IMMIGRATION POLICY AND THE AGRICULTURAL LABOR MARKET: SPECIALTY CROPS IN THE UNITED STATES

Nobuyuki Iwai
International Agricultural Trade and Policy Center
Food and Resource Economics Department
PO Box 110240
University of Florida
Gainesville, FL 32611
niwai@ufl.edu

Robert D. Emerson
International Agricultural Trade and Policy Center
Food and Resource Economics Department
PO Box 110240
University of Florida
Gainesville, FL 32611
remerson@ufl.edu

Orachos Napasintuwong
International Agricultural Trade and Policy Center
Food and Resource Economics Department
PO Box 110240
University of Florida
Gainesville, FL 32611
onapasi@ufl.edu

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Nobuyuki Iwai, Robert D. Emerson, and Orachos Napasintuwong
University of Florida

ABSTRACT

Mode 4 of the GATS addresses the temporary movement of natural persons. Since WTO negotiations have yet to significantly address unskilled labor mobility, our analysis is based on U.S. immigration policy and labor market rules with the presumption that it will be informative for forthcoming WTO negotiations. Of course, it is also directly relevant for U.S. immigration policy and labor markets. The paper examines the expected job length for farm workers under varying degrees of legal status. The National Agricultural Workers Survey (NAWS) found that 78 percent of agricultural workers in the U.S. were foreign-born for the years 2000-01. Of those, 68 percent were undocumented. Current immigration reform proposals vary in how they would alter the mix of workers under varying forms of work authorization. The concern for agricultural employers, and producers of highly labor-intensive specialty crops in particular, is the potential effect on the availability of labor for time sensitive activities. Our micro-econometric analysis of the NAWS data addresses the problem in two steps. First, an ordered probit model is estimated to characterize workers by legal status (unauthorized, authorized, permanent resident, or citizen), and then a duration model is estimated conditionally upon legal status of the worker. This approach permits asking the relevant "what if" question such as the expected job duration in agriculture for an unauthorized worker, were the worker to become legal as a result of legislative or policy changes. First, our findings are that job duration is affected by legal status and that the estimation adjustment via the ordered probit model is relevant. Second, we find that expected job duration would generally be longer for unauthorized workers if they were working in a legal status rather than in their unauthorized status.

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Introduction

WTO negotiations can potentially impact the terms of availability of labor for agriculture. Mode 4 of the General Agreement on Trade in Services specifically addresses the temporary movement of natural persons. While most existing commitments have addressed skilled labor, there is considerable potential for pressure from developing countries to liberalize less-skilled labor migration. Winters, et al., estimate that:

> … if quotas [on temporary inflows of workers] were increased by an amount equal to 3 per cent of developed countries' labour forces, there would be an increase in world welfare of $156 billion per year. Both developed and developing countries would share in these gains and they would owe more to greater mobility of less skilled workers than to that of more skilled workers. (p. v)

Clearly, these gains arise from economy-wide gains, not just agriculture. However, agriculture has a long history of employment of unskilled foreign labor, and the pattern has been towards the increased employment of foreign workers. This paper addresses the current situation for specialty crop agriculture, and focuses specifically on legal status of the worker and its relation to the expected duration of a person's job in specialty crop agriculture. Our emphasis is toward U.S. immigration policy and labor market rules rather than WTO negotiations since the former is what is currently in existence, but may be informative for forthcoming WTO negotiations.

Specialty crop agriculture, which faces intense competition from international markets, is a very labor-intensive industry. The nature of most specialty crops is that they are highly labor-intensive in contrast to most U.S. agricultural production. Labor expenditures account for

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37% of U.S. fruit, vegetable and horticultural crop production expenditures (*2002 Census of Agriculture)*. By contrast, labor expenditures for all of U.S. agriculture represent only 13% of production expenditures. Moreover, the labor requirements often vary considerably among the various specialty crops, and even within a crop there can be considerably different labor requirements depending on whether the crop is produced for the fresh or for the processing market (processing tomatoes are mechanically harvested while fresh market tomatoes are hand-harvested). Mechanical harvesting is the norm among major U.S. farm commodities such as corn, wheat, soybeans, and cotton, for example; however, a substantial portion of U.S. fruit and vegetable acreage remains totally dependent on hand harvesting (Sarig, et al.).

