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IMMIGRATION POLICY AND THE AGRICULTURAL LABOR MARKET: THE EFFECT ON JOB DURATION

By

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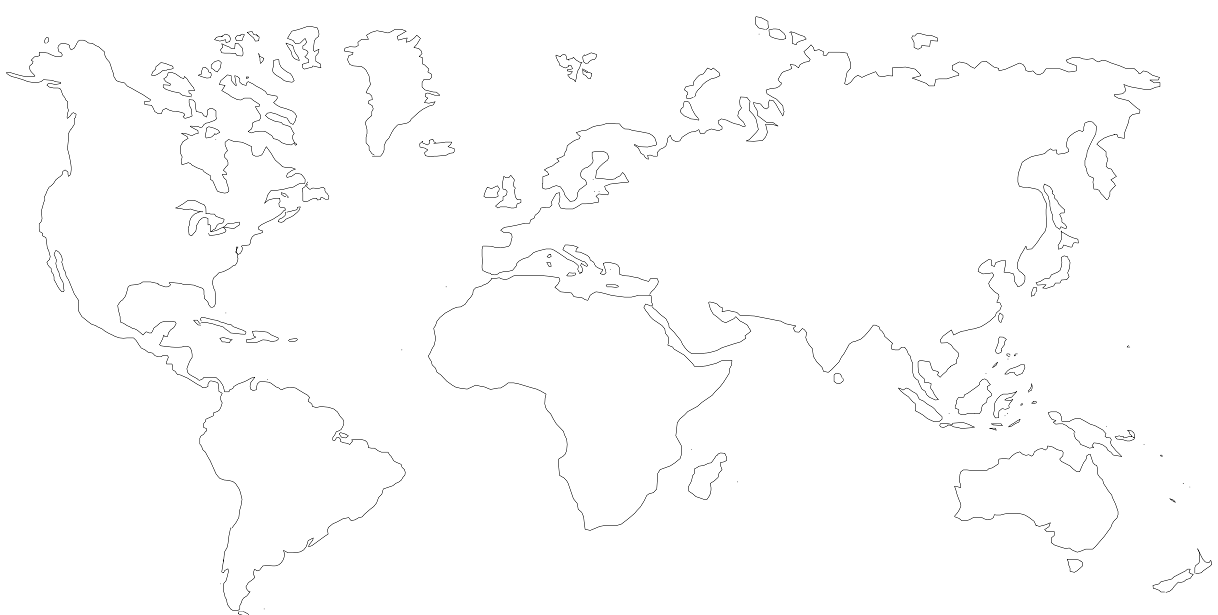

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IMMIGRATION POLICY AND THE AGRICULTURAL LABOR MARKET: THE EFFECT ON JOB DURATION*

Introduction

The effects of immigration policy change on the agricultural labor market have received much attention both economically and politically. The most important immigration policy change in recent years for the agricultural labor market was the Immigration Reform and Control Act (IRCA) of 1986. IRCA granted amnesty to a substantial number of undocumented agricultural workers, entitling them to work legally in the United States. Just before the passage of IRCA, many farmers and legislators expressed concern about its possible effect on the agricultural labor market. Their prediction was that undocumented agricultural workers who received amnesty would leave agriculture for other employment opportunities, which would lead to serious labor shortages and wage increases in agriculture.[1]

Limited empirical work has been done on the relationship between legal status and farm work duration (Hashida and Perloff 1996, Tran and Perloff 2002, and Emerson and Napasintuwong 2002). Generally, these studies conclude that estimated durations for documented, in contrast to undocumented, workers are significantly longer. Among these, the most comprehensive study is Tran and Perloff (2002). Using the National Agricultural Workers Survey (NAWS) data for the years 1987-91, Tran and Perloff estimate a stationary, first-order Markov model of employment turnover, and calculate the steady-state probability for each demographic group to work in agriculture. They conclude that "Predictions made when the 1986

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[1] See Tran and Perloff (427-28) for a detailed discussion of industry and legislative concerns.

