Job Displacement from Agriculture

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Abstract:

We analyze jobless spell duration and re-employment wages for workers displaced from agricultural employment. Unlike manufacturing employers, "small" agricultural employers are not required to participate in the Unemployment Insurance system, which leaves many farm workers ineligible to receive unemployment benefits. Theory implies that displaced workers who are ineligible for unemployment benefits experience shorter jobless spells and lower re-employment wages. With data from the Displaced Workers' Supplement to the Current Population Survey, we show that, compared to workers displaced from manufacturing, displaced agricultural workers face significantly shorter unemployment duration and lower re-employment wages. Our estimates imply that displaced agricultural workers experience 25 percent (5.6 weeks) less time unemployed and upon re-employment earn 10 percent less than displaced manufacturing workers. Further, we present evidence that in states where "small" agricultural employers are required to participate in the Unemployment Insurance program, displaced agricultural workers face unemployment duration and re-employment wages of similar magnitude to those of manufacturing workers. As expected, in states where the "small" farm employer exemption applies, workers displaced from agriculture face longer jobless spells and lower re-employment wages compared to workers displaced from the manufacturing sector.

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I. Introduction

While employment in agriculture has been steadily declining in the past few centuries, labor is still a critical input in the agricultural sector (see Findeis et al., 2002). In comparison to other labor markets, the agricultural labor market is quite different – it is characterized by higher turnover, larger fraction of seasonal employees and undocumented workers, and lower rates of unemployment insurance eligibility (see Tran and Perloff, 2002; Findeis et al., 2002; Bassi and McMurrer, 1997; Isé and Perloff, 1995). While much research has been done on agricultural workers' turnover, health insurance status, wages and employment, as well as welfare in general, not much work has investigated the impact of job displacement from agriculture on workers' post-displacement outcomes (see Zheng and Zimmer, 2008; Katchova, 2008; Kandilov and Kandilov, 2008; Tran and Perloff, 2002, McNamara and Ranney, 2002; Isé and Perloff, 1995, Dunn, 1985). This study attempts to fill that gap.

Using data on displaced workers from the agricultural and manufacturing sectors of the U.S. economy, we show that consistent with the less stringent unemployment insurance provisions for agricultural employers, workers dislocated from the agricultural industry have lower duration of unemployment and lower re-employment earnings than workers displaced from the manufacturing sector. To guide intuition, we develop a simple job search model which implies that workers who are displaced from the agricultural sector and therefore have much lower eligibility for unemployment benefits will also face a lower unemployment duration and lower re-employment earnings compared to workers displaced from the manufacturing sector. To test these predictions and compare the post-displacement outcomes of workers displaced from the agriculture to those of workers displaced from manufacturing we employ data from the Displaced Workers' Supplement to the Current Population Survey (CPS) from the ten largest (in terms of employment) agricultural states from 1979 to 2001. Our empirical analysis reveals that, on average, workers displaced from agriculture experience a 25 percent lower jobless spell duration and 11 percent lower re-employment wages. Further, we show that consistent with the job search model, in states where a larger fraction of agricultural employers are required to participate in the unemployment insurance program, i.e. in states where a larger fraction of displaced farm workers are eligible for unemployment benefits, displaced agricultural workers face a smaller decrease in unemployment duration and re-employment wages.

The rest of that paper is organized as follows. In the next section, we detail the differences for agricultural and manufacturing employers in the provisions for employer participation in the unemployment insurance system across states. Section III presents the job search model that guides our empirical specification. We describe our data in section IV, and detail our econometric strategy in section V. In section VI, we present and discuss the results. Section VII concludes.

II. Unemployment Insurance Provisions for Agricultural Employers

The unemployment insurance (UI) system in the U.S. is a federal-state program, and it is jointly financed with federal and state employer payroll taxes. Generally, employers pay both state and federal unemployment taxes if (1) they pay wages to employees totaling $1,500, or more, in any quarter of a calendar year, or (2) they had at least one employee during any day of a week during 20 weeks in a calendar year, regardless of whether or not the weeks were consecutive.[1]

[1] In order to qualify for unemployment benefits, workers also need to meet (1) state requirements for wages earned or time worked during a "base period" – in most states, this is usually the first

However, there are a few exceptions to the general rule above. Notably, "small" agricultural employers are not required to participate in the UI program.

The Federal Unemployment Tax Act (FUTA), authorizes the Internal Revenue Service to collect a federal employer tax used to fund state workforce agencies.[2] The federal regulations stipulate that agricultural employers become subject to the federal unemployment tax only if (1) they pay cash wages to employees of $20,000, or more, in any calendar quarter, or (2) in each of 20 different calendar weeks in the current or preceding calendar year, there was at least 1 day in which they had 10 or more employees performing service in agricultural labor.[3] These provisions exclude a large fraction of workers employed in the agricultural sector. For example, the 2002 Census of Agriculture reports that nearly 70 percent of farms in the U.S. have annual

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four out of the last five completed calendar quarters prior to the time the claim is filed, and (2) they must have become unemployed through no fault of your own.

[2] The FUTA tax rate is 6.2 percent of taxable wages, where the first $7,000 paid in wages to each employee during a calendar year constitutes the taxable wage base. Employers who pay the state unemployment tax on a timely basis receive a credit of up to 5.4 percent regardless of the rate of tax they pay the state. The net FUTA tax rate is then generally 0.8percent. State laws determine individual state's unemployment insurance tax rates. The federal employer tax covers the costs of administering the UI program in all states. In addition, it pays half of the cost of extended unemployment benefits during recessions, and provides for a fund from which states borrow, when necessary, to pay unemployment benefits. The state unemployment tax, which is paid to state workforce agencies, is used entirely to finance benefits for eligible unemployed workers.

[3] The 20 weeks do not have to be consecutive weeks, nor must they be the same 10 employees, nor must all employees be working at the same time of the day.

sales of less than \$20,000. If farm payroll is generally smaller than farm sales, this statistic implies that workers in nearly 70 percent of all farms in the U.S. may not be covered. Additionally, the 2002 Census of Agriculture reports that about 90 percent of all farms in the U.S. employ fewer than 10 workers, and employment in such farms represents more than 40 percent of total agricultural employment in 2002.