Florida, for example, is a major producing state for fruits, vegetables, and horticultural products. Labor expenditures represented 33% of Florida agricultural production expenditures across all commodities in the 2002 Census of Agriculture. Vegetable and melon producers devoted 40% of production expenditures to labor, and the fruits and tree nuts group allocated 38% of production expenditures to labor in 2002. These two commodity groups are major Florida commodities, and dominate agricultural employment in the state. The two commodity groups accounted for 46% of Florida agricultural labor expenditures in 2002: fruit and tree nut farms, 28%; and vegetables and melons, 18%. Both commodity groups are dominated by intense seasonal harvesting labor requirements.

Not only is the product side of the specialty crop market operating subject to strong international competitive forces, but the labor market is also an international labor market. The agricultural labor market is heavily dependent on foreign-born workers. According to the National Agricultural Workers Survey (NAWS) report (Carroll, et al., 2005), 78 percent of agricultural workers were foreign-born for the years 2000-01. In addition, most of the foreign-born workers were undocumented: 68 percent for the same period (Carroll, et al., 2005).[1] Therefore, 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.[2]

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). In addition, no duration study has focused on specialty crop agriculture whose production depends more on physical labor than other agricultural sectors. The previous studies which address the overall agricultural sector 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 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

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[1] Carroll, et al. (2005) note that 53 percent of all workers were unauthorized for work in the U.S. Since only foreign-born workers can be unauthorized, it follows that 68 percent of the foreign-born were unauthorized.
[2] See Tran and Perloff (427-28) for a detailed discussion of industry and legislative concerns.

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. According to the NAWS data, the portion of undocumented workers in specialty crop production rose from 16% for the years 1989-92, to 43% for the years 1996-98 and to 53% for the years 2002-2004. This implies that there has been a large-scale inflow of undocumented workers into the specialty crop agricultural labor market and a large-scale outflow of documented workers from it. The latter might mean that documented workers tend to leave specialty crop 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.[3] 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."[4]

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

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

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).[5] 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

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

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

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

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

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

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

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[6]

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

Note that we can use the result of the ordered probit model in the first stage for $\dfrac{\mu_j - x_i' \alpha}{\sigma_\varepsilon}$ and $\dfrac{\mu_{j-1} - x_i' \alpha}{\sigma_\varepsilon}$. 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.

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

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

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 incomplete. On the other hand, the legal status model (ordered probit) does not have any restrictions on the 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. In addition, we use only the data of laborers who worked in specialty crop agriculture. Specialty crops include all crops except corn, wheat, barley, oats, rice, rye, cotton/cottonseed, sugar beets, tobacco, soybeans, sugarcane, and multiple field crops.

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 statuses: 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 of speaking English, and does not include English reading skills. The variable is a discrete variable ranging from 1 to 4, where 1= not able to speak English at all, 2 = speaks a little English, 3 = somewhat able to speak English, and 4 = speaks English well.

Hispanic is a dummy variable including 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/Alaskan 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 earlier years when age was asked directly in the questionnaire.

Education is the highest grade level for education, and it ranges from 0 to 20. *Experience* is the number of years 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. The *Before 1993* dummy variable is for all years prior to 1993 when the majority of IRCA legalization was granted, and *After 2001* is the years after the 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 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 the estimates for parameters and asymptotic standard errors (given in the parentheses) using 29,194 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[7] 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, and non-hispanic are statistically significantly more likely to have a more advantageous legal status, all else being the same. We also find that both age and US farm experience have a significant nonlinear effect on legal status. US farm experience has a positive effect on legal status throughout a person's working life (up to 161 years), and age has positive effect on legal status up to 71 years, both of which are positive over the entire relevant range. 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 the 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 twelve percent more likely and those interviewed after

[7] The marginal effect for variables with a 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.

2001 are fourteen percent less likely to be legal, compared to those interviewed between these two periods.

Finally, Table 2 shows the 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 79 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 21 percent of citizens are correctly predicted in their legal status. Our ordered probit model does a very good job in distinguishing status 0 workers from legal workers, but many of status 1 workers and status 3 (citizen) workers are mistakenly predicted to be status 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 39,907 observations, status 1 (authorized) workers have 14,988 observations, status 2 (permanent resident) workers have 33,346 observations, and status 3 (citizen) workers have 15,012 observations. Based on asymptotic standard errors using a 0.05 significance criterion, the coefficients on the selectivity variable, λ, are all significant except for permanent resident workers. That is, using ordinary least squares without correcting for selectivity would lead to bias in all equations except for permanent resident workers.

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, 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 other status workers. That is, legal workers leave agricultural work earlier as their English speaking ability improves, all else being the same.

