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A COMPARISON OF THE LABOR FORCE BEHAVIOR OF MARRIED WOMEN IN THE UNITED STATES AND CANADA, WITH SPECIAL ATTENTION TO THE IMPACT OF INCOME TAXES[1]

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Estimation results are presented for the probability of working, the hourly wage rate, and the annual hours of work for wives in seven different age groups in both the U.S. and Canada. Federal and state or provincial taxes are incorporated into the analysis. An iterative estimation method is employed to circumvent the statistical problems resulting from the dependence of the hours of work on the tax rate, and the dependence of the tax rate on the hours of work.

1. INTRODUCTION

THE EXISTING CROSS-SECTIONAL LITERATURE on labor force behavior tells us that as wages rise American wives work more and more [**14, 16, 17, 19, 31**], while American husbands decrease their annual hours of work, presumably by doing less moonlighting and overtime work, bargaining for shorter work weeks, and so forth [**3, 10, 14, 32**]. Nor is there any collective discomfort evident in the literature about attributing this behavioral dissimilarity simply to a sex difference. Little effort has been made, for instance, to ascertain whether only certain types of wives display this work-loving behavior, with other wives behaving more like their husbands.

Yet in a recent study of the labor supply of Canadian wives [**27**], we find that working wives in Canada tend to work fewer hours per year the more they are paid per hour. Moreover, we find that the uncompensated wage elasticities of hours of work for Canadian wives are similar to those reported by other researchers for American men. Unemployment is a bigger problem and there are fewer job opportunities for women in Canada than in the U.S. Canadian wives are less educated, are married to men with lower employment incomes, and have different childbearing patterns than their American counterparts. Also the U.S. and Canada have quite different income tax laws. Unlike the U.S. situation, working husbands and wives in Canada must file separate tax returns. Thus all working Canadians face the same tax tables, with the exceptions of differences in provincial tax rates and allowable deductions. In the U.S., on the other hand the first dollar of a wife's earnings is taxed at the marginal rate which would apply to an additional dollar earned by the husband.

Given these differences in circumstances, substantial differences in the wage rates and aggregate labor supply of Canadian and American wives are to be expected. It can be seen from Table I, for instance, that although U.S. wives are generally paid more per hour than their Canadian counterparts, the average net

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TABLE I

WAGE RATES[a] AND LABOR SUPPLY OF U.S. AND CANADIAN WIVES

Age Group	Offered Wage U.S.	Offered Wage Can.	Net Offered Wage For First Hour of Work (after tax at $h = 1$)[b] U.S.	Net Offered Wage For First Hour of Work (after tax at $h = 1$)[b] Can.	Net Offered Wage For Last Hour of Work (after tax at h = actual hours worked)[b] U.S.	Net Offered Wage For Last Hour of Work (after tax at h = actual hours worked)[b] Can.	Average Net Offered Wage For Actual Hours Worked[c] U.S.	Average Net Offered Wage For Actual Hours Worked[c] Can.	Employment Rate[d] U.S.	Employment Rate[d] Can.	Mean Hours of Work for All Wives[b] U.S.	Mean Hours of Work for All Wives[b] Can.	Mean Hours of Work for Working Wives[b] U.S.	Mean Hours of Work for Working Wives[b] Can.
25–29	3.36	2.98	2.50	2.98	2.14	2.08	2.43	2.38	.51	.48	537	598	1,412	1,246
30–34	3.13	3.09	2.38	3.08	1.66	2.16	2.26	2.48	.47	.40	535	467	1,407	1,168
35–39	3.20	2.82	2.42	2.82	1.96	1.64	2.30	2.28	.50	.38	604	443	1,437	1,165
40–44	3.20	2.72	2.37	2.72	2.08	1.70	2.24	2.21	.53	.41	718	506	1,528	1,234
45–49	3.15	2.84	2.39	2.84	2.11	1.99	2.26	2.30	.54	.42	767	555	1,598	1,322
50–54	3.04	2.66	2.30	2.66	1.81	1.91	2.18	2.19	.52	.38	746	500	1,622	1,317
55–59	3.19	2.66	2.49	2.66	2.06	2.08	2.31	2.19	.46	.33	664	427	1,620	1,295

SOURCE: Calculated from the 1 per cent subsample from the 5 per cent primary State Public Use Sample of Basic Records from the 1970 U.S. Census; and from the 1 per cent Family File of the Public Use Sample from the 1971 Canadian Census.

[a] The U.S. wage figures are given in 1969 U.S. dollars. The Canadian wage figures are given in 1970 Canadian dollars, and are shown in this Table only in terms of 1969 U.S. dollars. See footnote 2 in text.

[b] See text and footnote 13 for details.

[c] The average tax rate on the wife's earnings is calculated as the tax on the family's income at the actual hours worked by the wife minus the tax at zero hours of work for the wife divided by the wife's actual earnings.

[d] We are referring here to the proportion of wives who earned at least one dollar of employment income in 1969 for the U.S. or in 1970 for Canada.

wage rates for the first hour of work are always lower for U.S. than for Canadian wives.[2] Also the employment rates for wives and the mean hours of work for working wives are consistently higher in the U.S. than in Canada. In the face of these differences, and other important differences between the data bases which are available for the two countries,[3] it would not be surprising to find that the labor supply responses of Canadian and U.S. wives to changes in their wage rates, and other variables as well, are different.

In this study, estimation results are presented for the probability of working, the offered (or market) wage rate, and the annual hours of work for wives in seven different age groups in both the U.S. and Canada. In making comparisons between the U.S. and Canada, it is essential to build taxes into the model to be estimated. Otherwise differences in the respective values of the estimated coefficients of the wage and income variables may simply reflect inter-country differences in the relationships of gross to disposable income. In previous studies incorporating income taxes, the endogeneity of the tax rate is ignored. Either the tax rate is calculated for the wife's actual hours of work [**14**] resulting in a least squares bias problem, or for an arbitrarily chosen number of hours of work [**15**, **31**]. Hausman and Wise [**15**, p. 434] note that, while the latter approach avoids the least squares bias problem inherent in using the actual hours of work, it is "rather ad hoc" since it ignores the dependence of the tax rate on the endogenously determined hours of work. The present model is more general in that the tax rate is fully endogenous. An iterative estimation method is employed to circumvent the statistical problems resulting from the dependence of the hours of work on the tax rate, and the dependence of the tax rate on the hours of work.

We first estimate our hours of work equation in a form which allows us to test the hypothesis that wives fully account for the impact of taxes on their earnings in choosing their hours of work, as found by Harvey Rosen [**31**] and by Hausman and

[2] As stated in the footnotes to the Table, all dollar figures in Table I are in terms of 1969 U.S. dollars. The 1969 U.S. dollar value of one 1979 Canadian dollar, the monetary unit in which the Canadian income figures we are using are reported, is found in two steps. First, using the average noon spot rate reported by the Bank of Canada for 1970, we find that one 1970 Canadian dollar was worth approximately 95.8 cents in 1970 U.S. currency. Second, using the Consumer Price Indices for 1969 and 1970 reported in the *1979 Economic Report of the President*, we find that in terms of purchasing power one 1970 U.S. dollar was worth approximately 94.0 cents in terms of 1969 U.S. dollars. Taking both these factors into account, we find that one 1970 Canadian dollar was worth about 90.0 cents in terms of 1969 U.S. currency.

Our estimation results presented in Tables VI–IX and the mean values shown in Table V, however, are all based on 1969 U.S. dollar figures for the U.S. and 1970 Canadian dollar figures for Canada. In the estimation portion of our paper, our primary goal is to capture the interrelationships between variables in determining the labor force behavior of wives in each country considered separately. Since spot exchange rates do not normally reflect the relative domestic balances between prices and wages within different countries, we felt we might introduce distortions into our behavioral relationships by converting all dollar figures to, say, 1969 U.S. dollars. Secondly, such a conversion of our Canadian dollar figures would have complicated the computation of the Canadian taxes.

We feel that, despite the value differences in the 1969 U.S. and 1970 Canadian dollars, it is still valid to note similarities between our U.S. and Canadian estimation results in terms of coefficient signs and which coefficients are significant. Finally in Table XI we summarize our most important results in terms of elasticities which are, of course, unit free.

[3] See footnote 13.

Wise [15]. These results then permit us to estimate a constrained hours of work equation, from which we obtain more efficient estimates of the response of the hours of work of wives to their net offered wage rates.

In agreement with our earlier findings for Canada, we find in this study that both American and Canadian wives with higher potential wage rates are more likely to work, and that both American and Canadian wives who do work tend to work fewer hours per year the more they are paid per hour. This latter result for American wives who work is a clear contradiction to the conclusions which other researchers have drawn from their studies of the labor supply of married women in the U.S. Possible explanations for this difference between our results and the results of other studies are discussed in Section 6, and our results are shown to be consistent with aggregate time series data on the labor force participation, hours of work per week, and weeks of work per year of women in the U.S. and Canada. The fact that our coefficient estimates are so similar for all variables in our model for the probability of working, our offered wage equation, and our annual hours of work equation for both countries within each one of our seven age groups increases our confidence in our findings.

Correcting for federal and state or provincial income taxes is imperative in this study because we wish to make comparisons between two countries with very different tax laws. Such a correction may also be important, however, in studying the labor supply of women (or men) within a single country. It is true that the hourly wage rate before taxes is highly correlated with the marginal after-tax wage rate in both the U.S. and Canada. Hence the observed wage rate is an excellent proxy for the marginal tax corrected wage rate [9]. Yet because of the progressive nature of the income tax, wives with the same before-tax wage rates will have different after-tax wage rates owing to different annual hours of work, differences in deductions and family asset income, and, in the U.S. case, differences in the earned incomes of their husbands. Hence for wives whose marginal tax rates are substantially different from the average for other wives with the same before-tax wage rates, the estimated response of their annual hours of work to changes in their wage rates will clearly be inappropriate if no account is taken of income taxes, and the estimated response may well be a biased estimate even for those wives whose marginal tax rates are close to the average for wives in any given offered wage category. Comparing our tax corrected results (both unconstrained and constrained) given in this paper for our hours of work equation with our earlier Canadian results [27], it can be seen that for the age groups 25–29, 30–34, and 35–39 where we have the best determination of the coefficient of our wage rate variable, this coefficient is consistently more negative and generally more significant when taxes are accounted for. Thus the failure to account for the effects of income taxes may result in estimation biases which partially mask the backward bending nature of the labor supply function for those who work. In the U.S. case, the failure to account for the marginal tax rate on the wife's earnings may also lead to biased estimates of the response of hours of work to changes in the employment income of the husband, since there is a strong correlation between this marginal tax rate and the husband's employment income.

When the husband's employment income and the asset income of the family are also corrected for income taxes (evaluated at zero hours of work for the wife) we find, in the case of Canada, that the associated coefficients for these variables are more negative compared with our earlier Canadian results which were obtained without adjusting these income variables for taxes. Based on these new results given in this paper, we see that in a period when inflation is pushing many husbands into higher and higher tax brackets even though their real earnings before tax are not increasing, there will be corresponding increases for the wives of these husbands in the probability of working and in their expected annual hours of work if they do work. This and other important interactions between the labor supply of wives, their net marginal wage rates, and the disposable incomes of their husbands cannot be analyzed within the context of models which ignore taxation.

Finally our local opportunity for jobs index, which was first introduced in our earlier Canadian study and which is similar to an index of the industrial composition of local employment proposed by Bowen and Finegan [**5**], is found to be just as important for U.S. as for Canadian wives in determining the probability that a wife will work and her expected offered wage rate if she does work.

2. THE MODEL

Assume a family maximizes a twice differentiable quasiconcave conditional utility function $U(x, T-h; E_HT, Z^*)$ subject to the time constraint $0 \leq h < T$ and a one period budget constraint[4]

$$px = I(h; E_HT, w) \tag{1}$$

where x is a Hicksian composite good representing the consumption of all goods other than leisure, h represents the market time (hours of work) of the wife at the offered (market) wage rate w, E_HT is the husband's earned income plus family asset income net of the income taxes which would be paid at zero hours of work for the wife, Z^* is a vector of constraints arising from previous events, T is the wife's total time, p is the price of the Hicksian composite good, and I is the total income of the family net of federal and state income taxes. The right hand side of (1) can be written as

$$I(h; E_HT, w) = E_HT + w\int_0^h (1 - TX_s)\,ds \tag{2}$$

where TX_s is the marginal tax rate on the wife's earnings at s hours of work.

The Lagrangian for this problem is

$$L = U(x, T-h; E_HT, Z^*) + \lambda\left\{E_HT + w\int_0^h (1-TX_s)\,ds - px\right\} + \gamma h \tag{3}$$

where λ and γ are, respectively, an unconstrained and a nonnegative dual variable. The necessary conditions for optimality [**25**, p. 173] are that there exist

[4] See the original version of this paper for further justification of this conditional utility function and budget constraint.

λ, γ, and $h(0 \leq h < T)$ such that $\gamma \geq 0$,

(4) $$U_x - \lambda p = 0,$$

(5) $$-U_l + \lambda\{w(1 - TX_h)\} + \gamma = 0,$$

and

(6) $$\gamma h = 0,$$

where $l = T - h$ is the nonmarket time (leisure) of the wife,

$$U_l = \partial U(x, l; E_H T, Z^*)/\partial l, \quad \text{and}$$

$$U_x = \partial U(x, l; E_H T, Z^*)/\partial x.$$

We will let

(7) $$w_h^n = w(1 - TX_h)$$

denote the wife's net offered wage rate at h hours of work. From (1), (2), and (7) we get

$$x = (E_H T + w \int_0^h (1 - TX_s)\, ds)/p$$

which is linearized around h to be

(8) $$x = (E_H T + w_h^n h)/p.$$

From (4) we get

(9) $$\lambda = U_x/p = \{\partial U(x, T - h; E_H T, Z^*)/\partial x\}/p.$$

Thus we see that $\lambda > 0$ since $p > 0$ and $U_x > 0$, and that, given any offered wage w, in equilibrium λ is a function of h, p, $E_H T$, $w_h^n h$, and Z^*.

