International Re-Migration Analysis: Evidence from Puerto Ricans

Yangi Li and Wallace E. Huffman*

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*Wallace E. Huffman
Iowa State University
478C Heady Hall
Ames, IA 50011-1070
515-294-6359
whuffman@iastate.edu

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Abstract

This study focuses on re-migration, where individuals return to their place of birth after living in a new location for several years. Puerto Rican male householders, born in Puerto Rico but residing on the U.S. mainland and re-migrating to Puerto Rico, is the sample for fitting a hazard rate model of re-migration. We find a strong quadratic effect of an individual's age on his hazard rate. Also, males having English proficiency, less schooling, and working disability are less likely to re-migrate. A higher predicted job growth rate for Puerto Rico (U.S. mainland) and unemployment rate for the mainland (Puerto Rico) have positive (negative) effects on the hazard rate. The hazard rate is negatively related to the Puerto Rican real minimum wage.

I. Introduction

Geographical human migration, including international, interregional, and interstate moves, is a very important form of human capital investment, and it has been receiving major popular and professional attention. We observe relatively high geographical migration rates near the time individuals complete their formal schooling. Also, significant migration follows episodes of unemployment and plans to retire from the workforce. Frequently individuals return to the area where they were born, especially in retirement.

The United Nations (1989, p.61) estimated that there are approximate 60 million people, or 1.2 percent of the world's population, now living in a country other than where they were born or in host (or foreign) countries. Over half of all immigrants go to the United States, Canada, or Australia. The U.S. has experienced a rapid increase in the size of immigrant flows and major changes in the national-origin composition of the immigrant population over time. These

* Yang Li is an Associate Professor in the Department of International Business Studies, National Chi-Nan University, Taiwan, R.O.C. Wallace E. Huffman is a Professor in the Department of Economics, Iowa State University.

changes are partly due to changes in the U.S. immigration policy.[1] Other factors are changes in political and economic conditions in other countries, for instance, Cuba and Mexico. Immigrant flows approximated 4.09 millions for 1931–1960, but it increased to 4.49 millions in the 1970s and 7.34 millions in the 1980s.

Most studies of international migration have focused on the impacts of immigrant flows in the host country and how immigrants have adapted to the host country. Borjas (1994) provided an excellent survey of the empirical results in this area. However, in an uncertain world, re-migration is part of an optimal migration strategy. For example, suffering a layoff or living in an unsteady political situation in a host country can trigger re-migration (Li 1996). Re-migration is a relatively new research topic, and a few studies have attempted to test the Roy model: Ramos (1992) studied Puerto Ricans moving between Puerto Rico and the U.S. mainland; Borjas and Bratsberg (1996) focused on how return migration affected the average earnings of the remaining immigrants in the U.S.; and Bratsberg (1993) examined return migration rates of U.S. foreign students. None of these studies, however, provides an empirical analysis of how social and economic factors influence re-migration behaviors.

This is a study of re-migration, where individuals return to their place of birth after living in a new location for several years. The objective is to provide econometric estimates of the contribution of personal and local-area and birth-area attributes to the hazard rate for re-migration of Puerto Rican born male householders during the 1980s, i.e. males living on the U.S. mainland and returning to Puerto Rico. The hazard rate model is a particular representation of the probability of an individual re-migrating during a particular time interval. By using the data on Puerto Ricans we are able to tap a relatively large and rich data set. We find that males have a low and decreasing hazard rate for re-migration as they age up to age 28, but after age 28, the hazard rate is increasing with each birthday and it becomes very large by age 64. Additional findings include males having more schooling, less English proficiency, and no work disability are more likely to re-migrate. Better Puerto Rican labor market conditions increase an individual's hazard rate for re-migration, but better mainland labor market conditions reduce it.

[1] Before 1965, the U.S. immigration policy was based on the national-origins quota system which favored Europeans. After the national-origins quota system was eliminated in 1965, U.S. legal immigrant flows shifted in favor of Asia. Furthermore, the 1986 Immigration Reform and Control Act (IRCA) gave amnesty to 2.7 million illegal aliens, and the 1990 Immigration Act permitted an additional 150,000 legal immigrants annually.

Also, a higher Puerto Rican real minimum wage reduces an individual's hazard rate for re-migration.

The outline of this study is as follows: Section 2 describes the pattern of migration for Puerto Ricans on the U.S. mainland. We are interested in the conditional probability of re-migrating, given that an immigrant has already lived in the host country for several years. A return migration econometric model is presented in section 3. Section 4 describes the microdata used in this study and empirical definitions of variables. The empirical results for the fitted hazard rate model of re-migration are presented in section 5. The conclusions are in the final section.

II. Puerto Ricans and Migration

Although Puerto Rico is a territory of the United States, migration between the island of Puerto Rico and the U.S. mainland resembles international migration in the sense of the areas having different cultures and languages and significant costs of moving. It is different from international migration in that individuals born in Puerto Rico carry a U.S. passport. Puerto Ricans can legally move between Puerto Rico and the U.S. mainland, so migration flows can, in effect, be attributed entirely to differences in social and economic factors.

For the past four decades, migration has been a critical ingredient in Puerto Rican economic development, and it has had some impacts on the mainland. Santiago (1992) suggests that migration to the U.S. mainland provides a "safety valve" to island population pressures including periods of surplus agricultural labor. Castillo-Freeman and Freeman (1992) emphasize that migration has been a key factor in the long-run growth of real earnings in Puerto Rico because it increased the average quality of workers and reduced the labor supply in Puerto Rico. The annual average outflow was 45,800 in the 1950s, 27,300 in the 1960s, and 24,300 in the 1970s; moreover, one-third of native born Puerto Ricans, aged 20–64, resided on the U.S. mainland in 1980 (Ramos 1992).

There are several reasons why large immigrant flows to the U.S. mainland occur. First, the Puerto Rican birth rate is relatively high. In 1980, the total population of the island was 3.1 million and 2 million Puerto Rican persons resided on the U.S. mainland. The average population density was over 900 people per square mile on the island which poses severe socioeconomic pressure on the island's resources. Thus, overcrowding and congestion in Puerto Rico is one push factor behind migration to the U.S. mainland where the average population density is 50.

Second, the Puerto Rican government has assisted moves to the mainland to ease over-population. The Migration Division of the Commonwealth Labor Department has established nine offices in eastern and mid-western American cities. They provide services such as negotiating contract which guarantee a fixed minimum wage, decent working conditions, and round-trip airfare to the mainland (Hauberg 1974, p. 103). Third, there is no legal restriction on island to mainland migration because of American citizenship. This is in contrast to the more stringent immigration quotas and qualifications on immigrants of other neighboring countries, like Mexico and those in Latin America.

Fourth, the minimum wage policy has been unsteady. Castillo-Freeman and Freeman (1992) found that imposing the U.S.-level minimum wage on Puerto Rico altered the earning distribution by creating marked spikes in the distribution of earning in the area of the minimum wage, and that most migrants from Puerto Rico to the mainland have been disemployed. Their regression results, based on time series data for the period 1951–1987, show: (1) the minimum wage has had a significantly negative effect on the employment-population rate; and (2) the unemployment effects of the minimum wage are positive, but not significant, because some workers may leave the island after displacement caused by the minimum wage.

The final strong pull factor is the familial tie. The family in Puerto Rico often sends the first member to discover opportunities. He/she lives with friends, learns the conditions of the country, and saves some money. Later other family members follow. These networks undoubtedly help explain the high population of Puerto Ricans living in the New York City area.

Return migration is also an important phenomenon in Puerto Rico. Re-migration generally takes place during periods of high unemployment on the U.S. mainland, for instance, in the late 1980s, and general economic recession, for example, in the mid-1970s.