Demographic variables tend to have various directions of influence on farm work duration for each legal status. Being female has a significantly negative effect on duration for authorized workers, while it has no significant effect for workers in any other legal status. Marriage has a significantly positive effect on duration for authorized and permanent resident workers, while it has a significantly negative effect for unauthorized and citizen workers. Permanent resident Hispanic workers tend to have longer farm work duration than non-Hispanic workers, while unauthorized and authorized Hispanic workers tend to have shorter 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 negative effect on the duration for citizen workers, but a significantly positive effect for workers in any other status. Age has a significant nonlinear effect on duration for all equations. The effect is positive up to an age of 84 years for unauthorized, up to 69 years for permanent residents, and up to 105 years for citizen workers. On the other hand, the effect is negative through 42 years for authorized workers.

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 set of simulations to examine 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 varying legal status. In addition, we vary the time period (1989-1992, 1993-2001, and 2002-2004), the location (California, Florida, and the rest of the U.S.), and the task (harvest or pre-harvest). We fix each continuous variable at the mean of unauthorized worker observations, and fix each remaining discrete variable at the category with the maximum number of observations of unauthorized workers. The profile of the "typical" unauthorized worker is illustrated in Table 5.

The expected durations for this "typical" unauthorized worker are shown in Table 6 (and Appendix Figure 1) using the equation estimates for each legal status, conditionally upon being an unauthorized worker. For 14 out of 18 of the simulations, unauthorized workers working as authorized workers would have longer expected job durations. The largest effects were for unauthorized workers working under a permanent resident status – all were non-negative, varying from zero to 19 percent. For 13 out of 18 of the simulations, unauthorized workers working under a citizen status would have *shorter* expected job durations. Combining the authorized, permanent resident and citizen categories into a single legal category, unauthorized workers working under a legal status would have longer expected job durations for 14 of the 18 simulations. Noteworthy, is that all of the simulations for the 2002-04 period indicated quite large positive effects on expected job duration for an unauthorized worker working under a legal status.

A striking result of the simulations is that expected job duration is markedly longer for work in Florida than for either California or for the rest of the United States, regardless of legal status (Figure 1). Moreover, working under a legal status in Florida has a considerably larger effect on expected job duration than elsewhere (last column of Table 6, and Figure 1 for 2002-2004).

Figures 2 and 3 illustrate the effects of legal status by time period for California and Florida. The relative differences between the 1989-92 and 1993-2001 were minimal for both California and Florida. California unauthorized workers would have been expected to have a marginally *shorter* job duration under a legal status than as unauthorized workers. In both states, expected job duration would have been longer following 2001, and the effect of working under a legal status would have been stronger in Florida.

Figures 4 and 5 contrast pre-harvest and harvest workers after 2001 for California and Florida. There is very little difference in the expected job durations for pre-harvest compared to harvest workers in either location.

Our estimated effect of a change in legal status from unauthorized to a legal status is roughly 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 as an 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 focused on specialty crop agriculture: 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 permanent resident workers. That is, using ordinary least squares would lead to inconsistent estimates in all equations except for permanent resident workers. Although most of the effects of switching from an unauthorized status to a legal status result in an increase in expected job duration, there are some instances where expected job duration is shorter under a legal status than an unauthorized status. The most common occurrence of this was an unauthorized worker working under a citizen status. We demonstrate large, positive differences in expected job duration between Florida and other states. Expected job durations appear not only to have notably increased following 2001, but the effect of having a legal status is also more positive following 2001.

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

	Parameter Estimate	Marginal Effect
Female	0.465	0.178
	(0.019)	
Married	0.206	0.079
	(0.017)	
English Speaking	0.364	0.139
	(0.010)	
Black	-0.163	-0.062
	(0.078)	
White	0.140	0.054
	(0.016)	
Hispanic	-0.638	-0.244
	(0.049)	
Age	0.034	0.008
	(0.004)	
Age^2	-0.0002	
	(0.00005)	
Education	0.033	0.015
	(0.007)	
$Education^2$	0.0004	
	(0.0005)	
Experience	0.150	0.040
	(0.003)	
$Experience^2$	-0.002	
	(0.00006)	
Before 1993	0.317	0.121
	(0.018)	
After 2001	-0.353	-0.135
	(0.020)	
μ_0	2.430	
	(0.093)	
μ_1	2.870	
	(0.093)	
μ_2	5.009	
	(0.096)	