Immigration Reform and Control Act was passed that granting people amnesty would induce most of them to leave agriculture were incorrect," (p. 427) and ". …. the steady state probability of working in agriculture is higher for someone with amnesty than for an undocumented worker, so that IRCA increased the long-run probability that people granted amnesty stayed in agriculture." (p. 437)

However, this conclusion is a little problematic. As the authors mentioned in their work, the portion of undocumented workers in the agricultural labor force grew substantially in the 1990s. In the sample (1987-91) used by Tran and Perloff, only 7% are undocumented workers. According to the NAWS data, the portion of undocumented workers rose to 46% for the years 1995-98, and 48% for the years 2002-2004. This implies that there has been a large-scale inflow of undocumented workers into the agricultural labor market and a large-scale outflow of documented workers from it. The latter might mean that documented workers tend to leave agriculture in the long-run: the opposite observation to their conclusion.

There are some concerns that might lead to statistical problems in their work. First, a data sample (1987-91) is taken in a transitional period in the sense that workers granted amnesty might not have had enough time to move to other industries. It is also a transitional period in another sense that the legal status of many workers changed. The study is unable to control for this status change using the observed status at the time of interview, the only legal status information available in the NAWS data. As a result of the 1987-91 sample used, the study cannot capture the major inflow of undocumented workers from foreign countries after IRCA and who have become a major component of the labor force in agriculture. The most serious problem, however, is that the study tries to estimate a probability matrix and a steady state for the whole migration process using data from only a small sector (the agricultural labor market). Most migration for any status of worker would be from non-agriculture to non-agriculture, and most would not work in agriculture at all. It may be difficult to estimate the whole migration pattern without data from all sectors. In this presentation, we present an alternative method (duration model with sample bias correction) to estimate the effect of the legal status of a worker on duration in farm work. Based on the existing studies which have used the duration model (Hashida and Perloff 1996, Emerson and Napasintuwong 2002), we develop the Heckman-type two-stage method, with the ordered probit model in the first stage and the duration model in the second stage.

The sample selection bias issue should be investigated first. Duration for a worker with a legal status is observed only if the worker is in that legal status. Each foreign-born worker chooses his/her legal status, considering conditions such as his/her individual demographic characteristics, cost of application, and benefit of the status. Without correcting for this selection process, the duration model will yield biased estimators. Hashida and Perloff (1996) correct selection bias using Lee's extension of Heckman's two-stage sample selection method (Lee 1983). In the first stage, the multinomial logit model is run to calculate a correction term assuming the error term has a Gumbel distribution. The second-stage duration model with this correction term does not generally yield consistent estimates with the normal distribution assumption of error term in the duration model.[2] We will use the ordered probit model in the first stage for two reasons: (1) this is consistent with the assumption about the error term in duration model in the second stage and (2) the multinomial logit does not account for the ordinal nature of the legal status. Considering the advantages in the labor market, they can be ordered as "citizen, permanent resident, authorized, and unauthorized workers."[3]

[2] Lee's method yields consistent estimator under very restrictive condition (Bourguignon et al. 2004).
[3] The definition of each legal status is given in the Data section.

Next, treatment of completed and uncompleted employment spells of workers should be considered. Hashida and Perloff (1996) and Tran and Perloff (2002) use only completed spells, while Emerson and Napasintuwong (2002) use only uncompleted spells. There are further distinctions in how spells have been defined in the literature. Hashida and Perloff (1996) define the duration variable as the average duration of completed spells of farm employment by a worker. Tran and Perloff (2002) work with employment transitions among three types of spells: agricultural employment, nonagricultural employment, and unemployed or abroad. They recorded a transition on a monthly basis over a two-year work history among the three above types of spells without regard to employer. Emerson and Napasintuwong (2002) define the duration variable as the number of years reported working in U.S. agriculture. At this point our estimation uses multiple completed spells per worker of agricultural employment at a single task. Our current definition is closest to the one used by Hashida and Perloff (2002), and specifically addresses variations in individual job duration by farm workers.

