Intertemporal Substitution in Labor Supply: Evidence from Micro Data

Author(s): Joseph G. Altonji

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# Intertemporal Substitution in Labor Supply: Evidence from Micro Data

Joseph G. Altonji

*Columbia University*

The sensitivity of the supply of labor to intertemporal variation in the wage is an important issue in macroeconomics, the analysis of social security and pensions, and the study of life-cycle patterns of work. This paper explores two approaches to the measurement of intertemporal substitution that have appeared in the literature. The first approach is to use consumption to control for wealth and unobserved expectations about future wages in the labor supply equation. The second approach is to estimate a first-difference equation for hours in which labor supply from the previous period serves as a control for wealth and wage expectations. The results indicate that the intertemporal substitution elasticity for married men is positive but small.

## I. Introduction

This paper examines the response of labor supply to variations in wage rates over time. Research on the intertemporal labor supply response deserves a high priority for several reasons. First, as Heckman and MaCurdy (1980) among others have emphasized, studies of the relationship between current labor supply and current wages and nonlabor income are difficult to interpret even as the response of individuals to permanent differences in wages and nonlabor income.

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Second, static models of labor supply cannot address issues that involve the response of labor supply to fluctuations over time in its price, which includes wages net of current taxes and the effect of working today on wages and transfers in the future. These issues include the role of intertemporal labor supply responses in cyclical fluctuations of employment and unemployment,[1] the effect of pensions and social security on hours worked and retirement,[2] and the link between the life-cycle profiles of wages and hours.

The pioneering studies of the intertemporal labor supply response include Mincer (1962), Lucas and Rapping's (1970) time-series analysis, Heckman's (1971) analysis of cross-section micro data, Ghez and Becker's (1975) and Smith's (1977) studies of cross-section data aggregated by age cohort, and the panel data analyses by Heckman and MaCurdy (1980) and MaCurdy (1981). For men, MaCurdy's estimates range from .10 to .45, Ghez and Becker's estimates range from −.068 to .44, and Smith obtains .322.

This paper presents new estimates of the intertemporal labor supply elasticity for married men.[3] The most important obstacle to implementation of intertemporal models of labor supply stems from the fact that, in principle, current labor supply depends on all past and expected future wage rates, while data on these variables are missing for most periods of the lifetime. The problem of missing wage data is compounded by measurement error in the wage and labor supply variables. (Estimates below suggest that response errors account for much of the variance across individuals in year-to-year changes in the standard wage measure from the University of Michigan Panel Study on Income Dynamics [PSID] data set, which is the data source for the present study.) On one hand, it is necessary to control for permanent differences in wages across individuals if the coefficient on the current wage is to represent the labor supply response to intertemporal variation in wages. On the other hand, the problem of measurement error is most serious in identifying the labor supply response to transitory wage movements. The approaches in this paper make use of various instrumental variables (IV) schemes to handle the measurement error problem and impose assumptions about preferences that restrict how wages in other periods are related to current labor supply and consumption. Much of the analysis involves use of data on family consumption behavior in the labor supply equation as a control for the effects of wealth and unobserved expectations about future wages

[1] See Lucas and Rapping's (1970) seminal paper and recent studies by Altonji and Ashenfelter (1980), Hall (1980), Altonji (1982), Andrews and Nickell (1982), Ashenfelter and Card (1982), Clark and Summers (1982), and Mankiw, Rotemberg, and Summers (1985).

[2] See the survey by Mitchell and Fields (1982).

[3] Killingsworth (1983) surveys the growing literature on life-cycle labor supply.

and real interest rates. This approach, which amounts to using a transformation of the familiar marginal rate of substitution condition for utility maximization as an estimating equation, was suggested by Heckman (1974*a*) and Metcalf (1974) and has been used recently by Altonji (1982) and MaCurdy (1983). This equation identifies the intertemporal labor supply response if preferences are separable between consumption and leisure. Since separability of preferences may be questioned, some evidence on its validity and on possible biases in the estimates is presented. In addition, the paper provides a reanalysis of MaCurdy's (1981) first-differenced labor supply equation, in which labor supply from the previous period serves as a control for wealth and wage expectations. The model is reestimated using a second independent wage measure as an instrumental variable for the wage change as well as using MaCurdy's human capital type wage equation.

The paper proceeds as follows. Section II provides a brief review of the theory of life-cycle labor supply and presents the methods used to estimate the intertemporal substitution parameter. Section III discusses the data, and Section IV presents the empirical analysis.

## II. The Life-Cycle Model of Labor Supply

This section begins with a review of the life-cycle theory of labor supply under uncertainty when preferences are additively separable over time.[4] At age $t$ the individual seeks to maximize utility over his remaining lifetime. The objective function is

$$\underline{U}_{it} = \sum_{k=t}^{T} \phi^k U(N_{ik}, C_{ik}; V_{nik}, V_{cik}), \tag{1}$$

where $\underline{U}_{it}$ is utility of individual $i$ over his remaining lifetime and is the discounted sum of the value for $U(\ )$ in the remaining periods, $\phi$ is a time discount factor, $N_{ik}$ is labor supply at age $k$, $C_{ik}$ is consumption of goods at age $k$, $V_{nik}$ and $V_{cik}$ are person-specific determinants of preferences in each period, and $T$ is the end of the planning horizon.

The budget and time constraints are

$$0 = A_{it} + \sum_{k=t}^{T} R_{tk}W_{ik}N_{ik} - R_{tk}C_{ik} \tag{2.1}$$

and

$$0 \leq N_{ik} \leq \bar{N}, \tag{2.2}$$

[4] See Killingsworth (1983) and MaCurdy (1983) for a detailed presentation. The recent paper by Browning, Deaton, and Irish (1983) presents the life-cycle model using duality theory.

where $A_{it}$ is the individual's wealth at age $t$, $R_{tk}$ is the real $(k - t)$-period interest rate factor, $W_{ik}$ is the real wage at age $k$, and $\bar{N}$ is the maximum hours constraint. The individual subscript $i$ is suppressed below unless it is needed for clarity.

The individual chooses $C_t$ and $N_t$ to maximize the expected value of (1) given the constraints (2.1) and (2.2) and given that he can choose optimal values of $C$ and $N$ in subsequent periods as new information about preferences, wages, and real interest rates becomes available. Wages and real interest rates are assumed not to depend on consumption and labor supply choices, which rules out from the problem the most important forms of human capital investment, credit rationing, progressive taxation, and overtime premiums. (More will be said about this later.)[5] On the assumption of an interior solution, the first-order conditions for utility maximization include (2.1) and the following:

$$\phi^t \frac{dU(N_t, C_t; V_{nt}, V_{ct})}{dN_t} = -W_t\lambda_t, \tag{3.1}$$

$$\phi^t \frac{dU(N_t, C_t; V_{nt}, V_{ct})}{dC_t} = \lambda_t, \tag{3.2}$$

$$\lambda_t = E_t\left(\frac{\lambda_{t+1}}{R_{t,t+1}}\right), \tag{3.3}$$

where $E_t$ is the expectations operator conditional on the information at age $t$, $d$ is the partial derivative operator, and $\lambda_t$ is the (expected value of) marginal utility of period $t$ income or wealth; $\lambda_t = E_t(d\underline{U}_t/dA_t)$.

Since $1/R_{t,t+1}$ is the effect on $A_{t+1}$ of saving an additional unit of $A_t$, the condition (3.3) for $\lambda_t$ simply states that the expected gain from an extra unit of wealth in period $t + 1$ must be equal to its cost in terms of utility in period $t$.[6]

After backdating (3.3) to obtain

$$\lambda_{t-1} = E_{t-1}\left(\frac{\lambda_t}{R_{t-1,t}}\right), \tag{4}$$

it is easy to show that $\lambda_t$ is related to $\lambda_{t-1}$ through the equation

$$\frac{\lambda_t}{R_{t-1,t}} = \lambda_{t-1} + e_{\lambda t}, \tag{5}$$

where $e_{\lambda t}$ is the expectation error relating $\lambda_t/R_{t-1,t}$ to $E_{t-1}(\lambda_t/R_{t-1,t})$. Since consumers choose $\lambda_{t-1}$ to satisfy (4) given the information available at age $t - 1$, $e_{\lambda t}$ will be uncorrelated with all information available

[5] MaCurdy (1983) shows how to incorporate some of these features. Abowd and Card (1983) incorporate an implicit wage contract in a perfect-foresight version of the life-cycle model.

[6] See Breeden (1980) for a derivation.

to the consumer at $t - 1$ if expectations are rational. This econometric implication of rational expectations has been exploited in a number of studies and is utilized in some of the empirical work below.[7]

Equations (3.1) and (3.2) demonstrate that, if preferences are additively separable over time, $N_t$ and $C_t$ depend on past wage rates and labor supply and consumption decisions and the distribution of future wages only through $\lambda_t$. These equations implicitly define the "$\lambda$-constant" demands for $C_t$ and $N_t$ as functions of the current wage $W_t$ and $\lambda_t$. If $\lambda_t$ were observed, all of the preference parameters could be estimated from (3.1) and (3.2). Of course, $\lambda_t$ is unknown, but one may eliminate it by working with combinations of the first-order conditions. Taking the ratio of (3.1) and (3.2) yields the marginal rate of substitution condition:

$$\frac{dU(N_t, C_t; V_{nt}, V_{ct})/dN_t}{dU(N_t, C_t; V_{nt}, V_{ct})/dC_t} = W_t. \tag{6}$$

The intertemporal optimality condition may be expressed in terms of the marginal utility of consumption and labor supply by combining (5) with (3.1) and (3.2), yielding

$$\begin{aligned} &(W_t R_{t-1,t})^{-1}\phi^t \frac{dU(N_t, C_t; V_{nt}, V_{ct})}{dN_t} \\ &= (W_{t-1})^{-1}\phi^{t-1}\frac{dU(N_{t-1}, C_{t-1}; V_{nt-1}, V_{ct-1})}{dN_{t-1}} + e_{\lambda t}, \end{aligned} \tag{7.1}$$

$$\begin{aligned} &(R_{t-1,t})^{-1}\phi^t \frac{dU(N_t, C_t; V_{nt}, V_{ct})}{dC_t} \\ &= \phi^{t-1}\frac{dU(N_{t-1}, C_{t-1}; V_{nt-1}, V_{ct-1})}{dC_{t-1}} + e_{\lambda t}. \end{aligned} \tag{7.2}$$

In principle, one may work with combinations of the first-order conditions to estimate the preference parameters. In practice, this may be done only if preferences are such that $N_t$, $C_t$, $V_{nt}$, and $V_{ct}$ enter the first-order conditions in a tractable manner given that much of the variance in $V_{nt}$ and $V_{ct}$ is unobserved and $N_t$ and $C_t$ may be measured with error.

A convenient preference specification that has been used by Heckman and MaCurdy (1980), MaCurdy (1981), and Altonji (1982), among others, is

$$\underline{U}_t = \sum_{k=t}^{T} \phi^k[(1 + B_c)^{-1}B_c V_{ck}(C_k)^{1+(1/B_c)} - V_{nk}(N_k)^{1+(1/B_n)}], \tag{8}$$

[7] Hall (1978) is the first to apply it to empirical analysis of consumption behavior over time. See also Hansen and Singleton (1983).

where the notation is the same as in (1).[8] Strict concavity of the utility function requires that $B_c < 0, B_n > 0$. For this utility function the first-order conditions (3.1) and (3.2) may be rearranged into the following loglinear equations for $N_t$ and $C_t$:

$$n_t = \text{constant} + B_n[w_t + \ln \lambda_t - t \ln(\phi) - v_{nt}], \tag{9.1}$$

$$c_t = \text{constant} + B_c[\ln \lambda_t - t \ln(\phi) - v_{ct}], \tag{9.2}$$

where $n_t$, $c_t$, $w_t$, $v_{ct}$, and $v_{nt}$ are the natural logs of $N_t$, $C_t$, $W_t$, $V_{ct}$, and $V_{nt}$ and ln is the natural log operator. The parameter $B_n$ is the elasticity of labor supply with respect to the current wage, with the marginal utility of income held constant. Since $\ln \lambda_t$ changes only when new information becomes available, $B_n$ is the response to an anticipated change in $w_t$, with expectations about wages in other periods also held constant. It is approximately equal to the response to an unanticipated temporary disturbance in the wage since a temporary change has little effect on $\ln \lambda_t$. Note that $w_t$ does not appear in (9.2) because (8) is additively separable between $C_t$ and $N_t$.

Accompanying (9.1) and (9.2) are equations characterizing the evolution of $\lambda_t$ over time. Since $\lambda_t$ enters (9.1) and (9.2) in log form, one may not use (5) directly. Taking logs of both sides of (5), performing a Taylor expansion of $\ln(\lambda_{t-1} + e_{\lambda t})$ around $e_{\lambda t} = 0$, and ignoring higher-order terms leads to

$$\ln \lambda_t = \ln \lambda_{t-1} + \ln R_{t-1,t} + u_t, \tag{10.1}$$

where $u_t$ is equal to $(1/\lambda_t)e_{\lambda t}$. The adequacy of the approximation depends on the variance in surprises about future wages, real interest rates, and preferences. *Assuming* that the approximation is adequate, then the log transformations of (7.1) and (7.2) for the preference specification (8) may be rearranged as

$$\begin{aligned} n_t - n_{t-1} = {} & B_n(w_t - w_{t-1}) - B_n(\ln \phi - \ln R_{t-1,t}) \\ & + B_n u_t - B_n(v_{nt} - v_{nt-1}), \end{aligned} \tag{11.1}$$

$$c_t - c_{t-1} = B_c u_t - B_c(\ln \phi - \ln R_{t-1,t}) - B_c(v_{ct} - v_{ct-1}). \tag{11.2}$$

Note that one may regard (11.1) as the result of using lagged labor supply to proxy $\ln \lambda_t$ in (9.1) since the equation is derived by using the relationship between $n_{t-1}$ and $\ln \lambda_{t-1}$ from the lag of (9.1) and the link between $\ln \lambda_{t-1}$ and $\ln \lambda_t$ from (10.1).

