Religious Intermarriage and Socialization in the United States

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This paper presents an empirical analysis of a choice-theoretic model of cultural transmission. In particular, we use data from the General Social Survey to estimate the structural parameters of a model of marriage and child socialization along religious lines in the United States. The observed intermarriage and socialization rates are consistent with Protestants, Catholics, and Jews having a strong preference for children who identify with their own religious beliefs and making costly decisions to influence their children's religious beliefs. Our estimates imply dynamics of the shares of religious traits in the population that are in sharp contrast with the predictions obtained by linear extrapolations from current intermarriage rates.

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I. Introduction

Since the 1950s a rich sociological literature has documented very low intermarriage rates along the religious dimension in the United States (see, e.g., Landis 1949; Thomas 1951) and predicted, as a consequence, a low rate of assimilation of immigrants in the United States. For instance, Will Herberg noted in his classic (1955) study of interfaith marriages that nothing seemed to suggest the assimilation of immigrants in the United States into a "melting pot" extending across the "three great faiths" (Protestant, Catholic, and Jewish). Ruby Kennedy (1944), facing similar evidence of low intermarriage rates in New Haven, Connecticut, in the period 1870–1950, introduced the metaphor of the convergence to a "triple melting pot" along the religious dimension.[1]

More recently, though, a documented rise in intermarriage unions for Jews (see the Council of Jewish Federation's 1990 National Jewish Population Survey) has spurred an intense debate about the prediction of the vanishing of American Jews (see Dershowitz [1997] for a detailed report on the debate and Chiswick [1999] for a critical discussion of such predictions in the context of a human capital accumulation model of religion).

Most predictions of this sort are obtained by simple linear extrapolation of demographic and sociological trends, assuming constant intermarriage rates in the future. The most sophisticated sociological analyses of the dynamic implications of data on interfaith marriage rates account for the distribution of the population by religious group. These analyses assume that a member of a minority religious community finds it more difficult to meet a spouse who shares his or her religious faith (Heer and Hubay [1975] and Johnson [1980] are examples of such analyses). By conditioning on the distribution of the population, these studies estimate, for members of each religious group, an unobserved component of their marriage choices, called "intrinsic homogamy" or "segregation effort," which drives homogamy rates. The studies then construct linear extrapolations of the dynamics of the distribution of the population maintaining these components constant.

But if intrinsic homogamy or segregation effort is the result of the choices of individual agents in the marriage market, then they should depend on the distribution of the population by religious trait. For instance, individuals in a minority religious community might compensate for their status by segregating in marriage more intensely. In this case, extrapolating from estimated measures of intrinsic homogamy or

[1] In fact, various historical evidence on the behavior of immigrants in the United States seems to imply that their assimilation into a "melting pot" has progressed very slowly along several dimensions other than the religious dimension (see, e.g., Glazer and Moynihan 1963; Gordon 1964; Mayer 1979).

segregation effort would underestimate the resilience of minority religious groups.

To evaluate the empirical relevance of the dependence of marriage choices on the distribution of the population by religious group, we construct a model of marriage segregation along religious lines (religious homogamy). Segregation effort is determined endogenously by the institutional characteristics of the marriage market and by the preference parameters of individuals by religious group. We estimate these parameters using U.S. survey data, over the period 1972–96, and simulate the dynamics of the distribution of the population by religious group at the estimated parameter values.

In light of our estimation and simulations, we conclude that linear extrapolations of intermarriage rates, even after one conditions on the distribution of the population, are severely misleading. The dependence of marriage rates on the distribution of the population by religious trait displays in fact substantial nonlinearities. Once such nonlinearities are taken into account, the simulations do not appear to support the triple melting pot hypothesis. Also, minorities do, in fact, segregate in marriage more intensely than majorities, and they socialize their children more strictly. As a consequence, Jews are not assimilated at the rapid rates implied by linear extrapolations.[2]

Our analysis of the determinants of marriage rates across the religious dimension is based on Becker's early contributions on the economics of marriage (Becker 1973, 1974, 1981). He shows that positive assortative marriages, or "marriage of the likes," arise as equilibria "when such pairings maximize aggregate … output over all marriages, regardless of whether the trait is financial …, biological …, or psychological" (1981, pp. 70–71). Many reasons can be given along these lines for the religious assortativeness of marriage in the United States, not the least of which is that homogamous marriages are more stable; that is, they have lower divorce rates (Becker, Landes, and Michael 1977; Heaton 1984; Lehrer and Chiswick 1993). We evaluate one particular explanation of the assortativeness of marriage along the religious dimension—an explanation that emphasizes the link between marriage choices and the socialization of children to their parents' religious beliefs.

In our model, parents have a taste and a technology for transmitting their own religious faith to their children. Moreover, families that are homogamous with respect to their religious beliefs are endowed with a

[2] The sociological literature contains several instances of the underestimation of the resilience of minorities. The transformation of ethnic neighborhoods in the 1960s has led, e.g., several sociologists to extrapolate from the demographic trend and predict the rapid and complete assimilation of Orthodox Jews to American cultural values. Such predictions have proved counterfactual already in the 1970s (see Mayer [1979] for a severely critical account of such predictions).

more productive technology to socialize children to such beliefs. Marriage choices are then motivated by the desire to socialize children and will result in assortative marriage pairs along religious lines. Intermarriage rates as well as socialization rates are therefore not only a consequence of social interactions of children but also in part a consequence of individual choices. Our empirical analysis confirms that the choice-theoretic framework is important to fit the observed intermarriage and socialization rates in the United States. An alternative model in which the marriage component is modeled as an economic decision problem and socialization is exogenously determined does not fit the data nearly as well. Nor does a model in which both marriage and socialization are exogenously determined.

The main structural parameters of the model are the "relative intolerance" parameters, which are preference parameters defined as the perceived utility gains parents of religious group i derive from offspring of religion i rather than j. We estimate such parameters by matching the empirical frequencies of religious intermarriages (e.g., Protestant-Catholic, Protestant-Jewish, etc.) and the empirical socialization rates with those implied by our model via a simple minimum distance procedure.

The observed marriage and socialization patterns are consistent with a strong preference by members of each religious group for having children who share their own religious trait. The estimated relative intolerance parameters are significant and in several cases asymmetric. For instance, the intolerance parameter of Protestants with respect to Catholics and the one of Catholics with respect to Protestants are not significantly different; for Jews, we estimate a much higher intolerance parameter with respect to Catholics than vice versa.

The socialization pattern in the data contains a bias in favor of the residual group, "Others," for those children of all religious groups who are not directly socialized in the family. Such bias is consistent with relatively high rates of conversion from the three major religious faiths into the group of individuals with a preference for "no religion" and for other religions.

Moreover, marriage segregation and socialization effort share the same qualitative nonlinear pattern in the simulations. When a group is a minority, marriage segregation and socialization efforts are increasing in the group's population share. The reason is that the estimated costs of socialization and marriage segregation are substantial for a minority. As a group grows toward being a majority, marriage segregation and socialization efforts become decreasing in the group's population share. The reason is that when a group population share is high, social interactions favor homogamy and socialization, independent of the explicit effort of individuals and parents.

As a consequence of such nonlinearities, extrapolations from demographic and sociological trends are potentially severely misleading in their conclusions about the religious dynamics of the population. Although with our data such extrapolations in fact partially reproduce the triple melting pot prediction (revised to account for the supposed vanishing of the Jewish population due to their recent intermarriage behavior), our simulations based on the parameter estimates of the structural model paint a very different picture.

We estimate very high intolerance toward Catholics for all other religious denominations, and as a consequence, the share of Catholics in the population decreases over time. We also show that the initial proportions of Protestants and Catholics, in large part, determine the dynamics of the share of Jews and of Others. In fact, Catholics have a very low estimated intolerance level toward Jews, and Protestants have lower intolerance toward Others than Jews. Thus, when Protestants are a high majority in the initial conditions, Jews tend to decline and Others gain a small but stable share of the population. When, instead, Catholics are well represented in the initial distribution, Jews are favored and their share rises.

More generally, our analysis is perhaps of some methodological value for empirical analyses of economies with social interactions.[3] Our implementation of the tests proposed by Vuong (1989) and Kitamura (2000) to compare nonnested models, for example, can be of general interest to evaluate the economic explanations of social phenomena versus non-choice-theoretic sociological analyses of the same phenomena. Also, we produce a general heuristic approach to the identification and computation difficulties of economic models in which social interactions give rise to multiple equilibria.

We deal in this paper with marriage and socialization patterns along only the religious dimension. We know of no other work in the economic and sociological literatures that aims at assessing, in a structural environment, the relevance of a cultural trait in the marriage market and simulates the dynamics of such traits in the population. Religious traits offer a particularly appropriate set of observations for the general analysis of cultural traits (e.g., ethnicity and race) because (i) religious traits are relatively well defined and measured (better, e.g., than ethnicity); (ii) they represent cultural traits that most families are keen to transmit to their children (see, e.g., Glazer 1997); and, finally, (iii) the families' incentives to transmit their religious trait are not much obscured by related economic incentives, since religious beliefs have relatively minor

[3] See Durlauf (1999), Glaeser and Scheinkman (2000, 2001), and Brock and Durlauf (2001) for general theoretical and empirical frameworks to study such economies.

effects per se in the United States on the agents' economic opportunities (but see Warren [1970]).

The paper is organized as follows. Section II presents a brief description of intermarriage and socialization patterns in the United States. Section III discusses the economic model. Section IV presents the empirical methodology and Section V the identification procedure. The estimation results are reported in Section VI. Section VII performs tests of alternative model specifications to verify the endogeneity of marriage and socialization. Section VIII analyzes the dynamic implications of our estimates for the long-run distribution of the population by religious group. We then discuss several potential issues in our analysis (such as migration and unobserved heterogeneity) in Section IX. Section X presents conclusions.

II. Intermarriage and Socialization in the United States

High homogamy rates along religious and ethnic dimensions in the United States are well documented. Using General Social Survey (GSS) data with U.S. states as geographic units, cumulated over the period 1972–96, figure 1 documents the sample probability that a member of a specific religious group marries homogamously. For all religious groups in the sample (Protestants, Catholics, Jews, and the residual Others), this probability is significantly higher than that implied by random matching (which would require all observations on the 45-degree line).[4] Such marriage patterns, strongly positively assortative along the religious dimension, are characteristic of the whole of the United States.

Religious socialization rates are also quite high in the United States, and especially so for homogamous couples. This is documented, for the GSS data set, in table 1. The probability that a Protestant parent has a Protestant child is, for instance, 92 percent in a homogamous marriage, whereas it is only 51 percent in a heterogamous marriage with a Catholic spouse.[5]

Can the high socialization rates associated with homogamous mar-

[4] Appendix table B1 reports marriage rates for 23 U.S. states, again from the same GSS data. For instance, in Tennessee the sample probability of homogamous marriage for Catholics is 66 percent, even though they represent only 4 percent of the sample population. This probability is 91 percent for Jews in Illinois, whose population represents less than 2 percent of the sample population of the state. Similarly, the probability that a Catholic marries a Protestant in South Carolina, where Protestants represent 87 percent of the population, is only 33 percent.

[5] The fact that homogamous marriages are more effective in socializing children along the religious dimension than heterogamous ones has been well documented in the sociological literature (see, e.g., Hoge and Petrillo 1978; Hoge, Petrillo, and Smith 1982; Heaton 1986; Ozorak 1989).

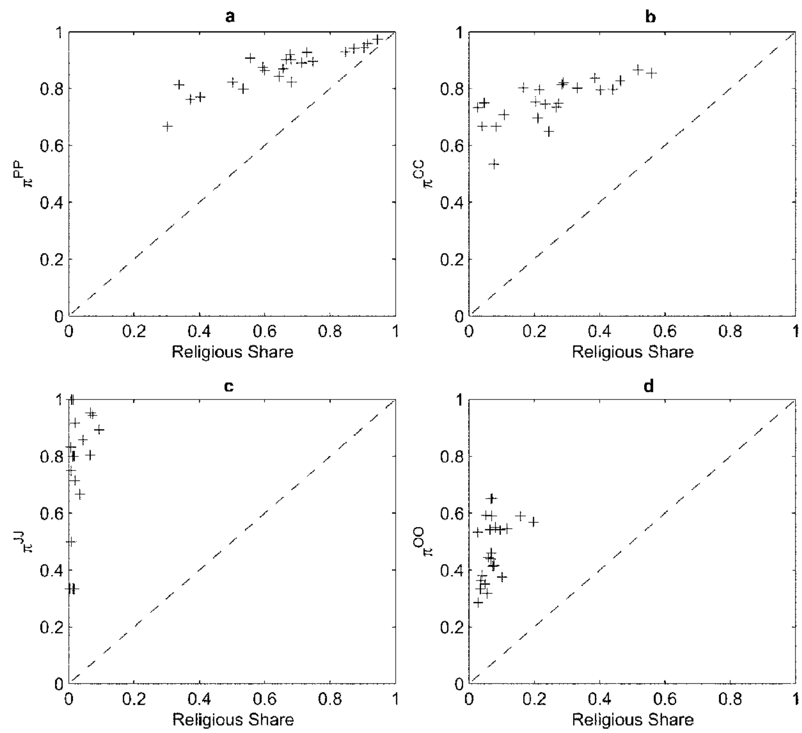

FIG. 1.—Probability of homogamous marriage. *a*, Protestants. *b*, Catholics. *c*, Jews. *d*, Others.

riages actually explain the high homogamy rates that are observed in the United States? This would be the case if parents valued having children who share their own individual beliefs. While several considerations other than the socialization of children affect actual marriage choices, substantial evidence points to the desire to socialize children as an important determinant of homogamy. Psychological studies of heterogamous couples consistently report the partners' concern about possible cultural attitudes of children when deciding to form a family (see, e.g., Mayer 1985; Smith 1996). Similarly, anthropological evidence points to the cultural identity of the children as a determinant of marriage choice (see, e.g., Riesman and Szanton 1992). Also, the documented fact that cohabitations are both much less fertile and less homogamous than marriages can be interpreted as evidence that homogamy matters mostly for fertile unions (see Rindefuss and VandenHeuvel [1990] for relative fertility of cohabitation and Schoen and Weinick [1993] for homogamy rates). Finally, most major religious

TABLE 1
Socialization Probabilities for Selected Marriage Types

	Protestants	Catholics	Jews	Others
PP marriage	.9179	.0284	0	.0537
CC marriage	.0850	.8571	.0034	.0544
JJ marriage	.0370	0	.9259	.0370
OO marriage	.3231	.0462	0	.6308
PC marriage	.5116	.3140	0	.1744
PO marriage	.7100	.1000	0	.1900
CO marriage	.1667	.5000	0	.3333

Note.—Each cell reports the sample probability that a child in the row marriage is a member of the column religious group. P = Protestants, C = Catholics, J = Jews, and O = Others.

denominations severely regulate intermarriages, often explicitly citing the difficulties of socialization as the main justification.[6]

Some indirect evidence from the GSS data set also supports our view that the desire for socialization explains in part the high homogamy rates along religious lines. In fact, homogamy rates are higher for young couples (less than 25 years of age at marriage), who are more fertile in expectation, and for effectively more fertile couples (with more than one child), who are also possibly more fertile in expectation when they get married. Fertility rates are also higher for homogamous couples, for any religious group (except the residual group, Others), as to be expected if socialization drives homogamy rates in a relevant way.