We also have by (5) that $\lambda = (U_l - \gamma)/w_h^n$, or $w_h^n = (U_l/\lambda) - (\gamma/\lambda)$, where the shadow price of the wife's time (asking wage) at h hours of work is defined by

(10) $$w_h^* = U_l/\lambda.$$

Hence the shadow price depends on h, p, $E_H T$, $w_h^n h$, and Z^* when $h > 0$, and on p, $E_H T$, and Z^* when $h = 0$. Moreover, since $\lambda > 0$ by (9) and since (6) implies that $\gamma = 0$ if $h > 0$ and $\gamma \geq 0$ if $h = 0$, we see that a wife will choose to work only if $w_h^n \geq w_h^*$ at $h = 0$, and that wives who work will adjust their hours of work such that $w_h^n = w_h^*$. We assume that a working wife's asking wage, w_h^*, is an increasing function of her hours of work, h, because the per hour costs of replacing her services in the home will presumably be higher the more time she devotes to market work. It should also be noted that in this model a working wife's asking wage depends on her net offered wage, w_h^n, because the mechanism by which leisure is traded for the increased consumption of other goods is through the relaxation of the household budget constraint and this constraint will be relaxed more rapidly as h is increased the higher the net offered wage is. When $h=0$, however, this income effect vanishes.[5]

3. SPECIFICATION OF THE MODEL

If we take the log of both sides of (10) and linearize it around Z^*, E_HT, $\ln w_h^n$, and h, we get a linear approximation of the log of the asking wage for the ith married woman:

$$(11)\qquad \ln w_h^* = \begin{cases} \beta_0 + Z^*\beta_1 + \beta_2 E_H T + \beta_3 \ln w_h^n + \beta_4 h + U^* & \text{if} \quad h>0, \\ \beta_0 + Z^*\beta_1 + \beta_2 E_H T + U^* & \text{for} \quad h=0, \end{cases}$$

where U^* denotes the disturbance term. The variable p does not appear in (11) because it is assumed to be the same for all families. The wife's net offered wage, $w_h^n = w(1-TX_h)$, is also really a price variable. It cannot be ignored in this way, however, since both w and TX_h differ systematically from one wife to another. We will assume that variations in the wife's offered wage w are explained by

$$(12)\qquad \ln w = \alpha_0 + Z\alpha_1 + R\alpha_2 + u,$$

where Z and R are, respectively, vectors of personal characteristics and regional macroeconomic variables, and u denotes the disturbance term.[6]

[5] Heckman assumes in his one-period models [**16**, **17**, **19**] that the asking wage is independent of the offered wage for $h \geq 0$. This allows him to identify the asking wage function. It is also equivalent to assuming that there are no income effects for those who work associated with a change in the offered wage rate.

[6] In specifying (12), we have implicitly assumed that the offered wage does not depend on the hours of work. There is no doubt that a woman's offered wage is affected by whether she has worked predominately full or part-time in recent years, although the age-earnings profiles for women are generally reported to be quite flat compared with those for men. However, it is not clear to us how short term fluctuations in a woman's hours of work will affect her offered wage.

Harvey Rosen [**31**, pp. 489–490] argues, following Lewis' work [**23**], that there are quasi-fixed costs to employers associated with each employee and that these costs imply both that the hours of work must appear in the offered wage function and that the derivative of this function with respect to the hours of work must be positive. This assumed dependence also implies that the average offered wage, which is observed for those who work, is no longer equal to the marginal offered wage which must be equated with the marginal asking wage. Hence Rosen [**31**, p. 493–494] must assume the form of the dependence of the marginal wage rate on the observable average wage rate and hours of work. The cost, therefore, of introducing hours into the offered wage function, in terms of additional assumptions which must be imposed on the model prior to estimation, is high, particularly if some of these assumptions are either wrong or overly simplified.

If a woman takes a second job, for instance, to meet short term financial needs it is likely that the offered wage for this second job will be less than for her primary job because of the types of jobs available at off hours and because she has no seniority in the new employment situation. The impact of short term decreases in hours of work due to illness or pregnancy on a woman's offered wage are frequently contractually determined. Moreover, because employers are often required by a variety of agreements and regulations to provide more expensive fringe benefits to full-time than to part-time workers, part-time workers are sometimes brought in at exceptionally high wage rates to meet short term needs for increased labor.

We are not convinced, therefore, that the observed offered wage must be an increasing function of hours of work. Nor do we feel that the evidence available to us is sufficient to justify any other such assumption concerning the dependence of the offered wage on hours of work.

It is not possible in our model to uniquely determine the coefficients of the asking wage equation (11). When $h > 0$, however, we can solve the equation $\ln w_h^* = \ln w_h^n$ for an expression for the wife's equilibrium number of hours of work:

$$(13) \qquad h = \frac{1}{\beta_4}[-\beta_0 + (1-\beta_3)\ln w_h^n - Z^*\beta_1 - \beta_2 E_H T] + v^*,$$

where $v^* = -(1/\beta_4)U^*$. Since TX_h and hence w_h^n depend on both w and h in a nonlinear manner, we see that (12) and (13) define a nonlinear interdependent system.

Denoting the marginal retention rate on the wife's earnings at h hours of work by

$$(14) \qquad RET = \begin{cases} 1 - TX_h(E_H + wh, A) & \text{for a joint return,} \\ 1 - TX_h(wh, A) & \text{for a separate return,} \end{cases}$$

where E_H and A are the husband's income and asset income before taxes, respectively, we see from (7) that (13) can be rewritten as

$$(15) \qquad h = \gamma_0 + \gamma_1 \ln w + (\gamma_1 \xi)\ln RET + Z^*\gamma_2 + \gamma_3 E_H T + v^*,$$

where $\gamma_0 = -\beta_0/\beta_4$, $\gamma_1 = (1-\beta_3)/\beta_4$, $\gamma_2 = -\beta_1/\beta_4$, $\gamma_3 = -\beta_2/\beta_4$, and RET has been replaced with $(RET)^\xi$. An estimate of ξ can be obtained by dividing the coefficient of $\ln RET$ by the coefficient of $\ln w$. If there is no tax illusion, ξ should be approximately equal to 1.

We define the vectors Z^*, Z, and R as follows:

Personal Characteristics Affecting a Wife's Asking Wage (Z^* and $E_H T$)

Z^*1. Number of children younger than 6. (+)
Z^*2. Number of children 6–14 years of age. (+)
Z^*3. Product of number of children younger than 6 and number 6–14 years of age. (−)
Z^*4. Number of children 19–24 years of age attending school full or part time. (−)
Z^*5. Number of children ever born. (+)
Z^*6. Language dummy (=1 if language of home is French, =0 otherwise; for Canada only). (+)
Z^*7. Employment income of husband plus asset income of family net of income taxes to be paid at zero hours of work for the wife, denoted as $E_H T$ above. (+)
Z^*8. Z^*7 divided by number of persons in family. (?)

Personal Characteristics Affecting a Wife's Offered Wage (Z)

$Z1$. Wife's years of education. (+)
$Z2$. Number of children younger than 6. (−)
$Z3$. Age of wife at first marriage. (+)
$Z4$. Race dummy (=1, if wife is Black, =0 otherwise; for U.S. only). (−)

Regional Economic Variables Affecting a Wife's Offered Wage (R)

$R1$. State (for U.S.) or provincial (for Canada) unemployed rate. (−)
$R2$. Local opportunity for jobs index. (+)

The plus and minus signs in parentheses following the variable names indicate the expected impact of each variable on the wife's asking or offered wage. We will discuss only certain variables included in Z^*, Z, and R which have not been included in previous studies, beginning with the vector Z^*.

Z^*3 has been included in Z^* to account for nonlinearities in time expenditures per child as both the total number and the number of older children increase. Z^*4 represents the financial burden of children attending post-secondary educational programs. Z^*8 is included to control for interactions between family size and

TABLE II

WOMEN AS PER CENT OF ALL WORKERS IN EACH OCCUPATION, FOR U.S., 1960 AND 1970, AND CANADA, 1961 AND 1971

Occupation	U.S. 1960	Can. 1961	U.S. 1970	Can. 1971	U.S.–Can. Difference 1960–61	1970–71
Managerial	14.8	10.4	17.0	15.7	4.4	1.3
Natural Sciences	4.8	4.8	7.8	7.3	−0.0	0.5
Social Sciences	32.8	29.4	39.4	37.4	3.4	2.0
Religion	15.4	28.9	10.3	15.7	−13.5	−5.4
Teaching	66.9*	64.4*	65.0*	60.4*	2.5	4.6
Medicine	67.5*	72.1*	72.9*	74.3*	−4.6	−1.4
Artistic	35.3	31.2	31.8	27.2	4.1	4.6
Clerical	68.9*	61.0*	74.6*	68.4*	7.9	6.2
Sales	32.1*	32.0*	33.6*	30.4*	0.1	3.2
Service	58.2*	46.7*	55.2*	46.2*	11.5	9.0
Farming	8.8	11.7	8.7	20.9	−2.9	−12.2
Other Primary	0.9	0.3	3.5	1.3	0.6	2.2
Processing	19.8	13.7	28.7	17.8	6.1	10.9
Machining and Fabricating	21.1*	17.9*	25.4*	18.7*	3.2	6.7
Construction	1.1	0.8	1.8	0.9	0.3	0.9
Transport	1.6	0.6	5.1	2.4	1.0	2.7
Other Occupations	14.4	13.6	16.8	15.7	0.8	1.1
Unknown	37.7*	26.0	40.1*	43.4*	11.7	−3.3
Total	32.7	27.3	37.8	34.3	5.4	3.5

SOURCE: Calculated from U.S. Census of Population: 1960, Final Report PC(2)7C, Table 2; U.S. Census of Population: 1970, Final Report PC(2)7C, Table 8; 1961 Census of Canada, Volume III, Part 1, Table 6; 1971 Census of Canada, Volume III, Part 2, Table 2.

* More than 5 per cent of the female labor force was in this occupation in the given year. See [**6**].

TABLE III

RATIOS OF EXPECTED JOBS FOR WOMEN TO NUMBER OF WOMEN 15 YEARS OF AGE AND OLDER, BY STATE AND PLACE OF RESIDENCE FOR 1969–1970

State	Urban	Rural	Entire State	State	Urban	Rural	Entire State
Alabama	.61	.44	—	Missouri	.65	.52	—
Alaska*	—	—	.64	Montana	.64	.55	—
Arizona	.66	.57	—	Nebraska	.73	.52	—
Arkansas	.58	.50	—	Nevada*	—	—	.76
California	.68	.58	—	New Hampshire	.72	.66	—
Colorado	.73	.56	—	New Jersey	.66	.65	—
Connecticut	.70	.70	—	New Mexico	.61	.53	—
Delaware*	—	—	.74	New York	.66	.65	—
District of Columbia	.73	—	—	North Carolina	.64	.54	—
Florida	.62	.55	—	North Dakota	.73	.49	—
Georgia	.66	.57	—	Ohio	.65	.58	—
Hawaii*	—	—	.63	Oklahoma	.64	.51	—
Idaho	.63	.60	—	Oregon	.69	.61	—
Illinois	.68	.59	—	Pennsylvania	.62	.56	—
Indiana	.68	.61	—	Rhode Island*	—	—	.62
Iowa	.70	.54	—	South Carolina	.62	.56	—
Kansas	.68	.60	—	South Dakota	.73	.47	—
Kentucky	.64	.45	—	Tennessee	.65	.51	—
Louisiana	.59	.46	—	Texas	.66	.49	—
Maine	.66	.57	—	Utah*	—	—	.69
Maryland	.70	.60	—	Vermont*	—	—	.62
Massachusetts	.67	.63	—	Virginia	.64	.57	—
Michigan	.67	.58	—	Washington	.69	.60	—
Minnesota	.72	.56	—	West Virginia	.58	.43	—
Mississippi	.63	.45	—	Wisconsin	.70	.56	—
				Wyoming*	—	—	.62

SOURCE: Calculated from the 1 per cent subsample from the 5 per cent primary State Public Use Sample of Basic Records from the 1970 U.S. Census.
* The distinction between urban and rural areas not available for these states.

TABLE IV

RATIOS OF EXPECTED JOBS FOR WOMEN TO NUMBER OF WOMEN 15 YEARS OF AGE AND OLDER, BY PROVINCE AND PLACE OF RESIDENCE FOR 1970–1971

	Urban		Rural	
Province	30,000 and over	Under 30,000	Non-farm	Farm
Newfoundland	.45	.34	.20	.19
Nova Scotia	.50	.38	.29	.30
New Brunswick	.48	.39	.28	.29
Quebec	.42	.35	.26	.17
Ontario	.49	.41	.35	.26
Manitoba	.50	.41	.32	.18
Saskatchewan	.48	.39	.27	.15
Alberta	.51	.43	.32	.17
British Columbia	.45	.40	.36	.33

SOURCE: Calculated from the 1 per cent Individual File of the Public Use Sample from the 1971 Canadian Census.

family income, excluding the income of the wife. Finally, for Canadian wives, $Z^{*}6$ is included to account for the alleged cultural conservatism of French Canadians toward wives working.

$Z2$ and $Z3$ have been included in the vector Z as proxies for the amount and recentness of a wife's previous work experience. And $R2$, our local opportunity for jobs index, has been included in the vector R to capture the effects of the availability of job opportunities for women on their offered wages.

Evidence is presented in [**6**, **27**, and **29**] indicating that there has been little change in the ratios of women to total workers within occupational groups in either the U.S. or Canada since the 1950's. Moreover, despite substantial differences in the age-specific participation rates for women in the two countries, as can be seen from Table II these ratios of women to total workers within occupations show relatively little between-country variation.[7]

Based on this insight, the values of our local opportunity for jobs index were calculated as follows. The expected numbers of jobs for women in each occupation, in each state or province and place of residence, were calculated by multiplying the national percentages of women in each occupation by the actual numbers of workers in each occupation, in each state or province and place of residence. These expected numbers were summed over all occupations, and the resulting totals were then divided by the total numbers of women 15 years of age and older in each state or province and place of residence. The resulting values for our index for the U.S. and Canada are shown in Tables III and IV, respectively.[8] Looking at these values, in both countries we find that the opportunities for paid employment for women are better in urban than in rural areas. It is also surprising, perhaps, to find that the values of our index are so much higher on the whole for the U.S. than for Canada, since the occupation-specific ratios of women to total workers in the two countries are similar. The explanation for the larger index values for the U.S. lies primarily in differences between the two countries in the total number of jobs in the different occupational groups relative to the potential numbers of female workers in the two countries.[9]

The basic U.S. data used in this study consists of 29,383 records for married couples living in the U.S., where the wife is 25–59 years old and no nonrelatives

[7] The recoding of the U.S. and Canadian occupational data according to the 1971 Canadian major codes used in this study is discussed in [**6**]. Canadian rather than U.S. major codes have been used because the occupational data provided in the Public Use Sample from the 1971 Canadian Census is not sufficiently detailed to permit reclassification according to the U.S. major codes.

[8] Virtually identical index values were calculated using an industrial rather than an occupational classification scheme. Our results in this paper are presented for the occupational codes, however, because these codes provide a more descriptive picture of the labor force segregation of women workers.

Subsequent to developing our opportunity for jobs index we found that Bowen and Finegan [**5**, pp. 772–776] calculated a similar index for the U.S. which differs from ours in that the denominator of their index for each geographical region is the total civilian employment, rather than the potential female labor force as in the case of our index. For purposes of examining the labor force behavior of married women we feel that our index is more appropriate, although use of either index probably would represent an improvement compared with the common practice in cross-sectional labor force studies of ignoring labor market conditions.

[9] See [**6**] for a more detailed discussion of this question.