[8] Use of labor supply rather than leisure as the argument of the utility function rules out a corner solution with zero hours of market work. This is not an important restriction given that the sample consists of married men. Note that one may add an additively separable argument in the wife's labor supply to the utility function without altering the analysis for men. This was done in earlier drafts of this paper. Complications arise in the nonseparable case. The empirical analysis of the labor supply of married women is left for a separate paper.

Finally, (10.2) below expresses the dependence of $\ln \lambda_t$ on wealth, current and future preferences, current and future wages, and real interest rates:

$$\ln \lambda_t = \ln \lambda(A_t, w_t, w_{t+1}, \dots, w_T, R_{t,t+1}, \dots, R_{t,T}, v_{nt}, \dots, v_{nT}, v_{ct}, \dots, v_{cT}). \tag{10.2}$$

The consumption argument in (8) may be replaced by separate arguments for different types of consumption goods if preferences are separable among them. Equations similar to (9.2) are obtained for each good. The price of the good relative to a consumption price index in period $t$ enters the equation with the coefficients $B_c$ of the particular good.

### A. *Approaches to Estimation*

The problem is to estimate the intertemporal labor supply elasticity $B_n$. Panel data are available on the measures $n_t^*$, $c_t^*$, and $w_t^*$. These are assumed to equal $n_t$, $c_t$, and $w_t$, respectively, plus the additive measurement error terms $v_{nt}^*$, $v_{ct}^*$, and $e_t^*$. The measurement error components are assumed to be uncorrelated with the true values. In this study $n_t^*$ is the log of annual hours worked, and $w_t^*$ is the log of total labor earnings divided by the product of annual hours worked and the GNP implicit price deflator. Thus $w_t^*$ is an average hourly wage and includes the effects of overtime. Since $n_t^*$ and $w_t^*$ depend on the same measure of hours worked, the measurement error components $v_{nt}^*$ and $e_t^*$ are likely to be negatively correlated with each other but uncorrelated with $v_{ct}^*$. Data are also available on a subset of the determinants of $v_{ct}$ and $v_{nt}$ and the determinants of $w_t$ and $A_t$. Finally, a second measure $w_t^{**}$ of $w_t$ is available. In practice, $w_t^{**}$ is based on the response to a direct question about the straight-time hourly wage, which is asked of those who are paid on an hourly basis. It refers to the survey date (typically March), while $n_t^*$ and $w_t^*$ are averages over the preceding calendar year.

#### Use of Past Labor Supply as a Proxy for $\ln \lambda_t$

This approach to estimation of the male labor supply equation is due to MaCurdy (1981) and is closely related to Heckman and MaCurdy's (1980) treatment of $\lambda_t$ as a fixed effect. MaCurdy works with the first-difference equation (11.1). After substituting in $Dw_t^*$ for $w_t - w_{t-1}$ and $Dn_t^*$ for $n_t - n_{t-1}$ one obtains

$$Dn_t^* = B_n Dw_t^* - B_n(\ln \phi - \ln R_{t-1,t}) + B_n u_t - B_n De_t^* - B_n Dv_{nt} + Dv_{nt}^*, \tag{12}$$

where $D$ is the first-difference operator, $De_t^*$ is the measurement error in $Dw_t^*$, and $Dv_{nt}^*$ is the measurement error in $Dn_t$. MaCurdy assumes explicitly that agents have perfect foresight, in which case $u_t$ is zero and $D \ln \lambda_t$ is simply $\ln R_{t-1,t}$, which MaCurdy assumes to be constant.[9] The wage measure $Dw_t^*$ is endogenous in (9) since at a minimum it is correlated with the measurement components $Dv_{nt}^*$ and $De_t^*$. MaCurdy uses education, age, interactions of these variables, and a set of year dummies as the principal instrumental variables for the change in wages (see below). These variables should be orthogonal to the measurement error components.[10] Also, MaCurdy's findings should not be very sensitive to the assumption of perfect foresight. Changes in $\ln \lambda_t$ occur because of new information on wages, the nonlabor income process, or preferences. With the exception of the year dummies, MaCurdy's instruments are known in advance and so should not induce changes in $\ln \lambda_t$.

A limitation of MaCurdy's approach is that the age and education terms have a very weak role in explaining wage changes. Furthermore, it is likely that the change in tastes for labor supply $Dv_{nt}$ is related to age, in which case age is invalid as an instrumental variable for $Dw_t^*$ in estimating (12). Finally, the theory of human capital investment suggests that the returns to investment in general training differ systematically over the life cycle, with the implication that the amount of time on the job devoted to investment in human capital declines with age. This suggests that the effect of age on wage growth will be negatively correlated with the effect of age on the change in future earnings power from an extra hour at the job today, which is part of the change in the price of leisure. The result is a downward bias in the elasticity estimate.[11] Consequently, it is useful to explore the sensitiv-

[9] Heckman and MaCurdy (1980) control for $\ln \lambda_t$ in the level eq. (9.1) by maintaining that consumers have perfect foresight, in which case (10.1) implies that $\ln \lambda_t$ is equal to $\ln \lambda_0$ plus $\ln(R_{0,t})$, where $\ln \lambda_0$ is the log of marginal utility of income in the initial period. This permits $\ln \lambda_0$ to be subsumed within a separate dummy variable for each individual. One may think of this procedure as the use of an average of labor supply behavior in other periods as a proxy for $\ln \lambda_t$. Heckman and MaCurdy assume that the real interest rate is constant, in which case the interest rate factor and $\ln(\phi)$ contribute to the equation a term involving $t$.

[10] Ashenfelter and Ham (1979) estimate labor supply equations for males based on an intertemporal Stone-Geary utility function under the assumption of perfect foresight using a fixed effect to control for $\ln \lambda$, although the focus of their discussion is on testing for constraints on choice of hours. The fact that their estimates of the intertemporal substitution elasticity are of the wrong sign (negative) may be due to measurement bias resulting from their use of the imputed wage without instrumental variables in estimating their equations.

[11] See Mincer (1974), who explains the quadratic effect of experience on the wage level by assuming a declining percentage of time spent in on-the-job training. The conclusion in the text may be altered if human capital raises productivity in producing human capital, as is often assumed in theoretical studies of human capital and life-cycle labor supply. Killingsworth (1983) surveys the work on this subject. Heckman (1976*a*,

ity of MaCurdy's results to other means of instrumenting the wage variable.

MaCurdy's approach is implemented below using the first difference of the second wage measure $Dw_t^{**}$ as an instrumental variable for the first measure $Dw_t^*$. The equation for $Dw_t^{**}$ is

$$Dw_t^{**} = a_0 + a_1 Dw_t + De_t^{**}. \tag{13}$$

Assume that (i) the measurement error $e_t^{**}$ is independent of the measurement errors $e_t^*$ and $v_{nt}^*$, (ii) $Dw_t$ is uncorrelated with the unobserved preference shift $Dv_{nt}$ (controlling for age and perhaps year effects), and (iii) $Dw_t$ is uncorrelated with $u_t$. Then $Dw_t^{**}$ is a valid instrumental variable for $Dw_t^*$. Assumption iii does not require perfect foresight. It will hold if individuals have rational expectations and one assumes that the real wage is known one period in advance. In this case $Dw_t$ is used to form the expectation of $\ln \lambda_t$ at time $t - 1$ and so is orthogonal to the revision $u_t$. This assumption is relaxed in some of the estimates below by using only prior values of $Dw_t^{**}$ as instrumental variables for $Dw_t^*$ (see the discussion of [5] above). However, because of the inconsistency in the timing of the $Dn_t^*$ and $Dw_t^{**}$, $Dw_{t-1}^{**}$ may incorporate some information about $u_t$ since $w_{t-1}^{**}$ is measured a few months after the start of calendar year $t$, and so this test is not entirely clean. One may also show that the use of annual averages ($n_t^*$ and $w_t^*$) rather than wages and hours at a point in time will lead to a downward bias in $B_n$ when $Dw_t^{**}$ (or $Dw_t^*$) are used as instruments for $Dw_t^*$ if the planning horizon for (10.1) and (11.1) is shorter than a year, unless consumers know wages 2 years in advance. Assumption ii of the independence of the wage changes and the preference changes is maintained throughout. The main justification for assumption i that $e_t^{**}$ is independent of $e_t^*$ and $v_{nt}^*$ is that $w_t^{**}$ is based on a question separate from those used to construct $w_t^*$ and $n_t^*$.[12] However, nonresponse measurement errors may also be present and result in a positive relationship between $e_t^*$ and $e_t^{**}$. A source of such errors is differences in the price level across locations that are reflected in nominal wages as a compensating differential. These geographical differences are likely to be very stable over time and so should not affect the analysis of first differences. Neither wage measure is adjusted for taxes, but to a first approximation this error component will be subsumed in the constant term of the first differences.[13]

1976*b*) appear to be the only studies exploring the empirical implications of human capital investment in studies of labor supply.

[12] Of course, it is possible that errors of memory that affect the responses to the questions about earnings and hours in the preceding year also affect the response to the question about the hourly wage.

[13] Bias might arise if changes in location are frequent. Comparison of the family codes for state and county of residence for a given year with the code from the previous

It should be kept in mind that the use of past values of $Dw_t^{**}$ might be regarded as a joint test of assumptions i, ii, and iii rather than simply iii.

### Use of Consumption as a Proxy for $\ln \lambda_t$

Equation (9.2) may be solved for $\ln \lambda_t$ as a function of $c_t$:

$$\ln \lambda_t = \text{constant} + \left(\frac{1}{B_c}\right)c_t + t\ln(\phi) + v_{ct}.$$

Substitution for $\ln \lambda_t$ in (9.1) and replacement of $n_t$, $c_t$, and $w_t$ with their respective measures leads to

$$\begin{aligned} n_t^* = \text{constant} + B_n w_t^* + \left(\frac{B_n}{B_c}\right)c_t^* \\ + B_n(v_{ct} - v_{nt}) + v_{nt}^* - B_n e_t^* - \left(\frac{B_n}{B_c}\right) v_{ct}^*. \end{aligned} \tag{14}$$

Variable $c_t^*$ must be treated as endogenous in the estimation of (14) for several reasons. First, $c_t$ depends on the preference variable $v_{ct}$ through equation (9.2). Furthermore, if $(v_{ct} - v_{nt})$ is correlated over time, as one would expect in a sample of heterogeneous individuals, then $c_t$ will be correlated with the error term because of the dependence of $\ln \lambda_t$ on the lifetime taste vectors. For example, those who "need" large quantities of consumption goods relative to their need for leisure will consume a great deal and work a great deal for any given profile of wages and nonlabor income. The result is to bias the estimate of the consumption coefficient in the positive direction. Finally, $c_t^*$ is also endogenous because of the measurement error $v_{ct}^*$, which is likely to bias the estimate of $B_n/B_c$ toward zero. The endogeneity of consumption may be handled through an IV scheme using variables that are related to wealth and to the distribution of lifetime wages. These affect $c_t$ through $\ln \lambda_t$ (see [10.2]). Note that $w_t^*$ must also be treated as an endogenous variable in estimating (14) since it is correlated with the measurement errors $v_{nt}^*$ and $e_t^*$.

The intuition underlying use of equation (14) is that both labor supply and consumption are determined as functions of current and

year indicated that only a small fraction of the effective sample for tables 1–3 and 5 changed location between years. Removal of observations for which a change of location occurred had little effect on the results. I also investigated the possibility that differences across regions in trends in prices or taxes affect the first-difference results by adding dummy variables for region, residence in a standard metropolitan statistical area (SMSA), and residence in a city larger than 500,000 people to some of the first-difference equations for labor supply and consumption. These made little difference.

expected future wages, real interest rates, wealth, and preferences. Consequently, the consumption choice provides information on these variables that may be useful in controlling for their effects on labor supply.[14] For the utility function (8) this approach may be implemented rigorously since it implies that consumption and labor supply depend on wealth and wages and real interest rates from other periods only through $\ln \lambda_t$. Furthermore, the within-period separability of preferences implies that consumption depends on the current wage only through $\ln \lambda_t$. As a result, (14), which is simply the log transformation of the marginal rate of substitution condition (6) above, is sufficient to estimate the intertemporal labor supply response *if the assumption of additive separability of preferences is correct,* although this assumption is not consistent with the available evidence (see below). The principle underlying use of consumption to eliminate $\ln \lambda_t$ is similar to the use of labor supply in other periods, and the intertemporal labor supply condition (12) and the labor supply–consumption condition (14) are complementary rather than competing approaches.[15] The main advantage of (14) is that consumption depends on wage rates only through $\ln \lambda_t$. Thus the substitution for $\ln \lambda_t$ using consumption does not introduce wage rates from other periods into the equation. Consequently, cross-sectional variation in wage profiles as well as intertemporal variation aids in identifying the intertemporal substitution parameter. Another advantage is that strong expectational assumptions such as perfect foresight or rational expectations are not needed to justify the procedures used to implement (14).[16] A third advantage is that the approach should not be very sensitive to the assumption of perfect credit markets that under-

[14] Altonji (1982) uses the approach of this section with aggregate time-series data. MaCurdy (1983) does so with a more general utility function and data from the Denver negative income tax experiment. The idea has many antecedents in the literature. Two of the most direct antecedents are Metcalf (1974) and Heckman (1974*a*). Metcalf (1974) uses observed savings behavior within a simple life-cycle model in his analysis of the extent to which the New Jersey income tax experiment measured responses to temporary rather than permanent wage changes. Heckman (1974*a*) explicitly discusses the possibility of using savings in the female labor supply equation to control for intertemporal aspects of the problem. The approach is closely connected to the life-cycle two-stage budgeting literature. See the discussion and references in Killingsworth (1983) and MaCurdy (1983) and the recent paper by Blundell and Walker (1983).