The analysis of marriage and socialization that follows will provide further evidence of the relationship between socialization and marriage along religious lines that we have suggested in this section.

III. The Model

The marriage and cultural transmission model we study is an extension of the model introduced by Bisin and Verdier (2000) to study the transmission of ethnic and religious traits.[7]

Parents have a taste and a technology for transmitting their own religious faith to their children. Compared to parents in heterogamous marriages, parents in homogamous marriages have a better technology to socialize their offspring to their own trait. As a consequence, ho-

[6] For example, the 1983 Code of Canon Law for the Catholic Church says that "Without the express permission of the competent authority, marriage is forbidden between two baptized persons, one of whom …[is] in the Catholic Church … and the other [is] in a … Church … which is not in full communion with the Catholic Church" (canon 1124). Moreover, the permission cannot be granted unless "the Catholic party … makes a sincere promise to … have all children … brought up in the Catholic Church" (canon 1125).

[7] Early cultural transmission models are discussed in Cavalli-Sforza and Feldman (1981) and Boyd and Richerson (1985). See the discussion in Bisin and Verdier (2000) for a comparison. The economic approach to the study of religion has been pioneered by Iannaccone; see his (1990) survey.

mogamous unions have a higher value than heterogamous ones, and agents are willing to spend effort to segregate into a restricted marriage pool in which they are more likely to meet prospective spouses of the same religious faith. The preference for socialization therefore drives the marriage choice of the agents.

We first introduce our modeling of the institutional structure of the marriage market (subsection *A*). We then introduce the socialization technology with which we assume families are endowed (subsection *B*). Finally, we study the decision of individuals regarding marriage segregation by religious group and the decision of parents regarding the socialization of children (subsection *C*).

A. The Marriage Market

Let $i, j, k = 1, \dots, n$ index different religious groups. All agents adhering to religious group i are ex ante (before marriage, i.e.) identical. Fix a geographic unit of reference, for example, a state in the United States. Let q^i denote the fraction of the population in the geographic unit (we do not keep track of the geographic unit in the notation for simplicity) adhering to religious group i. Clearly, $\sum_{i=1}^n q^i = 1$. Also let $\mathbf{q} = [q^1, \dots, q^n]$. Let π^{ij} denote the probability that a member of religious group i in the geographic unit is married to a member of religious group j. Let α^i denote the probability that an agent of religion i in a geographic unit marries homogamously (with another member of religious group i) in a restricted (religiously segregated) marriage pool, where α^i is chosen by each agent of group i. Also let $\boldsymbol{\alpha} = [\alpha^1, \dots, \alpha^n]$.

The matching process can be defined as follows. Agents of an arbitrary religious group i first have a marriage draw in the restricted marriage pool. With probability α^i, they are married there (all marriages in the restricted pool are homogamous). With probability $1 - \alpha^i$, they are not married in the restricted pool, and hence they marry in the common pool, which is formed of all agents who are not married in their religious groups' respective restricted pools. Marriage in the common pool occurs by random matching, and hence, for instance, the probability that an agent of group i is married homogamously in the common pool (conditionally on not having found a marriage partner in the restricted pool) is

$$\frac{(1-\alpha^i)q^i}{\sum_{j=1}^n (1-\alpha^j)q^j}.$$

We then write

$$\pi^{ii} = \alpha^i + (1 - \alpha^i)\frac{(1 - \alpha^i)q^i}{\sum_{k=1}^{n}(1 - \alpha^k)q^k}, \quad i = 1, \ldots, n, \tag{1}$$

and

$$\pi^{ij} = (1 - \alpha^i)\frac{(1 - \alpha^j)q^j}{\sum_{k=1}^{n}(1 - \alpha^k)q^k}, \quad i \neq j. \tag{2}$$

While very stylized, this marriage model does represent a rich statistical model of marriage and hence does not impose many restrictions on the data per se.[8] Estimating with the GSS data the probability of marrying homogamously in the restricted pool of each religious group i, α^i, for each U.S. state, so as to match the marriage rates (the realized homogamy and heterogamy rates in Appendix table B1), produces an almost perfect fit; see figure 2 for homogamy rates. The p-value of the Sargan test for this estimate is .9966. The restrictions imposed by our analysis on the data are those implied by our modeling of the decision to enter the restricted pool as made by rational agents rather than those implied by the postulated institutional structure of the marriage market.

B. The Socialization Technology

Socialization and cultural transmission occur in the family and in society at large (as a consequence of social contact with peers, role models, etc.). Families are indexed by pairs ij, where i and j indicate the religious group of each parent. Let P_{ij}^i denote the probability that a child of a family of type ij has religious trait i. The socialization (cultural transmission) mechanism is as follows:

1. A child from a religious homogamous family of type ii is directly socialized to the trait of the family with probability $\tilde{\tau}^i \equiv \tau^i + m$, where τ^i is chosen by the parents and m is an exogenous probability independent of parents' effort (exogenous direct socialization). With probability $1 - \tilde{\tau}^i$, he or she is not directly socialized by the family and picks a trait by matching in the population with a cultural parent who socializes the child to the parent's own religious trait.

[8] Johnson (1980) also estimates a statistical model of religious intermarriage.

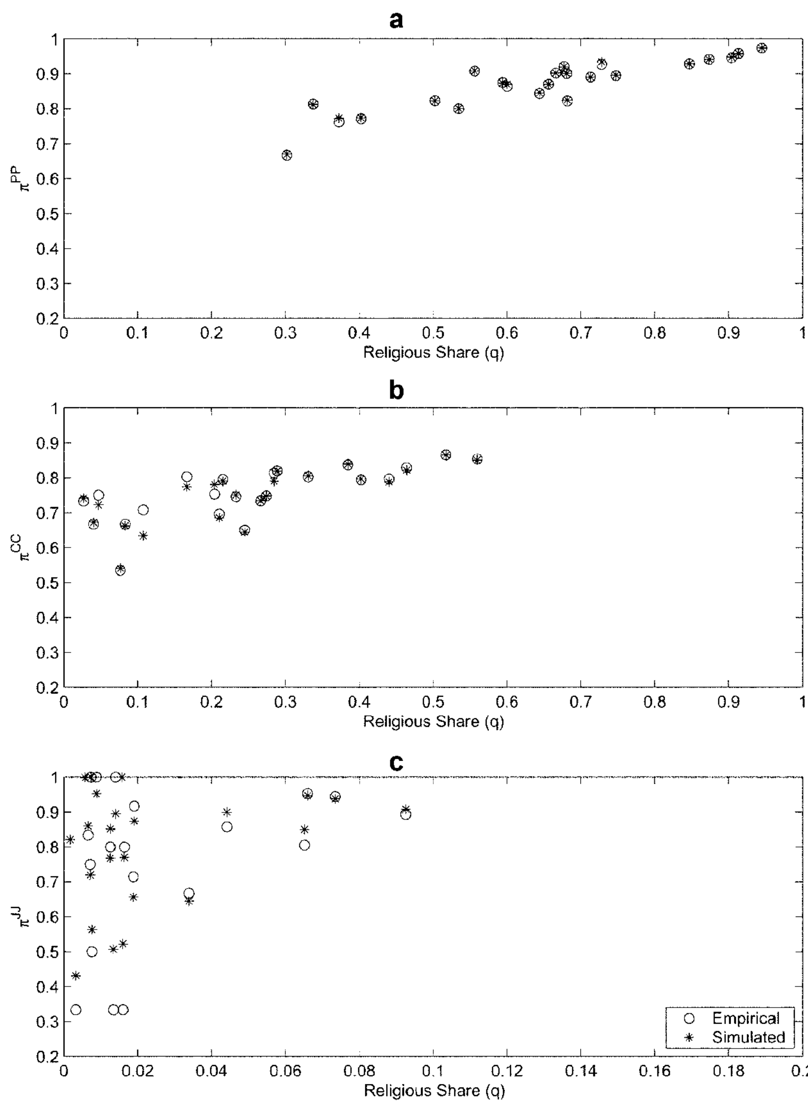

FIG. 2.—Statistical model of the marriage market: homogamous marriage probability. *a*, Protestants. *b*, Catholics. *c*, Jews.

2. A child in a heterogamous family of type ij, $i \neq j$, does not have a well-defined reference religious trait to be socialized to, and as a consequence he or she picks each parent's trait independently with some exogenous probability $m/2$. If not socialized by either parent, a child in a heterogamous family picks a trait by matching in the population with a cultural parent who socializes the child to the parent's own religious trait.

This mechanism embeds two assumptions. One is that parents in homogamous unions have a better socialization technology than parents in heterogamous ones. The other is that children can acquire a given trait either through a "vertical" socialization process from their parents or through an "oblique" socialization process from society at large. Substantial sociological evidence supports in fact such assumptions regarding the socialization mechanism. See Hoge et al. (1982), deVaus (1983), Clark, Worthington, and Danser (1988), Cornwall (1988), Ozorak (1989), Erickson (1992), and Hayes and Pittelkow (1993) for some direct evidence on religious socialization and Wilson (1987) for the sociological literature on adult role models.

The process of matching with cultural parents in the population, which determines the oblique socialization process, is not directed by agents' choices and might in principle be biased in favor of some particular religious groups, essentially because of conversions. While Catholics and some Protestant denominations do actively proselytize (and Jews do not), it turns out that this phenomenon has a substantial role in our analysis only for the residual group, Others, which includes (in large part) individuals with a preference for "no religion" and members of different religious sects.

Let Q^i denote the probability that any child not socialized in the family meets with a cultural parent of religious group i in the population and hence is socialized to religion i. Obviously, $\sum_i Q^i = 1$. Note that random matching requires $Q^i = q^i$; when $Q^i > q^i$, conversions to religion i affect the matching process of children with cultural parents. The distance of such a matching process from random matching will be estimated.

For any $i, j, k = 1, \ldots, n$, the socialization equations for homogamous families can be written as

$$P_{ii}^i = \tilde{\tau}^i + (1 - \tilde{\tau}^i)Q^i \tag{3}$$

and

$$P_{ii}^j = (1 - \tilde{\tau}^i)Q^j, \quad i,\ j \text{ distinct}, \tag{4}$$

and the socialization equations for heterogamous families as

$$P_{ij}^{i} = \tfrac{1}{2}m + (1 - m)Q^{i}, \quad i,\ j \text{ distinct}, \tag{5}$$

and

$$P_{ij}^{k} = (1 - m)Q^{k}, \quad i,\ j,\ k \text{ distinct}. \tag{6}$$

C. *Marriage Segregation and Socialization Choices*

Each agent of religion i chooses α^i, the probability of being matched in the restricted pool (where all mates have trait i). The cost associated with α^i when the share of religious group i in the population is q^i is $M(\alpha^i, q^i)$. The advantage of marrying homogamously is that it gives one the option of directly socializing one's children. The direct socialization rate in a homogamous marriage of religion i, τ^i, is chosen by parents. The cost associated with direct socialization τ^i when the share of religious group i in the population is q^i is denoted $S(\tau^i, q^i)$. The benefit of socialization derives from the fact that parents want their children to share their own religious faith. The value for a type i parent of a type j child, V^{ij}, is exogenously given and needs to be estimated. In this regard, we postulate $V^{ii} \geq V^{ij}$.

For clarity's sake, we analyze the model backward. We first study the choice of direct socialization τ^i by homogamous parents of religion i. We then study α^i, the marriage choice of agents of religion i.

The direct socialization of one child of a homogamous family of type ii, τ^i, is the solution of the following maximization problem:

$$\max_{0 \leq \tau^i \leq 1} P_{ii}^{i}V^{ii} + \sum_{j \neq i} P_{ii}^{j}V^{ij} - S(\tau^i,\ q^i) \tag{7}$$

subject to (3)–(6).

We do not write any explicit endogenous fertility problem for the agents, essentially because one extra optimization problem would make the model intractable. We assume that agents take as given a constant fertility rate of a type ij marriage, denoted n^{ij}. Then, for a type i agent, let $(n^{ij})^{\xi}W^{ij}$ denote the value of a marriage with a type j agent, given that the marriage produces n^{ij} children. The parameter ξ denotes the dependence of the parents' preferences on the number of children in the marriage. For a type i agent, the value per child of a marriage with a j spouse, W^{ij}, is the expected value of a child in such a marriage:

$$W^{ii} = P_{ii}^{i}V^{ii} + \sum_{j \neq i} P_{ii}^{j}V^{ij} - S(\tau^{*i},\ q^i) \tag{8}$$

and

$$W^{ij} = P_{ij}^{i}V^{ii} + \sum_{k \neq i} P_{ij}^{k}V^{ik}, \quad i \neq j. \tag{9}$$

We are implicitly assuming that all children in a marriage are socialized to the same trait. This is done just for simplicity and does not change any of our results.

When choosing the probability of being married in the restricted marriage pool, all agents take the composition of the common pool as given, since each agent is infinitesimal and hence does not affect the composition. In equilibrium the composition of the common pools will be required to be consistent with all the agents' choices.

Let A^{ii} denote a type i agent's probability of marrying homogamously in the common pool, and let A^{ij} denote his or her probability of marrying a type j agent in the common pool. Let $M(\alpha^{i}, q^{i})$ denote marriage segregation costs. The marriage problem of a type i agent is

$$\max_{0 \leq \alpha^{i} \leq 1} \pi^{ii}(n^{ii})^{\xi} W^{ii} + \sum_{j \neq i} \pi^{ij}(n^{ij})^{\xi} W^{ij} - M(\alpha^{i}, q^{i}) \tag{10}$$

subject to

$$\pi^{ii} = \alpha^{i} + (1 - \alpha^{i})A^{ii}$$

and

$$\pi^{ij} = (1 - \alpha^{i})A^{ij}, \quad j \neq i,$$

given A^{ii} and A^{ij}, for all $j \neq i$.

In equilibrium,

$$A^{ii} = \frac{(1 - \alpha^{i})q^{i}}{\sum_{k=1}^{n} (1 - \alpha^{k})q^{k}}$$

and

$$A^{ij} = \frac{(1 - \alpha^{j})q^{j}}{\sum_{k=1}^{n} (1 - \alpha^{k})q^{k}}.$$

The reduced-form equations of the structural model just introduced are reported and studied for clarity in Appendix A; a more detailed analysis is contained in the working paper version of the paper (Bisin, Topa, and Verdier 2003).

