The Effect of Income Taxation on Labor Supply in the United States

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ABSTRACT

The effect of income taxation on the labor supply of prime age married men and women in the United States is examined using econometric methods similar to those used in the influential work of Jerry Hausman. Male labor supply is found to be relatively invariant to the net wage and virtual income in all specifications estimated. The estimated impact of taxation on the labor supply of married women depends critically on the method used to estimate the labor supply function. Estimated net wage and virtual income elasticities are considerably larger when data on nonparticipants are included than when the estimation is conducted conditional on hours being greater than zero.

I. Introduction

The influential work of Jerry Hausman (1981) challenged the prevailing wisdom that the U.S. income tax system has little effect on hours of work by prime age married men. Like many previous analysts, Hausman estimated an uncompensated wage elasticity for males which was very close to zero. However, unlike previous analysts, Hausman estimated a large negative income elasticity. This finding, combined with his correctly noting that progressive income taxation combines reductions

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in the net wage with implicit lump sum subsidies for upper bracket taxpayers, resulted in his producing estimates of the excess burden and reduction in labor supply due to income taxation which were much larger than had previously been found. Hausman's estimates for women also showed large reductions in labor supply due to taxation, but were closer in magnitude to previous research than were his estimates for men.

Although Hausman's estimates have been both influential and controversial, there has been very little work investigating how sensitive his results are to the particular specification chosen. The primary purpose of this paper is to perform such a sensitivity analysis. Using a more recent (1983 rather than 1975) wave of the same dataset used by Hausman, I find that male labor supply is very price and income inelastic in all variants of the model estimated, although the uncompensated wage elasticity is positive and somewhat larger in magnitude than that estimated by Hausman. The results for women are mixed. Elasticity estimates are similar to those estimated by Hausman when the model is estimated using the entire sample, but labor supply is estimated to be considerably less elastic when the sample is truncated.

Section II of this paper presents the version of the labor supply model which I estimate. Section III briefly describes the individual income tax system existing in the United States in 1983, the year of the data set which I use in estimation. In section IV, a brief description of the data is provided. Section V presents the parameter estimates.

II. The Basic Model

The model in this section is essentially the one estimated by Hausman (1981), although several variants of this model were estimated. We consider only the case of married couples in a single period setting. Husbands and wives maximize separate utility functions defined over consumption and his or her own hours of work. All individuals are assumed to have a fixed gross wage, and a fixed amount of unearned income. Each individual's optimization problem is of the form:

$$(1) \quad \max_{\{C,H\}} \; u(C, H)$$
$$\text{s.t. } C = Y + w \cdot H - R(I)$$

where C is consumption (which is the numeraire), H is hours of work, $u(C, H)$ is the utility function, w is the gross wage, Y is unearned income, and R is tax payments. The partial derivative of u with respect to H is negative. It is assumed that husbands ignore the earnings potential and labor supply decisions of their wives in deciding on their own hours of work. Wives, in contrast, are assumed to take their husbands' labor supply decisions as given in making their own decisions. Therefore, for wives the unearned income variable is equal to the sum of asset income and their husbands' earnings. Taxable income, I, is equal to the sum of earned and unearned income ($wH + Y$) minus the values of deductions (D) and exemptions (E) allowed to the filing unit. Tax payments are a piecewise linear function of taxable income:

$$(2) \quad R(I) = R(I_j) + t_j(I - I_j)$$

where j is the index of the tax bracket for someone with taxable income I, t_j is the marginal tax rate in bracket j, and I_j is the lower taxable income limit for bracket j. Substituting the tax function and the definition of taxable income ($I = wH + Y - D - E$) into the budget constraint yields the linearized budget constraint for bracket j:

$$(3) \quad C = w(1 - t_j)H + (1 - t_j)Y + t_j(D + E) + [t_jI_j - R(I_j)].$$

Although each spouse's choice of H and C determines the segment of the budget constraint he is on, behavior is locally equivalent to that which would arise from maximizing utility subject to a linear budget constraint with a relative price of leisure equal to $w(1 - t_j)$ and "virtual income" equal to $(1 - t_j)Y + t_j(D + E) + [t_jI_j - R(I_j)]$. The third term in virtual income is an adjustment for the fact that $t_j(wH + Y)$ overstates the actual amount of taxes paid. That term, $[t_jI_j - R(I_j)]$, is the difference between the amount of taxes that an individual on segment j would pay if he faced a proportional tax with rate t_j, t_jI_j, and the taxes he actually pays, $R(I_j)$. On a consumption-hours graph, virtual income for a person on a given segment is equal to the vertical intercept (at zero hours of work) of the person's budget constraint linearized through that segment. Thus, the optimization process results in a labor supply (leisure demand) function which is locally a function of the net wage and virtual income.

Preferences, as represented by the indirect utility function, are given by:

$$(4) \quad v(w, Y) = e^{\beta w}\left(Y + \frac{\alpha}{\beta}w - \frac{\alpha}{\beta^2} + \frac{\gamma + \varepsilon}{\beta}\right)$$

where α, β, and γ are parameters, and ε is a random variable.[1] When the budget constraint is linear, the implied desired labor supply function is:

$$(5) \quad h^*(w, Y) = \gamma + \alpha w + \beta Y + \varepsilon.$$

1. In the estimation, γ is specified to be a linear function of demographic variables.

Heterogeneity in preferences which are due to factors unobserved by the econometrician are represented by ε. Although ε is fixed for any given individual, it is assumed to have a mean zero normal distribution ($N(0, \sigma_\varepsilon^2)$) in the population. The wage coefficient (α) being non-negative and the income coefficient (β) being non-positive are sufficient conditions for the compensated labor supply wage elasticity to be non-negative. When the budget constraint is nonlinear, Equation (5) holds conditional on location on a given budget segment, with the appropriate net wage and virtual income replacing the gross wage and unearned income. The lower limit of desired hours of work is zero; I have assumed, somewhat arbitrarily, that the upper limit of desired hours is 8,760 (the number of hours in a year).[2]

Observed hours of work (h) is assumed to be equal to desired hours of work plus an additive stochastic term (ν) representing either sample measurement error or the failure of the worker to find a job requiring hours exactly equal to the desired quantity:

$$(6)\quad h = h^* + \nu.$$

The deviation between observed and desired hours is assumed to be independent of ε and to have a mean zero normal distribution ($N(0, \sigma_\nu^2)$). Following Hausman (1981), I assumed that observed hours is equal to zero whenever desired hours is zero.