[15] One may think of the demand system involving $\ln \lambda_t$ and the measurement equations for $n_t$, $c_t$, and $w_t$ as a multiple indicator, multiple cause model. Subsets of the indicators (e.g., $n_t^*$, $c_t^*$, and $w_t^*$) are used to eliminate $w_t$ and $c_t$ from the labor supply eq. (9.1). The equation is estimated using a second set of indicators (e.g., $w_t^{**}$) as instrumental variables for the first. The approach is similar to Madansky's (1964) IV estimator for factor-analytic models. I am indebted to James Heckman for the Madansky reference.

[16] Rational expectations must hold to justify some of the inferences about within-period separability between consumption and labor supply discussed below.

lies the first-order conditions (3.1) and (3.2) and on which the $\lambda$-constant labor supply and consumption functions are based.[17]

The advantages of the consumption–labor supply approach are accompanied by several disadvantages that arise from the use of cross-sectional information. First, wages may be related to unobserved permanent differences in preferences for labor supply or consumption. If this is the case, then both $w_t^{**}$ and a measure of permanent differences across individuals in lifetime wages, which serve as instrumental variables for $w_t^*$ and $c_t^*$, will be correlated with the error term in the supply equation. The result is bias in the estimate of $B_n$. The direction of this bias is unclear even if one assumes a positive correlation between the wage component and labor supply preference component. Second, since the method uses the levels of the variables, it is likely to be more sensitive than the first-difference approaches to measurement error in real after-tax wages and in consumption that arises from the fact that regional variation in taxes and consumer prices may be reflected in nominal wages (as compensating differentials) and in nominal consumption expenditures.[18] On the other hand, the method should be less sensitive to failure to measure the returns to human capital investment in measuring the wage. The growth of wages associated with age and experience plays a much smaller role in identifying the intertemporal substitution parameter than in MaCurdy's analysis using a human capital type wage equation and a somewhat smaller role than in the analysis using the alternative wage change measure as an instrument.

Finally, it is important to assess the sensitivity of intertemporal labor supply estimates based on the first-difference equation (12) and the labor supply–consumption equation (14) to the assumption of within-period separability of the utility function (8) on which they are based.[19] If the assumption is false, then changes in the current price

[17] This point was made by Mankiw et al. (1985). Under the assumption that $n_t$ and $c_t$ affect the discount rates in proportion to their effects on $A_t$, one may show that (14) is unaffected. Note, however, that with imperfect capital markets knowledge of the preference parameters $B_n$ and $B_c$ that underlie the labor supply and consumption equations is not enough to determine the response of labor supply to a transitory wage disturbance even if it is anticipated. In this case, the change in labor income resulting from the transitory wage disturbance affects the level of $A_t$ and thus alters the shadow price of $n_t$ and $c_t$ by altering their marginal impact on future interest rates.

[18] To minimize this problem, dummies for region and urban residence are added to the equation. Dummies for each year will control for mismeasurement in the nationwide price variables.

[19] It would also be desirable to relax the assumption of intertemporal separability of preferences. Hotz, Kydland, and Sedlacek (1982) do so using the nonseparable utility function presented in Kydland and Prescott (1982) but assume that wages, labor supply, and consumption are measured without error and that there is no unobserved heterogeneity in preferences. The intractability of nonseparable preference structures, the problem of distinguishing between state dependence and heterogeneity in prefer-

of one good affect the demand for the other goods even when $\ln \lambda_t$ is held constant. Thus a simple way to analyze the empirical consequences of failure of the assumption is to add cross-substitution terms to the $\lambda_t$-constant demand equations (9.1) and (9.2), which leads to

$$n_t = \text{constant} + B_n w_t + (B_n + B_{nc}) \ln \lambda_t + B_2 t \ln(\phi) + B_3 v_{nt}, \quad (15.1)$$

$$c_t = \text{constant} + B_{cn} w_t + (B_c + B_{cn}) \ln \lambda_t + B_4 t \ln(\phi) + B_5 v_{ct}. \quad (15.2)$$

Equations (15.1) and (15.2) may be regarded somewhat loosely as a loglinear approximation to the true demand system in the nonseparable case. Symmetry of the $\lambda$-constant cross-substitution effects implies that the elasticity $B_{cn}$ is approximately equal to $B_{nc}(N_t W_t / C_t)$, and so $B_{cn}$ and $B_{nc}$ have the same sign. I assume that the within-period preferences are strictly concave and that both consumption and leisure are normal goods, which implies that $B_c + B_{cn} < 0$ and $B_n + B_{nc} > 0$ (see Heckman 1974*b*, p. 191).

Following through on the algebraic substitutions used to derive the empirical labor supply models leads to the first-difference equation for hours, which may be analyzed in the same way as equation (12):

$$\begin{aligned} Dn_t^* = {} & B_n Dw_t^* + (B_n + B_{nc}) \ln R_{t-1,t} + B_2 \ln \phi \\ & + (B_n + B_{nc}) u_t - B_n De_t^* + B_3 Dv_{nt} + Dv_{nt}^*. \end{aligned} \quad (16)$$

However, the effects of violation of separability are more serious for the consumption–labor supply approach. Equation (14) is replaced by

$$\begin{aligned} n_t^* = {} & \text{constant} + (B_n - \Gamma B_{cn}) w_t^* + \Gamma c_t^* - (\Gamma B_4 - B_2) t \ln \phi \\ & - \Gamma B_5 v_{ct} + B_3 v_{nt} + v_{nt}^* - (B_n - \Gamma B_{cn}) e_t^* - \Gamma v_{ct}^*, \end{aligned} \quad (17)$$

where $\Gamma$ is $(B_n + B_{nc})/(B_c + B_{cn})$, which is less than zero. In (17) the coefficient on $w_t^*$ is $B_n - B_{cn}\Gamma$, where $B_n$ continues to be the intertemporal labor supply response parameter. If the assumption that $B_{cn} = 0$ is false (i.e., consumption depends on the wage even when $\ln \lambda_t$ is controlled for), then the coefficient on $w_t^*$ will be biased as an estimate of the intertemporal substitution effect by the quantity $-B_{cn}\Gamma$. If $B_{cn} > 0$ (leisure and consumption are substitutes), then the

ences, and the problem of distinguishing between state dependence due to nonseparable preferences and state dependence due to failure to condition properly on past wages and wealth make the assumption of intertemporal separability a very difficult one to relax. My view is that, for periods of a year, intertemporal separability is a reasonable assumption and that the major sources of state dependence in labor supply are related to costs of job search and specific human capital considerations, which make job mobility and labor force transitions costly. Unfortunately, the empirical work below does not deal with these issues either. It is worth noting that Flinn and Heckman (1982) do not find much evidence of state dependence in their study of transitions by young men from employment to nonemployment, although this is consistent with a fixed cost of moving between the states, which does not vary with spell duration.

estimate of $B_n$ is biased upward given that $\Gamma$ is negative. If $B_{cn} < 0$, then the estimate of $B_n$ is biased downward. The household production literature, in which leisure and goods are viewed as inputs in the production of commodities such as food consumption, provides a strong presumption against the hypothesis $B_{cn} = 0$ (see Ghez and Becker 1975). The evidence from the limited number of studies of consumption and labor supply that have been conducted in a life-cycle setting indicates that $B_{cn}$ is positive. Ghez and Becker (1975, p. 62) report estimates based on synthetic cohorts that suggest that $B_{cn}$ is .35 for food and .53 for consumption as a whole. Smith (1977, p. 243) obtains .72 for consumption as a whole, and Thurow (1969, table 1) also uses cross-section data to document the fact that consumption and income vary together with age.[20] Sample evidence on the assumption that $B_{cn} = 0$ is obtained below by analyzing the following first-difference equation:

$$\begin{aligned} Dc_t^* &= B_{cn}Dw_t^* - B_{cn}De_t^* + (B_c + B_{cn})(u_t + \ln R_{t-1,t}) \\ &\quad + B_4 \ln \phi + Dv_{ct}^* + B_5 Dv_{ct}. \end{aligned} \tag{18}$$

This equation is easily obtained from (10.1) and (15.2). The sensitivity of estimates of $B_n$ based on the consumption–labor supply approach to alternative values of $B_{cn}$ is explored below.[21]

[20] The results of these studies reflect the empirical fact that consumption closely follows the concave age profile of wages and earnings. Biases would arise if consumption preferences vary over the life cycle in a nonlinear way. Also, for a number of reasons these estimates may be influenced in part by income effects. Thurow (1969) interprets his findings as evidence for a consumption function with liquidity constraints. Ghez and Becker (1975, p. 73) acknowledge that their overall findings are consistent with the absolute income model of consumption, although they reject this view. In the other direction, liquidity constraints cannot easily explain the fact that consumption and income decline together later in life (as Thurow is aware), and Heckman (1974*b*) shows that Thurow's evidence is fully consistent with the life-cycle model once the possibility that $B_{cn}$ is positive is taken into account. Furthermore, the results should not be discounted simply because they are based on synthetic cohorts since MaCurdy's (1981) panel data estimates for labor supply are reasonably close to Ghez and Becker's (1975) and Smith's (1977) synthetic cohort estimates.

[21] As was discussed above and MaCurdy (1983) and Mankiw et al. (1985) show, one may relax within-period additive separability directly by deriving the first-order conditions (7.1) and (7.2) along with (6) using a specific nonseparable utility function and estimating the utility parameters from these conditions. Unfortunately, because of nonlinearities in the way that $C_{t-1}$, $C_t$, $N_{t-1}$, and $N_t$ as well as the unobserved components of preferences enter these conditions, the method is likely to be very sensitive to variation in preferences and especially to measurement error in $C$, $N$, and $W$ of the type considered below. Given evidence that measurement error is very important in the data set used, I employ the approach outlined above to analyze the additive separability assumption rather than work with a nonseparable version of the utility function (8). Browning et al. (1983) work with nonseparable preferences but use grouped data and add on error terms to their $\lambda$-constant demand equations rather than build unobserved heterogeneity into the preference structure.

With the above as background the paper turns to a brief discussion of the data set followed by the empirical analysis.

## III. The Data Set

The data are from the first 14 (1968–81) panels of the Michigan Panel Study of Income Dynamics (PSID). The PSID is one of the few data sets that are longitudinal and contain information on both labor supply and consumption.[22] The sample was selected as follows. First, it is limited to men married to the same wife for the years 1968–79 from the 12-year PSID family tape. Second, individuals below age 25 in 1968 or above age 60 in 1979 are excluded to minimize the complications associated with schooling and retirement. Third, men are excluded if the wife was above age 63 in 1968 or if the husband was retired. Fourth, information for 1980–81 (from the 14-year tape) on the men is used if the marriage remained intact and the husband was not over 60, retired, or disabled in these years. Finally, observations for a particular year are excluded if the individual worked zero hours or if data are missing for the variables used in a particular procedure.[23] The sample size for the labor supply equations varies considerably with the specification. A major source of the variation is the unavailability of data on $c_t^*$,[24] and especially $w_t^{**}$ or $w_{t-1}^{**}$, which are required for some procedures but not for others. The reported hourly wage $w_t^{**}$ is available only for hourly rated workers, which means that salaried and self-employed workers are excluded from most of the analysis. About 56 percent of the observations that satisfy the other criteria for inclusion in the sample for the first-difference labor supply equation are excluded as a result. (The corresponding figure for the consumption–labor supply equation is 53 percent.) This limits the generality of the findings of the study. Also, the parameter estimates are not corrected for the possibility that selection bias results from the various sample selection criteria.

[22] As noted earlier, MaCurdy (1983) uses the consumption and labor supply data collected in the Denver negative income tax experiment. Another possible data source is the Retirement History Survey. Hamermesh (1982) has used the consumption data from this study.

[23] The imputed wage $w_t^*$ data were treated as missing if the wage measures increased by 250 percent or more than \$13 or fell by 60 percent or more than \$13 from one year to another. They were also treated as missing if the real wage was less than \$0.40 in 1972 dollars. The same criterion was applied to $w_t^{**}$. The 250 percent, 60 percent limits were also used for consumption and labor supply. In addition, the labor supply variable was treated as missing if annual hours exceeded 4,860.

