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David Card

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ABSTRACT

The lifecycle labor supply model has been proposed as an explanation for various dimensions of labor supply, including movements over the business cycle, changes with age, and within-person variation over time. According to the model, all of these elements are tied together by a combination of intertemporal substitution effects and wealth effects. This paper offers an assessment of the model's ability to explain the main components of labor supply, focusing on microeconomic evidence for men.

David Card
Department of Economics
Princeton University
Princeton, NJ 08544

Intertemporal Labor Supply: An Assessment

The systematic study of intertemporal labor supply began only two decades ago.[1] In a remarkably short time the lifecycle model of individual hours choice has moved to the forefront of both micra- and macro-econometric research. This paper begins with a look at the original questions that first lead to interest in the lifecycle approach. I then present a selective review of the evidence on various dimensions of intertemporal labor supply. I limit my discussion to microeconometric studies of male labor supply. making no attempt at an exhaustive survey of even this branch of the literature.[2] Rather, my goal is to offer an assessment of the success and/or failure of the lifecycle model in providing a useful framework for understanding the major components of individual labor supply

I conclude that the lifecycle labor supply literature sheds very little light on the questions that first generated interest in a lifecycle approach: what determines the shape of the lifecycle hours profile? how does labor supply respond to aggregate wage changes? what is the source of idiosyncratic changes in year-to-year labor supply? Part of the reason for this stems from a tendency in the literature to concentrate on one aspect of intertemporal hours variation -- the response to wage growth along a known lifecycle trajectory •• and to ignore another, namely, the response to wage innovations that lead to revised expectations about future wage

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[1] Lucas and Rapping (1970) seem to be the first authors to use an explicit intertemporal model to describe short and long run labor supply phenomena. although Mincer (1962) distinguished between the effects of short run unemployment and long run wage increases in explaining the behavior of female labor supply.

[2] Excellent surveys are available in Killingsworth (1983) and Pencavel (1986).

rates. More seriously perhaps. much of the effort devoted to studying lifecycle labor supply has taken the position that average hourly earnings during the year is a "sufficient statistic" for hours choices within the year. There is considerable evidence against this narrow reading of the lifecycle model.

I. The Questions

A series of substantive questions motivated the original interest in lifecycle labor supply. Lucas and Rapping (1970). following an original suggestion of Friedman (1976, pp.206-207), posited a lifecycle framework as a way to reconcile an elastic short-run labor supply curve with an inelastic or even backward-bending long-run labor supply curve. Lucas and Rapping's idea was to model cyclical hours variation as a response to a temporary wage change. Subsequently, much debate in the macroeconomics literature has focused on the size of this intertemporal wage elasticity.

A second motivation for studying lifecycle labor supply arose from interest in human capital theory, and the recognition that the pattern of lifecycle hours is influenced by the pattern of lifecycle wage rates. This goal is clearly articulated by Heckman (1976, page S12), who notes that a model with endogenous labor supply can potentially reconcile differences in the lifecycle profiles of earnings and hourly wage rates.[3]

A related question is whether wage growth over the lifecycle can explain the parallel profiles of consumption and earnings. The simplest permanent income model predicts no systematic relation between earnings and

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[3] Lucas and Rapping (1970. footnote 11) also noted in passing that their model had "...lifecycle as well as business-cycle implications."

consumption. The finding that individuals with steeper lifecycle profiles of earnings have steeper lifecycle profiles of consumption has therefore bee" used as evidence of credit constraints or other impediments to an optimal lifecycle allocation (Thurow (1969), Ghez in Ghez end Becker (1975). Carrol and Summers (1989)). As Heckman (1974) pointed out, however. a model with endogenous labor supply can explain the parallel age profiles of consumption and earnings, if leisure and consumption are complements.[4]

Other questions have also emerged: what (if anything) can we conclude about the interpretation of measured unemployment (Ashenfelter and Ham (1979))? how does a lifecycle perspective affect the interpretation of the responses measured in the Negative Income Tax experiments? how doer a" intergenerational transfer system (such as Social Security) affect the hours of young and old workers? Finally, and perhaps fundamentally, how can we explain the enormous year-co-year variation in individual-specific labor supply that appears in virtually every available panel data set?

The power of the lifecycle framework. and the extent of economists' faith in the model, are illustrated by considering a simple decomposition of individual labor supply into aggregate time effects, systematic age effects, permanent person-specific effects, and person-and-year specific effects. The lifecycle labor supply model has bee" proposed as a" explanation for all four components! Lucas and Rapping (1970) proposed that a lifecycle model could explain aggregate year-to-year m ovements in labor supply (the "time effects" in a components-of-variance model).

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[4] This same idea can potentially explain the "excess" covariation of income and consumption growth in aggregate data.

Heckman (1974, 1976), Ghez and Becker (1975), and others proposed that the lifecycle model could explain systematic age effects in hours of work, and also differences across people in their amount of market work over the lifecycle (i.e. the person-specific constants), Finally, models used by MaCurdy (1981), Altonji (1986) and others link person-and-year-specific changes in hours to the corresponding changes in wages.

II. The Basic Model

A prototypical lifecycle labor supply model begins with a time-reparable utility function defined over consumption (c_{it}) and hours of work (h_{it}) of individual i in each of a sequence of periods t=0,1,2,...:

$$(1) \qquad \Sigma_t \beta^t U(c_{it}, h_{it}, a_{it}).$$

Here, $\beta=(1+\rho)^{-1}$ measures subjective time discounting and a_{it} is a sequence of "taste shifters" that capture heterogeneity across individuals and over time. In models with uncertainty, preferences are assumed to be additive over states and time (with the same U( ) function) so that the consumer's objective function is simply the expectation of (1). conditional o" current information.

The second element of the model is the intertemporal budget constraint. which describes the change in the value of assets (A_{it}) between periods:

$$(2) \qquad A_{it+1}/p_{t+1} = (1 + r_t)(A_t/p_t + w_{it}h_{it} - c_t).$$

Here, p_t is the price of consumption goods in t, r_t is the real interest rate in period t (assumed to be known). and w_{it} is the real wage of individual i for hours worked in period t.

An interior solution for maximization of the expectation of (1), subject to (2) and a" appropriate terminal condition on assets. is characterized by first-order conditions for consumption and hours in period t, together with a" intertemporal optimality condition for the marginal utility of wealth in period t (λ_{it}):[5]

$$(3a) \qquad U_c(c_{it},h_{it},a_{it}) - \lambda_{it} = 0$$

$$(3b) \qquad U_h(c_{it},h_{it},a_{it}) + w_{it}\,\lambda_{it} = 0$$

$$(3c) \qquad \lambda_{it} - \beta\,(1 + r_t)\,E_t[\lambda_{it+1}] = 0.$$

Equations (3a) and (3b) can be solved for consumption and hours in terms of w_{it} and the current marginal utility of wealth. It is conventional to refer to the implied solution for hours as the "intertemporal labor supply function". With a" appropriate transformation of the taste shift variable a_t, write the log-linear approximation[6] of this function as:

$$(4) \qquad \log h_{it} = a_{it} + \eta \log w_{it} + \delta \log \lambda_{it}.$$

The parameter η represents the elasticity of hours in period t with respect to wages in t, holding constant the marginal utility of wealth. Following the literature. I shall refer to η as the intertemporal substitution elasticity. This elasticity is necessarily positive. and is strictly greeter than the (Hicksian) compensated labor supply elasticity associated with the same preferences. if leisure is a normal good. The parameter δ represents the elasticity of hours with respect to the marginal utility of

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[5] See MaCurdy (1985) for example

[6] Of course one could start with a specification of U that implies the log-linear intertemporal labor supply function (4). Issues of functional form are discussed in Browning, Deaton, and Irish (1965).

wealth, and also must be positive if leisure is a normal good. The two elasticities are related by the simple condition

$$\eta - \delta = \frac{c_{it}}{w_{it} h_{it}} \cdot \frac{\partial \log c_{it}}{\partial \log w_{it}}$$

If consumption is independent of wages, holding constant the marginal utility of wealth (as is implicitly assumed in the permanent income consumption model), then $\eta = 6$.

Note the convenient form of the lifecycle labor supply function (4). As a consequence of the additive structure of preferences, the effects of asset income and future wages are completely summarized in the value of λ_{it}. With perfect foresight and constant real interest rates, (3a) implies that $\lambda_{it} = \lambda_{i0} g^{t}$, where g is greater or less than 1 depending on the gap between the real interest rate and the rate of time preference ρ. In this case, apart from taste changes and a geometric trend. the lifecycle profile of labor supply is completely determined by the profile of wages.

The implications of the lifecycle model under uncertainty are most easily seen by combining equation (3c) with equation (4) to describe the change in hours between periods t-l and t:

$$(5) \quad \Delta \log h_{it} = \Delta a_{it} + \eta \, \Delta \log w_{it} - \delta \cdot (r_{t-1} \rho) + \delta \, \phi_{it} + \delta \, \xi_{it},$$

where $\phi_{it} = \log \lambda_{it} - E_{t-1} \log \lambda_{it}$ is the one-period-ahead forecast error in the logarithm of the marginal utility of wealth. and $\xi_{it} = -E_{t-1}[\exp(\phi_{it})]$.[7] The letter term is a constant if the (prior) distribution of ϕ_{it} is constant. Thus, the change in labor supply consists of a component due to changes in tastes (Δa_{it}), a component due to

[7] I have simplified (5) slightly using the approximations $\log(1+\rho)=\rho$ and $\log(1+r_t)=r_t$.

variation in wages, a component due to the difference between the real interest rate and the rate of time preference, and a component due to any updating in the logarithm of the marginal utility of income.