The contribution of each observation with observed hours of work greater than zero to the likelihood function is the density of h, $f(h)$:

$$(7)\quad f(h) = \sum_{i=1}^{m} \left[\int_{\varepsilon_{li}}^{\varepsilon_{ui}} \frac{1}{\sigma_\nu} \phi\left(\frac{[h - (\gamma + \alpha w_i + \beta Y_i + \varepsilon)]}{\sigma_\nu} \middle| \varepsilon \right) \frac{1}{\sigma_\varepsilon} \phi\left(\frac{\varepsilon}{\sigma_\varepsilon}\right) d\varepsilon \right]$$
$$+ \sum_{i=1}^{m-1} \left[\Phi\left(\frac{\varepsilon_{l(i+1)}}{\sigma_\varepsilon}\right) - \Phi\left(\frac{\varepsilon_{ui}}{\sigma_\varepsilon}\right) \right] \frac{1}{\sigma_\nu} \phi\left(\frac{h - h_i}{\sigma_\nu}\right)$$
$$+ \left[1 - \Phi\left(\frac{\varepsilon_{8760}}{\sigma_\varepsilon}\right) \right] \frac{1}{\sigma_\nu} \phi\left(\frac{h - 8760}{\sigma_\nu}\right)$$

where $\phi(\cdot)$ is the standard normal density function, $\Phi(\cdot)$ is the standard normal cumulative distribution function, w_i is the net wage on segment i of the budget constraint, Y_i is virtual income on segment i of the budget constraint, and h_i is hours of work at the kink point between segments i and $i + 1$ of the budget constraint. The first term is the joint density of

2. Blomquist (1983) also sets the maximum value of desired hours to 8,760.

desired hours being in the interior of one of the segments of the budget constraint and observed hours being equal to h, the second term is the joint density of desired hours being at one of the kink points and observed hours equal to h, and the last term is the joint density of desired hours being equal to 8,760 (the maximum possible value) and observed hours being equal to h. Each ε_{ui} is the maximum value of ε which results in desired hours of work being on segment i of the budget constraint, while each ε_{li} is the minimum value of ε which results in desired hours of work being on segment i (ε_{8760} is the minimum value of ε which results in desired hours being 8760). The values of ε_{ui} and ε_{li} depend upon the parameters of the labor supply function and hours at the kink points of the budget constraint:

$$(8) \quad \varepsilon_{ui} = h_i - (\gamma + \alpha w_i + \beta Y_i)$$

$$\varepsilon_{li} = h_{i-1} - (\gamma + \alpha w_i + \beta Y_i)$$

Hours of work at the kink points depends on the structure of the tax system and the values of the person's gross wage and unearned income:

$$(9) \quad h_i = \frac{(I_i - Y)}{w}$$

Moffitt (1986) shows that each element of the summation in the first part of Equation (7) can be transformed into the difference between two evaluations of a normal distribution function times a single evaluation of a normal density function; this transformation greatly facilitates estimation.

The truncation of hours of work at zero was handled in two different ways. One set of estimates was obtained by conditioning the above probability calculations on the event that observed hours is greater than zero. In addition to this truncated specification, a censored version was also estimated in which the zero hours observations were included. Since wage data were not available for those reporting zero hours of work, a wage imputation equation was first estimated. The imputed wages (for both workers and those with zero hours) were then treated as though they were the true wages in estimating the labor supply function.[3] The contribution to the likelihood of an observation with observed hours of work equal to zero is the probability that $h = 0$:

3. As noted by Heckman and MaCurdy (1981), a better procedure is to integrate over the wage distribution in estimating the labor supply function. That approach is not taken here due to its computational complexity. Details of the procedure used to impute gross wages are presented in the appendix.

$$(10)\quad \text{prob}(h = 0) = \int_{-\infty}^{\varepsilon_{l1}} \frac{1}{\sigma_\varepsilon}\phi\left(\frac{\varepsilon}{\sigma_\varepsilon}\right) d\varepsilon$$

$$+ \sum_{i=1}^{m}\left[\int_{\varepsilon_{li}}^{\varepsilon_{ui}}\int_{-\infty}^{-(\gamma+\alpha w_i+\beta Y_i+\varepsilon)} \frac{1}{\sigma_\nu}\phi\left(\frac{\nu}{\sigma_\nu}\right)\frac{1}{\sigma_\varepsilon}\phi\left(\frac{\varepsilon}{\sigma_\varepsilon}\right) d\nu\, d\varepsilon\right]$$

$$+ \sum_{i=1}^{m-1}\left[\Phi\left(\frac{\varepsilon_{l(i+1)}}{\sigma_\varepsilon}\right) - \Phi\left(\frac{\varepsilon_{ui}}{\sigma_\varepsilon}\right)\right]\Phi\left(\frac{-h_i}{\sigma_\nu}\right)$$

$$+ \left[1 - \Phi\left(\frac{\varepsilon_{8760}}{\sigma_\varepsilon}\right)\right]\Phi\left(\frac{-8760}{\sigma_\nu}\right).$$

The first term is the probability that desired hours of work is zero. The remaining three terms correspond to the three terms in the earlier expression, with observed hours of work now equal to zero and the measurement error term integrated over from minus infinity to minus desired hours of work. Each double integral in the second term can be evaluated as the difference between two evaluations of a bivariate normal distribution function after a change of variable from ν to $(\nu + \varepsilon)$. In the truncated case, using only observations with positive hours of work, the log likelihood function is the sum over all observations of the log of (7) divided by one minus (10). In the censored case, where all observations are included, the log likelihood function is the sum of the logs of (7) and (10), as appropriate.

In addition to maximum likelihood estimation of the full two error model, two simpler variants were also estimated. In one version of the model estimated, the only source of stochastic variation allowed for was measurement/optimization error in hours of work (ν). Under this specification, which was first introduced by Wales and Woodland (1979), we are assuming that desired hours of work for each individual could be predicted without error if we knew the true parameter values. While this assumption is not very appealing, this version of the model has the advantage of being computationally much simpler than the two error specification. In order to avoid a deterministic specification for those with desired hours predicted to be zero, I specify that observed hours is equal to desired hours plus ν even when desired hours is zero. Because of this, the measurement-error-only specification is not strictly nested within the two error model. Censored and truncated versions of the measurement error model were estimated using the maximum likelihood method.

A heterogeneity-only variant of the model was also estimated. In kinked budget constraint models, there is a positive probability of each individual desiring to locate at one of the convex kink points. If observed and desired hours are equal, as they are assumed to be in the heterogeneity-only model, then we should observe some piling up of observations exactly at the convex kink points. Inspection of Equation (7) reveals that the contribution to the likelihood collapses to that of ordinary least squares when there is no measurement error and no observations observed at the kink points (or the upper limit of hours). Since we very rarely observe sample members exactly at kink points, in practice maximum likelihood estimation of the heterogeneity-only model would be identical to least squares with no attempt to control for the endogeneity of taxes. In one application of a heterogeneity-only specification, Moffitt and Nicholson (1982) avoid this problem by classifying sample members observed within a band around a kink point as being at the kink.