Table 2. Actual-Predicted Legal Status Table

	Predicted Legal Status				Total
Actual Legal Status	0	1	2	3	
0	79%	10%	11%	0%	100%
1	43%	21%	36%	0%	100%
2	14%	15%	70%	1%	100%
3	8%	8%	63%	21%	100%

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

	Unauthorized	Authorized	Permanent Resident	Citizen
Constant	3.367	3.816	3.479	2.999
	(0.026)	(0.037)	(0.034)	(0.038)
λ	0.117	0.038	0.003	-0.013
	(0.006)	(0.008)	(0.007)	(0.005)
Female	0.002	-0.018	-0.007	0.109
	(0.007)	(0.007)	(0.007)	(0.007)
Married	-0.014	0.035	0.012	-0.043
	(0.006)	(0.006)	(0.006)	(0.006)
English Speaking	0.047	-0.018	-0.008	-0.013
	(0.003)	(0.004)	(0.004)	(0.004)
Hispanic	-0.200	-0.179	0.090	-0.015
	(0.011)	(0.012)	(0.010)	(0.012)
Age	0.021	-0.003	0.003	0.025
	(0.001)	(0.001)	(0.001)	(0.001)
Age^2	-0.0003	0.00007	-0.00004	-0.0002
	(0.00002)	(0.00002)	(0.00002)	(0.00002)
Education	0.004	0.008	0.006	0.017
	(0.0009)	(0.001)	(0.0008)	(0.001)
Experience	0.009	0.004	0.002	-0.002
	(0.0004)	(0.0007)	(0.0005)	(0.0005)
Task	0.058	0.054	0.064	0.107
	(0.002)	(0.002)	(0.002)	(0.002)
Free Housing	-0.056	-0.057	-0.097	-0.183
	(0.007)	(0.007)	(0.006)	(0.007)
California	0.227	0.067	0.203	0.104
	(0.005)	(0.006)	(0.005)	(0.006)
Florida	0.567	0.536	0.663	0.647
	(0.008)	(0.008)	(0.008)	(0.009)
After 2001	0.130	0.343	0.201	0.234
	(0.006)	(0.007)	(0.006)	(0.007)

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

Unauthorized	58.0
Authorized	58.4
Permanent Resident	57.2
Citizen	56.1

Table 5. Profile of the "Typical" Unauthorized Worker

Constant	1
Female	0
Married	1
English Speaking	1.502
Hispanic	1
Black	0
White	0
Age	28.484
Age^2	811.338
Education	6.018
Experience	5.821
Free Housing	0

Table 6. Simulated Changes in Job Duration by Legal Status[a]

Time Period	Location	Task	Legal Status: Unauthorized	Legal[b]: Authorized	Legal[b]: Permanent Resident	Legal[b]: Citizen	Legal[b]: Legal[c]
1989-92	California	Harvest	52.8	48.5 (-8.1)	54.2 (2.5)	45.2 (-14.5)	51.9 (-1.7)
1989-92	Florida	Harvest	74.2	77.6 (4.6)	85.8 (15.6)	77.8 (4.8)	82.6 (11.3)
1989-92	Rest of US	Harvest	42.1	45.4 (7.8)	44.2 (5.0)	40.7 (-3.3)	44.6 (6.0)
1993-01	California	Harvest	53.9	49.1 (-8.9)	54.2 (0.5)	45.0 (-16.5)	51.9 (-3.8)
1993-01	Florida	Harvest	75.8	78.5 (3.7)	85.9 (13.3)	77.5 (2.3)	82.6 (9.0)
1993-01	Rest of US	Harvest	43.0	45.9 (6.8)	44.2 (2.9)	40.6 (-5.6)	45.0 (4.6)
2002-04	California	Harvest	62.7	70.1 (12.0)	66.3 (5.8)	56.6 (-9.6)	68.2 (8.8)
2002-04	Florida	Harvest	88.0	112.2 (27.4)	106.0 (19.3)	97.5 (10.8)	108.5 (23.3)
2002-04	Rest of US	Harvest	49.9	65.6 (31.3)	54.1 (8.3)	51.1 (2.2)	49.6 (19.3)
1989-92	California	Pre-Harvest	49.8	46.0 (-7.8)	50.8 (2.0)	40.6 (-18.5)	48.9 (-1.9)
1989-92	Florida	Pre-Harvest	70.0	73.5 (5.0)	80.5 (15.0)	69.9 (-0.2)	77.8 (11.0)
1989-92	Rest of US	Pre-Harvest	39.7	43.0 (8.2)	41.5 (4.4)	36.6 (-7.9)	42.0 (5.7)
1993-01	California	Pre-Harvest	50.9	46.5 (-8.6)	50.9 (0.0)	40.4 (-20.5)	48.9 (-3.9)
1993-01	Florida	Pre-Harvest	71.5	74.4 (4.1)	80.6 (12.7)	69.6 (-2.6)	77.8 (8.8)
1993-01	Rest of US	Pre-Harvest	40.6	43.5 (7.3)	41.5 (2.3)	36.5 (-10.1)	42.4 (4.4)
2002-04	California	Pre-Harvest	59.1	66.4 (12.4)	62.2 (5.3)	50.9 (-13.9)	64.3 (8.8)
2002-04	Florida	Pre-Harvest	83.0	106.2 (27.9)	98.6 (18.7)	87.6 (5.5)	102.3 (23.2)
2002-04	Rest of US	Pre-Harvest	47.1	62.1 (31.9)	50.8 (7.8)	45.9 (-2.7)	56.2 (19.2)

