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International migration under incomplete information: a re-migration analysis

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International migration under incomplete information: A re-migration analysis

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ABSTRACT

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 professsional 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. These attributes of geographical migration suggest that finite human life, economic uncertainty/imperfect information, and wage and quality of life of a home location weight into individual/family migration/emigration decisions.

This study focuses on re-migration, where individuals return to their place of birth after living in a new location for several years. The objectives of the study are (1) formulate a multiperiod finite-life utility maximization model of an individual's migration decisions, given incomplete information about wage rates and quality of life in new locations, (2) provide econometrics evidence about the contribution of personal and local-area attributes to the hazard rate for re-migration of Puerto Rican born male householders living on the U.S. mainland and returning to Puerto Rico during the 1980s, and (3) provide econometrics evidence on structural differences/changes in wage/earning equations between and within major U.S. ethnic/racial groups in the 1980s.

The following is accomplished. First, a highly stylized five period life-time model of individual decision making is developed where utility depends on the wage and/or quality of life and retirement always occurs in the fifth period. Re-migration is triggered by unfulfilled wage or quality of life expectations, unemployment, and approaching retirement. Second, the U.S. PUMS data set for Puerto Rican male householders, age 18 to 64 in 1990, born in Puerto Rico but residing on the U.S. mainland and re-migrating to Puerto Rico during the 1980s is the sample for fitting a hazard rate model of re-migration. We find a strong quadratic effect of an individual's age, which supports the finite-life conceptual model. 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 Puerto Rican real minimum wage is negatively related to the hazard rate, which is an obvious public policy instrument. Third, the 1980 and 1990 PUMS for U.S. male householders provides the data set for testing for significant change in wage structure within and across major U.S. racial/ethnic groups (Hispanic, black, and white) between 1980 and 1990. We find that the cohort effect (year of immigration) was the only part of the return on personal attributes that has changed significantly during the 1980s. There is, however, evidence of significant change in contributions of local labor market conditions to earnings of all groups during the 1980s. Test of structural equality across racial/ethnic groups are rejected.

CHAPTER 1 INTRODUCTION

The factor price equalization theorem implies that there is no incentive for international migration because all individuals can earn the same wage rate when measured in the same currency. However, trade barriers, transportation cost, heterogeneous workers, and technology frequently cause the theorem to fail. On the other hand, workers have concerns about wage rates and local quality of life, including climate, education system, political system, public infrastructure, etc. Hence, international migration continues.

Immigration has been an important source of changes in labor supply for many countries; moreover, it has been the single most important contributor to the changes in the labor supply for some countries recently. 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. Table 1.1 shows the size of the immigrant population in selected OECD countries in 1981 and 1991. The U.S. has the largest immigrant population but Luxembourg has the highest immigrant share. Other countries with significant proportions of foreign born persons are Australia, Canada, and Switzerland.

Immigrant policies differ greatly across countries. Eligibility for immigration in Canada and Australia is based on skill qualifications, but in the U.S. it is based on kinship ties. Also, all individuals born in the U.S. are automatically granted citizenship.

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

Table 1.1 Immigrant Population in Selected OECD Countries

Countries	1981[a] Total immigrants (1000s)	1981[a] % of total population	1991[b] Total immigrants (1000s)	1991[b] % of total population
Australia	3,003.8	20.6	3,753.3	22.3
Belgium	885.7	9.0	922.5	9.2
Canada	3,843.3	16.1	4,342.9	16.1
France	3,714.2	6.8	3,596.6	6.3
Germany	4,629.8	7.5	5,882.3	7.3
Luxembourg	95.4	26.1	109.1	28.4
Switzerland	909.9	14.3	1,163.2	17.1
U.S.	14,079.9	6.2	19,767.3	7.9

Source: OECD, Trends in International Migration (1994).

[a]It is 1982 for France and 1980 for the U.S.
[b]It is 1990 for the U.S., France, and Luxembourg.

changes are partly due to changes in the U.S. immigration policy. Other factors are changes in political and economic conditions in other countries, for instance, Cuba and Mexico. 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.

Table 1.2 shows the contribution of immigrants to U.S. population growth during

Table 1.2 The U.S. Immigrant Flow from 1931 to 1990

Immigrant Flow	1931 to 1940	1941 to 1950	1951 to 1960	1961 to 1970	1971 to 1980	1981 to 1990
Total (1000s)	528.4	1,035.0	2,515.5	3,321.7	4,493.3	7,338.1
% of Europe	65.8	60.0	52.7	33.8	17.8	10.4
% of Asia	3.1	3.6	6.1	12.9	35.3	37.3
% of America	30.3	34.3	39.6	51.7	44.1	49.3
% of Canada	20.5	16.6	15.0	12.4	3.8	2.1
% of Mexico	4.2	5.9	11.9	13.7	14.3	22.6
% of Africa	0.3	0.7	0.6	0.9	1.8	2.4
% of Oceania	0.5	1.4	0.5	0.8	0.9	0.6
% of change in population	5.9	5.3	8.7	13.7	20.7	33.1
% of population that is foreign-born at end of decade	8.8	6.9	5.4	4.7	6.2	7.9

Source: Adapted from Borjas 1994.

1931–1990. Immigrant flows do not exceed 8.7 percent of the total population for 1931–1960, but their share of population growth increased to 20.7 percent for 1971–1980 and a larger 33.1 percent for 1981–1990. As to country or region, during 1931–1960 over two-thirds of legal immigrants entered the U.S. from Europe or Canada, less than 6 percent from Asia, and less than 10 percent from Mexico. However, during 1981–1990 only 12.5 percent of legal immigrants admitted originated in Europe or Canada, 37.3 percent originated in Asia, and 22.6 percent originated in Mexico.

What are the economic effects of immigrant flows on the U.S.? First, they increase the supply of labor. Huffman (1995) estimated that immigrant flows (both legal and illegal) accounted for approximately 26 percent of net growth of the U.S. population during 1961–70 and 39 percent during 1981–90. Second, new immigrants increase the rate of growth of the demand for U.S. goods and services. Immigrant flows (Huffman 1995) accounted for about 8 percent of the growth in aggregate demand for U.S. nontradeable goods and services in 1960 and 14 percent in 1990. Although immigrant flows do not affect the aggregate demand for pure public goods, they do cause a disproportionately large increase in the demand for impure public goods. Third, immigrant labor slows the rate of growth in the U.S. real wage. Even if a change in immigrant flows does not have significant long-run wage effects, they do affect the national average real wage rate of U.S. low-skilled workers in general. Borjas, Freeman, and Katz (1992) estimated that they can account for 30 to 50 percent of the decline in real wage of high school dropouts for the 1980s. The other major cause is increase of U.S. trade. Hence, immigration is an increasingly important component in the economy of the United States.

One question of interest is how an individual makes a decision to emigrate to the host country or to re-migrate to the home country. Hicks (1932) argued that wages are the main cause of migration, but Galor (1986) demonstrated that the incentives for international migration are the differences in time preferences across countries. Galor and Stark (1987) emphasized that international migration is induced by technological differences between countries. The above studies are based on a full information assumption. As we know, a potential migrant generally has incomplete information about the relevant variables especially for unfamiliar foreign countries. McCall and McCall (1987) showed that more variability in the relevant variables increases the attraction of a country. Berninghaus and Seifert-Vogt (1991) illustrated that when incomplete information exists, migration can occur even among identical countries.

Literature Review

Several studies have tried to model the economic decisions facing potential migrants. In the last few decades, international labor migration has been given significant consideration. On the other hand, empirical research on immigration has been an important topic. The first part of this section presents a review of search theory as applied to migration decisions. The second part discusses selected international migration theories.

Search Theory

Search theory is an important part in the economics of information and uncertainty. It is useful for sequential decision models where a decision maker must acquire and use information to take rational action in an ever changing and uncertain environment. The theory originally was developed in labor economics to help explain unemployment duration, searching on the job, and the allocation of the working life between market work and alternative activities (Mortensen 1986).

In the full information earnings analysis, earnings differences can be attributed to different human capital investments, differences in labor force participation, as well as other variables such as gender, race, and the non-pecuniary advantage and disadvantage of jobs. However, these factors fail to explain all wage differences, i.e., earnings seem to differ across similar quality workers at the same job. It is natural to use incomplete information to describe earnings differences. The idea is that firms find it costly to evaluate workers' reservation wages even holding worker quality constant; similarly, given an evaluation of working conditions, it is difficult for workers to know the best offer firms are willing to make for a given job. Hence, information gaps may exist in the labor market.

When we analyze how workers search for jobs and choose among job offers, the workers are assumed to know the distribution of wage offers and attempt to choose the best job offer, given the search cost. The worker is assumed to continue to search for new offers as long as the marginal benefit of extra search exceeds the marginal cost. Stigler in his pioneering articles (1961, 1962) formulated the job search problem as one of an optimal sample size. In his model, the job searcher not only knows the distribution of wage offers, but also is able to search and accumulate job offers. Hence, the decision problem for the worker is to choose the sample size so as to maximize the expected net return to search.

However, if a job seeker receives an extraordinarily high wage offer relative to the wage offer distribution, it may be more efficient to accept this offer without incurring additional search cost. Subsequent search models, developed by Gronau (1971), McCall (1970), and Mortensen (1970), are based on this argument. In this model, the worker samples a wage in a fixed distribution in one period and decides whether to accept employment at the observed wage or to wait, sampling further during the next period. Therefore, the worker's optimal policy is to choose a reservation wage and to accept the first offer exceeding this reservation wage. This strategy generally dominates the optimal sample size strategy in the sense that it has higher maximal expected present value of future income (Mortensen 1986). Moreover, when a sequential strategy is used, the random sample size can be interpreted as a distribution of lengths of the random search spells which provide a way to empirically implement search models.

Assume that (1) workers maximize expected presented value of income over an infinite horizon, discounted at a constant rate r; (2) the income flow of searching job per unit period, net of search cost, is a constant b over time; (3) offers arrive according to a Poisson process with parameter δ, called the offer arrival rate; (4) a job offer is characterized by wage rate w and the job will last forever when accepted; (5) once rejected, an offer cannot be recalled; (6) job offers are independent draws from a distribution function $F(w)$, known to the workers and invariant over time. Furthermore, denote $p(n, h), n = 0, 1, \ldots,$ to represent the probability that the worker will receive n offers during a period of length h ,and let $G(w, n)$ be the probability that the best of n offers is less than or equal to w for $n \geq 1$.

Let $V(w)$ be the present value of accepting the best offer received, w, and V_S be the value of searching during the next period of length h, given the worker's information. Then, V_S can be written as:

$$V_S = bh + e^{-rh}\left[\sum_{n=1}^{\infty} p(n, h)\int_0^{\infty} \max\{V_S, V(w)\}\, dG(w, n) + p(0, h)V_S\right] \qquad (1.1)$$

or equivalently,

$$\left(1 - e^{-rh}\right) V_S = bh + e^{-rh}\left[\sum_{n=1}^{\infty} p(n, h)\int_0^{\infty} \max\{0, V(w) - V_S\}\, dG(w, n)\right] \qquad (1.2)$$

The first term on the right-hand side of equation (1.1) is the net cost of job search during a period of length h, while the second term is the expected present value of job search. The fixed-point theorem for contraction mappings[1] guarantees that there always exists a unique solution for V_S in equation (1.1), provided that the mean of the wage offer distribution is finite (Denardo 1967). The reservation wage, w^*, is uniquely determined by $V(w^*) = V_S$. Furthermore, the Poisson process implies:

$$\lim_{h \to 0} \frac{p(1, h)}{h} = \delta \quad \text{and} \quad \lim_{h \to 0} \frac{p(n, h)}{h} = 0 \quad \text{for} \quad n \geq 2.$$

[1]Let $f : \mathbf{S} \longrightarrow \mathbf{S}$ be a function from a metric space $(\mathbf{S}, d)$ to itself. The function f is called a contraction of $\mathbf{S}$ if there is a positive number $c < 1$ such that $d(f(x), f(y)) \leq c\,d(x, y)$ for all $x, y \in \mathbf{S}$ (Apostol 1974, p. 92).

Hence, dividing both sides of the equation (1.2) and taking $h \longrightarrow 0$, we obtain:

$$rV_S = b + \delta \int_0^\infty \max\{0, V(w) - V_S\}\, dF(w). \tag{1.3}$$

rV_S is represented as the imputed income, derived from V_S per unit time period.

With assumption (4), $V(w) = w / r$. Hence, the reservation wage, w^*, is obtained by solving

$$w^* = b + \frac{\delta}{r} \int_0^\infty \max\{0, w - w^*\}\, dF(w) \tag{1.4}$$

or equivalently,

$$r(w^* - b) = \delta \int_{w^*}^\infty (w - w^*)\, dF(w). \tag{1.5}$$

The left-hand side of equation (1.5) represents the marginal cost of rejecting an offer, w^*. The right-hand side is the marginal expected return from continued search using reservation wage w^*. Therefore, the reservation wage w^* is the wage rate such that the marginal benefit of search equates with the marginal cost of search. It can be shown that w^* is negatively related to r, while positively related to b, δ, and the mean of the wage offer distribution (Devine and Kiefer 1991, p. 16).

The sequence of a worker's reservation wage is the central part of theoretical search models. Because a reservation wage can not be directly observed, the empirical research is based on responses to a question such as "What is the lowest wage you would be willing to accept?". Kasper (1967) hypothesizes that reservation wages decline with duration of a spell of unemployment because of decreasing marginal utility of leisure as leisure accumulates, and increasing marginal utility of income as assets dissipate. His sample consists of unemployed workers who respond to the question "What wage are you seeking?" (Minnesota Department of Employment Security between April and September 1961). The regression result shows a 0.3 percent decline in the asking wage when the length of unemployment increases one month.

Miller and Volker (1987) performed a empirical analysis of the youth labor market in Australia. The sample drawn from the 1985 interviews of the Australian Longitudinal Survey comprised workers, aged 19–24, who were seeking a full-time job and responding to the question "What is the lowest weekly pay you would accept to work in any (full-time/part-time) job?". They reported following the regression results: (1) the reservation wage declines initially, but changes little thereafter; (2) male workers with a university degree have about a 50 percent higher reservation wage, but no such education effect appears in females; and (3) there is no significant effect of labor market experience in the reservation wage.

The above two studies focus on testing the constant reservation wage hypothesis. Warner, Poindexter, and Fearn (1980) are interested in how the reservation wage is affected by the unemployment benefit, the offer distribution, and methods of search, keeping the constant reservation wage assumption. Their data come from workers in four U.S. cities (Baltimore, Boston, Chicago, and Cleveland), collected in the 1970 Census of Employment, who experienced unemployment in the year preceding the survey and responded to the question "lowest acceptable hourly wage". The regression results show that the mean of the offer distribution and the unemployment benefit are significantly positively related to the reservation wage (p-value < 0.001), consistent with search theory. The methods of labor market search, in the Census survey, consist of newspapers, private employment agencies, state employment agencies, direct applications to employers, friends or relatives, and unions. Their study indicated that the use of friends or relatives and unions has a significantly positive effect on the reservation wage.

Empirical studies report that the reservation wage of unemployed workers falls over time. This suggests that we should extend the search model to allow non-constant reservation wage rates. One approach is to impose a liquidity constraint to assume that the worker can self-finance the out-of-packet cost of search only for a finite time period (Mortensen 1986).

Search theory also has implications for the distribution of the duration of a completed spell. Denote λ to be the instantaneous probability of exiting search or the hazard rate. Then, λ can be written as:

$$\lambda = \delta(1 - F(w^{*})). \tag{1.6}$$

The first term is the probability of receiving an offer at an instantaneous moment. The second term is the conditional probability that once an offer is received, it will be accepted or equivalently that the offer is larger than the reservation wage, w^{*}. Hazard rate λ is positively related to r and the mean of the wage offer distribution, but negatively related to b (Devine and Kiefer 1991, p. 18).

For constant δ and w^{*}, the completed duration has an exponential distribution with parameter λ and expected duration $1 / \lambda$. If one wants to specify a non-constant hazard rate, for instance, increasing hazard rate for constant δ and decreasing w^{*}, one way to modify the model is to specify the distribution of duration as a member of the family of Weibull distributions (see Chapter 3).

Lancaster (1979) focuses on the use of data on the duration of unemployment to make inferences about behavior in search theory. He pointed out that since reservation wage sequence is not observable, we can approach the econometric specification of a search model using duration data by modeling the variation in the hazard function. He considered both exponential and Weibull specification of the duration distributions and used a sample of 479 British, unskilled, unemployed people, collected by Political and Economic Planning in 1973. Furthermore, Lancaster argues that we shall attempt to allow for the specification error in the systematic sources of variation of the hazard function due to the omitted relevant regressors. He suggests a gamma mixture model to capture unobserved heterogeneity[2].

[2] Detail explanation can be found in Chapter 3.

Feridhanusetyawan (1994) applies search theory on internal migration, viewing migration decisions as sequential decisions made over time under uncertainty. He studies the migration behavior of male householders, aged 19–45 in 1968, for 20 years, based on an individual data set from the Panel Study of Income Dynamics. The empirical analysis consisted of estimating a probit model based on the point-in-time discrete choice model, and two types of hazard rate models, constant hazard and time-dependent hazard models, based on the search theory. For the time-dependent hazard model, he considers several procedures to treat left-censored migration starting time problems: remove all left-censored observations; utilize all spells with perfertly observed length; and predict the unobserved spell lengths. Feridhanusetyawan concluded that the search theory model performed better than the point-in-time standard discrete choice model in explaning migration behavior.

International Migration

The Gittins-index theory can be used to derive optimal migration policies for the multi-armed bandit (MAB) model of international migration. The MAB model of international migration can be characterized as follow:

1. a migrant has, in each period, to choose only one of N countries to live in and then, receives a bounded reward,
2. these countries depend on the state space and transition probabilities which control the state transition from one period to another,
3. the objective of the migrant is to choose the sequence of countries to live in by maximizing his/her expected discounted reward,
4. the state of the country in which the individual is living at time t is the only one which will change from period $[t, t+1)$ to the next; hence, the states of remaining

countries are kept frozen during period $[t, t+1)$.

The optimal migration strategy is to assign each country a Gittins index and then, at each decision time point, choose that country with the largest index. One way of formulating the Gittins index is to compute the value of a fallback position which would make the migrant indifferent to living in country j with state $s_j \in \mathbf{S}_j$ one further period and then proceed with an optimal strategy, and retiring from this decision problem with a terminal payment $Z_j \in \mathbf{R}$. The main advantage of the Gittins-index method is that it collapses an extremely complicated N-dimensional problem into N relatively simple one-dimensional problems.

Let $V_j(Z_j, s_j)$ be the bounded expected present discounted value of being in state s_j of country j and following an optimal stopping rule when the terminal payment is Z_j. Then, we have the following functional equation

$$V_j(Z_j, s_j) = \max\left\{Z_j,\, U_j(s_j) + \beta_j \int V_j(Z_j, Y_j)\, \Psi_j(dY_j|s_j)\right\} \tag{1.7}$$

where $\beta_j \in (0,1)$ is the discounted factor, $U_j(s_j)$ is the migrant's per period utility functions for country j with state s_j, and $\Psi_j(B_j|s_j)$ is the migrant's subjective probability that the state of country j will be in $B_j \in \mathcal{B}(\mathbf{S}_j)$[3] during period t, given that the state of country j was s_j during period $t-1$. Note that we have only to compare the value of stopping immediately with the value of continuing at least one further period. The Gittins index, $Z_j^*(s_j) \in \mathbf{R}$, is the value of Z_j such that the migrant is indifferent to staying for a further period or retiring from the decision problem. Hence, we get a mapping $Z_j^* : \mathbf{S}_j \longrightarrow \mathbf{R}$ where $Z_j^*(s_j)$ is given by

$$Z_j^*(s_j) = V(Z_j^*(s_j), s_j) = U_j(s_j) + \beta_j \int V_j(Z_j^*(s_j), Y_j)\, \Psi(dY_j|s_j). \tag{1.8}$$

The fixed-point theorem for contraction mappings guarantees that there always exists a unique solution for the equation (1.8). Therefore, $Z_j^*(s_j)$ is well-defined and can be

[3] $\mathcal{B}(\mathbf{S}_j)$ denote the σ-algebra of all Borel subset of $\mathbf{S}_j$

calculated in principle from the equation (1.8). Roberts and Weitzman (1980) and Ross (1983, pp. 131–41) demonstrated that the optimal migration strategy is moving to country j if

$$Z_j^*(s_j) = \max_{i \in A} \; Z_i^*(s_i). \tag{1.9}$$

In other words, the migrant will move in each period into the country which has the largest Gittins index. This is an application of a corresponding theorem by Gittins (1979) and Whittle (1980).

McCall and McCall (1987) applied the Gittins-index method for describing migration decision behavior in the MAB framework. They assumed that individuals have incomplete information about city attributes and must incur cost to learn the true values, wage rates for each city are, however, known. The study showed that more variability in the non-pecuniary return increased the attraction of a city, and thus, cities with larger in-migration also have larger out-migration since the distribution of the city attributes is symmetric about zero. Furthermore, this also implied that re-migration, which may be part of the optimal migration policy, is negatively related with information prior to migration because more variability means less prior information.

Berninghaus and Seifert-Vogt (1991) generalized the McCall's model by assuming that a potential migrant has incomplete information on both pecuniary and non-pecuniary aspects of life in foreign countries associated with more complicated subjective transition probabilities. The economic implications are similar to McCall's model.

It is intuitively clear that observed migration behavior will not only be the result of individual migration decisions but also depend on opportunities the labor market offers to individuals; furthermore, all economic implications of aggregate migration behavior should be based on an underlying equilibrium concept. Galor (1986) utilizes a two-country overlapping generations (OG) framework in a dynamic general equilibrium model, and focuses on steady-state equilibria. All agents are supposed to have complete information. He demonstrated that migration is induced by differences in the time preference rate of the households. The steady-state welfare implications of international labor mobility suggests that non-migrants in the source country are at least as well off, but the population of the host country is worse off. However. bilateral migration improves each country's welfare.

Berninghaus and Seifert-Vogt (1991) use a two-country OG framework in a temporary equilibrium model, which isolates a particular period and asks for equilibrium during this period in a dynamic model. Their model is different from that of Galor by assuming that migrants are incompletely informed about the relevant variables in the foreign country, and by ignoring all goods markets. The main results are that migration can be induced by incomplete information for potential migrants even in identical countries, and that more uncertainty over the relevant variables in the foreign country might increase the incentives for migration.

The self-selection model (or called the Roy's model), developed by Borjas (1987), tries to characterize the immigrant flow. The author argued that immigrants may be systematically different from people who decide not to migrate and thus, the immigrant flow is not randomly selected from the population of the source country. He introduced three types of selection: positive selection implies that immigrants have above-average earnings both in the source and the host countries; negative selection means that immigrants have below-average earnings both in the source and the host countries: and refugee sorting suggests that immigrants have below-average earnings in the source country and above-average earnings in the host country. The empirical analysis revealed that immigrants in the United States were relatively less skilled, or negatively selected, in the 1970s, partly due to the 1965 elimination of the national-origins quota system. The other reason may be the source countries having a relatively higher rate of return to skills than the United States, seen for example in Mexico and Puerto Rico.

The Roy's model can also be extended to characterize return migration flows. Ramos (1992) studied persons who moved between Puerto Rico and the U.S. mainland. Although Puerto Rico is a territory of the United States, migration between the island and the mainland is more similar to international migration than internal migration because of different cultures and languages. He used the data drawn from 1980 U.S. Census Public Use Samples for Puerto Rico and the U.S. to test the self-selection model. The empirical analysis indicated the following results: Puerto Ricans who migrated to the mainland were relatively unskilled or negatively selected; and re-migrant flow, including both U.S.-born and Puerto Rico-born Puerto Ricans, was characterized by positive selection. This is consistent with the facts that a lot of displaced workers, because of minimum wage policy, migrated to the U.S., and that the island has relatively unequal earning distribution compared to the mainland.

Immigration policies are centered around the potential adverse effect on the host country's labor market (immigrants compete with natives in the labor market, take jobs away from them, or bid down wages), and immigrant adaption to the host country. Theoretical predictions of the impact of immigration on wages of natives, discussed by Friedberg and Hunt (1995), depend on whether the host country is an open or closed economy.

When the host country is a closed economy, immigrant flows will lower the price of perfectly substituted factors, have an ambiguous effect on the price of imperfectly substituted factors because substitution and scale effects are opposite, and raise the price of complementary factors. For example, if immigrants are unskilled workers, then the wage in the host country for unskilled workers will fall while it is ambiguous for skilled workers.

In the Heckscher-Ohlin open economy model, if countries' factor endowments are not too different such that the factor price equalization theorem holds, then immigrant flows will not affect factor prices and will only cause the host country to export more labor in the form of goods. When countries' factor endowments are quite different such that at least one country is specialized and thus, differences in factor prices exist, the impact of immigration on wages of natives is the same as in the closed economy.

These models do not directly predict unemployment resulting from immigration. In a closed economy, if the wage is made rigid by institutional arrangements such as unions, immigrant flows will be accompanied by increasing unemployment rather than declining wages.

In the efficiency wage model (Shapiro and Stiglitz 1984), the equilibrium unemployment rate is negatively related to wages. Immigration, increasing the size of the labor force in the host country, allows firms to lower wages and increase employment. However, unemployment rates will increase because of lower wages. Hence, immigration may result in a rise in the host country's unemployment rate.

The study by Borjas (1990) indicated the following impacts of immigrant flows on the United States. First, immigrant flow has little impact on the wage rates and employment opportunities of natives, but it has a significant effect on foreign-born persons already residing in the U.S. The empirical result showed that a 10 percent increase in the number of immigrants decreases the average wage at most 0.2 percent for natives, while at least 2 percent for foreign-born persons. Second, more recent immigrant waves are relatively unskilled, characterized by less schooling, lower earning, lower labor force participation rates, and higher poverty rates. As a consequence, more recent immigrant cohorts have a relatively slower assimilation than earlier cohorts, revealed in the empirical evidence that more recent immigrant waves will remain economically disadvantaged throughout their working lives.

Borjas (1995) stresses a positive theory of immigration policy, based on a benefit-cost analysis. The major benefit to the host country from immigration is production complementarities between immigrant workers and other factors of production. The more difference between immigrants and the stock of native productive inputs, the more the host country gains from immigration. The cost in the host country consists of structural unemployment and social welfare expenditures etc. For example, less skilled workers are inclined to qualify for and participate in public assistance programs and thus, increase the fiscal cost of immigration. He indicated that in order to increase the immigration surplus, U.S. immigration policy should attract a more skilled immigrant flow.

Statement of the Problems

The search theory and the Gittins-index method are powerful approaches in describing individual migration behaviors. However, the assumption that potential migrants have infinite life causes problems in illustrating re-migration behavior. For instance, it fails to reflect the fact that re-migrants contain a significant share of persons who are retired or near retirement. They may prefer their home country's quality of life and old family friends and relatives to the host country because of cultural and language differences. The reason for emigration in their early life is attraction to the host country's employment opportunities or higher wage rates. Hence, in later life, they may find the optimal strategy is to re-migrate and retire. In making the decision to re-migrate or not, they may know that their working life has, or almost has, ended and their expected remaining life is not too long.

On the other hand, in an uncertain world, re-migration is always a part of optimal migration strategy throughout the potential migrant's life given incoming information, for example, suffering a layoff or living in an unsteady political situations in the host country etc. The infinite life assumption is clearly not suitable when potential migrants decide whether to move during their later life. Therefore, a plausible model in describing re-migration behaviors should abandon the infinite life assumption.

In empirical research, most studies 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. But a few papers have studied return migration. The objective of these limit empirical studies was to test the Roy's 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) was interested in the different return migration rate of foreign students in the U.S. Nevertheless, none of them provides an empirical analysis of how social and economic factors influence re-migration behaviors. Because of finite life assumption, we expect a quadratic life-cycle effect on our empirical re-migration analysis.

Objectives of the Study

The objectives of this work follow

1. Present a consistent sequential decision framework for individual migration decisions under incomplete information. We applied a life-cycle model where individuals migrate only when they expect to have greater utility in a new location. The model tries to explain how immigrants make decisions to re-migrate in face of incoming information and finite life.

2. Provide a econometric analysis of how social and economic factors affect return migration behavior. The hazard rate approach, the instantaneous re-migration probability given that an immigrant has lived in the host country for t years, will be used in the empirical analysis. We estimate the conditional probability that Puerto Ricans born in Puerto Rico re-migrated during the 1980s, given that they have already lived on the U.S. mainland for several years.

3. Provide econometric evidence on structural differences/changes in wage equations between and within major ethnic/racial groups in the 1980s.

Organization of the Study

The outline of this study is as follows: Chapter 2 develops the sequential decision framework for individual migration decisions under incomplete information. This model focuses on describing re-migration behaviors. A return migration econometric model will be discussed in chapter 3. We are interested in the conditional probability of re-migrating, given that the immigrant has already lived in the host country for several years. Chapter 4 describes the pattern of return migration for Puerto Ricans on the U.S. mainland, and the microdata used in this study. The empirical results of the hazard function for re-migration are presented in chapter 5. Chapter 6 discusses the fitted wage equations. The last chapter summarizes and concludes the study.

CHAPTER 2 THE SEQUENTIAL MIGRATION MODEL

This chapter will present a consistent model of how a potential migrant makes a migration decision under incomplete information. The potential migrant's objective is to make a decision on where to live in each period in order to maximize his/her expected life-time utility in a two-country world, the home country and the host country. We assume that an individual's life-time can be well approximated by five periods, and he/she retires at the beginning of the last period. Furthermore, because of incomplete information about future states of both countries, the potential migrant will gather information through time and revise his/her decision with new information at each decision time point.

First, we will introduce a general framework for sequential decision-making over a five-period planning horizon. The key factors influencing the subjective well-being of an individual are the wage rate and the quality of life. We assume that the potential migrant has full information about his/her home country, but incomplete information about the host country; he/she can become well or at least better informed only by spending some time living in the host country. Then, the optimal migration strategy for the individual at each decision time point is to choose the country that provides him/her the greater expected life-time utility. Finally, some economic implications are presented.

Description of the Model

We consider a two-country model, the home country and the foreign country. Suppose that a potential migrant has five periods of life and that he/she retires at the beginning of the last period. The potential migrant's objective is to choose a place to live in each period in order to maximize his/her expected present value of utility over the expected future life span, discounted to the present at a constant discount rate $\beta \in (0, 1)$. Therefore, the action space is given by $\mathbf{A} \equiv \{h, f\}$.

Let $s_j(t)$ represent the state of information in country j in period t for $j = h, f$ and $t = 1, 2, \ldots, 5$. The state space for the decision process is then given by

$$\mathbf{S} \equiv \mathbf{S}^h \times \mathbf{S}^f$$

where "$\times$" is the Cartesian product, and the state space of country j, $\mathbf{S}^j$, is a subset of some finite or infinite dimensional vector space depending on the economic interpretation. We assume that the state space $\mathbf{S}$ is a complete, separable metric space. Given a state of information $s(t) \in \mathbf{S}$ at decision time point t, which describes the state of information of both countries, the potential migrant will take an action in $\mathbf{A}$ according to a decision rule. Then, the state of information $s(t)$ might change from period t to period $t + 1$ according to a stochastic transition law, and the individual has to take another action in $\mathbf{A}$ based on the state of information $s(t + 1)$.

Suppose that the state of information of all countries may change from one period to another according to a Markovian decision process. In other words, the stochastic law of motion is given by a sequence of transition probabilities $\{\mathbf{P}_t(\cdot|\cdot)\}_{t=2}^{5}$, where $\mathbf{P}_t(B|s, a(s))$ denotes the probability that the state of information will be in $B \in \mathcal{B}(\mathbf{S})$ [1] during period t, provided the system was in the state of information $s \in \mathbf{S}$ during period $t - 1$ and action $a(s) \in \mathbf{A}$ has been taken during this period. Since the state transition probability will primarily be based on subjective probability assessments of the potential migrant,

[1] $\mathcal{B}(\mathbf{S})$ denote the σ-algebra of all Borel subset of $\mathbf{S}$

it seems reasonable to assume that only the state of information of country j that the individual lives in can change because he/she gains no information about another country. Furthermore, this work only deals with a homogeneous Markov process, so the time index t can be omitted, say $\mathbf{P}(\cdot|\cdot)$. We assume that $\mathbf{P} : \mathbf{S} \longrightarrow \mathcal{P}(\mathbf{S})$ is continuous w.r.t. the weak topology on $\mathcal{P}(\mathbf{S})$ where $\mathcal{P}(\mathbf{S})$ denotes the set of all probability measures on the measurable space $(\mathbf{S}, \mathcal{B}(\mathbf{S}))$.