The simple form of equation (5) has considerable appeal. and variants of if are used in many recent microeconometric studies of labor supply. In a stochastic environment, however. it is important to keep in mind that the response of individual hours to a change in wages has two parts. The first of these is $\eta\ \Delta w_{it}$, as in the perfect foresight model. The second is the change in labor supply generated by the change in the marginal utility of wealth. The realization of w_{it} provides new information that generates an update in the distribution of future wages and brings about a revision in the forecast of λ_{it}. Unfortunately, there are no closed-form expressions for λ_{it} in an uncertain environment.[8] Thus, the component of the change in labor supply attributable to wealth effects is usually treated as a "nuisance", and is eliminated by an instrumental variables procedure. This is not to say that the wealth effects associated with observed wage changes are small. Indeed, my reading of the evidence suggests they are potentially significant. However, the difficulty of deriving a formal or even approximate expression for λ_{it} has lead most researchers to concentrate on the intertemporal substitution effect.

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[8] In fact, closed form expressions for λ_t under perfect foresight are not easily obtained. One case chat can be solved uses an LES-form for the within-period utility function U. See Ashenfelter and Ham (1979).

III. Empirical Implications and Evidence

a. The Lifecycle Profile of Hours

The first and most direct implication of the lifecycle mod.1 concerns the shape of the lifecycle hours profile. As pointed out earlier, with perfect foresight and constant real interest rates. the model implies that the lifecycle profile of hours consists of a taste component, a trend, and a component that is strictly proportional to wages. The presence of uncertainty adds other components with mean zero over a large sample of lifecycles. To see this, re-write the lifecycle labor supply function as:

$$\text{(4.3)}\quad \log h_{it} = a_{it} + \eta \log w_{it} + \delta \{ \log \lambda_{i0} + \sum_{j=0}^{t-1} (\rho \cdot r_{t-j-1} + \phi_{it-j}) \}$$

$$= E0 \log h_{it} + \eta (\log w_{it} - E0 \log w_{it})$$

$$+ \delta \{ \sum_{j=0}^{t-1} (r \cdot r_{t-j-1} + \phi_{it-j}) \},$$

where E0 denotes expectations at the beginning of the lifecycle, and r is the expected real interest rate in period 0 (assumed to be constant). Hours at age t differ from hours planned at the beginning of the lifecycle by a term representing the forecast error in wages, plus another representing the cumulative forecast errors in interest rates and the marginal utility of income. Over a large sample of lifecycles (spanning different periods of calendar time), the estimated age profile of lifecycle hours therefore converges to the mean of the planned profiles.[9]

The typical shapes of the lifecycle profiles of wages and hours for male workers are illustrated in Figures 1 and 2. The underlying data for

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[9] Obviously, it may not be possible to recover an unbiased estimate of the planned lifecycle profile of hours from a sample of individuals in the same cohort, since these individuals share the same aggregate-level shocks in each year of their life.

these figures are taken from the 1977-1989 March Current Population Surveys (CPS), and pertain to annual hours and average hourly earnings (annual earnings divided by annual hours) for calendar years 1976-88.[10] Figure 1 shows annual averages of log wages for 6 single-year age cohorts. Each distinct line in the figure tracks the wage profile of a single cohort over the 13 year sample period. Figure 2 shows the corresponding profiles of average annual hours.

The data in Figure 1 suggest char successive cohorts face similar expected wage profiles: real wages rise quickly between the ages of 20 and 30. and then grow more slowly to a peak around age 50. Nevertheless. there are obvious year effects in average hourly earnings.[11] and important cohort effects, During the 1980's, later cohorts tended to earn lower real wage rates than earlier ones. This negative wage growth provides a" interesting opportunity to test Lewis' (1956) influential interpretation of the trend toward lower hours of work during the first half of the 20th century. Levis (1956, p. 197) argued that the decline reflected a" income effect, driven by higher average wages for successive cohorts of workers. If this interpretation is correct, one should detect a" increase in hours for the mast recent cohorts.

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[10] The samples for each year consist of men age 16-70, excluding those who are classified as self employed and those with allocated wage and salary earnings. Individuals who report positive wage and salary earnings. positive weeks of work, and positive usual hours per week for the previous year are counted as working. Individuals who were working and who report average hourly earnings less than $1.00 or greater than $75 (in 1983 dollars) are deleted from the sample. The sample sizes in each year range from 36,000 to 42.000.

[11] Average real wage rates declined sharply between 1979 and 1981. For the youngest cohort in Figure 1, this effect appears as a slowdown in the rate of growth of wages. For older cohorts, real wages actually declined.

The lifecycle profiles of hours in Figure 2 have a rather different shape than rhe profile of wages. Per-capita hours of work reach their peak in the early 30s. are roughly constant co age 40, fall slightly to age 50, and then decline sharply. The pattern of hours among chose who actually work is similar. reaching a peak of about 2100 hours at age 30, remaining arable to age 50, falling to 1900 hours at age 60, and then declining sharply. The growth in hours at the beginning of the lifecycle coincides with a gradual withdrawal from school. Thirty percent of all 20 year olds in the March CPS (1977-89) report their main activity in the previous veek as "in school". This fraction falls to 11 percent by age 23 and to 2 percent by age 30.[12] Much of the decline in per-capita hours at the other end of the lifecycle reflects withdrawal from the labor force. By age 62, only 50 percent of men are still working any hours. Lifecycle patterns in enrollment and employment probabilities are illustrated in Figure 3, which graphs the average probabilities by age for men in the 13 year CPS sample.[13]

The hours profiles in Figure 2 indicate strong year effects, with all cohorts showing a downturn in hours in 1982. In contrast to the profiles of wages. however, the hours profiles of the younger cohorts are not systematically different from those of the older cohorts. Thus, there is no evidence for the inter-cohort income effects underlying Lewis' explanation for the earlier decline in per-capita hours.

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[12] The CPS does not ask "weeks in school" during the previous year, or give any breakdown of hours per week into work and school time.

[13] The employment and enrollment rates in Figure 3 are not adjusted for any cohort effects. 'However, adjusted races are very similar.

Ho" well does the lifecycle model explain the lifecycle profile of hours? Between the ages of 20 and 30, "ages grow by 40-45 percent, per capita annual hours grow by 55 percent, the employment rate grows by 10 points, and hours conditional on working grow by 45 percent. Between the ages of 30 and 50, "ages rise another 10-15 percent. conditional hours are constant. and the probability of working falls 5 points. Finally, between the ages of 50 and 60, "ages fall 5 percent, conditional hours fall 5-10 percent, and the employment rate falls by over 20 points. Clearly, the degrees of "curvature" in the lifecycles profiles of "ages and hours are different. Of course this does not refute the lifecycle model, because tastes may vary systematically with age. and it is also possible that the intertemporal substitution elasticity varies with the number of hours worked.[14]

A stronger test is provided by the data in Figures 4 and 5, which represent "age and conditional hours profiles for men in 3 education classes: O-0 years of schooling. exactly 12 years of schooling, and 16 or more years of schooling.[15] Between the ages of 30 and 50 the "age profiles of these three groups differ dramatically. Wages of college graduates grow some 40 percent, "ages of highschool graduates grow about 20 percent. and wages of individuals with minimal schooling grow only 10 percent. However,

[14]The "age profiles are also potentially biased estimates of the "age profiles for the "hole population, since we only observe "ages for workers. One way to evaluate the size of this bias is to assume that "ages for those not working would be at some lower bound (say. the minimum "age) and then to re-calculate the average "age. This procedure suggests that rhe bias in the "age profiles up to age 50 is trivial.

[15]These profiles are estimated age coefficients from regressions of average log wages and average log hours on age effects, year effects, and a set of broad (10-year interval) cohort effects.

for all three groups. hours (conditional on working) are constant between age 30 end 50. In fact, the hours profiles of the different education groups are very similar. To explain these data with a simple lifecycle model requires a fairly elaborate set of taste parameters.[16] A simpler interpretation is that the shape of the wage profile bears no causal relation to the shape of the hours profile.

It also is interesting to compare the three education classes in terms of their average lifetime hours and average lifetime wages. For simplicity, assume that individuals with 0-8 years of schooling begin work at age 16, vhile highschool graduates begin work at 18 and college graduates begin work at 22.[17] Then average hours worked per year between the ages of 16 and 69 for the three education groups are es follows:

Years Education	Hours/Year	Hours/Year. if working
0 – 8	1265	1756
12	1537	1809
16+	1638	1833

Given the wage differentials between the 3 groups. these data suggest that higher lifetime wages are associated with higher lifetime hours. This

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[16] One could also appeal to models with endogenous human capital accumulation. Evidence presented by Holtz-Eakin, Newey and Rosen (1988) however, indicates that legged hours have no influence on future wages. This seems to rule out simple capital accumulation models.

[17] These assumptions clearly understate the total labor supply of the more-educated workers. First, many students work part-time or part of the year. Second, actual time spent in school is arguably closer to work than leisure.

positive association calls into question the conventional view that long-run labor supply is negatively associated with wages.[18]

b. Economy-Wide Fluctuations

Much of the initial interest in lifecycle labor supply facussed on its potential value in explaining cyclical changes in employment and/or hours. Since cyclical variation in real wages is limited, en equilibrium model with a stable aggregate labor supply function requires a relatively high elasticity of labor supply to generate large swings in employment or hours. In principle, a lifecycle framework can reconcile relatively elastic labor supply responses over the business-cycle with inelastic (or even negatively sloped) 'long run" Labor supply, Recall that the intertemporal substitution elasticity (η in equation (4)) is necessarily larger than the elasticity of hours holding constant either utility or wealth. Thus the intertemporal substitution effect of a given change in wages $\eta\ \Delta w_{it}$, is certainly positive and is potentially large

To see the implications of the lifecycle model at the aggregate level, consider forming the average change in labor supply between periods t-1 and t for a sample of individuals. Equation (5) implies that

$$\text{(5a)}\quad \Delta \log h_t = \Delta a_t + \eta\ \Delta \log w_t - \delta\ (r_{t-1} - \rho) + \delta \phi_t,$$

[18] Finegan (1962) examined data on wages and weekly hours in different occupation and industry classes, and found a generally negative relation between them. On the other hand, Finegan's results indicate a positive association by level of education. However, he dismisses this evidence. asserting that wage differentials by education class include premia for training costs that should be netted out. I have attempted an analysis similar to Finegan's using data an 483 3-digit occupations for men inthe March 1988 CPS. These data show a strong positive association between average hours and average wages in different occupations. even controlling for education and other demographic factors.

where $\Delta \log h_t$ represents the average change in log hours in the sample, Δa_t represents the average change in the taste variable, $\Delta \log w_t$ represents the average change in log wages. and ϕ_t represents the mean of the forecast errors in log λ_{it}. In principle it is possible to estimate (5a) on aggregate-level data. Something like this is actually carried out in Lucas and Rapping (1970). Altonji (1982). and Mankiw, Rotemberg and Summers (1985). Here I vish to discuss the implications of (5e) for the "time effects" that emerge in microeconometric studies of labor supply. This idea was suggested by Ashenfelter (1984) and is pursued by Angrist (1989. 1990).

