Migration of Seasonal Agricultural Workers

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Nearly half of all seasonal farm workers migrate at least 75 miles in a given year. An expected earnings differential from migration weakly induces migration: a 10% earnings differential raises the probability of migrating by slightly more than 1%. This result indicates that there are substantial costs to migrating and that employers must offer large earnings premia to induce a substantial number of workers to move to their jobs. Some demographic groups earn substantially higher earnings by migrating. These higher earnings from migration are primarily due to higher wages rather than more hours of work.

Key words: earnings, hours, migration, wages, workers.

Which hired agricultural workers are most likely to migrate? Are workers without legal status more or less likely to migrate than others? What are the returns to migrating? Do migrants' earnings rise due to higher wages, more hours, or both? To answer these questions, we estimate the impact of legal status, other worker attributes, and other factors on migration decisions using a "mover-stayer" model of migration, which takes into account possible systematic, unobserved differences in those who migrate and those who do not. This model is estimated using the U.S. Department of Labor's National Agricultural Workers Study (NAWS) cross-sectional, longitudinal data base.

With the passage of the 1986 Immigration Reform and Control Act (IRCA), many economists, farmers, and worker advocates questioned whether workers would be as willing to migrate to obtain farm work. IRCA extended legal status to many workers under an amnesty program and provided for the fining of employers who hire unauthorized workers (though enforcement of immigration rules did not change radically after passage of IRCA). If IRCA and subsequent government actions eventually restrict the supply of undocumented immigrant labor in the United States, many farmers and legislators fear large wage increases that will lead to significant crop losses (at least in the short-run) or to noncompliance with the law.

In the recent debates on immigration, many growers continued to contend that a guest worker program is necessary to avoid labor shortages, especially at harvest time. In March 1996, a new agricultural guest worker program that would bring as many as a quarter million foreign workers into the country to harvest crops was proposed but defeated.

As immigration, particularly undocumented immigration, becomes politically charged and more efforts are made to close the borders, a better understanding of seasonal workers is needed to make informed policy decisions about the future of agricultural labor supply and to examine the consequences of previous legal reform.

We hypothesize that workers who stood to gain the most from migrating are more likely to have migrated within the last year, all else the same. In our model, migration decisions depend on the pecuniary returns to migrating and on tastes, which we hypothesize are related to family relations and other personal characteristics. For example, workers who are unmarried or otherwise living separately from their families may be more likely to travel to short-term agricultural jobs. The major innovation in this model is that workers base their migration decisions on differences in both wages and hours at two locations. Policies that affect legal status and other factors have

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different impacts on wages and hours and hence more complex effects on migrations than have been previously studied.

Although several well-done studies of general migration exist, few previous studies have examined the migration of agricultural workers. Most of these general migration studies look at rural to urban migration, migration from country to country, or other similar "permanent moves." Itinerant workers or seasonal workers—as are common in agriculture—have a different choice as they move more frequently to find jobs that are available in only certain seasons. The availability of agricultural workers at the needed time determines whether a grower prospers.

Emerson, in the best existing study of agricultural workers' migration decisions within the United States, also used a mover-stayer model.[1] He found that workers with the greatest potential for higher earnings from migrating are more likely to migrate. Emerson's study covers only men who are legally authorized to work in this country. His data came from a 1970 survey of Florida farm workers that would have included few unauthorized immigrants.

Because we have a richer data base and a more general model, we are able to look at four issues that were not examined by Emerson. First, by using a national sample, we can examine geographic differences in hours, wages, and the probability of migrating. Because proximity to Mexico and other Latin American countries affects the supply of immigrant labor, geography matters. Moreover, these regional differences in migration may partially capture the effects of policy changes in the ease of border crossings.

Second, we include both males and females in our sample. As a result, we can examine the hypothesis (based on anecdotal evidence) that male workers are more likely to travel for short-term agricultural jobs than are females with similar characteristics. Women agricultural workers tend to be concentrated in certain crops and tasks, possibly because of their reluctance to migrate.

Third, we decompose differences in weekly earnings due to differences in workers' characteristics to look at the separate and combined effects of wage and weekly hours effects. It is possible that workers move to a new location with lower wages if they can work more hours a week. Previous studies considered only wage or only earnings effects, so this study is the first to examine the effects of differences in hours. By generalizing in this way, we can examine the effects of changes in overtime laws and other legal limits on hours.