[24] The consumption measure $c_t^*$ is missing for all families for 1968 and 1973 and is not reported by some families in other years. The sample sizes for each procedure are reported in the tables. The App. analyzes the effects of the various selection criteria on the composition of the samples.

A few of the variables require discussion. Variable $w_t^*$ is the log of earnings divided by the product of annual hours working for pay and the GNP price deflator for consumption. The data on $w_t^{**}$ were obtained as follows. For the survey years 1970–81, workers who said that they were paid on an hourly basis were asked about their hourly wage rate as well as about labor earnings and hours.[25] The reported wages were deflated using the GNP price deflator for consumption. Unfortunately, for the years 1970–77 hourly wage responses above \$9.98 per hour are coded as \$9.98. Observations on $w_t^{**}$ affected by this ceiling were not used. This will not bias the labor supply estimates under the assumption that the error component in the wage equation is independent of the labor supply error.[26] Both $c_t^*$ and $w_t^{**}$ refer to the time of the interview (typically in March), while $n_t^*$ and $w_t^*$ are based on annual hours worked and labor earnings during the preceding calendar year. It was mentioned earlier that the inconsistency in the timing of $w_t^*$ and $w_t^{**}$ may affect the appropriate interpretation of the first-difference results when $Dw_{t-1}^{**}$ is used as an instrumental variable for $Dw_t^*$ for purposes of testing assumption iii that $Dw_t$ is uncorrelated with $u_t$. The inconsistency in timing is probably not a serious issue for the other procedures used below.[27] However, it does suggest that $a_1$, which plays a role in the discussion of the importance of measurement error in $Dw_t^*$ but not in the labor supply estimates, is less than one. The truncation of $Dw_t^{**}$ will also make $a_1$ less than one.

Since the consumption data are limited to information on food consumption, $c_t^*$ is the log of the sum of the family's food expenditures at home and outside the home, deflated by the food component of the consumer price index. It is natural to ask if the consumption data are sufficiently accurate to be usable and whether the fact that

[25] For the survey years 1977–81, the PSID contains an hourly wage for salaried workers. However, it is usually imputed from information on salary per year, per month, per week, etc., using a standard number of hours per pay period (e.g., 2,000 hours per year). The imputation process introduces a positive correlation between the wage measure and the true labor supply disturbance. The data for salaried workers are not used because of this and other doubts about the imputation process.

[26] The fact that the means of $w_t^*$ and $w_t^{**}$ are close (1.437 vs. 1.392) suggests that truncation is not a serious problem. Furthermore, variance decompositions based on data for 1978–81, which were not truncated, are very similar to those reported below. The means of $w_t^*$ and $w_t^{**}$ for these years are 1.494 and 1.448, respectively.

[27] Use of the same subscript ($t$) for both sets of variables despite the differences in dates is a bit misleading but seems preferable to introducing a second subscript. Since $c_t^*$ and $w_t^{**}$ may be correlated with changes in the marginal utility of income that occur in the months following the end of the calendar year, an error is introduced in using $c_t^*$ to proxy $\lambda_t$ in the equation for $n_t^*$. However, the problem is minor provided that innovations in $\lambda$ resulting from information in $w_t^{**}$ that is not incorporated in $n_t^*$ are small relative to the total variance in $\lambda_t$. This seems likely. Estimates very similar to those in table 4 are obtained when $w_{t-1}^{**}$ and $c_{t-1}^*$ are used in place of $w_t^{**}$ and $c_t^*$ in the consumption–labor supply analysis.

they are limited to food consumption has an important effect on the results. The answer to the first question is yes, since the first-stage equations for consumption have substantial explanatory power and are generally reasonable. The assumption that the response error in consumption is independent of the response errors for the wage and hours variables is reasonable given that these variables are based on independent questions. The answer to the second question hinges on the degree of separability between food consumption and consumption of other goods and the variability in the relative price of food (recall the composite good theorem). The use of food consumption may be justified rigorously if the period utility function depends on the sum of food consumption raised to an exponent and a separate argument for consumption of other goods.

## IV. Results

### A. *Estimation of the First-differenced Labor Supply Equation Using Alternative Wage Measures*

This section reports IV estimates of MaCurdy's labor supply equation (12) using $Dw_t^{**}$ in the first-stage equation for $Dw_t^*$.[28] Columns 1–4 of table 1 report the estimates of (12). The estimate of the first-stage equation for $Dw_t^*$ is

$$\underset{(.004)}{Dw_t^* = .0207} + \underset{(.026)}{.295}Dw_t^{**};\ \text{S.E.} = .227;$$

$$R^2 = .031;\ \text{obs.} = 4{,}004;\ F = 129.06.$$

The estimated standard errors of the parameter estimates are enclosed in parentheses. In column 1 the estimate of the intertemporal labor supply elasticity $B_n$ is .067 with a standard error of .080. The estimates in table 1 fall slightly with the addition of age and year dummies to the basic equation.[29]

Overall, the results in table 1 suggest an estimate of $B_n$ in the neigh-

[28] Correcting standard errors of the parameter estimates of the first-difference equations made little difference, and so the uncorrected standard errors are reported in the interest of simplicity. This reflects the fact that the correlation of the residuals over time for a given individual is close to zero after one lag. For example, the corrected and uncorrected standard errors of the wage coefficient in table 1, col. 1, are .086 and .080, respectively, and so the uncorrected standard error is understated by only 7.5 percent. The correction is more important when the data are in levels and is made for the estimates of the labor supply–consumption eq. (14) reported below (see n. 36). Note that three-stage least squares was not used to estimate (12) because it is inconsistent. This is because $u_{t-1}$ will be correlated with $Dw_t$ unless individuals have perfect foresight.

[29] Note that the low $R^2$'s mean little in a measurement error model of this type.

TABLE 1

FIRST-DIFFERENCE EQUATIONS FOR LABOR SUPPLY (Dependent Variable = $Dn_t^*$)

| Explanatory Variable | Instrumental Variables for $Dw_t^*$: $Dw_t^{**}$ | | | | Instrumental Variables for $Dw_t^*$: $Dw_{t-1}^{**}$, $w_{t-1}^{**}$ | | | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Constant | −.0087 | −.0138 | −.0350 | −.0309 | −.0079 | .0092 | −.0320 | −.0202 |
| | (.0039) | (.0199) | (.0098) | (.0213) | (.0104) | (.0241) | (.0156) | (.0293) |
| $Dw_t^*$ | .0663 | .0673 | .0432 | .0428 | .0556 | .0387 | .0181 | .0142 |
| | (.079) | (.0795) | (.0787) | (.0787) | (.4573) | (.454) | (.450) | (.449) |
| Age | ... | .0001 | ... | −.0001 | ... | −.0004 | ... | −.0002 |
| | | (.0004) | | (.0005) | | (.0005) | | (.0005) |
| Year dummies? | no | no | yes | yes | no | no | yes | yes |
| $F$-ratio | .70 | .38 | 4.84 | 4.44 | .01 | .31 | 4.44 | 4.07 |
| $R^2$ | .0002 | .0002 | .0132 | .0132 | .0000 | .0002 | .013 | .014 |
| Observations | 4,004 | 4,004 | 4,004 | 4,004 | 3,269 | 3,269 | 3,269 | 3,269 |

NOTE.—Standard errors are in parentheses. The first-stage equations are presented in table A1.

borhood of .05 with a standard error of .08.[30] For purposes of comparison, table 2 presents estimates of (12) using MaCurdy's human capital type first-stage equation for $Dw_t^*$.[31] The sample for columns 5–7 is essentially the same as for table 1. The estimate of $B_n$ is .100 when age is included as a control and .083 when it is excluded. The standard errors are about .16. The estimate rises to .306 when year dummies are added but is subject to a large standard error. The imprecision of the estimates is partially a reflection of the weak performance of the first-stage equation for $Dw_t^*$. (The $R^2$ is .0080. The only significant variables are the year dummies.) Columns 8–10 of table 2 report results using $Dw_t^{**}$ as the wage change measure in the labor supply and wage change equations. The estimate of $B_n$ is .4508 with a standard error of .19 in the basic equation. This falls to .2937 when age and year effects are added but is subject to a huge standard error.

Columns 1–4 of table 2 report results over the entire 1969–81 sample using MaCurdy's wage specification. This sample includes observations that lack data on $Dw_t^{**}$ and thus is not restricted to workers who are paid by the hour. It differs from MaCurdy's sample primarily in that it includes both whites and nonwhites, is not restricted to individuals who have complete data and worked positive hours in all years, includes households drawn from the Survey of Economic Opportunity sample of low-income areas, and extends three additional years. Despite these differences, the estimates of $B_n$ in columns 1–4 center around .28. These results compare with MaCurdy's two-stage least-squares (2SLS) estimates of .23 when year dummies are excluded and .15 when they are included. Exclusion of blacks from the sample has little effect on the results. If anything, then, the present sample results in slightly larger estimates using the human capital wage equation than does MaCurdy's sample. The fact that the estimates for the full sample lie in the middle of the estimates for the subsample with two wage measures in columns 5–10 suggests that the limitation of the subsample to hourly workers is unimportant. Given the large standard errors of these estimates and the contradictory

[30] Similar estimates and standard errors were obtained using both $Dw_t^{**}$ and $Dw_{t-1}^{**}$ as instrumental variables for $Dw_t^*$. This will improve efficiency if changes in $Dw_t$ are correlated over time but results in the loss of a substantial number of the observations because of missing data.

[31] This equation is reported in table A1. The regressors are years of schooling, schooling squared, age, interactions between age and schooling, year dummies, the education of the husband's mother and father, and the economic status of the head's family during his childhood. Since the latter three variables play almost no role in the first-stage equation and result in a substantial loss of observations because of missing data, they were eliminated from the first-stage equation for the subsample of observations that contain data on $Dw_t^*$ and $Dw_t^{**}$ to make the samples underlying cols. 1–4 of table 1 and 5–10 of table 2 as close as possible. In practice this makes little difference.

## TABLE 2

First-Difference Equations for Labor Supply, Instrumental Variables for $Dw_t^*$ and $Dw_t^{**}$: Human Capital, Family Background, Year Dummies (Dependent Variable = $Dn_t^*$)

| Explanatory Variable | Full Sample (1) | (2) | (3) | (4) | Sample with Data on $Dw_t^*$, $Dw_t^{**}$ (5) | (6) | (7) | (8) | (9) | (10) |
|---|---|---|---|---|---|---|---|---|---|---|
| Constant | −.0147 | −.0025 | −.0180 | .0035 | −.0092 | −.0173 | −.0491 | −.0165 | −.0439 | −.0471 |
| | (.1258) | (.0180) | (.0154) | (.0338) | (.0054) | (.0223) | (.0484) | (.0052) | (.0238) | (.108) |
| $Dw_t^*$ | .2817 | .2317 | .4782 | .2666 | .0834 | .1001 | .3057 | ... | ... | ... |
| | (.1258) | (.1434) | (.3255) | (.4155) | (.1587) | (.1661) | (.654) | | | |
| $Dw_t^{**}$ | ... | ... | ... | ... | ... | ... | ... | .4508 | .5425 | .2937 |
| | | | | | | | | (.1893) | (.2079) | (1.78) |
| Age | ... | −.0002 | ... | −.0004 | ... | .0002 | −.0001 | ... | .0006 | .0000 |
| | | (.0004) | | (.0005) | | (.0005) | (.00005) | | (.0005) | (.001) |
| Year dummies? | no | no | yes | yes | no | no | yes | no | no | yes |
| Standard error | .237 | .2293 | .2720 | .2347 | .2066 | .2084 | .2333 | .2074 | .2114 | .2013 |
| $F$-ratio | 5.01 | 2.90 | 1.48 | 1.88 | .28 | .21 | 3.32 | 5.67 | 3.43 | 4.43 |
| $R^2$ | .0005 | .0006 | .0019 | .0026 | .0001 | .0001 | .0099 | .0014 | .0017 | .0132 |
| Observations | 10,036 | 10,036 | 10,036 | 10,036 | 3,996 | 3,996 | 3,996 | 3,996 | 3,996 | 3,996 |

Note.—Standard errors are in parentheses. The first-stage equations for $Dw_t^*$ and $Dw_t^{**}$ are reported in table A1.

point estimates, it would be a mistake to overstate the differences in the results in tables 1 and 2.

### Least-Squares Estimates and Measurement Error in the Wage and Hours Data

Are the various instrumental procedures really necessary given that the low $R^2$'s from the first-stage equations suggest that they involve a substantial loss of efficiency? The answer appears to be yes. The ordinary least squares (OLS) estimate of $B_n$ is $-.319$ with the tiny standard error of .013, and the hypothesis that $Dw_t^*$ is exogenous in the equation is overwhelmingly rejected by a Wu (1973)–Hausman (1978) test. (The OLS estimate is .0195 with a standard error of .023 if $Dw_t^{**}$ is directly substituted for $Dw_t^*$ in the equation.) A common measurement error component in $Dn_t^*$ and $Dw_t^*$ would explain their strong negative relationship, which drives home the importance of accounting for measurement error in estimating (12).[32]

This section digresses from the labor supply analysis to examine the role of measurement error in the wage and hours data. The sample covariances and correlations (in parentheses) that underlie the labor supply estimates are reported in table 3. Let cov( , ) and var( ) denote the covariance and variance of their respective arguments. Cov($Dw_t^*$, $Dw_t^{**}$) is equal to .0049, which provides an estimate of $a_1$ var($Dw_t$) under the assumption that the measurement errors in the two wage indicators are independent. The sample variance of the reported wage measure $Dw_t^{**}$ is .0177. If one assumes that $a_1 = 1$, then the implied estimate of var($De_t^{**}$) for the reported wage measure is (.0177 − .0049) = .0128, which is 72.2 percent of the variance of $Dw_t^{**}$. The sample variance of $Dw_t^*$ is .0498. The estimate of var($De_t^*$) is (.0498 − .0049) = .0449. This calculation suggests that $Dw_t^*$ is an extremely noisy series, with measurement error accounting for 90.1 percent of its variance.

