Why So Little Trade: Externalities, Asymmetry of Information, and Welfare

by

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Submitted to the Department of Economics
in partial fulfillment of the requirements for the degree of

Doctor of Philosophy

at the

MASSACHUSETTS INSTITUTE OF TECHNOLOGY

May 1997
[June 1997]

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Abstract

Chapter 1 presents two simple unified approaches to the classical problem of the impossibility of implementation without an external subsidy of ex-post efficient trade when agents have private information about their valuations of a tradeable good. The first approach consists in *directly* integrateing the Bayesian IC constraint to get an easily interpretable expression for the required subsidy, which relates to the implementation in dominant strategies. The second, shows that dominant strategy implementation requires an external subsidy, and that this is necessarily the same as for Bayesian implementation. The classical result by Myerson and Satterthwaite (1983) becomes a corollary of the much earlier result by Vickrey (1961).

Chapter 2 provides a simple and easy method to prove Arrow's impossibility theorem and some of its extensions. The procedure starts by easily showing neutrality over the set of strict-preference profiles. Then, it is almost immediate to show that a pivotal voter is a dictator.

Chapter 3 develops a model of double matching in the labor market and the social environment in order to explain the different behaviors, emerging in communities, toward migration in response to local economic shocks. This approach does not explain the different behaviors of the workers in the US and in Europe in an *exogenous* manner by introducing regulations or unemployment benefits. but rather explains it in an *endogenous* way (showing the existence of multiple equilibria). On the other hand, different policies about unemployment benefits can be the result ex post of these different equilibria.

The model is estimated for the US using the PSID. Both economic variables and social variables are significant. The data show that the model is useful to explain the different migration behavior between Blacks and Whites, as well as across macro regions.

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Acknowledgments

I thank my supervisors Olivier Blanchard and Lones Smith for their encouragement and comments. I also want to thank Antonio Spilimbergo with whom I wrote chapter 3.

I would also like to acknowledge the advice, encouragement, or teachings of the following professors: Daron Acemoglu, Josh Angrist, Susan Athey, Jushan Bai, Abhijit Banerjee, Alberto Bisin, Ricardo Caballero, Peter Diamond, Rudiger Dornbusch, Richard Dudley, Glenn Ellison, Stan Fischer, Frank Fisher, David Genesove, Jon Gruber, Jeffrey Harris, Jerry Hausman, Bengt Holmström, Paul Joskow, Michael Kremer, Paul Krugman, Whitney Newey, Thomas Piketty, Michael Piore, Steve Pischke, Jim Poterba, Nancy Rose, Richard Schmalensee, Peter Temin, Jaume Ventura, William Wheaton, Dan McFadden, Drew Fudenberg, Oliver Hart, Jean Tirole, Paul Ruud, Jean-Luc Vila, Roland Benabou, Bob Solow, John Heaton, Chi Fu Huang, Andy Lo, David Scharfstein, Vijay Krishna, Tomas Sjöström, Eric Maskin, John Dinardo, Dick Eckaus, Martin Helwig, and John Geanakoplos.

I am also grateful to Montse Luengo, Kevin Lee, Chun Ping Liao, Yulin and Yimin Li, Yha and Jie Wei, Vlad Bina, Vincent Chow, Edson Miyamoto, and Kuo Shen Chen. They have made my life nicer and more interesting during these years at MIT.

I want to thank my classmates and other students at the economics department, from whom I think I have learned a lot.

Finally, I want to aknowledge the help I have received from staff members of the Department of Economics, as well as throughout MIT and Harvard.

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Chapter 1

Impossibility of Efficient Trade under Asymmetry of Information

1.1 Introduction.

At least since Vickrey's (1961) classical paper on auctions, economists have realized that when agents have asymmetric private information, mutually benefitial (efficient) trade may not take place. Asymmetry of information constitutes a natural barrier to efficient trade, which adds to the more classical barriers posed by governments.

Vickrey (1961) analyzed how private information limits efficient trade using the game-theoretic concept of dominant-strategy equilibrium. 22 years later Myerson and Satterthwaite reconsidered this issue using a different equilibrium concept: Bayes equilibrium. Despite the conceptual proximity of both results, the arguments used to establish them were far away, making it difficult to understand the source of the problem.

In this paper I present two proves of the Bayesian result. Both draw very close to the dominant-strategy result. The first proof directly shows the Bayesian result. Although it does not use the dominant-strategy result, it shows their intimate connection. Moreover, it provides a measure of the amount of inefficiency due to the asymmetry of information.

The second proof is conceptually more interesting, and considerably less involved algebraically, as it goes mainly just in words. It first shows the dominant-strategy result, and then shows how the Bayesian result immediately follows from the dominant-strategy result. This way, Myerson-Satterthwaite (1983) is a corollary of Vickrey's (1961) old result. This proof shows how close both results are: not only conceptually but formally, and that the results are really easy to obtain.

1.2 *Two* theorems on the impossibility of ex-post efficient trade.

There are two agents: a seller s who owns a unit of a good, and a buyer b. Each agent $i \in \{s, b\}$ has private valuation (type) θ_i which is common knowledge that is independently distributed on the interval $[\underline{\theta_i}, \overline{\theta_i}]$ with continuous distribution function F_i and density f_i, strictly positive on $(\underline{\theta_i}, \overline{\theta_i})$. We assume $\cap_i(\underline{\theta_i}, \overline{\theta_i}) \neq \emptyset$. Agent i has quasilinear VNM utility $u(x_i, y_i, \theta_i) = x_i\theta_i + y_i$, for an allocation which gives agent i x_i units of the object and a transfer of y_i. Each agent i has the option of not participating in trade and then getting $u_i^0 = e_i\theta_i$, where e_i is the initial endowment ($e_s = 1, e_b = 0$). Trade is ex post efficient if $\forall i$, $x_i = 1$ whenever $\theta_i > \theta_{-i}$.

Theorem DS. (Vickrey.) There is no individually rational mechanism that implements in dominant strategies ex-post efficient trade without an external subsidy.

Theorem B. (Myerson and Satterthwaite.) There is no individually rational mechanism that Bayesian implements ex-post efficient trade without an external subsidy.

1.3 Two unifying approaches.

This section presents shows the intimate relationship between theorem B and theorem DS. We do this task in the form of two formal proves of theorem B. The first proof does not use theorem DS, but finds an expression for the required expected subsidy for Bayesian implementation of ex-post efficient trade which is well apparent that it is as well the expected external subsidy required for implementation in dominant strategies. The proof also provides us with a measure of the magnitude of the problem in relation with the total expected gains of trade.

The second proof first presents a standard argument that proves theorem DS. Then it points out that theorem B follows immediately from theorem DS because the Bayesian IC constraint uniquely determines the slope of the expected revenue curves, which together with binding individual rationality constraints implies that the minimal external subsidy in Bayesian implementation is the same as the expectation of the minimal external subsidy for implementation in dominant strategies. This proof gives the clearest single argument to prove both impossibility theorems. Formally it makes theorem B an immediate corollary of the much older result in theorem DS. Besides of the advantage of an essentially unique argument for both results, the proof is considerably simpler than the standard proof of theorem B, which is so pervassive in advanced textbooks of microeconomics and game theory, a proof that does not make any connection with the theorem DS despite its conceptual proximity.

First proof of Theorem B. Without loss of generality we can restrict attention to direct revelation mechanisms. Let $Y_i(\theta_i)$ be the expected transfer for type θ_i. Incentive compatibility (IC) requires

$$\forall i, \forall \theta_i, \theta_i' \in [\underline{\theta_i}, \overline{\theta_i}], \quad F_{-i}(\theta_i)\theta_i + Y_i(\theta_i) \geq F_{-i}(\theta_i')\theta_i + Y_i(\theta_i'). \tag{1.1}$$

Thus,

$$-\theta_i[F_{-i}(\theta_i) - F_{-i}(\theta_i')] \leq Y_i(\theta_i) - Y_i(\theta_i') \leq -\theta_i'[F_{-i}(\theta_i) - F_{-i}(\theta_i')]. \tag{1.2}$$

Divide all across by $\theta_i - \theta_i'$ and let $\theta_i' \to \theta_i$, to get for all i and θ_i,

$$Y_i'(\theta_i) = -\theta_i f_{-i}(\theta_i). \tag{1.3}$$

Individual rationality (IR) requires $\forall i$, $\forall \theta_i$, $u_i(\theta_i) \geq u_i^0$. So,

$$Y_b(\underline{\theta_b}) \geq -F_s(\underline{\theta_b})\underline{\theta_b}, \;\; Y_s(\overline{\theta_s}) \geq (1 - F_b(\overline{\theta_s}))\overline{\theta_s}. \tag{1.4}$$

Then, the required outside subsidy S,

$$S \equiv E\sum_i Y_i(\tilde{\theta}_i) = E\,[\,Y_s(\overline{\theta_s}) - \int_{\tilde{\theta}_s}^{\overline{\theta_s}} -tf_b(t)dt + Y_b(\underline{\theta_b}) + \int_{\underline{\theta_b}}^{\tilde{\theta}_b} -tf_s(t)dt\,] \tag{1.5}$$

$$\geq E\,[\,(1 - F_b(\overline{\theta_s}))\overline{\theta_s} + E\tilde{\theta}_b 1_{\{\overline{\theta_s} > \tilde{\theta}_b > \tilde{\theta}_s\}} - F_s(\underline{\theta_b})\underline{\theta_b} - E\tilde{\theta}_s 1_{\{\tilde{\theta}_b > \tilde{\theta}_s > \underline{\theta_b}\}}\,] \tag{1.6}$$

$$= E(\tilde{\theta}_b \wedge \overline{\theta_s} - \tilde{\theta}_s \vee \underline{\theta_b}) 1_{(\tilde{\theta}_b > \tilde{\theta}_s)} > 0. \tag{1.7}$$

□

Before proceeding to the second proof, we make three remarks. First, note that (1.4) is equivalent to IR, because, by (1.3), $\frac{\partial}{\partial\theta_i}[u_i(\theta_i) - u_i^0] = F_{-i}(\theta_i) + f_{-i}(\theta_i)\theta_i + Y_i(\theta_i) - e_i = F_{-i}(\theta_i) - e_i \leq 0$ for $i = s$ and ≥ 0 for $i = b$. Second, the minimal expected subsidy required to Bayesian implement IR and ex post efficient trade, S^*, is then obtained for (1.4) held as equalities, and it is equal to (1.7). But this is exactly the same as the minimal required subsidy for implementation in dominant strategies, which may be seen as a generalized second price auction where each party pays (or is paid) the value in her support which is closest to the value of the other party, i.e., the buyer pays $\tilde{\theta}_s \vee \underline{\theta_b}$ and the seller receives $\tilde{\theta}_b \wedge \overline{\theta_s}$, whenever there is trade.

Finally, note that in the case that the suppports of both distributions of types coincide, $\tilde{\theta}_b \wedge \overline{\theta_s} = \tilde{\theta}_b$ and $\tilde{\theta}_s \vee \underline{\theta_b} = \tilde{\theta}_s$, so that,

$$S^* \equiv \min S = E(\tilde{\theta}_b - \tilde{\theta}_s) 1_{(\tilde{\theta}_b > \tilde{\theta}_s)} \equiv G. \tag{1.8}$$

i.e., Bayesian implementation of efficient trade requires an external subsidy equal to the maximum expected gains from trade.

Our second proof is less involved algebraically. It starts by showing (as in Vickrey (1961)) that a positive subsidy is required for implementation in dominant strategies—i.e., it proves theorem DS. Then it easily shows that the same subsidy (in expectation) is required for Bayesian implementation.

Second Proof of theorem B. Consider the minimum expected subsidy that allows for implementation in dominant strategies. When there is trade, in order for truth revelation to be dominant strategy, the transfer $y_i(\theta_s, \theta_b)$ for agent i can only depend on the other agent's type. And in order to minimize the external subsidy and yet have all types participate, the buyer will pay $\theta_s \vee \underline{\theta}_b$, and the seller will receive $\theta_b \wedge \overline{\theta}_s$. When there is not trade, the transfers will be zero. The resulting subsidy is clearly positive as the amount paid to the seller is always larger than the amount paid by the buyer. Obviously, this mechanism also implements effcient trade as a Bayesian equilibrium. The Bayesian IR constraint (1.4) is binding for the buyer with lowest valuation and for the seller with largest valuation.

On the other hand, in Bayesian implementation, the IC condition uniquely determines the slope of the expected transfers (as in (1.3)). This together with binding IR constraints (for subsidy minimization), uniquely determines the expected transfers for Bayesian implementation. Thus the same positive expected subsidy is required for Bayesian implementation. □

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Chapter 2

Social Choice Theory with Proofs at a Glance

We provide a simple and easy method to prove Arrow's impossibility theorem and some of its extensions. The procedure starts by easily showing neutrality over the set of strict-preference profiles. Then, it is almost immediate to show that a pivotal voter is a dictator.

2.1 Arrow's Impossibility Theorem

We first state and prove Arrow's general possibility theorem for social welfare functions. We easily prove it by first showing neutrality, from which it will shortly follow that a pivotal voter is a dictator.

Let A be a set of at least 3 alternatives. R_A is the set of complete and transitive binary relations (preferences) on A, and L_A, the subset of antisymmetric relations (strict preferences). Let $N = \{1, \ldots, n\}$. R_A^N is the set of functions from N into R_A (preference profiles). L_A^N is defined similarly. We'll henceforth assume $a, b, c, x, y, z \in A$, $\rho, \sigma \in R_A$, $p, q, r, s \in R_A^N$, and $i, j \in N$. $a\rho^{>}b$ means not $b\rho a$. $\rho_{ab} = \sigma_{xy}$ means that $(a\rho b \Leftrightarrow x\sigma y)$ & $(b\rho a \Leftrightarrow y\sigma x)$; similarly, $r_{ab} = q_{xy}$ means that $\forall i$, $r(i)_{ab} = q(i)_{xy}$.

