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International Migration, Sex Ratios, and the Socioeconomic Outcomes of Nonmigrant Mexican Women

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Abstract This article assesses whether international migration from Mexico affects the marital, fertility, schooling, and employment outcomes of the Mexican women who remain behind by exploiting variation over time as well as across Mexican states in the demographic imbalance between men and women. I construct a gauge of the relative supply of men for women of different age groups based on state-level male and female population counts and the empirically observed propensity of men of specific ages to marry women of specific ages. Using Mexican census data from 1960 through 2000, I estimate a series of models in which the dependent variable is the intercensus change in an average outcome for Mexican women measured by state and for specific age groups, and the key explanatory variable is the change in the relative supply of men to women in that state/age group. I find that the declining relative supply of males positively and significantly affects the proportion of women who have never been married as well as the proportion of women who have never had a child. In addition, states experiencing the largest declines in the relative supply of men also experience relatively large increases in female educational attainment and female employment rates. However, I find little evidence that women who do marry match to men who are younger or less educated than themselves.

Keywords Migration · Mexico · Fertility · Human capital

Introduction

Between 1970 and 2007, the foreign-born Mexican population residing in the United States increased more than 14-fold, from approximately 820,000 to 11.9 million. Over the comparable period, the resident population of Mexico increased 2.2 times,

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from 50.6 to 108.7 million. Thus, net Mexico-U.S. migration has increased well beyond what would be expected based on Mexican population growth alone.

Mexican migrants are hardly a cross section of the national Mexican population. In particular, the migrant population is disproportionately male. In 2007, the ratio of males to females among Mexican nationals residing in the United States was roughly 1.43 among those aged 16–20, 1.56 among those aged 21–25, and 1.49 for those aged 26–30.

The large and disproportionately male migratory flow from Mexico to the United States has lowered the ratio of males to females in Mexico, especially among those who are of prime working age. The consequent increasing relative scarcity of men may affect the behavioral choices of nonmigrant Mexican women along a number of dimensions. Young women may delay marriage and childbearing, invest more in schooling to ensure future economic self-sufficiency, and increase participation in the formal labor markets. Alternatively, the scarcity of men may induce women to marry men to whom they are less suitably matched.

Many have studied the impacts of immigration on various outcomes in the receiving country, such as labor markets (Borjas 2003; Card 2001, 2005; Ottaviano and Peri 2007), economics assimilation (Borjas 1995; Lubotsky 2007; Raphael and Smolensky 2009), crime (Butcher and Piehl 2007), and public expenditures (Smith and Edmonston 1997). Less effort has been devoted to studying the effects of mass migration on the sending country. In this article, I assess whether migration from Mexico impacts the marital, fertility, schooling, and employment outcomes of the Mexican women who remain behind.

I exploit variation over time and across Mexican states in the imbalance between men and women. The contribution of individual Mexican states to northern migration is quite uneven. For example, in 2000, less than 1 % of households in Campeche had a migrant abroad, compared with more than 15 % of households in Michoacán. Central Mexican states and a few northern states have historically contributed disproportionate numbers of migrants. Consequently, there is considerable heterogeneity across states in the time path of sex ratios.

I measure the relative supply of men for women of different age groups based on state-level male and female population counts. Potential male spouses are allocated across female age groups based on the empirically observed propensity of men of specific ages to marry women of specific ages. Using Mexican census data covering 1960 through 2000, I estimate a series of models in which the dependent variable is the intercensus change in an average outcome for Mexican women, measured by state and for specific age groups, and the key explanatory variable is the change in the corresponding relative supply of men. To address possible bias from selective out-migration of women in response to the scarcity of men, I also present results of analyses in which the supply measure is instrumented using a similar gauge calculated based on one's state of birth.

I find that the declining relative supply of males positively and significantly affects the proportion of women who have never been married as well as the proportion of women who have never had a child. In addition, states experiencing the largest declines in the relative supply of men also experience relatively large increases in female educational attainment and female employment rates. I find little evidence that women who do marry match to men who are relatively younger or less educated.

The Effect of the Relative Supply of Men on Female Socioeconomic Outcomes

There is a long history in sociology and economics of modeling the process by which men and women match to one another, marry, and/or have children in the context of search theory. In such a framework, those searching for spouses receive opportunities at irregular intervals, face considerable uncertainty regarding actual and future character traits of potential suitors, and must form some minimum acceptable standards in a complex multivariate manner that may or may not allow for substitutability along alternative dimensions (see, e.g., the seminal work of Becker 1991; Lichter et al. 1995; and Oppenheimer 1988).