TABLE V

MEAN VALUES OF EXPLANATORY VARIABLES FOR ALL WIVES, WIVES WHO DID NOT WORK, AND WIVES WHO WORKED

Variables	25–29 U.S.	25–29 Can.	30–34 U.S.	30–34 Can.	35–39 U.S.	35–39 Can.
1. Years of education	12.2[a]	10.7[a]	11.9	10.1	11.7	9.6
	12.0[b]	9.9[b]	11.8	9.7	11.6	9.3
	12.6[c]	11.5[c]	12.1	10.6	11.9	10.1
2. # of children <6	1.2	1.1	0.8	0.9	0.4	0.5
	1.5	1.4	1.0	1.1	0.6	0.6
	0.7	0.8	0.5	0.6	0.2	0.3
3. # of children 6–14	0.8	0.6	1.8	1.7	1.8	2.0
	0.8	0.8	1.9	1.9	2.0	2.1
	0.7	0.4	1.7	1.5	1.6	1.8
4. Product of # of children <6	0.8	0.6	1.3	1.3	0.9	1.1
and # 6–14	1.0	0.8	1.7	1.7	1.2	1.3
	0.5	0.3	0.8	0.7	0.4	0.6
5. # of children 19–24	0.0	0.0	0.0	0.0	0.0	0.0
attending school	0.0	0.0	0.0	0.0	0.0	0.0
	0.0	0.0	0.0	0.0	0.0	0.0
6. # of children ever born	2.0	1.7	2.8	2.6	3.1	3.1
	2.4	2.2	3.0	3.0	3.4	3.3
	1.5	1.2	2.4	2.1	2.8	2.8
7. Employment income of husband	7,660	6,240	8,543	6,868	9,001	7,085
+ asset income of family net	8,155	6,380	9,174	7,073	9,868	7,229
of income taxes at 0 hours of	6,876	6,094	7,526	6,566	7,832	6,857
work of wife						
8. Variable 7 divided by # of	2,128	1,890	2,011	1,655	2,076	1,581
persons in family	2,031	1,646	2,023	1,565	2,123	1,536
	2,282	2,156	1,991	1,791	2,014	1,654
9. Age of wife at first	19.8	21.4	20.1	21.8	20.7	22.3
marriage	19.5	20.8	20.2	21.5	20.9	22.5
	20.2	22.1	20.1	22.2	20.4	22.1
10. Dummy variable = 1 if	—	0.27	—	0.26	—	0.25
language of home is French;	—	0.31	—	0.29	—	0.30
0 otherwise	—	0.23	—	0.20	—	0.15
11. Dummy variable = 1 if wife	0.08	—	0.08	—	0.07	—
is Black; 0 otherwise	0.05	—	0.05	—	0.04	—
	0.11	—	0.13	—	0.10	—
12. State or provincial	3.5	6.0	3.4	6.0	3.5	5.9
unemployment rate	3.5	6.2	3.5	6.1	3.5	6.1
	3.4	5.8	3.4	5.7	3.4	5.6
13. Local opportunity for jobs	0.64	0.41	0.64	0.41	0.64	0.41
index	0.63	0.40	0.64	0.40	0.64	0.39
	0.64	0.43	0.63	0.42	0.64	0.43

SOURCE: Calculated from the 1 per cent subsample from the 5 per cent primary State Public Use Sample of Basic Records from the 1970 U.S. Census; and from the 1 per cent Family File of the Public Use Sample from the 1971 Canadian Census.

[a] Mean value for all wives in this age group.

[b] Mean value for wives in this age group who did not work.

[c] Mean value for wives in this age group who worked.

TABLE V

MEAN VALUES OF EXPLANATORY VARIABLES FOR ALL WIVES, WIVES WHO DID NOT WORK, AND WIVES WHO WORKED

40–44		45–49		50–54		55–59	
U.S.	Can.	U.S.	Can.	U.S.	Can.	U.S.	Can.
11.6	9.4	11.3	9.2	10.9	9.1	10.5	8.7
11.5	9.0	11.0	8.6	10.6	8.6	10.2	8.2
11.7	9.9	11.5	9.8	11.3	9.8	11.0	9.2
0.2	0.3	0.1	0.1	0.0	0.0	0.0	0.0
0.3	0.3	0.1	0.1	0.0	0.0	0.0	0.0
0.1	0.2	0.0	0.0	0.0	0.0	0.0	0.0
1.1	1.5	0.6	0.9	0.2	0.4	0.0	0.1
1.3	1.7	0.7	1.1	0.3	0.4	0.1	0.1
1.0	1.2	0.5	0.6	0.2	0.3	0.0	0.1
0.3	0.5	0.1	0.2	0.0	0.0	0.0	0.0
0.5	0.7	0.1	0.2	0.0	0.0	0.0	0.0
0.2	0.3	0.0	0.1	0.0	0.0	0.0	0.0
0.1	0.2	0.1	0.3	0.1	0.2	0.1	0.2
0.1	0.2	0.1	0.3	0.1	0.2	0.0	0.1
0.1	0.2	0.1	0.3	0.1	0.2	0.1	0.2
3.1	3.3	2.9	3.2	2.6	3.0	2.4	2.9
3.2	3.6	3.1	3.5	2.7	3.3	2.6	3.0
2.9	2.9	2.6	2.7	2.4	2.7	2.2	2.6
9,232	7,096	8,726	6,980	8,044	6,410	7,301	5,795
10,219	7,131	9,612	7,078	8,651	6,438	7,804	4,794
8,124	7,054	7,793	6,855	7,348	6,369	6,571	5,798
2,393	1,655	2,795	1,950	3,127	2,137	3,182	2,297
2,506	1,571	2,967	1,860	3,322	2,094	3,383	2,266
2,266	1,774	2,612	2,073	2,902	2,206	2,887	2,361
21.1	23.0	21.9	23.7	22.7	24.6	23.2	26.0
21.3	23.1	21.8	23.6	22.6	24.3	23.0	25.9
20.8	22.8	22.0	23.7	22.9	25.1	23.5	26.4
—	0.25	—	0.23	—	0.23	—	0.22
—	0.31	—	0.30	—	0.28	—	0.22
—	0.16	—	0.14	—	0.24	—	0.13
0.07	—	0.06	—	0.07	—	0.06	—
0.06	—	0.06	—	0.06	—	0.06	—
0.08	—	0.07	—	0.08	—	0.07	—
3.4	5.9	3.5	5.9	3.5	5.9	3.4	5.9
3.5	6.2	3.5	6.1	3.5	6.1	3.5	6.1
3.4	5.6	3.4	5.5	3.5	5.6	3.4	5.6
0.64	0.41	0.64	0.41	0.64	0.40	0.63	0.40
0.64	0.39	0.63	0.39	0.63	0.39	0.63	0.39
0.64	0.43	0.64	0.43	0.64	0.43	0.64	0.42

are present, which are contained in the 1 per cent subsample from the 5 per cent primary State Public Use Sample of Basic Records from the 1970 U.S. Census. The basic Canadian data consists of 30,412 records for married couples living in Canada, where the wife is 25–59 years old and no non-relatives are present, which are contained in the 1 per cent Family File of the Public Use Sample from the 1971 Canadian Census. The records for each country were divided into seven groups according to the age of the wife: 25–29, 30–34, 35–39, 40–44, 45–49, 50–54, and 55–59.

The mean values of the explanatory variables used in this study are presented in Table V for these seven age groups for (i) the entire sample, (ii) the subsample of married women who did not work, and (iii) the subsample of married women who worked. From this table we see that in all classifications the average U.S. wife has more years of formal education, married younger, and lives in a family where the combined income of the husband and family asset income are higher, than the average Canadian wife.[10] The younger U.S. wives have more children and the older U.S. wives have fewer children than their Canadian counterparts. Finally, we see that the mean U.S. unemployment rates are always lower, while the U.S. means for our local opportunity for jobs variable are always higher than the Canadian means.

Given these differences in the average characteristics of U.S. and Canadian wives, inter-country differences in labor force behavior are to be expected. One question to be explored though is whether these differences in the characteristics of wives in the two countries account for the observed differences in labor force behavior, or if there are also inter-country differences in responses as measured by the estimated coefficients of our model.

4. EMPIRICAL RESULTS FOR THE PROBABILITY OF WORKING

Values of the net offered wage rate cannot be observed for wives who do not work. However, from our maximization problem we have

$$(16) \qquad P(D_i = 1) = P(h_i > 0) = P(w_h^n - w_h^*|_{h=0} > 0) = \frac{1}{\sqrt{2\pi}} \int_{-\infty}^{\phi_i} e^{-(t^2/2)}\, dt,$$

where D_i is defined to be one or zero depending on whether or not the ith married woman worked for pay or profit in 1969 for the U.S. or in 1970 for Canada. In practice, D_i has been set equal to one if the ith married woman earned at least one dollar of employment income during the relevant year. Linearizing $\ln(RET)$ at $h = 0$ around $E_H T$ as

$$\ln(RET) = \alpha_0' + \eta(E_H T) + u'$$

[10] In the case of Canada it should be noted that the values for the variable for age at first marriage are unreliable for wives over approximately 50 years of age, since all wives who were first married 35 years ago or more are reported as having been first married 35 years ago. Also all dollar figures for the U.S. are in 1969 U.S. dollars while those for Canada are in 1970 Canadian dollars. See footnote 2.

for a joint return where u' denotes a disturbance term, we get by (12) and (14) that (see original version of paper for further details)

$$\ln w_0^n = \ln w + \ln (RET)|_{h=0}$$
$$= (\alpha_0 + \alpha_0') + Z\alpha_1 + R\alpha_2 + \eta(E_H T) + (u + u').$$

Redefine $\alpha_0 \equiv \alpha_0 + \alpha_0'$ and $u \equiv u + u'$. Since

$$P(w_0^n - w_0^* > 0) = P(\ln w_0^n - \ln w_0^* > 0),$$

we have by (11)

$$\phi_i = \frac{1}{\sigma}[(\alpha_0 - \beta_0) + Z_i\alpha_1 - Z_i^*\beta_1 + R_i\alpha_2 - (\beta_2 - \eta)(E_H T)_i], \tag{17}$$

where $(\alpha_0 - \beta_0)/\sigma$, α_1/σ, $-(\beta_1/\sigma)$, α_2/σ, and $-(\beta_2 - \eta)/\sigma$ are the probit coefficients to be estimated by maximum likelihood, and σ^2 is the variance of the random term $U_i^* - u_i$. We note that for a separate return $\ln (RET)_{h=0} = 0$ since $TX_h(wh, A)|_{h=0} = 0$, and hence $\alpha_0' = \eta = 0$. Estimates of the probit coefficients for the U.S. and Canadian wives are shown in Table VI.

The coefficients are very similar for both countries. Also for those coefficients which are significant with at least an 80 per cent confidence level, the coefficient signs are generally in agreement with our expectations. The exceptions are the positive signs of the coefficients of the number of children ever born for the 35–39 and 40–44 year old age groups for the U.S. and the negative coefficients for the age at first marriage for the 35–39 and 40–44 year old age groups for Canada and the 30–34 and 35–39 year old age groups for the U.S.

The consistently positive coefficients for the race variable included in the U.S. equations are also disturbing in the sense that we are not able to offer any explanation for them within the context of our model. Our hope had been that after controlling for education, child-status, the husband's income, family asset income, tax effects, and job opportunities for women the coefficients of this race variable would turn out to be insignificant, or even significantly negative because of the negative effects of discrimination on the offered wage rates of Black wives.[11]

Having noted the similarity of our results for the U.S. and Canada, and the general conformity of these results with our initial expectations, we will conclude

[11] Heckman [**18**] notes that Olsen [**28**] and Smith [**35**] find that for certain groups of women, lower wage women are the ones more likely to participate. Heckman [**18**, p. 205] notes also that: "It is significant that the perverse association between wage rates and participation status is found in demographic groups with the greatest volume of lifetime labor supply—such as married black women." Heckmann argues that these phenomena are manifestations of lifetime income effects associated with the intertemporally correlated offered wage rates of potential labor force participants. This is only one possible source of lifetime income effects, however. It is well known that the age-income profiles for men with little education peak earlier, and then drop more precipitously, than the age-income profiles for men with more education. Also the unemployment rates are higher for men with little education compared with those with more education, and for Black men compared with Whites. Thus the current earned income, or the wage rate, of the husband may tend to be associated with lower lifetime incomes in the case of the husbands of low wage and poorly educated, or Black women than in the case of other husbands.

TABLE VI
Probit Estimates for Married Women in U.S. and Canada[b,c]

Explanatory Variables	25–29 U.S.	25–29 Can.	30–34 U.S.	30–34 Can.	35–39 U.S.	35–39 Can.
1. Constant	−1.293**	−1.168**	.014	−.384*	−.728**	−.489**
	(.34)	(.25)	(.32)	(.23)	(.31)	(.21)
2. Years of education	.084**	.083**	.080**	.055**	.090**	.062**
	(.01)	(.01)	(.01)	(.01)	(.01)	(.01)
3. # of children <6	−.446**	−.399**	−.556**	−.421**	−.527**	−.407**
	(.05)	(.04)	(.06)	(.05)	(.06)	(.05)
4. # of children 6–14	−.018	−.157**	−.107**	−.101**	−.097**	−.088**
	(.05)	(.05)	(.03)	(.03)	(.03)	(.03)
5. Product of # of children <6	0.56**	.097**	.047**	.023*	.027*	.040**
and # 6–14	(.03)	(.02)	(.02)	(.02)	(.02)	(.02)
6. # of children 19–24					.050	.324**
attending school					(.15)	(.12)
7. # of children ever born					.042**	.016
					(.02)	(.02)
8. Employment income of husband	−.00017**	−.00017**	−.00009**	−.00012**	−.00010**	−.00009**
+ asset income of family net	(.00002)	(.00002)	(.00001)	(.00002)	(.00001)	(.00001)
of income taxes at 0 hours of						
work of wife						
9. Variable 8 divided by # of	.00032**	.00037**	.00010**	.00023**	.00016**	.00013**
persons in family	(.00006)	(.00006)	(.00005)	(.00006)	(.00004)	(.00005)
10. Age of wife at first	−.003	.022**	−.014*	.005	−.021**	−.011*
marriage	(.01)	(.01)	(.01)	(.01)	(.01)	(.01)
11. Dummy variable = 1 if wife	.461**		.580**		.440**	
is Black; 0 otherwise	(.08)		(.08)		(.09)	
12. Dummy variable = 1 if		−.198**		−.179**		−.332**
language of home is French;		(.05)		(.05)		(.06)
0 otherwise						
13. State or provincial	−.055**	−.007	−.106**	−.036**	−.039*	−.027**
unemployment rate	(.03)	(.01)	(.03)	(.01)	(.03)	(.01)
14. Local opportunity for jobs	1.949**	1.612**	.622*	1.562**	1.291**	1.919**
index	(.43)	(.25)	(.41)	(.25)	(.43)	(.25)
Combined grouped R^2 = .858 for U.S. and .890 for Canada.[a]						
Pseudo R^2	.2170	.2348	.1603	.1511	.1485	.1195
Maximum R^2 for model	.7363	.7495	.7353	.7395	.7442	.7351
Pseudo R^2 for model (Pseudo R^2 divided by maximum R^2 for model)	.2947	.3133	.2180	.2044	.1995	.1626
# of married women in sample	4,761	5,541	4,281	4,762	4,255	4,613
Proportion of married women who worked	.51	.48	.47	.40	.50	.38
Final value of log of likelihood function	−2,081	−3,094	−2,119	−2,812	−2,239	−2,771

SOURCE: Calculated from the 1 per cent subsample from the 5 per cent primary State Public Use Sample of Basic Records from the 1970 U.S. Census; and from the 1 per cent Family File of the Public Use Sample from the 1971 Canadian Census.

