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AUTHOR	Perloff, Jeffrey M.	
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ABSTRACT

This report examines the wage differential necessary to induce nonagricultural workers to work in agriculture. The responsiveness of labor supply to wage changes is important because the Immigration Reform and Control Act may reduce the supply of immigrant farm labor in the United States. A random sample of 931 males with a ninth grade education or less, living outside of major metropolitan areas, was selected from the Bureau of Labor Statistics Current Population Survey. Nineteen percent were hired agricultural workers. Subjects are described in terms of region, state, ethnicity, marital status, job characteristics, number of children, years of schooling (tables 2 and 3), years of experience, earnings, and weekly hours of work. Compared to nonagricultural workers, agricultural workers averaged 1 year less of education and 6 years less of work experience, and they were seven times as likely to be Mexican and about twice as likely to be Mexican-American or "other Hispanic." An empirical model was used to estimate the change in the number of agricultural workers given an increase in the agricultural wage. Results indicate that inducing nonagricultural workers to switch to agriculture may not be as costly as some have suggested. A 10 percent increase in wages may increase the numbers in agriculture of rural male workers with a ninth grade education by nearly 25 percent. However, in California, and especially in certain crops, a large percentage of the agricultural work force has traditionally been undocumented workers, so larger wage increases may be required to produce the same effects. (LP)

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The Effect of Wage Differentials on Choosing to Work in Agriculture

CALIFORNIA AGRICULTURAL STUDIES 90-4

Employment Development Department

The Effect of Wage Differentials on Choosing to Work in Agriculture: Implications for the Immigration Control and Reform Act

Jeffrey M. Perloff

University of California, Berkeley

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Executive Summary

This study estimates the likelihood of nonagricultural workers joining the agricultural work force in response to an increase in the agricultural wage. The responsiveness of labor supply to wage changes is important because the Immigration Reform and Control Act of 1986 (IRCA) is supposed to restrict the supply of ineligible immigrant labor in the United States. Many farmers and legislators predicted that if IRCA significantly reduces the immigrant agricultural work force, then large wage increases, significant crop losses (at least in the short-run), or mass noncompliance with the law will result. This study is designed to help assess how realistic are these fears that large wage adjustments will be required to equilibrate the hired farm worker labor market.

The study uses 1988 data from the U.S. Department of Labor, Bureau of Labor Statistics' (BLS) Current Population Survey (CPS). It simultaneously determines choice of sector (agriculture and nonagriculture) and wages in each sector. An empirical model is used to estimate the change in the number of agricultural workers given an increase in the agricultural wage.

Results indicate that inducing non-agricultural workers to switch to agriculture may not be as costly as some have suggested. A ten percent increase in wages may increase the share in agriculture of rural male workers with no more than a ninth grade education by nearly a quarter. Nonetheless, in California, and especially in certain crops, a large percent of the agricultural work force has traditionally been undocumented workers, so larger wage increases may be required to produce the same effects.

The Effect of Wage Differentials on Choosing to Work in Agriculture: Implications for the Immigration Control and Reform Act

This study estimates the likelihood of nonagricultural workers joining the agricultural work force in response to an increase in the agricultural wage. The responsiveness of supply to wage changes is important because the Immigration Reform and Control Act of 1986 (IRCA) is supposed to restrict the supply of ineligible immigrant labor in the United States.[1] Many farmers and legislators predicted, if IRCA significantly reduces the immigrant agricultural work force, large wage increases, significant crop losses (at least in the short-run), or mass noncompliance with the law. This study is designed to help assess how realistic are these fears that large wage adjustments will be required to equilibrate the hired farm worker labor market.

The study uses 1988 data from the U. S. Department of Labor, Bureau of Labor Statistics' (BLS) Current Population Survey (CPS). The study simultaneously determines choice of sector (agriculture and nonagriculture) and wages in each sector, controlling for choice of sector. Then the empirically estimated model is used to simulate the increase in the share of agricultural workers from a given increase in the relative agricultural wage.

The first section below describes the data set and presents summary statistics for the key variables. The second section develops the basic modeling methodology. Empirical results are discussed in the third section. The next section simulates the likely response of workers to higher wages in the agricultural sector. The final section draws inferences and conclusions from the analyses.