Even migrant workers who are employed on large farms most of the time may lose unemployment benefit eligibility as a result of the "small" farm exclusion. If agricultural worker's wages on "small" farms are not covered, using only their wages from "large" farms in determining monetary eligibility may disqualify them from receiving unemployment benefits even when their total wages would allow them to qualify. "Small" farms are not required to participate in the UI system for a number of reasons, such as small farmer's financial hardship, difficulties in insuring workers with multiple employers, and a large percentage of undocumented aliens (see Bassi and McMurrer, 1997).[4] Further, agricultural workers hired by a farms labor contractor, and not directly by the farm owner, could be considered employees of the contractor, and not employees of the farm. In these instances, worker advocates claim, the contractor, who is delegated with the responsibility of filing and paying for unemployment insurance, frequently neglects to comply. Compounding this problem is the more widespread use of farm labor contractors in recent years – for example, Mehta et al. (2000) report that 19 percent of all hired farm workers were employed by a farm labor contractor in 1997 (see also Martin, 1994).

[4] For example, analyzing crop workers with data from the National Agricultural Workers Survey, Mehta et al. (2000) find that 52 percent of hired farm workers lack work authorization in 1997-1998.

While at the federal level, the FUTA does not require “small” farms to participate in the UI program, some states have expanded their agricultural coverage provisions. For example, Florida considers agricultural employers liable to pay state UI taxes if (1) they pay cash wages to employees of $10,000, or more, in any calendar quarter, or (2) in each of 20 different calendar weeks in the current or preceding calendar year, there was at least 1 day in which they had 5 or more employees. Florida’s provisions are twice as stringent as the federal requirements outlined in FUTA when it comes to agricultural employers’ obligation to participate in the UI system. In Texas, UI program criteria are even more stringent – farm employment is covered by the UI program if the farm owner pays cash wages of $6,250 or more in a calendar quarter, or if (s)he employs 3 or more workers in 20 weeks. Finally, in California, the rules for coverage for agricultural workers are exactly the same as those for manufacturing workers, i.e. there is no “small” farm exclusion. While, some states do have expanded UI programs, many others still employ the federal requirements and exclude “small” farms from UI tax liability. In our sample of ten states with the largest agricultural number of workers, five states (Kentucky, North Carolina, Minnesota, Michigan, and Iowa) follow the federal guidelines and do not require “small” agricultural employers to participate in the UI system, while the other five states (California, Washington, Texas, Oregon, and Florida) have significantly expanded coverage for “small” agricultural employers. We call the first group of states less generous (when it comes to agricultural worker’s UI eligibility), and the second group – more generous.

III. Theoretical Framework

We consider a simple continuous time version of the standard job search model.[5] Assume job

[5] See Devine and Kiefer (1991), as well as Burdett and Ondrich (1985).

offers arrive to a searching unemployed worker at random intervals according to a Poisson process. Therefore, the probability of obtaining an offer in a given, short, interval of time is proportional to the length of that interval. To keep the model tractable, we make a number of simplifying assumptions. First, workers are assumed to maximize the expected present value of income over an infinite horizon at a known and constant (both in time and across individuals) discount rate r.[6] Second, we assume that the net income flow for an unemployed worker, the net unemployment benefit, b, is time-invariant throughout the duration of any given spell of unemployment. This delivers a *constant* reservation wage policy.[7] The third assumption has to do with the offer arrival process while unemployed. As standard in the job-search literature, we assume that offers arrive according to a Poisson process with arrival rate δ. Thus, the probability of receiving at least one offer within a short interval of length h is $\delta h + o(h)$, where $o(h)$ is the

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[6] This assumption is not crucial to the general implications of the model but it leads to more easily interpreted derivations than the utility maximization case.

[7] If b declines with unemployment spell duration, the reservation wage will also decline (see Van Den Berg, 1990). If workers displaced from agriculture have shorter unemployment benefit exhaustion duration compared to workers displaced from manufacturing, their reservation wages would be declining faster, hence unemployment duration will tend to be shorter ceteris paribus. Another issue, which we do not model here, that comes up in the policy debate is search intensity. Longer or more generous unemployment insurance benefits to workers displaced from manufacturing (for example as a result of increased international trade) may reduce their search intensity and increase unemployment duration as the affected workers may take some of this "windfall" as leisure.

probability of receiving more than one offer in the interval h.[8] The Poisson arrival rate assumption is convenient and makes the model tractable.

Next, we assume that an offer is summarized by a wage rate w. If accepted, this wage will be received continuously over the tenure of employment on the job. Successive job offers received over the course of a spell of unemployment are independent realizations from a known wage offer distribution with a finite mean, μ, and variance, σ, with cumulative distribution function $F(w)$, and density $f(w)$. Once rejected, an offer cannot be recalled. Moreover, given the assumptions we have mode so far, a recall option will not be exercised if available.[9] When accepted, a job lasts forever.[10]

As the net unemployment benefit is constant, wage offers are independent and identically distributed, the offer distribution and the arrival rate are known and time-invariant, the value function for an unemployed worker, V^U, is constant over the duration of a spell and it is implicitly defined by the following Bellman equation (1) below:

$$V^U = \frac{1}{1+rh} bh + \frac{\delta h}{1+rh} E_W[\max\{V^E(w), V^U\}] + (1-\delta h)\frac{1}{1+rh} V^U + o(h)K \qquad (1),$$

[8] Note that $o(h)/h \to 0$ as $h \to 0$. Therefore the probability of receiving more than one offer in a short period of time is negligible.

[9] With a constant reservation wage, if an offer is unacceptable today, it will be unacceptable tomorrow as well.

[10] Incorporating Poisson layoffs, for example, changes the model trivially. Incorporating retirement at a fixed date in the future changes the calculation in principle, but in practice, the change is trivial if the horizon is long and the discount rate is positive.

where h is a short time interval. The first term on the right hand side is the discounted present value of the net unemployment income over the interval h. The second term is the product of the probability of receiving an offer in the interval h and the discounted expected value of following the optimal policy if an offer w is received, where $V^E(w)$ denotes the present value of accepting the offer. The third term is the probability of no offer in the interval h times the discounted value of optimal search thereafter. The last term accounts for the probability of receiving more than one offer during h, with K being the optimal policy if that happens.