[a] Explained in text.

[b] Numbers in parentheses are standard deviations.

[c] Coefficients with two asterisks are significant at a 95 per cent level. Coefficients with one asterisk are significant at an 80 per cent level.

TABLE VI

PROBIT ESTIMATES FOR MARRIED WOMEN IN U.S. AND CANADA[b,c]

40–44		45–49		50–54		55–59	
U.S.	Can.	U.S.	Can.	U.S.	Can.	U.S.	Can.
−.873**	−.574**	−1.712**	−.983**	−1.693	−1.464**	−1.739**	−1.729**
(.30)	(.20)	(.30)	(.19)	(.30)	(.20)	(.34)	(.23)
.071**	.053**	.081**	.067**	.082**	.064**	.083**	.075**
(.01)	(.01)	(.01)	(.01)	(.01)	(.01)	(.01)	(.01)
−.614**	−.402**						
(.08)	(.06)						
−.179**	−.156**	−.143**	−.121**	−.154**	−.099**	−.040**	−.110*
(.03)	(.02)	(.03)	(.02)	(.05)	(.03)	(.11)	(.08)
.127**	.063**						
(.03)	(.02)						
.052	−.002	.118**	.108**	.085	.064*	.208*	.252**
(.07)	(.05)	(.06)	(.04)	(.08)	(.05)	(.11)	(.06)
.045**	−.000	−.008	−.018*	−.004	−.009	−.039**	.004
(.02)	(.02)	(.01)	(.01)	(.01)	(.01)	(.01)	(.01)
−.00007**	−.00005**	−.00007**	−.00008**	−.00005**	−.00005**	−.00006**	−.00008**
(.00001)	(.00001)	(.00001)	(.00001)	(.00001)	(.00001)	(.00002)	(.00002)
.00005*	.00003	.00003*	.00009**	.00001	.00002	.00003	.00008**
(.00003)	(.00004)	(.00002)	(.00003)	(.00003)	(.00003)	(.00003)	(.00004)
−.005	−.008*	.003	−.001	.001	.011**	−.001	.005
(.01)	(.01)	(.005)	(.004)	(.004)	(.004)	(.004)	(.005)
.168**		.282**		.198**		.113	
(.08)		(.09)		(.09)		(.10)	
	−.252*		−.286**		−.270**		−.331**
	(.06)		(.06)		(.07)		(.08)
−.048*	−.032**	−.033*	−.025**	−.016	−026*	−.023	−.016
(.03)	(.01)	(.03)	(.01)	(.03)	(.01)	(.03)	(.02)
1.380**	2.143**	2.173**	2.253**	1.832**	2.080**	1.804**	2.143**
(.40)	(.24)	(.40)	(.25)	(.44)	(.26)	(.47)	(.30)
.0963	.1096	.0799	.1171	.0594	.0854	.0604	.0902
.7491	.7426	.7498	.7442	.7487	.7364	.7411	.7180
.1285	.1476	.1066	.1574	.0793	.1160	.0815	.1256
4,567	4,570	4,525	4,476	3,813	3,509	3,181	2,941
.53	.41	.54	.42	.52	.38	.46	.33
−2,612	−2,836	−2,640	−2,773	−2,251	−2,182	−1,850	−1,722

this discussion of our probit results by noting that the coefficients of the education and the local opportunity for jobs variables are generally slightly more positive for the U.S. than for Canada; while the coefficients for the number of children younger than 6, the number of children 6–14, and the unemployment variable are persistently more negative for the U.S. than for Canada. One possible explanation of this observed pattern of coefficient differences is that U.S. wives, who live on the average in more affluent homes than Canadian wives, are freer to choose the circumstances under which they will work.

The pseudo R^2s for our probit equations, shown in Table VI, range from .0793 to .2947 for the U.S. and from .1160 to .3133 for Canada. Thus we cannot use these estimated equations to make accurate predictions about whether any particular wife will choose to work. These equations do explain a fairly large proportion of the variability in group behavior, however. The estimated probit coefficients in Table VI were used to compute the normal probability of working for each wife in each of our two data bases (all U.S. wives 25–59 years old, and all Canadian wives 25–59 years old). The work decision of each wife was then simulated by comparing her estimated probability of working with a random number between 0 and 1 from a uniform distribution. Next we grouped the wives of each nationality according to their age (25–34, 35–44, 45–59), their education (less than 12 years, complete high school, bachelor or first professional degree), their child status (no children younger than 14 years of age, children younger than 6 only, children 6–14 only, both children younger than 6 and children 6–14), the earned income of the husband plus family asset income net of taxes at zero hours of work for the wife (<\$3,000, \$3,000–\$5,999, \$6,000–\$8,999, . . . , ≥\$15,000), place of residence (urban, rural), and region (West, Northeast, North Central, and South for the U.S.; Maritimes, Quebec, Ontario, Prairies, and British Columbia for Canada). For each country then we regressed the simulated on the actual proportions of women working within each of these groups (1,728 for the U.S. and 2,160 for Canada) using generalized least squares to control for differences in group size. The R^2s from these two regressions, referred to in Table VI as the combined grouped R^2s, indicate that the estimated relationships shown in Table VI explain approximately 86 per cent for the U.S. and 89 per cent for Canada of the variation among our groups in the proportions of wives working.

5. EMPIRICAL RESULTS FOR THE OFFERED WAGE EQUATION

Because wage rates and hours of work are observed only for those wives in our sample who worked, we have a selection bias problem [**13**, **24**]. Since our equilibrium hours equation (15) involves the endogenous variables h_i and w_i on the right-hand side, we cannot include a selection bias term in this equation[12] as is

[12] We are indebted to a referee for noting this problem in an earlier version of the paper. A related error in exposition remains uncorrected in our original Canadian study [**27**], since the equilibrium hours equation for that study involves the endogenous variable w_i. For this reason, the consistency arguments given in our earlier paper are also incorrect; rather consistency may be shown as a special case of the proof given in the Appendices of this paper. These errors in exposition in our earlier paper do not affect either the computational results or the interpretation of these results, however.

done in [**17**]. In order to account for the selection bias, we rather rewrite the hours equation (15) in its reduced form (see Appendix 1, equations (A7)–(A11)):

$$h_i = \gamma_0 + \gamma_1 \overline{\ln w_i} + (\gamma_1 \xi) \ln \{1 - TX_h((E_H)_i + \bar{w}_i h_i^*, A_i)\} + Z_i^* \gamma_2 + \gamma_3 (E_H T)_i + u_i^*, \tag{18}$$

where h_i^* is the solution to

$$h_i^* = \gamma_0 + \gamma_1 \overline{\ln w_i} + (\gamma_1 \xi) \ln \{1 - TX_h((E_H)_i + \bar{w}_i h_i^*, A_i)\} + Z_i^* \gamma_2 + \gamma_3 (E_H T)_i, \tag{19}$$

$$\overline{\ln w_i} = \alpha_0 + Z_i \gamma_1 + R_i \alpha_2, \tag{20}$$

and

$$\bar{w}_i = e^{\{\alpha_0 + Z_i \alpha_1 + R_i \alpha_2\}} = e^{\overline{\ln w_i}}. \tag{21}$$

Since h_i^* does not depend on h_i by construction (see Appendix 1), we can include a selection bias term [**17**] in equations (12) and (18) calculated for each wife who worked as $\lambda_i = f(\phi_i)/F(\phi_i)$, where $f(\phi)$ and $F(\phi)$ are, respectively, the density and cumulative density functions of the standard normal distribution, ϕ_i is given by (17), and where u_i and u_i^* are assumed to be bivariate normal with mean zero, variances σ_1^2 and σ_2^2, respectively, and covariance σ_{12}. The resulting equations to be estimated are therefore

$$\ln w_i = \alpha_0 + Z_i \alpha_1 + R_i \alpha_2 + (\sigma_{12}/\sigma_2)\lambda_i + V_i \tag{22}$$

and

$$h_i = \gamma_0 + \gamma_1 \overline{\ln w_i} + (\gamma_1 \xi) \ln \{1 - TX_h((E_H)_i + \bar{w}_i h_i^*, A_i)\} + Z_i^* \gamma_2 + \gamma_3 (E_H T)_i + \sigma_2 \lambda_i + V_i^*, \tag{23}$$

where the covariance structure of the bivariate normal disturbance terms V_i and V_i^* is given in [**17**]. Since true λ_i cannot be observed in practice, we use its consistent estimate $\hat{\lambda}_i = f(\hat{\phi}_i)/F(\hat{\phi}_i)$ instead in equations (22) and (23). Least squares and generalized (weighted in our case) least squares will still provide consistent estimates of the parameters for equation (22) and also for equation (23). (The consistency proof is given in Appendices 1–3.)

The values of w_i for the ith wife were calculated by dividing her employment income for 1969 for the U.S. or 1970 for Canada by an estimate of her annual hours of work for the same year. In the case of Canada, annual hours of work, h_i, were computed by multiplying the number of weeks the wife worked in 1970 times her usual number of hours worked per week for the job held in the reference week for the 1971 Canadian Census, or otherwise for the job of longest duration held since January 1 of the previous calendar year. For the U.S. the values of h_i were computed by multiplying the number of weeks the wife worked in 1969 times the actual number of hours she worked at all jobs in the reference week for the 1970 U.S. Census if she was "at work" during this week. For U.S. wives who were not at work during this reference week, we have no measure of hours worked per week and hence no way of directly computing annual hours of work.

These U.S. wives in each of our age groups who earned employment income in 1969 but were not at work in the reference week were grouped by weeks worked in 1969 (<13, 14–26, 27–39, 40–47, 48–49, 50–52), and mean values for the variables included in our probit analysis were calculated for each group. These means were found, in general, to be insignificantly different with a 90 per cent level of confidence from the comparable means for U.S. wives who earned employment income in 1969 and were at work in the reference week. However, when the wives in each age and weeks worked category who worked in both 1969 and the reference week were grouped more finely according to our variables for education, child status, the employment income of the husband plus family asset income, and local job opportunities, the standard deviations for hours of work in the reference week were found to be quite large. Still finer groupings did not substantially improve this problem. Moreover the R^2s for various functions of variables which we fit within groups were found to be uniformly low. In order to impute hours of work for U.S. wives who earned employment income in 1969 but were not at work in the reference week, based on data for U.S. wives who were at work in the reference week, we would have had to ignore this individual variability. Instead, therefore, these U.S. wives with missing hours data were dropped at this point in our analysis from our U.S. data base. Since there is some tendency for the distribution of weeks worked in 1969 by these wives with missing hours data to be positively skewed, the result of dropping these wives from our analysis may well be to introduce some additional heteroscedasticity into the disturbance terms for our U.S. offered wage and hours of work equations. It should be borne in mind that this particular problem does not affect our Canadian results.[13]

Generalized least squares estimates of the coefficients of (22) are shown in Table VII. Again we find that, for those coefficients which are significant with at least an 80 per cent confidence level, the coefficient signs generally agree with our expectations.[14] The exceptions are the positive coefficients for the number of children younger than 6 for the 30–34 year old age group for Canada, and for the

[13] See [**36**] and [**37**] for further details. While this difference between the U.S. and Canada in the measurement of hours of work per week is clearly important, we have not been able to identify any systematic effects of this difference on our coefficient estimates shown in Tables VII–IX. Other sources of data for the U.S., such as the 1967 National Longitudinal Survey used by Heckman [**16**, **17**, **19**], provide what appear to be better measures of the offered wage rate. However, there are problems with these other data sources as well. For instance, when Heckman [**19**] computes the annual hours of work by dividing 1966 earnings by a questionnaire wage rate, he obtains one outlying observation as high as 5,473 hours per year. (This figure could only be obtained in reality by working 15 hours per day, 7 days per week, and 52 weeks per year.) Garfinkel [**11**], among others, has shown that a small number of such extreme observations can seriously distort estimation results even when they are based on several thousand observations. Moreover the alternative data sources available to us for the U.S. would not allow us to make even the limited comparisons we have attempted between estimation results for the U.S. and Canada.

[14] The standard errors shown in Table VII for the offered wage equation were computed using the usual weighted least squares formula. In the case of Table VIII, the estimates of the standard error of the regression used in computing the standard errors shown for the coefficients of the offered wage variable have been computed by substituting the actual for the estimated values of the log of the

unemployment rate for the 40–44 and 45–49 year old age groups for the U.S. and for the 40–44 year old age group for Canada. As is the case for Gronau's [**13**] results and our earlier results for Canada [**27**], but in marked contrast to Heckman's [**17, 19**] results for U.S. wives, we find the selectivity bias in the offered wage equation to be significant with at least an 80 per cent level of confidence for all age groups except 25–29 for the U.S. and 30–34 for Canada. This result is of some importance, since Heckman's finding of no selectivity bias suggests that wage data for working wives can be directly used to impute the potential offered wage rates of wives who are not working, as researchers such as Hall [**14**] and Harvey Rosen [**31**] have done in their studies. Our results, like Gronau's, suggest on the other hand that this procedure should not be used.[15]

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offered wage into our estimated unconstrained annual hours of work equation. This adjustment has been made to account for the fact that the estimated values of the log of the offered wage are orthogonal to the least squares residual series corresponding to these estimates. Thus these standard errors were computed just as one normally computes the standard errors for the coefficients in the second stage when performing two stage least squares estimation. (Miller and Modigliani [**26**, footnote 39], for instance, make this same sort of adjustment in some of their estimated coefficients in a cost of capital model which, like our model, is not fully simultaneous.) Similar adjustments were also made in the standard errors (which are not shown) for the coefficients of the number of children younger than 6 and of λ, since both of these variables also appear in our offered wage equation and hence are orthogonal to the residuals from this equation.

Similar adjustments were made too in the standard errors shown in Table IX for the coefficients of the number of children younger than 6 and of λ, since the residual from the offered wage equation is one component of the true error term for our constrained annual hours of work equation for all iterations in our estimation procedure. No such adjustments have been made in the standard errors for the coefficients of the log of the wife's net offered wage, however, since this composite variable is not orthogonal to the residuals from the offered wage equation. Thus these standard errors, and all the remaining standard errors shown in Tables VIII and IX, have been computed using the usual weighted least squares formulas. (It is perhaps worth noting that the above mentioned adjustments which we have made in the standard errors shown for some of our coefficients in Tables VIII and IX turned out to be numerically trivial.)