The Data

This study uses the entire 1988 calendar year data from the Current Population Survey (CPS) of the Bureau of Labor Statistics (BLS).[2] The BLS attempts to obtain a representative sample of the entire U. S. work force by randomly selecting household locations.[3] To prevent double counting particular individuals, only one interview with each worker (the fourth-month interview) is used in this study.

Because preliminary experiments indicated that the wage equations for males and females differ significantly, gender-specific analyses were used initially. Unfortunately, there are not enough women in the sample to estimate an agricultural wage equation for females. As a result, the following analyses are based on a data set containing only males.

1 Further, undocumented farm workers who gain legal residence status under IRCA are entitled (and many are expected) to leave agriculture.

2 In the following analyses, only data for subjects not missing any of the key variables are used.

3 Some critics have argued that the BLS undersamples migrant and illegal agricultural workers. Although this complaint is valid, this study uses the CPS data set for two reasons. First the CPS is the only (nearly) random sample that includes both agricultural and nonagricultural workers in sufficient quantities to conduct such a study. Second, we are primarily concerned with the industry choices of workers with legal status.

The CPS data set was further restricted to only a subset of individuals who were relatively likely to consider agricultural employment. It would be a nonproductive exercise to calculate the wage differential necessary to induce a brain surgeon or other skilled worker to start working as an agricultural field hand. For this reason, the data set was restricted to only those workers with no more than a ninth grade education.[4] Similarly, it is unlikely that workers in the middle of Manhattan are likely to switch to farm work with any plausible agricultural wage increase. Because we are concerned with the short-run effect of a wage differential on choice of working in agriculture, the sample has been further restricted to those workers who live outside of major metropolitan areas.[5]

Table 1 presents the means and standard deviations (for continuous variables) of several key variables for our sample of 931 men, of which, 19.4 percent are hired agricultural workers.[6] This table shows that more than a quarter (26.2 percent) of our sample live in the South Atlantic region; whereas, only slightly more than a fifth (21.0 percent) of the agricultural workers live in that region. Similarly, only 7.7 percent of the sample lives in the Pacific region; whereas, over a quarter (27.5 percent) of the agricultural workers live there. Moreover, only 5.7 percent of the total sample, but 27.1 percent of the agricultural workers live in California.

There are relatively more blacks and other nonwhites in the agricultural subsample than in the overall sample. Agricultural workers are less likely to be married and living with their spouses than nonagricultural workers (52.5 versus 77.9 percent), perhaps reflecting the migratory nature of many agricultural jobs and the relative youth of agricultural workers.

On average, nonagricultural workers are 7 years older than agricultural workers.[7] Although nearly a fifth of the nonagricultural workers are union members (17.9 percent), only 1.7 percent of the agricultural workers are. Nonagricultural workers are more likely than agricultural workers to be paid by the hour (78.5 percent versus 64.6 percent).

Even in this sample that excluded workers with more than nine years of education, agricultural workers average one fewer year of school than nonagricultural workers.

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4 While it is true that there are some highly educated farm workers, the majority of agricultural workers have less than a high school education. The choice of a ninth grade cut-off was arbitrary. Experiments suggest that our results are not sensitive to this decision. Qualitative results are relatively unchanged if the cut-off point is shifted to eighth grade, tenth grade eleventh grade, or the twelfth grade (short of a diploma).

One could argue whether restricting our sample to relatively uneducated workers creates a sample selection bias; yet there are several reasons why we need not worry about such bias. First, we are primarily interested in the short-run response of relatively uneducated male workers. Second, in the short run, education must be viewed as a predetermined variable. Third, Heckman sample selection tests (based on a larger sample of all workers) do not indicate a sample selection bias.

5 While some agricultural workers with limited education, especially in California, live in cities (e.g., Fresno), over two-thirds do not.

6 Agricultural workers are those who identify themselves as hired agricultural workers (including managers and foreman) who work in agricultural production crops or livestock.

7 This sample is relatively old. Nonhispanics with relatively little formal education tend to be older than the rest of the population. the average age in agriculture is 42, the average age of nonhispanics is 49 (and the median is 50), but the average age of Hispanics is only 35 (and the median is 32).

