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The Geographic Mobility of Displaced Workers: Do Local Employment Conditions Matter?

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Abstract

Using data drawn from the National Longitudinal Survey of Youth 1979, the Bureau of Labor Statistics, and County Business Patterns, this study investigates whether displaced workers adjust their job search strategies in response to local market conditions to favor migration out of declining labor markets. Empirical results from a Cox partial-likelihood proportional hazards model are supportive. A low density of local employment and low average wage levels are associated with shorter wait times to migration. Conversely, high local employment growth rates, high wages, and low unemployment rates correlate with an increased likelihood of obtaining local employment following displacement.

Keywords: Migration; Unemployment duration; Job search

JEL classification: J61; J64; R23

1. INTRODUCTION

With firms announcing nearly two million layoffs in 2001, the U.S. economy sank into its first recession in a decade. Although modest by historical standards, the 5.7 percent unemployment rate reported in December 2001 nevertheless represented more than eight million unemployed American workers. Even more troubling, as the economy entered the incipient stage of recovery, many businesses continued to shed jobs. The Bureau of Labor Statistics (BLS) reported that in the subsequent twelve-month period non-farm employment fell by an additional 400,000 jobs. More than two years after the recession trough, the unemployment rate stood at a nearly identical 5.6 percent. In contrast to previous recessions, part of the lag in employment recovery appears to be due to the fact that layoffs are more likely to be permanent with firms not planning to rehire those laid off (Schreft and Singh 2003). This has contributed to a particularly acute problem of long-term unemployment among the ranks of the displaced. This distinctive lack of job growth has prompted market analysts and economists to describe the current recovery as "jobless."

The absence of sustained job creation in the wake of mass layoffs has been a trying point of concern for many displaced workers and their families. According to the BLS, in February 2004 the average spell of unemployment was measured at 20.3 weeks without work, marking the highest average duration of joblessness in over 20 years. Because long-term unemployment typically occurs in markets where wages and employment are generally below the national average, much of the subsequent economic hardship may be attributed to a dearth of employment opportunities in the local labor market (Carrington 1993). As a consequence, geographically mobile workers may be in a position to substantially mitigate the economic loss following displacement by migrating to healthier labor markets. By extended job search efforts geographically, workers are able to access a wider range of employment opportunities, an essential part of any optimal job search strategy when new jobs are being created in other parts of the country. Such mobility also affords a displaced worker the opportunity to be more selective when evaluating reemployment opportunities, allowing for improved matching between workers and firms.

Researchers have long attempted to assess the role of migration in unemployment outcomes. Although it is well established that the unemployed are significantly more likely to undertake migration than employed workers (DaVanzo 1978; Herzog and Schlottmann 1984), there is considerable ambiguity in the empirical literature on the relationship between migration and the *duration* of unemployment (Herzog, Schlottmann, and Boehm 1993). The most recent lines of inquiry typically include a binary indicator of migrant status in a hazard model specification of unemployment duration.[1] Bailey (1994) finds inter-county migration to be associated with longer spells of unemployment. Goss, Paul, and Wilhite (1994) estimate a two-stage model of unemployment duration that corrects for migrant selectivity with similar results. In contrast, Boehm, Herzog, and Schlottmann (1998) find some evidence that interstate migration reduces the duration of unemployment and leads to more successful reemployment outcomes.

Although this modeling approach offers some merit, such "single-risk" specifications are unsatisfactory to the extent that workers change search strategies throughout the duration of unemployment. It is the contention of this paper that the appropriate modeling approach is that of a competing-risks specification that explicitly models search across multiple markets. In this setting, migration can be associated with either reduced or extended spells of unemployment. Individuals searching simultaneously over several labor markets may reduce the duration of unemployment by increasing the number of available jobs from which to sample during a given period of time. This not only increases the likelihood of migration but also may substantially reduce the length of time required to find an acceptable offer. In contrast, extended geographic search may only be optimal after displaced workers have exhausted local search possibilities, meaning that migrations will typically be associated with longer spells of unemployment when compared to those reemployed locally (McCall and McCall 1987). Hence, the establishment of simple correlations between migration and unemployment duration provides scant information on the actual search strategies of unemployed workers or the "success" of unemployed search.