However, $a_1$ is likely to be less than one for two key reasons, and so these estimates may substantially overstate the importance of measurement error in $Dw_t^*$. First, $w_t^*$ is an average hourly wage over the entire calendar year and reflects overtime, multiple job holding, and job changes during the year, while $w_t^{**}$ is the straight-time hourly wage on the main job at a point in time. Second, the timing of the two

[32] Many recent studies of labor supply have dealt with the problem of measurement error by using human capital type earnings functions to form instruments for the wage. Borjas (1980) makes use of the two wage measures available in another data set and simply replaces one measure with the other. Measurement error in the second wage measure will bias the results based on this procedure toward zero. The bias is likely to be small, however, because Borjas works with levels rather than first differences.

TABLE 3

SAMPLE COVARIANCES AND CORRELATIONS OF WAGES, HOURS, AND CONSUMPTION

A. FIRST DIFFERENCES*

| | $Dw_t^*$ | $Dw_t^{**}$ | $Dw_{t-1}^*$ | $Dw_{t-1}^{**}$ | $Dn_t^*$ | $Dc_t^*$ |
|---|---|---|---|---|---|---|
| $Dw_t^*$ | .04978<br>(1.0) | .00491<br>(.165) | −.01029<br>(−.205) | .00108<br>(.036) | −.01701<br>(−.396) | .00022<br>(.0037) |
| $Dw_t^{**}$ | | .01769<br>(1.0) | −.00140<br>(−.047) | −.00619<br>(−.347) | .00017<br>(.006) | .00231<br>(.064) |
| $Dw_{t-1}^*$ | | | .05035<br>(1.0) | .00493<br>(.164) | .00665<br>(.154) | −.00164<br>(−.027) |
| $Dw_{t-1}^{**}$ | | | | .01799<br>(1.0) | −.00017<br>(−.006) | .00018<br>(.005) |
| $Dn_t^*$ | | | | | .03600<br>(1.0) | .00188<br>(.036) |
| $Dc_t^*$ | | | | | | .07238<br>(1.0) |

B. LEVELS†

| | $w_t^*$ | $w_t^{**}$ | $w_i^*$ | $n_t^*$ | $c_t^*$ |
|---|---|---|---|---|---|
| $w_t^*$ | .1962<br>(1.0) | .1347<br>(.824) | .1193<br>(.672) | −.0150<br>(−.150) | .0407<br>(.251) |
| $w_t^{**}$ | | .1363<br>(1.0) | .1112<br>(.752) | .0001<br>(.002) | .0370<br>(.273) |
| $w_i^{**}$ | | | .160<br>(1.0) | −.0016<br>(−.017) | .0245<br>(.1667) |
| $n_t^*$ | | | | .0508<br>(1.0) | .0043<br>(.0524) |
| $c_t^*$ | | | | | .1346<br>(1.0) |

NOTE.—Correlations are in parentheses.
* The sample for the first five cols. has 3,250 observations. It is almost identical to the sample for table 1, cols. 5–8, but differs from the sample for table 1, cols. 1–4, because of missing data on $Dw_{t-1}^*$. The sample for the last col. has only 2,418 observations because of missing data on $Dc_t^*$. The estimated covariances are not sensitive to changes in the sample.
† The sample contains 4,367 observations and is identical to the sample used in table 4.

wage measures is inconsistent since $w_t^{**}$ is the wage at the survey date, while $w_t^*$ is the average wage over the preceding calendar year.

To address the first problem, I analyze the variances and covariances of the two wage measures using only observations on persons who did not work overtime, did not hold an extra job, did not report bonuses or commissions, and did not change jobs during the two survey years prior to the survey on which $w_t^*$ is based or the survey on which $w_{t-1}^*$ is based. To increase the sample size for the analysis, the sample was drawn from males who were heads of household in 1981 and are present on the 14-year (1981) PSID tape. Furthermore, observations on persons who were married and between the ages of 25 and 60 at the time of surveys from which current or lagged values of $w_t^{**}$ or $w_t^*$ are drawn are included in the analysis even if the person was not continuously married and between the ages of 25 and 60 from 1968 to 1981. Variables $w_t^*$ and $w_t^{**}$ were screened for outliers using the procedures described in note 23 and the Appendix. For the resulting sample of 4,132 person-years with good data on $Dw_t^*$ and $Dw_t^{**}$, $\text{var}(Dw_t^{**}) = .0127$, $\text{var}(Dw_t^*) = .0368$, and $\text{cov}(Dw_t^{**}, Dw_t^*) = .0031$. The implied estimate of the true variance of the wage under the extreme assumption that $a_1 = 1$ is .0031, which is less than the estimate of .0049 mentioned above. The estimate of the variance of measurement error falls from .0449 to .0337, but this is still 67.7 percent of the variance of $Dw_t^*$ for the sample used in the labor supply analysis. One would expect the true variance in the change in the straight-time wage to be smaller for persons who did not change jobs. But one might also expect a lower measurement error variance in $Dw_t^*$ for those who worked on only one job and did not work overtime since it may be easier for such workers to keep track of hours and income.

To analyze the problem of inconsistent timing, one must work with an empirically relevant model of the wage process. Let $w_{t \cdot q}$ denote the wage rate in quarter $q$ of year $t$. Assume that, apart from an individual-specific constant (which is irrelevant to the analysis of differences below), $w_{t \cdot q}$ is the sum of years since entry into the labor market $\text{EXP}_{t \cdot q}$ times an individual-specific trend $\pi$, a stationary transitory disturbance $\eta_{t \cdot q}$, and a random walk component $\epsilon_{t \cdot q}$, with

$$w_{t \cdot q} = \pi \text{EXP}_{t \cdot q} + \eta_{t \cdot q} + \epsilon_{t \cdot q}, \tag{19}$$

$$\epsilon_{t \cdot q} = \epsilon_{t \cdot q-1} + \nu_{t \cdot q}. \tag{20}$$

The disturbance $\nu_{t \cdot q}$ is serially uncorrelated and has variance $\sigma_\nu$. Assume that $\eta_{t \cdot q}$ is well approximated by a moving average (MA) process with autocovariance $\sigma_{\nu l}$ at a lag of $l$ quarters and that this autocovariance is zero at all lags greater than or equal to $K - 1$ years. Assume also that the processes for $\eta_{t \cdot q}$ and $\epsilon_{t \cdot q}$ are independent of $\pi$ and each other. Let $\sigma_\pi$ denote the variance of $\pi$ and $w_t$ denote the average wage in year $t$. Then

$$w_t = .25(w_{t \cdot 1} + w_{t \cdot 2} + w_{t \cdot 3} + w_{t \cdot 4}). \tag{21}$$

Equations (19), (20), and (21) imply that

$$\begin{aligned} w_t - w_{t-K} = {} & .25(\eta_{t \cdot 1} + \eta_{t \cdot 2} + \eta_{t \cdot 3} + \eta_{t \cdot 4} - \eta_{t-K \cdot 1} \\ & - \eta_{t-K \cdot 2} - \eta_{t-K \cdot 3} - \eta_{t-K \cdot 4}) \\ & + \left( \sum_{t'=t-K+1}^{t-1} \sum_{q=1}^{4} \nu_{t' \cdot q} \right) + \nu_{t \cdot 1} + .75\nu_{t \cdot 2} + .5\nu_{t \cdot 3} \\ & + .25\nu_{t \cdot 4} + .75\nu_{t-K \cdot 4} + .5\nu_{t-K \cdot 3} + .25\nu_{t-K \cdot 2} + \pi K \end{aligned} \tag{22}$$

and that

$$w_{t\cdot 1} - w_{t-K\cdot 1} = \eta_{t\cdot 1} - \eta_{t-K\cdot 1} + \left( \sum_{t'=t-K+1}^{t-1} \sum_{q=1}^{4} \nu_{t'\cdot q} \right) + \nu_{t\cdot 1} + \nu_{t-K\cdot 2} + \nu_{t-K\cdot 3} + \nu_{t-K\cdot 4} + \pi K. \tag{23}$$

Equations (22) and (23) and the assumptions above about $\eta_{t\cdot q}$ and $\nu_{t\cdot q}$ imply that

$$\mathrm{var}(w_t - w_{t-K}) = .5\sigma_{\eta 0} + .75\sigma_{\eta 1} + .5\sigma_{\eta 2} + .25\sigma_{\eta 3} + (4K - 4 + 2.75)\sigma_\nu + K^2\sigma_\pi, \tag{24}$$

$$\mathrm{var}(w_{t\cdot 1} - w_{t-K\cdot 1}) = 2\sigma_{\eta 0} + 4K\sigma_\nu + K^2\sigma_\pi, \tag{25}$$

$$\mathrm{cov}(w_t - w_{t-K}, w_{t\cdot 1} - w_{t-K\cdot 1}) = .5(\sigma_{\eta 0} + \sigma_{\eta 1} + \sigma_{\eta 2} + \sigma_{\eta 3}) + (4K - 4 + 2.5)\sigma_\nu + K^2\sigma_\pi. \tag{26}$$

Equations (24) and (26) imply that $\mathrm{cov}(w_t - w_{t-K}, w_{t\cdot 1} - w_{t-K\cdot 1}) - \mathrm{var}(w_t - w_{t-K}) = .25(\sigma_{\eta 3} - \sigma_{\eta 1}) - .25\sigma_\nu$. The term $-.25\sigma_\nu$ in this difference must be less than the fraction $.25/[4(K - 1) + 2.75]$ of $\mathrm{var}(w_t - w_{t-K})$ and is negligible even if $K$ is only 3. The difference is negative under the reasonable assumptions that the autocovariances of $\eta_{t\cdot q}$ are positive and declining at short lags but will be small if the decay in the autocovariance function is slow between lags of 1 and 3 quarters. In any case, one may easily show that the difference is at most ⅕ of $\mathrm{var}(w_t - w_{t-K})$ even if $\sigma_\nu$ and $\sigma_\pi$ are both zero, since $\sigma_{\eta 3} \geq 0$ and from the Cauchy-Schwarz inequality $\sigma_{\eta 0} \geq \sigma_{\eta 1}$.

It is necessary to relate the discussion above to the wage measures $w_t^{**}$ and $w_t^*$. Since $w_t^{**}$ is the average wage over the calendar year prior to the survey, it is equal to $w_t$ plus measurement error. Since interviews typically occur in March, the reported wage from the previous survey, $w_{t-1}^{**}$, corresponds to $w_{t\cdot 1}$ plus measurement error. (The calculations are not affected very much by the fact that some of the interviews occur during the late spring and early summer.) MaCurdy's (1982, tables 1, 3) finding that $Dw_t^*$ has a zero autocorrelation at lags of more than 2 years and is well approximated by a second-order MA process and the fact that the autocorrelations of both $Dw_t^*$ and $Dw_t^{**}$ for the sample above are also essentially zero after two lags implies that $K$ may be set to 3 (and that $\sigma_\pi$ is small). For the subsample of 1,552 person-years with data on $w_{t-1}^{**}$, $w_{t-4}^{**}$, $w_t^*$, and $w_{t-3}^*$, which are not affected by job changes, overtime, and so forth, $\mathrm{cov}(w_{t-1}^{**} - w_{t-4}^{**}, w_t^* - w_{t-3}^*) = .0061$, $\mathrm{var}(w_{t-1}^{**} - w_{t-4}^{**}) = .0203$, and $\mathrm{var}(w_t^* - Dw_{t-3}^*) = .0500$. The implied estimate of the measurement error in $\mathrm{var}(w_t^* - Dw_{t-3}^*)$ is .0439. When $K = 6$ (observations $= 690$), $\mathrm{cov}(Dw_{t-1}^{**} - Dw_{t-4}^{**}, Dw_t^* - Dw_{t-3}^*) = .0136$, $\mathrm{var}(Dw_{t-1}^{**} - Dw_{t-4}^{**}) = .0261$, and $\mathrm{var}(Dw_t^* - Dw_{t-3}^*) = .0559$. The implied estimate of the measurement error in $\mathrm{var}(w_t^* - Dw_{t-3}^*)$ is .0423. If one assumes that measurement error in $w_t^*$ is serially uncorrelated, then these are alternative estimates of the measurement error in $Dw_t^*$. They may be understated if measurement error is smaller for persons who do not work overtime, and so forth. They may be overstated (understated) if measurement error in $w_t^*$ is positively (negatively) correlated at intervals of less than 2 years. Finally, $\text{var}(w_{t-1}^{**} - w_{t-K-1}^{**})$ provides an upper-bound estimate for $\text{var}(w_{t\cdot 1} - w_{t-K\cdot 1})$ under the assumption that $w_t^{**}$ is measured without error. From (24) and (25) and the fact that $\sigma_{\eta 0} \geq \sigma_{\eta l}$, $\text{var}(w_{t\cdot 1} - w_{t-K\cdot 1}) \geq \text{var}(w_t - w_{t-K})$. Thus subtraction of $\text{var}(Dw_{t-1}^{**} - Dw_{t-K-1}^{**})$ from $\text{var}(w_t^* - w_{t-K}^*)$ yields a "lower-bound" estimate of the measurement error in $(w_t^* - w_{t-K}^*)$ equal to .0297 when $K = 3$ and .0298 when $K = 6$. Again, this lower bound may be overstated if measurement error is positively correlated at short time intervals.