The failure to observe observations exactly at the kink points suggests that the heterogeneity-only model is misspecified. Either deviations between observed and desired hours of work, as posited in the two error term model presented above, or errors in imputing the kink points of sample members' budget constraints may be the source of the problem. Much of the information needed to impute the location of the kink points, such as the amount of itemized deductions, must be imputed itself. For this reason, we may incorrectly classify an observation which is actually at a kink point as being on the interior of a segment. The maximum likelihood estimator of the heterogeneity model is particularly vulnerable to this problem, since it relies on accurate classification of the kink point observations to correct for the endogeneity of the marginal tax rate.

An alternative to the maximum likelihood estimator for the heterogeneity-only model which does not depend as heavily on correct classification of the kink point observations is instrumental variables. Instrumental variables is an appropriate estimation technique for models of labor supply with nonlinear budget constraints as long as we can correctly impute the net wage and virtual income arguments corresponding to desired hours of work. In the case of labor supply, the gross wage and unearned income can serve as instruments for the net wage and virtual income. As MaCurdy (1983) points out, instrumental variables is not consistent when there is measurement error in hours of work since this can induce measurement error in the net wage and virtual income (with a mean which depends on the true value of desired hours).[4] However, some Monte-Carlo experiments suggest that a limited amount of measurement error is not a serious problem (Triest 1987b). Thus, instrumental variables estimation of the heterogeneity-only model is an interesting, and compu-

4. This is true as long as hours of work are used in imputing the kink points of the budget constraint.

tationally simple, alternative to maximum likelihood estimation of the two error term specification. In estimating the heterogeneity-only model for women, the generalized method of moments was used to estimate an instrumental variables version of the Tobit model (Avery and Hotz 1985).

III. The 1983 U.S. Individual Income Tax System

In 1983, the U.S. federal personal income tax consisted of a progressive 14 bracket system. Table 1 presents the taxable income ranges, marginal tax rates, and implicit lump sum transfers associated with each bracket for married couples filing joint returns. In addition to the income tax, workers contributed 6.7 percent of their earnings (up to $35,700) in social security payments; employers made a matching contribution of 6.7 percent. Since an individual's retirement benefits are linked to past social security tax payments, incorporation of social security into the budget constraint is problematic. I treated the employee contribution as a pure tax, and ignored the employer contribution. Additional segments were added to the budget constraints of both husbands and wives to incorporate this tax.

Several special features of the tax code need to be incorporated into the budget constraint. I generally follow the procedures followed by Hausman (1981), although there are some minor differences. The earned income tax credit, which was designed to reduce the regressivity of the social security tax, gave couples filing jointly (or single heads of households) who were entitled to claim an exemption for a dependent child, and who had adjusted gross income of less than $10,000, a credit equal to 10 percent of the first $5,000 of earned income. The credit was reduced at a rate of 12.5 percent for adjusted gross income over $6,000. Extra brackets were added to the husbands' budget constraints to take account of the credit. Since, under the assumption that women take their husbands' earnings as given, few wives were eligible for the credit, the budget constraints of wives have not been adjusted.

The budget constraints of wives were adjusted to incorporate a special deduction available to married couples when both spouses work. This deduction was equal to 10 percent of the first $30,000 of the earned income of the spouse with the smaller earnings.

In 1983, married couples were entitled to an exemption of $2,000 plus $1,000 per dependent. I assumed that all children less than 18 years old could be claimed as dependents. If sample members indicated that they used the standard deduction, a deduction of $3,400 (the zero bracket amount in 1983) was built into their budget constraints. If they instead indicated that they itemized deductions, they were assigned a deductions

Table 1
1983 U.S. Personal Income Tax

Taxable Income Range (dollars)	Marginal Tax Rate	Implicit Lump Sum Subsidy due to the Tax System
0–3,400	0	0
3,400–5,500	.11	374
5,500–7,600	.13	484
7,600–11,900	.15	636
11,900–16,000	.17	874
16,000–20,200	.19	1,194
22,200–24,600	.23	2,002
24,600–29,900	.26	2,740
29,900–35,200	.30	3,936
35,200–45,800	.35	5,696
45,800–60,000	.40	7,986
60,000–85,600	.44	10,386
85,600–109,400	.48	13,810
over 109,400	.50	15,998

Note: This table is correct only for married couples filing jointly who do not itemize deductions. Taxable income in column one should be interpreted as being gross of deductions. The third column shows the $[t_j I_j - R(I_j)]$ adjustment to unearned income which must be made in linearizing the budget constraint. In 1983 employees paid a social security tax of 6.7 percent on earnings of up to \$35,700. For workers with earnings over this amount, \$2,391.90 (the maximum FICA payment) needs to be subtracted from unearned income in calculating virtual income. Other adjustments to the tax rates and virtual income are explained in the text.

value equal to the average itemized deductions (excluding the state tax payments deduction) within their adjusted gross income class in published Internal Revenue Service tables (Internal Revenue Service 1985, p. 50).[5] This is similar to the procedure followed by Hausman (1981), although he also had to impute itemization status due to lack of data.

All but ten states imposed their own income taxes in 1983.[6] State marginal tax rates ranged as high as 16 percent (the top rate in Minnesota),

5. In another paper (Triest 1987a), I allow for the endogeneity of deductions and itemization status.

6. The 1982–83 edition of *Significant Features of Fiscal Federalism* (U.S. Advisory Commission on Intergovernmental Relations 1984) was used as the source of information on the state tax systems.

although the possibility of deducting state income tax payments from federal taxable income (and in some states deducting federal tax payments from state taxable income) reduce the effective marginal tax rates considerably. I assumed that couples who itemized deductions on their federal returns also itemized on their state returns, and claimed the same amount of deductions. Several states link the value of the standard deduction to adjusted gross income; I ignored this feature of the tax system and assumed that the maximum value of the standard deduction was used for all nonitemizers. In the United States, state tax payments may be deducted from federal taxable income by those who itemize. The effective marginal tax rate decreases from $(t_f + t_s)$, where t_f is the federal marginal tax rate and t_s is the state marginal tax rate, to $(t_f + t_s - t_f t_s)$ for itemizers due to this deduction. In 1983, sixteen states allowed a deduction for federal tax payments. Although this deduction usually applies to federal taxes on the previous year's income, I have treated it as though it applies to taxes on the current year's income; this is necessary to fit this deduction into a single period framework. For itemizers in states which allow the deductibility of federal tax payments, the marginal tax rate is $(t_f + t_s - 2t_f t_s)/(1 - t_f t_s)$ (Feenberg and Rosen 1986). Following Hausman (1981), rather than add additional segments to the sample members' budget constraints, I averaged the state tax rates over the segments created by the federal tax. The maximum taxable income limit of the highest tax brackets was set at $999,999 in doing this averaging.