Fourth, we examine the role of legal status on migration decisions. Isé and Perloff found that legal status affects earnings. Documented workers earn more per hour and work more hours per week. These results, however, may stem from differences in documented workers' willingness to migrate to a higher-wage region. Two offsetting hypotheses concerning legal status are proposed. On the one hand, unauthorized workers may be hesitant to migrate due to concerns that traveling increases their chance of apprehension and deportation. On the other hand, unauthorized workers may have fewer ties to a particular community and hence be more willing to migrate.

We start by discussing the model and the data. We present results for the migration model. Simulations based on the models and a discussion of the results follow. The final section contains conclusions and a discussion of policy implications.

Model

We hypothesize that workers choose whether to migrate by comparing the expected pecuniary and nonpecuniary costs and benefits of a move. If E_m is the amount a worker that migrates earns, and E_s is the amount the worker earns by staying at the worker's current location, then the monetary benefit from moving is $B = E_m - E_s$. If C is the psychic and monetary costs of moving, a worker benefits from migrating only if $B > C$. Thus, to determine which workers migrate, we need to estimate E_m, E_s, and C.

Unlike previous studies, we examine both components of earnings—wages and hours—so that we can separately capture the effects of various policy and other variables on earnings and migration. In our data set, a worker who traveled 75 or more miles to obtain a perishable-crop agricultural job within the previous year is said to have migrated. The logarithm of wages if the worker migrates, w_m, and if the worker stays put, w_s, are each functions of observed exogenous variables

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[1] Most of the other important articles on agricultural migration are by Taylor, who covers slightly different issues than we do. One of his papers is discussed below.

(**X**, which is a vector of individual demographic characteristics, geographic dummies, and policy variables) and unobserved individual differences, η_m and η_s, which we assume are normally distributed:

$$(1) \quad \ln w_m = \beta_m \mathbf{X} + \eta_m$$

$$(2) \quad \ln w_s = \beta_s \mathbf{X} + \eta_s.$$

Similarly, the logarithm of hours worked if a worker migrates, h_m, and the hours worked by a nonmigrant, h_s, are also functions of the observed exogenous variables, **X**, and unobservable difference, ν_m and ν_s, which we assume are normally distributed:

$$(3) \quad \ln h_m = \alpha_m \mathbf{X} + \nu_m$$

$$(4) \quad \ln h_s = \alpha_s \mathbf{X} + \nu_s.$$

The logarithm of weekly earnings is the sum of the logarithm of wages and the logarithm of weekly hours: $\ln E_i = \ln w_i + \ln h_i$, where $i = m$ or s. The expectation of a worker's relative monetary benefit or loss, B, from migrating is the difference in the expectations of the earnings from migrating and staying in place:

$$(5) \quad B = (\beta_m \mathbf{X} + \alpha_m \mathbf{X}) - (\beta_s \mathbf{X} + \alpha_s \mathbf{X}) = (\beta_m + \alpha_m - \beta_s - \alpha_s)\mathbf{X}.$$

The relative disutility or utility from migrating is not observed but is hypothesized to vary with a worker's observed characteristics, **Z**, and with the worker's unobserved characteristics, ε:

$$(6) \quad C = \gamma \mathbf{Z} + \varepsilon.$$

If the difference between the benefit of moving, B, and the utility or disutility (cost) effect, C, is positive, then benefits exceed costs and the worker migrates ($I = 1$); otherwise, the worker stays put ($I = 0$):

$$(7) \quad I = 1 \text{ if } B > C, \quad I = 0 \text{ if } B \leq C.$$

Assuming that the disturbance terms η_m, η_s, ν_m, ν_s, and ε are jointly normally distributed, equation (7) can be estimated using probit. By substituting for B and C in equation (7) using equations (1)–(6), we obtain a reduced-form probit equation:

$$(8) \quad I = 1 \text{ if } (\beta_m + \alpha_m)\mathbf{X} - (\beta_s + \alpha_s)\mathbf{X} - \gamma\mathbf{Z} - \varepsilon > 0,$$
$$I = 0 \text{ if } (\beta_m + \alpha_m)\mathbf{X} - (\beta_s + \alpha_s)\mathbf{X} - \gamma\mathbf{Z} - \varepsilon \leq 0.$$

Equation (8) can be estimated using a maximum-likelihood (ML) probit, where the exogenous right-hand-side variables are all those in **X** and **Z**.

This reduced-form specification allows us to predict migration conditional on various demographic characteristics. Alternatively, we can estimate a structural-form probit based on the difference between the benefits and costs of migration, equation (7).