Heterogeneous men and women seek out potential mates in a world where search is costly and new prospects present themselves in some time-delayed manner. From the female perspective (or the male perspective, for that matter), prospective partners present themselves at a rate that increases with the relative supply of men. Assuming that prospective spouses can be ranked according to some gauge of quality, a woman searching for a partner will have a minimum quality standard below which a prospect will be rejected. For a prospect that clears the threshold, a match could form.

Although certainly simplistic, this framework is useful for thinking through how declining Mexican sex ratios may impact nonmigrant Mexican women. Gender-biased migration reduces the relative supply of men, which in turn should reduce the rate at which women encounter prospective suitors. For a given reservation-quality threshold, a specific woman must search longer to find a suitable spouse, thus delaying time until marriage. Alternatively, women searching for spouses may lower their standards and on average marry lower-quality men, a derivative behavioral response that will at least partially offset the direct effect of male scarcity on the likelihood of marriage.[1]

In addition, changes in the relative supply of men may also impact the gender balance of power within marriages and have consequent implications for female expectations that may affect educational and career choices. Chiappori et al. (2002) presented a theoretical model of household decision making in which the bargaining position of spouses within the household is influenced by "distribution factors" external to the household. Distribution factors—such as the ease with which one can divorce, or the ratio of men to women—influence the fallback position of each spouse should the marriage dissolve and thus determine bargaining power over any welfare surplus generated within the marriage. South and Lloyd (1995) indeed presented empirical evidence for the United States indicating that marriages are more likely to dissolve in regions where the availability of alternative partners is particularly high. Hence, decision making within formed and officially sanctioned unions regarding human capital investment, specialization within the household, and also fertility are likely to be made with an eye on the likely stability of the union.

[1] It is conceivable that the indirect effect on the likelihood of marriage of lowering one's standards may overwhelm the negative direct effect of a decline in male availability operating through the offer arrival rate. This would be the case if a slight drop in standards greatly increases the pool of available men. Of course, this will depend on the actual form of the empirical quality distribution of male suitors. Regardless, these opposing direct and indirect effects suggest that in theory, the effect of relative male scarcity on female marriage probabilities can go in either direction. As I will soon show, however, most studies reviewed in this article have found that mate availability positively covaries with female marriage rates.

A scarcity of male suitors may also improve the bargaining position of men when it comes to negotiating personal relationships outside marriage. For example, men may find it easier to find sexual partners, may have the upper hand in negotiations over whether and which birth control to use, and may be required to demonstrate less loyalty in personal relationships when they are scarce relative to women. Indeed, researchers have found significant inverse relationships between sex ratios and the rates of teen pregnancy (Sampson 1995), syphilis (Kilmarx et al. 1997), and gonorrhea (Thomas and Gaffield 2003).

One's marriage prospects are clearly contemporaneously and dynamically related to one's labor supply and human capital choices. The expectation of a lengthier period until marriage, or perhaps an increased probability of never marrying, is likely to induce women to make human capital investments that will ensure their future economic self-sufficiency (Becker 1991). When men are relatively scarce, it would be rational to invest more in formal education and to acquire experience in the formal labor market.

Finally, the relative supply of males may affect fertility through several channels. First, to the extent that marriage is a social precondition for childbearing, poorer marriage prospects may result in both lower marriage rates as well as lower age-adjusted fertility. Second, women in marriage markets with terms of trade that are relatively unfavorable toward women may be in less-secure relationships and, as a result, may be less willing to have children either within or outside marriage. Such effects may surface in either lifetime fertility or the proportion of women who have never had a child.

A host of studies of the United States have sought to test the theoretical predictions of search theory as applied to the market for marriage. Prompted in part by William Julius Wilson's declining marriageability hypothesis (Wilson 1987), much research has been devoted to understanding whether a paucity of men explains the relatively poor marital outcomes for African American women. Using metropolitan area level data from the 1980 census, South and Lloyd (1992) found positive partial correlations in cross-sectional data between mate availability and the proportion of women who are married for white women but not for black women. Using data from the 1980 census matched to the 1979 National Longitudinal Survey of Youth, South (1996) found that mate availability elevates the marriage hazard rate for young women as well as the probability of nonmarital childbearing for white women. For black women, however, the gauge of mate availability affects only the likelihood of an out-of-wedlock birth. The differential results for black women are particularly puzzling, since black women are considerably more likely to marry within race than white women and, thus, historically have not cast a wider net across racial and ethnic groups (Rosenfeld 2008).

