THE EFFECT OF HOME-BASED WORK ON EARNINGS

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ABSTRACT

This paper examines the effect of working at home on the hourly wages of men and women in the United States, using data from the May 1997 Current Population Survey. A human-capital model of hourly earnings is specified which incorporates the location of work as an explanatory variable. The model is estimated both by ordinary least squares (OLS) and by a maximum-likelihood (ML) procedure that treats work location as endogenous. The empirical results reveal a statistically significant and substantial wage premium for both men and women who work at home relative to their observationally equivalent, office-located counterparts, with the ML estimate of this premium exceeding the OLS estimate. Several explanations for these findings are discussed, and suggestions for future research are offered.

INDEX WORDS: Home-Based Work, Earnings, Endogenous-Treatment Models

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1. Introduction

According to the U.S. Bureau of Labor Statistics (1998), more than twenty million persons work at home full-time or part-time on their primary job. About one-half of those working at home are wage-and-salary workers or employees of a family-owned business who are not expressly paid for their work time at home, and approximately one-third are self-employed proprietors of home-based enterprises. However, the number of wage-and-salary employees who are paid for their work at home has doubled in the past ten years to almost four million, comprising seventeen percent of all home-based workers. Much of this increase in home-based employment has been attributed by Oettinger (2004) to rapid advancements in computer technology and the spread of high-speed data-transmission infrastructure that facilitate telecommuting in many clerical, professional, and managerial occupations.

Employees working at home have, in the recent past, attracted the attention of union activists and workplace-safety advocates, who lobbied for special protective legislation on behalf of these workers. The main argument for such protection was that the isolated character of home-based employment adversely affects the level and mix of compensation and working conditions received by such workers since they are less able than on-site employees to engage in the collective "voice" with management that is emphasized by Freeman and Medoff (1984).

There are at least two reasons to expect home-based workers to earn less than employees who work in a factory or office setting. First, the theory of compensating wage differentials predicts that home-based workers are willing to accept lower compensation for performing a given set of tasks because of the greater convenience and flexibility of working at home. For example, by working at home commuting costs are reduced or eliminated and opportunities arise to combine childcare and other productive activities in the home with job-related responsibilities. As a result, Edwards and Field-Henry (2002) argue that female home-based workers face lower wage offers and have lower reservation wages than women who work on site. Second, firms are predicted to offer home workers a lower wage because they are assumed to be less productive on the job. In particular, employees who work at home are more costly to monitor and supervise and are less integrated into the teamwork environment that characterizes efficient, modern workplaces [Becker and Murphy (1992)].

The general conclusion of an older empirical literature is consistent with the prediction that performing home-based work has a negative effect on hourly wages. Christensen (1988) found that women who work at home earn less than women who perform the same job in an office or factory. Specifically, employers were determined to have saved, on average, 30 to 50 percent in salaries on homeworkers. Moreover, firms did not offer home-based workers the same opportunities to improve their skills or receive promotions. A negative effect of working at home on the hourly wage was reported by Kraut (1988), and Kraut and Gerson (1988) estimated that homeworkers earned about $1800 less per year than on-site workers. Additionally, office workers were found to have benefit packages (except for pension contributions) that were superior to those of homeworkers.

More recently, Oettinger (2004) reported substantial wage penalties for U.S. home-based workers in all thirty-two Census sex-education-age categories in 1980 and 1990. However, Oettinger estimated small wage premia for working at home among younger, relatively well-educated males and females by 2000. Similarly, Pabilonia (2005) found that Canadian employees who performed some paid work at home during normal working hours ("telecommuters") earned about 13 percent more than their on-site counterparts.

In light of this more recent evidence, it is not surprising to find that both male and female wage-and-salary employees who perform some paid work at home now earn substantially <u>more</u>, on average, than on-site workers. Specifically, in 1997 home-based men earned $21.23 per hour (versus $14.43 per hour for office- or factory-located males), while women who were employed at home received $17.07 an hour (compared to $11.10 per hour for on-site female workers) [U.S.