Methodology

The basic structure of the Heckman-type two-stage method is specified with the ordered probit model for the first stage and the duration model for the second stage. The ordered probit model is used to explain the legal status of worker *i* as a function of the individuals' demographic and policy variables (denoted as vector x_i). A foreign-born worker's legal status (J_i) takes on four values: 0=unauthorized, 1=authorized, 2= permanent resident (green card holder), 3=citizen. With the familiar argument of latent regression (Greene 2003), we can assume that an unobserved variable J_i^* is censored as follows:

$$
\begin{aligned}
J_i &= 0 \quad \text{if } J_i^* \le \mu_0, \\
J_i &= 1 \quad \text{if } \mu_0 < J_i^* \le \mu_1, \\
J_i &= 2 \quad \text{if } \mu_1 < J_i^* \le \mu_2, \\
J_i &= 3 \quad \text{if } \mu_2 < J_i^*.
\end{aligned}
$$

where $J_i^* = x_i'\alpha + \varepsilon_i$; x_i is a vector of exogenous characteristics of individual *i*; and ε_i is a disturbance term. The characteristics include gender, marital status, English speaking ability, race (black, white, and other), ethnicity (Hispanic and other), age, age squared, education, education squared, US farm experience, US farm experience squared, and the year of interview (before 1993, after 2001, and in-between).[4] We assume that ε_i is normally distributed with a mean of zero and a standard deviation of σ_ε. Then the likelihood function can be expressed as

$$
L(\alpha, \sigma_\varepsilon, \mu_j \mid data) = \left\{ \begin{array}{l} \prod_{J_i=0} \left[\Phi\left(\frac{\mu_0 - x_i'\alpha}{\sigma_\varepsilon} \right) \right] \prod_{J_i=1} \left[\Phi\left(\frac{\mu_1 - x_i'\alpha}{\sigma_\varepsilon} \right) - \Phi\left(\frac{\mu_0 - x_i'\alpha}{\sigma_\varepsilon} \right) \right] \\ \prod_{J_i=2} \left[\Phi\left(\frac{\mu_2 - x_i'\alpha}{\sigma_\varepsilon} \right) - \Phi\left(\frac{\mu_1 - x_i'\alpha}{\sigma_\varepsilon} \right) \right] \prod_{J_i=3} \left[1 - \Phi\left(\frac{\mu_2 - x_i'\alpha}{\sigma_\varepsilon} \right) \right] \end{array} \right\}, \qquad (1)
$$

where $\Phi(\cdot)$ indicates the cumulative distribution for the standard normal.

Suppose the cumulative distribution function of farm work duration (t_{ij}) for person *i* with legal status *j* is given as

$$F_{ij}(t) = \Pr(t_{ij} < t).$$

We denote its density function as $f_{ij}(t)$. The probability for the spell to be of length of at least *t*, usually called the survival function, is given as

$$S_{ij}(t) = 1 - F_{ij}(t).$$

[4] See the Data section below for additional detail.

Suppose that the log of the spell is normally distributed with mean $\ln \tau_{ij}$ and variance σ_j. Then, the survival function is expressed as

$$S_{ij}(t) = 1 - \Phi\left(\frac{\ln t - \ln \tau_{ij}}{\sigma_j}\right).$$

The hazard rate, the rate at which the spell is completed after duration *t*, is

$$h_{ij}(t) = \frac{\phi\left(\frac{\ln t - \ln \tau_{ij}}{\sigma_j}\right)}{t\sigma_j S_{ij}(t)},$$

where $\phi(\cdot)$ is the probability density for the standard normal distribution. Next, we assume that the mean duration of a spell ($\ln \tau_{ij}$) depends on independent variables z_i (gender, marital status, age, age squared, education, US farm experience, English speaking ability, race, ethnicity, availability of free housing, task, region (California, Florida, and other), the year of the interview (after 2001 or not), dummy variable for seasonal workers) so that

$$\ln \tau_{ij} = z_i^{'}\beta_j .$$

Then, the duration can be expressed as $\ln t_{ij} = z_i^{'}\beta_j + u_{ij}$ where $u_{ij} \sim N(0, \sigma_j)$. However, duration t_{ij} is observed only if person *i* has legal status *j*. This is a typical case for selection bias. Assuming e_i and u_{ij} are bivariately normally distributed with correlation coefficient ρ, the mean of the log of the duration conditioned on the legal status of person *i* is corrected as

$$E\left[\ln t_{ij} \mid \ln t_{ij} \text{ is observed}\right] = z_i^{'}\beta_j + \rho\sigma_j\lambda_{ij}$$

where λ_{ij} is the correction term for the selection bias which is given as[5]

