Migration, Fixed Costs, and Location Specific Amenities: A Hazard Rate Analysis

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Wallace E. Huffman and Tubagus Feridhanusetyawan

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Abstract

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This paper presents a model of long-distance migration for a finite-life individual who consumes leisure, purchased goods, and local amenities and incurs significant fixed costs of moving and empirical results from a hazard rate analysis of residence spells/internal migration using data on adult males in the Panel of Income Dynamics Study over a 20-year period. The migration hazard for interstate moves is concave in residence duration, with the hazard of migration peaking at six years in residence. The hazard of migrations is quite sensitive to the wage difference between home and host locations, but variables that proxy fixed costs of internal migration, e.g., having less education, being married, having children in school, being self-employed or a union member, significantly lower the hazard of internal migration and increase duration at the home location or origin. An increase in the crime rate at the home/origin increases the hazard of internal migration and shortens duration, but other amenities do not matter. Although geographical mobility is highly responsive to real wage differences, migration cannot be expected to fully equalize real wage rates across internal labor markets because of the significant fixed costs associated with long-distance internal migration and importance of amenities at the origin.

Key words: internal migration, United States, interstate moves, hazard function, fixed costs, local amenities.
Migration, Fixed Costs, and Location-Specific Amenities: A Hazard Rate Analysis

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Since Schultz' important paper on human capital (Schultz 1961), human migration has been an important topic of economic research. Much of the human capital literature on internal migration has emphasized the expected net earnings benefit as the major factor driving human migration decisions, e.g., see Schwartz (1976), Schlottman and Herzog (1981), Herzog and Schlottman (1984), Sandefur (1985), Pissarides and Wadsworth (1989), Borjas, Bronars, and Trejo (1992), Detang-Dessendre and Molho (1999). Much less emphasis has been given to the fixed costs associated with migration and role of location-specific amenities (Greenwood 1997).

The objective of this paper is to present a model of long-distance migration for a finite-life individual who consumes leisure, purchased goods, and local amenities and incurs significant fixed costs of moving and empirical results from a hazard rate analysis of residence spells/internal migration using data on adult males in the Panel of Income Dynamics Study (PSID) over a 20 year period. The econometric results show a strong negative effect of the wage difference between the home/origin location and host locations and of fixed costs associated with a move and a positive effect of the local crime rate on the hazard of internal migration. The story unfolds in the following sections.

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The Conceptual Model of Internal Migration

An individual receives utility from leisure time, purchased goods, and local amenities/disamenities. The latter is not a choice at the margin, but is an indicator of location-specific culture and climatic, topographic (e.g., near seacoast, mountains, plains), and environmental conditions that affect the translation of leisure and goods into satisfaction. An individual considers staying in his home (h) area or migrating to a host (f) area, and he is uncertain about future wage and amenity outcomes at these locations. Assume inter-temporal separability utility function, risk neutral individuals, and denote the expected indirect utility function for each year as $V_j(w_j, x_j)$, where $w_j$ is the expected real wage in location $j$ and $x_j$ is expected local amenities in location $j$, $j=h,f$ (Greenwood 1997, p. 668, 677). Let all expected relocation costs associated with moving from h to f in t, except for the foregone earnings, be represented by $c_n$, and to simplify, assume that $c_n$ is fixed and invariant with the distance moved. Also for simplicity, assume that local amenities and relocation costs can be measured in real wage units.

The individual chooses his location so as to maximize the expected discounted reward. Assume an n period planning horizon (i.e., length of remaining life), migrate in the first period, and ignoring discounting (or the real discount rate is zero), then the individual will migrate if

(1) \[ \sum_{t=1}^{n} (w_{ht} + x_{ht}) < -c_{n} + \sum_{t=2}^{n} (w_{ft} + x_{ft}). \]

Clearly equation (1) implies that $\sum_{t=2}^{n} (w_{ht} + x_{ht}) < \sum_{t=2}^{n} (w_{ft} + x_{ft})$ and $\sum_{t=2}^{n} (w_{ht} - x_{ht}) < \sum_{t=2}^{n} (w_{ft} - x_{ht})$. Hence, the sum over remaining life of the difference in the expected value of local amenities between the home and host location is less than the sum over the remaining life of the difference in expected value of host and home real wage rates.
However, only if $\sum_{t=2}^{n} (x_{ht} - x_{ht}) = 0$ can we say for certain that the average expected real wage difference between the host and home location will be positive. For example, if the host area gives unusually high expected local amenity value relative to the home area, it may be lifetime utility maximizing for the individual to migrate from h to f even when the average expected wage difference after migration between the host and home areas is negative.

Define $D$ as an indicator variable taking a value of 1 if $\sum_{t=1}^{n} (w_{ht} + x_{ht}) < -c_{fl} + \sum_{t=2}^{n} (w_{rt} + x_{rt})$ and 0 otherwise. Then the probability of migrating from h to f can be represented as the following probability statement:

$$
(2) \quad P_r(D=1) = \mathbb{P} \left[ \sum_{t=1}^{n} (w_{ht} + x_{ht}) < -c_{fl} + \sum_{t=2}^{n} (w_{rt} + x_{rt}) \right].
$$

Thus, the following comparative static results hold:

$$
(3) \quad \frac{\partial P_r(D=1)}{\partial w_{jt}} = \begin{cases} < 0 & \text{for } j = h \\ > 0 & \text{for } j = f 
\end{cases}
$$

$$
(4) \quad \frac{\partial P_r(D=1)}{\partial x_{jt}} = \begin{cases} < 0 & \text{for } j = h \\ > 0 & \text{for } j = f 
\end{cases}
$$

$$
(5) \quad \frac{\partial P_r(D=1)}{\partial c_{fl}} < 0.
$$

Hence, the probability of migrating from h to f is increasing in $w_f$ or $x_f$ and decreasing in $w_h$ or $x_h$, other things equal. It is also decreasing in the fixed cost of migration.

In a static environment, a strong economic incentive exists for a finite life individual to migrate from h to f early, perhaps in the first period, or to stay at h. However, given a move
from \( h \) to \( f \) is completed, it is possible for additional moves to be optimal, e.g. to return to the home location.

