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FEMALE LABOR FORCE PARTICIPATION IN URBAN JAPAN:

A TRICHOTOMOUS LOGIT MODEL

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FEMALE LABOR FORCE PARTICIPATION IN URBAN JAPAN: A TRICHOTOMOUS LOGIT MODEL

M. Anne Hill*

The steady increase in the labor force participation of married women in the U.S. over the last few decades has stimulated considerable interest in the economic analysis of a woman's decision to work. As this body of literature has grown, the economic models developed within it have been implemented to analyze the labor force behavior of women in other countries. However, there is no consistent international pattern in the behavioral trends of the female labor force. For Japan in particular, the overall female labor force participation rate declined slightly from 47.4 percent in 1948 to 45.8 percent in 1976 after peaking at 56.7 percent in 1955. If dissimilarities in the aggregate trends provide evidence of international differences in the underlying economic, social, and cultural framework within which individuals act, then the economic models developed in response to behavior of women in the U.S. may be in some sense "culture bound."

One aspect of the labor force in which there is a great deal of international variation is the composition of the labor force by employment status. In the U.S., the labor force participation decision involves two choices: working and not working. Almost all female workers (92.8 percent in 1975) work as paid employees. In Japan, as in many less-developed countries, women may be considered economically active (i.e., workers) even though they are not engaged in the formal or so-called

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"paid" sector; they may work in the "informal" sector of the labor market on a farm, in a family business, or at home -- producing goods for market sale. Most of the workers in this sector are either self-employed (which generally includes home-handicraft workers) or family workers.

Table 1 presents the distribution of the labor force by employment status for twelve countries. More than 90 percent of female workers in the U.S., Canada, and Sweden are paid employees. This share is lowest in Thailand and Korea, with 20 and 30 percent respectively. Japan ranks between these extremes with 59 percent of its female workers in paid employment. Given its advanced economic development, a surprisingly large fraction (39 percent) of Japan's female labor force comprises self-employed and family workers.

If individuals regard the decision to enter the labor force as a paid employee as distinct from the choice to enter the labor force as a family worker, then economic models of labor force participation which treat these choices as identical will incorporate an aggregation bias. Individual behavioral responses may differ between these two sectors for several reasons. First, the wage offers may differ by sector. Second, while entering the informal sector may be virtually frictionless, there may be fixed costs (among them commuting time and child care) associated with working in the paid sector. Third, family workers may face more flexible working schedules than paid employees; the latter group may be subject to contractual working hours.

In order to treat the labor force decisions in Japan, this paper generalizes the standard labor force participation model by expanding the set of labor force alternatives to include working in the informal market

Table 1: Distribution of the Female Labor Force By Employment Status in Selected Countries

(percent)[a]

Country	Paid Employees	Family Workers	Self-Employed
United States, 1976	92.8	1.6	4.3
Sweden, 1976	92.8	2.4	2.8
Canada, 1977	91.1	2.5	5.5
Federal Republic of Germany, 1975	84.0	10.2	5.0
Hungary, 1976	77.7	5.8	4.4
Israel, 1976	76.9	6.1	12.2
Italy, 1976	67.9	13.5	13.3
Venezuela, 1975	67.1	5.6	19.7
Mexico, 1975	66.6	7.1	26.2
Japan, 1975	59.2	26.9	11.7
Korea, 1976	30.6	43.7	23.7
Thailand, 1976	20.1	50.8	28.4

Source: ILO Labor Yearbook, 1978

[a]Totals may not sum to 100. The table excludes "others" and "unknown."

as a choice distinct from working in the formal market. The result is a labor force participation model in which the dependent variable is trichotomous (working as a paid employee, working as a family worker, and not working) rather than dichotomous (working and not working). If women in fact do not differentiate between working in these sectors, the responses to explanatory variables will be identical and the trichotomous model will collapse to the more standard dichotomous model.

Section I of this paper briefly discusses the labor supply literature for the U.S. and Japan. Section II outlines the theoretical model, illuminates the relationship between the trichotomous and the dichotomous specifications of the model, and relates a test to ascertain whether the trichotomous model collapses to the simpler model. Section III describes the data set and the explanatory variables. Section IV presents the empirical results of both the trichotomous model and the dichotomous model. The final section offers a comparison of the Japanese results with the results of two dichotomous models estimated with U.S. data.

I. Economic Models of Labor Supply in the U.S. and Japan

Pioneering work by Mincer and Cain has served as a theoretical and empirical foundation for numerous studies of labor force behavior. These models have treated a woman's current labor force status (measured as dummy variable for the individual and as a rate for the population) alternately (1) as a measure of permanent labor supply which is an interior solution to maximizing lifetime utility and (2) as the result of a discrete choice integral to a larger model of labor supply. The discrete choice models generally base the individual choice of labor force status on a comparison of the wage offered in the market place and the marginal value of time at full leisure -- termed the reservation wage.

Two exceptions are Heckman (1978) and Hausman who formulate models in which the participation decision relies on a comparison of the maximum utility attainable given each choice. In the empirical application of these studies, a woman's labor force status is re-
lated systematically to economic and demographic characteristics such as female earnings, male earnings, non-earnings income, schooling, work experience, age, number of children, and health status, among others. (See also Bowen and Finegan, Cogan (1975, 1977), Gronau, Heckman (1974, 1977), and Schultz.) While the theoretical models and statistical techniques implemented in the empirical analysis of labor force decisions have been refined considerably (especially by Lewis, Ben-Porath, and Heckman),the estimated coefficients have remained fairly robust. Also, there is substantial agreement among micro-level and aggregate results. (See Cain and Bowen and Finegan for comparisons.) These results indicate a strong positive relationship between the probability of entering the labor force and a woman's wage and her level of schooling, and a strong negative relationship between her husband's earnings and her propensity to work. Jones and Long estimate a trichotomous probit model in which the decision to work part time is treated as distinct from the decision to work full time. Although the estimated coefficients do differ by work status, their results are consistent with the body of estimates for the simpler model.