Table 1
Means and Standard Deviations

Variable	All	Agriculture	Non-Agriculture
Number of Observations	931	181	750
Binary (0-1) variables (percent)			
Region			
New England (CT,MA,ME,NH,RI,VT)	8.8	2.8	10.3
Mid Atlantic (NJ,NY,PA)	4.4	0.6	5.3
East North Central (IL,IN,MI,OH,WI)	7.7	3.3	8.8
West North Central(IA,KS,MN,MO,ND,NE,SD)	10.2	14.4	9.2
South Atlantic (DC,DE,FL,GA,MD,NC,SC,VA,WV)	26.2	21.0	27.5
East South Central (AL,KY,MI,TN)	13.2	7.2	14.7
West South Central (AR,LA,OK,TX)	12.6	12.7	12.5
Mountain (AZ,CO,ID,MT,NM,NV,UT,WY)	9.2	10.5	8.9
Pacific (AK,CA,HI,OR,WA)	7.7	27.5	2.8
States			
California	5.7	27.1	0.5
Texas	6.2	8.3	5.7
Florida	3.7	3.9	3.6
Demographic characteristics			
Black	13.1	14.4	12.8
Other Nonwhites	2.1	1.7	2.3
Mexican (ethnicity)	13.2	43.1	6.0
Mexican American (ethnicity)	4.9	7.7	4.3
Other Hispanics (ethnicity)	1.4	2.2	1.2
Married, Living Together	72.9	52.5	77.9
Job Characteristics			
Union Member	14.7	1.7	17.9
Paid by the Hour	75.8	64.6	78.5
Agricultural Manager or Foreman	1.1	5.5	0.0
Agriculture (Hired Farm Worker)	19.4	100.0	0.0
Continuous variables [mean (s.d.)]			
Number of Children	0.7	1.07	0.6
	(1.2)	(1.5)	(1.1)
Years of Schooling	7.5	6.6	7.7
	(2.1)	(2.2)	(2.0)
Years of Experience	34.3	29.3	35.6
	(13.9)	(16.9)	(12.8)
Earnings per hour ("wage")	7.13	4.8	7.69
	(3.7)	(2.2)	(3.7)
Usual weekly hours	40.5	42.3	40.1
	(8.6)	(10.8)	(7.9)

Agricultural workers also have 6 years less work experience (calculated as age minus 6 years of formal school).

The average hourly earnings of agricultural workers ($4.80) is only 62 percent as high as that of nonagricultural workers ($7.69). They work, on average, 2.2 more hours a week, however, so that the average weekly earnings of agricultural workers are 66 percent of that of nonagricultural workers ($203 versus $308).

Methodology

This study uses a version of a model that has been employed by Lee (1978), Willis and Rosen (1979), Nakosteen and Zimmer (1980), and Robinson and Tomes (1982), among others. In this model, the natural logarithms of the hourly earnings ("wages") in the agricultural (w_a) and nonagricultural (w_n) sectors are a function of demographic and individual characteristics (X_j, $j = a,n$) and unmeasured sources of individual differences (ε_j); but the impact of individual characteristics on wages may vary across the sectors:[8]

$$w_a = X_a'\beta_a + \varepsilon_a \qquad (1)$$

and

$$w_n = X_n'\beta_n + \varepsilon_n\,. \qquad (2)$$

An individual's wage in a given sector is only observed if the individual is working in that sector at the time of the CPS interview. Agricultural work is often more physically taxing and dangerous than many other types of work. On the other hand, many people prefer working outside in agriculture to an indoor job. Let c be the cost or benefit (disutility or utility) of working in agriculture relative to working in another industry. This cost is modeled as a function of a worker's characteristics (Z):

$$c = Z'\delta + \varepsilon_c\,. \qquad (3)$$

The worker compares this cost or benefit to the relative wage in agriculture. The wage ratio (R) between agriculture and nonagricultural sectors is approximated by the difference in the natural logarithms of these wages:

$$R = w_a - w_n\,. \qquad (4)$$

The worker chooses to work in agriculture (industry choice, i, equals 1) only if the total benefits to working in agriculture, R - c, is positive:

$$i = 1 \quad \text{if } R - c > 0$$

8 Many agricultural workers receive piece rate payments (35 percent receive only piece rate compared to 11 percent of nonagricultural workers). The following equations use hourly earnings, which are calculated by dividing the reported weekly earnings by the reported usual weekly hours. For workers who receive time rate pay, this calculated hourly earnings number is usually identical or close to the reported wage.

and

$$i = 0 \quad \text{if } R - c \leq 0. \tag{5}$$

If i = 1, then the observed wage, w, is w_a. Otherwise, the observed wage is w_n.