Heckman [**19**] has shown that the standard errors which we present in Tables VII–IX might be downward biased, owing to the sample selection and the fact that λ must be estimated. Heckman [**19**] presents expressions for both ordinary and weighted least squares which correct for this problem. His corrected expressions for weighted least squares appear to be computationally intractable, however. In fact, Heckman himself resorts to ordinary least squares estimates throughout [**19**] despite his arguments in previous papers [**16, 17**] concerning the importance of heteroscedasticity in this model. In other words, it would appear that better estimates of the coefficient standard errors have been gained at the expense of poorer coefficient estimates. It should also be pointed out, however, that his corrected expression for standard errors would reduce to the ones used in this paper if there were no selection bias.

We have chosen to correct for heteroscedasticity in obtaining our coefficient estimates, while ignoring the biasedness of our estimated standard errors for these coefficients. Thus our statements in the text with respect to the significance of the coefficients shown in Tables VII–IX must be viewed with caution. Following the methodology espoused by Simon Kuznets [**22**, p. 233], however, we believe that our substantive conclusions are strongly supported by the fact that in most cases we obtain similar results for two countries and multiple age groups.

[15] One of the referees questioned whether perhaps our selection bias term is accounting here for the effects of other child status variables included in our probit and hours of work equations, but omitted from our offered wage equation, which may be systematically related to a wife's work experience. We note first of all that the number of children younger than 6 is by far the most powerful of the child status variables included in either our probit or hours of work equations. Yet this variable proved to be generally insignificant in our offered wage equation. Secondly, in exploratory regressions we in fact introduced our other child status variables into our offered wage equation, both singly and in various combinations, with even less impressive results.

TABLE VII

GLS ESTIMATES FOR LOG OF OFFERED WAGE EQUATION FOR MARRIED WOMEN IN U.S. AND CANADA[b,c]

Explanatory Variables	25–29 U.S.	25–29 Can.	30–34 U.S.	30–34 Can.	35–39 U.S.	35–39 Can.
1. Constant	−.857**	.257*	−1.191**	−.210	−1.791**	−.738**
	(.319)	(.194)	(.323)	(.239)	(.396)	(.353)
2. Years of education	.078**	.070**	.068**	.067**	.087**	.067**
	(.008)	(.005)	(.008)	(.006)	(.008)	(.007)
3. # of children <6	.005	.001	−.047	.050*	.004	−.083**
	(.042)	(.031)	(.047)	(.040)	(.049)	(.044)
4. Age of wife at first marriage	.012**	.008*	.012**	.005	.010**	.008**
	(.006)	(.004)	(.005)	(.004)	(.005)	(.005)
5. Dummy variable = 1 if wife is Black; 0 otherwise	−.044		−.193**		−.156**	
	(.057)		(.059)		(.061)	
6. State or provincial unemployment rate	.022	.003	.026	.004	.009	.004
	(.022)	(.008)	(.026)	(.012)	(.023)	(.013)
7. Local opportunity for jobs index	.211	−.242*	.388*	.479**	.960**	.646**
	(.279)	(.182)	(.271)	(.231)	(.328)	(.256)
8. Selection bias (λ)	.114	−.218**	.386**	−.070	.299**	.277*
	(.138)	(.105)	(.148)	(.161)	(.108)	(.174)
Combined grouped R^2 = .804 for U.S. and .739 for Canada.[a]						
R^2	.2706	.3378	.1456	.2118	.1195	.1115
Standard error of regression	1.00	1.00	1.01	1.01	1.00	1.00
# of married women in sample	1,472	2,651	1,399	1,899	1,580	1,755

SOURCE: Calculated from the 1 per cent subsample from the 5 per cent primary State Public Use Sample of Basic Records from the 1970 U.S. Census; and from the 1 per cent Family File of the Public Use Sample from the 1971 Canadian Census.

[a] Explained in text.

[b] Numbers in parentheses are standard deviations.

[c] Coefficients with two asterisks are significant at a 95 per cent level. Coefficients with one asterisk are significant at an 80 per cent level.

The only systematic national differences in the coefficients are that the education and age at first marriage coefficients are slightly larger in magnitude for the U.S. than for Canada for all age groups except 45–49 for the education variable and 55–59 for the marriage variable.

The U.S. coefficients for the dummy variable set equal to 1 if the wife is Black are negative for all age groups, and are significant at at least an 80 per cent level of confidence for all age groups except 25–29 and 50–54. These results support the widespread belief that Black wives receive lower wages than other wives with seemingly similar characteristics. Also the coefficients for the local opportunity for jobs index are positive and significant at at least an 80 per cent confidence level for all age groups except 25–29 for both the U.S. and Canada.

Combined grouped R^2s were also calculated for our offered wage equations using the same groups on which the combined grouped R^2s for our probit equations are based. These R^2s indicate that, for these groups, the estimated relationships presented in Table VII explain approximately 80 per cent for the U.S. and 74 per cent for Canada of the variation in the mean offered wage rate.

6. EMPIRICAL RESULTS FOR THE HOURS OF WORK EQUATION

Our hours equation (23) cannot be estimated directly, since hours of work depends on the deterministic part of the wage and the tax rate which depends on

TABLE VII

GLS ESTIMATES FOR LOG OF OFFERED WAGE EQUATION FOR MARRIED WOMEN IN U.S. AND CANADA[b,c]

40–44		45–49		50–54		55–59	
U.S.	Can.	U.S.	Can.	U.S.	Can.	U.S.	Can.
−2.551**	−.497**	−2.055**	−1.020**	−2.631**	−.567*	−3.214**	−1.748**
(.186)	(.208)	(.443)	(.168)	(.474)	(.311)	(.632)	(.601)
.071**	.052**	.065**	.066**	.094**	.056**	.096**	.079**
(.007)	(.005)	(.006)	(.005)	(.007)	(.007)	(.010)	(.010)
−.128**	−.060						
(.060)	(.057)						
.010**	.001	.007**	.003	.008**	.007*	.007*	.011**
(.004)	(.004)	(.003)	(.003)	(.003)	(.004)	(.004)	(.004)
−.163**		−.170**		−.044		−.112*	
(.054)		(.057)		(.063)		(.089)	
.044**	.016*	.060**	−.006	.016	−.012	−.006	.009
(.021)	(.012)	(.020)	(.012)	(.023)	(.013)	(.030)	(.016)
.692**	.818**	1.161**	1.045**	1.325**	1.183**	2.417**	1.265**
(.197)	(.186)	(.308)	(.193)	(.343)	(.286)	(.483)	(.355)
1.323**	.214*	.507**	.524**	.519**	.901*	.583**	.437*
(.156)	(.169)	(.133)	(.118)	(.120)	(.521)	(.186)	(.267)
.4278	.2712	.0954	.3445	.1458	.0679	.1459	.2063
1.05	1.02	1.00	1.01	1.00	1.00	1.00	1.00
1,915	1,896	1,968	1,900	1,583	1,352	1,167	965

the deterministic parts of both the offered wage and hours of work. We will first discuss the estimation of this equation for a U.S. joint tax return.

In this case, replacing $\overline{\ln w_i}$ by its predicted values $\widehat{\ln w_i}$ from (22) and $\bar{w}_i$ by $\hat{w}_i = e^{\widehat{\ln w_i}}$, the equation to be estimated is

$$(24)\qquad h_i = \gamma_0 + \gamma_1 \widehat{\ln w_i} + (\gamma_1 \xi) \ln \{1 - TX((E_H)_i + \hat{w}_i h_i^*, A_i)\} + Z_i^* \gamma_2 + \gamma_3 (E_H T)_i + \sigma_2 \hat{\lambda}_i + V_i^*,$$

where the subscript h has been dropped from TX_h for notational convenience. Note that $\widehat{\ln w_i}$ and $\hat{w}_i$ in equation (24) are both functions of exogenous variables. Also for notational convenience, in the following argument we will denote the right-hand side of (24) by $F_1(\Gamma, h_i^*) + V_i^*$ where Γ corresponds to the collection of parameters $(\gamma_0, \gamma_1, (\gamma_1 \xi), \gamma_2, \gamma_3, \sigma_2)$. Since $F_1(\Gamma, h_i^*)$ depends on the unknown Γ and h_i^*, we estimate Γ using the following iterative process (where the proof of convergence and consistency for this estimation process is given in Appendices 2 and 3): (i) Set $k = 0$ and $\hat{h}_i^{(k)} = 1$. (ii) Calculate $RET_i^{(k)} = 1 - TX((E_H)_i + \hat{w}_i \hat{h}_i^{(k)}, A_i)$ for each married woman i. (iii) Regress h_i on $\widehat{\ln w_i}$, $\ln RET_i^{(k)}$, Z_i^*, $(E_H T)_i$, and $\hat{\lambda}_i$ using GLS, and let $\Gamma^{(k)}$ denote the estimated regression coefficients. (iv) Let the predicted values of h_i be $\hat{h}_i^{(k+1)} = F_1(\Gamma^{(k)}, \hat{h}_i^{(k)})$. (v) Set $k = k + 1$, and go to step (ii). The iterative process is terminated when, for example, two successive sets of estimates for $\Gamma^{(k)}$ and $\Gamma^{(k+1)}$ are sufficiently close to each other in terms of percentage changes.[16]

[16] See Section 5 of text and footnote 13 for details concerning the computation of annual hours of work.

In the case of separate returns for Canada the only change to be made in this iterative process is to set $RET_i^{(0)} = 1$ (or $\ln RET_i^{(0)} = 0$); hence for $k = 0$, h_i is regressed on $\widehat{\ln w_i}$, Z_i^*, $(E_H T)_i$, and $\hat{\lambda}_i$. The rest is identical to the case of a joint return.

Let $t_i(h_i)$ denote the sum of state and federal taxes paid by the ith U.S. family given that the wife works h_i hours.[17] These taxes are computed by a computer program which incorporates most of the basic provisions of the 1969 U.S. federal and individual state tax tables. First the taxable income for each family is computed based on the earned income of the husband, the asset income of the family, the estimated income of the wife (her estimated offered wage times her estimated hours of work), and the standard demographic deductions (for a nonworking spouse, dependent children, and old age) to which the family is entitled. Then, based on the amount of this taxable income, the appropriate federal and (where applicable) state tax rates are applied.

Then $RET_i^{(k)}$ is given by

$$RET_i^{(k)} = 1 - TX((E_H)_i + \hat{w}_i \hat{h}_i^{(k)}, A_i) \tag{25}$$

where

$$TX((E_H)_i + \hat{w}_i \hat{h}_i^{(k)}, A_i) = \frac{t_i(\hat{h}_i^{(k)} + 1) - t_i(\hat{h}_i^{(k)})}{\hat{w}_i}. \tag{26}$$

For Canadian separate returns the calculation of the values of $RET_i^{(k)}$ is similar to the procedure for the U.S.: The values of $t_i(h_i)$, the sum of the provincial and federal income taxes paid by the ith Canadian family given that the wife works h_i hours, are calculated based on demographic deductions (the number and age distribution of children and the ages of the couple), the province in which the family resides, the estimated income of the wife, the income of the husband, and the asset income of the family. Deductions for children are applied to the husband's or wife's income, whichever is higher. Asset income is attached to the smaller of the husband's or the wife's income for tax purposes.[18] The basic way in which Canadian separate returns differ from a U.S. joint return is that in the Canadian case the federal and provincial income taxes that the wife must pay on her earnings are independent of her husband's income (except that the husband cannot claim a deduction for his wife for tax purposes).[19] For Canada, then,

[17] The U.S. federal tax tables for 1969 [**20**] and the state income tax tables [**1**] are closely followed to compute $t_i(h_i)$ for each U.S. family. Asset income is handled differently from earned income in certain states for state income tax purposes, and these rules are followed in our procedure. It is assumed where applicable in computing state and federal income taxes that the family maximizes its tax advantage by assigning asset income to either the husband's or wife's earned income, whichever gives the smaller tax value. We do assume for computational convenience, however, that asset income cannot be split for tax purposes between the husband and wife.

[18] Canadian federal and provincial tax tables, excluding Quebec, for 1970 [**8**] and the Quebec income tax table for 1970 [**12**] are closely followed to compute the values for $t_i(h_i)$.

[19] The U.S. Social Security tax and the Canada Pension Plan payments have not been included in this study, since these payroll taxes represent a type of forced savings and hence are quite different from federal, and from state and provincial income taxes.

$RET_i^{(k)}$ is given by

$$(27) \qquad RET_i^{(k)} = 1 - TX(\hat{w}_i \hat{h}_i^{(k)}, A_i)$$

where

$$(28) \qquad TX(\hat{w}_i \hat{h}_i^{(k)}, A_i) = \frac{t_i(\hat{h}_i^{(k)}+1) - t_i(\hat{h}_i^{(k)})}{\hat{w}_i}.$$

Our iterative estimation process converged satisfactorily for all age groups for the U.S. by the third iteration, and for four out of seven age groups for Canada by the fourth iteration.[20] For the age groups 30–34, 35–39, and 50–54 for Canada, all regression coefficients except those for the log of the wife's offered wage and the retention rate variable converged by the fourth iteration.[21]

In Table VIII we report coefficient estimates for the log of the offered wage and the log of the retention rate variable in equation (23), which was estimated using the iterative GLS procedure described above. Also shown in Table VIII are the ratios of the coefficients of the retention rate variable divided by the coefficients of the offered wage variable. These ratios are estimates of the tax perception coefficient, ξ. A consistent pattern of estimates of ξ exceeding 1 would suggest that wives on the average underestimate the impact of taxation on their net earnings in choosing their hours of work, while estimates of ξ less than 1 would suggest that wives tend to overreact to tax losses. The estimates reported in Table VIII for ξ do not support either of these scenarios. Moreover the t statistics shown in Table VIII for the linear restriction that the coefficients of the offered wage and retention rate variables are equal allow rejection of the null hypothesis that the coefficients are equal, and hence that ξ equals 1, with a 95 per cent level of confidence only for the 45–49 and 55–59 year old groups for the U.S. and the 45–49 year old age group for Canada.[22]

These results thus provide limited support for the conclusion of Harvey Rosen [**31**] and Hausman and Wise [**15**] that individuals do fully account for the impact

[20] Our estimation procedure gives consistent estimates of the wage and hours equations. (See Appendices 2 and 3 for convergence and consistency proof.) We have not proceeded to take one Gauss-Newton step toward maximizing the likelihood which would give asymptotically efficient estimators as proposed, for instance, in [**33**] for the following reasons: (i) To take such a step, linearization of our retention rate variable (*RET*) in terms of endogeous variables h and w as well as exogenous variables is required and such linearization clearly destroys the fundamental nonlinear dependence of the marginal tax rate on these interdependent variables; (ii) computational results of such a step toward maximizing the likelihood reported in [**17**] for a model simpler than ours do not appear to provide sufficient evidence for statistically meaningful improvement. In discussing these results, Heckman [**17**, p. 490] concludes that, "the first step iterate of the initial consistent estimator, an asymptotically efficient estimator, is numerically close to the maximum likelihood estimator but for most coefficients is not as close as the initial consistent estimator." Nor have we used instrumental variable methods [**2**, **21**] for nonlinear systems for the reason similar to (i) above that our retention rate variable must be regressed linearly on instruments which are, for example [**21**], polynomials of exogenous variables, thus resulting in arbitrary linearization of *RET* and the loss of direct dependence of *RET* on h.