[5] Correction term is set to zero for native-born citizen.

$$\lambda_{ij} = -\frac{\phi\left(\frac{\mu_j - x_i'\alpha}{\sigma_\varepsilon}\right) - \phi\left(\frac{\mu_{j-1} - x_i'\alpha}{\sigma_\varepsilon}\right)}{\Phi\left(\frac{\mu_j - x_i'\alpha}{\sigma_\varepsilon}\right) - \Phi\left(\frac{\mu_{j-1} - x_i'\alpha}{\sigma_\varepsilon}\right)}$$

Note that we can use the result of the ordered probit model in the first stage for $\frac{\mu_j - x_i'\alpha}{\sigma_\varepsilon}$ and $\frac{\mu_{j-1} - x_i'\alpha}{\sigma_\varepsilon}$. Also note that $\mu_{-1} = -\infty, \mu_3 = \infty$ from the assumption of normal distribution. In the second stage of this Heckman type two-stage method, we estimate equation (2) below by ordinary least squares with only completed spells.

$$\ln t_{ij} = z_i'\beta_j + \rho\sigma_j\lambda_{ij} + e_{ij} = z_i'\beta_j + \beta_{\lambda_j}\lambda_{ij} + e_{ij} \tag{2}$$

We can also show that the conditional variance of the log of the duration would be

$$\text{var}\left[\ln t_{ij} \mid \ln t_{ij} \text{ is observed}\right] = \sigma_j^2\left[1 - \rho^2\delta_{ij}\right],$$

where

$$\delta_{ij} = \left[\frac{\left(\frac{\mu_j - x_i'\alpha}{\sigma_\varepsilon}\right)\cdot\phi\left(\frac{\mu_j - x_i'\alpha}{\sigma_\varepsilon}\right) - \left(\frac{-x_i'\alpha}{\sigma_\varepsilon}\right)\cdot\phi\left(\frac{-x_i'\alpha}{\sigma_\varepsilon}\right)}{\Phi\left(\frac{\mu_j - x_i'\alpha}{\sigma_\varepsilon}\right) - \Phi\left(\frac{-x_i'\alpha}{\sigma_\varepsilon}\right)}\right] + \left[\frac{\phi\left(\frac{\mu_{j-1} - x_i'\alpha}{\sigma_\varepsilon}\right) - \phi\left(\frac{-x_i'\alpha}{\sigma_\varepsilon}\right)}{\Phi\left(\frac{\mu_{j-1} - x_i'\alpha}{\sigma_\varepsilon}\right) - \Phi\left(\frac{-x_i'\alpha}{\sigma_\varepsilon}\right)}\right]^2.$$

Then a consistent estimator of σ_j^2 is given by $\hat{\sigma}_j^2 = \sum_{i=1}^{n_j}\left[\hat{e}_{ij}^2 + \hat{\beta}_{\lambda_j}^2\hat{\delta}_{ij}\right]/n_j$. We can obtain the estimator of the asymptotic covariance matrix for $[\hat{\beta}_j, \hat{\beta}_{\lambda_j}]$ by substituting these results in the formulation in Greene (2003).

The difficulty in the farm worker duration model is that it has two sources of inconsistency. The observations are censored in the sense that the duration of a person with a particular legal status is observed only if the person has that status. Some observations are also censored in the sense that they are uncompleted. On the other hand, the legal status model (ordered probit) does not have restriction on observations, so that it should be a consistent estimator. The above method takes care of the selection bias by using correction terms for the mean duration. The current estimation approach drops uncompleted spells from the data set, introducing an unknown extent of bias in the estimation. However, given the size of the data set, the bias is believed to be minimal.

Data

The data used in this study are obtained from the National Agricultural Workers Survey (NAWS) (Office of Assistant Secretary for Policy). We used the study period from 1989, when the NAWS was first available, to the most recent year, 2004. This section will describe the definitions of each variable we used in our model.

