LOVE AND MONEY: A THEORETICAL AND EMPIRICAL ANALYSIS OF HOUSEHOLD SORTING AND INEQUALITY*

RAQUEL FERNÁNDEZ
NEZIH GUNER
JOHN KNOWLES

This paper examines the interactions between household formation, inequality, and per capita income. We develop a model in which agents decide to become skilled or unskilled and form households. We show that the equilibrium sorting of spouses by skill type (their correlation in skills) is an increasing function of the skill premium. In the absence of perfect capital markets, the economy can converge to different steady states, depending upon initial conditions. The degree of marital sorting and wage inequality is positively correlated across steady states and negatively correlated with per capita income. We use household surveys from 34 countries to construct several measures of the skill premium and of the degree of correlation of spouses' education (marital sorting). For all our measures, we find a positive and significant relationship between the two variables. We also find that sorting and per capita GDP are negatively correlated and that greater discrimination against women leads to more sorting, in line with the predictions of our model.

I. INTRODUCTION

With a few notable exceptions, the analysis of household formation has played a relatively minor role in macroeconomics. The vast majority of macroeconomic models tend to assume the existence of infinitely lived agents (with no offspring) or a dynastic formulation of a parent with children.[1] While this may be a useful simplification for understanding a large range of phenom-

* We thank Michael Kremer and Torsten Persson for stimulating discussions of our paper in the NBER Economic Fluctuations and Growth Research Meeting and in the CEPR Public Policy Conference, respectively. We thank Alberto Alesina and three anonymous referees for valuable suggestions. We also wish to thank Daron Acemoglu, Orazio Attanasio, Oriana Bandiera, Jere Behrman, Jess Benhabib, Alberto Bisin, Jason Cummins, Paolo Dudine, William Easterly, Luca Flabbi, Christopher Flinn, Mark Gertler, David Green, Sydney Ludvigson, Jonathan Portes, Neil Wallace, David Weil, and participants in numerous seminars for helpful comments. We thank Miguel Szekely and Alejandro Gaviria at the Inter-American Development Bank for assistance with the Latin-American data and the Luxembourg Income Study and the statistical agencies of the Latin American countries in our sample for access to the surveys. The first and third authors acknowledge financial support from their respective National Science Foundation grants, and the first author from the CV Starr Center at New York University as well.

1. Even Becker and Tomes' [1979, 1986] pioneering work on intergenerational transmission of inequality assumes a one-parent household.

The Quarterly Journal of Economics, February 2005

ena, it can also lead to the neglect of potentially important interactions between the family and the macroeconomy. This is especially likely to be the case in those areas in which intergenerational transmission plays a critical role, as in human capital accumulation, income distribution, and growth.

The objective of this paper is to examine how household formation ("marriage"), inequality, and per capita output may interact in an economy. The main idea that we wish to explore, theoretically and empirically, is the potentially reinforcing relationship between the strength of assortative matching and the degree of inequality. In particular, we want to examine the thesis that greater inequality may tend to make matches between different classes of individuals less likely, as the cost of "marrying down" increases. In an economy in which borrowing constraints can limit the ability of individuals to acquire optimal levels of education, this private decision of whom to marry may have important social consequences. In particular, it can lead to inefficiently low aggregate levels of human capital accumulation and thus higher wage inequality and lower per capita income. Thus, inequality and marital sorting are two endogenously determined variables that potentially reinforce one another.

To explore the ideas sketched above, we develop a model in which individuals obtain utility both from household consumption and from the match-specific quality of their partner. Individuals are either skilled or unskilled, according to education decisions made when young. Once education is completed, there is a matching process in which individuals have opportunities to form households. When a skilled individual meets an unskilled individual with high match quality (love), there will be a trade-off between forming a household with relatively lower consumption but high match quality and continuing the search for a match. Once a household is formed, individuals have children. The latter, in turn, must again decide whether to invest in skills or not. Becoming a skilled (or, equivalently, educated) agent is costly. To finance education, young individuals need to borrow in an imperfect capital market in which parental income plays the role of collateral. Thus, parental income as well as the net return to being a skilled versus an unskilled worker, including the expected utility from one's future match, determine the proportion of children who on aggregate become skilled.

We show that the steady state to which this economy converges will in general depend upon initial conditions. In particular, it is possible to have steady states with a high degree of sorting (skilled agents form households predominantly with others who are skilled; unskilled form households predominantly with unskilled), high inequality, and low per capita income. Alternatively, there can be steady states with a low degree of sorting, low inequality, and high per capita income. We also extend the model to incorporate gender discrimination and examine an alternative search model that generates multiple equilibria. We show that it is likely that societies with more gender discrimination will also have greater sorting, as women will tend to marry more for money than for love.

Our empirical analysis examines the main implication of our model: a positive correlation between the skill premium and marital sorting. To do this, we assemble a total of 34 country household surveys from the Luxembourg Income Study (LIS) and the Inter-American Development Bank (IDB) and use them to construct a sample of households for each country. From these samples we construct several measures of the skill premium as well as a measure of marital sorting—the correlation of spouses' years of education. For all our measures of the skill premium, we find a positive and significant relationship with marital sorting, even after controlling for other possible sources for this correlation, such as urbanization and ethnic fractionalization. Furthermore, we show that marital sorting and per capita income are negatively correlated across countries. We also explore the effect of gender discrimination and find, as predicted by our model, that greater gender discrimination is associated with more sorting.

Our work can be seen as integrating two literatures. A first, rapidly growing, literature is concerned with the intergenerational transmission of inequality in models with borrowing constraints. These models, though, either assume a dynastic formulation (e.g., Becker and Tomes [1986], Loury [1981], Ljungqvist [1993], Galor and Zeira [1993], Fernández and Rogerson [1998], Bénabou [1996], Dahan and Tsiddon [1998], Durlauf [1995], Owen and Weil [1998], Knowles [1999], Kremer and Chen [1999], and De La Croix and Doepke [2003]), or consider a two-parent household in which the degree of sorting is exogenously specified (e.g., Kremer [1997] and Fernández and Rogerson [2001]). The last two papers are particularly relevant as they are concerned with whether an (exogenous) increase in marital sorting can lead to a quantitatively significant increase in inequality. In our model, on the other hand, sorting and inequality are endogenously determined.

The second literature focuses on the determinants of who matches with whom, but basically abstracts from the endogeneity of the income distribution in the economy. The seminal paper in this literature is Becker [1973] and more recent contributions include Cole, Mailath, and Postlewaite [1992] and Burdett and Coles [1997, 2001]. Hess [2004] is a recent paper that analyzes, both theoretically and empirically, how the desire for income insurance affects the duration of marriage.[2]

Our paper, therefore, can be seen as trying to integrate both literatures in a simple, analytical framework. Some recent work that also shares our concerns, but that is more focused on fertility, marriage, and divorce, are Aiyagari, Greenwood, Guner [2000], Greenwood, Guner, and Knowles [2003], and Regalia and Ríos-Rull [1999]. The models, not surprisingly, are more complicated and rely on computation to obtain solutions for particular parameter values.

Last, there is a small empirical literature that is related to our work. As reviewed by Lam [1988], the general finding in the literature is the existence of positive assortative matching across spouses. Mare [1991] documents the correlation between spouses' schooling in the United States since the 1930s. Using a large cross section of countries, Smits, Ultee, and Lammers [1998] find that the relation between marital sorting and some indicators for development (such as per capita energy consumption and the proportion of the labor force not in agriculture) has an inverted-U shape. Dahan and Gaviria [1999] report a positive relation between inequality and marital sorting for Latin American countries. Boulier and Rosenzweig [1984] document assortative matching with respect to schooling and sensitivity to marriage market variables using data from the Philippines. Finally, Gould and Paserman [2003] examine the effect of male inequality on female marriage rates. They show that higher male inequality in a city is associated with lower marriage rate of women, a finding that they interpret as resulting from women searching longer for mates in cities with higher wage inequality.

Our paper is organized as follows. The second section pre-

2. Laitner [1979], on the other hand, endogenizes bequests (but not labor earnings) and hence the income distribution. He assumes, however, that matches are randomly determined. See Bergstrom [1997] and Weiss [1997] for a survey of the literature on theories of the family and household formation.

sents a dynamic model of endogenous matching and inequality and explores different variants of the matching model. The third section examines the empirical evidence, considers alternative candidates for the basic results, and explores causality. The last section concludes.

II. The Model

We assume that the economy is populated by overlapping generations who live for three periods. At the beginning of the first period, young agents decide whether to become skilled or unskilled. Once educated, they meet in what we call a "household matching market." Here they find another agent with whom to form a household, observing both the agent's skill type (and hence are able to infer that agent's future income) and a match-specific quality. In the second period, the now adult agents have children, work, and pay their education debt (if any). In the last period, households consume their income net of debt repayment and act as monitors to ensure that their children repay their education debts.

II.A. A Simple Model of Marital Sorting

We start by describing the matching problem faced by agents. The choice of whom to match with is, of course, driven by many factors: tastes, environment (e.g., who interacts with whom and the distribution of characteristics of individuals), and the prospects for one's material and emotional well-being. The simple model we develop below allows all these factors to interact to produce a household. To focus on the main features of our analysis—the interplay between inequality and household formation—we simplify in various dimensions. In particular, we assume that the two adult agents in a household share a common joint utility function, and we abstract from any differences between women and men, either exogenous (e.g., childbearing costs), or cultural and institutional (e.g., the degree of wage discrimination or the expected role of woman in the home relative to the workplace).[3] We discuss the implications of allowing men and women to differ in subsection II.C.

An agent is assumed to derive utility from consumption in

3. For models that focus on intrafamily bargaining problems, see, for example, Bergstrom [1997] and Weiss [1997].

the last period of life and from the quality of her match. Assuming that an agent is in a household (we will ensure later that this is the case), the indirect utility function for a couple with match quality q and household income I is given by

$$V(I, q) = u(I) + q, \tag{1}$$

where u is a continuous and strictly increasing function of I.

Individuals are either skilled (s) or unskilled (u). If skilled, they earn wages w_s in the second period and, in the third period after repaying their education debt, have net income $\tilde{w}_s \equiv w_s - d$ to contribute to the household, where d is the (constant) monetary cost of becoming skilled. Unskilled agents earn w_u in the second period, and this is the income they contribute to the household in the third period.

Wages are the outcomes of competitive labor markets for both factors of production—skilled and unskilled workers. The economy produces a single aggregate consumption good with constant returns to scale technology using skilled and unskilled workers. Thus, given a decomposition of the labor force L into skilled or unskilled workers ($L = L_s + L_u$), and denoting by λ the proportion of skilled workers in the population, full employment and constant returns to scale imply that output is given by

$$F(L_s, L_u) = LF(\lambda, 1 - \lambda) = L_u F\left(\frac{\lambda}{1-\lambda}, 1\right) \equiv L_u f(k),$$

where $k \equiv \lambda/(1 - \lambda)$. Hence wages depend only on λ:

$$w_s = f'(k) \text{ and } w_u = f(k) - f'(k)k. \tag{2}$$

Note that $\tilde{w}_s$ is decreasing in λ, w_u is increasing in λ, and thus that the skill premium is a decreasing function of the proportion of skilled workers.

Households can be categorized by the skill types of its two partners. Let I_{ij} denote the net household income for a couple composed by skill types $i, j \in \{s,u\}$. Thus,

$$I_{ij} = \begin{cases} 2\tilde{w}_s, & \text{if } ij = ss \\ \tilde{w}_s + w_u, & \text{if } ij = su \\ 2w_u, & \text{if } ij = uu. \end{cases} \tag{3}$$

We assume that in the first period, once their education decisions have been made, agents have two opportunities to match and form a household. In the first round, all agents meet randomly and draw a random match-specific quality q. This match can be accepted by both agents resulting in a "marriage" or rejected by at least one of the agents whereupon both agents enter the second round of matching. In the second round, agents are matched only with their own skill group and draw a new random match quality. We assume that qualities are match-specific i.i.d. draws from the same continuous cumulative distribution function Q (with its pdf denoted by Q'), and with expected value μ and support $[0,\bar{q}]$. An agent who remains single obtains utility only from her own income. Thus, the assumption of $q \geq 0$ is sufficient to ensure that all agents form a two-agent household in the second round.[4]

To solve the matching problem faced by agents, note first that since a skilled agent's second-round option dominates that of an unskilled agent (given $\tilde{w}_s \geq w_u$—a necessary condition in order for any individual to choose to become a skilled worker), the skilled agent will determine whether a match between a skilled and an unskilled agent is accepted in the first round. In particular, a skilled agent who encounters an unskilled agent in the first round and draws a high q will face a trade-off between forming a lower-income household with a high quality match, or waiting an additional period and forming a higher-income household with another skilled agent but with an expected quality level equal to μ; i.e., this agent faces a trade-off between love and money.

Letting $V_{ij}(q) \equiv u(I_{ij}) + q$, note that a skilled agent is indifferent between accepting a first-round match with an unskilled agent and rejecting that match and proceeding to the second round if $V_{su}(q) = V_{ss}(\mu)$. Solving for the level of q at which this occurs, q^*, yields a threshold quality of

$$q^* = u(2\tilde{w}_s) - u(\tilde{w}_s + w_u) + \mu. \tag{4}$$

The intuition underlying (4) is clear. The expected quality differential in the matches, $q^* - \mu$, must compensate for the loss in utility in matching with a lower income agent; i.e., $u(2\tilde{w}_s) - u(\tilde{w}_s + w_u)$. Of course, the threshold quality for two agents of the same type to match in the first round is μ as this is the expected

4. Although unrealistic, this allows us to abstract from the issue of how inequality affects the decision to remain single, which is not the focus of the analysis here. This is for simplicity only, as is the assumption that match quality is drawn from a distribution that is independent of the types. In our comparative static analysis, we will assume that $\bar{q}$ is sufficiently large so that in equilibrium some matches occur between skilled and unskilled individuals.

value of next round's match quality and there is no difference in household income.

Now that we have solved for the threshold qualities at which different households form, given a distribution of individuals into skilled and unskilled at time t, λ_t, we can find the equilibrium distribution of household types. The equilibrium distribution of households of each type depends only on the probability of meeting in the first round and on q^*. Both of these are only a function of λ_t since this variable determines both household incomes and first-round matching probabilities.

Denoting by ρ_{ij} the proportion of households formed between agents of skill types i and j, i, $j \in \{s,u\}$ with $\rho_{su} = \rho_{us}$, the equilibrium distribution of households is given by

$$\rho_{ij}(\lambda_t) = \begin{cases} \lambda_t^2 + \lambda_t(1-\lambda_t)Q(q^*(\lambda_t)), & \text{if } ij = ss \\ 2\lambda_t(1-\lambda_t)(1-Q(q^*(\lambda_t))), & \text{if } ij = su \\ (1-\lambda_t)^2 + \lambda_t(1-\lambda_t)Q(q^*(\lambda_t)), & \text{if } ij = uu. \end{cases} \tag{5}$$

Each line in (5) is obtained by calculating the probability that in the first round types i and j meet, multiplying it by the proportion of those types of meetings that will be accepted and, if $i = j$, adding the matches that occur in the second round. For example, meetings between a skilled and an unskilled agent occur with probability $2\lambda_t(1 - \lambda_t)$ in the first round, and are accepted with probability $1 - Q(q^*)$, with no additional matches between these two types in the second round.

It is important to note that $Q(q^*)$ measures the degree of marital sorting in this economy. If individuals were not picky and simply matched with whomever they met in the first round, then q^* would equal zero, and ρ_{su} would equal the probability of a skilled and an unskilled individual meeting, i.e., $2\lambda_t(1 - \lambda_t)$. If individuals only cared about match quality and not about income, then q^* would equal μ. Last, if individuals cared only about income and not about match quality, then $Q(q^*)$ would equal one, and there would be no matches between skilled and unskilled agents.

As shown below, $Q(q^*)$ is the correlation coefficient between skill types of spouses across households. This is most easily seen by setting Q equal to zero or one in (5) and noting that in the first case the distribution of households would be the same as that generated by purely random matching (and hence would have a zero correlation), whereas in the second case there would be perfectly assortative matching (i.e., no mixed couples and a correlation equal to one).

LEMMA 1. $Q(q^*)$ is the correlation between skill types of spouses across households.

Proof of Lemma 1. Let X be the random variable associated with the skill distribution of "women" (i.e., $x = 1$ if $x = s$, and $x = 0$ if $x = u$). Similarly, let Y be the random variable associated with the skill distribution of "men." Note that $\mu_x = \mu_y = \lambda$, $\sigma_x = \sigma_y = (\lambda(1 - \lambda))^{1/2}$, and that $E(XY) = \rho_{ss}$. Recalling that the correlation coefficient $\rho = (E(XY) - \mu_x\mu_y)/\sigma_x\sigma_y$ yields $\rho = Q(q^*)$.‖

The observation above will be very useful when we examine the data as although the proportions of matches of each type that form may have ambiguous comparative statics with respect to λ (as we discuss below), this is not true for the degree of sorting (i.e., for the correlation coefficient).

II.B. Sorting and Inequality

Key to our analysis is the effect of greater inequality on household formation. We start by analyzing how sorting is affected by an exogenous increase in inequality, e.g., technological change that increases $\tilde{w}_s$ relative to w_u. Taking the appropriate derivatives of q^* in (4) yields

$$\frac{\partial q^*}{\partial \tilde{w}_s} = 2u'(I_{ss}) - u'(I_{su}), \tag{6}$$

$$\frac{\partial q^*}{\partial w_u} = -u'(I_{su}) < 0,$$

and thus, whereas an increase in the unskilled wage unambiguously decreases q^* and therefore decreases sorting, the effect of an increase in the skilled wage depends on the degree of concavity of the utility function. Henceforth we assume that the latter is not too concave, i.e., that A1 below holds:

$$2u'(2x) \geq u'(x). \tag{A1}$$

Note that A1 is a sufficient condition for $\partial q^*/\partial \tilde{w}_s > 0$ and that it holds for a variety of utility functions such as linear, log, and more generally, for CRRA with $\sigma \leq 1$.

THEOREM 1. An increase in the skill premium, w_s/w_u, increases marital sorting.

Proof of Theorem 1. Recall that the degree of sorting is given by $Q(q^*)$. Noting that $I_{ss} \geq I_{su}$, it follows from A1 and (6) that q^*, and hence sorting, increases with the skill premium.‖

The theorem above implies that an exogenous increase in inequality increases sorting by making skilled workers less willing to form households with unskilled workers. We now turn to examining the effect of a change in λ.

Unlike an exogenous increase in the skill premium, a change in λ affects not only the return to being a skilled or an unskilled worker through its effect on wages, but also directly affects the probability of a skilled and an unskilled worker meeting by changing the distribution of agents. Thus, even if agents did not change their sorting behavior, a change in λ would alter the equilibrium distribution of households. The correlation coefficient is extremely useful as a measure of sorting since it takes into account that the underlying distribution of agents has changed. Consequently, although as we will see below a change in λ has potentially ambiguous consequences for changes in the equilibrium distribution of households, its implications for the degree of marital sorting are unambiguous.

