# IMMIGRATION REFORM, JOB SELECTION AND WAGES IN THE U.S. FARM LABOR MARKET

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## Introduction

For much of the last decade, U.S. agricultural employers have hired a largely immigrant workforce: the *Findings from the National Agricultural Workers Survey for 2001-2002* reported that approximately 78 percent of all U.S. crop workers were foreign-born and that 53 percent of the workforce was unauthorized for U.S. employment (Carroll et al. 2005). Statistics such as these have undoubtedly contributed to increase national interest in immigration reform which has become quite a contentious issue on the political landscape. Immigration issues regained prominence mainly following the events of September 2001, and since then, much of the U.S. public and Congress have clamored for increased border and interior enforcement in an effort to crackdown on illegal immigration.

Judging from the basic stipulations of the immigration reform bill (H.R. 4437)[1] that was passed in the U.S. House of Representatives in December 2005, it is evident that some lawmakers favor legislation that may be more restrictive in scope than the 1986 Immigration Reform and Control Act (IRCA). The strong enforcement provisions of H.R. 4437 are in stark contrast to the pro-immigration measures outlined in the Comprehensive Immigration Reform Act of 2006 (S. 2611) approved by the Senate on May 25, 2006. Whatever the final outcome, a compromise on immigration reform could very well include greater border and interior enforcement, earned legalization and guest worker programs as the final measures which may be approved by the U.S. Congress.

[1] HR. 4437: Border Protection, Antiterrorism & Illegal Immigration Act of 2005.

The specialty crop sector is the most labor intensive sector of U.S. agriculture, and is highly vulnerable to immigration reform since much of its workforce is foreign-born and unauthorized. For this reason, if the U.S. Congress finally approves legislation that is inherently more stringent than adopted in the past (i.e. IRCA), the lack of low cost labor alternatives may pose significant challenges for the sector. Similar to the alternatives debated before IRCA's passage in 1986, legalization for unauthorized workers is once again being considered as an alternative to undocumented status. Proponents of legalization argue that the lack of legal status hinders the labor market options of unauthorized immigrant workers and that their wages and job opportunities would improve with an adjustment to legalized status – the expectation being that employers would be pressed to increase wages and improve working conditions in order to retain a stable core of workers. Employers, however, have expressed concern that labor availability and cost may be affected if the supply of immigrant labor were restricted by immigration reform[2]. This may have important implications for the viability of sectors that are dependent on immigrant labor.

Previous work has examined the extent to which legal status determines wage differentials and whether it affects the types of jobs[3] for which workers are hired. Taylor (1992) explained wages separately for primary (skilled) and secondary (unskilled) jobs in agriculture, arguing that there was self-selectivity into the two types of work. Legal status of the worker entered the earnings equations as an exogenous influence, argued to affect earnings differently for the two types of jobs. Isé and Perloff (1995) explained

[2] Although the unauthorized immigrant workforce are more likely to be affected by immigration reform, authorized workers, such as those on guest permits, may be impacted if the new legislation is more stringent than that which currently exists.
[3] Job type is designated by skill (skilled/unskilled).

farm wages based on a model with self-selectivity into legal status, and specified separate earnings equations for each status. Job type was not a consideration in their model. Using an ordered probit model to account for workers' self-selectivity into legal status, Iwai et al (2006) examined farm wage differentials and simulated how the wages of unauthorized workers change with adjustment to legal status. Selection on job type was not considered.

This research attempts to contribute to the body of literature by accounting for self-selectivity on job type as well as legal status. We specify an earnings model for farm workers using a double selection framework to represent the likely non-random selection of workers into the separate legal status and type of work categories. The distinction from the earlier work is that the legal status choices and type of work choices are treated jointly, reflecting potential joint choices by workers into various combinations of legal status and type of work. Clearly, specifying one of these choices as exogenous, or ignoring it, leads to biased and inconsistent estimates and creates selectivity bias in the estimated wage equations. We adopt an approach outlined by Tunali (1986) to introduce the double selection criteria into the specification. The data used to estimate the model are from the National Agricultural Worker Survey (NAWS) public use data set, including data from 1993-2002. The NAWS is a rich nationally representative data set on farm workers in crops in the U.S., including approximately 2,500 workers each year. The data set includes the key legal status and job type variables as well as the standard variables that are typically included in an earnings equation.