It is obvious that both pecuniary aspects and quality of life are key factors influencing the migration behavior. The quality of life contains all factors influencing the subjective well-being of an individual except the pecuniary aspects, e.g., the cultural climate, working conditions, the education system, and acceptance by the local inhabitant etc. Suppose that the state of information of country j (without uncertainty) in period t can be interpreted by the real wage rates, denoted by w_j, and the quality of life , denoted by x_j. For simplicity, we assume that the quality of life can be represented by a real number, in the sense that the quality of life can be interpreted as compensated wage rates, and that the state of information of country j for the last period only depends on the quality of life by postulating that both countries provide no pension or the same amount of pensions after people retire. Therefore, the state of information for country j (without uncertainty) at time t can be represented by the vector

$$s_j(t) = \begin{cases} (w_j, x_j) \in \mathbf{R}_+ \times \mathbf{R} & \text{for } t \;=\; 1, 2, 3, 4, \text{ and,} \\ x_j \in \mathbf{R} & \text{for } t \;=\; 5. \end{cases}$$

Because of uncertainty about future states, the potential migrant will be modeled as the agent who gathers information and then makes the decision between the options migration and non-migration according to the expected utility maximizing approach. Let $U_j : \mathbf{S}_j \longrightarrow \mathbf{R}$ represents the migrant's individual preferences about the possible states $s_j(t)$ in country j. This specification implies that the potential migrant may feel differently in different countries. If we consider the state of information $s_j(t)$ to be a probability distribution over (w_j, x_j), a nature candidate for the per period utility function $U_j(\cdot)$ over the states in country j would be the von Neumann-Morgenstern expected utility

$$U_j(s_j(t)) = \int_{\mathbf{R}_- \times \mathbf{R}} \hat{U}(w_j, x_j)\, ds_j(t) \tag{2.1}$$

where $\hat{U}(\cdot)$ is the indirect utility function over wage rates and quality of life. Then, the bounded reward function $U(\cdot)$ of the Markovian decision process can be constructed by $U : \mathbf{S} \times \mathbf{A} \longrightarrow \mathbf{R}$ where $U(s, j) \equiv U_j(s_j)$ for $j \in \mathbf{A}$. Hence, the objective of the migrant can be restated by finding a migration strategy, a sequence of action $a(s(t)) \in \mathbf{A}$, to maximize the expected discounted total reward

$$\sum_{t=1}^{5} \beta^t U(s(t), a(s(t))) \tag{2.2}$$

where the state of information $s(t)$ evolves according to the stochastic transition law $\mathbf{P}(\cdot|s(t-1), a(s(t-1)))$ in time for $t = 2, 3, 4, 5$.

It is reasonable to assume that the potential migrant has full information about his/her home country, but is only partially informed about the pecuniary aspects and the quality of life in a potential country before migration. The potential migrant can become well or at least better informed only by spending some time living there. Let's assume s_0^f represents the potential migrant's no-information state for host country attributes at the beginning of the first period, but he/she does have his/her own personal subjective probability distribution about w_f and x_f. The migrant will be completely informed about the wage he can earn in the host country after being there one period to search for the best employment offer, and about the quality of life in the host country after one further period of living there. We assume that the migrant only cares about the wage rate during the second period of residence in the host country. After works for one peiod, he/she is then concerned the quality of life as well as the wage rate. Denote $k_j \in \mathbf{R}_-$, which is assumed to be known to the agent, as the migration cost for moving from country i to country j. Both k_f and k_h are generally different, because k_f contains the search cost and the moving cost but k_h only includes the moving cost. Note that if an individual decides to stay in his/her home country at the beginning of the first period, he/she will not migrate to the host country in his/her subsequent periods because of the learning cost and finite life assumptions. We assume further that there are constant probabilities p_h and p_f, known to the agent, of being unemployed in the home country and the host country respectively in the fourth period. To simplify the model, we assume that the immigrant is not allowed to migrate to the host country after he/she re-migrates; this is because we are mainly interested in one-cycle re-migration behavior.

According to the model structure, the state spaces for the home country and the the host country can be regarded as the subsets:

$$
\begin{aligned}
\mathbf{S}^h &\subset (\mathbf{R}_+ \times \mathbf{R}) \cup \mathbf{R}, \text{ and} \\
\mathbf{S}^f &\subset \{s_0^f\} \cup \mathbf{R}_+ \cup (\mathbf{R}_+ \times \mathbf{R}) \cup \mathbf{R}.
\end{aligned}
$$

That is, the state of information for the home country will be $(w_h, x_h) \in \mathbf{R}_+ \times \mathbf{R}$ until the last period when it reduces to $x_h \in \mathbf{R}$; moreover, if the individual decides to stay in the home country at the beginning of the first period, his/her state of information for the host country will be frozen at s_0^f. However, if he/she decides to emigrate to the host country at the beginning of the first period, the migrant's state of information for the host country will change from s_0^f to $w_f \in \mathbf{R}_+$, and then, the state of information will change from $w_f \in \mathbf{R}_+$ to $(w_f, x_f) \in \mathbf{R}_+ \times \mathbf{R}$, provided the migrant decides to stay one more period in the host country; after the agent retires, his/her relevant state of information reduces to $x_f \in \mathbf{R}$. Furthermore, the time-dependent action space for the potential migrant can be described by

$$
\begin{aligned}
\mathbf{A}_1 &= \{h, f\}, \text{ and} \\
\mathbf{A}_t &= \begin{cases} \{h\} & \text{when } s(t) \in \mathbf{S}^h \\ \{h, f\} & \text{when } s(t) \in \mathbf{S}^f \end{cases} \quad \text{for } t = 2, 3, 4, 5.
\end{aligned}
$$

Finally, the individual's per period utility function $U_j : \mathbf{S}^j \longrightarrow \mathbf{R}$ for country j with the state of information $s^j(t)$ can be defined by

$$U_f(s^f(t)) = \begin{cases} -k_f & \text{for } t = 1 \\ w_f & \text{for } t = 2 \\ w_f + x_f & \text{for } t = 3,4 \\ x_f & \text{for } t = 5 \end{cases}$$

and

$$U_h(s^h(t)) = \begin{cases} w_h - x_h & \text{for } t = 1,2,3,4 \\ x_h & \text{for } t = 5. \end{cases}$$

If the migrant re-migrates to his/her home country at the beginning of period t, then the utility function for the individual at the beginning of period t in the home country is defined by $w_h - x_h - k_h$ (or $x_h - k_h$).

The Optimal Migration Strategy

We have already introduced the basic structure of the model in the previous section. What is the optimal migration strategy for the potential migrant? The basic assumption is that the individual at each decision time point chooses the country which provides the larger expected presented value of remaining life-time utility. We will apply the general backward solution to solve this migration problem, i.e., solve the re-migration problem at the beginning of period $t+1$ first, and based on this solution solve the re-migration problem at the beginning of period t.

Because an individual can not emigrate to the host country except during the first period, we will only consider the re-migration behavior for the last four periods. Denote V_t^f and V_t^h to be the migrant's life-time utilities in the foreign country and the home country in period t, respectively. Re-migration is the optimal strategy at the beginning of period t if and only if

$$RM_t = V_t^h - V_t^f > 0 \qquad t = 2, 3, 4, 5. \tag{2.3}$$

is satisfied. Therefore, RM_t can be interpreted as the net benefit from re-migration in period t, given emigration in the first period. Similarly, emigration is an optimal choice at the beginning of the first period if and only if

$$M_1 = V_1^f - V_1^h > 0 \tag{2.4}$$

is satisfied. Hence, M_1 is the net benefit from emigration in the first period. To complete the description of the migrant's strategy, we have to find V_t^j for $j = h, f$ and $t = 1, 2, \ldots, 5$.

The individual is forced to retire at the beginning of the fifth period; hence, the key factor influencing his/her well-being is the quality of life. Therefore, we have

$$V_5^h = (x_h - k_h), \text{ and} \tag{2.5}$$

$$V_5^f = x_f. \tag{2.6}$$

Here, we try to capture the phenomena of re-migration near the retirement point.

If the immigrant re-migrates at the beginning of the fourth period, he/she must retire in the home country. However, if he/she decides to stay in the host country in the forth period, he/she has a choice to stay or to re-migrate in the next period. It can be shown, however, that re-migration is not a rational strategy at the beginning of the forth period if the agent does not suffer a layoff. The reason is that the migrant knows all relevant information at the beginning of the third period, and hence, if the host country can provide a larger remaining life-time utility, it should also have a larger remaining life-time utility in the fourth period, provided he/she is not unemployed.

Proposition 2.1 *The immigrant will stay in the host country during the forth period (the year before retirement) if he/she is not unemployed, provided he/she has lived in the host country during the third period (previous year).*

Proof: See Appendix A.

Because the immigrant faces the uncertainty of being unemployed in the fourth period, his/her presented value of remaining life-time utility for both countries at the beginning of the fourth period can be written as:

$$
\begin{aligned}
V_4^f &= \begin{cases} x_f + \beta \max\{x_h - k_h, x_f\} & \text{if unemployed} \\ (w_f + x_f) + \beta \max\{x_h - k_h, x_f\} & \text{otherwise} \end{cases} \\
&= \begin{cases} x_f + \beta (x_h - k_h) + \beta(-RM_5)^+ & \text{if unemployed} \\ (w_f + x_f) + \beta (x_h - k_h) + \beta(-RM_5)^+ & \text{otherwise} \end{cases} && (2.7) \\
V_4^h &= [(1 - p_h)w_h + x_h - k_h] + \beta x_h. && (2.8)
\end{aligned}
$$

We define $(X)^+$ to be $\max\{X, 0\}$. The $\max\{x_h - k_h, x_f\}$ in V_4^f captures the fact that the agent can stay or re-migrate in the next period if he/she decides to live in the host country in the current period. According to proposition 2.1, the agent will re-migrate if and only if he/she suffers a layoff and $V_4^h - [x_f + \beta(x_h - k_h) + \beta(-RM_5)^+] > 0$, since $V_4^h - [(w_f + x_f) + \beta(x_h - k_h) + \beta(-RM_5)^+]$ is always negative. Therefore, the net value of re-migration in the fourth period (the year before retirement) is defined by

$$RM_4 = V_4^h - \left[x_f + \beta(x_h - k_h) + \beta(-RM_5)^+\right]. \tag{2.9}$$

Note that even though the agent is unemployed, he/she need not re-migrate because the host country may still provide a higher remaining life-time utility. Thus, the model describes why some immigrants choose to stay in a host country but others decide to re-migrate when they are unemployed in the foreign country.

At the beginning of the third period, the immigrant knows the quality of life x_f; he/she has all relevant information about the host country. Then, the expected life-time utility at the beginning of the third period is described by

$$
\begin{aligned}
V_3^f &= (w_f + x_f) + \beta\Big[(1 - p_f)[(w_f + x_f) + \beta \max\{x_h - k_h, x_f\}] \\
&\quad + p_f \max\left\{x_f + \beta \max\{x_h - k_h, x_f\}, V_4^h\right\}\Big]
\end{aligned}
$$

$$
\begin{aligned}
&= (w_f + x_f) + \beta\Big[(1-p_f)\Big((w_f + x_f) + \beta(x_h - k_h) + \beta(-RM_5)^+\Big) \\
&\quad + p_f V_4^h + p_f(-RM_4)^+\Big], \quad \text{and} \qquad (2.10) \\
V_3^h &= (w_h + x_h - k_h) + \beta\left[(1-p_h)w_h + x_h\right] + \beta^2 x_h. \qquad (2.11)
\end{aligned}
$$

V_3^h represents the value for an individual who will stay in the home country for the rest of his/her life. On the other hand, V_3^f represents the value when the immigrant decides not to re-migrate. He/she continues to live in the host country if he/she does not suffer a layoff in the fourth period, and the value of remaining life-time utility is

$$(w_f + x_f) + \beta(x_h - k_h) + \beta(-RM_5)^+.$$

If he/she is unemployed in the fourth period, he/she will choose a country that provides a higher remaining life-time utility, and the value is represented by

$$V_4^h + (-RM_4)^+.$$

Theoretically, there exists a reservation quality of life, x_f^* which is determined by $V_3^h = V_3^f$, such that the immigrant will re-migrate if $x_f \leq x_f^*$.

The migrant only has the information about the wage rate w_f and his/her own personal subjective distribution function of x_f at the beginning of the second period. Denote $G(x_f)$ as the migrant's subjective cumulative distribution function of x_f given w_f. Therefore, we can define the expected remaining life-time utility at the beginning of the second period by

$$
\begin{aligned}
V_2^f &= w_f + \beta \int_{\mathbf{R}} \max\{V_3^h, V_3^f\}\, dG(x_f) \\
&= w_f + \beta V_3^h + \beta \int_{\mathbf{R}} (-RM_3)^+\, dG(x_f), \quad \text{and} \qquad (2.12) \\
V_2^h &= (w_h + x_h - k_h) + \beta(w_h + x_h) + \beta^2\left[(1-p_h)w_h + x_h\right] + \beta^3 x_h \qquad (2.13)
\end{aligned}
$$

where $\max\{V_3^h, V_3^f\}$ in V_2^f represents the options of the individual living in the host country in the current period to stay or return in the next period. Moreover, because of uncertainty about x_f, we use the mean of $\max\{V_3^h, V_3^f\}$ to represent the future utility of the host country. Consequently, we can find a reservation wage, w_f^* which is determined by $V_2^h = V_2^f$, such that the individual will re-migrate if $w_f \leq w_f^*$. Therefore, we can show that some immigrants re-migrate because the wage rate and/or the quality of life in the host country are not good enough.

Finally, the potential migrant has no information about the host country at the beginning of the first period except his/her own personal subjective distribution functions of w_f and x_f. Define $F(w_f)$ to be the conditional distribution function of w_f given s_0^f. Hence, V_1^h and V_1^f are given by

$$\begin{aligned} V_1^f &= -k_f + \beta \int_{\mathbf{R}_+} \max\{V_2^h, V_2^f\}\, dF(w_f) \\ &= -k_f + \beta V_2^h + \beta \int_{\mathbf{R}_+} (-RM_2)^+\, dF(w_f), \quad \text{and} \qquad (2.14) \\ V_1^h &= \sum_{t=1}^{3} \beta^{t-1}(w_h - x_h) + \beta^3[(1-p_h)w_h - x_h] - \beta^4 x_h. \qquad (2.15) \end{aligned}$$

Equations (2.14) and (2.15) are similar to (2.12) and (2.13) except $\max\{V_2^h, V_2^f\}$ illustrates options to a potential migrant if he/she emigrates in the first period. The mean of $\max\{V_2^h, V_2^f\}$ represents the future life time utility in the host country because of uncertainty about w_f.

Suppose the potential migrant is at the beginning of the decision process. His/her optimal migration strategy can be summarized by the migration decision tree in Figure 2.1. The H and F represent a decision to live in the home country (H) or the host country (F) in period t. He/she will re-migrate if the net values of re-migration is positive, except for the first period, where he/she will emigrate if the net value of emigration is greater than zero. Note that once the migrant has returned to the home country, he/she will stay there for the rest of his/her life.

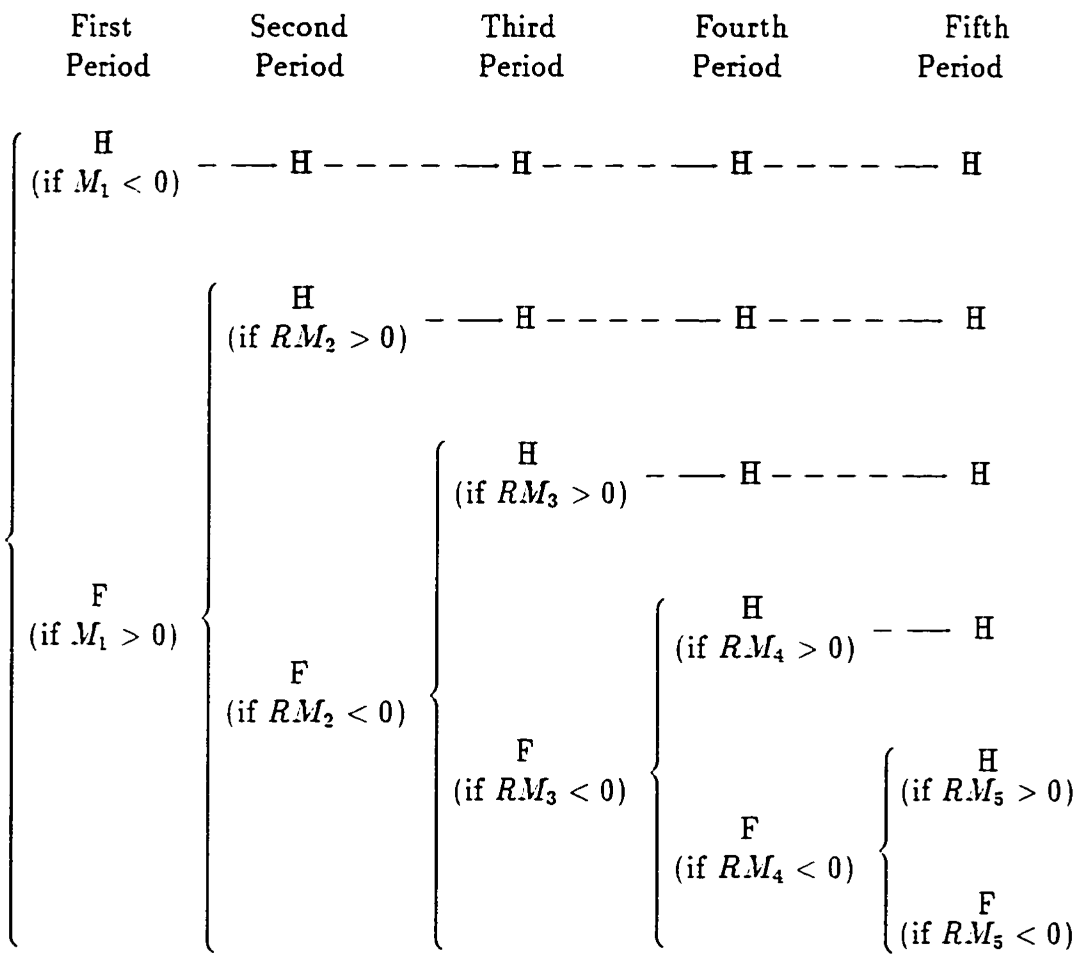


Figure 2.1 Migration Decision Tree

Some Economic Implications

Let $\mu_w = \int w_f \, dF(w_f)$ and $\mu_x = \int x_f \, dG(x_f)$ be the expected wage rate and expected quality of life, respectively. For analytical convenience, we assume that $M_1 \neq 0$ and $RM_t \neq 0$ for $t = 2, 3, 4, 5$. Table 2.1 shows the outcome of the comparative static analyses with the re-migration model.

We can see that the net present values of the emigration decision is increasing in p_h, μ_w, or μ_x, but decreasing in p_f, k_h, k_f, w_h, or x_h. These results are intuitively reasonable.

Table 2.1 The Comparative Static Effects of p_h, p_f, k_h, k_f, μ_w, μ_x, w_h, and x_h on the Net Value of Emigration and Re-migration

Net Values of Migration	Exogenous Variables[a]							
	p_h	p_f	k_h	k_f	μ_w	μ_x	w_h	x_h
M_1	+	−	−	−	+	+	−	−
RM_2	−	+	−	0	0	−	+	+
RM_3	−	+	−	0	0	0	−	+
RM_4	−	0	−	0	0	0	+	+
RM_5	0	0	−	0	0	0	0	+

[a]Detailed derivations can be found in Appendix A.

For increasing moving cost k_h or k_f, the host country should be less attractive because the expenditure on emigration goes up, but increasing in μ_w or μ_x, or decreasing in p_f will inflate the expected life-time utility of the host country residency at the beginning of the first period. Moreover, the expected life-time utility of living in the home country at the beginning of the first period will decline when p_h, w_h, or x_h increases. The table also shows that the net value of re-migration is decreasing in p_h, k_h, or μ_x, but increasing in p_f, w_h, or x_h. These results seem also to be intuitively plausible. The argument is similar to the one above. When increases μ_x or decreases p_f, the expected life-time utility from being in the host country at the beginning of period t will increase, but if p_h or k_h increases, the expected life-time utility of the home country at the beginning of period t will decrease. The same argument can also be applied to w_f^* and x_f^*.

Proposition 2.2 *1. The reservation wage w_f^* is an increasing function of x_h, w_h, or p_f, but a decreasing function of k_h, p_h, or μ_x.*

2. *The reservation quality of life x_f^* is an increasing function of x_h, w_h, or p_f, while a decreasing function of k_h, or p_h.*

Proof: See Appendix A.

Proposition 2.3 *Suppose $\hat{F}(\cdot) \succ_{\text{MPS}} F(\cdot)$* [2] *and/or $\hat{G}(\cdot) \succ_{\text{MPS}} G(\cdot)$.* [3] *Denote the corresponding net values of migration by $\hat{M}_1$, $R\hat{M}_2$, and M_1, RM_2, respectively. Then, we have*

$$\begin{aligned} \hat{M}_1 &\geq M_1, \quad \textit{and} \\ R\hat{M}_2 &\leq RM_2. \end{aligned}$$

Proof: See Appendix A.

The interpretation is that an increase in the variability in the w_f and/or x_f distribution increases the attraction of the host country. This result is based on risk neutrality and the characteristic of the sequential decision problem so that the decision-maker can revise a decision when outcomes become bad. Therefore, greater variability may improve his/her expected payoff. Berninghaus and Seifert-Vogt (1991) showed that this result is still valid if the degree of risk-aversion of migrants is small in a well defined sense. Hence, we may conclude that a country becomes more attractive if less information is known about it before migration. This might help explain immigrant flows from Europe to the U.S. during the past centuries. On the other hand, this may imply that countries with high immigration rates may have high emigration if the attractability of a country is based on a large "mean preserving spread". This result can be found in many empirical studies on internal migration (e.g. Greenwood 1975). However, no empirical study of international migration has been found that supports this economic implication. The possible explanation is that there are many institutional restrictions on

[2] *$x \succ_{\text{MPS}} y$ means that x has a larger "mean preserving spread" than y.*

[3] *In other words, $\hat{F}(\cdot)$ and/or $\hat{G}(\cdot)$ display more variability than $F(\cdot)$ and $G(\cdot)$, respectively.*

international migration or that the potential migrant may generally choose the country which is objectively superior to the home country to emigrate.

CHAPTER 3 THE ECONOMETRIC MODEL OF INDIVIDUAL RE-MIGRATION DECISIONS

The econometric model will capture the re-migration decision of immigrants, who have lived in a host country for several years, then return to their home country. The interesting question is what is the conditional probability that the immigrant will re-migrate given that he/she has already lived in the host country for n years.

We will first introduce a discrete-time version of the econometric model, which is directly derived from the theoretical model presented in chapter 2. Then, a continuous-time version of the model, called the hazard rate approach, will be presented. The hazard rate approach in which decision making is at a random interval rather than once each period is essentially the same as the discrete time version model, but frequently more appropriate in the empirical analysis. Hence, this work will use the hazard rate approach. Usually, the covariates included in the model can not completely control heterogeneity across individuals; the mixture model will be used to treat heterogeneity situations. Finally, we will discuss time dependent covariates.

The Discrete Time Version Model

Let's assume the key factors influencing re-migration behavior are the wage rates and the quality of life. Denote w_{ht} and w_{ft} to be the wage rate in the home country and the host country at time t, respectively, and z_t (which is interpreted as the net value of the quality of life) to be the quality of life that the immigrant may experience in the host country at time t. Therefore, the state space at time t is given by

$$\mathbf{S}(t) = \mathbf{S}^h(t) \times \mathbf{S}^f(t)$$

where $s^h(t) = \{w_{ht}\} \in \mathbf{S}^h(t)$ and $s^f(t) = \{w_{ft}, z_t\} \in \mathbf{S}^f(t)$ for $t = 1, 2, \cdots T$.

Assume that the immigrant is completely informed about w_{ht} and w_{ft}, $t = 1, 2, \cdots T$, in the sense that he/she can forecast them based on personal characteristics and/or local market conditions, denoted by X. Let $\{z_t\}_{t=1}^T$ be a sequence of realizations, that the immigrant may experience in the host country, of a sequence $\{Z_t\}_{t=1}^T$ of stochastically independent, identically distributed random variables Z having c.d.f. $G(\cdot\,; \boldsymbol{\alpha})$, which is supposed to be known up to a parameter vector $\boldsymbol{\alpha}$.

Denote $U(w_{ht})$ ($U(w_{ft}, z_t)$) to be the immigrant's period t utility function if he/she re-migrates (stays in the host country). Then, we have the life-time utility of living in the home country starting in period t

$$\bar{U}_t^h = \sum_{i=t}^{T} \eta^{i-t} U(w_{hi}) \tag{3.1}$$

and the life time utility of living in the host country starting in period t

$$\bar{U}_t^f = U(w_{ft}, z_t) + \eta \int V_{t+1}(w_{ft+1}, z)\, dG(z\,; \boldsymbol{\alpha}) \tag{3.2}$$

where $\eta \in (0, 1)$ is the discount factor and $V_t(S) = \max\{\bar{U}_t^h, \bar{U}_t^f\}$. Hence, there is a sequence $\{\hat{z}_t\}_{t=1}^T$ of reservation quality of life indices such that the immigrant will re-migrate in period t if and only if $z_t \leq \hat{z}_t$ where the $\hat{z}_t$ is determined by the equation

$$\sum_{i=t}^{T} \eta^{i-t} U(w_{hi}) = U(w_{ft}, \hat{z}_t) + \eta \int V_{t+1}(w_{ft+1}, z)\, dG(z\,; \boldsymbol{\alpha}). \tag{3.3}$$

According to the equation (3.3), $\hat{z}_t$ is a function of X, $\boldsymbol{\beta}$, and $\boldsymbol{\alpha}$ where $\boldsymbol{\beta}$ is a unknown coefficient vector. Let n denote exactly the number of years the immigrant plans to stay in the host country before re-migrating to his/her home country. We describe n as realizations of a random variable $N \in \mathbf{N}$. Then, the conditional probability that the immigrant has lived in the host country exactly n years and then re-migrates in the next year is given by

$$\begin{aligned}\Pr(N = n) &= G(z_{n+1}; \boldsymbol{\alpha}) \prod_{t=1}^{n} (1 - G(z_t; \boldsymbol{\alpha})) &(3.4)\\ &= h(n, X, \boldsymbol{\beta}, \boldsymbol{\alpha}). &(3.5)\end{aligned}$$

Therefore, given a random sample of size m, the likelihood function can be constructed by

$$L(\boldsymbol{\beta}, \boldsymbol{\alpha} \mid X_1, X_2, \ldots, X_m, n_1, n_2, \ldots, n_m) = \prod_{i=1}^{m} h(n_i, X_i, \boldsymbol{\beta}, \boldsymbol{\alpha}). \qquad (3.6)$$

Then, we can maximize equation (3.6) w.r.t. $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$ to find the maximum likelihood (ML) estimators of $\boldsymbol{\beta}$ and $\boldsymbol{\alpha}$.

The Hazard Rate Approach

The previous econometric model is a discrete time version which is essentially the same as a continuous time version where decision making is at a random interval rather than once each period. However, two arguments favor continuous time models over discrete-time models (Heckman and Singer 1986). First, most economic models do not have a natural decision period in which agents make decisions and take actions. Hence, it is natural and analytically convenient to model the agent who makes decision and takes action in continuous time. Second, even though there is a natural time unit, there is no reason to presume that the discrete periods are synchronized across agents or that the natural time periods correspond to the typical form (like annual or quarterly) data are available for empirical analysis. Fortunately, continuous time models are invariant to the time unit used to record the available data, and are moreover often mathematically simpler. Empirically the actual data are collected for discrete time intervals (years).

Let T be the duration of a completed spell[1] 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

$$\lambda(t) = \lim_{h \to 0} \frac{\Pr(t < T \leq t + h \mid T > t)}{h} = \frac{f(t)}{1 - F(t)} = \frac{f(t)}{S(t)} \tag{3.7}$$

where $\lambda(t)$, 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 survivor function. The integrated hazard function is defined by

$$\Lambda(t) = \int_0^t \lambda(v)\, dv. \tag{3.8}$$

Although $\Lambda(t)$ does not have a conventional interpretation, it is a useful function in practice. Note that $\lambda(t)$ is also equal to $-d \ln S(t)/dt$; by differential equation, we have

$$S(t) = \exp\left(-\int_0^t \lambda(v)\, dv\right) = \exp(-\Lambda(t)). \tag{3.9}$$

The hazard function provides a convenient definition of duration dependence. Positive (negative) duration dependence, $d\lambda(t)/dt > 0$ (< 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 fall (Lancaster 1990). Furthermore, if we know $\lambda(t)$ for all t, then we will know $S(t)$ via equation (3.11); similarly, $\lambda(t)$ could be derived from $S(t)$ because $\lambda(t) = -d \ln S(t)/dt$. Thus, $\lambda(t)$, $f(t)$, and $S(t)$ are alternative ways to describe the distribution function of a completed spell over the time axis, i.e., if we know one, we can deduce the other.

[1] 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.

We are mainly interested in the effect of the regressors X on the hazard rate and the duration. Hence, the hazard function can be written as

$$\lambda(t, X, \boldsymbol{\beta}) = \lim_{h \to 0} \frac{\Pr(t < T \leq t + h \mid T > t, X, \boldsymbol{\beta})}{h} = \frac{f(t \mid X, \boldsymbol{\beta})}{1 - F(t \mid X, \boldsymbol{\beta})} \tag{3.10}$$

where $\boldsymbol{\beta}$ is a unknown coefficient vector. Then,the corresponding survivor function can be written as

$$S(t, X, \boldsymbol{\beta}) = \Pr(T > t \mid X, \boldsymbol{\beta}) = \exp\left(-\int_0^t \lambda(v, X, \boldsymbol{\beta})\, dv\right), \tag{3.11}$$

and the conditional density function for duration is given by

$$f(t \mid X, \boldsymbol{\beta}) = \lambda(t, X, \boldsymbol{\beta})\, S(t, X, \boldsymbol{\beta}). \tag{3.12}$$

The next question is how the explanatory variables affect the distribution of the duration or the hazard rate.

Covariates can affect the distribution of durations or the hazard rate in many ways. The sign of a coefficient of a covariate indicates the direction of the effect of the variable on the conditional probability of completing a spell. Nevertheless, the coefficients of explanatory variables, generally, do not have clear interpretations as a partial derivative (analogous the linear regression model) of the hazard rate. The interpretation of the coefficients depends on the specification. In two special cases, the coefficients can be given partial derivative interpretations. The proportional hazard model, which is popular and simple to interpret, specifies that the effect of covariates is as a multiplicative effect on the hazard function. On the other hand, the accelerated lifetime model, which has been less used in economics but is also easy to interpret, is characterized by the multiplicative effect of regressors on time rather on the hazard function.

The hazard function in the proportional hazard model is specified by

$$\lambda(t, X, \boldsymbol{\beta}) = \lambda_0(t)\, \phi(X, \boldsymbol{\beta}) \tag{3.13}$$

where $\lambda_0(t)$ is a "baseline" hazard function corresponding to an agent for whom $\phi(X, \boldsymbol{\beta}) = 1$. In practice, we can measure the covariates such that $\phi(X, \boldsymbol{\beta}) = 1$ at the mean value of the covariates; then, $\lambda_0(t)$ can be interpreted as the hazard function for the mean individual in the sample. The baseline hazard is a function of time only and its functional form is determined by the distribution of the spell. The corresponding survivor function is given by

$$S(t) = \exp\{-\Lambda_0(t)\,\phi(X, \boldsymbol{\beta})\} \tag{3.14}$$

where $\Lambda_0(t) = \int_0^t \lambda_0(v)\,dv$ is the integrated baseline hazard.

A general specification of $\phi(X, \boldsymbol{\beta})$ is

$$\phi(X, \boldsymbol{\beta}) = \exp\left(-X'\boldsymbol{\beta}\right). \tag{3.15}$$

This specification does not impose restriction on $\boldsymbol{\beta}$ because $\phi(X, \boldsymbol{\beta})$ is nonnegative, and then, estimation and inference are straightforward. Note that given $\phi(X, \boldsymbol{\beta}) = \exp(-X'\boldsymbol{\beta})$, $\lambda_0(t)$ must be nonnegative and $\Lambda_0(t)$ has to diverge to obtain a proper distribution of the duration.[2] This implies that each individual must eventually complete his/her spell. Hence, we assume that every immigrant will eventually re-migrate. The partial derivative of $\ln \lambda(t, X, \boldsymbol{\beta})$ w.r.t. X is

$$\frac{\partial \ln \lambda(t, X, \boldsymbol{\beta})}{\partial X} = -\boldsymbol{\beta} \tag{3.16}$$

so $\boldsymbol{\beta}$ can be interpreted as the (negative) constant proportional effect of X on the hazard rate, or on the conditional probability of completing a spell. Furthermore, it exists as the linear model interpretation by changing variables $\varepsilon = \ln \Lambda_0(t) - X'\boldsymbol{\beta}$ (Kiefer 1988); then, the model can be rewritten as:

$$\ln \Lambda_0(t) = X'\boldsymbol{\beta} + \varepsilon \tag{3.17}$$

[2] If $F(\cdot)$ is a proper distribution function, then $\lim_{t\to\infty} F(t) = 1$, or equivalently, $\lim_{t\to\infty} S(t) = 0$. Hence, it requires $\Lambda_0(t) = \infty$ by equation (3.14).

where ε is the error term with type 1 extreme value distribution.[3] Note that the proportional hazard model postulates no direct relationship between covariates and duration itself.