While the problems of comparability of $w_t^*$ and $w_t^{**}$ make it impossible to provide specific estimates of the relative importance of true wage variation and measurement error in $Dw_t^*$ for the full sample in the text, the analysis above strongly suggests that measurement error is an important problem. In addition, the fact that the correlation of $Dn_t^*$ and $Dw_t^*$ in table 3 is strong and negative while the correlation of $Dn_t^*$ and $Dw_{t-1}^*$ is about 40 percent as strong and positive is also consistent with the presence of measurement error in $Dw_t^*$ arising from measurement error in hours of work, especially in the light of the very small correlations between $Dn_t^*$ and both $Dw_t^{**}$ and $Dw_{t-1}^{**}$. Abowd and Card's (1983) analysis of the covariance structure of annual hours and earnings suggests that there is substantial measurement error in both. More direct evidence on the importance of measurement error is presented by Duncan and Hill (1984), who matched responses by employees of a single large firm to questions similar to those in the PSID to the records of the firm. Duncan and Hill claim that the firm's records are very accurate. Their results indicate that errors in annual hours and in the imputed average hourly wage based on the employees' responses are very large, while the earnings data are somewhat better. Their findings also suggest, however, that the measurement errors in some variables are correlated with the true values, which would raise problems for the labor supply analysis in the present paper (and many others).

In summary, the evidence strongly suggests that measurement error is a substantial problem that must be explicitly accounted for in making inferences about the dynamics of labor supply and wages using the PSID.

### B. *Using Consumption as a Proxy for the Marginal Utility of Income*

Before I turn to the estimates of (14), it is necessary to discuss the instrumental variables for $w_t^*$ and $c_t^*$. The principal instrumental variables for $w_t^*$ are $w_t^{**}$ and an estimate of $w_i$, which is assumed to be an individual-specific permanent component of the wage. The estimate is labeled $w_i^{**}$. The $w_i^{**}$ are obtained from a regression of $w_t^{**}$ on dummy variables for each individual and variables that fluctuate over time.[33] The first-stage equation for $w_t^*$ is reported in table A1, column 10. It contains $w_t^{**}$ and $w_i^{**}$ plus a set of additional variables that enter the labor supply equation, many of which should be related to $w_t^*$. These consist of age, age squared, year dummies, health status, region and urban residence dummies, race, family size, number of children less than 6 years old, and number of children in the family unit. They are added in the hope of improving efficiency of the estimation procedure rather than to avoid bias in the labor supply coefficients on $w_t^*$ and $c_t^*$. In the absence of measurement error and the effects of overtime, multiple job holding, and timing on the comparability of $w_t^*$ and $w_t^{**}$, the coefficient on $w_t^{**}$ would be one and the coefficient on $w_i^{**}$ (along with the other variables) would be zero. In practice, they are .51 and .50, respectively.

The principal variables that identify the effect of consumption are the fixed wage component $w_i^{**}$ and $w_t^{**}$. Variable $w_i^{**}$ measures $w_i$, which is an important determinant of the lifetime wages and thus affects $c_t$ through ln $\lambda_t$. Variable $w_t^{**}$ will capture variation in the profile of wages around the experience profile for persons with a given $w_i$. These deviations may be caused by the evolution of union status, health, the degree of success of job matches, and the many other unobserved factors that affect wage rates. Wife's education and financial status of the husband's parents when the husband was growing up are also added since they are determinants of the budget constraint and affect ln $\lambda_t$. In addition, the equation for consumption contains a number of control variables that are used in the labor supply equation as well. These may be related to tastes for consumption and labor supply, regional differences in the price level or in the relative price of food, as well as (in some cases) the lifetime wage profile of the husband or wife.[34] The instrumental regression for

[33] The latter include controls for region (North Central, South, West, and Northeast), urban location (residence in SMSA, residence in city with population above 500,000), health status, dummy variables for each year, and human capital measures (experience squared and the product of schooling and experience). The experience measure is (age − schooling − 5). The linear experience term is excluded from the equation because it is perfectly collinear with $w_i^{**}$, the schooling-experience interaction term, and the year dummies. Dummy variables for each year capture the effects on real wages of additional experience and of aggregate changes. The wage equation is estimated using all observations with good data on the current value of hours worked and on the current values of the variables in the wage equation.

[34] Since the expected value of wage growth associated with experience in the labor force should not affect the consumption path, age and age squared of the husband are added separately to the consumption equation to prevent this factor from biasing the coefficient on $w_i^{**}$ toward zero. They also control for the rate of time preference minus

consumption is presented in table A1. The equation is basically sensible, which provides some reassurance that the consumption data are adequate for the problem at hand. The fixed effects and the current wage variables have a strong effect on $c_t^*$. The total elasticity with respect to $w_i^{**}$ is .325, which is obtained as the sum of the coefficients on $w_i^{**}$ and $w_t^{**}$. The current wage $w_t^{**}$ has a coefficient of .111, which is below that of $w_i^{**}$, as expected. However, error in $w_t^{**}$ and the fixed effects will bias (probably downward) the coefficients on the wage variables, although this bias does not affect the consistency of the second-stage equation for labor supply.[35] Discussion of the other variables is omitted to save space since they are not of central interest. Many of the variables do not have a clean interpretation since they capture a mixture of income and taste effects and regional differences in wages and prices.

The labor supply estimates are reported in table 4. For the moment, I maintain the assumption of additive separability in interpreting the results. All the equations contain age, age squared, year dummies, health status, region and urban residence dummies, race, and the family composition variables. These variables serve as controls for consumption preferences, labor supply preferences, and price differences across areas. They do not have a simple interpretation in terms of labor supply preferences and will not be discussed further. Column 3 of the table reports IV estimates of the equation. The estimated standard errors have been corrected for the effects of correlation in the errors across observations on each individual.[36]

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the interest rate and for age-related changes in preferences for consumption. Consequently, they must be added to the labor supply equation as well.

[35] Other studies that have used the food data from the PSID are Hymans and Shapiro (1976) and Hall and Mishkin's (1982) recent investigation of the permanent income hypothesis.

[36] Since the sample is unbalanced, with missing data for some individuals in certain years, each element of the covariance matrix of the errors across years was estimated as the average of the cross products of the residuals from individuals with nonmissing data for the pair of years corresponding to the particular element. The covariance matrix of the parameter estimates of the labor supply equation may be estimated as follows. Let $\mathbf{Z}_{ik}$ be a row vector containing the values of the right-hand-side variables from the labor supply equation for individual $i$ in survey year $k$, where the individual subscript $i$ is made explicit. Note that $\mathbf{Z}_{ik}$ contains the predicted values for $w_k^*$ and $c_k^*$ from the first-stage equations rather than the actual values. Replace all the elements of $\mathbf{Z}_{ik}$ with zeros if any of the elements of $\mathbf{Z}_{ik}$ or $n_{ik}^*$ are missing. Stack the observations on $\mathbf{Z}_{ik}$ for different values of $k$ into the matrix $\mathbf{Z}_{i\cdot}$. Let $S$ denote the estimate of the covariance of the errors for a given individual across years, and let $I$ equal the number of individuals in the sample. Then the estimate of the covariance of the parameter estimates is

$$\left(\sum_{i=1}^{I} \mathbf{Z}_{i\cdot}'\mathbf{Z}_{i\cdot}\right)^{-1}\left(\sum_{i=1}^{I} \mathbf{Z}_{i\cdot}'S\mathbf{Z}_{i\cdot}\right)\left(\sum_{i=1}^{I} \mathbf{Z}_{i\cdot}'\mathbf{Z}_{i\cdot}\right)^{-1}.$$

The uncorrected standard errors of the wage and consumption coefficients in col. 3 of the table are .071 and .227. In col. 4 they are .042 and .142.

TABLE 4

LABOR SUPPLY ESTIMATES USING FOOD CONSUMPTION AS A PROXY FOR $\lambda_t$ (See Eq. [14])

| | ESTIMATION METHOD | | | |
|---|---|---|---|---|
| | OLS (1) | OLS (Reduced Form)* (2) | IV (3) | IV (4) |
| Intercept | 7.528 | 7.416 | 8.386 | 7.995 |
| | (.155) | (.157) | (.644) | (.359) |
| $w_t^{*\dagger}$ | −.1126 | . . . | .1721 | .0943 |
| | (.014) | | (.119) | (.057) |
| $c_t^{*\ddagger}$ | .0788 | . . . | −.5341 | −.2972 |
| | (.015) | | (.386) | (.202) |
| $w_t^{**}$ | . . . | −.019 | . . . | . . . |
| | | (.025) | | |
| $w_i^{**}$ | . . . | −.031 | . . . | . . . |
| | | (.032) | | |
| Black | −.021 | −.0028 | −.0449 | −.0315 |
| | (.015) | (.015) | (.036) | (.025) |
| Health | −.079 | −.0683 | −.0702 | −.0703 |
| | (.015) | (.015) | (.020) | (.0168) |
| Age | .008 | .0089 | .0220 | .0177 |
| | (.007) | (.007) | (.013) | (.0101) |
| Age$^2$ | −.0001 | −.0001 | −.0003 | −.0239 |
| | (.0001) | (.0001) | (.0002) | (.0116) |
| Size of family unit | −.0187 | −.0060 | .0424 | .0192 |
| | (.006) | (.006) | (.037) | (.0194) |
| Number of children | .0025 | .0020 | −.0171 | −.0088 |
| under 6 | (.008) | (.008) | (.018) | (.0120) |
| Number of children | .0110 | .0074 | .0424 | .0029 |
| in family | (.006) | (.006) | (.011) | (.0079) |
| SMSA | −.0079 | −.0198 | .0026 | −.0033 |
| | (.015) | (.014) | (.025) | (.0197) |
| City ≥ 500,000 | .0100 | .0021 | .0384 | .0250 |
| | (.017) | (.016) | (.034) | (.0246) |
| South | .0198 | .0527 | −.0116 | .0038 |
| | (.019) | (.021) | (.042) | (.0300) |
| West | .0284 | .0417 | .0025 | .0060 |
| | (.022) | (.024) | (.035) | (.0287) |
| North Central | .0327 | .0542 | −.0151 | .0053 |
| | (.019) | (.024) | (.045) | (.0299) |
| Year dummies? | yes | yes | yes | yes |
| Standard error | .217 | .220 | .292 | .250 |
| $R^2$ | .0809 | .0544 | .0258 | .034 |

NOTE.—Observations = 4,367. The standard errors (in parentheses) have been corrected for correlation over time in the observations on a given individual (see n. 36).

* The reduced-form equation also contains wife's education and dummy variables for whether the individual's parents were average or rich while he was growing up. The coefficients on these variables are .0112, .0086, and .0044. The standard errors are .0028, .0136, and .0224.

† Treated as endogenous in cols. 3 and 4. The instrumental variables for col. 3 are $w_t^{**}$, $w_i^{**}$ plus the other variables in the labor supply equation. The instrumental variables for col. 4 are $w_t^{**}$ and $w_i^{**}$ only (see table A1, cols. 9 and 10).

‡ Treated as endogenous in cols. 3 and 4. The first-stage equation is in table A1, col. 11.

The estimate of the intertemporal labor supply elasticity $B_n$ is .172 with a standard error of .119. This value is very close to those obtained using MaCurdy's approach but is more tightly estimated. The coefficient on $c_t$ is also of correct sign and is statistically significant. It is equal to $-.534$ but with the relatively large standard error of .386. Since the coefficient on consumption is $B_n/B_c$, the estimate of $B_c$ is the ratio of the wage and consumption coefficients. This equals $-.322$. To get a sense of whether this number is reasonable, note that, if male labor supply does not respond to a shift in the permanent wage component, then the total differential of $n_t$ with respect to $w_i$ is zero. From (14) this implies that $B_n = -(B_n/B_c)(dc_t/dw_i)$, which is equivalent to the condition $-B_c = dc_t/dw_i$. Consequently, if male labor supply is not responsive to permanent wage differences, than $-B_c$ provides an estimate of the elasticity of consumption with respect to a permanent wage change. Viewed in this light, the point estimate of $B_c$ seems a bit large in absolute value. Column 4 corresponds to column 3 but is estimated using only $w_t^{**}$ and $w_i^{**}$ as instrumental variables for the wage. The estimates of $B_n$ and $B_n/B_c$ are .094 and $-.297$ with standard errors of .073 and .251. The implied estimate of $B_c$ is $-.317$. The difference is attributed to sampling error since the additional instrumental variables used in estimating column 3 are included in the labor supply and consumption equations.