continued

Table 6. Simulated Changes in Job Duration by Legal Status, *continued*

Notes:

[a] All other worker characteristics are as in Table 5.

[b] Numbers in parentheses are percentage changes from expected duration in an unauthorized status.

[c] The *legal* category combines the three previous categories – authorized, permanent resident, and citizen. The calculation is

$$\frac{E[\ln(t_{.1}) \mid authorized]\Pr[authorized] + E[\ln(t_{.2}) \mid perm\,res]\Pr[perm\,res] + E[\ln(t_{.,3}) \mid citizen]\Pr[citizen]}{1 - \Pr[unauthorized]}$$

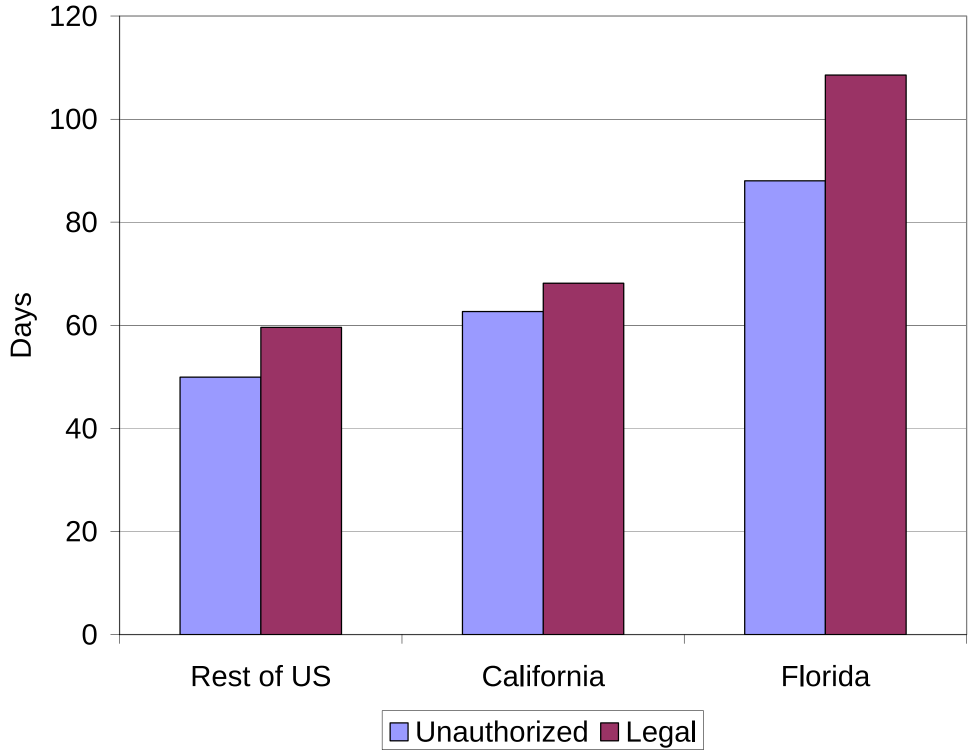