In contrast, the straightforward application of the standard labor force participation model to Japanese data has yielded somewhat anomolous results. Several studies use aggregate Japanese prefectural data to estimate models of labor force participation. In the first, Obuchi regresses the 1960 and 1965 female labor force participation rates on two variables: male earnings and the fraction of the labor force in agriculture. His analysis is performed for ten five-year age groups of women over 20. Male income is negative although significant for only five age groups. The fraction in agriculture has a positive effect and is significant for each group except the youngest (20 to 24). Umetani regresses the aggregate 1965 prefectual participation rate of married women on male income, female income, schooling, children, the fraction in agriculture, and a measure of labor market tightness. He finds negative and statistically significant effects of both the own wage and schooling. The fraction of employment in agriculture is significant, positive, and alone explains 80 percent of the total variation in the dependent variable. Male income and the number of children have negative and significant effects on labor force participation. Umetani reduces his regression sample by excluding those prefectures in which more than 30 percent of the labor force is employed in agriculture. Again, he finds negative, significant effects of both schooling and the own wage. In the third study, Hamilton uses similar aggregate data for 1960 to estimate a simultaneous equations model for children ever born, female wages, and labor force participation. Hamilton's model yields a positive, though insignificant, effect of female schooling and a negative effect of female wages on the propensity to work.

Unlike in the U.S., few researchers can access survey data for Japanese househol One such researcher, Obi, uses household data from Japan's Survey of Family Income and Expenditure for 1961-1964 to estimate a model of female force participation. Although he has no information on the wife's wage or her employment status, he regresses a zero-one participation variable on the husband's income and the number of children under six. As expected, both coefficients are negative and statistically significant.

The apparent incongruities in the estimates of the prefectural models may result from two aggregation problems. First, while SMSA's in the U.S. may be considered fairly homogeneous, prefectures in Japan may not. These areas range from being predominantly rural and agricultural to highly urbanized and industrial. Hence the characteristics of the underlying data base may work against a reasonable comparison of the Japanese results with those for the U.S. Second, the dependent variable in the Japanese models is the sum of the labor force participation rates in family work and in paid employment. (Hamilton does exclude agricultural workers.) If, as suspected, the measured prefectural wage better represents the wage in paid employment than that in family work, the negative wage effect is not surprising. An increase in this wage will raise participation in the formal sector, reduce participation in the informal sector, with the net effect being ambiguous. Using household level data in conjunction with treating the decision to participate in paid employment and the decision to participate in family work separately resolves these ambiguities.

II. The Labor Force Participation Model

This model assumes that each individual may select among three alternatives: working in the formal sector as a paid employee (indexed p), working in the informal sect as a family worker (indexed f), and not working (indexed n).[1] The individual compares the maximum utility attainable <u>given</u> each participation alternative and selects that alternative which yields the maximum maximorum.

Preferences are assumed to be described by a well behaved utility function (U) that is maximized subject to time and wealth constraints with no uncertainty:

(1) $U = U(L_w, C, X)$,

where L_w is the household time of the wife, C is the number of children, X is the Hicksian composite commodity which is taken as the numeraire. The time and wealth constraints are:

(2) $W_w H_w + I_h + Y = p_x X + p_c C$

(3) $H_w + L_w = T$

where W_w is the market wage, H_w is hours worked, I_h is the husband's income, Y is non-wage income, P_x is the price of X, P_c is the fixed price of children, and T is the time endowment.

The budget constraints will differ by employment status since the wage offers differ by sector. Also, there may be greater time and money costs incurred on entrance into the formal sector than the informal sector. The maximum utility attainable <u>given</u> each alternative will be a function of the

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[1] Initially, self-employment and home production for market sale were treated as two additional, distinct alternatives yielding five choices altogether. Excluding these categories resulted in no loss of statistical precision and a significant reduction in computational expense .

wage offer, the husband's income, non-wage income, and the price of children:

$$(4)\quad V_p = V(W_{wp}, P_c, I_h, Y)$$

$$V_f = V(W_{wf}, P_c, I_h, Y)$$

$$V_n = V(P_c, I_h, Y)$$

where V ($\cdot$) is the indirect utility or, in other words, the direct utility function U ($\cdot$) evaluated at the optimal demands for L_w, C, and X. The individual then compares V_f, V_p, and V_n and selects that participation status for which V ($\cdot$) is the maximum.[2] The labor force decision then depends upon the comparative value of these indirect utilities. Variables which raise the utility of one employment status relative to the others will raise the selection probability of that employment status.

Let V_{ji} be the maximum utility attainable for individual i if she chooses participation status j = p, f, n, and suppose that this indirect utility function can be decomposed into a non-stochastic components (S) and a stochastic component (ε):

$$(5)\quad V_{ji} = S_{ji} + \varepsilon_{ji}\ ,\ j = p, f, n, \text{ where}$$

S_{ji} is a function of observed variables and ε_{ji} is a function of unobserved variables. The probability that the ith woman selects the jth participation status is then given by:

$$(6)\quad P_{ji} = \Pr\ [V_{ji} > V_{ki}],\ j \neq k,\ j = p, f, n.$$

or, substituting in from (5),

$$(7)\quad P_{ji} = \Pr\ [S_{ji} - S_{ki} > \varepsilon_{ki} - \varepsilon_{ji}]\ .$$

If the stochastic components have independent Weibull

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[2]This utility comparison decision rule leads to the same outcome as the more standard reservation wage offered wage comparison.

distributions, then the difference between the errors $(\varepsilon_{ki} - \varepsilon_{ji})$ has a logistic distribution and the choice model is multinomial logit (McFadden (1974)).[3] Then (2) may be expressed:

$$(8)\quad P_{ji} = \frac{e^{S_{ji}}}{e^{S_{pi}} + e^{S_{fi}} + e^{S_{ni}}} .$$

In order to estimate this model, we must first specify a functional form for the non-stochastic component of the indirect utility function S_{ji}. This component is approximated in linear form $(S_{ji} = \beta_j{}'X_i)$, yielding

$$(9)\quad P_{ji} = \frac{e^{\beta_j X_i}}{e^{\beta_p{}'X_i} + e^{\beta_f{}'X_i} + e^{\beta_n{}'X_i}} \quad , \text{ where}$$

X_i is a vector of independent variables explaining labor force participation and β_j is the parameter vector.

The dichotomous model is of the form:

$$(10)\quad P_{wi} = \frac{e^{\beta_w{}' X_i}}{e^{\beta_w' X_i} + e^{\beta'_{nw} X_i}} ,$$

where w subscripts "working" and nw subscripts "not working."

It is of considerable interest to test whether the trichotomous model simply collapses to the dichotomous model. The simpler model effectively restricts the parameters for family workers (β_f) to equal those for paid employees (β_p). The dichotomous model then misspecifies the underlying choice framework unless these coeffecients are in fact equal. This can be seen most clearly in terms of the log-odds ratio.[4] In the logit model, the log-odds ratio of two probabilities is a linear-in-parameters function of the explanatory variables. Consider the log-odds

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[3]The Weibull distribution has a unimodal bell shape roughly similar to the normal distribution (See McFadden).