IV. The Empirical Implementation

We consider the following religious groups: Protestants, Catholics, Jews, and the residual group, Others. The latter category includes individuals with a preference for "no religion" as well as individuals with preferences for other religious faiths. Moreover, we assume that the geographic unit of reference for marriage and socialization processes coincides with the state. In other words, we assume that the composition of the population by religious group, which is relevant for each agent in the marriage market and for each family when choosing the direct socialization levels, is the composition of the population in the U.S. state in which the agent or the family resides. The definitions of both the religious groups and the geographic unit of reference in the analysis are arbitrary and are determined essentially by the available data. We discuss the possible problems associated with such definitions, and the several attempts we made at assessing the robustness of our empirical results to different definitions of groups and geographical units, in Section IX.

In this section, we first make operational the model described in the previous section by introducing relevant assumptions and the necessary functional form parameterization. We discuss identification of the model parameters, and we briefly present the data we use in the estimation. We then construct a map from the parameters of the model, the composition of the population, and fertility rates by religious group into intermarriage and socialization rates, to be matched with those that we observe in the data. (Some mathematical properties of this map are studied in App. A.) Finally, we introduce an appropriate estimation procedure.

Index each state by s. Let $\Delta V^{ij} \equiv V^{ii} - V^{ij}$, for any i, j (obviously, $\Delta V^{ii} = 0$). The term ΔV^{ij} measures the perceived increment in utility for a type i agent associated with having a child of type i rather than j: we refer to it as the "intolerance" of type i agents toward group j. We parameterize the cost functions by[9]

$$S(\tau_s^i,\ q_s^i) \equiv [\sigma_\tau + \epsilon_\tau(1 - q_s^i)^2] \cdot \left\{\lambda_\tau \frac{(\tau_s^i)^2}{2} + (1 - \lambda_\tau)\left[\exp\left(\frac{\tau_s^i}{1 - \tau_s^i}\right) - 1\right]\right\} \tag{11}$$

and

$$M(\alpha_s^i,\ q_s^i) \equiv [\sigma_\alpha + \epsilon_\alpha(1 - q_s^i)^2] \cdot \left\{\lambda_\alpha \frac{(\alpha_s^i)^2}{2} + (1 - \lambda_\alpha)\left[\exp\left(\frac{\alpha_s^i}{1 - \alpha_s^i}\right) - 1\right]\right\}. \tag{12}$$

[9] The parameterization satisfies the convexity and Inada conditions that are necessary for our analysis; see assumption 1 in App. A.

We also parameterize the matching probabilities in the oblique socialization as

$$Q_s^O = \frac{q_s^O + o}{1 + o}, \quad Q_s^i = \frac{q_s^i}{1 + o}, \quad i = \text{P, C, J}. \tag{13}$$

The parameter o represents the deviation away from pure random matching, in the matching process that determines the socialization of children in society at large. This deviation could be explained as the effect of the conversion rate into the residual group. We find no evidence in the data of such conversions into our main religious groups, Protestants, Catholics, and Jews.

The parameters of the model consist of the intolerance parameters ΔV^{ij}, for any i and $j \neq i$; the parameters that describe the cost functions $(\sigma_c, \epsilon_c, \lambda_c)$, where $c = \{\alpha, \tau\}$; the bias due to the conversions to Others in the oblique socialization process, o; the exogenous direct socialization rate, m; and the preference for fertility, ξ. Let $\boldsymbol{\theta}$ denote the vector of parameters.

We use data from the GSS, covering the period 1972–96, on the composition of marriages by religious affiliation of the spouses for each state; the composition of the population by religious group for each state; the socialization rates by religious affiliation of the spouses aggregated over the United States; and fertility rates by religious group of the spouses aggregated over the United States. According to our notation, we have data on π_s^{ij}, for all i, j, and s; q_s^i, for all i and s; $P_{ij}^k = \sum_s \omega_s^{ij} \cdot P_{ij,s}^k$, where ω_s^{ij} are sample weights representing the percentage of respondents in an ij marriage that live in state s, for all i, j, and k; and n^{ij}, for all ij. (We do not have enough data to construct accurate empirical frequencies of socialization rates, $P_{ij,s}^k$, for each state s.)

Given the values of homogamous and heterogamous marriage unions (W^{ii} and W^{ij}, respectively), an equilibrium in the marriage market is a solution to the fixed-point problem of (A8)–(A10) in Appendix A and (12). While an equilibrium always exists, there is no guarantee that the equilibrium is unique for general cost functions $M(\alpha^i, q^i)$ as in (12). Multiple equilibria are the consequence of the coordination problem implicit in our formulation of the marriage market. Suppose that under the parameters of the model two religious groups aim at segregating in the marriage market. The same segregation pattern can be achieved if agents of group i choose high α^i and agents of group j choose low α^j, as well as if, vice versa, group i chooses low α^i and group j chooses high α^j. In the first case, agents of group j can segregate in the residual pool, which is composed mainly of agents of group j thanks to the high segregation effort of the other group; in the second case, it is agents

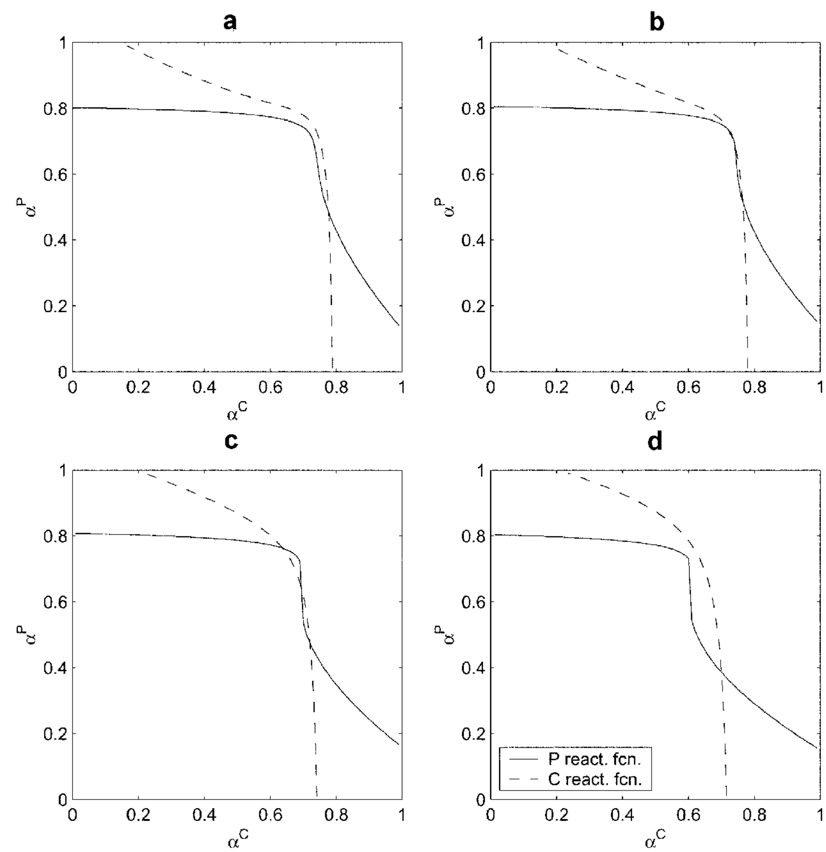

FIG. 3.—Marriage probability reaction functions. *a*, Protestant share = .54, Catholic share = .36. *b*, Protestant share = .58, Catholic share = .32. *c*, Protestant share = .69, Catholic share = .21. *d*, Protestant share = .73, Catholic share = .17.

of group i who can segregate in the residual pool. Such different segregation patterns have important distributional effects (the group segregating in the residual pool is favored, since the costs to enter the restricted pool are saved), but homogamy rates for the two groups can remain unaffected (see fig. 3).[10]

Given q_s^i for all i and s and n^{ij} for all i and j, the structural model

[10] In fig. 3 we plot a pair of best-reply functions for the restricted pool marriage probabilities, α^i. We fix Jews' and Others' religious shares at their mean values and plot the Protestant and Catholic best-reply functions to each other's α^i, while keeping the marriage segregation probabilities for Jews and Others at their equilibrium levels. We repeat the exercise for different combinations of Protestant and Catholic religious shares. The plots clearly indicate the presence of nonconvexities in the best-reply functions that generate multiple equilibria for at least some values of religious shares.

(represented by eqq. [A1]–[A10] in App. A and [11]–[13]) defines a mapping, $\tilde{\mathbf{\Pi}}(\boldsymbol{\theta})$, from $\boldsymbol{\theta}$ into π_s^{ij} for all i, j, and s and into P_{ij}^k for all i, j, and k. We use a minimum distance estimation procedure that matches the vector $\hat{\mathbf{\Pi}}$ of empirical moments $(\hat{\pi}_s^{ij}, \hat{P}_{ij}^k)$ from the data with the vector $\tilde{\mathbf{\Pi}}(\boldsymbol{\theta})$ of moments implied by the model for a given choice of $\boldsymbol{\theta}$. Formally, given a square weighting matrix $\mathbf{\Omega}_N^*$ (where N denotes the total sample size), the minimum distance estimator $\hat{\boldsymbol{\theta}}$ minimizes

$$J_N(\boldsymbol{\theta}) \equiv [\hat{\mathbf{\Pi}} - \tilde{\mathbf{\Pi}}(\boldsymbol{\theta})]^\top \mathbf{\Omega}_N^* [\hat{\mathbf{\Pi}} - \tilde{\mathbf{\Pi}}(\boldsymbol{\theta})]. \tag{14}$$

Possible discontinuities of the map $\tilde{\mathbf{\Pi}}(\boldsymbol{\theta})$ may be problematic for various reasons. First, standard consistency proofs usually require continuity of the criterion to be minimized (and hence of $\tilde{\mathbf{\Pi}}(\boldsymbol{\theta})$). However, it is easy to show that *local* continuity at the global minimum of $J_N(\boldsymbol{\theta})$ is sufficient for consistency.

Second, in order to compute standard errors, one needs to ensure that $\tilde{\mathbf{\Pi}}(\boldsymbol{\theta})$ is locally smooth at $\hat{\boldsymbol{\theta}}$ and hence that the partial derivatives $\partial\tilde{\mathbf{\Pi}}(\boldsymbol{\theta})/\partial\boldsymbol{\theta}^\top$ are well defined. We check that this requirement is satisfied in a neighborhood of $\hat{\boldsymbol{\theta}}$. As long as $\hat{\boldsymbol{\theta}}$ is indeed the global minimizer of the criterion, this is sufficient for local continuity as well.

Finally, discontinuities in $\tilde{\mathbf{\Pi}}(\boldsymbol{\theta})$ typically make it much harder for standard minimization algorithms to find the global minimum of $J_N(\boldsymbol{\theta})$. However, we use a simulated annealing algorithm that is especially well suited for problems in which the objective function may have various discontinuities or several distinct local optima. A more detailed description of the data and the estimation methodology, as well as the simulated annealing algorithm, is contained in Appendix B.

V. Identification

Our crucial identifying assumptions are that (i) cost functions are not specific to any religious group, nor to the geographic units of reference (the states); (ii) the intolerance parameters, ΔV^{ij}, are naturally specific to the religious groups (i and j) but independent across states; and (iii) the socialization bias due to conversions, o, as well as the exogenous direct socialization rate, m, and the demographic preference parameter, ξ, are constant across states and across religions. Under these assumptions, we are able to identify independently the intolerance of group j with respect to group i, ΔV^{ji}, as well as the intolerance of group i with respect to group j, ΔV^{ij}, out of data on intermarriages.

In fact, even though the marriage unions between individuals of groups i and j are also unions between individuals of groups j and i, in our model realized intermarriage rates depend on the segregation efforts of the two groups, which depend nonlinearly on the shares of the population by religious groups. Such nonlinearities, together with the variation in the population distribution of religious groups across U.S. states, can be exploited to identify asymmetric intolerance parameters.

Identification might instead fail, for example, if marriage segregation and socialization costs were allowed to differ across religious groups. In this case, in fact, the dependence of realized intermarriage rates on the shares of the population by religious groups could be due equivalently to variation in intolerance levels or in costs. In other words, any difference in costs across religious groups in reality would be captured in our estimation procedure by differences in intolerance levels. The limitations of our data set, and in particular the lack of independent direct observations of marriage segregation and socialization costs, do not allow us to address this problem.

The following simple example illustrates more precisely our identification procedure. Consider an economy with only two religious groups, for example, Catholics (C) and Protestants (P). Assume that fertility rates are constant across all family types (so that the model is independent of fertility rates). Also, assume that the exogenous direct socialization rate, m, is zero and that cost functions are quadratic: $S(\tau^i, q^i) = \frac{1}{2}(\tau^i)^2$ and $M(\alpha^i, q^i) = \frac{1}{2}(\alpha^i)^2$. With equations (A1) and (A2), in this special case we can solve for α_s^i in each state s:

$$\alpha_s^i = (1 - A_s^{ii})[(1 - q_s^i)\Delta V^{ij}]^2.$$

In this example our estimation procedure would need to match only one moment, for instance π_s^{CC}, for each state s. In fact, given π_s^{CC} and q_s^C, we can solve for π_s^{CP} using $\pi_s^{CP} = 1 - \pi_s^{CC}$, then for π_s^{PC} using $\pi_s^{CP} q_s^C = \pi_s^{PC} q_s^P$, and finally for π_s^{PP} using $\pi_s^{PP} = 1 - \pi_s^{PC}$.

When equations (A9) and (A10) are written in implicit form, $A_s^{CC} = A^{CC}(q_s^C;\ \alpha_s^C, \alpha_s^P)$ and $A_s^{PP} = A^{PP}(q_s^P;\ \alpha_s^P, \alpha_s^C)$, the model is reduced to three equations in each state s:

$$\pi_s^{CC} = \alpha_s^C + (1 - \alpha_s^C) A^{CC}(q_s^C;\ \alpha_s^C, \alpha_s^P),$$

$$\alpha_s^C = [1 - A^{CC}(q_s^C;\ \alpha_s^C, \alpha_s^P)] \cdot [(1 - q_s^C)\Delta V^{CP}]^2,$$

and

$$\alpha_s^P = [1 - A^{PP}(q_s^P;\ \alpha_s^P, \alpha_s^C)] \cdot [(1 - q_s^P)\Delta V^{PC}]^2,$$

and four unknowns: α_s^C, α_s^P, ΔV^{CP}, and ΔV^{PC}. Therefore, the parameters ΔV^{CP} and ΔV^{PC} cannot be identified independently with data on q_s^C and π_s^{CC} for a particular state s. But since we restrict the "intolerance" parameters ΔV^{CP} and ΔV^{PC} to be independent of the state s and exploit the variability of the observations of q_s^C and π_s^{CC} across Y states, we face $3Y$ independent equations and $2Y + 2$ unknowns, ΔV^{CP}, ΔV^{PC}, and

(α_s^C, α_s^P) for each state (for $Y > 2$ states, the system is therefore overidentified).[11]

Finally, if π_s^{CC} nontrivially depends on both α_s^P and α_s^C, then the system of equations is locally independent.