**Research Methodology**

We adopt the double selection model proposed by Tunali (1986) to jointly model foreign-born workers' self-selectivity into legal status and farm jobs, and the subsequent implications for farm wages. Previous empirical investigations assumed only one potential source of selectivity bias due to a single decision; however, we contend that selection bias could potentially arise from two decisions – in this case, legal status and type of work. We specify a bivariate probit model to reflect the two decisions in the first stage from which selectivity parameters are derived and included as explanatory variables in the second stage. Thus, we assume that the decisions made by the $i^{th}$ individual regarding legal status and job type are specified as follows:

$$y_{1i}^{*} = x_{1i}'\beta_1 + u_{1i} \qquad \textit{Legal status decision} \qquad (1)$$

$$y_{2i}^{*} = x_{2i}'\beta_2 + u_{2i} \qquad \textit{Job type decision} \qquad (2)$$

The log wage regression equation is specified as:

$$\ln W_{3i} = x_{3i}'\beta_3 + \sigma_3 u_{3i} \qquad (3)$$

Note that the individual's decisions (denoted by $y_{1i}$ and $y_{2i}$) are unobserved but that log earnings ($W_{3i}$) are observed in accordance with those decisions. The explanatory variables and unknown coefficients are represented by $x$ and $\beta$, respectively, and the error terms $(u_{1i}, u_{2i}, u_{3i})$ have zero mean and covariance matrix $\Sigma = \begin{bmatrix} 1 & \rho & \rho_{13} \\ \rho & 1 & \rho_{23} \\ \rho_{13} & \rho_{23} & 1 \end{bmatrix}$. With respect to legal status, explanatory variables that may affect a worker's decision include such characteristics as gender, marital status, ethnicity, English speaking ability, education, age, and U.S. farm work experience. Similarly, a worker's decision on the

type of work may be shaped by certain job characteristics such as the number of years of employment with his current employer, U.S. farm work experience, weeks of farm work in the previous year, whether or not he is paid by piece rate, employer type, and type of crop production he is involved with (i.e. specialty crop or otherwise). Additionally, demographic characteristics such as age, English speaking ability, the number of years since migration to the U.S. may also affect the work decision. With respect to the wage equations, relevant explanatory variables include U.S. farm work experience, education, age, English speaking ability, gender and ethnicity, payment scheme (piece rate) and seasonality of employment. Since we expect wages to be affected by the legal status and job type decisions made by the worker, we also include selectivity variables reflecting each decision.

Since $y_{1i}^{*}$ and $y_{2i}^{*}$ are unobserved, we therefore observe only dichotomous variables $D_1$, indicating whether the farm worker is authorized or not, and $D_2$ indicating whether he selects a skilled job or not. The outcomes of the selection rules are indicated by the dichotomous variables $D_1$ and $D_2$:

$$D_1 = \begin{cases} 1 & if \quad y_1^{*} > 0 \\ 0 & if \quad y_1^{*} \le 0 \end{cases} \qquad D_2 = \begin{cases} 1 & if \quad y_2^{*} > 0 \\ 0 & if \quad y_2^{*} \le 0 \end{cases} \tag{4}$$

Four subgroups $G_j \ (j = 1,...,4)$ are generated. The elements of $G_j$ are combinations of $D_1$ and $D_2$., i.e. $G_1 = (0,0), G_2 = (0,1), G_3 = (1,0)\ and\ G_4 = (1,1)$ where $G_1$ denotes foreign-born farm workers who are *unauthorized* and *unskilled*, and $G_2$, $G_3$ and $G_4$ are foreign-born farm workers who are *unauthorized* and *skilled, authorized* and *unskilled,* and

*authorized* and *skilled*, respectively[4]. Provided that all four subgroups are distinct and are completely classified, the probability $S_j$ that an individual is assigned to the $j^{th}$ subgroup is given by:

$$\begin{aligned} S_1 &= Pr(D_1 = 0, D_2 = 0) = Pr(y_{1i}^{*} \le 0, y_{2i}^{*} \le 0) \\ &= Pr(u_{1i} \le -C_1, u_{2i} \le -C_2) \\ &= \Phi_2(-C_1, -C_2; \rho) \end{aligned} \tag{5}$$

