17 is not particularly relevant to my calculations unless it introduces a differential selection bias with respect to ability and willingness to work at child ages 15 and 19. In an intertemporally separable model, it seems that the selection biases should be similar for aged 15 and aged 19 work measures.

This may not be the case if there is some willingness to substitute over time *and* labor supply is not intertemporally separable. To the extent that work (leisure) capital is accumulated by working (not working), the sample of adults not working much in the year prior to an anticipated increase in the market value of time (ie, when the child is aged 17) is a sample of adults who, for one reason or another, do not anticipate working much following the increase. If instead, work (leisure) capital is accumulated by not working (working) as in Kydland and Prescott [1982], the sample of adults not working much in the year prior to an anticipated increase in the market value of time is a sample of adults who are “resting” in anticipation of an above average amount of work in the following year or two . Thus intertemporal nonseparabilities introduce a differential selection bias on the aged 15 and aged 19 work variables but, without saying more about the form of the nonseparabilities, the direction of the differential bias cannot be determined.

While the “tastes” or nonmarket value of time of a typical woman might not change substantially at the 18th birthday of her youngest child, my AFDC sample might - because of AFDC rules regarding

marital status - be one where tastes do change. Marriage decreases the likelihood of AFDC eligibility (Rosenzweig [1995]) so my AFDC sample may select women who, because of program incentives or for other reasons, are more likely to get married after their youngest child's 18th birthday. If marriage increases a woman's nonmarket value of time, then this is a force *discouraging* work after age 18 which would bias my estimates downward.⁴⁹

IV.F. "Income" Effects

Since the termination of AFDC benefits at date t^* is fully anticipated, it cannot be a "wealth" or "income" effect in a model in which each period's budget is somehow tied together. In such a model, the only determinant of the *relative* labor supply before and after t^* is the relative value of time (including, if appropriate, time discounting) and the rate at which resources can be transferred across periods.

If budgets between periods were not tied together in any way - perhaps because borrowing and lending involve a substantial fixed cost or because individuals do not realize that such transactions are possible and advantageous - then, despite its anticipation, the termination of benefits would have an "income" effect as well as the usual substitution effect. Both effects would tend to promote work after AFDC. One way to test for this possibility is to see whether consumption falls at t^* . Using Skinner's [1987] proxy for nondurable consumption (a weighted average of food purchased at home, food purchased away from home, and rent or housing value), real family consumption decreases 8% (\$250 a year, in 1967 \$) from age 15 to 18 in the sample from column (5) of Table 4 with valid consumption indicators. Annual family food consumption (including food stamps) falls \$100. Given that AFDC/ADC annual income falls by an average of \$1121, these decreases are economically insignificant. But it is a decrease, and it should be noted that an decline in consumption together with an increase in work is not found in other life cycle studies (Ghez and Becker [1975], Heckman [1974]).

There are two reasons why even a large reduction in consumption at t^* might be consistent with the dynamic labor supply view. First, consumption and leisure may not be separable. Second consumption as I measure it is excessively sensitive to household size since it measures rent and food expenditures. If, for example, part of the effect of AFDC were to keep teenagers living at home until their benefits terminate at which time the teenager moves out, then household rent and food expenditures would fall even though each individual's standard of living might be unchanged. Unfortunately, there are not enough households

⁴⁹Note that there is not an "income effect" of marriage in the life cycle model unless the expectations of marriage by women in my sample deviate from the fraction of them who actually get married.

in my AFDC sample with stable composition to test this second hypothesis.

V. Conclusions

Following Mulligan [1998], I build a model of labor force participation and time aggregation which, in reduced form, looks very much like the empirical models used by Lucas and Rapping [1969], MaCurdy [1981], or Altonji [1986]. The model predicts that labor supply is substituted over time in response to *anticipated* wage changes, although the magnitude of the response depends on the parameter σ which in theory can vary from 0 to ∞ . The model is quite explicit about the measurement of wages and hours including the treatment of on-the-job-training, time spent searching for work, nonlinear tax rules, and the aggregation of various components of labor supply.

It might be argued that people cannot substitute over time because they are unable to “choose their hours” or because they cannot borrow or lend as freely as the model implies. Although my model does include the idea that people are less than free to choose their hours - labor is “indivisible” - it is true that the present value budget constraints (3) are crucial for deriving the main empirical specifications. While the validity of the assumptions of the theory is open to debate, I have shown that a reasonable analysis of various age-related changes in wages and labor supply confirms one main prediction of the theory - labor supply growth is positively correlated with anticipated wage growth.