Ashenfelter (1984) observed that aggregate changes in labor supply for a fixed cohort take a particularly simple form if (i) there are no aggregate components of taste variation. (ii) the real interest rate equals the rate of rime preference, and (iii) individuals have perfect foresight. In this case equation (5a) reduces to

$$\Delta \log h_t = \eta \, \Delta \log w_t.$$

Apart from sampling error. the mean change in hours is strictly proportional to the mean change in wages. This specification can be freed up by assuming that the taste components of individual labor supply follow a systematic lifecycle trend. For example, suppose that

$$a_{it} = a_i + b \, Age_{it} + c/2 \, Age_{it}^2,$$

where a_i is a permanent person-specific component of tastes, Age_{it} denotes the age of individual i in period t, and b and c are common population parameters. Then equation (5a) implies

$$(6) \quad \Delta \log h_t = -b \cdot c/2 + c \cdot t + \eta \, \Delta \log w_t$$

Since (by assumption) the only aggregate components of labor supply are taste and wage variation. equation (6) should fit the mea" changes in hours end wages exactly. apart from sampling error in the estimated means. Therefore, as the number of individuals in the panel increases, the R^2 associated with (6) should tend to unity.[19]

Estimates of this equation are presented in Angrist (1989) using the means of wages and hours for a panel of males in the Panel Study of Income Dynamics (PSID).[20] Corresponding estimates based on cohort-level data from consecutive CPS samples are presented in Angrist (1990). In analyzing the CPS samples, Angrist (1990) divides the available data into two subsamples -- 1963-74 and 1975-87 -- and follows me" age 25-50 in 1964 in the first subsample, and me" age 25-50 in 1976 in the second. Angrist's estimates qf the intertemporal substitution elasticity (with their estimated standard errors in parentheses) are as follows:

[19] These implications are unchanged if one adds a person- and time-specific component of taste variation to the model.

[20] Actually, Angrist estimates the aggregated labor supply function in level form.

	Sample:		
	PSID 1969-79	CPS 1963-74	CPS 1975-87
Preference Trend:			
None	-.13	-.01	.61
	(.04)	(.01)	(.09)
Linear	.56	.25	.58
	(.12)	(.08)	(.09)
Quadratic	.63	-.04	.94
	(.21)	(.10)	(.14)

Angrist also reports a specification test based on the R^2 of the fitted models. The specifications that include either linear or quadratic taste components yield test statistics below conventional significance levels in the PSID sample. In the CPS sample, all of the test statistics are above their .5 percent critical values. although it must be recognized that the sample sizes are large -- 7,000 to 10,000 per year. Interestingly, none of the CPS results is substantively different when the analysis is repeated on samples of men with a fixed age distribution in each year.

These results suggest that there is a systematic positive relation between mean wages and mean hours, particularly in the more recent sample period. The relationship is illustrated in Figure 6, which plot= two measures of average annual labor supply together with a measure of mean log wages for men age 20-50 in 1976. Wages and hours for these men (and for other cohorts) rose between 1976 and 1978, fell in the early 1980's. and then recovered. The timing of the post-1980 upturn differs between wages (which grew between 1981 and 1982) and hours (which continued to fall until 1982). The covariation of wages and hours is also weak in the last 4 years

of the data. Nevertheless. wage and hours changes from 1976 to 1988 are very highly correlated

Should we conclude that intertemporal labor supply does a good job of explaining the time effects that emerge in an microeconometric model? My belief is that such a conclusion is premature. The reason is that the assumption of perfect foresight regarding the aggregate changes that occurred in the late 1970's and early 1980's is surely false. In the 3 decades before 1976, average real wages in the U.S. economy grew fairly steadily at 2-3 percent per year.[21] After 1975. real wage growth essentially stopped. This sharp downward adjustment in trend, coupled with the actual losses in real wages in the early 1980s, suggests that many individuals suffered unexpected reductions in their lifetime wealth According to the Lifecycle model, these changes should have affected hours decisions, and therefore should be modelled in the aggregate labor supply equation.

The difficulty is that very little is known about the evolution of the marginal utility of income or the size of the wealth elasticity 6. One approach is to write down a" intuitively plausible or econometrically convenient model for λ_{it}. For example. Lucas and Rapping (1970) specified a labor supply function of the form

$$(7) \quad \log h_{it} = a_{it} + \eta (\log w_{it} - \log w^{*}_{it}) + \theta \log w^{*}_{it}$$

where

$$\log w^{*}_{it} = \sum_{j=0}^{T-t} b_j \, E_t \log w_{t+j} \qquad (\Sigma b_j = 1)$$

[21] Between 1947 and 1976, for example, real average hourly earnings of "on-supervisory workers rose at a" average annual rate of 2.38 percent.

This is equivalent to replacing 6 log λ_{it} with $-(\eta + \theta) \log w^{*}_{it}$ in the labor supply function (4).[22] As Altonji and Ashenfelter (1980) pointed out, the labor supply effects of aggregate wage changes in this model depend critically on the degree of persistence in innovations to the real wage. I" fact, it is difficult to reject the hypothesis that the aggregate real wage rate is a random walk with drift. If workers assume that the "year effects" in individual wages have the same property. then the labor supply effect of a change in the aggregate component of wages depends only on the "long run" elasticity 8.[23] If this is negative (as Lewis (1956) and many subsequent authors have assumed) and if individuals expect aggregate-level changes in real wages to persist indefinitely (as is perhaps true for changes in economy-wide real wage rates) then the predicted correlation between the year effects in hours and wages from a panel of individual data is negative!

The only evidence in the microeconometric literature pertaining to the sign of the "long run" labor supply elasticity (i.e., the elasticity of hours with respect to a parallel shift in wage profiles) is from MaCurdy (1981, 1985). MaCurdy (1985) suggests a less restrictive specification for the marginal utility of income than Lucas-Rapping:

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[22] One can derive a" intertemporal labor supply function that is approximately equivalent to the Lucas-Rapping function (with θ=0) using the within-period preference function

$$U(c,h) = c \cdot a\ h^{(1+\eta)/\eta}.$$

However, this is only valid in the absence of uncertainty.

[23] To see this, decompose log w_{it} into a permanent person effect. a year effect v_t, and a person and year specific effect. and suppose $E_t(v_{t+j}) = v_t$. Then (7) implies log $h_t = a_t + \theta v_t$.

$$\delta \log \lambda_{it} = \gamma_a A_{it} + \sum_{j=0}^{T-t} \gamma_j \quad E_t \log w_{it+j} \quad .$$

This specification implies that the elasticity of mean hours with respect to a permanent change in wages is $\eta + \kappa_{it}\bar{\gamma}$, where $\bar{\gamma}$ is the mean of the γ_j coefficients over the remaining lifecycle and

$$\kappa_{it} = 1 - \frac{\gamma_a A_{it}}{\delta \log \lambda_{it}}$$

varies with the share of current assets in lifetime wealth. MaCurdy (1985) presents estimates for $\bar{\gamma}$ centering on -.07 for individuals at the start of their lifecycle. This is an upper bound on the absolute magnitude of the wealth effect of a permanent innovation in wages for older workers, since these individuals have a larger share of lifetime wealth in assets. MaCurdy's estimates, then, suggest that the wealth effect of a permanent change in wages is small. and that a permanent 10 percent increase in wages is associated with a roughly 1 percent increase in hours. [24]

In my opinion, much more work needs to be done on measuring the wealth effects of expected future wages before we can conclude that the lifecycle model provides an adequate description of the year-to-year changes in average labor supply observed in a panel. One useful exercise that has not yet been carried out is to combine information on mean levels of consumption and hours for a panel such as the PSID. The assumption of perfect foresight implies that changes in mean consumption are described by an equation of the form

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[24] MaCurdy's estimates of the intertemporal substitution elasticity center on .15 -- see below.

$$\Delta \log c_t = \Delta a_t^c + e\, \Delta \log w_t,$$

where a_t^c represents the mea" across individuals of a taste shifter in consumption, and e (which is approximately equal to η - 6) measures the degree of complementarily or substitutability between leisure and consumption, holding constant the marginal utility of wealth. At a minimum, the goodness-of-fit of this equation provides a" indication of the magnitude of aggregate changes in the marginal utility of wealth.

c. Individual-Specific Comp ents of Wage and Hours Variation

In addition Co its implications for the age and time effects in microeconometric studies of labor supply, the lifecycle model offers a potential explanation for individual and period-specific hours variation Specifically. suppose that individual log wages are determined by a" equation of the form:

$$(8) \qquad \log w_{it} = \omega_i + v_t + u_{it},$$

where ω_i is a person-specific constant, v_t is an aggregate effect, and u_{it} is a person and time-specific effect. Then equation (5) implies

$$(9) \qquad \Delta \log h_{it} - \Delta \log h_t = (\Delta a_{it} - \Delta a_t) + \eta u_{it} + (\phi_{it} - \phi_t).$$

The person-specific component of hours variation in period t consists of person-specific taste variation, a person-specific intertemporal substitution effect ηu_{it}, and the difference between the person-specific forecast error in log λ_{it} end the average forecast error over the entire sample.