$$\begin{aligned} S_2 &= Pr(D_1 = 0, D_2 = 1) = Pr(y_{1i}^{*} \le 0, y_{2i}^{*} > 0) \\ &= Pr(u_{1i} \le -C_1, u_{2i} > -C_2) \\ &= \Phi_2(-C_1, C_2; -\rho) \end{aligned} \tag{6}$$

$$\begin{aligned} S_3 &= Pr(D_1 = 1, D_2 = 0) = Pr(y_{1i}^{*} > 0, y_{2i}^{*} \le 0) \\ &= Pr(u_{1i} > -C_1, u_{2i} \le -C_2) \\ &= \Phi_2(C_1, -C_2; -\rho) \end{aligned} \tag{7}$$

$$\begin{aligned} S_4 &= Pr(D_1 = 1, D_2 = 1) = Pr(y_{1i}^{*} > 0, y_{2i}^{*} > 0) \\ &= Pr(u_{1i} > -C_1, u_{2i} > -C_2) \\ &= \Phi_2(C_1, C_2; \rho) \end{aligned} \tag{8}$$

where $C_1 = x'_{1i}\beta_1$ and $C_2 = x'_{2i}\beta_2$, $\Phi_2$ is the standard bivariate normal distribution function and $\rho$ is the correlation coefficient (Tunali, 1986). Thus, for each subgroup with complete observations, we have $E(W_{3i} \mid x_{3i}, \theta) = x_{3i}'\beta_i + \sigma_3 E(u_{3i} \mid x_{3i}, \theta)$, where $\theta$ denotes the joint outcome of the double selection process. Selectivity bias arises if $E = (u_{3i} \mid x_{31}, \theta) \neq 0$ (Tunali, 1986; Vella, 1998).

The likelihood function (9) is maximized to yield consistent estimates of the parameters of the equations on legal status and skill in the first stage:

[4] Per the NAWS dataset, pre-harvest, harvest and post-harvest jobs are classified as unskilled positions, whereas semi-skilled and supervisory jobs are classified as skilled.

$$L = \begin{Bmatrix} \prod_{S_1} \Phi_2(-C_1,-C_2;\rho) . \prod_{S_2} \Phi_2(C_1,-C_2;-\rho) . \prod_{S_3} \Phi_2(-C_1,C_2;-\rho) . \\ \prod_{S_4} \Phi_2(C_1,C_2;\rho) \end{Bmatrix} \quad (9)$$

The inverse Mills ratios corresponding to each subgroup are calculated shown in equations (10.1) through (10.4). The $S_j$'s denote the probability that individuals are assigned to the $j^{th}$ subgroup, and $\phi(.)$ and $\Phi(.)$ represent the standard univariate normal density and distribution functions, respectively (Tunali, 1986).

(i) *For* $i \in G_1$ *(i.e. $D_1$=$D_2$=0)*:

$$E(u_{3i} \mid u_{1i} \le -C_1, u_{2i} \le -C_2) = \rho_{13}\lambda_{11} + \rho_{23}\lambda_{12}$$

$$\lambda_{11} = -\left[\frac{\phi(C_1)\Phi(-C_2^*)}{(S_1)}\right],\ \lambda_{12} = -\left[\frac{\phi(C_2)\Phi(-C_1^*)}{(S_1)}\right]$$
$$where\ C_1^* = \frac{C_1 - \rho C_2}{\sqrt{1-\rho^2}};\ C_2^* = \frac{C_2 - \rho C_1}{\sqrt{1-\rho^2}} \quad (10.1)$$

(ii) *For* $i \in G_2$ *(i.e. $D_1$=0; $D_2$=1):*

$$E(u_{3i} \mid u_{1i} \le -C_1, u_{2i} > -C_2) = \rho_{13}\lambda_{21} + \rho_{23}\lambda_{22}$$

$$\lambda_{21} = -\left[\frac{\phi(C_1)\Phi(C_2^*)}{(S_2)}\right],\ \lambda_{22} = \left[\frac{\phi(C_2)\Phi(-C_1^*)}{(S_2)}\right] \quad (10.2)$$

(iii) *For* $i \in G_3$ *(i.e. $D_1$=1; $D_2$=0):*

$$E(u_{3i} \mid u_{1i} > -C_1, u_{2i} \le -C_2) = \rho_{13}\lambda_{31} + \rho_{23}\lambda_{32}$$

$$\lambda_{31} = \left[\frac{\phi(C_1)\Phi(-C_2^*)}{(S_3)}\right],\ \lambda_{32} = -\left[\frac{\phi(C_2)\Phi(C_1^*)}{(S_3)}\right] \quad (10.3)$$