Puerto Rican population figures both in Puerto Rico and on the mainland in 1980 and 1990 are shown in Table 1. There were 2.7 (2) million Puerto Ricans residing on the mainland in 1990 (1980), 41.6 (46.2) percent were born in Puerto Rico and 55 (50.4) percent born on the mainland. The total population in Puerto Rico in 1990 (1980) was 3.5 (3.1) million persons, of whom 3.2 (2.9)

Table 1. Puerto Rican Population Living in Puerto Rico and on the U.S. Mainland in 1980 and 1990

	Total Population	Born in Puerto Rico	Born on Mainland U.S.	Other
Puerto Ricans living in the U.S. in 1980 [a]	2,014,000	930,600	1,014,500	68,900
Percentage [a]	100.0	46.2	50.4	3.4
Population of Puerto Rico in 1980 [a]	3,196,520	2,881,641	199,524	70,768
Percentage [a]	100.0	93.3	5.7	1.0
Persons 5 years old and over residing in the U.S. mainland more than 6 consecutive months during the 1970s	395,708	283,223	112,485	
% of total population of Puerto Rico in 1980	12.4	9.8	3.4	
Puerto Ricans living in the U.S. in 1990 [b]	2,727,754	1,134,746	1,500,265	92,743
Percentage [b]	100.0	41.6	55.0	3.4
Population of Puerto Rico in 1990	3,522.037	3,200,940	229,304	91,793
Percentage	100.0	90.9	6.5	2.6
Persons 5 years old and over residing in the U.S. mainland more than 6 consecutive months during the 1980s	398,143	292,516	105,627	
% of total population of Puerto Rico in 1990	11.3	9.1	3.0	

Source: Census of Population and Housing, 1980 and 1990.

[a] Adopted from Table 2.1 of Ramos (1992).

[b] Estimated from 1-perecnt sample of PUMS for the U.S.

million were born in Puerto Rico. Moreover, 398 (396) thousand persons, who are 5 years old and over, have ever resided on the U.S. mainland for 6 or more consecutive months during the 1980s (1970s), of whom 293 (283) thousand were natives of Puerto Rico. In other words, about 11.3 (12.4) percent of the population, or about 9.1 (9.8) percent of natives of Puerto Rico in 1990

(1980) have lived on the U.S. mainland for 6 or more consecutive months during 1980s (1970s).

What is the pattern of re-migration for Puerto Ricans born in Puerto Rico? Table 2 show statistics for Puerto Ricans, who were 5 years old and over, born in Puerto Rico, and re-migrated to Puerto Rico from the U.S. mainland during 1970s and 1980s. The re-migration flows were relatively stable during the 1970s with a range between 21,802 and 29,928 persons per year. Nevertheless, the re-migration flows grew in the late 1980s to more than 35,000 per year. Part of the reason for the increased re-migration flows to Puerto Rico after 1986 might be the large (legal and illegal) Mexican and Central American immigration flows including the impact of the Immigration Reform and Control Act of 1986 (IRCA).

III. The Econometric Model

A. Hazard Rate Models

Let T be the duration of a completed spell[2] of immigration with c.d.f. $F(t)$ and p.d.f. $f(t)$. Then, the hazard function for re-migration, or the limiting probability that a spell will be completed in a short time period *h*, is defined by

(1) $$l(t) = \lim_{h \to 0} \frac{\Pr(t < T \le t + h \mid T > t)}{h} = \frac{f(t)}{1 - F(t)} = \frac{f(t)}{S(t)}$$

where $l(t)$, called the instantaneous re-migration probability or the hazard rate at time t, can be interpreted as the rate at which a spell will be completed at duration t, given that it lasts until time t, and $S(t) = \Pr(T > t)$ is the survival function.

[2] A spell is complete if the time when the spell began and ended can be observed from the data. When the beginning time is not observable, it is called left-censored. Similarly, when the ending time is not observable, it is called right-censored.

Table 2. Puerto Rico-born, 5 Years Old and Over, Re-Migrated to Puerto Rico from the U.S. Mainland during the 1970s and 1980s

Year	Number of years	Male	Female	Total
1970 ~ 1972	3	34,404	31,000	65,404
1973 ~ 1974	2	24,496	22,565	47,061
1975	1	14,652	13,747	28,079
1976 ~ 1977	2	27,460	26,619	54,079
1978	1	15,176	14,752	29,928
1979 ~ 1980 [a]	2	28,318	24,675	52,993
1980 ~ 1982 [a]	3	26,831	25,535	52,366
1983	1	9,598	9,348	18,946
1984	1	11,796	10,835	22,631
1985	1	15,718	14,738	30,456
1986	1	14,258	13,043	27,301
1987	1	17,864	17,773	35,637
1988	1	18,989	17,969	36,958
1989	1	27,858	24,813	52,671

Source: Census of Population and Housing, 1980 and 1990.

[a] Note that 1980 appears twice in this table.

The hazard function provides a convenient definition of duration dependence. Positive (negative) duration dependence, $d\,l(t)/dt > 0\,(< 0)$, means the probability that a spell will end shortly increases (decreases) when the spell increases in length. For example, decreasing hazard functions are commonly found in data on unemployment duration, but in some models of employment duration, the shape of the hazard function rises to a peak before starting to decline (Lancaster 1990).

We are mainly interested in the effect of a set of explanatory variables (X) on the hazard rate and the duration. Hence, the hazard rate could be rewritten as:

(2) $$l(t,X,b) = \lim_{h\to 0}\frac{\Pr(t < T \le t+h | T > t, X, b)}{h} = \frac{g(t|X,b)}{1-G(t|X,b)} = \frac{g(t|X,b)}{S(t|X,b)}$$

where b is an unknown coefficient vector. The corresponding survival function is

(3) $$S(t,X,b) = \Pr(T > t | X, b) = \exp\left(-\int_0^t l(z,X,b)dz\right),$$

and the conditional density function of the duration is given by

(4) $$g(t|X,b) = l(t,X,b)S(t,X,b).$$

Unlike the ordinary least squares method, the coefficients of explanatory variables in the hazard function do not have clear interpretation as a partial derivative of the hazard or the duration. The interpretation depends on the specification. Two special cases, the proportional hazard rate and the accelerated life-time models, are easy to interpret. In the proportional-hazard model specification, the effect of explanatory variables (X) enters as a multiplicative effect on the hazard function. The accelerated lifetime model, which is the class of log-linear model for T (Kalbfleisch and Prentice 1980, ch. 2), is characterized by the multiplicative effect of explanatory variables on time rather on the hazard function.

The Weibull regression model is both the proportional hazard rate and the accelerated life-time models (Kalbfleisch and Prentice 1980, ch. 2). The corresponding hazard and survival functions are

(5) $$l(t,X,b) = \frac{1}{s}t^{\frac{1}{s}-1}\exp(-X'b)^{\frac{1}{s}},$$ and

(6) $$S(t,X,b) = \exp\left\{-\left[t\exp(-X'b)\right]^{1/s}\right\}$$

where $(1/s) > 0$ is the shape parameter. The partial derivative of $l(t,X,b)$ w.r.t. t is

(7) $$l(t,X,b) = \frac{1}{s}\left(\frac{1}{s}-1\right)t^{\frac{1}{s}-2}\exp(-X'b)^{\frac{1}{s}}$$

This result means that the hazard rate is constant, increasing, or decreasing over time as $s = 1$, $s < 1$, or $s > 1$, respectively.