Legal status is a discrete variable ranging from 0 to 3. Status 0 = "unauthorized" workers means that the worker is undocumented (did not apply to any legal status or application was denied) and also includes those who had no work authorization even if they were documented. Status 1 = "authorized" workers or documented workers; these workers must have a work authorization and may fall into any of the following status: having border crossing card/commuter card, with pending status, or temporary residents holding a non-immigrant visa. Status 2 = "permanent residents or green card holders" who have the right to reside and work in the U.S., and status 3 = "citizens" who are a citizen by birth or a naturalized foreign born citizen.

The variable *English* measures the capability to speak English, and does not include English reading skills. The variable is a discrete variable ranging from 1 to 4, where 1= not speaking English at all, 2 = speak a little English, 3 = somewhat able to speak English, and 4 = speaking good English.

Hispanic is a dummy variable for Hispanic which includes Mexican-American, Mexican, Chicano, Puerto Rican, and other Hispanic ethnic groups. *Black* (or African American) and *White* are also dummy variables derived from a question regarding their race which may also be American Indian/Alaka Native, Indigenous, Asian, Native Hawaiian or Pacific Islander, or others. *Age* was calculated from the difference between the date of interview and the date of birth, except for the questionnaire in the earlier years when age was asked directly.

Education is the highest grade level for education, and it ranges from 0 to 20. *Experience* is the number of years of doing farm work in the U.S. (not including farm work experience abroad). *Task* is the task at the time of interview. Although task is also asked for each period of work in the past two years, we use only the task at the time of interview for each duration. Although the original questions have over 100 task codes, tasks are grouped into six categories as follows: 1 = pre-harvest, 2 = harvest, 3 = post-harvest, 4 = semi-skilled, 5 = supervisor, and 6 = others. They are argued to be ordered by increasing skill requirements. *Seasonal Worker* is a dummy for workers who were working on a seasonal basis for the employer at the time of interview. *Free housing* is a dummy variable for workers (or workers and their family) who receive free housing from their current employer. It does not include those who own the house or live for free with friends or relatives. It also excludes those who pay for housing provided by employers or by the government or charity.

The dummies for *Florida* and *California* are the state for each work duration, and not necessarily the state at the time of interview. *Before 1993* dummy variable is for all the years prior to 1993 when the majority of IRCA legalization was granted, and *After 2001* is the years post-September 11, 2001 event.

Duration or farm work spells is a variable created from the work grid in the questionnaire. It is the difference between the ending dates and starting dates for each "farm work" spell, and only includes the completed spells (all spells completed at the time of interview).

Ordered Probit Model for Legal Status

Here we estimate the ordered probit model for legal status for foreign-born farm workers using NAWS data. Table 1 shows estimates for parameters and asymptotic standard errors (given in the parentheses) using 30912 observations of foreign-born farm workers. Using a 0.05 significance criterion, we find that all coefficients except education squared are statistically significant.

The third column of Table 1 shows the marginal effect of each variable on the probability of a worker being legal. The probability of worker *i* being legal is given by $\Pr ob(J_i^* > \mu_0) = 1 - \Phi(\mu_0 - x_i'\alpha)$. Then the marginal effect of variable *k* evaluated at the mean is $\phi(\mu_0 - \bar{x}'\alpha)\alpha_k$ for the continuous variables[6] and $\Phi(\mu_0 - \bar{x}'_{-k}\,\alpha_{-k}) - \Phi(\mu_0 - \bar{x}'_{-k}\,\alpha_{-k} - \alpha_k)$ for the dummy variables, where $\bar{x}'_{-k}$ and α_{-k} are variables and coefficients excluding *k*. Females, married, workers with higher English speaking ability, non-black, white, non-hispanic are statistically significantly more likely to have more advantageous legal status all else being the

[6] Marginal effect for variables with squared term is given by $\phi(\mu_0 - \bar{x}'\alpha)(\alpha_k + 2\alpha_{k\_sq}\bar{x}_k)$ where α_{k_sq} is coefficient for the squared variable. Also, we treated English speaking ability as a continuous variable.

same. We also find that both age and US farm experience have a significant nonlinear effect on legal status. US farm experience has positive effect on legal status up to thirty-five years. Age has positive effect on legal status up to eighty years. Education has a significantly positive linear effect on legal status. We find that the greatest positive marginal effect is from the female dummy followed by English speaking ability and before 1993 dummy. The greatest negative marginal effect is from the Hispanic dummy followed by the after 2001 dummy. Note that, holding all other characteristics the same, the workers interviewed before 1993 are eleven percent more likely and those interviewed after 2001 are fourteen percent less likely to be legal compared to those interviewed between these periods.