COROLLARY 1. A decrease in λ increases marital sorting.

Proof of Corollary 1. To determine the effect of a change in λ on marital sorting, we need to determine only the sign of $dq^*/d\lambda$. Note that $d\tilde{w}_s/d\lambda = f''(dk/d\lambda) < 0$, $dw_u/d\lambda = -f''k(dk/d\lambda) > 0$, where $dk/d\lambda = 1/(1-\lambda)^2$. It follows immediately from Theorem 1 above that $dq^*/d\lambda < 0$ and hence that a decrease in λ increases marital sorting.‖

The intuition for the result above is that a decrease in λ increases the skill premium and thus makes skilled workers less willing to form matches with unskilled workers. One might ask, however, how does a change in proportion of skilled workers in the population affect the fraction of households of each type? An increase in λ will unambiguously decrease the fraction of couples who are uu as, for any given q^* they are less likely to end up in uu households. Furthermore, the effect on wages implies that q^* will decrease, thereby increasing the probability that a first round match between a high and low skilled worker results in a household. The effect on us and ss households, on the other hand, is ambiguous (although the aggregate fraction of the population that is in one of these two types of households must, of course, increase). For any given q^*, the fraction of ss households increases, but as a skilled individual is now more willing to match with an unskilled one, this will work to decrease the fraction of ss households. The effect on us households is positive if $\lambda \leq 1/2$ (as both the likelihood of s and u individuals meeting in the first round and the probability that the match will be accepted increase) and ambiguous otherwise. This ambiguity is not troubling, however, as the theory leads to unambiguous results with respect to sorting—the focus of our empirical analysis. It also indicates why studying correlations (rather than proportions of types) is important, as the former does not depend on the distribution of agents, except endogenously.

As the analysis above indicates, λ is the key variable in our model. It determines wages, sorting, and the equilibrium distribution of households. Before turning to an analysis of how λ is itself determined in the model, it is useful to clarify the role of some of the simplifying assumptions in our model and discuss how they contribute to our results.

A useful feature of our matching model is that, in any given period, λ only affects the degree of sorting through its effect on the skill premium (since that is what determines q^*). This is due to the way in which we modeled the rounds of search: only two rounds and the second round with one's own type. A more general model of matching would have the proportion of individuals of different types in the matching market evolve endogenously as a result of matches made in previous rounds. This would then produce a dependence of the correlation on the initial fraction of skilled individuals, independently of the latter's effect on the skill premium. In general, however, this type of search model with nontransferable utility also generates multiple equilibria.[5] In Appendix 1 we develop an alternative model of matching in which the degree of sorting is no longer independent of λ. Although this alternative model can give rise to multiple equilibria, for all locally stable equilibria this model nonetheless generates the same qualitative relationships between sorting, inequality, and λ as in our original simpler specification. Our results, therefore, do not require a very specific matching structure but rather one in which, in response to an increase in λ, the increased incentive for

5. See Burdett and Coles [1997].

a skilled individual to match with an unskilled one due to the now lower wage differential is stronger than the increased incentive to search longer for the now relatively more abundant skilled partner. Our simpler matching model highlights only the first force since it assumes that a skilled individual can always match with another one by rejecting a first-period match. Our alternative model allows both forces to work and obtains similar results, but at the cost of multiple equilibria. In our empirical work we examine the effect of both λ and of inequality on sorting.

Another key feature of our matching model is that it is essentially genderless; there are only skilled and unskilled agents (or alternatively, equal numbers of skilled males and females and equal numbers of unskilled males and females, and both sexes' market earnings depend only on their skill type). The cost of this simplification is that it does not allow us to analyze questions about gender discrimination. We next turn to a more complicated model that permits an exploration of some of these issues before returning to our simpler model for the dynamic analysis.

II.C. Sorting and Gender

Our model in essence has only one gender. This enormously simplifies the analysis since it allows us to deal with only two groups—skilled and unskilled—rather than the four that would result if we further differentiated our agents into females and males. An interesting question, therefore, about which our simple model is unable to shed any light is how discrimination against women may affect the trade-off between love and money, i.e., sorting.[6] Although doing full justice to this question would require us to significantly modify our model and the main focus of our analysis (earnings inequality and sorting) and thus deserves an independent paper, we think that it is nonetheless illuminating to attempt to examine how two factors—female wages and the ease with which women work outside the home—may affect sorting. Below we present a partial analysis.

Consider an economy with two genders and, as before, two skill levels. Agents are endowed with a unit of time. Men spend their entire unit of time working in the market whereas women decide how much time to devote to raising their children and how

6. We thank Alberto Alesina for suggesting that we explore the implications of discrimination for sorting.

much time to spend in the formal labor market. Time spent raising children contributes to the quality of the latter.

The household utility function is given by

$$U(w_{mi}, w_{fj}) = \ln (w_{mi} + w_{fj}t_j) + \beta \ln [(1 - t_j)\psi_j], \tag{7}$$

where $\beta < 1$, and $i, j \in \{s,u\}$ indicate the skill type of the agent, m and f denote male and female, respectively, t_j is the amount of time a woman of type j spends in the formal labor market, and $(1 - t_j)\psi_j$ is the quality of the child that results if a woman of type j spends $1 - t_j$ units of time raising the child.[7] Note that wages are indexed by sex as well as skill type.

The first-order condition for t_j is given by

$$\frac{w_{fj}}{w_{mi} + w_{fj}} - \beta \frac{1}{1 - t_j} + \gamma = 0, \tag{8}$$

with

$$\gamma \geq 0,\ t_j \geq 0,\ \gamma t_j = 0.$$

Hence,

$$t_j = \begin{cases} \dfrac{w_{fj} - \beta w_{mi}}{w_{fj}(1 + \beta)}, & \text{if } w_{fj} > \beta w_{mi} \\ 0, & \text{otherwise.} \end{cases} \tag{9}$$

Thus, the amount of time a woman spends working outside the home is increasing in her own wage and decreasing in the wage of her partner.

We assume that matching takes place over two rounds, both at random. Agents who decide to marry in the first round do not participate in the second. Hence, as before, a skilled agent who meets an unskilled agent in the first round and draws a high q will have to decide between accepting that match or going on to the next round where the expected quality of the matches is given by μ, but where there is also a positive probability that another skilled agent will be met and thus that household income will be greater.

7. This formulation assumes that the quality of the child depends only on the skill level of the female and the amount of time she devotes to raising it. A more general specification would allow both spouses to determine the quality of the child and allow men to also choose how much time to devote to child rearing. Assuming that men have a comparative advantage in market work then yields similar results (i.e., men work in the formal labor market, and women are either housewives or spend some time in each activity), but requires more algebra.

Now that we differentiate between men and women, it also makes sense to allow the perceived quality of a match to differ across both potential spouses, otherwise, if a man of type i and a woman of type j meet, the same gender-type would always be decisive in accepting the match for all meetings between these two gender-types. Hence the utility from a match is now given by $U(w_{mi}, w_{fj}) + q_{mi}$ for the man and by $U(w_{mi}, w_{fj}) + q_{fj}$ for the woman, where q_{mi} and q_{fj} are (perhaps correlated) draws from a distribution Q.

To examine some of the possible effects of discrimination against women, a first interesting exercise is to compare the relative "pickiness" of skilled women in two societies that share the same parameters but differ in the wages that they pay skilled women (e.g., wage discrimination).[8] Thus, consider the choice of a skilled female who has met an unskilled male in the first round and must decide whether to match with him or to proceed instead to the second round in which she faces a probability λ' of obtaining a match with a skilled male and a probability $1 - \lambda'$ of matching with an unskilled male. Assume that $w_{fs} > \beta w_s$, so that skilled women spend at least some of their time in the formal labor market.

The cutoff level of love required for a woman not to want to continue her search is given by equating her payoff from marrying this unskilled male ($U(w_{mu}, w_{fs}) + q^*_{fs}$) to the expected value of the match from the second period $\lambda'(U(w_{ms}, w_{fs}) + \mu) + (1 - \lambda')(U(w_{mu}, w_{fs}) + \mu)$, yielding

$$q^*_{fs} = \mu + \lambda'[U(w_{ms}, w_{fs}) - U(w_{mu}, w_{fs})] \tag{10}$$

$$= \mu + \lambda'\left[(1 + \beta)\ \ln\left(\frac{w_{ms} + w_{fs}}{w_{mu} + w_{fs}}\right)\right],$$

where we have made use of (9) to obtain the household utilities.

It is straightforward to show that $\partial q^*_{fs}/\partial w_{fs} < 0$, i.e., skilled women who live in a society in which, ceteris paribus, they earn higher wages will favor "love" over money by more than similar women who live in a society in which their work is compensated at a lower rate. The intuition for this is clear: in a society in which skilled women are less well compensated, they will value on the margin more the additional income of a higher paid spouse.

8. See Galor and Weil [1996] for a model in which exogenous differences between women and men lead to a large wage gap at low levels of capital, which is then reduced as capital accumulates.

Hence, ceteris paribus, we might expect to see more sorting in societies in which skilled women are paid less.[9]

We next compare the effect of participating actively in the formal labor market on women's marital choices. Women may work less in one society than in another because of low wages, discriminatory hiring, or social norms. To avoid unnecessary algebra, we examine two extreme societies: one in which females choose their work time freely (and $w_{fs} > \beta w_{ms}$) and one in which they do not work outside the home; all other parameters are assumed to have the same values in both societies. As before, we consider the choice faced by a skilled woman who has met an unskilled man in the first round and must decide whether to form a household or to proceed to the second round in which the probability in which she meets a skilled man is given by λ'. We will show that when women are free to work outside the house, ceteris paribus, they will favor love over money by more than women who are restricted to being housewives.

Letting a tilde denote the value of a variable when a woman does not work outside the home, after some algebra we obtain

$$q^*_{fs} - \tilde{q}^*_{fs} = \lambda' \left[(1 + \beta) \ln \left(\frac{w_{ms} + w_{fs}}{w_{mu} + w_{fs}} \right) - \ln \frac{w_{ms}}{w_{mu}} \right], \tag{11}$$

so that if women are less picky in a society in which they work, $q^*_{fs} - \tilde{q}^*_{fs}$ should be negative. This is shown below, establishing that skilled women who are able to work freely are more inclined to choose love over money than women who can only work at home.

PROPOSITION 1. $q^*_{fs} - \tilde{q}^*_{fs} < 0$.

Proof of Proposition 1. See Appendix 2.‖

The analysis above gives an indication that there may be greater marital sorting in societies in which there is more employment discrimination against women whether this occurs through wages, through a lower probability of finding a job, or through social norms that make it more difficult for women to be employed in the formal labor market. The analysis is incomplete, however, as we confined our derivation to comparing the choices of skilled women who have a given probability of meeting a skilled man in the second round. This probability, however, is

9. Note that while we are attributing the lower wages of women to discrimination, these could be instead the effect of different technologies.

itself endogenous and depends, as well, on how men perceive the trade-off between skilled and unskilled women in both societies. In Appendix 2 we show results from numerical simulations of this model that lead to a similar conclusion for a range of parameters. We now return to our simple genderless model to analyze the dynamic evolution of the economy.

II.D. Education Choice

We assume that households have n_{ij} children in their second period of life, and for simplicity, treat this as an exogenous and continuous variable with $n_{ss} \leq n_{su} \leq n_{uu}$.[10] Young agents make their education decisions in the first period. Becoming a skilled agent is costly; it requires an expenditure of d which young agents must borrow and later repay in the second period. The relative attractiveness of being a skilled worker depends both on net wages and on the expected return to matching. The expected utility from being a skilled worker given that a fraction λ_{t+1} of the population also becomes skilled is given by

$$(12)\quad V^s(\lambda_{t+1}) = \lambda_{t+1}\int_0^{\bar{q}} \max\,[V_{ss}(x;\,\lambda_{t+1}),\,V_{ss}(\mu;\,\lambda_{t+1})]\,dQ(x) + (1-\lambda_{t+1})\int_0^{\bar{q}} \max\,[V_{su}(x;\,\lambda_{t+1}),\,V_{ss}(\mu;\,\lambda_{t+1})]\,dQ(x),$$

whereas the expected utility of being an unskilled worker is

$$(13)\quad V^u(\lambda_{t+1}) = \lambda_{t+1}\left[\int_0^{q^*} V_{uu}(\mu;\,\lambda_{t+1})\,dQ(x) + \int_{q^*}^{\bar{q}} V_{su}(x;\,\lambda_{t+1})\,dQ(x)\right] + (1-\lambda_{t+1})\int_0^{\bar{q}} \max\,[V_{uu}(x;\,\lambda_{t+1}),\,V_{uu}(\mu;\,\lambda_{t+1})]\,dQ(x).$$

We assume that in addition to a monetary cost of d, becoming a skilled worker entails an additive nonpecuniary cost of

10. See the working paper version of this paper [Fernández, Guner, and Knowles 2001] for an analysis with endogenous fertility.

$\delta \in [0,\infty]$.[11] This cost can be thought of as effort, and it is assumed to be identically and independently distributed across all young agents with cumulative distribution function Φ. Thus, an agent with idiosyncratic cost δ_i will desire to become skilled if $V^s - V^u \geq \delta_i$.

We define by $\delta^*(\lambda)$ the skilled-unskilled payoff difference generated when a fraction λ of the population is skilled; i.e.,

$$\delta^*(\lambda_{t+1}) \equiv V^s(\lambda_{t+1}) - V^u(\lambda_{t+1}). \tag{14}$$

Note that given δ^*, all agents with $\delta_i \leq \delta^*$ would want to become skilled.

Let us first consider what would happen if all agents were able to borrow freely. Note that in this case, contingent on their value of δ_i, young agents would make the same decision irrespective of their parents' household type. Hence in equilibrium a fraction $\Phi(\delta^*)$ of each family would become skilled yielding $\lambda_{t+1} = \Phi(\delta^*)$. In order for this to be a rational expectations equilibrium, it must be the case that

$$\delta^*(\Phi(\delta^*)) \equiv V^s(\Phi(\delta^*)) - V^u(\Phi(\delta^*)). \tag{15}$$

However, if parental income is a factor that influences a child's access to capital markets (either in terms of the interest rate faced or in determining whether they are rationed in the amount they are able to borrow), then children of different household types may make different education decisions although they have the same δ_i. In this case, the fraction of children of different household types that become skilled will depend on the parental household income distribution, and thus on λ_t.[12]

In particular, we assume that children within a family with household income I can borrow on aggregate up to $Z(I)$, $Z' > 0$. One way to think about this constraint is that parents can act as monitoring devices for their children in an incentive-compatible fashion by putting their own income up for collateral (in period 2

11. This assumption simply ensures that not all unconstrained agents will want to become skilled, and thus allows the skill premium to vary with the severity of borrowing constraints.

12. It is important to note that this constraint should not be interpreted literally as the inability to borrow freely to, for example, attend college. It could also reflect parental inability to borrow against their children's future human capital so as to live in a neighborhood in which the quality of primary and secondary public education is high or to opt out of public education for a high-quality private education. It is the quality of this earlier education that then determines the probability of an individual attending college even if the latter is free.

of their lives). This ensures that the children will use the funds to become educated rather than for consumption. Hence, a family with income I and n children can at most afford to educate at a cost d per child a fraction $\Phi(\hat{\delta}(I,n))$ of their children, where $\Phi(\hat{\delta}(I,n))$ is implicitly defined by[13]

$$\frac{Z(I)}{n\Phi(\hat{\delta}(I,n))} = d. \tag{16}$$

Note that, as indicated in (16), children from families with low household income are hampered in their ability to become skilled because of the lower aggregate amount that can be borrowed.[14]

Thus, given λ_t (and hence family income), in equilibrium a proportion

$$\pi_{ij}(\lambda_t,\ \lambda_{t+1}) \equiv \min\left[\Phi(\delta^*(\lambda_{t+1})),\ \Phi(\hat{\delta}(I_{ij}(\lambda_t),\ n_{ij}))\right] \tag{17}$$

of each family type become skilled, yielding in aggregate

$$\lambda_{t+1} = \frac{\Sigma_{ij}\ \pi_{ij}(\lambda_t,\ \lambda_{t+1})n_{ij}\rho_{ij}(\lambda_t)}{\Sigma_{ij}\ n_{ij}\rho_{ij}(\lambda_t)}, \tag{18}$$

where the numerator (times L_t) gives the number of skilled individuals at time $t + 1$ and the denominator (times L_t) gives the population at time $t + 1$.

II.E. Equilibrium

An equilibrium is a sequence $\{\lambda_t\}_{t=1}^{\infty}$ such that (i) given λ_t, a skilled and unskilled wage pair $(w_s(\lambda_t),\ w_u(\lambda_t))$, a threshold match quality (between skilled and unskilled agents) $q^*(\lambda_t)$, and a distribution of households $\rho_{ij}(\lambda_t)$ are determined by (2), (4), and (5), respectively, for all $t \geq 1$; (ii) $\{\lambda_t\}_{t=1}^{\infty}$ satisfies (18) where the fraction of children from each family type becoming skilled $\pi_{ij}(\lambda_t,\ \lambda_{t+1})$ is determined by (17), and threshold nonpecuniary cost levels δ^* and $\hat{\delta}$ are given by (14), and (16), respectively.

Figure I depicts the equilibrium λ_{t+1} generated by a given λ_t. The upward-sloping line, $\delta = \Psi(\lambda_{t+1};\ \lambda_t)$, is derived in the following fashion. For a given λ_t, it gives the value of δ such that the proportion of young individuals who both have $\delta_i \leq \delta$ and are

13. We are implicitly normalizing the gross interest rate to equal one. Note that as we are not endogenizing the supply of funds for loans, it is best to think of loans being provided on a world market (in which this country is small).

14. If lower-income households have more children than higher-income households, this constraint is even more binding.

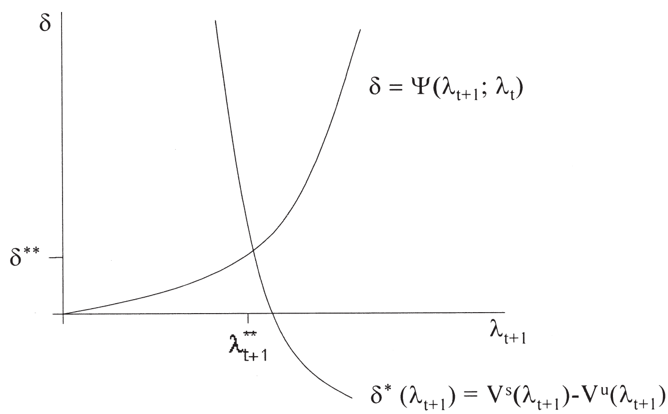

FIGURE I
Equilibrium Determination of λ_{t+1}

able to borrow to pay the cost of becoming skilled equals λ_{t+1}. As δ increases, so does the proportion of the population with $\delta_i \leq \delta$ and so λ_{t+1} must increase. Note that the domain of this function in general may be smaller than 1 since, if not all individuals are able to borrow, then λ_{t+1} will be below one even if $\delta \rightarrow \infty$. In the absence of borrowing constraints, $\Psi(\lambda_{t+1}; \lambda_t)$ does not depend on λ_t and its inverse equals $\Phi(\delta)$. Note that the unconstrained $\Psi(\cdot)$ curve is the lower envelope of the family of curves parameterized by different values of λ_t since for a given δ, the proportion of agents with $\delta_i \leq \delta$ that can afford to become skilled is highest in the absence of borrowing constraints.