The accelerated lifetime model is the class of log-linear model for T (Kalbfleisch and Prentice 1980, ch. 2). Suppose that $\ln T$ and covariates X has a linear relationship by

$$\ln T = X'\boldsymbol{\beta} + W \tag{3.18}$$

where W is a random term with p.d.f. $f_W(w)$. This specification assumes that the effect of the covariates is to accelerate (or decelerate) the time to complete a spell. Taking the exponential of equation (3.18), we have $T = \exp\left(X'\boldsymbol{\beta}\right) Y$ where $Y = \exp(W) > 0$ has a hazard function $\lambda_0(y)$ which is independent of X. Then, the p.d.f. and c.d.f. of T can be written in terms of Y as follows:

$$\begin{aligned} f_T(t, X, \boldsymbol{\beta}) &= f_Y(t\exp(-X'\boldsymbol{\beta}))\exp(-X'\boldsymbol{\beta}), \quad \text{and} && (3.19)\\ F_T(t, X, \boldsymbol{\beta}) &= F_Y(t\exp(-X'\boldsymbol{\beta})). && (3.20)\end{aligned}$$

Hence, the hazard function for T in the accelerated lifetime model can be written in terms of $\lambda_0(\cdot)$ as follows:

$$\begin{aligned} \lambda(t, X, \boldsymbol{\beta}) &= \frac{f_T(t, X, \boldsymbol{\beta})}{1 - F_T(t, X, \boldsymbol{\beta})} = \frac{f_Y(t\exp(-X'\boldsymbol{\beta}))}{1 - F_T(t\exp(-X'\boldsymbol{\beta}))}\exp(-X'\boldsymbol{\beta}) \\ &= \lambda_0(t\exp(-X'\boldsymbol{\beta}))\exp(-X'\boldsymbol{\beta}). && (3.21)\end{aligned}$$

The corresponding survivor function is

$$S(t, X, \boldsymbol{\beta}) = \exp\left(-\int_0^{t\exp(-X'\boldsymbol{\beta})} \lambda_0(v)\,dv\right). \tag{3.22}$$

With $\exp(-X'\boldsymbol{\beta}) = \phi(X, \boldsymbol{\beta})$, we have

$$\begin{aligned} \lambda(t, X, \boldsymbol{\beta}) &= \lambda_0(t\,\phi(X, \boldsymbol{\beta}))\,\phi(X, \boldsymbol{\beta}), \quad \text{and} && (3.23)\\ S(t, X, \boldsymbol{\beta}) &= \exp\left(-\int_0^{t\,\phi(X,\boldsymbol{\beta})} \lambda_0(v)\,dv\right). && (3.24)\end{aligned}$$

[3]The p.d.f. for ε is $g(\varepsilon) = \exp\{\varepsilon - \exp(\varepsilon)\}, -\infty < \varepsilon < \infty$.

Furthermore, the convenient interpretation of the coefficients can be given by

$$\frac{\partial \ln t}{\partial X} = \boldsymbol{\beta}. \tag{3.25}$$

That is, $\boldsymbol{\beta}$ is the constant proportional effect of the covariates on the time of completing a spell.

The exponential distribution is frequently used for duration data. Its corresponding hazard function and survivor function are

$$\lambda(t) = \alpha, \quad \text{and} \tag{3.26}$$

$$S(t) = \exp(-\alpha t). \tag{3.27}$$

We see that the hazard function is independent of time so that no duration dependence exists. In other words, the conditional probability of completing a spell in a time interval of specific length is the same regardless of how long the spell has been. This is referred to as the "memoryless" property. Let $\alpha = \exp(-X'\boldsymbol{\beta})$. Then, we obtain a proportional hazard model by letting $\phi(X, \boldsymbol{\beta}) = \alpha = \exp(-X'\boldsymbol{\beta})$ and $\lambda_0(t) = 1$. Because $\Lambda_0(t) = t$, we have a linear model interpretation as:

$$\ln t = X'\boldsymbol{\beta} + \varepsilon \tag{3.28}$$

Which is exactly the same as the linear model interpretation of the accelerated lifetime model. Hence, the exponential regression model is both the proportional hazard model and the accelerated lifetime model. The advantages of the exponential regression model are its simplicity to work with and its interpretation. Frequently, it is an adequate model for duration data when duration has little variation. The main disadvantage is that if the sample contains both very long and short durations, it is unlikely to be an adequate description of the data because the family of distributions obtained by varying the one parameter α is not very flexible.

In the two-parameter Weibull distribution, a scale parameter $\alpha > 0$ and a shape parameter $p > 0$ maintain simplicity but add some flexibility; this is widely used as a model for duration data. For the Weibull distribution, the hazard function and the survivor function are:

$$\lambda(t) = \frac{1}{\sigma} t^{\frac{1}{\sigma}-1} \alpha^{\frac{1}{\sigma}}, \quad \text{and} \tag{3.29}$$

$$S(t) = \exp\left\{-(\alpha t)^{\frac{1}{\sigma}}\right\} \tag{3.30}$$

where $\frac{1}{\sigma} = p$. The partial derivative of $\lambda(t)$ w.r.t. t is

$$\frac{\partial\lambda(t)}{\partial t} = \frac{1}{\sigma}\left(\frac{1}{\sigma} - 1\right) t^{\frac{1}{\sigma}-2} \alpha^{\frac{1}{\sigma}}. \tag{3.31}$$

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

The Weibull-based model can be extended to a regression model by allowing α and σ to depend on covariates. If we let α depend on X, $\alpha = \exp(-X'\boldsymbol{\beta})$, and σ is a proportional factor, then the partial effect of X on $\ln \lambda(t, X, \boldsymbol{\beta})$ is

$$\frac{\partial \ln \lambda(t, X, \boldsymbol{\beta})}{\partial X} = -\left(\frac{\boldsymbol{\beta}}{\sigma}\right). \tag{3.32}$$

Hence, $\boldsymbol{\beta}$ is the (negative) constant proportional effect of X on the hazard rate. The only difference between equation (3.32) and equation (3.16) is that the partial derivative is re-scaled by the parameter σ. In other words, the marginal effect of covariates on the hazard rate is smaller (bigger) when the hazard rate is decreasing (increasing) over time.

It is obvious that the Weibull regression model is a proportional hazard model if we note that

$$\phi(X, \boldsymbol{\beta}) = \alpha^{\frac{1}{\sigma}} = \left[\exp\left(-X'\boldsymbol{\beta}\right)\right]^{\frac{1}{\sigma}}, \quad \text{and} \tag{3.33}$$

$$\lambda_0(t) = \frac{1}{\sigma} t^{\frac{1}{\sigma}-1}. \tag{3.34}$$

Nevertheless, it is also an accelerated lifetime model because

$$\phi(X, \boldsymbol{\beta}) = \alpha = \exp\left\{-X'\boldsymbol{\beta}\right\}, \quad \text{and} \tag{3.35}$$

$$\lambda_0(t\phi(X, \boldsymbol{\beta})) = \frac{1}{\sigma}(t\alpha)^{\frac{1}{\sigma}-1} = \frac{1}{\sigma}\left(t \exp\left\{-X'\boldsymbol{\beta}\right\}\right)^{\frac{1}{\sigma}-1}. \tag{3.36}$$

Actually, the Weibull (exponential is a special case of Weibull with $\sigma = 1$) regression model is the only model belonging to both the accelerated lifetime model and the proportional hazard model(Kalbfleisch and Prentice 1980, ch. 2). We can also see this from the linear model interpretation. The corresponding linear model interpretation from the proportional hazard model is given by

$$\ln \Lambda_0(t) = \frac{1}{\sigma}\ln t = \frac{1}{\sigma}X'\boldsymbol{\beta} + \varepsilon. \tag{3.37}$$

Equivalently,

$$\ln t = X'\boldsymbol{\beta} + \sigma\varepsilon \tag{3.38}$$

which is exactly the same as the linear model interpretation from the accelerated lifetime model by letting $w = \sigma\varepsilon$.

Once the family of duration distributions under consideration has been specified, say the Weibull distribution, the parameters $\boldsymbol{\beta}$ and σ that are of interest can be estimated by maximum likelihood (ML) method. The ML estimators often have several desirable properties, like invariance, consistency, asymptotic normality, and asymptotic efficiency. If a sample of m independently completed spells $\{t_i\}_{i=1}^{m}$ and the corresponding explanatory variables $\{X_i\}_{i=1}^{m}$ are available, the log-likelihood function is constructed as usual:

$$\ln L(\boldsymbol{\beta}, \sigma) = \sum_{i=1}^{m} \ln f(t_i, X_i, \boldsymbol{\beta}, \sigma). \tag{3.39}$$

In other words, the likelihood function is the joint distribution of the sample as a function of parameters $\boldsymbol{\beta}$ and σ.

When a spell is right-censored at duration t_i, the only information available is that the spell was at least t_i. Therefore, the contribution to the likelihood function from this observation is the value of the survivor function $S(t_i, X_i, \boldsymbol{\beta}, \sigma)$, the probability that the duration is longer than t_i. Let $\delta_i = 1$ if the *ith* spell is completed and $\delta_i = 0$

otherwise. Then, the log-likelihood function for a sample of m independent spells $\{t_i\}_{i=1}^m$ with corresponding covariates $\{X_i\}_{i=1}^m$ is given by[4]

$$\ln L(\boldsymbol{\beta},\sigma) = \sum_{i=1}^m \delta_i \ln f(t_i,\, X_i,\, \boldsymbol{\beta},\, \sigma) + \sum_{i=1}^m (1-\delta_i) \ln S(t_i,\, X_i,\, \boldsymbol{\beta},\, \sigma). \tag{3.40}$$

That is, completed spells contribute to the density term $f(t, X, \boldsymbol{\beta}, \sigma)$ and censored spells contribute to the probability $S(t, X, \boldsymbol{\beta}, \sigma)$. Based on the fact that the density function is the product of the hazard function and the survivor function, the log-likelihood function can be rewritten as:

$$\ln L(\boldsymbol{\beta},\, \sigma) = \sum_{i=1}^m [\delta_i \ln \lambda(t_i,\, X_i,\, \boldsymbol{\beta},\, \sigma) + \ln S(t_i,\, X_i,\, \boldsymbol{\beta},\, \sigma)]. \tag{3.41}$$

The ML estimator $\hat{\boldsymbol{\pi}} = (\hat{\boldsymbol{\beta}}',\, \hat{\sigma})'$ can be found by maximizing the equation (3.41). The necessary condition is

$$\frac{\partial \ln L(\hat{\boldsymbol{\pi}})}{\partial \boldsymbol{\pi}} = \sum_{i=1}^m \left(\delta_i \frac{\partial \ln \lambda(t_i,\, X_i,\, \hat{\boldsymbol{\pi}})}{\partial \boldsymbol{\pi}} + \frac{\partial \ln S(t_i,\, X_i,\, \hat{\boldsymbol{\pi}})}{\partial \boldsymbol{\pi}} \right) = \mathbf{0}. \tag{3.42}$$

These equations are generally nonlinear in $\hat{\boldsymbol{\pi}}$ so iterative procedures will be used to find the maximum of the likelihood function.

The Fisher information matrix is defined by

$$\mathbf{I}(\boldsymbol{\pi}) = -\mathbf{E}\left[\frac{\partial^2 \ln L(\boldsymbol{\pi})}{\partial \boldsymbol{\pi} \partial \boldsymbol{\pi}'}\right]. \tag{3.43}$$

If certain regularity conditions are satisfied,[5] the ML estimator $\hat{\boldsymbol{\pi}} = (\hat{\boldsymbol{\beta}}',\, \hat{\sigma})'$ is consistent, asymptotically normal with mean $\boldsymbol{\pi}$ and variance $(\mathbf{I}(\boldsymbol{\pi}))^{-1}$, and asymptotically efficient.

[4]There is another way to construct likelihood function *via* the log-linear regression model (Kalbfleisch and Prentice 1980, ch. 3.6). Let $W = (\ln T - X'\boldsymbol{\beta})/\sigma$. Then, the p.d.f. for $\ln T$ can be written as $\sigma^{-1} f_W(w)$. Hence, the log likelihood function is given by

$$\ln L(\boldsymbol{\beta},\, \sigma) = \sum_{i=1}^m [\delta_i \sigma^{-1} \ln f_W(w_i) + (1-\delta_i) \ln S_W(w_i)]$$

where $S_W(w) = \int_w^\infty f_W(v)\, dv$.

[5]Roughly speaking, $\boldsymbol{\pi}$ is an interior point, $L(\boldsymbol{\pi})$ is thrice differentiable, and certain boundedness conditions on the third derivatives are satisfied.

Nevertheless, if some spells are left-censored, the likelihood function we have constructed can be applied only to the exponential distribution. It is not correct for Weibull distribution or any other distribution. 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). We can ignore the duration before the beginning of the study period when the distribution of completing a spell is exponential because of its "memoryless" property. If we solve the left-censored problem by assuming that the distribution of T is exponential, Heckman and Singer (1985) showed that the corresponding ML estimator is biased and inconsistent when this assumption is false. Some expedients will be discussed in the empirical specification.

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 heterogeneity. Economic data, however, are seldom homogeneous. They generally consist of measured, and possibly unmeasured, systematic differences between economic agents. When we include explanatory variables in the econometric model, it is not only to find the value of interested parameters, but also to control for heterogeneity. If we do not control for heterogeneity, the results will generally base the estimates of the hazard function toward negative duration dependence (Heckman and Singer 1985). Lancaster (1985) found that the effect of heterogeneity in the Weibull model causes the ML estimator $\hat{\sigma}$ to be biased upward, i.e. to bias downward the estimated hazard function duration dependence, and the ML estimator $\hat{\beta}$ 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 people, omitted regressors will cause heterogeneity.

The following mixture model was suggested by Lancaster (1979). Let's assume the effect of heterogeneity is represented by a positive random variable V, which is independently distributed with X and T, multiplying the hazard function. Let's consider the Weibull version of the model with $\alpha = \exp(-X'\boldsymbol{\beta})$. Then, the conditional hazard function and the conditional survivor function, both depending on X and V, are written as:

$$\begin{aligned}
\bar{\lambda}(t, X, \boldsymbol{\beta}, \sigma, v) &= v\,\frac{1}{\sigma}\,t^{\frac{1}{\sigma}-1}\left[\exp\left(-X'\boldsymbol{\beta}\right)\right]^{\frac{1}{\sigma}}, \quad \text{and} && (3.44)\\
\bar{S}(t, X, \boldsymbol{\beta}, \sigma, v) &= \exp\left\{-v\left[t\,\exp(-X'\boldsymbol{\beta})\right]^{\frac{1}{\sigma}}\right\}. && (3.45)
\end{aligned}$$

The three arguments discussed above can be verified by this modification (Lancaster 1990, ch. 4). Consider the first case of error in recorded durations. Suppose that the true or correctly measured duration S differs from T by a random multiplicative measurement error Z which is distributed independently of S, i.e., $T = S \times Z$. Then, we have the conditional hazard function expressed in equation (3.44) when $V = Z^{\frac{-1}{\sigma}}$. If error in recorded regressors occurs, then assume $X = X_1 + Z$ where X_1 is the true covariate vector and Z is the measurement error vector distributed independently of X_1. The corresponding conditional hazard function is the same as equation (3.44) when $V = \exp(Z'\boldsymbol{\beta})$. Finally, let $X^* = X + X_2$ where X^* is the regressors accounting fully for the true differences among individuals, X is included covariates, and X_2 is omitted covariates. We will have the desired conditional hazard function by letting $V = \exp(-X_2'\boldsymbol{\beta}_2)$.

If V is distributed as gamma with unit mean,[6] the mixture survivor function $S_m(\cdot)$,

[6]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).

depending only on X, is given by

$$\begin{aligned} S_m(t, X, \boldsymbol{\beta}, \sigma, \theta) &= \int \tilde{S}(t, X, \boldsymbol{\beta}, \sigma, v)\, f_V(v)\, dv \\ &= \left(1 + \theta \left[t \exp(-X'\boldsymbol{\beta})\right]^{\frac{1}{\sigma}}\right)^{\frac{-1}{\theta}} \qquad (3.46) \end{aligned}$$

where θ is the variance of V. The mixture hazard function $\lambda_m(\cdot)$, depending only on X, is derived by differentiating $-\ln S_m(\cdot)$ w.r.t. t giving

$$\begin{aligned} \lambda_m(t, X, \boldsymbol{\beta}, \sigma, \theta) &= [S_m(t, X, \boldsymbol{\beta}, \sigma, \theta)]^{\theta} \left(\frac{1}{\sigma}\right) t^{\frac{1}{\sigma}-1} \left[\exp(-X'\boldsymbol{\beta})\right]^{\frac{1}{\sigma}} \qquad (3.47) \\ &= [S_m(t, X, \boldsymbol{\beta}, \sigma, \theta)]^{\theta}\, \lambda(t, X, \boldsymbol{\beta}, \sigma) \qquad (3.48) \end{aligned}$$

Another way to find the mixture hazard function is given by

$$\begin{aligned} \lambda_m(t, X, \boldsymbol{\beta}, \sigma, \theta) &= \int \tilde{\lambda}(t, X, \boldsymbol{\beta}, \sigma, v)\, f_V(v \mid T \geq t)\, dv \\ &= \lambda(t, X, \boldsymbol{\beta}, \sigma) \int v\, f_V(v \mid T \geq t)\, dv \\ &= \lambda(t, X, \boldsymbol{\beta}, \sigma) \frac{1}{1 + \theta\,[t \exp(-X'\boldsymbol{\beta})]^{\frac{1}{\sigma}}} \qquad (3.49) \end{aligned}$$

where

$$f_V(v \mid T \geq t) \propto v^{\frac{1}{\theta}-1} \exp\left\{-v\left(\frac{1}{\theta} + \left[t \exp(-X'\boldsymbol{\beta})\right]^{\frac{1}{\sigma}}\right)\right\}. \qquad (3.50)$$

Hence, the mixture hazard function at t can be interpreted as the mean of the conditional hazards at t, average with respect to the distribution of V over the survivor to that date (Lancaster 1990, ch. 4). Furthermore, θ can be interpreted to capture the sensitivity of the hazard function to heterogeneity. Note that

$$\lim_{\theta \to 0} \lambda_m(t, X, \boldsymbol{\beta}, \sigma, \theta) = \left(\frac{1}{\sigma}\right) t^{\frac{1}{\sigma}-1} \left[\exp(-X'\boldsymbol{\beta})\right]^{\frac{1}{\sigma}} = \lambda(t, X, \boldsymbol{\beta}, \sigma) \qquad (3.51)$$

which is the hazard function with homogeneity. Moreover, if θ increases, then the deviation of mixture hazard function from the hazard function with homogeneity is greater. Therefore, the effect of heterogeneity is larger when θ is further away from zero. We have already identified the mixture hazard function and the mixture survivor function, both depending only on X, from the mixture model. Hence, the likelihood function can be constructed to find the ML estimators $\hat{\boldsymbol{\beta}}$, $\hat{\sigma}$, and $\hat{\theta}$.

Time Dependent Covariates

An appropriate way to specify the hazard function ought to depend on what is meaningful and interesting for the phenomenon under study. The hazard rate approach we discussed so far assumed that the covariates X are constant over spells. This is true, for example, with sex and race, but it is not appropriate for age and the predicted rate of unemployment. Actually, most economic covariates vary over time. The interpretation of the coefficient $\boldsymbol{\beta}$ for the proportional hazard model with $\phi(X(t), \boldsymbol{\beta}) = \exp(-X(t)'\boldsymbol{\beta})$ is that it measures the negative effect on the log hazard rate of a unit change in the value of covariates at time t. Similarly, $\boldsymbol{\beta}$ for the accelerated lifetime model with $\phi(X(t), \boldsymbol{\beta}) = \exp(-X(t)'\boldsymbol{\beta})$ measures the effect on the log duration of a unit change in the value of covariates at time t. Nevertheless, if we include time dependent covariates into the econometric model, it raises computational problems. Consider a hazard function conditional on the time dependent covariates $X(t)$, $\lambda(t, X(t), \boldsymbol{\beta})$. The corresponding survivor function is given by

$$S(t, X(t), \boldsymbol{\beta}) = \exp\left\{-\int_0^t \lambda(v, X(v), \boldsymbol{\beta})\, dv\right\}$$

which does not have a closed-form expression, except the special time path of the covariates, and requires numerical integration to evaluate it.

One of two expedients is often adopted to overcome this difficulty.

1. Replace covariates $X(t)$ by $\bar{X}(t) = \frac{1}{t}\int_0^t X(v)\, dv$, the average of $X(t)$ within a spell.

2. Use the beginning of spell values $X(0)$ instead of $X(t)$.

The likelihood function for these two cases is the same as derived in equation (3.41), except $\bar{X}(t)$ or $X^e(t)$ replaces X. However, the first treatment $\bar{X}(t)$ often results in undesirable effects of building spurious relationships between duration length and regressors; for example, $X(v) = c + dv$ implies $\bar{X}(t) = c + dt$, causing a linear dependence between $\bar{X}(t)$ 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 (Petersen 1986), i.e., the time dependent covariates are assumed to be constant within each period, but may change from one period to the next. Figure 3.1 plots the relationship between time dependent covariates and step functions. The solid line is the true values of an explanatory variable. The horizontal dash lines represent the step function used to approximate the time dependent explanatory variable.

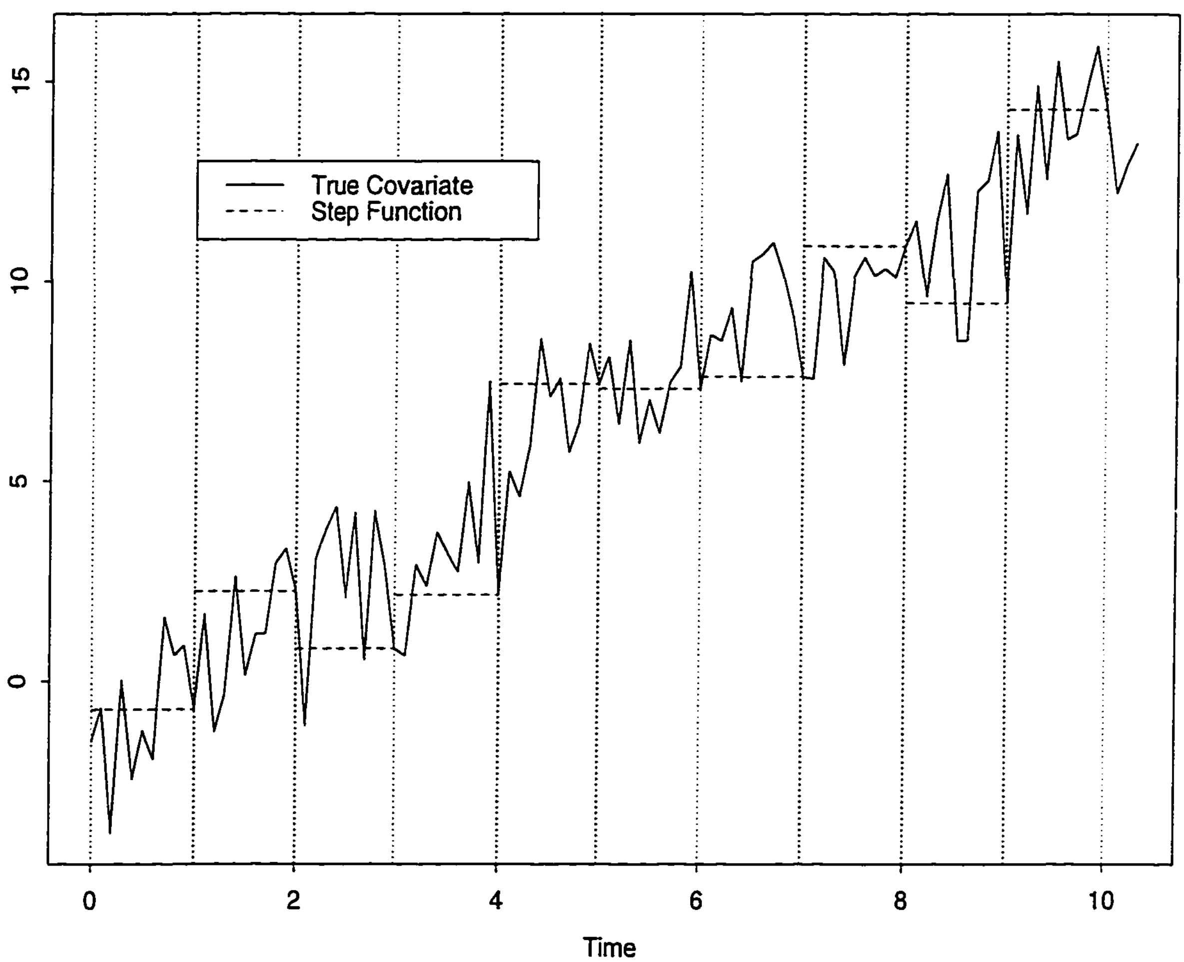


Figure 3.1 The Relationship between Time-Dependent Covariates and Step Functions

The likelihood function is formulated as follow: let's partition the time interval $[0, t]$ into k non-overlapping periods, say $0 = t_0 < t_1 < \cdots < t_k$. The covariates are assumed to stay constant within each period. Let $\lambda(t, X_j, \boldsymbol{\beta})$ be the hazard function from time t_{j-1} to t_j. Then, the corresponding survivor function can be written as:

$$S(t, X_k, \boldsymbol{\beta}) = \exp\left\{ -\sum_{j=1}^{k} \int_{t_{j-1}}^{t_j} \lambda(v, X_j, \boldsymbol{\beta})\, dv \right\}. \tag{3.52}$$

Hence, the log-likelihood function for individual i is

$$\begin{aligned} \ln L_i(\boldsymbol{\beta}) &= \delta_i \ln \lambda(t, X_k, \boldsymbol{\beta}) + \ln S(t, X_k, \boldsymbol{\beta}) \\ &= \delta_i \ln \lambda(t, X_k, \boldsymbol{\beta}) - \sum_{j=1}^{k} \int_{t_{j-1}}^{t_j} \lambda(v, X_j, \boldsymbol{\beta})\, dv \end{aligned} \tag{3.53}$$

where $\delta_i = 1$ if the *ith* spell is completed and $\delta_i = 0$ otherwise. We believe that this method is better because it contains more information and is closer to the theoretical econometric model.

CHAPTER 4 DATA DESCRIPTION AND EMPIRICAL SPECIFICATION

Puerto Rico is a territory of the United States. However, migration between the island and the mainland resembles international migration in the sense of different cultures and languages. 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. Our work focuses on the re-migration behavior of Puerto Ricans from the U.S. mainland during the 1980s.

Our empirical re-migration analysis will be based on a data set drawn from the 1990 5-percent sample of the Public Use Microdata Sample (PUMS) for Puerto Rico and the U.S. mainland. We restrict our sample to the male Puerto Rico-born householder of age 18 to 64 who resided on the U.S. mainland or resided on Puerto Rico but re-migrated from the U.S. mainland during the 1980s.

The empirical specification for the hazard rate model will be discussed. Also, the Immigration Reform and Control Act of 1986 (IRCA) allowed approximately 2.7 million persons to legally work in the U.S. The new workers legalized under IRCA were largely Spanish speaking (about 74-percent from Mexico). Hence, IRCA might have impacted Puerto Ricans residing on the U.S. mainland, for instance, increasing the competition in the mainland labor market. One possibility is that IRCA changed the wage structure for Hispanics on the U.S. mainland. To pursue this issue, data drawn from the 1980 and

1990 1-percent samples of the PUMS for the U.S. will be used.

Puerto Rico

Puerto Rico is a small island of 3,435 square miles, roughly 55 percent of the island's surface lies between sea level and altitudes up to 500 feet. San Juan is its capitol city. The island was a Spanish colony for 405 years, from 1493 to 1898. The Spanish–American War brought Puerto Rico to the U.S. in 1898. Furthermore, Puerto Rico was the first Latin American property obtained by the U.S. since the 1823 Monroe Doctrine (in which the United States declared no more European intervention was to occur in Latin America).

From 1898 to 1917, Puerto Rico changed from a Spanish to an American colony. The people of Puerto Rico wanted U.S. citizenship, self-government through statehood or territorial status, and free trade. Puerto Rico was shut out of the U.S. market by U.S. protective tariffs (Perusse 1990, p. 15). The 1917 Jones Act gave Puerto Ricans U.S. citizenship and more local self-government, but they cannot vote in U.S. elections, including the presidential election, unless they have legal residence on the mainland. The people of Puerto Rico could not elect their own governor until the 1947 Crawford–Butler Act. The 1950 Public Law 600 enabled the people of Puerto Ricans to form their own government under their own constitution. In 1952, Puerto Rico was granted a new type of government and renamed the Commonwealth of Puerto Rico, equivalent to the Associated State formula granted former British colonies (Baver 1993, p. 1). The Constitution of the Commonwealth of Puerto Rico increased the degree of local self-government in Puerto Rico, but fundamental political and economic relationship with the U.S. remained unchanged (Perusse 1990, p. 35). Moreover, Puerto Ricans are exempt from federal taxes and most revenue derived from U.S. customs and excise taxes.

During much of the twentieth century, a major issue to the island has been the nature of the relationship between the U.S. and Puerto Rico. The choices are: full integration into the U.S. as a state, a continuation of the present relationship, or complete political independence (Morris 1995, p. 7). Even through more than half of the Puerto Rican politicians have used the words "nation" or "nationality" rather than "state" in describing Puerto Rico, most Puerto Ricans are not seeking independence; for instance, no candidate who advocates independence has acquired more than 12.5 percent of the vote in the last thirty years and the independence option obtained just 4.4 percent of the vote in a 1994 ballot on the island's future (Morris 1995, p. 15).

Economic Performance

The island has a large share of fertile land. In pre-colonial day, the Indians of Puerto Rico raised corn, yucca, sweet potatoes, yams, peanuts, pineapple, guava, tamarind, papaya, and other exotic products. The key crops now are sugar, coffee, tobacco, and fruit. The island's natural beauty attracts many tourists, especially since U.S. restrictions were imposed in 1962 on travel to Cuba. Tourist facilities are centered in and about San Juan, the capitol. In the late 1960s, the income of tourism was more than 200 million dollars, even exceeding that of agriculture (Hauberg 1974, p. 6).

Puerto Rican industrialization started in 1947. Under the island's industrialization strategy called Operation Bootstrap, the public sector factories were sold to private investors and tax exemption was provided to attract private investment. In the Puerto Rican post-war industrialization process, it is possible to identify three stages (Santiago 1992).

The first stage, 1947 to the late 1950s, is labeled the phase of "resource extractive surplus labor", and is characterized by extracting the redundant agricultural labor pool from rural areas. Light manufacturing activities grew rapidly because of low wage labor, and emigration to the U.S. mainland reduced population pressure (Santiago 1992).

The second stage, roughly the 1960s, is labeled the phase of "resource expansive capital growth", and is characterized by increasing the output of capital-intensive factories. The island's government tried to reduce the cyclical sensitivity of the island economy to changing conditions on the mainland. Emigration to the U.S. continued, but at a far less rapid rate than during the 1950s. The labor force is characterized by increasing education attainment, augmented skill, increasing productivity, and raising real wages (Santiago 1992). GNP increased from $755 million in 1950 to $3.7 billion in 1968 with a growth rate of 9.3 percent per year. Per capita income was $279 in 1950 and grew to $1,129 in 1968.

The third stage, starting in the early 1970s and continuing to the present, is labeled the phase of "marginalized labor sluggish growth". Puerto Rico suffered during the recession of the mid-1970s, due to oil shocks and experienced a negative growth rate roughly of −2.4 percent in 1974 and 1975. A major factor was a reduction in private investment. The unemployment rate rose from 12.3 percent in 1974 to 20 percent in 1976, partly due to the return of Puerto Ricans affected by the recession on the mainland after fiscal 1972. The 1978 Incentive Act was designed to encourage specific types of capital-incentive and knowledge-incentive firms to reinvest in the island economy. As a consequence, the output of labor-intensive factories fell from 63 percent in 1970 to 31 percent in 1980, while the capital-incentive output rose from 35 to 67 percent.

Puerto Rico experienced another economic downturn during the 1981–1983 period, even more severe than that of the mid-1970s: the unemployment rate reached 23 percent, the labor force participation rate dropped to 41 percent, and approximately 142,000 people migrated from the island to the mainland between 1983 and 1984. The reasons seem to have been the mainland recession associated with President Reagan's New Federalism and significant cuts in federal funds to the island (Baver 1993, p. 39). The economy began an upturn with a 5.5 percent growth rate in 1987, primarily due to economic revival on the mainland. Other factors were the 30 percent decline in oil prices, a dramatic reduction in interest rate, and the 1987 act that maintained tax exemptions at a 90 percent maximum rate for the entire exemption period, reduced the tollgate tax, and the small fee paid on corporate profit repatriated to the State. Nevertheless, a high unemployment rate, low labor force participation, and declines in real investment still caused fundamental problems for Puerto Rico.

Migration

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) points out that migration to the U.S. mainland provides a "safety valve" to island population pressures including surplus agricultural labor. Castillo-Freeman and Freeman (1992) indicate 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 of outflow was 45,800 in the 1950s, 27,300 in the 1960s, and 24,300 in the 1970s; moreover, one-third of native Puerto Ricans, aged 20–64, resided on the mainland in the 1980 (Ramos 1992).

There are several reasons why there exists a large immigrant flow to the U.S. mainland. First, the birth rate is 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 population density was over 900 people per square mile on the island which poses severe socioeconomic pressure on the island's resources. Thus, overcrowding in Puerto Rico is one push factor behind migration to the U.S. mainland.

Second, the Puerto Rican government has promotions to ease the problem of overpopulation. The Migration Division of the Commonwealth Labor Department 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 air fare to the mainland (Hauberg 1974, p. 103).

Third, there is no legal restriction on 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. Empirical study (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 the short-run elasticities of employment to the minimum wage range from −0.15 (standard error 0.07) to −0.1 (0.05), suggesting that the U.S.-level minimum wage resulted in massive job losses; (2) the unemployment effects of the minimum wage are positive, but not significant with range from 0.21 (0.14) to 0.27 (0.2), 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 undoubtly help explain the high population of Puerto Ricans living in the New York 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.