The person-specific component of year-to-year changes in labor supply is large. For example, Altonji and Paxson (1985) estimate that the cross-sectional standard deviation of the change in log annual hours between

consecutive years is 0.35 for me" age 18 to 60 in the Panel Study of Income Dynamics (PSID). Using data constructed from survey information gathered every 4 months in the Survey of Income and Program Participation (SIPP), I estimate that the standard deviation of the change in log annual hours for men age 22-59 who worked in 1984 end 1985 is 0.54 (Card(1990)). some of this variation is clearly attributable to measurement error. Evidence reported by Duncan and Hill (1985) suggests that the signal-to-noise ratio in the measured change in lag annual hours in the PSID is 1.22. [25] Applying this correction factor. the estimated standard deviation of true hours changes for continuously employed me" in the PSID is 0.26, and eve" larger for me" in the SIPP panel. [26]

Nevertheless. virtually "one of this variation is explained by the person-specific intertemporal substitution effect. Altonji (1986, Tables 1,2) reports measures of R^2 for labor supply equations like (9) that instrument the individual-specific component of wage variation and treat the other two components (person-specific changes in taste and person-specific updating in the marginal utility of income) as residuals. The proportions of explained variance are essentially 0.

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[25] This estimate is based on a sample of individuals working for a single employer over two years. and is surely an upper bound on the signal/noise ratio.

[26] I suspect that a retrospective survey on annual hours in the previous year probably understates the true variation in average hours per week. since many individuals with substantial within-year variation in hours per week are likely to report a simple "umber like "40 hours per week". This is especially problematic in the CPS survey, because interviewers are instructed to gather modal (rather than average) hours per week from such individuals. However. I have bee" unable to ascertain if the more frequent interview schedule in the SIPP accounts for the higher variation in annual hours changes.

One reason for this low proportion of explained variance is the very small magnitude of the estimated intertemporal substitution elasticities that typically emerge from microeconometric studies.[27] MaCurdy's (1981) estimates from the PSID range from 0.10 to 0.45. Altonji's (1986) estimates, also based on PSID data, range from 0 to 0.5, with the more precise estimates clustered near the bottom of this interval. A similar range of estimates emerges from other studies of the PSID, including Ham (1986), and from the detailed study of cohort-level data from the British Family Expenditure Survey by Browning. Deaton, and Irish (1985). Taken together, the literature suggests that the elasticity of intertemporal substitution is surely no higher than 0.5, and probably no higher than 0.20. Given such small elasticities, the component of individual hours changes attributable to intertemporal substitution effects is tiny.

This leads to the question of whether there is <u>any</u> explanation for individual-specific hours variation. One source of systematic hours variation that is described by the labor supply model, but ignored in most studies, is idiosyncratic variation in the marginal utility of wealth. Some of this is potentially explainable by observed wage changes, particularly if person-specific wage innovations are highly persistent. To collect some evidence on the persistence of idiosyncratic wage shocks. I fit a very simple version of the components-of-variance model (8) to data on log wages for men in the PSID. Specifically. the model assumes that the measured log wage of individual i in period t is given by:

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[27] One exception is MaCurdy's (1983) study using a sample of males in the control group of the Denver Income Maintenance Experiment. MaCurdy does not parameterize preferences in such a way as to imply a constant intertemporal substitution elasticity. However. his estimates imply that the intertemporal elasticity is high: in the neighborhood of 2.0.

$$(10) \qquad \log w_{it} = \omega_i + v_t + u_{it} + \mu_{it},$$

where $u_{it} = \alpha u_{it-1} + \zeta_{it}$,

$$\text{var}(\zeta_{it}) = \sigma_t^2, \quad \text{cov}(\zeta i_s, \zeta_{it}) = 0, \quad t \neq s,$$

$$\text{var}(\omega_i) = \sigma_\omega^2, \quad \text{var}(\mu_{it}) = \sigma_\mu^2,$$

$$\text{cov}(\zeta_{it}, \omega_i) = \text{cov}(\zeta_{it}, \mu_{it}) = \text{cov}(\omega_i, \mu_{it}) = 0.$$

In this model the person- and period-specific wage shock consists of two components: a first-order serially correlated component with a time-varying variance (u_{it}), and a serially uncorrelated component (μ_{it}). One interpretation of the latter is as a white noise survey measurement error. However, this is indistinguishable from a "purely transitory" wage component. I have fit this model (using minimum distance techniques) to the covariance matrix of individual wage data for 1374 men who worked in each year between 1969 and 1979.[28] For convenience in estimation I have used the wage data for 1971-78 only.

The covariances of the wage data are presented in Table 1, together with their estimated standard errors and the average autocovariances at each lag. The autocorrelations decline from 0.78 (et lag 1) Co 0.59 (at lag 7). There is some evidence of nonstationarity in the data, with the variances and covariances rising in the last years of the panel. The sample excludes 105 individuals who otherwise meet the data requirements but who are eliminated by virtue of reporting an hourly wage less than $0.75 or greater than $100 (in 1967$) in one or more years. When these

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[28] Specifically. I estimated the vector of parameters β by minimizing $(\mathbf{m} - \mathbf{f}(\beta))'V^{-1}(\mathbf{m} - \mathbf{f}(\beta))$, where $\mathbf{m}$ is the vector of 36 second moments of the wage data, $\mathbf{f}(\beta)$ is the vector of fitted moments, and V is the estimated variance matrix of the second moments.

individuals are included. the variances and covariances are 25 percent larger but the autocorrelations are very similar.

This very simple model fits the wage data surprisingly well. The overall goodness-of-fit statistic is 35.31, which has a probability value of 8 percent. The parameter estimates and their implications are summarized in Table 2.[29] One half of the cross-sectional variance in wages is attributed to permanent person-effects. Another 15 percent is attributable to the pure white noise component. This variance share is actually much lower than the share of measurement error reported in the PSID validation study (Bound et.al (1989), Table 2), suggesting that all of μ_{it} can easily be attributed co measurement error. The remaining component of variance is highly persistent: the estimated AR(1) coefficient α is 0.89.

To see the implications of this persistence. consider the effect of a unit innovation in the person-specific wage component on the simple discounted average of expected future wages

$$(1 - \beta) \sum_{j=0}^{\infty} \beta^j E_t \log w_{it+j}$$

In the first case. suppose that μ_{it} is all measurement error. so that a unit innovation in wages is purely an innovation in u_{it}. Then, assuming $\beta=.9$ (i.e., a discount race of 11.1 percent), the effect on the discounted average of expected future wages is $(1 - \beta)/(1 - \alpha\beta) = 0.494$. On the ocher hand, suppose that there is no measurement error in wages. The" a unit innovation in the wage shock implies a 0.69 innovation in u_{it} and a 0.31

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[29] There are a total of 12 parameters in the model, including the 8 period-specific variances σ_t^2 end the variance of the pre-sample shock u_{i0}.

innovation in μ_{it}.[30] In this case. the discounted average of expected future log wages rises by 0.34.

The results of this exercise suggest that a typical person-specific wage innovation results in a significant revision to lifetime wealth. Of course, it is possible that individuals have better information with which to forecast future wages than is available to an outside data analyst. In this case. wage innovations in the statistical model (LO) do not necessarily represent new information. Clearly, we need much further evidence before we can use the lifecycle model to model the wealth effects of person-specific wage shocks.

One possible approach is to combine consumption and hours information to obtain direct estimates of λ_{it}, and then to consider projections of the forecast errors in log λ_{it} on wages and other information.[31] To see how this might work. write the log-linear version of the intertemporal consumption function implied by the first-order conditions (3a) and (3b) as:

$$\log c_{it} = e \log w_{it} + f \log \lambda_{it}.$$

(For simplicity I will ignore any components of taste variation, although these can be handled). This consumption function can be combined with the labor supply function (4) to give:

$$\log h_{it} = (\eta + \delta e/f) \log w_{it} + \delta/f \log c_{it}. \quad (11)$$

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[30] This uses the linear projection $E(a|a+b) = (a+b) \cdot var(a)/var(a+b)$.

[31] This approach follows up on MaCurdy' (1983) method for estimating the parameters of the lifecycle model. MaCurdy's procedure is used by Blundell (1990).

This equation is the within-period optimality condition implied by setting the marginal rate of substitution between goods and leisure equal to the wage.[32] An instrumental variables procedure can be applied to (11) to estimate the coefficients $(\eta \cdot be/f)$ and δ/f. Similarly, the intertemporal labor supply elasticity η can be estimated by conventional means, for example by applying instrumental variables to (8). Then, using the (approximate.) restriction that $e = \eta \cdot \delta$, it is possible to recover estimates of the coefficients e and f. (Alternatively, one can estimate the coefficient e in the intertemporal consumption function directly .. see Altonji (1986) for example). Finally, these can be used to form an estimate of log λ_{it} from the observed consumption and wage data for each person.

Given estimates of log λ_{it} it should be possible to estimate the relation between the marginal utility of income end observable information. such as current assets and current and lagged wages. One could then test a specific modal for log λ_{it}, such as the one implied by the Lucas-Rapping labor supply function, or the one implied by perfect foresight. It would also be useful to estimate components-of-variance models for the change in the marginal utility of income. A recent paper by Altug and Miller (1990) shows that the assumption of complete contingent markets imposes a simple factor structure on λ_{it}: $\lambda_{it} = \lambda_i \cdot \lambda_t$. If this is correct, the idiosyncratic component of the estimated change in log λ_{it} should be orthogonal to individual-specific information. controlling for a

[32] Notice that if one maintains the assumption e-0 (i.e., that wages have no effect on consumption, holding constant λ), then one can obtain estimates of the intertemporal substitution elasticity from cross-sectional data| This procedure is used by Altonji (1986). and seems to give estimates of η about the same size as those obtained by estimating (8).

homogeneous time effect. Altug and Miller's results suggest that this set of restrictions may be acceptable.