Bureau of Labor Statistics (1998)]. These simple descriptive statistics may be misleading, of course, since comparisons of sample means do not control for the observable characteristics of workers, such as overall work experience, tenure on the current job, and years of schooling, that are positively correlated with on-the-job productivity and wages. Moreover, the location of workers on site or at home is not random; individuals choose whether or not to work at home, in part on the basis of their place of residence, number and ages of children, and occupation [Edwards and Field-Henry (2002)]. These considerations suggest that a better understanding of the effect of home-based employment on wages requires an analysis that not only controls for the human capital of workers which helps determine labor-market productivity, but also allows for the endogenous determination of the location of work.

This paper re-examines the effect of working at home on the hourly wage for men and women, using data from a special supplement to the May 1997 Current Population Survey (CPS). A human-capital model of hourly earnings is specified which incorporates the location of work as an explanatory variable. The model is estimated both by ordinary least squares and by a maximum-likelihood procedure that allows work location to be endogenously determined.

In the next section, we discuss the models used to estimate the effect of the location of work on the hourly wage. In the third section, we describe the data, present the empirical results, and provide several interpretations. Finally, we summarize our findings and conclude with caveats about these results and suggestions for future research. To preview our results, we find that home-based workers earn substantially more than their observationally equivalent, site-based counterparts. This conclusion holds for both men and women, and is a robust to estimation which treats the location of work as endogenous.

2. The Model

To analyze the effect on individual earnings of working at home, one could specify and estimate a standard human-capital equation

$$y_i = X_i\beta + \gamma D_i + \varepsilon_i, \tag{1}$$

where i indexes observations on individuals ($i = 1,...,N$), y_i is the natural logarithm of annual, weekly, or hourly earnings, X_i is a vector of exogenous explanatory variables, β is a vector of population parameters, γ is a scalar parameter, D_i is a dichotomous explanatory variable indicating whether the individual works at home ($D_i = 1$) or on site ($D_i = 0$), and ε_i is an i.i.d. error term. Under the assumption that the explanatory variables, including work location, are exogenous, $E(\varepsilon_i|X_i,D_i) = 0$ and the model in equation (1) can be consistently estimated by ordinary least squares (OLS). However, the decision to work at home or on site is affected by observed individual characteristics (for example, age, number of children, and marital status), as well as unobserved traits such as ability, that also affect wages. Therefore, it seems more appropriate to treat the location of work as an endogenous variable.

Accordingly, we specify an "endogenous-treatment-effect" model consisting of a probit equation determining the location of work and a normal regression equation determining the natural logarithm of the hourly wage, conditional on work location. This system of equations can be written

$$y_{1i}^* = X_{1i}\beta_1 + \varepsilon_{1i} \tag{2}$$

$$y_{1i} = 1 \qquad if\ y_1^* > 0$$

$$y_{1i} = 0 \qquad if\ y_1^* \leq 0$$

$$y_{2i} = X_{2i}\beta_2 + \delta y_{1i} + \varepsilon_{2i} \tag{3}$$

where y_1^* is the unobserved variable "preference for working at home," y_1 is an observed dichotomous variable indicating whether or not the individual works at home, y_2 is the natural logarithm of the hourly wage, X_1 and X_2 are vectors of explanatory variables, β_1 and β_2 are vectors of population parameters to be estimated, δ is a scalar parameter, and ε_1 and ε_2 are random errors which capture unobserved, individual-specific attributes and are assumed to have constant variances $\sigma_1^2 = 1$ and σ_2^2, respectively, and covariance σ_{12}.

When y_1 in (3) is endogenous, the conditional mean of ε_2 $[E(\varepsilon_2|X_2,y_1)]$ is not zero unless $\sigma_{12} = 0$, in which case consistent estimation of the wage equation can be achieved by OLS. If $\sigma_{12} \neq 0$, then consistent estimates are obtained by the two-step estimator described by Heckman (1979). These two-step estimates can serve as starting values for an iterative procedure for obtaining asymptotically efficient maximum-likelihood (ML) estimates. The ML estimator for this model is discussed in detail by Maddala (1983, pp. 121-122). Identification of the parameters of the linear wage equation (3) is achieved through the nonlinear functional form of the work-location equation (2). If over-identifying restrictions are available [for example, by the exclusion of exogenous explanatory variables from X_2 in (3) that are included in X_1 in (2)], their validity can be assessed by a likelihood-ratio test.