Table 4.1 shows Puerto Rican population figures in Puerto Rico and on the mainland in 1980 and 1990. There were 2.7 (2) million Puerto Ricans residing on the mainland in 1990 (1980), 41.6 (46.2) percent 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

Table 4.1 Puerto Rican Population Living in Puerto Rico and in the U.S. in 1980 and 1990

	Total Population	Born in Puerto Rico	Born in The 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	

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

[a] Adopted from Table 2.1 of Ramos (1992).
[b] Estimate from 1990 1-percent sample of PUMS for the U.S.

persons, of whom 3.2 (2.9) million were born in Puerto Rico. Moreover, 398 (396) thousand persons, who are 5 years old and over, have ever resided in the U.S. mainland 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 4.2 and Figure 4.1 show statistics for Puerto Ricans, who are 5 years old and over, born in Puerto Rico, re-migrating 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 range of re-migration flows in the 1980s grew to between 18,946 and 52,671 after 1986. This is indicated by the vertical dotted line in Figure 4.1. The flows increased to above 35,000 individuals per year and reached 52,671 in 1989. Part of the reason for the increased re-migration flows after 1986 might be the large Hispanic immigration flow (legal and illegal) and the impact of the Immigration Reform and Control Act of 1986 (IRCA).

Census of Population and Housing

The U.S. census of population began in 1790, and it has been performed every 10 years in the years ending in "0", as required by the Constitution, adopted in 1787, "*Representatives and direct taxes shall be apportioned among the several States which may be included within this Union, according to their respective number...*". Since the purpose of the 1790 census was to enumerate population, the first census was very simple and only asked the name of the head of the household and the number of persons in each household with some description. Because the Nation's needs and interests shifted through years, the questions included in censuses have greatly increased both in number

Table 4.2 Puerto Rico-Born, 5 Years Old and Over, Re-Migrated to Puerto Rico from the U.S. 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	28,399
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.

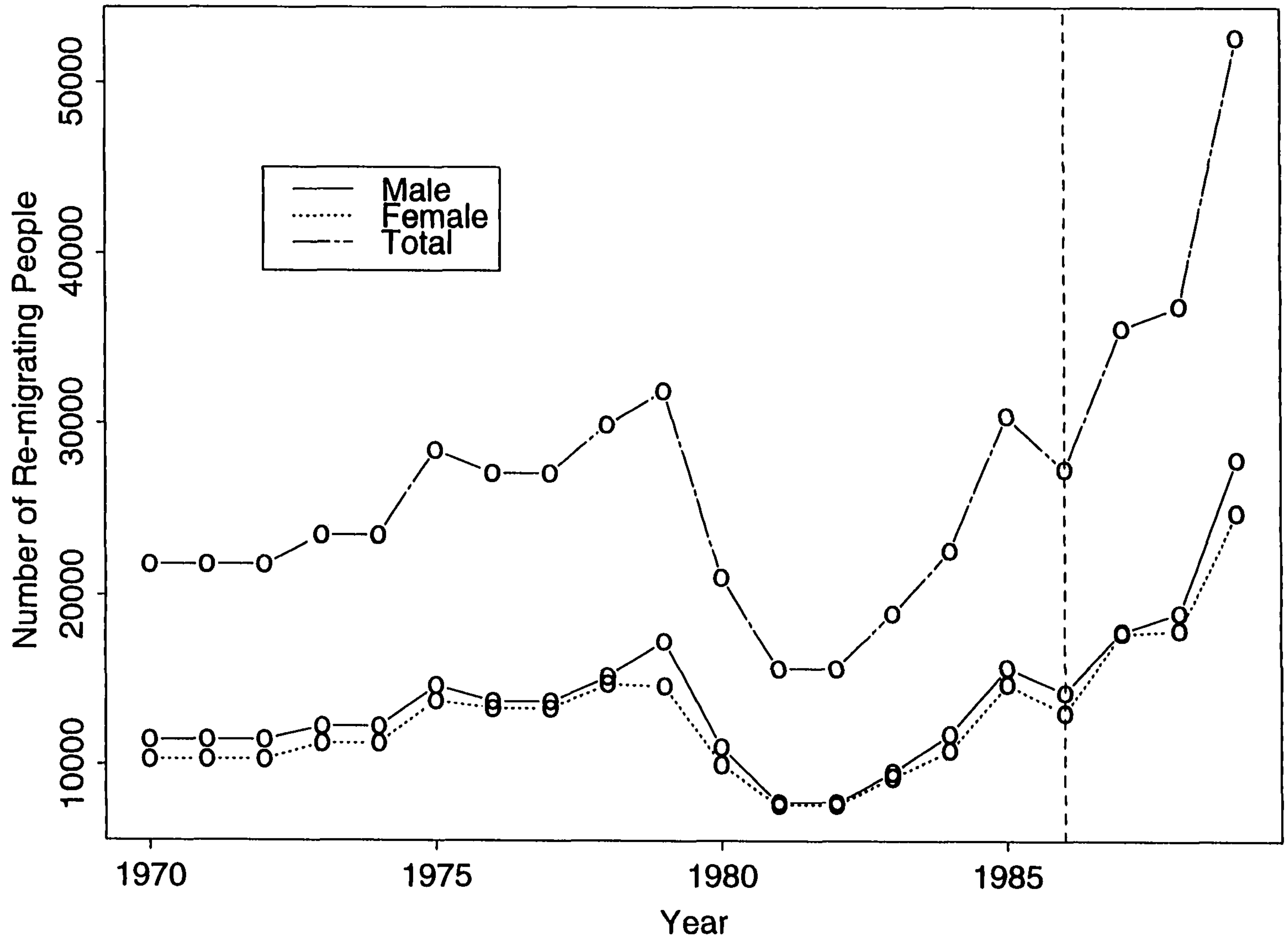


Figure 4.1 Puerto Ricans, 5 Years Old and Over Born in Puerto Rico, Re-Migrated from the U.S. Mainland during the 1970s and 1980s

and type. A full-fledeged housing census was first taken in the 1940 census.

The first nine censuses were conducted by marshals of the U.S. judicial districts. Then, marshals and their assistants were replaced by specially appointed agents to collect technical data, supervisors, and enumerators in 1880. The census organization was temporary from the 1840 through the 1900 censuses; a temporary office was established before each decennial census and disbanded after the work was finished. A permanent Bureau of the Census was established in 1902 in the Department of the Interior and later. moved to the new Department of Commerce and Labor in 1903. When the Department of Labor was split off in 1913, the Bureau of the Census continued in the Department of Commerce. Since then, the census of population has been supervised by the Bureau of the Census. The legal provision of the Census of Population and Housing was made in the Act of Congress of August 31, 1954 (Title 13, U.S. Code). The individual is required to answer the census questions and no one can see the information which may identify an individual under any circumstances except Census Bureau employees for 72 years.

The data of the Census of Population and Housing come from two primary versions of questionnaires, a short-form questionnaire and a long-form questionnaire. The short-form questionnaire, consisting of a limited number of basic population and housing questions, was asked of all persons and housing units, often referred to as the 100-percent questions. The long-form questionnaire, containing both the 100-percent questions and a number of additional questions, was asked of a sample of housing units; a sampling procedure was use to choose which housing units received the long-form questionnaire. Since the 1970 census, the Bureau of the Census not only published population and housing census data in several series of reports, either on paper and/or computer summary tapes (CD-ROM in 1990), but also issued 5-percent and 1-percent Public Use Microdata Samples (PUMS) on computer tapes (CD-ROM in 1990), which consist of records for a sample of long-form housing units, with information on the characteristics of each unit and each person in it.

PUMS

The PUMS 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 consists of information about when the individual came to the U.S. mainland; the Puerto Rico sample contains information about the individual's residence on the U.S. mainland (for more than 6 consecutive months during 1980s), and if the individual has done so, then it gives how long he/she had stayed and which activity he/she performed. With the duration and activity information on an individual residing on the U.S. mainland and the information about personal characteristics, we can 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 at age 18 to 64 during the 1980s. We further excluded those people in the armed forces, self-employed, and enrolled in school. Each record will include 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 etc.

The sample for this study consists of 12,108 observations, 2,544 from the Puerto Rico sample and 9,564 from the U.S. sample. If we could know the starting time point of the 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 chapter can be used to locate the ML estimators of the parameters. Unfortunately, some individuals do not record when a migration spell began. Our sample contains 1,896 left-censored obser-

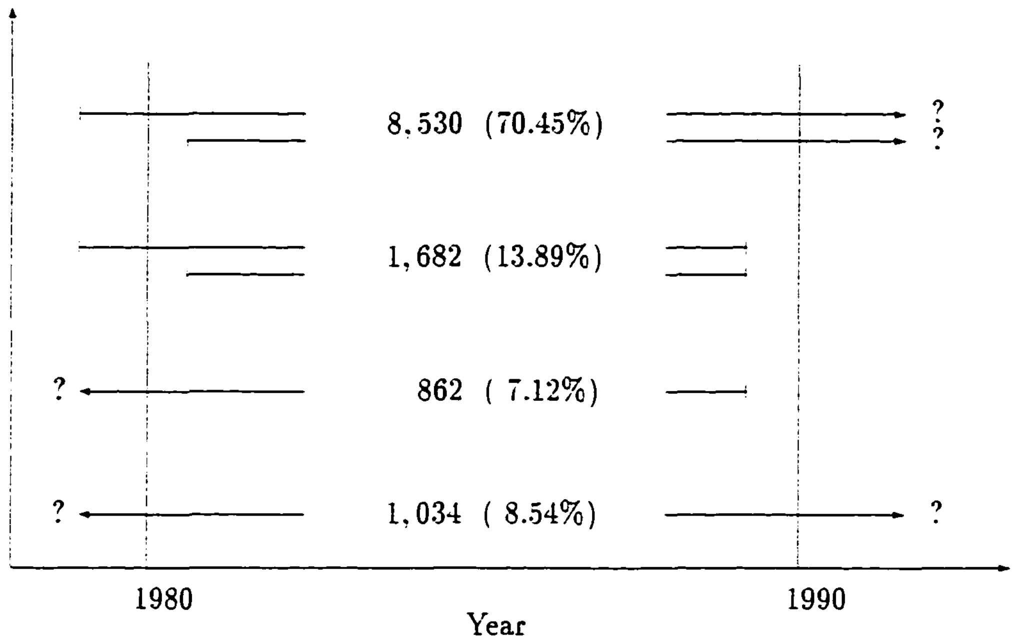


Figure 4.2 Distribution of Completed and Censored Observations from the Sample of Size 12,108 Records

vations. Figure 4.2 shows that there are 1,682 completed spells and 862 left-censored spells (from the Puerto Rico sample) and 8,530 right-censored spells and 1,034 both left- and right-censored spells (from the U.S. sample) in our data set.

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 to be $\min\{T^s, \text{ (ENDAGE} - 6 - \text{ED}), \text{ (ENDAGE} - 18)\}$ 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 individuals, 808 from the Puerto Rico sample and 375 from the U.S. sample. Figure 4.3 contains more detailed information; the sample consists of 1,736 completed

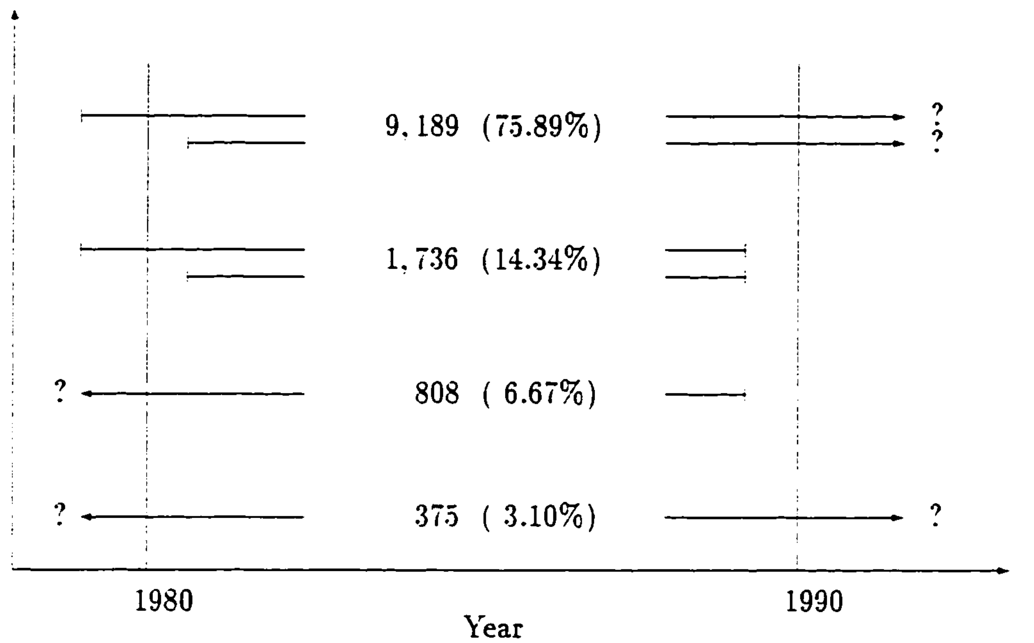


Figure 4.3 Distribution of Completed and Censored Observations from the Sample of Size 12,108 Records after Adjustment

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).

Empirical Specification for Hazard Function

The main endogenous variable is the duration of the working spell on the U.S. mainland. The theoretical model introduced in Chapter 2 indicates that the wage rate is one of the key factors expected to influence a immigrant's re-migration behavior. Since the wage rates are not truly exogenous variables, the predicted wage rates will be considered instead of the actual wage rates. However, some Puerto Ricans have lived in the U.S. mainland more than forty years. It is clear that the coefficients of the wage equations in both the U.S. mainland and Puerto Rico are not constant during this period. We do not have enough information to estimate them. Hence, the predicted wage rates are replaced by a set of variables which are used to forecast the wage rates. By applying this methodology, we can solve several problems. There is, however, no free lunch and the price we pay is that we can not get the direct impact of the wage rate on the hazard rate for re-migration.

The reduced-form hazard function for individual i with spell t_i is specified as a function of personal and local characteristics. Denote these explanatory variables as $X_i(t_i)$. Therefore, the hazard function and the survival function for the Weibull model with homogeneity, for instance, are:

$$\lambda(t_i, X_i(t_i), \boldsymbol{\beta}, \sigma) = \frac{1}{\sigma} t^{\frac{1}{\sigma}-1} \left[\exp\left(-X_i(t_i)'\boldsymbol{\beta}\right)\right]^{\frac{1}{\sigma}}, \quad \text{and} \tag{4.1}$$

$$S(t_i, X_i(t_i), \boldsymbol{\beta}, \sigma) = \exp\left\{-\left[t \exp\left(-X_i(t_i)'\boldsymbol{\beta}\right)\right]^{\frac{1}{\sigma}}\right\} \tag{4.2}$$

where σ is the time dependence parameter and the hazard rate is constant, increasing, or decreasing over time as $\sigma = 1$, $\sigma < 1$, or $\sigma > 1$, respectively.

Education represents the general human capital that not only affects the individual's earning, but also influences the likelihood of migration. Empirical studies have shown that people who have more education have higher wage rates on average. Education impacts the hazard rate through wage effects in both the U.S. mainland and Puerto Rico. The non-wage effects of education are due to acquiring and processing information (Huffman 1985). Age is used to capture individual and family life-cycle effects and the amount of potential experience. Investment in human capital often continues for much of an individual's work life so it is suitable to distinguish separately the wage effects of education and experience. On the other hand, given that life is finite, when the individual's age increases the time period remaining to capture returns from migration shrinks. Hence, age can affect the likelihood of migration directly and indirectly. We further allow for the non-linear effect of age by adding the square term.

Other personal characteristics expected to be of significance include an individual's English proficiency and disability status. English proficiency is a form of human capital and poor English proficiency is a disadvantage for working on the U.S. mainland. Disability status is expected to affect the size of earning differences associated with migration variables. Individuals with disability are less likely to migrate. We do not include family status in the model because it is not clearly exogenous to the decision to migrate (Sandefur and Tuma 1987).

Labor market conditions on the mainland and Puerto Rico may affect migration decision. 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 variables rather than 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 the U.S. mainland and Puerto Rico as stationary autoregressive (AR) processes. In other words, the stochastic process $\{y_t\}$ with constant mean μ, say the U.S. annual unemployment rate, is generated by

$$(y_t - \mu) = \phi_1(y_{t-1} - \mu) + \phi_2(y_{t-2} - \mu) + \ldots + \phi_p(y_{t-p} - \mu) + u_t \tag{4.3}$$

where u_t is white noise with zero mean and finite variance σ_u^2, and the roots of $1 - \phi_1 z - \phi_2 z^2 - \ldots - \phi_p z^p = 0$ are outside the unit circle. Then, the predicted values are constructed by the one-step-ahead predicted values. The corresponding coefficients are estimated by the maximum likelihood method, based on the annual data for 1945–93 in the U.S. mainland and for 1950–93 in Puerto Rico.[1] Furthermore, the order of these AR processes is chosen by minimizing the Akaike information criterion (AIC).

Table 4.3 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 AR of order 2, while the other two series are generated by AR of order 1. Note that we implicitly assume that the processes of the annual employment growth rates and

[1] The data in Puerto Rico are not available fro 1941–49.

Table 4.3 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	ϕ_1	ϕ_2
Job growth rates 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.

the annual unemployment rates in both the U.S. mainland and Puerto Rico are weakly stationary and time invariant.

As we mentioned before, the minimum wage policy has a strong impact in Puerto Rico. Castillo-Freeman and Freeman (1992) indicated that imposing the 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 have been disemployed. Hence, we include the real minimum wage in Puerto Rico into our model. 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. Therefore, we also add the interaction between the predicted unemployment for Puerto Rico and the real minimum wage in

Puerto Rico into the model.

Some explanatory variables vary over a migration spell. One approach to accommodate this factor is to approximate the hazard function by step-functions, i.e., the time dependent covariates are assumed to be constant within each year, but may change from one year to the next.

The empirical hazard function $\lambda(\cdot)$ is formed by specified $X_i(t_i)'\boldsymbol{\beta}$ as follow:

$$
\begin{aligned}
X_i(t_i)'\boldsymbol{\beta} &= \beta_0 + \beta_1\, AGE_i(t_i) + \beta_2\, AGESQ_i(t_i) + \beta_3\, ED_i + \beta_4\, ENG_i + \beta_5\, DISAB_i \\
&+ \beta_6\, PJRUS_i(t_i) + \beta_7\, PURUS_i(t_i) + \beta_8\, PJRPR_i(t_i) + \beta_9\, PURPR_i(t_i) \\
&+ \beta_{10}\, PRMIN_i(t_i) + \beta_{11}\, PRMINUR_i(t_i). \qquad (4.4)
\end{aligned}
$$

This specification implicitly assumes that ED, ENG, and DISAB are constant over the migration spells. The definition and sample mean of the variables in equation (4.4) are presented in Table 4.4. Sample means are derived from two groups, re-migrated during 1980s and remained in the U.S. mainland at the end of the study period (1990). The re-migrating sample consists of completed and left-censored spells, while the remaining group contains right- and both right- and left-censored spells. Note that 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 remaining group is calculated from the value in 1990.

Economic theories and previous studies provide some indication of expected signs of the parameters. We expected age to have a negative effect on the the hazard rate to re-migrate, since time periods over which to capture the discounted returns from migration shrink as individuals age and this negative effect with age decreases. Hence, the expected sign of β_1 and β_2 is positive and negative respectively. Education represents a prior investment which affects economic returns. Castillo-Freeman and Freeman (1992) demonstrated that for Puerto Rico-born men, the economic return of weekly earnings on schooling in Puerto Rico is higher than that in the U.S. mainland in 1970 and 1980; moreover, the gap was increasing from 1970 to 1980. On the other hand, people with higher

Table 4.4 Variable Definitions and Sample Means for the Hazard Function

Variable	Description	Sample Mean Remaining	Sample Mean Re-migrated
T	Duration of the spell	20.28	5.12
AGE	Age. in years.	43.29	38.56
AGESQ	Square of AGE divided by 100.	19.95	16.41
ED	Highest grade of school completed.	10.43	9.44
ENG	1 if respondent reported speaking English well or very well; 0 otherwise.	0.82	0.49
DISAB	1 if reported a health condition that limited the kind of work or amount of work he would do; 0 otherwise.	0.14	0.18
PJRUS	Predicted job growth rates for the U.S. mainland.	1.64	1.59
PURUS	Predicted unemployment rates for the U.S. mainland.	5.31	7.08
PJRPR	Predicted job growth rates for Puerto Rico.	1.53	1.78
PURPR	Predicted unemployment rates for Puerto Rico.	14.49	18.35
PRMIN	Real minimum wages on Puerto Rico in 1990 dollars	3.80	4.11
PRMINUR	Interaction of PURPR and PRMIN	55.07	75.77

education are likely to migrate because of the 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 and supports a negative β_3. Because less English proficiency is a disadvantage for work on the U.S. mainland, individuals with good English ability will reduce the hazard rate to re-migrate. Hence, we expect a positive β_4. Individuals having disability status are less likely to migrate, so the expected sign of β_5 is positive.

A higher net predicted job growth rate and/or lower net predicted unemployment rate for the U.S. mainland means the U.S. mainland is a more attractive labor market. Therefore, we will expect a positive β_6 and a negative β_7. Similarly, a higher net predicted job growth rate and/or lower net predicted unemployment rate for Puerto Rico should raise the conditional probability to re-migrate. Hence, We expect a positive effect of the predicted unemployment rate for Puerto Rico on the working spells of Puerto Rico-born men on the U.S. mainland, while a negative effect of the predicted unemployment rate for the U.S. mainland on the duration.

Time Dependent Hazard Model

The time dependent hazard model requires that all 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. If the duration of completed spells is exponentially distributed, i.e., we have a constant hazard rate or time-independent hazard model; we can treat the left-censored (or both right- and left-censored) spells to completed (or right-censored) spells by ignoring the period from the starting of the spells to when we begin to observe them. This is derived from the "memoryless" property of exponential distribution. We would have 2,544 completed spells and 9,564 right-censored spells; the estimation of the hazard function based on the covariate expressed in equation (4.4) will be straightforward. Nevertheless, if we solve the left-censored problem by assuming that the completed duration T is a family of exponential distribution, Heckman and Singer (1985) demonstrate that the corresponding ML estimators are biased and inconsistent when this assumption is false. There is no standard approach to deal with left-censored problem. Therefore. two expedients to overcome the left-censored problem will be considered in this study.

For the first expedient, we assume that those persons having left-censored spells began their work life on the U.S. mainland. In other words, individuals having left-censored spells either never worked outside the U.S. mainland if they were in the U.S. in 1990 or only worked on the U.S. mainland before re-migrating. Furthermore. the individual is assumed to be able to perform a full time job only if he is over 18 years of age and has completed his education. Hence, if the individual having a left-censored spell has completed more than 12 years of education, his working spell started in the year when he completed his education. Otherwise, it started in the year when he was 18 years of age for the others having left-censored spells. The advantage of this treatment is its simplicity. The drawback is that we might overstate some people's duration of working spells on the U.S. mainland.

The second expedient 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 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 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 why Puerto Rico-born males worked on the U.S. mainland more than 10 years[2] 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 relevant personal characteristics and local characteristics. The detail variable definition used in the probit model and the fitted equation are reported in Appendix B.

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.

[2] Left-censored spells in the Puerto Rican sample are those people who worked in the U.S. mainland more than 10 years.

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 individuals' starting-age of spells on a set of explanatory variables including the inverse Mill's ratio to generate consistent estimates of this equation's parameters. The detailed variable definitions and the fitted equation can be found in Appendix B. Then, we predict the starting-age of 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 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 lengths of spells when it is 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 durations; 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 going to fit the predicting equation. The only difference is that the binary probit model tries to explain why Puerto Rico-born men worked on the U.S. mainland more than 40 years with dependent variable equal to 1 if people had worked in the U.S. mainland more than 40 years, and 0 otherwise. This is because left-censored spells in the U.S. sample come from people who have lived in the U.S. mainland more than 40 years.

IRCA

The United States is the world leader in agriculture. This leadership is expressed in terms of share of household income spent on food in 1991: only 8.3-percent for Americans, versus 19.1-percent in Germany, 16.3-percent in France, and 53.1-percent in India (Statistical Abstract of the U.S., 1995, p. 858). Furthermore, farming is one of the few sectors having a trade surplus for decades in the U.S. economy, which exports about 25-percent of annual farm production. On the U.S. mainland, labor-intensive agriculture has a long history of using an immigrant work force, especially in California. Moreover, farmers not only admit their dependence on illegal immigrant workers but also assert that alien workers are necessary to their survival.

The goal of the Immigration Reform and Control Act of 1986 (IRCA), the latest federal policy meant to alter the farm labor market, was to get control of the problem of illegal immigration, particularly that from Mexico. IRCA created two legalization programs for illegal aliens to become temporary and then, permanent residents of the United States:

1. A general program that granted legal status for illegal aliens (legalization applicants) who have been in the U.S. since January 1, 1982.

2. The Special Agricultural Worker (SAW) program that granted legal status for illegal aliens (SAW applicants) who were employed in seasonal agricultural work for a minimum of 90 days in the 12 months ending May 1, 1986.

Table 4.5 shows that there are more than 3 million applicants filed under the IRCA legalization programs, 1,759,705 legalization applicants and 1,272,143 SAW applicants. Hispanic aliens are the major group in both programs, accounting for 88.4 percent of total legalization applicants and 88.7 percent of total SAW applicants. Especially, Mexico is the predominant country of citizenship, accounting for 69.8 percent of total legalization applicants and 81.6 percent of total SAW applicants.

A total of 2,669,968 aliens had been approved for temporary status under the IRCA provisions (1,528,200 legalization applicants and 1,037,349 SAW applicants), of whom 88.9 percent is Hispanic (74.16 percent is Mexican). Since these new workers legalized under IRCA were largely Spanish speaking, we hypothesize that the IRCA impacted Puerto Ricans residing in the U.S. mainland, for instance, increasing the competition in labor markets. From Table 4.2 and Figure 4.1, we can see that between 1970 and 1986 Puerto Rican individuals born in Puerto Rico returned to Puerto Rico from the U.S. mainland below 30,500 for each year. However, after 1986 the flow increased to above 35,000 individuals per year and reached 52,671 in 1989. These trends are consistent with our hypothesis.

We want to examine whether the IRCA altered the real wage of Hispanics on the U.S. mainland. On the other hand, most new workers legalized under IRCA provisions are less-skilled. Therefore, this policy may also affect the earnings of black men, defined on the basis of race and as non-Hispanic, because there is very high percentage of less-skilled black males. Furthermore, Hispanics as well as blacks are a disadvantaged group in the U.S. labor market, based on comparisons of labor market outcomes between Hispanic men and white non-Hispanic men. Hence, we treat white men, defined on the basis of race and as non-Hispanic, as a reference group.

Empirical Specification

Our earning analysis is based on the 1980 and 1990 1-percent sample of the PUMS for the U.S. The sample consists of male householders aged 18–64. We restrict the samples to those whose earnings were from wages and salaries only; thus, we excluded individuals in the military, self-employed, enrolled in school, working without pay in family or farm, not in the labor force, or with positive wage but zero weeks or hours worked. We

Table 4.5 Aliens Legalized under IRCA

	Legalization Applicants	SAW Applicants	Total under IRCA
Total applicants	**1,759,705**	**1,272,143**	**3,031,848**
% of Mexicans	69.8	81.6	74.8
% of Hispanic[a]	88.4	88.7	88.5
Media age at entry	23	24	- -
% of male	57.2	82.3	- -
% of married	41.2	42	- -
Approved applicants			
Total	**1,582,200**	**1,087,768**	**2,669,968**
% of Mexicans	- -	- -	74.2
% of Hispanic	- -	- -	88.9

Source: Statistical Yearbook of the immigration and naturalization Service (1993)

[a]Persons of Hispanic origin, according to the definition of CPS, are those whose origin was Mexican-American, Chicano, Mexican, Mexicano, Puerto Rican, Cuban, central or south American, or other Hispanic

further delete observations who lived in Alaska and Hawaii, not on the U.S. mainland. The 1980 (1990) sample consists of 21,144 (25,626) Hispanics, 29,416 (24,176) blacks, and 297,881 (293,768) whites. Since the sample size of white men is huge, we reduce the sample size of white men by randomly selecting one tenth in our earning analysis. Therefore, the sample size of white men in 1980 (1990) is 29,377 (29,788) in our data set. All variables used in the earning analysis are described in Table 4.6. The dependent variable is the natural logarithm of real hourly wage. and thus, the estimated coefficients can be interpreted as a rate of return. Since hourly wage rates are not reported directly, we compute from reported annual earning divided by the product of total weeks worked and usual hours worked per week in 1979 (1989) for the 1980 (1990) sample.

The explanatory variables include the individual's characteristics and local market characteristics. Education represents the general human capital affecting the individual's earning and people with higher education tend to have a higher wage profile. However, investment in human capital often continues for much of an individual's work life so we will overstate the returns on education if we fail to distinguish separately the return to subsequent on-the-job-training (OJT). We assume that OJT is proportional to years of job experience after leaving school. Hence, experience, defined by age minus education minus 6, is used to capture the return to OJT. We further allow for the non-linear effect by adding the square of experience. Because we are interested in the way education levels affect experience-wage profiles, we add interaction terms of education and experience.

English proficiency is a form of human capital and deficiency is a disadvantage for work in the U.S. Hence, it should have a positive effect on earnings. Furthermore, because English ability is likely to be lower for people who are foreign-born, our model also controls for place of birth by operationalizing with a dummy variable. Disability status is matter because it explains a systematic difference in wage rates. The dummy variables US8590, US8084, US7579, US7074, US6569, US6064, and USBE50 are used to control the effect of successive cohorts of immigrant men. Borjas (1990, p. 20) found that the recent immigrant waves have relatively less skill and slower assimilation than earlier waves.

Employment growth rates and unemployment rates are key factors in local markets. Previous studies (Tokle and Huffman 1991, and Topel 1986) suggest that economic agents respond to expected values of local variables rather than their actual values. We model the annual employment growth rates and the annual unemployment rates in state k as the stationary autoregressive (AR) processes. In other words, the stochastic process $\{y_{kt}\}$ with constant mean μ_k, say the annual employment growth rates of state k, is

Table 4.6 Definition of Variables Included in the Earning Analysis

Variables	Definition
Endogenous Variables	
W	Hourly wage
D	Wage-work participation index. 1 if the individual works for a wage; 0 otherwise.
Exogenous Variables	
AGE	Age, in years
AGESQ	Square of AGE, divided by 100
EXP	Potential post-schooling experience; AGE−ED−6.
EXPSQ	Square of EXP, divided by 100
ED	Highest grade of school completed
EDEXP	Interaction between ED and EXP
EDEXPSQ	Interaction between ED and EXPSQ
ENG	1 if respondent reported speaking English well or very well; 0 otherwise
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.
MARRIED	1 if married; 0 otherwise
KID06	Number of children at home that are 0–6 years of age.

Table 4.6 (Continued)

Variables	Definition
Exogenous Variables	
KID618	Number of children at home that are 6–18 years of age.
FORB	1 if born outside the U.S. mainland; 0 otherwise
US8590	1 if born outside U.S. mainland and immigrated to U.S. 1985–1990; 0 otherwise
US8084	1 if born outside U.S. mainland and immigrated to U.S. 1980–1984; 0 otherwise
US7579	1 if born outside U.S. mainland and immigrated to U.S. 1975–1979; 0 otherwise
US7074	1 if born outside U.S. mainland and immigrated to U.S. 1970–1974; 0 otherwise
US6569	1 if born outside U.S. mainland and immigrated to U.S. 1965–1969; 0 otherwise
US6064	1 if born outside U.S. mainland and immigrated to U.S. 1960–1964; 0 otherwise
USBE50	1 if born outside U.S. mainland and immigrated to U.S. before 1950; 0 otherwise
OTHINC	Family other (non-wage) income
PLAND	Land price
URBAN	Percentage of urban population
JAN	Normal January average temperature (degree F)
JANSQ	Square of JAN, divided by 100

Table 4.6 (Continued)

Variables	Definition
<u>Exogenous Variables</u>	
PJOBGR	Predicted state job growth rates
PURATE	Predicted sate unemployment rates
ESHOCK	Relative state labor demand shock
RURATE	Residual sate unemployment rates
NC	North central regional dummy variable
SOUTH	South regional dummy variable
WEST	West regional dummy variable
$\hat{\lambda}$	Inverse of Mill's ratio
P	The implicit GNP deflator for personal consumption expenditure with 1992 = $1.00

generated by

$$(y_{kt} - \mu_k) = \phi_1(y_{kt-1} - \mu_k) + \phi_2(y_{kt-2} - \mu_k) + \ldots + \phi_p(y_{kt-p} - \mu_k) + u_{kt} \qquad (4.5)$$

where u_{kt} is white noise with zero mean and finite variance $\sigma^2_{u_k}$, and the roots of $1 - \phi_1 z - \phi_2 z^2 - \ldots - \phi_p z^p = 0$ are outside the unit circle. Then, the predicted values are constructed by the one-step-ahead predicted values. The corresponding coefficients are estimated by the maximum likelihood method, based on the annual data for 1967–93 for each state. Furthermore, the proper order of these AR processes is chosen by minimizing the Akaike information criterion (AIC). The summary of these AR models are reported in Appendix C. Since the predicted state unemployment rate measures the forecasted state equilibrium unemployment rate, we use one-step-ahead prediction error for unemployment rates to capture the unanticipated unemployment rate.