[4]I would like to thank T. Paul Schultz for passing along John Begg's suggestion of this exposition.

ratio of working and not working derived from the trichotomous model in (9):

$$(11)\quad \ln \frac{P_p + P_f}{P_n} = \ln \left(e^{\beta_p' X_i} + e^{\beta_f' X_i} \right) - \beta_n' X_i \quad .$$

Then if $\beta_p = \beta_f = \beta_w$, a simple dichotomy appropriately specifies the choice model:

$$(12)\quad \ln \frac{P_w}{P_n} = \ln 2 + \beta_w' X_i - \beta_n' X_i \quad ,$$

where $p_f + p_p = p_w$. The right-hand side of (12) is in linear form. If these vectors are not equal, the right-hand side of (11) is nonlinear and will be misspecified with a dichotomous dependent variable. The null hypothesis is then that $\beta_f = \beta_p$, which can be tested using a likelihood ratio test. The test statistic is:

$$(13)\quad \lambda = -2 \left[L(\hat{\beta}_r) - L(\hat{\beta}_u) \right] \quad ,$$

where $L(\hat{\beta}_r)$ is the log-likelihood function evaluated under the restrictions and $L(\hat{\beta}_u)$ is the unrestricted log-likelihood function. Under the null hypothesis, λ is distributed asymptotically as a chi-square variate with k degrees of freedom, where k is the number of restrictions.

III. Description of the Data

A 1975 survey of women between the ages of 20 and 59 living in the Tokyo Metropolitan Area provides the data base for estimating the labor force participation model. The National Institute for Vocational Research sponsored the survey and conducted the interviews, the sample was drawn from 177 urban area (shi) and 7 rural areas (gun) within 50 kilometers of the center of Tokyo. This area includes parts of the prefectures of Ibaraki, Saitama, Chiba, Tokyo, and Kanagawa, and all of the major cities of Tokyo and Yokohama.[5] The population in these areas represents roughly 23 percent of the relevant population for the entire country.[6] The surveyed population is primarily urban. Only 3.2 percent of the 1405 respondents live in rural areas as against 23.2 percent of the population of all Japan.[7] Therefore, the rual areas and consequently, the agricultural sector, are underrepresented in this survey. In 1975, 17.8 percent of all employed females worked in the agricultural sector as against 2.8 percent of the employed females in the surveyed area.[8]

The list of independent variables is restricted to a fairly standard one to enhance the comparability of these estimates with estimates for the U.S. These variables are: the wife's wage (W_w), the husband's income (I_h), a home ownership variable which serves as a proxy for non-wage income (Y), the wife's years of schooling, the husband's level of schooling, the number of children under six, and a dummy variable that equals one if the family owns a business. While most of these variables are conventional, several warrant additional discussion.

As in U.S. data, wages are not reported for women who are not working.

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[5]See Hill, Chapter IV, for a more complete description of the survey and relevant variables.

[6]The Population Census of Japan, 1975, Table 4.

[7]Ibid.

[8]op. cit., Table 7.

The wage in family work is predicted for all women in the sample based on the wages reported by women who are family workers. An anologous procedure is used to predict the wage in paid employment. These wages are specified to be a function of the woman's years of schooling, her potential labor market experience (age-years of schooling-6), and an urban residence variable. The wages are estimated with and without a correction for sample selection bias. The correction is based on work by Hay who adapted Heckman's (1978) Mill's ratio correction for logit models. The estimating procedure and wage estimates are detailed in the appendix. Also included in this appendix are predictions for the husband's income.

The number of children under six is included as a regressor for comparison with U.S. estimates. The appropriate variable is the price of children P_c which we unfortunately do not observe. The number of young children substitutes in attempt to capture exogenous variation in P_c. As C enters the utility function, its inclusion in the indirect utility function is inappropriate. (See especially Rosenzweig and Wolpin.) The models are estimated both with and without this variable.

One could argue that the husband's income, the ownership of a home, and the ownership of a business are each endogenous. To focus the wife's participation decision, we ignore these complications and treat these as exogenous variables.

For estimation, the sample is restricted to currently married women who are neither self-employed nor home handicraft workers. Women with missing values for any of the relevant variables are excluded. The resulting sample includes 1038 women. Table 2 presents the means and standard deviations of the independent variables. The overall participation rate is 28.6 percent, with 60.3 percent of these workers employed in the paid sector and 39.7 percent engaged as family workers.

The means are included for the entire sample and for the subsamples of paid employees and family workers. The mean years of schooling for the

Table 2: Explanatory Variables and Their Means

(standard deviations in parentheses)

Independent Variable	All Women	Employees	Family Workers
Years of Schooling	11.2302 (2.2367)	11.0559 (2.3385)	10.7719 (1.8829)
Experience (Age-schooling-6)	20.5212 (10.4202)	21.6257 (10.1067)	24.1780 (8.5920)
ℓn Husband's Income (Predicted, yen per year)	14.2688 (.2060)	14.2089 (.2079)	14.3148 (.1573)
Husband's schooling (Dummy variable = 1 if husband completed junior high school)	.7235 (.4475)	.6816 (.4672)	.5424 (.5003)
Family Business (Dummy variable = 1 if husband is self-employed or a shop owner)	.2293 (.4206)	.0894 (.2861)	.8475 (.3611)
Home Ownership (Dummy variable = 1 if the family owns their house or apartment)	.5992 (.4903)	.4413 (.4979)	.8305 (.3768)
Children Under Six	.6108 (.8019)	.2570 (.6189)	.4237 (.7090)
Urban Residence (Dummy Variable = 1 if residence is a city with a population greater than 100,000)	.7659 (.4236)	.7542 (.4318)	.7797 (.4162)
Number of Observations	1038	179	118

sample is roughly eleven years with little variation across employment status. Potential experience is highest for family workers with a mean of 24.2 years, implying that their mean age is also highest. The geometric mean of the husband's income is about 6,500 U.S. dollars, and also varies little by employment status. Three variables which do vary significantly by employment status are family business, home ownership, and young children. As expected, a large fraction of family workers (.85) have self-employed spouses. In marked contrast, the husbands of one-tenth as many paid employees are self-employed. Also, nearly twice as many family workers as paid employees own their own homes (.83 as against .44). Paid employees are characterized by relatively few young children. Each family has on average .61 children under six. Paid employees have .26 on average while family workers have slightly more with .42.