The downward-sloping curve shows $\delta^*(\lambda_{t+1}) \equiv V^s(\lambda_{t+1}) - V^u(\lambda_{t+1})$ as a function of λ_{t+1}. Note that this curve does not depend on λ_t. The intersection of the two curves gives the equilibrium pair $(\delta^{**},\lambda^{**}_{t+1})$ for a given λ_t.

Existence of an interior equilibrium (for any initial λ_t) is guaranteed if we assume that $\tilde{w}_s(\lambda) < w_u(\lambda)$ for some $\lambda \in (0,1)$ (i.e., there exists λ such that for any λ greater than it, no one wishes to become skilled) and that for some other $\lambda \in (0,1)$ the inequality is reversed.[15] Note that the Ψ curve is continuous, upward sloping, starts at zero, and becomes vertical once all household types are constrained. Thus, this and the fact the

15. This is guaranteed, for example, by imposing Inada conditions on the production function.

$\delta^*(\lambda_{t+1})$ is a continuous function defined over the entire range of [0,1] and goes from strictly positive to strictly negative numbers, guarantees the existence of an interior equilibrium.

Uniqueness of equilibrium (for any given λ_t) is guaranteed if δ^* is monotonically decreasing in λ. This may not be the case, however, as an increase in λ also increases the probability of meeting a skilled agent. The resulting potential ambiguity is discussed in Appendix 3 in further detail. Hereafter we simply assume that[16]

$$\text{(A2)} \qquad \frac{d\delta^*(\lambda)}{d\lambda} < 0,$$

as this type of potential multiplicity is not the focus of our analysis.

II.F. Steady State

The state variable is the proportion of skilled agents in the economy, λ_t. The dynamic evolution of this variable is given by equation (18). A steady state is defined as a $\lambda_t = \lambda^*$ such that $\lambda_{t+1}(\lambda^*) = \lambda^*$. Note that if λ is constant, so are wages, and so is the cutoff quality for a skilled agent to match with an unskilled agent and the education decisions of children.

If the economy had perfect capital markets, then independently of the initial value of λ, the ability of individuals to borrow would imply that a proportion $\tilde{\lambda} = \Phi(\tilde{\delta})$ of them will choose to become skilled, i.e., $\pi_{ij} = \tilde{\lambda}$, $\forall ij$, $\forall \lambda_t$ such that $\delta^*(\tilde{\lambda}) = \tilde{\delta}$. Thus, the economy would converge immediately to the unique steady state.

In the absence of perfect capital markets, the initial distribution of individuals into skilled and unskilled determines the dynamic evolution of the economy. With borrowing constraints, for those household types who are constrained, a proportion smaller than $\Phi(\delta^*)$ will be able to become skilled. Aggregating over all family types, therefore, a proportion smaller than $\Phi(\delta^*)$ will become skilled next period. Obviously, the first family type to be constrained will be the *uu* type, followed by the *us* type, and last by the *ss* type, as lower family income implies both more binding borrowing constraints and a larger number of children who wish to borrow.

16. Simulation of the model for various functional forms and parameter values always resulted in a unique equilibrium.

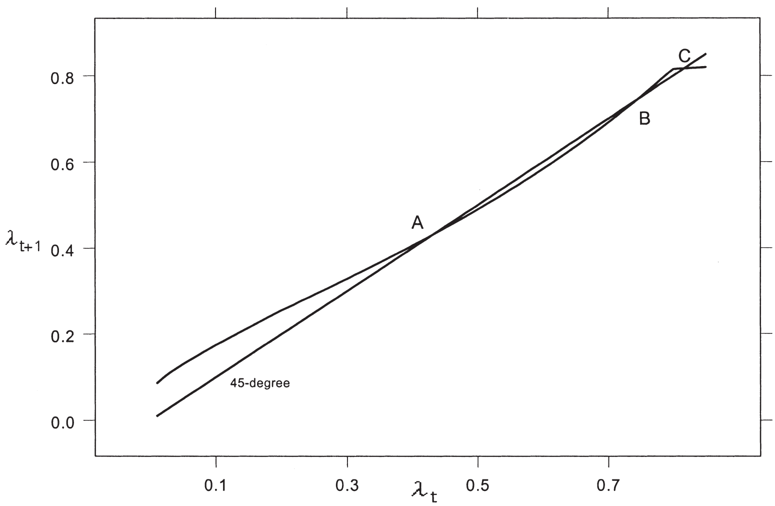

FIGURE II
An Example with Multiple Steady States

As shown in Figure II for a particular CES production function, this economy can easily give rise to multiple steady states, here given by the intersections of λ_{t+1} with the 45 degree line.[17] As depicted in the figure, the steady states A and C are locally stable.[18] The steady state in A is characterized by a low proportion of skilled individuals, high inequality between skilled and unskilled workers, and a high degree of sorting in household formation (i.e., skilled individuals predominantly marry other skilled ones; unskilled individuals predominantly marry other unskilled). In the steady state C, the opposite is the case: there is a large fraction of skilled individuals, low inequality, and low sorting.

Across steady states and indeed across any equilibrium at a point in time, higher inequality is associated with higher sorting.

17. The functional forms used to generate this figure are a production function given by $F(L_s,L_u) = (\alpha L_s^{\gamma} + (1 - \alpha)L_u^{\gamma})^{1/\gamma}$, and a limit on aggregate borrowing by children within a family of a fraction θ of household income; i.e., $Z(I) = \theta I$. Last, we assume that δ is distributed uniformly and that q is distributed with a triangular density function. The parameter values used are $\alpha = 0.25$, $\gamma = 0.5$, $\theta = 0.1$, $\bar{\delta} = 0.2$, $\bar{q} = 10$, $d = 0.1$, and $n_{ss} = 1$, $n_{su} = 2$, and $n_{uu} = 4$.

18. Note that the number of locally stable steady states can be greater than two since this depends on the change in the fraction of children of different family types that are constrained at different values of λ.

This follows simply from the static analysis in which we showed that greater wage differentials imply greater sorting (Theorem 1). We now turn to our empirical analysis in which we investigate whether this positive relationship between sorting and inequality holds.[19]

III. Empirical Analysis

Our model predicts that countries with higher skill premiums should have higher degrees of household sorting. This relationship should hold independently of whether countries have the same technology or whether they are converging to the same or different steady states. Indeed, this positive association follows from the static part of our theoretical analysis, in which greater inequality in the incomes of skilled relative to unskilled individuals causes the former to reject a higher proportion of potential matches with unskilled individuals.[20] Furthermore, across steady states, the relationship between sorting and inequality is mutually reinforcing. That is, higher degrees of household sorting should be associated with higher skill premiums and vice versa.

The purpose of this section is to investigate the relationship between marital sorting and the skill premium across countries. To our knowledge, this is the first paper that has attempted to study the interaction of these variables in a systematic fashion for a relatively heterogeneous set of countries.[21] We examine the main implications of our model using household surveys from 34 countries in various regions of the world. For each country, we assemble a sample of households with measures of the education and earnings of both spouses, and construct several measures of

19. One might also ask whether economies that start out with greater inequality necessarily end up in a steady state with at least as much inequality (and sorting) than an economy that starts out with lower inequality. We have shown this to be so for a large number of simulations, but it is potentially ambiguous since the proportion of households that are type *su* first increases and then decreases with λ which implies that λ_{t+1} may have a range in which it is a decreasing function of λ_t. This does not, however, affect the prediction which we will examine in the data: the existence of a positive correlation between sorting and the skill premium.

20. One can think of the initial source of variation in the skill premium as arising from different initial conditions or, more generally, from country-level shocks with no serial correlation across generations. Under these assumptions, the initial-conditions problem discussed in Heckman [1981], does not arise.

21. See Dahan and Gaviria [2001] for descriptive evidence on the positive correlation between marital sorting and inequality for Latin American countries.

the skill premium and a measure of the correlation of education between spouses (marital sorting).

We find a positive and significant relation between the skill premium and marital sorting across countries. We show that the relationship is robust to alternative specifications of the model, to alternative measures of the skill premium and sorting, and to various concerns regarding heterogeneity in the way the data are reported across countries. We examine other candidate explanations for this positive correlation and attempt to establish causality. We also investigate the effect of gender discrimination and find that, as in our model, lower discrimination is associated with less sorting. Our model also predicts that, for countries with similar technologies, marital sorting and GDP per capita should be negatively correlated, which we also show to be so in the data. Altogether, we take our findings as providing agreement of our basic hypotheses with the data.

III.A. The Sample

The data consist of twenty household surveys assembled from the Luxembourg Income Study (LIS) and thirteen Latin-American (LA) household surveys. In addition, we include the British Household Panel Study (1997).[22] These surveys were carried out by the governments of each country, and so a major issue for our analysis is ensuring that variables are comparable across countries. The LIS countries are largely European, but also include Australia, Canada, Israel, Taiwan, and the United States.[23] The years of these surveys range from 1990 to 1996, while LA surveys date from 1996–1997. We provide a more detailed discussion of these household surveys in Appendix 4, where we also list the names, coverage, and sample sizes of the surveys by country.

The way we select our sample varies somewhat with the variable we are measuring. To measure sorting, for each country we first construct a sample of couples where the husband is between 36 to 45 years old.[24] We do not restrict the definition of

22. We use the BHPS rather than the data from the LIS because the LIS reports education for Britain as the age at which an individual completes her education, a variable that is hard to map into years of schooling.

23. Education and earnings data for Russia are also available in the LIS, but the low quality of the data resulted in negative estimates of the skill premium. As a result we exclude it from our sample.

24. We restrict our attention to this age group for our measure of sorting, since younger cohorts presumably are less stable regarding their marriage pat-

a spouse to legally married couples, but for convenience we refer to them as “wives” and “husbands.” We include households in this sample of couples, if, in addition to various age requirements, there is a spouse present and education variables are available for both spouses.

To calculate skill premium measures, we construct a sample of husbands from a wider age group than 36 to 45, since presumably what individuals care about is some measure of the lifetime income of their spouses rather than earnings at a particular point of the life cycle (the age requirements are reported in the variable section that follows). We include husbands in the sample for our skill premium calculations, if, in addition to age requirements, there is a spouse present and husbands’ wages are nonmissing (including zeros). For both sorting and skill-premium calculations, we restrict the sample to single-family households and to those with male household heads. Observations are weighted using the household weights provided by the country survey.

II.B. Variables

We use labor income from all sources as our measure of the return to education. The exact definition of reported income differs by country. Some LIS countries report gross annual labor earnings, all forms of cash wage and salary income, and some report these net of taxes, which is closer in spirit to earnings in the model. The Latin American surveys report gross monthly labor income from all sources. The fact that some countries report gross income while others report net income could distort our cross-country comparisons, as net income will be more equally distributed than gross income in those countries with progressive taxation. We discuss our attempt to deal with this problem later on in the paper. Appendix 4 provides the details of our income measures.

We construct four measures of the skill (education) premium for each country. Three of the measures rely on a definition of a skilled individual. We define individuals as skilled if they have more years of education than those required to complete high

terns. Ideally we would want to examine a population for which we could observe both marital decisions and the expectations of lifetime wage inequality at the time of the marriage decision. The latter consideration argues for younger rather than older cohorts since presumably the observed wage inequality corresponds more closely to the expected one than is the case for older individuals.

school, creating a skill-indicator variable that equals one if an individual is skilled and equals zero otherwise. The fourth definition requires knowing an individual's years of education.

Education measures also differ across countries. Almost all LIS countries report education in terms of levels, rather than years, and often these levels are stated in terms of the nation's education system, so some standardization is required.[25] The Latin American data also required standardization because the number of years required for high-school completion varies across countries. For countries that report attainment together with years of schooling, our skill indicator equals one if some postsecondary education was reported for an individual. For countries that do not report attainment level, our skill-indicator equals one if the years of schooling exceeded the standard time required to complete high school in that country. The mapping of reported education measures into years of schooling and into an indicator for high school completion is summarized in Table XIV in Appendix 4.

Our four measures of the skill premium are defined as follows. The first is the ratio of earnings for skilled male workers to unskilled ones in our sorting sample, i.e., husbands between ages 36 and 45.[26] This measure is very simple and intuitive, and has a direct counterpart in our model. A potential drawback of using the wage ratio as described above is that it reflects income at a particular stage in the life cycle, and the mapping from this variable to lifetime income is likely to differ across skill groups. It also ignores information other than education that could also affect earnings, such as age or labor market experience. We control for such effects by constructing another measure of the skill premium; this is the coefficient on an indicator for being skilled (i.e., having at least some post-high-school education) in the following regression:

$$\log(e_i) = a_0 + a_1 I_i + a_2(age - s_i - 6) + a_3(age - s_i - 6)^2 + \varepsilon_i,$$

where e_i is earnings, I_i is an indicator for being skilled, s_i is years of schooling, and $(age - s_i - 6)$ is potential experience for individual i. This regression is estimated for each country by OLS

25. Furthermore, while most countries report the highest level of education undertaken, Italy reports only the highest level completed.

26. We focus primarily on the male skill premium as women's labor supply decision is more likely to depend on her spouse's earnings. This is discussed more at length farther on in the paper.

for all husbands aged 30–60 who have positive earnings rather than solely for those aged 36–45. Given that we have controlled for experience, this measure may be able to better capture potential lifetime labor earnings inequality than the simple ratio of earnings for our smaller sample.[27] We will refer to this measure as the *skill indicator* measure of inequality and to the previous one as the *wage ratio* measure of inequality. These two measures will differ as the skill indicator uses a larger sample, omits zero-earnings and controls for experience.

Although these two measures of the skill premium have clear counterparts in our model and hence are easy to interpret, both of these measures depend on our definition of being skilled. Since this definition, i.e., going beyond high school, can be considered rather arbitrary we also use a measure that does not depend on this threshold—the *Mincer coefficient.* The Mincer coefficient is the coefficient b_1 on years of schooling, s_i, in the following regression:

$$\log (e_i) = b_0 + b_1 s_i + b_2(age - s_i - 6) + b_3(age - s_i - 6)^2 + \varepsilon_i.$$

We estimate this regression for all husbands aged 30–60 in our samples, as we did with our skill indicator measure.

Finally, note that our analysis so far has been based on inequality in annual incomes. A better measure, were it available, would be that in expected lifetime incomes, as presumably that is what an individual cares about when making a trade-off between quality and income across matches. In the absence of panel data, we cannot observe lifetime labor incomes. We can, however, construct crude estimates based on the standard "synthetic cohort" method [Ghez and Becker 1975]. We create projections of lifetime income using observations on older cohorts to predict the future income of the young. Our simplest measure does this by dividing the life cycle into five-year intervals, from 25–30 up to 60–65, and then computing average labor income over five-year intervals for skilled and unskilled individuals separately. We take the present value of the predicted income profiles as the measure of lifetime labor income assuming an annual discount factor of 0.96.[28] The ratio of these lifetime-income mea-

27. How good this measure is of lifetime labor earnings inequality depends on how well the earnings of different cohorts at a point in time represents the life-cycle earnings of an individual (i.e., on the stability of the earnings profile).

28. We exclude higher ages because some of the age-country-skill cells are empty for particular countries.

sures for skilled relative to unskilled workers constitutes our fourth measure of the skill premium, which we call the *lifetime income* measure.[29]

Our measure of sorting is the Pearson correlation coefficient between husband's and wife's years of education across couples in our sample. Note that we use education rather than income (in our model the two are synonymous). We follow this strategy because in reality a female's labor force participation decision is often dependent on her spouse's earnings and social norms. Note that, as shown in our section on sorting and gender, as long as skilled women produce higher quality children than their unskilled counterparts, men will still want to marry more educated women.

Table I reports the measures of the skill premium and marital sorting for each country. The first column gives, for each country, the means and standard deviation for the proportion of skilled men (those with more than high school) in our 36–45 years old husbands sample. The sample mean across countries is around 25 percent with a standard deviation of 12.5 percent. The next four columns report means and standard deviations, by country, of the different skill premium measures. The sample mean across countries of both the wage ratio and of the lifetime income ratio is around 2 with a standard deviation close to 0.8 for the wage ratio and 1 for the lifetime income ratio. The same statistics are 0.63 with a standard deviation of 0.30 for the skill indicator measure and 0.09 with a standard deviation of 0.03 for the Mincer coefficient. The countries with the lowest skill premiums are Australia (wage ratio and the Mincer coefficient), Norway (wage ratio), and Denmark (Mincer coefficient), while Colombia and Brazil (wage ratio) and Paraguay (Mincer coefficient) have the highest skill premiums. As shown in Table II, these four

29. As a further robustness check, we also compute an analogous measure of lifetime income that controls for age variation within cohorts. We estimate the following equation on husbands aged between 30 and 60 years:

$$\log(y_{it}) = \beta_0 + \beta_1(age - s_i - 6) + \beta_2(age - s_i - 6)^2 + \varepsilon_i,$$

where y_{it} is earnings, s_i is years of schooling, and $(age - s_i - 6)$ is years of potential experience. We estimate this separately for each skill group. We then compute the present value of predicted earnings over this age range, discounted back to age 30 using an annual discount factor of 0.96 for each educational class, and as before, take the ratio of skilled to unskilled earnings as the measure of lifetime labor income inequality. This measure is highly correlated with the first measure (the correlation coefficient is about 0.94).