Search theory predicts that local labor market conditions will affect both the duration of unemployment and the likelihood of migration. When searching over multiple labor markets, displaced workers must allocate search resources between alternative markets. When local job prospects are good, an optimal search strategy dictates that more resources will be devoted to search in the local labor market, with migration becoming a less likely outcome of the search process. Conversely, when local reemployment conditions are poor, workers are predicted to allocate greater search intensity to alternative markets, resulting in an increased likelihood of migration. I test this proposition by estimating a competing-risks hazard model for exits from unemployment in an effort to determine whether displaced workers adjust their job search strategies to favor migrating out of declining labor markets. Using data drawn from the National Longitudinal Survey of Youth 1979, the Bureau of Labor Statistics, and County Business Patterns, I find that a low density of local employment and low average wage levels are associated with shorter wait times to migration. In contrast, high local employment growth rates, high wages,

[1] An alternative research vein examines the impact of unemployment duration on migration propensity. Among these studies, Goss and Schoening (1984) find that the likelihood of migration increases with the duration of unemployment. Applying a non-linear specification, McHone and Rungeling (1993) find that likelihood of migration increases with duration up to a certain point, but declines thereafter.

and low unemployment rates are associated with an increased likelihood of obtaining local employment following displacement.

2. THEORETICAL FRAMEWORK

Consider a displaced worker searching for employment in period t. In standard models of job search, the individual is posited to choose between continued unemployed search and accepting a wage offer in each period (Mortensen 1986). Let wage offers be drawn from a known distribution, $F(w)$, with offer arrival rate $\alpha(t)$. Assume the length of a period is sufficiently small such that the probability of receiving more than one job offer in a single period is zero. Define the reservation wage $w^r(t)$ as the wage that equates the discounted present value of accepting a wage offer and of continuing search in period t. The probability of exiting from unemployment at time t is equal to the product of the offer arrival rate and the probability that the offer is acceptable. In a model of single market job search, the optimal strategy then is to accept the first wage offer that exceeds the reservation wage.

As illustrated by Fallick (1993), search becomes more complicated when workers draw wage offers from multiple markets.[2] Let the worker now be endowed with a certain capacity of search "intensity" (e.g., time, resources, or effort), with the proportion of intensity devoted to search in each location j denoted by $s_j(t)$. The probability that the worker will receive a job offer in period t is then $\alpha(t)\phi(s(t))$, where ϕ is an increasing concave function with $\phi(0) = 0$ and $\phi(1) = 1$. The hazard rate from unemployment to employment in market j can then be written as

$$(1) \qquad h_{i,j}(t) = \alpha_j(t, X_i)\phi\left(s_j(t, X_i)\right)\left[1 - F_j\left(w_j^r(t, X_i); X_i\right)\right]$$

where X_i is a vector of individual characteristics. Search strategies now consist of location-specific reservation wages and search intensity allocation. The key prediction of the competing-risks model is that unemployed workers set their job search strategies to favor finding a job in those markets offering the highest returns to search. That is, although search is conducted simultaneously across labor markets, workers orient search intensity towards those markets offering the most favorable employment prospects.

Conditions in the current labor market will affect the search process in two ways. First, they will directly affect the distribution of search outcomes from any given search strategy. Second, they will also affect the choice of search strategy. Notice when search is conducted over multiple markets, the rate of escape from unemployment via migration

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[2] Fallick (1993) develops a model of simultaneous job search over multiple industrial sectors of the economy. He finds empirical support for the notion that unemployed workers adjust search strategies to favor finding employment in those industries where the more promising conditions prevail. The implications of the model have not been tested in the context of extended geographic search and migration.

(i.e., the "migration hazard") is unaffected by the conditions in the current labor market except through their effect on the choice of search strategy. Thus, any impact of local employment conditions found in the migration hazard can be inferred to result from a given choice of search strategy. I exploit this notion in the subsequent analysis by observing whether variation in local economic conditions affect the hazard rate of escapes from unemployment via migration.

3. ECONOMETRIC SPECIFICATION

In order to implement the search model empirically, I employ a competing-risks application of the Cox partial-likelihood proportional hazards model. In this model, the instantaneous hazard rate of escape from unemployment at time *t*, conditional on survival to time *t*, can be written as

$$(2) \qquad h(t;Z) = h_0(t)\exp(Z\beta)$$

where $h(t)$ is the hazard of escape from unemployment at time t for an individual described by a vector of coefficients β associated with covariates that characterize the personal and location-specific attributes of individuals in the sample, and $h_0(t)$ is the baseline hazard rate. This approach is attractive because it incorporates both the market-specific components that may favor relocating to an alternative labor market and the person-specific components that tend to inhibit migration. The estimated parameters and their standard errors provide information on the direction and statistical significance of the partial effect of each covariate specified in Z. The risk of exit from unemployment is allowed to vary over time and with variation in the covariates. Positive coefficients on the covariates are indicative of increasing hazard rates and are thus associated with a reduction in the expected time until reemployment.