For state employment growth rates, the one-step-ahead prediction error, e_{kt}, captures time varying local demand conditions in state k and year t. Furthermore, the predicted aggregate employment growth rates were obtained by the one-step-ahead predicted value of the U.S. aggregate employment growth rates; the corresponding one-step-ahead prediction error, e_t, explains the aggregate labor demand disturbance in year t. Hence, the relative local demand shocks are defined by RSHOCK$_{kt} = e_{kt} - e_t$, expressing the current labor demand shocks as a deviation from the aggregate labor demand shock.

The wage equation for individual i of Hispanics in state k in 1989 is specified as follows[3]:

$$\begin{aligned}
\ln\left(\frac{W_{ik}}{P_k}\right) &= \theta_0 + \theta_1\, EXP_i + \theta_2\, EXPSQ_i + \theta_3 ED_i + \theta_4\, EDEXP_i + \theta_5\, EDEXPSQ_i \\
&+ \theta_6\, ENG_i + \theta_7\, DISAB_i + \theta_8\, RACE_i + \theta_9\, FORB_i + \theta_{10}\, USBE50_i \\
&+ \theta_{11}\, US6064_i + \theta_{12}\, US6569_i + \theta_{13}\, US7074_i + \theta_{14}\, US7579_i + \theta_{15}\, US8084_i \\
&+ \theta_{16}\, US8590_i + \theta_{17}\, PJOBGR_k + \theta_{18}\, PURATE_k + \theta_{19}\, ESHOCK_k \\
&+ \theta_{20}\, RURATE_k + \epsilon_{ik} \qquad (4.6)
\end{aligned}$$

[3]RACE is not in white and blacks, and US8084, and US8590 are not in 1979

where P_k is the local price index in state k and ϵ_{ik} is a random disturbance term with zero mean, constant variance, and $E[\epsilon_{ik}\,\epsilon_{jl}] = 0$ for all $i \neq j$ and/or $k \neq l$. Since P_k is not observable, we will represent P_k by a state-specific hedonic price index. Huffman and Tokle (1995) suggested that the hedonic price index for state k consists of two kinds of goods, homogeneous and tradeable across state boundaries, and state-specific and nontradeable across state boundaries. For the first one, the implicit GNP deflator for personal consumption expenditures P and the regional dummy variables are good indicators, while for the other one, land prices, extent of urbanization, and seasons normal temperature are good indicators. Hence, the state specific-hedonic price index is:

$$\begin{aligned} \ln P_k &= \alpha_0 + \alpha_1 \ln P + \alpha_2 \ln PLAND_k + \alpha_3 URBAN_k + \alpha_4 JAN_k \\ &+ \alpha_5 JANSQ_k + \alpha_6 NC_k + \alpha_7 SOUTH_k + \alpha_8 WEST_k + \omega_k \end{aligned} \tag{4.7}$$

where ω_k is a random disturbance term with zero mean and constant variance. Then replace $\ln P_k$ in equation (4.6) by equation (4.7).

Because the data which can be used to estimate the wage equation are those with positive wage earning, there exists a non-random selection bias and the OLS estimator is biased and inconsistent. To correct this selection bias problem, we have to estimate labor force participation equation by the probit model, and then calculate the inverse of Mill's ratio which will be added to the wage equation. It can be shown that the corresponding estimator is consistent (Heckman 1979). The labor force participation equation is specified as:

$$\begin{aligned} \Pr(D_i = 1) &= F(\beta_0 + \beta_1 AGE_i + \beta_2 AGESQ_i + \beta_3 ED_i + \beta_4 ENG_i + \beta_5 EDEXP_i \\ &+ \beta_6 EDEXPSQ_i + \beta_7 MARRIED_i + \beta_8 KID06_i + \beta_9 KID618_i \\ &+ \beta_{10} DISAB_i + \beta_{11} RACE_i + \beta_{12} FORB_i + \beta_{13} USBE50_i + \beta_{14} US6064_i \\ &+ \beta_{15} US6569_i + \beta_{16} US7074_i + \beta_{17} US7579_i + \beta_{18} US8084_i + \beta_{19} US8590_i \end{aligned}$$

$$
\begin{aligned}
&+ \beta_{20}\ln(OTHINC_i/P) + \beta_{21}\ln(PLAND_k/P) + \beta_{22}\,URBAN_k \\
&+ \beta_{23}\,JAN_k + \beta_{24}\,JANSQ_k + \beta_{25}\,PJOBGR_k + \beta_{26}\,PURATE_k \\
&+ \beta_{27}\,ESHOCK_k + \beta_{28}\,RURATE_k + \beta_{29}\,NC_k + \beta_{30}\,SOUTH_k \\
&+ \beta_{31}\,WEST_k) \qquad (4.8)
\end{aligned}
$$

where $F(\cdot)$ is the normal distribution function.

Hence, the wage equation is estimated by the following form:

$$
\begin{aligned}
\ln\left(\frac{W_{ik}}{P}\right) &= \beta_0 + \beta_1\,EXP_i + \beta_2\,EXPSQ_i + \beta_3\,ED_i + \beta_4\,EDEXP_i \\
&+ \beta_5\,EDEXPSQ_i + \beta_6\,ENG_i + \beta_7\,DISAB_i + \beta_8\,RACE_i + \beta_9\,FORB_i \\
&+ \beta_{10}\,USBE50_i + \beta_{11}\,US6064_i + \beta_{12}\,US6569_i + \beta_{13}\,US7074_i \\
&+ \beta_{14}\,US7579_i + \beta_{15}\,US8084_i + \beta_{16}\,US8590_i + \beta_{17}\ln(PLAND_k/P) \\
&+ \beta_{18}\,URBAN_k + \beta_{19}\,JAN_k + \beta_{20}\,JANSQ_k + \beta_{21}\,PJOBGR_k \\
&+ \beta_{22}\,PURATE_k + \beta_{23}\,ESHOCK_k + \beta_{24}\,RURATE_k + \beta_{25}\,NC_k \\
&+ \beta_{26}\,SOUTH_k + \beta_{27}\,WEST_k + \beta_{28}\,\hat{\lambda}_i + v_{ik} \qquad (4.9)
\end{aligned}
$$

where v_{ik} is a random disturbance term with zero mean, constant variance, and $E[v_{ik}\,v_{jl}] = 0$ for all $i \neq j$ and/or $k \neq l$.

CHAPTER 5 EMPIRICAL RESULTS OF RE-MIGRATION

This chapter presents the empirical results from the hazard function of re-migration. The results are divided into four parts. Part one reports the estimation results for exponential regression models. We assume that the completed duration has an exponential distribution. Part two presents Weibull regression models. The distribution of duration is specified as a member of the family of Weibull distributions. The third part discusses the marginal effect of explanatory variables on the hazard rate of return migration. The last part presents some policy implications.

Exponential Regression Models

The exponential regression model treats the left-censored (both right- and left-censored) spells to completed (right-censored) spells by ignoring the period from the starting of the spells to when we begin to observe them. This is based on the memoryless property of the exponential distribution. Maximum likelihood estimates are reported in Table 5.1, containing two equations. The first equation gives the estimates without considering heterogeneity effect. The second incorporates gamma heterogeneity represented by the parameter θ, referred to as a mixture model. The dependent variable is the natural logarithm of working spells in years on the U.S. mainland. The signs of the parameter estimates are as expected across equations. Furthermore, all estimated coefficients are statistically significantly different from zero at the 0.1 percent significant level, except for disability status.

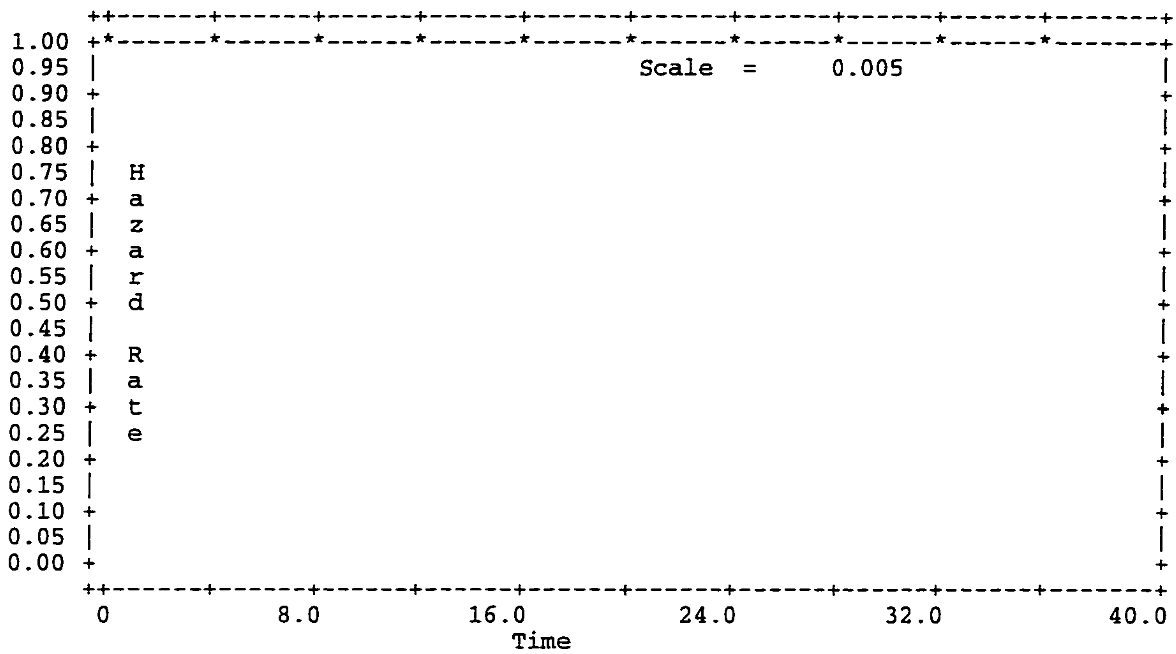
Table 5.1 The Hazard Function for Re-Migration of Puerto Rico-Born Male Householders during the 1980s from the 1990 Census Data in Exponential Regression Models

Explanatory Variables[a]	Without Heterogeneity		With Heterogeneity	
Intercept	34.150	(10.41)[b]	68.996	(11.24)
AGE	0.213	(18.67)	0.085	(3.74)
AGESQ	−0.259	(−18.72)	−0.173	(−5.68)
ED	−0.034	(−5.85)	−0.038	(−3.54)
ENG	1.580	(35.54)	2.091	(23.13)
DISAB	0.014	(0.25)	0.176	(1.55)
PJRUS	1.919	(20.39)	3.266	(21.25)
PURUS	−0.786	(−8.92)	−1.393	(−7.94)
PJRPR	−2.539	(−15.06)	−4.529	(−18.98)
PURPR	−1.697	(−8.50)	−3.572	(−9.76)
PRMIN	−7.318	(−8.54)	−14.205	(−9.17)
PRMINUR	0.436	(7.83)	0.902	(8.84)
θ			2.249	(23.12)
Total Spells	12,108		12,108	

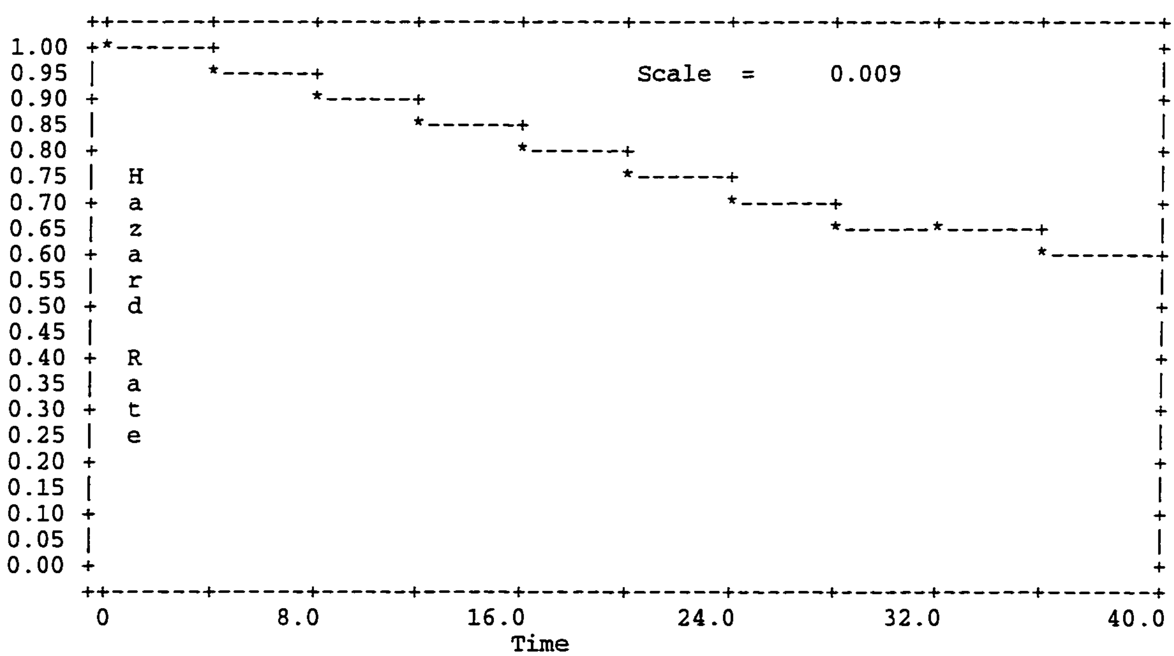
[a]Dependent variable is the natural logarithm of working spells in years on the U.S. mainland.
[b]t-values are in parentheses.

Figure 5.1 shows the relation between the duration on the U.S. mainland and the predicted hazard rate of re-migration, evaluated at the sample means of explanatory variables. Note that the vertical axis should be re-scaled by a scale factor which is placed in the middle of the figure. Part (a) presents the exponential regression model. The predicted hazard rate is constant at 0.005. Part (b) is the exponential regression model with gamma heterogeneity. The predicted hazard rate is negatively related to the duration on the U.S. mainland with a range from 0.009 to 0.0054. This result shows that the hazard rate of re-migration is smaller when Puerto-born males stay on the mainland longer.

We find that the coefficients of age and age square are statistically significantly different from zero at 0.1 percent significant level. The life-cycle effect shows that the spell duration on the U.S. mainland increases when people get older, but at a diminishing rate. In other words, there is a inverted U-shaped form of the age-spell duration profile. This quadratic life-cycle effect is consistent for the model with a finite life assumption. When an individual becomes older, time periods over which to capture the discounted returns from return migration shrink. The inclusion of heterogeneity reduces the effect. The spell duration on the U.S. mainland achieves a maximum at 41.1 years of age in the model without heterogeneity, and at 24.8 years of age in the model with gamma heterogeneity. The finding is consistent with previous studies(Tienda and Wilson 1992, Pissarides and Wadsworth 1989, and Feridhanusetyawan 1994). Furthermore, there exists an equivalent interpretation on the hazard rate of re-migration. The effect of age on the hazard rate of re-migration displays a U-shaped form. The hazard rate decreases when people get older, but at a diminishing rate. Furthermore, the hazard rate of re-migration achieves a minimum at 41.1 (24.6) years of age in the model without (with gamma) heterogeneity. This suggests that at a younger age, the conditional probability of return migration decreases as age increases, while at an older age, the probability increases as age increases.



(a) Without Heterogeneity



(b) With Heterogeneity

Figure 5.1 The Predicted Hazard Rate for Re-Migration of Puerto Rico-Born Male Householders during the 1980s from the 1990 Census Data in Exponential Regression Models

Education has a strong negative effect on the working spell of Puerto Rica-born males on the U.S. mainland. The coefficients of education in both equations are statistically significantly different from zero at the 0.1 percent significant level. The model with gamma heterogeneity increases the size of the effect slightly. The result shows that increasing one year of schooling reduces by 3.4 to 3.8 percent the duration of the working spells for Puerto Rico-born males on the mainland. This supports the prediction that individuals with higher education are likely to migrate because of their efficiency in acquiring and processing information. Previous migration studies, for instance by Mincer (1978), Schwartz (1976), Tienda and Wilson (1992), and Feridhanusetyawan (1994), have presented similar finding. The empirical result is also consistent with the studies by Ramos (1992), and Castillo-Freeman and Freeman (1992). Ramos found that Puerto Ricans who return to Puerto Rico tend to be more skilled than those who remain on the U.S. mainland. Castillo-Freeman and Freeman showed that for Puerto Rico-born men, the economic return of weekly earnings on schooling in Puerto Rico is higher than that on the U.S. mainland.

We also find that English proficiency plays a significant role in return migration decisions for Puerto Rico-born men. The sign of the coefficient is strongly positive and statistically significantly different from zero at the 0.1 percent significant level. The inclusion of gamma heterogeneity strengthens the size of the effect. The spell duration on the mainland for Puerto Rico-born males with English proficiency is 1.6 to 2.1 times larger than for those with poor English proficiency. This empirical result supports the hypothesis that English proficiency is a form of human capital for working on the U.S. mainland.

Disability status is the only variable which is not statistically significantly different from zero at the 5 percent significant level in the exponential regression models. The sign of the coefficient, however, is negative and as expected in both equations. The mixture model tightens up the effect. The results indicate that disability raises the duration on the U.S. mainland by about 1.4 to 17.6 percent. This supports the prediction that individuals with disability are less likely to move.

The empirical results yield strong evidence for the effect of labor market conditions on re-migration decisions. All estimated coefficients of labor market conditions not only have expected signs, but also are statistically significantly different from zero at the 0.1 percent significant level. The inclusion of gamma heterogeneity tends to increase the size of the effect.

The 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 job growth rate for the U.S. mainland augments the spell duration on the mainland by 19 to 33 percent, while a similar predicted job growth rate for Puerto Rico reduces the spell duration by 25 to 45 percent. This supports the prediction that people are attracted by higher job growth rates.

We find that the predicted unemployment rate for the U.S. mainland is negatively related to the duration of working spells of Puerto Rico-born men on the U.S. mainland. The empirical results exhibit that a 0.1 percentage point increase in the predicted unemployment rate for the U.S. mainland reduces about 8 to 14 percent the duration of working spells. This also advocates the hypothesis that a lower unemployment rate increases the attraction of a country.

The estimated coefficient of the interaction between the predicted unemployment rate for Puerto Rico and the real minimum wage in Puerto Rico is positive and statistically significantly different from zero at the 0.1 percent significant level. This implies that the unemployment rate for Puerto Rico is positively related to the real minimum wage in Puerto Rico, which is consistent with previous studies (Castillo-Freeman and Freeman 1992).

The proportional effect of the predicted unemployment for Puerto Rico on the duration of working spells of Puerto Rico-born males on the U.S. mainland is β_9 –

$\beta_{11}\,PRMIN$. The effect is positive when the real minimum wage in Puerto Rico is greater than 3.89 (3.96) for the model without heterogeneity (with gamma heterogeneity), and negative otherwise. When we evaluate this effect at the mean of PRMIN from 1980 to 1990, 4.163, the results show that a 0.1 percentage point increase in the predicted unemployment rate for Puerto Rico amplifies the spell duration by 1.2 to 1.8 percent.

Similarly, the proportional effect of the real minimum wage in Puerto Rico on the spell duration on the U.S. mainland is $\beta_{10} + \beta_{11}\,PURPR$. The effect is positive when the predicted unemployment rate for Puerto Rico is greater than 16.78 (15.75) for the model without (with gamma) heterogeneity, and negative otherwise. When we evaluate this effect at the mean of PURPR from 1980 to 1990, 18.636, the results show that a 10 cents increase in the real minimum wage in Puerto Rico augments the spell duration by 8.1 to 26.1 percent.

We find significant heterogeneity effects in the model. The estimated parameter θ, measuring the sensitivity of the hazard rate and the survivor function to heterogeneity, is 2.249 and statistically significantly different from zero at the 0.1 percent significant level. The inclusion of θ in the model changes the magnitude of the estimated coefficients of the explanatory variables, but not the signs. We may conclude that the model with gamma heterogeneity may perform better than the model without considering heterogeneity.

Weibull Regression Models

The advantages in using the exponential regression model are its simplicity to work with and its interpretation. Frequently, it is an adequate model for duration data when duration has little variation. The main disadvantage is that if the sample contains both very long and short durations, this model is unlikely to be an adequate description of the data because the family of distributions obtained by varying the one parameter α is not very flexible.

The Weibull regression model maintains simplicity but adds some flexibility. The use of a Weibull distribution for the duration of completed spells adds only one parameter σ, measuring the effect of the duration of residence spells on the hazard rate of re-migration. The coefficient vector $\boldsymbol{\beta}$ of explanatory variables is still interpreted as the constant proportional effect of the covariates on the time of completing a spell, which is the same as the exponential regression model. However, the marginal (proportional) effect of explanatory variables on the hazard rate of re-migration is $-\boldsymbol{\beta} / \sigma$ in Weibull regression models instead of $-\boldsymbol{\beta}$ in exponential regression models.

In Weibull regression models, we have the left-censored problem. Based on discussion in previous chapter, we applied two expedients or research strategies. As a result, the empirical results from the two treatments are very similar. Furthermore, the signs of the estimated coefficients from both expedients are the same as those from the exponential regression model.

First Expedient

The first expedient or research strategy is to assume that those Puerto Rico-born males having left-censored spells began their working life on the U.S. mainland. Moreover, the migrant is assumed to be able to perform a full time job only if he is over 18 years of age and has completed his education. Therefore, if the Puerto Rican having a left-censored spell has completed more than 12 years of education, the duration of the spell is the age at the end of the observed spell minus 6 minus his highest grade of school completed. Otherwise, it is the age at the end of the observed spell minus 18 for those having left-censored spells.

The fitted equations are presented in Table 5.2, consisting of models without and with gamma heterogeneity. The estimated parameter θ, measuring the effect of heterogeneity, is statistically significantly different from zero at the 0.1 percent significant level. This may indicate that there does exist heterogeneity in hazard and survivor functions across

Table 5.2 The Hazard Function for Re-Migration of Puerto Rico-Born Male Householders during the 1980s from the 1990 Census Data in Weibull Regression Models, Based on the First Expedient

Explanatory Variables[a]	Without Heterogeneity		With Heterogeneity	
Intercept	59.024	(11.43)[b]	58.932	(12.71)
AGE	0.144	(7.65)	0.099	(4.67)
AGESQ	−0.223	(−9.65)	−0.177	(−6.31)
ED	−0.034	(−3.53)	−0.046	(−4.67)
ENG	2.200	(24.18)	2.217	(24.11)
DISAB	0.118	(1.30)	0.179	(1.73)
PJRUS	2.503	(15.34)	2.816	(19.24)
PURUS	−0.614	(−10.38)	−0.871	(−13.93)
PJRPR	−3.763	(−12.60)	−4.027	(−16.36)
PURPR	−2.966	(−9.56)	−3.101	(−11.08)
PRMIN	−12.455	(−9.36)	−12.143	(−10.41)
PRMINUR	0.719	(8.72)	0.754	(10.29)
σ	1.621	(32.33)	0.946	(23.98)
θ			1.962	(13.23)
Total Spells	12,108		12,108	

[a] Dependent variable is the natural logarithm of working spells in years on the U.S. mainland.
[b] t-values are in parentheses.

individuals. 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.

For the model without gamma heterogeneity, the estimated parameter σ, 1.621, is statistically significantly different from unit at the 0.1 significant level, with t-statistics of 12.39. This implies that there is a decreasing effect of length of duration on the U.S. mainland on the hazard rate of re-migration. Other things being equal, the longer a Puerto Rican resides on the mainland, the lower is the hazard rate of re-migration. Part (a) of Figure 5.2 presents the relationship between the spell duration on the U.S. mainland and the predicted hazard rate of re-migration, evaluated at the sample means of explanatory variables. Note that the vertical axis should be re-scaled by a scale factor which is placed in the middle of the figure. For the first 4.6 years of spell duration, the predicted hazard rate is at the peak, 0.31. After that, it suddenly drops close to zero.

The estimated parameter σ in the mixture model is 0.942 and cannot reject the null hypothesis $H_o : \sigma = 1$ at the 5 percent significant level, with t-statistics of 1.39. The pattern of the predicted hazard rate of re-migration, evaluated at the sample means of covariates, associated with the spell duration on the mainland appears as an inverted U-shaped curve in part (b) of Figure 5.2. The predicted hazard rate increases and then decreases after it climbs to a peak of 0.008 over the spell duration. The plot shows that the maximum hazard rate of re-migration occurs when Puerto Rico-born men worked on the U.S. mainland for 4.6–13.8 years.

Compared to the results in the exponential regression models, the signs of the coefficients are the same. All estimated coefficients are statistically significantly different from zero at the 0.1 percent significant level, except disability status. But the coefficient of disability is significantly different from zero at the 10 percent level in the model with gamma heterogeneity. Hence, we conclude that the results do not overstep the prediction of economic theories.

```
     ++------+------+------+------+------+------+------+------+------+-------+
1.00 +*------+                                                               +
0.95 |       |                   Scale  =     0.310                          |
0.90 +       |                                                               +
0.85 |       |                                                               |
0.80 +       |                                                               +
0.75 |  H    |                                                               |
0.70 +  a    |                                                               +
0.65 |  z    |                                                               |
0.60 +  a    |                                                               +
0.55 |  r    |                                                               |
0.50 +  d    |                                                               +
0.45 |       |                                                               |
0.40 +  R    |                                                               +
0.35 |  a    |                                                               |
0.30 +  t    |                                                               +
0.25 |  e    |                                                               |
0.20 +       |                                                               +
0.15 |       |                                                               |
0.10 +       |                                                               +
0.05 |       |                                                               |
0.00 +       *------*------*------*------*------*------*------*------*-------+
     ++------+------+------+------+------+------+------+------+------+-------+
      0            9.2          18.4          27.6          36.8          46.0
                                     Time
```

(a) Without Heterogeneity

```
     ++------+------+------+------+------+------+------+------+------+-------+
1.00 +       *------*------+                                                 +
0.95 |       |             *------+      Scale  =     0.008                  |
0.90 +       |                    *------*------+                            +
0.85 |       |                                  *------+                     |
0.80 +       |                                         *------+              +
0.75 |  H    |                                                *------+       |
0.70 +  a    |                                                       *-------+
0.65 |  z    |                                                               |
0.60 +*-a----+                                                               +
0.55 |  r                                                                    |
0.50 +  d                                                                    +
0.45 |                                                                       |
0.40 +  R                                                                    +
0.35 |  a                                                                    |
0.30 +  t                                                                    +
0.25 |  e                                                                    |
0.20 +                                                                       +
0.15 |                                                                       |
0.10 +                                                                       +
0.05 |                                                                       |
0.00 +                                                                       +
     ++------+------+------+------+------+------+------+------+------+-------+
      0            9.2          18.4          27.6          36.8          46.0
                                     Time
```

(b) With Heterogeneity

Figure 5.2 The Predicted Hazard Rate for Re-Migration of Puerto Rico-Born Male Householders during the 1980s from the 1990 Census Data in Weibull Regression Models, Based on the First Expedient

The inverted U-shaped form of the age-spell duration profile indicates that the working spell on the U.S. mainland attains the maximum at 32.3 (28.0) years of age in the model without (with) gamma heterogeneity. An additional year of schooling is associated with declines of 3.4 to 4.6 percent of the spell duration on the U.S. mainland for Puerto Ricans. English proficiency tends to increase the spell duration about 2.2 times. The estimated spell duration on the U.S. mainland for an individual having a disability is about 11.8 to 17.9 percent longer.

The mixture model tends to increase the size of the effect of the variables of local market conditions slightly, except the effect of the the real minimum wage in Puerto Rico. A 0.1 percentage point increase in the predicted job growth rate for the U.S. mainland raises the spell duration on the mainland by 25 to 28 percent, while the predicted job growth rate for Puerto Rico decreases the spell duration about 38 to 40 percent. When the predicted unemployment rate increases 0.1 percentage point, the estimated spell duration on the U.S. mainland for Puerto Rican males shrinks about 6 to 9 percent. If we calculated the effect of the predicted unemployment rate at the sample mean of the real minimum wage in Puerto Rico for 1980–1990, 4.163, an increase in one percentage point augments the length of working spells of Puerto Ricans on the U.S. mainland by about 2.7 to 3.8 percent. Finally, a 10 cents increase in the real minimum wage in Puerto Rico prolongs the spell duration on the mainland by 9.4 to 19.1 percent, evaluated at the sample mean of the predicted unemployment rate for Puerto Rico for 1980–1990, 18.636.

Second Expedient

For the second expedient or search strategy, we predict the duration of working spells on the U.S. mainland for those Puerto Rico-born men having left-censored spells. By using the sample of completed spells (or right-censored spells), we regress the natural logarithm of individuals' starting-age of spells on a set of explanatory variables. When the inverse of Mill's ratio is added into the multiple regression model to correct the non-random selection bias, the ordinary least square (OLS) estimates are consistent. Then we predict the starting-age of the working spells on the U.S. mainland for those Puerto Ricans having left-censored spells. Finally, the length of migration spells for individuals with left-censored spells can be derived as the ending-age of a spell minus the predicted starting-age. Note that we only use these predicted lengths of spells when they 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. When this condition is violated, the predicted lengths of spells are based on the first expedient.

Table 5.3 reports the empirical results, containing the model with and without gamma heterogeneity. The pattern of the predicted hazard rate of re-migration, evaluated at the sample means of explanatory variables, associated with the spell duration on the U.S. mainland is presented in Figure 5.3. Because the empirical results between two expedients are very similar, the discussion is omitted. Furthermore, we may conclude that the two research strategies provide no differences for Weibull regression models especially in our data.

Hazard Rate of Re-Migration

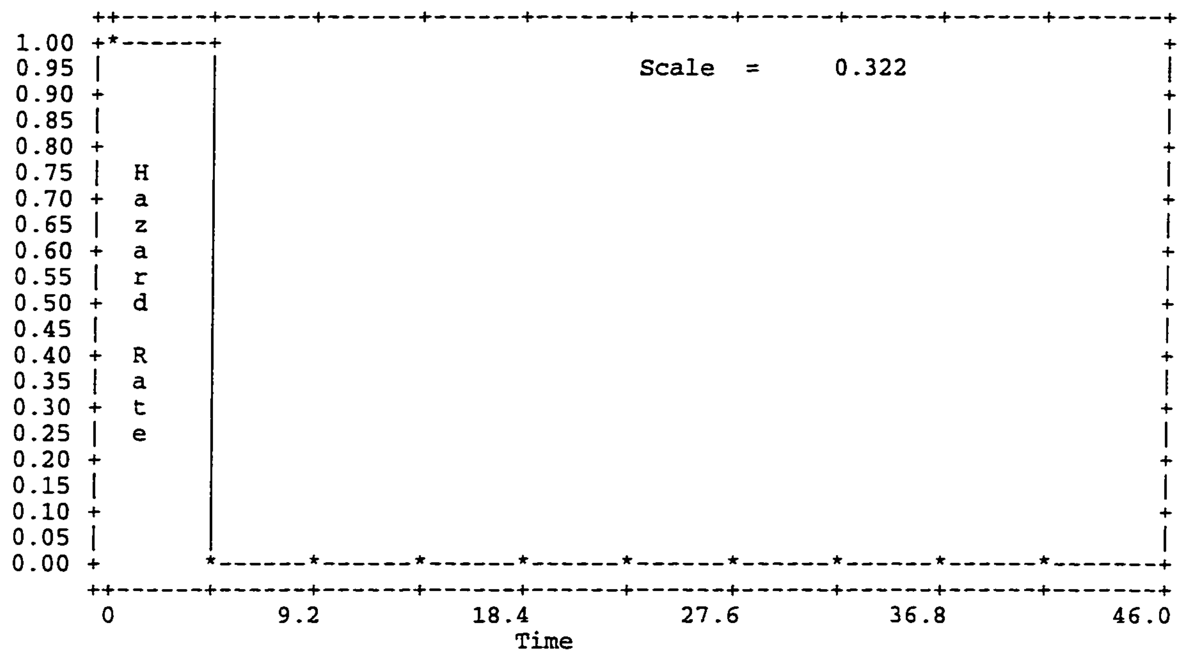
As we discussed in Chapter 3, the coefficient vector $\boldsymbol{\beta}$ of explanatory variables has an equivalent interpretation on the hazard rate. The marginal proportional effect of explanatory variables on the hazard rate of re-migration is $-\boldsymbol{\beta}$ for the exponential regression model, while it is $-\boldsymbol{\beta} / \sigma$ for the Weibull regression model.