[21] See Appendices 1–3 for the condition of convergence for our algorithm.

[22] We are grateful to Christopher A. Sims for suggesting that we test this linear restriction, and that we then reestimate our hours equation incorporating this constraint.

TABLE VIII

UNCONSTRAINED GLS ESTIMATES FOR ANNUAL HOURS OF WORK EQUATION FOR MARRIED WOMEN IN U.S. AND CANADA[a,b]

	25–29		30–34		35–39	
	U.S.	Can.	U.S.	Can.	U.S.	Can.
Log of Offered Wage	−563.88**	−509.36**	−391.47**	−265.86**	−202.37*	−400.16**
	(133.94)	(126.82)	(114.34)	(136.17)	(112.55)	(143.20)
Log of Retention Rate	−338.56*	−662.88*	−437.88*	−239.21**	−440.59**	−1,109.70**
	(210.19)	(425.57)	(178.36)	(91.63)	(185.63)	(396.44)
Tax Perception	.60	1.30	1.12	.90	2.18	2.77
Test Statistic for Null Hypothesis Coefficients Are Equal	.89	.40	.22	.17	1.12	1.79*

SOURCE: Calculated from the 1 per cent subsample from the 5 per cent primary State Public Use Sample of Basic Records from the 1970 U.S. Census; and from the 1 per cent Family File of the Public Use Sample from the 1971 Canadian Census.

NOTE: The other variables included in these regressions for each age group are the same variables for which coefficient estimates are shown in Table IX.

[a] Numbers in parentheses are standard deviations.

[b] Coefficients with two asterisks are significant at a 95 per cent level. Coefficients with one asterisk are significant at an 80 per cent level.

of taxes on their net earnings in choosing their hours of work. These results also suggest that the efficiency of our parameter estimates might be improved by reestimating our hours equation (23) subject to the constraint that the coefficients of the offered wage and retention rate variables are equal. This is equivalent to estimating the equation

$$h_i = \gamma_0 + \gamma_1 \overline{\ln w_h^n} + Z_i^* \gamma_2 + \gamma_3 (E_H T)_i + \sigma_2 \lambda_i + V_i^*, \tag{29}$$

where

$$\overline{\ln w_h^n} = \overline{\ln w_i} + \ln \{1 - TX((E_H)_i + \bar{w}_i h_i^*, A_i)\},$$

using the same iterative procedure used to estimate (23). The empirical equation actually used is (24) in which the log of the wage and log of the tax terms are combined. (The same convergence and consistency proof as before applies.) The resulting coefficient estimates are shown in Table IX.

As expected, the estimated standard errors of the estimates of γ_1 are, in general, substantially smaller for the constrained than for the unconstrained hours equation. The coefficient estimates shown in Table IX for the other variables in equation (29) are almost identical to the unreported coefficient estimates for these same variables in our unconstrained hours equation. However, the estimated standard errors of these coefficients are almost always smaller for the constrained equation.

All coefficients which are significant with at least an 80 per cent level of confidence have the expected signs with the exception of the coefficient for the number of children 19–24 attending school for the 50–54 year old age group for the U.S. In particular, the coefficients of the selection bias term are found to be positive and significant with at least an 80 per cent level of confidence for all age groups for both the U.S. and Canada. This is reassuring since the specifications of

TABLE VIII

UNCONSTRAINED GLS ESTIMATES FOR ANNUAL HOURS OF WORK EQUATION FOR MARRIED WOMEN IN U.S. AND CANADA[a,b]

40–44		45–49		50–54		55–59	
U.S.	Can.	U.S.	Can.	U.S.	Can.	U.S.	Can.
−33.85	96.31	−31.08	−108.33	197.58	315.85**	722.56**	114.17
(104.73)	(185.33)	(113.13)	(150.86)	(180.21)	(145.26)	(243.02)	(252.40)
−55.66	263.39	399.46**	−906.53**	97.65	−27.27	50.55	119.27
(191.99)	(349.44)	(123.60)	(322.60)	(99.42)	(328.89)	(226.92)	(112.67)
1.64	2.73	−12.85	8.37	.49	−.09	.07	1.04
.11	.52	2.86**	2.66**	.49	.83	2.20**	.02

our model require the coefficients for this variable to be positive (see equation (29)).

The generally negative coefficients of the offered wage variable for both the U.S. and Canada, which are significant with at least an 80 per cent level of confidence for the three younger age groups for the U.S. and for the four younger groups for Canada, are in marked contrast to the findings of other researchers of a strong positive relationship, however. Harvey Rosen [**31**, p. 503], for instance, concludes his analysis of his hours of work equation by concurring with Hall [**14**, p. 131] that, "those (wives) with higher wages work substantially more than those with lower wages in the same income group." This difference between our findings and those of other researchers is believed to be due to our corrections for income taxes, differences in the form in which the labor supply function is estimated, the choice of variables used to control for the child status of wives, and, in the case of Heckman's studies, peculiarities in the procedure used to estimate the impact of the offered wage rate on hours of work. The latter three of these issues will now be considered one at a time. (The first of these issues is discussed in the introduction.)

A common practice in studying the labor supply of wives is to first impute offered wage rates to wives who did not work, or to all wives, based on data for the wives who did work. An hours of work, or overall labor supply, equation is then estimated for all wives, including those who worked zero hours.[23] Besides the selection bias problems noted by Gronau [**13**], Lewis [**24**], and Heckman [**16**], an additional drawback of this procedure is that it implicitly assumes that the labor supply responses of wives to changes in their offered wage rates will be the same whether or not these wives are already working. Ben-Porath [**4**, p. 702] argues to the contrary that: "For those who are out of the labor force there is no income effect in a wage rise, and only substitution works . . . ." Thus, on the average, the probability of working must increase as the offered wage rate increases. Those who are already working, however, will experience both substitution and income effects, and may actually reduce their hours of work, though probably not to zero.

[23] See, for instance, [**14** and **31**].

TABLE IX

CONSTRAINED GLS ESTIMATES FOR ANNUAL HOURS OF WORK EQUATION FOR MARRIED WOMEN IN U.S. AND CANADA[b,c]

	25–29		30–34		35–39	
	U.S.	Can.	U.S.	Can.	U.S.	Can.
1. Constant	2659.19**	2234.74**	2409.76**	2008.95**	2610.40**	2128.44**
	(174.62)	(152.83)	(133.09)	(207.36)	(153.87)	(200.52)
2. Log of wife's net offered wage	−550.85**	−460.68**	−342.84**	−315.28**	−237.78**	−355.12**
	(123.54)	(121.42)	(109.30)	(160.66)	(110.04)	(156.58)
3. # of children <6	45.00	−143.99**	−94.80*	−103.34*	−212.19**	−158.85**
	(53.70)	(50.59)	(62.26)	(72.33)	(73.08)	(64.59)
4. # of children 6–14	23.33	−1.25	−9.48	22.87	−99.81**	−16.38
	(38.25)	(36.67)	(27.97)	(28.90)	(25.68)	(26.48)
5. Product of # of children <6 and # 6–14	−25.64	−11.39	8.85	−18.63	72.10**	21.05*
	(20.61)	(18.47)	(18.37)	(15.71)	(20.87)	(15.90)
6. # of children 19–24 attending school					94.02	33.27
					(119.82)	(109.20)
7. # of children ever born					−7.54	−9.46
					(16.44)	(17.50)
8. Employment income of husband + asset income of family net of income taxes at 0 hours of work of wife	−.052**	−.045**	−.067**	−.088**	−.009	−.032**
	(.019)	(.017)	(.013)	(.018)	(.011)	(.014)
9. Variable 8 divided by # of persons in family	.090*	.113**	.154**	.202**	−.032	.104**
	(.050)	(.048)	(.040)	(.056)	(.031)	(.047)
10. Dummy variable = 1 if language of home is French; 0 otherwise		81.68**		−28.49		55.81
		(36.38)		(52.78)		(70.74)
11. Selection bias (λ)	262.63**	288.18**	475.67**	464.86**	374.74**	289.06**
	(108.68)	(102.89)	(98.90)	(154.52)	(104.61)	(118.94)
Combined grouped R^2 = .969 for U.S. and .961 for Canada.[a]						
R^2	.0433	.0735	.0393	.0568	.0357	.0401
Standard error of regression	1,175	1,085	1,208	1,266	1,169	1,337

SOURCE: Calculated from the 1 per cent subsample from the 5 per cent primary State Public Use Sample of Basic Records from the 1970 U.S. Census; and from the 1 per cent Family File of the Public Use Sample from the 1971 Canadian Census.

[a] Explained in text.

[b] Numbers in parentheses are standard deviations.

[c] Coefficients with two asterisks are significant at a 95 per cent level. Coefficients with one asterisk are significant at an 80 per cent level.

Direct estimation of the labor supply function for working men has typically resulted in the finding of a negative relationship between the offered wage and hours of work. Moreover, Da Vanzo, De Tray, and Greenberg [7] find for White husbands aged 25–54 that, whereas the response of hours worked per week or per year to either an observed or imputed offered wage is always significantly negative when the sample is restricted to men who worked, this response is generally *positive* when men who did not work are included in the sample. Similar results are also reported by Garfinkel [**11**, pp. 215–217].

By first estimating a function for the probability of working, and then estimating a conditional hours of work equation for working wives, the labor supply response differences noted by Ben-Porath are at least partially accounted for in this study.

The second major difference between our study and previous cross-sectional studies lies in the choice of variables used to control for child status. There are at least four different dimensions of the child status of married women which may be of importance in examining their labor force behavior. These are (i) the presence

TABLE IX

CONSTRAINED GLS ESTIMATES FOR ANNUAL HOURS OF WORK EQUATION FOR MARRIED WOMEN IN U.S. AND CANADA[b,c]

40–44		45–49		50–54		55–59	
U.S.	Can.	U.S.	Can.	U.S.	Can.	U.S.	Can.
2365.15**	1917.43**	2169.91**	1858.79**	2218.50**	1365.47**	1396.02**	2114.85**
(142.36)	(129.70)	(128.72)	(105.27)	(188.66)	(79.42)	(314.85)	(279.30)
−71.60	−105.69*	−25.73	−112.58	−8.60	188.33*	330.99*	−65.81
(102.79)	(66.01)	(105.57)	(104.52)	(113.97)	(103.93)	(170.14)	(130.73)
−63.72	−271.83**						
(85.39)	(57.06)						
−46.39*	−69.66*	−54.78**	−41.80*	−35.07	−28.96	−37.12	−23.14
(27.22)	(24.56)	(25.23)	(25.17)	(40.39)	(31.84)	(84.60)	(82.97)
−8.25	92.72**						
(29.05)	(19.08)						
−45.13	24.30	10.73	−4.17	−99.80*	36.34	−35.42	−8.56
(52.38)	(43.28)	(45.00)	(37.61)	(58.88)	(45.49)	(95.05)	(71.02)
11.43	−8.12	1.77	3.84	3.91	−16.08	−2.67	−12.14
(11.78)	(13.66)	(9.81)	(12.65)	(9.19)	(13.75)	(14.53)	(14.34)
−.022**	−.047**	−.038**	−.040**	−.019**	−.056**	−.011	−.017
(.009)	(.010)	(.007)	(.010)	(.009)	(.012)	(.017)	(.019)
.056**	.077**	.082**	.070**	.042**	.119**	−.008	.025
(.021)	(.038)	(.017)	(.029)	(.019)	(.031)	(.030)	(.043)
	10.21		35.35		134.18**		299.20**
	(56.78)		(57.03)		(62.21)		(91.73)
241.93**	544.31**	489.86**	427.01**	380.06**	581.63**	809.77**	186.93*
(114.57)	(94.06)	(77.78)	(102.70)	(117.51)	(345.04)	(167.74)	(115.60)
.0151	.0611	.0430	.0381	.0340	.0338	.0670	.0427
1,125	1,328	1,061	1,121	1,058	817	1,128	1,464

of children in different age groups, (ii) the numbers of children in different age groups, (iii) interaction effects resulting from the presence of children in two or more different age groups, and (iv) the total number of children cared for. Full incorporation of all these aspects of child status into a study focused on labor force behavior might well result in the introduction of an unacceptably large number of closely related variables. Thus in our own study we have included variables for the number of children younger than 6, the number of children 6–14 years of age, the product of the numbers of children in these two age groups, and the number of children ever born. Some further control over the age distributions of children is also gained by carrying out our analysis separately for wives in five-year age groupings.

Other researchers have generally taken a more parsimonious approach than we have to the incorporation of child status information. For instance, Hall [**14**] includes separate dummy variables for the presence of children younger than 7 only, the presence of the children 7–13 years of age only, and the presence of children in both age groups. His omitted category is no children younger than 14. Heckman [**16, 17, 19**] includes a linear variable for the number of children younger than 6. And Harvey Rosen [**31**] includes separate dummy variables for one child younger than 6, two children younger than 6, and three or more children younger than 6. Rosen's omitted category is no children younger than 6. While limiting the amount of information about child status incorporated into their studies, most researchers have also included in their data samples older wives who would be expected both to exhibit a wide range of different child status configurations and to have relatively few children younger than 6. For instance, Hall's study includes wives 20–59 years of age, while the studies by Heckman and Harvey Rosen examine the labor force behavior of married women 30–44 years of age.

In Table X the working wives in our samples for the U.S. and Canada have been cross classified by their estimated net offered wage rate and number of children younger than 6. For each cell we show the mean number of children ever born and the number of observations. From this table it can be seen that after controlling for the number of children younger than 6, there is still a strong predominantly negative correlation between the number of children ever born and the estimated offered wage rate. Similar results were also obtained by controlling for child status as Hall [**14**] did. Nor is it intuitively surprising that wives with fewer children ever born also tend to work more hours than wives with more children ever born, other factors being held equal.