(iv)*For* $i \in G_4$ *(i.e.* $D_1$*=1;* $D_2$*=1):*

$$E(u_{3i} \mid u_{1i} > -C_1, u_{2i} > -C_2) = \rho_{13}\lambda_{41} + \rho_{23}\lambda_{42}$$

$$\lambda_{41} = \left[\frac{\phi(C_1)\Phi(C_2^{*})}{(S_4)}\right], \lambda_{42} = \left[\frac{\phi(C_2)\Phi(C_1^{*})}{(S_4)}\right] \tag{10.4}$$

The inverse Mills ratios are used as covariates in log wage equations for each subgroup of the foreign-born farm workers with the legal status and skill classifications outlined previously. This can be illustrated with the regression function for subgroup $G_1$:

$$\begin{aligned} W &= x'\beta + \sigma_3\rho_{13}\lambda_{11} + \sigma_3\rho_{23}\lambda_{12} + \sigma_3\nu \\ &= x'\beta + \beta_1\lambda_{11} + \beta_2\lambda_{12} + \sigma_3\nu \end{aligned} \tag{11}$$

where $\beta_1 = \sigma_3\rho_{13}$, $\beta_2 = \sigma_3\rho_{23}$, and $\nu = u_{3i} - \rho_{13}\lambda_{11} - \rho_{23}\lambda_{12}$.

**Data**

The data utilized in this study were obtained from National Agricultural Workers Survey (NAWS) public use data set for the period 1993-2002. The NAWS is an employment-based, random survey of the demographic and employment characteristics of the U.S. crop labor force, which samples crop workers in three cycles each year in January, April and May, and October to reflect the seasonality of agricultural production and employment (DOL, 2005). Table 1 shows the variable definitions for bivariate probit and wage equation models. Table 2 reports the summary statistics for the variables identified. As per the mean wages reported for each of the subgroups, workers who choose to be *authorized & unskilled* earn $7.33, followed by $7.15 by *authorized & skilled* workers, $6.76 by *unauthorized & unskilled* workers and $6.41 by *unauthorized &*

*skilled workers*, respectively. These mean wages do not take selectivity bias into account and are not likely to have resulted from random samples.

**Bivariate Probit Model: Selection on Legal Status & Job Type**

Table 3 shows the estimated coefficients and asymptotic standard errors for foreign-born workers' legal status (*auth*) and job type (*skill*) decisions that are jointly estimated by the bivariate probit model. Based on a 0.05 significance criterion, all of the coefficients of the *legal status equation* are statistically significant. Holding all other factors constant, authorized status is more likely for foreign-born workers who are female, married, English-speaking, non-hispanic, educated and experienced in U.S. farmwork. As per *job type,* all of the coefficients except for *farmwork weeks* and *age* are statistically significant at the 10% level of significance or better. Foreign-born workers who have completed several years of employment with their current employer and have U.S. farmwork experience, who can speak English and are not paid a piece rate, employed with a grower, or involved in specialty crop production are more likely to be skilled than workers who do not fit this profile. The correlation (ρ) between the errors of the two equations is positive and significant, which implies that the two decisions are interrelated. This further signals that foreign-born workers who are authorized for U.S. employment are more likely to be skilled.

Table 4 shows the marginal effects of the bivariate probit estimates of selection on legal status and job type. The model generates four possible outcomes for the joint legal status and job type decisions and their probabilities which indicate the classification of foreign-born workers into the following subgroups: *authorized & skilled; authorized & unskilled; unauthorized & skilled; unauthorized and unskilled*. In the *authorized &*

*skilled* subgroup, English-speaking ability has the largest positive marginal effect, followed by the female dummy variable. Conversely, the largest negative marginal effect is produced by the Hispanic variable. The direction of effect is similar for the *authorized & unskilled* subgroup of foreign-born workers, in that the largest positive marginal effects arise from the female dummy and the English-speaking ability variable, respectively, whereas the Hispanic variable has the largest negative marginal effect. The female dummy has the largest negative marginal effect in the *unauthorized & skilled* subgroup, followed by the piece rate dummy. Conversely, the Hispanic variable has the largest positive marginal effect. Lastly, as shown in the final column of Table 2, the Hispanic variable has the largest positive marginal effect on the *unauthorized & unskilled* subgroup: all other characteristics held constant, foreign-born workers who are Hispanic are 17 percent more likely to be unauthorized and unskilled compared to those who are non-Hispanic. On the other hand, workers who are female are 13 percent less likely to be unauthorized and unskilled, followed by those who speak English (11%).