A social welfare function (SWF) for the society N over choice set A is a function from R_A^N into R_A. A SWF f is independent of irrelevant alternatives (IIA) if

$\forall p, q, a, b,\ p_{ab} = q_{ab} \Rightarrow f(p)_{ab} = f(q)_{ab}$; Pareto (P), if $\forall p, a, b$, such that $\forall i\ ap(i)^{>}b$, $af(p)^{>}b$; dictatorial (D), if $\exists i$ such that $\forall p, a, b\ \ ap(i)^{>}b \Rightarrow af(p)^{>}b$; neutral on L_A^N (N), if $\forall a \neq b, x \neq y, \forall r, q \in L_A^N,\ \ r_{ab} = q_{xy} \Rightarrow f(r)_{ab} = f(q)_{xy}$. Since we can get from a, b to x, y by replacing only one alternative at a time, it is clear that f is neutral on L_A^N iff for all distinct a, b, c and $\forall r, q \in L_A^N,\ \ r_{ab} = q_{ac} \Rightarrow f(r)_{ab} = f(q)_{ac}$.

Theorem 1 (Arrow). Let f be an IIA and P social welfare function for a finite society N over a choice set A with at least three alternatives, then f is dictatorial.

The proof first shows neutrality on L_A^N and then dictatorship. [1]

Proof. (i) Let $a \neq b \neq c \neq a$, and $r, q \in L_A^N$ with $r_{ab} = q_{ac}$. We show $f(r)_{ab} = f(q)_{ac}$, i.e., f is neutral on L_A^N. Say, $af(r)b$. Take $p \in L_A^N$ so that $\forall i,\ ar(i)b \Rightarrow ap(i)bp(i)c$ and $br(i)a \Rightarrow bp(i)c \,\&\, bp(i)a$. Then $af(p)bf(p)c$, so $af(p)c$. Thus, $af(q)c$. (ii) Let $a \neq b$. $\forall l \in \{0, \ldots, n\}$ take $q^l \in L_A^N$ with $aq^l(j)b \Leftrightarrow j \leq l$. Let $D = \{l \in \{0, \ldots, n\} : af(q^l)b\}$. By P, $n \in D$ and $i \equiv min(D) > 0$. For any s, x, y with $xs(i)^{>}y$, take p and $z \notin \{x, y\}$ such that: $p_{xy} = s_{xy}$, $p_{zy} = q^i_{ab}$, and $p_{zx} = q^{i-1}_{ab}$. Then $xf(p)^{>}zf(p)y$, so $xf(p)^{>}y$. Thus $xf(s)^{>}y$. □

2.2 Impossibility Theorems without Pareto

We provide simple proofs of two Wilson's impossibility theorems. These theorems extend Arrow's result by first weakening and then completely eliminating the Pareto principle in order to study the effect of IIA alone.

For $\rho \in R_A$, $x\rho^{=}y$ means $x\rho y\rho x$, while $I_\rho(x) \equiv \{z : z\rho^{=}x\}$, and $-\rho$ is the inverse of ρ, i.e., $x(-\rho)y \Leftrightarrow y\rho x$. A SWF f is null, if $\forall p, x, y,\ \ xf(p)^{=}y$; CS, if $\forall x, y\ \ \exists p$ such that $xf(p)y$.[2]

Theorem 2 (Wilson). Let f be a non-null, IIA, and CS SWF for a finite society N

[1] (i) also implies monotonicity and $f(L_A^N) \subseteq L_A$ (a nondesirable property) but we don't use them in the proof.

[2] This is a weak version of citizens' sovereignty.

over a choice set A with at least 3 alternatives, then either f or $-f$ is dictatorial.

The proof starts again by showing neutrality on L_A^N. Then it shows that $f(L_A^N) \subseteq L_A$. These two facts imply that either f or $-f$ is Pareto, so that Arrow's theorem applies.

Proof. (i) Let $a \neq b \neq c \neq a$, and $r, q \in L_A^N$ with $r_{ab} = q_{ac}$. We show $f(r)_{ab} = f(q)_{ac}$, i.e., f is N on L_A^N. Say, $af(r)b$. By CS, $\exists p$ such that $bf(p)c$. Take s so that $s_{bc} = p_{bc}$ and $s_{ab} = r_{ab}$. By IIA, $af(s)bf(s)c$. So $af(s)c$ and $af(q)c$. (ii) Note that in (i), if $bf(p)^{>}c$ then $af(q)^{>}c$. By f non-null, $\exists y \neq z$, & p' with $yf(p')^{>}z$. Thus, by passing through (y, z) in a replacement chain, we get $af(r)^{>}b$. So, $f(L_A^N) \subseteq L_A$. (iii) By (ii) and (i), either f or $-f$ is P, hence dictatorial. □

The next theorem drops CS, keeping only IIA.

Theorem 3 (Wilson). Let f be an IIA SWF for a finite society N over a choice set A. Let ρ be the binary relation on A defined by $x\rho y \Leftrightarrow \exists q : xf(q)y$. Then: (i) $\rho \in R_A$. (ii) For any x, if $I_\rho(x)$ has at least three elements, then $f_{I_\rho(x)}$ or $-f_{I_\rho(x)}$ is dictatorial or null. Where $f_{I_\rho(x)}$ is the restriction of f to $R^N_{I_\rho(x)}$.

Proof. Clearly, ρ is complete. If for some p, q, x, y, z, $xf(p)y$ and $yf(q)z$, take s with $s_{xy} = p_{xy}$ and $s_{yz} = q_{yz}$ to get $xf(s)z$. So, ρ is transitive. (ii) is theorem 2 applied to $f_{I_\rho(x)}$. □

2.3 Arrow's Impossibility Theorems for 3-Free-Triple Domains

In this section we consider the three-free-triple domain restriction introduced by Kelly (1994).

For any r, x, y, z, let $r_{\{xyz\}}$ be the restriction of r to $\{x, y, z\}$. We define the domain $D_3(A, N) \subseteq R_A^N$ by $D_3(A, N) \equiv \{r \in R_A^N : \forall x, y, z, |\{r(i)_{\{xyz\}} : i \in N\}| \leq 3\}$.

For any $a, b, x, y, \rho, \sigma, r, q$, $\rho_{ab} \geq \sigma_{xy}$ means that $(x\sigma y \Rightarrow a\rho b)$ & $(x\sigma^{>}y \Rightarrow a\rho^{>}b)$;

similarly, $r_{ab} \geq q_{xy}$ means that $\forall i,\ r(i)_{ab} \geq q(i)_{xy}$. Finally, $r_{ab} = \mathbf{1}$ means $\forall i, ar(i)^{>}b$.[3]

For any $D \subseteq R_A^N$, a (generalized) SWF is a function from D into R_A. D is the domain of f. We'll assume in the remainder of this section that $p, q, r, s \in D$, and $D_3(A, N) \subseteq D$. Note that this does not impose any restriction on orderings of pairs of alternatives in any profile in D.

An IIA SWF f is monotonic on L_A^N if $\forall r \in L_A^N$ and $\forall q, a, b,\ q_{ab} \geq r_{ab} \Rightarrow f(q)_{ab} \geq f(r)_{ab}$. Note that f is neutral and monotonic on L_A^N (N&M) if $\forall r \in L_A^N$ and $\forall q, a \neq b \neq c \neq a$ with $q_{ac} \geq r_{ab},\ af(r)b \Rightarrow af(q)^{>}c$.

The following is a generalization of Arrow's impossibility theorem.

Theorem 1' (Kelly). Let f be an IIA and P SWF for a finite society N over a choice set A with at least three alternatives. If the domain of f includes $D_3(A, N)$ then f is dictatorial.

Proof. We need to slightly modify the prove of Theorem 1. Notice that (i) proves also monotonicity on L_A^N if we take $q_{ac} \geq r_{ab}$ and only for one i $q(i)_{ac} \neq r(i)_{ab}$. In part (ii) take s with $\forall j \neq i, ys(j)^{>}x$ to conclude $xf(s)^{>}y$. Thus, by monotonicity, i is a dictator. □

Theorem 2 extends as well to this domain restriction. However theorem 3 does not.

Theorem 2'. Let f be a non-null, IIA, and CS SWF for a finite society N over a choice set A with $|A| \geq 3$. If the domain of f includes $D_3(A, N)$ then either f or $-f$ is dictatorial.

Proof. In proof of Theorem 2, replace L_A^N by its diagonal, i.e., $\{r \in L_A^N : \forall i, j,\ r(i) = r(j)\}$. □

[3] In other words, $\rho_{ab} = 1$ if $a\rho^{>}b$, $\rho_{ab} = -1$ if $b\rho^{>}a$, and $\rho_{ab} = 0$ otherwise. Then, $r_{ab} = (r(1)_{ab}, \ldots, r(n)_{ab})$. Finally, $r_{ab} \geq q_{xy}$ iff $\forall i, r(i)_{ab} \geq q(i)_{xy}$.

2.4 Gibbard-Satterthwaite Impossibility Theorem for Non-Manipulable Social Choice Functions

We now state and prove a fundamental theorem on the strategic aspects of social choice.

In this section we assume $p, q, r, s \in L_A^N$. For any $B \subseteq A$ and p, define $p^B \in L_A^N$ by $\forall b, b' \in B$, $\forall a, a' \in A - B$, $p^B_{aa'} = p_{aa'}$, $p^B_{bb'} = p_{bb'}$ & $\forall i, bp^B(i)a$. For any profile p and individual i, we use the standard notation $p = (p^i, p^{-i})$, where $p^i = p(i)$ and $\forall j \neq i$, $p^{-i}(j) = p(j)$. Recall that for any $a, b, x, y, \rho, \sigma, r, q$, $\rho_{ab} \geq \sigma_{xy}$ means that $(x\sigma y \Rightarrow a\rho b)$ & $(x\sigma^> y \Rightarrow a\rho^> b)$; similarly, $r_{ab} \geq q_{xy}$ means that $\forall i$, $r(i)_{ab} \geq q(i)_{xy}$.

A social choice function (SCF) for society N over choice set A is a function from L_A^N into A. A SCF f is nonmanipulable (NM) if $\forall i, p, q$, $f(p)p^i f(q^i, p^{-i})$; dictatorial, if $\exists i$ such that $\forall a \in f(L_A^N)$ $\exists p^i \in L_A^{\{i\}} : \forall r^{-i} \in L_A^{N-\{i\}}$ $f(p^i, r^{-i}) = a$. f is monotonic (M) if $\forall a \neq b$ & p, q with $q_{ab} \geq p_{ab}$, $f(p) = a \Rightarrow f(q) \neq b$.

Theorem 4 (Gibbard-Satterthwaite). Any nonmanipulable SCF f for a finite society N over a choice set A with $|f(L_A^N)| \geq 3$ is dictatorial.

The proof first shows monotonicity of f, second, a Pareto-like condition, and finally extends f into an IIA and P SWF g. By Arrow's Impossibility Theorem (AIT), g is dictatorial, which implies that f is dictatorial as well.

Proof. (i) For any i, p, q and $a \neq b$ suppose $f(p) = a$, $p^{-i} = q^{-i}$ and $q^i_{ab} \geq p^i_{ab}$. Then, by NM, $a = f(p)p^i f(q)$ and $f(q)q^i f(p) = a$. Thus, by $q^i_{ab} \geq p^i_{ab}$, $f(q) \neq b$. So, f is M. (ii) Let $a \in f(L_A^N)$, b and r with $\forall i, ar(i)^> b$. We show that $f(r) \neq b$. Take q with $f(q) = a$. By (i), $f(q^{\{a\}}) = a$, and $f(r) \neq b$. (iii) For each r, define a complete binary relation on $f(L_A^N)$ by $ag(r)b \Leftrightarrow f(r^{\{ab\}}) = a$. By (i), g is IIA. By (ii), g is P. For some $a \neq b \neq c \neq a$ and r, suppose $ag(r)bg(r)c$. Then, by (ii) $f(r^{\{abc\}}) = a$ because by (i), $f(r^{\{abc\}}) = b \Rightarrow f(r^{\{ab\}}) = b$ and $f(r^{\{abc\}}) = c \Rightarrow f(r^{\{bc\}}) = c$. So, by (i), $f(r^{\{ac\}}) = a$, so $ag(r)c$, i.e., g is transitive. By AIT, g is dictatorial, and so is f. $\square$

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Chapter 3

Multiple Equilibria in Geographic Labor Mobility.

3.1 Introduction

> "Americans are still full of get-up-and-go. Witness the mass exodus from California in the past few years... The average American moves roughly twice as often as the average Briton. In the year to March 1991, the last period covered by the Bureau of Census, 41.5 millions Americans, or 17 percent of the population moved home, and 7 millions of them moved out of state. One result is that unemployment differential between regions are far less persistent in America than in Europe. ...**American mobility is partly a consequence of culture** in a country that was built by immigrants and has thrived on economic opportunism... (*The Economist*, October 16, 1993).

This article in *The Economist* uses cultural differences to explain the observed differences in mobility patterns between the US and Europe. How relevant are non-economic reasons in explaining moving decisions? More specifically, how much do family and friends influence the decision of moving? Are the observed differences due solely to different "taste for friends" and "economic opportunism", or due to different equilibria? These issues are addressed in this paper.

Welfare systems and regulations are often invoked to explain scarce labor mobility in Europe (e.g. the unemployment benefits system in Britain discourages the unemployed from looking for jobs in other regions). These explanations are valid at the individual level, but they suffer serious problems of endogeneity at the aggregate level. Why do European governments maintain this system? The citizens of these countries, who apparently do not like to move, voted for a "social insurance" that guarantees them income without the need to move. The question that follows is: why are communities willing to pay more taxes to have some sort of "insurance" against moving?