Although the model is dynamic, the data are recorded in discrete intervals. As a result, there are numerous unemployment spells of the same duration. Duration ties are handled using the Peto-Breslow approximation procedure described in Kalbfleisch and Prentice (1980). This implies the likelihood function can be approximated as

$$(3) \qquad L = \prod_{i=1}^{k} \frac{\exp\left(\sum_{j\in D_i} Z_{ij}\beta\right)}{\left[\sum_{l\in R_i}\exp(Z_l\beta)\right]^{m_i}}$$

where i indexes the (ordered) failure times $t(i)$, (i = 1,…, k), D_i is the set of observations j that fail at $t(i)$, m_i is the number of individuals who exit at $t(i)$, and R_i is the set of all observations l that are at risk to exit at time $t(i)$. Estimates of the parameters β can be obtained by maximizing the associated partial log-likelihood function.[3]

4. DATA

To estimate the model described in the preceding section, I use data drawn primarily from the National Longitudinal Survey of Youth 1979 (NLSY79). The NLSY79 is an extensive longitudinal survey that follows young workers between the ages of 14 and 22 in 1979 on a yearly basis continuous through 1994 and biennially thereafter. Data from the Bureau of Labor Statistics (BLS) and County Business Patterns (CBP) are used to construct measures reflecting the economic conditions in the current labor market.

In constructing a sample suitable for study I use data collected in annual interviews between 1979 and 1994. The sample is restricted to non-military male workers displaced from full-time jobs having lasted at least six months.[4] In addition, I exclude workers who fail to exit formal schooling or report valid schooling information (such as the highest grade completed). The NLSY79 data is well suited for this study because it provides detailed information about each respondent's employment behavior. Each year, respondents are asked to provide start and stop dates for each job held (up to a maximum of five) as well as a multitude of characteristics describing each job. Using the NLSY79 work history files, I link all jobs held subsequent to the date of exit from formal schooling and identify those jobs ending in displacement due to layoff or plant closing. The duration measure used in this study is the number of weeks in which the individual is without a job from the time of displacement until either a new job is found or survey attrition (with the latter case treated as a censored duration). In order to maximize the sample size, respondents are allowed to contribute multiple observations (spells) to the sample. In all, I analyze 2,419 spells of unemployment (contributed by 1,478 men) lasting on average 26.6 weeks including those spells censored due to attrition or incomplete information.

The NLSY79 data files contain information on a variety of personal, family, and job-related characteristics for each respondent. Personal and family controls include race, education level (*HGC*), marital status, children, home ownership, residence in a rural

[3] A highly desirable feature of the Cox model is that it allows for non-parametric treatment of the baseline hazard, relaxing the need for assumptions regarding the shape of the baseline hazard. Unfortunately, this specification does not explicitly account for the potential effect of unobserved heterogeneity on the hazard rate. To account for individual heterogeneity, I include a broad array of covariates in the hazard model meant to control for personal, job, and family-related characteristics. Moreover, empirical work by Meyer (1990) suggests that explicitly modeling unobserved heterogeneity has little effect on the estimated coefficients in a model in which the baseline hazard rate is allowed to be non-parametric.

[4] This restriction is introduced because women are much more likely to leave the labor force as a result of job displacement. Maxwell and D'Amico (1986) find that some 40 percent of women but only 3.8 percent of men exit the labor force by the first survey post-displacement. Mass attrition from the displaced sample unrelated to economic conditions may bias the results.

area, and percentile scores for the Armed Forces Qualifications Test (*AFQT*).[5] Job-related controls include previous job tenure, experience, union membership, and public sector employment. *UI Benefit* is the average weekly state-specific unemployment benefit for which the respondent qualifies at the time of displacement. In addition, selected model specifications include a complete set of industry and occupation controls.

Integral to this study is the influence of local employment characteristics on the search behavior of displaced workers. Data from the Bureau of Labor Statistics (BLS) and County Business Patterns (CBP) are used to construct variables reflecting the economic conditions in the current labor market. In all, I construct four alternative measures: *Employment Level* is the state employment level in the year of displacement as measured by the BLS. The employment level is used to proxy for that part of the offer arrival rate common to all workers searching in the current labor market. A superior offer arrival rate should make the need for migration (i.e., the need to search alternative labor markets) less likely, thus reducing the length of time searching in the current market. Hence, *Employment Level* is expected to have a negative effect on the migration hazard and a positive effect on the local employment hazard. *Employment Growth* is the percentage change in the state employment level during the year of displacement as measured by the BLS. This also proxies the offer arrival rate and so provides a measure of the robustness of the current market. The effect should be in the same direction as *Employment Level*. *Unemployment* is the state unemployment rate in the year of displacement as measured by the BLS. This variable is meant to provide a measure of slackness in the current labor market. Finally, *Wage Level* is the average weekly earnings in the state of residence at the time of displacement as provided in the CBP release. This is meant to proxy that part of the expected wage offer that is common to all workers within the current labor market. A superior wage offer distribution in the current labor market will reduce the length of time searching in the current market while extending the duration of search in alternative markets.