The upper limit to the social security tax and the phase out of the earned income tax credit create nonconvex regions of the budget set. Since Hausman (1981) found that his results were relatively insensitive to using a convex approximation of the budget set, I followed his procedure of treating the convex hull of the budget set as the effective budget set in estimating the model.

IV. The Dataset Used

Wave XVII of the Panel Study of Income Dynamics (PSID) is the source of data for the empirical work described in this paper (Institute for Social Research 1986). Data for this wave were collected in 1984, but pertain to calendar year 1983. Observations which were part of the Survey of Economic Opportunity were excluded (eliminating 3,189 of 6,918 family records). Only married couples were selected in the sample used for estimation (eliminating 1,495 family records). Other sample restrictions, in the order in which they were applied, include restricting the husband's age to between 25 and 55 (670 records eliminated), eliminating observations where the husband reported that he was disabled (22 records), restricting the wife's age to between 25 and 55 (104 records), eliminating observations where the wife reports that she is disabled (9 records), eliminating observations where the couple reports self-employment or farm income (289 records), and elimination of observations with missing data (133 records). Finally, observations were eliminated if average hourly earnings for either spouse was greater than $50 or less than $1 (28 records eliminated), or if unearned income was negative (1 record). This selection procedure resulted in a dataset with 978 observations.

Hours of work is measured in terms of hours on all jobs held in 1983. This variable was constructed by summing weeks worked by average hours worked per week over all jobs. Average hourly earnings, calculated by dividing labor income by hours of work, is the measure of the wage rate used in the estimation. Unearned income was calculated by summing income from rent, dividends, interest, trust funds, and royalties.[7]

Descriptive statistics for the variables used in the labor supply estimation are presented in Table 2. Since the data is fairly recent (1983), the female labor force participation rate, .73, is higher than typically seen in female labor supply studies using U.S. data. It is also interesting to note that the subsample of couples in which the wife works does not seem to vary significantly from the full sample in terms of number of young children, family size, or age of either spouse.

V. Estimation Results

Results from estimation of the models described in Section II are presented in Table 3 for husbands. The results for wives are presented in Tables 4, 5, and 6. The elasticities reported in the tables are valid only for local movements along a given budget segment. The demographic variables included in all specifications include the number of children less than six years old, family size, and a dummy variable set equal to one if the sample member reported a health problem which limited the amount or type of work he or she could perform. The age variable is set to zero for those less than 45 years old, and is equal to years of age minus 45 for those who are 45 or older.[8] Approximately 1.3 percent of the husbands

7. Hausman (1981) reports that he calculated nonlabor income by attributing an 8 percent return to financial assets. He also includes housing equity as a separate variable in the labor supply function. These differences in the treatment of unearned income may account for some of the differences between our estimates.

8. In choosing the demographic variables, I tried to maintain a reasonable degree of comparability to Hausman (1981).

Table 2
Sample Descriptive Statistics

	Full Sample		Subsample with Working Women	
Variable Name	Mean	Std. Dev.	Mean	Std. Dev.
Unearned income	1,187	3,661	957	3,177
Children < 6 years	.57	.77	.50	.71
Family size	3.72	1.16	3.61	1.14
Husband's				
Hours of work	2,113	602	2,097	586
Participation rate	.99		.99	
Hourly earnings	12.2	6.1	11.4	5.5
Age	37.6	8.13	37.3	8.02
Age–45 (45 = 0)	1.15	2.65	1.03	2.48
Bad health	.08		.07	
Wife's				
Hours of work	1,057	879	1,445	703
Participation rate	.73			
Hourly earnings			7.5	4.6
Age	35.4	7.7	35.0	7.6
Age–45 (45 = 0)	.67	1.98	.61	1.85
Bad health	.09		.08	
Sample size	978		715	

worked zero hours during 1983. Since this percentage is very small, these observations were discarded in estimating the male labor supply functions.

Maximization of the log likelihood functions for both men and women was performed using the modified scoring algorithm developed by Berndt, Hall, Hall, and Hausman (1974).[9] A variety of starting values were tried to

9. The "Gauss" software for IBM compatible personal computers was used for the maximum likelihood computations. Some of the iterations used the Broyden-Fletcher-Goldfarb-Shanno algorithm rather than BHHH. Standard errors were computed using the outer product of the score vector approximation of the information matrix. The instrumental variables standard errors were computed using the heteroskedasticity consistent estimator proposed by White (1982).

The truncated instrumental variables computations were performed using the "Hotztran" software package. An asymptotically optimal weighting matrix (based on initial estimates) was used in this estimation.

guard against estimating parameter values associated with local optima (which were found in some cases). Achieving convergence often proved to be fairly difficult. A principal reason for this appeared to be the difficulty in distinguishing between the effects of the heterogeneity and measurement error stochastic terms.

The first column of Table 3 shows results from instrumental variables estimation of the heterogeneity-only model for husbands. The wage coefficient is negative, although small in magnitude and statistically insignificant, while the income coefficient is small and positive. The compensated wage elasticity is negative, although small in magnitude. Reestimating with the income coefficient set to zero (Column 2) results in an estimate of the wage coefficient which is small but positive.

In estimating the other variants of the model, it was necessary to constrain the virtual income coefficient (β) to be nonpositive in order to avoid situations where the likelihood function is not well defined.[10] This constraint turned out to be binding in both the measurement-error-only model (Column 3) and in the two error specification (Column 4). The appropriateness of the restriction can be investigated by constructing a Lagrange multiplier test of the hypothesis that the virtual income coefficient is equal to zero. This yields a test statistic of .625 (χ_1^2) for the dual error specification. Thus, at any significance level less than .43, we fail to reject the hypothesis that the virtual income coefficient is equal to zero. It is very surprising that the income effect is close to zero, since Hausman (1981) found that on average the income effect was quite large in magnitude.

The coefficients resulting from estimation of the measurement-error-only and two error models are virtually identical. They are also both quite close to the constrained instrumental variables estimates. The uncompensated (and compensated, given that the income effect is equal to zero) wage effect is small but positive. A one-dollar increase in the net wage rate would result (locally) in approximately 14 additional hours of work (in the two error case). The other parameter estimates have reasonable values.