The expected present value of $V^E(w)$ here is simply the present value of expected lifetime income at that wage, hence

$$V^E(w) = \frac{w}{r} \qquad (2)$$

Note that $V^E(w)$ is continuous and strictly increasing in w and V^U does not depend on the offer w, therefore the optimal strategy for the worker is a constant in time reservation wage policy, such that the worker would always accept if $w \geq w^r$, where w^r is the minimum acceptable wage offer defined implicitly by

$$V^E(w^r) = \frac{w^r}{r} = V^U \qquad (3)$$

The time invariance of the worker's optimal policy is a result of the stationary of the environment.

By substituting eq.(2) and eq.(3) in eq.(1) we obtain

$$\frac{w}{r} = \frac{1}{1+rh} bh + \frac{\delta h}{1+rh} E_W[\max\{\frac{w}{r}, \frac{w^r}{r}\}] + \frac{(1-\delta h)}{1+rh} \frac{w^r}{r} + o(h)K \qquad (4)$$

After rearranging and taking the limit as $h \rightarrow 0$, the optimality condition becomes

$$w^r = b + \frac{\delta}{r} \int_{w^r}^{\infty} (w - w^r) dF(w) \qquad (5)$$

The Bellman equation (5) produces w^r which in turn defines the optimal policy for the job seeking individual. If a wage higher than w^r is offered, employment should be accepted, if not, the optimal action is to continue searching. Eq. (5) also allows us to investigate the impact of unemployment benefit on the reservation wage, the expected re-employment wage, and the expected duration of unemployment. Differentiating the Bellman equation (5) with respect to the unemployment benefit, b, shows that the reservation wage, w^r, falls as b declines:

$$\frac{dw^r}{db} = \frac{r}{r + \delta(1 - F(w^r))} > 0 \qquad (6).$$

Further, equation (6) implies that the expected re-employment wage, $E_w[w \mid w \geq w^r]$, is an increasing function of the unemployment benefit, b:

$$\frac{dE_w[w \mid w \geq w^r]}{db} > 0 \qquad (7).$$

Workers displaced from the agricultural sector are much less likely to qualify for unemployment benefits and if they do qualify, the benefit may have lower value and shorter potential duration. Table 2 presents evidence that only 37 percent of displaced agricultural workers have collected unemployment benefits, whereas 63 percent of displaced manufacturing workers received unemployment benefit during the sample period. Hence, eq. (7) implies our first testable hypothesis:

(H1): Workers displaced from the agricultural sector face a lower re-employment wage than do workers displaced from the manufacturing sector.

Without any additional assumptions, the search model also implies that the distributions of the unemployment spells are exponential (Devine and Kiefer 1991). The transition rate (also known as the hazard or escape rate) between unemployment and employment is given by

$$\tau = \delta \int_{w^r}^{\infty} f(w)dw = \delta(1 - F(w^r)) \qquad (8).$$

And since the completed durations have an exponential distribution with parameter τ, the expected length of a completed spell is

$$E(T) = \frac{1}{\tau} \qquad (9).$$

Differentiating equation (8) with respect to the unemployment benefit, b, shows that the hazard rate, τ, rises as b declines:

$$\frac{d\tau}{db} = \frac{-r\delta f(w^r)}{r + \tau} < 0 \qquad (10).$$

Hence, equation (10) implies that the expected unemployment duration, $E(T)$ is an increasing function of the unemployment benefit, b:

$$\frac{dE[T]}{db} > 0 \qquad (11).$$

As we already discussed, workers displaced from the agricultural sector are much less likely to qualify for unemployment benefits than workers in manufacturing. Eq. (11) therefore implies our second testable hypothesis:

(H2): Workers displaced from the agricultural sector experience a shorter spell of unemployment than do workers displaced from the manufacturing sector.

To summarize, because displaced agricultural workers are less likely to qualify for unemployment benefits than displaced manufacturing workers, the theory developed here predicts that agricultural workers would face lower re-employment wages (H1), but would also experience shorter jobless spell duration (H2) than would workers displaced from manufacturing. The next section describes the data we use to test these two hypotheses.

IV. Data

To compare post-displacement outcomes between workers displaced from the agricultural sector and those displaced from manufacturing, we use data from the only large-scale and multi-state data source – the Displaced Workers' Supplement (DWS) to the January or February Current Population Survey (CPS), which is jointly administered by the Bureau of Labor Statistics and the Census Bureau. The first DWS was instituted in January of 1984 and on a biennial basis afterwards. We use all the surveys through year 2002, which supplies us with data on displaced workers from 1979 to 2001. DWS is intended for all workers who have been displaced from their jobs in the 3 (or 5) years prior to the survey. In addition to personal characteristics found in the regular monthly CPS, DWS collects information on both old and new employment for displaced workers – previous and current wages, hours, current sector of employment, sector of displacement, reason for displacement, and duration of unemployment, among other things.[11] To

[11] Workers can be displaced due to six different reasons (see Table 2): (1) establishment closed, (2) insufficient work, (3) position abolished, (4) seasonal job ended, (5) self-employed business failed, and (6) other reasons. Prior to the 1994 DWS, information on old and new employment, including the duration of unemployment, was collected from workers displaced for any of these six reasons. Starting with the 1994 DWS, information on old and new employment, including the duration of unemployment, was collected for workers displaced for reasons (1), (2), or (3), but not (4), (5), or (6). Note that agricultural workers are proportionately more likely to be displaced for reasons (4), (5), or (6). In our empirical analysis, we use all the DWS surveys from

control for local labor market conditions in the year of displacement, we also use data on the annual state unemployment rate (from the Bureau of Labor Statistics) matched to the displaced worker's state of residence and year of displacement.