A major limitation to this line of research is the absence of panel data sets with information o" consumption expenditures. The leading panel data source, the PSID, only contains information on food expenditures. Some progress may be possible using the cohort level data in the British Family Expenditure Survey, although the labor supply information contained in this survey is limited to weekly hours.

d. Other Sources of Variation in Individual Labor Supply

Although careful ❑ odelling of wealth effects may go some way toward improving our understanding of the determinants of individual labor supply, I am not optimistic that a conventional lifecycle model can ever explain more than a tiny fraction of the year-to-year variation in the data.[33] One may be tempted to attribute the unexplained changes to tastes or measurement error. There is a graving body of evidence, however, which suggests that idiosyncratic changes in labor supply are systematically related to conditions on the demand side of the labor market. There are two complementary explanations for this link. On one hand, individuals may be unable to sell all their offered labor supply. On the other, some form of fixed costs may enter into either the supply or demand sides of the labor market.

Ashenfelrer and Ham (1979) and Ham (1986) present models of intertemporal labor supply which assume that reported unemployment contains

[33] For example, Altonji's (1986) use of observed food consumption as a control for the marginal utility of income results in only a small increase in the explanatory power of his fitted labor supply equations.

information on hours constraints faced by workers. Specifically. these authors assume that desired hours of work are described by an equation such as (4). In the presence of labor market disequilibrium. actual hours sold may be lower. Following Ashenfelter (1978) suppose that a fraction θ of reported weeks of unemployment represent weeks in which an individual was unable to sell his or her labor. This leads to a specification of the lifecycle labor supply function that includes measured unemployment (or its first difference) on the right-hand-side. with a coefficient of 8. Estimates of this coefficient reported in Ashenfelter and Ham (1979) and "am (1986) are positive and significant. Furthermore, the inclusion of measured unemployment leads to a significant increase in the explanatory power of the labor supply equation.

The interpretation of such an augmented labor supply function is an issue of considerable dispute. Heckman and MaCurdy (1989). following Lucas and Rapping (1970), argue that measured unemployment is simply another component of leisure. Because of the hours constraint, the sum of leisure and unemployment is necessarily negatively correlated with hours of work. According to Heckman and MaCurdy then. individuals with longer hours of unemployment are simply those who are consuming more leisure.

Evidence presented by Ham (1986) sheds some interesting light on the interpretation of reported unemployment, and also on the underlying question of whet causes individual hours of work to vary from year to year. To see the nature of this evidence, consider the following (simplified) intertemporal labor supply function:

$$(12) \qquad \Delta \log h_{it} = \eta \, \Delta \log w_{it} + \ell \, \Delta D_{it} + 6 \, \phi_{it}$$

where D_{it} is a vector of demand conditions in an individual's local labor market, industry, and/or occupation. There is no mechanical connection between the measurement of hit and the measurement of D_{it}. If the labor supply model is correctly specified, however, the" P-O, since market-level information should be irrelevant to individual hours decisions, controlling for individual-specific wages. Although he does not report direct estimates of ℓ in his 1986 paper, Ham's results using ΔD_{it} as instrumental variables for individual unemployment indicate that ℓ is far different from zero. A" earlier unpublished version of the paper (Ham (1984)) presents direct tests for the exclusion of industry. occupation, and local unemployment rates from a" individual labor supply equation. The test statistics are highly significant, indicating a" explicit role far demand-side variables in the determination of individual labor supply. When Ham uses the demand-side variables to instrument reported unemployment in the labor supply equation, he continues to find evidence of a negative and significant effect of unemployment on hours of work. This is evidence against a strict labor supply model, and in favor of a model in which reported unemployment conveys information about the demand conditions facing a" individual worker.

One need not appeal to Keynesian-style labor market constraints to rationalize Ham's findings. however. An alternative explanation is that. labor supply decisions are made at a higher frequency rime unit than the year (for example the week), and that there are significant fixed costs on either the worker's side or the employer's side of the labor market. A model along the latter lines is presented in Rosen (1986) and Card

(1990).[34] In this model, effective labor input from a pool of N workers is Ng(h), where g is an S-shaped function of hours worked per person. The optimal employment policy of a firm with this technology consists of a two-part rule: if product demand is sufficiently low, lay off a fraction of the labor force and employ the remainder at some minimum threshold level of hours. If product demand is sufficiently high, employ all available workers at hours above the threshold.

The implications of this firm-level behavior for individual labor supply data are two-fold. First. some component of annual hours variation will occur at a fixed hourly wage rate. In particular. individuals working at firms with relatively low product demand will vary their number of weeks worked. but in each week of employment they will supply the same number of hours, and (presumably) earn the same weekly wage. For these individuals. hours of work will vary directly with measures of the firm's product demand. Second, weekly hours will be observed to fluctuate above a (parson-specific) minimum threshold. Evidence presented in Card (1990) indicates that the latter prediction is surprisingly close to the truth. In a sample of 2800 men observed working for the same employer over a two year period. reported hours per week in each of 8 interviews were observed to fall below 35 hours par week in only 11 percent of cases.[35]

A simple fixed cost model of this kind suggests that employer demand conditions should affect weeks of employment per year. Predictions on the connection between employer demand and hours per week depend on the assumed

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[34] A class of models with similar properties are analyzed in a macro context by Hansen (1985) and Rogerson (1988). In these papers, labor supply within tha week is assumed to be either 0 or 1.

[35] See Card (1990, Table 3).

form of employment contract. My paper (Card (1990)) presents a case in which. conditional on working, hours per week lie on a conventional supply schedule. I" this case, controlling for the wage, employer demand should have no effect on hours per week. Some simple evidence on this prediction is presented in Table 3, which shows the results of estimating the augmented labor supply function (12) a" three measures of labor supply: hours per week, weeks per quarter. and total quarterly hours.

The data summarized in Table 3 pertain to me" in the 1984 SIPP panel. The sample is restricted to individuals who are observed working for at most one employer over the 9 quarters of the available sample period. Demand-side conditions are measured by the logarithm of employment in the individual's one-digit industry. Thus, ΔD_{it} refers to the percentage change in employment in an individual's industry in the most recent quarter. The equations are estimated by instrumental variables. using as an instrument for wages the change in wages observed for the same person 4 quarters in the past or 4 quarters in the future. There is a small but highly significant seasonal correlation in individual wage changes that gives this instrumental variable its power.

The estimates suggest that measures of employment demand are significantly correlated with both hours per week and weeks per quarter. I" comparison, the estimated intertemporal substitution elasticities are small and relatively imprecise.[36] One could easily conclude from this evidence that changes in labor supply are directly connected to employer

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[36] OLS estimates of the equation result in negative and significant wage coefficients. presumably as a consequence of measurement error in average hourly earnings. Further results are reported in Card (1990).

demand conditions, and that wages play little or no role in the short-run labor-leisure decision.

The relatively weak connection between hours per week and wages illustrated in columns 1 and 4 of Table 3 may seem puzzling. give" that the Fair Labor Standards Act mandates overtime payments for individuals in many occupations who work over 40 hours per week. Some additional evidence o" the relation between weekly hours and wages is provided by data in the May 1985 CPS. This survey gathered information on usual hours per week, actual hours worked in the previous week, and whether or not the individual received any overtime payments. The responses suggest that there is substantial variation in actual weekly hours around "usual" weekly hours: 13 percent of me" indicate that they worked lass than their usual hours. while another 19 percent indicate that they worked more.[37] Individuals in the letter group report 10 extra hours per week on average, bringing their weekly total to 51 hours. However, only 47 percent of these men report receiving any additional overtime compensation. For the majority. weekly hours are higher then usual but weekly earnings are fixed.[38]

Table 4 provides more detailed information on a very narrow subset of individuals •• those who usually work 35-40 hours per week and who report 41 or more hours in the survey week.[39] Sixty-two percent of all workers normally work 35-40 hours per week. Of these, 13.5 percent worked 41 or

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[37] These statistics pertain to me" sge 16-64 who hold only one job and who are not self-employed. Variation in weekly hours among the excluded group is even larger.

[38] Unfortunately, the survey does not ask about reduced compensation for individuals who worked less than usual hours.

[39] In s" effort to obtain a reasonably large sample, this table includes both ma" and women.

more hours in the survey week, and are summarized in the Table. The fraction receiving overtime compensation among this group is 59 percent. Interestingly. however, extra hours worked are actually slightly higher for the group with no overtime pay.

These data suggest that even within the week, a simple labor supply model is inadequate for a large fraction of the population. Many individuals appear to be working extra hours for no extra pay. "he" this behavior is added to the phenomenon of weekly layoffs, it becomes clear how a simple model of labor supply can easily fail to explain movements in annual hours.

Further work is obviously needed to isolate the systematic components of individual labor supply, and to describe the links between employer demand and employee hours choices. While such work falls outside the "arrow realm of a conventional lifecycle model. it seems to me that further understanding of individual hours outcomes will require a broader perspective than the standard model can provide. As it stands. the lifecycle model provides essentially no insight into the year-to-year variation in individual hours.

IV. Conclusions

I" principle. the lifecycle labor supply model offers a" explanation for the four main aspects of individual hours choices: mea" hours over the lifecycle; the age profile of hours; aggregate movements in hours; and individual-specific variation in hours around the lifecycle profile. All of these components are tied together by a combination of intertemporal substitution effects and wealth effect. In this paper I have tried to

gauge the success of lifecycle model in explaining the various dimensions of male labor supply. My assessment is hardly positive: the only real success for the model has come as a description of aggregate patterns in wages and hours during the post-1970 period. Even here, my suspicion is that a careful consideration of wealth effects will undermine the success of the model.