Table 1 presents the mean and standard deviation for each covariate used to estimate the model. Columns 1 and 2 refer to all spells of unemployment, including those spells censored due to attrition or incomplete information. Columns 3 and 4 pertain only to those unemployment spells ending in migration, as defined by a change in state of residence.[6] Columns 5 and 6 refer only to those spells ending in reemployment in the current (local) labor market. At the bottom of the table are both the total number of unemployment spells (*Number of Spells*) and the mean duration of unemployment measured in weeks (*Unemployment Duration*). Notice that the mean spell of unemployment is con-

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[5] Approximate AFQT scores are constructed from selected scores from the Armed Services Vocational Aptitude Battery, which was administered in 1980 to 94 percent of the original NLSY sample. AFQT scores are considered to be useful measures of skills valued in the workplace (as well as in the armed forces). Because AFQT scores are unavailable for about 5 percent of men, for those missing I set the AFQT variable equal to the sample mean and set an "AFQT missing" variable equal to one.

[6] Blanchard and Katz (1992) suggest that out-of-state migration is the primary adjustment mechanism by which states recover from employment shocks.

siderably longer when ending in reemployment involving migration. For many displaced workers, migration is likely to be a strategy of last resort, occurring only after local reemployment efforts are exhausted. Moreover, migration necessarily entails a higher reservation wage due to moving expenditures and high search costs. A high reservation wage exerts a negative effect on the hazard rate from unemployment. It is important to remember, however, that the migrant sample may have disproportionately left labor markets with the worst reemployment opportunities, and thus might have experienced even longer spells of unemployment had they not moved. This, in essence, is the crux of the study.

TABLE 1

Means of Variables Included in the Hazard Models

	All Spells		Spells Ending in Migration		Spells Ending Locally	
Variables	Mean	Standard Deviation	Mean	Standard Deviation	Mean	Standard Deviation
Black	0.322	0.467	0.323	0.468	0.316	0.465
Hispanic	0.202	0.402	0.223	0.417	0.193	0.395
HGC < 12	0.401	0.490	0.424	0.495	0.389	0.488
HGC = 12	0.365	0.481	0.378	0.485	0.365	0.482
13 ≤ HGC ≤ 15	0.167	0.373	0.149	0.357	0.173	0.378
HGC ≥ 16	0.067	0.251	0.048	0.214	0.073	0.260
Experience	6.960	4.100	6.460	3.980	6.910	4.020
Married	0.267	0.443	0.261	0.440	0.268	0.443
Children	0.258	0.438	0.241	0.428	0.261	0.439
Home Owner	0.141	0.141	0.092	0.290	0.153	0.360
AFQT	30.150	26.060	28.900	25.170	30.790	26.330
AFQT Missing	0.035	0.184	0.039	0.196	0.034	0.181
Rural	0.203	0.403	0.225	0.418	0.198	0.399
Tenure† (weeks)	89.250	83.610	90.440	80.980	89.250	82.440
Union†	0.199	0.399	0.201	0.401	0.200	0.400
Public Sector†	0.113	0.317	0.072	0.258	0.114	0.318
UI Benefit	125.530	15.240	125.110	15.180	125.650	15.280
Emp. Level‡	4602.500	3343.300	4105.600	3179.400	4700.000	3349.300
Emp. Growth‡	1.690	2.280	1.580	2.340	1.740	2.280
Unemployment‡	7.090	2.010	7.040	2.140	7.120	1.980
Wage Level‡ (weekly)	279.570	38.390	271.690	37.180	281.280	38.310
Unemp. Duration	26.640	43.270	35.900	58.770	23.760	38.040
Spells	2419		502		1838	

† Data pertains to pre-displacement job.
‡ Data measured at the start of unemployment spell.

To further explore this difference, Table 2 presents Kaplan-Meier empirical hazard estimates of the wait time until reemployment for both local and migratory reemployment.[7] Both curves take the familiar hyperbolic shape for unemployment duration with the empirical hazard for local reemployment everywhere higher than the migration hazard. The likelihood of exiting the unemployed state via local employment is four times more likely than by migratory reemployment in the first three weeks post migration. In the intermediate intervals, the likelihood of local reemployment is only three times greater than that for migration. In the later intervals, the local hazard is roughly twice the migration hazard.