In a broad sample of men, I find that age groups of men with 1% more age-related growth in their after-tax market value of time have 1-2% more age related growth in group hours worked. This response is an order of magnitude greater than those found in other life cycle studies of men. Equation (11) summarizes what I believe to be my essential departures from that literature: “labor supply” is defined to include the “employment” and all other margins, the market value of time is adjusted for taxes, average hourly earnings are adjusted for self-financed OJT, and CPS hours are adjusted for systematic reporting bias.

Because my inferences from the life cycle data are so different from those made in previous studies, I expose my empirical analysis through a few simple graphs, difference estimators, and regressions. More efficient estimators are certainly available, but I want it to be clear that my point of departure from previous studies is not in the econometric details but in the very basic economic and statistical issues which are arguably of primary relevance to the problem.

Studies of individual panel or synthetic cohort data - mine included - cannot do much about the fact that the market value of time when not working is not observed. Someone who is not working at an instant in time may have a market value of time that is low when compared with his average hourly

earnings at some other point in time or with the average hourly earnings of an otherwise similar person who is working at that instant. My study of the termination of AFDC benefits upon the 18th birthday of the youngest child is an attempt to mitigate this bias since I believe that the observable change in the after-tax market value of time dominates any unobserved changes. IES estimates from this study are quite similar to those found with the synthetic cohort data.

Social security rules create another life cycle event that, like the termination of AFDC benefits, is a quantitatively important anticipated change in the after-tax market value of time. This event has been included in my analysis of synthetic cohorts although, for students of intertemporal substitution, retirement is worth its own study. And there have been a number of studies of the effects of social security rules on retirement decisions with a variety of results. Cross-country studies (eg., Gruber and Wise [1997], Modigliani and Sterling [1983]) have typically enjoyed large and fairly obvious differences in benefit rules and have found large differences in retirement. Some time series studies (eg., Burtless [1986], Hausman and Wise [1985], and Krueger and Pischke [1992]) have enjoyed less variation in benefit rules and found even smaller labor supply responses.⁵⁰ However, there are two disadvantages of social security rules as a source of an “exogenous” change in the anticipated rate of wage growth. First, it is not at all clear that workers are fully aware of all of the subtleties of social security rules and their changes over time, including those changes that are exploited by some of the time series studies. Second, it can be argued that social security rules and their changes are less likely than other policies to be “exogenous” with respect to the situation of the people affected by the policy. The AARP is America's most powerful lobby (*Fortune* December 8, 1997) and its preferences are certainly reflected in social security legislation. The endogeneity of social security policy might also explain the relatively large cross-country correlation between labor supply and social security rules. Nevertheless, I eagerly await a study of retirement designed to measure the magnitude and importance of intertemporal substitution and perhaps even to overcome some of the disadvantages I've mentioned.

Another avenue for future life cycle research may be to exploit some of the differences in wage growth across occupation and schooling categories. However, doing so is a difficult exercise if the issues I emphasize in this paper are important. It seems that the level and life cycle rate of change of OJT varies across occupation and schooling categories (although none can be detected in my fairly small data set - see Stafford and Duncan [1985]) so that the more rapid average hourly earnings growth in some categories cannot be fully attributed to more rapid growth in the market value of time. Also, the currently available

⁵⁰Several time series studies also utilize unanticipated policy changes which are less relevant for the ISH.

time diary data sets seem too small to correct for hours reporting biases separately by occupation and schooling categories.

I study synthetic cohorts of men mainly because there is a greater consensus in the previous literature about the lack of response of male labor supply to temporary wage fluctuations. But macro data is an aggregate of men and women so a differential response by gender would be one reason why my synthetic cohort estimates might not directly apply to macro modeling. I do study women in Section IV and obtain similar estimates of σ , but it is unclear whether the AFDC women are more like the representative man or representative woman. Like men, the AFDC women are typically heads of household and might for that reason have a lower σ than the representative woman. On the other hand, AFDC women are women. AFDC women might even have a higher σ than other women because a higher σ makes AFDC participation more attractive.

There is another reason that caution should be exercised in the application of my estimates to macro modeling. The OJT, reporting, composition, and other biases which are important in the life cycle data may also be present in the macro data. There is no easy solution to this since the biases may be of different magnitudes or even different signs in the two sources, but one can try to similarly correct macro data or rely mainly on earnings data (rather than hours and wage data) where we suspect some of the biases to be less important.