**The Empirical Model**

At a point in time, individuals are observed to have lived in their current area of residence for various lengths of time. In a free society, individuals are observed to occasionally move internally to a new location and take up a new residence location. There are good reasons to expect duration dependence in migration behavior. Individuals may form attachments to their home area which grows over time. This will occur as they build local human capital in knowledge of local shopping opportunities, restaurants, recreation, and schools, and social and economic ties develop. However, economic conditions may change locally (determinate) or in a potential host location (improve) and cause one to move geographically or migrate. As local human capital accumulates with time living in an area and significant fixed cost of migrating exist, we expect the migration hazard to rise as local employment opportunities are exhausted relative to new developments in other location but then to decline (Greenwood 1997). Hence, econometric methods based on time-dependent hazard functions provide one natural approach for an analysis of internal migration.

**The Hazard Model**

Define \( T \) as the duration or length of a completed residence spell in a location (home area) with associated c.d.f. of \( F(t) \) and p.d.f. of \( f(t) \) where \( t \) is a realization of \( T \). An individual's hazard function for migration can be represented as the limiting probability that a residence spell is completed in \( t \), given that he has stayed in the home area until time \( t \).

\[
H(t) = \lim_{h \to 0} \frac{P(t < T \leq t + h \mid T > t)}{h} = \frac{f(t)}{1 - F(t)} = \frac{f(t)}{S(t)}
\]
where \( H(t) \) is the individual’s migration rate at \( t \), and \( S(t) = P_r(T>t) \) is the individual’s survival function for the home area. Hence, \( S(t) \) expresses the probability that a residency spell is of length at least \( t \) (Kiefer 1988; Lancaster 1990; Greene 2000, pp. 939).

In our model, we wish to test for effects of a set of explanatory variables \( X \) on the migration hazard rate, i.e., a proportional hazard rate specification, and to consider possible heterogeneity in the residency spells (individuals). Heterogeneity can arise as (i) individual-specific unmeasured effects, e.g., intensity of psychic costs of moving, (ii) measurement error in \( X \), or (iii) measurement error in the duration of a residency spell. If we impose the Weibull distribution on residency duration \( t \) and let \( v \), distributed as gamma with unit mean and variance \( \theta \), represent the individual-specific unmeasured heterogeneity, the mixed resident survivor function is

\[
S(t, X, \beta, \sigma, \theta) = \left\{ 1 + \theta \left[ t \exp(X\beta) \right]^{\frac{1}{\sigma}} \right\}^{-\frac{1}{\theta}}.
\]

The associated migration hazard function is:

\[
H(t, X, \beta, \sigma, \theta) = \left[ S(t, X, \beta, \sigma, \theta) \right]^{\theta} \left( \frac{1}{\sigma} t \right)^{-1} \left[ \exp(X\beta) \right]^{\frac{1}{\sigma}}.
\]

In particular, the effect of unmeasured resident heterogeneity is increasing in \( \theta \), but as \( \theta \) goes to zero, heterogeneity vanishes. If \( \theta \) is not significantly different from zero, the hazard for migration is monotone in duration.

Some variables in \( X \) for the \( i \)-th individual, say \( X_{ij} \), change over time and are jointly determined with duration. When this is the case, \( X_{ij} \) is typically assigned its value at the beginning of the residency spell, say \( X_{ij0} \) (Greene 2000, pp. 943, Lancaster 1990). Some other variables change over time and are not endogenous to duration, e.g., an individual’s age and location-specific attributes.
The Data

The individuals in this study are adult males who were surveyed for the Panel Study of Income Dynamics (PSID). The PSID has major advantages over other micro-data sets in that multiple moves can be distinguished and individual characteristics are known at the time of the move and do not have to be approximated. Individuals were selected into the panel in 1968 or added as new adult males because they either grew into adult males or joined the household of an original 1968 data panelist. Individuals who leave an original PSID household are followed and re-interviewed for as long as they can be located. Every panel household member for our sample was 19-45 years of age in 1968, followed from the year after he completed school to the time when he retired, died, or was lost from the survey. Data for these individuals are supplemented with data for their residence area.

Migration over longer distance is generally associated with changes in employment or retirement; whereas short distance migration is frequently associated only with only a change in housing. The latter is not of interest to us. Some important amenity/disamenity attributes of areas (climate, crime, topography, environment, and public schools) and occupational licensing are associated with major political jurisdictions. The most frequently used area grouping for the United States is states, and other interstate migration studies include Pashigan (1979), Borjas, Bronars, and Trejo (1992), Brinig and Buckley (1996), Sandefur (1985), Clark, Knapp, and White (1996). Hence, for our study an individual is defined as a long-distance migrant if he moves across a state boundary. We, however, are treating all inter-state moves the same, i.e., we are not distinguishing between interstate moves that vary in distance from the home to the new host location. This is consistent with our conceptual model where migration costs are dominated
by fixed costs associated with making a decision to move, saying good-by and loading a family's possessions at the original location, and finding a new place to live, unpacking the family's possessions, finding new shopping and business establishments, and making new friends in the host location.\(^3\)

Also, since host locations and their location-specific amenities are a choice, all potential migrants will be facing a similar common fixed amenity-effect of potential host locations. Hence, given they live at different locations, the local amenity attributes that differ among them is the amenity attributes of their home/origin location.

Using the PSID, the state of residence of each individual can be identified. We focus on the period 1968 to 1987. Table 1 summarized the information on frequency of interstate moves for the PSID participants. Among the 915 individuals in the sample, 722 (78.9 percent) stayed in the same state throughout the period. Among the 193 individuals who had at least one interstate move, 97 (more than 50 percent) moved only once. The frequency of interstate movers decreased as the number of moves increased. The full number of PSID participating males 19 to 45 years of age in 1968 is 915 individuals, and we derived a total of 865 residence spells having known starting dates for each spell and 1,268 residence spells when adjusted starting dates replaced missing dates.\(^4\) See Appendix A for details. Because young individuals have stronger incentives to move, young participants, males who were 12-24 years of age in 1968, were used to derive a set of resident spells of young potential migrants.\(^5\)

**The Empirical Hazard Function**

The symbols used to define the variables in the empirical hazard rate model are presented in table 2, and the systematic part of the hazard rate function is specified as:
\[ X(t_{is}) \beta = \beta_0 + \beta_1 W(\text{ht}_{is}) + \beta_2 W(\text{ft}_{is}) + \beta_3 \text{AGE}(t_{is}) + \beta_4 \text{AGE}^2(t_{is}) + \beta_5 \text{EDU}(t_{is}) + \beta_6 \text{AVUNHR}(t_{is}) + \beta_7 \text{UNION}(t_{is}) + \beta_8 \text{DSLFEMP}(t_{is}) + \beta_9 \text{MARR}(t_{is}) + \beta_{10} \text{CHILD}(t_{is}) + \beta_{11} \text{CRIME}(t_{is}) + \beta_{12} \text{PARK}(t_{is}) + \beta_{13} \text{JAN}(t_{is}) + \beta_{14} \text{JULY}(t_{is}) + \beta_{15} \text{DWHITE}(t_{is}) \]

where:

- \( i = 1, 2, \ldots, n \) denotes individuals,
- \( s = 1, 2, \ldots, C_i \) denotes the spells, and
- \( C_i \) = maximum number of spells for individual \( i \).