These results stand in sharp contrast to the OLS estimates of the wage and consumption coefficients in column 1, which equal $-.113$ and .079, respectively, with very small standard errors. The negative sign on $w_t^*$ when OLS is used may be attributed to a negative association between the measurement errors $v_{nt}^*$ and $e_t^*$ in (14). The positive sign on $c_t^*$ is due to correlation with consumption and labor supply preference terms that are not controlled for by the other variables in the equation.[37]

The results indicate that the use of consumption as a proxy for the marginal utility of income is a workable approach to estimating the intertemporal labor supply parameter and also confirms that, given the characteristics of existing data and the endogeneity of consumption, a simultaneous equations approach to estimation is required.[38]

[37] The $F_{2,4,341}$ statistic (uncorrected for correlation across observations on each individual) for the Wu-Hausman test of joint exogeneity in (14) of $w_t^*$ and $c_t^*$ is 108.8. The hypothesis that $w_t^*$ alone is exogenous is rejected with a $t$-statistic of 8.8. The $t$-statistic to test the hypothesis that $c_t^*$ alone is exogenous is a more modest but still significant 3.18. It is conceivable, of course, that the differences in the OLS and IV results reflect differences in the sensitivity of the two procedures to violations of the maintained assumption that the $w_t$ and $w_i$ are exogenous in (14).

[38] The log of the sum of hours unemployed plus 2,000 was added to some of the equations to test the sensitivity of the labor supply results to treatment of unemployment in the labor supply equation in the spirit of Ashenfelter and Ham (1979), Ashen-

These results suggest an estimate of $B_n$ of about .15. Thus they are consistent with the conclusions of the analysis based on first differences. It should be kept in mind that the specific point estimates are somewhat sensitive to choice of instruments for consumption and labor supply, and so the estimated standard errors understate the uncertainty about them.[39]

### C. *Bias Caused by Unanticipated Wage Changes and Nonseparable Preferences*

To sum up, the estimate of $B_n$ based on (12) with $Dw_t^{**}$ used as an instrumental variable is near .05. The results based on (14) suggest an estimate of .15. In view of the standard errors of the estimates, these results and those of MaCurdy are basically compatible and together suggest a small, positive estimate for $B_n$. However, it was mentioned earlier that the estimates in table 1 will be biased if assumption iii, that consumers know the current wage change one period in advance, is false. In this case, $Dw_t$ will be correlated with the expectational error $u_t$ in (12). Its regression coefficient will equal $B_n + B_n\theta$ if preferences are separable and $B_n + (B_n + B_{nc})\theta$ if they are not, where $\theta$ is the regression coefficient of $u_t$ on $Dw_t$. As a result, the estimate of $B_n$ will have a negative bias because the labor supply response to the wage change will include an income effect associated with the surprise in the wage. The size of $\theta$ will depend on the extent to which wage changes are associated with shifts in the expectation of the profile of future wages. If most wage changes persist over time (e.g., the real wage process is a random walk) and the nonstochastic variation (associated, e.g., with

felter (1980), and Ham (1982, 1983). The change in this variable was added to the first-difference equations for labor supply discussed earlier. Since the variable was included without instrumental variables and is subject to a number of interpretations (see, e.g., Heckman, Killingsworth, and MaCurdy 1981), the results are not presented in detail. However, it is worth reporting that its coefficient is in the neighborhood of $-1.2$ and is highly significant in both the first-difference and level equations. The $R^2$ of the first-difference equations rises to .1746 (from .0132) when the change in the unemployment measure is added to the equation in col. 4 of table 1. The estimate of $B_n$ is unaffected in the first-difference equations, and the change in the unemployment variable is almost orthogonal to the wage change variables. However, the estimates of $B_n$ based on the labor supply–consumption equation corresponding to those in cols. 3 and 4 of table 4 are .077 and .041, respectively. Note that Ham (1983) finds that the change in unemployment remains highly significant in (12) even when various IV schemes are used to correct for the endogeneity of this variable.

[39] One may also estimate $B_n$ by using $n_t^*$ as a proxy for $\ln \lambda_t$ in the consumption equation (i.e., inverting [14]). However, this normalization is likely to produce poor results if $B_n$ is small, in which case the link between $n_t^*$ and $\ln \lambda_t$ is weak. Estimates of the model in this form are qualitatively consistent with those in cols. 3–4 of table 4 but are very unstable. This is consistent with a small value for $B_n$ and a reflection of the weak role played by $w_t^{**}$ and $w_t^{**}$ in the reduced-form hours equation reported in col. 2 of the table. The coefficients on these variables are $-.031$ and .019, respectively.

the quadratic term in the wage-experience profile) is small, then the bias may be substantial (if $B_n + B_{nc}$ is also large). The estimates based on the labor supply–consumption relationship are not sensitive to this issue. However, equation (17) shows that they are biased by the quantity $-B_{cn}\Gamma$ if the assumption of additive separability is false and $B_{cn}$ is nonzero, where $\Gamma$ equals $(B_n + B_{nc})/(B_c + B_{cn})$ and is the coefficient on consumption in that equation. Thus it is important to address the possibility that $B_n$ is in fact large and the estimates based on (12) and (14) are both biased toward zero by these separate factors.

As a check on bias in the estimates from the first-differenced equation, the labor supply equation was estimated using $Dw^{**}_{t-1}$ and $w^{**}_{t-1}$ as instrumental variables for $Dw^*_t$. This is a valid IV procedure if expectations are rational and wages dated $t - 1$ or earlier are known at $t - 1$. Unfortunately, the results of this procedure are not very informative. The point estimates of $B_n$ range from .056 to .014 in table 1, columns 5–8, and thus are fully consistent with those reported earlier (see cols. 1–4). However, they are subject to a standard error of .45. The imprecision is caused by the fact that $Dw^{**}_{t-1}$ and $w^{**}_{t-1}$ are very poor predictors of $Dw^*_t$. (The $R^2$ in the first-stage regression is only .001.) This reflects the importance of error in both wage measures as well as the fact that the correlation of $Dw_t$ and $Dw_{t-1}$ is small and most of the changes in $w_t$ are permanent.[40] Evidence that the innovations in the wage process are persistent leaves open the possibility that failure of assumption iii might result in a substantial negative bias.

The estimates of the first-difference equation for consumption (17) shed light on the values of both $B_{cn}$ and $\theta$. Columns 9–11 of table 5

[40] Estimation of the process for $Dw_t$ is complicated by the presence of measurement error. One may regress $Dw^*_t$ on $w^*_{t-1}$ and $w^*_{t-1}$ using $w^{**}_{t-1}$ and $w^{**}_{t-1}$ as instrumental variables (although this ignores the impact of inconsistency in the timing of the two wage measures discussed above). The estimated equation is

$$\underset{}{Dw^*_t} = \underset{(.018)}{.0307} + \underset{(.110)}{.1802}w^*_{t-1} - \underset{(.110)}{.1898}w^*_{t-2};\ R^2 = .001,\ F = 1.73,\ \text{obs.} = 3{,}274.$$

If one replaces $Dw^*_t$, $w^*_{t-1}$, and $w^*_{t-2}$ in the equation above with $Dw^{**}_t$, $w^{**}_{t-1}$, and $w^{**}_{t-2}$ and estimates the resulting equation using $w^*_{t-1}$ and $w^*_{t-2}$ as instrumental variables, one obtains a coefficient of $-.239$ on the first lag and .2138 on the second with standard errors of about .09. The OLS estimate of the equation above for $Dw^*_t$ for the same sample is

$$Dw^*_t = \underset{(.013)}{.1788} - \underset{(.017)}{.3943}w^*_{t-1} + \underset{(.016)}{.2901}w^*_{t-2};\ R^2 = .1514,\ F = 291.9,\ \text{obs.} = 3{,}274.$$

This equation is basically comparable with those obtained by MaCurdy (1982) for $Dw^*_t$, although he found that an MA(2) fits the data better than the process above. The coefficient on the lagged wage has a strong negative bias due to association between the measurement errors in the left- and right-hand-side variables. The important point for present purposes is that the evidence from all three equations suggests that much of the stochastic variation in wages over the lifetime is permanent rather than transitory. This is consistent with MaCurdy's conclusion for the wage process.

TABLE 5

FIRST-DIFFERENCE EQUATIONS FOR CONSUMPTION (Dependent Variable = $Dc_t^*$)

| EXPLANATORY VARIABLE | OLS ESTIMATION | | IV ESTIMATION: INSTRUMENTAL VARIABLES ARE | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | $Dw_t^{**}, w_{t-1}^{**}$ | | | $Dw_t^*, w_{t-1}^*$ | | | $Dw_{t-1}^*, w_{t-1}^*$ | | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) | (10) | (11) |
| Intercept | .087 | .0307 | −.028 | .0116 | .0078 | −.0168 | −.0763 | .0821 | −.0147 | .0807 | .0863 |
| | (.018) | (.031) | (.0059) | (.031) | (.033) | (.0037) | (.0180) | (.0201) | (.0054) | (.0186) | (.0206) |
| $Dw_t^*$ | .016 | ... | .366 | .362 | .365 | ... | ... | ... | ... | ... | ... |
| | (.012) | | (.126) | (.126) | (.126) | | | | | | |
| $Dw_t^{**}$ | ... | .099 | ... | ... | ... | .249 | .219 | .154 | .144 | .059 | −.023 |
| | | (.035) | | | | (.118) | (.118) | (.117) | (.242) | (.242) | (.242) |
| Age | −.0019 | −.0010 | ... | −.0009 | −.0014 | ... | −.0021 | −.0019 | ... | −.0021 | −.0019 |
| | (.0004) | (.0007) | | (.0007) | (.0007) | | (.0004) | (4.29) | | (.0004) | (.0004) |
| Year dummies? | yes | yes | no | no | yes | no | no | yes | no | no | yes |
| Standard error | .267 | .267 | .281 | .281 | .280 | .299 | .269 | .267 | .269 | .269 | .267 |
| $F$-ratio | 15.7 | 5.31 | 8.44 | 5.09 | 4.01 | 4.44 | 16.3 | 15.4 | .36 | 14.6 | 15.3 |
| $R^2$ | .0184 | .0170 | .0027 | .0033 | .0116 | .0005 | .0039 | .0199 | .0000 | .0035 | .0197 |
| Observations | 8,353 | 3,083 | 3,083 | 3,083 | 3,083 | 8,353 | 8,353 | 8,353 | 8,353 | 8,353 | 8,353 |

NOTE.—Standard errors are in parentheses. The first-stage equations for $Dw_t^*$ and $Dw_t^{**}$ are reported in table A1.

present IV estimates of the link between $Dc_t^*$ and $Dw_t^{**}$ using the predicted value of $Dw_t^{**}$ based on $Dw_{t-1}^*$ and $w_{t-1}^*$ as an instrument.[41] Under the rational expectations assumption, the wage coefficient is an estimate of $B_{cn}$, the effect of $w_t$ on $c_t$ with $\ln \lambda_t$ constant. When age is controlled for, the point estimate is $-.023$. This value is consistent with the hypothesis that $B_{cn}$ is close to zero but is subject to a standard error of .242. Columns 2–5 report estimates using $Dw_t^*$ and $w_{t-1}^*$ as instrumental variables. When age and year effects are controlled for, the coefficient estimate is .154 with a standard error of .12. This is an estimate of $B_{cn} + (B_c + B_{cn})\theta$. Since $(B_c + B_{cn})\theta$ must be greater than or equal to zero on theoretical grounds (see the discussion around [15]), this is an estimate of the upper bound for $B_{cn}$. If $(B_c + B_{cn})\theta$ is approximately zero and $B_{cn} = .154$, then the .05 estimate of the wage coefficient in the first-difference equation for labor supply is the estimate of $B_n$. The .17 estimate of $B_n$ using (14) to interpret the consumption–labor supply estimates (which is the largest of the IV estimates reported) may be corrected for nonseparability by subtracting the product of the estimates of $B_{cn}$ and $\Gamma$ ($.154 \times -.534$), leading to a revised estimate of .09. This value is in the same range as those discussed earlier. If one uses Ghez and Becker's (1975) estimate of $B_{cn} = .35$, the revised estimate of $B_n$ from the consumption–labor supply equation is essentially zero. Thus, under the assumption that $(B_c + B_{cn})\theta = 0$, the evidence from both the first-difference equation and the consumption–labor supply relation suggests that $B_n$ is between 0 and .1.

Since the evidence in table 5 and Ghez and Becker's results indicate that it is unlikely that $B_{cn}$ is negative, I take as the other extreme the possibility that $B_{cn} = 0$ and the coefficient of .154 is an estimate of $B_c\theta$. If $B_{cn} = 0$, $B_{nc} = 0$ also (by symmetry), and equations (12) and (14) may be used to interpret the results. The estimates of $B_n$ from the first-difference equation should be adjusted by adding $-(B_n/B_c)B_c\theta$ ($= .534 \times .154$) to the estimates of $B_n$ in table 1, columns 1–4. This suggests a revised estimate of .135, which is fully consistent with the estimates based on the labor supply–consumption equation.