The disturbances in the equations are assumed to be jointly normally distributed. As a result, a probit estimation technique can be used to estimate the choice of industry equation (5). Substituting for R in equation (5) using equations (4), (1), and (2) and for c using (3), we can rewrite the industry choice equation as:

$$I = 1 \quad \text{if } X_a'\beta_a - X_n'\beta_n - Z'\delta - \varepsilon_c > 0$$

and

$$I = 0 \quad \text{if } X_a'\beta_a - X_n'\beta_n - Z'\delta - \varepsilon_c \leq 0. \tag{6}$$

That is, the reduced-form probit specified in equation (6) can be estimated using all the exogenous (X_a, X_n, and Z) variables in equations (1), (2), and (3). Conditional on industry choice, as determined by equation (6), the hourly earnings equations (1) and (2) can be estimated using Heckman's (1979) technique to compensate for sample selection bias. Were we to estimate equations (1) and (2) using standard ordinary least squares techniques, the estimates would be biased because workers are not randomly assigned to the agricultural and nonagricultural sectors. That is, the unobserved individual difference or disturbance terms (ε_a and ε_n) would not be normally distributed if we examine data for only those workers observed in each sector.

Based on the estimated wage equations, (1) and (2), the wage differential is calculated as $\hat{R} = X_a'\hat{\beta}_a - X_n'\hat{\beta}_n$. Substituting this estimated value for R in equation (5), the structural probit for industry choice can then be estimated.

The key exogenous variables are geographic dummy variables (variables that take on values of one or zero) and demographic characteristics. Agricultural and nonagricultural wages differ geographically (reflecting differences in labor demand and supply). That is, the geographic dummies are included in X_a and X_n. Years of schooling and years of experience are also hypothesized to influence wages and are included in X_a and X_n. Racial and ethnic characteristics may influence wages and the costs of working in agriculture due to discrimination or because they reflect language skills and legal status. They may, for similar reasons, also affect the costs of working in agriculture, hence they are included in X_a, X_n, and Z.

Because of the relatively high variance in agricultural wage, having a working spouse may make one relatively more likely to work in agriculture; on the other hand, to the degree that agricultural jobs require migration, living with a spouse may be difficult. Similarly, having children may make migratory living relatively unattractive; however, whole families may work together in agriculture. Thus, marital status and children are included in Z, but their signs are uncertain.

The Empirical Results

As just described, the first step of the analysis is to estimate a reduced form probit equation describing how choice of industry depends on exogenous geographic and demographic variables. Then, conditional on industry choice, wage (hourly earnings) equations are estimated for each sector. Finally, a structural probit equation is estimated.

Reduced-Form Probit Equation

The reduced form probit coefficient and asymptotic standard error estimates are shown in the second and third columns of Table 2. Because it is a reduced form equation, the coefficients reflect both wage differential and cost factors, as described above. Compared to the Pacific region, living in the West North Central region makes one more likely to be in agriculture (all else the same). Similarly, Californians in the sample are much more likely to work in agriculture than others in the Pacific region.

One more year of experience makes a worker more likely to work in agriculture if the individual has at least 32 years of experience, and less likely otherwise. One more year of formal schooling makes one less likely to work in agriculture if one has had at least 5 years of schooling. Married men living with their spouses are less likely to be agricultural workers; however, the more children one has, the more likely one works in agriculture.

Workers who report their ethnicity as Mexican (as opposed to Mexican-American, Chicano, or other non-Mexican Hispanic) are much more likely to work in agriculture than other groups. Workers who report that they are non-Mexican Hispanics (including Mexican-American and Chicano) are also more likely to work in agriculture than other groups, but not as likely as the Mexicans.

The equation correctly predicts 87.3 percent of the observations. The various R^2 measures (see Maddala, 1983, Hensher and Johnson, 1981, and Chow, 1983) range from 0.29 to 0.46.[9] A log-likelihood test strongly rejects the hypothesis that only the constant term matters.

Wage Equations

The wage equations (the regression of the natural logarithm of hourly earnings on demographic and geographic variables) reported in Table 3 were estimated using Heckman's two-stage technique to control for the probability that one works in the relevant sector. The correlations between the disturbance in the regression and the selection criteria are very high (nearly one). On the basis of a Heckman-test (the t-statistic on the selectivity parameter), we cannot reject the hypothesis that sample selectivity matters in both the wage equations.