Expectations about the signs of the \( \beta \)'s are as follows. A wage increase at the original/home location will reduce the hazard of migration and a wage increase at the host location will increase it. Although there is a stronger incentive for young males to migrate than older pre-retirement males, the hazard of migration may not peak at the youngest age (Greenwood 1997, pp. 655-6). Because finite life and the life cycle plays an important role in the timing of human capital and other family decisions, the marginal effect of an individual's age seems unlikely to be linear, and \( \text{AGE}^2(t_{is}) \) is also included as a regressor. We expect \( \beta_3 > 0 \) and \( \beta_4 < 0 \).

The next seven variables are associated with the costs of migration. A potential migrant's education is expected to be positively related to the hazard rate for interstate migration (i.e., \( \beta_5 > 0 \)). An individual's education has been shown to be associated with the ability to acquire, process, and make efficient decisions (Huffman 1977; Schwartz 1976; Schultz 1975). A large amount of uncertain is associated with moving to a new location, and additional education and information can greatly reduce it. Also, Detang-Dessendre and Molho (1999, 2000) have shown that an individual's schooling increases his/her hazard of internal long distance migration. When a potential migrant experiences unemployment, the opportunity cost of his time for searching for
or/arranging new location and employment options is reduced (Herzog and Schlottmann 1988; Pissarides and Wadsworth 1989; Goss and Paul 1990), and we expect his hazard rate for migration to increase (i.e., $\beta_6 > 0$).

Being a union member, self-employed, in a profession or trade association, is associated with non-transferable rights, revenue generating clients, or location-specific information largely tied to a potential migrant's current location (Goss and Paul 1990; Pashigian 1979). These attributes of his employment are expected to increase a potential migrant's utility at his current location relative to a new location and to reduce his hazard rate for migration (i.e., $\beta_7$ and $\beta_8 < 0$).

Having school age children and being married increases the psychological and monetary costs of moving because of the larger number of personal ties that must be broken, and greater volume of personal possessions and larger number of individuals to get ready to move and to get settled in a new location (Mincer 1978; Greenwood 1997, pp. 701-705). They tend to tie a potential migrant to his current location and reduce his expected gain from internal migration, other things equal, e.g., $\beta_9$ and $\beta_{10} < 0$.

Amenities/disamenities at a potential migrant's current location relative to potential host locations are expected to affect the hazard rate of internal migration (Greenwood 1997, pp. 676-7). We focus on the crime rate, area in parks, and normal July and January temperatures. A higher area crime rate against persons and real property provides a negative local public good to residents (and others) by imposing psychic and self-protection costs on all residents (and others) and lowers their utility, other things equal. Hence, we expect a larger value of CRIME to increase the hazard of migration, i.e., $\beta_{11} > 0$, or to reduce residence duration. In contrast, resident area parks provide a positive local public good, and can be expected to reduce the hazard
for migration (i.e., $\beta_{12} < 0$). January and July temperatures play an important role in determining the types of winter and summer season outdoor recreational opportunities that are in an area (and winter heating and summer cooling costs). If these weather variables are not fully reflected in real wage rates, then $\beta_{13}$ and $\beta_{14}$ will be statistically significant.

Whites are expected to have a higher hazard rate for migration than nonwhites (i.e., $\beta_{15} > 0$). Filler (1992) has shown that whites in the United States have many geographical locations where they can potentially move to and approximately maintain their well being relative to the opportunities available to nonwhites. His finding suggests whites will experience shorter duration at a location than nonwhites, other things equal.

The Wage at Host Locations

An individual’s expected real wage at potential host locations is an important variable in the hazard of migration analysis, but it is not directly observable. If we assume that the U.S. labor market functions relatively well in valuing heterogeneous attributes of workers, locations, and jobs, a well specified hedonic wage equation may provide useful information about the expected wage in potential host locations (Roback 1988, 1982; Rosen 1986; Topel 1986; Tokle and Huffman 1991; Kenny and Denslow 1980; Hoch and Drake 1974). In particular, state units provide valuable variation for identifying the parameters of a wage or individual labor demand equation. Consider the following hedonic hourly real wage equation for an individual:

$$\ln(W_{ik}/P_k) = \beta_X X_i + \beta_z Z_k + \beta_A A_k + \epsilon_i$$

where

$W_{ik} =$ the nominal hourly wage of individual $i$ living in state $k$, $P_k =$ the price level at location $k$. 

...
\( P_k \) = price index for purchased inputs used to produce indirect utility in households of individuals living in state \( k \),

\( X_i \) = personal characteristics,

\( Z_k \) = state characteristics which affect labor productivity,

\( A_k \) = hedonic state amenity/disamenity attributes,

\( \varepsilon_i \) = random disturbance representing luck, \( E(\varepsilon_i) = 0 \),

\( \beta_x \) = return to personal characteristics,

\( \beta_2 \) = return to local productivity characteristics,

\( \beta_a \) = return to hedonic local attributes.

When markets work well, the prices of traded goods are approximately the same across states, adjusted for transport costs. The prices of nontraded goods, e.g., local amenities, seem likely to differ across states. State cost of living indexes do not exist, but assume that the following linear state consumer cost of living index provides a good approximation:

\[
\ln(P_k) = \alpha_1 \ln(P) + (1 - \alpha_1) \ln(PLAND_k) + \alpha_2 URBAN_k \\
+ \alpha_3 CLIMATE_k + \alpha_4 REGION_k + \delta_k
\]

\[
= \ln(P) + (1 - \alpha_1) [\ln(PLAND_k) - \ln(P)] + \alpha_2 URBAN_k \\
+ \alpha_3 CLIMATE_k + \alpha_4 REGION_k + \delta_k
\]

where

\( P \) = national price index for variable goods and services purchased by households,

\( PLAND_k \) = nominal price of land (proxy for home-site values) in state \( k \),