Higher transaction costs (selling and buying a house costs more in Europe than in the US on average) are often quoted as another cause for lower regional mobility. However, again in this case, it is not clear which is the cause and which is the effect. People move less because there are high costs of moving, but the costs of moving depend on the number of people who move [1]

The previous two examples suggest that an explanation for different attitudes toward migration must go back to a micro-economic model. The first requirement for this model is that different attitudes must be explained without the need of "pseudo-exogenous" reasons. The second requirement is that the model must abstain from postulating different utility functions across communities and, instead, derives different aggregate behavior from the different interactions of agents within the community.

The basic idea is that potential movers care about both economic factors (wages, unemployment rate, etc.) and social factors (presence of family, proximity to the friends, etc.). The utility of an individual depends on the action of other individuals (e.g. I cannot enjoy my friends if they decide to move away), and this generates the possibility of multiple equilibria. Therefore, if an individual becomes unemployed, his decision to migrate or not depends on whether his family is present. If the family is present (and will be present in the future), the individual will accept the pain of a longer unemployment spell or of a substantial drop in income; if the individual lives alone, he will be more likely to follow only economic incentives.

The paper is structured in three parts. The first part places the problem within

[1] We are thinking about the costs of selling or buying a house in the US compared to Europe, or the existence of a second hand market for furniture.

the context of recent literature; both micro- and macro- aspects are considered. The micro literature traditionally has been more focused on the determinants of migration; macro-economists, on the other hand, are more interested in the consequences of migration. We review recent findings on the causes of migration. Since our model shows the possibility of two equilibria in labor mobility pattern, we proceed by contrasting empirical facts about regional labor mobility in Europe and in the US

The second section develops a model to formalize our hypothesis about the role of friends and family in the migration decision. In the last part, we construct an index of family attachment and we test the role of family in the decision of migrating. An attempt is made to measure the relative sizes of the economic and non-economic factors. In the last part, we use the framework of our model to explain peculiarities in the migration pattern of different communities in the US (across states and across ethnic groups).

3.2 Background

3.2.1 Determinants of migration

The micro-economic literature on migration has been dominated by the "orthodox approach": the decision to move is treated as a decision to invest in human capital (Sjaastad (1962)). Potential migrants compare the costs of investment (cost of moving) with the returns given by the flow of higher discounted wages (and other benefits) in the new location.

The empirical validity of this theory in its simplest form is quite difficult to assess because the costs of moving (both pecuniary and non- pecuniary), the proper discount rate and the forecasts of future wages are each difficult to estimate [2].

Beyond these theoretical problems, in the past, there was the problem of data availability: to properly test the hypothesis of migration as a form of human capital

[2]The estimation becomes even more difficult if a general equilibrium model is taken into account: the agent should consider that the present migration toward high wage regions will decrease the wage differential.

investment, we should have a database which keeps track of the individual movers, their incomes, and their costs of moving. Unfortunately, before the 1970's, few data sets of this type were available and most of the studies used aggregate data (Greenwood (1985), Cebula (1979), Clark (1983) and Bowles (1970)). [3]

A major change in the study of the determinants of migration came at the end of the 70's, when micro-data bases began to be used. [4] The careful study of these data pointed out that migration is quite a complex phenomenon and that the "orthodox approach" needs substantial corrections in order to explain some empirical evidence. In particular, the new data sets allowed for a more careful analysis of the life-cycle and personal histories which play an important role in the decision to migrate.

The subsequent literature concentrates on discovering the effects of these micro determinants of migration; several factors found relevant include: marital status (Mincer (1978) and Sandell (1977)), non-tradeable goods (Graves and Linneman (1979)) and climatological amenity variables.

Interestingly, micro data also add to understanding the behavior of repeat migration initially observed in the aggregate data (for instance in Kao and Sirmans (1977) and later DaVanzo (1981)).

The availability of new time series data have also changed the theoretical interest from a merely static modeling of the causes and the consequences of migration to the development of more complex models to explain the observed dynamics. Among the dynamics observed, the lagged responses are fundamental in the migration process, thus "pure cross-sectional models are unlikely to encompass fully the complex nature of these relationships" (Molho (1984)). This kind of approach has been extended to many other countries where time series data are available (Greenwood (1985), p.537).

[3] Even with aggregate data, it was possible to find that social groups in the US differ significantly in their migration habits. For instance Bowles in (1970) shows that blacks and whites differ significantly in the responsiveness to expected income gain from moving.

[4] Notably for the US, the PSID (Panel Study of Income Dynamic), NLS (National Longitudinal Surveys) and PUMS (Census Public Use Microdata Samples) are mostly used.

3.2.2 Consequences of migration

Regional labor mobility has recently attracted new research because it is a key factor in evaluating the size of an optimal currency area (Mundell (1961)).

Recent papers (Blanchard (1992)) on the effects of regional shocks in the United States have pointed out that a reduction in the level of unemployment in a specific region in the United States is due more to internal out-migration than to job creation. This finding implies that the level of average unemployment depends crucially on the mobility of the workers as a response to wage differentials and/or level of unemployment.

Eichengreen (1993) finds that "in Europe, workers are less sensitive to economic incentives (the elasticity of migration with respect to unemployment differential is twice as large in the United States than in the United Kingdom or in Italy.") Eichengreen argues that "low levels of labor mobility in Europe reflect not only legal restrictions, but also culture, language, and history."

If unemployment reduction is mostly due to out-migration (Blanchard (1992)), and if Europeans respond less for economic incentives, the result is the persistence of high regional unemployment rates in Europe result[5]; moreover, in Europe, unemployment rates show higher regional variation.

3.3 The Model

There is a great deal of evidence that both purely economic factors and social factors are fundamental in the decision to migrate. This section develops a model that relates geographical mobility, employment status and family-and-friends (F&F) attachment.

Each worker at each period t has an employment status L_t, a F&F status F_t and lives in a region R_t. We consider the simplest case in which L and F can take only two values each, i.e., each period a worker can be either employed (e) or unemployed

[5]Using time-series techniques, Decressin and Fatas (1992) find that in the short run, even within countries, there is little migration in response to region-specific labor demand shocks.

Country	Correlation between 74 and 87	varinace ($\frac{u_i}{u}$) in 86
Australia (8)	-0.11	1.8
Canada (10)	0.67	8.2
France (22)	0.50	2.8
Germany (11)	0.83	8.3
Finland (12)	0.91	20.5
Italy (20)	0.84	13.6
Japan (20)	0.91	5.8
Sweden (24)	0.69	14.8
Britain (10)	0.92	5.1
United States (51)	-0.33	6.6

Source: Padoa Schioppa (1990) and OECD.
Note: number of regions in each country is given parentheses.

Table 3.1: Correlation between regional unemployment in 1974 and 1987, and regional variance in 1986

(u), and either married (m) or single (s).

At the beginning of each period, a worker chooses whether or not stay in the current region or move to another one next period. Whatever his current working and family situation is, if a worker decides to change region, he'll be single and employed next period. We assume that he can find a region where a job is offered to him, but he cannot find a partner before moving.

With this framework we try to capture in a simple way both economic and social factors in the decision to migrate. The literature has already shown theoretically and empirically the importance of economic factors such as the employment status of the individual, region-of-origin and region-of-destination unemployment rates, wage gain upon moving, etc. In our model all these effects are captured by a wage differential w —the wage earned by an employed worker.

The fundamental effect of family and friends as a migration determinant has been pointed out often in the literature, at least since Greenwood (1969). Two main F&F effects have been considered in the literature. First, Greenwood (1969) considers that if family and friends of an individual have migrated in the past from region i —current region of the individual— to region j, it is more likely that the individual, when moving out of region i, decides to go to region j rather than to another region.

The evidence on the importance of this effect came from the fact that the coefficient of the stock of past migration from i to j was strongly significant in a regression with dependent variable the current amount of migration from i to j.[6] In order to keep the model as simple as possible, we do not include this F&F effect. However, as it will become clear later, the main result in this section is robust to the inclusion of this effect.

More important for the assumptions of our model is the discouraging effect on migration due to attachment to family and friends living in the individual's region. The importance of this F&F effect as a fundamental determinant of internal migration has been emphasized since the 1970s (Mincer (1978), Sandell (1977)). Using panel data, these and later papers show strong evidence of the importance of this effect. This effect is captured in our model assuming that a married individual obtains a utility M from living with his/her spouse, and that if the individual migrates, then ceases to obtain this utility.

We restrict our study to symetric steady state equilibria. If a worker with status (L, F) does not move, the probability of changing of employment status is π_L and the probability of changing of family status is π_F. These two events are assumed independent. In this way a Markov process with four states is defined.

The behavior of each individual is assumed to correspond to the maximization of an expected utility function. This utility is the discounted per period sum of wage plus family satisfaction.

$$\sum_{t=0}^{\infty}(1+r)^{-t}[W(L_t) + S(F_t)]$$

Where $W(e) = w > 0$, $S(m) = M > 0$ and $W(u) = S(s) = 0$.

For three of the four choice situations a worker can face, it is clear what his optimal decision is. Imagine a worker who knows that if he doesn't move, he will be single and unemployed next period. Clearly, he will choose to move to another region so

[6] Using census data this effect has been tested in this very indirect way for a number of countries including US (Greenwood (1969)), India (Greenwood (1973)) and Venezuela (Levy and Wadycki (1973)). However, it is not at all clear why a strong significance of the coefficient of past migration should be attributed to this F&F effect.

that he'll be employed and single. The case of a worker who knows that if he doesn't move he'll be married and employed is also clear, he will decide to stay. A worker who knows that if he doesn't move he will be employed and single next period, will be indifferent between moving or not. Whatever he chooses he will be employed and single.

The choice of a worker who knows that if he doesn't move, he will be married but unemployed, depends on what he expects his current and future partners will do when faced with this same decision. Imagine that a worker believes that the probability that a person who faces this decision chooses to move is q. Notice that in this case $\pi_m = \pi_e q$, $q \in \{0, 1\}$.

To solve this problem we use the following lemmas.

Lemma 1.

Let A and B be policies from time 0 to ∞. Assume that $r > 0$. If policy A is preferred to policy B, then there exists some time $\bar{t}$ such that for all $t > \bar{t}$ and for any policies C and D from time t on, the compound policy "follow A until t and then adhere to C" is preferred to the compound policy "follow B until t and then adhere to D".

Proof. It follows immediately from the form of the utility function as a *discounted* sum and from the boundedness of the value from following any policy.

Lemma 2.

Consider a stationary problem with two alternative actions a and b each period. Let A_t be the policy of choosing every period from t on action a and B_t be the policy of choosing every period from t on action b. Let $x(y_t)$ be the best action to take at $t-1$ if from t on policy y_t is compulsory, for $y = A, B$. Then $x(A_t) = x(B_t)$. I.e., the best action at $t-1$ is independent of which policy is compulsory from t on.

Proof. By stationarity, the optimal policy is "*always* choose a" or "*always* choose b". Then it is not possible that simultaneously $x(A_t) = b$ and $x(B_t) = a$.

Suppose that for some $\bar{t}$, $x(A_{\bar{t}}) = a$ and $x(B_{\bar{t}}) = b$. Then for all t, $x(A_t) = a$ and $x(B_t) = b$. Thus for any t, if policy A_t is compulsory, the best policy from 0 up to $t-1$ is "choose a from 0 to $t-1$", while if B_t is compulsory, the best policy from 0 to $t-1$ is "choose b from 0 to $t-1$". This contradicts the previous lemma. Q.E.D.

Using these facts we can solve the choice problem of the married and unemployed worker just by determining his optimal choice at time t assuming that after t he will always decide to change of region whenever faced with the prospect of being unemployed and married for the incoming period.

Let V_{FL} be the value of the policy "change of region whenever married and unemployed" for a worker who knows that if he doesn't move to another region the next period, he'll be in employment status L and in family situation F. And let V'_{mu} be the value of a worker who is obliged next period to be married and unemployed, and after next period he follows the above policy of moving whenever married and unemployed. Note that $V_{su} = V_{se} = V_{mu}$.

The following relations hold:

$$V_{se} = (1+r)^{-1}[w + \pi_s(1-\pi_e)V_{me} + (1-\pi_s(1-\pi_e))V_{se}] \tag{3.1}$$

$$V_{me} = (1+r)^{-1}[w + M + (1-\pi_e q)(1-\pi_e)V_{me} + (1-(1-\pi_e q)(1-\pi_e))V_{se}] \tag{3.2}$$

$$V'_{mu} = (1+r)^{-1}[M + (1-\pi_e q)\pi_u V_{me} + (1-(1-\pi_e q)\pi_u)V_{se}] \tag{3.3}$$

We are interested in the sign of $y \equiv V_{se} - V'_{mu}$. From equations (3.1) and (3.2) we obtain

$$(1+r)(V_{me} - V_{se}) = M + [(1-\pi_e q)(1-\pi_e) - \pi_s(1-\pi_e)](V_{me} - V_{se}) \tag{3.4}$$

Thus

$$V_{me} - V_{se} = \frac{M}{1 + r - (1 - \pi_e q)(1 - \pi_e) + \pi_s(1 - \pi_e)} \tag{3.5}$$

From (3.1) and (3.3),

$$(V_{se} - V'_{mu}) = \frac{w - M + [\pi_s(1 - \pi_e) - (1 - \pi_e q)\pi_u](V_{me} - V'_{se})}{1 + r} \tag{3.6}$$

Substituting,

$$y \equiv V_{se} - V'_{mu} = \frac{a + bq}{c(q)} \tag{3.7}$$

Where,

$$a = w[r + \pi_e + \pi_s(1 - \pi_e))] - M[r + \pi_e + \pi_u], \tag{3.8}$$

$$b = [(w - M)(1 - \pi_e) + M\pi_u)]\pi_e \tag{3.9}$$

$$c(q) = (1 + r)[r + \pi_e + (1 - \pi_e)(\pi_s + \pi_e q)] > 0, \quad \textit{for all } q. \tag{3.10}$$

A worker who knows that next period he will be married and unemployed, unless he changes region, chooses to move if y is positive, and stays if it is negative.