To estimate a structural model, however, we need to construct a consistent estimate of B, the benefits of migrating. We cannot directly calculate B because earnings are only observed at the worker's original location if the worker did not migrate, and at the new location only if the worker did migrate. We overcome this problem by using estimates of the wage and hours equations to consistently estimate the worker's unobserved earnings.

First, we estimate the reduced-form probit to obtain a consistent estimate of equation (8). Second, we consistently estimate the log wage and log hours equations (1)–(4) using a technique introduced by Heckman that employs the consistent estimate of equation (8).[2] Third, using the resulting consistent estimates of the log wages and log hours equation parameters, we calculate $\hat{\beta} = (\hat{\beta}_m + \hat{\alpha}_m)\mathbf{X} - (\hat{\beta}_s + \hat{\alpha}_s)\mathbf{X}$, which is a consistent estimate of the expectation of the logarithmic earning differential from migrating. By substituting this estimate of B into equation (7), we can estimate the structural probit equation for the migration model.

Both **X** and **Z** include year, season, and geographic dummies, age, legal status, gender, language skills, ethnicity, and race. Because agricultural wages differ geographically (reflecting difference in labor demand and supply), regional dummies are included. Language skills may affect workers' job per-

[2] The methodology we use to avoid sample selection problems has been used to study union members' wages (Lee), agricultural workers' wages (Perloff), migration (Robinson and Tomes, Emerson), and many other topics. Were we to use ordinary least squares to estimate wage equations, (1) and (2), or hours equations, (3) and (4), the results may be inconsistent if workers' decision to migrate is systematic rather than random. For example, sample selection biases would be observed if unobserved characteristics, such as ambition or drive, affect both the probability of migrating and earnings. In such cases, η_m, η_s, ν_m, and ν_s are not normally distributed with a mean of zero if the equations are estimated using data only for workers who did or did not migrate.

formance (increase productivity), their knowledge of other job opportunities, and the cost of migrating and the fear of apprehension. Gender and ethnicity (and possibly age) may be proxies for discrimination, strength, or work ethic. Unauthorized workers may be hesitant to migrate for fear that they will increase their probability of being apprehended; on the other hand, they may have fewer ties to their current community. Taylor finds that unauthorized immigrants are less likely than other workers with similar characteristics to be observed in relatively high-skill, "primary" farm jobs that pay high wages. Employers may pay unauthorized workers lower wages to compensate for the fact that they might be fined if caught. In addition, these workers may be afraid that their employers will turn them over to immigration authorities if they quit, or they may be less aware of better jobs elsewhere than other workers.

Other predetermined and exogenous variables in **Z** that affect the migration decision are age, marital status (= 1 if married or living together), number of young children under fourteen years of age (to capture constraints on a worker's ability or willingness to move), and whether worker was born in Mexico. Place of birth may capture willingness to work hard at relatively unpleasant work and or being part of the local job-information network.

The structural migration equation also includes B, whereas the reduced-form migration equation replaces B with demographic variables that affect wages and hours (variables in **X** that are not in **Z**). Other predetermined and exogenous variables in **X** that affect the logarithmic wages (and hours) also include years of education and farm work experience. Experience should have a positive effect on hours and wages. We tested whether education and experience affect migration directly. The likelihood-ratio test statistic was 18.4 with 4 degrees of freedom. As a consequence, we included them in the structural probit.

To test hypotheses about the impact of gender, we included interactions between gender and many of the variables. For example, the number of children may affect a woman's cost of migrating differently than that of a man. Similarly, a woman may receive a different return on her investment in education than would a man.

Based on a likelihood-ratio test, we could not reject the hypothesis that the slope coefficients in equations for men and women are the same. As a result, the final equation includes a female dummy variable but no interaction terms with that female dummy.

There are several identifying restrictions: variables in **X** that are not in **Z** and vice versa. For example, family characteristics and place of birth are included in the migration equation but not in the wage and hours equations. We tested whether having children affects wages and hours and found that it does.[3] Thus, the number-of-children variable is included in the wage and hours equations.

Data

Our data are from the U.S. Department of Labor's National Agricultural Worker's Survey (NAWS). The NAWS is a national random sample of seasonal agricultural service workers (SAS), a group that includes most field workers in perishable crops and includes field tasks in fruit and vegetables, nursery crops, field crops, and cash grains. To ensure that a variety of types of workers and crops are covered, interviewers surveyed random samples of workers not previously interviewed three times a year in January, April and May, and October (Mines, Gabbard, and Boccalandro). The number of interviews within a cycle is in proportion to the amount of SAS activity at the time of year.