\( URBAN_k \) = percentage urban population in state \( k \) (congestion index),

\( CLIMATE_k \) = climatic characteristics of state \( k \),

\( REGION_k \) = regional dummy variables, and,

\( \delta_k \) = random disturbance, \( E(\delta_k) = 0 \).
Substituting equation (11) into equation (10), obtain the following wage equation:

\[
\ln(W_{ik}/P_i) = \beta_x X_{ik} + \beta_z Z_{ik} + \beta_A A_{ik} + (1 - \alpha_i)[\ln(\text{PLAND}_k/P)] + \alpha_2 \text{URBAN}_k \\
+ \alpha_3 \text{CLIMATE}_k + \alpha_4 \text{REGION}_k + \epsilon_i + \delta_{ik}
\]

By incorporating PLAND, URBAN, CLIMATE, and REGION, which control for "unpriced" local attributes in the national consumer cost of living index, into equation (12), the $\beta_x$, $\beta_z$, $\beta_A$, and $\beta_s$'s can be interpreted as the average U.S. real price/value of an attribute. More importantly for this study, we expect to obtain better predictions of the real wage at host locations.

The complete empirical specification of the real hourly wage equation is:

\[
\ln(W_{iky}/P_y) = \alpha + \alpha_1 \text{EDU}_{iy} + \alpha_2 \text{EXP}_{iy} + \alpha_3 \text{EXP}^2_{iy} + \alpha_4 \text{RACE}_i + \alpha_5 \ln(\text{PLAND}_{ky}/P_y) \\
+ \alpha_6 \text{URBAN}_{ky} + \alpha_7 \text{CRIME}_{ky} + \alpha_8 \text{JAN}_k + \alpha_9 \text{JULY}_k + \alpha_{10} \text{PJOBGR}_{ky} \\
+ \alpha_{11} \text{URATE}_{ky} + \alpha_{12} \text{RSHOCK}_{ky} + \alpha_{13} \text{RURATE}_{ky} + \alpha_{14} \text{DS}_{ky} + \alpha_{15} \text{DW}_{ky} + \\
\alpha_{16} \text{DNC}_{iy} + \alpha_{17} \text{TIME}_y + \alpha_{18} \text{TIME}_y^2 + \epsilon_{iy}
\]

where the local labor market variables follow closely Topel's (1986) and Adams' (1985) definitions (See Appendix B for details). TIME and TIME squared are included to allow for a possible long term negative trend in male real wage rates (Mishel and Bernstein 1993, pp. 142). All variables are defined in table 3.

The Empirical Results

Using the PSID and supplemental data on attributes of states, empirical results are reported for the hedonic wage equation and for the hazard of internal migration.

The Wage Equation

The wage equation (equation 13) is fitted to all the observations for the 915 PSID individuals pooled over 20 years (1968-87). The performance of the fitted wage equation is generally in agreement with results reported in other studies. A one-year increase in an
individual’s schooling increased his real wage by about 7.5 percent. An increase in his experience has a positive effect on his real wage but at a diminishing marginal rate. The maximum effect of EXP occurs at 26 years of experience (approximately 45 years of age). All other measured variables held equal, white males earn 11 percent more relative to nonwhites. These results are consistent with those reported by Neal and Johnson (1996) and smaller than by Topel (1986).

Wage rates also differ because of local cost of living and amenity differences. Both the real price of land (PLAND/P) and congestion as reflected in URBAN have positive effects on the real wage rate. The elasticity of the wage with respect to the land price is about 0.05. A one percentage point increase in the proportion of a state's population that is urban increases the real wage by about 30 percent. The land price effect on the wage compares favorable to the findings of Kenny and Denslow (1980) and Tokle and Huffman (1991) but the effects of URBAN is larger in this study.

Local amenities have significant affects on the real wage. A one percentage point increase in a state's crime rate increases the real wage by about 1.5 percent (significant at the 1 percent level), which is consistent with finding reported by Roback (1982, 1988). A higher average January or July temperature for an area reduces the real wage by 4 percent or 17 percent per 10 degree increase, respectively, suggesting positive amenity value or reduced labor demand, other things equal, including Census Region. Hence, the effects of average January and July temperatures are not fully captured by other variables.

State labor market characteristics have a significant effect on individuals' wage rates. A one percentage point increase in the predicted state job growth rate (PJOBGR) increases the real
wage by 6.0 percent, and a one percentage point increase in the predicted state unemployment (PURATE) rate increases the real wage by 4.3 percent. The latter result is greater by a factor of three than those obtained by Topel (1986) and Tokle and Huffman (1991). The signs of the coefficients for both RSHOCK and RURATE, which are unanticipated outcomes, are consistent with the results in Topel (1986) and Tokle and Huffman (1991).

Historically the U.S. has had some broad regional differences in real wage rates, and our results show that these have not gone away, even after controlling for land prices, urbanization, crime rates, and climate. Compared to the Northeastern region, the real male wage rate in the South is 8.9 percent lower and 13.7 percent lower in the West. However, the male real wage rate in the North Central region is not significantly different from those in the Northeast region. Consistent with other evidence, the results show a statistically significant negative trend in the male real wage rate of slightly less than 1 percent per year.

**The Hazard of Migration**

The empirical hazard function for migration is fitted to the data on residence spells using the maximum likelihood estimation procedure (see Greene 2000; Kiefer 1988; Lancaster 1990). In fitting the model, we have combined the information on real wage at the home/original location with that of the potential host locations to define $\Delta W$ as the average annual difference between the log actual wage at the home location and predicted log wage for host locations over a residence spell. Six sets of estimates are reported in table 4 and two sets of marginal effects are reported in table 5. The results in columns (3) and (5) show that the pure-Weibull shape parameter $\sigma$ is significantly different from one (t value of 5.45) and the heterogeneity parameter $\theta$ is significantly positive. Furthermore, these results imply that the hazard of migration is
concave in residence duration. For the full sample, the maximum hazard rate for internal migration occurs when an individual has been in place for 6 years (see figure 2).

We focus on the results in column (5), the full sample of residence spells, and in column (6), the residence spells of young potential migrants. In general, the results are strongly supportive of the hypothesis that real wage differences, length of remaining life, local amenities, and costs of migration are important determinants of the hazard of male internal migration. The estimated coefficient of $\Delta W$ is negative and significantly different from zero at the 1 percent level. For the full sample, a one percent increase in the actual real male wage at the home/origin location, one percent decrease of the predicted real male wage at host locations; or one percent increase in the difference between the two real wage rates decreases his hazard of internal migration by 82 percent. For the young sample, the responsiveness is higher, a 158 percent decrease.