TABLE I
SUMMARY STATISTICS

				Skill premium		Marital sorting			Gender Inequality	
Country	Statistic	Proportion of skilled husbands (1)	Wage ratio (2)	Skill indicator (3)	Mincer coef. (4)	Lifetime ratio (5)	Pearson corr. coef. (6)	Raw wage gap (7)	Gender-related dev. index (8)	Female/male enrollment ratio (9)
Argentina	mean	0.218	2.297	0.746	0.072	2.394	0.522	0.890	0.839	1.106
	std.	(0.409)	(0.149)	(0.045)	(0.004)					
Australia	mean	0.303	1.224	0.301	0.043	1.461	0.322	0.850	0.938	1.045
	std.	(0.460)	(0.011)	(0.034)	(0.005)					
Belgium	mean	0.296	1.472	0.446	0.073	1.594	0.654		0.931	1.067
	std.	(0.457)	(0.060)	(0.027)	(0.004)					
Bolivia	mean	0.190	2.405	0.951	0.115	4.377	0.696	0.630	0.663	0.909
	std.	(0.395)	(1.057)	(0.044)	(0.004)					
Brazil	mean	0.107	3.908	1.373	0.162	4.682	0.710	0.670	0.770	1.043
	std.	(0.312)	(1.900)	(0.016)	(0.001)					
Britain	mean	0.468	1.603	0.291	0.071	1.369	0.388	0.760	0.928	1.133
	std.	(0.490)	(0.288)	(0.029)	(0.006)					
Canada	mean	0.597	1.368	0.322	0.089	1.426	0.499	0.750	0.934	1.055
	std.	(0.490)	(0.284)	(0.015)	(0.003)					
Chile	mean	0.208	3.799	1.199	0.145	3.798	0.703	1.010	0.821	0.877
	std.	(0.355)	(0.267)	(0.020)	(0.002)					
Colombia	mean	0.153	4.025	1.211	0.137	4.669	0.764	0.700	0.774	1.043
	std.	(0.356)	(11.176)	(0.022)	(0.002)					
Costa Rica	mean	0.156	2.472	0.846	0.097	2.257	0.643	0.910	0.824	1.015
	std.	(0.337)	(0.802)	(0.043)	(0.004)					
Czech Republic	mean	0.153	1.824	0.531	0.084	1.925	0.595	0.730	0.857	1.013
	std.	(0.360)	(0.303)	(0.014)	(0.002)					
Denmark	mean	0.268	1.341	0.359	0.049	1.383	0.530	0.870	0.928	1.074
	std.	(0.443)	(0.040)	(0.038)	(0.005)					

Ecuador	mean	0.300	2.069	0.733	0.087	2.236	0.749	0.660	0.716	0.973
	std.	(0.447)	(0.172)	(0.033)	(0.003)					
Finland	mean	0.297	1.579	0.477	0.063	1.575	0.402		0.928	1.091
	std.	(0.457)	(0.036)	(0.051)	(0.007)					
France	mean	0.219	1.797	0.574	0.084	1.815	0.585	0.820	0.923	1.033
	std.	(0.414)	(0.029)	(0.035)	(0.004)					
Germany	mean	0.278	1.489	0.379	0.077	1.471	0.520	0.700	0.924	0.979
	std.	(0.448)	(0.073)	(0.031)	(0.005)					
Hungary	mean	0.162	2.026	0.763	0.143	2.346	0.660	0.800	0.834	1.038
	std.	(0.369)	(0.025)	(0.071)	(0.012)					
Israel	mean	0.418	1.415	0.310	0.063	1.243	0.644	0.730	0.900	1.045
	std.	(0.493)	(0.032)	(0.031)	(0.004)					
Italy	mean	0.084	1.537	0.512	0.079	1.596	0.614	0.840	0.910	1.037
	std.	(0.278)	(0.326)	(0.033)	(0.003)					
Luxembourg	mean	0.187	1.593	0.518	0.087	1.800	0.532		0.920	1.028
	std.	(0.391)	(0.168)	(0.041)	(0.005)					
Mexico	mean	0.196	3.162	1.058	0.126	3.261	0.734	0.850	0.790	1.000
	std.	(0.379)	(0.472)	(0.036)	(0.003)					
Netherlands	mean	0.265	1.403	0.346	0.067	1.442	0.494	0.760	0.934	0.990
	std.	(0.442)	(0.068)	(0.029)	(0.005)					
Norway	mean	0.288	1.310	0.385	0.068	1.354	0.477	0.760	0.941	1.085
	std.	(0.453)	(0.060)	(0.043)	(0.008)					
Panama	mean	0.253	3.018	0.953	0.144	3.387	0.683	0.800	0.781	1.068
	std.	(0.412)	(0.472)	(0.041)	(0.004)					
Paraguay	mean	0.090	2.939	1.106	0.151	2.873	0.709		0.739	1.000
	std.	(0.286)	(1.148)	(0.079)	(0.006)					
Peru	mean	0.240	2.594	0.852	0.140	2.621	0.689	0.840	0.734	0.876
	std.	(0.423)	(0.555)	(0.061)	(0.006)					
Poland	mean	0.098	1.642	0.448	0.058	1.285	0.617	0.750	0.839	1.058
	std.	(0.297)	(0.143)	(0.014)	(0.002)					

(continued on next page)

TABLE I
(CONTINUED)

Country	Statistic	Proportion of skilled husbands (1)	Wage ratio (2)	Skill premium: Skill indicator (3)	Skill premium: Mincer coef. (4)	Marital sorting: Lifetime ratio (5)	Marital sorting: Pearson corr. coef. (6)	Marital sorting: Raw wage gap (7)	Gender Inequality: Gender-related dev. index (8)	Gender Inequality: Female/male enrollment ratio (9)
Slovakia	mean	0.151	1.510	0.426	0.055	1.601	0.600		0.834	1.028
	std.	(0.359)	(0.019)	(0.018)	(0.002)					
Spain	mean	0.213	1.644	0.528	0.064	1.772	0.689		0.912	1.056
	std.	(0.409)	(0.098)	(0.019)	(0.002)					
Sweden	mean	0.315	1.444	0.376	0.052	1.384	0.457	0.810	0.940	1.194
	std.	(0.464)	(0.135)	(0.035)	(0.004)					
Taiwan	mean	0.265	1.546	0.472	0.074	1.716	0.706	0.620		
	std.	(0.441)	(0.023)	(0.015)	(0.002)					
Uruguay	mean	0.337	1.675	0.379	0.101	2.010	0.630	0.740	0.935	1.078
	std.	(0.474)	(0.224)	(0.020)	(0.002)					
United States	mean	0.604	1.743	0.485	0.108	1.654	0.625	0.700	0.830	1.127
	std.	(0.489)	(0.221)	(0.013)	(0.002)					
Venezuela	mean	0.148	2.125	0.621	0.071	2.340	0.629	0.930	0.767	1.077
	std.	(0.348)	(0.364)	(0.031)	(0.003)					
Sample	mean	0.251	2.029	0.626	0.091	2.180	0.602	0.781	0.856	1.038
	std.	(0.125)	(0.782)	(0.303)	(0.034)	(0.987)	(0.110)	(0.094)	(0.080)	(0.066)

TABLE II
CROSS-COUNTRY CORRELATIONS

	Skill premium				Marital sorting Pearson corr. coef.
	Wage ratio	Skill indicator	Mincer coef.	Lifetime ratio	
Skill premium					
Wage ratio	1.000				
Skill indicator	0.966 (0.000)	1.000			
Mincer coef.	0.842 (0.000)	0.861 (0.000)	1.000		
Lifetime ratio	0.928 (0.000)	0.934 (0.000)	0.810 (0.000)	1.000	
Marital sorting					
Pearson corr. coef.	0.636 (0.000)	0.684 (0.000)	0.651 (0.000)	0.626 (0.000)	1.000

Significance levels are shown in parentheses.

measures of the skill premium are highly correlated with each other.

The sixth column in Table I reports the correlation measure of marital sorting. The correlation ranges from 0.322 for Australia to 0.764 for Colombia. On average, across countries the correlation between spouses' years of schooling is around 0.60 with a standard deviation of 0.11.

As a check, we can compare our calculation of the percentage of skilled men with similar calculations performed by others. Barro and Lee [2000] provide a data set on educational attainments for a large set of countries. The correlation across countries between the proportion of skilled males in our sample and proportion of men aged 25 and over with postsecondary education in Barro and Lee is 0.85.[30] If we do not restrict our sample to husbands, we also obtain 0.85 for the correlation. Another check comes from the OECD [2002] which reports the proportion of men between the ages of 35 and 44 with at least tertiary (postsecondary) educational attainments in 2001. Again the correlation across countries between the two samples is very high, 0.87 (and

30. Based on 31 countries, since the data for Czech Republic, Luxembourg, and Slovakia are not available in Barro and Lee [2000].

0.84 if we do not restrict our sample of men to husbands).[31] These results leave us confident that our measures of the male skilled population is consistent with available estimates that are commonly used in the literature.

We can also compare our skill premium estimates with those reported by other studies. Szekely and Hilgert [2000] estimate the Gini index for hourly wages for 15 to 65 year old males for a large set of countries. Although this is not the measure of inequality that we are using, the correlation between the Gini index for hourly wages and our skill premium measures are quite high: between .76 and .79 depending on our skill premium measure.[32] Another study is Bils and Klenow [2000], who provide a collection of Mincer coefficient estimates. The correlation between our Mincer coefficient estimates is around .6, which is quite high considering that they are using a large variety of sources and a wide range of years—from 1970 for Britain to 1990 for Denmark, Peru, and Spain.[33] Thus, our skill premium estimates are also in line with available estimates.

III.C. Results

This subsection reports the main results of our empirical analysis and conducts several robustness checks.

Note first from Table II that marital sorting is positively and significantly correlated with all our measures of the skill premium (between 0.6 and 0.7 in each case). Table III reports the results of our baseline regressions of marital sorting on the skill premium, for all measures of the latter. We find that the relationship is positive and significant at the 1 percent level for all measures of the skill premium (the standard errors of the OLS regression are based on the Eicker-White robust covariance matrix in order to correct for heteroskedasticity). While the coefficients on the skill premium vary with the definition used, it is interesting to note that they all imply that an increase in skill premium by one standard deviation is associated with an increase

31. Based on 20 countries: Australia, Belgium, Britain, Canada, Czech Republic, Denmark, Finland, France, German, Hungary, Italy, Luxembourg, Mexico, Netherlands, Norway, Poland, Slovakia, Spain, Sweden, and the United States.

32. Szekely and Hilgert's [2000] estimates are based on household surveys from 1994–1997. The correlations are based on 28 countries. The Gini index is not available for Belgium, Czech Republic, Denmark, Israel, Slovakia, and Spain.

33. Based on 26 countries, since estimates for Belgium, Czech Republic, Finland, France, Luxembourg, Norway, Slovakia, and Taiwan are not available in Bils and Klenow [2000].

TABLE III
MARITAL SORTING AND THE SKILL PREMIUM

Explanatory variable	Wage ratio (1)	Wage ratio (2)	Skill indicator (1)	Skill indicator (2)	Mincer coef. (1)	Mincer coef. (2)	Lifetime ratio (1)	Lifetime ratio (2)
	Dependent variable—marital sorting. Skill premium measure (married men)							
Constant	0.421	0.448	0.447	0.456	0.408	0.438	0.450	0.474
	(0.041)***	(0.043)***	(0.036)***	(0.040)***	(0.046)***	(0.048)***	(0.037)***	(0.036)***
Skill premium	0.089	0.067	0.248	0.220	2.127	1.559	0.070	0.050
	(0.015)***	(0.022)***	(0.042)***	(0.067)***	(0.389)***	(0.457)***	(0.012)***	(0.016)***
LA dummy		0.041		0.022		0.058		0.051
		(0.041)		(0.042)		(0.030)*		(0.040)
Number of obs.	34	34	34	34	34	34	34	34
Adjusted R^2	0.386	0.382	0.451	0.437	0.405	0.427	0.373	0.374
	Skill premium measure (all men)							
Constant	0.436	0.465	0.451	0.463	0.416	0.451	0.449	0.473
	(0.039)***	(0.041)***	(0.037)***	(0.040)***	(0.049)***	(0.050)***	(0.037)***	(0.036)***
Skill premium	0.080	0.057	0.240	0.203	1.985	1.322	0.069	0.050
	(0.013)***	(0.022)**	(0.042)***	(0.065)***	(0.397)***	(0.454)***	(0.012)***	(0.016)***
LA dummy		0.048		0.029		0.071		0.052
		(0.046)		(0.042)		(0.030)**		(0.040)
Number of obs.	34	34	34	34	34	34	34	34
Adjusted R^2	0.365	0.362	0.428	0.416	0.353	0.394	0.374	0.376

Standard errors reflect Eicker-White correction for heteroskedasticity.
* Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent

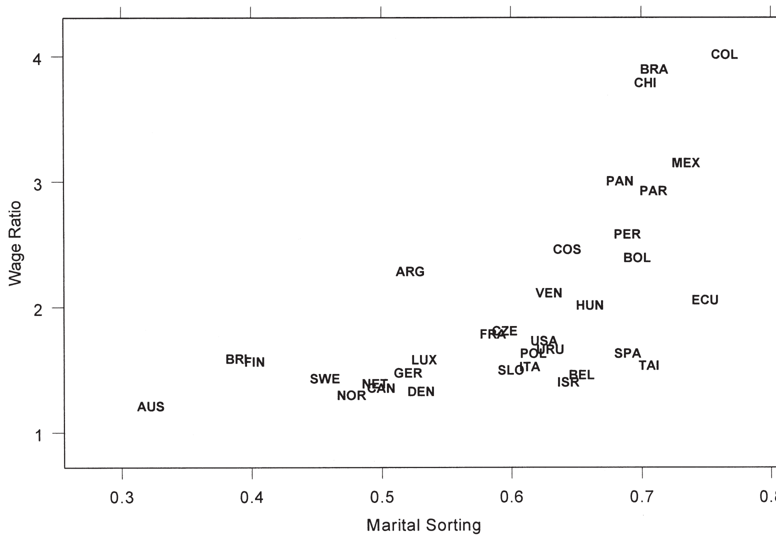

FIGURE III
Inequality (Wage Ratio) and Marital Sorting

in sorting of around .06 to .07, i.e., around 60 to 70 percent of one standard deviation of marital sorting. Thus, variation in the skill premium can be associated with a large fraction of the variation in marital sorting. The implied elasticity is between 0.25 and 0.32, depending on the skill premium concept used. We conclude that our baseline empirical test agrees with the basic prediction of our theory: a positive, statistically significant, economically strong relationship between the skill premium and marital sorting.

Figures III and IV show the data used in the regressions of Table III for the wage ratio and the Mincer coefficient measures of the skill premium (pictures with the skill indicator and lifetime income ratio measures are similar). As is clear from these figures, Latin American countries tend to have a greater degree of inequality and marital sorting than the rest of our sample. One possible interpretation of this finding is that Latin American countries are in a steady state with high inequality and high sorting, whereas the rest of our sample (predominantly European countries) are in a low inequality-low sorting steady state with the variation within these subsamples being explained by country-specific factors (e.g., labor-market institutions, education and

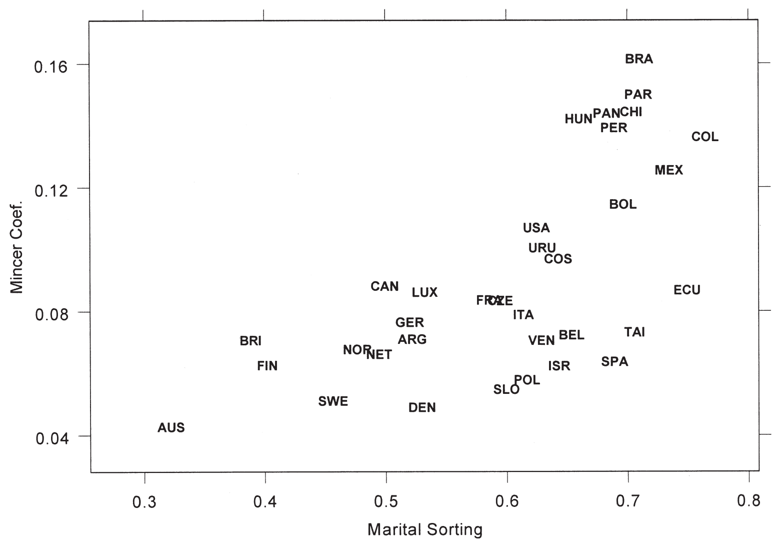

FIGURE IV
Inequality (Mincer Coefficient) and Marital Sorting

tax policy, credit markets, etc.). To make sure that our results are not driven by some factor that is common to Latin American countries, specification 2 in Table III introduces a Latin American dummy variable. The results from this specification show that the skill premium is still significant at the 1 percent level, suggesting that the correlation between sorting and inequality is not driven by some difference between the LA countries and the rest of the sample.[34]

To further explore whether there is something different about the Latin American countries that may also be responsible for our results, we reran our basic regression for our LA and LIS subsamples separately. Despite the significantly reduced sample sizes, the relationship between sorting and the skill premium remains positive and significant in each subsample in Table IV. It is interesting to note that the magnitude of the relationship doubles or triples for the LIS subsample for almost all specifications. The standard deviation of the skill premium is significantly smaller, however, for the LIS subsample. Consequently, a one-

34. We also checked for outliers that shifted the estimated coefficient on the skill premium by more than one standard deviation; there were no such outliers.

TABLE IV
MARITAL SORTING AND THE SKILL PREMIUM: SUBSAMPLES

	Dependent variable—marital sorting							
	LIS subsample Skill premium measure (married men)				LA subsample Skill premium measure (married men)			
Explanatory variable	Wage ratio (1)	Skill indicator (2)	Mincer coef. (1)	Lifetime ratio (2)	Wage ratio (1)	Skill indicator (2)	Mincer coef. (1)	Lifetime ratio (2)
Constant	0.137	0.324	0.412	0.276	0.553	0.558	0.530	0.571
	(0.133)	(0.082)***	(0.069)***	(0.114)**	(0.064)***	(0.054)***	(0.081)***	(0.059)***
Skill premium	0.268	0.520	1.902	0.175	0.046	0.133	1.270	0.035
	(0.082)***	(0.170)***	(0.738)**	(0.066)**	(0.019)**	(0.047)**	(0.591)*	(0.015)**
Number of obs.	21	21	21	21	13	13	13	13
Adjusted R^2	0.202	0.280	0.115	0.143	0.223	0.257	0.316	0.215
	Skill premium measure (all men)				Skill premium measure (all men)			
Constant	0.127	0.354	0.446	0.330	0.564	0.558	0.530	0.567
	(0.169)	(0.093)***	(0.071)***	(0.112)**	(0.070)***	(0.054)***	(0.083)***	(0.060)***
Skill premium	0.277	0.448	1.390	0.139	0.040	0.133	1.256	0.036
	(0.107)**	(0.185)**	(0.706)*	(0.065)**	(0.020)*	(0.047)**	(0.599)*	(0.015)**
Number of obs.	21	21	21	21	13	13	13	13
Adjusted R^2	0.166	0.195	0.049	0.103	0.176	0.254	0.306	0.240

Standard errors reflect Eicker-White correction for heteroskedasticity.
* Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent

standard-deviation increase in the skill premium is associated with an increase of between .4 and .5 of a standard deviation in marital sorting in the LIS sample and with an increase of between .5 and .6 of a standard deviation in marital sorting in the LA sample.[35]

We performed several other robustness checks. Since our model is silent about single men, we recalculated the skill premiums using all men (married and single), with the appropriate age restrictions, for each measure. Our results are very similar to the ones obtained for husbands, and they are reported in the second panel of Tables III and IV.