IV. Empirical Results

This section presents the multinomial logit estimates for both the trichotomous and the dichotomous forms of the labor force participation model. For n choices in this model, only n - 1
distinct parameter vectors may be identified. This linear dependence requires some normalization of the parameters. For both the trichotomous and the dichotomous models, the parameters are normalized:

$$(14) \qquad \sum_j \beta_j = 0.$$

The coefficients are reported for paid employment and family work. The coefficients for non-participation may be recovered $\beta_n = -(\beta_f + \beta_p)$. Similarly for the dichotomous model, $\beta_{nw} = -\beta_w$.

For comparison among these empirical results it is fruitful to calculate the elasticities of the dependent variables, the probability of entering the paid labor force (P_p) and the probability of engaging in family work (P_f), with respect to each independent variable. The partial derivatives are:

$$(15) \qquad \frac{\partial P_j}{\partial x} = P_j(1 - P_j)\frac{\partial S_j}{\partial x} - P_jP_k\frac{\partial S_k}{\partial x} - P_jP_n\frac{\partial S_n}{\partial x},$$

$$= P_j(1 - P_j)\beta_j - P_jP_k\beta_k + P_jP_n(\beta_j + \beta_k), \quad j,k = p, f, j \neq k,$$

taking derivatives of (9) and substituting in from (14).

For the dichotomous model, $\beta_p = \beta_f = \beta_w$ and $P_p + P_f = P_w$:

$$(16) \qquad \frac{\partial P_w}{\partial x} = 2\,(P_w\,(1 - P_w))\,\beta_w \quad .$$

Using these formulae for the derivatives, we may then evaluate the elasticities at the sample means. The standard errors of these elasticities may be calculated in a straghtforward manner using the variance-covariance matrix of the estimated parameters as the elasticities are simply linear combinations of the parameters.

McFadden suggests several measures of goodness-of-fit for the estimated logit model. Among them is a likelihood ratio statistic:

$$(17) \qquad -2\ [L(\hat{\beta}_o) - L(\hat{\beta})] \sim \chi^2_k \quad ,$$

which, under the null hypothesis that all parameters equal zero, is asymptotically distributed as a chi-square variate with k degrees of freedom, where k is the number of estimated parameters. $L(\cdot)$ is the log likelihood which is evaluated at $\hat{\beta}$, the maximum likelihood estimate of the parameter vector and $\hat{\beta}_o$, a vector of zeroes.
The likelihood ratio index (McFadden, p. 121) is analogous to a least squares multiple correlation coefficient:

$$(18) \qquad \rho^2 = 1 - \frac{L(\hat{\beta})}{L(\hat{\beta}_o)} \quad .$$

Each of the logit models is estimated in reduced form, then reestimated, including the predicted wages (both uncorrected and corrected for sample selection bias). Each model excludes the number of young children (specification (1)), then, for comparison with other estimates, includes this variable (specification (2)).

Table 3 presents the estimates of the trichotomous model in reduced form. Instead of the predicted wage, this model includes potential labor market experience and urban residence as regressors. Years of schooling is hypothesized to influence both the market wage and the value of time out of the market.

Table 3: Maximum Likelihood Logit Estimates of the Trichotomous Participation Model--Reduced Form

(1038 Observations)[a]

Independent Variable	(1) Employee (β_p)	(1) Family Worker (β_f)	(2) Employee (β_p)	(2) Family Worker (β_f)
Years of Schooling	-.0052	.0626	-.0355	.0696
	(.0386)	(.0502)	(.0397)	(.0507)
	[.3319]	[1.0933]	[-.1454]	[1.2153]
	(.4525)	(.7302)	(.4685)	(.7373)
Experience	.0184	-.0049	-.0045	-.0018
	(.0077)	(.0097)	(.0090)	(.0116)
	[.5217]	[.0436]	[-.1645]	[-.1091]
	(.1690)	(.2561)	(.4063)	(.1875)
ℓn Husband's Income	.8352	-2.842	.8559	-2.882
	(.6579)	(1.089)	(.6351)	(1.033)
	[-.4183]	[-4.0955]	[-.4104]	[-4.1483]
	(.5309)	(1.4368)	(.5342)	(1.3639)
Husband's Schooling	-.0866	.1572	-.0828	.1593
	(.2153)	(.2776)	(.2145)	(.2700)
	[-.0283]	[.1481]	[-.0232]	[.1520]
	(.1625)	(.2598)	(.1652)	(.2527)
Family Business	-1.748	3.010	-1.752	3.024
	(.2575)	(.3455)	(.2546)	(.3343)
	[-.2036]	[.8874]	[-.2031]	[.8920]
	(.0597)	(.1027)	(.0603)	(.0942)
Home Ownership	-.7785	.7633	-.7773	.7645
	(.1513)	(.2052)	(.1539)	(.2060)
	[-.4446]	[.4793]	[-.4430]	[.4808]
	(.0919)	(.1601)	(.0947)	(.1607)
Children Under Six	-	-	-.6671	.1633
	-	-	(.1313)	(.1418)
	-	-	[-.5682]	[_.0610]
	-	-	(.0893)	(.1083)
Urban Residence	.0140	-.1082	.0430	-.1182
	(.1594)	(.2016)	(.1631)	(.2025)
	[-.0332]	[-.1268]	[-.0036]	[-.1270]
	(.1290)	(.1995)	(.1337)	(.2002)
Intercept	-11.34	37.24	10.55	37.63
	(9.11)	(15.14)	(8.08)	(14.37)
ℓn likelihood		-667.6		-638.1
ρ^2		.4146		.4405
χ^2		945.6		1005

[a]Standard errors are in parentheses and elasticities are in brackets.

The magnitudes and, in some instances, the signs of the elasticities differ by participation status. Both years of schooling and experience increase the probability of working in each sector, although experience has a slightly larger effect than schooling on paid employment (with elasticities of .5 and .3 respectively), while schooling has much greater effect than experience for family work (the elasticities are 1.1 and .04 respectively). The husband's income has a strong negative influence on labor force participation although this effect is ten times as great for family workers as employees ; the probability of engaging in family work is very responsive to the husband's wage with an elasticity of -4.1. The level of the husband's schooling has opposite effects for each employment status. An increase in the husband's level of schooling actually increases the chance of working as a family worker while decreasing the chance of working in paid employment. (This variable incorporates both wealth and wage effects). As expected, owning a family business significantly reduces the probability of working in the paid sector while raising the probability of engaging in family work, with elasticities of -.20 and .89 respectively. Surprisingly, home ownership increases the propensity to work in the informal sector. Residence in an urban area decreases the probability of working in both sectors.