The relative magnitude of the OLS and ML estimators of the coefficient on the work-location variable is theoretically ambiguous. The OLS estimator will be biased upwards (downwards) if omitted – i.e., unobserved – factors that increase both an individual's productivity and earnings are positively (negatively) correlated with performing paid work at home; in this circumstance, there would be positive (negative) selection into home-based work. Suppose that, as Brown (1980) claims, workers with extensive unobserved productive abilities have high "full incomes" which they use to purchase desirable job traits. Then, if the opportunity to work at home is considered desirable, employees with greater unobserved earnings ability will be more likely to select into home-based work. On the other hand, if workers who face higher fixed costs of commuting to on-site work or are more easily able to combine home production with paid work at home have fewer unobserved market skills, then the OLS estimator of the true compensating wage differential for home-based work will be biased downward [Oettinger (2004)].

Conceptually, the proper framework for incorporating the endogeneity of the location of work is Heckman's (1978) treatment-effects model with dichotomous endogenous variables rather than the sample-selection model of Heckman (1979). The canonical sample-selection model [Greene (2000, pp. 928-930)] consists of two equations: an outcome equation (for example, determining the wage), which is of primary interest, and a sample-selection equation, typically determining employment status. In that model, the wage will be zero if the variable indicating employment status is equal to zero; that is, if the individual is not employed. However, in the present context the observed wage will not be zero if the dichotomous variable indicating whether or not the individual is working at home equals zero, because the individual will also have positive earnings if he or she works on site. Nevertheless, as is well-known [Maddala (1983, pp. 120-121)], two-step estimation of the sample-selection model will also generate consistent estimates of the focal parameter, δ.

3. Data and Empirical Results

3.1 Descriptive statistics

The data were taken from the May 1997 Current Population Survey (CPS). This survey contains two sets of questions, the basic CPS and the May special supplement. The May 1997 supplement provides extensive data on the compensation, socio-economic and demographic characteristics, and work location for approximately 60,000 persons employed in nonagricultural industries who worked during the survey reference week. The nonresponse rate for the May 1997 basic CPS was 6.6 percent, and there was an additional 11.9 percent nonresponse rate for the supplement (U.S. Bureau of the Census, 1999). Previous May supplements that provided data on working at home were conducted in 1991, 1985, and 1981. However, data from these prior surveys are not strictly comparable because of changes over time in survey questions and definitions.

We restricted the sample to persons who were between the ages of 16 and 65 at the time of the survey in order to include only work-eligible individuals and to exclude retired persons. The sample was further limited to persons who were not self-employed. The exclusion of self-employed persons is a very important feature of this study that contrasts it with most of the previous literature. One reason for this sample restriction is that self-employed workers cannot be subjected to opportunistic behavior by firms in the way wage-and-salary employees may be. In contrast to wage-and-salary workers, self-employed persons determine how many hours they want to spend on the job and under what conditions they work (for example, if they have their own office, work inside or outside the home, and which tools or equipment they use). Additionally, we restricted the sample to those who were paid for their work, thereby excluding volunteer workers and unpaid family members from the data. Finally, only civilian workers are included in the data. The resulting sample contains observations on 7,397 persons, with an almost equal proportion of men (50.76 percent) and women (49.24 percent).

A person is considered to work at home if he or she is paid for working at least one hour per week at home on a primary job. Under this definition, 987 individuals (or about 13 percent) in the sample work at home. Table 1 presents the percentages of men and women working at home in the three major occupational categories. The vast majority of men and women working at home in 1997 belonged to the Managerial & Professional Specialty, Technical, and Sales & Administration Support group. Those jobs, like computer system analyst, operations and system researcher, author, technical writer, designer, editor and reporter, computer programmer, insurance agent, and real estate agent, are usually relatively well-paid. Low-paid service occupations, like launderers, ironers, barbers, hairdressers, and cosmetologists, comprise only 2.5 percent of male and 4 percent of female homeworkers. The fraction of men and women in the sample who were working in Precision Production, Craft & Repair, and Operators is also very low (6 percent and 0.8 percent, respectively).