The partial effect of X on $\ln\ l(t,X,b)$ is

(8) $$\frac{\partial \ln\ l(t,X,b)}{\partial X} = -\left(\frac{b}{s}\right)$$

Hence, b is the (negative) constant proportional effect of X on the hazard rate, but re-scaled by the parameter s . In other words, the marginal effect of covariates on the hazard rate is smaller (bigger) when the hazard rate is decreasing (increasing) over time. The corresponding linear model interpretation is

(9) $$\ln t = X'b + se$$

where e is the error term with type 1 extreme value distribution.[3] The partial effect of X on $\ln t$ is

(10) $$\frac{\partial \ln t}{\partial X} = b \ .$$

Therefore, b is also the constant proportional effect of the covariates on the length of time for completing a spell.

The parameters b and s of equations (5) and (6) can be estimated by maximum likelihood (ML) methods. Let $d_i = 1$ if the i^{th} spell is completed and $d_i = 0$ otherwise. Then, the log-likelihood function for a sample of n independent spells $\{t_i\}_{i=1}^{n}$ with corresponding covariates $\{X_i\}_{i=1}^{n}$ is given by

(11) $$\ln L(b, s) = \sum_{i=1}^{n} d_i \ln f(t_i, X_i, b, s) + \sum_{i=1}^{n} (1 - d_i) \ln S(t_i, X_i, b, s).$$

That is, completed spells contribute to the density term $f(t, X, b, s)$ and censored spells contribute to the probability $S(t, X, b, s)$. Given that the density function is the product of the hazard function and the survivor function, the log-likelihood function can be rewritten as:

(12) $$\ln L(b, s) = \sum_{i=1}^{n} \left[d_i \ln l(t_i, X_i, b, s) + \ln S(t_i, X_i, b, s)\right].$$

Nevertheless, if some spells are left-censored, the exponential distribution leads easily to the likelihood function. Other distributions, including the Weibull are problematic. The reason is that the distribution of *duration* before and after the beginning of the study period are different, except for the exponential distribution (Heckman and Singer 1985). When the distribution of completed spells is exponential, we can ignore the duration distribution before the beginning of the study period because of the ``memoryless'' property. However, if we solve the left-censored problem by assuming that the distribution of T is exponential, Heckman and Singer (1985) have shown that the corresponding ML estimator of the parameters is biased and inconsistent when this assumption is false. Because there is no standard approach to deal with left-censored problem, we will apply Heckman's procedure to predict the starting year of the left-censored spells to overcome the left-censored problem.

B. Heterogeneity and Mixture Models

In the parametric methods discussed so far, we have assumed that the hazard function and the survival function are homogeneous across individuals, or that explanatory variables included in the model can completely control for heterogeneity. Economic data, however, are seldom homogeneous. They generally consist of measured, and possibly unmeasured, differences between economic agents. When we include explanatory variables in an econometric model, it is not only to help identify the effect of a particular variable, but also to control for heterogeneity. Heckman and Singer (1985) have shown that if heterogeneity is present and we do not control for it, the parameter estimates of the hazard function will be biased toward negative duration dependence (Heckman and Singer 1985). Lancaster (1985), also, found that the effect of heterogeneity in the Weibull model causes the ML estimator $\hat{s}$ to be biased upward, i.e. to bias downward the estimated hazard function duration dependence, and the ML estimator $\hat{b}$ is biased toward zero.

There are several arguments which lead us to consider models with heterogeneity, called mixture models (Lancaster 1990, ch. 4). First, we may record the duration, the realization of a random variable T, with error. Second, there may exist a measurement error in covariates X. Third, if the covariates X fail to account fully for the true differences among individuals, omitted regressors will cause heterogeneity.

Lancaster (1979) suggested the following mixture model. Let V, distributed as gamma with unit mean[4], represent the effect of heterogeneity, which is independently distributed with X and T, multiplying the hazard function. For the Weibull distribution of duration, the mixture survivor function is given by:

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[3] The p.d.f. for e is $g(e) = \exp\{e - \exp(e)\}, -\infty < e < \infty$ (Kalbfleisch and Prentice 1980).

[4] The unit mean is not necessary. If we do not have the unit mean instead of the finite mean, we can always assimilate the deviation from the unit mean into the rest of the hazard function (Lancaster 1990, ch. 4).

$$\textbf{(13)}\qquad S_m(t, X, b, s, q) = \int \exp\left(-\int_0^t v\, l(z, X, b, s)dz\right) f_V(v)dv$$

$$= \int \exp\left(-v[t\exp(-X'b)]^{1/s}\right) f_V(v)dv = \left(1 + q\,[t\exp(-X'b)]^{1/s}\right)^{\frac{-1}{q}}$$

where q is the variance of V. The mixture hazard function is derived by differentiating $-S_m(\cdot)$ w.r.t. t giving

$$\textbf{(14)}\qquad l_m(t, X, b, s, q) = [S_m(t, X, b, s, q)]^q \left(\frac{1}{s}\right) t^{(1/s)-1} [\exp(-X'b)]^{1/s}.$$

Note that

$$\textbf{(15)}\qquad \lim_{q \to 0} l_m(t, X, b, s, q) = \left(\frac{1}{s}\right) t^{(1/s)-1} [\exp(-X'b)]^{1/s} = l(t, X, b, s)$$

which is the hazard function without heterogeneity. Hence, q can capture the sensitivity of the hazard function to unmeasured heterogeneity; and the effect of heterogeneity is larger when q is further away from zero. Given the mixture hazard function and the survival function, the likelihood function can be constructed to find the ML estimators $\hat{b}$, $\hat{s}$, and $\hat{q}$. Note that the value of s cannot simply indicate the time dependence of the hazard rate.

C. Time Dependent Covariates

In the hazard rate models we have discussed so far, the covariates X are constant over a spell. This is true, for example, with gender and race, but it is not an appropriate assumption for an individual's age and labor market conditions. Actually, most economic covariates vary over time. Nevertheless, if time dependent covariates are included in the econometric model, a new set of issues arise. Consider a hazard function conditional on the time dependent covariates $X(t)$, $l(t, X(t), b)$. The corresponding survivor function is given by

$$(16)\qquad S(t, X(t), b) = \exp\left(-\int_0^t l(z, X(z), b)dz\right)$$

which does not have a closed-form expression, except for the special constant-value time-path of the covariates, and requires numerical integration to evaluate it. The latter is a difficult task.

One of two expedients is often adopted to overcome this difficulty. Consider replacing covariates $X(t)$ by $\bar{X}(t) = (1/t)\int_0^t X(z)dz$ or the average of $X(t)$ within a spell. Alternatively, consider replacing *X(t)* by it *X(0)* or the value of *X* at the beginning of a spell. The likelihood function for these two cases is the same as derived before, except $\overline{X}(t)$ or $X(0)$ replaces $X(t)$. However, the first treatment $\overline{X}(t)$ can have the undesirable effect of building in spurious relationships between duration length and regressors; for example, $X(z) = c + dz$ implies $\overline{X}(t) = c + dt$, which causes a linear dependence between $\overline{X}(t)$ and t (Heckman and Singer 1985). The second treatment $X(0)$ ignores the time heterogeneity in the environment (Heckman and Singer 1985).

We will approximate the hazard function by step-functions (Petersen1986), i.e., the time dependent covariates are assumed to be constant within each period, but may change from one period to the next. Figure 1 plots the relationship between time dependent covariates and the step functions. The solid line plots the true value of an explanatory variable. The horizontal dash lines represent the step function used to approximate a time dependent explanatory variable.