We also investigated whether the way in which variables are reported may significantly affect our results. As noted previously, some countries report earnings net of taxes and some report gross earnings. Since, due to progressive taxation, gross earnings will in general tend to be more unequal than the net ones, this can create differences in the measured skill premium and affect our results. In order to control for the way income is reported, we introduce a dummy variable that takes a value of one if the country reports net earnings and zero otherwise. As can be seen in Table V, the interaction of this dummy with the skill premium is positive and (with one exception) significant. Thus, an increase in the skill premium has a larger effect, not surprisingly, for those countries in which earnings are reported net of taxes since when individuals compare alternatives they should care about the after-tax household income of their potential partners rather than their gross income.[36]

How education is reported also varies across countries. As we noted previously, some LIS countries report years of education whereas others report only the highest formal level attained, such as high-school diploma or undergraduate degree. As a result, for some countries the years of education or skilled category includes only those who have completed college or the appropriate degree and excludes those who have not obtained the pertinent degree but may have progressed beyond high school. In order to check whether this feature of our data affects our results, we ran our

35. The standard deviation of sorting is 0.1 and .06, respectively, in the LIS and LA samples. The standard deviations of the inequality measures are given by .19 (.75) for wage ratio, .11 (.27) for skill indicator, .02 (.03) for Mincer, and .26 (.96) for lifetime ratio, with values for LA in parentheses.

36. We also reran our regressions using Szekely and Hilgert's estimate of the Gini index for hourly wages. This had a positive and significant effect on sorting, both with and without a LA dummy.

TABLE V
MARITAL SORTING AND THE SKILL PREMIUM: SOME ROBUSTNESS CHECKS

	Dependent variable—marital sorting							
	Skill premium measure							
	Wage ratio		Skill indicator		Mincer coef.		Lifetime ratio	
Explanatory variable	(1)	(2)	(1)	(2)	(1)	(2)	(1)	(2)
Constant	0.391	0.437	0.433	0.455	0.404	0.445	0.427	0.461
	(0.047)***	(0.041)***	(0.039)***	(0.035)***	(0.048)***	(0.047)***	(0.041)***	(0.035)***
Skill premium	0.097	0.055	0.255	0.166	2.124	1.172	0.075	0.039
	(0.017)***	(0.020)***	(0.043)***	(0.059)***	(0.394)***	(0.532)**	(0.013)***	(0.015)**
Net earnings dummy * skill premium	0.040	0.055	0.085	0.134	0.250	0.758	0.035	0.051
	(0.017)**	(0.020)**	(0.049)*	(0.064)**	(0.384)	(0.427)*	(0.017)**	(0.020)**
LA dummy		0.090		0.072		0.097		0.097
		(0.047)*		(0.049)		(0.043)**		(0.048)*
Number of obs.	34	34	34	34	34	34	34	34
Adjusted R^2	0.429	0.467	0.466	0.481	0.393	0.459	0.406	0.505
	Dependent variable—rank correlation measure of marital sorting							
Constant	0.395	0.423	0.427	0.438	0.394	0.430	0.430	0.454
	(0.046)***	(0.048)***	(0.041)***	(0.043)***	(0.053)***	(0.054)***	(0.041)***	(0.040)***
Skill premium	0.095	0.074	0.260	0.224	2.144	1.460	0.074	0.053
	(0.017)***	(0.024)***	(0.048)***	(0.070)***	(0.448)***	(0.519)***	(0.014)***	(0.018)***
LA dummy		0.044		0.027		0.070		0.054
		(0.045)		(0.043)		(0.034)**		(0.044)
Number of obs.	34	34	34	34	34	34	34	34
Adjusted R^2	0.369	0.361	0.413	0.399	0.342	0.370	0.348	0.348

	Dependent variable—marital sorting							
Constant	0.470	0.498	0.463	0.478	0.479	0.491	0.499	0.522
	(0.051)***	(0.059)***	(0.047)***	(0.053)***	(0.048)***	(0.052)***	(0.043)***	(0.046)***
Skill premium	0.081	0.059	0.239	0.205	1.938	1.532	0.063	0.043
	(0.015)***	(0.024)**	(0.044)***	(0.071)***	(0.380)***	(0.490)***	(0.011)***	(0.017)**
Proportion skilled	−0.130	−0.130	−0.044	−0.054	−0.213	−0.182	−0.134	0.051
	(0.116)	(0.123)	(0.103)	(0.111)	(0.098)**	(0.106)*	(0.116)	(0.122)
LA dummy		0.044		0.025		0.045		0.044
		(0.043)		(0.043)		(0.033)		(0.042)
Number of obs.	34	34	34	34	34	34	34	34
Adjusted R^2	0.386	0.381	0.435	0.422	0.445	0.450	0.373	0.374

Standard errors reflect Eicker-White correction for heteroskedasticity.
* Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

regressions including a dummy variable that takes the value of one for countries which report the finer classifications of education and zero otherwise. This variable was not significant, and it did not affect the results. We also attempted to correct for errors introduced by erroneous mapping of education attainment categories to years of education and for the use of high school as a perhaps arbitrary dividing line between skilled and unskilled by weighing each country's estimate of the skill indicator and Mincer coefficient by the inverse of the square of standard deviations of the estimate. This alternative also yielded similar results.

Last, a related concern is that although we studied each country's education system to understand how it progresses, the actual number of years of schooling that we assign to each attainment level may affect our measure of marital sorting. A possible check is to use the Spearman rank correlation between years of schooling of husbands and wives as an alternative measure of sorting as this is a purely ordinal measure. The Spearman rank correlation and Pearson correlation measures of sorting are highly correlated (0.98). Not surprisingly, we obtain similar results as in Tables III and IV as can be seen in the second panel of Table V.

Before investigating whether our results are driven by some third variable, it is useful to repeat our basic regressions but this time control as well for the proportion of the sample of husbands who are skilled, i.e., control for λ_t. Recall that, although in our basic model the effect of the skill premium on marital sorting is independent of the proportion of the population that is skilled, this would not be the case in a variety of alternative matching models (for example, the one developed in Appendix 1). The last panel of Table V reports the results from this regression. In all specifications, the effect of λ_t is negative, although significant only for the Mincer measure of the skill premium. In all cases, the effect of the skill premium on sorting is positive and significant, at the 1 percent level without a Latin American dummy, and at 1 or 5 percent with it.[37]

We conclude that our results hold for our entire sample of countries and within the LIS and LA subsamples separately, and furthermore they are robust to a series of measurement issues. We next turn to the issue of causality.

37. We also controlled for a quadratic relationship in λ_t with similar results.

III.D. Causality

In our model both the current skill premium and marital sorting are endogenously determined variables. Marital sorting ($Q(q^*)$) is determined by the current skill premium. The latter, on the other hand, is not determined by sorting. Rather, the current skill premium is a function of the proportion of individuals who decided to become skilled earlier (λ_t), i.e., of education decisions that precede any decision about whom to marry. Sorting instead affects next generation's skill premium (i.e., λ_{t+1}) since it, along with the skill premium, determines the proportion of next generation's population who will become skilled. In terms of our analysis, this means that the skill premium faced by our sample of 35–45 years-old married men is not simultaneously determined with their sorting patterns. Rather, the skill premium is the outcome of decisions made a couple of decades ago when this generation was making its education decisions.

The argument above indicates that causality runs from the expected skill premium to sorting. It could be argued, however, that technology shocks that affect the skill premium tend to be serially correlated. If technology shocks are neutral and the production function is constant returns to scale, however, then a measure such as the wage ratio will not be affected by these shocks. On the other hand, if shocks are not neutral, then this concern is valid since there will be a bias in the coefficient estimate resulting from the correlation of the explanatory variable with the error term. To correct for this endogeneity bias, we would like to find a variable that is highly correlated with the explanatory variable but not with the error term in the regression equation.

Ex ante, an excellent candidate as an instrument for the skill premium is the amount of capital per worker, since presumably this variable is positively correlated with the skill premium (if capital and skilled labor are more complementary than capital and unskilled labor), and there is no reason to believe that it would have an independent effect on the degree of marital sorting. We attempted to use both capital per worker from the Penn World Tables developed by Heston, Summers, and Aten [2002], and investment in machinery as a percentage of GDP from Flug and Hercowitz [2000] to instrument for the skill premium. In both cases, the measures were unable to capture sufficient variation across countries beyond those between countries from Latin

America and the LIS sample; i.e., they did not survive the inclusion of a Latin American dummy.

As an alternative, we also explored using labor union strength as an instrument, since the latter could affect the skill premium but should not have any direct effect on marital sorting. We used total trade union membership as a percentage of the total labor force (union density) and workers covered by collective bargaining as a percentage of total salaried workers as two different proxies for union strength (both of these measures are from Rama and Artecona [2000]). As an instrument, however, these measures had the same problem as capital per worker and investment in machinery. While the strength of labor unions for the 1990–1994 period is negatively and significantly correlated with our skill premium measures, the significant relation disappears once we control for Latin America.[38] Consequently, we decided to use data over a longer period of time (average union density from 1950–1994) since, as suggested by Oskarsson [2002], this may be a better indicator of labor union strength. This longer series (again from Rama and Artecona), has the disadvantage, however, of being available only for a small number of OECD countries. The two-stage-least square (2SLS) estimation results for these countries, together with OLS results, are shown in Table VI.[39]

As expected, Table VI shows that union density is negatively correlated with the skill premium. With the 2SLS estimation procedure, the effect of the skill premium on sorting is smaller for the Mincer coefficient (2.82 versus 3.59), but it is larger for other two measures (0.43 versus 0.27 for the wage ratio measure and 0.43 versus 0.41 for the lifetime income measure), indicating that endogeneity is not necessarily a problem that creates an upward bias in our estimates. These results show that the skill premium has a large and positive effect on marital sorting. The estimate of 0.43 for the wage ratio, for example, implies that going from being a more equal country like Germany to a more unequal country like the United States leads to a 0.10 points increase in marital sorting.

Another possible instrument for the skill premium is the

38. The union strength measures in Rama and Artecona [2000] are reported by five-year intervals.

39. The results in Table VI are based on 13 countries: Australia, Belgium, Britain, Canada, Denmark, Finland, France, Germany, Italy, Netherlands, Norway, Sweden, and the United States. The mean value of union density in the sample is about 39 percent with a standard deviation of about 16 percent.

TABLE VI
IV REGRESSION OF MARITAL SORTING ON SKILL PREMIUM: UNION DENSITY

	Dependent variable—marital sorting Skill premium measure					
	Wage ratio		Mincer coef.		Lifetime ratio	
Explanatory variable	Second stage (1)	OLS (2)	Second stage (1)	OLS (2)	Second stage (1)	OLS (2)
Constant	−0.128	0.105	0.305	0.250	−0.143	−0.103
	(0.323)	(0.196)	(0.091)***	(0.086)**	(0.294)	(0.197)
Skill premium	0.426	0.269	2.822	3.591	0.431	0.405
	(0.209)*	(0.132)*	(1.131)**	(1.083)***	(0.190)**	(0.133)**
	First stage		First stage		First stage	
Constant	1.689		0.102		1.703	
	(0.111)***		(0.009)***		(0.083)***	
Total union density	−0.005		−0.001		−0.005	
	(0.003)*		(0.0002)***		(0.002)**	
Number of obs.	13	13	13	13	13	13
Adjusted R^2	0.197	0.133	0.508	0.346	0.325	0.251

Standard errors reflect Eicker-White correction for heteroskedasticity.
* Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

average years of schooling in a country, since one might expect countries with a higher skilled share of the population (and hence a lower skill premium according to our model) to have higher average years of schooling. Furthermore, this is a variable that evolves slowly over time. According to our model, this variable would not have an effect on sorting other than through its effect on the skill premium.

Table VII shows our 2SLS estimates of the correlation between marital sorting and three skill premium measures (the wage ratio, the skill indicator measure, and the lifetime income ratio).[40] We use the measures of average years of schooling for the population aged 25 and above given in Barro and Lee [2000]. In the first-stage regressions, with or without an additional Latin American dummy, average years of schooling is negatively and significantly correlated with the skill premium.

40. The results in Table VII are based on 33 countries since the average years of schooling is missing for Luxembourg. Years of schooling does not have a significant effect on the Mincer coefficient estimates of the skill premium in the first stage of 2SLS.

TABLE VII
IV REGRESSION OF MARITAL SORTING ON SKILL PREMIUM: AVERAGE YEARS OF SCHOOLING

Explanatory variable	Dependent variable—marital sorting											
	Wage ratio				Skill indicator				Lifetime ratio			
	Second stage		OLS		Second stage		OLS		Second stage		OLS	
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
Constant	0.332	0.242	0.423	0.449	0.406	0.376	0.449	0.457	0.377	0.320	0.453	0.475
	(0.060)***	(0.180)	(0.043)***	(0.044)***	(0.046)***	(0.095)***	(0.037)***	(0.040)**	(0.055)***	(0.114)***	(0.038)***	(0.036)***
Skill premium	0.133	0.202	0.089	0.068	0.315	0.406	0.247	0.222	0.104	0.149	0.069	0.050
	(0.029)***	(0.116)*	(0.016)***	(0.022)***	(0.064)***	(0.211)*	(0.043)***	(0.068)***	(0.024)***	(0.070)**	(0.012)***	(0.016)***
LA dummy		−0.126		0.043		−0.071		0.019		−0.106		0.048
		(0.155)		(0.042)		(0.114)		(0.043)		(0.108)		(0.041)
	First stage				First stage				First stage			
Constant	4.114	2.715			1.503	1.017			4.851	3.158		
	(0.390)***	(0.428)***			(0.137)***	(0.152)***			(0.487)***	(0.544)***		
Average years of schooling	−0.266	−0.133			−0.112	−0.066			−0.342	−0.181		
	(0.049)***	(0.047)***			(0.017)***	(0.017)***			(0.060)***	(0.060)***		
LA dummy		0.922				0.320				1.115		
		(0.200)***				(0.071)***				(0.253)***		
Number of obs.	33	33	33	33	33	33	33	33	33	33	33	33
Adjusted R^2	0.477	0.685	0.380	0.374	0.569	0.736	0.449	0.433	0.492	0.681	0.369	0.368

Standard errors reflect Eicker-White correction for heteroskedasticity.
* Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

In the second-stage regressions, the effect of the skill premium on marital sorting is positive and highly significant in the specification without an LA dummy, though the significance decreases once the LA dummy is introduced (though the latter is not significant in these regressions).[41] The estimate of 0.315 for the skill indicator implies that going from a more equal country like Sweden, for example, to a more unequal one like Chile results in an increase in the correlation coefficient for spouses' education of about 0.26 points. The IV regressions have larger estimates of the effect of the skill premium on sorting than our OLS estimates, suggesting that downward bias due to measurement error in the skill premium might be larger than the potential reverse causality problem.[42]

It should be noted before concluding that a potential problem with our instrument is that one could argue that average years of schooling may have an independent (presumably negative) effect on sorting by allowing people to mix for a greater amount of time before finally separating across skill lines. This consideration suggests that our results in Table VII should be approached with some caution.[43]

III.E. Alternative Explanations

A natural concern is that our finding of a positive and significant relationship between sorting and the skill premium is driven by some third factor that is positively correlated with our variables. We now turn to an examination of various possible candidate variables that could be driving our results. We first discuss several possibilities and then report the results of including them in our baseline regression in Table VIII.

A possible (presumably exogenous) variable that could affect both sorting and the skill premium is the country's degree of ethnic fractionalization. Some authors have argued, for example, that greater ethnic fractionalization leads to greater inequality, through various political economy mechanisms (see, e.g., Alesina, Baqir, and Easterly [1999] and Easterly [2001]). This, coupled

41. Furthermore, the coefficient on the LA dummy is now negative rather than the positive effect we had found in our baseline regressions.

42. As we discussed previously, an alternative matching model would imply that the proportion of the population that is skilled also affects sorting. We attempted to include this variable in our IV estimation, but as it was not significant in the first stage of our procedure, we decided to omit this specification.

43. We thank Torsten Persson for bringing this point to our attention.

TABLE VIII
MARITAL SORTING AND THE SKILL PREMIUM: ALTERNATIVE EXPLANATIONS

	Dependent variable—marital sorting								
Explanatory variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Constant	0.362	0.395	0.506	0.548	0.579	0.579	0.536	0.589	1.406
	(0.044)***	(0.043)***	(0.095)***	(0.093)***	(0.046)***	(0.049)***	(0.073)***	(0.111)***	(0.412)***
Skill premium (Wage ratio)	0.097	0.072	0.088	0.063	0.050	0.057	0.055	0.070	0.039
	(0.015)***	(0.020)***	(0.015)***	(0.021)***	(0.013)***	(0.018)**	(0.018)***	(0.020)***	(0.020)*
Ethnic fractionalization	0.001	0.001					0.001	0.001	−0.0007
	(0.000)**	(0.000)**					(0.0006)*	(0.0006)*	(0.0007)
Urbanization			−0.001	−0.001					
			(0.001)	(0.001)					
GDP per capita					−8.01E−06	8.62E−06	−7.51E−06	−1.21E−05	−2.89E−06
					(2.45E−06)***	−2.74E−06)***	(2.98E−06)**	(6.42E−06)*	(6.53E−06)
Raw wage gap [World Bank 2001a]									−0.208
									(0.148)
Gender-related development index [Human Development Report 2003]									−0.492
									(0.411)
Enrollment ratio, female/male [Human Development Report 2003]									−0.201
									(0.215)
Percentage of labor force female									−0.001
									(0.004)
LA dummy		0.048		0.050		−0.019		−0.784	
		(0.045)		(0.042)		(0.039)		(0.083)	
Number of obs.	30	30	34	34	32	32	30	30	24
Adjusted R^2	0.449	0.444	0.386	0.386	0.484	0.469	0.501	0.499	0.590

Standard errors reflect Eicker-White correction for heteroskedasticity.
* Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

with the fact that people tend to marry individuals within their own ethnic group, could be responsible for our result.

We examine the above possibility by introducing a variable that captures the degree of ethno-linguistic fractionalization within the country. This variable, which take values between 0 and 100 with higher values indicating more fractionalization, is taken from the World Bank Growth Network (WBGN) data set.[44] For our set of countries (excluding the Czech Republic, Hungary, Poland, and Slovakia for which data were not available, the degree of ethno-linguistic fractionalization ranges from a minimum of 3 (Germany) to a maximum of 75 (Canada) with a mean value of 26.3 and a standard deviation of 20.8.