We use data on workers from the ten largest agricultural states in terms of number of workers employed in the agricultural sector – California, Washington, Texas, Oregon, Florida, Kentucky, North Carolina, Minnesota, Michigan, and Iowa (see Table 1). Information on the state's UI system generosity towards agricultural workers in each of the ten states considered in the analysis was collected from the state's Department of Labor (or its equivalent) website and, on a number of occasions (North Carolina and Michigan), via a direct phone conversation with the appropriate liaison at the state's labor affairs agency.

Descriptive statistics based on the DWS data for workers displaced from the agricultural and the manufacturing sectors in those ten states are reported in Table 2. Note that not everyone is re-employed at the date of the survey, in part because some have been displaced just a few weeks before the interview – about 73 percent (3,153) of all workers (4,347) hold a job at the time of the interview. As expected, higher fraction of workers employed in (and displaced from) agriculture are seasonal workers (12 percent in agriculture vs. 1 percent in manufacturing). Both pre-displacement and re-employment earnings of agricultural workers are lower than those of manufacturing workers. Consistent with the less stringent UI participation requirements for agricultural employers, the summary statistics show that only 36 percent of displaced agricultural workers receive unemployment benefits, whereas this statistic is 60 percent for workers displaced from manufacturing sector. Because agriculture tends to be low-skill intensive

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1984 to 2002, however very similar results to those presented here obtain if we employ the DWS surveys from 1984 to 1992 alone.

industry compared to manufacturing, we also see a larger fraction of workers with no high school diploma employed in agriculture. The average job tenure among agricultural workers (3.50 years) is lower than that of manufacturing workers (5.30 years), reflecting the larger fraction of migrant workers employed in the agricultural sector. Not surprisingly, as we already discussed, 27 percent of all (displaced) farm workers are Hispanic, whereas only 13 percent of manufacturing workers are Hispanic.[12] Given that they have found employment, about two-thirds (66 percent) of displaced agricultural workers have landed a job outside of agriculture; only 47 percent of manufacturing workers leave the sector upon re-employment.

V. Econometric Specification

V.1 Unemployment Duration

To deal with right-censored unemployment duration observations present in the DWS data, we employ maximum likelihood (ML). More formally, we estimate a Weibull model maximizing the following log-likelihood function

$$logL = \sum_{i=1}^{N} \{d_i log[f(t_i \mid \boldsymbol{Z}_i, \boldsymbol{\eta})] + (1-d_i) log[1-F(t_i \mid \boldsymbol{Z}_i, \boldsymbol{\eta})]\} \qquad (12),$$

[12] The previously cited statistics of 52 percent of unauthorized (undocumented) workers from Mehta et al. (2000) does imply that about 50 percent (or more) of crop workers are Hispanic, as almost all undocumented crop workers are Hispanic. Note that Mehta et al. (2000) use the National Agricultural Worker Survey, which collects data on hired crop workers and uses a different survey methodology from the DWS. For more details on the differences between the DWS and the NAWS see Larson et al. (2002). One notable difference between the two surveys is that the DWS has no information on the worker's legal status, whereas the NAWS does.

where the Weibull distribution has the following conditional density

$$f(t_i \mid \boldsymbol{Z}_i, \boldsymbol{\eta}) = \exp(\boldsymbol{Z}_i, \boldsymbol{\eta})\varphi t_i^{\varphi-1} \exp[-\exp(\boldsymbol{Z}_i, \boldsymbol{\eta}) t_i^{\varphi}] \qquad (13),$$

and d_i is a censoring indicator equal to unity if the unemployment duration of displaced worker i is uncensored, and N is the number of displaced workers included in the analysis. The vector of parameters to be estimated is $\boldsymbol{\eta}$, and $\boldsymbol{Z}_i$ is the matrix of personal characteristics. The Weibull hazard (of leaving the unemployment pool) at time, or unemployment duration, t_i, is given by

$$\phi(t, \boldsymbol{Z}_i, \boldsymbol{\eta}) = \exp(\boldsymbol{Z}_i, \boldsymbol{\eta})\varphi t^{\varphi-1} \qquad (14).$$

It captures a monotonically increasing or monotonically decreasing (in unemployment duration) hazard – if $\varphi > 1$, the hazard exhibits positive duration dependence, and if $\varphi < 1$, it exhibits negative duration dependence. This specification accommodates for the negative duration dependence visually found in the data. When presenting the results, we employ the Weibull model accelerated failure time (AFT) representation, allowing the interpretation of the estimated coefficients as semi-elasticities of the expected unemployment duration with respect to a given covariate in $\boldsymbol{Z}_i$.

The matrix of personal characteristics, $\boldsymbol{Z}_i$, can be written as

$$\boldsymbol{Z_i} = [\mathbf{X}_{ikjst} \mid U_{st}^{RATE} \mid \sigma_s \mid \tau_t \mid \delta_k \mid Ag_{ikjst}],$$

which includes $\boldsymbol{X}_{ikjst}$ – a vector of personal characteristics for individual i, surveyed in year k (k = 1984, 1986, …, 1992) displaced from sector j (j = agriculture, manufacturing) in year t (t = 1979, 1980, …, 1992) and residing in state s. Personal characteristics included are education, current age, current age squared, tenure on the lost job, the natural logarithm of the lost job weekly wage rate, and dummies for race, gender, marital status, metropolitan area residence status, and Hispanic origin. We use six education categories – no high school, high-school dropout, high-school graduate, some college, college graduate, and advanced degree. The omitted category is high-school graduate. The state unemployment rate at the time of displacement, U_{st}^{RATE}, is included as a proxy for the local labor market condition, which affects the likelihood of re-employment. To control for time-invariant state of residence characteristics, such as unemployment benefits generosity, $\boldsymbol{Z}_i$ includes state of residence dummies, σ_s. Year of displacement and year of the survey dummies, τ_t and δ_k, are added to absorb annual economy-wide shocks in the year of displacement and year of the survey. Finally, $\boldsymbol{Z}_i$ also includes an indicator variable Ag_{ikjst} that indicates if worker i was displaced from the agricultural sector as opposed to from the manufacturing sector. Based on the theory developed in section II, in particular hypothesis (H2), we would expect that the Ag_{ikjst} would have a negative impact on unemployment duration.