In the context of the more general model, an identification argument can be sketched along the following lines. From (A7) we can write the optimal τ^i as

$$\tau_s^i = S_1^{-1}\left[\sum_h q_s^h \Delta V^{ih} | \sigma_\tau, \epsilon_\tau, \lambda_\tau\right] \quad \forall i, s. \tag{15}$$

The parameters m and o are identified from the equations for the socialization rates, (A3)–(A6). The observed π_s^{ij} pin down the α_s^i and the A_s^i. Then one can use the first-order conditions for α_s^i, (A8), and (15) to write down a list of nonlinear equations in the observed (or estimated) (α_s^i, A_s^{ii}, n^{ij}, q_s^i, m, and o) and the remaining unknown parameters (V^{ii}, ΔV^{ij}, σ_τ, ϵ_τ, λ_τ, σ_α, ϵ_α, λ_α, and ξ). They constitute nY equations in $n^2 + 7$ parameters (where n is the number of religious groups).[12] The order condition for identification is satisfied in our case, with $n = 4$ and $Y = 23$, and the rank condition can be checked locally.

A. *Multiplicity of Equilibria*

Because of the possibility of multiple equilibria, our identification procedure must jointly identify the parameters of the model and the equilibrium selection. In terms of the minimum distance criterion, for our estimate to be consistent we need to find the value of $\boldsymbol{\theta}$ that minimizes the lower envelope of the multiple $J_N(\boldsymbol{\theta})$ surfaces generated by the different equilibria.

No standard procedure for identification is available in the face of multiple equilibria.[13] We therefore use a heuristic approach to locally identify the equilibrium selection and the parameters of the model. In the course of the minimization of $J_N(\boldsymbol{\theta})$, for each candidate value $\bar{\boldsymbol{\theta}}$, we let the algorithm randomly pick several distinct starting values for the iteration that yields the equilibrium, in order to try to generate several possible equilibria. We then compute J_N for each of these equilibria and use the lowest value as the value $J_N(\bar{\boldsymbol{\theta}})$ for that particular value of $\boldsymbol{\theta}$.

[11] Note that in this identification procedure we crucially exploit the assumed independence of cost functions from state and religious group when solving for α^i.

[12] In the actual estimation we restrict V^{ii} to be the same across all groups i; therefore, the number of parameters is reduced to $n(n-1) + 8$.

[13] But see Dagsvik and Jovanovic (1994). Jovanovic (1989) discusses identification in a general framework. Moro (2003) has ingeniously introduced a procedure that allows the local identification of a specific model of statistical discrimination with multiple equilibria. Moro's procedure cannot be simply adopted in our setup.

TABLE 2
STRUCTURAL ESTIMATION RESULTS

	Estimate	Standard Error	p-Value
Value of same-religion child	526.2563	102.3493	0
Intolerance of:			
P toward C	125.2046	2.3046	0
P toward J	121.3225	9.7949	0
P toward O	31.9216	4.7295	0
C toward P	152.9794	2.9070	0
C toward J	14.9936	5.0554	.0015
C toward O	12.1115	2.6230	0
J toward P	501.2928	82.7201	0
J toward C	526.2551	222.0905	.0089
J toward O	525.9072	77.8761	0
O toward P	106.0829	9.1286	0
O toward C	165.5395	21.0772	0
O toward J	.7108	173.2716	.4984
Cost parameter:			
σ_τ	1.9227	.4263	0
σ_α	5.9062	.4156	0
ϵ_τ	69.4675	5.8155	0
ϵ_α	67.1701	2.7009	0
λ_τ	.6773	.0417	0
λ_α	.9996	.0001	0
Exogenous direct socialization: m	.3457	.0201	0
Conversions to others: o	.2062	.0155	0
Fertility parameter: ξ	.0108	.0093	.1230
J-test:			
Marriage only (116 degrees of freedom)	134.7098		.1129
Overall (137 degrees of freedom)	174.2282		.0173

NOTE.—P = Protestants, C = Catholics, J = Jews, and O = Others.

Since we cannot compute all possible equilibria for each $\boldsymbol{\theta}$ because of computational limitations, this procedure is at least a step in the direction of searching for the equilibrium selection, as well as the parameter estimate, that minimize the criterion $J_N(\boldsymbol{\theta})$.

VI. Estimation Results

Table 2 presents the estimation results of the structural marriage and socialization model introduced in Section III with the GSS data.

The model fits the intermarriage data quite well, whereas it fits the socialization data less well. The p-value of the Sargan test of the overidentifying restrictions is quite high (.11) when one considers the intermarriage moments alone but drops to about .017 when one considers the socialization moments as well. In order to get a visual impression of the fit, see figure 4, which compares the empirical homogamous

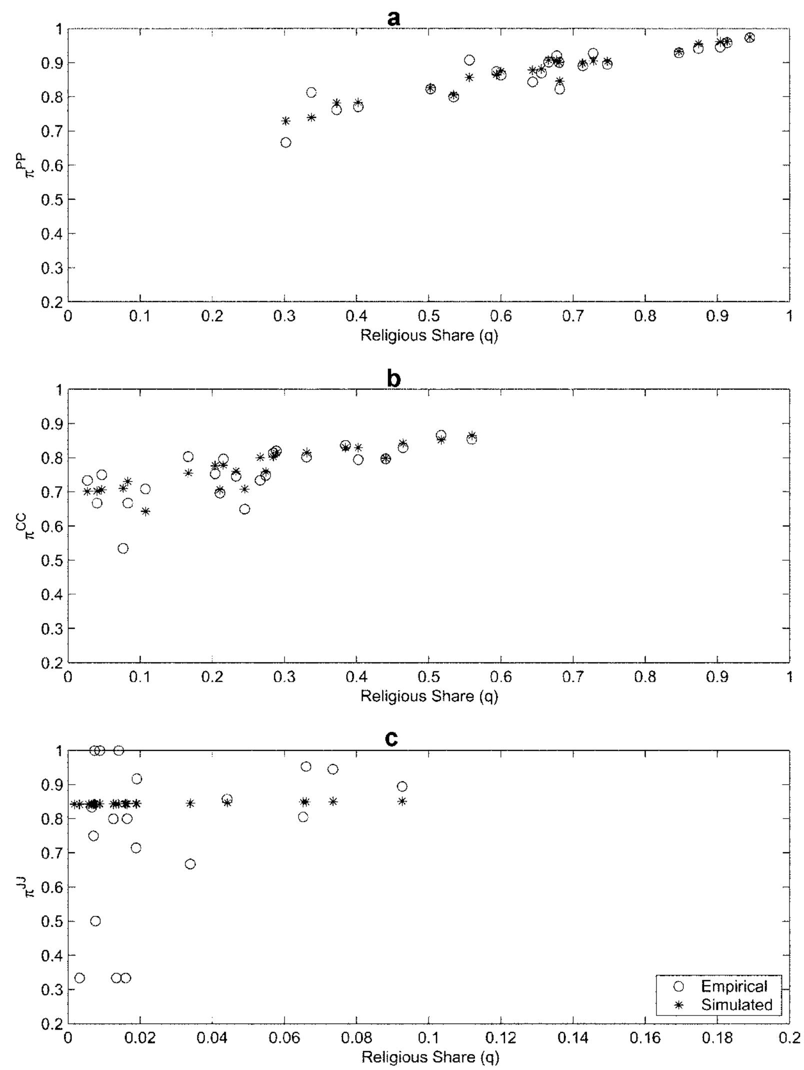

FIG. 4.—Simulated vs. empirical homogamous marriage probability. *a*, Protestants. *b*, Catholics. *c*, Jews.

marriage frequencies, $\hat{\pi}^{ii}$, to those generated by the model, $\tilde{\pi}^{ii}(\boldsymbol{\theta})$, at the estimated parameter values. Table 3 compares the empirical socialization frequencies with those implied by the model. We are able to match the homogamous socialization rates quite well, whereas we do less well in matching the heterogamous ones.

The low empirical frequencies of religious intermarriages are the

TABLE 3
SOCIALIZATION PROBABILITIES: DATA VS. MODEL

	Protestants	Catholics	Jews	Others
	A. Empirical Frequencies			
PP marriage	.9179	.0284	0	.0537
CC marriage	.0850	.8571	.0034	.0544
JJ marriage	.0370	0	.9259	.0370
OO marriage	.3231	.0462	0	.6380
PC marriage	.5116	.3140	0	.1744
PO marriage	.7100	.1000	0	.1900
CO marriage	.1667	.5000	0	.3333
	B. Simulated Frequencies from the Model			
PP marriage	.9227	.0349	.0031	.0394
CC marriage	.1078	.8293	.0065	.0564
JJ marriage	.0308	.0220	.9291	.0180
OO marriage	.1472	.0712	.0078	.7738
PC marriage	.4855	.3409	.0165	.1571
PO marriage	.5168	.1378	.0131	.3323
CO marriage	.3051	.3425	.0192	.3333

NOTE.—Each cell reports the sample probability that a child in the row marriage is a member of the column religious group. P = Protestants, C = Catholics, J = Jews, and O = Others.

consequence of a strong estimated preference by members of each group for having children who share their own religious faith, that is, of high intolerance parameters.[14] We estimate significant positive intolerance parameters (with the exception of the parameter describing attitudes toward Jews of the residual group, Others). The most striking estimates are those describing the intolerance parameters of Jews, which are about four times as high as those of any other religious group.[15]

The parameter estimates for the cost functions reveal a strong dependence of both socialization costs and marriage costs on the proportion of one's religious group in the state, ϵ_τ and ϵ_α. The more a given

[14] Since cost functions are assumed to be independent of the specific religious group, intolerance levels can be meaningfully compared across groups. In particular, the implicit unit of measure can be identified as follows. The estimated socialization cost function of any religious group i, when the group represents half of the population, $q^i = \frac{1}{2}$, can be easily computed and takes the value one for $\tau^i = .12$. Therefore, both costs and intolerance levels are measured as multiples of the cost of increasing the probability of socializing a child, for a homogamous family of a religious group that constitutes half the population, by 12 percentage points. While such a cost is not pinned down by the analysis, if we roughly identify it with the opportunity cost of spending two hours with the child per week for 10 years, it lies in the $20,000–$40,000 range (for hourly wages in the $20–$40 range).

[15] From the rough computations in n. 14, our estimates of the intolerance parameters for Jews imply a relative value of a child with maintained Jewish identity in the $10–$20 million range. We should note, though, that such a high estimate of intolerance levels for Jews could partly be a consequence of lower marriage segregation costs for Jews with respect to other religious groups. As noted in Sec. V, in fact, we would not be able to separately identify costs and intolerance levels that varied by religious group; see also the discussion in Sec. IX.

religious faith is a minority in the population of reference, the harder it is to socialize one's children to that particular faith or to segregate in marriage.

For instance, because of the small median share of Jews in the population, the cost of directly socializing a child with probability one-half is about twice as large for Jews as for Catholics.

The matching probabilities Q_s^i in the oblique socialization are biased in favor of the residual group Others. The estimated bias parameter o induces a sizable distortion: the implied probability of becoming Other in society at large, once direct family socialization fails, exceeds, on average, the share of Others in the population by about 16 percentage points. Such bias can be accounted for by conversions. The residual group Others includes a majority of individuals with no religious preference (over 70 percent, on average, in the United States of our residual group in the sample) and a minority that includes major religious faiths not largely represented in the United States (Islam, Buddhism, Hinduism, etc.) and religious sects. Proselytizing activities (broadly speaking) could account for the high conversion rate implied by our analysis of the socialization data, at least in the case of individuals with no religious preferences and individuals belonging to religious sects.

Parents' preferences are not very sensitive to the number of children in a given marriage. The estimated ξ is very close to zero, implying that parents seem to care only about their average child. Since we do not explicitly model endogenous fertility, this result must be interpreted with particular caution.

The estimated choices of direct socialization of homogamous families, τ^i, which is the differential probability of direct socialization with respect to the exogenous direct socialization rate, m, and the choice of marriage segregation in the restricted pools, α^i, are quite instructive about the implications of our results for socialization and for the marriage market. Figure 5 (respectively, fig. 6) presents the estimated τ^i (α^i) for Protestants, Catholics, and Jews as a function of q^i.[16]

For Protestants and Catholics, the direct socialization levels of homogamous families when fully minority ($q^i = 0$) are significantly positive: $\tau^i > .3$ in both cases, and $m = .35$, giving a probability of direct socialization for homogamous families when they are minorities above two-thirds. If socialization costs were independent of q^i, socialization levels would decrease with q^i. As a result of the estimated strong dependence of socialization costs on q^i, direct socialization for both Protestants and Catholics first increases and then decreases, peaking at about

[16] These plots are constructed by fixing the religious shares for two groups (e.g., Jews and Others) to their means and letting the religious share of a specific group (e.g., Protestants) increase and the share of the residual group (Catholics) decrease in order to satisfy $\sum_{i=1}^{n} q^i = 1$.

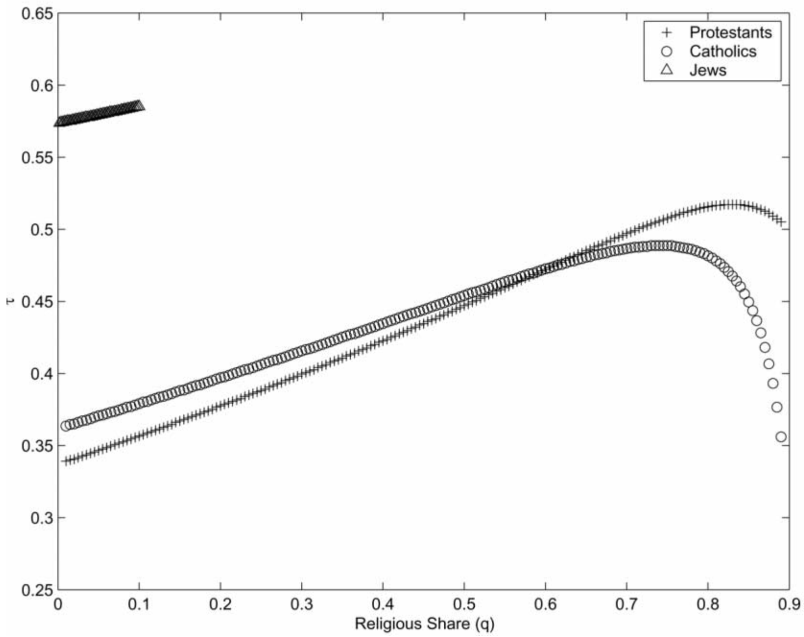

FIG. 5.—Socialization probability

$q^i = .75$. Jews socialize much more than Catholics and Protestants in the whole relevant range of q^i. For example, when Jews are a small minority, the probability of direct socialization levels for homogamous Jewish families is roughly 90 percent.