Tables 2 and 3 provide descriptive statistics for individuals who work at home and on site, by gender. Table 2 reveals that men who work at home are, on average, older, much better educated, and more likely to be married (with spouse present) than men who work on site. Male homeworkers are also less likely to be nonwhite, a union member, or to reside in a rural area than men who do not work at home. Most importantly, men who perform paid work at home supply more hours per week, earn substantially more per hour, and thus receive higher weekly earnings than males who work on site. These descriptive contrasts are largely confirmed by the signs of the ML-estimated coefficients in the equation determining the location of work for men that are reported in Table 5.

Table 3 provides these same descriptive statistics for women. As is the case for men, women homeworkers are more likely to be older, better educated, married (with spouse present), and white than women who do not work at home. However, in contrast to men, women who perform paid work at home are more likely to be members of a union (or, if not a union member, more likely to be covered by a union contract) than women who do not work at home. Like men, however, women homeworkers are less likely to live in a rural area than their counterparts who work on site. Finally, as was true for men, weekly hours worked, the hourly wage, and weekly earnings are substantially higher for women who work at home than for women who do not perform home-based work. Once again, these descriptive comparisons are consistent with the signs of the estimates of the work-location equation for women reported in Table 5.

3.2 Determinants of the hourly wage: OLS estimates

Table 4 reports the results of estimating by OLS a (log) hourly-wage equation separately for men and women, treating work location as an exogenous variable. The t-statistics are calculated as the ratio of the coefficient estimates to the associated Huber-White (robust) standard errors. This equation includes the following explanatory variables: age, age squared, race, education, marital status, region, metropolitan status, main occupational category, union-membership status, union-contract coverage, sector of employment, and location of work. The race categories are white, black, and other. We divided educational attainment into high-school dropout, high-school graduate, some college, college degree, and post college. Marital status is categorized as either married-with-spouse-present, divorced, separated, married-with-spouse-absent, widowed, or never married. The categorical variables representing region are Northeast, Midwest, South, and West. Metropolitan status has three categories: central city, suburban, and rural. The categories for main occupation are Managerial & Professional Specialty, Technical, Sales, and Administrative Support, Service occupations, and Precision Production, Craft, and Repair, and Operators. We also distinguish between private-sector and government-sector workers, and among individuals who are union members, those who are covered by a union contract without belonging to a union, and nonunion workers. The omitted categories in the model, then, are: white, high-school degree, never married, service occupations, residing in the Midwest, living in a suburb, government-sector employment, and nonunion worker who is not covered by a union contract.

The OLS estimates of the coefficients on the demographic, human-capital, geographic, occupational, and union-status variables are unremarkable in the sense that they are broadly consistent with the results routinely reported for wage regressions in past research. Age (a proxy for work experience and job tenure, neither of which is observed in the CPS) increases earnings but at a decreasing rate. Black males earn 14 percent less than white males, other things equal, while there is no statistically significant difference between the earnings of black and white females, ceteris paribus. Relative to high-school graduates, high-school dropouts earn less, and the return to schooling rises at successively higher educational categories for both men and women. Married men (with spouse present) and, somewhat surprisingly, married women earn more than their single counterparts, although the earnings premium for married men is much larger and more precisely estimated than for married women. The small, but statistically significant, positive effect of marriage on earnings for women might arise from assortive mating in which men are more likely to marry women with attributes that are unobservable to the researcher but that are rewarded in the labor market. Interestingly, divorced, separated, or married-with-spouse-absent men also earn more than never-married men, but this finding does not hold for women. Earnings are marginally lower in the West and South regions than in the Midwest, but are not significantly different in the Northeast. Compared to males residing in the suburbs, residents of central cities and rural areas earn less, other things equal. For women living in a central city, however, there is no earnings disadvantage. Both men and women who are members of a union earn a substantial wage premium, while nonunion workers of both genders who are covered by a union contract enjoy smaller, but statistically significant, positive earnings differentials relative to their uncovered counterparts. Finally, there is no difference in the earnings of private-sector and government-sector workers, ceteris paribus.

Of most interest for present purposes is the estimated coefficient on the dummy variable indicating the location of work. The results imply that male employees who perform paid work at home earn about 12 percent more than observationally equivalent men who work on site, while the earnings premium for female at-home workers is approximately 14 percent. This finding is somewhat surprising in light of the earlier empirical literature which emphasized the relative earnings disadvantage of working at home. It is consistent, however, with the OLS results reported by Oettinger (2004) and Pabilonia (2005). Nevertheless, the results from estimating this model should be viewed cautiously since the location of work is treated as exogenous.