The NAWS uses site-area sampling to obtain a national representative cross section. First, seventy-three counties in twenty-five states from twelve distinct agricultural regions were selected. For each interviewing cycle, interviews were conducted in a subsample of thirty randomly selected counties using weights based on the size of the seasonal agricultural payroll in each county.

Employer names are obtained from the Bureau of Labor Statistics, the Agricultural Soil and Conservation Service, and Farm Labor Contractor Registration lists, as well as from other sources. NAWS regional coordinators contact randomly selected employers, explain the purpose of the survey, and obtain access to the work site to schedule worker interviews. Workers are then selected randomly and interviewed outside of work hours at the worker's home or at another location selected by the worker.

For the federal fiscal years 1989 through

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[3] The likelihood-test statistic for the two-wage equation is 110.9 with 2 degrees of freedom. The comparable statistic for the hours equation is 148.5.

1991, 4,718 people were interviewed. Our model is estimated using a subset, 3,343 hired farm workers, of this sample for which we have no missing variables.[4]

The means and standard deviations of the variables used in this study are reported in table 1. In the sample, nearly half (1,599 workers, or 48%) migrated by moving 75 or more miles for a new perishable-crop agricultural job at least once during the preceding year.

Over half (56%) of the sample reported their race as white. The nonwhite classification includes Asians, American Indians, and many Latinos who did not specify a race. The dominant ethnicity is Latino (87% of the entire sample).

We use four legal-status categories. In our regressions, the residual category is the set of workers who are not authorized to work in the United States (12% of the sample).[5] The other legal-status categories include U.S. citizens (17%), permanent residents ("green-card holders," 21%), and farm workers who were granted amnesty under IRCA (50%).[6]

Farmwork experience in the United States is measured in years. Education is the highest grade completed in the United States or abroad. English language skills are measured dichotomously: Either the worker does not speak English or the worker speaks at least some English. Another measure of language skills is whether a worker is a native English speaker. Wages are calculated as the hourly equivalent for both piece-rate and hourly workers. The wage is for the current job at the time of the interview. The weekly hours are for the week preceding the interview. Weekly hours worked ranged from ten to eighty-four, with a mean of forty-one hours a week in the field.

[4] We dropped workers from the sample who were missing any relevant variable, reported working fewer than ten hours (part-time workers) or more than eighty-five hours a week (which we believe to be implausible), were younger than fifteen, or said they worked in U.S. farm work for more than sixty-five years (the oldest worker was seventy-one).

[5] Legal status was self-reported; however, internal checks of consistency of responses about legal status were used to control for obviously false answers. Some workers may have reported that they had legal status because they possessed forged documents. Of course, to the degree that employers accepted those documents at face value, such workers are indistinguishable from workers with true legal status.

[6] Under the Immigration Reform and Control Act of 1986, seasonal agricultural workers who could establish they had worked for ninety days continuously in field work between 1 May 1985 and 1 May 1986 could obtain amnesty under the law if they applied by 30 November 1988. They received temporary work authorization status, then legal temporary resident status, and then legal permanent resident (green-card) status. This entire process took at least one year and normally two or more years. Aliens with temporary or permanent status could live and work anywhere and at any job within the United States.

Migration Model

The migration model is based on whether the worker actually migrated within the last year. We first discuss the reduced-form and structural probit equations. Then we report separate wage and hours equations for migrants and nonmigrants. Using likelihood-ratio tests, we rejected the hypothesis that the wage and hours equations are identical for migrants and nonmigrants.

Migration Probit Equations

We report parameter estimates for both reduced-form and structural probit equations in table 2. The structural equation includes the

Table 1. Means and Standard Deviations

	Migrated (1,599)	Did Not Migrate (1,744)
Female	0.147	0.292
Green card	0.181	0.236
Citizen	0.076	0.251
Amnesty	0.586	0.429
Latino	0.956	0.792
White	0.519	0.596
Nonnative English speaker	0.972	0.811
Speaks at least some English	0.154	0.336
Born in Mexico	0.858	0.623
Spouse	0.637	0.688
Northeast	0.052	0.029
Southeast	0.024	0.004
Midwest	0.020	0.081
Northwest	0.116	0.072
Southwest	0.008	0.019
Arizona	0.068	0.078
Texas	0.052	0.063
Florida	0.218	0.257
1990	0.502	0.490
1991	0.383	0.423
Winter	0.269	0.269
Spring	0.432	0.444
Number of children	1.460	1.403
	(1.77)	(1.58)
Age	32.120	34.261
	(11.35)	(11.97)
Education	5.534	6.321
	(3.60)	(4.00)
U.S. farmwork experience	9.043	11.221
	(8.27)	(9.27)
Wages	\$5.59	\$5.36
	(2.70)	(2.03)
Hours	41.03	40.75
	(13.45)	(12.41)
Earnings	\$227.76	\$218.19
	(129.56)	(102.86)