Much of the microeconometric research over the past two decades has concentrated on the magnitude of the intertemporal substitution effect, and in particular on modelling the intertemporal substitution effect of individual-specific wage variation. As Pencavel noted in his 1986 survey, the available evidence suggests that this effect is of second-order importance. My view is that a similar conclusion holds with respect to the intertemporal substitution effect in the age profile of hours. With respect to the permanent component of hours, there is much ambiguity in the literature. A fairly wide-spread belief among labor economists is that a permanent increase in wages leads to a reduction in hours. Using modern panel data it is surprisingly hard to verify this hypothesis, and in fact the preponderance of the evidence suggests to me a positive association between long-run wages and average hours.

Two major avenues for further work are suggested. one involves a detailed effort to estimate the wealth effects in intertemporal labor supply. Existing methods can be used to estimate the marginal utility of wealth, and test its properties. Progress in this direction will depend on the quality of available data linking individual consumption and hours choices. A second involves a re-evaluation of the premise that average hourly earnings are a "sufficient statistic" for current labor market

opportunities. A variety of models suggests that individual hours are influenced directly by employer-specific demand conditions. Limited empirical evidence confirms this suspicion. If true, our basic notions of labor supply, and in particular our notions about the degree of substitutability between current and future leisure. may be incomplete.

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Table 1
Auto-covariance Structure of Individual Wages
Continuously Employed Male Household Heads in PSID
1971-78

(estimated standard errors in parentheses)

	Auto-covariance of Wage in:							
	1971	1972	1973	1974	1975	1976	1977	1978
With Wage in:								
1971	0.243							
	(0.012)							
1972	0.196	0.252						
	(0.010)	(0.011)						
1973	0.185	0.199	0.251					
	(0.010)	(0.011)	(0.012)					
1974	0.180	0.188	0.194	0.240				
	(0.010)	(0.010)	(0.010)	(0.011)				
1975	0.173	0.187	0.187	0.193	0.253			
	(0.010)	(0.010)	(0.010)	(0.010)	(0.011)			
1976	0.168	0.178	0.180	0.188	0.206	0.271		
	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.012)		
1977	0.163	0.177	0.180	0.183	0.185	0.205	0.254	
	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.011)	
1978	0.151	0.166	0.164	0.170	0.180	0.186	0.205	0.293
	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)	(0.011)	(0.011)	(0.016)
Average Auto-covariance at Lag:								
	7	6	5	4	3	2	1	0
	0.151	0.165	0.169	0.175	0.182	0.188	0.200	0.257
	(0.010)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)

Note: Sample consists of 1374 male household heads from households with no change in head between 1969 and 1979, who earned positive labor earnings and worked positive hours in each year between 1969 and 1979, and whose hourly wage was between $0.75 and $100 (in constant 1967 dollars) in all years.

Table 2
Summary of Estimated Components-of-variance Wage Model

A. Parameter Estimates
(estimated standard errors in parentheses)

Parameter	Estimate
1. Variance of Permanent Effect (σ_ω^2)	0.124 (0.040)
2. Variance of Measurement Error/ Purely Transitory Component (σ_μ^2)	0.039 (0.003)
3. AR(1) Coefficient (α)	0.886 (0.077)
4. Average variance of Wage Innovations (Average of σ_t^2)	0.027
5. Goodness of Fit (24 degrees freedom)	35.314

Note: Model is fit by optimal minimum distance to the 36 wage covariances displayed in Table 1. The model is

$$\log w_{it} = \omega_i + u_{it} + \mu_{it}, \quad u_{it} = \alpha u_{it-1} + \zeta_{it},$$

with $\text{var}(\zeta_{it}) = \sigma_t^2$ (t=1,2,...8) and $\text{var}(u_{i0}) = \sigma_0^2$.

B. Implications of Estimates

1. Average Variance of Wages	0.249
2. Share Attributable to Permanent Effect	0.500
3. Share Attributable to Measurement Error/Purely Transitory Effect	0.157
4. Effect of Unit Wage Shock on Average Expected Future Wage:[a]	
(i) Assuming μ_{it} is measurement error	0.494
(ii) Assuming no measurement error	0.340

Note: 'Change in discounted average of expected future log wages, assuming a" infinite life and a .11 discount rate see text.

Table 3

Estimated Labor Supply Functions for
Quarterly Hours Outcomes: SIPP Sample of Hen

(standard errors in parentheses)

	Dependent Variable (All in First-Differences)					
	Log Hours/Wk	Log Wks	Log Total Hrs	Log Hours/Wk	Log Wks	Log Total Hrs
1. Log wage	0.10	0.05	0.16	0.10	0.05	0.14
	(0.14)	(0.13)	(0.22)	(0.14)	(0.13)	(0.22)
2. Industry Employment	--	--	--	0.21	0.24	0.46
				(0.06)	(0.06)	(0.10)
3. R-squared	0.00	0.00	0.00	0.00	0.00	0.00

Notes: Sample consists of 19566 observations on quarterly changes in labor supply of 4814 men age 16-64 with same employer over 9 quarters (1983-W to 1985-I) in 1983 SIPP panel. All equations are estimated in first-difference form, and include 9 unrestricted quarterly dummies as well as potential experience. Log wage is instrumented by the change in log wages of the same person 4 quarters in the past (or 4 quarters in the future. for observations from the first 3 quarters of the sample). The standard deviations of the dependent variables are: log hours per week -- 0.142; log weeks per quarter -- 0.147; log quarterly hours -- 0.234.

Table 4

Wages, Hours, and Overtime Premiums for Individuals Working 40 or More Hours: nay 1985 Current Population Survey

(standard errors in parentheses)

	Paid Overtime?	
	No	Yes
1. Number of Individuals	1651	2416
2. Average Hours Last Week	48.58 (0.18)	47.82 (0.13)
3. Usual Weekly Hours	39.78 (0.02)	39.86 (0.01)
4. Hours Paid Overtime	..	8.10 (0.16)
5. Percent Paid Time-and-a-half	—	92.34
6. Percent Female	41.67	32.37
7. Percent Paid by Hour	38.10	05.67
a. Average Hourly wage	10.65 (0.27)	a.97 (0.16)

Note: Sample consists of 4067 individuals age 16-64 in Hay 1985 CPS who reported usual weekly hours between 35 and 40 and who reported working 41 or more hours in the survey week. Dual-job holders and self-employed workers are excluded. In the May 1985 CPS 62.4 percent of all individuals report usual weekly hours between 35 and 40 (62.3 percent of me", 62.5 percent of women). Of these, 13.5 percent reported working 41 or more hours last week.