While allowing for unobserved heterogeneity in our hazard specification is not necessary – it will not change the mean effects of the observed covariates, which is our main interest, but only the error distribution (see Wooldridge 2002) – we do so to check for robustness. In particular, we add unobserved Gamma heterogeneity to the Weibull specification (14), resulting in a hazard of the following form

$$\phi(t;\mathbf{Z}_i,\boldsymbol{\eta}) = \nu_i \exp(\mathbf{Z}_i\boldsymbol{\eta})\varphi\, t^{\varphi-1} \qquad (15)\,,$$

where v_i ~ Gamma (θ, θ), with E(v_i)=1 and Var(v_i)=1/θ. For inference, we calculate robust standard errors.

Unlike the re-employment wage, the jobless spell duration is observed for all displaced workers both re-employed and those still looking for a job at the date of the interview. For the latter group, we only observe interrupted (right-censored) spells, which were accommodated in the likelihood function. Hence, problems associated with selection based on worker's re-employment status do not arise in the analysis of unemployment duration. Recall bias, on the other hand, may potentially bias the estimates both in the jobless spell duration regressions as well as in the re-employment wage regressions. If erroneous recall is assumed to behave as a classical measurement error, in the context of the linear re-employment wage regressions in the next section, the effects of such a bias would be toward zero, which implies that the magnitudes of the estimated coefficients may be biased downward and the true effects are even more pronounced than the estimates indicate.

V.2 Re-employment Wage

Specification (16) below estimates the impact of displacement from the agricultural sector relative to the impact of displacement from the manufacturing sector on the worker's post-displacement wage. The dependent variable is the logarithm of the weekly re-employment wage for a worker *i* surveyed in year *k* (*k* = 1984, 1986, …, 1992) displaced from sector *j* (*j* = agriculture, manufacturing) in year *t* (*t* = 1979, 1980, …, 1992) and residing in state *s*.

$$ln(w_{ikjst}^{re-employment}) = \beta_0 + \mathbf{X}_{ikjst}\boldsymbol{\beta}_1' + \beta_2 U_{st}^{RATE} + \sigma_s + \tau_t + \delta_k + \beta_3 Ag_{ikjst} + \varepsilon_{ikjst} \qquad (16),$$

where, as before, $\mathbf{X}_{ikjst}$ is a vector of personal characteristics that includes the same set of covariates used in the unemployment duration analysis. Similarly, equation (16) includes state of residence, σ_s, fixed effects. Year of displacement and year of the survey dummies, τ_t and δ_k, absorb annual economy-wide shocks in the year of displacement and year of the survey. Note that as we include both year of displacement and year of the survey dummies, we effectively control for the time since displacement.

Because not every displaced worker is re-employed by the date of the survey, we do not have information on the re-employment wage for those who are still unemployed at that date. As a result, we estimate the re-employment equation (16) for those who are employed at the time of the interview, but we also show that the censoring and the potential selection problem do not appear to affect the results much. Note that we include both year of the survey and year of displacement dummies in the re-employment wage equation (16). Because they control for the length of time between the date of the survey and the date of displacement, which is intrinsically associated with the re-employment censoring (selection) mechanism, these dummies alleviate selection concerns.[13] Also, to avoid potential selection issues, instead of estimating equation (16) by OLS, we employ the Heckman correction procedure (see Heckman 1979) which delivers consistent estimates in the presence of selection. Because they are intrinsically associated with the censoring mechanism, to satisfy the exclusion restriction requirement in the Heckman procedure, we follow Addison and Portugal (1989) and drop the year of displacement dummies from the re-employment wage equation, but use them in the probit equation to predict the likelihood of re-employment by the date of the survey.[14] We estimate both equations, the sample

[13] The difference between the year of the survey and the year of displacement is the length of the period (in years) since displacement, which is associated with the censoring mechanism as those who were more recently displaced would have had less time to locate a job by the time of the DWS interview.

[14] We also estimated the Heckman model using reasons for displacement as additional explanatory variables in the first stage probit equation, but excluded it from the second stage (the

selection equation (the probit re-employment equation) and the re-employment wage equation jointly using maximum likelihood. Given the difficulty of finding a good exclusion restriction, however, the Heckman model estimates should be interpreted with caution. Nonetheless, the results from this correction procedure are very similar to the results from the equation (16) using displaced workers who are employed at the date of the survey.

VI. Results and Discussion

Previous research on displacement has shown that workers with lower skill and less education also have lower probabilities of leaving the unemployment pool, resulting in longer jobless spell duration (Farber 2005; Addison and Portugal 1989). Our results are consistent with these earlier studies; we find that displaced workers with less than a high school education spend significantly more time unemployed compared to high school graduates, while workers with undergraduate or graduate degrees spend much less time out of work following displacement (see Table 3, column 1). Displaced workers who are non-white have jobless spells that are 42 percent longer than those of displaced white workers.[15] Being married is associated with a

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re-employment wage equation). The reasons for displacement plausibly affect the probability of re-employment at the time of the interview, but they should not affect the re-employment wage rate. The results with this alternative Heckman specification are nearly identical to the Heckman model results reported in the text.

[15] The reported coefficients from the Weibull model of unemployment duration are semi-elasticities. To calculate the percentage change in unemployment duration resulting from a discrete change in an indicator variable (such as the education dummies), one needs to

shorter jobless spell for displaced males, but displaced married females spend more time unemployed than displaced single females, as evidenced by the large and positive coefficient (0.40) on the interaction term between Female and Married. The economic conditions in the state of residence significantly affect the jobless spells of displaced workers – as state unemployment rate increases, so does the duration of unemployment. Displaced workers of Hispanic origin experience about 14 percent longer jobless spells. These results are very similar when the Weibull specification is modified to allow for gamma heterogeneity (Table 3, column 2).