Table 2. Probit Equations

	Reduced Form		Structural	
	Coefficient	ASE	Coefficient	ASE
Constant	–0.514	0.28	–0.418	0.28
Female	–0.207*	0.06	–0.255*	0.06
Green card	–0.002	0.09	–0.130	0.09
Citizen	–0.207	0.13	0.317	0.14
Amnesty	0.119	0.08	–0.107	0.08
Latino	0.383*	0.10	0.050	0.12
White	–0.170*	0.05	–0.133*	0.05
Nonnative English speaker	0.303*	0.12	0.482*	0.14
Speaks at least some English	–0.070	0.07	–0.029	0.07
Born in Mexico	0.262*	0.08	0.396*	0.09
Children	–0.009	0.01	0.046*	0.02
Spouse	–0.023	0.03	–0.094	0.06
Age	0.004	0.01	–0.001	0.01
Age^2	–0.000	0.0002	0.000	0.0002
Education	–0.011	0.02	–0.019	0.02
$Education^2$	0.001	0.001	0.001	0.0015
U.S. farmwork experience	–0.042*	0.01	0.005	0.01
U.S. farmwork $experience^2$	0.001*	0.0002	–0.0003	0.0003
Northeast	0.863*	0.14		
Southeast	0.522	0.30		
Midwest	0.424*	0.21		
Northwest	0.191*	0.075		
Southwest	0.088	0.22		
Arizona	0.165*	0.075		
Texas	–0.041	0.10		
Florida	0.166*	0.06		
1990	–0.159*	0.076		
1991	–0.216*	0.09		
Winter	–0.014	0.07	0.061	0.06
Spring	–0.123*	0.067	0.072	0.06
Earnings differential			0.910*	0.13
Maddala R^2	0.16		0.15	
Cragg-Uhler R^2	0.21		0.20	
McFadden R^2	0.13		0.12	
Chow R^2	0.16		0.15	
Percentage correctly predicted	66%		65%	

		Predicted			
		0	1	0	1
Actual	No (0)	1,099	645	1,071	673
	Yes (1)	501	1,098	503	1,096

Note: An asterisk indicates that, on the basis of an asymptotic t-test, we reject the null-hypothesis that the coefficient is zero using a 0.05 criterion.

estimated difference of the log of earnings, $\hat{B}$, as a right-hand-side variable. This variable was constructed using the estimated wage and hours equations reported below.

Because of the nonlinearity of the probit equation, the size of an estimated coefficient does not directly show the effects of a change in the corresponding variable directly. We want to determine the effects of a change in a variable on the probability that a particular type of worker (who has a particular set of exogenous variables) migrates. We examine the effects for a "typical" worker, who is a thirty-three-year-old Latino nonwhite male born in Mexico, who is unauthorized to work in the United States. He speaks little or no English. The worker has no spouse, one child, and six years of education. He has ten years of U.S. farmwork experience and was interviewed in California in the fall of 1990. In the structural model, we use the average earnings differential for the sample and the "typical" worker characteristics.

The coefficient on the female dummy is negative and asymptotically statistically significant (we can reject the null hypothesis that this coefficient is zero) at the 0.05 level in both the reduced-form and structural equations. Were our typical worker a female instead of a male, the probability that she migrates would be 9% lower based on the structural probit equation and 8% lower based on the reduced-form equation.

None of the legal-status coefficients is statistically significantly different from zero at the 0.05 level. In both equations, whites are statistically significantly less likely to migrate (4% less likely according to the structural equation and 7% less likely according to the reduced-form equation) than other races.

According to the reduced-form estimates, a Latino is 15% more likely to migrate than a non-Latino. In the structural equation, the coefficient on the Latino dummy is not statistically significantly different than zero. That is, once we control for the expected earnings differential, Latinos are no more likely to migrate than others. According to both equations, those workers born in Mexico are more likely to migrate (14% more in the structural equation, 10% more in the reduced form).

U.S. farm work experience decreases the probability of migrating for the first twenty-four years according to the reduced-form equation. Experience does not have a statistically significant effect in the structural equation, after controlling for the expected earnings differential. Education does not have a statistically significant effect in either equation.