Although the life cycle data suggest that the temporal pattern of work responds to the temporal pattern of the market value of time, the data do not necessarily support the hypothesis that all or even most of the variation in work across workers or over time for a given worker can be explained as a response to temporary wage fluctuations. In fact, my model allows for a number of other potential determinants of work through the “indivisibility” of labor as well as cross-sectional and intertemporal variation in the taste parameter γ . I only suggest that, *if* and *when* temporary fluctuations in the after-tax market value of time occur, aggregate labor supply will respond and respond in a quantitatively important way. And, regardless of the number and importance of other determinants of labor supply, my claim has important implications for macroeconomics.

VI. Appendix I: Notes on Social Security Benefits and Work Incentives

Because Social Security benefits are a function of a worker's life history of earnings and labor force participation, they can affect incentives to work at all ages. These incentives can be parsed into two categories: (i) the effect of work on earnings and the “primary insurance amount” (PIA) that will be used to compute benefits when retirement occurs, and (ii) the effect of participation and earnings at age 62 and

older on the fraction of the PIA to be received as a social security benefit. This appendix draws heavily on the House Ways and Means Committee [1996, Section 1] and Myers [1993, chapters 2 and 3].

VI.A. *Work before age 62 and the PIA.*

A person's PIA is a concave function of his lifetime average earnings where each year's earnings is indexed to reflect nationwide wage growth (with the exception of earnings after age 60 which are unindexed), the average caps each year's earnings according to the payroll tax cap, and the average drops the lowest indexed earnings years. The second provision means that no additional benefits accrue for men earning above the cap. If the number of potential work sessions in a year (K) is small, the last provision is particularly relevant for workers who participate most of their lives because the marginal participation decision has no affect on PIA. To see this notice that, unless the labor supply trend term is very different in magnitude from the rate of nationwide wage growth, the marginal participation decision occurs in a year with low indexed earnings conditional on working.⁵¹ I refer to this as the "PIA fixed" case.

If K is large, the marginal work session will typically affect computation of the PIA and this is the case analyzed by Feldstein and Samwick [1992]. For single men and married men whose wives earn enough to have substantial social security benefits, they use the following formula to compute the effect of an additional dollar of earnings at age a on discounted expected social security benefits:

$$\frac{1}{35} e^{.01 \max\{a_R - a - 5, 0\}} \frac{\text{PIA}}{\text{AIME}} \left(1 - \frac{\tau^{\text{IT}}}{2} \right) S(a_R, a, r)$$

The first term reflects the use of 35 years to compute lifetime average indexed earnings (AIME). The second term reflects the indexing, where a_R is the normal retirement age and a is age (Feldstein and Samwick assume individual and national earnings growth of 1%/year). PIA/AIME depends on AIME; Feldstein and Samwick assume the median ratio of 0.32. τ^{IT} is the marginal individual income tax rate to be paid in the retirement years and S is the factor used to compute the discounted expected value of an annuity that begins paying at age a_R . Feldstein and Samwick do their calculations for 1992 under the assumption that benefit rules (a_R , PIA/AIME , τ^{IT}), will never change from the values dictated by current

⁵¹Even if the marginal participation decision did not occur in a low index earnings year (conditional on working), nonparticipation would produce zero earnings, remove the year from the average and add another year to the average so the effect of the decision on the lifetime average would be limited to the difference between what might can be earned in the year in question and what was earned in the best year not included in the average.

law. In fact, rules did change between 1976 and 1992 and probably will change again. Rather than making guesses about what each age group in 1976 knew or expected about the policy and mortality parameters, I use Feldstein and Samwick's parameters. I use their highest discount rate of 6%/year to reflect (a) expectations of the benefit cuts in one form or another that so many experts and laymen have been forecasting and (b) the associated political risk premium.

Feldstein and Samwick also perform calculations for married men with wives without substantial lifetime earnings. I do not review their calculations here, but assume that roughly half of all men earning below the cap expect to be in this category when they retire and therefore take a simple average of the formula above and the increment to discounted expected benefits enjoyed by this second category of men.

In summary, my "PIA endogenous" calculations for men earning less than the social security cap average the increment to discounted expected benefits implied by Feldstein and Samwick's [1992] columns 2 and 4 of their Table 1. This increment is subtracted from the 1976 payroll tax rate.⁵²

VI.B. *Workers aged 62-71*

The same PIA effects act as a subsidy to work for those 62 and over who have not elected to begin receiving their social security benefits. I extend Feldstein and Samwick's calculations by adjusting their annuity factor $S(a_R, a, r)$ for 6% interest as age nears 65 and, after age 65, for mortality as indicated below. Those working after age 61 are also still liable for payroll taxes.