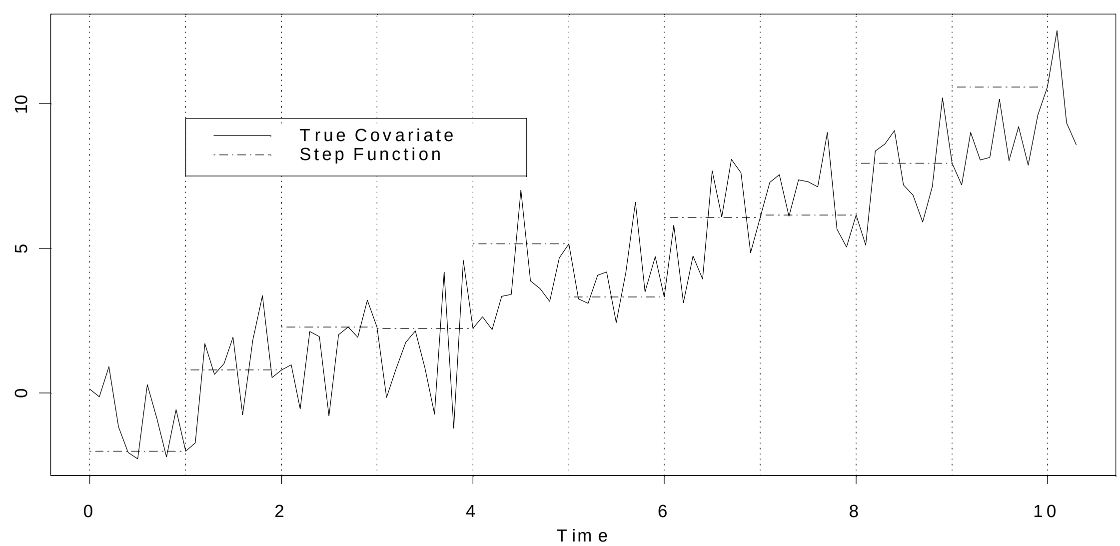

Figure 1
The relationship between Time-dependent covariates and Step function

IV. Data Description and Empirical Specification

A. Data Description

The Public Use Microdata Samples (PUMS), issued by the Bureau of the Census, provides a wide variety of information at the household and the individual level. The sample is stored in a hierarchical file structure in which household and individual information make up each household unit record. Most importantly, the U.S. sample contains information for the year when an individual came to the U.S. mainland; and the Puerto Rico sample contains information about the individual's residence on the U.S. mainland (for more than 6 consecutive months during 1980s). If a Puerto Rican lived on the U.S. mainland for more than 6 consecutive months, then it gives his length of time on the mainland and which activity he performed. With the duration and activity information on an individual residing on the U.S. mainland and the information about personal characteristics, we have the data needed to be able to fit the empirical hazard rate models.

The data we utilize in the empirical re-migration analysis are drawn from the 1990 5-percent samples of the PUMS for Puerto Rico and the United States. We restrict our sample to Puerto Rican male householders, 18 to 64 years of age in 1990, born in Puerto Rico, residing on the U.S. mainland or residing in Puerto Rico, but who had re-migrated from the U.S. mainland during the 1980s. We further excluded those individuals who were in the armed forces, self-employed, and enrolled in school. Each record includes personal characteristics (age, education, etc.), local market conditions, how long the individual has lived on the U.S. mainland, and when he re-migrated to Puerto Rico.

The sample for this study consists of 12,108 observations on mainland working spells of Puerto Rican born males, 2,544 from the Puerto Rico sample and 9,564 from the U.S. sample. If we knew the starting time for a migration spell for each individual, we would have 2,544 completed spells and 9,564 right-censored spells. Then, the likelihood function introduced in the previous section could be used to locate the ML estimates of the parameters. Unfortunately, for some individuals, data are missing on the date when a spell began. Our sample contains 1,896 left-censored observations. From the Puerto Rico sample, there are 1,682 completed spells and 862 left-censored spells. From the U.S. sample, there are 8,530 right-censored spells and 1,034 both left- and right-censored spells. However, we are only interested in the working spells of individuals, so we exclude any period in school. We assume further that the individual can perform a full-time job only if he is older than age 18 and has completed his education. More precisely, the duration of the spell is defined empirically to be

$$\mathbf{min}\left\{T^{S}\mathbf{,}\quad \mathbf{(}ENDAGE-6-ED\mathbf{),}\quad \mathbf{(}EDNAGE-18\mathbf{)}\right\}$$

where T^{S} is the duration recorded in the sample, *ED* is the years of formal schooling completed, and *ENDAGE* is the age in 1990 for the U.S. sample and the age when the last migration spell was complete for the Puerto Rico sample. After adjustments, we still have 1,183 left-censored spells; the sample consists of 1,736 completed spells and 808 left-censored spells (both from the Puerto Rico sample), and 9,189 right-censored spells and 375 both left- and right-censored spells (both from the U.S. sample).

Recall that the Weibull distribution for the spells is adversely affected by left-censored spells. There is some advantage of staying with the Weibull rather than shifting to the exponential distribution because the exponential distribution is a special case of Weibull. We, however, must fix the left-censored spells. We follow Heckman's procedure of predicting the individual's age at which these spells started and then computing the spell length using these predicted ages. Predicting the starting age rather than the length of the left-censored spells is less likely to distort the distribution of the spell lengths or to cause biased parameters in the hazard rate model. See Appendix A for more details.

B. Empirical Specification

The endogenous variable is the length of time a Puerto Rican born male spent working on the U.S. mainland or duration of the working spell. The wage structure on the mainland and Puerto Rico are expected to be important factors affecting an immigrant's re-migration decision. Because wage rates seem unlikely to be exogenous to migration and re-migration decisions, consider the predicted wage from an equation fitted to cross-sectional data as an instrument for the actual wage. This approach, however, provides only a snapshot of the wage structure at one point in time. The coefficients in hedonic wage equations may reasonably be constant, when adjusted for inflation, over modest length time periods but seem unlikely to be constant over forty years which is the length of time spent on the mainland by the Puerto Rican migrant with the longest duration in the sample. An alternative approach is to substitute the regressors of the wage equation directly into the re-migration equation replacing the predicted wage. By applying this methodology, we can solve several problems. There is, however, no free lunch and the price we pay is that we cannot obtain direct estimates of the impact of the U.S. and Puerto Rican wage rates on the hazard rate for re-migration.

Prime candidates for wage equation regressors are an individual's own human capital and local- and birthplace labor market conditions. An individual's years of schooling completed represents general human capital that is valuable in the labor market on the U.S. mainland and Puerto Rico and is a key wage determinant. An individual's education also affects information processing skills that can reduce transactions costs associated with the decision to relocate (Huffman 1985). An individual's age captures individual and family life-cycle effects on consumption and work. Given that human life is finite, an individual's age is an indicator of length of remaining life over which returns from re-migration could be obtained. An individual's age is also strongly correlated with potential labor market experience. The later generally has a quadratic marginal effect on the wage or earnings of an individual. For these reasons, it seems important to permit nonlinear effects of an individual's age on his re-migration decision.

An individual's English proficiency and disability status are two qualitative human capital variables that seem likely to affect the re-migration decision. English proficiency is a valuable skill for working on the U.S. mainland but of limited value for working in Puerto Rico. Greater English proficiency is expected to reduce the tendency to re-migrate. When an individual with a disability earns less in the market than able-bodied individuals, being disabled reduces the size of the mainland – Puerto Rico wage differential. This seems likely to weigh against mobility.[5]

Labor market conditions on the mainland and in Puerto Rico will affect both hours of work and wage rates, given employment. In addition, these labor market outcomes seem likely to be affected by the Puerto Rican minimum wage policy. We focus on the employment growth rate and the unemployment rate. Previous studies (Tokle and Huffman 1991, and Topel 1986) suggest that economic agents respond to expected values of local labor market variables rather than the actual values. The predicted employment growth rate (unemployment rate) measures the anticipated equilibrium employment growth rate (unemployment rate).