3.3 Determinants of the hourly wage: ML estimates

To allow for the endogeneity of work location, maximum likelihood was used to estimate jointly the probit and normal-regression equations (2) and (3). Both the hourly-wage and work-location equations contain many of the same exogenous variables. However, the equation determining work location includes, additionally, two overidentifying exogenous variables measuring, separately, other family income (defined as household nonlabor income plus labor income not earned by the survey respondent) and the number of family members. [Nakamura and

Shaw (1994) summarize a body of research on female wage determination indicating that number of children does not affect wages directly but, rather indirectly through work experience.]

The resulting ML estimates of the wage equation are presented in Table 6. In almost every case, the size and significance levels of the ML-estimated coefficients on the control variables are very close to those obtained with OLS. Interestingly, however, the ML estimate of the coefficient on the variable denoting the location of work is substantially larger than the corresponding OLS estimate. That is, ignoring the endogeneity of the location of work would result in an underestimate of the return to working at home. Specifically, the ML estimates of the coefficients of the work-location variable reveal an additional 5 percent wage premium for both men and women working at home, relative to the OLS estimates. This finding is similar to the OLS and instrumental-variables results reported by Pabilonia (2005) for Canadian telecommuters. The evidence that the ML estimates exceed the OLS estimates, as well as the negative sign of the estimate of σ_{12}, suggest that employees who perform paid work at home possess relatively low amounts of unobserved skills that increase market productivity and wages; that is, there appears to be negative selection into home-based work.

Both OLS and ML estimates of the model reveal a statistically significant and sizeable earnings premium for working at home rather than on site. Nevertheless, these estimates differ substantially, and so it is worthwhile to evaluate the relative merits of the two estimators in the present context. We examine, in turn, the strength of the association between the predicted and actual values of the dummy variable indicating the location of work, the statistical validity of the over-identifying restrictions on the earnings equations, and the correlation between the error terms in the earnings and work-location equations. Each of these investigations provides some evidence regarding the desirability of OLS versus ML estimation.

Bound, et al. (1995) argued that simultaneous-equations estimators may be seriously biased, even in large samples, if the instruments are only weakly correlated with the potentially endogenous variable. To examine this correlation, we calculate the appropriate coefficient of determination for the reduced-form probit equation determining work location, as well as the proportion of predicted values that correctly identify the location of work. For the sample of males, the Kullback-Leibler (pseudo) R^2 for the reduced-form equation determining work location is 0.24, and the fraction of correct predictions is 87%. For the sample of women, the pseudo R^2 for the reduced-form probit equation is 0.21, and the fraction of correct predictions is, again, 87%. These two test statistics, then, provide conflicting evidence on the validity of the instrument for work location, since the R^2 is somewhat low but the fraction of correct predictions is high.

Another useful test statistic for assessing the validity of the ML estimator is the partial-R^2, F, or χ^2 statistic for the statistical significance of the exogenous variables that are included in the work-location equation but excluded from the wage equation. A likelihood-ratio (LR) test of the null hypothesis that these (over-) identifying, exclusion restrictions are invalid can be performed by calculating the LR statistic $-2(L_R - L_U)$, where L_R is the value of the likelihood function with the exclusion restrictions imposed and L_U is the value of the unrestricted likelihood function. With two zero restrictions imposed on the coefficients of the excluded exogenous variables, under the null hypothesis the LR statistic has a $\chi^2_{2,\alpha}$ distribution at the α–level of significance; with α = 0.05, the critical value for $\chi^2_{2,.05}$ = 5.991. For the sample of men, the calculated LR statistic is 5.58 < 5.991, so we cannot reject the null hypothesis that the exclusion restrictions are invalid. For the female sample, the LR statistic for the test of the null hypothesis that the exclusion restrictions are invalid is 9.04 > 5.991 = $\chi^2_{2,.05}$ so we conclude that the over-identifying restrictions are valid for the sample of women. Again, this criterion for judging the ML estimator provides mixed results, with the evidence from the sample of males providing no support for the exclusion restrictions, and the estimates from the female sample supporting them.