Emerson suggests that household composition may have offsetting effects on migration. Having dependent children increases the cost of migrating, as does the desire to keep the children in school. Yet children may also contribute to family earnings in migratory work activities, increasing the wage differential obtained from migrating. Our model shows that workers who are married or living together ("spouse" = 1) are not statistically significantly different from workers who are not cohabitating. The reduced-form equation does not show a statistically significant effect based on the number of children. According to the structural equation, a worker with more dependent children is statistically significantly more likely to migrate, but the effect is small: A 10% increase in the number of children leads to a 0.2% increase in the probability of migrating.

In the structural estimates, the expected earnings differential has the expected positive effect on migration (the same as Emerson observes) and is statistically significantly different from zero at the 0.05 level. According to our estimates, for every 10% increase in the expected earnings differential, the probability of migrating rises by 1.2%. That is, earnings differentials are an order of magnitude higher than the probability of migrating response. This result suggests that, in order to increase the number of workers willing to travel over 75 miles to find work, an employer must offer a substantially higher wage than the worker is currently receiving.

Migration and Wages

The consistently estimated log-wage equations for workers who migrated and those who remained in one location are shown in table 3. Based on asymptotic *t*-statistics, we can strongly reject the null hypothesis that the errors in these equations are orthogonal to the error in the reduced-form migration probit equation (the "no-sample selectivity" effect hypothesis). The asymptotic *t*-statistic on the correlation between the error terms in the two equations in the maximum-likelihood estimate, ρ, is over 29 for both equations, and estimated values of ρ are 0.98 and 0.91.

Gender, legal status, race, ethnicity, place of origin, and farmwork experience statistically significantly affect wages for migrants. Females earn 9% ($\approx e^{-0.096} - 1$) less than comparable males. Amnesty workers earn 10% more than unauthorized workers. Whites earn 8% less than nonwhites. Latinos earn 10% more than others. Workers born in Mexico receive a 15% wage premium. Surprisingly, wages fall with farm work experience for the first eighteen years and then rise.[7]

For nonmigrants, legal status, ethnicity, education, and experience statistically significantly affect wages. Unlike migrants, nonmi-

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[7] A referee suggests that the initial fall in wages with experience is due to a "matching" or "sorting" process where "good" workers move out of migrant work and "bad" workers continue to migrate. Thus, anyone who is still a migrant after several years may be a "bad" worker and earns a low wage.

Table 3. Log Wage Equations

	Migrated		Did Not Migrate	
	Coefficient	ASE	Coefficient	ASE
Constant	1.049*	0.13	1.204*	0.07
Female	−0.096*	0.03	0.017	0.02
Green card	0.037	0.04	0.113*	0.03
Citizen	−0.038	0.06	0.162*	0.04
Amnesty	0.098*	0.04	0.056*	0.025
Latino	0.098*	0.05	−0.062*	0.03
White	−0.084*	0.02	−0.006	0.02
Nonnative English speaker	0.067	0.06	−0.039	0.03
Speaks at least some English	−0.055	0.03	0.019	0.02
Children	−0.009	0.007	0.010*	0.004
Born in Mexico	0.141*	0.04	−0.002	0.02
Age	0.002	0.006	−0.003	0.003
Age^2	−0.0001	0.0001	0.000	0.00
Education	0.0007	0.01	0.019*	0.006
$Education^2$	0.0006	0.0007	−0.001*	0.0004
U.S. farmwork experience	−0.011*	0.004	0.032*	0.002
U.S. farmwork $experience^2$	0.0003*	0.0001	−0.0007*	0.0001
Northeast	0.302	0.05	−0.267*	0.04
Southeast	0.074	0.12	−0.414*	0.15
Midwest	0.106	0.10	−0.147*	0.04
Northwest	0.105*	0.03	−0.012	0.02
Southwest	−0.077	0.13	−0.147*	0.04
Arizona	0.032	0.03	−0.146*	0.02
Texas	−0.210*	0.05	−0.238*	0.03
Florida	0.005	0.03	−0.115*	0.02
1990	0.010	0.04	0.036	0.02
1991	−0.022	0.04	0.098*	0.03
Winter	0.030	0.03	−0.041*	0.02
Spring	0.017	0.03	−0.001	0.02
ρ	0.98*	0.002	0.91*	0.01
R^2	0.10		0.15	

Note: An asterisk indicates that, on the basis of an asymptotic t-test, we reject the null-hypothesis that the coefficient is zero using a 0.05 criterion.

grants do not have a gender wage differential. Citizens earn 18% more than comparable unauthorized, settled workers. Latinos earn 6% less than other settled workers—in stark contrast to the positive differential for migrants. Wages rise with education through eighth grade and then fall. Wages rise with U.S. farm work experience for twenty-three years and then fall.