Figure 1. Expected harvest worker job duration after 2001

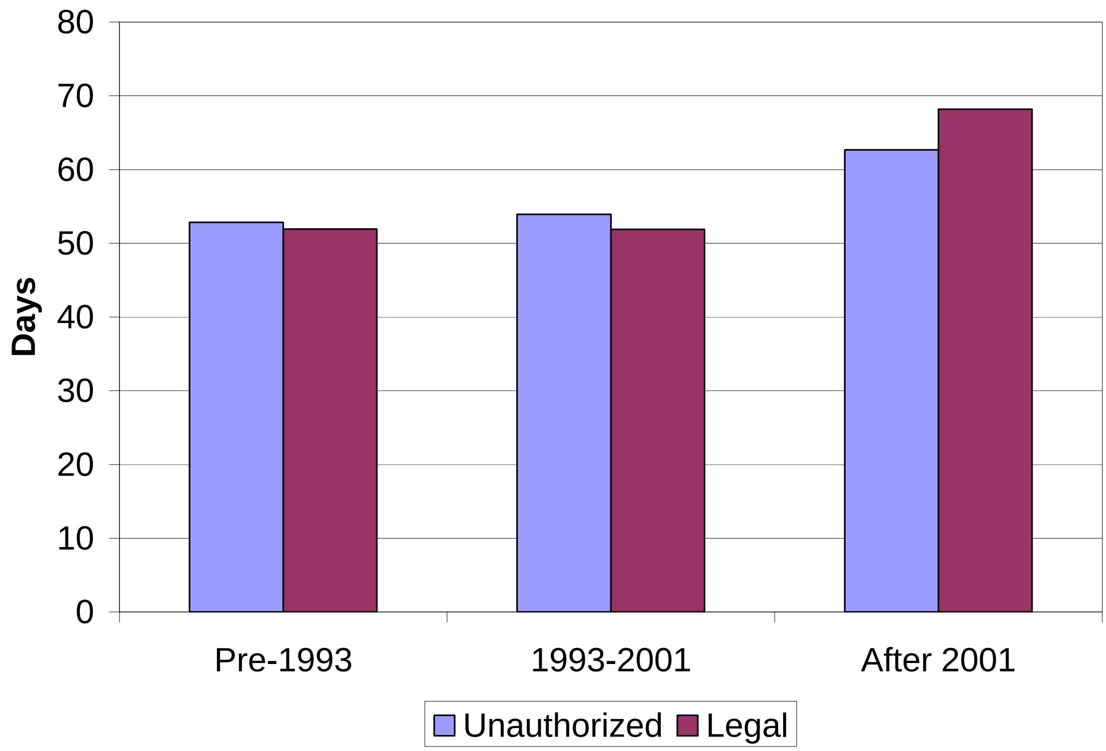

Figure 2. Expected harvest worker job duration: California

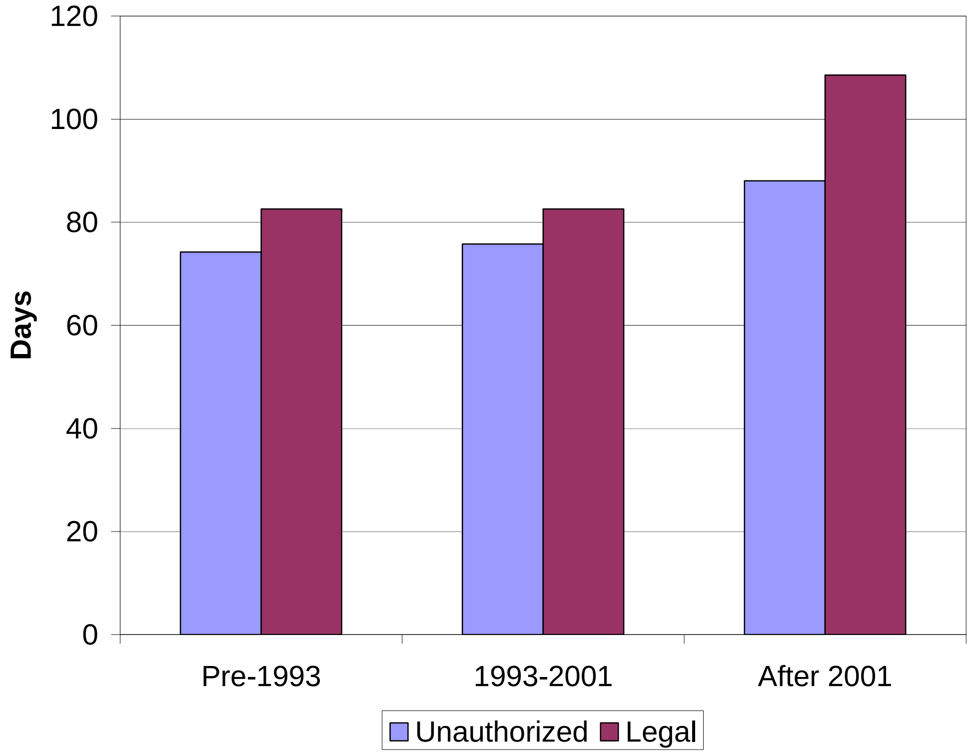

Figure 3. Expected harvest worker job duration: Florida

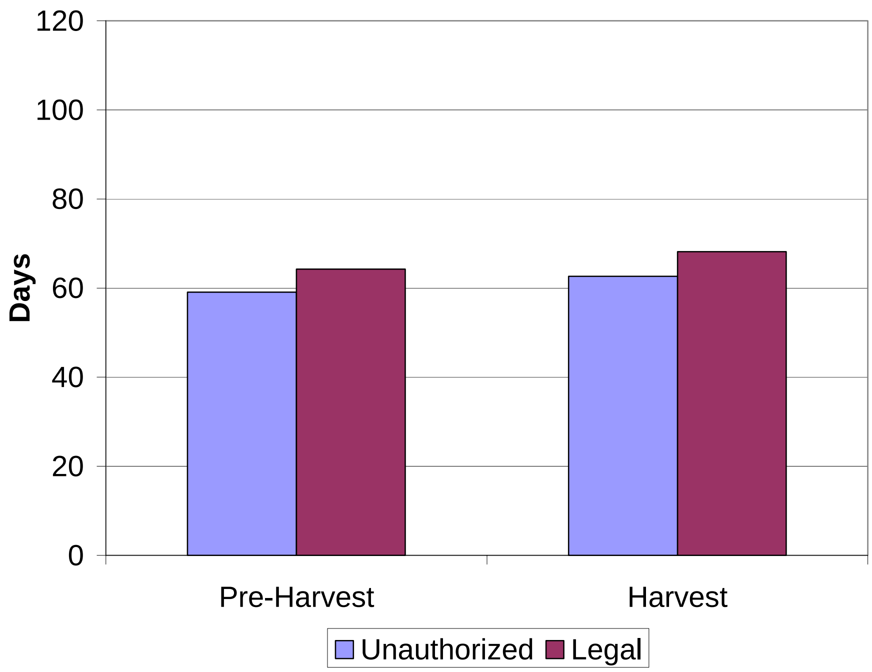