Exploiting variation across ethnic/racial/age groups within a single large metropolitan area (Los Angeles), Catanzarite and Ortiz (2002) found that never-married women in demographic groups with a relative scarcity of men are more likely to have had children relative to never-married women in demographic groups among which men are relatively abundant. Fossett and Kiecolt found strong associations between male socioeconomic status (SES), female marriage prevalence, and family structure among African Americans using both cross-metropolitan area variation (Fossett and Kiecolt 1993) as well as variation across counties within a single southern state (Fossett and Kiecolt 1990).

All studies reviewed thus far rely on cross-sectional geographic variation, cross-sectional interracial and interethnic variation, or both sources of cross-sectional variation in mate availability to estimate the relationship between marriage market conditions and female outcomes. Although all these studies make attempts to control for possible confounding variables, such efforts are limited to what can be measured with census data and other large microdata surveys. One might certainly contend that geographic variation in sex ratios may be correlated with unobservable factors that affect both sex ratios and the female outcomes of interest. This is certainly a plausible alternative interpretation in studies that rely on cross-racial and ethnic variation. In an attempt to methodologically address omitted variables bias, a recent analysis of marital outcomes by Charles and Luoh (2010) for African American women exploited variation occurring within demographic groups over time in marriage market conditions. Moreover, the authors identified the specific institutional source of variation in the change in mate availability over time.

Charles and Luoh analyzed the effect of variation in male incarceration rates on the marital outcomes of African American women in the United States. Using data from several years of the decennial census, the authors estimated average marital outcomes that vary by race, age, and state, and tested for an effect of male incarceration rates relying solely on variation occurring within demographic/state groups over time. The authors found that in states experiencing particularly large increases in black male incarceration rates, women are less likely to marry and more likely to marry men who are less educated than themselves. To be sure, cross-group variation in the change in incarceration rates may itself be endogenous. For example, if demographic groups experiencing large increases in incarceration experience a contemporaneous worsening in employment prospects or a differential change in illicit behavior either economically motivated or otherwise, marital outcomes as well as male incarceration rates may simultaneously worsen. However, the use of within-group change in marriage market conditions and adult female outcomes does effectively control for any factors that are specific to demographic groups and that affect a time-invariant effect on the outcome in question.

Several recent studies have analyzed the effect of the ratio of men to women on the marital and economic outcomes of women in a number of alternative social and historical contexts where there are clear identifiable sources of exogenous variation in sex ratios. Angrist (2002) tested for an effect of immigrant-induced variation in sex ratios in the United States on the marital outcomes of the second-generation U.S. children of immigrants. Based on an analysis of 11 ethnic groups defined by country of origin with a large immigrant presence in the United States, the author found that second-generation women are more likely to marry, less likely to work, and more likely to have high incomes when immigration increases the sex ratio of their ethnic group. Abramitzky et al. (2011) assessed the impact of geographic variation in sex ratios across French regions caused by variation in World War I casualties on the marital outcomes of men and women. The authors found that in those regions suffering the most war casualties, men were more likely to marry women from a higher social class than their own.

Rao (1993) presented a particularly novel assessment of the impact of marriage market conditions on the surplus that women derive from marriage. The author investigated the determinants of dowry payments from the family of the bride to the family of the groom for a set of villages in central India. Rao found an association between the relative supply of women and the size of the dowry, with a glut of women associated with higher dowries.

One particularly relevant study analyzed whether marriage market conditions affects the quality of existing relationships. Harknett (2008) matched data from the U.S. census to data from the multicity Fragile Families and Child Wellbeing Study (hereafter, Fragile Families study) to analyze the effects of sex ratios on the quality of relationships among couples who have had children out of wedlock. Although Harknett's analysis relied on cross-sectional variation in mate availability and thus is subject to the critique of such studies raised earlier, the detailed data in the Fragile Families study does permit controlling for many factors beyond the usual set of covariates from census data and also permits the researcher to analyze several novel outcome variables.

Harknett found little evidence of an effect of marriage market conditions on the labor market prospects of mates, but did find a significant effect of marriage market conditions favorable to women on the likelihood that male partners are at least as educated as the mothers of their children. Harknett also found that high sex ratios are positively associated with various gauges of relationship quality, including measures of mutual supportiveness, lack of within-couple conflict, and whether the father visits the mother in the hospital after birth. Finally, Harknett found that in areas with higher sex ratios, the probability that relationship results in a marriage after the birth of a child is higher.