A different hypothesis is that the degree of urbanization may be driving our results. Cities, it can be argued, are places where greater mixing may occur among different types than in the countryside, which may be characterized by a more uniform degree of skill acquisition. Thus, if countries whose population is more concentrated in cities tend to have also lower skill premiums (perhaps due to better borrowing opportunities in cities), this might be responsible for our finding of a positive correlation between sorting and the skill premium.

To examine the hypothesis above, we introduce an urbanization variable (which we take from the WBGN data set as well). This variable measures the proportion of a country's population that in 1990 lived in urban areas, as reported to the United Nations by each country. In our sample, the lowest value for urbanization is 47.1 (Costa Rica), and the highest value is 96.5 (Belgium). Overall, the mean value for urbanization is 73.3, with a standard deviation of 13.2.

A last concern is that our results are driven by the level of GDP per capita. Although GDP per capita is an endogenous variable in our model, and hence treating it as an exogenous variable is problematic in the absence of a valid instrument (see the discussion in the next section), we nonetheless add it as a control variable in our regression where it can also be thought of as representing that part of income that is not captured by the variables in the model. This variable could be responsible for our observed positive correlation if, for example, countries with low income are more unequal (maybe due to technology or to some

44. This variable is available at http://www.worldbankorg/research/growth/GDNdata.htm.

political economy reason) and if not marrying "down" matters more when income is low than when it is high.[45] Thus, countries with low levels of GDP should also see higher levels of sorting.

To evaluate the argument above, we incorporate a measure of real per capita GDP (its 1997 value from the WBGN) into our regression analysis. The poorest country in our sample has real GDP per capita of $1896 (Bolivia), and the richest one has $21,974 (Luxembourg); the average for the entire sample is $9897, with a standard deviation of $5941.[46]

Table VIII reports the results of introducing each of these variables separately in a regression of marital sorting on the skill premium, using the wage ratio as the measure of the skill premium.[47] As one might expect from our discussion above, ethnic fractionalization has a positive and significant effect on marital sorting. The effect of urbanization is negative, as expected, but insignificant. Last, GDP per capita is negative and highly significant. In each specification (both with and without the additional control for Latin American), the coefficient on the skill premium remains positive and significant at the 1 or 5 percent level. In columns (7) and (8) we simultaneously include the variables that had a significant effect on sorting in the simple regressions. The effect of the skill premium remains positive and significant.

The results from the regressions reported above indicate that the positive correlation between marital sorting and the skill premium is not an artifact of some obvious third factor.

III.F. Gender Inequality and Sorting

In this subsection we analyze the relationship between gender inequality and marital sorting. As we discussed in subsection II.C, lower participation by women in the formal labor market (due either to low wages, lower probability of employment, or social norms that discourage women from working outside the home) is likely to be associated with greater marital sorting. In line with our theory, we examine how two different sets of measures that relate to how women fare across countries—gender differences in pay and wider measures of gender inequality—affect sorting. Columns (7)–(9) in Table I report several of our

45. This argument more generally depends on the sign of the third derivative of the utility function.

46. The data for Germany are from 1992.

47. We obtain similar results if we use the other measures of the skill premium.

measures of gender equality by country. Our main results are presented in Table IX.

The World Bank [2001a] uses both published and unpublished sources to provide an estimate of a "raw gender gap," which is defined as female earnings or wages as a fraction of male earnings or wages. Using the most recent available estimates for the countries in our sample, Chile and Venezuela have the highest numbers (women earn relatively well), and Taiwan and Bolivia have the lowest.[48] The sample mean is .78 with a standard deviation of 0.09. The World Bank also provides a measure of the "unexplained gender gap" which is the female-male earning difference adjusted for differences in education and experience. This variable was not significant, however, and so we omit it from the table.[49]

The UNDP publishes the Human Development Report 2003 which provides two alternative measures of gender inequality: the "gender-related development index" and gross enrollment differences in education between females and males. The gender-related development index is constructed by adjusting the average achievement of countries on life expectancy, adult literacy rate, combined primary, secondary, and tertiary education gross enrollment rate, and income by the degree of gender disparities in these dimensions.[50] The second measure isolates the education dimension used to construct the gender-related development index, i.e., the combined primary, secondary, and tertiary education gross enrollment ratio of females to males. Both measures are available for all the countries in our sample, except Taiwan. Norway has the highest value for the gender development index, and Sweden has the highest enrollment ratio, whereas Bolivia has the lowest score in the gender-development index, and Chile has the lowest enrollment ratio.

We also examined the effect of the percentage of the labor

48. Belgium, Finland, Luxembourg, Paraguay, and Slovakia, are not in the sample. Furthermore, we did not include the estimates for Spain since they were based on a study using only private sector, white-collar workers.

49. We also experimented with constructing a similar measure using our LIS and LA data sets by regressing earnings on years of schooling, potential experience (and its squared value), and a female dummy variable. We likewise obtained insignificant results when we used the coefficient on the female dummy in our sorting regression. This leads us to suspect that the more relevant features of discrimination may lie in what a woman is able to achieve rather than in her earnings given her education.

50. A detailed description of how the gender-related development index is calculated is provided in the technical notes section of the Human Development Report 2003, pages 343–344.

TABLE IX
MARITAL SORTING AND THE GENDER GAP

	Dependent variable—marital sorting								
Explanatory variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Constant	0.661	0.719	1.091	1.120	0.858	0.850	0.703	0.710	1.300
	(0.112)***	(0.107)***	(0.157)***	(0.216)***	(0.215)***	(0.212)***	(0.141)***	(0.171)***	(0.265)***
Skill premium (wage ratio)	0.096	0.067	0.045	0.050	0.079	0.060	0.064	0.065	0.052
	(0.020)***	(0.023)***	(0.012)***	(0.021)**	(0.016)***	(0.023)**	(0.019)***	(0.025)**	(0.016)***
Raw wage gap	−0.331	−0.361							−0.138
[World Bank 2001a]	(0.157)*	(0.149)**							(0.119)
Gender-related development index			−0.683	−0.721					−0.592
[Human Development Report 2003]			(0.175)***	(0.258)***					(0.195)***
Enrollment ratio, female/male					−0.406	−0.376			−0.190
[Human Development Report 2003]					(0.199)*	(0.196)*			(0.151)
Percentage of labor force female							−0.006	−0.006	
							(0.003)**	(0.004)	
LA dummy		0.063		−0.015		0.040		−0.005	
		(0.036)*		(0.030)		(0.039)		(0.064)	
Number of obs.	28	28	33	33	33	33	32	32	27
Adjusted R^2	0.488	0.504	0.543	0.529	0.463	0.458	0.469	0.451	0.621

Standard errors reflect Eicker-White correction for heteroskedasticity.
* Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

force that is female to see whether this variable would be able to summarize different types of discrimination. We used 1990 numbers, available from the World Bank [2001b]. In this sample, Luxembourg and Taiwan are excluded since the data were not available. The lowest value for the fraction of labor force that is female is 27.7 (Ecuador), and the highest is 48 (Finland and Sweden).

As can be seen in Table IX, all measures of female status enter negatively and significantly in explaining sorting, as predicted by our theory. Note that a higher value of the "discrimination" variable is in all cases associated with a better position for women, i.e., earning a greater fraction of male wages, higher female to male enrollment, and higher gender development. Only the percentage of labor force that is female is no longer significant once we control for Latin America. In all cases, furthermore, the effect of the (male) skill premium remains positive and significant at the 1 or 5 percent level. The last column in Table IX enters all the gender-related variables that were significant on their own (and survived the inclusion of a Latin American dummy). As can be seen, in this specification only the gender-development index and the skill premium remain significant.[51]

We also included the gender discrimination variables as a final specification in Table VIII, omitting the LA dummy as this variable had no explanatory power once the alternative candidates were included. As can be seen, none of these variables is significant in the larger regression (including GDP per capita), although the skill premium remains positive and significant (though, not surprisingly, the significance is reduced as there are only 24 observations and 7 explanatory variables).

From this section we conclude that discrimination against women appears to be positively and significantly related to sorting, which is in line with the predictions of our model.

III.G. Per Capita Income, Skilled Population, and Sorting

We now turn to an examination of a different prediction of our model: the existence of a negative relation between marital sorting and per capita income across countries. Note that our model implies that across steady states, economies with the same technology are characterized by a positive relationship between

51. Throughout this table we used the wage ratio measure of the skill premium. We obtain similar results using the other measures.

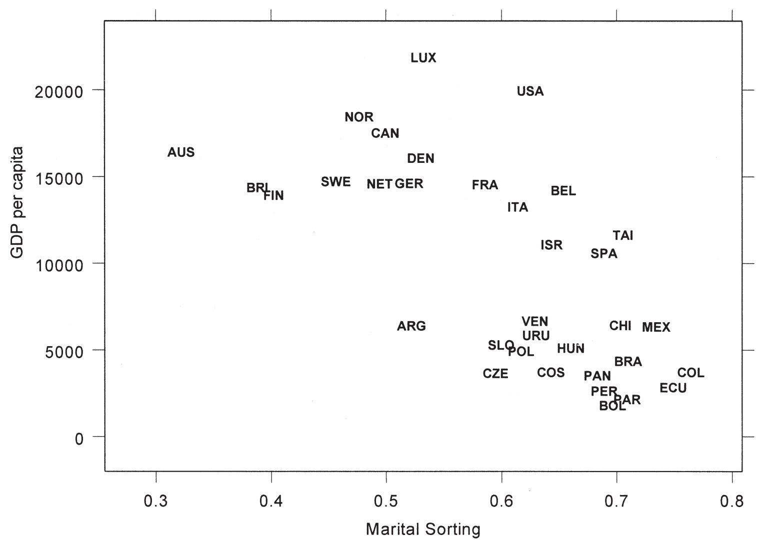

FIGURE V
GDP per Capita and Marital Sorting

sorting and the skill premium and a negative relationship of these with the proportion of skilled individuals. Consequently, ceteris paribus, we expect economies with similar technologies but with greater sorting to have lower per capita income as their level of human capital will be further from the efficient level.[52] We turn to an examination of this relationship in the data.[53]

Figure V shows the relation between marital sorting and per capita income where per capita income is real GDP per capita in 1997. Table X reports the regression results for specifications in which the dependent variable is the skilled proportion of the male population in our sample (columns (1) and (2)) and per capita

52. We also attempted in a variety of ways to investigate the effect of credit market imperfection on our results, but we found the effect to be negligible (we thank Michael Kremer for suggesting this exercise). We think that this is likely to be due to the fact that the various measures of credit market imperfections (e.g., M3/GDP, deposit money bank assets/GDP, private credits by money banks and other financial institutions/GDP, concentration of the banking sector, overhead costs in the banking sector, and share of foreign banks in the banking sector), are not really able to capture how difficult it is to borrow against future human capital nor to distinguish among different countries' education policies which would be the more relevant variables.

53. Note that this prediction of our model need hold only across steady states and for countries with the same technology, unlike our other predictions which require neither assumption.

TABLE X
IMPLICATIONS OF MARITAL SORTING FOR GDP PER CAPITA AND THE PROPORTION SKILLED

	Dependent variable			
	Proportion skilled		GDP per capita	
Explanatory variable	(1)	(2)	(3)	(4)
Constant	0.513	0.466	31476.25	23978.05
	(0.513)***	(0.102)***	(3396.62)***	(2628.19)***
Marital sorting	−0.436	−0.332	−35843.56	−19352.62
	(0.141)***	(0.192)*	(5397.69)***	(5395.02)***
LA dummy		−0.04		−6354.79
		(0.048)		(1534.31)***
Number of obs.	34	34	34	34
Adjusted R^2	0.121	0.110	0.422	0.601

Standard errors reflect Eicker-White correction for heteroskedasticity.
* Significant at 10 percent; ** Significant at 5 percent; *** Significant at 1 percent.

income (columns (3) and (4)). In both cases, the explanatory variable is marital sorting with and without a Latin American dummy. As shown, sorting has a significant and negative effect in each specification as predicted by our model.

IV. CONCLUSIONS

This paper has examined the relationship between marital or household sorting and income inequality. Using a general equilibrium model in which potentially borrowing constrained individuals make decisions over education and matching, we find that there is a positive relationship between household sorting and inequality. In particular, whether at a point in time, or across steady states, economies with greater skill premiums should also display a greater degree of sorting. Our model also predicts that if economies have identical technologies, more sorting should be associated with lower per capita income and a lower proportion of skilled workers. Furthermore, in a variant of our model that differentiates between men and women, we find that economies with less discrimination against women should have less sorting.

Our empirical work is based on household surveys for 34 countries. From these surveys we create our main sample for each country, a measure of marital sorting, and various measures of the skill premium. Our regression analysis supports our central prediction of a positive relationship between sorting and inequality across countries even after controlling for other possible sources of this correlation, such as urbanization and ethnic fractionalization. We also show that our results are not driven by GDP per capita and that less female discrimination is associated with less sorting. Last, we also find evidence in favor of a negative relationship between sorting and per capita income.

It should be noted that our story of greater pickiness with respect to household partners in the face of an increased skill premium is of course not the only one compatible with a positive correlation between these two variables. An alternative story, with similar mechanics, would be of greater sorting into communities or schools in response to greater inequality (say, in response to fear of more crime). This could then lead to fewer opportunities to interact between individuals of different skill groups and consequently to a greater correlation of spouses in education. We do not see this mechanism as being very different. Once again, private decisions that are sensitive to the degree of inequality (e.g., where to live, where to go to school) would feed through to marital choices, which would then have important social consequences as a result of borrowing constraints.

Appendix 1: An Alternative Model of Sorting

In this alternative framework, sorting arises from the efforts of skilled agents to segregate themselves from the rest of society. One can think of this effort as the cost that a skilled agent is willing to bear in order to live in a more segregated neighborhood, attend a more segregated school, or belong to a more exclusive social club. A simple way to model this is to assume that each skilled agent chooses an effort level e. This effort results in a probability $p(e)$ of being perfectly segregated and thereby obtaining a match with another skilled agent for sure, and a probability $1 - p(e)$ of not being able to segregate and hence obtaining a match at random. The benefit associated with segregation is that the skilled agent obtains $u(I_{ss})$, whereas otherwise the agent obtains a random draw with expected utility given by $\hat{\lambda}u(I_{ss}) + (1 - \hat{\lambda})u(I_{su})$, where $\hat{\lambda}$ is the proportion of skilled agents in the nonsegregated (i.e., random matching) pool. The cost of making an effort to segregate is $c(e)$, where $c(e)$ is an increasing, continuous, convex twice-differentiable cost function with $c(0) = 0$. To simplify, we assume that all matches end in marriages.

A skilled agent chooses e to maximize

$$(19)\qquad p(e)u(I_{ss}) + (1 - p(e))[\hat{\lambda}u(I_{ss}) + (1 - \hat{\lambda})u(I_{su})] - c(e),$$

taken as given the (simultaneous) effort levels of all other skilled agents and thus $\hat{\lambda}$. We assume that $p(e)$ is an increasing, concave, continuous, twice-differentiable function with $p(0) = 0$, $\lim_{e\to\infty} p(e) = 1$, and that it satisfies the Inada conditions thus ensuring an interior solution.

The first- and second-order conditions associated with this maximization problem are, respectively,

$$(20)\qquad p'(e)(1 - \hat{\lambda})(u(I_{ss}) - u(I_{su})) - c'(e) = 0$$

and

$$(21)\qquad p''(e)(1 - \hat{\lambda})(u(I_{ss}) - u(I_{su})) - c''(e) < 0.$$

Equation (20) defines a relationship $e^*(\hat{\lambda})$. Note that e^* is decreasing in $\hat{\lambda}$: the greater the proportion of skilled agents in the nonsegregated population, the less effort a skilled agent will make to segregate. This is a source of potential multiple equilibria since if other skilled agents are making a larger (smaller) effort to segregate, an individual skilled agent will likewise find it optimal to make a larger (smaller) effort as well. Note that since all skilled agents face the same marginal cost and benefit decision, they will all choose the same level of e in equilibrium.[54]

$\hat{\lambda}$ is a function both of the proportion of skilled agents, λ, and of the effort that these make to segregate themselves. We can write $\hat{\lambda}$ as

$$(22)\qquad \hat{\lambda} = \frac{(1 - p(e))\lambda}{(1 - p(e))\lambda + (1 - \lambda)} = \frac{(1 - p(e))\lambda}{1 - p(e)\lambda}$$

since if skilled agents choose effort level e this will result in a

54. Alternatively, we could assume that skilled agents can segregate with probability one, but that they face heterogeneous costs in doing so. Thus, the greater the benefit from segregation, the larger is the cost the marginal skilled agent is willing to bear to do so, and thus the larger is the proportion of skilled agents that choose to segregate. As before, this can generate multiple equilibria since the relative benefit of segregation is increasing in the number of skilled agents who do so, as this decreases the proportion of skilled agents in the random sorting pool.

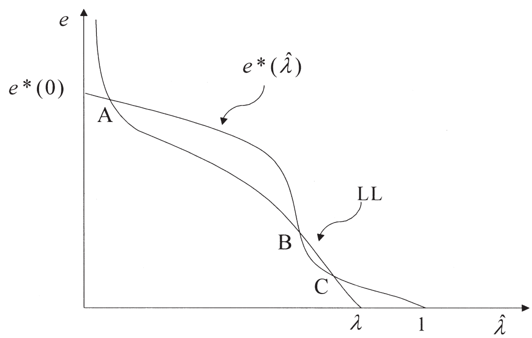

FIGURE VI
Equilibrium Determination of $\hat{\lambda}$

measure $p(e)$ of them segregating and leave $1 - p(e)$ of them to be matched randomly in a population of $(1 - p(e))\lambda + (1 - \lambda) = 1 - p(e)\lambda$ agents. Equation (22) defines a relationship $\hat{\lambda}(e;\, \lambda)$. Note that $\hat{\lambda}$ is a decreasing function of e.

Figure VI depicts equilibrium in this model as the intersection of two loci, for a given λ. The $e^*(\hat{\lambda})$ locus gives the value of e that an individual skilled agent would choose if she expected the probability of being matched at random with another skilled agent to be $\hat{\lambda}$. The LL locus depicts the values of e and $\hat{\lambda}$ that satisfy equation (22) for a given λ. Note that, as shown in the figure, there may be multiple equilibria. We will focus only on those equilibria that are locally stable such as point A or C in Figure VI. Point B is not locally stable since an increase in the effort of a small measure of skilled agents would lead to a sufficiently large decline in $\hat{\lambda}$ that, faced with that new level of $\hat{\lambda}$, agents would now choose an even greater level of effort.