**Wage Equation Models with Selectivity Bias Corrections**

Table 5 reports the estimated coefficients and asymptotic standard errors for the four wage equation models corrected for selection bias. The selectivity variables were created from the results of the bivariate probit in the first stage. The selectivity variable pertaining to legal status, $\lambda_1$, accounts for possible selection bias from the legal status decisions of foreign born workers. The estimated coefficients on $\lambda_1$ are statistically significant in the *authorized & unskilled* and the *unauthorized & unskilled* worker subgroups only; thus, selection bias would occur if the corresponding wage models were estimated using ordinary least squares and selectivity bias was not accounted for. The

selectivity variable on skill, $\lambda_2$, measures possible selection bias stemming from the job type decisions made by foreign born farm workers. The estimated coefficients on $\lambda_2$ are significant for the *authorized & unskilled* and the *unauthorized & unskilled* subgroups, respectively. These results imply that selectivity bias is present in parts of the system and that using ordinary least squares on each of the wage equations would be inappropriate.

With respect to the direction of influence on wages, workers who are paid by piece rate, educated, experienced and speak English are likely to be paid a higher wage. Piece rate and education have significantly positive effects on wages across all legal status and job type categories. English speaking ability has a significantly positive effect on wages on all categories except for *authorized & unskilled* workers. In contrast, experience has a significantly positive nonlinear effect on wages of the *unauthorized & skilled* and *unauthorized & unskilled* workers, respectively. The significantly positive nonlinear effect of age on wages is evident for all workers except those who are *unauthorized & unskilled*. Having Hispanic ethnicity appears to have a positive and significant effect on wages for workers who are *authorized & skilled* only. Workers who are employed on a seasonal basis are statistically significantly more likely to have a lower wage. Female workers are also likely to earn lower wages, though this is statistically significant in the case of workers who are *authorized & unskilled* and *unauthorized & unskilled*, respectively.

Table 6 shows the predicted wages for foreign born farm workers for each legal status and job type category. The average predicted wage is highest for workers who choose to be *authorized & unskilled* ($7.41), followed by that for workers who choose to be *authorized & skilled* ($7.11). Workers who are *unauthorized & unskilled* earn an

average wage of $6.71, and workers who are *unauthorized & skilled* earn $6.43. Foreign-born workers who select into the *authorized & skilled* subgroup earn higher wages (~6% greater) than those who choose to be otherwise, i.e. *unauthorized & unskilled*, thus there is an expected gain associated with selecting into *authorized & skilled* status. Noting the differences with the mean wages reported earlier in Table 2, we conclude that those mean wages are inconsistent. Per our results, we know that selectivity bias is present in parts of the system. Hence, wages must be conditioned on the selectivity variables in order to derive consistent estimates.

Overall, the direction of influence compares well with the findings of Isé and Perloff (1995) and Iwai et al. (2006), in that authorized status is associated with higher wages. The largest expected gain over *unauthorized & unskilled* status is associated with *authorized & unskilled* status (~10%). There is no expected gain from selecting into *unauthorized & skilled* status from *unauthorized & unskilled* status. Taylor (1992) concluded that unauthorized workers would realize no (significant) earnings gain moving from secondary (unskilled) jobs to primary (skilled) jobs; our results seem to be suggestive of the same.

## Conclusions

Based on data from the National Agricultural Workers Survey public use data set, an earnings model for foreign born farm workers was specified and estimated. A double selection framework was used to represent their likely non-random selections into separate legal status and job type categories. A bivariate probit model was employed in the first stage of the analysis from which selectivity variables were generated for the two

decisions; these were included as covariates in the wage equation model that was estimated in the second stage.