Similar considerations hold for the estimated marriage segregation probabilities in the restricted pool. As we have mentioned earlier, the marriage game that determines marriage segregation levels in the restricted pools exhibits multiple equilibria, for the estimated parameters. As a consequence, the estimated equilibrium marriage segregation levels are a discontinuous selection of the equilibrium set. For Protestants and Catholics the marriage segregation level first increases (because of the high socialization and marriage costs) and then decreases as a function of q^i. When a religious group is a small minority (i.e., when its fraction in the population is close to zero), marriage segregation in the restricted pool is about 65 percent for Catholics and about 55 percent for Protestants. Catholics have a higher α^i than Protestants for most of the range of q^i. Jews' marriage segregation does not display much variation in the relevant range of the proportion of Jews in the population; the marriage segregation level is very high in the whole range (including when fully minority), above 80 percent.

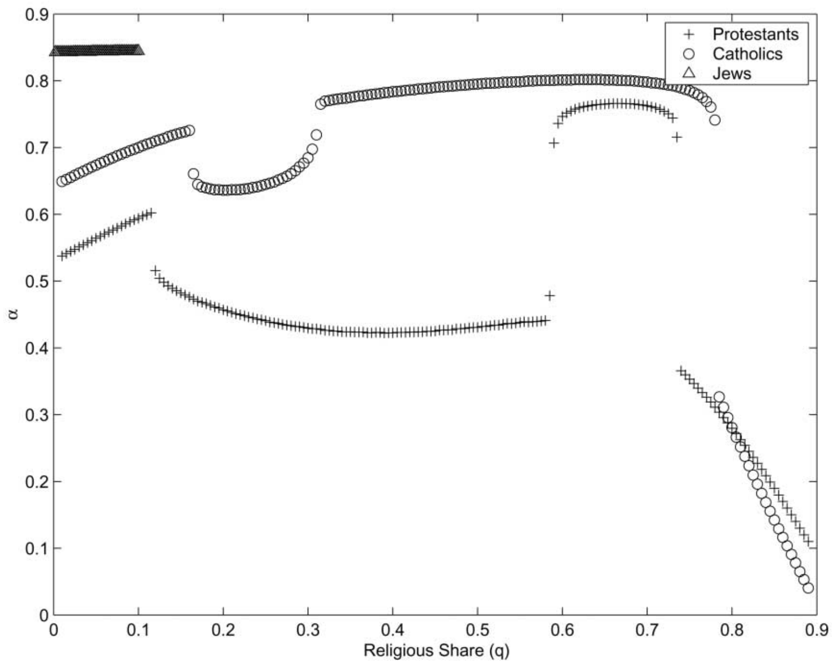

Fig. 6.—Marriage segregation probability

Finally, out-of-sample simulations of homogamous marriage probabilities, π^{ii}, implied by our parameter estimates as a function of religious share q^i (not reported; see the working paper version of the paper [Bisin et al. 2003]), indicate that when q^i is close to zero, the probability of homogamous marriage is well above that implied by random matching (which lies on the 45-degree line): for Protestants it is around .55, for Catholics it is about .65, and for Jews it is above .8. This is due to the strictly positive socialization and marriage segregation levels implied by our estimates. The simulated π^{ii} is increasing in q^i and becomes close to the probability implied by random matching only when the share of religious group i in the population approaches 90 percent.[17]

[17] It is worth noting that the discontinuities in α^i generate only small jumps in the implied homogamous marriage probabilities π^{ii}. Multiple equilibria correspond, in fact, to different equilibrium distributions of segregation costs across the different religious groups without affecting the implied homogamy and heterogamy rates much. The segregation of one group in its own restricted pool has in fact a positive externality on the other groups as they gain higher implied homogamy rates without the need of segregating in their own restricted marriage pools.

VII. Are Marriage and Socialization Endogenous?

The model of marriage and socialization we estimate is based on the behavioral assumption that marriage and socialization are endogenously determined as economic decisions of agents who have preferences for children with their own religious attitudes. In this section we aim to assess the relevance of economic behavior to explain the observed socialization and marriage rates. To this end, we conduct some tests on our baseline estimates and compare the performance of our model to several alternative specifications that make different behavioral assumptions.

Results are reported in table 4 (col. 1 reproduces our baseline estimate to simplify comparisons). Column 2 reports parameter estimates for a model in which marriage segregation choices are endogenous but socialization is exogenous. In particular, it reports the results of the estimation of the direct socialization effort, τ^i, assuming that it varies across religions but not across states. Column 3 examines instead a model in which both marriage and socialization are exogenous. Both α^i and τ^i are estimated to vary across religions but not across states. Finally, column 4 reports estimation results for a model in which the value of a homogamous marriage ($W^{ii} - W^{ih}$) is exogenous and independent of own group's share in the geographic state. This is an attempt to capture alternative explanations of the high prevalence of homogamous marriages, where the benefits of homogamy are intrinsic in homogamous marriage unions and therefore constant across states. Examples of some alternative explanations are that the benefit of homogamous marriages resides in the possible advantages (consumption value) of sharing the same cultural representation of life and society or in the homogamous marriages' inherent stability with respect to divorce (see Becker et al. 1977; Heaton 1984; Lehrer and Chiswick 1993).

The rankings of the Sargan test of the overidentifying restrictions suggest that all three alternative models do not fit the data nearly as well as our baseline model: p-values vary between .02 and .0017, compared with .11 in our estimate. However, a formal test comparing our baseline model with the alternative specifications we are interested in is not straightforward since the models are nonnested. We adopt, therefore, a procedure to compare nonnested models first introduced by Vuong (1989) and further developed by Kitamura (2000). The procedure consists of determining the distance of each model from the true distribution that generates the data, where the distance is measured by the Kullback-Leibler information criterion (KLIC); see the working paper version of this paper (Bisin et al. 2003) for a more detailed discussion. The results of the test are reported at the bottom of table 4. Our model performs better, in the KLIC sense, than each of the three al-

TABLE 4
Alternative Specifications

	Baseline Model		Exogenous Socialization		Exogenous Socialization and Marriage		No Socialization	
Value of same-religion child	526.2563	(102.3493)	590.4655	(38.0190)				
Intolerance of:								
P toward C	125.2046	(2.3046)	84.8239	(4.1833)				
P toward J	121.3225	(9.7949)	55.1652	(13.9895)				
P toward O	31.9216	(4.7295)	14.9304	(1.8718)				
C toward P	152.9794	(2.9070)	91.7336	(3.0832)				
C toward J	14.9936	(5.0554)	13.8521	(3.0555)				
C toward O	12.1115	(2.6230)	4.2515	(1.6202)				
J toward P	501.2928	(82.7201)	487.1561	(89.4113)				
J toward C	526.2551	(222.0905)	583.8977	(110.2248)				
J toward O	525.9072	(77.8761)	448.0000	(84.5392)				
O toward P	106.0829	(9.1286)	65.5840	(11.1779)				
O toward C	165.5395	(21.0772)	125.5341	(19.1617)				
O toward J	.7108	(173.2716)	220.5076	(54.0039)				
Cost parameter:								
σ_τ	1.9227	(.4263)						
σ_α	5.9062	(.4156)	1.8331	(.2895)			4.4286	(1.9803)
ϵ_τ	69.4675	(5.8155)						
ϵ_α	67.1701	(2.7009)	69.7436	(2.8974)			75.6543	(15.0092)
λ_τ	.6773	(.0417)						
λ_α	.9996	(.0001)	.9996	(.0001)			.9990	(.0004)
Exogenous direct socialization: m	.3457	(.0201)	.3239	(.0207)	.3366	(.0781)		
Conversions to Others: o	.2062	(.0155)	.1866	(.0394)	.1877	(.0446)		
Fertility Parameter: ξ	.0108	(.0093)	.0158	(.0084)			.0002	(.0033)
τ:								
P			.4740	(.0140)	.4778	(.0833)		
C			.5004	(.0253)	.4648	(.0834)		

J		.6088	(.0466)	.5960	(.0927)		
O		.2711	(.0355)	.2598	(.1064)		
α:							
P				.6412	(.0388)		
C				.7149	(.0148)		
J				.8630	(.0167)		
O				.4753	(.0428)		
Marriage value:							
PP						376.6728	(10.2635)
PC						332.5983	(15.0042)
PJ						309.0455	(54.4309)
PO						367.7591	(13.1458)
CC						457.7011	(14.4514)
CP						395.9555	(16.9939)
CJ						457.6879	(55.7026)
CO						457.6584	(23.1344)
JJ						599.1997	(163.2456)
JP						.3271	(270.6800)
JC						1.3832	(116.6423)
JO						569.5244	(1,310.2941)
OO						500.6260	(22.9707)
OP						482.7180	(24.0713)
OC						423.6760	(31.8818)
OJ						307.5558	(66.2610)
p-value of *J*-test:							
Marriage only	.1129	.0239		.0017		.0147	
Overall	.0173	.0047		.0001			
Kitamura test*		−18.8738		−15.4801		−4.8318†	

Note.—Standard errors are in parentheses. P = Protestants, C = Catholics, J = Jews, and O = Others.
* The test statistic is distributed as a standard normal under the null; a negative value indicates that the baseline is better than the alternative.
† For comparability, the baseline model was reestimated to match empirical intermarriage rates only.

ternative specifications examined here.[18] The p-values of the tests are practically zero.[19] The estimates reported in column 4 of table 4 refer to a model in which the subjective benefits from marrying homogamously stay constant across states and are estimated to match the observed homogamous and heterogamous marriage rates. The good fit of this model suggests that the benefits of homogamy might also contain components that are intrinsic to homogamous marriages and therefore invariant with respect to the distribution of the population by religious group.

We take our analysis of alternative models to suggest that endogenous socialization and marriage segregation are indeed an important part of marriage and socialization mechanisms with respect to religious trait.

VIII. Long-Run Dynamics of the Distribution of Religious Groups

Given the distribution of the population by religious group at some time t, the marriage and socialization mechanisms we estimated determine the distribution of the population in the successive generation, at time $t+1$.

The difference equation ruling the dynamics of the distribution of religious traits in the population is[20]

$$q_{t+1}^{i} = \frac{N_t}{N_{t+1}} \sum_j q_t^j \sum_h \pi_t^{jh} \frac{n^{jh}}{2} P_{jh,t}^{i}, \tag{16}$$

where N_t denotes the total number of adults at time t. The evolution of N_t can be obtained by studying the evolution of the number of adults for each religious group, N_t^i:

$$N_t = \sum_i N_t^i \tag{17}$$

and

$$N_{t+1}^{i} = \sum_j N_t^j \sum_h \pi_t^{jh} \frac{n^{jh}}{2} P_{jh,t}^{i}. \tag{18}$$

[18] In order to properly compare our baseline model to the model in col. 4, we reestimated our model to match only the empirical intermarriage rates $\hat{\pi}$, since the alternative model does not have implications for the socialization rates P_{ij}^{k}.

[19] It is interesting to note that the ranking of the four models is different from that implied by comparing the Sargan specification test results. In particular, the model in col. 4 performs best among the alternative specifications.

[20] The dynamics and local stability properties of stationary states are studied, for a simple version of this economy, in Bisin et al. (2000).

A. Simulations of the Dynamics

Using the estimated structural parameters and the empirical religious composition of several U.S. states as initial conditions, we can simulate the evolution of the distribution of the population by religious group, q_s^i, over time. It is worth noting that such simulation exercises are necessarily based on the assumption that the parameters estimated are stable and therefore constant over time.[21]

A time period in the simulation is a generation, and therefore, stability of the parameters over the 40 or so generations that the distribution of the population takes to reach a stationary state is impossible to maintain. As a consequence, the simulations we report are aimed at illustrating the implications of our estimation results and should not be interpreted as direct forecasts of the future prevalence of the different religious denominations.

We use the current composition of California, Illinois, New York, and Texas as initial conditions to illustrate the dynamic paths implied by our estimates. Results are reported in figure 7.[22]

We find two different stationary distributions of the population by religious trait, which are attractive for different sets of initial conditions: one has a large majority of Protestants (about 90 percent) and a minority of the residual group, Others (about 10 percent); the other is uniquely composed of Jews. The stationary state in which only Jews are represented is attractive, for instance, for the initial composition of the populations of Illinois and New York. The stationary state composed of Protestants and Others is attractive for the initial conditions of California and Texas. The population settles into a stationary distribution in at most 45 periods (a period should be interpreted as a generation, i.e., 25–30 years).

The dependence of the dynamics on the initial conditions is interesting and complex. For instance, even though the initial proportion of Jews is higher in California than in Illinois, their share rises exponentially in Illinois whereas it declines quickly in California. The reason is that in Illinois Catholics are well represented in the initial distribution (about 40 percent) and Catholics have an estimated low intolerance level toward Jews. Protestants, on the other hand, have lower intolerance toward Others than Jews (by a factor of four), and as a consequence, when Protestants are a large majority in the initial conditions, Jews tend

[21] Since we estimate the deep preference and technology parameters of the marriage and socialization model, stability is less severe an assumption here than in the case in which behavioral rules are directly estimated and the simulations of the population composition dynamics are obtained by linear extrapolations from such rules.

[22] The results of the simulations are robust with respect to variations of the estimated parameters in their confidence interval; see the working paper (Bisin et al. 2003) for details.

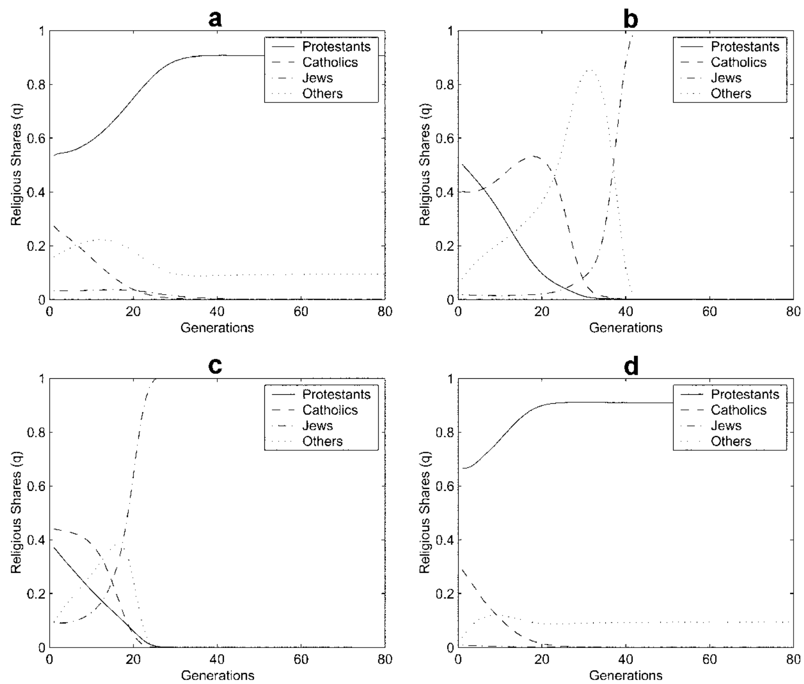

FIG. 7.—Long-run dynamics. *a*, California. *b*, Illinois. *c*, New York. *d*, Texas.

to decline and Others gain a small but stable share of the population; this happens, for instance, for the initial conditions represented by the present composition of California and Texas.