TABLE 2

Kaplan-Meier Empirical Hazard Estimates of Time to Reemployment

		Local Reemployment Hazard			Migratory Reemployment Hazard		
Elapsed Interval (weeks)	Spells at Risk	Spells Ending Locally	Censored Spells[a]	Empirical Hazard	Spells Ending in Migration	Censored Spells[b]	Empirical Hazard
1-3	2,419	555	122	0.081	112	565	0.020
4-6	1,742	177	37	0.035	34	180	0.007
7-9	1,528	127	42	0.028	40	129	0.010
10-12	1,359	111	34	0.028	32	113	0.009
13-18	1,214	185	65	0.027	56	194	0.009
19-24	964	121	43	0.022	37	127	0.007
25-30	800	132	37	0.029	30	139	0.008
31-36	631	82	31	0.023	22	91	0.007
37-48	518	101	48	0.018	35	114	0.007
49-60	369	76	26	0.018	19	83	0.006
61-72	267	41	21	0.014	16	46	0.006
73-108	205	64	32	0.010	28	68	0.006
109-144	109	27	16	0.008	16	27	0.005
145+	66	39	27	0.003	25	41	0.002
Spells		1,838	581		502	1,917	

[a] Censored unemployment spells in the local reemployment hazard include both spells lost due to sample attrition and spells lost due to migration.

[b] Censored unemployment spells in the migration reemployment hazard include both spells lost due to sample attrition and spells lost due to local reemployment

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[7] In general, a censored duration is an unemployment spell that terminates for some reason other than reemployment (like attrition from the sample). Thus, reemployment is the hazard "risk." When there are competing risks, an unemployment duration is treated as censored when a spell terminates for any reason other than the specific risk being considered. Hence, in the migration hazard, unemployment spells ending in local reemployment are treated as censored in the same way as a spell lost to attrition, and vice-versa.

5. RESULTS

5.1 Hazard Estimates Without Labor Market Controls

Table 3 presents the results obtained from the Cox proportional hazards models before including any labor market variables. The table provides parameter estimates, standard errors, and hazard ratios from four alternative model specifications.[8] Specifications (1) – (4) refer to models of escapes from unemployment in the local labor market. Specifications (5) – (8) refer to models of escapes from unemployment via migration. In each of the competing-risks models, unemployment spells ending in the alternative reemployment category are treated as censored observations.[9]

In all, four alternative specifications are estimated. Columns (1) and (5) pertain to a specification without previous job characteristics of any kind. Columns (2) and (6) refer to a specification that includes previous job characteristics such as completed job tenure, union status, and public sector employment. Columns (3) and (7) add industry and occupation controls. Columns (4) and (8) refer to a specification identical to (3) and (7) but omits the multiple unemployment spell indicator.

Looking first at the local reemployment hazards in Columns (1) – (4), both black and Hispanic workers typically experience extended periods of displacement-induced unemployment (lower hazards) relative to their white counterparts. Workers living in rural areas also tend to experience longer spells of unemployment. On the other hand, homeowners and public sector employees are the most likely candidates to find jobs locally. Home ownership is a good proxy for location-specific capital as well as an indicator of a particular attachment to the community. Such investments might also include information about the quantity and quality of jobs available in the surrounding area that may facilitate efficient job search. Government employment may also involve a certain amount of location-specific capital that would be lost upon migration.

Consistent with theory, higher education levels (*HGC*) are found to be associated with higher local hazard rates. Podgursky and Swaim (1987) suggest that greater educational attainment reduces the potential for earnings loss from displacement while significantly increasing the likelihood of obtaining full-time employment post displacement. *Ceteris paribus*, this implies an unambiguous positive effect on the local reemployment hazard. Interestingly, those respondents contributing multiple unemployment spells

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[8] The hazard ratio is interpreted as the proportional change in hazard resulting from a one-unit increase in the associated variable.

[9] Because this specification of the Cox model treats exits into the alternative market as a censored duration, the estimation procedure does not explicitly control for stochastic dependence between the two hazard rates. In this case, the estimated hazard rates should be interpreted as transition rates conditional on having not yet exited into either market (Fallick 1993).