If y is negative whenever $q = 0$, and y is positive whenever $q = 1$, then there are two stable equilibria: $q = 0$ in which no married worker moves, and $q = 1$ in which every would be unemployed worker moves so that everybody is employed.

From (3.7) and (3.10) it follows that there are two equilibria if simultaneously $a < 0$ and $a + b > 0$. From (3.8) and (3.9) this happens when

$$\frac{r + \pi_e + (1 - \pi_e)(\pi_u + \pi_e)}{r + \pi_e + (1 - \pi_e)(\pi_s + \pi_e)} < \frac{w}{M} < \frac{r + \pi_e + \pi_u}{r + \pi_e + (1 - \pi_e)\pi_s} \tag{3.11}$$

Let ρ_1 and ρ_2 be the left hand side and right hand side terms in (3.11), and $\rho = \frac{w}{M}$.

For (3.11) to hold it is necessary that the range $\rho_2 - \rho_1$ be positive. This is equivalent to:

$$\pi_u(1+r) > (1-\pi_e)\pi_s(1-\pi_e-\pi_u) \tag{3.12}$$

This condition is satisfied in particular if $\pi_u > \pi_s$, i.e., if the probability of an unemployed getting a job is greater than or equal to the probability of a single getting married. It is apparent that if the range $\rho_2 - \rho_1$ were not positive for some parameter values, by increasing any of the parameters r, π_u or π_e, or by decreasing π_s, the range would eventually become positive [7].

3.3.1 The model with an exogenous moving cost

The model can be easily extended to the case where there is an exogenous cost of moving. If the worker must pay a cost k each time he moves then the value of a single and unemployed worker is no longer equal to that of a single and employed worker. If the moving cost were very high, no worker would ever want to move. Let's then assume that at least some workers move, i.e., that k is not so large to deter from migrating even to the single and unemployed workers.

From lemmas 1 and 2, to determine the optimal choice of a married and unemployed worker we can just consider his current period choice assuming that from next period on he'll have to change of region whenever presented with the prospect of becoming married and unemployed again. Thus, defining the values V_{FL} and V'_{mu} as above, the worker will move if and only if $V_{se} - k > V'_{mu}$. Equations (3.1), (3.2) and (3.3) must now be changed to take into account the cost of moving.

$$V_{se} = (1+r)^{-1}[w + \pi_s(1-\pi_e)V_{me} + (1-\pi_s(1-\pi_e))V_{se} - \pi_e k] \tag{3.13}$$

[7]However, the range $\rho_2 - \rho_1$ is not increasing in r. For r large enough, the range is decreasing in r. In particular, if $\pi_u > \pi_s$, for any parameter values, the smaller r, the larger the range $\rho_2 - \rho_1$. The derivation is presented in Appendix A.

$$V_{me} = (1+r)^{-1}[w+M+(1-\pi_e q)(1-\pi_e)V_{me}+(1-(1-\pi_e q)(1-\pi_e))V_{se}-\pi_e k] \quad (3.14)$$

$$V'_{mu} = (1+r)^{-1}[M + (1-\pi_e q)\pi_u V_{me} + (1-(1-\pi_e q)\pi_u)V_{se} - (1-\pi_u)k] \quad (3.15)$$

Now, as in our previous analysis of zero exogenous moving cost, $V_{me} - V_{se}$ is given by (3.5). However, instead of (3.6) now

$$(V_{se} - V'_{mu}) = \frac{w + (1-\pi_u-\pi_e)k - M + [\pi_s(1-\pi_e) - (1-\pi_e q)\pi_u](V_{me} - V'_{se})}{1+r} \quad (3.16)$$

holds. Note that equation (3.16) can be written also in the form,

$$(V_{se} - V'_{mu} - k) = \frac{w - (r+\pi_u+\pi_e)k - M + [\pi_s(1-\pi_e) - (1-\pi_e q)\pi_u](V_{me} - V'_{se})}{1+r} \quad (3.17)$$

Which allows us to reduce this problem to the case previously treated where $k = 0$. Define $y(w, M, k) \equiv V_{se} - V'_{mu} - k$. This can be computed plugging (3.5) into (3.17). This is similar to what we did to derive the value of y in (3.7), but there we used (3.6) rather than (3.17). By comparing (3.6) and (3.17) it is clear that

$$y(w, M, k) \equiv V_{se} - V'_{mu} - k = y(w - (r+\pi_u+\pi_e)k, M, 0) \quad (3.18)$$

We conclude that the introduction of an exogenous moving cost k is equivalent to a decrease in wage differential of $(r+\pi_u+\pi_e)k$. It reduces the interest in moving in $y(r+\pi_u+\pi_e, M, 0)k$.

Therefore, there are two stable equilibria if the following inequalities hold,

$$\frac{r+\pi_e+(1-\pi_e)(\pi_u+\pi_e)}{r+\pi_e+(1-\pi_e)(\pi_s+\pi_e)} < \frac{w-(r+\pi_e+\pi_u)k}{M} < \frac{r+\pi_e+\pi_u}{r+\pi_e+(1-\pi_e)\pi_s}$$

3.3.2 Some Empirical Implications of the Model

Consider the two equilibria case. In both equilibria there is migration, but in the $q = 1$ equilibrium people move more than in the $q = 0$ equilibrium. In the $q = 1$ equilibrium all unemployed workers move, while in the $q = 0$ equilibrium only single and unemployed workers move. If $q = 0$, family status predicts perfectly whether an unemployed worker moves. While if $q = 1$, family status does not affect the moving decision of anyone.

In the next section, we move into the task of testing the empirical validity of these results: whether those groups (Europeans vs. Americans, US Blacks vs. US Whites, North-East US vs. West US, etc.) a with larger moving rate show a smaller effect of F&F and a stronger effect of purely economic factors on the determination of the moving decision. For this task we estimate a migration equation using US micro-data and focus on the differences across ethinic groups and across macro-regions.

3.4 Multiple equilibria across ethnic groups and regions

In this section we first estimate a migration equation for the US including both variables accounting for the economic reasons to migrate as well as variables accounting for non-economic reasons. Secondly, we discuss how the observed different propensities to move between Blacks and Whites can be explained using the framework of our model. Finally, we consider how the differences in migration habits across macro-regions in US can be framed in terms of our model.

3.4.1 Data and Variables

The empirical tests are based on the 1968-1988 waves of the University of Michigan's Panel Study of Income Dynamics (PSID), that give twenty one years of data on over 5,000 US families. "Person-year" observations, each representing one year in which a person is at risk to migrate, are used as units of analysis. [8] Before transforming the data in "person-year" observations, the data have been elaborated to exploit the panel nature of the information. To understand the way in which we constructed our variables, the structure of PSID must first be explained.

PSID, which started in 1968, interviewed 4,802 households across 40 states. Each individual was assigned a "family number" and a "person number". These two numbers, used in combination, identify a person throughout the years. If a person leaves the original family and forms a new family, he keeps the original "68 family number" for identification purposes [9] ; so, after the 1968 wave, sharing the same "68 family number" does not necessarily mean living in the same household, e.g. brothers-in-law will share the same "68 family number" even if they live in different households.

We define "extended family" as the collection of individuals who share the same "68 family numbers". The "extended families", that coincide with the households in 68, later diverge as families split and as new individuals join the survey, etc. As a result, the "extended families" are more numerous than the households. In table 3.2 there is a comparison between the two concepts of families and a description of how many members of the "extended family" live in the same metropolitan area. The data refer to the 1980's.

We use the concept of "extended family" to construct the variable *fff*. *fff* is defined as the fraction of members aged more than 16 excluding the individual that live in

[8] DaVanzo (1978) includes in her sample individuals who are head of the family in the considered year. This gives her more information about the individual (since PSID provides additional information on the heads of the family) but substantially reduces the number of obsevations available (about 1 person out of 6 is head of the family). We chose to extend the analysis to all individuals, giving us many more observations (approximately 50,000) at the cost of less detailed information.

[9] Also, new members of the family who join the PSID sample later are given the same "68 family number" of the person they are associated with. Just under one third of the 37,528 have joined panel families through marriage or cohabitation. About one fifth of the 37,528 were born into sample families

Type of family	Average Size	Variance	Maximum
Household	3.46	3.14	16
Extended Family	5.61	12.12	22
Ext. Fam. in Same Location	4.36	10.12	20
Sources: Authors' calculations based on PSID data (1979-1988).			

Table 3.2: Description of the families

the same metropolitan area [10] over the total number of individuals in the "extended family" excluding the individual. This measure is constructed so that it is 0 if the individual lives alone (i.e. without any "extended relative" in his area) and 1 if all the family lives with himself.[11] *fff* is a proxy to measure family attachment.

3.4.2 The migration equation for the US (white men)

Given the discrete nature of the choice of moving we use a logit setting. The dependent variable *move* refers to period t, the independent variables to period $t+1$; all variables are fully defined in Appendix B. Table 3.3, as all the other following tables, includes: the logit coefficients coming from the estimation of the previous equation, the marginal effects calculated at the average value of the independent variables[12], the t statistics for the marginal effects, and their significance. In appendix D, we give a graphical interpretation of the results of table 3.3.

The regressions are limited to White men, unless otherwise specified.

The estimated equation contains the usual control variables (head of the family's education , dummy variables for young people, marital status, and ownership of a house [13]). All these coefficients are significant and have the appropriate signs: a higher degree of education increases the probability of moving by 1 %, being young (younger than 24) increases the probability by 1.7 %, being married and the ownership

[10]If the individual lives in a rural area, members of the same family living in rural areas of the same state are considered.

[11]*fff* is not defined if an "extended family" is formed by only one individual (this is the case for very few individuals inn our sample). In this case *fff* was given the value of 0.

[12]The marginal effects are calculated by multiplying the logit coefficients by the number $\Lambda(\beta'x)(1-\Lambda(\beta'x))$ (where Λ is the logistic function calculated at the average value of the independent variables).

[13]This variable is included as a proxy of location-specific capital.

	Logit Coeff.	Marg. Effect	t-stat	> |t|
CONSTA	-1.442	-0.053	-13.40	0.000
FFF	-1.649	-0.060	-21.06	0.000
MV1	1.033	0.038	13.558	0.000
OWN	-0.839	-0.031	-15.26	0.000
EDH	0.264	0.010	4.350	0.000
finc	0.209	0.008	3.497	0.000
AGE1	0.462	0.017	6.995	0.000
INC1	-0.198	-0.007	-2.855	0.004
INC2	-0.282	-0.010	-3.620	0.000
INC3	-0.226	-0.008	-2.485	0.013
MAR	-0.116	-0.004	-1.807	0.071
UNEM	0.425	0.016	3.826	0.000
Number of observations: 31,191 tolerance: 4.44e-16				
Likelihood ratio chi-sqare 1,490.069 with prob 0.000				
percentage of correct predictions 94.652				

Table 3.3: LOGIT for US White males (moved 5.33 %)

of a house discourage people from migrating. The dummy variables for the level of income are significantly negative, such that having a range of income between \$ 10,000 and \$ 40,000 significantly decreases the probability of moving.

DaVanzo (1981) showed that the phenomenon of repeat migration is quite relevant: people who have migrated in the year before tend to migrate the following year with high probability (12,6 % returned to the previous residence and 15 % moved onward in her sample). To deal with the repeat migration pattern we introduced the dummy variable *MV1*, which is very significant in the same manner as in DaVanzo (1981). This confirms that a large part of migration is due to repeat migration.

To deal with the economic motivations for migration, we introduce two variables: *unem* and *finc*. The idea behind *unem* is that: if the individual is unemployed and in the county the unemployment rate is bigger than 4 %, the individual has a strong incentive to move. For simplicity in our model the only reason for moving is that it is easier to find a job in another town [14]. In reality, there are plenty of other economic

[14]When asked, people give many resons for moving. The ranking of these reasons itself is not clear. In the 1987 interview, 18 % of the heads of the family, who thought probable a future move, indicated getting another job as the probable cause. One year later only 11 % of the heads who actually moved indicated getting another job as the cause why they moved in the past. On the

reasons to move. Even though the individual is employed, he may want to move to get a better job. Ideally, we would like to know how the individual assesses his future income if he stays in the same town versus the possibilities offered in other locations; however, this information is not available. To account for the possibility that an abrupt fall in income can change the willingness of an individual to move for economic reasons, we introduced the variable *finc* for people who are employed.

Both employment status and a fall in income are significant. In both cases, the individuals are more likely to move. This question is taken up later when we specifically consider unemployed people.

The value which measures the family attachment (*FFF*) is significant and has a negative effect on the probability of moving. The relative size of the family attachment effect in comparison with the economic motivations for moving is quite large [15]. The

other hand, personal non-economic or involontary reasons accounted for approximately 40 % of the reasons for moving.

Reasons for Moving

Reasons	**Why might move**	**...and...**	**why moved**
Take another job	6 .0		2.4
Get nearer to work	0.9		1.3
Need more space	6.9		3.0
Need less space or less rent	2.8		1.9
Want own home (got married)	7.7		4.6
Closer to school (better heighborhood)	2.4		2.0
Involontary reasons (e.g. health)	2.6		4.4
Ambiguous reasons	3.7		1.7
Not available	0.6		0.8
Not moved	66.5		77.8

Source: PSID 1987 question "*why might move*" in the near future, and PSID 1988 question "*Why moved,*" which is the actual reason for having moved. In this table move means moving to another house, not necessarely to another metropolitan area.