As I demonstrate in the following section, migration between Mexico and the United States has certainly lowered the ratio of men to women, consequently shifting the "terms of trade" in the Mexican marriage market decisively in favor of Mexican men. Whether this change has affected average outcomes for Mexican women is an empirical question to which I now turn.

Data Description and the Link Between International Migration and Sex Ratios

The analysis presented in this article is based on microdata from the 1960 (1.5 %), 1970 (1 %), 1990 (10 %), and 2000 (10.5 %) Mexican censuses. With the exception of the 1960 census, which was drawn from a sample of individuals, all censuses are based on household samples and permit matching between spouses within households, which is an important feature for constructing the relative supply of males index in subsequent sections. Data from census year 1980 is unavailable because large portions of the database from that year were destroyed in the 1985 Mexico City earthquake (Rabell 2001). All data for this project were downloaded from the Integrated Public Use Microdata Series (IPUMS) International webpage at the University of Minnesota (http://www.ipums.org/).

The population of Mexican nationals residing in the United States is disproportionately working age and male. Moreover, this migrant population has grown considerably during the past four decades. Here, I document the consequent effect on the ratio of men to women in Mexico and explore the relationship within Mexico between the proportion of households that send migrants abroad and the relative supply of men.

Figure 1 presents the ratio of males to females for five-year age groups in each census year between 1960 and 2000. Even in 1960, the ratio of males to females declined between the age groups of 11–15 and 21–25, which is likely reflective of the presence of migrant men in the United States and elsewhere. From 1960 to 2000, however, these ratios declined further, especially for age groups older than 20. For example, the ratio of men to women aged 21–25 declined from 0.93 to 0.88 between 1960 and 2000. Among those aged 31–35, the sex ratio declined from 0.98 to 0.90.

The changing sex ratios in Mexico mirror changes in the sex ratio among Mexican nationals residing in the United States. Figure 2 presents comparable sex ratios for the Mexican-born Mexican population in the United States tabulated from the 1970 and 1990 U.S. Census of Population and Housing as well as from the 2007 American Community Survey.[2] The 1970 census reveals a fair degree of balance between Mexican-born men and women residing in the United States for all age groups. In 1990 and 2007, however, there were notable increases (to more than 1.2, but in some instances more than 1.4) in the ratio of men to women, especially among those in their 20s.

One can ascertain the relative importance of migration in explaining low sex ratios in Mexico by combining data from the United States and Mexico. Using data from both countries, I calculate sex ratios within each age group when combining Mexican nationals living in the United States and Mexicans living in Mexico, and then compare these ratios with those using only the resident population of Mexico. The results from this exercise are presented in Fig. 3. For 1970 and 2000, the figure presents the difference between the hypothetical sex ratio combining the U.S. foreign-born Mexican and Mexican populations and the sex ratio using the Mexico population only. Interestingly, in 1970, migration to the United States had little effect on sex ratios in the home country (as can be seen by the negligible differences for this year in 1970). In 2000, however, migration to the United States contributed considerably to lower sex ratios. Specifically, Fig. 3 shows that the sex ratio for those in their 20s and early 30s would be higher by roughly 0.05 if all Mexican nationals returned to Mexico from the United States.

Unfortunately, the Mexican census does not contain detailed information on migrants abroad; thus, it is not possible to depict the gender composition of the migratory outflow. Moreover, the early census years do not contain information on whether the household has a family member living abroad. However, the 2000 Mexican census does include a question at the household level inquiring whether the household currently has one of its members residing in a foreign country. With this variable, it is possible to estimate a migration rate defined specifically as the proportion of households with a migrant abroad. In conjunction with some simple theorizing, the variation within Mexico in the proportion of households with a migrant abroad can be used to establish the relationship between migration and sex ratios.

To the extent that migration is gender-biased, one should observe a negative correlation between a state's overall migration rate and the state's sex ratio. To formally identify the source of this correlation, I define M_i^b as the population total at the time of

[2] The microdata for these tabulations were downloaded from the IPUMS webpage at the University of Minnesota (http://www.ipums.org/).