Finally, Table 2 shows actual-predicted legal status table. A worker is predicted to be status 0 (unauthorized) if $x_i'\hat{\alpha} < \hat{\mu}_0$, and is predicted to be status 1 (authorized) worker if $\hat{\mu}_0 < x_i'\hat{\alpha} < \hat{\mu}_1$ and so on. Table 2 shows that 80 percent of unauthorized workers are correctly predicted to be unauthorized. In the same way, 21 percent of authorized workers, 70 percent of permanent resident and 26 percent of citizens are correctly predicted in their legal status. Our ordered probit model does a very good job in distinguishing type 0 workers from legal workers, but many of type 1 workers and type 3 (citizen) workers are mistakenly predicted to be type 2 (permanent resident) workers.

Duration Model with Selection Bias Correction

Here we estimate the duration model with selection bias correction using the results from the ordered probit legal status model in the first stage. Table 3 shows estimates for parameters and asymptotic standard errors (given in the parentheses) for farm workers with each legal status. Status 0 (unauthorized) workers have 33,865 observations, status 1 (authorized) workers have

12,560 observations, status 2 (permanent resident) workers have 30,240 observations, and status 3 (citizen) workers have 18,307 observations. Based on asymptotic standard errors using a 0.05 significance criterion, the coefficients on the selectivity variable, λ, are all significant except for citizen workers. That is, using ordinary least squares without correcting for selectivity would lead to bias in all equations except for citizen workers. Actually, the selection bias correction term is set to zero for majority of citizen workers, because they are native born. So, the selection bias does not have a significant effect for this equation as it does for the other legal status equations.

Many variables have a statistically significant effect on duration in a common direction for all equations. Regardless of the legal status, workers in tasks requiring higher skill, non-seasonal workers, workers without free housing from employers, workers in California, workers in Florida, and workers interviewed after 2001 are statistically significantly more likely to have a longer duration farm job. Most of the signs of these coefficients are reasonable, except for the availability of free housing offered by the employer, which we expected to have a positive effect on duration. This may be because workers offered free housing are often migratory, seasonal workers with low skill and whose length of *contract* is generally short.

An interesting result is for English speaking ability. For unauthorized workers, higher English speaking ability is more likely to lengthen the duration in farm work. However, English speaking ability tends to shorten the duration in farm work for authorized and permanent resident workers. That is, legal workers leave agricultural work earlier as their English speaking ability improves, all else being the same. This variable does not have a significant effect on duration of citizen workers most of whom (77 percent) can speak English well, so that the variable has little variation.

Demographic variables tend to have various directions of influence on farm work duration for each legal status. Being female has a significantly positive effect on duration for authorized and citizen workers, while it has a significantly negative effect for permanent resident workers and no significant effect for unauthorized workers. Marriage has a significantly positive effect on duration for authorized and permanent resident workers, while it has a significantly negative effect for citizen workers. Permanent resident and citizen Hispanic workers tend to have shorter farm work duration than non-Hispanic workers, while unauthorized and authorized Hispanic workers tend to have longer farm work duration than non-Hispanic workers. Education has a significantly positive effect on the duration for all legal status, and experience has a significantly positive effect on the duration for unauthorized and authorized workers, but no significant effect on permanent resident or citizen workers. Age has a significant nonlinear effect on duration for all equations. The effect is positive up to an age of 87 years for unauthorized, and up to 105 years for citizen workers. On the other hand, the effect is negative through 39 years for authorized, and continuously for permanent resident workers (not turning positive until age 199 years).

Next, using estimates of each equation, we calculate the predicted durations of farm work by legal status by averaging the predictions over all observations for each equation (Table 4). The results indicate that the average predicted duration for unauthorized workers is not necessarily shorter than those for legal workers (authorized, permanent resident, or citizen). Actually, its average predicted duration is the second longest, and longer than for permanent resident and citizen workers.