[15]In comparing the relative sizes one must be careful that *unem* is a discrete variable, while *fff* ranges from 0 to 1.

result is robust in terms of specification of the model [16].

The number of moves is not distributed uniformly across the individuals since a high percentage of moves come from very few individuals who move many times (table B.1). A plausible doubt is that the coefficient on the variable *fff* picks up omitted variables. The individuals more likely to move live in towns with few relatives so their *fff* is, on average, lower. Since these individuals move more frequently as compared to the rest of the population, there will be a strong correlation between moving and low values of *fff*. In principle, it is very difficult to get rid of this problem because it is problematic to identify the possible types of individuals. As an attempt to address this problem, we divide our sample in 4 sub-samples using as a criterion the number of times that an individual moved in the 21 years of the sample. We run the same regression for these sub-samples. In each of these regressions, *fff* remains highly significant [17]. (The results are in appendix C).

Usually, to get rid of the individual effect in a logit contest, the conditional probability is constructed (Green, 1993); the condition to apply this method is that the probability of moving a certain year is indipendent of having moved in the past years. We know that this is not the case in the present context because, if a person moved in the past, even controlling for all the other variables, the probability of another move is much higher. For this reason, we present regressions with individual effects just in a linear setting. In tables C and C, there are the results for fixed and random effects. In both cases the relevant parameters remain significant with the expected sign.

We are especially interested in the behavior of unemployed people. In our sample, the propensity to move among white men excluding those unemployed is 5.15 % each year, if only the unemployed are considered, this propensity increases to 9.01 % each year. Table 3.4 shows the results of a regression which only considers White unemployed. The additional variable *unc* (the local unemployment rate) is introduced

[16] A probit version of the same regession gives similar results (see appendix). A linear version gives same results.

[17] In the appendix we give the complete results for n =2,3,4,5. The regression for n=1 gives spurious results because we analyze the subsample of people who moved from their origin family and never moved again.

	Logit Coeff.	Marg. Effect	t-stat	> $|t|$
CONSTA	-1.560	-0.105	-3.399	0.001
FFF	-1.033	-0.070	-3.220	0.001
MV1	1.393	0.094	5.285	0.000
OWN	-0.545	-0.037	-2.582	0.010
EDH	0.608	0.041	2.250	0.025
FINCM	-0.193	-0.013	-0.897	0.370
AGE1	0.017	0.001	0.070	0.944
INC1	-0.015	-0.001	-0.058	0.954
INC2	-0.137	-0.009	-0.339	0.734
INC3	0.295	0.020	0.579	0.563
MAR	-0.163	-0.011	-0.735	0.463
UNC	1.781	0.120	0.556	0.578
Number of observations: 1,410 tolerance: 8.32e-13				
Likelihood ratio chi-sqare 87.502 with prob 0.000				
percentage of correct predictions 90.780				

Table 3.4: LOGIT for the US white unemployed (move 9.01 %)

to control for the condition of the local labor market.

This suggests that, if an individual is unemployed, all income-related variables are not relevant. The local level of unemployment is not significant as well. This does not come as a surprise because other studies have shown that it is difficult to find aggregate macro-economic variables which are significant for individual decisions. The other variables remain significant and roughly have the same size as in the previous regression.

3.4.3 Two equilibria: Blacks and Whites

Blacks are far less prone to move than whites; using the definition of moves given in this paper, 3.68 % of black males move each year versus 5.31 % of White males. Table 3.5 reports the statistics for the variables for the two groups.

Black men, on average, show many characteristics that should increase migration rates such as higher unemployment rates, a lower percentage of married people, lower incomes, lower levels of ownership of homes and higher level of family attachment. Samuel Bowles (1970) found that, even controlling for the usual demographic variables, there is an unexplained difference in the migration patterns between blacks and

Variables	Mean Blacks	Std Dev	Whites	Std Dev
FFF	0.8589	0.2561	0.7756	0.3071
OWN	0.4616	0.4985	0.7076	0.4549
EDH	0.0542	0.2264	0.2371	0.4253
finc	0.2989	0.4578	0.2824	0.4502
AGE1	0.3380	0.4730	0.2236	0.4167
INC1	0.2529	0.4347	0.2091	0.4066
INC2	0.1324	0.3389	0.2089	0.4065
INC3	0.0527	0.2235	0.1524	0.3594
MAR	0.4997	0.5000	0.7184	0.4498
UNEM	0.1299	0.3361	0.0410	0.1983

Source: Authors' elaboration using the PSID waves (1979-1980)

Table 3.5: Statistics for Blacks and Whites

whites. Blacks seem to be less responsive to economic incentives (Bowles (1970)).

To address this problem in the context of our model, we pose the following question: Are blacks more sensitive to family attachment and less sensitive to economic incentives? Is there a significant difference in the parameters for family attachment and for employment status? The table 3.6 show the results of a regression restricted to blacks. [18]

The main difference with the analogous table for whites is that the economic variables now play a less important role in migration decisions.

The following exercise focuses on the role of economic versus non economic causes of migration in the two communities, constraining all the parameters of the demographic variables to be the same. In the pooled regression (blacks and whites), three variables are included: *black*, which is a dummy variable for blacks, *fffb* which is the value of *fff* if the individual is black and *unemb*, which is assigned the value of 1 if the individual is black and unemployed (table 3.7).

The results show a significant difference in the parameters for family attachment (*fffb*) and for unemployment *unemb*; once this difference is taken into account the dummy variable for blacks is not significant. Blacks appear to be more sensitive to family attachment and less sensitive to economic incentives. In appendix D, we give

[18]In the regression restricted to blacks we do not include the variable *mv1* to avoid problems of convergence in the estimation.

	Logit Coeff.	Marg. Effect	t-stat	> $|t|$
CONSTA	-1.034	-0.024	-6.732	0.000
FFF	-2.879	-0.067	-21.17	0.000
OWN	-0.796	-0.019	-8.626	0.000
EDH	0.264	0.006	1.777	0.076
finc	0.126	0.003	1.250	0.211
AGE1	0.618	0.014	6.362	0.000
INC1	0.129	0.003	1.171	0.242
INC2	-0.058	-0.001	-0.376	0.707
INC3	0.142	0.003	0.688	0.492
MAR	-0.281	-0.007	-2.810	0.005
UNEM	0.099	0.002	0.734	0.463

Number of observations: 17,242 tolerance: 4.44e-16
Likelihood ratio chi-sqare 676.197 with prob 0.000
Percentage of correct predictions 96.317

Table 3.6: LOGIT for US black males (move 3.68 %)

	Logit Coeff.	Marg. Effect	t-stat	> $|t|$
CONSTANT	-1.513	-0.047	-15.560	0.000
FFF	-1.618	-0.050	-20.285	0.000
MV1	1.160	0.036	17.347	0.000
OWN	-0.782	-0.024	-16.424	0.000
EDH	0.266	0.008	4.727	0.000
finc	0.191	0.006	3.700	0.000
AGE1	0.472	0.015	8.618	0.000
INC1	-0.127	-0.004	-2.160	0.031
INC2	-0.224	-0.007	-3.214	0.001
INC3	-0.154	-0.005	-1.844	0.065
MAR	-0.164	-0.005	-3.043	0.002
BLACK	0.134	0.004	1.326	0.185
UNEM	0.431	0.013	3.898	0.000
fffb	-0.839	-0.026	-6.030	0.000
unemb	-0.336	-0.010	-2.035	0.042

Number of observations: 48,348 tolerance: 7.44e-13
Likelihood ratio chi-sqare 2,350.348 with prob 0.000
Percentage of correct predictions 95.239

Table 3.7: LOGIT for US males (move 4.74 %)

	Logit Coeff.	Marg. Effect	t-stat	$>\|t\|$
CONSTA	0.078	0.002	0.157	0.875
FFF	-2.840	-0.076	-7.114	0.000
OWN	-0.805	-0.022	-2.774	0.006
EDH	0.213	0.006	0.356	0.722
FINCM	-0.042	-0.001	-0.169	0.866
AGE1	-0.055	-0.001	-0.212	0.832
INC1	0.328	0.009	0.989	0.323
INC2	0.884	0.024	1.576	0.115
INC3	1.196	0.032	1.289	0.198
MAR	-0.187	-0.005	-0.710	0.478
UNC	-9.769	-0.262	-2.115	0.035
Number of observations: 2,340 tolerance: 5.15e-08				
Likelihood ratio chi-sqare 88.779 with prob 0.000				
Percentage of correct predictions 96.197				

Table 3.8: LOGIT for unemployed Blacks (move 3.89 %)

a graphical interpretation of the results of table 3.7.

This result can be explained, in the framework of our model, as a two equilibria situation. Comparing the equilibrium where people move more (q=1) with the equilibrium where people move less (q=0), we obtained that, in explaining the decision of moving, family situation is less significant, and the unemployment status is more significant. This is precisely what we find for these two groups. In this way, we do not need to postulate different preferences for these two groups [19].

Table 3.8 presents the results of a regression that uses as sample only unemployed Blacks. It is interesting to note that there is not much of an increase in the migration rate due to unemployment (from 3.68 to 3.89 %), unlike the increase seen among Whites (from 5.31 to 9.01 %).

All the coefficients have plausible signs with the exception of the local unemployment rate [20].

[19]Bowles (1970) writes: "this finding is consistent with the notion that the stability of an unequal income distribution may be explained in part by sociallly generated attributes - risk aversion and high rates of time preference, for example - which inhibit black people from taking advantage of those avenues for higher incomes, such as education and geographical mobility".

[20]Probably, in this regression, there is an issue of endogeneity since the low migration rate in very poor areas (i.e. inner cities) implies a very high level of unemployment, hence the positive correlation found.

	North East	North Central	South	West
Observations	6,476	8,480	9,880	6,154
MOVE	0.0375	0.0520	0.0563	0.0647
FFF	0.8076	0.7702	0.7868	0.7369
MV1	0.0324	0.0445	0.0668	0.0720
OWN	0.6677	0.7635	0.7153	0.6601
EDH	0.2855	0.2349	0.2036	0.2447
FINC	0.2644	0.2871	0.2845	0.2922
AGE1	0.2302	0.2256	0.2260	0.2119
INC1	0.1961	0.1925	0.2480	0.1817
INC2	0.2123	0.2097	0.2084	0.2065
INC3	0.1467	0.1681	0.1269	0.1776
MAR	0.6875	0.7261	0.7398	0.7030
UNEM	0.0460	0.0372	0.0371	0.0470

Table 3.9: Statistics

3.4.4 Regional Differences within the United States

Macro regions [21] (North-East, North-Central, South and West) show substantial differences in mobility patterns. In the years 1979-1988 only 3.53 % of the population in the North-East moved, while in the West, 6.58 % of the population moved. The South and North-Central show intermediate values (4.37 % and 4.89 %) [22]. Blanchard and Katz (1992) found that there are substantial differences in the effect of a local shock across macro regions.

These facts call for a disaggregated analysis of the determinants of migration across regions. In the following regression (table 3.10) we constrain all the control variables to have the same coefficient across regions. The South, where the larger number of individuals is concentrated, is used as base: we allow for regional effects of the family variable and of the unemployment status.

With the notable the exception of the unemployment status in the North East, the dummy variables for the regions are not significant. The effects of family and employment status are similar across South, North Central and West. The only major

[21] In the appendix, we give the definitions of the 4 macro regions used in this paper.

[22] These data come from authors' elaborations using PSID waves 1979-1988. We use a conservative definition of move: if an individual moved from a rural county to another rural county within the same state. Since the states in the West are on average larger than the states in North-East, the difference in the migration rates is probably even larger.

	Logit Coeff.	Marg. Effect	t-stat	> $\|t\|$
CONST	-1.46	-0.05	-13.4	0.000
FFF	-1.57	-0.05	-16.3	0.000
MV1	1.029	0.037	13.35	0.000
OWN	-0.84	-0.03	-15.2	0.000
EDH	0.283	0.010	4.638	0.000
FINC	0.218	0.008	3.630	0.000
AGE1	0.475	0.017	7.157	0.000
INC1	-0.19	-0.00	-2.75	0.006
INC2	-0.26	-0.00	-3.38	0.001
INC3	-0.21	-0.00	-2.33	0.019
MAR	-0.13	-0.00	-2.05	0.040
UNEM	0.447	0.016	2.188	0.029
fff * NE	-0.46	-0.01	-3.90	0.000
unem * NE	-0.10	-0.00	-0.30	0.757
fff * NC	0.087	0.003	0.872	0.383
unem * NC	-0.09	-0.00	-0.33	0.737
fff * W	0.048	0.002	0.439	0.661
unem * W	0.080	0.003	0.276	0.782
Number of observations: 31,093 tolerance: 4.28e-10				
Likelihood ratio chi-sqare 1497.877 with prob 0.000				
Percentage of correct predictions 94.703				

Table 3.10: LOGIT for US white males by regions

exception is given by the family variable in the North East. People in the North East appear to weigh family attachment more heavily than the average American. [23]

3.5 Conclusion

At the beginning of the paper we asked whether American labor mobility can be explained as the aggregate consequence of different equilibria. We developed a model of double matching both in the labor market and in the social environment, that can generate two equilibria. Using micro data, we found that social variables do play an important role in the migration decisions. We also found that blacks and whites differ significantly in the migration patterns.

In the first section about the determinants of migration, we referred to the "standard view" of considering the decision to move as a decision of the individual to invest in human capital. As other theoretical and empirical papers have shown, among the costs of moving there is the "family detachment".

The point stressed in this paper is that this cost is not constant but depends on the other people's behavior. The resulting externality can generate the two equilibria of the model. In this manner, we can explain the behavior of different communities as well as the persistence of migration patterns.