In addition to the payroll taxes, two major provisions of Social Security have discouraged work for pay by the elderly: earnings limits and delayed retirement credits. The earnings limit is the maximum earnings an elderly worker can earn without losing some of his old age benefits. The limit does not apply to those aged 70 or higher. For earnings above this amount (\$2760 per year in 1976 for those aged 62-71), old age benefits are reduced (employment status does not and did not affect Medicare eligibility, although since 1980 employment status can affect the health insurance premiums paid by an elderly person. House Ways and Means Committee [1996, Section 3, p. 223]). The rate of reduction during the 1970's was \$1 for each \$2 over the limit.

For those *workers* aged 62-71 who are receiving social security, I compute the additional tax on work as 0 if he earns below the limit and 0.5 if he earns at or above it. It is important to note that these men are typically part-time workers earning a lower *hourly* pretax wage than they would have earned if they had worked full time, so the use of their pretax wage in my calculations in Subsection III.B already

⁵²Feldstein and Samwick report the increment in 5-year intervals which I geometrically interpolate.

includes some of the disincentive effects of social security.

For workers aged 72+, no additional tax or PIA subsidy is computed although the low hourly earnings of these men is to some degree a consequence of Social Security's encouraging them to work fewer hours in their 60's.

If a worker elects not to receive any old age (or disability) benefits, he is credited with some additional old age benefits when he later retires. These credits are called "delayed retirement credits," but are small enough (1%) that most people aged 65-71 were effectively penalized for delaying retirement.⁵³ A person aged 65-71 of average health planning to retire in 1976 lost over 90% of a year's benefits by delaying retirement an additional year. However, only if K is small is this provision a *marginal* tax on work.

For someone who first retired in 1976, it is straightforward to compute the amount of these benefits, the implicit tax on nonretirement for that year and (assuming K is small) the implicit marginal tax on work. However, too little is known about the earnings histories of CPS men who have not yet retired to compute exactly the benefits they would enjoy if they had retired in 1976. If these potential benefits were known, the implicit tax on work (in addition to the IIT, payroll taxes, and the accumulation of SS wealth considered previously) could be computed as:

$$[0.92 + (a - 62)0.005] \frac{b_i}{e_i}$$

where b_i/e_i is person i 's ratio of potential benefits to actual earnings. The term in brackets is greater than zero and a function of age because the delayed retirement credits were actuarially unfair.⁵⁴ I estimate the ratio b_i/e_i as 0.4 for single men, which is a fairly typical "replacement rate" reported by the House Ways and Means Committee [1996, Table 1-14] for cohorts retired by 1976. 0.6 is used for married men over 65 under the assumption that their retirement also entitles their spouse to benefits equal to half of the amount of their own benefits.

⁵³Hurd and Boskin [1984] and others argue that the decision to retire at age 65 is not distorted because the delayed retirement credits are "actuarially fair." This may have been true at the time of writing their paper for a person of average mortality risk, but not before 1979 when the DRC was only 1% and before 1975 when there was no DRC (House Ways and Means Committee [1996, Section 1]). The DRC for those aged 62-64 was 8.4%, which I treat as actuarially fair.

⁵⁴Details of my derivation of the term in brackets - which equals $[1 - (0.01)S(a_R, a, r)]$ - are available upon request.

VII. Appendix II: Two Models of Reporting Errors

Two of my four corrections for CPS hours reporting bias rely on a model of reporting errors. The first is a “compression model” (A-1) where each worker reports hours that are closer to a stereotypical number \tilde{n} than is his actual hours. The second is a “liars model” (A-2) where some apparently random workers choose to report the stereotypical number \tilde{n} .

$$\ln \hat{n} - \ln \tilde{n} = \theta (\ln n - \ln \tilde{n}) \quad n > 0 \quad (\text{A-1})$$

$$\hat{n} = \begin{cases} n & \text{w.p. } 1 - \theta \\ \tilde{n} & \text{w.p. } \theta \end{cases} \quad (\text{A-2})$$

where n is true hours and \hat{n} is reported hours.