TABLE 3

Hazard Model Parameter Estimates (No Labor Market Variables)

	Hazard Rate Analysis of Escapes from Unemployment in Local Labor Market				Hazard Rate Analysis of Escapes from Unemployment via Migration			
Covariates	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Black	-0.213***	-0.212***	-0.186***	-0.167***	-0.071	-0.069	-0.032	-0.005
	(0.062)	(0.063)	(0.064)	(0.064)	(0.122)	(0.122)	(0.124)	(0.124)
	[0.808]	[0.809]	[0.830]	[0.846]	[0.932]	[0.934]	[0.969]	[0.995]
Hispanic	-0.148**	-0.145**	-0.134**	-0.129*	0.102	0.096	0.115	0.116
	(0.068)	(0.068)	(0.068)	(0.068)	(0.128)	(0.128)	(0.130)	(0.130)
	[0.863]	[0.865]	[0.874]	[0.879]	[1.108]	[1.101]	[1.122]	[1.123]
HGC	0.052***	0.052***	0.050***	0.048***	-0.020	-0.021	-0.021	-0.023
	(0.015)	(0.015)	(0.015)	(0.015)	(0.028)	(0.028)	(0.029)	(0.029)
	[1.053]	[1.053]	[1.051]	[1.049]	[0.980]	[0.979]	[0.979]	[0.977]
Experience	0.004	0.002	-0.002	0.002	-0.011	-0.011	-0.019	-0.013
	(0.006)	(0.007)	(0.007)	(0.007)	(0.013)	(0.014)	(0.015)	(0.014)
	[1.004]	[1.002]	[0.998]	[1.002]	[0.989]	[0.989]	[0.981]	[0.987]
Married	0.032	0.043	0.037	0.032	0.156	0.135	0.115	0.115
	(0.073)	(0.074)	(0.074)	(0.074)	(0.139)	(0.140)	(0.143)	(0.143)
	[1.032]	[1.044]	[1.038]	[1.032]	[1.168]	[1.144]	[1.122]	[1.122]
Children	-0.050	-0.045	-0.046	-0.048	-0.128	-0.138	-0.125	-0.129
	(0.073)	(0.073)	(0.073)	(0.073)	(0.141)	(0.142)	(0.145)	(0.145)
	[0.952]	[0.956]	[0.955]	[0.953]	[0.880]	[0.871]	[0.882]	[0.879]
Home Owner	0.220***	0.218***	0.207***	0.219***	-0.318*	-0.319*	-0.322*	-0.309*
	(0.076)	(0.077)	(0.077)	(0.077)	(0.174)	(0.175)	(0.177)	(0.177)
	[1.246]	[1.243]	[1.230]	[1.245]	[0.727]	[0.727]	[0.725]	[0.734]
Rural	-0.138**	-0.144**	-0.134**	-0.128**	0.078	0.084	0.132	0.139
	(0.062)	(0.063)	(0.065)	(0.065)	(0.116)	(0.116)	(0.121)	(0.121)
	[0.871]	[0.866]	[0.875]	[0.880]	[1.082]	[1.088]	[1.141]	[1.149]
Professional	0.083	0.064	0.022	-0.028	0.382	0.396*	0.664**	0.588*
	(0.121)	(0.122)	(0.155)	(0.155)	(0.241)	(0.241)	(0.327)	(0.326)
	[1.087]	[1.066]	[1.022]	[0.972]	[1.465]	[1.485]	[1.942]	[1.800]

	Hazard Rate Analysis of Escapes from Unemployment in Local Labor Market				Hazard Rate Analysis of Escapes from Unemployment via Migration			
Covariates	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
AFQT	-0.086	-0.071	-0.093	-0.112	0.112	0.097	0.102	0.081
	(0.120)	(0.120)	(0.121)	(0.122)	(0.236)	(0.237)	(0.237)	(0.238)
	[0.918]	[0.932]	[0.912]	[0.894]	[1.118]	[1.101]	[1.107]	[1.085]
AFQT Missing	0.112	0.111	0.099	0.102	0.155	0.165	0.175	0.188
	(0.133)	(0.133)	(0.134)	(0.134)	(0.241)	(0.241)	(0.245)	(0.245)
	[1.118]	[1.117]	[1.104]	[1.108]	[1.167]	[1.179]	[1.191]	[1.206]
Weekly	-0.001	-0.001	-0.001	-0.000	-0.003	-0.003	-0.003	-0.003
UI Benefit	(0.002)	(0.002)	(0.002)	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)
	[0.999]	[0.999]	[0.999]	[0.999]	[0.997]	[0.997]	[0.997]	[0.997]
Multiple Spells	0.238***	0.238***	0.235***		0.285***	0.301***	0.297***	
	(0.051)	(0.052)	(0.052)		(0.098)	(0.099)	(0.099)	
	[1.268]	[1.268]	[1.265]		[1.330]	[1.351]	[1.346]	
Tenure		-0.010	-0.007	-0.017		0.038	0.041	0.026
		(0.016)	(0.016)	(0.016)		(0.029)	(0.030)	(0.029)
		[0.990]	[0.993]	[0.983]		[1.039]	[1.041]	[1.027]
Union		-0.032	-0.021	-0.003		-0.026	-0.032	-0.009
		(0.059)	(0.060)	(0.060)		(0.113)	(0.116)	(0.116)
		[0.969]	[0.979]	[0.997]		[0.974]	[0.968]	[0.991]
Public Sector		0.128*	0.156*	0.143*		-0.218	-0.011	-0.032
		(0.076)	(0.087)	(0.087)		(0.183)	(0.194)	(0.194)
		[1.136]	[1.169]	[1.154]		[0.804]	[0.989]	[0.969]
Industry	No	No	Yes	Yes	No	No	Yes	Yes
Occupation	No	No	Yes	Yes	No	No	Yes	Yes
Log Likelihood	-12654.1	-12652.4	-12643.9	-12654.5	-3309.4	-3307.8	-3294.1	-3298.7
No. of Spells	2419	2419	2419	2419	2419	2419	2419	2419
No. of Failures	1838	1838	1838	1838	502	502	502	502