The details of the calculations above depend on the specific point estimates chosen, but the analysis of possible biases in the two sets of estimates does not alter the conclusion that the intertemporal labor

[41] The instrumental regressions are shown in table A1 (cols. 3–5) and are based on a subsample of the observations used to estimate the consumption equation because of missing data on $Dw_t^{**}$. One may use the imputed wage variables to form instrumental variables (and dramatically increase the sample) when estimating the consumption equation because the measurement errors in $Dw_t^*$, $Dc_t^*$ are independent. Estimates obtained using $Dw_t^*$ in the consumption equation with $Dw_{t-1}^{**}$ and $w_{t-1}^{**}$ as instrumental variables are reported in cols. 3–6 and are in the neighborhood of .35. They are not discussed in the text because the inconsistency in the timing of $Dw_t^*$ and $Dc_t^*$ has the effect of biasing the coefficient of $Dw_t^*$ upward in this equation by the factor $1/a_1$.

supply elasticity is positive but small. An upward revision in both sets of estimates would be called for only if $B_{cn}$ is a large negative number. This seems unlikely, although it cannot be ruled out without further research. There is in fact a remarkable consistency between the first-difference results and the labor supply–consumption estimates despite the fact that they are based on very different information.

## V. Concluding Remarks

With allowances for sampling error and variation across specifications, the analysis of MaCurdy's first-difference equation for labor supply using the reported wage change as an instrument for the imputed wage change and the estimates obtained using consumption to proxy the marginal utility of income all suggest an intertemporal labor supply elasticity between 0 and .35. These are in the same range as MaCurdy's estimates. The marginal rate of substitution condition for consumption and labor supply proves to be a useful way of dealing with the problem of missing data on past and expected future real wages and interest rates in studying the intertemporal labor supply response. However, it must be combined with information on intertemporal consumption behavior, unless one is willing to maintain the empirically questionable assumption of separable preferences between consumption and labor supply.

Many limitations of the study have been mentioned in the text. Among the most important are the assumptions that workers freely choose hours and face exogenous wages, that the unobserved components of wages and labor supply preferences are independent, and that there is separability between preferences for husband's and wife's leisure. Biases might also arise from the use of an average wage measure. Also, since the sample is restricted to continuously married prime-age men and most of the evidence is for hourly rated workers, no general conclusions about the intertemporal labor supply response of the aggregate work force should be drawn from this study alone. One finding of the present study is that it is very important for studies of wage and income dynamics and intertemporal labor supply to take account of measurement error. Simultaneous use of multiple indicators for a given set of variables is one approach.

## Appendix

### Selection of the Samples

This Appendix lists the selection criteria for the three main samples used in the empirical analysis. Exclusion criteria common to all three samples are listed first, followed by the additional criteria affecting a particular sample. Marginal losses of both persons and person-years are given for each criterion.

# TABLE A1

## First-Stage Equations

| Explanatory Variable | Dependent Variable: $Dw_t^*$ (1) | $Dw_t^*$ (2) | $Dw_t^*$ (3) | $Dw_t^{**}$ (4) | $Dw_t^{**}$ (5) |
|---|---|---|---|---|---|
| Intercept | .0207 | .0276 | .0031 | .0276 | .0493 |
| | (.004) | (.0161) | (.0169) | (.0078) | (.0076) |
| $Dw_t^{**}$ | .295 | ... | .2967 | ... | ... |
| | (.026) | | (.0303) | | |
| $Dw_{t-1}^{**}$ | ... | .0570 | ... | ... | ... |
| | | (.0295) | | | |
| $Dw_t^*$ | ... | ... | ... | .0948 | ... |
| | | | | (.0097) | |
| $Dw_{t-1}^*$ | ... | ... | ... | ... | −.0163 |
| | | | | | (.0095) |
| $w_{t-1}^{**}$ | ... | −.0055 | .0114 | ... | ... |
| | | (.0114) | (.0118) | | |
| $w_{t-1}^*$ | ... | ... | ... | −.0066 | −.0199 |
| | | | | (.0053) | (.0052) |
| $F$-statistic | 129.1 | 1.88 | 48.1 | 56.6 | 10.19 |
| $R^2$ | .031 | .0011 | .0303 | .0276 | .0051 |
| S.E. | .227 | .225 | .228 | .135 | .137 |
| Observations | 4,004 | 10,036 | 3,083 | 3,994 | 3,994 |

| | Dependent Variable: $Dw_t^*$ (6) | $Dw_t^*$ (7) | $Dw_t^{**}$ (8) | $w_t^*$ (9) | $w_t^*$ (10) | $c_t^*$ (11) |
|---|---|---|---|---|---|---|
| Intercept | .0572 | −.0581 | .0457 | −.0629 | −.2788 | 1.556 |
| | (.1006) | (.119) | (.0712) | (.0191) | (.115) | (.145) |
| $w_t^{**}$ | ... | ... | ... | .8785 | .5109 | .1111 |
| | | | | (.015) | (.028) | (.035) |
| $w_i^{**}$ | ... | ... | ... | .1346 | .503 | .2143 |
| | | | | (.0143) | (.031) | (.0387) |
| Wife's schooling | ... | ... | ... | ... | ... | −.0019 |
| | | | | | | (.0023) |
| Parents average? | −.0046 | ... | ... | ... | ... | −.0417 |
| | (.009) | | | | | (.0109) |
| Parents rich? | −.0046 | ... | ... | ... | ... | .0008 |
| | (.009) | | | | | (.0193) |
| Year dummies? | yes | yes | yes | ... | yes | yes |

| | | | | | | |
|---|---|---|---|---|---|---|
| Black | ... | ... | ... | ... | −.0371 | −.0909 |
| | | | | | (.0101) | (.0126) |
| SMSA | ... | ... | ... | ... | .0235 | .0448 |
| | | | | | (.0094) | (.0117) |
| City ≥ 500,000 | ... | ... | ... | ... | .0394 | .0845 |
| | | | | | (.0107) | (.0133) |
| South | ... | ... | ... | ... | −.1851 | −.1702 |
| | | | | | (.0162) | (.0204) |
| West | ... | ... | ... | ... | −.2122 | −.1392 |
| | | | | | (.0174) | (.0217) |
| North Central | ... | ... | ... | ... | −.1829 | −.1858 |
| | | | | | (.0179) | (.0225) |
| Size of family | ... | ... | ... | ... | −.0079 | .0908 |
| | | | | | (.0043) | (.0054) |
| Number of children under 6 | ... | ... | ... | ... | −.0003 | −.0365 |
| | | | | | (.0067) | (.0084) |
| Number of children in family unit | ... | ... | ... | ... | .0062 | −.0148 |
| | | | | | (.0051) | (.0063) |
| Health | ... | ... | ... | ... | −.0017 | .0055 |
| | | | | | (.0119) | (.0147) |
| Age | −.0003 | .0020 | −.0003 | ... | .0141 | .0274 |
| | (.0017) | (.0021) | (.0001) | | (.0053) | (.0066) |
| Age$^2$ | ... | ... | ... | ... | −.0003 | −.0004 |
| | | | | | (.0001) | (.0000) |
| Schooling | .0043 | .0150 | ... | ... | ... | ... |
| | (.016) | (.020) | | | | |
| Schooling$^2$ | −.0001 | −.0003 | ... | ... | ... | ... |
| | (.0007) | (.0009) | | | | |
| Age · schooling | −.00008 | −.0002 | −.00001 | ... | ... | ... |
| | (.0002) | (.0002) | (.00014) | | | |
| Age · schooling$^2$ | .00000 | −.00000 | −.00000 | ... | ... | ... |
| | (.0000) | (.0000) | (.0000) | | | |
| Father's schooling | −.00053 | ... | ... | ... | ... | ... |
| | (.00095) | | | | | |
| Mother's schooling | −.00063 | ... | ... | ... | ... | ... |
| | (.0009) | | | | | |
| $F$-statistic | 2.61 | 2.13 | 4.24 | 4,742.3 | 435.7 | 83.6 |
| $R^2$ | .0054 | .0080 | .0157 | .685 | .707 | .342 |
| S.E. | .254 | .230 | .1374 | .2487 | .2406 | .2985 |
| Observations | 10,036 | 3,996 | 3,996 | 4,367 | 4,367 | 4,367 |

NOTE.—Columns 1 and 2 are the first-stage equations for $Dw_t^*$ in table 1, cols. 1–4 and cols. 5–8, respectively. Column 3 is the first-stage equation for $Dw_t^*$ in table 5, cols. 3–5. Column 4 is used to form an instrument for $Dw_t^{**}$ in table 5, cols. 6–8. Column 5 is used for $Dw_t^{**}$ in table 5, cols. 9–11. Columns 6 and 7 are used for $Dw_t^{**}$ in table 2, cols. 1–4 and cols. 5–7, respectively. Column 8 is the first-stage equation for $Dw_t^{**}$ in table 2, cols. 8–10. Columns 9 and 10 are used for $w_t^*$ in table 4, cols. 3 and 4, respectively. Column 11 is the first-stage equation for $c_t^*$ in table 4, cols. 3 and 4. Standard errors are not corrected for possible correlations among the error terms for each individual.

The sample is drawn from the 4,539 male heads of household in 1979 who are part of the 12-year Michigan Panel Study of Income Dynamics (PSID) data tape, and the analysis of losses of observations refers to these data only. In principle, each individual can contribute 11 person-years to the sample, with the first survey year (1968) excluded because of missing data on consumption and lags of the wage and hours measures. In practice, many of the 4,539 persons were not in the survey or were not heads of household in the earlier years. As is indicated in Section III, matching data on the individuals for the 1980 and 1981 survey years are obtained from the 14-year PSID data tape if the person remained in the survey. The additional observations are used if the criteria for inclusion in the sample for a particular procedure are met. The number of additional observations on person-years for 1980 and 1981 is listed below for each procedure.

The selection criteria common to all of the subsamples are as follows. (1) Only men from 36 to 60 years of age in 1979 are included. Loss: persons, 2,961. (2) Only men married to the same wife from 1968 to 1979 are included. Loss: persons, 549; person-years, 6,039. (3) Men are excluded if their wives were older than 63 in 1968. Loss: persons, 1; person-years, 11. (4) Men are excluded if they retired prior to 1979. Loss: persons, 24; person-years, 264. The remaining criteria apply to a particular person-year. (5) The man must be employed, on temporary layoff, or unemployed at the time of the survey. Loss: persons, 7; person-years, 391. (6) Annual hours worked must be positive. Loss: persons, 0; person-years, 29. (7) Annual hours worked must be less than 4,860, must not rise by more than 250 percent or fall by more than 60 percent from the preceding year, and the absolute change in hours must be less than 3,000. Loss: persons, 1; person-years, 285. (8) The antilog of the real imputed hourly wage $w_t^*$ must exceed \$0.40, must not rise by more than 250 percent or fall by more than 60 percent from the preceding year, must not change from the preceding year by more than \$13 in absolute value, and must not be missing for any other reason. Loss: persons, 1; person-years, 302.

Additional sample criteria for table 1, columns 1–4, include the following. (A.1) Observations prior to 1971 are lost because of missing data on the lag of the reported hourly wage ($w_{t-1}^{**}$), which is needed to form $Dw_t^{**}$. Loss: persons, 8; person-years, 1,875. (A.2) Both the current value and the first lag of the imputed hourly wage and annual hours worked must satisfy criteria 6, 7, and 8. Loss: persons, 1; person-years, 12. (A.3) In a given year the person must be paid by the hour and report an hourly wage ($w_t^*$). Loss: persons, 319; person-years, 4,190. (A.4) The antilog of the reported hourly wage ($w_t^{**}$) must not be affected by the upper bound of \$9 in nominal terms for surveys prior to 1978. Loss: persons, 12; person-years, 69. (A.5) The antilog of the reported wage must exceed \$0.40 and must not rise by more than 250 percent, fall by more than 60 percent, or change by more than \$13 in absolute value from the preceding year. Loss: persons, 1; person-years, 7. (A.6) Both the current value and the first lag of the reported wage must satisfy criteria A.3, A.4, and A.5. Loss: persons, 60; person-years, 451.

The surviving sample consists of 594 persons and 3,463 person-years. An additional 541 person-years from the survey years 1980 and 1981 bring the total sample to 4,004 person-years.

Additional sample criteria for table 2, columns 1–4, include the following. (B.1) Both the current value and the first lag of the imputed hourly wage and annual hours worked must satisfy criteria 6, 7, and 8. Loss: persons, 1; person-years, 12. (B.2) Data must be available for years of schooling, parents' financial status during the individual's childhood, father's schooling, and mother's schooling. Loss: persons, 131; person-years, 1,289.

The surviving sample contains 863 persons and 8,766 person-years. An additional 1,270 person-years from 1980 and 1981 bring the total sample to 10,036 person-years.

Additional sample criteria for table 4 include the following. (C.1) All observations from 1969 are excluded because of missing data on the reported hourly wage. Observations from 1973 are excluded because of missing data on food consumption. Loss: persons, 4; person-years, 1,863. (C.2) Data must be available for years of schooling, age, father's schooling, mother's schooling, economic status of the person's family during his childhood, wife's schooling, and the other demographic characteristics used in the first-stage equations for the wage and consumption (see table A1, cols. 10 and 11). Loss: persons, 4; person-years, 73. (C.3) The reported hourly wage $w_t^{**}$ must be available and satisfy criteria A.2, A.3, and A.4. As a result of experiments with estimation of the wage fixed effects using dynamic wage models (not reported in the paper), persons lacking valid data on both the current value and the first lag of the reported wage for at least one survey year are also excluded. Loss: persons, 390; person-years, 4,265. (C.4) Real food consumption must not rise by more than 250 percent or fall by more than 60 percent from the preceding year, and it must not be missing for any other reason. Loss: persons, 0; person-years, 66.

The surviving sample contains 597 persons and 3,800 person-years. An additional 567 person-years from 1980 and 1981 bring the total to 4,367 person-years.

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