The second specification of the model is superior in terms of its explanatory power. Although excluding the number of young children increases the explanatory power of the equation (the ρ^2 rises from .41 to .44), it produces **sign changes in several coefficients.** As in the U.S., the presence of young children reduces the probability of working. Children have a stronger and statistically more significant impact on the propensity to work in the formal sector than in the informal sector. The inclusion of young children in the Japanese model has two effects which are unexpected given the results for the U.S. First, the effect of schooling on paid employment becomes negative, changing from .52 to -.16. Second, the effect of experience on working in either sector becomes negative. These two results are inconsistent with our expectations and provide additional evidence to support the contention that fertility should be excluded from labor force participation equations.

Tables 4 and 5 present estimates of the trichotomous model including instrumental variables predictions for the wages. Potential labor market experience and urban residence are excluded from these equations. The wages in Table 4 are not corrected for sample selection bias while those in Table 5 are. Overall, correcting for selectivity bias in the wage estimates results in surprisingly little change. When comparing the results in these two tables, one can readily see that this correction raises neither the explanatory power of the models nor the statistical precision with which the coefficients are estimated.

The estimated wage elasticities for the first specification are 2.63 and .70 (uncorrected wages) and 2.58 and .68 (wages corrected for selectivity bias) for paid employment and family work respectively. The remaining coefficients change only marginally in size. One exception is elasticity of paid employment with respect to schooling, which decreases from .33 to -1.7 when the predicted wage is included.

Table 4: Maximum Likelihood Logit Estimates of the Trichotomous Participation Model Including Predicted Wages

(1038 Observations)[a]

Independent Variables	(1) Employee (β_p)	(1) Family Worker (β_p)	(2) Employee (β_p)	(2) Family Worker (β_f)
Years of Schooling	-.1389	.1010	.0023	.0940
	(.0459)	(.0569)	(.0499)	(.0581)
	[-1.7238]	[.9704]	[.6734]	[1.7116]
	(.6423)	(.8836)	(.7223)	(.9109)
ln Husband's Income	.8393	-2.836	.8650	-2.885
	(.6006)	(1.089)	(.6276)	(1.033)
	[-.4084]	[-4.0837]	[-.3982]	[-4.148]
	(.5318)	(1.4348)	(.5253)	(1.3682)
Husband's Schooling	-.0915	.1626	-.0832	.1590
	(.2155)	(.2775)	(.2130)	(.2698)
	[-.0314]	[.1524]	[-.0237]	[.1515]
	(.1624)	(.2596)	(.1643)	(.2531)
Family Business	-1.741	3.002	-1.747	3.017
	(.2570)	(.3443)	(.2533)	(.3333)
	[-.2022]	[.8853]	[-.2023]	[.8901]
	(.0597)	(.1027)	(.0603)	(.0942)
Home Ownership	-.7683	.7452	-.7803	.7654
	(.1465)	(.1974)	(.1500)	(.1996)
	[-.4416]	[.4652]	[-.4455]	[.4803]
	(.0905)	(.1545)	(.0935)	(.1561)
Children Under Six	-	-	-.6698	.1683
	-	-	(.1231)	(.1209)
	-	-	[-.5689]	[-.0615]
	-	-	(.0879)	(.0918)
ln Wage	1.769	-.1621	-.5212	-.1220
	(.5419)	(.1885)	(.6227)	(.1932)
	[2.6296]	[.6985]	[-.8767]	[-.4775]
	(.8117)	(.3357)	(.9371)	(.3680)
Intercept	-19.56	37.48	-8.206	37.97
	(9.434)	(15.14)	(9.134)	(14.37)
ln Likelihood	-667.6		-638.1	
ρ^2	.4145		.4404	
χ^2	945.5		1004	

[a]Standard errors are in parentheses and elasticities are in brackets.

Table 5: Maximum Likelihood Logit Estimates of the Trichotomous Participation Model Including Wages Corrected for Selectivity Bias

(1038 Observations)[a]

Independent Variable	(1) Employee (β_p)	(1) Family Worker (β_f)	(2) Employee (β_p)	(2) Family Worker (β_f)
Years of Schooling	-.1377	.1009	.0059	.0934
	(.0457)	(.0570)	(.0511)	(.0583)
	[-1.7037]	[.9759]	[.7317]	[1.7143]
	(.6388)	(.8845)	(.7430)	(.9131)
ln Husband's Income	.8539	-2.836	.8571	-2.887
	(.6683)	(1.121)	(.6661)	(1.089)
	[-.3859]	[-4.0759]	[-.4116]	[-4.1557]
	(.5282)	(1.4801)	(.5419)	(1.4334)
Husband's Schooling	-.0912	.1615	-.0845	.1601
	(.2158)	(.2815)	(.2182)	(.2778)
	[-.0316]	[.1512]	[-.0247]	[.1522]
	(.1619)	(.2639)	(.1659)	(.2597)
Family Business	-1.722	3.002	-1.759	3.019
	(.2587)	(.3511)	(.2591)	(.3449)
	[-.1955]	[.8879]	[-.2063]	[.8893]
	(.0596)	(.1046)	(.0606)	(.1026)
Home Ownership	-.7559	.7468	-.7857	.7644
	(.1457)	(.1974)	(.1488)	(.1996)
	[-.4296]	[.4708]	[-.4508]	[.4780]
	(.0895)	(.1545)	(.0916)	(.1559)
Children Under Six	-	-	-.6829	.1682
	-	-	(.1274)	(.1209)
	-	-	[-.5813]	[-.0615]
	-	-	(.0930)	(.0924)
ln Wage	1.737	-.1604	-.5775	-.1198
	(.5333)	(.1882)	(.6431)	(.1937)
	[2.5813]	[.6839]	[-.9621]	[-.5044]
	(.7987)	(.3337)	(.9686)	(.3749)
Intercept	-19.60	37.48	-7.804	37.99
	(9.548)	(15.57)	(9.753)	(15.14)
ln Likelihood		-667.7		-638.1
ρ^2		.4145		.4405
χ^2		945.4		1005

[a]Standard errors are in parentheses and elasticities are in brackets

This schooling effect is in some sense "net" of the wage effect. Schooling increases the value of time in the home as well as in the market so it is not surprising that schooling decreases the probability of working when the wage is held constant.