The effects of AGE and AGE-squared are significantly different from zero in the migration hazard equations for the full sample and young sample. In the full sample, which is not unduly constrained by age, the results imply that the marginal effect of a male’s AGE on the hazard of internal migration is positive up to 42 years of age; but then it declines for additional years of age. The hazard of internal migration is affected by the expected costs of migration. A male who has more schooling has a higher hazard rate for internal migration. For the full sample, a one-year increase in a potential migrant’s education increases his hazard of internal migration by 6.6 percent (24 percent for the young-male sample). These results support the findings of Detang-Dessendre and Molho (1999, 2000). A potential migrant who has experienced unemployment at his home/origin is more likely to move. For the whole sample, the coefficient
of UNEMP is, however, significantly different from zero only at the 7 percent level, but it is
significant at the one percent level for the young sample.

Union membership and self-employment reduce a potential migrant’s hazard for
interstate migration. The estimated coefficient of self-employment is significantly different from
zero at the 1 percent level, but for UNION at only the 9 percent level (full sample). The marginal
effect of a male being self-employed (or a farmer) is to reduce his hazard for interstate migration
by 179 percent for the full sample and by an even larger amount in the young sample. Being
married and having school age children reduces significantly the hazard of internal migration in
the full sample. Being married (for the whole time in a residence spell) causes a dramatic
reduction in the male’s hazard of internal migration, by 194 percent, and an additional school age
child reduces the migration hazard by 22 percent. As expected, the effects of being married is
smaller and weaker statistically in the young-male sample.

When interpreting the effects of local amenities on the hazard of internal migration, it is
important to recall that the real wage rates at the home and host locations have been adjusted for
the market’s valuation of these attributes. Hence, the remaining direct effects on the migration
hazard rate of local amenities is capturing effects that are not fully reflected in real wage rates.
The crime rate at the home location/origin has a positive and significant effect at the 1 percent
level on the migration hazard rate in the full sample. The marginal effect of a one percentage
point increase in CRIME is to increase the hazard for internal migration by 7 percent. However,
the direct effect of PARK, JAN, and JULY on the hazard of internal migration is not statistically
significant, suggesting that these local amenities are adequately reflected in spatial real wage
differences. The young male sample, however, is also less sensitive to CRIME than the full sample.

As expected, a white male has a significantly higher hazard rate for internal migration than for a nonwhite male. In the full sample, the hazard of migration is 77 percent higher.

Conclusion

This paper has presented an analysis of human long-distance internal migration for adult males. For a finite-life utility maximizing individual, we showed that an increase in the expected home/origin area wage, value of local amenities there, or fixed costs associated with moving to a host location decrease the probability of internal migration. Similarly, an increase in the expected real wage of host locations or value of local amenities there increase the probability of internal migration.

The empirical results were obtained using data on adult males in the PSID over the 20 year period 1968 to 1987. In the first part, we showed that the local deflated land price and amenities reflected in congestion, crime rate, and average January and July temperatures have significant effects on the expected real male wage rate. Human capital variables and state labor market conditions were also shown to be important determinants of the male wage.

In the second part, the migration hazard for adult males was shown to be concave in residence duration, with the hazard of long-distance internal migration peaking after six years in residence. A larger positive real wage difference between the home and host locations was shown to lower the migration hazard, and the migration hazard was shown to be concave in a potential migrant's age. Variables that proxy fixed costs of internal migration, e.g., having less education, being married, having children in school, being self-employed or a union member,
were shown to significantly lower the hazard of internal migration and to increases duration at the home location or origin.

The local crime rate is a disamenity—a negative local public good which imposes psychic and self-protection costs on all residents. An increase in the crime rate at the home/origin area increases the hazard of internal migration and shortens duration. However, other local amenity variables, e.g., area in parks, average January and July temperatures, were shown to have no significant effect on the hazard of migration, suggesting that real wage rates accurately compensate for these amenities. Overall, strong support is found for fixed costs and local amenities at the origin in long-distance internal migration decisions of adult males.

The empirical results imply movement across internal labor markets is responsive to real wage differences. However, our results also imply that fixed costs associated with migration and local amenities, e.g., a low crime rate at the origin, are a major drag on internal migration. Because of significant fixed costs associated with long distance internal migration and importance of local amenities at the origin, migration cannot be expected to fully equalized real wage rates across internal labor markets.
References


APPENDIX A: Treatment of Censored Spells

For a time-dependent hazard model, estimation requires that every spell be completed or right-censored. In this study, the existence of left-censored spells, however, is unavoidable because all individuals resided in a state for some time prior to the start of the survey in 1968. However, the PSID does not have information about how long participants lived in a state before 1968. Because there is no standard procedure for dealing with the issue of left-censoring, this study presents three empirical strategies, or treatments, to close some of the spells. Each treatment has its advantages and disadvantages.

For the first treatment, we selected only completed and right-censored spells that started in 1968. In other words, we ignore the spells starting before 1968. This selection led to 207 completed and 184 right-censored spells. The advantage of this treatment is that all information for estimating the hazard rate, namely, the duration and explanatory variables, are readily observable. The problem, however, is nonrandom selection.

For the second treatment, we tried to utilize all completed and right-censored spells, including those starting before 1968. To proceed, assume that those residing in 1968 in the same state where they grew up had lived in that state continuously before 1968. This assumption means that the starting time for the spell is the year when the individual turned age 19. When an individual moved before 1968 from the state where he grew up, the starting time of the first spell and thus the duration of the spell remained unknown. In this second treatment, we do not include these unobservable spells in the sample used for estimation. The second treatment has 284 completed and 581 right-censored spells. The 284 completed spells consist of 207 spells starting in 1968 or after, and 77 spells starting before 1968. Compared to the first treatment, the number
of completed spells increased significantly with the second treatment, and the problem of nonrandom selection associated with the first treatment is reduced.

For the third treatment, we predicted the starting time for the left-censored residence spells and then used all spells in fitting the time-dependent hazard function. The starting date is unknown for some spells because, prior to 1968, some individuals had migrated from the state where they grew up. In other words, the starting date for some first spells were observed and some were not. When data are censored, we can apply Heckman's procedure to predict the unobserved starting date for those first spells (Heckman 1980). The idea of Heckman's procedure is to predict the unobserved length of spells based on the observed group, adjusted to the sample selectivity. The selection criteria is the migration which took place before 1968.