Turning our attention now to Heckman's most recent results [**19**], we find that the sign of the response of hours of work to changes in the offered wage rate is inferred from the sign of the coefficient of an "experience" variable in his hours of work equation. This experience variable could just as well be called an indicator of hours of work in previous years, however, since it is defined as the number of years since leaving school that a woman has worked six months or longer.[24] When entered into an hours of work equation one would expect this variable to capture not only some of the effects on hours of work of experience related differences in the offered wage rate, but also unmeasured tastes and preferences for work reflected in the work histories of each wife. As such, it would be surprising if the coefficient of this variable in Heckman's hours of work equation were not positive. Nor can we agree with Heckman's [**16**, p. 681] assertion that, "Historical time-series ... suggest that there is a monotonic *positive* relationship between wage rates and labor supply for married women ... so that excluding the 'backward bending' case is not objectionable ...." Looking at historical data for the U.S. we find that, while both real wages for women and the female labor force participation rate have clearly risen over time, the percentage of all female workers working 40 weeks per year or more fell from 69.4 per cent in 1950 to 67.2 per cent in 1960 and then rose slightly to 70.1 per cent in 1970, and the percentage

[24] The procedure followed in Heckman's earlier studies [**16**, **17**] is similar in this respect with the exception that the experience variable is defined as the number of years the woman worked full time since marriage. Also the results presented in [**16**] are erroneous because the data on this experience variable was incorrectly coded by the primary data source. One further feature which inhibits comparisons between [**16**, **17**] and [**19**] is that the annual hours of work in the first two studies were computed as the product of weeks worked in 1966 and "usual hours worked," while in [**19**] the annual hours of work were computed by dividing annual earnings in 1966 by a questionnaire wage for early 1967. While Heckman [**19**] argues that this latter measure of the annual hours of work is preferable, most other data sources do not contain the wage information necessary to calculate this measure.

TABLE X

MEAN NUMBERS OF CHILDREN EVER BORN FOR WORKING WIVES CLASSIFIED BY THEIR ESTIMATED NET OFFERED WAGE RATES AND NUMBER OF CHILDREN YOUNGER THAN 6

Estimated Net Offered Wage[a]	U.S. 0	U.S. 1	U.S. 2	U.S. 3+	Can. 0	Can. 1	Can. 2	Can. 3+
25–29								
<$1.75	1.48[b]	2.23	2.71	4.32	1.18	1.77	2.62	4.12
	(227)[c]	(253)	(101)	(19)	(214)	(237)	(115)	(17)
≥$1.75	.39	1.55	2.45	4.00	.29	1.32	2.16	3.33
	(461)	(276)	(109)	(26)	(978)	(727)	(327)	(36)
30–34								
<$1.75	2.28	3.14	3.87	6.14	2.07	2.68	3.70	4.00
	(506)	(209)	(45)	(7)	(591)	(285)	(64)	(9)
≥$1.75	1.69	2.67	3.17	4.00	1.38	2.08	2.76	3.80
	(371)	(171)	(80)	(10)	(426)	(368)	(136)	(20)
35–39								
<$1.75	2.59	4.17	6.67		2.57	3.78	4.67	6.00
	(651)	(75)	(9)		(495)	(165)	(27)	(6)
≥$1.75	2.48	3.66	3.63	5.33	2.51	3.01	3.35	3.13
	(645)	(154)	(40)	(6)	(778)	(216)	(60)	(8)
40–44								
<$1.75	2.68	4.08	8.64	10.25	2.79	4.16	4.78	7.00
	(1111)	(90)	(11)	(4)	(698)	(130)	(18)	(4)
≥$1.75	2.77	3.71	4.92	6.00	2.72	3.88	4.33	2.33
	(617)	(68)	(13)	(1)	(950)	(84)	(9)	(3)
45–49								
<$1.75	2.45	5.33			2.62	4.83	7.00	5.00
	(948)	(9)			(751)	(47)	(1)	(3)
≥$1.75	2.75	4.38	6.00		2.71	4.81	2.00	
	(970)	(40)	(1)		(1081)	(16)	(1)	
50–54								
<$1.75	2.46	2.33			2.68	3.22	7.00	
	(775)	(3)			(658)	(9)	(2)	
≥$1.75	2.43	3.00			2.65	4.43		
	(802)	(3)			(676)	(7)		
55–59								
<$1.75	2.26	1.50			2.98			
	(556)	(2)			(448)			
≥$1.75	2.16	1.25			2.29			
	(605)	(4)			(517)			

SOURCE: Calculated from the 1 per cent subsample from the 5 per cent primary State Public Use Sample of Basic Records from the 1970 U.S. Census; and from the 1 per cent Family File of the Public Use Sample from the 1971 Canadian Census.

[a] Calculated using coefficient estimates shown in Table XII.

[b] Mean number of children ever born.

[c] Number of observations.

of all female workers working 35 hours per week or more fell from 80.8 per cent in 1950 to 72.3 per cent in 1960 to 69.3 per cent in 1970.[25] Likewise for Canada we find that both real wages for women and the labor force participation rate have risen over time, while the percentage of all female workers working 40 weeks per year or more has fallen from 74.4 per cent in 1961 to 67.0 per cent in 1971 and the percentage of all female workers working 35 hours per week or more has fallen from 90.0 per cent in 1951 to 81.7 per cent in 1961 to 71.3 per cent in 1971.[26]

In Table XI we show the uncompensated wage and net wage elasticities of hours of work, evaluated at the mean hours of work. These wage elasticities all lie between $-.390$ and .204 when the computations are based on our constrained annual hours of work equation and between $-.409$ and .446 when the computations are based on our unconstrained hours equation, with positive values occurring only for older wives. All these uncompensated elasticities are well below the range of positive uncompensated elasticities reported by other researchers for married women. For instance, Harvey Rosen [**31**] reports an uncompensated wage elasticity of 2.3, while Heckman [**19**] considers his best estimate of this elasticity to be 4.5. On the other hand, our uncompensated wage elasticities seem broadly consistent with Ashenfelter and Heckman's [**3**] estimate of $-.15$ from individual data, Sherwin Rosen's [**32**] estimates of $-.30$ to $-.07$ from interindustrial data, Finegan's [**10**] estimates of $-.35$ to $-.25$ from interoccupational data, and Owen's [**30**] estimates of $-.24$ to $-.11$ from U.S. time-series data, where all of these estimates are for men.

In addition to these uncompensated elasticities, we also report income elasticities and compensated wage elasticities in Table XI. As can be seen from this

[25] The figures shown for the percentages of all female workers working 40 weeks per year or more are calculated from the 1950 United States Census of Population, Special Report P-E No. 1B, Table 17; the 1960 United States Census of Population, Special Report PC(2)7A, Table 17; and the 1970 U.S. Census of Population, Special Report PC(2)7A, Table 14. The data for all three years is for wage and salary earners reporting the number of weeks worked in the previous year.

The figures shown for the percentages of all female workers working 35 hours per week or more are calculated from the 1950 United States Census of Population, Special Report P-E No. 1B, Table 15; the 1960 United States Census of Population, Special Report PC(2)7A, Table 13; and the 1970 United States Census of Population, Special Report PC(2)7B, Table 48. The data for 1950 and for 1970 are for wage and salary earners, and the data for 1960 is for employed persons, all reporting the number of hours worked at all jobs in the week prior to enumeration.

[26] The figures shown for the percentages of all female workers working 40 weeks per year or more are calculated from the 1961 Census of Canada, Volume II—Part 3, Table 22; and from the 1971 Census of Canada, Public Use Sample—Individual File. The 1951 data on weeks worked are not comparable to the data for 1961 and 1971 since, in that year, part-time employment was converted to a full-time weekly basis. The data for both 1961 and 1971 are for wage and salary earners reporting the number of weeks worked in the previous year.

The figures shown for the percentages of all female workers working 35 hours per week or more are calculated from the 1951 Census of Canada, Volume V, Table 8; the 1961 Census of Canada, Volume III—Part 3, Table 21; and the 1971 Census of Canada, Volume III—Part 7, Table 32. The 1951 data are based on wage and salary earners 14 years of age and older who reported the number of hours worked in the week prior to enumeration. The data for 1961 and 1971 are based on wage and salary earners 15 years of age and older who reported the usual number of hours worked per week for the job held in the week prior to enumeration, or otherwise for the job of longest duration held since January 1 of the previous calendar year.

TABLE XI

ESTIMATED WAGE AND INCOME ELASTICITIES

Age Group	Uncompensated Wage Elasticities[a] U.S.	Can.	Income Elasticities[b] U.S.	Can.	Compensated Wage Elasticities[a] U.S.	Can.
25–29	−.390[c]	−.370	−.253	−.220	−.137	−.150
	(−.403)[d]	(−.409)	(−.219)	(−.225)	(−.184)	(−.184)
30–34	−.244	−.270	−.358	−.495	.114	.225
	(−.278)	(−.228)	(−.346)	(−.478)	(.068)	(.250)
35–39	−.165	−.305	−.049	−.188	−.116	−.117
	(−.141)	(−.344)	(−.071)	(−.140)	(−.070)	(−.204)
40–44	−.047	−.086	−.117	−.269	.138	.183
	(−.022)	(.078)	(−.130)	(−.273)	(.108)	(.351)
45–49	−.016	−.085	−.185	−.207	.169	.100
	(−.019)	(−.082)	(−.101)	(−.282)	(.082)	(.200)
50–54	−.055	.143	−.086	−.271	.081	.414
	(.122)	(.240)	(−.134)	(−.285)	(.256)	(.525)
55–59	.204	−.051	−.045	−.076	.249	.025
	(.446)	(.088)	(−.208)	(−.109)	(.654)	(.197)

SOURCE: Calculated from the 1 per cent subsample from the 5 per cent primary State Public Use Sample of Basic Records from the 1970 U.S. Census; from the 1 per cent Family File of the Public Use Sample from the 1971 Canadian Census; and from the coefficient estimates of the offered and net offered wage variables, and our combined variable for the employment income of the husband plus the asset income of the family net of income taxes at 0 hours of work of the wife, for our unconstrained and constrained annual hours of work equations.

[a] Evaluated at mean of wage variable and mean hours of work for each age group.

[b] Evaluated at mean of our combined income variable and mean hours of work for each age group.

[c] The top figures are calculated using the coefficient estimates shown in Table IX for our constrained annual hours of work equation.

[d] The figures in parentheses are calculated using the coefficient estimates for our unconstrained annual hours of work equation.

Table, the compensated wage elasticities based on both our constrained and unconstrained hours equations and for both the U.S. and Canada are slightly negative for the age groups 25–29 and 35–39, while the remaining compensated wage elasticities are all positive in conformity with accepted economic theory.

Combined grouped R^2s were calculated for the constrained hours equation using the same groups on which the combined groups R^2s for our probit and offered wage equations are based. These R^2s indicate that, for these groups, the estimated relationships presented in Table IX explain approximately 97 per cent for the U.S. and 96 per cent for Canada of the variation in the mean annual number of hours worked. Using the estimated values of both the offered wage rate and the annual hours of work, we also calculated combined grouped R^2s of .669 and .683 for the mean estimated annual incomes of wives in the U.S. and Canada, respectively.

7. CONCLUSION

We have analyzed the labor force behavior of U.S. and Canadian wives in three stages: the probability of working, the offered wage rate, and the annual hours of work. We find that the offered wage rates of wives are positively related to an index of local job opportunities for women. We also find that a wife is more likely to work the higher her potential offered wage rate. However, in contrast to previous research of others on the labor supply of U.S. wives, we find that the hours of work and the offered wage rate are negatively related for working wives. The resulting uncompensated wage elasticities evaluated at the mean hours of work are found to be similar to those reported by other researchers for working men.

Income taxes have been built into our model to account for differences in income tax laws between the U.S. and Canada. An iterative procedure has been used to estimate our hours of work equation in order to overcome the statistical problems resulting from the dependence of the hours worked on the tax rate, and the nonlinear dependence of the tax rate on the hours of work and the offered wage rate. We first treat the retention rate, which is 1 minus the tax rate, as a separate variable in our hours of work regressions. These results allow us to conclude, in agreement with the findings of Harvey Rosen [**31**] and Hausman and Wise [**15**], that wives do fully account for the impact of income taxes on their earnings in choosing their hours of work. Based on this finding, we then estimate a constrained hours of work equation in which the offered wage and retention rate variables are replaced by a single variable for the net offered wage rate. The resulting estimated relationships could be used to predict the impact of various proposed changes in the income tax laws on the labor supply of wives.

It should be noted that, despite the agreement of our results with the tax perception results of Rosen and of Hausman and Wise, our study and their studies predict opposite effects on the hours of work of working wives of a decrease in their income tax rates. Their studies find that wives would work more hours if the income tax burden on their earnings were less. We find that working wives, like working husbands, would tend to spend more time at home with their families and in other nonmarket work activities if their income tax rates were lower. Since the procedure developed in this study allows the direct incorporation of information on tax rates into the analysis of hours of work, the procedure might also be useful in the examination of data from negative income tax experiments. Although the tax rate schedule is exogenously manipulated in these experiments, researchers have generally found it difficult to incorporate this information into their studies [**15**].

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APPENDIX

REDUCED FORM, AND THE CONVERGENCE AND CONSISTENCY PROOF OF THE ITERATIVE ESTIMATION METHOD FOR THE HOURS OF WORK EQUATION

1. *Reduced form equation for h_i*

The offered wage equation (12) is already in reduced form. The reduced form for the hours equation is derived as follows. If we write the tax retention function defined by (14) as $RET_i = 1 - TX(h_i, u_i)$

noting its dependence on u_i, the disturbance term of the wage equation, then (15) becomes

(A1) $$h_i = F(\Gamma, h_i, u_i) + v_i^*$$

where Γ corresponds to the collection of coefficients $\{\gamma_0, \gamma_1, (\gamma_1\xi), \gamma_2, \gamma_3\}$ and

(A2) $$F(\Gamma, h_i, u_i) = \gamma_0 + \gamma_1 \ln w_i + (\gamma_1\xi) \ln \{1 - TX(h_i, u_i)\} + Z_i^*\gamma_2 + \gamma_3(E_H T)_i.$$

We first assume that (i) $TX(h_i, u_i)$ is differentiable in h_i over the region of our interest. Then by the implicit function theorem there exist a differentiable function $\phi_i(\Gamma, u_i, v_i^*)$ and a neighborhood Ω_i such that for $(\Gamma, h_i, u_i, v_i^*) \in \Omega_i$

(A3) $$\phi_i - F(\Gamma, \phi_i, u_i) - v_i^* = 0$$

and

(A4) $$h_i = \phi_i(\Gamma, u_i, v_i^*).$$

In the following we will assume that (ii) $(\Gamma, h_i, 0, 0) \in \Omega_i$ for any h_i and Γ. Since ϕ_i is differentiable in Ω_i, applying the mean value theorem, we get

(A5) $$\phi_i(\Gamma, u_i, v_i^*) = \phi_i(\Gamma, 0, 0) + a_{1i}u_i + a_{2i}v_i^*$$

where

(A6) $$a_{1i} = \left.\frac{\partial \phi_i}{\partial u_i}\right|_{u_i = \tilde{u}_i, v_i^* = \tilde{v}_i^*}, \quad a_{2i} = \left.\frac{\partial \phi_i}{\partial v_i^*}\right|_{u_i = \hat{u}_i, v_i^* = \hat{v}_i^*}$$

for some $(\Gamma, h_i, \tilde{u}_i, \tilde{v}_i^*) \in \Omega_i$ and some $(\Gamma, h_i, \hat{u}_i, \hat{v}_i^*) \in \Omega_i$. By assuming that (iii) a_{1i} and a_{2i} are independent of i over the product region $\Pi_i \Omega_i$, i.e., $a_1 \equiv a_{1i}$ and $a_2 \equiv a_{2i}$, we have from (A5)

(A7) $$h_i = h_i^* + u_i^*$$

where

(A8) $$u_i^* = a_1 u_i + a_2 v_i^*$$

and

(A9) $$h_i^* = \phi_i(\Gamma, 0, 0).$$

Clearly h_i^* thus defined does not depend on h_i; h_i^* corresponds to h_i defined by (15) in which both u_i and v_i^* are set equal to zero. Thus, by setting u_i and v_i^* equal to zero in both sides of (A1) and using (12), we get

(A10) $$h_i^* = \gamma_0 + \gamma_1\{\alpha_0 + Z_i\alpha_1 + R_i\alpha_2\} + (\gamma_1\xi) \ln \{1 - TX(h_i^*, 0)\} + Z_i^*\gamma_2 + \gamma_3(E_H T)_i,$$

where, by (14), and (12),

(A11) $$TX(h_i^*, 0) = TX((E_H)_i + e^{\{\alpha_0 + Z_i\alpha_1 + R_i\alpha_2\}}(h_i^*), A_i).$$

Throughout this paper we will assume that (iv) only Γ and h_i^* satisfy (A3) when $u_i = v_i^* = 0$, i.e., Γ and h_i^* uniquely satisfy $h_i^* = F(\Gamma, h_i^*, 0)$ or (A10) in Ω_i.