We model the annual employment growth rate and the annual unemployment rate for both

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[5] We do not include family status in the model because it is not clearly exogenous to the decision to migrate (Sandefur and Tuma 1987).

the U.S. mainland and Puerto Rico as stationary autoregressive (AR) processes. In other words, the stochastic process $\{y_t\}$ with constant mean m, say the U.S. annual unemployment rate, is assumed to be generated by

(17) $$(y_t - m) = f_1(y_{t-1} - m) + f_2(y_{t-2} - m) + \cdots + f_p(y_{t-p} - m) + u_t$$

where u_t is white noise with zero mean and finite variance s_u^2, and the roots of $1 - j_1 z - f_2 z^2 - \cdots - f_p z^p = 0$ are outside the unit circle. Then, the predicted values are constructed by the one-step-ahead predicted values. The parameters in equation (17) are estimated by the maximum likelihood method, using annual data for 1945–93 in the U.S. mainland and for 1950–93 in Puerto Rico.[6] Furthermore, the order of these AR processes is chosen by minimizing the Akaike information criterion (AIC).

Table 3 shows the summary of AR models for these four series. The job growth rate for the U.S. mainland and the unemployment rate for Puerto Rico are represented by an AR process of order 2, while the other two series are best represented by an AR process of order 1. Note that we implicitly assume that the processes of the annual employment growth rates and the annual unemployment rates in both the U.S. mainland and Puerto Rico are weakly stationary and time invariant.

Castillo-Freeman and Freeman (1992) have showed that imposing the U.S. minimum wage on Puerto Rico created marked spikes in the distribution of earnings in the area of the minimum wage, and most migrants from Puerto Rico to the mainland came when they were locally unemployed. Furthermore, the minimum wage policy is one of the reasons for higher unemployment rates in Puerto Rico (Ramos 1992, and Castillo-Freeman and Freeman 1992), and Reynolds and Gregory (1965) and Castillo-Freeman and Freeman (1992) found that the minimum wage resulted in substantial employment loss. Therefore, we also add an interaction term between the Puerto Rican predicted unemployment rate and the real minimum wage.

[6] The data in Puerto Rico are not available for 1941–1949.

Table 3. Summary of AR models for the annual Employment Growth Rates and the Annual Unemployment Rates for the U.S. Mainland and Puerto Rico

	Mean	Order	f_1	f_2
Job growth rate For the U.S. mainland [a]	1.6975	2	0.2280	-0.2604
Unemployment rates For the U.S. mainland [a]	5.5837	1	0.9783	--
Job growth rates For Puerto Rico [b]	1.6238	1	0.1081	--
Unemployment rates For Puerto Rico [b]	15.2128	2	1.2726	-0.2792

[a] Contain the annual data for 1945–1993.
[b] Contain the annual data for 1950–1993.

Thus, the empirical specification of the $X_i(t_i)'b$ part of the hazard function $l(\cdot)$ is:

(18) $$X_i(t_i)'b = b_0 + b_1 AFE_i(t_i) + b_2 AGESQ_i(t_i) + b_3 ED_i + b_4 ENG_i + b_5 DISAB_i + b_6 PJRUS_i(t_i) + b_7 PURUS_i(t_i) + b_8 PJRPR_i(t_i) + b_9 PURPR_i(t_i) + b_{10} PRMIN_i(t_i) + b_{11} PRMINUR_i(t_i)$$

where ED, ENG, and DISAB are constant over the migration spells. The definitions of variables in equation (18) are presented in Table 4.

Table 4. Variable Definitions and Sample Means for the Hazard Function

Variables	Description	Sample Mean		
		Remaining	Re-migrated	Total
T	Duration of mainland working spell (years).	20.28	5.12	12.24
AGE	Individual's age, in years.	43.29	38.56	42.29
AGESQ	Square of AGE divided by 100.	19.95	16.41	19.20
ED	Highest grade of school completed.	10.43	9.44	10.23
ENG	1 if respondent reported speaking English well or very well; 0 otherwise.	0.82	0.49	0.75
DISAB	1 if individual reported a health condition that limited the kind of work or amount of work he would do; 0 otherwise.	0.14	0.18	0.15
PJRUS	Predicted job growth rates for the U.S. mainland.	1.64	1.59	1.63
PURUS	Predicted unemployment rates for the U.S. mainland.	5.31	7.08	5.68
PJRPR	Predicted job growth rates for Puerto Rico.	1.53	1.78	1.59
PURPR	Predicted unemployment rates for Puerto Rico.	14.49	18.53	15.30
PRMIN	Real minimum wages on Puerto Rico in 1990 dollars	3.80	4.11	3.86
PRMINUR	Interaction of PURPR and PRMIN	55.07	75.77	59.36

A brief discussion of the expected signs of the b_js in equation (18) follows. While an individual's age has several different types of effects on the hazard rate of re-migration or expected length of the mainland working spell, we expect the net effect to give $b_1 < 0$ and $b_2 > 0$.

Castillo-Freeman and Freeman (1992) have shown that for Puerto Rico-born men, the economic return in 1970 and 1980 to schooling from working in Puerto Rico is higher than working on the U.S. mainland in 1970 and 1980; moreover, the gap was increasing from 1970 to

1980. On the other hand, individuals having more education are likely to migrate because of greater efficiency in acquiring and processing information. Both wage and non-wage effects suggest that education has a positive effect on the conditional probability to re-migrate. Furthermore, the study by Ramos (1992) found that Puerto Ricans who return to Puerto Rico tend to be more skilled than those who remain on the U.S. mainland. We expect b_3 to be negative. Because poor English proficiency is a disadvantage for work on the U.S. mainland, we expect b_4 to be positive. Individuals having work disability status are less likely to migrate, and we expected b_5 to be positive.

We assume that a higher predicted job growth rate and/or lower predicted unemployment rate is a more attractive labor market for local workers. Therefore, we expect $b_6 > 0$, $b_7 < 0$, $b_8 > 0$, and $b_9 < 0$. We expect a negative effect of the real minimum wage in Puerto Rico on the hazard rate for re-migration. Furthermore, the minimum wage policy is one of the reasons for higher unemployment rates in Puerto Rico (Ramos 1992, and Castillo-Freeman and Freeman 1992). Reynolds and Gregory (1965) and Castillo-Freeman and Freeman (1992) also found that the minimum wage resulted in substantial employment loss. Thus, for the minimum wage-unemployment interaction variable $RMINUR_i(t_i)$, we expected the sign of b_{11} to be positive.

Table 4 reports sample mean values for the regressors. Sample means are derived for those who re-migrated during 1980s and those who remained on the U.S. mainland at the end of the study period (1990). The re-migrating group consists of completed and adjusted left-censored spells, while the group remaining contains right- and both right- and left-censored spells. Note that the sample mean for variables of the re-migrating group is based on the year when the individual re-migrated, but the mean for variables of the group remaining is calculated for 1990.

V. Empirical Results

The empirical hazard rate model without and with heterogeneity effect is fitted by the maximum likelihood method and reported in Table 5. The dependent variable is the natural logarithm of the length of an individual's working spell on the U.S. mainland, measured in years. The heterogeneity effect (θ) is represented by gamma heterogeneity, giving a mixed model. In the Weibull version of the hazard rate model, the coefficient vector b of explanatory variables is interpreted as the constant proportional effect of the covariates on the length of a completed spell. However, the marginal (proportional) effect of an explanatory variable on the hazard rate of re-migration is in general $-\ b_j/s$. In Table 5, the signs of the estimated coefficients are as expected across the two specifications of the hazard rate model, and all the coefficients are significantly two different from zero at the 1 percent significance level, except for the coefficient for disability.