There are 874 spells starting before 1968 in the sample. Among them, 585 are right-censored and 289 are right-closed spells. In the right-censored group, the starting time for 127 spells is unknown. Thus, 127 spells are both right- and left-censored. In the right-closed group, the starting time of 138 out of 289 spells is unknown. These 138 spells are left-censored but right-closed. Therefore, 265 potentially unknown spells out of the total 1,268 are to be predicted/estimated. Because the right-closed and the right-censored spells came from different populations, we applied Heckman's procedure twice, first for the group of right-closed spells and second for the group of right-censored spells. Note that the right-censored first spells are derived from individuals who never moved until the end of the period of observation.

A problem remains, however, because some of the explanatory variables, especially the wage rates, were not observed before 1968. The best information available is the value of explanatory variables in the first spell derived from the data in 1968 or earlier. The use of
partially observed explanatory variables for some spells that started before 1968 could lead to measurement error problems. To reduce the measurement error problem, we will fit the model twice. First, we used the spells derived from all individuals in the sample. Second, we selected only individuals who were 19 to 24 years of age in 1968. It is important to note that spells cannot start until the individual is 19 years old. When the young male sample is used in the analysis, the starting time of the first spell could go back as far as 1963. This approach decreases the number of partially observed spells and reduces measurement error problems without causing nonrandom selection error, and leads to 137 completed and 141 right-censored spells.
APPENDIX B: Definition of Labor Market Conditions

Predicted job growth in state k in year y (PJOBGR\textsubscript{ky}) is the difference between the forecasted value of the natural logarithm of the state's private sector employment in years y and y-1. The forecasts were obtained from a regression of the natural logarithm of non-agricultural employment for 1968-91 on a quadratic trend. The residuals from these regressions, \(e_y\), are indexes of time varying local demand conditions in state k in year y. Next, the natural logarithm of national (U.S.) aggregate employment was regressed on a quadratic trend. The residual from this regression, \(e_y\), captures the aggregate labor demand disturbance in year y. The relative local labor disturbance of state k in year y (RSHOCK\textsubscript{ky}) is defined as \(RSHOCK\textsubscript{ky} = e_y - e_y\). This variable expresses the current labor demand shock as a deviation from the aggregate labor demand shock.

The predicted state unemployment rate in state k in year y (PURATE\textsubscript{ky}) measures the anticipated state equilibrium unemployment rate. This rate is obtained by regressing the state's annual unemployment rate for 1968-91 on a quadratic trend. The unanticipated unemployment rate is captured by the residual unemployment rate (RURATE\textsubscript{ky}).
ENDNOTES

1. The hazard function $H(t)$ is obtained from the survival function as $H(t) = -\frac{d\ln S(t)}{dt}$; so there is a sign reversal of coefficients in going between the survival and hazard functions. The Weibull distribution is monotone (constant, increasing, or decreasing) in duration, but employment duration is generally non-monotonic concave in duration (Lancaster 1990, pp. 9; Gritz 1993). We permit this pattern in residence duration by adopting a mixture distribution—Weibull and gamma. An alternative distribution with this pattern is log-logistic (Keifer 1988; Lancaster 1990; Greene 2000, pp. 940-41). All are distributions for a nonnegative random variable.

2. Heterogeneity will arise when a population of residents (residence spells) has potentially different distributions of duration after controlling for the effects of observable variables. The gamma distribution is frequently used for representing the distribution of $v$ associated with this unobserved heterogeneity. Chamberlain (1985) and Heckman and Singer (1985) have shown that failure to include heterogeneity when it is present causes significant bias in the estimated coefficients of the regressors in the hazard function. Han and Hausman (1990) have shown that a parametric gamma distribution of unobserved heterogeneity leads easily to estimable models and is not unduly restrictive.

3. Borjas, Bronars, and Trejo (1992) make the opposite assumption that migration costs are dominated by variable costs associated with the distance moved. This assumption seems most applicable to never married, young individuals, and their empirical results are for the National Longitudinal Survey of Youth (waves 1979-1986).

4. Among 915 individuals in the PSID in 1968, 56 (6.12 percent) died, 144 (15.74 percent) were lost, and 132 (14.4 percent) retired during 1968-87. Individuals who refused to participate in the survey at any time during 1968-87 were deleted from the sample. Individuals who were classified as 'lost' consist of those who joined the army (30.5 percent), and those who moved out from the United States and/or were really lost (69.5 percent).

5. These males were 38 to 43 years of age in 1987.

6. REGION also controls for regional differences in the wage structure in the United States (e.g., see Roback 1988).

7. Some positive attributes of $\Delta W$ are as follows. First, the wage differences represents luck, which can explain why two persons having the same measured attributes may have different wage rates. Second, the wage difference captures the unmeasured state-specific returns to factors affecting wage rates, which might differ from the average return across the United States. Third, the wage difference incorporates tenure effects due to firm- or location-specific human capital associated with the current residence.
Table 1. The distribution of the number of interstate moves by males, 1968-87

<table>
<thead>
<tr>
<th>Number of move</th>
<th>Frequency</th>
<th>Percent</th>
<th>Frequency (Cumulative)</th>
<th>Percent</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>722</td>
<td>78.9</td>
<td>722</td>
<td>78.9</td>
</tr>
<tr>
<td>1</td>
<td>97</td>
<td>10.6</td>
<td>819</td>
<td>89.5</td>
</tr>
<tr>
<td>2</td>
<td>59</td>
<td>6.4</td>
<td>878</td>
<td>96.0</td>
</tr>
<tr>
<td>3</td>
<td>23</td>
<td>2.5</td>
<td>901</td>
<td>98.5</td>
</tr>
<tr>
<td>4</td>
<td>9</td>
<td>1.0</td>
<td>910</td>
<td>99.5</td>
</tr>
<tr>
<td>5</td>
<td>3</td>
<td>0.3</td>
<td>913</td>
<td>99.8</td>
</tr>
<tr>
<td>6</td>
<td>2</td>
<td>0.2</td>
<td>915</td>
<td>100.0</td>
</tr>
</tbody>
</table>

Source: PSID files
Table 2. Variable names and sample means for the hazard function