Figure 1

Lifecycle Wages for Six Cohorts
CPS Data 1976–88

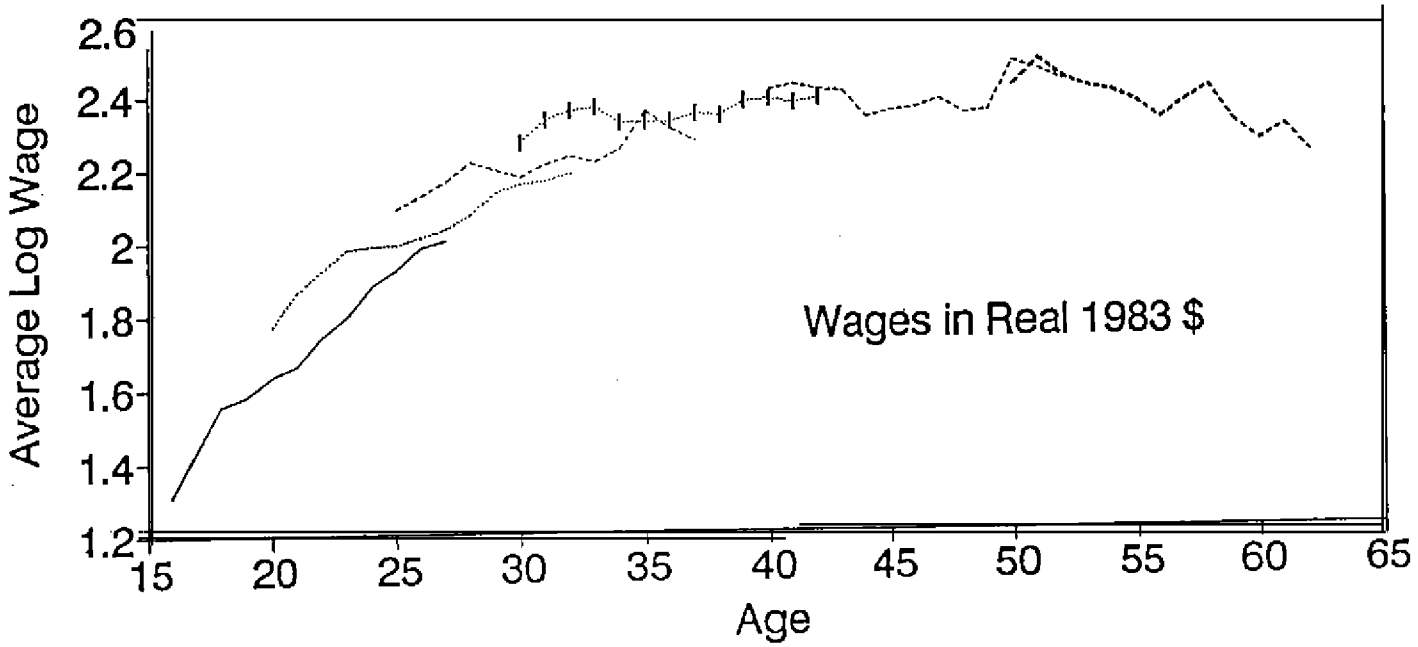


Figure 2

Lifecycle Hours for Six Cohorts
CPS Data 1976–88

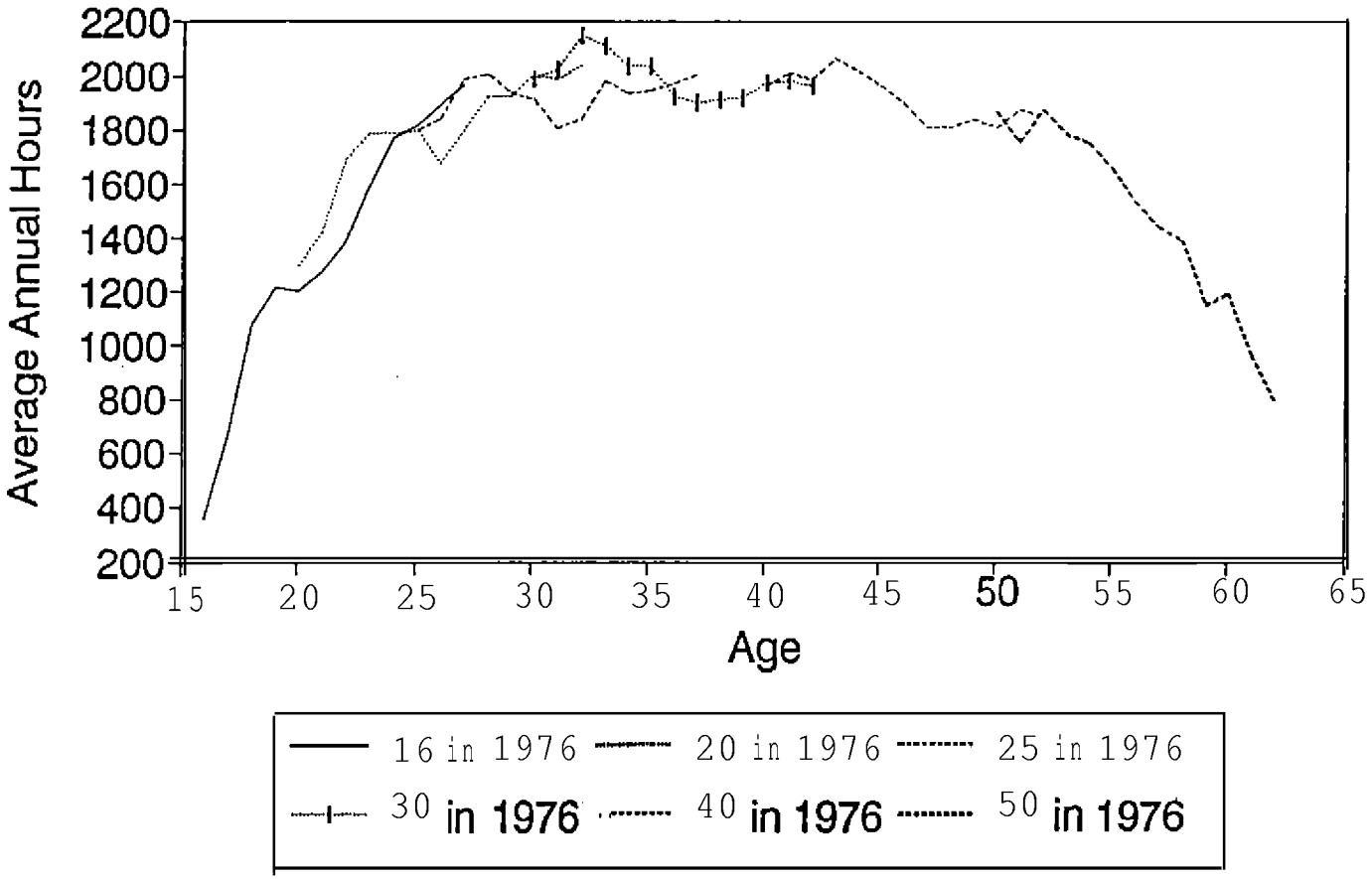


Figure 3

Employment and School Enrollmen
By Age: Men in CPS 1977-89

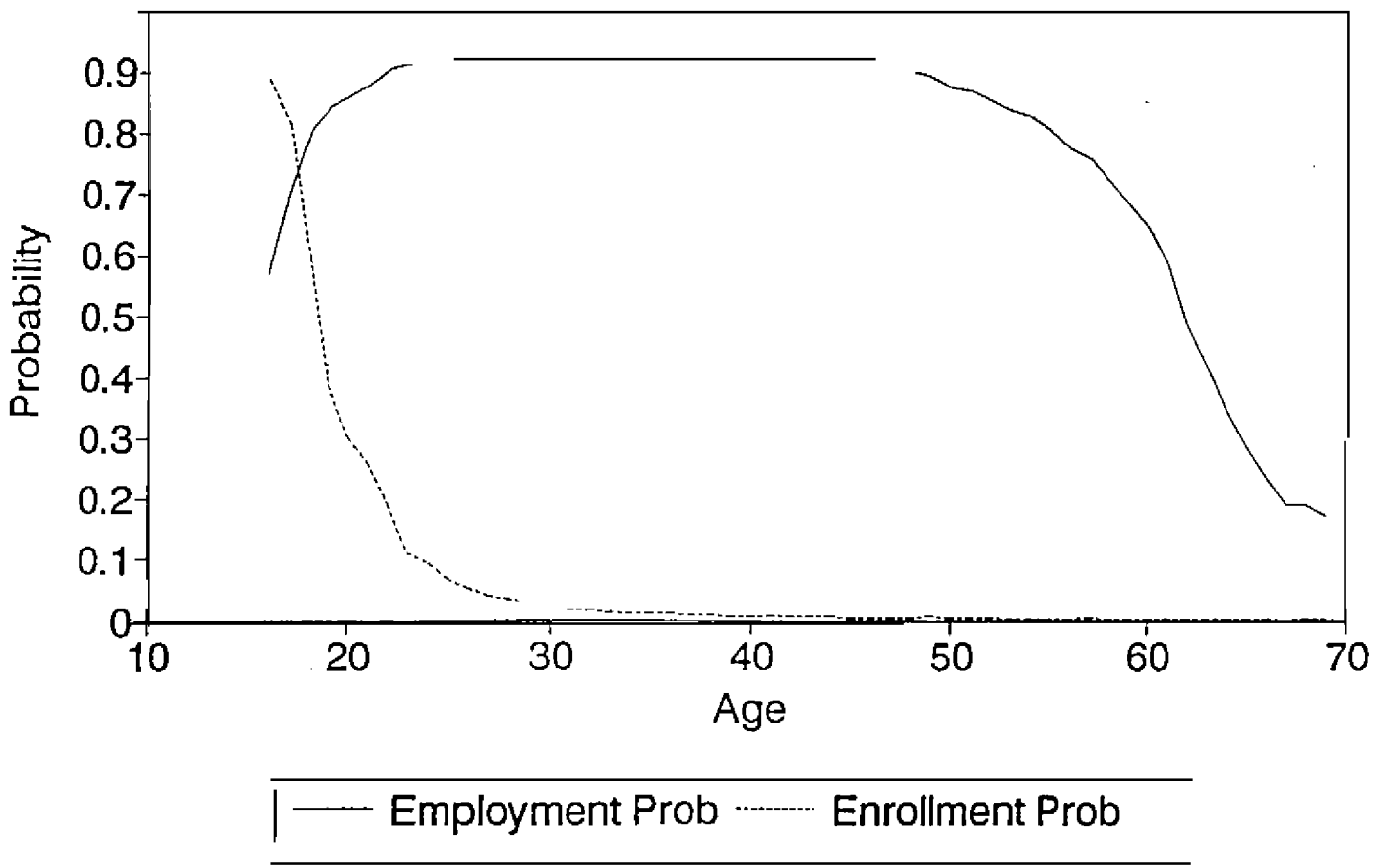


Figure 4
Lifecycle Profiles of Wages