For the model without heterogeneity, the estimate for $\hat{s}$ is 1.623, and it is statistically significantly different from one at the 1 percent significant level. This estimate implies a decreasing marginal effect of length of mainland employment duration on the hazard rate of re-migration. Other things being equal, as length of time a Puerto Rican resides on the mainland increases, his hazard rate for re-migration decreases. In Figure 2, part a, the relationship between the mainland spell duration and the predicted hazard rate of re-migration is graphed. Explanatory variables are set equal to sample means values. This graph suggests the hazard rate decreases exponentially with mainland residency.

The hazard rate model with gamma heterogeneity gives a different picture. In Figure 2, part b, the pattern of the predicted hazard rate of re-migration for the mixture model[7] is graphed. The relationship is now an inverted U-shaped curve. The maximum hazard rate for re-migration (0.8%) occurs when a Puerto Rico-born man has worked on the U.S. mainland for 3.75 years. Thereafter, the hazard-rate declines for additional duration on the mainland. The relationship in part b of Figure 2 is much more economically appealing than in part a. The graph with heterogeneity suggests that individuals who migrate tend to stay a while on the mainland before feeling a strong incentive to re-migrate. This provides time whereby they can capture returns on the Puerto Rico-to-mainland move.

[7] The value of s can not simply indicate the time dependence of the hazard rate in the mixture model.

Table 5. The Hazard Function for Re-migration of Puerto Rico-Born Male Householders during the 1980s from the 1990 Census Data in Weibull Regression Models

Explanatory Variables	Without Heterogeneity: Coefficients (t-value)		Marginal Effect On Hazard Rate of Re-Migration	With Heterogeneity: Coefficients (t-value)		Marginal Effect On Hazard Rate of Re-Migration
Intercept	59.178	(11.46)		58.701	(12.71)	
AGE	0.140	(7.45)	0.028 [a]	0.098	(4.67)	0.055 [a]
AGESQ	–0.219	(–9.44)		–0.177	(–6.30)	
ED	–0.034	(–3.52)	0.021	–0.045	(–4.59)	0.048
ENG	2.202	(24.17)	–1.357	2.215	(24.08)	–2.351
DISAB	0.118	(1.29)	–0.073	0.183	(1.76)	–0.194
PJRUS	2.498	(15.35)	–1.539	2.803	(19.24)	–2.976
PURUS	–0.614	(–10.37)	0.378	–0.863	(–13.83)	0.916
PJRPR	–3.747	(–12.63)	2.309	–4.004	(–16.37)	4.251
PURPR	–2.973	(–9.59)	–0.018 [b]	–3.092	(–11.09)	–0.041 [b]
PRMIN	–12.489	(–9.39)	–0.584 [c]	–12.103	(–10.42)	–2.092 [c]
PRMINUR	0.721	(8.75)		0.752	(10.30)	
s	1.623	(32.25)		0.942	(23.95)	
q				1.984	(13.25)	
Total Spells	12,108			12,108		

[a] Evaluated at the sample mean of AGE, 42.295.
[b] Evaluated at the sample mean of PRMIN from 1980 to 1990, 4.163.
[c] Evaluated at the sample mean of PURPR from 1980 to 1990, 18.636.

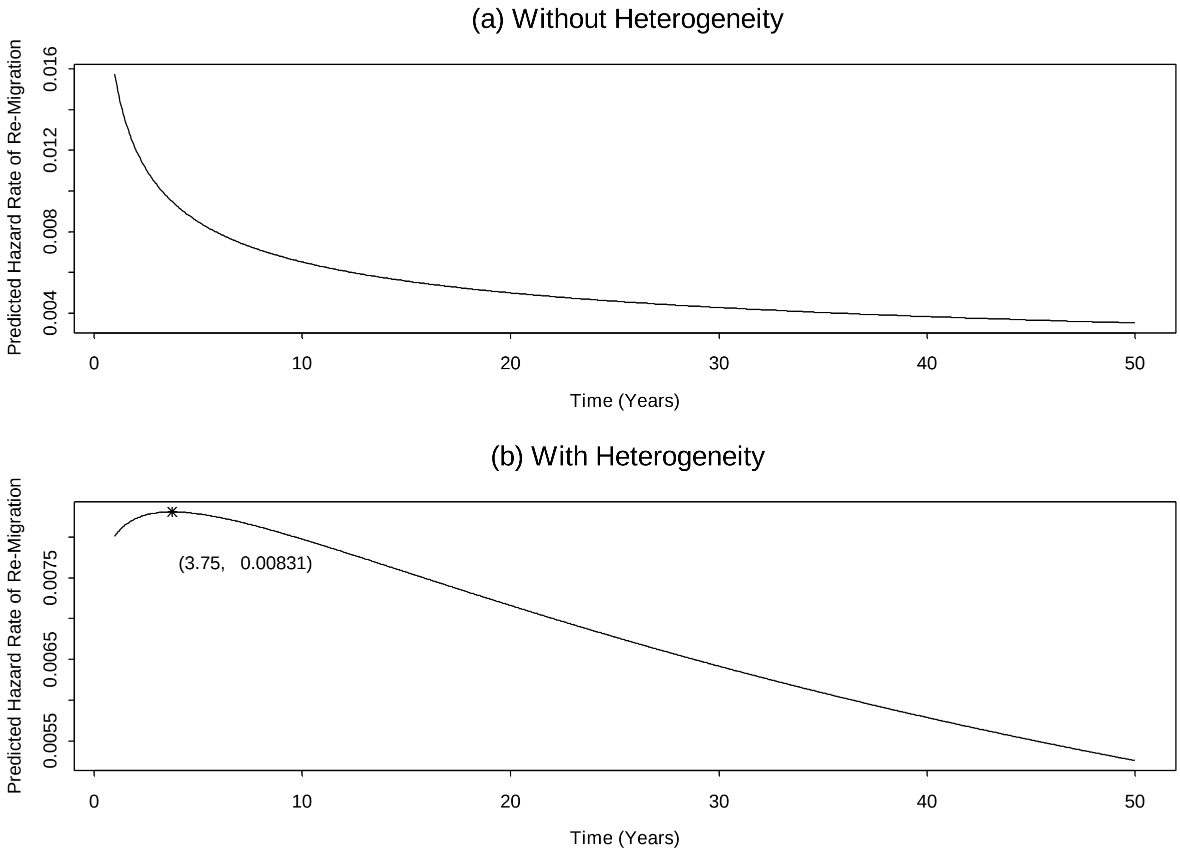

Figure 2
The Predicted Hazard Rate of Re-Migration for the Puerto Rico-Born Male Householder, Evaluated evaluated at the sample means of explanatory variables.

The estimated parameter $\hat{q}$, measuring the effect of heterogeneity, is 1.984 and statistically significantly different from zero at the 1 percent significant level. Hence heterogeneity across individuals seems to exist in hazard and survivor functions. The inclusion of gamma heterogeneity tends to reduce the effect of age, age square, and the real minimum wage in Puerto Rico, while it strengthens the effect of the other coefficients. We conclude that the model with gamma heterogeneity performs better than the model without heterogeneity.

We focus the remainder of the discussion of our empirical results for the model with heterogeneity. The quadratic effects of AGE on the expected duration of a mainland spell or on the hazard rate implies that the hazard rate for re-migration decreases as an individual gets older up to 28 years of age and thereafter it increases. Thus, when a Puerto Rican on the mainland is less than 28 years of age, the conditional probability of return migration decreases with each birthday, but when he is older, say ≥ 28 years of age, the probability of return migration increases with each birthday and becoming relatively large by age 60 (see figure 3).[8] The quadratic lifecycle effect is consistent with finite life affecting re-migration decisions (and with returning to one's "roots" in retirement). The results are consistent with Tienda and Wilson (1992) and Pissarides and Wadsworth (1989).