<table>
<thead>
<tr>
<th>Symbol</th>
<th>Variable Description</th>
<th>Sample Means</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td>Completed</td>
</tr>
<tr>
<td></td>
<td></td>
<td>Spells</td>
</tr>
<tr>
<td>(W(h_{ti}))</td>
<td>The average real hourly wage in the home location or origin over the residence spell</td>
<td>-0.199</td>
</tr>
<tr>
<td>(W(f_{ti}))</td>
<td>The average real hourly predicted wage for potential host locations over the residence spell</td>
<td>34.473</td>
</tr>
<tr>
<td>(\Delta W(t_{is}))</td>
<td>The average percentage annual differences between actual and predicted wage in residence spell (see text)</td>
<td>14.043</td>
</tr>
<tr>
<td>(AGE(t_{is}))</td>
<td>Age at the beginning of the spell (year)</td>
<td>70.474</td>
</tr>
<tr>
<td>(EDU(t_{ie}))</td>
<td>Level of education at the beginning of the spell</td>
<td>0.135</td>
</tr>
<tr>
<td>(AVUNHR(t_{is}))</td>
<td>Average annual unemployment hours in the spell (hr/year)</td>
<td>0.150</td>
</tr>
<tr>
<td>(UNION(t_{ie}))</td>
<td>Dummy variable, equals to 1 if an individual is a union member at the beginning of the spell, or 0 otherwise</td>
<td>0.713</td>
</tr>
<tr>
<td>(DSLFEMP(t_{is}))</td>
<td>Dummy variable, equal to 1 if self-employed or a farmer and 0 otherwise</td>
<td>0.766</td>
</tr>
<tr>
<td>(MARR(t_{is}))</td>
<td>Share of time being married in a residence spell</td>
<td>0.195</td>
</tr>
<tr>
<td>(CHILD(t_{ie}))</td>
<td>Number of children who are school age at the beginning of a spell</td>
<td>-0.116</td>
</tr>
<tr>
<td>(CRIME(t_{is}))</td>
<td>Average annual crime rate in the state corresponding to residence spell, relative to the U.S. average</td>
<td>4.781</td>
</tr>
<tr>
<td>(PARK(t_{ie}))</td>
<td>Share of state area in state and national parks corresponding to residence spell, relative to U.S. average</td>
<td>1.382</td>
</tr>
<tr>
<td>(RACE(t_{is}))</td>
<td>Dummy variable, equal to 1 if white and 0 otherwise</td>
<td>0.986</td>
</tr>
</tbody>
</table>

Number of spells 207 1,061
Table 3. Variable names, sample means, and the coefficient in the wage equation (t-values are in parentheses)

<table>
<thead>
<tr>
<th>Symbol</th>
<th>Variable Description</th>
<th>Sample mean</th>
<th>coefficient t-values</th>
</tr>
</thead>
<tbody>
<tr>
<td>ln(Wiy/Py)</td>
<td>Log of real hourly wage of individual i in year y (ln $/hr)</td>
<td>2.431</td>
<td></td>
</tr>
<tr>
<td>EDUiy</td>
<td>Education of individual i in year y (in years)</td>
<td>12.709</td>
<td>0.075 (34.07)</td>
</tr>
<tr>
<td>EXPiy</td>
<td>Experience of individual i in year y (Age-EDU-6 in yrs)</td>
<td>23.597</td>
<td>0.051 (21.83)</td>
</tr>
<tr>
<td>EXPiy^2</td>
<td>Experience squared/100</td>
<td>6.581</td>
<td>-0.098 (-19.62)</td>
</tr>
<tr>
<td>RACEi</td>
<td>Dummy variable, equal to 1 if individual i is white and 0 otherwise</td>
<td>0.929</td>
<td>0.110 (4.73)</td>
</tr>
<tr>
<td>ln(PLAN_Dk/Py)</td>
<td>Log of real price of land in state k where individual lives in year y ($/acres)</td>
<td>1,118.55</td>
<td>0.048 (3.07)</td>
</tr>
<tr>
<td>URBAN_k</td>
<td>Proportion of urban population in state k in year y (percent)</td>
<td>71.771</td>
<td>0.295 (4.10)</td>
</tr>
<tr>
<td>CRIME_k</td>
<td>Crime rate in state k in year y (percent)</td>
<td>8.090</td>
<td>0.015 (5.94)</td>
</tr>
<tr>
<td>JAN_k</td>
<td>Thirty-year average of January temperature in state k (degrees F.)</td>
<td>33.167</td>
<td>-0.004 (-3.39)</td>
</tr>
<tr>
<td>JULY_k</td>
<td>Thirty-year average of July temperature in state k (degrees F.)</td>
<td>75.648</td>
<td>-0.017 (-10.42)</td>
</tr>
<tr>
<td>PJOBGR_k</td>
<td>Predicted job growth in state k between years y and y-1 (see Appendix B)</td>
<td>0.213</td>
<td>0.060 (7.46)</td>
</tr>
<tr>
<td>PURATE_k</td>
<td>Predicted unemployment rate in state k in year y (see Appendix B)</td>
<td>6.357</td>
<td>0.043 (7.15)</td>
</tr>
<tr>
<td>RSHOCK_k</td>
<td>Relative employment shock in state k in year y (see Appendix B)</td>
<td>0.000</td>
<td>0.009 (2.92)</td>
</tr>
</tbody>
</table>
Table 3. (continued)

<table>
<thead>
<tr>
<th>Symbol</th>
<th>Variable Description</th>
<th>Sample mean</th>
<th>Wage equation coefficient t-values</th>
</tr>
</thead>
<tbody>
<tr>
<td>RURATE$_{ky}$</td>
<td>Residual unemployment rate in state k in year y (see Appendix B)</td>
<td>0.040</td>
<td>-0.005 ( -1.06)</td>
</tr>
</tbody>
</table>

Regional dummies and Trend:

<table>
<thead>
<tr>
<th>Symbol</th>
<th>Variable Description</th>
<th>Coefficient</th>
<th>t-values</th>
</tr>
</thead>
<tbody>
<tr>
<td>DS$_i$</td>
<td>Dummy variable equals 1 if individual i lives in the South and 0 otherwise</td>
<td>0.290</td>
<td>-0.089 ( -3.03)</td>
</tr>
<tr>
<td>DW$_i$</td>
<td>Dummy variable equals 1 if individual i lives in the West and 0 otherwise</td>
<td>0.226</td>
<td>-0.137 ( -4.47)</td>
</tr>
<tr>
<td>DNC$_i$</td>
<td>Dummy variable equals 1 if individual i lives in the North Central region and 0 otherwise</td>
<td>0.272</td>
<td>-0.018 ( -0.92)</td>
</tr>
<tr>
<td>TIME$_y$</td>
<td>Time indicator, 1968 = 1, ... 1987 = 20</td>
<td>10.123</td>
<td>-0.012 ( -1.88)</td>
</tr>
<tr>
<td>TIME$_{y^2}$</td>
<td>Time squared/100</td>
<td>1.357</td>
<td>0.008 ( -0.29)</td>
</tr>
</tbody>
</table>