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9 The probit equation and summary statistics were estimated using Kenneth J. White's Shazam program version 6.1. The sample selection adjusted wage equations, reported below, were estimated using William H. Greene's Limdep program version 5.1.

Table 2
Probit Equation: Probability of Working in Agriculture

	Reduced-Form Equation Coefficient	Reduced-Form Equation Asymptotic S.E.	Structural Equation Coefficient	Structural Equation Asymptotic S.E.
Constant	-1.6850	0.7919	-0.8320	0.6336
New England	0.8507	0.6578	1.7278	0.6718
Mid Atlantic	0.2749	0.7558	1.8039	0.8124
East North Central	0.7512	0.6572	1.9637	0.7020
West North Central	1.6973	0.6319	2.1878	0.6269
South Atlantic	0.7273	0.6270	2.2353	0.6831
East South Central	0.8943	0.6425	2.3763	0.7039
West South Central	0.7172	0.6432	2.2399	0.7095
Mountain	0.5270	0.6298	1.9857	0.6958
California	2.5634	0.6735	4.2979	0.7616
Texas	-0.3190	0.3340	0.1003	0.3459
Florida	0.1140	0.3002	-0.1936	0.2986
Mexican	1.3387	0.2331	0.4290	0.3167
Non-Mexican Hispanic	0.9263	0.2601	0.0284	0.3384
Black	0.5120	0.1741	-0.1448	0.2335
Other Non-White	0.3833	0.4249	-0.7369	0.4859
Married, Living with Spouse	-0.6666	0.1385	-0.6085	0.1298
Number of Children	0.0899	0.0537	0.0825	0.0495
Years of School	0.3610	0.1400	-	-
Years of School Squared	-0.0358	0.0120	-	-
Experience	-0.0561	0.0160	-	-
Experience Squared	0.0009	0.0002	-	-
$R = \ln(w_a) - \ln(w_n)$	-	-	1.7273	0.3742

	Reduced-Form Equation	Structural Equation
Number of Observations	931	931
Log-Likelihood (Constant only)	-458.57	-458.57
Log-Likelihood	-300.27	-301.65
Likelihood ratio test	316.61 (21 d.f.)	313.85 (18 d.f.)
R^2 Measures:		
Maddala R^2	0.29	0.29
Cragg-Uhler R^2	0.46	0.46
McFadden R^2	0.35	0.34
Chow R^2	0.38	0.37
Percentage of Correct Predictions	87.3	87.5

Prediction Success Table (Reduced-Form Equation)

		Actual 0	Actual 1
Predicted	0	731	99
	1	19	82

Prediction Success Table (Structural Equation)

		Actual 0	Actual 1
Predicted	0	733	99
	1	7	82

The equations show that there is a different relationship in the two sectors with respect to geographic and demographic variables. In the nonagricultural sector, controlling for demographic characteristics, wages do not differ statistically significantly across regions or states with the sole exception that wages are higher in California. In contrast, there are pronounced regional differences in agricultural wages. Compared to the Pacific region, wages are significantly lower in the New England, Mid Atlantic, East North Central, South Atlantic, East and West South Central, and Mountain regions. For this sample, and unexpectedly, agricultural wages are lower in California than in the rest of the Pacific region, controlling for other factors. Similarly, the wages in Texas are significantly lower than in the Pacific region.

Surprisingly, at least for this relatively uneducated group, extra education does not statistically significantly increase one's nonagricultural wage. In agriculture, however, extra education is positively related to wages up to 5 years of school, and negatively for more years.[10] Agricultural wage differences due to extra years of school are small (only a few pennies per hour), however.

At least until one has 33 years of experience, extra experience increases the nonagricultural wage; whereas, extra experience (or age) does not have a statistically significant effect on the agricultural wage. In the nonagricultural sector, having 20 years of experience instead of 10 is worth 18¢ more per hour.

In the nonagricultural sector, blacks earn 15¢ less per hour than whites; whereas there is no statistically significant wage differential between whites and other nonwhites or either group of Hispanics. In the agricultural sector, workers who report their ethnicity as Mexican earn 60¢ more an hour than whites, other Hispanics earn 50¢ more, blacks earn 25¢ more, and other nonwhites earn 73¢ more. Thus, all else the same, agriculture pays nonwhites and Hispanics relatively more compared to whites than does the nonagricultural sector. These wage differentials may reflect either discrimination or other effects not otherwise captured by the exogenous variables in this equation. For example, nonwhites and Hispanics may be more likely to work in jobs that pay premium wages such as dangerous jobs or certain migrant jobs.[11]

The R^2 measure in the agricultural equation is more than twice as high as that in the nonagricultural sector (but R^2 measures need to be viewed with caution in non-ordinary least squares estimates such as these). The hypothesis that only the constant term matters can be strongly rejected, of course, in both equations.