As in most applications of the two error term model (as surveyed by Moffitt 1986), the estimated standard deviation of the heterogeneity error is larger than the estimated standard deviation of the measurement/optimization error term. Neither standard deviation is estimated very

10. In the measurement-error-only model, there may not be a unique value of desired hours when the compensated wage elasticity is negative. In the two error model, there may be negative probabilities of locating at the kink points. When $\beta > 0$, ε_{1i} may be less than $\varepsilon_{u(i-1)}$ unless α is sufficiently large. Since α always remained positive in the course of maximizing the log likelihood, the issue of constraining α did not arise. The $\beta > 0$ constraint was imposed by reparameterizing β as $\beta = -\eta^2$.

Table 3
Male Labor Supply Estimation Results (standard errors in parentheses)

Dependent Variable: Husband's annual hours of work

Variable	Instrumental Variables: Unconstrained	Instrumental Variables: Constrained	Maximum Likelihood: Measurement Error Only	Maximum Likelihood: Dual Additive Errors	Maximum Likelihood: Random Income Coefficient
Constant	2,175.6	2,151.5	2,122.2	2,092.9	2,121.7
	(88.3)	(89.6)	(65.9)	(66.9)	(65.0)
Net wage	−3.8	6.3	11.5	14.1	11.6
	(7.7)	(6.2)	(5.0)	(6.2)	(6.2)
Virtual income ($1,000's)	11.8	0.0**	0.0**	0.0**	
	(5.0)				
Age–45 (45 = 0)	−8.5	−6.3	−7.4	−7.6	−7.5
	(7.2)	(7.4)	(7.2)	(7.1)	(7.1)
Children < 6	−5.3	−9.6	−9.5	−10.1	−9.3
	(26.9)	(27.5)	(23.9)	(24.3)	(24.0)
Family size	−6.9	−7.8	−11.1	−9.0	−11.3
	(18.3)	(18.8)	(15.7)	(15.6)	(15.5)

Bad health	−283.4	−284.9	−280.5	−279.1	−282.4
	(72.5)	(76.4)	(64.1)	(67.1)	(64.6)
σ_ε	535.1	549.5		498.5	
				(223.4)	
σ_ν			545.6	234.5	546.2
			(8.4)	(465.0)	(8.7)
μ_β					0.147
					(*)
σ_β					0.013
					(*)
Mean β (truncated)					−0.0009
Standard deviation of β (truncated)					0.0060
Uncompensated wage elasticity***	−0.02	0.03	0.05	0.06	0.05
Income elasticity***	0.07	0**	0**	0**	0**
Log likelihood			−7,452.1	−7,451.0	−7,452.1
Number of observations: 965					

* This standard error could not be accurately computed due to near perfect collinearity between the rows corresponding to μ_β and σ_β in the approximation to the information matrix.

** This parameter was constrained to be less than or equal to zero.

*** Elasticities are evaluated at the observed mean net wage ($9.07) and virtual income ($11,662) values.

Table 4
Female Labor Supply Estimation Results: Censored Specifications (standard errors in parentheses)

Dependent Variable: Wife's annual hours of work

Variable	Maximum Likelihood		
	Measurement Error Only	Dual Additive Errors	Random Income Coefficient
Constant	1,003.6	1,322.1	994.5
	(222.7)	(23.1)	(265.6)
Net wage	306.6	235.2	265.6
	(40.8)	(14.2)	(19.7)
Virtual income	−21.9	−22.1	
($1,000's)	(3.6)	(2.6)	
Age−45 (45 = 0)	−66.6	−68.0	−63.4
	(19.1)	(15.9)	(13.9)
Children < 6	−374.3	−338.9	−333.9
	(59.9)	(32.2)	(32.6)
Family size	−213.9	−199.6	−154.3
	(37.0)	(15.7)	(20.9)
Bad health	−350.8	−206.4	−144.6
	(129.9)	(47.8)	(79.6)
σ_ε		777.9	
		(26.0)	
σ_ν	1,010.6	663.6	984.4
	(35.9)	(18.7)	(29.4)
μ_β			−23.1
			(2.7)
σ_β			9.1
			(2.8)
Mean β (truncated)			−23.2
Standard deviation of β (truncated)			8.9
Uncompensated wage elasticity*	1.12	0.86	0.97
Income elasticity*	−0.31	−0.31	−0.33
Log likelihood	−6,203.4	−6,185.0	−6,196.9
Number of observations: 978			

* Elasticities are evaluated at the observed mean net wage ($5.30) and virtual income ($20,449) of participants.

Table 5
Female Labor Supply Estimation Results: Truncated Specifications Using Reported Gross Wages (standard errors in parentheses)

Dependent Variable: Wife's annual hours of work

		Maximum Likelihood		
Variable	Truncated Instrumental Variables	Measurement Error Only	Dual Additive Errors	Random Income Coefficient
Constant	2,194.3	1,891.7	1,870.4	1,882.2
	(136.5)	(198.3)	(200.4)	(200.0)
Net wage	7.0	70.9	73.1	76.7
	(12.7)	(31.9)	(32.1)	(32.3)
Virtual income	−13.2	−10.5	−11.1	
($1,000's)	(3.1)	(2.8)	(2.9)	
Age–45 (45 = 0)	−30.5	−30.6	−31.4	−33.6
	(15.3)	(17.5)	(17.9)	(18.5)
Children < 6	−293.2	−297.5	−301.9	−300.6
	(53.4)	(45.4)	(47.3)	(45.8)
Family size	−109.4	−119.2	−112.3	−114.1
	(28.8)	(27.6)	(28.2)	(28.4)
Bad health	−138.1	−124.6	−129.3	−118.0
	(116.9)	(102.9)	(106.0)	(104.9)
σ_ε	714.3		295.4	
	(23.2)		(724.7)	
σ_ν		710.3	651.3	700.0
		(26.2)	(285.6)	(36.4)
μ_β				−12.1
				(3.2)
σ_β				5.1
				(7.1)
Mean β (truncated)				−12.1
Standard deviation of β (truncated)				4.9
Uncompensated wage elasticity*	0.03	0.26	0.27	0.28
Income elasticity*	−0.19	−0.15	−0.16	−0.17
Log likelihood		−5,638.5	−5,638.5	−5,638.1
Number of observations: 715				

* Elasticities are evaluated at the observed mean net wage ($5.30) and virtual income ($20,449) of participants.