The relative success of Others in the simulations is made even more striking when one notices that the average fertility rate of this group is below reproduction (less than two) and is particularly low for homogamous marriages (less than 1.7).

Catholics are never present in the stationary distributions. The reason is that we estimate very high intolerance levels toward Catholics for all other religious denominations, including Others (see table 2).[23]

Our simulations are in striking contrast with the triple melting pot prediction as well as the predictions concerning the vanishing of the U.S. Jewish population derived from the National Jewish Population data. These predictions are derived from linear extrapolations of inter-

[23] Simulations reported in the working paper version of this paper (Bisin et al. 2003) show that even increasing the fertility rate of Catholics to account for the Hispanic Catholic migration does not generate a transition toward a path converging to a stationary state in which Catholics are present. Rather, accounting for such migration flows into the Catholic population has the effect of favoring Jews in the long run.

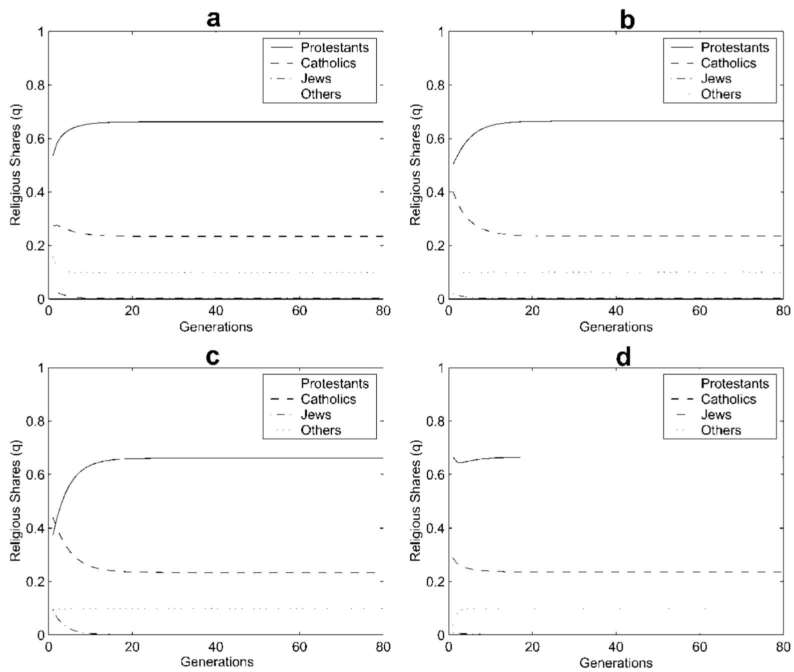

FIG. 8.—Long-run dynamics: extrapolations from constant (π, P). *a*, California. *b*, Illinois. *c*, New York. *d*, Texas.

marriage data and hence, contrary to our methodology, do not account for the nonlinearities in the way marriage and socialization rules depend on the distribution of the population by religious groups.

To clarify this point, we have also simulated the dynamics of the distribution of the population by religious group by linear extrapolation under two alternative assumptions: (i) the average marriage and socialization rates in the United States are constant (and are set equal to the observed rates), and (ii) the behavioral rules for marriage segregation and socialization of the different groups are constant (and use the estimated values of α^i and $\tau^i + m$, for each i, in the alternative model with exogenous socialization and marriage introduced in Sec. VI and col. 3 of table 4). Results for the initial conditions of California, Illinois, New York, and Texas are reported in figures 8 and 9. In accordance with the triple melting pot prediction, revised to account for the recent increase in the intermarriage rates of Jews, the simulations obtained with constant marriage and socialization rates predict a limit population composed mainly of Catholics and Protestants, the major religious groups in the initial conditions. A statistical artifact of random matching is that such groups have low intermarriage rates. The Jewish population

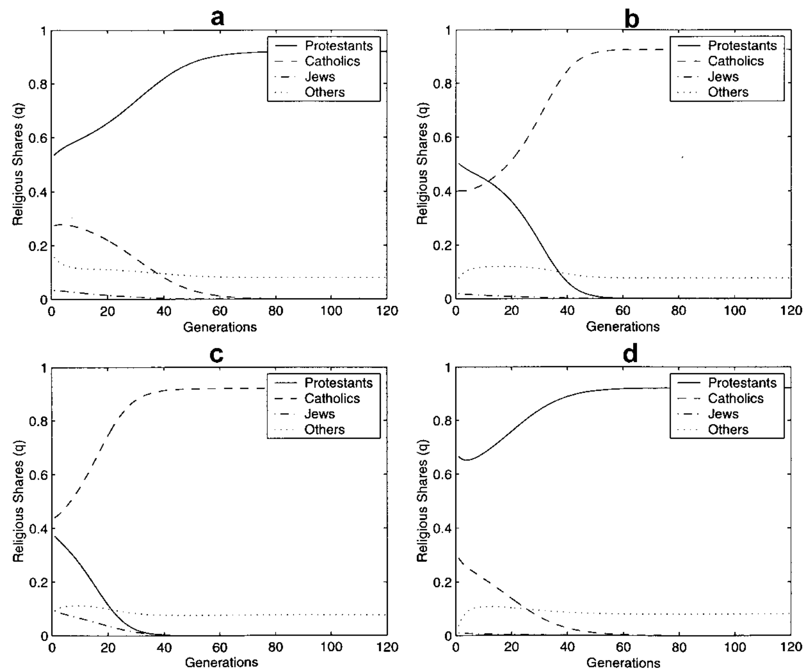

FIG. 9.—Long-run dynamics: extrapolations from estimated constant (α, τ). *a*, California. *b*, Illinois. *c*, New York. *d*, Texas.

in fact vanishes, whereas the conversion bias guarantees a small population of Others in the limit (Protestants account for 66 percent of the unique stationary distribution in this case, Catholics for 24 percent, and Others for 10 percent; see fig. 8). If, on the other hand, we extrapolate from the constant estimated behavioral rules for marriage segregation and socialization of the different groups, we already find evidence against the triple melting pot: Catholics and Protestants cannot coexist in the limit, and which group is represented in the stationary distribution depends on the initial conditions. But in all these simulations, Jews still vanish (and Others maintain a 10 percent share of the population; see fig. 9).

IX. Discussion

In order to empirically estimate the marriage and socialization processes along the religious dimension in the United States, we assume that the geographical unit of reference coincides with the state. Moreover, while our analysis treats marriage as an economic decision of each agent, motivated by his or her preferences regarding the socialization of children, we treat the distribution of agents by state as fixed. In other words, we do not consider the agents' moving decisions as endogenous. This is done for obvious data limitations, since the GSS survey records the residence of the individual respondent only at the moment of the survey and whether the respondent ever moved in the past. The endogeneity of moving decisions might be problematic for our estimates, in principle, if these decisions were motivated in part by marriage and socialization and if they caused some unobserved heterogeneity that could otherwise explain our results. We turn, therefore, to addressing these issues in the context of our analysis of marriage and socialization by religious traits in the United States.

Small religious communities are often concentrated in religiously homogeneous city neighborhoods or counties, for instance, the Orthodox Jews of Boro Park, Brooklyn, New York, or the Amish of Lancaster County, Pennsylvania. The relevant marriage pools of the members of such "enclaves" are not the state, as our model postulates, but rather the enclave itself. However, this is less of a problem for our estimates than it might appear. Consider the Jews of Boro Park as an illustration (see Mayer [1979] for a sociological analysis of this community). Modeling Boro Park as a geographical unit of our analysis would imply considering the Jews living in this neighborhood as a majority in which marriages are homogamous and children are socialized with minimal effort by families. But while it is true that living in Boro Park essentially ensures homogamy and socialization, we claim it is incorrect to conclude that this is achieved with minimal effort on the part of families. This argument in fact disregards the effort and costs associated with moving into such small close-knit communities and the cost associated with not moving out of them (see, e.g., Borjas [1995] for an analysis of such costs for ethnic communities). In the analysis of this paper, instead, the Jews of Boro Park are a small minority in New York State; it is costly for them to marry homogamously and socialize their children, exactly because it requires some form of segregation. Their high homogamy and socialization rates in the data therefore are attributed to the intensity of their preferences for transmitting their own religious faith to their children and, hence, contribute to the estimation of high rates of intolerance.

As we noted when discussing identification in Section V, without independent data on segregation and socialization costs, our estimation procedure attributes any difference in such costs across religious groups to differences in intolerance levels. The existence of religious *enclaves* such as Boro Park could therefore be also due to differentially low segregation costs for Jews. In particular, suppose that for any reason it is particularly difficult for a religious group to match within the common pool (e.g., in the case of Jews, anti-Semitism may imply—in an extension of the model—a lower probability of marriage in the common pool).

In this case, the opportunity cost to enter the restricted pool for the members of this group is low, and the formation of *enclaves* is facilitated.[24]

Another potentially serious problem, which is a consequence of not considering the endogeneity of moving decisions, is due to the unobserved heterogeneity regarding the intensity of religious preferences. Small religious communities could, for instance, have more intense religious preferences because those individuals with limited religious identity and attachment to the "land of the ancestors" have left over the generations. In this case, our estimation procedure would erroneously attribute high marriage and socialization rates of minorities to high intolerance rates of the religious denomination as a whole rather than to the religious intensity of the minority itself. We would then overestimate the intolerance levels of religious groups for which minorities have a relatively higher intensity of religious preferences. The reasons are that (i) our identification procedure requires identical preferences within religious groups, and therefore it disregards any heterogeneity in religious intensity; and (ii) it is exactly the homogamy and socialization rates of minorities, and hence their intolerance levels, that determine in our estimation procedure the intolerance levels attributed to the whole religious group (homogamy and socialization rates of majorities are due mostly to random matching).

Within the limitations of the data, we have attempted to address these issues, using data on "church attendance" from the GSS survey over the period 1972–96 as a measure of religious intensity (for Others, we obviously exclude individuals who express no religious preference). While there is some evidence of unobserved heterogeneity with regard to religious intensity, it turns out that it is in larger communities that members display higher religious intensity, rather than in smaller communities. Therefore, if anything, we are in effect underestimating intolerance levels in our analysis.

In particular, for Jews, Catholics, and Others, average attendance does not depend on the religious share of the state of residence (the GSS does not record the state of origin); this is true for both movers (the subset of agents who moved across states prior to the GSS interview) and nonmovers separately. For Protestants instead, attendance is positively correlated with the share of Protestants in the state of residence, for both movers and nonmovers (see panel A of table 5). By grouping agents into low and high intensity and then looking at the distribution of movers by religious share in the state of residence, one finds essentially the same pattern (see panel B of table 5). For Protestants, the ratio of high-intensity to low-intensity movers is higher for states in which

[24] We owe thanks to an anonymous referee for this remark.

TABLE 5
A. Religious Intensity and Religious Share

	Movers	Nonmovers	All
Protestants	.173	.260	.277
	(.081)	(.037)	(.052)
Catholics	−.024	−.032	−.026
	(.061)	(.037)	(.061)
Jews	.004	−.004	.001
	(.004)	(.004)	(.005)

B. Ratio of High- to Low-Intensity Movers, by Religious Share

	1st Quartile	2d Quartile	3d Quartile	4th Quartile
Protestants	.588	.561	.647	.840
Catholics	.904	.834	1.351	.855
Jews	.130	.053	.131	.186
Others	.127	.138	.120	.096

Note.—In panel A, each cell reports the ordinary least squares regression coefficient of average attendance on religious share in the current state of residence (q). Standard errors are in parentheses. In panel B, each cell reports the ratio of high- to low-intensity movers, by quartiles of the distribution of q (religious share in the current state of residence). High intensity corresponds to individuals who attend religious services nearly every week or more frequently.

Protestants are a majority; the same is true for Jews, but no such relationship appears in the data for Catholics or Others.

Because the estimates of the intolerance levels of the different religious groups are so central in our analysis, we have further pursued the analysis of the implication of unobserved heterogeneity with respect to religious intensity. We have estimated our structural model of marriage and socialization under the assumptions that, given a baseline intolerance level for religion i, ΔV^i, the intolerance level of minorities (defined as communities in the lowest quartile of the distribution of q^i) is either $0.75 \cdot \Delta V^i$ or $1.25 \cdot \Delta V^i$. Consistent with the analysis we just reported of the correlation of average attendance and religious share, we find that the model in which the intolerance level of minorities is smaller ($0.75 \cdot \Delta V^i$) fares better, in the KLIC sense, than the model in which it is larger.[25] Neither model fits the data quite as well as our baseline model, in which intolerance levels are assumed invariant across religious shares, that is, for both minorities and majorities.[26] We conclude, therefore, that this study of religious intensity, albeit necessarily coarse, supports our baseline model, which restricts the preferences of individuals of the same religious group to be identical in terms of intensity. If a bias due to the unobserved heterogeneity of religious intensity is present in our

[25] The Kitamura test statistic is −14.625, with a p-value very close to zero.

[26] The Kitamura test statistic when we test our baseline model against the two specifications in which minorities are less (more) intolerant is −7.796 (−20.219), with p-values practically zero.

analysis, it leads us to underestimate the importance of intolerance levels in determining the high observed homogamy and socialization rates.[27]

Alternative explanations of the evidence regarding the high socialization and marriage segregation levels of small religious communities can be developed that rely on various other possible instances of unobserved heterogeneity across agents. High homogamy rates of minorities could also be rationalized, for instance, if small religious communities are more homogeneous in some dimension along which marriage would be assortative. In this case, in fact, high religious homogamy rates in small religious communities would be a statistical artifact of the assortativeness of marriages along dimensions other than religious faith. Natural examples might consist of race and education levels. It is well known, for instance, that individuals tend to marry spouses with a similar education level (respectively, of the same race). Hence, if small religious communities are more homogeneous in terms of education (race) than larger communities, we would observe disproportionately high religious homogamy rates in small religious communities. While the correlation of the coefficient of variation of race between the members of religious group i and the population share of religion i is zero or negative for all groups except Others, the correlation of the coefficient of variation of education between the members of religious group i and the population share of religion i is in fact positive for Protestants and Catholics (.5 and .37, respectively). We conclude that the homogeneity of education levels (but not of race) could contribute to explaining the socialization and marriage behavior of minorities. Considering religious faith and education levels as joint determinants of the assortativeness of marriage rates is potentially very important (see Sec. X).