Table 5.4 reports the marginal proportional effects of explanatory variables on the hazard rate of re-migration for Puerto Rico-born males. The first two columns present the results from the exponential regression models, taken from Table 5.1. The third and fourth columns exhibit the results from the Weibull regression models using the

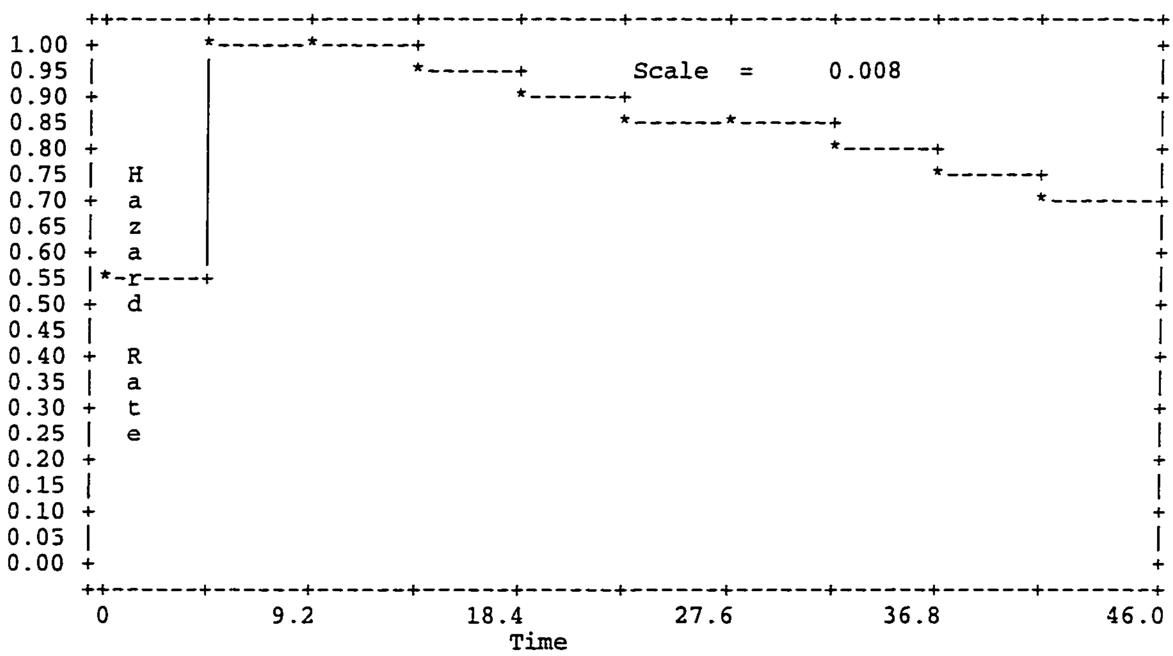
Table 5.3 The Hazard Function for Re-Migration of Puerto Rico-Born Male Householders during the 1980s from the 1990 Census Data in Weibull Regression Models, Based on the Second Expedient

Explanatory Variables[a]	Without Heterogeneity		With Heterogeneity	
Intercept	59.178	(11.46)[b]	50.701	(12.71)
AGE	0.140	(7.45)	0.098	(4.67)
AGESQ	−0.219	(−9.44)	−0.177	(−6.30)
ED	−0.034	(−3.52)	−0.045	(−4.59)
ENG	2.202	(24.17)	2.215	(24.08)
DISAB	0.118	(1.29)	0.183	(1.76)
PJRUS	2.498	(15.35)	2.803	(19.24)
PURUS	−0.614	(−10.37)	−0.863	(−13.83)
PJRPR	−3.747	(−12.63)	−4.004	(−16.37)
PURPR	−2.973	(−9.59)	−3.092	(−11.09)
PRMIN	−12.489	(−9.39)	−12.103	(−10.42)
PRMINUR	0.721	(8.75)	0.752	(10.30)
σ	1.623	(32.25)	0.942	(23.95)
θ			1.984	(13.25)
Total Spells	12,108		12,108	

[a]Dependent variable is the natural logarithm of working spells in years on the U.S. mainland.
[b]t-values are in parentheses.



(a) Without Heterogeneity



(b) With Heterogeneity

Figure 5.3 The Predicted Hazard Rate for Re-Migration of Puerto Rico-Born Male Householders during the 1980s from the 1990 Census Data in Weibull Regression Models, Based on the Second Expedient

Table 5.4 Marginal Proportional Effects of Explanatory Variables on the Hazard Rate of Re-Migration for Puerto Rico-Born Male Householders

Explanatory Variables	Exponential Regression Models		Weibull Regression Models			
			First Expedient		Second Expedient	
	Without Heterogeneity	With Heterogeneity	Without Heterogeneity	With Heterogeneity	Without Heterogeneity	With Heterogeneity
AGE[a]	0.006	0.061	0.028	0.054	0.028	0.055
ED	0.034	0.038	0.021	0.049	0.021	0.048
ENG	−1.580	−2.091	−1.357	−2.344	−1.357	−2.351
DISAB	−0.014	−0.176	−0.073	−0.189	−0.073	−0.194
PJRUS	−1.919	−3.266	−1.544	−2.977	−1.539	−2.976
PURUS	0.786	1.393	0.379	0.921	0.378	0.916
PJRPR	2.539	4.528	2.322	4.257	2.309	4.251
PURPR[b]	−0.118	−0.183	−0.017	−0.041	−0.018	−0.041
PRMIN[c]	−0.807	−2.605	−0.583	−2.018	−0.584	−2.029

[a]Evaluated at the sample mean of AGE, 42.295.
[b]Evaluated at the mean of PRMIN from 19080 to 1990, 4.163.
[c]Evaluated at the mean of PURPR from 1980 to 1990, 18.636.

first expedient, taken from Table 5.2. The last two columns show results from the Weibull regression models using the second expedient, taken from Table 5.3. The signs of the marginal proportional effects are consistent across models. Moreover, the models including gamma heterogeneity have stronger effects.

The marginal effect of a one year increase in age on the hazard rate of re-migration ranges from 0.6 to 6.1 percent, evaluated at the sample mean of age, 42.3. Increasing one year of schooling augments the hazard rate by 2.1 to 4.9 percent.

People having English proficiency and disability status are less likely to re-migrate. The marginal proportional effect of English proficiency on the hazard rate ranges from −1.36 to −2.35. Disability status decreases the hazard rate of re-migration by 1.4 to 19.4 percent.

Predicted unemployment rates for the U.S. mainland or predicted job growth rates for Puerto Rico increase the hazard rate of re-migration. The marginal effect of a 0.1 percentage point increase in predicted unemployment rates for the U.S. mainland on the hazard rate ranges from 3.8 to 13.9 percent. A 0.1 percentage point increase in predicted job growth rates for Puerto Rico amplifies the hazard rate by 23.1 to 45.3 percent.

Predicted job growth rates for the U.S. mainland, predicted unemployment rates for Puerto Rico, or real minimum wages for Puerto Rico have negative effects on the hazard rate of return migration. Increasing 0.1 percentage point in predicted job growth rates for the U.S. mainland reduces the hazard rate about 15.4 to 32.7 percent. The marginal effect of a 0.1 percentage point increase in predicted unemployment rates on the hazard rate ranges from −0.2 to −1.8 percent, evaluated at the sample mean of real minimum wages for Puerto Rico for 1980–1990, 4.16. We can see that the effects of predicted job growth rates are stronger than those of predicted unemployment rates.

Finally, a 10 cents increase in real minimum wages for Puerto Rico diminishes the hazard rate by 5.8 to 26.1 percent, evaluated at the sample mean of predicted unemployment rates for Puerto Rico for 1980–1990, 18.64. This may reflect the fact that the minimum wage policy for Puerto Rico resulted in massive job losses. A significant portion of these unemployed persons migrated to the U.S. mainland. Hence, from their viewpoint, a higher real minimum wage might mean a lower chance of finding a job in Puerto Rico.

Previous graphs of estimated hazard rates versus duration may be misleading. Figure 5.2 (b), for example, shows an inverted U-sharped form, while the corresponding coefficients of age and age square revealed a U-sharped form on the hazard rate of re-migration. This is because in Figure 5.2 (b) age is held constant at its sample mean. Hence, we have chosen to simulate the life-cycle effect on the predicted hazard rate of re-migration, based on estimated coefficients from Table 5.2 with gamma heterogeneity. We held everything constant at sample means, except spell duration and 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. Figure 5.4 shows that the hazard rate for re-migration is higher when people get older. Furthermore, the marginal effect of age on the predicted hazard rate increases rapidly when the immigrant is approaching retirement.

Policy Implications

The estimated coefficients of the hazard functions can give us some guidance in how some government policies will affect the return migration decision of Puerto Rico-born men on the U.S. mainland. In particular, we will focus on the official English amendments and the minimum wage policy.

Although the majority of U.S. citizens speak English as their first or second language[1]

[1] According to the 1990 census, 94 percent of U.S. residents above 5 years of age reported that they spoke English "very well", and 13.8 percent of U.S. residents above 5 years of age spoke other than English in households.

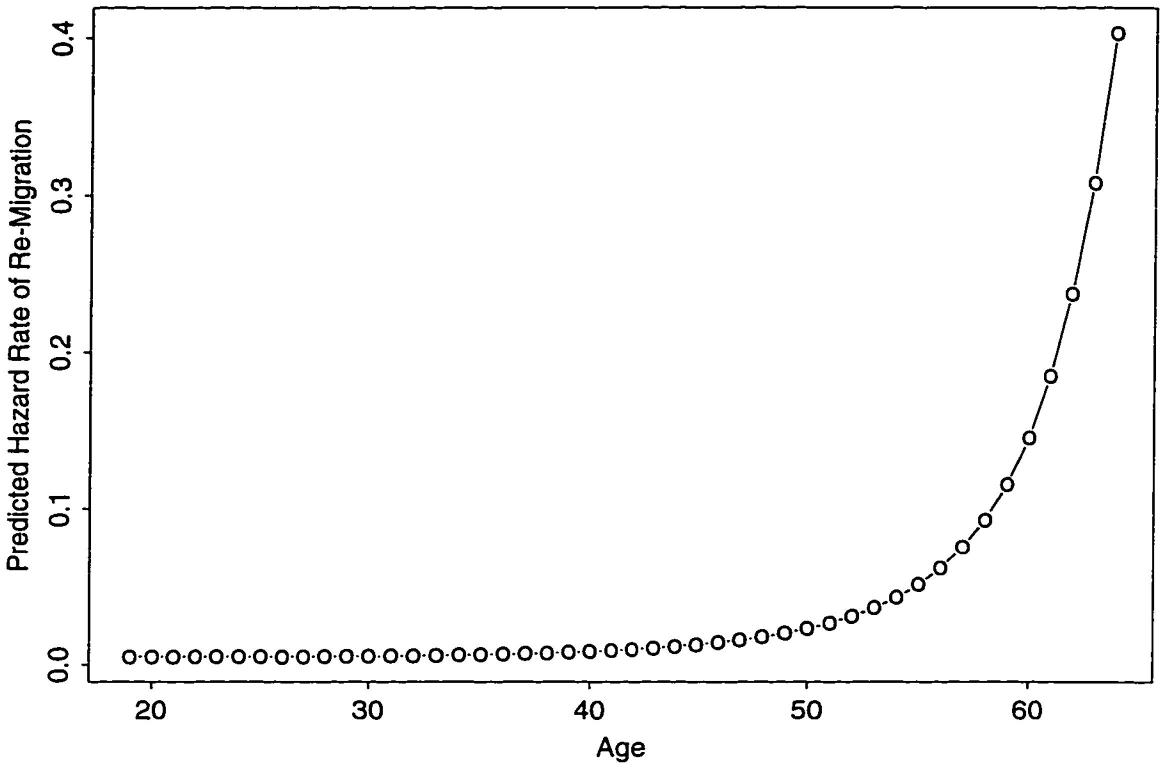


Figure 5.4 The Simulated Life-Cycle Effect on the Hazard Rate of Re-Migration for the Puerto Rico-Born Male Householder, Evaluated on Estimated Coefficients from Table 5.2 with Gamma Heterogeneity and Holding Everything Constant at Sample Means except Spell Duration and Age

and it is the language of the laws, constitution, and government on the U.S. mainland, it is not the official language of the country. Since the early 1980s, the official status of English has received much attention on the mainland. U.S. English is the primary group promoting laws to establish English as the official language of the United States. They see official language legislation as a matter of national unity and a means of resolving conflicts in a nation of diverse racial, ethnic, and religious groups. However, the English Plus Information Clearinghouse (EPIC), established in 1983, opposes official language legislation but favors strong English language proficiency plus mastery of second or even multiple languages. It views official English as synonymous with "English Only" which takes away the right to free speech as well as insulting the heritage of cultural minorities, including Mexican Americans and American Indians whose roots in the U.S. go deeper than those of English speakers.

Two major reasons for supporters demanding official English are dissatisfaction with bilingual programs, and the increase of Hispanic and Asian immigrants. Supporters have argued that the Hispanic and Asian population, aided by bilingual programs that maintain native language, do not want to learn English and do not tend to assimilate into mainstream American culture; an emotional response is that those immigrants who are different and want to remain different should go back to where they are comfortable and everyone speak their language (Stalker 1988). Furthermore, some anti-Hispanic arguments could be found in U.S. English's newsletters; for example, a lot of articles in U.S. English's newsletters assault Hispanics as the source of language difficulties in the U.S. (Adams 1992).

There are 18 states which have established official language laws (see Table 5.5). Some states have simply declared that their official language is English. Chapter 1, Section 3005 of Illinois Annotated Statutes (1969), for instance, reads: "The official language of the State of Illinois is English". Other states have added some enforcement provisions. California's official English amendment requires the use of English in all official state business, and prohibits printing ballots in any language other than English. The law in Arizona has the most severe enforcement provisions. It prohibits all government entities and employees from using any language other than English during the performance of government business.

Our estimated result shows that Puerto Rican men having poor English proficiency are more likely to re-migrate. This may suggest that, other things being equal, the greater number of states, like Arizona or California, enforce official English, the more Puerto Rico-born males will decide to re-migrate.

Table 5.5 States with Official "English" Language Legislation

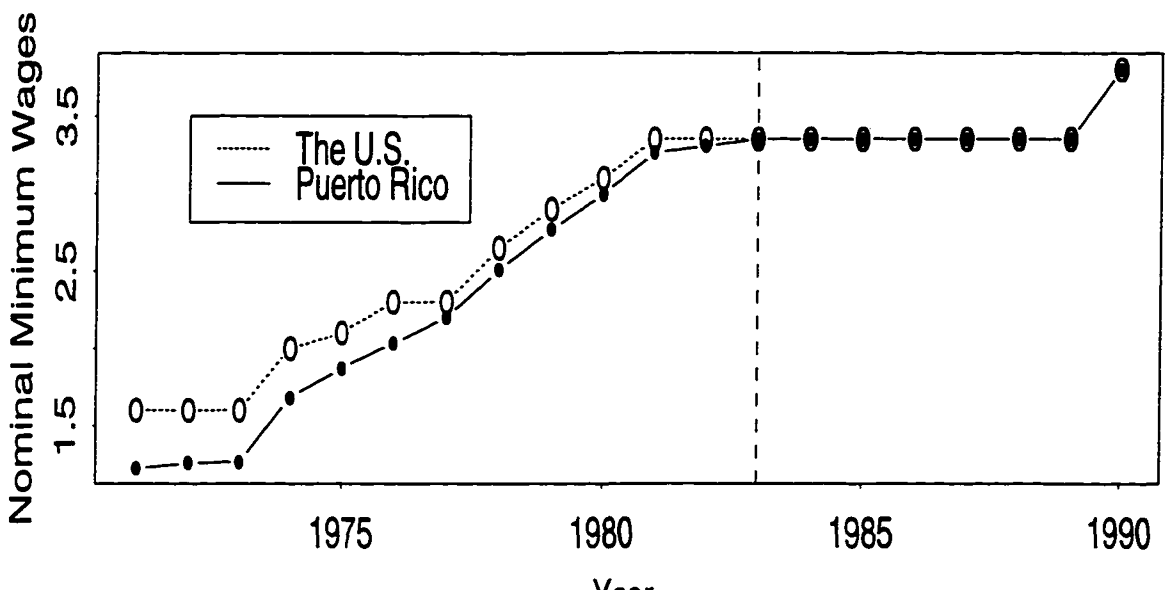
Alabama* (1990)[a]	Arizona** (1988)	Arkansan* (1987)
California** (1986)	Colorado* (1988)	Florida* (1988)
Georgia* (1986)	Hawaii** (1978)	Illinois* (1969)
Indiana* (1984)	Kentucky* (1984)	Mississippi* (1987)
Nebraska** (1920)	North Carolina* (1987)	North Dakota* (1987)
South Carolina** (1987)	Tennessee** (1984)	Virginia** (1981)

Source: Adopted from Table 1 of Tatalovich (1995).

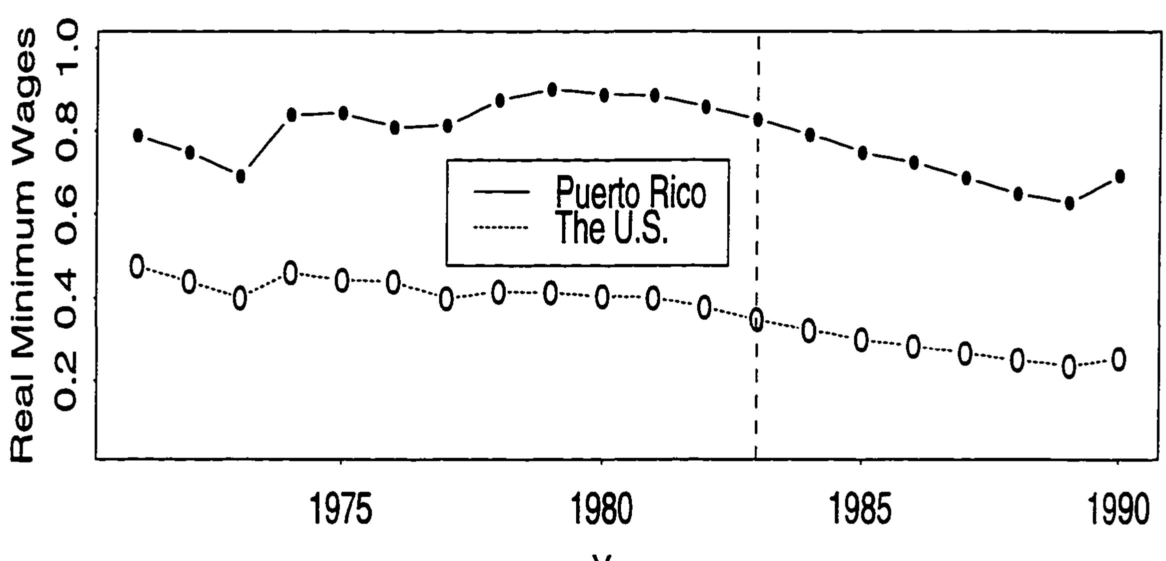
[a]The year of established official English legislation is in parentheses.
* States just declared that their official language is English.
** States added some enforcement provisions.

The U.S.-level minimum wage has been completely enforced in Puerto Rico since 1983. Figure 5.5 shows nominal minimum wages and minimum wages in terms of per capita personal consumption expenditures (PCE). Before 1983, indicated by the vertical dotted line, the nominal minimum wage for Puerto Rico was lower than that for the mainland; since then, both are identical. However, this could be misleading because of significantly different living cost between the mainland and the island. The per capita PCE is treated as a proxy for living cost. The minimum wage in the lower figure is derived from the minimum wage times 1,000 and then, divided by per capita personal consumption expenditures. It shows that minimum wages in terms of per capita PCE are much higher in Puerto Rico than those on the U.S. mainland. Hence, it is obvious that the minimum wage policy has much stronger impact on the island. Reynolds and Gregory (1965) and Castillo-Freeman and Freeman (1992) showed that the minimum wage resulted in a substantial number of unemployed people.

We find that a 10 cents increase in real minimum wages for Puerto Rico (in constant



(a) Nominal Minimum Wages



(b) Minimum Wages in terms of Per Capita Personal Consumption Expenditures

Figure 5.5 Nominal Minimum Wages (a) and Minimum Wages in terms of Per Capita Personal Consumption Expenditures (b) for the U.S. Mainland and Puerto Rico

1990 dollars) reduces the hazard rate of re-migration by 5.8 to 26.1 percent for Puerto Rico-born men on the U.S. mainland. Therefore, when the federal government raises the minimum wage , our model predicts that the number of re-migrating Puerto Ricans will fall, other things being equal.

CHAPTER 6 EMPIRICAL RESULTS OF EARNING ANALYSIS

This chapter discusses the empirical results of earning functions. First, we present the estimated reduced-form probit equations for the probability for being a wage earner. and estimated wage equations, corrected from the sample-selected bias. Then, we focus on the Chow test for parameter equality and stability.

Wage Determination

The coefficients in Table 6.1 provide estimates of the impact of the variables on the probability of wage work for men. In general, the signs of the estimated coefficients are mostly as expected and consistent across equations, at least when they are statistically significant. The life-cycle effect on probability of being a wage or salary earner is quadratic, revealed an inverted U-sharped form.

As expected, people with higher schooling, English proficiency, and who are married. are much more likely to be wage earners. Additional children under 18 years of age and having a disability have the expected effect, decreasing the probability. For Hispanics. more recent immigrants have a lower probability to be employees. Other income, urbanization, and predicted state unemployment rates have negative coefficients, while the real price of land raises the likelihood of being in the wage sample.

The sample means of variables used in the wage equations are reported in Table 6.2. The real hourly wage was much higher for whites than for Hispanics or for blacks. For

Table 6.1 Estimated Coefficients of Reduced-Form Probit Equations for the Probability with Positive Wage Earnings for U.S. Male Householders

	Hispanics		Whites		Blacks	
Variables	1979	1989	1979	1989	1979	1989
AGE	0.031*	0.035**	0.024	0.052	0.025	0.048*
	(1.68)[a]	(1.97)	(0.77)	(1.54)	(1.34)	(1.74)
AGESQ	−0.033	−0.043*	−0.028	−0.048	−0.014	−0.027
	(−1.38)	(−1.92)	(−0.68)	(−1.08)	(−0.58)	(−0.79)
ED	0.034***	0.032***	0.066***	0.082***	0.065***	0.095***
	(3.92)	(3.55)	(5.05)	(5.24)	(7.16)	(6.68)
EDEXP	−0.001	−0.0001	−0.0004	−0.001	−0.001	−0.002*
	(−0.58)	(0.07)	(−0.34)	(−0.86)	(−0.62)	(−1.93)
EDEXPSQ	0.0002	−0.001	−0.0003	−0.001	−0.001	0.0005
	(0.09)	(−0.68)	(−0.07)	(−0.19)	(−0.70)	(0.18)
ENG	0.082*	0.078*	0.659***	−0.444	0.265	0.296*
	(1.73)	(1.84)	(3.31)	(−1.30)	(1.25)	(1.78)
MARRIED	0.263***	0.194***	0.493***	0.471***	0.333***	0.354***
	(5.15)	(4.61)	(8.16)	(8.32)	(9.21)	(8.60)
KID06	−0.043*	−0.042*	−0.038	−0.041	−0.022	−0.038
	(−1.95)	(−1.91)	(−0.90)	(−1.08)	(−1.02)	(−1.46)
KID618	−0.034**	−0.024	−0.011	−0.034	−0.009	0.014
	(−2.30)	(−1.52)	(−0.44)	(−1.24)	(−0.73)	(0.75)
DISAB	−0.690***	−0.896***	−0.660***	−0.744***	−0.714***	−0.777***
	(−11.74)	(−17.04)	(−11.44)	(−13.21)	(−17.78)	(−16.75)
RACE	0.058	0.043	- -	- -	- -	- -
	(1.63)	(1.29)	- -	- -	- -	- -

[a] t-values are in parentheses.
* P-value ≤ 0.1 ** P-value ≤ 0.05 *** P-value ≤ 0.01

Table 6.1 (Continued)

Variables	Hispanics 1979	Hispanics 1989	Whites 1979	Whites 1989	Blacks 1979	Blacks 1989
FORB	0.175**	0.106	−0.008	−0.097	−0.290	−0.033
	(2.05)	(1.25)	(−0.05)	(−0.46)	(−1.51)	(−0.11)
USBE50	−0.154	0.135	−0.023	−0.272	−0.102	−0.244
	(−1.09)	(0.79)	(−0.09)	(−0.80)	(−0.40)	(−0.40)
US6064	−0.037	0.073	0.303	−0.393	0.620*	0.140
	(−0.34)	(0.64)	(0.75)	(−1.38)	(1.76)	(0.32)
US6569	0.0002	−0.015	0.367	−0.208	0.264	0.098
	(0.00)	(−0.15)	(0.96)	(−0.64)	(1.11)	(0.27)
US7074	−0.007	−0.182*	0.186	−0.108	0.143	0.355
	(−0.07)	(−1.94)	(0.55)	(−0.30)	(0.65)	(0.98)
US7579	−0.402***	−0.143	−0.484*	−0.272	0.245	0.182
	(−4.04)	(−1.49)	(−1.76)	(−0.79)	(1.01)	(0.52)
US8084	--	−0.185**	--	0.004	--	0.030
	--	(−1.97)	--	(0.01)	--	(0.09)
US8590	--	−0.368***	--	−0.776***	--	−0.029
	--	(−3.81)	--	(−2.66)	--	(−0.08)
$\ln(\frac{OTHINC}{P})$	−0.522***	−0.006	−0.744***	−0.298***	−0.428***	−0.084
	(−6.31)	(−0.11)	(−7.88)	(−4.45)	(−6.30)	(−1.50)
$\ln(\frac{PLAND}{P})$	0.098	0.050	0.155***	−0.003	0.064*	0.145***
	(1.59)	(1.08)	(2.82)	(−0.06)	(1.72)	(2.68)
URBAN	−0.005	−0.003	−0.005**	0.001	−0.003	−0.005***
	(−1.53)	(−0.91)	(−2.20)	(0.51)	(−1.63)	(−2.85)

Table 6.1 (Continued)

Variables	Hispanics 1979	Hispanics 1989	Whites 1979	Whites 1989	Blacks 1979	Blacks 1989
JAN	−0.010	0.027	0.001	0.008	−0.002	−0.026
	(−0.49)	(1.49)	(0.09)	(0.61)	(−0.13)	(−1.52)
JANSQ	0.006	−0.034	−0.002	−0.016	−0.002	0.035*
	(0.22)	(−1.57)	(−0.13)	(−0.95)	(−0.11)	(1.67)
PJOBGR	−0.021	0.027	0.068**	−0.028	0.012	−0.010
	(−0.74)	(0.86)	(2.44)	(−0.91)	(0.60)	(0.35)
PURATE	−0.107***	−0.089*	0.014	−0.047*	−0.056**	−0.095***
	(−2.64)	(−1.91)	(0.44)	(−1.68)	(−2.24)	(−4.13)
ESHOCK	−0.005	0.029	0.081***	−0.002	−0.034*	0.008
	(−0.13)	(1.17)	(2.98)	(−0.06)	(−1.72)	(0.32)
RURATE	0.005	−0.100	0.141***	−0.015	−0.065**	−0.047
	(0.10)	(−1.41)	(3.60)	(−0.33)	(−2.29)	(−1.45)
NC	−0.108	0.138	−0.083	0.065	−0.055	0.075
	(−0.62)	(1.13)	(−0.82)	(0.75)	(−0.80)	(1.01)
SOUTH	0.194*	0.057	−0.172	0.077	0.032	0.230**
	(1.87)	(0.33)	(−1.50)	(0.64)	(0.48)	(2.48)
WEST	0.510***	0.100	0.008	0.074	0.085	0.235**
	(4.39)	(0.60)	(0.07)	(0.55)	(0.93)	(2.04)
Intercept	6.912***	0.677	7.165***	3.795***	4.853***	0.355
	(6.32)	(0.82)	(5.80)	(3.51)	(5.35)	(0.39)
N	21,114	25,626	29,377	29,788	29,416	24,176

Table 6.2 Means of Variables in the Sample of Wage Earners for U.S. Male Householders

	Hispanics		**Whites**		**Blacks**	
Variables	1979	1989	1979	1989	1979	1989
W/P	11.9879	11.1625	14.8729	14.6599	12.7810	11.9801
EXP	21.2515	22.0353	20.9342	21.5191	22.0154	22.6432
EXPSQ	6.1290	6.3109	6.0126	5.8987	6.5498	6.4663
ED	10.1732	10.4264	12.9790	13.3793	11.5526	12.3151
EDEXP	191.0125	205.8754	257.7862	278.7496	230.6131	263.9058
EDEXPSQ	49.8153	53.3169	71.0271	73.9078	63.0003	71.0483
ENG	0.7859	0.7824	0.9961	0.9959	0.9977	0.9945
DISAB	0.0390	0.0350	0.0501	0.0474	0.0527	0.0472
RACE	0.5920	0.5331	- -	- -	- -	- -
FORB	0.4068	0.5632	0.0365	0.0353	0.0385	0.0631
USBE50	0.0205	0.0140	0.0069	0.0030	0.0032	0.0006
US6064	0.0631	0.0502	0.0047	0.0049	0.0038	0.0033
US6569	0.0864	0.0679	0.0046	0.0043	0.0089	0.0088
US7074	0.1081	0.0914	0.0045	0.0037	0.0116	0.0125
US7579	0.0760	0.0977	0.0044	0.0041	0.0079	0.0128

Table 6.2 (Continued)

Variables	Hispanics 1979	Hispanics 1989	Whites 1979	Whites 1989	Blacks 1979	Blacks 1989
US8084	--	0.1140	--	0.0036	--	0.0147
US8590	--	0.0812	--	0.0035	--	0.0081
ln(PLAND/P)	6.9140	6.8584	7.1461	6.9388	7.1809	7.0008
URBAN	85.4647	86.1201	75.8112	76.6493	76.5893	76.4888
JAN	42.5092	43.8115	33.7994	34.1715	37.1175	38.7952
JANSQ	19.8186	20.7549	12.8726	13.1628	15.1026	16.2785
PJOBGR	3.6321	2.2329	3.0532	2.1540	2.9961	2.2293
PURATE	6.4667	6.1004	6.1717	5.7829	6.2583	5.9566
ESHOCK	0.5361	0.7024	−0.7278	0.2266	−0.2407	0.3522
RURATE	0.1631	−0.4079	1.0759	−0.2667	0.9471	−0.3800
NC	0.0898	0.0733	0.2750	0.2761	0.1969	0.1563
SOUTH	0.2897	0.3401	0.2977	0.3259	0.5086	0.5993
WEST	0.4367	0.4538	0.1769	0.1792	0.0923	0.1006
$\hat{\lambda}$	−0.0752	−0.0695	−0.0319	−0.0317	−0.0924	−0.0829

whites. the real hourly wage was \$14.78 (\$14.66) in 1979 (1989), while for Hispanics it was \$11.99 (\$11.16) and for blacks \$12.78 (\$11.98). Although the real hourly wage for all three ethnic groups declined, Hispanics' and blacks' real wage rates decreased 7 and 6 percent respectively, in comparison with whites' real wage rates just 1 percent. Average education levels had increased from 1979 to 1989 for all ethnicities: 10.2 to 10.43 for Hispanics, 13 to 13.4 for whites. and 11.6 to 12.3 for blacks. Average education attainments reveal that Hispanics are disadvantaged by possessing less human capital on average than whites or blacks.

The significantly lower percentage of Hispanics responding speaking English well or very well mirrors their predominantly foreign origin. It was 41 (56) percent of foreign born for Hispanics in 1979 (1989) in contrast to 3.7 (3.5) and 3.9 (6.3) for whites and blacks respectively. This also reflects the fact that there was a large immigrant flow of Hispanics during the 1980s.

The regional dummy variables show that most Hispanics lived in the west and south, while blacks concentrated in the south. In contrast, whites lived relatively uniformly throughout the U.S. This also results in some local variables having substantially different sample means among them.

The estimated coefficients of wage equations, corrected for sample-selected bias. are reported in Table 6.3; two for Hispanic men, two for white men. and two for black men. The first equation is the earning analysis of 1979 and the second is for 1989. The inverse of Mill's ratio $\hat{\lambda}$ is added to each equation to correct for possible sample-selection bias in order to get consistent estimates of the parameters of the wage-offer function facing each ethnic group.