Finally, a Hausman-type specification test of the exogeneity of the work-location variable y_1 in the hourly-wage equation (3) is equivalent to a test of the null hypothesis $\sigma_{12} = 0$ against the two-sided alternative $\sigma_{12} \neq 0$. Rejection of the null hypothesis would amount to rejection of the exogeneity of the location of work in determining the hourly wage, providing support for the use of ML rather than OLS for estimating the wage equation. Under the null hypothesis, the LR statistic -2 (L_r - L_u) has a $\chi^2_{1,\alpha}$ distribution at the α-level of significance, where L_r is the value of the likelihood function when the restriction $\sigma_{12} = 0$ is imposed and L_u is the value of the unrestricted likelihood function. For the sample of men, the calculated LR is $1.63 < 3.84 = \chi^2_{1,.05}$ so that we cannot reject the null hypothesis that work location is exogenous. Similarly, for the sample of women the calculated LR is $1.11 < 3.84 = \chi^2_{1,.05}$, and we conclude again that the location of work is exogenous. The results of the specification test, then, suggest that ML estimation provides no advantage over OLS. Combined with the evidence from the goodness-of-fit and exclusion-restrictions tests, support for the use of ML (and treating work location as endogenous) is tenuous, at best.

One explanation for the results in Tables 4 and 6, and specifically the stark contrast to earlier estimates of a negative earnings effect of working at home, is the changing nature of home-based work. Advances in technology such as increased computerization, improvements in the speed of data transmission, and the rapid expansion of the Internet have all greatly enhanced the ability to work productively at home while staying connected to one's employer and colleagues. To the extent that these technological changes have increased the productivity of home-based employment relative to on-site work over time, there will be a growing (declining) earnings premium (discount) for working at home.

Another possible interpretation of the recent appearance of a wage premium earned by home-based workers is that it reflects a widening compensating differential for the work-related infrastructure capital normally provided by firms on site. While employers of homeworkers often supply personal computers, cable modems, and unfinished goods for use at home, they avoid having to provide these workers with office space, parking lots, dining rooms, and restrooms. If such workplace capital is more costly to provide and/or more highly valued by workers now than in the past, firms will be willing to pay (and workers will demand) a rising wage premium for working at home.

4. Concluding Remarks

The findings of most previous studies suggested that working at home reduces one's hourly wage. However, descriptive statistics from the May 1997 CPS reveal that the average hourly earnings of home-based workers are higher than those of their on-site counterparts. This conclusion is confirmed by more recent estimates of models that control for a lengthy set of individual characteristics and allow for work location to be endogenous. One possible explanation for these latter results is that recent technological changes have enhanced the relative productivity work at home. Another interpretation is that the emergence of a wage premium for home-based work represents a rising compensating differential for employee-provided workplace infrastructure.

This paper is exploratory, and several issues should be addressed in future work. First, omitted personal traits that are correlated with the location of work may be at least partially responsible for the positive estimated effect of working at home on earnings. Panel data could, in principle, be used to control for such unobserved heterogeneity. Unfortunately, data from the Census of Population and the CPS containing information on work location are cross-sectional. Second, it would be interesting to explore whether the estimated premium for working at home is related to usual hours of work. However, such an investigation would require the estimation of a model in which, along with work location, hours of work are endogenous. Finally, although both OLS and ML estimates point to a positive and substantial effect of home-based work on earnings, there are reasons to be cautious about the size – and even the sign – of this effect. Single-equation estimates of the earnings equation ignore the endogeneity of the work-location variable, and the resulting bias can in principle be reduced by the ML approach like the one followed here. On the other hand, simultaneous-equations estimates are notoriously fragile with respect to the specification and strength of the instrumental variables. Unfortunately, CPS (and, even more so, Census) data typically contain very few exogenous variables that are valid instruments for an earnings equation. Ideally, one would like to observe the outcomes of an experiment in which firms randomly assigned employees to identical on-site and home-based work tasks. In the absence of such experimental data, further progress in overcoming identification issues in the analysis of the earnings implications of working at home awaits the appearance of data sets with a richer set of exogenous variables or the discovery of an appropriate "natural" experiment.