Another dimension in which our analysis of marriage homogamy and socialization by religious denomination in the United States could be problematic consists of our definition of the religious groups, which aggregates several potentially very different subgroups. The residual group, Others, is obviously the most heterogeneous in terms of beliefs and cultural characteristics, but substantial heterogeneity also characterizes Protestants and even Jews (perhaps less so Catholics). Our assumption that preferences are identical within any religious groups requires, therefore, some scrutiny: our analysis might be mistakenly aggregating over different preference parameters of, say, black and white Protestants or Catholics of Hispanic and non-Hispanic ethnic origin. To address this issue, we reestimate our model after expanding the religious

[27] It is worth mentioning that we have also estimated our baseline model using data for nonmovers only, and the point estimates are quite similar to those obtained using the whole sample.

TABLE 6
Structural Estimation Results: Five Groups

	Estimate	Standard Error	p-Value
Value of same-religion child	597.4714	46.3771	.0000
Intolerance of:			
W toward B	593.8036	170.3606	.0002
W toward C	.1430	130.9795	.4996
W toward J	23.3231	196.0469	.4527
W toward O	453.3800	90.5955	.0000
B toward W	124.3143	6.0569	.0000
B toward C	.0002	84.3807	.5000
B toward J	.0563	10.1429	.4978
B toward O	.0001	25.4890	.5000
C toward W	211.0727	27.3175	.0000
C toward B	525.5415	286.4389	.0333
C toward J	458.3937	336.2965	.0864
C toward O	19.4319	56.1801	.3647
J toward W	597.4630	127.7992	.0000
J toward B	597.3089	325.6189	.0333
J toward C	597.4679	104.8939	.0000
J toward O	591.5346	191.4764	.0010
O toward W	287.5699	10.5128	.0000
O toward B	60.2092	87.6056	.2460
O toward C	285.7005	18.3575	.0000
O toward J	3.6374	25.1521	.4425
Cost parameter:			
σ_τ	4.6251	.5346	.0000
σ_α	.0018	.0718	.4898
ϵ_τ	70.6795	5.3932	.0000
ϵ_α	2.9946	.2658	.0000
λ_τ	.3897	.0567	.0000
λ_α	.6424	.0318	.0000
Exogenous direct socialization: m	.3582	.0179	.0000
Conversions to Others: o	.1154	.0352	.0005
Fertility parameter: ξ	.9654	.0698	.0000
J-test:			
Marriage only (240 degrees of freedom)	1,028.5		.00
Overall (261 degrees of freedom)	1,075.6		.00

Note.—W = white Protestants, B = black Protestants, C = Catholics, J = Jews, and O = Others.

groups to include black and white Protestants separately.[28] The structural estimation results in this case with five religious groups are reported in table 6.

We interpret the results here, even though a word of caution is nec-

[28] We have also reestimated the model treating separately black and white Protestants as well as Hispanic and non-Hispanic Catholics. We do not discuss here such estimates (but see the working paper version of the paper [Bisin et al. 2003]) because they are too imprecise as a result of the small sample size in several cells (the overall number of usable observations is reduced to 8,147 in 16 states). However, most of the point estimates agree with our baseline estimates: interestingly, the distinction between Hispanic and non-Hispanic Catholics seems to matter very little.

essary since the point estimates have fairly large standard errors (owing to the relatively large number of moments used in the estimation and the relatively small sample size of several cells). First, black and white Protestants appear in fact as two quite distinct groups, with elevated estimated intolerance levels for each other. This is especially so for white Protestants toward blacks rather than vice versa. Second, the elevated intolerance levels of Jews and the small intolerance levels of Catholics we estimated in our baseline model with four groups hold with the extension to five groups; similarly, the conversion parameter in favor of the residual group, Others, remains stable.

Since the intolerance parameters are difficult to interpret per se, we find it useful to further assess the robustness of our analysis by running the simulations of the dynamics of the distribution of the population by religious trait for the estimated model with five groups. For all initial conditions, a unique stationary state appears, composed of a majority of white Protestants and a minority of Others. In contrast to our baseline estimates with four groups, a stationary state with a majority of Jews does not appear in the version with five groups. However, a slight perturbation of the parameter values allows the stationary state with a majority of Jews to reemerge and the dynamics of the population distribution to converge to it from the initial conditions of New York and Illinois.[29] We conclude that our baseline analysis of marriage and socialization with four groups is relatively robust to the extension to five groups, that is, to separating Protestants into blacks and whites. The race distinction is not irrelevant for religious segregation, though, and while we cannot pursue it further because of data limitations, our analysis suggests the presence of interesting phenomena related to the interaction of race and religion in marriage and socialization.

X. Conclusions

We have concentrated our analysis on a simple dimension of assortativeness in marriage: religious homogamy. It is well documented, however, that marriages are assortative in several other dimensions, including education (see, e.g., the survey of Mare [1991]) and race. The possible correlation between education, as well as other characteristics of spouses, and religious faith can, therefore, have important implications for our analysis of marriage and socialization. As we noticed, for instance, the relative homogeneity with respect to education of small

[29] In particular, the critical parameter for the survival of Jews in the simulations is the intolerance of white Protestants toward Others, ΔV^{WO}: it is sufficient to reduce it slightly (well inside its 99 percent confidence interval) to observe a majority of Jews in the stationary state with the initial conditions of New York and Illinois.

religious communities could at least in part explain the observed high socialization and homogamy rates of minorities.

We have conducted a preliminary analysis of a model in which assortative marriage can occur along both the education and religious dimensions. This indicates that, overall, our socialization-based interpretation of intermarriage rates is robust to the inclusion of a preference for educated spouses. However, the many interesting interactions between the preferences for education and the preferences for religious homogamy require an independent treatment.[30] The same, we argue, can be said for the interaction of religious homogamy, socialization, and race.

Similarly, a more detailed analysis of the endogenous determinants of fertility is bound to be of great value, especially with regard to the simulations of the dynamics of the distribution of the population. For instance, a significantly higher estimated sensitivity of parental preferences with respect to fertility might have an important effect on the long-run distribution of the population by religious groups.

Also, we have dealt only marginally with the issue of conversions. Our analysis indicates that the issue has substantial relevance for the dynamics of the relative shares of the religious groups. A substantial literature has stressed the role of conversions in marriage for religious socialization (see Iannaccone 1990).

Finally, data limitations do not allow us to fully disentangle intolerance levels from segregation and socialization costs, nor to consider several other important issues related to marriage and socialization: the effects of mobility in the determination of the relevant marriage pools, gender asymmetries in socialization, and the hierarchical representation of different religious groups in the social psychology of the United States, to cite only some of such issues.

Appendix A

The Structural Model and Structural Equations

The structural model is introduced in Section III. After some algebraic manipulations and after we take explicit account of the geographic state index, the socialization and marriage model introduced in the paper is simply represented by the following system of equations:

$$\pi_s^{ii} = \alpha_s^i + (1 - \alpha_s^i)A_s^{ii} \quad \forall i, \tag{A1}$$

$$\pi_s^{ij} = (1 - \alpha_s^i)A_s^{ij} \quad \forall i \neq j, \tag{A2}$$

[30] Sacerdote and Glaeser (2001) study empirically the relationship between education and religion but center on religious attendance rather than on marriage and socialization rates.

$$P^{i}_{ii,s} = \tilde{\tau}^{i}_{s} + (1 - \tilde{\tau}^{i}_{s})Q^{i}_{s}, \tag{A3}$$

$$P^{j}_{ii,s} = (1 - \tilde{\tau}^{i}_{s})Q^{j}_{s}, \quad i,\ j \text{ distinct}, \tag{A4}$$

$$P^{i}_{ij,s} = \tfrac{1}{2}m + (1 - m)Q^{i}_{s}, \quad i,\ j \text{ distinct}, \tag{A5}$$

$$P^{k}_{ij,s} = (1 - m)Q^{k}_{s}, \quad i,\ j,\ k \text{ distinct}, \tag{A6}$$

$$\frac{\partial S}{\partial \tau^{i}_{s}}(\tau^{i}_{s},\ q^{i}_{s}) = \sum_{j} Q^{j}_{s}\Delta V^{ij} \quad \forall i, \tag{A7}$$

$$\begin{aligned}\frac{\partial M}{\partial \alpha^{i}_{s}}(\alpha^{i}_{s},\ q^{i}_{s}) = {} & \left[(1 - A^{ii}_{s})(n^{ii})^{\xi}\tilde{\tau}^{i}_{s} - m\sum_{j\neq i} A^{ij}_{s}(n^{ij})^{\xi}\right]\sum_{h} q^{h}_{s}\Delta V^{ih} \\ & + \frac{m}{2}\sum_{j\neq i} A^{ij}_{s}(n^{ij})^{\xi}\Delta V^{ij} + \left\{\sum_{j\neq i} A^{ij}_{s}[(n^{ii})^{\xi} - (n^{ij})^{\xi}]\right\}\sum_{h} q^{h}_{s}V^{ih} \\ & - (1 - A^{ii}_{s})(n^{ii})^{\xi}S(\tau^{i}_{s},\ q^{i}_{s}),\end{aligned} \tag{A8}$$

$$A^{ii}_{s} = \frac{(1 - \alpha^{i}_{s})q^{i}_{s}}{\sum_{k=1}^{n}(1 - \alpha^{k}_{s})q^{k}_{s}} \quad \forall i, \tag{A9}$$

and

$$A^{ij}_{s} = \frac{(1 - \alpha^{j}_{s})q^{j}_{s}}{\sum_{k=1}^{n}(1 - \alpha^{k}_{s})q^{k}_{s}} \quad \forall i \neq j. \tag{A10}$$

We make the following assumption.

Assumption 1. The cost functions $M(\alpha^{i},\ q^{i})$ and $S(\tau^{i},\ q^{i})$ are differentiably strictly convex and satisfy the Inada conditions for interiority: for all $q^{i} \in (0, 1)$,

$$\lim_{\alpha^{i}\to 1} M(\alpha^{i},\ q^{i}) = \infty, \quad \lim_{\tau^{i}\to 1} S(\tau^{i},\ q^{i}) = \infty.$$

We can now prove the following theorem (see Bisin et al. 2003).

Theorem 1. The solution of equations (A1)–(A10) and (11)–(13), given q^{i}_{s} for all i and s and n^{ij} for all i and j, defines a mapping, $\tilde{\boldsymbol{\Pi}}(\boldsymbol{\theta})$, from $\boldsymbol{\theta}$ into π^{ij}_{s} for all i, j, and s and into P^{k}_{ij} for all i, j, and k. Under assumption 1, such a map is an upper-hemicontinuous correspondence and is smooth except at points of discontinuity. Moreover, it is sufficient for costs to be independent of q^{i}_{s} for the map to be a continuously differentiable function.

Appendix B

Data and Methodology

The data for the empirical exercise come from the General Social Survey. We report religious shares and intermarriage rates by state in table B1.[31] In the survey, respondents each report their religious affiliation and that of their spouse.[32] Respondents are considered a representative sample of the religious affiliations of individuals in each state of residence. In the empirical implementation, we consider four religious groups (Protestants, Catholics, Jews, and the residual group, Others), $i, j = \mathrm{P, C, J}$, and O. The GSS survey does distinguish between individuals who prefer "no religion" and those with religious faiths other than Protestant, Catholic, and Jewish. The dimension of the sample, though, forces us to aggregate the last two groups of individuals into one group, which we call Others.[33]

The total number of respondents is 35,284. We eliminate respondents who are not married at the time of the survey or are divorcees or those for whom we lack information about own or spouse religion. This leaves us with 16,722 observations. The 23 states we consider are California, Colorado, Connecticut, Florida, Georgia, Illinois, Indiana, Maryland, Massachusetts, Michigan, Minnesota, Missouri, New Jersey, New York, North Carolina, Ohio, Oregon, Pennsylvania, South Carolina, Tennessee, Texas, Virginia, and Wisconsin. For the remaining states we do not have sufficient data to estimate all the q_s^i religious shares. This final filter brings the number of observations down to 13,790.

We also estimated the model excluding respondents who changed residence at some time before the survey, since we do not know the state in which they resided when they got married. This reduces the number of observations to 7,286 and reduces the number of states for which we have sufficient data to 15. While the fit is much worse and the estimates are quite imprecise in this case, the point estimates are very similar to those with the whole sample.

Since respondents are the sampling unit in the GSS survey, we construct our measure of the religious composition of marriages and our measure of the religious composition of the population on the basis of *respondents* rather than marriages.[34] We report in table B1 the shares of each religious group as well as the sample probability that a member of each religious group i in our subdivision marries homogamously or with a member of religious group $j \neq i$, by state.

[31] Most of the data are publicly available at http://www.icpsr.umich.edu/, with the exception of the state geocodes for all respondents, which are available on request only from the National Opinion Research Center. The GSS is a nearly annual national survey of U.S. residents that focuses on attitudes, perceptions, and social trends, in addition to more conventional socioeconomic and demographic characteristics of respondents. An incomplete list of topics covered over the years includes class, religion, politics, sex, and health issues. The cumulative data set covers the period 1972–96.

[32] The exact text of the question in the GSS is "What is your religious preference? Is it Protestant, Catholic, Jewish, some other religion, or no religion?"

[33] Overall, individuals with no religious preference are the majority of our residual group Others, accounting for about 77 percent of Others, on average, over the United States, with a maximum of 86 percent in Tennessee and a minimum of 60 percent in Maryland.

[34] Independent measures of religious affiliation by state do exist, e.g., the American Religion Data Archive (ARDA) data set. Such measures grossly underestimate the members of each denomination as a fraction of the population, since they are based on church membership. We have replicated our estimates with the ARDA measures of religious affiliation. While the model's fit is worse, the qualitative structure of the parameter estimates remains unchanged.