Migration and Hours

In table 4, we reject the null hypothesis of no sample selection in the migrant hours equation ($\hat{\rho}$ = 0.49 and the asymptotic *t*-statistic = 2.28), but not in the hours equation for settled workers ($\hat{\rho}$ = 0.07 and the asymptotic *t*-statistic = 0.13). For migrants, hours vary statistically significantly with legal status, race, ethnicity, and gender. Some of these effects are relatively large. For example, a female works 19% ($\approx e^{-0.21} - 1$) fewer hours per week than a male worker with otherwise identical characteristics. A citizen works 12% fewer hours than an unauthorized worker.

For settled workers, gender and age have statistically significant effects on hours worked. Females work 14% fewer hours per week than comparable males. Hours worked per week rise with age, though at a declining rate.

Simulations

How great are the differences between those who migrate and those who do not? To answer this question, we calculate the expected

Table 4. Log Hours Equations

	Migrated		Did Not Migrate	
	Coefficent	ASE	Coefficient	ASE
Constant	3.365*	0.17	3.450*	0.16
Female	−0.206*	0.04	−0.154*	0.04
Green card	−0.028	0.04	0.035	0.04
Citizen	−0.125*	0.06	0.080	0.05
Amnesty	0.009	0.03	0.036	0.03
Latino	0.161*	0.07	0.011	0.05
White	−0.074*	0.03	−0.025	0.03
Nonnative English speaker	−0.003	0.08	−0.053	0.06
Speaks at least some English	−0.002	0.04	−0.018	0.03
Children	0.004	0.01	0.007	0.006
Born in Mexico	−0.005	0.06	0.004	0.07
Age	0.008	0.01	0.011*	0.005
Age^2	−0.0001	0.0001	−0.0001*	0.0001
Education	−0.0006	0.01	−0.009	0.008
Education^2	−0.0001	0.001	0.0006	0.001
U.S. farmwork experience	−0.007	0.006	−0.002	0.01
U.S. farmwork experience^2	0.0001	0.0001	0.0001	0.0001
Northeast	−0.018	0.07	−0.287*	0.07
Southeast	0.159	0.09	0.113	0.17
Midwest	−0.100	0.10	−0.017	0.05
Northwest	−0.037	0.04	−0.011	0.04
Southwest	−0.183*	0.09	0.113	0.07
Arizona	−0.007	0.05	−0.006	0.04
Texas	−0.173*	0.04	−0.150*	0.04
Florida	−0.192*	0.03	−0.002	0.03
1990	0.029	0.04	0.019	0.04
1991	−0.004	0.05	0.045	0.05
Winter	−0.017	0.04	0.045	0.03
Spring	0.044	0.03	0.070*	0.03
ρ	0.487*	0.21	0.065	0.49
R^2	0.07		0.07	

Note: An asterisk indicates that, on the basis of an asymptotic t-test, we reject the null-hypothesis that the coefficient is zero using a 0.05 criterion.

wages, hours, and earnings for various demographic groups based on our point estimates, which control for selectivity bias.[8]

Tables 5 and 6 show our simulation results. In both tables, the first row shows the simulations for our typical worker. Workers in the other rows have the same characteristics as the typical worker except for the characteristics in the first column.

For example, the typical worker has a 49.4% probability of migrating (first row, first column of table 5). A similar worker who is a citizen, however, has only a 41.2% probability of migrating (third row, first column).

The typical worker who migrates earns $264 a week, whereas one who stays put earns only $212 (first row, second and third columns of table 5). Thus, the typical worker earns 19.7% [= (264 − 212)/264] more by migrating, as shown in table 6 (first row, first column).

Why does this worker earn more by migrating? As table 6 shows, 13.4% of this increase in earnings is due to an increase in the wage, while only 6.2% is due to extra hours of work.[9]

Not all workers gain by migrating, however. For example, citizens with these characteristics are expected to earn less if they migrate. The returns to migrating are larger for younger,

[8] These simulations should be viewed with some caution as the underlying estimated equations explain only a fraction of the variations in wages, hours, and probability of migrating.

[9] The earnings differential, ΔE, approximately equals $H\Delta w + w\Delta H \equiv \Delta\tilde{E}$. The change in earnings due to wages reported in the table is $H\Delta w$ times an adjustment factor, $\Delta E/\Delta\tilde{E}$, which insures that the change due to wages plus the change due to hours add to the total earnings differential.