2. *Convergence (as $k \to \infty$)*

The hours of work equation (23) can be written as

(A12) $$h_i = F(\Gamma, h_i^*) + v_i^*$$

where Γ is redefined to represent the collection of parameters $(\gamma_0, \gamma_1, (\gamma_1\xi), \gamma_2, \gamma_3, \sigma_2)$, and F is redefined to be

(A13) $$F(\Gamma, h_i^*) = \gamma_0 + \gamma_1 \overline{\ln w_i} + (\gamma_1\xi) \ln \{1 - TX((E_H)_i + \bar{w}_i h_i^*, A_i)\} + Z_i^*\gamma_2 + \gamma_3(E_H T)_i + \sigma_2\lambda_i,$$

where $\overline{\ln w_i}$ and $\bar{w}_i$ are redefined as follows:

(A14) $$\overline{\ln w_i} = \alpha_0 + Z_i\alpha_1 + R_i\alpha_2 + (\sigma_{12}/\sigma_2)\lambda_i$$

and

(A15) $$\bar{w}_i = e^{\{\alpha_0 + Z_i\alpha_1 + R_i\alpha_2 + (\sigma_{12}/\sigma_2)\lambda_i\}} = e^{\overline{\ln w_i}}.$$

We write the tax function as $TX(h) = TX((E_H) + \bar{w}h, A)$, and prove convergence for the case of a joint return since the proof for the case of a separate return is identical. We assume that (v) $TX(h)$ is monotonically increasing and concave over the region of our interest. Then, noting that $0 < TX(h) < 1$, the function $1-TX(h)$ is monotonically decreasing and convex in h. Then $\ln(1-TX(h))$ is monotonically decreasing and convex in h. This is because

$$\frac{d\ln(1-TX(h))}{dh}=\frac{d\ln(q)}{dq}\frac{dq}{dh}<0$$

since $d\ln(q)/dq>0$ and $dq/dh<0$ where $q=1-TX(h)$ and

$$\frac{d^2\ln(1-TX(h))}{dh^2}=\frac{d^2\ln(q)}{dq^2}\frac{dq}{dh}+\frac{d\ln(q)}{dq}\frac{d^2q}{dh^2}>0$$

since $d^2\ln(q)/dq^2<0$, $dq/dh<0$, $d\ln(q)/dq>0$, and $d^2q/dh^2>0$. We also assume (vi)

$$\sup_h|\gamma_1\xi|\left|\frac{d\ln(1-TX(h))}{dh}\right|<1$$

for any h satisfying $0\leq h<T$.

We first show that, for a fixed vector of Γ, $F(\Gamma, h^*)$ defined by (A13) is a contraction mapping (see, for instance, [**34**]), i.e., for any h_1^* and h_2^* satisfying $0\leq h_1^*, h_2^*<T$ there exists a K such that $0<K<1$ and

$$\rho(F(\Gamma,h_1^*),F(\Gamma,h_2^*))\leq K\rho(h_1^*,h_2^*)$$

where $\rho(a,b)$ defines the distance between a and b. This is seen as follows. We assume without loss of generality that $\rho(a,b)=|a-b|$ and $h_1^*>h_2^*$. Then

$$\begin{aligned}\rho(F(\Gamma,h_1^*),F(\Gamma,h_2^*))&=\rho\{(\gamma_1\xi)\ln(1-TX(h_1^*),(\gamma_1\xi)\ln(1-TX(h_2^*))\}\\&\leq|\gamma_1\xi|\rho\{\text{in}(1-TX(h_1^*),\ln(1-TX(h_2^*))\}\\&\leq|\gamma_1\xi|(\ln(1-TX(h_2^*)-\ln(1-TX(h_1^*)))\\&\leq|\gamma_1\xi|\left|\frac{d\ln(1-TX(h_2^*))}{dh}\right||h_1^*-h_2^*|\\&\leq K|h_1^*-h_2^*|\end{aligned}$$

where

$$K=\sup_{0\leq h^*<T}|\gamma_1\xi|\left|\frac{d\ln(1-TX(h^*))}{dh^*}\right|$$

and where the monotonicity of $\ln(1-TX(h))$ and the inequality for convex functions

$$\ln(1-TX(h_1^*))\geq\ln(1-TX(h_2^*))+\left(\frac{d\ln(1-TX(h_2^*))}{dh^*}\right)(h_1^*-h_2^*)$$

were used.

Given some $h_i^{(k)}$ such that $0\leq h_i^{(k)}<T$ we define $h_i^{(k+1)}$ by

(A16) $$h_i^{(k+1)}=F(\Gamma,h_i^{(k)}).$$

If we define $\delta_i^{(k)}$ by

$$\delta_i^{(k)}=h_i^{(k)}-h_i^{(k+1)}$$

then we have

(A17) $$h_i^{(k)}=F(\Gamma,h_i^{(k)})+\delta_i^{(k)}.$$

Since F is a contraction, we have $\lim_{k\to\infty}h_i^{(k)}=\tilde{h}_i$ for some $\tilde{h}_i$ and hence $\lim_{k\to\infty}\delta_i^{(k)}=0$. Thus $\tilde{h}_i$ satisfies

(A18) $$\tilde{h}_i=F(\Gamma,\tilde{h}_i).$$

(A18) is, however, exactly the same as equation (A10) satisfied by h_i^* (by (20) and (21)). Hence $\tilde{h}_i=h_i^*$, and[27]

(A19) $$\lim_{k\to\infty}h_i^{(k)}=h_i^*.$$

[27] Note that all previous assumptions (i)–(vi) still hold for $F(\Gamma, h_i^*)$ defined by (A13)–(A15).

3. *Consistency of* $\lim_{k\to\infty} \Gamma^{(k)}$

For each step k, we estimate by least squares $\Gamma^{(k)}$ from the equation (see equations (23) and (24) in the text)

(A20) $$\begin{aligned} h_i = {} & \gamma_0 + \gamma_1 \widehat{\ln w_i} + (\gamma_1 \xi) \ln\{1 - TX((E_H)_i + \hat{w}_i \hat{h}_i^{(k)}, A_i)\} + Z_i^* \gamma_2 + \gamma_3 (E_H T)_i + \sigma_2 \hat{\lambda}_i \\ & + \gamma_1 \{\overline{\ln w_i} - \widehat{\ln w_i}\} \\ & + (\gamma_1 \xi)[\ln\{1 - TX((E_H)_i + \bar{w}_i h_i^*, A_i)\} - \ln\{1 - TX((E_H)_i + \hat{w}_i \hat{h}_i^{(k)}, A_i)\}] \\ & + \sigma_2 \{\lambda_i - \hat{\lambda}_i\} \\ & + V_i^*, \end{aligned}$$

where $\overline{\ln w_i}$ and $\bar{w}_i$ are defined by (A14) and (A15), respectively,

(A21) $$\hat{\lambda} = f(\hat{\phi}_i)/F(\hat{\phi}_i),$$

and $\hat{\phi}_i$ is given by (see equation (17) for ϕ_i)

(A22) $$\hat{\phi}_i = \left(\widehat{\frac{\alpha_0 - \beta_0}{\sigma}}\right) + Z_i\left(\frac{\hat{\alpha}_1}{\sigma}\right) - Z_i^*\left(\frac{\hat{\beta}_1}{\sigma}\right) + R_i\left(\frac{\hat{\alpha}_2}{\sigma}\right) - \left(\widehat{\frac{\beta_2 - \eta}{\sigma}}\right)(E_H T)_i,$$

where $(\widehat{\alpha_0 - \beta_0/\sigma})$, $(\widehat{\alpha_1/\sigma})$, $(\widehat{\beta_1/\sigma})$, $(\widehat{\alpha_2/\sigma})$, and $(\widehat{\beta_2 - \eta/\sigma})$ are maximum likelihood (consistent) estimates of the respective probit coefficients, and

(A23) $$\hat{w}_i = e^{\widehat{\ln w_i}}$$

where $\widehat{\ln w_i}$ is the least squares predicted value of $\ln w_i$ of (22), i.e.,

(A24) $$\widehat{\ln w_i} = \hat{\alpha}_0 + A_i \hat{\alpha}_1 + R_i \hat{\alpha}_2 + (\widehat{\sigma_{12}/\sigma_2})\hat{\lambda}_i.$$

Let the number of observations be N. Then

(A25) $$\operatorname*{plim}_{N\to\infty} (\overline{\ln w_i} - \widehat{\ln w_i}) = 0,$$

(A26) $$\operatorname*{plim}_{N\to\infty} \hat{w}_i = e^{\operatorname{plim} \widehat{\ln w_i}} = e^{\alpha_0 + Z_i \alpha_1 + R_i \alpha_2 + (\sigma_{12}/\sigma_2)\lambda_i} \equiv \bar{w}_i,$$

and

(A27) $$\operatorname*{plim}_{N\to\infty} (\lambda_i - \hat{\lambda}_i) = \lambda_i - \operatorname*{plim}_{N\to\infty} \frac{f(\hat{\phi}_i)}{F(\hat{\phi}_i)} = \lambda - \frac{f(\phi_i)}{F(\phi_i)} = 0.$$

Note that $F_1(\Gamma^{(k)}, \hat{h}_i^{(k)})$ defined in Section 6 in fact corresponds to the following regressor terms in equation (A20):

(A28) $$\begin{aligned} F_1(\Gamma^{(k)}, \hat{h}_i^{(k)}) = {} & \gamma_0^{(k)} + \gamma_1^{(k)} \widehat{\ln w_i} + (\gamma_1 \xi)^{(k)} \ln\{1 - TX((E_H)_i + \hat{w}_i \hat{h}_i^{(k)}, A_i)\} \\ & + Z_i^* \gamma_2^{(k)} + \gamma_3^{(k)} (E_H T)_i + \sigma_2^{(k)} \hat{\lambda}_i \end{aligned}$$

and $h_i^{(k+1)}$ is defined by

(A29) $$\hat{h}_i^{(k+1)} = F_1(\Gamma^{(k)}, \hat{h}_i^{(k)})$$

where superscript k on each coefficient corresponds to the kth iterate estimate.

For a fixed positive integer K, consider

(A30) $$\hat{h}_{i,K}^{(k+1)} = F_1(\Gamma^{(K)}, \hat{h}_{i,K}^{(k)}),$$

and take the limit and probability limit of both sides of (A30) as follows:

(A31) $$\begin{aligned} \lim_{k\to\infty} \operatorname*{plim}_{N\to\infty} \hat{h}_{i,K}^{(k+1)} &= \lim_{k\to\infty} \operatorname*{plim}_{N\to\infty} F_1(\Gamma^{(k)}, \hat{h}_{i,K}^{(k)}) \\ &= \lim_{k\to\infty} F(\Gamma^{(k)}, \hat{h}_{i,K}^{(k)}), \end{aligned}$$

since by (A25)–(A28), $\operatorname{plim}_{N\to\infty} F_1(\tilde{\Gamma}, \tilde{h}) = F(\tilde{\Gamma}, \tilde{h})$ for any $\tilde{\Gamma}$ and $\tilde{h}$. Since F is a contraction in h, (A31) provides

(A32) $$\hat{h}_{i,K}^* = F(\Gamma^{(K)}, \hat{h}_{i,K}^*)$$

where

(A33) $$\hat{h}^*_{i,K} = \lim_{k\to\infty} \operatorname{plim}_{N\to\infty} \hat{h}^{(k+1)}_{i,K}.$$

Since (A32) holds for any positive integer K, and since the $\hat{h}^*_{i,K}$ form a bounded sequence in K, there is a convergent subsequence, i.e., for some $K = s_1, s_2, s_3, \ldots$, there exists[28]

(A34) $$h^*_{i,\infty} = \lim_{j\to\infty} \hat{h}^*_{i,s_j}$$

such that

(A35) $$h^*_{i,\infty} = F(\Gamma^{(\infty)}, h^*_{i,\infty}).$$

By Assumption (iv), (A35) cannot hold in the region of our interest except for Γ and h^*_i. Thus we have

(A36) $$h^*_{i,\infty} = h^*_i.$$

By renaming the subsequence $s_1, s_2, s_3, \ldots$, as $1, 2, 3, \ldots$, and by (A30)–(A36), we have from (A29)

(A37) $$\lim_{k\to\infty} \operatorname{plim}_{N\to\infty} \hat{h}^{(k)}_i = h^*_i = \lim_{k\to\infty} \operatorname{plim}_{N\to\infty} F_1(\Gamma^{(k)}, \hat{h}^{(k)}_i) = F(\Gamma, h^*_i).$$

Thus, $\Gamma^{(k)}$ converges in the limit as k goes to infinity to Γ in probability, hence $\lim_{k\to\infty} \Gamma^{(k)}$ is consistent. By taking the limit with respect to k and probability limit with respect to N of both sides of (A20) and by using (A25)–(A27) and (A37), we derive equation (23), the hours equation of our interest. Thus our iterative least squares method and obviously the weighted (GLS) least squares version converge to consistent estimates.

Although there can be regions of h over which some or all of Assumptions (i)–(vi) fail to hold, and hence this algorithm may fail to converge, our empirical results suggest that this algorithm converges in most cases. It converged for all seven age groups for U.S. married women in three iterations and for four out of seven age groups for Canadian married women in four iterations. For those Canadian age groups where convergence did not occur, all regression coefficients except those for $\ln w_i$ and $\ln RET_i$ have converged. Finally we note that for practical use we can define the convergence of the process by a variety of criteria. In this study we used the percentage change in each regression coefficient in Γ over two successive iterations as the criterion of convergence.

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[28] In practice convergence of a subsequence is often the most one can expect to get in nonlinear programming. Convergence theorems in Zangwill [38], for example, are all stated in terms of convergent subsequences.

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