Increasing an individual's schooling by one year reduces by 4.5 percent the duration of mainland working spells for Puerto Rico-born males. This supports the prediction that individuals with higher education are likely to migrate because of their efficiency in acquiring and processing information and with Ramos' (1992) finding that Puerto Ricans who return to Puerto Rico tend to be more skilled than those who remain on the U.S. mainland and with Castillo-Freeman and Freeman's (1992) finding that for Puerto Rico-born men, the economic return to schooling in Puerto Rico is higher than on the U.S. mainland.

An individual's English proficiency plays a significant role in return migration decisions for Puerto Rico-born men. It reduces the hazard-rate for re-migration. The spell duration on the mainland for Puerto Rico-born males with English proficiency is 2.22 times longer than for those with poor English proficiency. This empirical result supports the hypothesis that English proficiency is an important form of human capital for working on the U.S. mainland.

The results suggest that that work disability lengthens the duration on the U.S. mainland by about 18.3 percent. Its coefficient is significantly different from zero only at the 10 percent significance level. Overall, the results supports the prediction that individuals with disability are less likely to make a second move.

The empirical results yield strong evidence for the effect of labor market conditions on re-migration decisions. Predicted job growth rates provide a strong factor in return migration decisions. The results show that a 0.1 percentage point increase in the predicted U.S. mainland job growth rate lengthens the mainland spell duration by 28 percent, while a similar increase of the predicted job growth rate for Puerto Rico reduces the mainland spell duration by 40 percent. This supports the prediction that people are attracted to high job growth rate areas. The empirical

[8] The graph in figure 2(b) is sensitive to the value of an individual's age. Because the sample mean of education is 10.23 years, the individual by assumption began his working life on the U.S. mainland when he was 18 years of age. Hence, the spell duration can be derived as an individual's age minus 18. Using this assumption, figure 3 shows that the hazard rate achieves the minimum when he is 29 years old; then the hazard rate for re-migration increases as he gets older.

results show that a 0.1 percentage point increase in the predicted unemployment rate for the U.S. mainland reduces by about 9 percent the duration of the mainland working spell. The proportional effect of the predicted unemployment for Puerto Rico on the duration of mainland working spells of Puerto Rico-born males on the U.S. mainland is $b_9 + b_{11} PRMIN$. When we evaluate this effect at the mean of PRMIN from 1980 to 1990, which is 4.163, the result is a one percentage point increase in the predicted unemployment rate for Puerto Rico lengthens the spell duration by 3.9 percent. Similarly, the proportional effect of the real minimum wage in Puerto Rico on the mainland spell duration is $b_{10} + b_{11} PURPR$. At the mean of PURPR from 1980 to 1990, which is 18.636, the results show that a 10 cents increase in the Puerto Rican real minimum wage lengthens the mainland spell duration by to 19.7 percent. The estimate of b_{11} which is the coefficient of the interaction variable for the Puerto Rican predicted unemployment rate and real minimum wage, is positive. Given that the estimates of b_9 and b_{10} are negative, this interaction effect has a moderating effect.

For each variable, its marginal effect on the hazard rate of re-migration is opposite the sign of its marginal effect on the length of a working spell on the mainland. The marginal

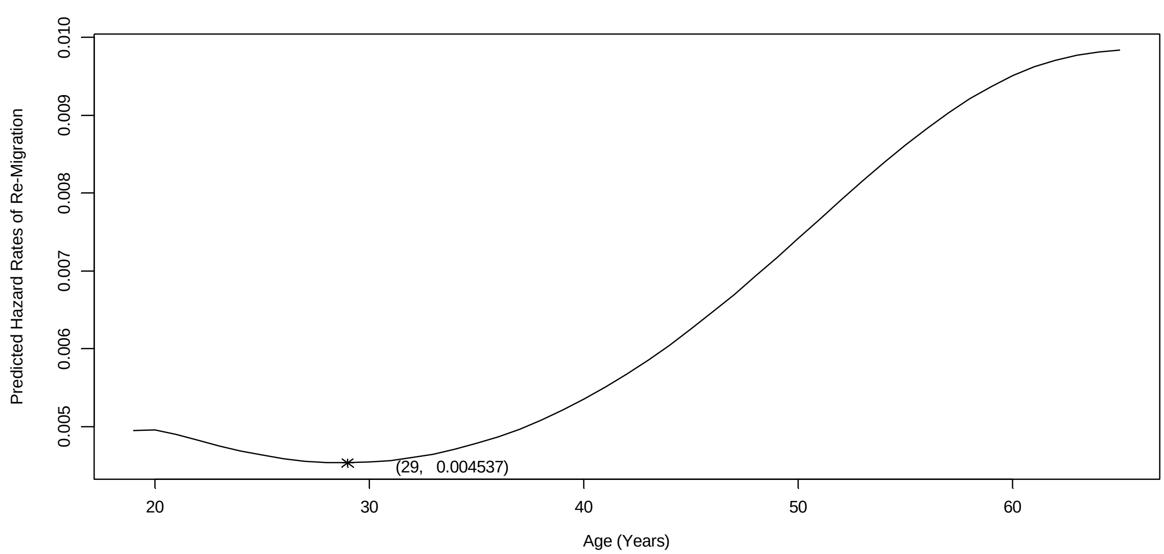

Figure 3
The Simulated Life-Cycle Effect on the Hazard Rate of Re-Migration for the Puerto Rico-Born Male Householder, Evaluated on Estimated Coefficients From Table 5 with Gamma Heterogeneity and Sample Means Except Spell Duration and Age

proportional effect of each explanatory variable on the hazard rate of re-migration for Puerto Rico-born males are also reported in Table 5, but they are not discussed further.

VI. Conclusion

This study has explored a largely new research area under the broad topic of migration—re-migration or returning to ones' birthplace after spending some time working in another region. In particular, the study examined the re-migration decisions of Puerto Rico-born males who migrated to the U.S. mainland and then returned to Puerto Rico. We showed that the hazard rate model was a useful tool for this empirical analysis. Although we included major personal attributes and U.S. and Puerto Rican labor market variables in the hazard rate model, unmeasured heterogeneity effects were shown to be significant and to lead to somewhat different results than a model without heterogeneity. In particular, the implication for the effect of an individual's age on the length of a mainland employment spell are quite different. The hazard rate model with heterogeneity gives the intuitively appealing result that the effect of an individual's age on his probability of re-migration is low and decreasing up to 28 years and it is increasing after age 29 and becomes relatively large by age 60. This pattern is consistent with migration and re-migration being human capital investment decisions which are strongly affected by an individual's finite work life (and with returning to one's roots in retirement). The model without heterogeneity gives the highest probability of re-migration during the first year on the mainland. This pattern is largely inconsistent with a motive of mainland migration for work.

Although we have chosen to fit the model to data for individuals who are moving as citizens from one region to another, the U.S. mainland and Puerto Rico have different languages and different cultures and a move between Puerto Rico and the mainland requires a significant resource investment. Thus, we believe that our results are generalizable to other re-migration situations where individuals are moving across country boundaries. In particular, our results seem easily generalizable to explain re-migration of other individuals from the U.S. mainland, e.g., Mexicans and Central Americans who have migrated to the U.S. for work.