The qualitative results from these wage equations are similar to those of many earlier studies that are based on CPS data. For example, the agricultural sector equation's

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[10] Most previous studies based on the CPS (without educational limits on the sample) find that education does not have a statistically significant effect in the agricultural sector (e.g., Perloff, 1985), but does in the non-agricultural sector.

[11] Within the CPS sample, unionized agricultural workers are almost always Hispanic, which may explain part of this effect. Estimating this system of equations including union as an independent variable in the wage equations and an extra probit equation for union status, which is correlated with the industry choice equation, proved impossible to estimate (possibly due to the sample). Using the specification reported here, but including union as an exogenous variable leaves the other coefficients in the wage equation relatively unaffected. As a result, union status was left to the residual term in the equations reported here.

Table 3
Logarithmic Wage Equations Adjusting for Sample Selectivity

	Agriculture		Non-Agriculture	
	Coefficient	S.E.	Coefficient	S.E.
Constant	1.3161	0.3586	1.4641	0.1777
New England	-0.5338	0.2736	0.0637	0.1235
Mid Atlantic	-0.9429	0.3073	-0.0006	0.1323
East North Central	-0.7060	0.2761	0.0578	0.1245
West North Central	-0.3139	0.2846	0.0540	0.1326
South Atlantic	-0.9556	0.2555	-0.0343	0.1184
East South Central	-1.0005	0.2658	-0.0745	0.1211
West South Central	-0.9481	0.2734	-0.0169	0.1248
Mountain	-0.8560	0.2648	0.0473	0.1211
California	-0.5710	0.2771	0.4837	0.2378
Texas	-0.3038	0.1215	-0.0474	0.0932
Florida	0.1633	0.1099	0.0314	0.0807
Years of School	0.1535	0.0555	0.0339	0.0395
Years of School Squared	-0.0149	0.0050	-0.0023	0.0034
Experience	-0.0061	0.0069	0.0325	0.0057
Experience Squared	0.0001	0.0001	-0.0005	0.0001
Mexican	0.6067	0.1136	0.0820	0.1004
Non-Mexican Hispanic	0.5082	0.1129	-0.0225	0.0840
Black	0.2511	0.0762	-0.1543	0.0521
Other Non-White	0.7284	0.1866	0.0755	0.1061
Selectivity Parameter	0.3867	0.0936	0.4607	0.1348
Number of Observations	181		750	
Mean of ln(wage)	1.50		1.94	
Standard Deviation of ln(wage)	0.35		0.44	
Standard Error of the Regression	0.27		0.40	
Standard Error Corrected for Selection	0.40		0.47	
Sum of Squared Residuals	13.62		119.85	
R^2	0.39		0.17	
F(20, 729)	5.15		7.71	
Log-Likelihood	-22.70		-376.50	
Log-Likelihood (Constant only)	-67.70		-448.41	
$\chi^2(20)$	90.00		143.82	
Squared Correlation of Disturbance in Regression and Selection Criterion	0.937		0.979	

result are similar to those in Perloff (1985). It should be noted, however, that none of these earlier studies controlled for possible simultaneity bias due to industry choice.

Structural Probit Equation

The structural probit equation is reported in the last two columns of Table 2. To estimate it, the wage ratio, R, is approximated by the difference in the natural logarithm of the wage one would earn in agriculture and in the nonagricultural sector. That is, the wage equations are used to calculate a wage ratio even though wages are actually observed only in one sector.

In the structural equation, controlling for demographic and wage ratios, workers are more likely to work in agriculture in most other regions of the country than in the Pacific region. They are also more likely to work in agriculture in California than in the rest of the Pacific region. Of course, given that wage ratios vary geographically, much of the geographic difference in choice of sector is captured by the wage ratio term.

The structural probit equation does not show a significant difference between choice of sector among whites and other racial or ethnic groups, after controlling for wage ratios. That is, the preference of working in agriculture of these groups shown in the reduced form equation is presumably captured in the structural equation by the wage ratio term, which reflects relatively high agricultural wages for this group.