Years of work experience and education produce the expected positive association with earnings. The negative coefficient of EXPSQ shows a diminishing marginal rate of work experience on wage earnings. Our estimated coefficients of EDEXP and EDEXPSQ measure on average not only how years of schooling affect the experience-wage profile,

Table 6.3 Estimated Coefficients of Wage Equations for U.S. Male Householders, Corrected for Sample-Selected Bias

	Hispanics		Whites		Blacks	
Variables	1979	1989	1979	1989	1979	1989
EXP	0.044***	0.054***	0.020***	0.033***	0.033***	0.046***
	(9.37)[a]	(14.93)	(4.18)	(6.75)	(6.10)	(8.48)
EXPSQ	−0.050***	−0.065***	−0.001	−0.013	−0.021**	−0.036***
	(−6.61)	(−11.49)	(−0.15)	(−1.52)	(−2.49)	(−4.16)
ED	0.068***	0.080***	0.047***	0.081***	0.073***	0.089***
	(12.87)	(18.79)	(11.83)	(18.60)	(12.55)	(14.79)
EDEXP	−0.0003	−0.001***	0.002***	0.001***	0.0003	−0.0004
	(−0.87)	(−3.59)	(5.31)	(2.61)	(0.66)	(−0.98)
EDEXPSQ	−0.002**	−0.001	−0.005***	−0.004***	−0.003***	−0.002***
	(−2.17)	(−1.01)	(−7.75)	(−6.71)	(−4.43)	(−3.54)
ENG	0.113***	0.117***	−0.022	0.135***	0.086	−0.125**
	(7.08)	(10.22)	(−0.35)	(2.56)	(0.89)	(−2.38)
DISAB	−0.095**	−0.198***	−0.029	−0.060**	−0.051	0.149***
	(−2.13)	(−3.85)	(−1.26)	(−2.42)	(−1.29)	(4.82)
RACE	0.032***	0.040***	- -	- -	- -	- -
	(2.85)	(5.12)	- -	- -	- -	- -
FORB	0.056**	0.040**	0.046	0.058	−0.062	0.003
	(2.29)	(2.09)	(1.35)	(1.58)	(−0.74)	(0.03)
USBE50	−0.094**	−0.002	0.009	0.021	0.073	−0.105
	(−2.19)	(−0.04)	(0.17)	(0.30)	(0.63)	(−0.59)
US6064	−0.039	0.038	0.074	0.062	0.014	0.076
	(−1.27)	(1.55)	(1.19)	(1.04)	(0.12)	(0.74)

[a]t-values are in parentheses.
* P-value $\leq$ 0.1 ** P-value $\leq$ 0.05 *** P-value $\leq$ 0.01

Table 6.3 (Continued)

	Hispanics		Whites		Blacks	
Variables	1979	1989	1979	1989	1979	1989
US6569	−0.082***	−0.046**	−0.180***	0.062	0.002	−0.024
	(−2.82)	(−2.01)	(−2.86)	(1.00)	(0.02)	(−0.28)
US7074	−0.155***	−0.110***	−0.150**	0.045	−0.023	−0.066
	(−5.43)	(−4.89)	(−2.33)	(0.70)	(−0.25)	(−0.78)
US7579	−0.223***	−0.140***	−0.099	−0.115*	−0.194**	−0.139*
	(−6.47)	(−6.23)	(−1.53)	(−1.82)	(−1.99)	(−1.65)
US8084	- -	−0.223***	- -	−0.070	- -	−0.184**
	- -	(−9.86)	- -	(−1.06)	- -	(−2.20)
US8590	- -	−0.326***	- -	0.031	- -	−0.219**
	- -	(−12.08)	- -	(0.45)	- -	(−2.48)
$\ln(\frac{PLAND}{P})$	0.057***	0.100***	0.023***	0.069***	0.012	0.058***
	(3.42)	(9.79)	(2.42)	(7.91)	(0.94)	(4.17)
URBAN	0.001	0.002***	0.003***	0.004***	0.004***	0.005***
	(1.20)	(3.42)	(6.44)	(14.40)	(7.20)	(11.75)
JAN	−0.011**	−0.014***	−0.001	−0.005**	0.009*	−0.005
	(−1.97)	(−3.22)	(−0.26)	(−2.31)	(1.86)	(−1.23)
JANSQ	0.015*	0.015***	−0.002	0.005*	−0.017***	−0.003
	(1.95)	(3.03)	(−0.60)	(1.84)	(−2.77)	(−0.61)
PJOBGR	−0.006	−0.022***	0.002	−0.044***	−0.002	−0.026***
	(−0.67)	(−3.23)	(0.40)	(−9.01)	(−0.33)	(−3.74)
PURATE	−0.016	−0.043***	0.026***	0.001	0.021**	0.025***
	(−1.40)	(−4.28)	(4.74)	(0.25)	(2.44)	(4.41)
ESHOCK	−0.018*	−0.014***	−0.017***	−0.019***	0.010	−0.006
	(−1.78)	(−2.66)	(−3.68)	(−4.70)	(1.52)	(−0.91)

Table 6.3 (Continued)

	Hispanics		**Whites**		**Blacks**	
Variables	1979	1989	1979	1989	1979	1989
RURATE	0.004	−0.077***	0.027***	−0.008	0.029***	0.019**
	(0.26)	(−5.26)	(4.00)	(−1.02)	(2.98)	(2.24)
NC	0.079*	0.060**	0.050***	0.049***	0.152***	0.010
	(1.79)	(2.21)	(3.02)	(3.63)	(6.43)	(0.52)
SOUTH	0.027	0.011	0.009	0.075***	0.002	−0.015
	(0.85)	(0.28)	(0.43)	(4.10)	(0.09)	(−0.62)
WEST	0.123***	0.088**	0.130***	0.180***	0.167***	0.191***
	(3.55)	(2.40)	(6.30)	(8.44)	(5.56)	(6.71)
$\hat{\lambda}$	−0.015	−0.282	1.831***	1.362***	0.229	1.369***
	(−0.06)	(−1.20)	(9.75)	(7.30)	(1.25)	(10.56)
Intercept	0.731***	0.367***	0.945***	0.070	0.262	0.250*
	(4.11)	(2.68)	(7.82)	(0.68)	(1.42)	(1.68)
R^2	0.1089	0.2171	0.1401	0.2178	0.0870	0.1988
Adj. R^2	0.1077	0.2162	0.1393	0.2171	0.0862	0.1979
N	20.412	24,807	28.985	29,393	28.117	23.218

but also the relative importance of the general human capital (education) and the specific human capital (experience).

The marginal effect of education on wage earnings is given by $\beta_3 + \beta_4 EXP - \beta_5 EXPSQ$. The positive β_4 and negative β_5 in white men suggest that the marginal returns of education on earnings is an inverted U-shaped form along the years of work experience. The interpretation may be that in early working life, general human capital (education) is a critical factor determining the wage earnings, but in latter working life, specific human capital (experience) plays a strong role explaining the real wages. This seems reasonable because in their early working life, workers have not accumulated much specific human capital. Hence, general human capital contributes a major part of human capital. Later, workers have accumulated considerable specific human capital, of which work experience provides an important element. The negative β_4 and β_5 in Hispanics imply that the marginal returns of schooling on real wages decrease at an increasing marginal rate when the years of work experience increase. The reason might be Hispanic men's lower average education attainments, slightly above 10 years. Hence, the specific human capital (work experience) plays a relatively more important role than general human capital (education) in explaining wage earnings. For people with 20 (30) years of work experience, the marginal effects of education in earnings are as follow: a one year increase of schooling augments the real wage by about 5.4 (4.1) percent in 1979 and 5.6 (4.1) percent in 1989 for Hispanic men, about 6.7 (6.2) percent in 1979 and 8.5 (7.5) percent in 1989 for white men, and about 6.7 (5.5) percent in 1979 and 7.3 (5.9) percent in 1989 for black males. The effects are increasing from 1979 to 1989 for all ethnic groups. This may be because the sample means of education attainments increased from 1979 to 1989 for all ethnic groups. Furthermore, Hispanic men show the lowest returns on schooling, while white men have the highest returns. The reason may be different average education attainments among ethnic groups. The other explanation could be differences in average schooling quality by ethnic groups. The positive effect of schooling on earnings is consistent with earlier work by Reimers (1985) and Rivera-Batiz (1991).

The negative coefficients of EDEXPSQ for all ethnic groups indicate that more educated men have more concave experience-wage profiles. The effects of EDEXP are not consistent across all ethnic groups: the experience-wage profile for more educated males is flatter for Hispanics, steeper for whites, and has no significant effect for blacks. In other words, more educated white men have higher returns of work experience but at a diminishing marginal rate, while the returns of experience is lower for more educated Hispanic men and even at an increasing marginal rate. However, for a given education level, people in a sample having a higher mean of schooling tend to achieve the peak of experience-wage profiles later. EXP$= 50(\beta_1 - \beta_4 \times \text{ED})/(\beta_2 - \beta_5 \times \text{ED})$ gives the value of EXP for which the wage is the highest. For men graduated from high school (4 years university), the peak effect of experience occurs at 27.3 (23.9) years in 1979 and 27.3 (23.5) in 1989 for Hispanics, 36.1 (32.1) years in 1979 and 36.9 (31.8) in 1989 for whites, and 32.1 (27.4) in 1979 and 34.3 (29.1) for blacks. The inverted U-sharped form of the experience-wage profile has been reported in many studies.

Lack of fluency in English does have a significant effect on Hispanic men's wages, but the effect is not clear for white and black men. Reimers (1985) found a similar result. The earnings of Hispanic males with English proficiency are rewarded by about 11.3 percent in 1979, slightly increasing to 11.7 percent in 1989. The effect is not consistent for white and black men. English proficiency raises white men's wage earnings by 13.5 percent in 1989, but insignificantly lowers their real wages by 2.2 percent in 1979. It reduces black men's earnings by 12.5 percent in 1989, but increases earnings about 8.6 percent in 1979.

The signs of estimated coefficients of the disability status are as expected, except for blacks in 1989. This result is similar to the finding of Reimers (1985). Disability status significantly lowers Hispanic men's earnings. It hurts the earning power of Hispanics much more than that of whites and blacks. The estimate of wage loss due to disability is 9.5 percent in 1979 and 19.8 percent in 1989. White males having disability incur an earning loss of 3 percent in 1979 and 6 percent in 1989. Although disabled black men's earnings are penalized by 5 percent in 1979, they are rewarded by 14.5 percent in 1989

We also find that race has a significant impact on the wage earnings of Hispanic men. The estimated coefficients on the RACE dummy variable are positive and significantly different from zero at the 1 percent significant level. White Hispanic males earn 3 percent to 4 percent more than non-white Hispanic men, slightly increasing from 1979 to 1989.

The dummy variables FORB, USBE50, US6064, US6569, US7074, US7579, US8084, and US8590, tell us how the wage earnings of successive cohorts of immigrant men compared with U.S.-born members of their ethnic group. For example, other things being equal, FORB + US6064 shows the wage differential of an immigrant who arrived in the U.S. between 1960 and 1964, compared with a U.S.-born Hispanic. The results indicate that Hispanic immigrant men will catch up with U.S.-born Hispanic men after they have been here at least 15 years for the 1980 sample, but it requires at least 25 years for the 1990 sample. Furthermore, both Hispanic samples show that the cohort that arrived after 1965 earns less than U.S. natives. White immigrant men do not match the U.S.-born white men until they have been here 15–19 years in the 1980 sample, while in the 1990 sample, new arrivals and those who have been here more than 15 years earn more than U.S. natives. Black immigrant men overtake U.S. natives when they arrived before 1950 in the 1980 sample and when they arrived between 1951 and 1964 in the 1990 sample.

Other things being equal, Hispanic immigrant men show that more recent immigrant cohorts have greater economic disadvantage. New arrivals (and those who have been here 5 to 10 years), for instance, earn 16.7 (9.9) percent less than U.S. natives in 1979, but 28.6 (18.3) percent less in 1989. This is consistent with the study by Borjas (1990). He found that more recent immigrant waves are less skilled than earlier waves. We did not find a similar pattern in whites and blacks perhaps because of the lower percentage of the foreign-born in these groups. Among Hispanics, 41 (56) percent were foreign-born in 1979 (1989), in contrast to 3.7 (3.5) and 3.9 (6.3) percent for whites and blacks respectively.

Differences in cost of living and local amenities also cause variable wage earnings. A test of the null hypothesis that the four coefficients of these variables (ln(PLAND/P), URBAN, JAN, JANSQ) are jointly equal to zero is rejected at the 1 percent significant level. The sample value of the F-statistic is 9.9 (115.9) in 1979 (1989) for Hispanics, 21.1 (143.5) for whites, and 37.3 (98.1) for blacks. The critical value of the F-statistic with 4 and infinite degrees of freedom is 3.32 at the 1 percent significant level.

The real price of land and proportion of urban population have positive effects on earnings. The magnitudes are increasing from 1979 to 1989. Moreover, Hispanic men have the highest elasticity of land price and the lowest effect of URBAN, while blacks have the lowest elasticity of land price and the highest effect of URBAN. All estimated coefficients of the real price of land are statistically significantly different from zero at 1 percent significant level, except blacks in 1979. The elasticities range from 0.057 to 0.1 for Hispanics, 0.023 to 0.069 for whites, and 0.012 to 0.058 for blacks. Tokle and Huffman (1991) and Feridhanusetyawan (1994) found similar effects: 0.060–0.073 for Tokle and Huffman and 0.041–0.044 for Feridhanusetyawan. The marginal effects of a one percentage point increase in URBAN on wage earnings range from 0.001 to 0.002 for Hispanics, 0.003 to 0.004 for whites, and 0.004 to 0.005 for blacks. Compared with Tokle and Huffman (1991), 0.180–0.255, and Feridhanusetyawan (1994), 0.283–0.303, our effects are much lower.

The average local temperature of January is treated as a proxy for local amenities. The estimated coefficients of JAN and JANSQ are not consistent across equations. The effects of January temperature on earnings are quadratic and statistically significant for Hispanics, whites in 1989, and blacks in 1979. The effects on the other two equations are neither quadratic nor statistically significant. Furthermore, the the quadratic effect is a U-shaped form for Hispanics and whites, while it is an inverted U-shaped form for blacks.

We also find that local labor market characteristics plays an important role in explaining the wage earnings. A joint test of no effect of local labor market characteristics (PJOBGR, PURATE, ESHOCK, RURATE) on wage earnings is rejected at the 1 percent significant level for all equations, except Hispanics in 1979, rejected at the 10 percent level. The sample value of the F-statistic is 2.1 (8.5) in 1979 (1989) for Hispanics, 9.5 (24.0) for whites, and 4.8 (14.5) for blacks. The critical value of the F-statistic with 4 and infinite degrees of freedom is 3.32 (1.94) at the 1 (10) percent significant level. Anticipated state unemployment rates have positive effects on earnings of whites and blacks. This result is consistent with the economic theory that a higher probability of unemployment is compensated by a higher wage rate. However, we find negative effects on Hispanic men. A one percentage point increase in predicted state unemployment rates cause 1.6 to 4.3 percent decline on wage earnings for Hispanics, 0.1 to 2.6 percent increase for whites, and 2.1 to 2.5 percent enhancement for blacks.

The estimated coefficients of predicted state job growth rates are negative, except for whites in 1979, but it is not significant. This suggests that localities with higher anticipated job growth rates will have lower equilibrium wage rates than other localities. This is in contrast to the positive effect from Tokle and Huffman (1991). Empirical results show that a one percentage point increase in PJOBGR results in 0.6 to 2.2 percent decrease on the real wage rate for Hispanics, and 0.2 to 2.6 percent decline for blacks. The marginal effect ranges from −4.4 to 0.2 percent for whites. The loss of earnings resulting from PJOBGR increases from 1979 to 1989.

The effect of unanticipated labor market shocks on real wages is not consistent across equations. The marginal effect of a one percentage point increase in RURATE on earnings ranges from −0.077 to 0.004 for Hispanics, −0.008 to 0.027 for whites, and 0.019

to 0.029 for blacks. There is a lower return of RURATE in 1989 than in 1979 for all ethnic groups. The marginal effect of a one percentage point increase in ESHOCK on real wages is −0.018 in 1979 and −0.014 in 1989 for Hispanics, 0.017 in 1979 and −0.019 in 1989 for whites, and 0.01 in 1979 and −0.029 in 1989 for blacks.

The significant estimates of regional dummy variables show the importance of unmeasured regional effects on wage earnings. The F-test of the null hypothesis that the coefficients of three regional dummies (NC, SOUTH, WEST) are jointly equal to zero is rejected at the 1 percent significant level. The sample value of the F-statistic is 6.3 (8.9) in 1979 (1989) for Hispanics, 25.2 (32.3) for whites, and 30.3 (50.6) for blacks. The critical value of the F-statistic with 3 and infinite degrees of freedom is 3.78 at the 1 percent significant level.

Wage rates in the western states are statistically significantly higher than those in the north-eastern states for all equations. Being in the western states rather than in the north-eastern states earns more by 8.8 to 12 percent for Hispanics, 16.7 to 18 percent for whites, and 16.7 to 19.1 percent for blacks. Furthermore, the western states have the highest earnings in the U.S. The north-central states also have higher wage earnings than the north-eastern sates: all equations are statistically significant except blacks in 1989. However, wage rates in the southern sates are not statistically significantly different from the north-eastern sates except for whites in 1989 with the estimated coefficient 0.075 with p-value < 0.01.

Parameter Equality and Stability

This section discusses the structure difference of earning functions. The Chow test procedure is used to examine the parameter equality of wage functions between ethnic groups, or the parameter stability between 1979 and 1989 for each ethnic group. We not only perform the null hypothesis of parameter equality or stability in the earning functions, but also do the modified Chow test, in which the null hypothesis is that the coefficients of a particular batch of explanatory variables are identical in the two earning functions.

Table 6.4 presents the Chow test for parameter equality of earning functions between each pair of ethnic groups. We roughly divide all explanatory variables into four categories, personal characteristics, cost of living and local amenities (consisting of ln(PLAND/P), URBAN, JAN, and JANSQ), local labor market characteristics (containing PJOBGR, PURATE, ESHOCK, and RURATE), and regional effect (including NC, SOUTH, and WEST). The personal characteristics consists of human capital (including EXP, EXPSQ, ED, EDEXP, EDEXPSQ), DISAB, ENG, and cohort effect which contains all dummies representing successive cohorts of immigrant men. The overall effect includes all explanatory variables in these four categories and the intercept.

The Chow tests for parameter equality between two wage functions are significantly rejected at the 0.01 percent significant level for each pair of ethnicities both in 1979 and 1989. This suggests that the economic returns of the earning function are significantly different between ethnic groups.

We find that the returns on personal characteristics in earnings are significantly different for any two ethnic groups at the 0.01 percent significant level. Of the personal characteristics, the coefficients of the human capital variables contribute the major part of the structural differences. The cohort effect is significantly different between whites and blacks both in 1979 and 1989, and between Hispanics and whites in 1989, at the 5 percent significant level, while it is not significant for the other three pairs. The return to English proficiency is insignificantly different for each pair in 1979, but it is stronger and significantly different in 1989 except for Hispanics and whites.

The null hypothesis that all parameters of cost of living and local amenities are equal between two wage functions is rejected at the 95 percent confidence level for each pair of ethnicities. The corresponding null hypothesis of parameter equality in the local labor

Table 6.4 The Chow Test for No structural Difference in the Wage Equations of Male Householders between Ethnic Groups

	1979			1989		
	d.f.	F-value	P-value	d.f.	F-value	P-value
Hispanics and Whites						
Overall effect[a]	(25, 49344)	12.147	0.0001	(27, 54143)	26.443	0.0001
Personal characteristics[b]	(13, 49344)	5.866	0.0001	(15, 54143)	20.374	0.0001
Human capital[c]	(5, 49344)	12.019	0.0001	(5, 54143)	47.681	0.0001
Cohort effect[d]	(6, 49344)	1.986	0.0639	(8, 54143)	6.582	0.0001
ENG	(1, 49344)	3.559	0.0592	(1, 54143)	0.105	0.7457
Cost of living & local amenities[e]	(4, 49344)	2.455	0.0436	(4, 54143)	3.338	0.0097
Labor market characteristics[f]	(4, 49344)	4.737	0.0008	(4, 54143)	10.673	0.0001
Regional effect[g]	(3, 49344)	0.336	0.7990	(3, 54143)	2.864	0.0352
Hispanics and Blacks						
Overall effect	(25, 48476)	4.120	0.0001	(27, 47968)	10.994	0.0001
Personal characteristics	(13, 48476)	2.729	0.0001	(15, 47968)	9.857	0.0001

[a]The overall effect includes all explanatory variables in personal characteristics, cost of living and local amenities, local labor market characteristics, regional effect, and the intercept.

[b]The personal characteristics consists of human capital, DISAB, ENG, and cohort effect.

[c]Human capital variables include EXP, EXPSQ, ED, EDEXP, and EDEXPSQ.

[d]Cohort effect contains all dummies representing successive cohorts of immigrant men.

[e]The variables of cost of living and local amenities consist of ln(PLAND/P), URBAN, JAN, and JANSQ.

[f]Local labor market variables contain PJOBGR, PURATE, ESHOCK, and RURATE.

[g]Regional dummy variables include NC, SOUTH, and WEST.

Table 6.4 (Continued)

	1979			1989		
	d.f.	F-value	P-value	d.f.	F-value	P-value
Human capital	(5, 48476)	5.109	0.0001	(5, 47968)	16.120	0.0001
Cohort effect	(6, 48476)	0.951	0.4568	(8, 47968)	0.555	0.8157
ENG	(1, 48476)	0.076	0.7823	(1, 47968)	19.934	0.0001
Cost of living & local amenities	(4, 48476)	6.473	0.0001	(4, 47968)	14.101	0.0001
Labor market characteristics	(4, 48476)	2.567	0.0362	(4, 47968)	12.024	0.0001
Regional effect	(3, 48476)	1.934	0.1217	(3, 47968)	10.430	0.0001
Whites and Blacks						
Overall effect	(25, 57050)	22.068	0.0001	(27, 52555)	19.896	0.0001
Personal characteristics	(13, 57050)	8.045	0.0001	(15, 52555)	7.509	0.0001
Human capital	(5, 57050)	18.136	0.0001	(5, 52555)	9.917	0.0001
Cohort effect	(6, 57050)	2.146	0.0451	(8, 52555)	4.371	0.0001
ENG	(1, 57050)	0.913	0.3393	(1, 52555)	12.247	0.0005
Cost of living & local amenities	(4, 57050)	4.433	0.0014	(4, 52555)	13.989	0.0001
Labor market characteristics	(4, 57050)	0.685	0.6020	(4, 52555)	3.411	0.0085
Regional effect	(3, 57050)	5.959	0.0005	(3, 52555)	10.508	0.0001

market characteristics between ethnic groups is also rejected at the 5 percent significant level, barring white and black men in 1979. Finally, the F-test statistic of the null hypothesis that two earning functions have the same regional effect is rejected at the 5 percent significant level in 1989, while it is insignificant in 1979, except for white and black males which is significant at the 99 percent confidence level.

In summary, the economic returns on personal characteristics provide a major source of wage differentials between ethnic groups. Of the personal characteristics, the most important part is the returns on human capital variables. This may reflect the fact that different ethnicities have different average education attainments. Furthermore, the other three categories of explanatory variables also serve important roles in explaining wage discrimination between ethnic groups for men.

The Chow test for parameter stability of earning functions from 1979 to 1989 for each ethnicity is reported in Table 6.5. The explanatory variables in each category are the same as in Table 6.4, except the cohort effect. The corresponding null hypothesis for the cohort effect is that the mean earning disadvantage of immigrant men compared with U.S. natives only depends on how long they have been in the U.S. Hence, under null hypothesis, FORB + US8084 in 1989, for example, is equal to FORB + US7074 in 1979. We consider four cohorts when they have been here 1–5, 6–10, 11–15, or 16–20 years. Furthermore, the overall effect and personal characteristics consists of RACE for Hispanic men.

We find that the F-value from the overall effect for parameter stability is greater than the critical value of the F-statistic at the 0.01 percent significant level for each ethnicity. This suggests that the economic returns of the earning function already changed from 1979 to 1989. However, there are different sources for these changes. The F-values show that the coefficients for personal characteristics, and cost of living and local amenities are statistically significantly different between 1979 and 1989 for all of them. Nevertheless, we do not reject the null hypothesis of parameter stability in local labor market

Table 6.5 The Chow Test for No Structural Change for the Wage Equations of Male Householders within an ethnic group during the 1980s

	Hispanics			Whites			Blacks		
	d.f.	F-value	P-value	d.f.	F-value	P-value	d.f.	F-value	P-value
Overall effect	(24, 45163)	7.170	0.0001	(23, 58324)	14.770	0.0001	(23, 51281)	7.590	0.0001
Personal characteristics[a]	(12, 45163)	3.271	0.0001	(11, 58324)	11.441	0.0001	(11, 51281)	6.055	0.0001
Human capital	(5, 45163)	0.823	0.5331	(5, 58324)	22.436	0.0001	(5, 51281)	6.330	0.0001
RACE	(1, 45163)	0.426	0.5140	--	--	--	--	--	--
Cohort effect[b]	(4, 45163)	7.092	0.0001	(4, 58324)	1.334	0.2546	(4, 51281)	1.225	0.2976
ENG	(1, 45163)	0.047	0.8277	(1, 58324)	3.617	0.0572	(1, 51281)	3.911	0.0480
Cost of living & local amenities	(4, 45163)	3.697	0.0052	(4, 58324)	14.661	0.0001	(4, 51281)	4.647	0.0009
Labor market characteristics	(4, 45163)	7.780	0.0001	(4, 58324)	12.898	0.0001	(4, 51281)	1.968	0.0964
Regional effect	(3, 45163)	0.204	0.8940	(3, 58324)	2.734	0.0420	(3, 51281)	7.691	0.0001

[a]Personal characteristics consists of RACE for Hispanic men.
[b]The null hypothesis for the cohort effect is that the mean earning disadvantage of immigrant men compared with the U.S. natives only depends on how long they have been in the U.S. We consider four cohorts when they have been here 1–5, 6–10, 11–15, or 16–20 years.

characteristics for black men, and the coefficients of regional dummies do not change significantly for Hispanic men.

Although the null hypothesis of parameter stability for personal characteristics is rejected at the 0.01 percent significant level for each ethnic group, the major source is different for them. The results show that the coefficients of human capital variables and ENG are the principal source of the change in the personal characteristics during the 1980s for whites and blacks. We find, however, that the cohort effect is the only part of the returns on personal characteristics having changed significantly for Hispanics from 1979 to 1989 (p-value = 0.0001); the other part of personal characteristics is insignificant at the 50 percent level. This may imply that the decrease in Hispanic men's average real wage rates mainly come from cost of living and local amenities, and local labor market characteristics. The result is consistent with our hypothesis that Hispanics suffered labor market competition during the 1980s.

In summary, the structural change for the earning function of Hispanic men during the 1980s mainly comes from the returns for cohort effect, cost of living and local amenities, and labor market characteristics. For whites, it is because of the change in coefficients for human capital, ENG, cost of living and local amenities, local labor market characteristics, and regional dummies. Blacks have a situation similar to whites, except for no significant change in the coefficients of local labor market characteristics.

CHAPTER 7 SUMMARY AND CONCLUSIONS

Sequential Migration Model

This study has developed a consistent sequential decision framework to describe international migration decisions under incomplete information. We applied a life-cycle model where individuals migrate only when they expect to have greater utility in a new location. In contrast to other international migration models, our model focuses on explaining how immigrants make decision to re-migrate when faced with new incoming information.

In our model, potential emigrants do not know the wage rate and/or the quality of life in a new host country, but they do have their own personal subjective probability distribution about these economic variables. This suggests that emigrants are attracted by higher expected wage rates and/or quality of life of new destinations. After individuals understand the wage rate they can earn and/or experience the true quality of life in a new country, they must make a decision about where to live in order to maximize their expected life-time utility. Hence, some imigrants re-migrate because the wage earnings and/or the quality of life in a new host country are not good enough to keep them.

In our model, even though the individual suffered a layoff in the host country, the optimal strategy is not necessarily to re-migrate because the host country may still provide a higher expected remaining life-time utility. Finally, when people approach retirement, the key factor influencing emigrants is the quality of life. This may explain why a lot of people re-migrate near retirement.

The model predicted that higher mean or variance of the personal subjective wage distribution (or quality of life distribution) increases the attraction of the host country, while the wage rate or the quality of life in the home country has a negative effect on the net value of emigration. Furthermore, a decrease of the probability of an individual being unemployed in the host country, a higher moving cost, or an increase in the probability of a layoff in the home country increases the likelihood of emigration.

On the other hand, we showed that a larger moving cost of re-migration or the probability of unemployment in the home country has a negative effect on the net value of re-migration, but a larger wage rate, better quality of life in the home country, or higher probability of unemployment in the host country encourages emigrants to re-migrate.

The Hazard Rate for Return Migration

The most important objective of this thesis was to examine econometrically how social and economic factors affect return migration behaviors of Puerto Rico-born men who are on the U.S. mainland. The hazard rate approach, the instantaneous re-migrating probability given that Puerto Rico-born men have been on the mainland for several years, was used in the empirical analysis. We used both the exponential and Weibull regression models in this work.

The sample consisted of 12,108 observations, where 1,682 observations had not recorded when a migration spell began. We assumed that an individual can perform a full-time job only if he is older than age 18 and has completed his education. After adjustments, we still had 1,183 left-censored spells. Because of the "memoryless' property of the exponential distribution, we ignored the period from the start of a residence spell to when we begin to observe individuals. Nevertheless, two expedients or research strategies were applied to the Weibull regression model.

In the first expedient or research strategy, we assumed that those Puerto Rico-born men having left-censored spells began their working life on the U.S. mainland. For the second expedient or research strategy, we predicted the length of the working spells on the U.S. mainland for those Puerto Rico-born males having left-censored spells. Instead of predicting the spell itself, we predicted the starting-age of a spell to avoid an inconsistent assumption about the distribution of the duration of a completed spell and to prevent overestimating the coefficients of some explanatory variables used in the predicted stage. With the sample of completed spells (or right-censored spells), we regressed the natural logarithm of an individual's starting-age of a residence spell on a set of explanatory variables. Equations were adjusted to correct possible non-random selection bias by adding the inverse of Mill's ratio. The length of migration spells for individuals with left-censored spells was derived as the ending-age of a spell minus the predicted starting-age.

Furthermore, possible heterogeneity resulted from measurement error in duration or explanatory variables, or omitted regressors. Therefore, gamma heterogeneity was added to both exponential and Weibull regression models in order to correct for this possible bias of the estimates.

The signs of estimated coefficients of explanatory variables in the hazard rate equation for re-migration were consistent across equations. The life-cycle effect on the hazard rate of return migration is quadratic, revealed in a U-sharped form. At a younger age, the conditional probability of return migration decreases as age increases. After achieving the minimum, a older age increases the hazard rate of re-migration.

Education has a positive effect on the hazard rate of re-migration, but people having English proficiency and disability status are less likely to re-migrate. One additional year of schooling of a migrant augments the hazard rate by 2.1 to 4.9 percent. The marginal proportional effect of English or disability on the hazard rate ranges from -1.36 to -2.35 and -0.014 to -0.194 respectively.

The predicted job growth rate for an area has a stronger effect on the hazard rate of re-migration than its predicted unemployment rate. A 0.1 percentage point increase in predicted job growth rates for the U.S. mainland (for Puerto Rico) reduces (amplifies) the hazard rate by 15.4 to 32.7 (23.1 to 45.3) percent. An increase of 0.1 percentage point of the predicted unemployment rates for the mainland augments the likelihood of re-migration by 3.8 to 13.9 percent. When we evaluated at the sample mean of the real minimum wage for Puerto Rico for 1980–1990, the marginal effect of a 0.1 percentage point increase in the predicted unemployment rate for the island on the hazard rate ranges from −0.2 to −1.8 percent.

The estimated coefficient of the real minimum wage on the hazard rate of re-migration was consistent with previous finding (Castillo-Freeman and Freeman 1992) showing that the minimum wage policy for Puerto Rico resulted in massive job losses, and the unemployed persons tended to migrate to the U.S. mainland. A 10 cents increase in the real minimum wage for Puerto Rico diminishes the likelihood of return migration by 5.8 to 26.1 percent, evaluated at the sample mean of predicted unemployment rates for the island for 1980–1990.

The public policy implications of the estimated coefficients of the hazard function are as follow: (1) holding other things equal, when the federal government raises the minimum wage, the number of re-migrating Puerto Rico-born men will be reduced; (2) holding other things equal, if more states like Arizona or California enforce official English, more Puerto Rican-born males will decide to re-migrate.

Earning Functions

The Immigration Reform and Control Act of 1986 (IRCA) created two legalization programs, a general program and the Special Agricultural Worker (SAW) program, for illegal aliens to become temporary and then permanent residents of the United States.

More than 3 million individuals applied. Hispanic aliens were the major group, accounting for about 88 percent of total applicants. A total of 2.7 million aliens have been approved for temporary status under the IRCA provisions, of whom 88.9 percent were Hispanics. Since new workers legalized under the IRCA were largely Spanish speaking, we examined possible effects of IRCA on the real wage rate of Hispanics on the U.S. mainland.

The Chow tests for parameter equality of earning functions across pairs of ethnic groups suggested that the economic returns for personal characteristics provide a major source of wage differentials across ethnic groups. Of the personal characteristics, the most important part was the difference in returns to the human capital variables. This may reflect the fact that different ethnicities have different average education attainments or different average schooling quality. Furthermore, the other three categories of explanatory variables also played an important role in explaining wage differences across ethnic groups.

From the The Chow tests for parameter stability of wage functions for 1979 and 1989 for each ethnicity, we concluded that the cohort effect was the only part of the returns to personal characteristics having significantly changed for Hispanics; the coefficients of other personal characteristics were equal at the 50 percent significant level. This may imply that the decrease in Hispanic men's average real wage rates from 1979 to 1989 was produced by changes in local labor market characteristics. The result is consistent with our hypothesis that Hispanics suffered adverse labor market competition during the 1980s. For whites, structural change of earning function was concentrated in coefficients of human capital, ENG, cost of living and local amenities, local labor market characteristics, and regional dummies. The evidence for structural change for blacks is very similar to that for whites, except there is no significant change in the coefficients of local labor market characteristics.

APPENDIX A PROOF OF PROPOSITIONS AND DERIVATION OF COMPARATIVE STATISTICS

Proof of Proposition 2.1

We want to show that

$$(w_f + x_f) + \beta \max\{x_h - k_h, x_f\} > V_4^{h} \tag{A.1}$$

where $V_4^{h} = [(1 - p_h) w_h + x_h - k_h] + \beta x_h$. Suppose that it is not true. Then, the individual's presented value of life time utility in the host country at the beginning of the third period is

$$\begin{aligned}
& (w_f + x_f) + \beta \max\{[(1 - p_f)(w_f + x_f) + p_f x_f] + \beta \max\{x_h - k_h, x_f\}, V_4^{h}\} \\
= \; & (w_f + x_f) + \beta \max\{[(1 - p_f) w_f + x_f] + \beta \max\{x_h - k_h, x_f\}, V_4^{h}\} \\
\leq \; & (w_f + x_f) + \beta \max\{(w_f + x_f) + \beta \max\{x_h - k_h, x_f\}, V_4^{h}\} \\
= \; & (w_f + x_f) + \beta V_4^{h}.
\end{aligned} \tag{A.2}$$

The individual's presented value of life time utility for re-migration at the beginning of the third period is

$$\begin{aligned}
& (w_h + x_h - k_h) + \beta [((1 - p_h)w_h + x_h) + \beta x_h] \\
= \; & [x_h + x_h - (1 - \beta)k_h] + \beta V_4^{h}.
\end{aligned} \tag{A.3}$$

However, according to the fact,

$$\beta \max\{x_h - k_h, x_f\} \geq \beta (x_h - k_h),$$

we have

$$\begin{aligned} (1-p_h)w_h + x_h - (1-\beta)k_h &\geq w_f + x_f \\ \Longrightarrow \qquad w_h + x_h - (1-\beta)k_h &> w_f + x_f \end{aligned}$$

when equation (A.1) is not true. This implies that equation (A.3) is greater than equation (A.2), violating the assumption that the individual has lived in the host country during the third period. Q.E.D.