Finally, we implement a simulation to test how farm work duration of a typical unauthorized worker would be expected to change with a change in legal status. This approach isolates the effect of legal status of the worker from differing characteristics of workers by holding the characteristics constant across different legal status. We fix each continuous variable at the mean of unauthorized worker observations, and fix each discrete variable at the category with the maximum number of observations of unauthorized workers, except for the "After 2001" dummy variable. Although observations after 2001 are approximately 24 percent of all unauthorized worker observations, the post-2001 period is more relevant for current policy purposes. The profile of the "typical" unauthorized worker is illustrated in Table 5.

The expected duration for this "typical" unauthorized worker is shown in Table 6 using the equation estimates for each legal status, conditionally upon being an unauthorized worker. The first row of Table 6 shows the typical unauthorized worker's expected duration under each legal status; the second row shows the percentage change from the unauthorized status. The result indicates that the duration of the "typical" unauthorized worker would be 4.4 percent longer if he were working as an authorized worker, and 3.9 percent larger if working as a permanent resident. Expected duration would decline by 4.7 percent were he to be working as a citizen, although this result is based on a statistically insignificant parameter estimate. In contrast to the results in Table 4, the Table 6 results hold worker characteristics constant across status whereas they vary across status in Table 4.

Setting aside the result for citizens that is based on a statistically insignificant estimate for the coefficient on the Mills ratio (λ), our estimated effect of a change in legal status from unauthorized to a legal status (either temporary authorization or permanent resident) is largely consistent with Tran and Perloff's result. In our case, expected duration is somewhat longer when working under a legal status; they report that "… IRCA increased the long-run probability that people granted amnesty stayed in agriculture." (p. 437) Hashida and Perloff's result is in the same direction, but larger. Emerson and Napasintuwong's result similarly suggested a longer duration for authorized rather than unauthorized workers. Their result referred to the number of years working in U.S. agriculture, rather than individual jobs as the above three analyses do. Their model did not directly address the sample selection issue, as the other three analyses do.

Conclusion

We have proposed and estimated a Heckman-type two stage model with legal status ordered probit model in the first stage and a duration model in the second stage. This methodology aims at overcoming two sources of inconsistency of farm work duration study: selection bias and the censoring problem. Our first methodology deals with the former problem adequately, but it takes only a rudimentary measure on the second problem: we have used only completed spells. Our current estimation result is based on this method.

The current estimation has significant coefficients on the selection bias correction term for all legal status equations except for that of citizen workers. That is, using ordinary least squares would lead to inconsistent estimates in all equations except for citizen workers. The most important finding from our estimation is that unauthorized workers do not necessarily have shorter farm work duration than legal workers. This is supported by two statistics. First, average predicted farm work duration for unauthorized workers is second longest. Second, the simulation analysis shows that the duration of the "typical" unauthorized worker will be longer when working under an authorized or permanent resident status.

Table 1. Orderd Probit Model for Legal Status for Foreign-Born Farm Workers

	Parameter Estimate	Marginal Effect
Female	0.463	0.162
	(0.019)	
Married	0.206	0.078
	(0.017)	
English Speaking	0.374	0.140
	(0.010)	
Black	-0.172	-0.066
	(0.077)	
White	0.151	0.056
	(0.015)	
Hispanic	-0.616	-0.206
	(0.048)	
Age	0.035	0.007
	(0.004)	
Age^2	-0.0002	
	(0.00005)	
Education	0.033	0.014
	(0.007)	
$Education^2$	0.0003	
	(0.0005)	
Experience	0.151	0.038
	(0.003)	
$Experience^2$	-0.002	
	(0.00006)	
Before 1993	0.313	0.113
	(0.018)	
After 2001	-0.356	-0.137
	(0.020)	
μ_0	2.487	
	(0.091)	
μ_1	2.912	
	(0.091)	
μ_2	5.074	
	(0.094)	

Table 2. Actual-Predicted Legal Status Table

	Predicted Legal Status				Total
Actual Legal Status	0	1	2	3	
0	80%	9%	11%	0%	100%
1	43%	21%	36%	0%	100%
2	14%	15%	70%	1%	100%
3	7%	7%	60%	26%	100%