TABLE 1: Summary Statistics; Location of work, by occupation, May 1997

Occupation	Work at Home		Work on Site	
	Men	**Women**	**Men**	**Women**
Managerial and Professional; Specialty, Technical, Sales and Administrative Support	442 (91.32%)	481 (95.63%)	1380 (42.19%)	2240 (71.36%)
Service Workers	12 (2.48%)	18 (3.58%)	339 (10.36%)	525 (16.73%)
Precision Production, Craft and Repair; Operators	30 (6.20%)	4 (0.80%)	1552 (47.45%)	374 (11.92%)
Subtotal	**484**	**503**	**3271**	**3139**
Total	**987**		**6410**	

TABLE 2: Variable Means, Men

Variables	Work at Home	Work on Site
Age	42.0372	37.9285
White	.9442	.8771
Black	.0289	.0789
Other	.0269	.0440
High School dropout	.0124	.1351
High School	.1198	.3718
Some college	.2128	.2651
College degree	.3698	.1645
Post college	.2851	.0636
Married – spouse present	.7810	.6154
Widowed	.0062	.0055
Divorced, separated, or married – spouse absent	.0764	.1110
Never married	.1364	.2681
Managers, Professionals, or Sales	.9132	.4219
Service workers	.0248	.1036
Production and Craft, or Repair; Operators	.0620	.4745
Northeast	.2107	.2033
Midwest	.2293	.2479
South	.2727	.2803
West	.2872	.2684
Central city	.2707	.2706
Suburban	.5599	.4766
Rural	.1694	.2528
Union member	.0971	.1899
Covered by union contract	.0227	.0156
Government	.1632	.1357
Private	.8368	.8643
Hours usually worked per week	48.2707	41.5622
Weekly earnings	1010.9829	608.4711
Hourly wage	21.2393	14.4250
Other income	1554.9953	968.5637
Number of family members	3.0682	3.2226

TABLE 3: Variable Means, Women

Parameters	Work at Home	Work on Site
Age	41.1988	38.0172
White	.9225	.8398
Black	.0457	.1105
Other	.0318	.0497
High School dropout	.0080	.0997
High School	.1332	.3756
Some college	.1968	.3195
College degree	.3996	.1650
Post college	.2624	.0401
Married – spouse present	.6521	.5581
Widowed	.0258	.0274
Divorced, separated, or married – spouse absent	.1690	.1803
Never married	.1531	.2342
Managers, Professionals, or Sales	.9563	.7136
Service workers	.0358	.1673
Production and Craft, or Repair; Operators	.0080	.1192
Northeast	.2326	.1972
Midwest	.2306	.2647
South	.2525	.3084
West	.2843	.2297
Central city	.2664	.2733
Suburban	.5408	.4552
Rural	.1928	.2714
Union member	.2127	.1137
Covered by union contract	.0418	.0166
Government	.3519	.1688
Private	.6481	.8311
Hours usually worked per week	41.0437	36.3549
Weekly earnings	706.9399	413.9756
Hourly wage	17.0676	11.0974
Other income	1384.8100	946.6590
Number of family members	2.8310	3.1446

TABLE 4: Hourly-wage equation; OLS estimates

Variable	**Men** **Estimate (t-statistic)**	**Women** **Estimate (t-statistic)**
Constant	1.1409 (12.07)	.9667 (12.38)
Work at Home	.1094 (4.14)	.1319 (5.38)
Age	.0456 (8.93)	.0472 (11.57)
Age squared	-.0005 (-7.43)	-.0005 (-10.18)
Black	-.1304 (-4.35)	.0015 (.06)
Other	-.0303 (-.78)	-.0137 (-.39)
High School dropout	-.2125 (-9.05)	-.1172 (-4.82)
Some college	.0455 (2.43)	.1393 (8.14)
College degree	.2956 (12.13)	.3685 (16.50)
Post college	.4265 (12.42)	.4701 (12.75)
Married – spouse present	.1803 (7.69)	.0416 (1.97)
Widowed	.1269 (1.45)	-.0301 (-.68)
Divorced, separated, or married – spouse absent	.1072 (3.36)	.0156 (.61)
Managers, Professionals, or Sales	.2749 (9.94)	.2345 (11.53)
Production and Craft, or Repair; Operators	.2243 (8.57)	.1239 (4.56)
Northeast	.0229 (1.08)	.0484 (2.24)
South	-.0286 (-1.40)	.0031 (.17)
West	-.0365 (-1.81)	.0308 (1.54)
Central city	-.0718 (-3.93)	-.0074 (-.42)
Rural	-.1420 (-8.00)	-.1567 (-9.07)
Union member	.1991 (10.19)	.1819 (7.87)
Covered by union contract	.0902 (1.32)	.0879 (1.75)
Private-sector employer	.0280 (1.17)	.0132 (.64)
Number of observations	**3755**	**3642**
R – squared	**0.3740**	**0.3543**