The effect ot including children under six in this model is again striking. The coefficients for schooling change significantly from -1.7 to .67 for paid employment and from .97 to 1.7 for family work. More surprising however is the effect on the wage elasticities. When children under six are included, both wage elasticities are negative, with elasticities of -.96 and -.50 for employees and family workers respectively.

Table 6 presents the test statistics for testing the parameter restrictions $\beta_p = \beta_f$ in the trichotomous model. Under the null hypothesis, these statistics are distributed as chi-square variates with the degrees of freedom equal to the number of restrictions. This test clearly indicates at a .005 level of significance are 18.5, 20.3, and 22.0 with six, seven, and eight degrees of freedom respectively. These statistics clearly indicate that one should reject the null hypothesis at any conventional level of statistical significance; the statistics range from 218.4 to 230.8.

Although the likelihood ratio test indicates that the dichotomous model misspecifies the underlying choice model, this simpler model is estimated for purposes of comparison. Tables 7 and 8 present these empirical results. The coefficients are reported along with their standard errors. The elasticies are presented in brackets.[9]

[9] In the dichotomous models, a significant coefficient implies a significant elasticity.

Table 6: Likelihood Ratio Test Statistics for Testing the Hypothesis that $\beta_f = \beta_p$

	Empirical Specification	
	(1)	(2)
Reduced Form Model (degrees of freedom)	218.4 (7)	229.6 (8)
Model with Predicted Wages (degrees of freedom)	226.2 (6)	230.0 (7)
Model with Predicted Wages Corrected for Selectivity Bias (degrees of freedom)	226.0 (6)	230.8 (7)

Table 7: Maximum Likelihood Logit Estimates of the Dichotomous Participation Model Reduced Form

(1038 Observations)[a]

Independent Variable	(1)	(2)
Years of Schooling	.0343	.0138
	(.0210)	(.0216)
	[.5450]	[.2213]
Experience	.0123	-.0054
	(.0042)	(.0050)
	[.3604]	[-.1582]
ℓn Husband's Income	-.8007	-.8144
	(.2867)	(.2914)
	[-1.1432]	[-1.1628]
Husband's Schooling	-.0271	-.0271
	(.1117)	(.1140)
	[-.0280]	[-.0280]
Family Business	.7311	.7657
	(.0992)	(.1021)
	[.2393]	[.2507]
Home Ownership	-.1925	-.1847
	(.0801)	(.0825)
	[-.1647]	[-.1580]
Children Under Six	-	-.4217
		(.0660)
		[-.3678]
Urban Residence	.0546	-.0379
	(.0866)	(.0891)
	[-.0597]	[-.0414]
Intercept	10.30	11.28
	(3.95)	(4.02)
ℓn Likelihood	-577.2	-553.3
ρ^2	.1978	.2309
χ^2	284.6	332.3

[a]Standard errors are in parentheses and elasticities are in brackets.

Table 8: Multinomial Logit Estimates of the Dichotomous Participation Model Including Predicted Wages

(1038 Observations)[a]

Independent Variable	Wages Not Corrected for Selectivity Bias (1)	(2)	Wages Corrected for Selectivity Bias (1)	(2)
Years of Schooling	-.0909	.0708	-.0892	.0760
	(.0396)	(.0465)	(.0393)	(.0511)
	[-1.4575]	[1.1352]	[-1.4303]	[1.2186]
ℓn Husband's Income	-.8179	-.8263	-.8430	-.8588
	(.2864)	(.2904)	(.2860)	(.2906)
	[-1.1678]	[-1.1798]	[-1.2036]	[-1.2262]
Husband's Schooling	-.0258	-.0267	-.0262	-.0270
	(.1118)	(.1140)	(.1116)	(.1141)
	[-.0266]	[-.0276]	[-.0271]	[-.0279]
Family Business	.7283	.7635	.7492	.7911
	(.0992)	(.1019)	(.0994)	(.1060)
	[.2384]	[.2500]	[.2453]	[.2590]
Home Ownership	-.1861	-.1800	-.1911	-.1870
	(.0794)	(.0818)	(.0800)	(.0804)
	[-.1592]	[-.1540]	[-.1635]	[-.1600]
Children Under Six	-	-.4224	-	-.4381
		(.0660)		(.0740)
		[-.3684]		[-.3821]
ℓn Wage	1.241	-.5724	1.231	-.6223
	(.4328)	(.5102)	(.431)	(.5599)
	[1.7719]	[-.8173]	[1.7576]	[-.8885]
Intercept	5.058	13.94	5.453	14.63
	(4.347)	(4.63)	(4.284)	(4.96)
ℓn Likelihood	-577.4	-553.4	-577.4	-553.5
ρ^2	.1975	.2308	.1974	.2308
χ^2	284.2	332.1	284.1	332.1

[a]Standard errors are in parentheses and elasticities are in brackets.

Again, the results are presented with and without the number of young children. Most of the estimated coefficients in the specification without fertility are as expected. Years of schooling and potential labor market experience increase the probability of working. The husband's income, his level of schooling, and owning a home each decrease the probability of working. Also, owning a family business significantly raises the propensity to work. As in the trichotomous models, urban residence decreases the likelihood of working. In specification (2), the effect of schooling and experience decline, the effect of experience in fact becomes negative. Table 8 presents estimates which include the predicted wage, both uncorrected and corrected for selectivity bias. These estimates differ from those of the reduced form in one major respect; the effect of schooling becomes negative when the presence of children is not held constant. The estimated wage effect is significant and positive with an elasticity of 1.8. In the empirical specification that includes children however, the wage elasticity changes sign to - .8. As in the general models, ther is little evidence of selectivity bias in the wage estimates.

VII. Summary and Comparison with U.S. Estimates

Selected elasticities from the dichotomous and trichotomous logit models are presented in Table 9 along with estimates from two U.S. models. The Schultz results draw from his maximum likelihood logit estimation of a dichotomous participation model for white wives from the 1967 Survey of Economic Opportunity. The regressors include instrumental variables predictions of both the wife's and the husband's hourly wages. His asset variable is annual nonemployment income. The Heckman (1977) estimates are for white wives from the 1967 National Longitudinal Survey of Work Experience of Women Age 30-44. He estimates a probit model, the elasticities of which are based on his reported sample means. The probit model includes the husband's reported hourly wage as a regressor, but does not include the wife's wage. Instead, it includes schooling and labor market experience (both actual and predicted). In contrast to Schultz, Heckman (1977) does include children under six as a regressor.