Our results indicate that the legal status and job type choices made by foreign born farm workers are strongly correlated, implying that these two factors are taken into joint consideration when selecting into U.S. farm work. The coefficients in the bivariate probit model for the legal status and job type decisions are all statistically significant except for farmwork weeks and age. With respect to the marginal effects and the subsequent direction of influence, the female dummy and the variable representing English speaking ability indicate the largest positive marginal effects on the probability of a foreign born worker being *authorized & skilled* and *authorized & unskilled.* Not surprising, the Hispanic dummy is associated with the greatest negative marginal effects, meaning that Hispanic workers are less likely to be observed in those subgroups. The situation is reversed in the *unauthorized & skilled* subgroup, in that the Hispanic dummy has the largest positive marginal effect on the probability of a worker being *unauthorized & skilled*. For this particular subgroup, the female dummy variable has the largest negative marginal effect. The findings are similar for the *unauthorized & unskilled* subgroup, in that all other factors held constant, Hispanic workers are 17 percent more likely to be *unauthorized & unskilled* in comparison to those who are non-Hispanic. In contrast, female workers are 13 percent less likely to be *unauthorized & unskilled*, as well as those who speak English (11 percent).

The results from our wage model for the different subgroups point to the presence of selectivity bias in the system, implying that the data observed with respect to farm worker earnings are not randomly generated. We therefore made the appropriate

corrections to the wage regressions by including selectivity variables to reflect workers' self-selections on legal status and job type. For most of the worker subgroups, we find that workers who are paid by piece rate, educated, experienced and speak English are likely to be paid a higher wage. Except for those who choose to be *authorized & unskilled*, workers who are *authorized & skilled* tend to earn more than those who self-select into the other subgroups (*unauthorized & unskilled* and *unauthorized & skilled*). Skill (job type) also does not appear to have as dramatic an effect on workers' earnings.

Although previous investigations did not explore worker self-selectivity arising from a joint decision framework as was done in this study, we find similar directions of influence. For example, Isé and Perloff (1995) and Iwai et al. (2006) also found that authorized workers earned more on average than unauthorized workers. Taylor (1992) explicitly considered skill and concluded that unauthorized workers would not realize an earnings gain by self-selecting into skilled positions; our results are also suggestive of this. From our perspective, the logical next step in the short term is to implement a set of simulations to determine what our findings may imply in the context of impending immigration reform.

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**Table 1: Explanatory Variables for Bivariate Probit & Wage Models**[5]

| Variable | Definition |
|---|---|
| **LnWage** | Natural logarithm of the real farm wage in 2002 dollars. Conversions from the nominal wage were made using the consumer price index for all urban consumers |
| **Authorized** | =1 if farm worker is authorized for U.S. employment (citizen, permanent resident, or has other work authorization)<br>= 0 if otherwise (i.e. unauthorized) |
| **Skill** | =1 if task is semi-skilled or supervisory job<br>=0 if otherwise (pre-harvest, harvest, post harvest jobs) |
| **Piece Rate** | = 1 if worker is paid by piece rate<br>= 0 if otherwise (by the hour, hour/piece combination, or salary) |
| **Seasonal Work** | =1 if worker is employed on a seasonal basis<br>= 0 if otherwise (year-round) |
| **Female** | =1 if female<br>=0 if male |
| **Hispanic** | = 1 if worker is Mexican-American, Chicano, Puerto-Rican, or a member of any other Hispanic ethnic group identified in the NAWS<br>=0 if otherwise |
| **Education** | Highest grade level of education completed by the farm worker, ranging from 0 to 16 |
| **English speaking ability** | = 1 if 'none at all'<br>= 2 if 'a little'<br>= 3 if 'somewhat'<br>= 4 if 'well' |
| **Married** | = 1 if 'married/living together'<br>=0 if otherwise |
| **Years with current employer** | Number of years of employment worker has completed with current employer. One year is measured as one or more days per year (NAWS) |
| **Farmwork weeks** | Farmwork weeks in the last year |

[5] Data were sourced from the National Agricultural Workers Survey. Definitions enclosed in quotation marks are as they appear in the NAWS *Codebook for Public Access Data*.