The fact that communities respond differently to economic incentives has important implications for macroeconomists. If a region has shown low labor regional mobility in the past, it is probable to expect low mobility also in the future. This is quite pessimistic about European regional unemployment. Europeans seem to be in the equilibrium where (q=0), so they are less responsive to economic incentives. Therefore, a policy which hinges on economic incentives will have a smaller effect.

[23]The case of the West is quite interesting: even if there is a much larger mobility than in the rest of the United States, the preferences do not appear significantly different from the American average. The difference is mostly due to the different average level of *fff* in the West respect to the average of *fff* in the United States.

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Appendix A

Computation of $\frac{\partial}{\partial r}(\rho_2 - \rho_1)$

This appendix shows how the range in (3.11) varies with r.

$$
\begin{aligned}
\frac{\partial}{\partial r}(\rho_2 - \rho_1) &= \frac{(1-\pi_e)\pi_s - \pi_u}{[r+\pi_e+(1-\pi_e)\pi_s]^2} - \frac{(1-\pi_e)(\pi_s-\pi_u)}{[r+\pi_e+(1-\pi_e)(\pi_s+\pi_e)]^2} \\
&= \{[(1-\pi_e)\pi_s - \pi_u][r+\pi_e+(1-\pi_e)(\pi_s+\pi_e)]^2 \\
&\quad -(1-\pi_e)(\pi_s-\pi_u)[r+\pi_e+(1-\pi_e)\pi_s]^2\} \\
&\quad \times[r+\pi_e+(1-\pi_e)\pi_s]^{-2}[r+\pi_e+(1-\pi_e)(\pi_s+\pi_e)]^{-2} \\
&= \{[(1-\pi_e)\pi_s - \pi_u] \\
&\quad \times\{[r+\pi_e+(1-\pi_e)\pi_s]^2 \\
&\quad +2[r+\pi_e+(1-\pi_e)\pi_s][(1-\pi_e)\pi_e)] + [(1-\pi_e)\pi_e]^2\} \\
&\quad -(1-\pi_e)(\pi_s-\pi_u)[r+\pi_e+(1-\pi_e)\pi_s]^2\} \\
&\quad \times\{[r+\pi_e+(1-\pi_e)\pi_s][r+\pi_e+(1-\pi_e)(\pi_s+\pi_e)]\}^{-2} \\
&= -\{\pi_u\pi_e[r+\pi_e+(1-\pi_e)\pi_s]^2 \\
&\quad +2[r+\pi_e+(1-\pi_e)\pi_s][\pi_u-(1-\pi_e)\pi_s](1-\pi_e)\pi_e \\
&\quad +[\pi_u-(1-\pi_e)\pi_s][(1-\pi_e)\pi_e]^2\} \\
&\quad \times\{[r+\pi_e+(1-\pi_e)\pi_s][r+\pi_e+(1-\pi_e)(\pi_s+\pi_e)]\}^{-2}
\end{aligned}
$$

The sign of $\frac{\partial}{\partial r}(\rho_2 - \rho_1)$ is that of the numerator. The numerator is a quadratic function in r, with negative coefficient on r^2. Thus, for all r above some value $\overline{r}$, $\frac{\partial}{\partial r}(\rho_2 - \rho_1) < 0$.

In particular, if $\pi_u > (1 - \pi_e)\pi_s$, the numerator is negative for any positive value of r. Therefore $\frac{\partial}{\partial r}(\rho_2 - \rho_1) < 0$ for all $r \geq 0$.

Appendix B

Definition of the Variables in Chapter 3

The variable *move* is measured at time *t+1*, the other variable are measured at time *t*.

- The PSID records each individual's county of residency annually. To take into account that many moves are across different counties, but within the same metropolitan area, counties are aggregated up to SMSAs (Standard Metropolitan Statistical Areas). Moves are defined by comparing residence information of two consecutive years according to the following criteria:

1. a movement from a SMSA to another SMSA is considered a move;
2. a movement from a rural area to a SMSA (or vice versa) is considered a move;
3. a movement from a rural area to another rural area is considered a move only if the two areas are in different states.

Out of 37,528 individuals, 24 per cent moved at least once. Table B.1 divides the individuals of the PSID sample according the number of moves in the period 1969-1988.

- *mv1* is a dummy variable that is assigned the value 1 if the individual moved in the previous year and 0 if otherwise.

- *edh* is a dummy variable that is assigned the value 1 if the head of the family as defined in PSID has a college degree, 0 if otherwise. This variable is meant to indicate the level of education of the family which previous studies found to be relevant (i.e. DaVanzo (1978)).

- *age1* is a dummy variable which is assigned the value 1 if the individual is less than 24 years old and 0 if otherwise [1].

- *frf* is the number of past moves of the members of the same extended family excluding the individual over the total numbers of past individual-years of the extended family excluding the individual. It is 1 if everybody in the extended family moved in every year in the past. It is 0 if nobody ever moved. This variable controls for the history of the family.

- *unem* is a dummy variable which is assigned the value 1 if the individual is unemployed and if the unemployment rate of the county of residence is more than 4 %; 0 if otherwise. [2] [3]

- *black* is a dummy variable which is assigned the value of 1 if the individual is Black.

- The maximum of the total income (deflated for inflation) of each individual in the last four years is computed. If the income in year *t* is less the 75 per cent of the maximum income in the last 4 years, and the individual is not unemployed, then *finc* is assigned the value of 1.

- *own* is assigned the value of 1 if the individual lives in a home that belongs to the family, 0 if otherwise.

[1] The use of a dummy variable for young people is the best way to encompass effect of age on migration decision

[2] Unemployed are more probable to move than unemployed, and that higher local unemployment rates encourage migration of the unemployed, but exert little influence on those that have a job.

[3] The state unemployment rate is used when the county unemployment rate is not available.

- *inc0, inc1, inc2* and *inc3* are three dummy variables for the income level. *inc0* (*inc1* , *inc2*, *inc3*) is a dummy variable which is assigned the value of 1 if the income of the individual measured in 1988 constant dollars is between \$ 0 and \$ 10,000 (or respectively between \$ 10000 and \$ 20,000, \$ 20,000 and \$ 30,000 , \$ 30,000 and \$ 40,000).

- *mar* is a dummy variable for the marital status: it is assigned the value of 1 if the individual is married or lives stably with a companion, 0 if otherwise.

- *fffb* is obtained by multiplying *fff* and *black*; so it is equal to the family value *fff* if the individual is Black, 0 if otherwise.

- *uneb* is obtained by multiplying *unem* and *black*; so it is equal to 1 if the individual is Black and unemployed and he lives in an area with unemployment larger than 4 %, 0 if otherwise.

The variables are not available for all the years. Unemployment status is available for each individual since 1979, so all the regressions which use this variable refer to the period 1979-1988. Moreover, just males greater than 16 years and less than 70 years are considered. Given these selections criteria the sample size is of 57,025 "person-years" observations. Table B.2 gives the statistics for the sample used:

Number of Moves	Number	percentage
0	28,644	76.33
1	4,405	11.74
2	2,647	7.05
3	1,000	2.66
4	447	1.19
5	239	0.64
6	77	0.21
7	35	0.09
8	30	0.08
9	3	0.00
10	1	0.00

Authors' calculations based on the definition of "move" given above.

Table B.1: Distribution of the moves

Variables	**Mean**	**Std. Dev.**	**Minimum**	**Maximum**
MOVE	0.0510	0.2200	0	1
FFF	0.8052	0.2927	0	1
MV1	0.0477	0.2130	0	1
OWN	0.6207	0.4852	0	1
EDH	0.1722	0.3776	0	1
f inc	0.2882	0.4529	0	1
AGE1	0.2643	0.4410	0	1
INC1	0.2247	0.4174	0	1
INC2	0.1818	0.3857	0	1
INC3	0.1171	0.3215	0	1
MAR	0.6410	0.4797	0	1
BLACK	0.3549	0.4785	0	1
UNEM	0.0726	0.2594	0	1
fffb	0.3048	0.4384	0	1
unemb	0.0461	0.2097	0	1

Table B.2: Summary Statistics

Appendix C

Additional Estimation Results

	Probit coeff.	St. error	t-value	$p > \|t\|$
CONSTANT	-0.90569	0.0486	-18.62	0.000
FFF	-0.80188	0.0389	-20.60	0.000
MV1	0.58075	0.0390	14.90	0.000
OWN	-0.40843	0.0268	-15.25	0.000
EDH	0.13472	0.0294	4.58	0.000
finc	0.10406	0.0288	3.61	0.000
AGE1	0.23038	0.0330	6.98	0.000
INC1	-0.10172	0.0340	-3.00	0.003
INC2	-0.13410	0.0371	-3.62	0.000
INC3	-0.10781	0.0426	-2.53	0.011
MAR	-0.05207	0.0317	-1.64	0.101
UNEM	0.21646	0.0559	3.87	0.000
Number of observations: 31,191				
Percentage of correct predictions 94.652				

Table C.1: PROBIT for US White males (move 5.33 %)

	Logit Coeff.	Marg. Effect	t-stat	> $\|t\|$
CONSTA	-0.907	-0.082	-4.463	0.000
FFF	-1.433	-0.129	-9.703	0.000
MV1	0.329	0.030	2.363	0.018
OWN	-0.396	-0.036	-3.612	0.000
EDH	-0.182	-0.016	-1.559	0.119
FINC	-0.039	-0.003	-0.321	0.748
AGE1	0.665	0.060	5.091	0.000
INC1	-0.044	-0.004	-0.320	0.749
INC2	-0.198	-0.018	-1.303	0.193
INC3	-0.324	-0.029	-1.751	0.080
MAR	-0.150	-0.013	-1.141	0.254
UNEM	0.140	0.013	0.614	0.539
Number of observations: 3,876 tolerance: 4.44e-16				
Likelihood ratio chi-sqare 193.375 with prob 0.000				
Percentage of correct predictions 88.287				

Table C.2: LOGIT for nn=2 (move 11.71 %)

	Logit Coeff.	Marg. Effect	t-stat	> $\|t\|$
CONSTA	-1.017	-0.147	-4.517	0.000
FFF	-0.694	-0.100	-3.826	0.000
MV1	0.053	0.008	0.342	0.732
OWN	-0.590	-0.085	-4.460	0.000
EDH	0.166	0.024	1.249	0.212
FINC	0.142	0.021	1.023	0.306
AGE1	0.372	0.054	2.302	0.021
INC1	0.170	0.025	1.036	0.300
INC2	0.143	0.021	0.803	0.422
INC3	-0.062	-0.009	-0.297	0.767
MAR	-0.084	-0.012	-0.529	0.597
UNEM	0.120	0.017	0.371	0.711
Number of observations: 1,866 tolerance: 7.51e-13				
Likelihood ratio chi-sqare 58.760 with prob 0.000				
Percentage of correct predictions 81.565				

Table C.3: LOGIT for nn=3 (move 18.44 %)

	Logit Coeff.	Marg. Effect	t-stat	$>\|t\|$
CONSTA	-0.921	-0.157	-2.881	0.004
FFF	-0.833	-0.142	-3.314	0.001
MV1	0.316	0.054	1.755	0.080
OWN	-0.534	-0.091	-3.180	0.002
EDH	0.120	0.020	0.635	0.525
FINC	0.167	0.028	0.941	0.347
AGE1	1.013	0.173	4.472	0.000
INC1	-0.157	-0.027	-0.746	0.456
INC2	-0.234	-0.040	-1.013	0.311
INC3	-0.759	-0.129	-2.549	0.011
MAR	0.222	0.038	0.975	0.330
UNEM	0.302	0.051	1.032	0.302
Number of observations: 971 tolerance: 2.22e-16				
Likelihood ratio chi-sqare 73.865 with prob 0.000				
Percentage of correct predictions 77.034				

Table C.4: LOGIT for nn=4 (move 23.69 %)

	Logit Coeff.	Marg. Effect	t-stat	$>\|t\|$
CONSTA	-0.529	-0.111	-1.943	0.052
FFF	-0.565	-0.118	-2.564	0.010
MV1	0.483	0.101	3.202	0.001
OWN	-0.400	-0.084	-2.560	0.011
EDH	-0.211	-0.044	-1.232	0.218
FINC	0.126	0.026	0.776	0.438
AGE1	0.334	0.070	1.705	0.088
INC1	0.094	0.020	0.483	0.629
INC2	0.054	0.011	0.247	0.805
INC3	0.172	0.036	0.725	0.469
MAR	-0.191	-0.040	-1.022	0.307
UNEM	0.499	0.104	1.836	0.067
Number of observations: 990 tolerance: 8.77e-15				
Likelihood ratio chi-sqare 48.099 with prob 0.000				
Percentage of correct predictions 69.293				

Table C.5: LOGIT for $nn >= 5$ (move 30.61 %)

North East	North Central	South	West
Connecticut	Illinois	Alabama	Arizona
Maine	Indiana	Arkansas	California
Massachusetts	Iowa	Delaware	Colorado
New Hampshire	Kansas	Florida	Idaho
New Jersey	Michigan	Georgia	Montana
New York	Minnesota	Kentucky	Nevada
Pennsylvania	Missouri	Louisiana	New Mexico
Rhode Island	Nebraska	Maryland	Oregon
Vermont	North Dakota	Mississippi	Utah
	Ohio	North Carolina	Washington
	South Dakota	Oklahoma	Wyoming
	Wisconsin	South Carolina	
		Tennessee	
		Texas	
		Virginia	
		Washington, D.C.	
		West Virginia	