Figure 4. Expected farm worker job duration: California after 2001

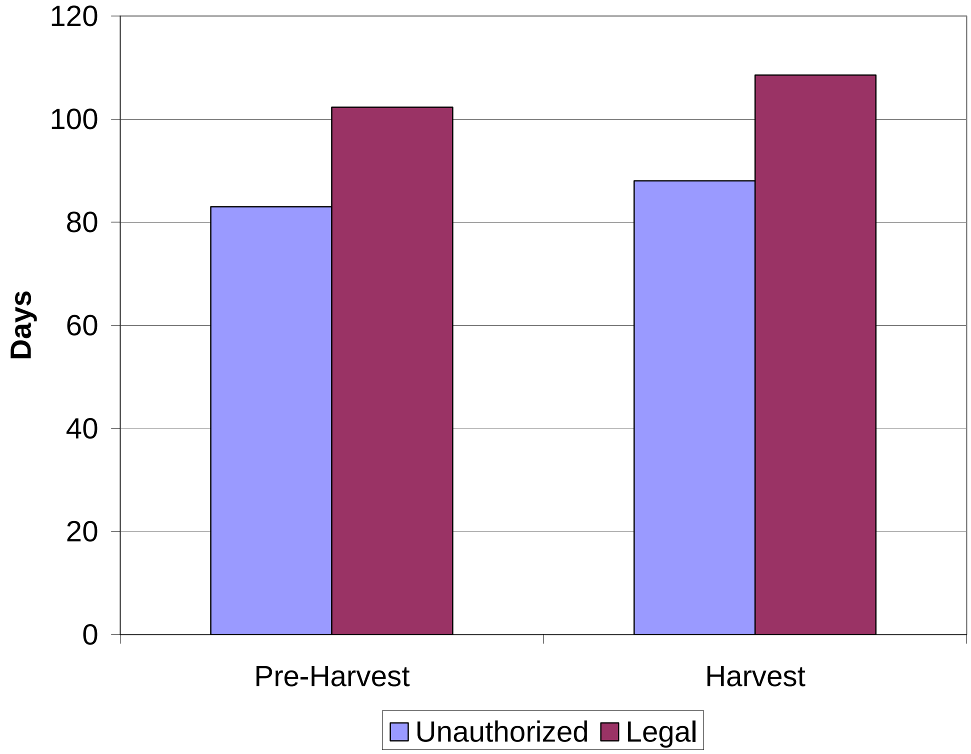

Figure 5. Expected farm worker job duration: Florida after 2001

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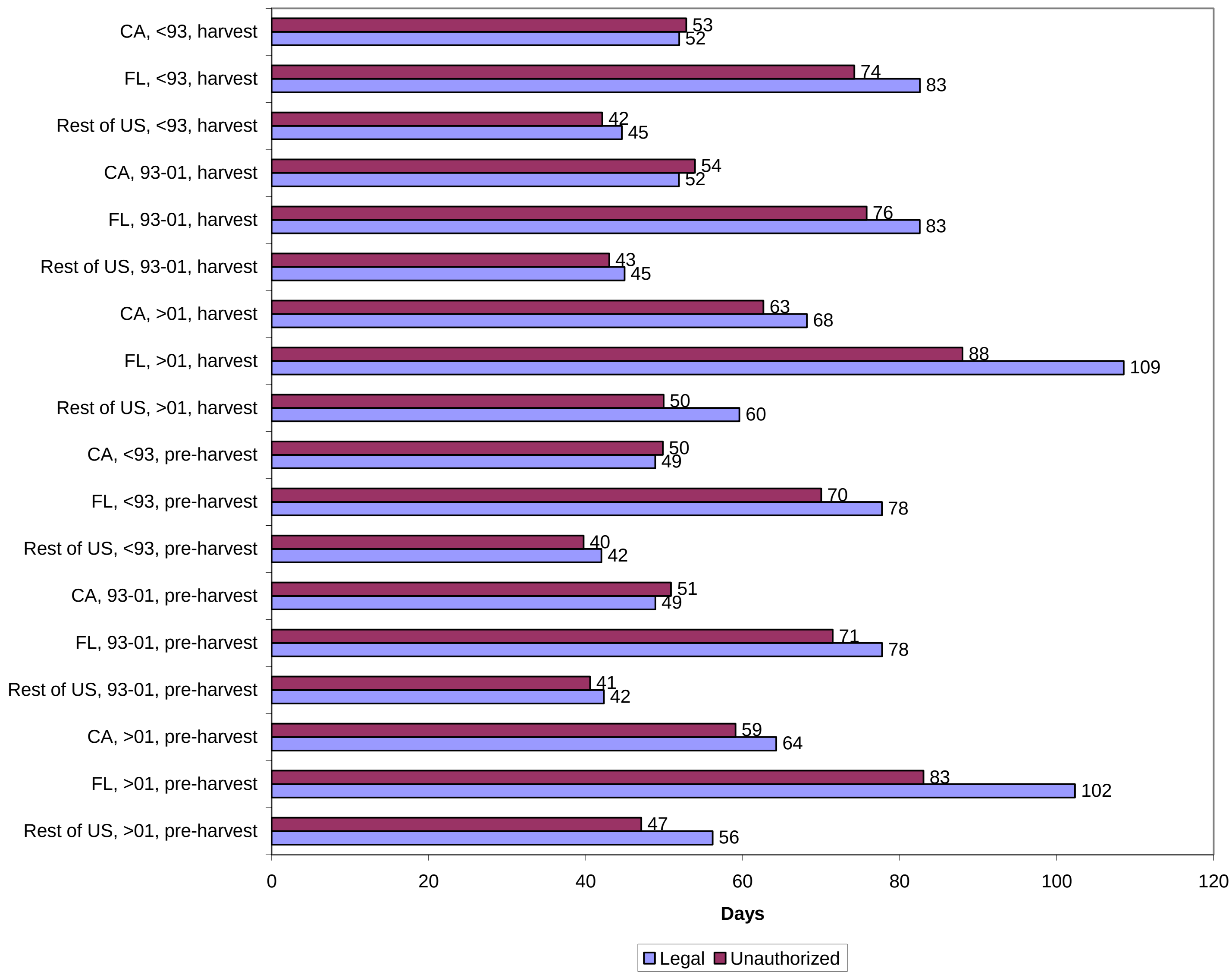

Appendix Figure 1. Farm worker job duration simulations