**Table 1: Explanatory Variables for Bivariate Probit & Wage Models (continued)**

| Variable | Definition |
|---|---|
| **Years since immigration** | Difference between the interview date and the year in which the farm worker first entered the U.S. to live or work |
| **Grower** | = 1 if employed by a grower<br>= 0 if employed by a farm labor contractor |
| **Specialty Crop** | = 1 if worker was employed in specialty crop production at the time of the interview<br>=0 if otherwise |
| **Age** | Respondent age in years |
| **Age$^2$** | Age squared |
| **Experience** | Years of U.S. farm work |
| **Experience$^2$** | Experience squared |
| $\lambda_1$ | Selectivity correction term from the legal status (authorized) decision equation |
| $\lambda_2$ | Selectivity correction term from the job type (skill) decision equation |

**Table 2: Summary Statistics for Explanatory Variables**

| | Authorized & Skilled Subgroup | | Authorized & Unskilled Subgroup | | Unauthorized & Skilled Subgroup | | Unauthorized & Unskilled Subgroup | |
|---|---|---|---|---|---|---|---|---|
| **Variable** | Mean | Standard Deviation | Mean | Standard Deviation | Mean | Standard Deviation | Mean | Standard Deviation |
| Real wage | 7.148 | 1.880 | 7.332 | 2.558 | 6.408 | 1.590 | 6.756 | 2.117 |
| Piece Rate | 0.143 | 0.351 | 0.207 | 0.405 | 0.165 | 0.371 | 0.251 | 0.434 |
| Seasonal Work | 0.717 | 0.451 | 0.720 | 0.449 | 0.822 | 0.383 | 0.772 | 0.419 |
| Female | 0.131 | 0.337 | 0.214 | 0.410 | 0.084 | 0.278 | 0.145 | 0.353 |
| Hispanic | 0.936 | 0.244 | 0.975 | 0.156 | 0.986 | 0.120 | 0.985 | 0.122 |
| Education | 5.749 | 3.332 | 5.730 | 3.475 | 6.088 | 3.118 | 6.027 | 3.242 |
| English speaking ability | 2.116 | 0.910 | 1.954 | 0.927 | 1.553 | 0.711 | 1.530 | 0.717 |
| Age | 39.117 | 11.205 | 38.340 | 11.218 | 28.161 | 9.716 | 28.510 | 10.085 |
| Experience | 16.183 | 8.766 | 14.765 | 8.641 | 5.758 | 6.061 | 5.162 | 5.642 |
| Married | 0.810 | 0.393 | 0.759 | 0.428 | 0.504 | 0.500 | 0.490 | 0.500 |
| Years with Current Employer | 6.364 | 5.593 | 5.607 | 5.129 | 2.641 | 2.662 | 2.426 | 2.391 |
| Farmwork Weeks | 37.584 | 12.351 | 36.037 | 13.535 | 33.149 | 15.103 | 32.741 | 16.269 |
| Years since Immigration | 18.151 | 9.023 | 17.393 | 9.270 | 6.637 | 7.018 | 5.972 | 6.777 |
| Grower | 0.804 | 0.397 | 0.798 | 0.401 | 0.744 | 0.437 | 0.770 | 0.421 |
| Specialty Crop | 0.803 | 0.398 | 0.852 | 0.355 | 0.694 | 0.461 | 0.774 | 0.418 |
| Sample size | 1854 | | 3841 | | 1449 | | 5719 | |

**Table 3: Bivariate Probit Model Estimates for Foreign-Born Workers' Decisions on Legal Status and Job Type[6]**

| Authorized | Parameter Estimate | Skill | Parameter Estimate |
|---|---|---|---|
| | 0.442** | Years with Current | 0.0097** |
| Female | (0.037) | Employer | (0.003) |
| | 0.229** | | 0.0009 |
| Married | (0.031) | Farmwork Weeks | (0.0008) |
| | 0.335** | | -0.261** |
| English | (0.018) | Piece Rate | (0.032) |
| | -0.595** | | -0.009** |
| Hispanic | (0.092) | Years since immigration | (0.003) |
| | 0.033** | | 0.026** |
| Education | (0.005) | Experience | (0.003) |
| | 0.203** | | 0.008 |
| Experience | (0.005) | Age | (0.006) |
| | -0.004** | | -0.0001* |
| $Experience^2$ | (0.0001) | $Age^2$ | (0.00007) |
| | 0.046** | | 0.107** |
| Age | (0.008) | English | (0.015) |
| | -0.0004** | | -0.107** |
| $Age^2$ | (0.0001) | Grower | (0.298) |
| | | | -0.214** |
| | | Specialty Crop | (0.029) |
| Sample size | | 12,863 | |
| Log-likelihood | | -12472.257 | |
| Rho (ρ) | | 0.142** | |

[6] Standard errors are given in parentheses.
*Asterisks (**, *) indicate that the estimated coefficient is statistically significant at 1% and 10% levels of significance, respectively.