More formally, for an equilibrium to be locally stable, the slope of the $e^*(\hat{\lambda})$ curve must be greater than the slope of the LL curve; i.e.,

$$S = [p''(1 - \lambda)(u(I_{ss}) - u(I_{su})) - (1 - p\lambda)c''] + p'^2\lambda \frac{1 - \lambda}{1 - p\lambda}(u(I_{ss}) - u(I_{su})) < 0. \tag{23}$$

Note that a stable equilibrium always exists since whereas the LL curve intersects the $\hat{\lambda}$ axis at $\hat{\lambda} = \lambda$ and asymptotes to infinity on the e axis, the $e^*(\hat{\lambda})$ curve intersects the $\hat{\lambda}$ axis at $\hat{\lambda} = 1$ and meets the e axis at some finite value, $e^*(0)$.

Given a proportion λ_t of skilled agents at time t, the equilibrium distribution of households is given by

$$\rho_{ij}(\lambda_t) = \begin{cases} p(e^*)\lambda_t + \dfrac{\lambda_t^2(1-p(e^*))^2}{1-p(e^*)\lambda_t}, & \text{if } ij = ss \\ \dfrac{2\lambda_t(1-\lambda_t)(1-p(e^*))}{1-p(e^*)\lambda_t}, & \text{if } ij = su \\ \dfrac{(1-\lambda_t)^2}{1-p(e^*)\lambda_t}, & \text{if } ij = uu. \end{cases} \tag{24}$$

LEMMA A1. The correlation of spouses in skill type is given by $p(e^*)(1 - \lambda)/(1 - p(e^*)\lambda)$.

Proof of Lemma A1. As in the proof of Lemma 1, let X be the random variable associated with the skill distribution of women (i.e., $x = 1$ if $x = s$, and $x = 0$ if $x = u$), and let Y be the random variable associated with the skill distribution of men. Solving for the correlation ρ yields $\rho = p(e^*)(1 - \lambda)/(1 - p(e^*)\lambda)$.‖

Note that, unlike the previous model of sorting, the proportion of skilled agents in the economy directly affects the correlation coefficient. Nonetheless, it is easy to show that, in this model as well, an increase in inequality will increase the amount of sorting. As before, we first examine the consequences of an exogenous increase in the skill premium and then analyze the effect of a change in λ.

THEOREM A1. An increase in the skill premium, w_s/w_u, increases marital sorting.

Proof of Theorem A1. Totally differentiating e^* with respect to w_s (using (20) and (22)) yields

$$\frac{de^*}{dw_s} = -\frac{p'(1-\lambda)(1-p\lambda)}{S}\,[2u'(I_{ss}) - u'(I_{su})] > 0,$$

which is positive by A1 and the stability condition (23). Similarly,

$$\frac{de^*}{dw_u} = \frac{p'(1-\lambda)(1-p\lambda)}{S}\,u'(I_{su}) < 0.$$

Recalling that $\rho = p(e^*)(1 - \lambda)/(1 - p(e^*)\lambda)$, and hence that $\partial\rho/\partial e^* > 0$, it follows that sorting is increasing in the skill premium.‖

Thus, in this alternative model, as in the previous one, an increase in inequality between skilled and unskilled agents causes the former to form a greater proportion of households with their own type, thereby increasing the degree of marital sorting in the economy. In terms of Figure VI, an increase in the skill premium shifts out the e^* locus, causing any stable equilibrium to be characterized by a higher e and a lower $\hat{\lambda}$.

COROLLARY A1. A decrease in λ increases marital sorting.

Proof of Corollary A1. Note that $\partial\rho/\partial\lambda = -p(e^*)(1 - p(e^*))/(1 - p\lambda)^2 < 0$ and that $\partial e^*/\partial\lambda < 0$. Furthermore, $d\tilde{w}_s/d\lambda < 0$, and $dw_u/d\lambda > 0$. Totally differentiating ρ, it then follows directly from the theorem above that a decrease in λ increases sorting.‖

The intuition for the result above is clear. A decrease in λ increases the effort that skilled agents make to segregate themselves both because a lower λ increases inequality and because it directly decreases the probability that random matching will result in contact with another skilled agent. In terms of Figure VI, the decrease in λ not only shifts out the e^* locus as before due to the increase in the skill premium, but it also shifts in the LL locus due to the decrease in $\hat{\lambda}$. The combined effect of these two shifts on any locally stable equilibrium is to increase e and decrease $\hat{\lambda}$.

APPENDIX 2: SORTING AND GENDER

Proof of Proposition 1. Note that $1 + \beta < 1 + w_{fs}/w_{ms}$. Hence, to show $q^*_{fs} < \tilde{q}^*_{fs}$, it is sufficient to show that $(w_{ms} + w_{fs}) \ln((w_{ms} + w_{fs})/(w_{mu} + w_{fs})) < w_{ms} \ln(w_{ms}/w_{mu})$. Noting that the expression on the right-hand side of the inequality is the same as the one on the left-hand side but with w_{fs} set to zero, we can prove the inequality by showing that $(w_{ms} + w_{fs}) \ln((w_{ms} + w_{fs})/(w_{mu} + w_{fs}))$ is decreasing in w_{fs}. Let $\partial(w_{ms} + w_{fs}) \ln((w_{ms} + w_{fs})/(w_{mu} + w_{fs}))/\partial w_{fs} = \ln((w_{ms} + w_{fs})/(w_{mu} + w_{fs})) + (w_{mu} - w_{ms})/(w_{mu} + w_{fs}) \equiv z(w_{ms}, w_{mu}, w_{fs})$. To show that this expression is negative, note that it is decreasing in w_{ms} (i.e., $\partial z/\partial w_{ms} = (w_{mu} - w_{ms})/(w_{mu} + w_{fs})(w_{ms} + w_{fs}) < 0$). Also note that $w_{ms} > w_{mu}$. Thus, evaluating z, at $w_{ms} = w_{mu}$, yields $z = 0$, and hence $q^*_{fs} < \tilde{q}^*_{fs}$.‖

Simulation Results

Let λ_m be the fraction of skilled males and λ_f be the fraction of skilled females in the population. If $t_j > 0$, after substituting equation (15) into equation (13), the value of being married between a type-i male and a type-j female with match quality q is given by

$$V_{ij} + q = \ln\left[w_{mi} + \frac{w_{fj} - \beta w_{mi}}{1+\beta}\right] + \beta \ln\left[\frac{\beta}{1+\beta}\frac{w_{fj} + w_{mi}}{w_{fj}}\psi_j\right] + q,$$

whereas for $t_j = 0$ it is given by

$$V_{ij} + q = \ln(w_{mi}) + \beta \ln(\psi_j) + q.$$

Consider the first round of matching, and let $q^*_{f,ij}$ be the threshold match quality level for a type-j female matched with a type-i male, and $q^*_{m,ij}$ be the reservation quality for a type-i male matched with a type-j female.

The equilibrium distribution of matches in the first round is thus given by

$$\tilde{\phi}_{ij} = \begin{cases} \lambda_m \lambda_f (1 - Q(q^*_{f,ss}))(1 - Q(q^*_{m,ss})), & \text{if } ij = ss \\ (1-\lambda_m)(1-\lambda_f)(1 - Q(q^*_{f,uu}))(1 - Q(q^*_{m,uu})), & \text{if } ij = uu \\ \lambda_m (1-\lambda_f)(1 - Q(q^*_{f,su}))(1 - Q(q^*_{m,su})), & \text{if } ij = su \\ \lambda_f (1-\lambda_m)(1 - Q(q^*_{f,us}))(1 - Q(q^*_{m,us})), & \text{if } ij = us. \end{cases}$$

This implies that in the second round, the new proportion of skilled males and females is given by

$$\lambda'_m = \frac{\lambda_m - \tilde{\phi}_{ss} - \tilde{\phi}_{su}}{1 - \tilde{\phi}_{ss} - \tilde{\phi}_{uu} - \tilde{\phi}_{su} - \tilde{\phi}_{us}}$$

and

$$\lambda'_f = \frac{\lambda_f - \tilde{\phi}_{ss} - \tilde{\phi}_{us}}{1 - \tilde{\phi}_{ss} - \tilde{\phi}_{uu} - \tilde{\phi}_{su} - \tilde{\phi}_{us}}.$$

Therefore, the equilibrium distribution of households is finally given by

$$\phi_{ij} = \begin{cases} \tilde{\phi}_{ss} + \lambda'_m \lambda'_f (1 - \tilde{\phi}_{ss} - \tilde{\phi}_{uu} - \tilde{\phi}_{su} - \tilde{\phi}_{us}), & \text{if } ij = ss \\ \tilde{\phi}_{uu} + (1-\lambda'_m)(1-\lambda'_f)(1 - \tilde{\phi}_{ss} - \tilde{\phi}_{uu} - \tilde{\phi}_{su} - \tilde{\phi}_{us}), & \text{if } ij = uu \\ \tilde{\phi}_{su} + \lambda'_m (1-\lambda'_f)(1 - \tilde{\phi}_{ss} - \tilde{\phi}_{uu} - \tilde{\phi}_{su} - \tilde{\phi}_{us}), & \text{if } ij = su \\ \tilde{\phi}_{us} + (1-\lambda'_m)\lambda'_f (1 - \tilde{\phi}_{ss} - \tilde{\phi}_{uu} - \tilde{\phi}_{su} - \tilde{\phi}_{us}), & \text{if } ij = us. \end{cases}$$

This yields a correlation between husbands' and wives' skill types of

TABLE XI

	$\psi_s = 1.5$		$\psi_s = 2.0$		$\psi_s = 2.5$	
	Work	No work	Work	No work	Work	No work
ρ	0.0990	0.1109	0.1216	0.1369	0.1382	0.1562
$q^*_{f,ss}$	0.7008	0.6007	0.6989	0.5979	0.6976	0.5957
$q^*_{f,su}$	0.6496	0.6007	0.6475	0.5979	0.6458	0.5957
$q^*_{f,us}$	1.0718	1.0905	1.0699	1.0877	1.0685	1.0855
$q^*_{f,uu}$	1.0840	1.0905	1.0819	1.0877	1.0803	1.0855
$q^*_{m,ss}$	0.8429	0.8760	0.7494	0.7834	0.6748	0.7093
$q^*_{m,su}$	1.0335	1.0277	1.0476	1.0429	1.0566	1.0523
$q^*_{m,us}$	0.7906	0.8760	0.6961	0.7834	0.6208	0.7093
$q^*_{m,uu}$	1.0446	1.0277	1.0578	1.0429	1.0660	1.0523

$$\rho = \frac{\phi_{ss} - \lambda_m\lambda_f}{\sqrt{\lambda_m(1-\lambda_m)}\ \sqrt{\lambda_f(1-\lambda_f)}}$$
$$= \frac{\tilde{\phi}_{ss} + \lambda'_m\lambda'_f(1 - \tilde{\phi}_{ss} - \tilde{\phi}_{uu} - \tilde{\phi}_{su} - \tilde{\phi}_{us}) - \lambda_m\lambda_f}{\sqrt{\lambda_m(1-\lambda_m)}\ \sqrt{\lambda_f(1-\lambda_f)}}.$$

We can now solve for the reservation qualities of each gender-type as

$$q^*_{f,ij} = \begin{cases} \mu - (V_{ss} - V_{us})(1 - \lambda'_m), & \text{if } ij = ss, \\ \mu - (V_{uu} - V_{su})\lambda'_m, & \text{if } ij = uu, \\ \mu - (V_{su} - V_{uu})(1 - \lambda'_m), & \text{if } ij = su, \\ \mu - (V_{us} - V_{ss})\lambda'_m, & \text{if } ij = us, \end{cases}$$

for females; and as

$$q^*_{m,ij} = \begin{cases} \mu - (V_{ss} - V_{su})(1 - \lambda'_f), & \text{if } ij = ss, \\ \mu - (V_{uu} - V_{us})\lambda'_f, & \text{if } ij = uu, \\ \mu - (V_{su} - V_{ss})\lambda'_f, & \text{if } ij = su, \\ \mu - (V_{su} - V_{uu})(1 - \lambda'_f), & \text{if } ij = us, \end{cases}$$

for males.

Note that $q^*{}'s$ are functions of λ'_m and λ'_f. Hence, an equilibrium is a fixed point of q^*'s and λ'_m and λ'_f. The exercise now is to compare two societies. In the first, women choose t_j freely. In the second society, women have $t_j = 0$ (either because of social norms or because $w_{fj} \leq \beta w_{m,i}$), and show that there is less sorting when women work.

In Tables XI and XII we report results from two sets of

TABLE XII

	$\psi_s = 1.5$		$\psi_s = 2$		$\psi_s = 2.5$	
	Work	No work	Work	No work	Work	No work
ρ	0.1924	0.2034	0.2297	0.2426	0.2562	0.2708
$q^*_{f,ss}$	0.3432	−0.0003	0.3383	−0.0079	0.3342	−0.0141
$q^*_{f,su}$	0.0701	−0.0003	0.0631	−0.0079	0.0573	−0.0141
$q^*_{f,us}$	1.0643	1.0950	1.0593	1.0874	1.0552	1.0812
$q^*_{f,uu}$	1.0910	1.0950	1.0840	1.0874	1.0781	1.0812
$q^*_{m,ss}$	0.8246	0.8648	0.7227	0.7649	0.6421	0.6854
$q^*_{m,su}$	1.0144	1.0141	1.0185	1.0203	1.0201	1.0229
$q^*_{m,us}$	0.5474	0.8648	0.4416	0.7649	0.3582	0.6854
$q^*_{m,uu}$	1.0371	1.0141	1.0373	1.0203	1.0361	1.0229

simulations. In both cases Q is the cumulative of a triangular probability distribution with $\bar{q} = 2$ and $\mu = 1$. We also set $\psi_u = 1$ and $w_{fu} = 1$ in both cases. Table XI shows results when the rest of the parameters are set to their average values in our LIS sample. In particular, we set $\lambda_m = 0.282$, $\lambda_f = 0.248$, $w_{fs} = 1.427$, $w_{mu} = 1.455$, and $w_{ms} = 2.331$. Given the wage levels, we set $\beta = 0.3743$, which implies that a skilled female who is matched with another skilled male works about 0.3 of her one unit time in the market. In Table XII we set the proportion of skilled individuals and the skill premium to their average values in our LA sample ($\lambda_m = 0.199$, $\lambda_f = 0.161$, $w_{fs} = 1.335$, $w_{mu} = 2.495$, and $w_{ms} = 3.992$), and following the same procedure as in Table XI choose $\beta = 0.3684$.[55]

Results in Tables XI and XII provide further support for the positive link between gender discrimination and sorting that we established in subsection II.C and documented empirically in subsection III.F. In both tables skilled females who are matches with unskilled females are more picky (i.e., $q^*_{f,us}$ values are higher) and sorting is higher when women are not allowed to work.

We obtained results similar to the above for a wide range of parameter values. It is not, of course, impossible to generate the

55. Since there are only two rounds of random matching, the correlation coefficients in Tables XI and XII are quite a bit lower than the ones in the data. Although adding more rounds of random matching will improve the match of the simulations in this dimension, this would unnecessarily complicate things without adding to the analysis.

opposite conclusion. For this to happen, children must not matter very much (i.e., β is low), the skill premium for women, w_{fs}/w_{fu} should be high, and the difference in child quality from a skilled relative to an unskilled woman should be small (i.e., ψ_s/ψ_u close to one). This implies that, in an economy in which women work freely, they will work a great deal, and the differences in female earnings by skill level will be high. Hence skilled men in this economy will be very picky, and sorting will be high. If women are not able to work, on the other hand, since men do not care much about the difference in child quality (which in any case is small), skilled men will not care very much whether they marry a skilled or an unskilled woman. This indifference can be sufficiently powerful so as to undo the greater pickiness of skilled women in this economy.[56] We think that our examples in which the greater pickiness of skilled women drive the result of greater sorting are perhaps more realistic, though a full-scale calibration is beyond the range of this paper.

Appendix 3: Uniqueness of Equilibrium

Note first that an increase in $\tilde{w}_s$ makes it more attractive to be a skilled agent since it increases the direct return to being skilled. It also makes skilled agents less likely to want to form a household with an unskilled agent; i.e., it increases q^*. An increase in $\tilde{w}_s$, on the other hand, has ambiguous effects on an unskilled agent's payoff since although it increases the value of being in a household with a skilled worker, it also makes these matches more unlikely:

$$\begin{aligned}(25)\quad \frac{d\delta^*}{d\tilde{w}_s} &= \frac{d[V^s - V^u]}{d\tilde{w}_s} \\ &= [\lambda + (1-\lambda)Q(q^*)]\frac{\partial V_{ss}}{\partial \tilde{w}_s} + [(1-Q(q^*))(1-2\lambda)]\frac{\partial V_{su}}{\partial \tilde{w}_s} \\ &\quad - \frac{dq_1^*}{d\tilde{w}_s}\lambda[V_{uu}(\mu) - V_{su}(q^*)]Q'(q^*) > 0,\end{aligned}$$

which is strictly positive as $\partial V_{ss}/\partial \tilde{w}_s = 2u'(2\tilde{w}_s) > \partial V_{su}/\partial w_s = u'(\tilde{w}_s + w_u)$, $\lambda + (1-\lambda)Q(q^*) > (1 - Q(q^*))(1 - 2\lambda)$,

56. For example, $\beta = 0.1$, $w_{um} = w_{uf} = 1$, $w_{sm} = w_{sf} = 2$, $\psi_s = 1.2$, $\psi_u = 1$, $\bar{q} = 2$, $\lambda_m = \lambda_f = .3$ result in greater sorting if women cannot work than if they can freely choose the amount they wish to work ($\rho = .082$ if women work as compared with $\rho = .073$ if they do not work).

$dq^*/dw_s > 0$ and $V_{uu}(\mu) - V_{su}(q^*) \equiv u(2w_u) - u(2\tilde{w}_s) < 0$ (with the latter following from the fact that skilled workers choose a higher cutoff quality level in their matches with unskilled individuals than what the latter find optimal).[57]

An increase in w_u, on the other hand, has a negative effect on the relative desirability of being a skilled agent as

$$\text{(26)} \quad \frac{d\delta^*}{dw_u} = [(1 - Q(q^*))(1 - 2\lambda)] \frac{\partial V_{su}}{\partial w_u} - [1 - \lambda + \lambda Q(q^*)] \frac{\partial V_{uu}}{\partial w_u} + \frac{dq^*}{dw_u} \lambda [V_{su}(q^*) - V_{uu}(\mu)] Q'(q^*) < 0$$

is strictly negative, since both the expression on the first line, which can be written as $[(1 - Q(q^*))(1 - 2\lambda)]u'(\tilde{w}_s + w_u) - [1 - \lambda + \lambda Q(q^*)]2u'(2w_u)$, and the one in the second line, which can be written as $dq^*/dw_u \lambda[u(2\tilde{w}_s) - u(2w_u)]Q'(q^*)$, are negative.