Table 5. Probability of Migrating, Earnings, and Wages

	Probability of Migrating	Weekly Earnings		Wage	
		Migrate	Stay	Migrate	Stay
Typical[a]	49.4	$264	$212	$5.52	$4.74
Green card	49.3	266	246	5.73	5.31
Citizen	41.2	224	270	5.31	5.57
Amnesty	54.1	293	232	6.09	5.02
Age = 25, experience = 5	55.2	278	183	5.80	4.22
Age = 45, experience = 25	42.3	250	243	5.35	5.42
Female	41.2	195	185	5.01	4.82

[a] Male Hispanic, unauthorized, speaks little or no English, born in Mexico, has no spouse and one child, thirty-three years old, six years of formal education, ten years of U.S. farmwork experience. He was interviewed in California in the fall of 1990.

Table 6. Percentage Change in Earnings from Migrating

	Change in Earnings	Due to Changes in	
		Wages	Hours
Typical[a]	19.7%	13.4%	6.2%
Green card	7.6	7.2	0.4
Citizen	–20.5	–5.1	–15.4
Amnesty	20.9	17.0	3.8
Age = 25, experience = 5	33.9	25.3	8.6
Age = 45, experience = 25	2.7	–1.4	4.1
Female	65.2	42.0	23.3

[a] Male Hispanic, unauthorized, speaks little or no English, born in Mexico, has no spouse and one child, thirty-three years old, six years of formal education, ten years of U.S. farmwork experience. He was interviewed in the California in the fall of 1990.

less experienced workers and females.

According to one view, unauthorized workers are less likely to migrate within the United States for fear of being caught and sent home. This prediction, however, is false. Unauthorized workers are as likely or more likely to migrate (table 5) than are permanent residents and citizens, though they are less likely to migrate than workers with amnesty. Apparently, the gains to migrating outweigh unauthorized workers' fear of detection. Unauthorized workers gain more from migrating than do citizens and green-card holders (table 6). Workers with amnesty have slightly higher returns to migrating than do unauthorized workers.

Discussion of Results

Nearly half (48%) of all seasonal farm workers migrate at least 75 miles in a given year. The decision to migrate varies substantially with workers' demographic characteristics.

Although our study uses more recent data and extends in several ways Emerson's study of authorized male workers in Florida, our results are consistent with his major finding that an expected earnings differential from migration induces migration. We find, however, that this effect is not very strong: A 10% earnings differential raises the probability of migrating by only slightly more than 1%. This result indicates that there are substantial costs to migrating and that employers must offer large earnings premia to induce a substantial number of workers to move to their job.

As our simulations show, some demographic groups earn substantially higher wages by migrating. For example, a typical unauthorized Mexican male Hispanic earns 20% more by migrating, while a comparable female earns 65% more. Some groups, particularly citizens, earn less by migrating and hence are unlikely to migrate. One possible explanation for this result is that relatively many citizens have longer-term jobs and would lose the associated long-term-employment earnings premium by migrating.

By using national data, we show that workers' hours, wages, and probability of migrating differ across the country. By using data for both men and women, we show there are pronounced gender differences. Strikingly, there is a sizeable gender wage differential for migrants but not for nonmigrants. Both migrants and nonmigrants have a large gender-hours differential. As a consequence, for both groups, men earn more than women. We find that women are less likely to migrate than are men, even though the rewards of migrating are larger.

By decomposing earnings into wages and hours, we show that the earnings increases due to migration are primarily due to wage differentials. Indeed, the actual number of hours worked per week by migrants was less than that of nonmigrants.

Since 1970, when Emerson's sample was interviewed, the share of immigrants in the farm labor force has increased substantially. According to the NAWS, in 1988, soon after IRCA went into effect and many workers received amnesty, unauthorized workers made up a trivial percentage of the hired-agricultural work force. Today, however, nearly a third of these workers are unauthorized.

Our examination of the impact of legal status on migration decisions shows that legal status plays an important role in migration decisions. We find that unauthorized workers are not only more willing to migrate but a greater proportion actually have migrated than authorized workers. Thus, fear that travel will increase the probability of being apprehended and deported cannot be the major concern of unauthorized workers when making their migration decisions.

Given this result, if federal efforts to prevent illegal immigration become more effective, farmers will have to offer increasingly large wage premia to induce migration to harvest crops. Farmers may see this result as supporting their argument for a guest worker program. Farmworker advocates, in contrast, may conclude from this result that offering higher wages can avoid crises at harvest time and raise the earnings of legal workers.

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