TABLE 5: Work-location equation; joint ML estimates

Variables	Men	Women
	Estimate (t-statistic)	Estimate (t-statistic)
Constant	-4.5808 (-10.19)	-3.6186 (-8.78)
Age	.1152 (5.17)	.0919 (4.49)
Age squared	-.0012 (-4.57)	-.0010 (-4.05)
Black	-.4774 (-3.15)	-.2661 (-1.97)
Other	-.4381 (-2.57)	-.3019 (-1.83)
High School dropout	-.3294 (-1.79)	-.3406 (-1.56)
Some college	.2399 (2.63)	.1949 (2.30)
College degree	.5128 (5.65)	.8658 (10.35)
Post college	.8581 (8.15)	1.3670 (13.21)
Married – spouse present	.2278 (2.38)	.1454 (1.51)
Widowed	.2348 (.68)	.1910 (.93)
Divorced, separated, or married – spouse absent	.0164 (.12)	.1119 (1.01)
Managers, Professionals, or Sales	.7003 (4.94)	.4347 (3.54)
Production and Craft, or Repair; Operators	-.3141 (-2.05)	-.4887 (-2.07)
Northeast	.0082 (.09)	.0848 (.97)
South	.0330 (.38)	-.0384 (-.47)
West	.1284 (1.50)	.2243 (2.67)
Central city	-.0530 (-.72)	-.1231 (-1.68)
Rural	-.1103 (-1.33)	-.1712 (-2.15)
Union member	-.2119 (-2.14)	.0547 (.63)
Covered by union contract	.0639 (.28)	.2341 (1.32)
Private-sector employer	.1805 (2.05)	-.1686 (-2.28)
Other income	.0001 (1.77)	.0001 (1.31)
Number of family members	-.0378 (-1.56)	-.0712 (-2.54)

TABLE 6: Hourly-wage equation; joint ML estimates

Variable	Men Estimate (t-statistic)	Women Estimate (t-statistic)
Constant	1.1512 (12.17)	.9735 (12.54)
Work at Home	.1586 (3.19)	.1799 (3.36)
Age	.0451 (8.84)	.0467 (11.50)
Age squared	-.0005 (-7.36)	-.0005 (-10.13)
Black	-.1278 (-4.27)	.0031 (.12)
Other	-.0269 (-.70)	-.0106 (-.30)
High School dropout	-.2122 (-9.07)	-.1168 (-4.81)
Some college	.0441 (2.36)	.1381 (8.06)
College degree	.2902 (11.84)	.3590 (14.83)
Post college	.4143 (11.63)	.4506 (10.72)
Married – spouse present	.1789 (7.64)	.0409 (1.94)
Widowed	.1254 (1.42)	-.0308 (-.70)
Divorced, separated, or married - spouse absent	.1072 (3.37)	.0150 (.59)
Managers, Professionals, or Sales	.2696 (9.69)	.2320 (11.44)
Production and Craft, or Repair; Operators	.2258 (8.67)	.1246 (4.60)
Northeast	.0226 (1.07)	.0478 (2.22)
South	-.0291 (-1.43)	.0033 (.18)
West	-.0377 (-1.88)	.0289 (1.43)
Central city	-.0712 (-3.91)	-.0063 (-.36)
Rural	-.1412 (-7.97)	-.1556 (-8.99)
Union member	.2005 (10.22)	.1809 (7.83)
Covered by union contract	.0898 (1.32)	.0848 (1.69)
Private-sector employer	.0263 (1.10)	.0151 (.73)
Log-likelihood	**-3372.4378**	**-3149.2506**
σ_{12}	**-0.0640** **(-1.28)**	**-0.0646** **(-1.05)**

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