The wage ratio term is statistically significant and has a large effect. A one percent increase in the relative wage in agriculture increases the probability that one works in agriculture by 3.37 percent at the sample mean. On average over the entire sample (Hensher and Johnson, 1981), a one percent increase in the relative agricultural wage increases the probability of working in agriculture by 1.3 percent .

The ratio of the estimated agricultural wage to the nonagricultural wage is 0.37 for those workers in agriculture and only 0.30 for those who are not in agriculture. Moreover, the ratio of the estimated wage in agriculture of those in agriculture is 1.24 times that of those who choose not to work in agriculture. Thus, choosing to work in agriculture appears, in large part, to be based on a comparison of wages between the two sectors.

The structural probit has virtually the same explanatory power as the reduced form equation. The R^2 measures range from 0.29 to 0.46 and the prediction success table shows that the correct sector is predicted for 87.5 percent of the sample.

Sensitivity Experiments

Other experiments were used to test the sensitivity of these results to the specifications used. In none of these experiments was the key result (the effect of the wage ratio in the structural probit) substantially affected.

As a sensitivity test on the assumed error structure, the system was reestimated using logit rather than probit equations (that is, the disturbances were modeled as logistic rather than normal). Although the key result was virtually the same as with the probit system, the correlation between the reduced form logit equation and the selectivity equations' disturbance terms were estimated to be greater than one. For that reason, only the probit equations are reported here.

As mentioned above, the results are not sensitive to the educational threshold or union status. In another experiment, seasonal dummy variables which were included in all the equations had coefficients that were not statistically significantly different from zero either individually or collectively in any equation. In yet another experiment, military veteran status was used as a proxy for legal status in the probits. Although its coefficient was significant in the reduced-form probit, it is not included in the equations reported here because of its ambiguous interpretation: for example, it may be a proxy for other factors such as age.

Finally, the dummy variable for non-Mexican Hispanic was divided into Mexican-American (or Chicano) and other Hispanics. These latter two variables had virtually identical coefficients, so that only their aggregate is used here.

The Response to Higher Agricultural Wages

The system of equations can be used to simulate the effect of higher wages on the supply of labor to agriculture. Table 4 and Figure 1 show the results of these simulations. As shown in Table 4, the estimated agricultural wage is 29.38 percent of the estimated nonagricultural wage when averaged across the sample. The simulations examine the effects of an across-the-board increase in the agricultural wage holding the nonagricultural wage constant. That is, the simulations increase the constant term in the regression on the logarithm of the agricultural wage, which is equivalent to a constant percentage increase in the agricultural wage for all workers.

In the table and figure, two methods are used to calculate the effect of the wage increase on the share of workers in agriculture. In the first method, a worker is assigned to the agricultural sector if the probability he works in agriculture (according to the structural probit equation) is at least 50 percent. Using this 50 percent rule, the model predicts that 10.63 percent of the workers will work in agriculture. In the second method, the probability of working in agriculture that the model predicts for each individual is averaged across all individuals (using equal weights). This average is 19.45 percent, which is virtually the same as the actual percent in the sample. Both methods are shown in Table 4 and in Figure 1.

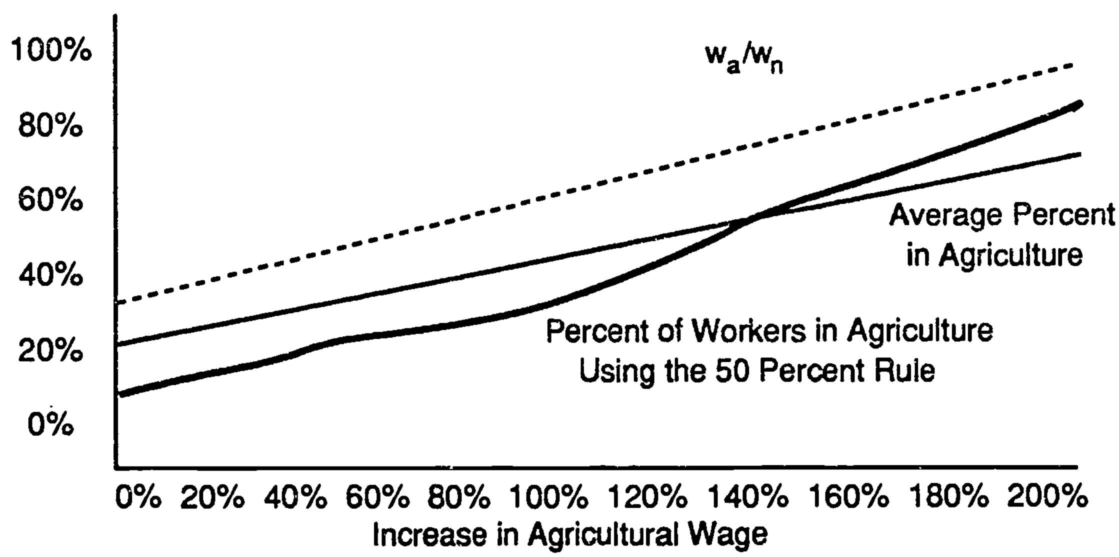
If the agricultural wage were increased by 2 percent, the wage ratio would increase by 2.2 percent. Using the 50 percent rule, the share of workers in agriculture would increase by 3.1 percent (or 0.33 percentage points from 10.63 to 10.96 percent); whereas, using the second method, the share would increase by 3.2 percent (or 0.62 percentage points from 19.45 to 20.07 percent).