**Table 4: Marginal Effects of Bivariate Probit Estimates of Selection into Legal Status and Job Type**

| Variable | Marginal Effect | | | |
|---|---|---|---|---|
| | Authorized & Skilled | Authorized & Unskilled | Unauthorized & Skilled | Unauthorized & Unskilled |
| Female | 0.0439 | 0.1305 | -0.0439 | -0.1305 |
| Married | 0.0229 | 0.0652 | -0.0229 | -0.0652 |
| English speaking | 0.0489 | 0.0814 | -0.0149 | -0.1155 |
| Hispanic | -0.0574 | -0.1759 | 0.0575 | 0.1759 |
| Education | 0.0032 | 0.0093 | -0.0033 | -0.0094 |
| Experience | 0.0241 | 0.0546 | -0.0157 | -0.0629 |
| Age | 0.0057 | 0.0119 | -0.0031 | -0.0147 |
| Years with Current Employer | 0.0014 | -0.0014 | 0.0017 | -0.0017 |
| Farmwork Weeks | 0.0001 | -0.0001 | 0.0002 | -0.0002 |
| Piece Rate | -0.0353 | 0.0353 | -0.0434 | 0.0434 |
| Years since Immigration | -0.0012 | 0.0012 | -0.0016 | 0.0016 |
| Grower | -0.0154 | 0.0154 | -0.0194 | 0.0194 |
| Specialty Crop | -0.0312 | 0.0312 | -0.0395 | 0.0395 |

**Table 5: Wage Equation Models for Each Legal Status and Job Type Worker Subgroup**[7]

| | Authorized & Skilled | Authorized & Unskilled | Unauthorized & Skilled | Unauthorized & Unskilled |
|---|---|---|---|---|
| Piece rate | 0.07645** | 0.28580** | 0.09565** | 0.19551** |
| | (0.02076) | (0.01410) | (0.01808)) | (0.01059) |
| Seasonal work | -0.02522* | -0.06671** | -0.02462* | -0.03389** |
| | (0.01187) | (0.00959) | (0.01309) | (0.00767) |
| Female | -0.01733 | -0.04902** | -0.01310 | -0.03891** |
| | (0.01628) | (0.01248) | (0.01947) | (0.01008) |
| Hispanic | 0.06413* | 0.03077 | 0.04674 | -0.01533 |
| | (0.02571) | (0.02766) | (0.04302) | (0.02701) |
| Education | 0.00436* | 0.00482** | 0.00445* | 0.00640** |
| | (0.00195) | (0.00149) | (0.00184) | (0.00111) |
| English speaking ability | 0.03252** | 0.00901 | 0.02914** | 0.02322** |
| | (0.00990) | (0.00811) | (0.01091) | (0.00673) |
| Age | 0.01425** | 0.00515* | 0.01015** | 0.00146 |
| | (0.00344) | (0.00268) | (0.00328) | (0.00171) |
| $Age^2$ | -0.00019** | -0.00007* | -0.00015** | -0.00002 |
| | (0.00004) | (0.00003) | (0.00005) | (0.00002) |
| Experience | 0.00675 | 0.00076 | 0.02000** | 0.01788** |
| | (0.00491) | (0.00438) | (0.00517) | (0.00302) |
| $Experience^2$ | -0.00007 | -0.00006 | -0.00031** | -0.00035 |
| | (0.00009) | (0.00008) | (0.00010) | (0.00006) |
| $\lambda_1$ | -0.02555 | -0.07301* | 0.04518 | 0.05376* |
| | (0.04582) | (0.03510) | (0.04550) | (0.02808) |
| $\lambda_2$ | -0.05197 | -0.34329** | 0.20545** | 0.09941 |
| | (0.06957) | (0.07339) | (0.05971) | (0.06998) |
| Sample size | 1854 | 3841 | 1449 | 5719 |

[7] Standard errors have not been corrected for the two-step estimation.

**Table 6: Average Predicted Wage for Each Legal Status & Job Type Subgroup**

| Legal Status and Job Type Subgroups | Wage[8] ($) |
|---|---|
| Authorized & skilled | 7.11 |
| Authorized & unskilled | 7.41 |
| Unauthorized & skilled | 6.43 |
| Unauthorized & unskilled | 6.71 |

[8] Average wages are conditioned on the selectivity variables for legal status and job type.