Appendix A

The time dependent hazard model requires that all working spells be either completed or right-censored to construct the true likelihood function. Unfortunately, our data drawn from the PUMS for Puerto Rico and for the U.S. do not provide enough information to complete spells for some individuals. There exist 2,111 left-censored spells in our data set. There is no standard approach to deal with left-censored problem. Therefore, the expedient we use to overcome the left-censored problem is to predict the starting year of the left-censored spells, based on all completed or right-censored spells. Because there exists a non-random selection, Heckman's procedure (Heckman 1979) will be used to correct for possible selection bias. However, if we try to predict the length of spells directly, i.e., treat spells or natural logarithm of spells to be the dependent variable, two potential problems arise. First, Heckman's procedure requires the duration (or natural logarithm of duration) of spells to be distributed normally, but we already assumed that completed duration T is a family of two-parameter Weibull distribution in the time dependent hazard model. Hence, some conflicts occur because we have inconsistent assumptions about the distribution of completed duration. Secondly, the natural logarithm of completed duration is the dependent variable in the time dependent hazard model. If we predict the length of duration based on a set of exogenous variables and use this predicted length of duration as the dependent variable for left-censored observations, ML estimates for hazard rate models using some of the same exogenous variables included in predicting the length of duration might overestimate the contribution of exogenous variables used in the first stage predictions. Alternatively, we try to predict the individual's age when he began working on the U.S. mainland, based on a set of exogenous variables. This variable can uniquely identify the starting year of mainland working spells for all individuals; more importantly, it does not have the above two problems.

We have 808 left-censored spells in the Puerto Rico sample and 375 both right- and left-censored spells in the U.S. sample. Hence, there are 1,285 spells of unknown length out of a total of 12,108 spells to be predicted. Since our data come from two different population, we will predict the unknown spells separately. For the Puerto Rico sample, there are 808 out of 2,544 left-censored spells which need a prediction of starting-age of working spells; hence, a 68.24-percent sample will be used to fit the predicting equation. We first fit a binary probit model, based on a normality assumption, to explain whether a Puerto Rico-born males worked on the U.S. mainland more than 10 years (left-censored spells in the Puerto Rican sample are those people who worked in the U.S. mainland more than 10 years.) with dependent variable equal to 1 if people had worked on the U.S. mainland more than 10 years, and 0 otherwise. Explanatory variables contain personal characteristics and local characteristics. The detailed variable definition used in the probit model and the fitted equation are reported in Table A 1.

Our purpose is to fit a multiple regression model in order to predict an individual's age when he began working on the U.S. mainland, based on a set of explanatory variables. However, the sample we can use to estimate the corresponding parameters only consists of completed spells. It seems likely that non-random selection bias exists. Based on Heckman's procedure, we can construct the non-random selection bias correction, known as the inverse of Mill's ratio, from the above fitted probit model. In other words, by using the sample of completed spells, we regress the natural logarithm of an individuals' starting-age for a mainland working spells on a set of explanatory variables including the inverse Mill's ratio to generate consistent estimates of this equation's parameters. The fitted equations can be found in Table A 2. Then, we predict the starting-age for an individual’s mainland working spells for those persons with left-censored spells, based on the multiple regression model (excluding the inverse of Mill's ratio) fitted from the sample of completed spells.

Finally, the length of migration or mainland working spells for individuals with left-censored spells can be derived as the ending-age of a spell minus the predicted starting-age. Nevertheless, we only use these predicted spell lengths when they are necessary. In other words, when predicted durations are less than 10 years, the corresponding spells used in the time dependent hazard model are 10 years instead of the predicted duration; moreover, even though the predicted durations are greater than 10 years, they need adjustment (if it is necessary) to satisfy the restriction that a individual is able to perform a full time job only if he is over 18 years of age and has completed his education.

Similar procedures are applied to the U.S. sample. There are 375 unknown migration spells out of 9,574, i.e., a 96.1-percent sample is used to fit the predicting equation. The only difference between Puerto Rico and U.S. sample missing data problems is that the binary probit model tries to explain whether a Puerto Rico-born men worked on the U.S. mainland more than 40 years with dependent variable equal to 1and 0 otherwise. This is because left-censored spells in the U.S.

sample arise only in individuals who have lived in the U.S. mainland more than 40 years.

Table A 1

Variable Definition

Variables	Definition
AGE	The ending-age of a spell, in years.
AGE^2	$AGE^2/100$.
AGE^3	$AGE^3/1000$.
AGE^4	$AGE^4/10000$.
ED	Highest grade of school completed.
ED^2	$ED^2/100$.
ENG	1 if respondent reported speaking English well or very well; 0 otherwise.
EDENG	Interaction of ED and ENG.
DISAB	1 if respondent reported a health condition that limited the kind of work or amount of work he would do; 0 otherwise.
RACE	1 if white; 0 otherwise.
PUSEM	Predicted job growth rates for the U.S. mainland.
PUSUN	Predicted unemployment rates for the U.S. mainland.
PPREM	Predicted job growth rates for Puerto Rico.
PPRUN	Predicted unemployment rates for Puerto Rico.
PRMIN	Real minimum wages in Puerto Rico in 1990 dollars.
PRMINUN	Interaction of PPRUN and PRMIN.
RUSEM	Residual of job growth rates for the U.S. mainland.
RUSUN	Residual of unemployment rates for the U.S. mainland.
RPREM	Residual of job growth rates for Puerto Rico.
RPRUN	Residual of unemployment rates for Puerto Rico.
$\hat{l}$	Inverse of Mill's ratio.

Table A.2 Estimated Coefficients of Binary Probit for Mainland Work and Regression Model for starting age (t-value are in parentheses).

Covariates	Puerto Rico Sample		U.S. Sample	
	Probit Model [a]	Regression Model [b]	Probit Model [a]	Regression Model [b]
INTERCEPT	−53.381***	6.560***	2795.617***	−4.435***
	(−4.88)	(11.59)	(5.38)	(−6.33)
AGE	4.494***		−292.798***	
	(4.32)		(−5.38)	
AGE^2	−14.112***		986.201***	
	(−3.89)		(5.35)	
AGE^3	1.936***		−136.554***	
	(3.53)		(−5.32)	
AGE^4	−0.098***		6.762***	
	(−3.20)		(5.29)	
ED	0.059**	−0.019***	0.146***	−0.039***
	(2.10)	(−6.19)	(3.96)	(−21.39)
ED^2	−0.559***	0.123***	−0.961***	0.213***
	(−3.20)	(7.21)	(−4.33)	(22.01)
ENG	1.142***	−0.732***	0.284	−0.072***
	(6.69)	(−29.70)	(1.33)	(−5.38)
EDENG	−0.055***	0.047***	0.018	0.001
	(−2.96)	(22.03)	(0.62)	(0.47)
DISAB	0.133*	−0.035***	−0.047	0.072***
	(1.79)	(−3.76)	(−0.54)	(11.78)
RACE			0.272***	0.010**
			(3.33)	(2.52)
PUSEM		−0.234***		−0.698***
		(−7.84)		(−7.13)
PUSUN		−0.141***		0.110**
		(−6.06)		(2.19)
PPREM		0.085**		0.822***
		(2.35)		(10.17)
PPRUN		−0.143***		0.426***
		(−4.37)		(9.02)
PRMIN		−3.412***		7.787***
		(−5.57)		(10.62)
PRMINUN		0.211***		−0.486***
		(5.34)		(−10.12)
RUSEM		−0.079***		−0.038**
		(−4.35)		(−2.48)
RUSUN		−0.205***		−0.054
		(−5.58)		(−0.69)
RPREM		0.010*		−0.274***
		(1.80)		(−18.52)
RPRUN		0.050***		−0.491***
		(4.48)		(−12.03)
$\hat{l}$		1.028***		0.712***
		(90.65)		(48.26)
Adjust R^2		0.8568		0.4340

[a] The dependent variable is the dummy variable with the value equal to1 if Puerto Rico –born males worked on the U.S. mainland more than 10 (40) years in the Puerto Rico (U.S.) sample; 0 otherwise.

[b] The dependent variable is the Natural Logarithm of the Starting-age of a spell.
* P-value = 0.1, ** P-value = 0.05, *** P-value = 0.01.

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