Table 6
Female Labor Supply Estimation Results: Truncated Specifications Using Imputed Gross Wages (standard errors in parentheses)

Dependent Variable: Wife's annual hours of work

Variable	Truncated Instrumental Variables	Maximum Likelihood: Measurement Error Only	Maximum Likelihood: Dual Additive Errors	Maximum Likelihood: Random Income Coefficient
Constant	1,896.7	1,895.5	1,886.6	1,899.3
	(210.1)	(187.7)	(162.2)	(194.3)
Net wage	57.1	67.5	69.5	72.7
	(31.5)	(28.0)	(27.6)	(29.2)
Virtual income	−15.2	−10.5	−10.9	
($1,000's)	(3.7)	(2.8)	(2.8)	
Age−45 (45 = 0)	−28.1	−32.2	−31.1	−32.5
	(16.2)	(16.9)	(17.6)	(18.0)
Children < 6	−323.1	−288.1	−303.4	−302.1
	(61.2)	(44.1)	(45.7)	(45.6)
Family size	−91.3	−115.6	−112.6	−111.0
	(30.0)	(27.4)	(26.1)	(28.2)
Bad health	−123.9	−130.5	−135.6	−139.9
	(120.0)	(100.8)	(104.2)	(102.9)
σ_ε	747.3		301.2	
	(44.1)		(188.8)	
σ_ν		708.9	648.5	697.9
		(26.4)	(89.7)	(37.0)
μ_β				−12.2
				(3.2)
σ_β				5.2
				(7.4)
Mean β (truncated)				−12.3
Standard deviation of β (truncated)				5.0
Uncompensated wage elasticity*	0.21	0.25	0.25	0.27
Income elasticity*	−0.22	−0.15	−0.15	−0.17
Log likelihood		−5,638.4	−5,638.1	−5,638.0
Number of observations: 715				

* Elasticities are evaluated at the observed mean net wage ($5.30) and virtual income ($20,449) of participants.

precisely, and the estimated correlation between the estimators of the two parameters is $-.99$. This is not very surprising; as Moffitt (1986) points out, the only thing distinguishing between the effects of the two error terms is the degree of clustering of observations near the kink points of the budget constraint. The heterogeneity error tends to produce clustering at the kink points of a convex budget set, while the measurement/optimization error results in observations with desired hours at one of the kink points being moved away from that point. Since the U.S. tax system produces budget constraints with many kink points spaced fairly close together, and the estimated value of the standard deviation of the measurement/optimization error term is fairly large, one would expect to have difficulty distinguishing between the two sources of stochastic variation at the margin.

Since Hausman (1981) modeled heterogeneity in preferences as a random income coefficient rather than as an additive disturbance, this specification was also estimated; the results from this estimation are reported in Column 5. Following Hausman (1981), β was specified as having a truncated (from above at zero) normal distribution.[11] The distribution of β is estimated to be nearly degenerate at zero. In effect, the data force this specification to collapse to the measurement-error-only model. It is difficult to determine why the results here differ so markedly from those of Hausman (1981). However, it should be noted that the data used here are from a different year than those used by Hausman. In addition, the specification is similar but not identical. Computational problems may also play a role. It is very difficult to accurately evaluate the likelihood and score functions when the mean of the untruncated β distribution is many times its variance. A further problem is that the likelihood function appears to be relatively insensitive to the parameters of the β distribution.

The overall impression conveyed by these estimates is that economic variables have relatively little effect on hours of work by prime age married males in the U.S. It is interesting to compare these results to those of Hausman (1981). Since Hausman used data from 1975, his coefficient estimates must be adjusted for the increase in the price level between that year and 1983 before making this comparison.[12] Multiplying Hausman's (1981, p. 51) estimates for husbands (convex budget set case) by the ratio of the price levels yields a net wage coefficient of approximately 0.1 and a mean virtual income coefficient of approximately -90. For the nonconvex budget set case, converting Hausman's estimates to 1983 dollars

11. See Blomquist (1983) for the likelihood function for this specification.
12. The U.S. consumer price index was 161.3 in 1975 and 298.4 in 1983. The ratio of the 1975 level to the 1983 level is approximately 0.54.

yields a net wage coefficient of about 6.1 and an average virtual income coefficient of approximately −83. While my estimates for the net wage coefficient are fairly close to those of Hausman, he estimates a substantial income effect while I find no evidence of one.

It is interesting to consider the magnitude of the reduction in labor supply caused by the U.S. (combined state and federal) tax system. Using the parameter estimates from the dual additive errors specification, the model was simulated (taking one draw from the error distributions for each sample member) under the actual budget constraints facing sample members, and then again with all tax effects removed from the budget constraints. The mean simulated value of hours of work in the no-tax world is 2,208 hours per year, while in the post-tax situation the mean simulated hours is 2,150. This implies that the tax system has resulted in approximately a 2.6 percent reduction in hours of work. While this reduction in labor supply is fairly small, it is not trivial. Analyses of tax reform proposals do need to take the labor supply response of married men into consideration.

While the results for husbands vary very little with the stochastic specification and estimation method employed, this does not hold true for wives. Results from estimation of the censored specifications are presented in Table 4. Table 5 contains results for the truncated (using the subsample with working wives) specifications using the reported gross wages. Table 6 differs from Table 5 in that imputed gross wages were substituted for the reported wages before performing the estimation. Comparison of Tables 5 and 6 provides an informal measure of the magnitude of the econometric problems caused by treating the imputed wages as though they are the true wage values (rather than integrating over the wage distribution). The differences in the estimates presented in the two tables appear to be quite small except for the truncated instrumental variables estimates, where a much smaller wage elasticity is estimated when using the reported gross wages.

Estimated wage elasticities are much larger when the censored specifications are estimated than when the truncated specifications are employed. Income elasticities are found to be fairly small in all specifications. As expected, the number of children has a much larger impact on the labor supply of wives than it does on husbands' labor supply.

It is again interesting to compare the results to those of Hausman (1981, p. 56). After adjusting for the change in the price level, Hausman's estimate of the net wage coefficient for wives (convex budget set case) is about 268, while his estimate of the average virtual income coefficient is approximately −68. In the censored dual errors specification, I obtain an estimated wage coefficient fairly close to Hausman's (226), but a virtual income coefficient which is only about a third as large (−23).

There does not appear to be the same problem in distinguishing between the effects of the heterogeneity and measurement/optimization error terms in estimating the censored specification for wives that there was in the case of husbands. The standard deviations of the two disturbances are estimated fairly precisely in the censored specification, and are correlated to a much lesser extent than was true for husbands. In estimating the censored specification for wives, the nonparticipants help to identify the model. The heterogeneity term seems better able to "explain" the piling up of hours at zero than is the measurement/optimization error. In the truncated specifications, there does appear to be a problem in distinguishing between the effects of the two error terms. When reported gross wages are used in the estimation the correlation between the estimators of the standard deviations of the error terms is nearly -1, while when imputed gross wages are used the correlation is approximately $-.95$.

Estimation of the random income coefficient (β) specification produces coefficient estimates very similar to those produced by the dual additive errors specification in both the censored and truncated cases. In each case, the distribution of β is estimated to have a mean slightly larger than the value estimated for β in the dual additive errors specification. The estimated standard deviations of the distribution of β are roughly 40 percent of the means.