A change in λ_{t+1} has two effects: (i) it changes wages and hence household incomes; (ii) it changes the probability with which an agent encounters skilled and unskilled agents in the first round of matching. The total effect on the payoff differential δ^* between skilled and unskilled agents is given by

$$\text{(27)} \quad \frac{d\delta^*(\lambda_{t+1})}{d\lambda_{t+1}} = \frac{\partial[V^s - V^u]}{\partial \tilde{w}_s} \frac{d\tilde{w}_s}{d\lambda_{t+1}} + \frac{\partial[V^s - V^u]}{\partial w_u} \frac{dw_u}{d\lambda_{t+1}} + \frac{\partial \delta^*(\lambda_{t+1})}{\partial \lambda_{t+1}}.$$

Note that we can rewrite δ^* as

$$\delta^* = \lambda \left[V_{ss}(\mu) Q(\mu) + \int_{\mu}^{\bar{q}} V_{ss}(x)\, dQ(x) - \int_{q^*}^{\bar{q}} V_{su}(x)\, dQ(x) - V_{uu}(\mu) Q(q^*) \right]$$

57. For notational convenience, we have suppressed everywhere the dependence of V_{ij} on λ.

$$+ (1-\lambda)\left[V_{ss}(\mu)Q(q^*) + \int_{q^*}^{\bar{q}} V_{su}(x)\, dQ(x) - V_{uu}(\mu)Q(\mu) - \int_{\mu}^{\bar{q}} V_{uu}(x)\, dQ(x)\right],$$

which after substituting in (1) and (3) yields

$$\delta^* = u(2\tilde{w}_s)[\lambda + (1-\lambda)Q(q^*)] - u(2w_u)[1-\lambda+\lambda Q(q^*)] \times (2\lambda - 1)\int_{\mu}^{q^*} (x-\mu)\, dQ(x) - (2\lambda-1)(1-Q(q^*))u(\tilde{w}_s + w_u).$$

Taking the derivative of δ^* with respect to λ yields (after some manipulation)

(28)

$$\frac{d\delta^*}{d\lambda} = \frac{dk}{d\lambda}\{u'(2\tilde{w}_s)2f''(k)(\lambda + (1-\lambda)Q(q^*)) - u'(2w_u)2(-f''(k)k) \times [1-\lambda(1-Q(q^*))] - u'(\tilde{w}_s + w_u)(f''(k)(1-k)) \times [(2\lambda-1)(1-Q(q^*))]\} + Q'\frac{\partial q^*}{\partial \lambda}\lambda[u(2\tilde{w}_s) - u(2w_u)] + [u(2\tilde{w}_s) + u(2w_u) - 2u(\tilde{w}_s + w_u)](1-Q(q^*)) + 2\int_{\mu}^{q^*} (x-\mu)\, dQ(x).$$

In order to sign $d\delta^*/d\lambda$, note that the sign of the expression in the first brace of (28) is negative. The expression in the third line is negative, since $\partial q^*/\partial\lambda < 0$ and the expression multiplying it is positive. The expression in the fourth line is also negative, if u is concave. The term in the last line, however, is strictly positive, i.e., $\int_{\mu}^{q^*} (x-\mu)dQ(x) > 0$, which makes the sign of the entire expression ambiguous.

Appendix 4: Data

Table XIII reports the years of the households surveys used in our empirical study. All surveys are nationally representative samples, except for Argentina and Uruguay for which we have only urban samples (70 percent of the population for Argentina and 90 percent for Uruguay). The income in the Latin American countries is gross monthly labor income from all sources. This definition includes income from both primary and secondary labor activities; the exact components vary somewhat across countries, but generally include wages, income from self-employment, and proprietor's income, as well as adjustment to reflect imputation of nonmonetary income. Some LIS countries report gross annual earnings and income, and some report these net of taxes.[58] We use gross labor earnings for LIS countries whenever it is available. The gross earnings measure for LIS countries include all forms of cash wage and salary income, including employer bonuses, thirteenth month bonus, etc. (gross of employee social insurance contributions/taxes but net of employer social insurance contributions/taxes). While most countries report gross earnings, the following countries report only net earnings: Belgium, France, Hungary, Italy, Luxembourg, Poland, and Spain. Since taxation tends to be progressive in the countries we are comparing, inequality of income is likely to be higher than reported in those countries for which pretax income is not reported.

For most Latin American countries we have both total years of schooling as well as highest level of educational attainments. The only country for which we do not have levels of educational attainment is Costa Rica, while Brazil, Paraguay, and Peru do not report years. We define as skilled all agents who went beyond high school education. When years of education is not reported, this indicator is constructed using the standard age-grade progression for that country. For Britain we would have preferred to define as skilled any individual with at least 2A levels passes (as in Fernández [2001] or Pissarides [1982]), but as the data did not permit us to distinguish among individuals with different numbers of A levels, we instead categorized them all as unskilled. Our results are robust to categorizing them all as skilled instead. Table XIV reports our mapping of education measures into years of schooling and into an indicator for high school completions. For most countries we were able to compare the percentage of adults with education beyond the high-school level with published

58. Some LIS countries are excluded because they do not report all of the variables required for the analysis. Ireland and Austria do not report individual labor income. Ireland also does not report the education of the spouse in the household sample. Education variables are not available for Switzerland.

TABLE XIII
SURVEY INFORMATION

Country	Year	Name	Agency	Number of households	Coverage
Argentina	1996	Encuesta permanente de hogares	Instituto nacional de Estadistica y Censos	3369	Greater Buenos Aires
Australia	1994	Australian Income and Housing Survey	Australian Bureau of Statistics	7441	National
Belgium	1992	Panel Survey of the Centre for Social Policy	Centre for Social Policy	3821	National
Bolivia	1997	Encuesta Nacional de Empleo	Instituto nacional de Estadistica y Censos	8461	National
Brazil	1996	Pequiso Nacional por Amostra de Domicilios	Fundaçao Instituto Brasileiro de Geografia e Estatistica	84947	National
Britain	1997	British Household Panel Study, Wave G	Institute for Social and Economic Research	4384	National
Canada	1994	Survey of Consumer Finances	Statistics Canada	39039	National
Chile	1996	Encuesta Nacional de Empleo	Instituto Nacional de Estadistica	30953	National
Colombia	1997	Encuesta Nacional de Hogares	Departamento Administrativo Nacional de Estadistica	31264	National
Costa Rica	1996	Encuesta Permanente de hogares de Propositos Multiples	Direccion General de Estadistica y Censos	9471	National
Czech Republic	1996	Microcensus	Czech Statistical Office	16234	National
Denmark	1992	Income Tax Survey	National Institute of Social Research	12895	National

Ecuador	1996	Encuesta Periodica de Empleo y Desempleo en el Area Urbana	Instituto nacional de Estadistica y Censos	8153	Urban
Finland	1995	Income Distribution Survey	Statistics Finland	9262	National
France	1994	Enquête Budget des familles	INSEE Division Conditions de vie des Ménages	11294	National
Germany	1994	German Social Economic Panel Study	DIW Berlin	6045	National
Hungary	1994	Hungarian Household Panel	Endre Sik/Istvan Toth	1992	National
Israel	1992	Family Expenditure Survey	Israeli Central Bureau of Statistics	5212	National
Italy	1995	Indagine Campionaria sui Bilauci Delle Famiglie	Ufficio Informazioni Statistiche	8135	National
Luxembourg	1994	Liewen zu Letzebuerg	Centre d'Etudes de Populations, de Pauvreté et de Politiques Socio-Economiques	1813	National
Mexico	1996	Encuesta Nacional de Increso Gasto de los Hogares	Instituto Nacional de Estadistica, Geografia e Informatica	14042	National
Netherlands	1994	Socio-Economic Panel	Centraal Bureau voor de Statistiek	5187	National
Norway	1995	Income and Property Distribution Survey	Statistics Norway	10127	National

(continued on next page)

TABLE XIII
(CONTINUED)

Country	Year	Name	Agency	Number of households	Coverage
Panama	1997	Encuesta de Hogares	Direccion de Estadistica y Censo	9897	National
Paraguay	1998	Encuesta Integrada de Hogares	Direccion General de Estadistica, Encuestas y Censos	4353	National
Peru	1997	Encuesta de Hogares	Instituto Nacional de Estadistica e Informatica	3843	National
Poland	1995	Household Budget Survey	Central Statistical Office	6602	National
Slovakia	1992	Slovak Microcensus	Statistical Office of the Slovak Republic Division of Social Statistics and Demography	17714	National
Spain	1990	Expenditure and Income Survey	Instituto Nacional de Estadistica	11294	National
Sweden	1995	Inkomstfördelningsundersokningen	Statistics Sweden Program for Income and Wealth	16260	National
Taiwan	1995	Survey of Personal Income Distribution	Academia Sinica	14706	National
Uruguay	1996	Encuesta Continua de Hogares	Instituto Nacional de Estadistica	19322	Urban
United States	1994	March Current Population Survey	Bureau of Labor Statistics	66014	National
Venezuela	1996	Encuesta de Hogares por Mustreo	Oficina Central de Estadistica e Informatica	16323	National

TABLE XIV
EDUCATION THRESHOLDS

Country	Years of schooling beyond which a person qualifies as skilled	The LIS education level beyond which a person qualifies as skilled
Australia	12	Basic/skilled vocational qualification
Argentina	12	N/A
Belgium	12	2nd level upper professional/ technical/general; other 2nd level upper
Bolivia	12	N/A
Brazil	11	N/A
Britain	13	A Level
Canada	12	Grade 11–13; high school grad.
Chile	11	N/A
Colombia	11	N/A
Costa Rica	11	N/A
Czech Republic	12	Secondary general/professional
Denmark	10	Level 2, 2nd stage
Ecuador	12	N/A
Finland*	12	N/A
France	12	Second stage of secondary
Germany	10	Secondary
Hungary	12	Secondary
Israel*	12	N/A
Italy	12	High school
Luxembourg	12	Higher secondary education
Mexico	12	N/A
Netherlands	12	Secondary higher
Norway*	12	N/A
Panama	12	N/A
Paraguay	12	N/A
Peru	11	N/A
Poland	12	Complete secondary
Slovakia	12	Secondary/secondary special/ skilled with leaving exam
Spain	12	Secondary education/basic tech. edu.
Sweden	12	Secondary school
Taiwan	10, 12	Senior high/vocational graduate
Uruguay	11	N/A
United States	12	High school diploma
Venezuela	11	N/A

* Finland, Israel, and Norway report years of education.

sources, and to reconcile our statistics with the previously published numbers.

NEW YORK UNIVERSITY, LONDON SCHOOL OF ECONOMICS, CENTRE FOR ECONOMIC POLICY RESEARCH, AND NATIONAL BUREAU OF ECONOMIC RESEARCH
PENNSYLVANIA STATE UNIVERSITY
UNIVERSITY OF PENNSYLVANIA

REFERENCES

Aiyagari, S. Rao, Jeremy Greenwood, and Nezih Guner, "On the State of the Union," *Journal of Political Economy,* CVIII (2000), 213–244.
Alesina, Alberto, Reza Baqir, and William Easterly, "Public Goods and Ethnic Divisions," *Quarterly Journal of Economics,* CXIV (1999), 1243–1284.
Barro, Robert J., and Jong-Wha Lee, "International Data on Educational Attainments: Updates and Implications," NBER Working Paper No. 7911, 2000.
Becker, Gary S., "A Theory of Marriage: Part I," *Journal of Political Economy,* LXXXI (1973), 813–846.
Becker, Gary S., and N. Tomes, "An Equilibrium Theory of the Distribution of Income and Intergenerational Mobility," *Journal of Political Economy,* LXXXVII (1979), 1153–1189.
Becker, Gary S., and N. Tomes, "Human Capital and the Rise and Fall of Families," *Journal of Labor Economics,* IV (1986), S1–S39.
Bénabou, Roland, "Heterogeneity, Stratification and Growth," *American Economic Review,* LXXXVI (1996), 584–609.
Bergstrom, Theodore C., "A Survey of Theories of the Family," in M. R. Rosenzweig and O. Stark, eds., *Handbook of Population and Family Economics* (Amsterdam, The Netherlands: Elsevier Science B.V., 1997), pp. 21–79.
Bils, Mark, and Peter J. Klenow, "Does Schooling Cause Growth?" *American Economic Review,* XC (2000), 1160–1183.
Boulier, Bryan L., and Mark R. Rosenzweig, "Schooling, Search, and Spouse Selection: Testing Economic Theories of Marriage and Household Behavior," *Journal of Political Economy,* XCII (1984), 712–732.
Burdett, Kenneth, and Melvyn G. Coles, "Marriage and Class," *Quarterly Journal of Economics,* CXII (1997), 141–168.
Burdett, Kenneth, and Melvyn G. Coles, "Transplants and Implants: The Economics of Self-Improvement," *International Economic Review,* XLII (2001), 597–616.
Cole, Harold L., George J. Mailath, and Andrew Postlewaite, "Social Norms, Savings Behavior and Growth," *Journal of Political Economy,* L (1992), 1092–1125.
Dahan, Momi, and Alejandro Gaviria, "Sibling Correlations and Intergenerational Mobility in Latin America," *Economic Development and Cultural Change,* XLIX (2001), 537–554.
Dahan, Momi, and Daniel Tsiddon, "Demographic Transition, Income Distribution and Economic Growth," *Journal of Economic Growth,* III (1998), 29–52.
Durlauf, Steven, "A Theory of Persistent Income Inequality," *Journal of Economic Growth,* I (1995), 75–93.
Easterly, William, "The Middle Class Consensus and Economic Development," *Journal of Economic Growth,* VI (2001), 317–336.
Fernández, Raquel, "Education, Segregation, and Marital Sorting: Theory and an Application to the UK," *European Economic Review,* XLVI (2001), 993–1022.
Fernández, Raquel, Nezih Guner, and John Knowles, "Love and Money: A Theoretical and Empirical Analysis of Household Sorting and Inequality," NBER Working Paper No. 8580, 2001.
Fernández, Raquel, and Richard Rogerson, "Public Education and Income Distribution: A Dynamic Quantitative Evaluation of Education Finance Reform," *American Economic Review,* LXXXVIII (1998), 813–833.

Fernández, Raquel, and Richard Rogerson, "Sorting and Long-Run Inequality," *Quarterly Journal of Economics,* CXVI (2001), 1305–1341.
Flug, Karnit, and Zvi Hercowitz, "Equipment Investment and the Relative Demand for Skilled Labor: International Evidence," *Review of Economic Dynamics,* III (2000), 461–485.
Galor, Oded, and David N. Weil, "The Gender Gap, Fertility and Growth," *American Economic Review,* LXXXVI (1996), 374–387.
Galor, Oded, and Joseph Zeira, "Income Distribution and Macroeconomics," *Review of Economic Studies,* LX (1993), 35–52.
Ghez, Gilbert R., and Gary S. Becker, *The Allocation of Time and Goods over the Life Cycle* (New York, NY: Columbia University Press for NBER, 1975).
Gould, Eric D., and M. Daniele Paserman, "Waiting for Mr. Right: Rising Inequality and Declining Marriage Rates," *Journal of Urban Economics,* LIII (2003), 257–281.
Greenwood, Jeremy, Nezih Guner, and John Knowles, "More on Marriage, Fertility and the Distribution of Income," *International Economic Review,* XLIV (2003), 827–862.
Heckman, James J., "The Incidental Parameters Problem and the Problem of Initial Conditions in Estimating a Discrete Time-Discrete Data Stochastic Process," in C. F. Manski and D. McFadden, eds., *Structural Analysis of Discrete Data with Econometric Applications* (Cambridge, MA: The MIT Press, 1981), pp. 179–195.
Hess, Gregory D., "Marriage and Consumption Insurance: What's Love Got to Do with It?" *Journal of Political Economy,* CXII (2004), 290–318.
Heston, Alan, Robert Summers, and Bettina Aten, *Penn World Table Version 6.1,* Center for International Comparisons at the University of Pennsylvania (CICUP), 2002.
Knowles, John, "Can Parental Decisions Explain U. S. Income Inequality?" Ph.D. thesis, University of Rochester, 1999.
Kremer, Michael, "How Much Does Sorting Increase Inequality?" *Quarterly Journal of Economics,* CXII (1997), 115–139.
Kremer, Michael, and Daniel Chen, "Income Distribution Dynamics with Endogenous Fertility," *American Economic Review Papers and Proceedings,* LXXXIX (1999), 155–160.
Laitner, John, "Household Bequests, Perfect Expectations, and the National Distribution of Wealth," *Econometrica,* XLVII (1979), 1175–1194.
Lam, David, "Marriage Markets and Assortative Mating with Household Public Goods: Theoretical Results and Empirical Implications," *Journal of Human Resources,* XXIII (1988), 462–487.
Ljungqvist, Lars, "Economic Development: The Case of a Missing Market for Human Capital," *Journal of Development Economics,* XL (1993), 219–239.
Loury, Glenn C., "Intergenerational Transfers and the Distribution of Earnings," *Econometrica,* XLIX (1981), 843–867.
Mare, Robert D., "Five Decades of Educational Assortative Mating," *American Journal of Sociology,* LVI (1991), 15–32.
OECD, *Education at a Glance: OECD Indicators 2002* (Paris, France: OECD, 2002).
Oskarsson, Sven, "Class Struggle in the Wake of Globalization: Union Organization and Economic Integration," mimeo, University of Uppsala, 2002.
Owen, Ann, and David N. Weil, "Intergenerational Earnings Mobility, Inequality and Growth," *Journal of Monetary Economics,* XLI (1998), 71–104.
Pissarides, Christopher A., "From School to University: The Demand for Post-Compulsory Education in Britain," *Economic Journal,* XCII (1982), 654–667.
Rama, Martin, and Raquel Artecona, "A Database of Labor Market Indicators across Countries," The World Bank, Development Research Group, 2000.
Regalia, Ferdinando, and José-Víctor Ríos-Rull, "What Accounts for the Increase in Single Households and the Stability in Fertility?" mimeo, University of Pennsylvania, 1999.
Smits, Jeroen, Wout Ultee, and Jan Lammers, "Educational Homogamy in 65 Countries: An Explanation of Differences in Openness Using Country-Level Explanatory Variables," *American Sociological Review,* LXIII (1998), 264–285.

Szekely, Miguel, and Marianne Hilgert, "What Drives Differences in Inequality across Countries?" Inter-American Development Bank, Working Paper No. 439, 2000.

United Nations Development Programme, *Human Development Report 2003* (New York, NY: Oxford University Press for the United Nations Development Programme, 2003).

Weiss, Yoram, "The Formation and Dissolution of Families: Why Marry? Who Marries Whom? and What Happens upon Divorce?" in M. R. Rosenzweig and O. Stark, eds., *Handbook of Population and Family Economics* (Amsterdam, The Netherlands: Elsevier Science B.V., 1997), pp. 81–123.

World Bank, *Engendering Development Through Gender Equality in Rights, Resources, and Voice* (New York, NY: Oxford University Press for The World Bank, 2001a).

World Bank, *World Development Indicators 2001* (Washington, DC: The World Bank, 2001b).