Derivation of Comparative Statistics

1. The net value of re-migration at the beginning of the fifth period is defined as:

$$RM_5 = V_5^h - V_5^f = (x_h - k_h) - x_f. \qquad (A.4)$$

Taking derivative of RM_5 with respect to k_h, and x_h,

$$\frac{\partial RM_5}{\partial k_h} = -1 < 0$$

$$\frac{\partial RM_5}{\partial x_h} = 1 > 0$$

It is clear that w_h, k_f, μ_x, μ_w, p_h, or p_f has no impact on RM_5. Hence, we have:

$$\frac{\partial RM_5}{\partial w_h} = \frac{\partial RM_5}{\partial k_f} = \frac{\partial RM_5}{\partial \mu_x} = \frac{\partial RM_5}{\partial \mu_w} = \frac{\partial RM_5}{\partial p_h} = \frac{\partial RM_5}{\partial p_f} = 0.$$

2. The net value of re-migration at the beginning of the fourth period is written as:

$$RM_4 = V_4^h - \left[x_f + \beta(x_h - k_h) + \beta(-RM_5)^+\right]. \qquad (A.5)$$

Taking derivative of RM_4 with respect to k_h, p_h, w_h, and x_h,

$$\frac{\partial RM_4}{\partial k_h} = \underbrace{-1-\beta}_{(-)} - \beta \underbrace{\frac{\partial(-RM_5)^+}{\partial k_h}}_{(-)} < 0$$

$$\frac{\partial RM_4}{\partial p_h} = -w_h < 0$$

$$\frac{\partial RM_4}{\partial w_h} = 1 - p_h > 0$$

$$\begin{aligned}\frac{\partial RM_4}{\partial x_h} &= (1+\beta) - \beta - \beta\frac{\partial(-RM_5)^+}{\partial x_h} \\ &= 1 - \beta\underbrace{\frac{\partial(-RM_5)^+}{\partial x_h}}_{(-)} > 0\end{aligned}$$

Because k_f, μ_x, μ_w, or p_f does not affect on RM_4,

$$\frac{\partial RM_4}{\partial k_f} = \frac{\partial RM_4}{\partial \mu_x} = \frac{\partial RM_4}{\partial \mu_w} = \frac{\partial RM_4}{\partial p_f} = 0.$$

3. The net value of re-migration at the beginning of the third period is:

$$\begin{aligned}RM_3 = {} & V_3^{*h} - \Big\{(w_f + x_f) + \beta\Big[(1-p_f)\big[(w_f + x_f) + \beta(x_h - k_h) + \beta(-RM_5)^+\big] \\ & - p_f V_4^{*h} + p_f(-RM_4)^+\Big]\Big\}\end{aligned} \tag{A.6}$$

Taking derivative of RM_3 with respect to p_h, k_h, p_f, x_h, and w_h,

$$\begin{aligned}\frac{\partial RM_3}{\partial p_h} &= (-\beta\, w_h) - \left[\beta^2(1-p_f)\frac{\partial(-RM_5)^+}{\partial p_h} - \beta\, p_f\, w_h + \beta\, p_f\frac{\partial(-RM_4)^+}{\partial p_h}\right] \\ &= \beta\, w_h\underbrace{(p_f - 1)}_{(-)} - \beta^2(1-p_f)\underbrace{\frac{\partial(-RM_5)^+}{\partial p_h}}_{=\,0} - \beta\, p_f\underbrace{\frac{\partial(-RM_4)^+}{\partial p_h}}_{(-)} < 0\end{aligned}$$

$$\begin{aligned}\frac{\partial RM_3}{\partial k_h} &= -1 - \Big[-\beta^2(1-p_f) + \beta^2(1-p_f)\frac{\partial(-RM_5)^+}{\partial k_h} + \beta\, p_f \\ &\quad - \beta\, p_f\frac{\partial(-RM_4)^+}{\partial k_h}\Big] \\ &= -1 + \underbrace{[p_f\,\beta - (1-p_f)\beta^2]}_{<\,\beta} - \beta^2(1-p_f)\underbrace{\frac{\partial(-RM_5)^+}{\partial k_h}}_{(-)} \\ &\quad - \beta\, p_f\underbrace{\frac{\partial(-RM_4)^+}{\partial k_h}}_{(-)} < 0\end{aligned}$$

$$\begin{aligned}
\frac{\partial RM_3}{\partial p_f} &= -\left\{-\beta\left[w_f + x_f + \beta(x_h - k_h) + \beta(-RM_5)^+\right] + \beta V_4^h + \beta(-RM_4)^+\right\} \\
&= \beta\overbrace{[(w_f + x_f) + \beta\max\{x_h - k_h,\, x_f\}]}^{(>\,V_4^h \text{ by Proposition 2.1})} \\
&\quad - \beta\max\left\{x_f + \beta\max\{x_h - k_h,\, x_f\},\, V_4^h\right\} > 0
\end{aligned}$$

$$\begin{aligned}
\frac{\partial RM_3}{\partial x_h} &= (1 + \beta + \beta^2) - \left\{\beta^2(1 - p_f) + \beta^2(1 - P_f)\,\frac{\partial(-RM_5)^+}{\partial x_h}\right. \\
&\quad \left. + p_f\,\beta(1 + \beta) - \beta\,p_f\,\frac{\partial(-RM_4)^+}{\partial x_h}\right\} \\
&= \underbrace{1 + \beta(1 - p_f)}_{(+)} - \beta^2(1 - p_f)\underbrace{\frac{\partial(-RM_5)^+}{\partial x_h}}_{(-)} - \beta\,p_f\underbrace{\frac{\partial(-RM_4)^+}{\partial x_h}}_{(-)} > 0
\end{aligned}$$

$$\begin{aligned}
\frac{\partial RM_3}{\partial w_h} &= [1 + \beta(1 - p_h)] - \left[\beta\,p_f(1 - p_h) + \beta\,p_f\,\frac{\partial(-RM_4)^+}{\partial w_h}\right] \\
&= \underbrace{1 + \beta(1 - p_h)(1 - p_f)}_{(+)} - \beta\,p_f\underbrace{\frac{\partial(-RM_4)^+}{\partial w_h}}_{(-)} > 0
\end{aligned}$$

Since k_f, μ_x, or μ_w, does not appear in RM_3,

$$\frac{\partial RM_3}{\partial k_f} = \frac{\partial RM_3}{\partial \mu_x} = \frac{\partial RM_3}{\partial \mu_w} = 0.$$

4. The net value of re-migration at the beginning of the second period is defined as:

$$\begin{aligned}
RM_2 &= V_2^h - \left[w_f + \beta V_3^h + \beta\int_{\mathbf{R}}(-RM_3)^+\,dG(x_f)\right] \\
&= w_h + x_h + (\beta - 1)k_f - w_f - \beta\int_{\mathbf{R}}(-RM_3)^+\,dG(x_f). \qquad \text{(A.7)}
\end{aligned}$$

Taking derivative of RM_2 with respect to k_h, x_h, w_h, p_f, and p_h,

$$\frac{\partial RM_2}{\partial k_h} = \underbrace{\beta - 1}_{(-)} - \beta\int_{\mathbf{R}}\underbrace{\frac{\partial(-RM_3)^+}{\partial k_h}}_{(+)}\,dG(x_f) < 0$$

$$\frac{\partial RM_2}{\partial x_h} = 1 - \beta\int_{\mathbf{R}}\underbrace{\frac{\partial(-RM_3)^+}{\partial x_h}}_{(-)}\,dG(x_f) > 0$$

$$\frac{\partial RM_2}{\partial w_h} = 1 - \beta \int_{\mathbf{R}} \underbrace{\frac{\partial(-RM_3)^+}{\partial w_h}}_{(-)} dG(x_f) > 0$$

$$\frac{\partial RM_2}{\partial p_f} = -\beta \int_{\mathbf{R}} \underbrace{\frac{\partial(-RM_3)^+}{\partial p_f}}_{(-)} dG(x_f) > 0$$

$$\frac{\partial RM_2}{\partial p_h} = -\beta \int_{\mathbf{R}} \underbrace{\frac{\partial(-RM_3)^+}{\partial p_h}}_{(+)} dG(x_f) < 0$$

Let $v = x_f - \mu_x$ with distribution function $G_v(v)$. Note that $G_v(v)$ does not depend on μ_x and $x_f = v + \mu_x$. Taking derivative of RM_2 with respect to μ_x,

$$\frac{\partial RM_2}{\partial \mu_x} = -\beta \int_{\mathbf{R}} \underbrace{\frac{\partial(-RM_3)^+}{\partial x_f}}_{(+)} \underbrace{\frac{\partial x_f}{\partial \mu_x}}_{=1} dG_v(v) < 0.$$

Because k_f or μ_w, does not appear in RM_2,

$$\frac{\partial RM_2}{\partial k_f} = \frac{\partial RM_2}{\partial \mu_w} = 0.$$

5. The net value of emigration at the beginning of the first period is written as:

$$\begin{aligned} M_1 &= V_1^{*f} - V_1^{*h} \\ &= -k_f + \beta \int_{\mathbf{R}_+} \left[(RM_2)^+ + V_2^{*f}\right] dF(w_f) - V_1^{*h} \\ &= -k_f + \beta V_2^{*h} + \beta \int_{\mathbf{R}_+} (-RM_2)^+ \, dF(w_f) - V_1^{*h} \\ &= -k_f + \beta \int_{\mathbf{R}_+} (-RM_2)^+ \, dF(w_f) - (x_h - w_h) - \beta \, k_h \end{aligned} \tag{A.8}$$

Taking derivative of M_1 with respect to k_f, k_h, x_h, w_h, p_h, p_f, and μ_x,

$$\frac{\partial M_1}{\partial k_f} = -1 < 0$$

$$\frac{\partial M_1}{\partial k_h} = \beta \int_{\mathbf{R}_+} \left[\frac{\partial(RM_2)^+}{\partial k_h} + \frac{\partial V_2^{*f}}{\partial k_h} V_2^{*f}\right] dF(w_f)$$

$$= \beta \int_{\mathbf{R}_+} \Big[\underbrace{\frac{\partial (RM_2)^+}{\partial k_h}}_{(-)} + \beta \int_{\mathbf{R}} \Big(\underbrace{\frac{\partial (RM_3)^+}{\partial k_h}}_{(-)} + \underbrace{\frac{\partial V_3^f}{\partial k_h}}_{(-)} \Big) \, dG(x_f) \Big] \, dF(w_f) \; < \; 0$$

because $$\frac{\partial V_3^f}{\partial k_h} = \beta^2 (1 - p_f) \underbrace{\frac{\partial (RM_5)^+}{\partial k_h}}_{(-)} + \beta p_f \underbrace{\frac{\partial (RM_4)^+}{\partial k_h}}_{(-)} + \beta^2 p_f \underbrace{\frac{\partial (RM_5)^+}{\partial k_h}}_{(-)} \; < \; 0$$

$$\frac{\partial M_1}{\partial x_h} = \beta \int_{\mathbf{R}_+} \underbrace{\frac{\partial (-RM_2)^+}{\partial x_h}}_{(-)} \, dF(w_f) - 1 \; < \; 0$$

$$\frac{\partial M_1}{\partial w_h} = \beta \int_{\mathbf{R}_+} \underbrace{\frac{\partial (-RM_2)^+}{\partial w_h}}_{(-)} \, dF(w_f) - 1 \; < \; 0$$

$$\frac{\partial M_1}{\partial p_h} = \beta \int_{\mathbf{R}_+} \underbrace{\frac{\partial (-RM_2)^+}{\partial p_h}}_{(+)} \, dF(w_f) \; > \; 0$$

$$\frac{\partial M_1}{\partial p_f} = \beta \int_{\mathbf{R}_+} \underbrace{\frac{\partial (-RM_2)^+}{\partial p_f}}_{(-)} \, dF(w_f) \; < \; 0$$

$$\frac{\partial M_1}{\partial \mu_x} = \beta \int_{\mathbf{R}_+} \underbrace{\frac{\partial (-RM_2)^+}{\partial \mu_x}}_{(+)} \, dF(w_f) \; > \; 0$$

Let $u = w_f - \mu_w$ with distribution function $F_u(u)$. Note that $F_u(u)$ does not depend on μ_w and $w_f = u + \mu_w$. Taking derivative of M_1 with respect to μ_w,

$$\frac{\partial M_1}{\partial \mu_w} = \beta \int_{\mathbf{R}_+} \underbrace{\frac{\partial (-RM_2)^+}{\partial w_f}}_{(+)} \underbrace{\frac{\partial w_f}{\partial \mu_w}}_{=1} \, dF_u(u) \; > \; 0.$$

Proof of Proposition 2.2

The reservation quality of life x_f^* is the quality of life in the foreign country such that $RM_3 = 0$. By the implicit function theory,

$$\frac{\partial x_f^*}{\partial a} = - \frac{\partial RM_3 \, / \, \partial a}{\partial RM_3 \, / \, \partial x_f^*}. \tag{A.9}$$

Taking derivative of RM_3 with respect to x_f^*

$$\begin{aligned}\frac{\partial RM_3}{\partial x_f^*} &= -\left\{1 + \beta\,(1-p_f) + \beta^2\,(1-p_f)\,\frac{\partial(-RM_5)^+}{\partial x_f^*} + \beta\,p_f\,\frac{\partial(-RM_4)^+}{\partial x_f^*}\right\} \\ &= \underbrace{-1-\beta\,(1-p_f)}_{(-)} - \beta^2\,(1-p_f)\,\underbrace{\frac{\partial(-RM_5)^+}{\partial x_f^*}}_{(+)} - \beta\,p_f\,\underbrace{\frac{\partial(-RM_4)^+}{\partial x_f^*}}_{(-)} < 0\end{aligned}$$

because both $\dfrac{\partial RM_5}{\partial x_f^*} = -1$ and $\dfrac{\partial RM_4}{\partial x_f^*} = \left[-1 - \beta\dfrac{\partial(-RM_5)^+}{\partial x_f^*}\right]$ are negative.

The reservation wage w_f^* is the wage rate in the foreign country such that $RM_2 = 0$. Based on the implicit function theory,

$$\frac{\partial w_f^*}{\partial b} = -\,\frac{\partial RM_2\;/\;\partial b}{\partial RM_2\;/\;\partial w_f^*}. \tag{A.10}$$

Taking derivative of RM_2 with respect to w_f^*

$$\frac{\partial RM_2}{\partial w_f^*} = -1 - \beta\int_{\mathbb{R}} \underbrace{\frac{\partial(-RM_3)^+}{\partial w_f^*}}_{(+)}\,dG(x_f) < 0$$

since $\dfrac{\partial RM_3}{\partial w_f^*} = -\left[\underbrace{1+\beta\,(1-p_f)}_{(+)} - \beta\,p_f\,\underbrace{\dfrac{\partial(-RM_4)^+}{\partial w_f^s}}_{=\,0}\right] < 0.$

Combining with the results of comparative statistics, we have the desired results. Q.E.D.

Proof of Proposition 2.3

RM_5 and $-RM_5$ are linear in x_f, so are both convex in x_f. Based on the fact that $\max\{f(x), g(x)\}$ is convex in x if $f(x)$ and $g(x)$ are convex in x, $(-RM_5)^+$ is convex in x_f. Furthermore, since

$$RM_4 = V_4^{*h} - \left[x_f - \beta(x_h - k_h) - \beta(-RM_5)^+\right] \tag{A.11}$$

is concave in x_f, both $-RM_4$ and $(-RM_4)^-$ are also convex in x_f. Similarly, $(-RM_3)^-$ is convex in x_f.

On the other hand, RM_3 is linear in w_f so it is both concave and convex in w_f. Hence, $-RM_3$ and $(-RM_3)^-$ are convex in w_f. By the same argument, $-RM_2$ and $(-RM_2)^-$ are convex in w_f. Applying the Theorem 2 of Lippman and McCall (1982), we have desired results:

$$
\begin{aligned}
R\hat{M}_2 &\leq RM_2, \text{ and} \\
\hat{M}_1 &\geq M_1.
\end{aligned}
$$

Q.E.D.

APPENDIX B PREDICTING LEFT-CENSORED SPELLS

Table B.1 Variable Definitions

Variable	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

Table B.1 (Continued)

Variable	Definition
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
$\hat{\lambda}$	Inverse of Mill's ratio

Table B.2 Estimated Coefficients of Binary Probit Equations for the Probability with Left-Censored Spells

Covariates[a]	Puerto Rico Sample		U.S. Sample	
AGE	4.494***	(4.32)[b]	−292.798***	(−5.38)
AGE^2	−14.112***	(−3.89)	986.201***	(5.35)
AGE^3	1.936***	(3.53)	−136.554***	(−5.32)
AGE^4	−0.098***	(−3.20)	6.762***	(5.29)
ED	0.059**	(2.10)	0.146***	(3.96)
ED^2	−0.559***	(−3.20)	−0.961***	(−4.33)
ENG	1.142***	(6.69)	0.284	(1.33)
EDENG	−0.055***	(2.96)	0.018	(0.62)
DISAB	0.133*	(1.79)	−0.047	(−0.54)
RACE	– –	– –	0.272***	(3.33)
INTERCEPT	−53.381***	(−4.88)	2795.617***	(5.38)

[a]The dependent variable is the dummy variable with the value equal to 1 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]t-values are in parentheses.

* P-value $\leq$ 0.1   ** P-value $\leq$ 0.05   *** P-value $\leq$ 0.01

Table B.3 Estimated Coefficients of Predicted Starting-Age Equations, Corrected for Sample-Selected Bias

Covariates[a]	Puerto Rico Sample		U.S. Sample	
ED	−0.019***	(−6.19)[b]	−0.039***	(−21.39)
ED^2	0.123***	(7.21)	0.213***	(22.01)
ENG	−0.732***	(−29.70)	−0.072***	(−5.38)
EDENG	0.047***	(22.03)	0.001	(0.47)
RACE	--	--	0.010**	(2.52)
DISAB	−0.035***	(−3.76)	0.072***	(11.78)
PUSEM	−0.234***	(−7.84)	−0.698***	(−7.13)
PUSUN	−0.141***	(−6.06)	0.110**	(2.19)
PPREM	0.085**	(2.35)	0.822)***	(10.17)
PPRUN	−0.143***	(−4.37)	0.426***	(9.02)
PRMIN	−3.412***	(−5.57)	7.787***	(10.62)
PRMINUN	0.211***	(5.34)	−0.486***	(−10.12)
RUSEM	−0.079***	(−4.35)	−0.038**	(−2.48)
RUSUN	−0.205***	(−5.58)	−0.054	(−0.69)
RPREM	0.010*	(1.80)	−0.274***	(−18.52)
RPRUN	0.050***	(4.48)	−0.491***	(−12.03)
$\hat{\lambda}$	1.028***	(90.65)	0.712***	(48.26)
INTERCEPT	6.560***	(11.59)	−4.435***	(−6.33)
R^2	0.8581		0.4350	
Adj. R^2	0.8568		0.4340	
N	1,736		9,189	

[a]The dependent variable is the natural logarithm of the starting-age of a spell.
[b]t-values are in parentheses.
* P-value $\leq$ 0.1 ** P-value $\leq$ 0.05 *** P-value $\leq$ 0.01

APPENDIX C SUMMARY OF AR MODELS FOR STATE JOB GROWTH RATES AND UNEMPLOYMENT RATES

Table C.1 State Job Growth Rates

State	Mean	Order	ϕ_1	ϕ_2	ϕ_3	ϕ_4
Alabama	2.281	2	0.472	−0.360	--	--
Arizona	4.860	2	0.645	−0.396	--	--
Arkansas	2.643	1	0.182	--	--	--
California	2.440	2	0.751	−0.645	--	--
Colorado	3.620	1	0.243	--	--	--
Connecticut	1.163	2	0.674	−0.469	--	--
Delaware	2.208	1	0.241	--	--	--
District of Columbia	0.766	3	0.617	−0.217	−0.331	--
Florida	4.298	2	0.627	−0.661	--	--
Georgia	3.106	1	-0.080	--	--	--
Idaho	3.244	1	0.545	--	--	--
Illinois	0.906	2	0.755	−0.470	--	--
Indiana	1.456	2	0.348	−0.329	--	--
Iowa	1.607	2	0.577	−0.316	--	--

Table C.1 (Continued)

State	Mean	Order	ϕ_1	ϕ_2	ϕ_3	ϕ_4
Kansas	2.114	1	0.390	--	--	--
Kentucky	2.339	1	0.386	--	--	--
Louisiana	1.856	1	0.474	--	--	--
Maine	1.896	1	0.373	--	--	--
Maryland	2.204	2	0.565	−0.329	--	--
Massachusetts	1.113	2	0.928	−0.614	--	--
Michigan	1.297	6[a]	0.349	−0.370	−0.137	−0.397
Minnesota	2.401	2	0.495	−0.556	--	--
Mississippi	2.412	2	0.703	−0.424	--	--
Missouri	1.549	2	0.429	−0.464	--	--
Montana	2.076	1	0.421	--	--	--
Nebraska	2.091	1	0.309	--	--	--
Nevada	5.390	2	0.761	−0.479	--	--
New Hampshire	2.807	2	0.570	−0.341	--	--
New Jersey	1.434	2	0.553	−0.338	--	--
New Mexico	3.180	1	0.326	--	--	--
New York	0.461	2	0.752	−0.508	--	--
North Carolina	2.785	2	0.190	−0.371	--	--

[a] $\phi_5 = 0.201$ and $\phi_6 = -0.496$

Table C.1 (Continued)

State	Mean	Order	ϕ_1	ϕ_2	ϕ_3	ϕ_4
North Dakota	2.469	5[a]	0.319	0.395	0.488	−0.093
Ohio	1.177	2	0.442	−0.431	--	--
Oklahoma	2.161	1	0.551	--	--	--
Oregon	2.690	2	0.799	−0.477	--	--
Pennsylvania	0.784	2	0.371	−0.331	--	--
Rhode Island	0.906	1	0.146	--	--	--
South Carolina	2.826	1	0.205	--	--	--
South Dakota	2.594	1	0.220	--	--	--
Tennessee	2.498	1	0.215	--	--	--
Texas	3.182	3	0.594	−0.365	0.361	--
Utah	3.477	1	0.309	--	--	--
Vermont	2.433	2	0.485	−0.306	--	--
Virginia	3.028	2	0.514	−0.359	--	--
Washington	2.946	2	0.809	−0.503	--	--
West Virginia	0.990	1	0.444	--	--	--
Wisconsin	1.992	2	0.439	−0.386	--	--
Wyoming	2.892	3	0.752	−0.405	0.421	--
United States	1.967	2	0.401	−0.373	--	--

[a] $\phi_5 = -0.503$

Table C.2 State Unemployment Rates

State	Mean	Order	ϕ_1	ϕ_2	ϕ_3	ϕ_4
Alabama	7.562	2	1.104	−0.330	- -	- -
Arizona	6.342	1	0.551	- -	- -	- -
Arkansas	7.031	1	0.728	- -	- -	- -
California	7.285	2	0.939	−0.527	- -	- -
Colorado	5.385	1	0.654	- -	- -	- -
Connecticut	5.800	3	1.274	−1.018	0.464	- -
Delaware	5.969	2	1.099	−0.353	- -	- -
District of Columbia	6.781	2	1.356	−0.550	- -	- -
Florida	6.189	2	0.893	−0.347	- -	- -
Georgia	5.700	1	0.684	- -	- -	
Idaho	6.439	1	0.696	- -	- -	- -
Illinois	6.662	1	0.793	- -	- -	- -
Indiana	6.477	1	0.761	- -	- -	- -
Iowa	4.769	2	1.133	−0.339	- -	- -
Kansas	4.404	2	0.627	−0.297	- -	- -
Kentucky	6.781	1	0.778	- -	- -	- -
Louisiana	8.131	2	1.128	−0.378	- -	- -
Maine	6.612	1	0.619	- -	- -	- -
Maryland	5.304	2	0.986	−0.367	- -	- -

Table C.2 (Continued)

State	Mean	Order	ϕ_1	ϕ_2	ϕ_3	ϕ_4
Massachusetts	6.262	2	1.047	−0.474	--	--
Michigan	8.800	2	0.867	−0.276	--	--
Minnesota	5.504	1	0.659	--	--	--
Mississippi	7.485	1	0.820	--	--	--
Missouri	5.885	1	0.710	--	--	--
Montana	6.439	1	0.648	--	--	--
Nebraska	3.573	1	0.709	--	--	--
Nevada	6.681	2	0.929	−0.473	--	--
New Hampshire	4.735	2	0.820	−0.338	--	--
New Jersey	6.542	2	1.165	−0.518	--	--
New Mexico	7.446	1	0.611	--	--	--
New York	6.739	2	1.073	−0.485	--	--
North Carolina	5.265	1	0.616	--	--	
North Dakota	4.627	1	0.607	--	--	--
Ohio	6.869	2	0.965	−0.321	--	--
Oklahoma	5.396	1	0.648	--	--	--
Oregon	7.408	2	1.077	−0.471	--	--
Pennsylvania	6.669	2	1.084	−0.392	--	--

Table C.2 (Continued)

State	Mean	Order	ϕ_1	ϕ_2	ϕ_3	ϕ_4
Rhode Island	6.473	1	0.587	--	--	--
South Carolina	6.235	2	0.928	−0.359	--	--
South Dakota	3.835	1	0.750	--	--	--
Tennessee	6.396	1	0.779	--	--	--
Texas	5.739	1	0.769	--	--	--
Utah	5.677	1	0.649	--	--	--
Vermont	5.646	2	0.981	−0.443	--	--
Virginia	4.831	2	0.895	−0.333	--	--
Washington	7.900	2	0.931	−0.460	--	--
West Virginia	7.746	4	0.480	−0.073	−0.574	0.338
Wisconsin	5.681	1	0.747	--	--	--
Wyoming	5.119	1	0.795	--	--	--
United States	6.473	2	0.901	−0.350	--	--

APPENDIX D DETAILED DATA

Table D.1 Job Growth Rates and Unemployment Rates for the U.S. mainland and Puerto Rico, Minimum Wages on Puerto Rico, and Puerto Rico Deflator

Year	Job Growth Rate[a]		Unemployment Rate[a]		Minimum Wage for Puerto Rico[b]	Puerto Rico Deflator[c]
	U.S.	P.R.	U.S.	P.R.		
1946	4.50	--	3.9	--	0.280	--
1947	4.53	--	3.9	--	0.280	--
1948	2.23	--	3.8	--	0.320	0.887
1949	−1.18	--	5.9	--	0.320	0.873
1950	2.24	--	5.3	12.8	0.328	0.859
1951	1.72	--	3.3	--	0.336	0.881
1952	0.41	--	3.1	--	0.344	0.953
1953	1.48	--	2.9	--	0.352	0.970
1954	−1.71	--	5.6	--	0.361	1.00

[a]Data from Statistical Abstract of the United States (various years)

[b]Data for 1950–1987 adapted from Castillo-Freeman and Freeman (1992) and data before 1950 from Investigation of Minimum Wages and education in Puerto Rico and the Vigin Islands (1950)

[c]Data for 1950–1987 adapted from Castillo-Freeman and Freeman (1992) and others from Informe economico al gobernador (various years)

Table D.1 (Continued)

Year	Job Growth Rate U.S.	Job Growth Rate P.R.	Unemployment Rate U.S.	Unemployment Rate P.R.	Minimum Wage for Puerto Rico	Puerto Rico Deflator
1955	3.32	--	4.4	15.2	0.369	1.003
1956	2.76	3.46	4.2	13.2	0.447	1.011
1957	0.47	−1.08	4.3	13.2	0.488	1.035
1958	−1.62	0.54	6.8	12.9	0.555	1.089
1959	2.49	−1.64	5.5	14.1	0.588	1.110
1960	1.66	2.17	5.6	12.1	0.616	1.138
1961	0.17	1.07	6.7	11.7	0.608	1.173
1962	1.56	3.66	5.6	12.5	0.707	1.216
1963	−0.12	3.53	5.7	12.8	0.723	1.247
1964	2.25	8.08	5.2	10.9	0.809	1.298
1965	2.54	4.61	4.5	10.9	0.834	1.327
1966	2.51	4.13	3.8	11.3	0.854	1.358
1967	2.01	−4.71	3.8	12.6	0.971	1.421
1968	2.06	2.46	3.6	12.1	1.104	1.500
1969	2.58	2.95	3.5	10.6	1.149	1.552
1970	0.99	−3.96	4.9	10.7	1.209	1.616
1971	0.56	8.43	5.9	11.6	1.224	1.708
1972	3.21	3.39	5.6	12.3	1.257	1.780

Table D.1 (Continued)

Year	Job Growth Rate		Unemployment Rate		Minimum Wage for Puerto Rico	Puerto Rico Deflator
	U.S.	P.R.	U.S.	P.R.		
1973	3.26	3.65	4.9	12.1	1.262	1.817
1974	1.79	−5.98	5.6	13.3	1.681	1.946
1975	−0.11	−6.50	8.5	18.2	1.871	2.082
1976	1.89	2.62	7.7	19.6	2.034	2.174
1977	3.44	2.16	7.0	20.0	2.198	2.240
1978	5.90	5.95	6.1	18.1	2.509	2.340
1979	2.85	−6.62	5.8	17.0	2.768	2.483
1980	0.48	1.73	7.1	17.0	2.997	2.716
1981	1.10	−2.13	7.6	19.8	3.264	2.954
1982	−0.87	−5.26	9.7	22.8	3.305	3.175
1983	1.31	2.53	9.6	23.4	3.350	3.321
1984	4.05	5.00	7.5	20.7	3.350	3.461
1985	2.02	2.22	7.2	21.8	3.350	3.548
1986	2.26	7.21	7.0	18.9	3.350	3.697
1987	2.56	5.37	6.2	16.8	3.350	3.787
1988	2.22	6.38	5.5	15.0	3.350	4.010
1989	2.04	1.48	5.3	14.6	3.350	4.149
1990	0.49	1.98	5.5	14.2	3.800	4.373

Table D.2 State Percentage of Urban Population. Land Prices. and Temperature

State	Urban Population (Percentage)		Land Prices ($ /acre)		Average Temp. of January (Degree F.)
	1980	1990	1979	1989	
Alabama	60.0	60.4	515	832	50.8
Arizona	83.8	87.5	134	276	52.3
Arkansas	51.6	53.5	691	784	39.9
California	91.3	92.6	936	1670	56.0
Colorado	80.6	82.4	332	369	29.5
Connecticut	78.8	79.1	2158	4463	25.2
Delaware	70.6	73.0	1725	2065	31.2
District of Columbia	100.0	100.0	1463	1969	35.2
Florida	84.3	84.8	930	1897	53.2
Georgia	62.4	63.2	609	1003	41.9
Idaho	54.0	57.4	485	601	29.9
Illinois	83.3	84.6	1786	1388	21.4
Indiana	64.2	64.9	1498	1251	26.0
Iowa	58.6	60.6	1458	1108	18.6
Kansas	66.7	69.1	437	438	29.6
Kentucky	50.9	51.8	792	923	32.5

Table D.2 (Continued)

State	Urban Population (Percentage) 1980	1990	Land Prices ($ /acre) 1979	1989	Average Temp. of January (Degree F.)
Louisiana	68.6	68.1	763	959	52.4
Maine	47.5	44.6	485	1029	21.5
Maryland	80.3	81.3	1799	2487	32.7
Massachusetts	83.8	84.3	1366	3802	29.6
Michigan	70.7	70.5	955	1000	23.4
Minnesota	66.9	69.9	854	749	11.2
Mississippi	47.3	47.1	520	718	45.7
Missouri	68.1	68.7	674	678	25.9
Montana	52.9	52.5	186	209	18.7
Nebraska	62.9	66.1	470	526	20.2
Nevada	85.3	88.3	104	234	32.2
New Hampshire	52.2	51.0	802	2260	19.9
New Jersey	89.0	89.4	2222	4644	31.8
New Mexico	72.1	73.0	100	193	34.8
New York	84.6	84.3	642	1053	21.1
North Carolina	48.0	50.4	819	1339	40.5
North Dakota	48.8	53.3	306	329	6.7

Table D.2 (Continued)

State	Urban Population (Percentage) 1980	1990	Land Prices ($ /acre) 1979	1989	Average Temp. of January (Degree F.)
Ohio	73.3	74.1	1516	1271	28.9
Oklahoma	67.3	67.7	442	523	35.9
Oregon	67.9	70.5	330	542	38.9
Pennsylvania	69.3	68.9	1245	1911	31.2
Rhode Island	87.0	86.0	2133	5080	28.2
South Carolina	54.1	54.6	582	949	44.7
South Dakota	46.4	50.0	257	293	12.4
Tennessee	60.4	60.9	669	1021	39.6
Texas	79.6	80.3	354	517	44.0
Utah	84.4	87.0	265	425	28.6
Vermont	33.8	32.2	657	1203	16.6
Virginia	66.0	69.4	864	1354	39.9
Washington	73.5	76.4	586	769	39.1
West Virginia	36.2	36.1	472	716	32.9
Wisconsin	64.2	65.7	807	867	18.7
Wyoming	62.7	65.0	119	143	26.1
United States	73.7	75.2	559	667	32.2

Source: Statistical Abstract of the United States (various years)

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