Given the problem of distinguishing between the effects of the heterogeneity and measurement error terms, it is not surprising that estimation of the measurement-error-only specification results in parameter estimates very close to those of the dual additive errors specification. As in the dual errors case, the wage coefficient is much larger when the censored version of the model is estimated than when the truncated version is used. Also similar to the dual error specification, there is virtually no difference between the coefficient estimates produced by the truncated specification using imputed gross wages and the truncated specification using reported gross wages.

The wage elasticity produced by the truncated instrumental variables estimator is considerably larger when the gross wage is treated as endogenous (Column 1 in Table 6) than when it is treated as exogenous (Column 1 in Table 5).[13] The wage coefficient obtained treating the gross wage as endogenous is somewhat smaller than that produced by the truncated

13. The truncated instrumental variables estimates reported in Table 6 used the same variables used in the wage imputation regression (reported in the appendix) as instruments (in addition to husband's earnings, unearned income, and the exogenous variables included in the labor supply function); the imputed wage was not explicitly used in the estimation. The truncated instrumental variables estimates reported in Table 5 used the reported gross wage, husband's earnings, and unearned income as instruments.

measurement-error-only specification, while the wage coefficient estimated when the reported gross wage is used as an instrument is much smaller than in the other specifications estimated.

Estimates of the reduction in labor supply due to the U.S. (combined state and federal) income tax system highlights the differences in the elasticities estimated in the censored and truncated cases. The same procedure used to simulate the reduction in husbands' labor supply due to the tax system was also used for wives. In the simulations, husbands' labor supply was held fixed at observed values, although husbands' earnings and unearned income were treated as untaxed income in simulating the labor supply of wives in the no-tax state. Using the parameter estimates from the censored version of the dual additive errors specification (Column 2 in Table 4) results in mean predicted labor supply of 1,558 hours in the no-tax world and 1,086 hours under the actual budget constraints. This implies a 30 percent reduction in wives' labor supply due to the tax. However, using the parameter estimates from the truncated version of the two-error specification (Column 3 in Table 5) suggests that the tax system causes a 10 percent decrease in labor supply. Mean predicted labor supply is 1,361 hours under the present system, and 1,509 hours under the no-tax scenario.

Note that while the mean simulated post-tax hours using parameter estimates from the censored specification is quite close to the mean observed hours, the mean simulated post-tax hours using parameter estimates from the truncated specification is considerably higher. This appears to be due to the truncated specification estimates underpredicting the number of women with zero hours of work. If the models were correctly specified, the parameter estimates and predictions from the truncated and censored specifications should have been the same. That they are not indicates that some misspecification is present.

VI. Conclusions

The results of this paper suggest that the labor supply of prime aged married men is relatively invariant to the net wage and virtual income. The results are remarkably similar across the various specifications considered. It makes virtually no difference if one assumes the stochastic variation in male labor supply is due to heterogeneity in preferences, measurement/optimization error, or some combination of the two. The finding that male labor supply is quite inelastic is surprising, since earlier work by Hausman (1981) using a specification similar to one estimated in this paper found that while the net wage effect was close to zero, there was evidence that the virtual income effect was large in magnitude for part of the population. Although the results reported here indicate that the virtual income elasticity is zero, the estimated net wage elasticities are positive and larger in magnitude than the elasticity estimated by Hausman.

The estimated impact of taxation on the labor supply of married women depends critically on the method used to estimate the labor supply function. When a censored estimator is used, the net wage elasticities are similar to that estimated by Hausman (1981). However, when a truncated estimator is used (conditioning on hours being greater than zero), the estimated wage elasticities are much smaller. The estimated virtual income elasticities are also smaller (in absolute value) in the truncated specifications than they are in the corresponding censored specifications. Whether imputed or reported gross wages are used in the estimation seems to make little difference. One interpretation of the results for women is that while hours of work conditional on participation are relatively inelastic with respect to the net wage and virtual income, the participation decision is quite elastic with respect to the net wage.

While the treatment of the truncation of hours at zero does make a difference, other elements of the stochastic specification appear to be relatively unimportant. The elasticity estimates are roughly the same whether optimization/measurement error, heterogeneity in preferences, or a combination of the two is specified to be the source of stochastic variation in hours of work. It is difficult to determine the generality of this result. The particular specification chosen may matter more in other applications of this method.

It seems safe to say that taxation causes fairly little reduction in the labor supply of prime-aged married males in the United States. It is much more difficult to state with any precision what effect the U.S. tax system has on the labor supply of married women. Further research into the effect of the tax system on labor force participation of married women will be needed before we can answer this question with greater confidence.

Appendix

Imputation of Wives' Wage Rates

This appendix provides details of the procedure used to impute wages for wives. Heckman's (1979) technique for correcting for sample selection bias was used to estimate a wage equation for women. While there is no reason to suppose that the distributional assumptions underlying this technique are correct in this case, in practice it is probably better to include the selection bias term in the wage regression than to ignore the

Table A-1
Reduced Form Probit: Wives' Labor Force Participation (standard errors in parentheses)

Variable	Maximum Likelihood Estimates
Constant	1.780
	(0.198)
Unearned income	−0.022
Plus husband's	(0.003)
earnings ($1,000's)	
Age–35	0.003
(35 = 0 & 45+ = 10)	(0.006)
Age–45 (45 = 0)	−0.066
	(0.042)
Children < 6	−0.270
	(0.066)
Family size	−0.127
	(0.045)
Bad health	−0.240
	(0.154)
College education	0.427
	(0.104)
Log likelihood	−514.4
Number of observations: 978	

issue of sample selection bias entirely. The results of the first step, estimation of a reduced form probit equation for wives' labor market participation, are presented in Table A-1. Independent variables not previously described in Section V of this paper include the sum of the couple's unearned income plus the husband's pre-tax earnings, an extra age variable which is equal to zero for women less than 35, equal to age − 35 for those between 35 and 45, and equal to 10 for those 45 or over, and a dummy variable set equal to one if the woman has completed more than 12 years of education ("college education").