Notes: Standard errors are in parentheses. Hazard ratios are in square brackets. Parameter estimates obtained from Cox partial-likelihood proportional hazards model. Escapes from unemployment to employment via migration are counted as censored spells in the local labor market hazard. Escapes from unemployment to employment in the current labor market are counted as censored spells in the migration hazard.

*** Significantly different from zero at the 1% level.

** Significantly different from zero at the 5% level.

* Significantly different from zero at the 10% level.

(*Multiple Spells*) to the sample have higher hazard rates in the local reemployment models. This may point to the likelihood that such workers become adept at unemployed search or are perhaps indicative of workers who select into industries with frequent terms of short-term displacement.

The results for the migration hazard are presented in Columns (5) – (8). Although relatively few covariates are statistically significant, each model shows a relatively good fit with the data with a likelihood ratio test statistic (distributed Chi-squared) significant at better than a five-percent confidence level (not shown). Consistent with expectations, homeowners are significantly less likely to take jobs requiring migration while professional workers are the most likely to migrate. As discussed earlier, *Home Owner* proxies for location-specific investments that would tend to impede migration. Workers in professional occupations are likely to have a greater proportion of their skills in general human capital and thus participate in more national labor markets. This makes migration a more likely outcome in the search process since the expected net benefits of search will be higher in alternative markets for professionals.

It is important to note that even some of the insignificant parameter estimates are also consistent with prior expectations. For example, the positive and significant impact of *HGC* on the local reemployment hazard would suggest a negative counter-effect on the migration hazard. However, higher levels of education imply greater stocks of general human capital, which exerts a positive influence on the migration hazard. Thus, the predicted effect of education on the migration hazard is theoretically ambiguous. One reaches a similar interpretation for *Married*. Workers tied by marital constraints may be less mobile if the spouse has a strong attachment to the current location for either economic or social reasons (Mincer 1978). On the other hand, moving to a new location with a partner can significantly reduce the psychic costs associated with migration (i.e., leaving behind friends and family). Therefore, this variable has an ambiguous effect on geographic mobility. In both cases, the insignificant parameter estimates are consistent with opposing effects canceling each other's impact.

5.2 Hazard Estimates of Local Labor Market Effects

Central to this study is the impact of the local labor market variables on the reemployment hazards. The key variables of interest are the state employment level (*Employment Level*), employment growth rate (*Employment Growth*), unemployment rate (*Unemployment*), and average wage level (*Wage Level*). In order to avoid issues relating to collinearity, each labor market variable is entered into the models separately (along with the complete array of baseline controls presented in Table 3). Table 4 presents parameter estimates, standard errors, and hazard ratios for each of the employment variables. Columns (1) – (4) pertain to the local reemployment hazard while Columns (5) – (8) pertain to the migration hazard.