Table C.6: Definition of Regions

	Logit Coeff.	Marg. Effect	t-stat	$> \|t\|$
CONSTA	-1.878	-0.041	-6.598	0.000
FFF	-1.740	-0.038	-8.153	0.000
MV1	1.172	0.026	5.528	0.000
OWN	-0.836	-0.018	-6.063	0.000
EDH	0.365	0.008	2.496	0.013
f inc	0.314	0.007	2.103	0.035
AGE1	0.628	0.014	3.745	0.000
INC1	-0.090	-0.002	-0.533	0.594
INC2	-0.411	-0.009	-2.031	0.042
INC3	-0.256	-0.006	-1.102	0.271
MAR	-0.068	-0.001	-0.425	0.671
BLACK	0.343	0.008	0.918	0.359
UNEM	0.386	0.008	1.401	0.161
fffb	-1.651	-0.036	-2.851	0.004
unemb	0.148	0.003	0.226	0.821

Number of observations: 7,667 tolerance: 2.22e-16
Likelihood ratio chi-sqare 300.078 with prob 0.000
Percentage of correct predictions 96.413

Table C.7: LOGIT for North East males (move 3.53 %)

	Logit Coeff.	Marg. Effect	t-stat	> $\|t\|$
CONSTA	-1.458	-0.047	-7.338	0.000
FFF	-1.334	-0.043	-8.332	0.000
MV1	1.277	0.042	8.907	0.000
OWN	-0.869	-0.028	-8.911	0.000
EDH	0.343	0.011	3.147	0.002
finc	0.081	0.003	0.775	0.438
AGE1	0.489	0.016	4.467	0.000
INC1	-0.222	-0.007	-1.770	0.077
INC2	-0.335	-0.011	-2.338	0.019
INC3	-0.087	-0.003	-0.552	0.581
MAR	-0.325	-0.011	-2.999	0.003
BLACK	0.761	0.025	3.012	0.003
UNEM	0.245	0.008	1.137	0.255
fffb	-1.644	-0.054	-4.978	0.000
unemb	-0.347	-0.011	-1.040	0.299
Number of observations: 11,553 tolerance: 5.55e-15				
Likelihood ratio chi-sqare 569.752 with prob 0.000				
Percentage of correct predictions 94.980				

Table C.8: LOGIT for North Central males (move 4.89 %)

	Logit Coeff.	Marg. Effect	t-stat	> $\|t\|$
CONSTA	-1.630	-0.046	-10.18	0.000
FFF	-1.651	-0.046	-12.00	0.000
MV1	1.294	0.036	12.471	0.000
OWN	-0.669	-0.019	-9.063	0.000
EDH	0.288	0.008	2.997	0.003
finc	0.257	0.007	3.191	0.001
AGE1	0.520	0.015	6.156	0.000
INC1	-0.061	-0.002	-0.684	0.494
INC2	0.025	0.001	0.231	0.817
INC3	-0.093	-0.003	-0.642	0.521
MAR	-0.185	-0.005	-2.180	0.029
BLACK	-0.127	-0.004	-0.884	0.377
UNEM	0.512	0.014	2.450	0.014
fffb	-0.537	-0.015	-2.748	0.006
unemb	-0.391	-0.011	-1.461	0.144
number of observations: 21,576 tolerance: 5.87e-13				
likelihood ratio chi-sqare 1,058.572 with prob 0.000				
percentage of correct predictions 95.620				

Table C.9: LOGIT for South males (move 4.37 %)

	Logit Coeff.	Marg. Effect	t-stat	> $\|t\|$
CONSTA	-1.192	-0.054	-6.115	0.000
FFF	-1.662	-0.075	-10.52	0.000
MV1	0.729	0.033	5.403	0.000
OWN	-0.903	-0.041	-8.429	0.000
EDH	0.154	0.007	1.283	0.199
finc	0.134	0.006	1.186	0.236
AGE1	0.264	0.012	2.171	0.030
INC1	-0.146	-0.007	-1.134	0.257
INC2	-0.408	-0.018	-2.771	0.006
INC3	-0.281	-0.013	-1.637	0.102
MAR	-0.046	-0.002	-0.392	0.695
BLACK	0.343	0.015	1.405	0.160
UNEM	0.470	0.021	2.211	0.027
fffb	-0.793	-0.036	-2.195	0.028
unemb	-0.413	-0.019	-1.137	0.256

number of observations: 7,448 tolerance· 8.41e-12
likelihood ratio chi-sqare 407.609 with prob 0.000
percentage of correct predictions 93.434

Table C.10: LOGIT for West males (move 6.58 %)

	Logit Coeff.	Marg. effect	t-stat	> $\|t\|$
CONSTA	-2.126	-0.072	-17.33	0.000
FRF	1.748	0.059	5.038	0.000
FFF	-1.351	-0.046	-16.38	0.000
MV1	0.904	0.031	11.430	0.000
MVX2	0.676	0.023	10.088	0.000
OWN	-0.733	-0.025	-12.76	0.000
EDH	0.294	0.010	4.721	0.000
FINC75	0.294	0.010	3.925	0.000
AGE1	0.566	0.019	7.927	0.000
MAR	-0.142	-0.005	-2.089	0.037
UNEM	0.422	0.014	3.666	0.000
INCO	0.074	0.003	1.087	0.277

number of observations: 31181 tolerance: 1.11e-16
likelihood ratio chi-sqare 1563.951 with prob 0.000
percentage of correct predictions 94.833

Table C.11: LOGIT for US (white males)

	Coeff.	t-stat	> $\|t\|$
CONSTANT	0.106	14.841	0.000
FFF	-0.070	-16.096	0.000
FRF	0.051	2.629	0.009
OWN	-0.039	-13.694	0.000
MAR	0.001	0.377	0.706
MV1	0.115	19.625	0.000
MVX2	0.059	13.991	0.000
INC0	0.011	2.106	0.035
INC1	-0.000	-0.161	0.871
INC2	-0.003	-0.657	0.511
INC3	-0.002	-0.535	0.592
INC4	0.002	0.385	0.700
EDH	0.011	3.810	0.000
AGE1	0.028	7.469	0.000
UNEMPLO	0.019	3.288	0.001
31,181 observations			
F(15,31166) = 158.631; P-value = 0.00			

Table C.12: OLS (white only)

	Coeff.	t-stat	> $\|t\|$
FFF	-0.086	-10.376	0.000
FRF	-1.281	-18.428	0.000
OWN	-0.057	-12.053	0.000
MAR	0.003	0.553	0.580
MV1	-0.061	-9.750	0.000
MVX2	-0.075	-14.653	0.000
INC0	0.022	2.705	0.007
INC1	0.018	2.377	0.017
INC2	0.013	1.762	0.078
INC3	0.007	1.070	0.284
INC4	0.005	0.690	0.490
EDH	0.001	0.178	0.859
AGE1	0.002	0.542	0.588
UNEMPLO	0.017	2.567	0.010
31,181 observations; 4,546 individuals			
F(14,26621) = 67.46; P-value = 0.00			
F-statistic for equality of dummy variables:			
F(4545, 26621) = 1.869; P-value: 0.00			

Table C.13: OLS with fixed effects (white only)

	Coeff.	t-stat	> $\|t\|$
CONSTANT	0.163	16.470	0.000
FFF	-0.092	-14.419	0.000
FRF	-0.145	-4.051	0.000
OWN	-0.055	-14.137	0.000
MAR	0.000	0.054	0.957
MV1	-0.017	-3.047	0.002
MVX2	-0.040	-8.676	0.000
INC0	0.015	2.243	0.025
INC1	0.007	1.118	0.263
INC2	0.003	0.560	0.575
INC3	0.000	0.140	0.888
INC4	0.002	0.353	0.724
EDH	0.013	2.764	0.006
AGE1	0.016	3.639	0.000
UNEMPLO	0.020	3.273	0.001
31,181 observations; 4,546 individuals			
F(15,31166) = 63.66; P-value = 0.00			

Table C.14: OLS with random effects (white only)

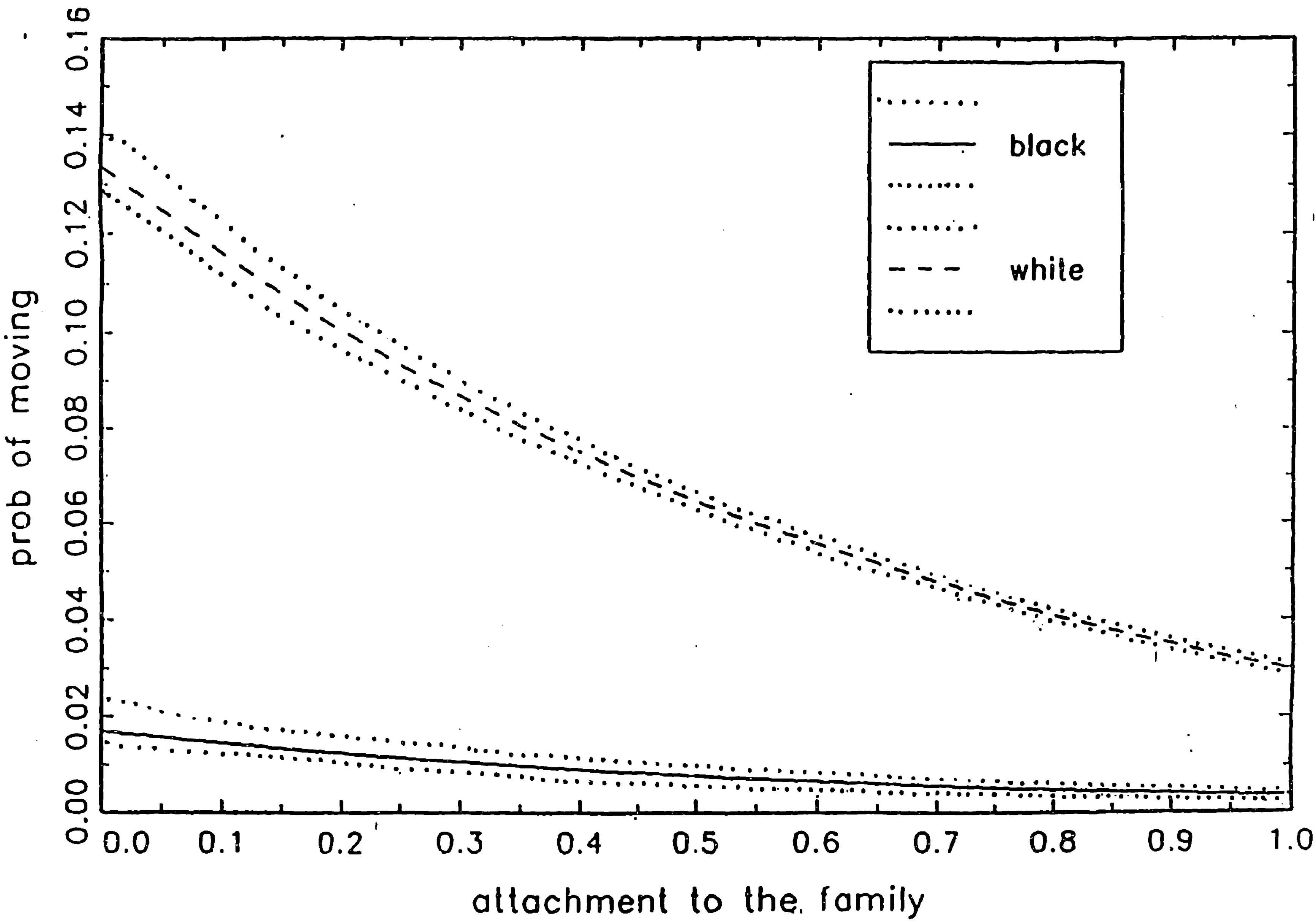
Appendix D

Figures

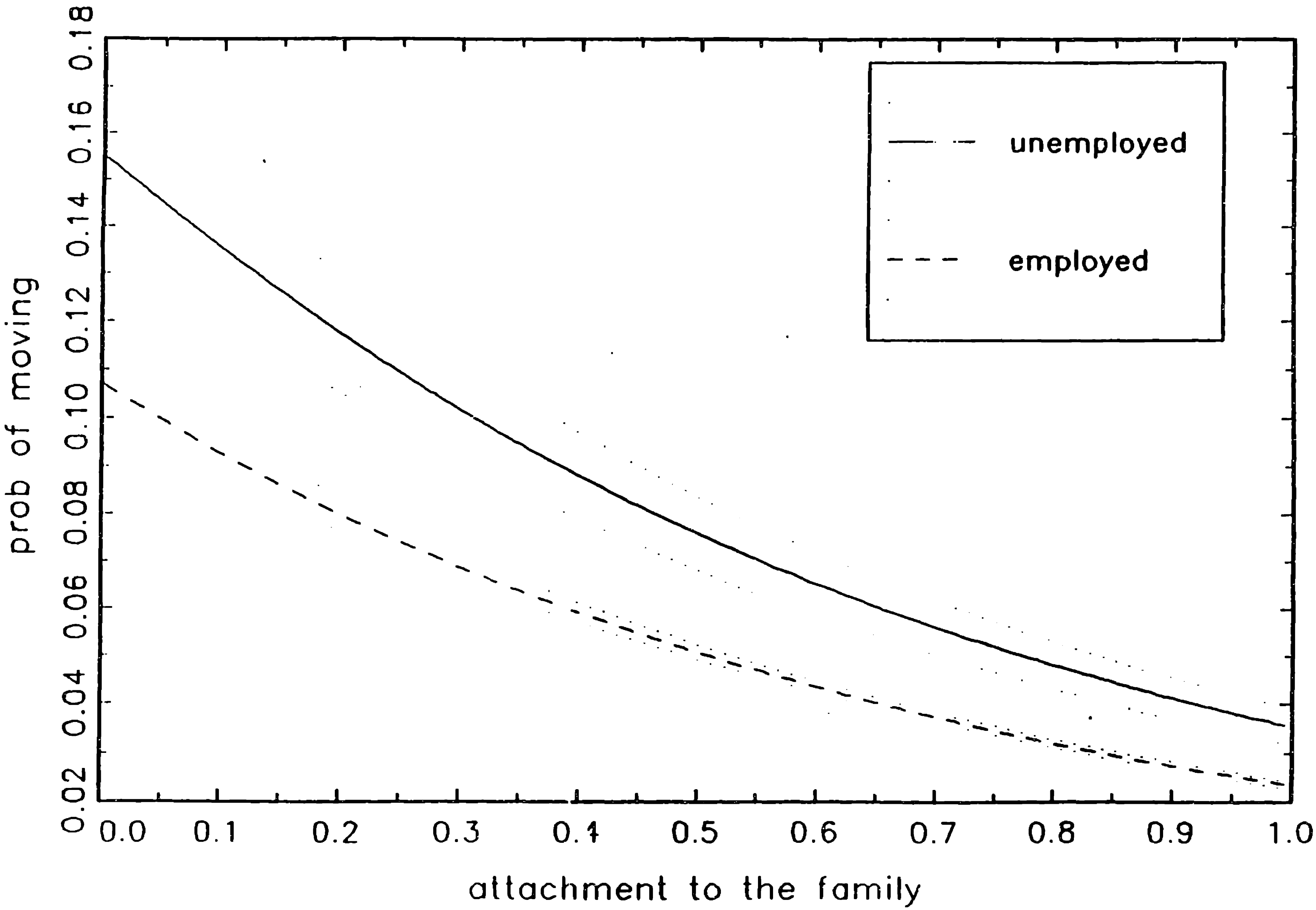
The first three figures represent the estimated marginal effect on the probability of moving of family attachment, race, unemployment, and a fall in income. These are based on the parameter estimates reported in tables 3.3 and 3.7. The logit equation is estimated at the average value of the independent variables.