Displacement from agriculture is associated with significantly shorter jobless spells. On average, workers displaced from agriculture spent about 25 percent ($e^{0.22} - 1 = 0.25$, see Table 3, column 1) less time unemployed following displacement than do manufacturing workers. Evaluated at the sample mean of 22.40 weeks of unemployment duration (11.20 two-week intervals, see Table 2), this means that agricultural workers have jobless spells that are 5.6 weeks shorter than those of manufacturing workers. The estimated impact of Ag_{ikjst} on the unemployment duration allowing for Gamma heterogeneity in the Weibull model is slightly more pronounced at 28 percent ($e^{0.25} - 1 = 0.28$, see Table 3, column 2), but still quite similar to the baseline specification of 25 percent (Table 3, column 1). These results confirm our theoretical hypothesis (H2), which asserts that workers displaced from the agricultural sector experience a shorter spell of unemployment than do workers displaced from the manufacturing sector.

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exponentiate the reported (semi-elasticity) coefficient and subtract 1. For the coefficient on Non-white (0.35) from Table 3, column 1, the percentage change is given by $e^{0.35} - 1 = 0.42$.

In addition to the duration of unemployment, the other post-displacement outcome of interest is the re-employment wage. Recall that it is only observed for the portion of the displaced workers who are re-employed at the time of the survey. Column 1 of Table 4 presents the results for the re-employment wage using OLS and controlling from time since displacement, one of the most important predictors of re-employment. Columns 2 and 3 of Table 4 present the results for the Heckman selection model, which combines two equations – the re-employment wage equation and the probit selection model that predicts the probability of re-employment at the time of the survey.

Educational attainment has a strong influence on re-employment wages; on average, those with less than a high school diploma have significantly lower re-employment wages, while those with college or advanced degrees have significantly higher re-employment wages than do high school graduates. An additional year of tenure on the previous job results in a one percent decrease in re-employment wages, highlighting the importance of job-specific and possibly industry-specific human capital. The pre-displacement wages are an important predictor of the re-employment wages in both specifications. Displaced females, non-whites, and Hispanics have lower re-employment wages than displaced white males. Married men have higher wages upon re-employment than do single men, but married women have lower re-employment wages than single women. As expected, local labor market conditions significantly affect re-employment wages and the likelihood of re-employment – higher state unemployment reduces both the re-employment wage and the probability of re-employment. The signs on all of these coefficients are the same across the two specifications, though most of the magnitudes are smaller when the Heckman selection correction is employed. In both specifications, workers displaced from the agricultural industry face re-employment wages that are 10 percent lower than those of manufacturing workers, confirming our theoretical hypothesis (H1).

While our findings – that workers displaced from agriculture have lower re-employment wages and shorter jobless spells – are consistent with our theoretical hypotheses (H1 and H2), the evidence in Tables 3 and 4 does not directly indicate that the lack of unemployment insurance for agricultural workers is the cause for these findings. To further explore the effect that unemployment benefits have on the post-displacement outcomes of agricultural workers, we take advantage of the policy experiment provided by cross-state variation in the requirements for agricultural employers to participate in the Unemployment Insurance program. We estimate the impact of displacement from the agricultural sector (compared to displacement from manufacturing) on the unemployment duration and re-employment wages separately for workers in states that require a larger proportion of farms to participate in the UI program and also for workers in states that enforce only the minimum federal requirements. Additionally, we use the entire sample but include an interaction variable for workers displaced from agriculture and living in the less generous states with only the minimum requirements. The results from these specifications are presented in Table 5. They are consistent with the theory that UI availability does have a significant effect on the duration of unemployment and re-employment wages for agricultural workers.

In more generous states such as California and Texas, where a greater proportion of farms are required to participate in the UI program, agricultural workers have jobless spells that are 15 percent shorter than those of manufacturing workers (Table 5, column 1), but in states where many more farms are exempt from the UI tax, unemployment duration is 51 percent shorter, a difference of about 8.06 weeks (when evaluated at the sample mean). For re-employment wages, our theory predicts that the presence of an unemployment benefits will increase the reservation wage, thus driving up the re-employment wage that the displaced workers will accept. Employing the cross-state variation in UI participation validates the re-employment wage prediction, as well. In states where smaller farms are required to pay UI premiums for their workers, such that displaced agricultural and manufacturing workers face more similar likelihood of unemployment benefits eligibility, the re-employment wages for displaced agricultural workers are not significantly different from those of manufacturing workers (Table 5, column 4). However, the re-employment wage penalty for displaced agricultural workers in less generous states, where only the farms meeting the federal requirements are participating in the UI program, is nearly 25 percent (Table 5, column 5). The lack of access to unemployment benefits drives down the reservation wage for agricultural workers, and leads them to accept re-employment wages that are significantly lower than those of manufacturing workers. Pooling all the data across more and less generous states reveals the same pattern (Table 5, columns 3 and 6) – in states with less strict UI tax provisions for agricultural employers, workers displaced from agriculture face longer jobless spells and lower re-employment wages compared to workers displaced from the manufacturing sector.

VII. Conclusion

We analyze post-displacement outcomes, both the unemployment duration and the re-employment wage, for workers displaced from agricultural employment. In general, unlike manufacturing employers, "small" agricultural employers are not required to participate in the Unemployment Insurance system, which leaves many farm workers ineligible to receive unemployment benefits. Our theoretical job search model suggests that displaced workers who face lower or no unemployment benefits also experience shorter unemployment duration spells and lower re-employment wages. We test these predictions with a sample of workers displaced from the agricultural and the manufacturing sector obtained from the Displaced Workers' Supplement to the Current Population Survey. Unlike workers displaced from manufacturing, workers who have lost an agricultural job are much less likely to qualify for and receive unemployment benefits, and so face significantly shorter unemployment duration spells and lower re-employment wages. In particular, our estimates imply that displaced agricultural workers experience 25 percent (5.6 weeks) less time unemployed and upon re-employment earn 10 percent less than otherwise similar manufacturing workers. Finally, we present evidence that in states with stricter UI tax provisions for agricultural employers, displaced agricultural workers face unemployment duration and re-employment wages of similar magnitude to those of displaced manufacturing workers. However, in states with less strict UI tax provisions for agricultural employers, workers displaced from agriculture face longer jobless spells and lower re-employment wages compared to workers displaced from the manufacturing sector.

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TABLES

Table 1. Number of Agricultural Workers and State UI Provisions for Agricultural Workers on "Small" Farms, States Ranked by No. of Ag. Workers.