TABLE 4

Hazard Model Parameter Estimates for Labor Market Variables

Covariates	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Hazard Rate Analysis of Escapes from Unemployment in Local Labor Market				Hazard Rate Analysis of Escapes from Unemployment via Migration			
Emp. Level	0.001	0.001	0.001	0.001	-0.006***	-0.006***	-0.007***	-0.007***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.002)
	[1.001]	[1.001]	[1.001]	[1.001]	[0.994]	[0.994]	[0.993]	[0.993]
Emp. Growth	0.033***	0.031***	0.030***	0.030***	0.008	0.010	0.003	0.001
	(0.010)	(0.010)	(0.011)	(0.010)	(0.019)	(0.019)	(0.019)	(0.019)
	[1.033]	[1.032]	[1.031]	[1.030]	[1.008]	[1.010]	[1.003]	[1.001]
Unemp.	-0.380***	-0.360**	-0.357**	-0.359**	-0.009	-0.026	0.086	0.097
	(0.150)	(0.151)	(0.152)	(0.152)	(0.280)	(0.282)	(0.286)	(0.286)
	[0.684]	[0.698]	[0.700]	[0.698]	[0.991]	[0.975]	[1.089]	[1.102]
Wage Level	0.042	0.045*	0.045*	0.046*	-0.244***	-0.250***	-0.258***	-0.256***
	(0.026)	(0.026)	(0.026)	(0.026)	(0.051)	(0.051)	(0.052)	(0.052)
	[1.043]	[1.046]	[1.046]	[1.048]	[0.784]	[0.779]	[0.773]	[0.774]

Notes: Each labor market variable is entered into each specification separately (but jointly with the respective controls specified in Table 3). Standard errors are in parentheses. Hazard ratios are in square brackets. Parameter estimates obtained from Cox partial-likelihood proportional hazards model. Escapes from unemployment to employment via migration are counted as censored spells in the local labor market hazard. Escapes from unemployment to employment in the current labor market are counted as censored spells in the migration hazard. *Employment Level* entered into model in hundred thousands (total employment / 100,000). *Wage Level* entered into model as an hourly wage (weekly wage level / 40).

*** Significantly different from zero at the 1% level.

** Significantly different from zero at the 5% level.

* Significantly different from zero at the 10% level.

Looking first at the local reemployment hazard in Columns (1) – (4) of the table, each of the local employment measures is of the anticipated sign with only the employment level not significant at conventional levels. The positive coefficients on both *Employment Growth* and *Wage Level* indicate that improvements in employment growth rates and wage levels are associated with shorter wait times until local reemployment. In contrast, the negative coefficient estimate for the unemployment rate implies a longer spell of unemployment. To give some idea of the meaning of these estimates, consider a one-unit change in each of the statistically significant variables. For easier interpretation, *Employment Growth* was entered into the model in hundred thousands whereas *Wage Level* was entered as an hourly wage. The coefficient estimates imply that an individual residing in a state with an employment growth rate of 2.75 percent (one unit above the sample mean) has a reemployment hazard that is 3 percent higher than an individual residing in an area with a 1.75 percent growth rate; a worker living in an area paying an average weekly wage of $8 per hour ($1 above the sample mean) has a reemployment hazard about 4.6 percent above a worker in an area paying an average wage of $7 per hour; and a worker whose state unemployment rate is 8.1 percent (one unit above the sample mean) has a reemployment hazard nearly 30 percent lower than a characteristically similar person in area with an unemployment rate at the sample mean.

The key test of the search model, however, is the impact of the local labor market variables on the migration hazard. Recall that when search is conducted over multiple markets, the migration hazard is unaffected by conditions in the current labor market *except* through their effect on search strategies. Thus, any observed impact of the local employment conditions on the migration hazard can be inferred to be in response to a spatially extended search strategy. Turning attention to Columns (5) – (8), both *Employment Level* and *Wage Level* are significantly and negatively related to the migration hazard. Consistent with the predictions of the multi-market search model, each variable implies a longer wait time until reemployment via migration (a lower hazard rate of exit from the unemployed state). Neither *Unemployment* nor *Employment Growth* is found to be statistically significant, and only the unemployment rate is of the anticipated sign (specifications (7) and (8)). The coefficient estimates for *Employment Level* suggests that an increase of 100,000 in state employment lowers the migration hazard by nearly 7 percent. Similarly, the coefficient estimates for *Wage Level* imply that worker living in an area paying an average weekly wage of $8 per hour ($1 above the sample mean) has a migration hazard more than 20 percent below a characteristically similar worker living in a market paying the sample average.

6. CONCLUSIONS

Using data drawn from the National Longitudinal Survey of Youth 1979, this study investigates whether displaced workers adjust their job search strategies in response to local market conditions to favor migrating out of declining labor markets. Empirical results from a Cox partial-likelihood proportional hazards model are supportive. A low density of local employment and low average wage levels are associated with shorter wait times to migration. Conversely, high local employment growth rates, high wages, and low unemployment rates are associated with an increased likelihood of obtaining local employment following displacement. By addressing the "competing-risks" nature of job search across multiple labor markets, this research marks another step in the application of formal search theory to the study of migration among the unemployed.

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