If the agricultural wage were raised by 10 percent (with no response in the nonagricultural wage), the first method indicates a 23.2 percent increase in the share of agricultural workers; and the second method predicts a 16.0 percent increase. The comparable figures for a 50 percent increase in the agricultural wage are a 138.5

Table 4
Effect of an Increase in the Agricultural Wage

Increase in the Agricultural Wage (%)	w_a/w_n (%)	Percent Agricultural Workers 50 % Rule	Average
0.00	29.38	10.63	19.45
2.00	30.04	10.96	20.07
4.00	30.63	11.82	20.69
6.00	31.21	12.35	21.31
8.00	31.80	12.78	21.94
10.00	32.39	13.10	22.57
20.00	35.34	15.47	25.76
30.00	38.28	18.69	28.96
40.00	41.23	21.91	32.15
50.00	44.17	25.35	35.30
100.00	58.90	46.51	49.84
150.00	73.62	62.73	61.71
200.00	88.34	79.91	70.93

Figure 1
Effect of an Increase in the Agricultural Wage



percent and a 81.5 percent increase. If agricultural wages were doubled, either method indicates that nearly half of these workers would work in agriculture, all else the same.

This approach implicitly assumes that geographical differences are additive in the logarithms (multiplicative), so that the effect of a wage differential is the same in California as in the rest of the country.[12] Unfortunately the CPS does not provide a large enough sample of Californians to estimate this system of equations just for California. Simple specification tests (California interaction terms), however, do not suggest a problem with the current specification.

While the results above are encouraging in general, they are less encouraging in California where much agricultural labor has been supplied by immigrants lacking legal status. One survey, Rosenberg and Perloff (1983), indicates that in 1987 one-third of new hires were undocumented workers and that the use of undocumented workers was greater in certain crops such as grapes and other fruit and tree-fruit crops. Thus, unless these formerly undocumented workers became documented workers under IRCA's amnesty program and stay in agriculture, at least some California crops will need to replace many workers.

The calculations above are based on the response of only relatively uneducated workers who live in rural areas. It is, of course, possible that other workers would be induced to take farm employment if agricultural wages are raised relative to those elsewhere. Thus, the simulations reported here may be lower bounds on the true (larger) response.

Conclusions

This paper presents a model of industry choice and wage determination. It shows that discrimination against blacks and other minority groups occurs in nonagricultural sectors; but not in agriculture.

The chief result of this analysis is that inducing more workers to switch to agriculture may not be as costly as some commentators have suggested. Indeed, a 10 percent increase in wages may increase the share in agriculture of rural male workers with no more than a ninth grade education by nearly a quarter. Nonetheless, in California, and especially in certain crops, a very large percent of the agricultural work force has traditionally been undocumented workers. Thus, although these results indicate that additional workers are available, fairly large wage increases may be required, at least in certain areas and crops.

Further work on agricultural labor supply remains to be done. For example, this report has focused on the role of higher wages in attracting agricultural labor. In general, though, better working conditions and other benefits (such as health insurance and housing) could also attract extra workers, holding wages constant.

12 It should also be noted that the CPS does not report the crop for agricultural workers, so that we cannot determine whether an individual is covered by IRCA's Special Agricultural Worker (SAW) program. Since the SAW program may attract workers to agriculture, our results may be special to that current period.

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