TABLE B1

Religious Shares and Intermarriage Rates by State

	Religious Shares				Marriage Probabilities: Protestants				Marriage Probabilities: Catholics				Marriage Probabilities: Jews				Marriage Probabilities: Others			
State	P	C	J	O	Pi-PP	Pi-PC	Pi-PJ	Pi-PO	Pi-CP	Pi-CC	Pi-CJ	Pi-CO	Pi-JP	Pi-JC	Pi-JJ	Pi-JO	Pi-OP	Pi-OC	Pi-OJ	Pi-OO
California	.5345	.2742	.0339	.1575	.7993	.0974	.0153	.0880	.1762	.7483	.0114	.0641	.1667	.1296	.6667	.0370	.2629	.1355	.0120	.5896
Colorado	.6560	.2107	.0160	.1173	.8699	.0935	.0000	.0366	.1646	.6962	.0000	.1392	.3333	.3333	.3333	.0000	.3182	.1136	.0227	.5455
Connecticut	.3375	.5174	.0442	.1009	.8131	.1308	.0000	.0561	.1159	.8659	.0000	.0183	.0000	.1429	.8571	.0000	.2500	.3125	.0625	.3750
Florida	.6806	.2332	.0189	.0674	.9010	.0574	.0079	.0337	.1850	.7457	.0116	.0578	.0714	.2143	.7143	.0000	.4200	.1400	.0000	.4400
Georgia	.9040	.0468	.0141	.0351	.9456	.0155	.0026	.0363	.2000	.7500	.0000	.0500	.0000	.0000	1.0000	.0000	.6000	.0667	.0000	.3333
Illinois	.5024	.4022	.0191	.0763	.8228	.1297	.0095	.0380	.1621	.7945	.0040	.0395	.0833	.0000	.9167	.0000	.3125	.2500	.0208	.4167
Indiana	.7475	.1667	.0126	.0732	.8953	.0574	.0000	.0473	.0606	.8030	.0000	.1364	.0000	.0000	.8000	.2000	.4138	.1034	.0690	.4138
Maryland	.6000	.2449	.0735	.0816	.8639	.1088	.0000	.0272	.2667	.6500	.0167	.0667	.0556	.0000	.9444	.0000	.3000	.1500	.0000	.5500
Massachusetts	.3019	.5597	.0660	.0723	.6667	.2500	.0104	.0729	.1292	.8539	.0000	.0169	.0000	.0000	.9524	.0476	.0870	.2174	.0435	.6522
Michigan	.6439	.2846	.0032	.0682	.8437	.0992	.0017	.0555	.1369	.8137	.0000	.0494	.3333	.3333	.3333	.0000	.3651	.1746	.0000	.4603
Minnesota	.6773	.2665	.0073	.0489	.9206	.0722	.0000	.0072	.2385	.7339	.0000	.0275	.0000	.0000	1.0000	.0000	.5000	.1500	.0000	.3500
Missouri	.7280	.2040	.0126	.0554	.9273	.0277	.0035	.0415	.1728	.7531	.0000	.0741	.0000	.2000	.8000	.0000	.5455	.1364	.0000	.3182
New Jersey	.4022	.4642	.0652	.0684	.7708	.1542	.0079	.0672	.1130	.8288	.0068	.0514	.1220	.0732	.8049	.0000	.1395	.1628	.0465	.6512
New York	.3724	.4404	.0926	.0946	.7619	.1772	.0053	.0556	.1320	.7964	.0112	.0604	.0106	.0532	.8936	.0426	.2500	.1875	.0208	.5417
North Carolina	.9447	.0267	.0018	.0267	.9736	.0057	.0038	.0170	.2000	.7333	.0000	.0667	.0000	.0000	1.0000	.0000	.4000	.0000	.0667	.5333
Ohio	.7132	.2154	.0066	.0648	.8906	.0616	.0000	.0478	.1735	.7959	.0000	.0306	.0000	.1667	.8333	.0000	.2712	.1695	.0169	.5424
Oregon	.6816	.1076	.0135	.1973	.8224	.0789	.0000	.0987	.2083	.7083	.0000	.0833	.0000	.6667	.3333	.0000	.3182	.1136	.0000	.5682
Pennsylvania	.5939	.3308	.0164	.0590	.8750	.0901	.0055	.0294	.1716	.8020	.0033	.0231	.0667	.0667	.8000	.0667	.3889	.1481	.0185	.4444
South Carolina	.8735	.0830	.0158	.0277	.9412	.0317	.0000	.0271	.3333	.6667	.0000	.0000	.0000	.0000	1.0000	.0000	.4286	.2857	.0000	.2857
Tennessee	.9135	.0404	.0058	.0404	.9579	.0105	.0000	.0316	.2857	.6667	.0000	.0476	.0000	.0000	1.0000	.0000	.6190	.0000	.0000	.3810
Texas	.6659	.2887	.0088	.0365	.9020	.0681	.0000	.0299	.1609	.8199	.0038	.0153	.0000	.0000	1.0000	.0000	.4545	.1515	.0303	.3636
Virginia	.8467	.0766	.0071	.0695	.9284	.0295	.0021	.0400	.3721	.5349	.0000	.0930	.0000	.0000	.7500	.2500	.3077	.1026	.0000	.5897
Wisconsin	.5562	.3848	.0076	.0514	.9075	.0616	.0000	.0308	.1287	.8366	.0000	.0347	.2500	.0000	.5000	.2500	.0741	.3333	.0000	.5926

Note.—Each column Pi-ik reports the sample probability that an individual of religion i marries an individual of religion k, by state. P = Protestants, C = Catholics, J = Jews, and O = Others.

The data on socialization come from a special module of the GSS, collected in 1988, that asks respondents to report the religious identity of their parents. Unfortunately, there are only 1,232 useful observations in this module. Therefore, it is possible only to estimate the empirical frequencies of the offspring's religious choices given the parents' identities for the whole United States and not for individual states.[35]

The structural parameters $\boldsymbol{\theta}$ are estimated via minimum distance by matching the vector $\hat{\boldsymbol{\Pi}}$ of empirical moments $(\hat{\pi}_s^{ij}, \hat{P}_{ij}^k)$ from the data with the vector $\tilde{\boldsymbol{\Pi}}(\boldsymbol{\theta})$ of moments $(\tilde{\pi}_s^{ij}(\boldsymbol{\theta}), \tilde{P}_{ij}^k(\boldsymbol{\theta}))$ implied by the model for a given choice of $\boldsymbol{\theta}$ (see theorem 1 in App. A).

Formally, given a square weighting matrix $\boldsymbol{\Omega}_N^*$ (where N denotes the total sample size), the minimum distance estimate $\hat{\boldsymbol{\theta}}$ minimizes

$$J_N(\boldsymbol{\theta}) \equiv [\hat{\boldsymbol{\Pi}} - \tilde{\boldsymbol{\Pi}}(\boldsymbol{\theta})]^\top \boldsymbol{\Omega}_N^* [\hat{\boldsymbol{\Pi}} - \tilde{\boldsymbol{\Pi}}(\boldsymbol{\theta})]. \tag{B1}$$

Each empirical probability $\hat{\pi}_s^{ij}$ is estimated by computing the empirical frequency of marriage ij in state s:

$$\hat{\pi}_s^{ij} \equiv \frac{\#(i \text{ married } j)_s}{\#(i)_s} = \frac{1}{B_s q_s^i} \sum_{b=1}^{B_s} X_b^{ij},$$

where the subscript b denotes an individual observation, B_s is the number of observations in state s, and X_b^{ij} is a dummy variable that is equal to one if individual respondent b is of religion i and is married to a j person, and zero otherwise.[36]

Each empirical probability $\hat{P}_{ij}^k$ is estimated by computing the empirical frequency of socialization to group k of children in families of type ij, aggregating over all states:

$$\hat{P}_{ij}^k \equiv \frac{\#(k \text{ children})}{\#(ij \text{ parents})}.$$

Thus the only source of randomness in the estimation is the sampling error present in each sample moment $\hat{\boldsymbol{\Pi}}$.

We do not use all the available moment conditions in the estimation. In particular, we match the moments π_s^{ij} only for ij = PP, PC, PJ, CC, CJ, and JJ. The reason we omit the residual moment conditions is that they are linearly dependent on the others. By definition of the probabilities π_s^{ij} and q_s^i, the following linear restrictions hold in the population for each state s:

$$\begin{aligned} \pi_s^{ij} q_s^i &= \pi_s^{ji} q_s^j \quad \forall i \neq j, \\ \sum_j \pi_s^{ij} &= 1 \quad \forall i. \end{aligned} \tag{B2}$$

In the estimation, we can therefore omit any 10 of the 16 available marriage moment conditions for each state (given four religious groups) since they are linearly dependent. A subset of these restrictions (eq. [B2]) do not hold exactly

[35] Moreover, since the individual respondents are the sampling units, the distribution by religion of the parents of the respondents we observe is not representative of the distribution of the population of the parents. While this is obviously problematic in principle, various attempts at correcting the distortion have not resulted in significantly different estimates.

[36] The unit of observation in the data is an individual respondent, not an individual marriage.

in the data, though, because of sampling error. Then the choice of which moment conditions to omit makes a difference in the estimation. We have omitted the moment conditions for the groups for which sampling error is likely to be more prevalent because of small sample size, that is, Jews and Others. Also, for any possible couple ij, only three socialization moments are considered since $\sum_k P_{ij}^k = 1$.

The optimal weighting matrix $\mathbf{\Omega}_N^*$ used in the minimum distance criterion (B1) is $\hat{\mathbf{\Omega}}_N^* = \hat{\mathbf{S}}^{-1}$, where $\mathbf{S}$ is the covariance matrix of the vector of moments $[\hat{\mathbf{\Pi}} - \tilde{\mathbf{\Pi}}(\boldsymbol{\theta})]$. We assume that the individual $\mathbf{\Pi}$ moments that we do include in the estimation are uncorrelated across religions (i.e., that the sampling error associated with the estimation of each $\hat{\pi}_s^{ij}$ is uncorrelated with that of any $\hat{\pi}_s^{kj}$, $i \neq k$). On the other hand, $\hat{\pi}_s^{ij}$ and $\hat{\pi}_s^{ij'}$, $j \neq j'$, are negatively correlated according to a multinomial distribution. Therefore, only the within-religion $V(\hat{\pi}_s^{ij})$ and $\mathrm{Cov}\,(\hat{\pi}_s^{ij}, \hat{\pi}_s^{ij'})$ terms of $\mathbf{S}$ are nonzero and can be easily estimated using the properties of multinomial distributions.

The estimation procedure requires that the map $J_N(\boldsymbol{\theta})$ be locally smooth in a neighborhood of the estimated value of the parameters $\boldsymbol{\theta}$. The map $J_N(\boldsymbol{\theta})$ inherits the properties of $\tilde{\mathbf{\Pi}}(\boldsymbol{\theta})$, which is only upper-hemicontinuous (theorem 1), in general, because of the possibility of multiple equilibria. An investigation of the properties of $J_N(\boldsymbol{\theta})$ in a neighborhood of our point estimates reveals that in fact the map is locally smooth, even though the estimates are sometimes close to critical points of the marriage market equilibrium set that generate the displayed discontinuities in the minimum distance criterion. The properties of the estimator $\hat{\boldsymbol{\theta}}$ are then the standard ones of minimum distance estimators. In particular, $\hat{\boldsymbol{\theta}}$ is consistent[37] and asymptotically normal, with variance equal to $V \equiv \{\mathbf{D}\mathbf{S}^{-1}\mathbf{D}^{\top}\}^{-1}$, where

$$\mathbf{D}^{\top} \equiv p\lim \left\{ \frac{\partial \tilde{\mathbf{\Pi}}(\boldsymbol{\theta})}{\partial \boldsymbol{\theta}^{\top}} \bigg|_{\boldsymbol{\theta}=\boldsymbol{\theta}_0} \right\}$$

is a matrix of partial derivatives evaluated at the true value $\boldsymbol{\theta}_0$.[38]

The $\tilde{\mathbf{\Pi}}(\boldsymbol{\theta})$ moments implied by the model are computed as follows. For a given value of the parameters $\boldsymbol{\theta}$, a choice of ΔV^{ij}, $i \neq j$ and o, together with the religious shares q_s^i, pins down the socialization probabilities τ_s^i through equations (A7) and (13). Given these equations and V^{ii}, we can compute (W_s^{ii}, W_s^{ih}) for an individual of type i in state s. Conditional on a set of (W_s^{ii}, W_s^{ih}), n^{ij}, and q_s^i, equation (A8) defines a mapping $f: A \rightarrow \alpha$. Likewise, equations (A9) and (A10) define a mapping $g: \alpha \rightarrow A$. Therefore, we need to find a fixed point of the mapping $A = g \circ f(A)$. This, in turn, yields a choice of α, the restricted pool-matching probabilities. Then the equilibrium values of α and A determine the vector of theoretical moments $\tilde{\mathbf{\Pi}}(\boldsymbol{\theta})$ implied by the model, through equations (A1)–(A6).

Finally, the $J_N(\boldsymbol{\theta})$ criterion is minimized by using a simulated annealing algorithm. Simulated annealing performs a random search over the parameter space and accepts not only downhill moves but also uphill moves. The probability of accepting an upward move depends positively on a "temperature" parameter

[37] It is easy to show that local smoothness of $J_N(\boldsymbol{\theta})$ is sufficient to ensure consistency of our estimator $\hat{\boldsymbol{\theta}}$.

[38] In our context, the distribution of religious shares is a discrete-time process. However, we make use of data only at a single time t. Therefore, the problem of initial conditions in estimating discrete time–discrete data stochastic processes (see Heckman 1981) does not arise.

that decreases as the search progresses. At the beginning of the search, the algorithm is allowed to make large upward moves and thus searches over the whole parameter space. As the temperature drops, the algorithm concentrates on more promising regions, but the random nature of the search still allows it to escape local minima.[39]

This algorithm is explicitly designed to find a global minimum of functions that may present multiple local optima or discontinuities. Because of computational limitations, we cannot compute all possible equilibria for each $\boldsymbol{\theta}$. To identify the equilibrium selection jointly with the parameters, in the face of possibly multiple equilibria, we adapt the algorithm so that, in the course of the minimization of $J_N(\boldsymbol{\theta})$, for each candidate value $\bar{\boldsymbol{\theta}}$, it randomly picks several distinct starting values for the iteration that yields the equilibrium. This procedure effectively searches for several possible equilibria, computes the criterion J_N for each of these equilibria, and picks the lowest value as the value $J_N(\bar{\boldsymbol{\theta}})$ for that particular value of $\boldsymbol{\theta}$. As a robustness check, given our estimates $\hat{\boldsymbol{\theta}}$, we are able to compute all the possible multiple equilibria for the set of α and A and, consequently, for $\tilde{\boldsymbol{\Pi}}$. We then evaluate the minimum distance criterion $J_N(\boldsymbol{\theta})$ over all possible equilibria to check that our estimate $\hat{\boldsymbol{\theta}}$ is indeed a local minimizer of $J_N(\boldsymbol{\theta})$ over the entire equilibrium set.

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[39] For a more detailed description of the algorithm, see Goffe, Ferrier, and Rogers (1994) and Goffe (1996). We are very grateful to Bill Goffe for providing us with his MATLAB simulated annealing routines.

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