Rank	State	No. Ag. Workers	Payroll ($1,000)	No. Farms	More Generous State UI Provisions for Ag. Workers on "Small" Farms ($LessGenAgUI_s$ =0)
1	California	535,256	4,317,078	34,342	Yes
2	Washington	262,528	987,399	13,598	Yes
3	Texas	166,117	969,979	49,206	Yes
4	Oregon	122,845	620,422	10,978	Yes
5	Florida	118,581	1,157,569	10,672	Yes
6	Kentucky	115,177	291,881	24,882	No
7	North Carolina	97,138	552,486	16,091	No
8	Minnesota	95,055	459,332	22,623	No
9	Michigan	86,855	469,731	12,279	No
10	Iowa	82,991	409,190	28,135	No
	United States	3,036,470	18,568,446	554,434	

Note: Data from the 2002 Census of Agriculture.

Table 2. Summary Statistics, Displaced Worker Supplement (DWS) to the Current Population Survey (CPS), 1979-1992.

Variable	Agriculture		Manufacturing		All	
	Mean	St. Dev.	Mean	St. Dev.	Mean	St. Dev.
Agriculture	1	0	0	0	0.07	0.26
Manufacturing	0	0	1	0	0.93	0.26
Reason for Displacement						
Establishment Closed	0.28	0.45	0.42	0.49	0.41	0.49
Insufficient Work	0.31	0.46	0.38	0.48	0.37	0.48
Position Abolished	0.13	0.34	0.12	0.33	0.12	0.33
Seasonal Job Ended	0.12	0.33	0.01	0.10	0.02	0.13
Self-employed Business Failed	0.05	0.21	0.00	0.07	0.01	0.09
Other Reason	0.11	0.31	0.06	0.24	0.07	0.25
Re-employment Weekly Wage (2003 dollars)	430.63	282.40	581.97	384.24	572.01	380.20
Pre-displacement Weekly Wage (2003 dollars)	451.54	311.26	658.53	401.48	644.01	399.31
Hours Last Week	35.72	18.31	36.69	16.47	36.62	16.60
Unemployment Benefit Receipt	0.36	0.48	0.60	0.49	0.58	0.49
Unemployment Duration (two-week intervals)	8.31	10.44	11.42	12.97	11.20	12.83
Education						
No High School	0.17	0.38	0.07	0.25	0.08	0.27
High School Drop-out	0.15	0.36	0.14	0.34	0.14	0.34
High School	0.31	0.46	0.39	0.49	0.38	0.49
Some College	0.24	0.43	0.26	0.44	0.26	0.44
College	0.09	0.29	0.10	0.30	0.10	0.30
Advanced Degree	0.03	0.16	0.05	0.22	0.05	0.21
Age	33.92	10.49	37.54	11.12	37.28	11.11
Tenure	3.50	4.82	5.30	6.60	5.17	6.51
Female	0.30	0.46	0.34	0.47	0.34	0.47
Non-white	0.08	0.27	0.14	0.35	0.14	0.34
Married	0.60	0.49	0.63	0.48	0.63	0.48
Metropolitan	0.62	0.49	0.74	0.44	0.73	0.44
State Unemployment Rate (percentage points)	7.31	2.18	7.21	2.35	7.22	2.34
Non-white Female	0.03	0.17	0.06	0.23	0.06	0.23
Married Female	0.18	0.38	0.19	0.39	0.19	0.39
Hispanic Origin	0.27	0.44	0.13	0.34	0.14	0.35
Moved to Find Current Job	0.28	0.45	0.22	0.42	0.23	0.42
Left Sector of Displacement	0.66	0.47	0.47	0.50	0.48	0.50
No. Obs.	306		4,041		4,347	

Table 3. Unemployment Duration.

Variable	Unemployment Duration: Weibull Model (1)	Unemployment Duration: Weibull Model with Gamma Heterogeneity (2)
No High School	0.16**	0.23***
	(0.08)	(0.08)
High School Drop-out	0.30***	0.29***
	(0.06)	(0.07)
Some College	-0.06	-0.10*
	(0.06)	(0.06)
College	-0.21**	-0.22**
	(0.08)	(0.09)
Advanced Degree	-0.25***	-0.36***
	(0.08)	(0.09)
Age	0.02	0.03**
	(0.01)	(0.01)
Age^2 x 1,000	-0.01	-0.11
	(0.14)	(0.13)
Tenure	0.004	0.003
	(0.003)	(0.003)
ln(Pre-displacement Wage)	-0.01	0.08*
	(0.03)	(0.04)
Female	-0.04	-0.01
	(0.06)	(0.05)
Non-white	0.35***	0.43***
	(0.07)	(0.08)
Married	-0.21***	-0.28***
	(0.05)	(0.05)
Metropolitan	0.11**	0.07
	(0.05)	(0.05)
State Unemployment Rate	0.11***	0.12***
	(0.02)	(0.02)
Non-white Female	0.02	-0.03
	(0.12)	(0.11)
Married Female	0.40***	0.42***
	(0.06)	(0.07)
Hispanic Origin	0.13**	0.19***
	(0.06)	(0.07)
Ag_{ikjst}	-0.22***	-0.25***
	(0.06)	(0.05)
Weibull Duration Dependence Parameter, φ	0.94***	1.35***
	(0.01)	(0.05)
Variance of Gamma Heterogeneity, $1/\theta$	-	0.87***
		(0.11)
Log Likelihood	-6,659.43	-6,558.30
No. Obs.	4,347	4,347

Note: All coefficients are semi-elasticities. All specifications include state, year of the survey and year of displacement dummies. Robust standard errors are reported in parentheses. *** Indicates significance at 1 percent, ** at 5 percent, and * at 10 percent.

Table 4. Re-employment Wage.