The Japanese estimates summarize the results from Tables 3, 4, 7, and 8. For comparison with the Schultz estimates, the first set of Japanese results are based on the models which include the wife's predicted wage and which exclude young children. As discussed earlier, the results for Japanese wives vary substantially by specification of the dependent variable. The wage elasticities for paid employees and all employees are 2.6 and 1.8 respectively and are nearly twice the magnitude of the Schultz estimates which range from 0.2 to 1.0. Surprisingly, the own wage response for family workers is much closer at 0.7. Family workers are much more sensitive to the husband's income than are employees. The elasticity

Table 9: Summary of Japanese Elasticities and Comparison with U.S. Elasticities

Independent Variable	Trichotomous Model[a]		Dichotomous Model	Schultz Estimates[c]			Trichotomous Model[d]		Dichotomous Model[e]	Heckman Estimates[f]
	Employees	Family Workers		Age 25-34	Age 35-44	Age 45-54	Employees	Family Workers		
Wife's Wage	2.6	0.7	1.8	1.0	0.2	0.8	-	-	-	-
Husband's Wage	-0.4	-4.1	-1.2	-1.2	-1.0	-0.9	-0.4	- 4.1	-1.2	-0.1
Assets	-0.4	0.5	-0.2	0.004	0.02	0.06	-0.4	0.5	-0.2	.001
Schooling	-	-	-	-	-	-	-0.1	1.2	0.2	1.9
Labor Market Experience	-	-	-	-	-	-	-0.2	-0.1	-0.2	1.6
Number of Children Under Six	-	-	-	-	-	-	-0.6	-0.06	-0.4	-0.6

[a] Table 5, specification (1).

[b] Table 8, specification (1).

[c] Schultz, Table 2, pp. 39-40. Additional regressors included farm residence and own disability.

[d] Table 4, specification (2).

[e] Table 7, specification (2).

[f] Heckman (1977), calculated from Table 1, Specification (1), p. 25, using means for the entire sample reported in Table B.2, p. 49.

for all workers is -1.2 which equals Schultz's estimated elasticity with respect to the husband's wage for women 25-34. The "asset" variable (home ownership in our data and nonemployment income in Schultz's) has the expected negative sign only for employees and all workers in Japan.

Table 9 also includes results from the "reduced form" model which comprises children as a regressor. These models compare roughly with that of Heckman (1977). The effect of the Japanese husband's income is robust with respect to the model specification. Heckman's estimate (-0.1) is much lower than those of Schultz, and is in fact lower than that for employees in Japan. Also, Heckman's asset elasticity is positive. When children under six are included in the Japanese models, all of the experience elasticities and the elasticity of paid employment with respect to schooling become negative. (Recall from Tables 3 and 7 that these elasticities are appropriately signed when the fertility variable is left out). The elasticity of participation in family work with respect to schooling is 1.2, reasonably close to Heckman's estimate of 1.9. Regarding children, the effect for paid employees equals Heckman's (-0.6). Young children have a weakly negative effect (-0.06) on the decision to enter the informal sector. The elasticity for all workers lies between these two at -0.4.

There are several obvious limitations to a comparison of this nature. The empirical specifications of each model and the age stratification of each sample differ. Regardless, there is a remarkable similarity among these elasticities (and certainly little greater diversity than one encounters when comparing the two sets of estimates for the U.S.). The Japanese results for both the dichotomous and trichotlomous models strongly contradict the previous findings that an increase in their wage rate decreases the likelihood that Japanese women will work. Even the "standard" labor force participation model performs well.

There are three notable results that Table 9 does not highlight specifically. First, although the results of the dichotomous model are very strong, the likelihood ratio test clearly indicates that this model misspecifies the underlying framework of choice. Individuals obviously are not indifferent between working in the informal sector and working in the formal sector. Second, fertility has been treated extensively as an endogenous variable subject to individual choice. Rosenzweig and Wolpin clearly demonstrate the bias introduced when the actual rather than a truly exogenous fertility variable is included as a regressor in a labor force participation equation. The models estimated here are extremely sensitive to the inclusion of a fertility variable: the wage elasticity becomes negative. The persistent inclusion of this variable in other Japanese models may have contributed to the anomolous estimates of the own wage elasticity. Third, correcting for selectivity bias in the wage estimates produces very little change in the empirical results of either the trichotomous or the dichotomous models. (It did, however, add substantially to the computational expense.)

The specification and estimation of a theoretical participation model that accommodates the Japanese labor force has generated some surprising results. The estimates derived from this model closely resemble those for the U.S. The surprise arises from the fact that Japan's labor force is considered radically different from the labor force in the U.S.--generally in terms of the cultural and social heritage and specifically in terms of the job opportunities available to women in the informal sector in small business and cottage industries. However, the similarities in the labor force behavior evidenced by the preceding comparison should not be accepted as the last word. While women in the Tokyo Metropolitan Area may behave like women in the U.S., their behavior may not represent that of women in all Japan. Our data set especially underrepresents agricultural workers. Unfortunately, until the Japanese national labor force surveys are made available for public use, we cannot ascertain the representativeness of these results.

Appendix

As discussed in the text, instrumental variables predictions for both the wife's wage and the husband's income are used to estimate the participation models. Wages for a particular employment status are reported only for those participants so employed, consequently wages must be predicted for the entire sample. Three wages are estimated for each individual: a wage in paid employment, a wage in family work, and an overall wage. The husband's income is reported only if he is the primary income earner. Also, his reported wage may incorporate transitory variation and/or measurement error, so his income is predicted as well. This appendix first discusses the estimation of the wife's wage then reports the results of predicting the husband's wage.

The equations for the wife's hourly wage are corrected for samples election bias. This correction is based on work by Hay which extends Heckman's (1979) correction for probit choice model for application to logit choice models.

If the specification for the wage function is:

$$\text{(A.1)} \quad \ell n\, W_{ji} = \beta_j{}'\, X_i + \eta_{ji} \qquad j = p,\ f;\ i = 1,\ \ldots,\ N.$$

then sample selection bias arises when the expected value of η_{ij} given X_i does not equal zero, violating the assumptions for ordinary least squares regressions. We have wage data reported only for those women who are working and this group of women may not be a randomly selected subsample of the population. Heckman (1979) treated this problem as one of specification error. If the expected value of the error depends upon the selection rule, i.e. the labor force participation decision, and we have information regarding this decision, then we can include a regressor that corrects the error term:

(A.2) $E(\ln W_{ji} \mid X_i, \text{sample selection rule}) = \beta_j{}'X_i + E(\eta_{ji} \mid \text{sample selection rule})$.