Table 3. Duration Model for Farm Workers with Each Legal Status

	Unauthorized	Authorized	Permanent Resident	Citizen
Constant	3.482	4.016	3.808	2.955
	(0.027)	(0.041)	(0.035)	(0.037)
λ	0.084	0.047	-0.067	0.001
	(0.006)	(0.009)	(0.007)	(0.004)
Female	0.002	0.019	-0.025	0.157
	(0.007)	(0.008)	(0.007)	(0.008)
Married	-0.006	0.052	0.004	-0.050
	(0.006)	(0.006)	(0.006)	(0.006)
English Speaking	0.039	-0.026	-0.035	0.005
	(0.003)	(0.005)	(0.004)	(0.004)
Hispanic	-0.113	-0.189	0.048	0.086
	(0.010)	(0.012)	(0.010)	(0.011)
Age	0.016	-0.003	-0.0035	0.023
	(0.001)	(0.001)	(0.001)	(0.001)
Age^2	-0.0002	0.00007	0.00002	-0.0002
	(0.00002)	(0.00002)	(0.00002)	(0.00002)
Education	0.002	0.003	0.003	0.014
	(0.0009)	(0.001)	(0.0009)	(0.001)
Experience	0.006	0.004	-0.00002	-0.0005
	(0.0004)	(0.0008)	(0.0004)	(0.0005)
Task	0.044	0.033	0.050	0.075
	(0.002)	(0.002)	(0.002)	(0.002)
Seasonal Worker	-0.159	-0.143	-0.042	-0.083
	(0.006)	(0.006)	(0.006)	(0.006)
Free Housing	-0.082	-0.078	-0.146	-0.163
	(0.006)	(0.007)	(0.006)	(0.007)
California	0.256	0.073	0.228	0.145
	(0.006)	(0.006)	(0.006)	(0.006)
Florida	0.596	0.498	0.676	0.688
	(0.009)	(0.009)	(0.008)	(0.009)
After 2001	0.191	0.314	0.289	0.275
	(0.006)	(0.008)	(0.006)	(0.006)

Table 4. Average Predicted Duration for Each Legal Status (Days)

Unauthorized	56.2
Authorized	59.3
Permanent Resident	54.1
Citizen	53.6

Table 5. Profile of the "Typical" Unauthorized Worker

Constant	1
Female	0
Married	0
English Speaking	1.470
Hispanic	1
Age	28.201
Age^2	795.296
Education	6.073
Experience	5.075
Task	2
Seasonal Worker	1
Free Housing	0
California	1
Florida	0
After 2001	1

Table 6. Change in Duration for the "Typical" Unauthorized Worker

	Unauthorized	Authorized	Permanent Resident	Citizen
Expected Duration (Days)	62.6	65.3	65.0	59.6
Percent Change		4.4%	3.9%	-4.7%

References

Bourguignon, F., M. Fournier, and M. Gurgand. “Selection Bias Corrections Based on the Multinomial Logit Model: Monte-Carlo Comparisons,” DELTA Working Papers 2004-20, DELTA (Ecole normale supérieure) (September 2004).

Emerson, R. D., and O. Napasintuwong. “Foreign Workers in Southern Agriculture.” Selected paper, Southern Agricultural Economics Association Annual Meeting, Orlando, Florida (February 2002).

Greene, W. H. *Econometric Analysis*, 5th Edition.Upper Saddle River, N.J.: Prentice Hall, 2003.

Hashida, E., and J. M. Perloff. “Duration of Agricultural Employment.” Working Paper No. 779, Department of Agricultural and Resource Economics (CUDARE), University of California, Berkeley (February 1996).

Lee, L. F. “Generalized Econometric Models with Selectivity,” *Econometrica*, 51(2) (March 1983): 507-12.

Office of the Assistant Secretary for Policy, U.S. Department of Labor. “The National Agricultural Workers Survey: What is the National Agricultural Workers Survey (NAWS)?” (May 2005). Available at www.dol.gov/asp/programs/agworker/naws.htm

Tran, L. H., and J. M. Perloff. “Turnover in U.S. Agricultural Labor Markets.” *American Journal of Agricultural Economics.* 84 (2) (May 2002): 427-37.