The last three figures present the distributions of family size, and family attachment for blacks and whites.

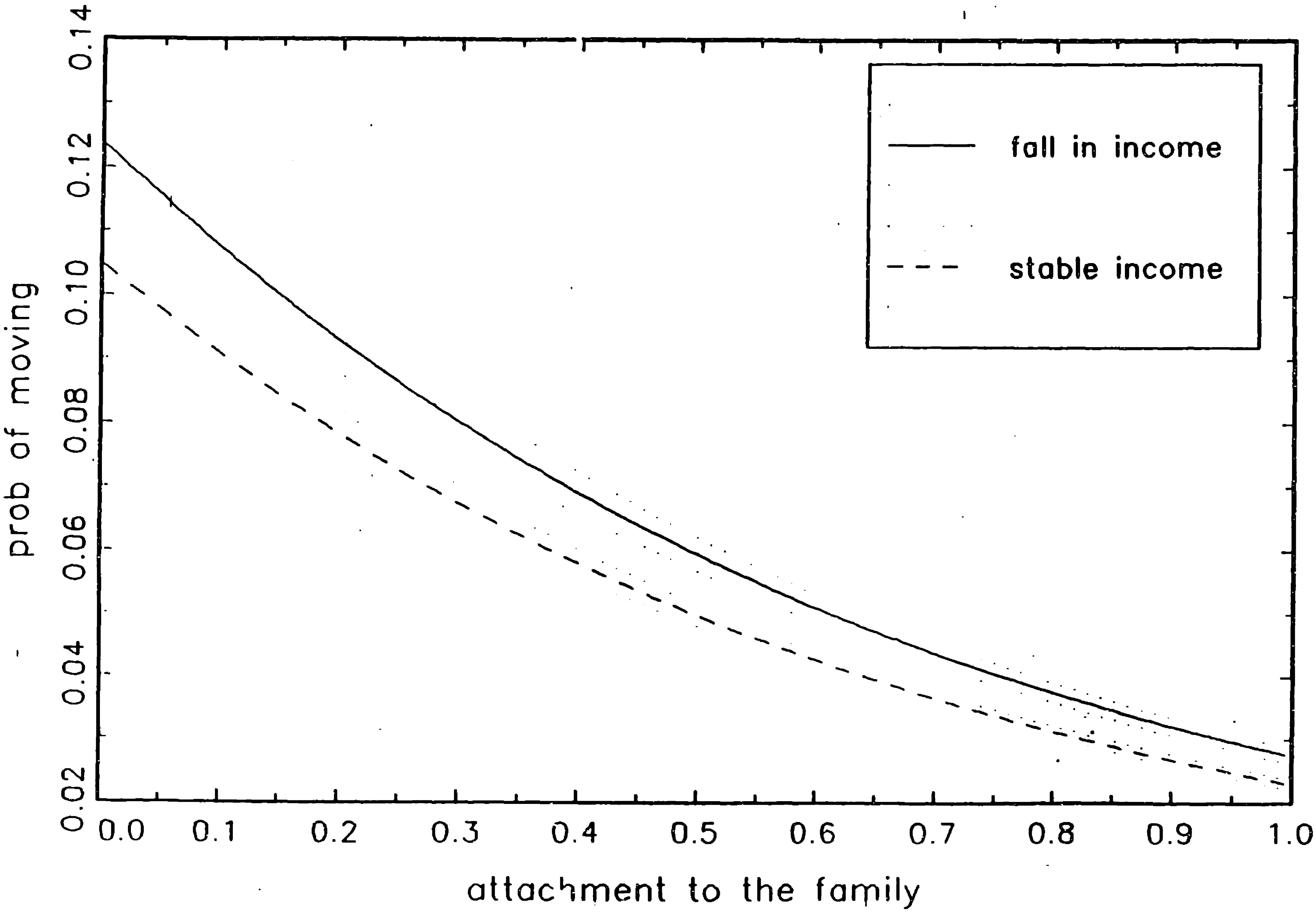
DIFFERENT BEHAVIOR AMONG SOCIAL GROUPS



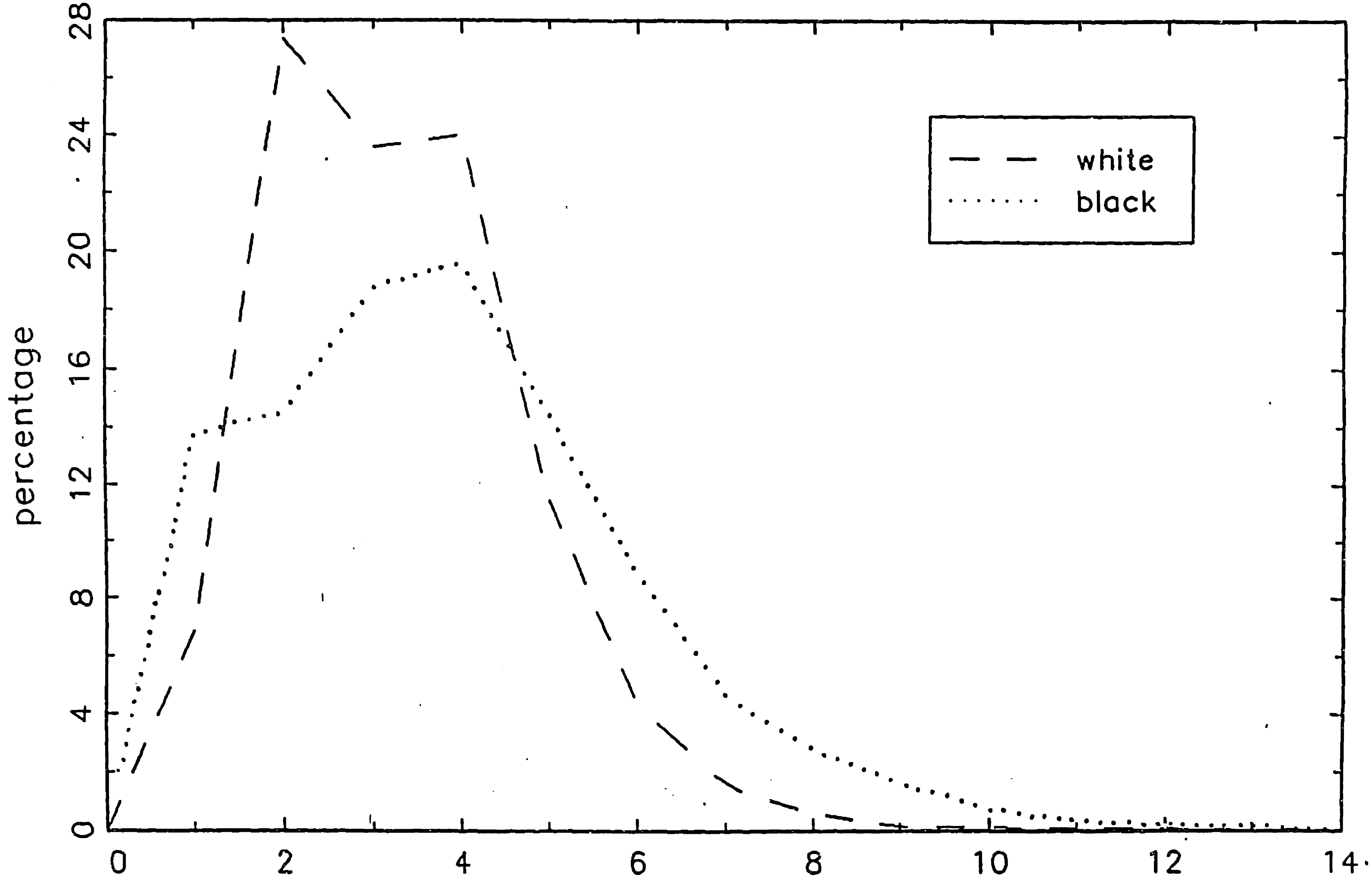
EFFECT OF UNEMPLOYMENT AND FAMILY ON THE PROBABILITY OF MOVING

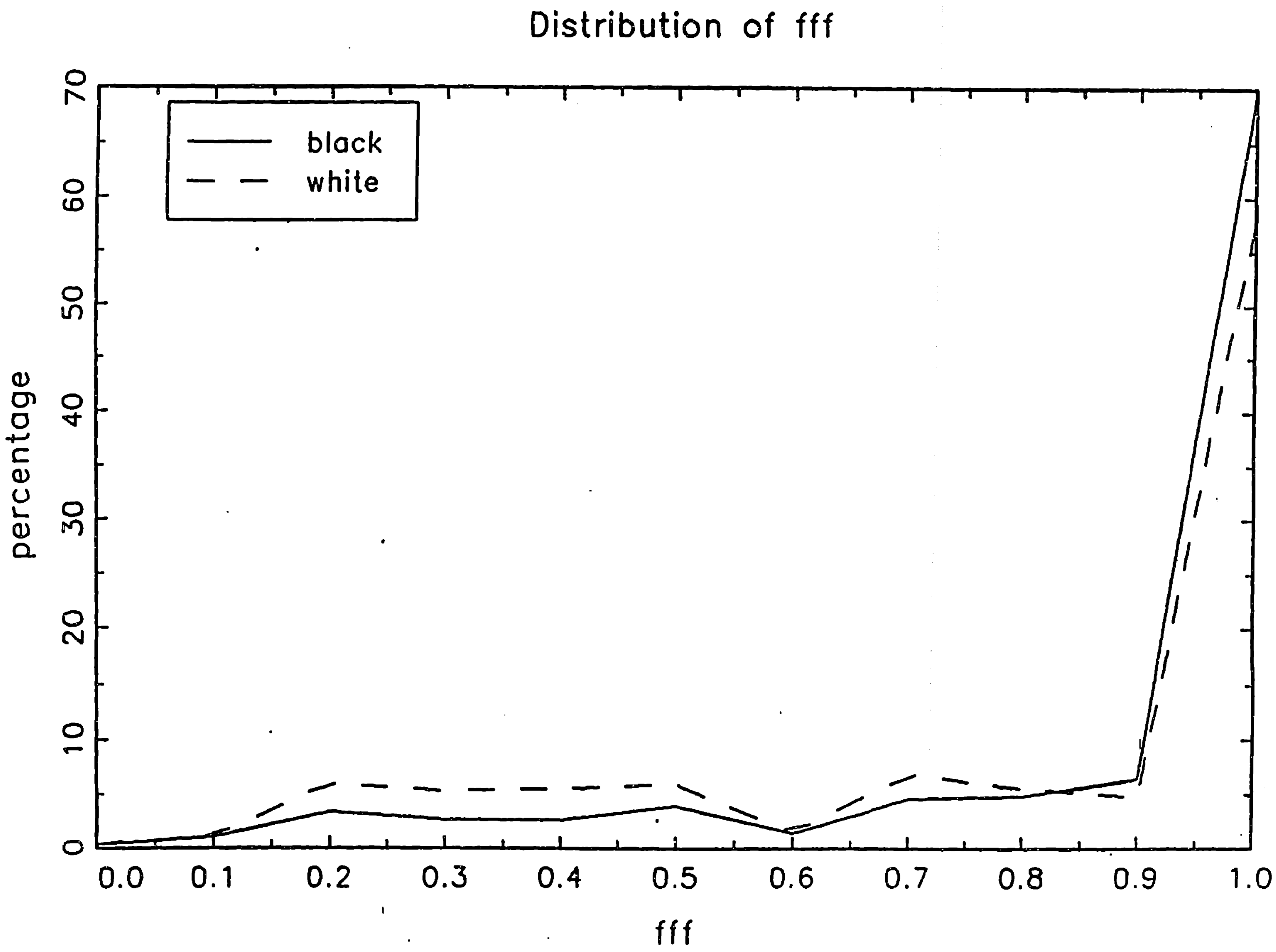


EFFECT OF A FALL IN INCOMF AND FAMILY ON THE PROBABILITY OF MOVIN(



Distribution of size




Distribution of fff
black
white
percentage
0
10
20
30
40
50
60
70
0.0
0.1
0.2
0.3
0.4
0.5
0.6
0.7
0.8
0.9
1.0
fff

Distribution of fff (White males)