Number of observations 15,367

$^a$Py = Implicit price deflator for personal consumption expenditure (1987 = 1.00)

$^b$From the 915 individuals observed over 20 years, we derived 15,367 observations.
<table>
<thead>
<tr>
<th>Variables</th>
<th>Constant Hazard Rate</th>
<th>Time Dependent Hazard Rate</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>Full Sample (1)</td>
<td>(2)</td>
</tr>
<tr>
<td></td>
<td>(-3.47)</td>
<td>(-2.09)</td>
</tr>
<tr>
<td>ΔW</td>
<td>-0.215</td>
<td>-0.286</td>
</tr>
<tr>
<td></td>
<td>(-2.40)</td>
<td>(-1.71)</td>
</tr>
<tr>
<td>AGE</td>
<td>-0.180</td>
<td>-0.199</td>
</tr>
<tr>
<td></td>
<td>(-4.16)</td>
<td>(-2.13)</td>
</tr>
<tr>
<td>AGE^2/100</td>
<td>0.269</td>
<td>0.308</td>
</tr>
<tr>
<td></td>
<td>(4.97)</td>
<td>(2.49)</td>
</tr>
<tr>
<td>EDU</td>
<td>0.135</td>
<td>0.212</td>
</tr>
<tr>
<td></td>
<td>(5.93)</td>
<td>(4.90)</td>
</tr>
<tr>
<td>UNEMP</td>
<td>0.001</td>
<td>0.002</td>
</tr>
<tr>
<td></td>
<td>(4.91)</td>
<td>(2.16)</td>
</tr>
<tr>
<td>UNION</td>
<td>-0.507</td>
<td>-0.760</td>
</tr>
<tr>
<td></td>
<td>(-2.31)</td>
<td>(-2.50)</td>
</tr>
<tr>
<td>DSLFBEMP</td>
<td>-1.344</td>
<td>-1.744</td>
</tr>
<tr>
<td></td>
<td>(-7.03)</td>
<td>(-5.99)</td>
</tr>
<tr>
<td>MARR</td>
<td>-1.139</td>
<td>-1.240</td>
</tr>
<tr>
<td></td>
<td>(-7.30)</td>
<td>(-3.78)</td>
</tr>
<tr>
<td>CHILD</td>
<td>-0.139</td>
<td>-0.167</td>
</tr>
<tr>
<td></td>
<td>(-2.32)</td>
<td>(-2.12)</td>
</tr>
<tr>
<td>CRIME</td>
<td>0.002</td>
<td>0.052</td>
</tr>
<tr>
<td></td>
<td>(0.11)</td>
<td>(1.59)</td>
</tr>
<tr>
<td>PARK</td>
<td>-0.060</td>
<td>-0.039</td>
</tr>
<tr>
<td></td>
<td>(-1.57)</td>
<td>(-0.74)</td>
</tr>
<tr>
<td>JAN</td>
<td>0.024</td>
<td>0.020</td>
</tr>
<tr>
<td></td>
<td>(3.41)</td>
<td>(1.82)</td>
</tr>
<tr>
<td>JULY</td>
<td>0.015</td>
<td>0.022</td>
</tr>
<tr>
<td></td>
<td>(1.28)</td>
<td>(1.12)</td>
</tr>
<tr>
<td>RACE</td>
<td>1.693</td>
<td>2.073</td>
</tr>
<tr>
<td></td>
<td>(2.98)</td>
<td>(3.17)</td>
</tr>
<tr>
<td>σ</td>
<td>0.632</td>
<td>0.184</td>
</tr>
<tr>
<td></td>
<td>(7.97)</td>
<td>(2.62)</td>
</tr>
<tr>
<td>θ</td>
<td>3.814</td>
<td>3.843</td>
</tr>
<tr>
<td></td>
<td>(4.47)</td>
<td>(3.83)</td>
</tr>
</tbody>
</table>

Log-likelihood: -793.5 -755.9 -795.0 -316.3 -1,287.2 -417.6
Completed Spells: 207 207 284 137 496 191
Total Spells: 1,268 1,268 865 278 1,268 350
(Table 4 continued)

*Full sample consists of males 19-45 years old in 1968, observed for 20 years (1968-87).
*Young sample consists of males 19-24 years old in 1968, observed for 20 years (1968-87).
Table 5. Marginal effect of selected explanatory variables on the hazard rate for interstate migration

<table>
<thead>
<tr>
<th>Explanatory Variable</th>
<th>Unit</th>
<th>Full sample (From Table 5, Column 5)</th>
<th>Young sample (From Table 5, Column 6)</th>
</tr>
</thead>
<tbody>
<tr>
<td>ΔW</td>
<td>percent</td>
<td>-0.823</td>
<td>-1.584</td>
</tr>
<tr>
<td>AGE</td>
<td>(0 - 1)*</td>
<td>0.947</td>
<td></td>
</tr>
<tr>
<td>EDU</td>
<td>Year</td>
<td>0.066</td>
<td>0.244</td>
</tr>
<tr>
<td>UNEMP</td>
<td>hour</td>
<td>0.001</td>
<td>0.006</td>
</tr>
<tr>
<td>UNION</td>
<td>Dummy: 0, 1</td>
<td>-0.328</td>
<td>-0.816</td>
</tr>
<tr>
<td>DSLFARM</td>
<td>Dummy: 0, 1</td>
<td>-1.786</td>
<td>-5.283</td>
</tr>
<tr>
<td>MARR</td>
<td>(0 - 1)</td>
<td>-1.940</td>
<td>-1.043</td>
</tr>
<tr>
<td>CHILD</td>
<td>0, 1, 2, ...</td>
<td>-0.225</td>
<td>0.492</td>
</tr>
<tr>
<td>CRIME</td>
<td>percentb</td>
<td>0.069</td>
<td>0.165</td>
</tr>
<tr>
<td>RACE</td>
<td>Dummy: 0, 1</td>
<td>0.768</td>
<td>0.451</td>
</tr>
</tbody>
</table>

*The share of time being in the twenties is used to calculate the marginal effect of age

bMeasured relative to U.S. average

1-16-01
Figure 1. Distributions of starting and ending times of observations for the 915 males in the sample.
Figure 2. Predicted hazard rate and the duration of stay in the time-dependent hazard model, using treatment 3 for the full sample.