The selection rule is simply the participation decision given by (7) in the text.[1] Then we may write:

(A.3) $E(\ln W_{ji} \mid X_i, (S_{ji} - S_{Ki}) > (\varepsilon_{Ki} - \varepsilon_{ji})) = \beta_j{}'X_i + \lambda_i \gamma,$

where in the probit model, λ_i is the inverse of Mill's ratio. Hay derives the appropriate form of λ_i when $(\varepsilon_{Ki} - \varepsilon_{ji})$ has a logistic distribution. Based on his derivation, we may estimate λ_i, include the estimate as a regressor, and obtain unbiased estimates of β_j using ordinary least squares.

If P_i is the probability of working, then the estimate for λ_i in the dichotomous choice model is relatively simple:

(A.4) $\lambda_i = \frac{1}{P_i} \left| P_i \log(P_i) + (1 - P_i) \log(1 - P_i) \right|$.

[1]The appropriate sample selection rule should perhaps be based on the decision to report earnings rather than the decision to participate in the labor force as those workers reporting earnings may be a non-ramdon sample of all workers.

The correction is more complicated in the trichotomous case since it is specific to the employment status:

$$(A.5)\quad \lambda_{ji} = \left(\frac{6}{\pi^2}\right)(-1)^{J+1}\left[\sum_{K \neq j} \frac{1}{J}\left(\frac{P_{Ki}}{1 - P_{Ki}}\right)\log(P_{Ki}) + \left(\frac{J-1}{J}\right)\log(P_{ji})\right].$$

The estimates for these corrections are based on the estimates of the appropriate P_{ji}. That is, for each individual, $\hat{P}_{ji}$ is calculated using the correct equation as given in the text. The estimated $\hat{\lambda}_{ji}$ is then included as a regressor for each wage equation.

The wages are formulated as standard earnings functions in which they depend on years of schooling, potential labor market experience, and an urban residence variable.[1] Tables A.1, A.2, and A.3 present these estimates. The first table (A.1) includes the estimates for the wage in paid employment. The first column contains the wage equation without $\hat{\lambda}_i$. The second column includes the correction and corresponds to the first specification of the trichotomous model (model (1) in Table 4). Similarly, the third column corresponds to the second specification of the trichotomous model (model (2) in Table 4). As is readily apparent, the inclusion of $\hat{\lambda}$ results in very little change in either the explanatory power of the wage function or in the estimated coefficients and their levels of statistical significance. The coefficient for λ_p is statistically insignificant.

Table A.2 includes comparable regressions for the wage in family work. Again, little change results when the selectivity correction is included.

[1]These variables are defined in the text, table 3.

Table A.1: Wage Equations for Paid Employees

(156 Observations)[a]

Independent Variable	Standard Wage Equation	Equation Corrected for Selectivity Bias (Model (1))	Equation Corrected for Selectivity Bias (Model (2))
Intercept	4.6698 (.3908)	4.6922 (.4460)	4.7104 (.4178)
Years of Schooling	.0731 (.0270)	.0733 (.0271)	.0733 (.0271)
Experience	.0092 (.0064)	.0091 (.0065)	.0091 (.0054)
λ_p	-	.0154 (.1460)	.0292 (.1042)
F	3.67	2.44	2.46
R^2	.0450	.0450	.0455

[a]Standard errors are in parentheses.

Table A.2: Wage Equations for Family Workers

(55 Observations)[a]

Independent Variable	Standard Wage Equation	Equation Corrected for Selectivity Bias (Model (1))	Equation Corrected for Selectivity Bias (Model (2))
Intercept	2.9167 (.8278)	2.9110 (.9038)	2.9409 (.9029)
Years of Schooling	.1925 (.0568)	.1928 (.0590)	.1915 (.0591)
Experience	.0090 (.0120)	.0090 (.0122)	.0089 (.0121)
Urban Residence	.6459 (.2657)	.6460 (.2682)	.6456 (.2682)
λ_f	-	-.0024 (.1459)	.0103 (.1458)
F	6.76	4.98	4.98
R^2	.2695	.2695	.2696

[a]Standard errors are in parentheses.

Table A.3 presents the wage estimates for the dichotomous participation models. The estimate for λ_1 is based on the logit models in Table 7. As in the wages estimated by employment sector, the overall average estimates are not influenced by the correction for selectivity bias.

In comparing the alternative wage estimates , only one noticeable difference arises. The effect of schooling is much higher in family work than in paid employment. For family workers, the elasticity of the wage with respect to schooling is 2.05 as against 0.77 for paid employees.

The predicting equation for the husband's income is based on a linear model in which the dependent variable, the natural logarithm of annual earnings, is regressed on three dummy variables measuring the highest level of schooling which the husband completed and a set of dummy variables describing the husband's occupation. Since the husband's age is not reported in the survey, the wife's age serves as a proxy. A dummy variable equal to one if the husband has ever been unemployed represents disruption in work experience. Table A.4 presents these results.

Table A.3: Wage Equations for All Workers

(215 Observations)[a]

Independent Variable	Standard Wage Equation	Equation Corrected for Selectivity Bias (Model (1))	Equation Corrected for Selectivity Bias (Model (2))
Intercept	4.3918 (.3573)	4.3538 (.4486)	4.4836 (.4076)
Years of Schooling	.1004 (.0248)	.1006 (.0249)	.0996 (.0249)
Experience	.0097 (.0057)	.0100 (.0061)	.0092 (.0058)
λ	-	-.0143 (.1018)	.0362 (.0769)
F	8.26	5.49	5.56
R^2	.0713	.0714	.0723

[a]Standard errors are in parentheses.

Table A.4: Equation for Predicting the Husband's Income

(918 Observations)[a]

Independent Variables	
Intercept	13.9613 (.0601)
Ever Unemployed (yes = 1)	-.1494 (.0513)
Schooling Dummies:	
High School	.1558 (.0319)
Technical School or Junior College	.2006 (.0465)
College	.3291 (.0424)
Occupation Dummies:	
Self-Employed	.3033 (.0359)
Professional	.1238 (.0520)
Manager	.4282 (.0553)
Salary Man, large firm	.2152 (.0423)
Salary Man, small firm	.1371 (.0448)
Blue collar, large firm	.0706 (.0468)
Wife's Age	.0077 (.0014)
F	29.34
R^2	.26

[a]Standard errors are in parentheses.

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