Remember that, by definition, aggregate labor supply is the product of average hours conditional on positive hours and the fraction working positive hours. In order to derive the effect on aggregate labor supply of reporting bias among those reporting positive hours during some time interval, we need to model both aggregate labor supply and the fraction working positive hours (the “employment rate”). My life cycle model provides a quite tractable model of aggregate labor but, unless K is known, not a model of the employment rate. Even if K were known, my model of the employment rate is not particularly tractable. So, as an approximation, I assume a log-linear model for true age group hours averaged across those with positive hours:

$$\ln n_t^a = \eta \ln \tilde{w}_t^a + u_t^a, \quad n_t^a > 0 \quad (\text{A-3})$$

with n_t^a the arithmetic mean of true hours among those with positive hours, and \tilde{w}_t^a the age group *geometric* average after-tax market value of time. n_t^a is assumed to be uncorrelated across age groups with \tilde{w}_t^a or with the error term in equation (10). Total age group hours N_t^a are still determined according to equation (10). It is assumed that measured average age group hourly earnings \hat{w}_t^a differ from true average age group hourly earnings \tilde{w}_t^a according to:

$$\hat{w}_t^a = \tilde{w}_t^a n_t^a / \hat{n}_t^a \quad (\text{A-4})$$

Up to an aggregation bias term, the compression model implies equation (A-1) for average

reported age group hours conditional on positive hours (\hat{n}_i^a) as a function of average true age group hours conditional on positive hours (n_i^a). Given the compression factor θ , the reported hours elasticity with respect to the measured wage \hat{w} , and the R^2 of the reported hours equation, we can infer the true hours elasticity η from equations (10), (A-1), (A-3), and (A-4):⁵⁵

$$\eta = \left[\frac{\theta}{\hat{w}} \frac{\theta - (1 - \theta)\hat{w}}{\theta - (1 - \theta)\hat{w} \frac{1}{R^2}} - (1 - \theta) \right]^{-1}$$

while the elasticity of true hours with respect to the reported wage is \hat{w}/θ .

Given also the reported total hours elasticity with respect to the measured wage $\hat{\sigma}$, we can infer the true total hours elasticity σ :

$$\sigma = \eta + (\hat{\sigma} - \hat{w}) \frac{\eta \theta}{[1 + (1 - \theta)\eta] \hat{w} \frac{\theta - (1 - \theta)\hat{w} \frac{1}{R^2}}{\theta - (1 - \theta)\hat{w}} - (1 - \theta)\eta \hat{w}}$$

The elasticity of true total hours with respect to the measured wage is $(\hat{\sigma} - \hat{w} + \hat{w}/\theta)$. Similar calculations (not shown here) can be made for the liars model.

Both the compression and liars model have a parameter θ which can be calibrated by comparing reported hours in the CPS and the time diaries. For example, the micro-level standard deviation of log hours is twice as large in the diaries as in the CPS and (appropriately correcting for the different sample sizes) the standard deviation across groups aged 25-55 is almost three times as large in the diaries. This suggests a compression parameter of $\theta=0.4$ or $\theta=0.5$. I use a conservative $\theta = 0.7$ for the compression model calculations reported in the text.

It is beyond the scope of this paper to build and formally test various models of reporting errors, but I do comment on their "realism." Four facts come out of an elementary comparison of the distribution of positive "hours last week" (or "usual hours last year") reported by men aged 25-55 from the March 1977 CPS with positive hours worked in the synthetic week reported by men aged 25-55 from the 1975-76 Time Diaries. First, close to 50% of CPS respondents report *exactly* 40 hours and 75% report working exactly 52 weeks in 1976 as compared to 19% of time diary synthetic weeks between 38.0 and 42.0 hours. When either CPS measure of weekly hours is multiplied by weeks worked in 1976, about 40% report exactly 40 hours as 1976 average weekly hours. Second, the standard deviation of log reported weekly

⁵⁵All of the estimated parameters are to be understood as probability limits.

hours is twice as large in the time diary as in the CPS. Third, the fraction of CPS workers reporting exactly 40 hours is very slightly increasing with age until age 63 after which it is about half of what it is for younger workers although the age group variance of log reported weekly hours increases steadily with age in both the CPS and diary studies. Fourth, the central tendency of the distribution of weekly diary hours is roughly 40 hours, with a majority of synthetic weeks between 30 and 50 hours.

The first fact is more consistent with the liars model since the compression model does not literally predict that many would report exactly 40 hours, although the compression model might be amended to account for the first fact by introducing some rounding to the nearest integer. Both models are consistent with the second fact. The third fact suggests that the liars model is more appropriate up to age 63 and the compression model after age 63 since the compression model predicts that a widening of the distribution of true hours would decrease the fraction of CPS respondents reporting hours within a given distance of 40. However, the fourth fact is more easily derived from the compression model since, in contrast to the liars model, the stereotype is related to the distribution of misreported hours even for those not reporting the stereotype.

Notice that the liars model also justifies the procedure used to produce results in the last two rows of Table 2 if we assume in addition that all of those reporting 40 are "liars." Actually, the procedure is justified even if liars are not random with respect to their wage as long as they are random with respect to other determinants of hours.

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