Parameter estimates from this step were used to compute the estimated value of the inverse of Mills' ratio for every working woman. A wage imputation linear regression was then estimated with the inverse Mills' ratio variable included as a regressor. Variables which are exogenous and likely to be good predictors of the wage were chosen as regressors. Re-

Table A-2
Wives' Wage Imputation Regression (standard errors in parentheses)[a]

Variable	Estimated Coefficients
Constant	6.673
	(10.897)
Age	0.007
	(0.329)
$Age^2/100$	−0.004
	(0.002)
Education	−1.015
	(1.021)
$Education^2/10$	0.028
	(0.0266)
(Education * age)/10	0.032
	(0.015)
Nonwhite	0.469
	(0.487)
County unemployment rate	−0.006
	(0.006)
Inverse Mills' ratio	0.209
	(0.822)
R^2: .17	
Number of observations: 715	

a. Standard errors have been corrected for use of an imputed inverse Mill's ratio.

sults of this estimation are shown in Table A-2. The last step was to use the estimated coefficients to compute imputed wage rates for all women in the sample. An adjustment for the expectation of the disturbance of the wage equation conditional on participation or nonparticipation (based on the probit estimates) was included in calculating the predicted wage rates.

References

Avery, Robert B., and V. Joseph Hotz. 1985. "Estimating Systems of Nonlinear Equations with Limited Dependent Variables." Economics Research Center/NORC discussion paper 85–19.

Berndt, Ernst R., Bronwyn A. Hall, Robert E. Hall, and Jerry A. Hausman. 1974. "Estimation and Inference in Nonlinear Structural Models." *Annals of Economic and Social Measurement* 3(4):653–66.

Blomquist, N. Sören. 1983. "The Effect of Income Taxation on the Labor Supply of Married Men in Sweden." *Journal of Public Economics* 22(2):169–97.

Burtless, Gary, and Jerry A. Hausman. 1978. "The Effect of Taxation on Labor Supply: Evaluating the Gary Negative Income Tax Experiment." *Journal of Political Economy* 86(6):1103–30.

Feenberg, Daniel R., and Harvey S. Rosen. 1986. "State Personal Income and Sales Taxes." In *Studies in State and Local Public Finance,* ed. Harvey S. Rosen. Chicago: University of Chicago Press.

Hausman, Jerry. 1981. "Labor Supply." In *How Taxes Affect Economic Behavior,* ed. Henry J. Aaron and Joseph A. Pechman. Washington, D.C.: Brookings Institution.

———. 1985. "The Econometrics of Nonlinear Budget Sets." *Econometrica* 53(6):1255–82.

Heckman, James J. 1979. "Sample Selection Bias as a Specification Error." *Econometrica* 47(1):153–62.

Heckman, James J., and Thomas MaCurdy. 1981. "New Methods for Estimating Labor Supply Functions: A Survey." In *Research in Labor Economics,* Vol. 4, ed. Ronald G. Ehrenberg, Greenwich, Conn.: JAI Press.

Institute for Survey Research. 1986. *A Panel Study of Income Dynamics: Procedures and Tape Codes, Wave XVII.* Ann Arbor: Institute for Survey Research.

Internal Revenue Service. 1985. *Individual Income Tax Returns.* Washington, D.C.: GPO.

MaCurdy, Thomas E. 1983. "A Simple Scheme for Estimating an Intertemporal Model of Labor Supply in the Presence of Taxes and Uncertainty." *International Economic Review* 24(2):265–89.

Moffitt, Robert. 1986. "The Econometrics of Piecewise-Linear Budget Constraints: A Survey and Exposition of the Maximum Likelihood Method." *Journal of Business and Economic Statistics* 4(3):317–27.

Moffitt, Robert, and Walter Nicholson. 1982. "The Effect of Unemployment Insurance on Unemployment: The Case of Federal Supplemental Benefits." *Review of Economics and Statistics* 64(1):1–11.

Triest, Robert K. 1987a. "The Effect of Income Taxation on Labor Supply When Deductions are Endogenous." Department of Political Economy working paper 105, Johns Hopkins University.

———. 1987b. "A Monte Carlo Test of the Robustness of Alternative Estimators of Labor Supply Functions." Department of Political Economy working paper 198, Johns Hopkins University.

U.S. Advisory Commission on Intergovernmental Relations. 1984. *Significant Features of Fiscal Federalism.* Washington, D.C.: GPO.

Wales, T. J., and A. D. Woodland. 1979. "Labor Supply and Progressive Taxes." *Review of Economic Studies* 46(1):83–95.

References

The Effect of Taxation on Labor Supply: Evaluating the Gary Negative Income Tax Experiment

Gary Burtless; Jerry A. Hausman

The Journal of Political Economy, Vol. 86, No. 6. (Dec., 1978), pp. 1103-1130.

Stable URL:

http://links.jstor.org/sici?sici=0022-3808%28197812%2986%3A6%3C1103%3ATEOTOL%3E2.0.CO%3B2-9

The Econometrics of Nonlinear Budget Sets

Jerry A. Hausman

Econometrica, Vol. 53, No. 6. (Nov., 1985), pp. 1255-1282.

Stable URL:

http://links.jstor.org/sici?sici=0012-9682%28198511%2953%3A6%3C1255%3ATEONBS%3E2.0.CO%3B2-R

Sample Selection Bias as a Specification Error

James J. Heckman

Econometrica, Vol. 47, No. 1. (Jan., 1979), pp. 153-161.

Stable URL:

http://links.jstor.org/sici?sici=0012-9682%28197901%2947%3A1%3C153%3ASSBAAS%3E2.0.CO%3B2-J

A Simple Scheme for Estimating an Intertemporal Model of Labor Supply and Consumption in the Presence of Taxes and Uncertainty
Thomas E. MaCurdy
International Economic Review, Vol. 24, No. 2. (Jun., 1983), pp. 265-289.
Stable URL:
http://links.jstor.org/sici?sici=0020-6598%28198306%2924%3A2%3C265%3AASSFEA%3E2.0.CO%3B2-S

The Econometrics of Piecewise-Linear Budget Constraints: A Survey and Exposition of the Maximum Likelihood Method
Robert Moffitt
Journal of Business & Economic Statistics, Vol. 4, No. 3. (Jul., 1986), pp. 317-328.
Stable URL:
http://links.jstor.org/sici?sici=0735-0015%28198607%294%3A3%3C317%3ATEOPBC%3E2.0.CO%3B2-4

The Effect of Unemployment Insurance on Unemployment: The Case of Federal Supplemental Benefits
Robert Moffitt; Walter Nicholson
The Review of Economics and Statistics, Vol. 64, No. 1. (Feb., 1982), pp. 1-11.
Stable URL:
http://links.jstor.org/sici?sici=0034-6535%28198202%2964%3A1%3C1%3ATEOUIO%3E2.0.CO%3B2-0

Labour Supply and Progressive Taxes
T. J. Wales; A. D. Woodland
The Review of Economic Studies, Vol. 46, No. 1. (Jan., 1979), pp. 83-95.
Stable URL:
http://links.jstor.org/sici?sici=0034-6527%28197901%2946%3A1%3C83%3ALSAPT%3E2.0.CO%3B2-C