Variable	*ln*(Re-employment Wage) OLS (1)	*ln*(Re-employment Wage) Heckman Selection Model (2)	Re-employed Heckman Selection Model (3)
No High School	-0.21***	-0.15***	-0.26***
	(0.03)	(0.05)	(0.09)
High School Drop-out	-0.11***	-0.03	-0.27***
	(0.03)	(0.04)	(0.06)
Some College	0.08**	0.06**	0.13**
	(0.03)	(0.03)	(0.05)
College	0.23***	0.16***	0.40***
	(0.05)	(0.04)	(0.09)
Advanced Degree	0.35***	0.27***	0.34***
	(0.05)	(0.06)	(0.12)
Age	0.04***	0.03***	-0.01
	(0.01)	(0.01)	(0.01)
Age^2 x 1,000	-0.49***	-0.41***	0.01
	(0.08)	(0.08)	(0.14)
Tenure	-0.01	-0.01***	0.01***
	(0.00)	(0.00)	(0.00)
ln(Pre-displacement Wage)	0.43***	0.42***	-0.12***
	(0.04)	(0.02)	(0.04)
Female	-0.09**	-0.14***	0.15**
	(0.04)	(0.04)	(0.07)
Non-white	-0.17***	-0.06	-0.38***
	(0.05)	(0.04)	(0.08)
Married	0.14***	0.06**	0.27***
	(0.03)	(0.03)	(0.06)
Metropolitan	0.07**	0.03	0.18***
	(0.03)	(0.03)	(0.06)
State Unemployment Rate	-0.02***	-0.00	-0.05***
	(0.01)	(0.01)	(0.02)
Non-white Female	0.08	0.06	0.05
	(0.06)	(0.07)	(0.12)
Married Female	-0.22***	-0.16***	-0.26***
	(0.04)	(0.05)	(0.09)
Hispanic Origin	-0.09***	-0.02	-0.30***
	(0.03)	(0.04)	(0.07)
Ag_{ikjst}	-0.10***	-0.10**	0.13
	(0.03)	(0.05)	(0.09)
Inverse Mill's Ratio	-	-0.60***	-
		(0.02)	
R^2	0.39	-	-
Log Likelihood	-	-4,641.66	
No. Obs.	3,153	4,347	

Note: Specifications (1) and (3) include state, year of the survey, and year of displacement dummies. Specification (2) only includes state and year of the survey dummies. Robust standard errors are reported in parentheses. *** Indicates significance at 1 percent, ** at 5 percent, and * at 10 percent.

Table 5. Unemployment Duration and Re-employment Wages – More and Less Generous UI States

Variable	Unemployment Duration			*ln*(Re-employment Wage)		
	More Generous	Less Generous	All	More Generous	Less Generous	All
	(1)	(2)	(3)	(4)	(5)	(6)
No High School	0.08	0.38***	0.16**	-0.23***	-0.14	-0.21***
	(0.09)	(0.11)	(0.08)	(0.04)	(0.09)	(0.03)
High School Drop-out	0.33***	0.23***	0.31***	-0.07*	-0.16**	-0.11***
	(0.08)	(0.09)	(0.06)	(0.04)	(0.06)	(0.03)
Some College	-0.06	-0.08	-0.06	0.07**	0.07	0.08**
	(0.07)	(0.07)	(0.06)	(0.04)	(0.05)	(0.03)
College	-0.12	-0.38**	-0.21**	0.23***	0.23***	0.23***
	(0.09)	(0.15)	(0.08)	(0.06)	(0.09)	(0.05)
Advanced Degree	-0.22**	-0.34**	-0.25***	0.32***	0.40***	0.35***
	(0.10)	(0.17)	(0.08)	(0.06)	(0.07)	(0.05)
Age	0.03**	0.01	0.02*	0.03***	0.05***	0.04***
	(0.01)	(0.02)	(0.01)	(0.01)	(0.01)	(0.01)
Age^2 x 1,000	-0.08	0.02	-0.01	-0.39***	-0.69***	-0.49***
	(0.14)	(0.23)	(0.14)	(0.09)	(0.14)	(0.08)
Tenure	0.01*	-0.01	0.01	-0.01	-0.01	-0.01
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
ln(Pre-displacement Wage)	-0.01	-0.02	-0.00	0.44***	0.40***	0.43***
	(0.04)	(0.06)	(0.03)	(0.05)	(0.05)	(0.04)
Female	-0.06	-0.02	-0.05	-0.05	-0.13*	-0.09**
	(0.07)	(0.09)	(0.06)	(0.05)	(0.07)	(0.04)
Non-white	0.24***	0.49***	0.35***	-0.12***	-0.26**	-0.17***
	(0.08)	(0.12)	(0.07)	(0.04)	(0.11)	(0.05)
Married	-0.22***	-0.19***	-0.22***	0.10***	0.21***	0.14***
	(0.06)	(0.07)	(0.05)	(0.03)	(0.06)	(0.03)
Metropolitan	0.09	0.16**	0.11**	0.12***	0.04	0.07**
	(0.07)	(0.08)	(0.05)	(0.04)	(0.03)	(0.03)
State Unemployment Rate	0.11***	0.08**	0.11***	-0.03**	-0.03	-0.02***
	(0.03)	(0.04)	(0.02)	(0.01)	(0.02)	(0.01)
Non-white Female	-0.14	0.16	0.02	0.01	0.18	0.08
	(0.13)	(0.17)	(0.12)	(0.06)	(0.12)	(0.06)
Married Female	0.49***	0.27***	0.40***	-0.20***	-0.27***	-0.22***
	(0.10)	(0.10)	(0.06)	(0.05)	(0.07)	(0.04)
Hispanic Origin	0.12**	0.16	0.13**	-0.10***	0.23	-0.09***
	(0.06)	(0.20)	(0.06)	(0.04)	(0.20)	(0.03)
Ag_{ikjst}	-0.14*	-0.41***	-0.15*	-0.03	-0.22***	-0.05*
	(0.08)	(0.11)	(0.08)	(0.04)	(0.08)	(0.03)
$Ag_{ikjst} * LessGenAgUI_s$	-	-	-0.26*	-	-	-0.17**
			(0.16)			(0.08)
R^2	-	-	-	0.41	0.36	0.39
Log Likelihood	-3,999.11	-2,622.02	-6,657.98	-	-	-
No. Obs.	2,631	1,716	4,347	1,863	1,290	3,153

Note: All specifications include state, year of the survey and year of displacement dummies.
*** Indicates significance at 1 percent, ** at 5 percent, and * at 10 percent.