Men Age 18-66 in CPS, 1977-89

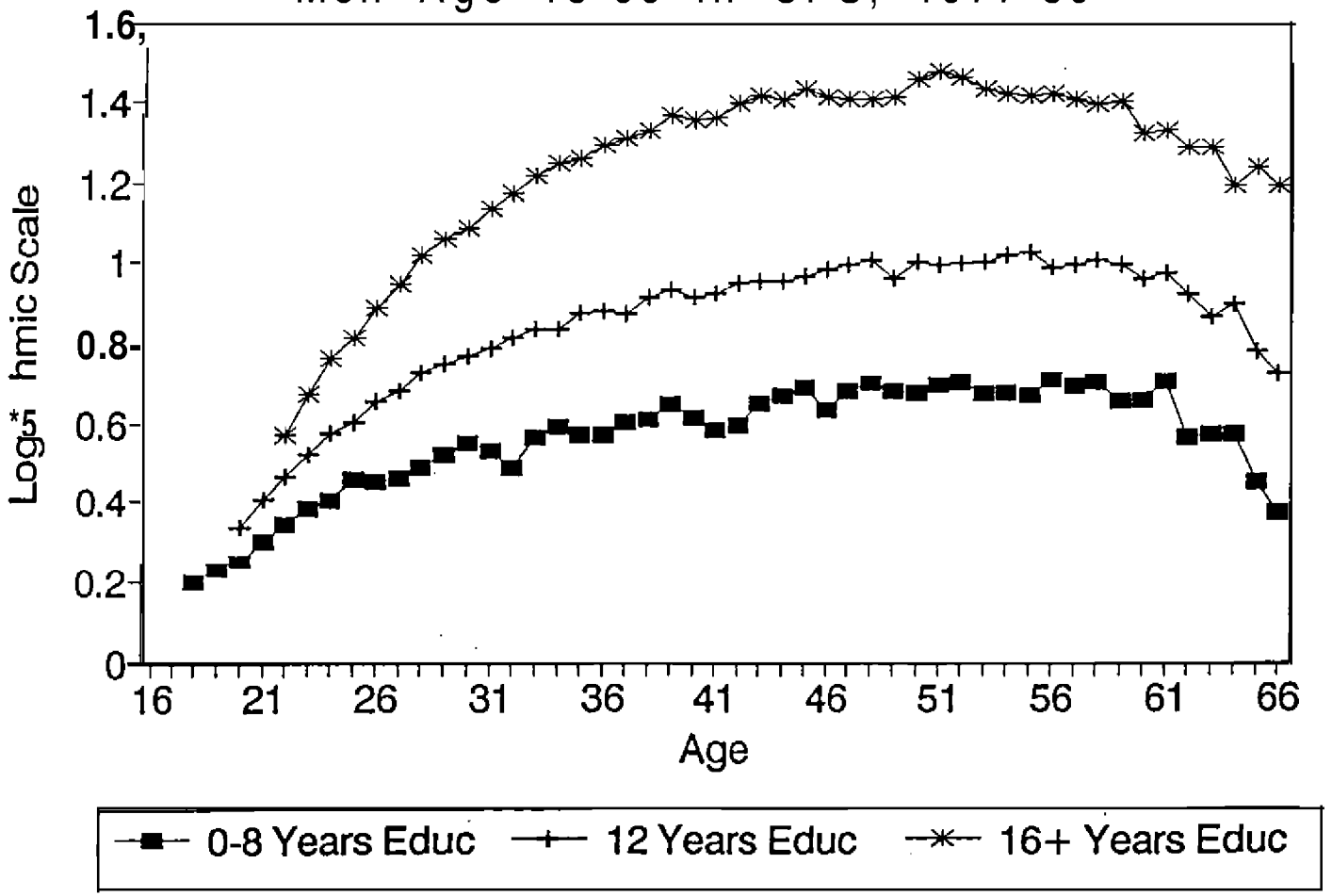


Figure 5

Lifecycle Profiles of Hours

Men Age 18-66 in CPS, 1977-89

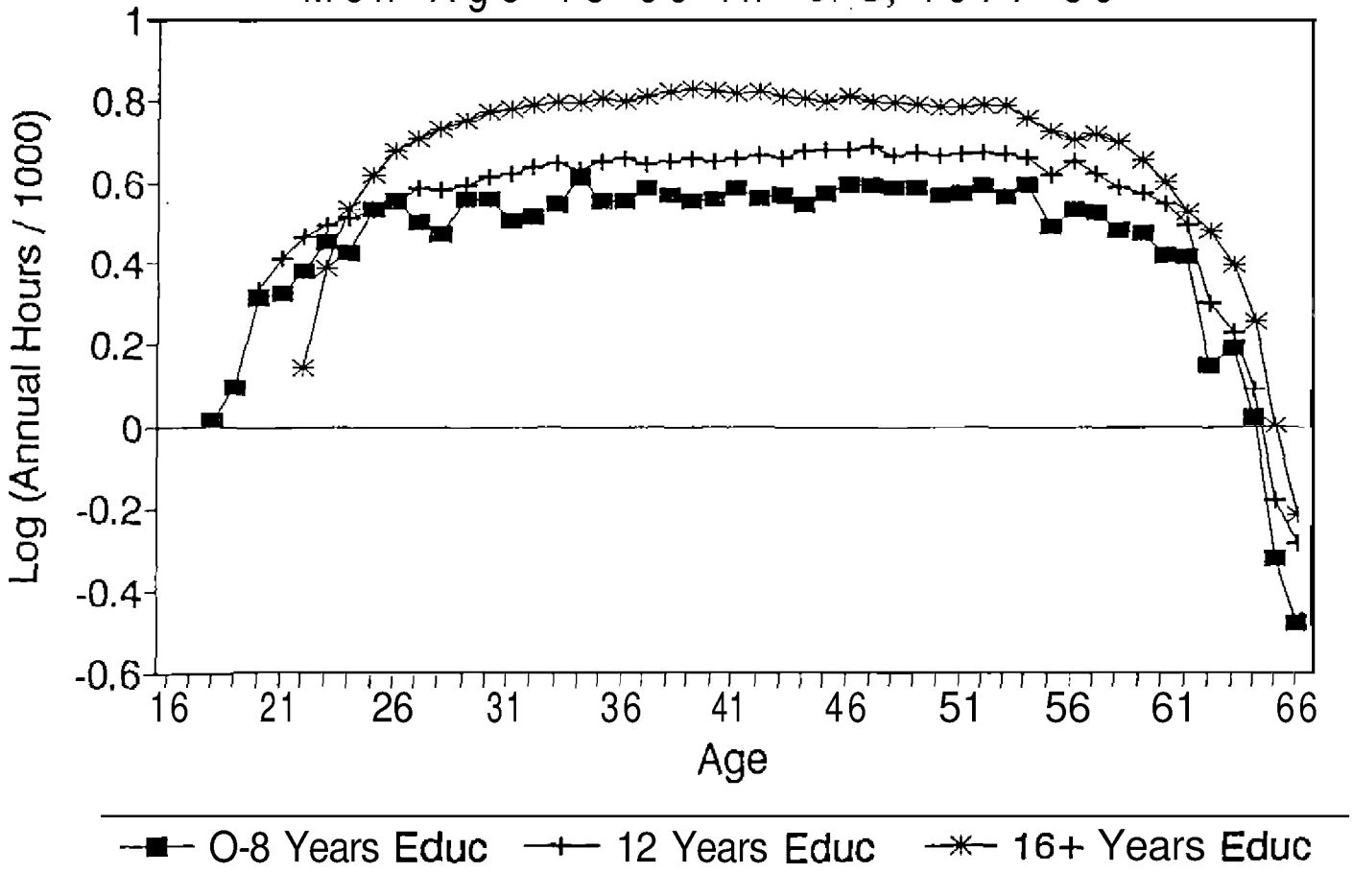
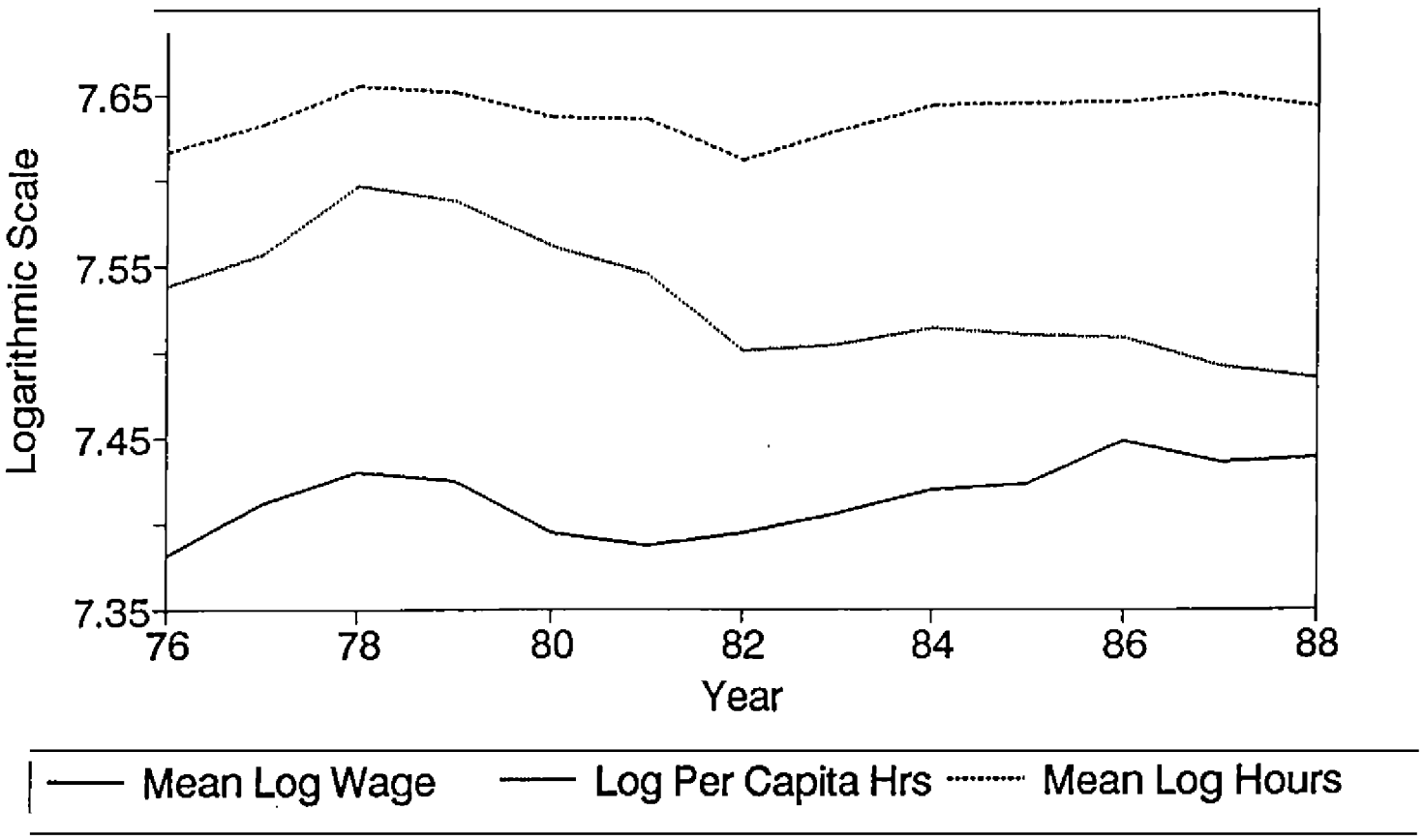


Figure 6

Aggregate Hours and Wages
Cohort Age 20–50 in 1976



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