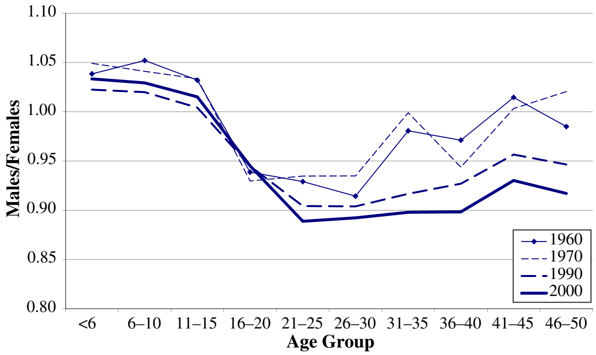

Fig. 1 Ratio of males to females in Mexico by age: 1960, 1970, 1990, and 2000

birth for a given male birth cohort in state i, and M_i^c as the current population of males from this birth cohort in state i. Assume that the difference between these two totals, $M_i^m = M_i^b - M_i^c$, is due entirely to international migration out of the state. Define the comparable totals, W_i^b, W_i^c, and W_i^m, for the corresponding birth cohort of women in state i. Assume for the moment that only men migrate internationally. Under this assumption, one can derive the following linear relationship between the sex ratio of the specific birth cohort and the overall migration rate as follows:

$$\begin{aligned}\frac{M_i^c}{W_i^c} &= \frac{M_i^b - M_i^m}{W_i^b}\\ \frac{M_i^c}{W_i^c} &= \frac{M_i^b}{W_i^b} - \frac{M_i^m/(M_i^b+W_i^b)}{W_i^b/(M_i^b+W_i^b)}\\ \frac{M_i^c}{W_i^c} &= \frac{M_i^b}{W_i^b} - \left[\frac{M_i^b}{W_i^b}+1\right] migration_i,\end{aligned} \tag{1}$$

where $migration_i$ is the total migration rate for the cohort under the assumption that only males migrate. Equation (1) shows that with perfectly gender-imbalanced migration, the current sex ratio for a given birth cohort in state i equals the sex ratio at birth minus 1 plus the sex ratio at birth times the overall migration rate. Assuming that males and females are born in proportion to one another,[3] Eq. (1) implies that a regression of sex ratios on the overall migration rate should yield a slope coefficient of approximately −2.

Similarly, it is easy to show that with gender-balanced migration, the current sex ratio should not depend on the migration rate. Specifically, assuming that both males and females of a given birth cohort leave state i at rate $migration_i$, the following condition will hold:

$$\frac{M_i^c}{W_i^c} = \frac{M_i^b - migration_i M_i^b}{W_i^b - migration_i W_i^b} = \frac{M_i^b(1-migration_i)}{W_i^b(1-migration_i)} = \frac{M_i^b}{W_i^b}. \tag{2}$$

[3] The natural odds ratio of a male birth is slightly greater than 1, with sex ratios at birth without selective aborting around 1.05 (Almond and Edlund 2008).

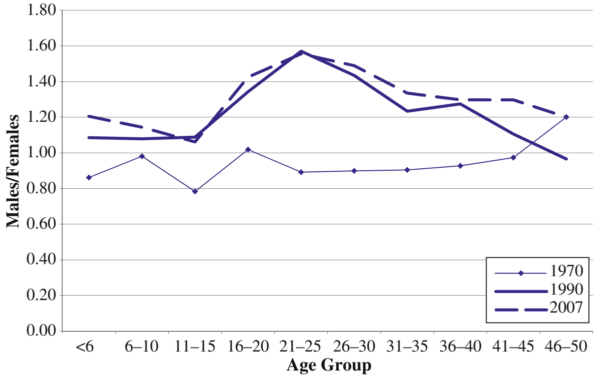

Fig. 2 Ratio of males to females by age among the Mexican-born population residing in the United States: 1970, 1990, and 2007. Sex ratios for 1970 and 1990 are tabulated using data from the 1 % Public Use Microdata Sample of the 1970 and 1990 U.S. Census of Housing and Population. The 2007 sex ratios are tabulated using data from the 2007 American Community Survey

That is to say, with gender-balanced migration, the current sex ratio will equal the sex ratio at birth. Consequently, a regression of state sex ratios against the state migration rate should yield a slope coefficient of zero.

Figures 4 and 5 present scatterplots of sex ratios measured at the state level in 2000 against the proportion of households in each state with a migrant abroad. Figure 4 presents a scatter for those aged 20–25, and Fig. 5 presents a similar scatter for those aged 31–35. The figures reveal a great degree of variation across states in the male-to-female ratio within the age groups depicted. Among those aged 31–35, the sex ratio varies from below 0.85 to nearly 1.05, and the values of this ratio among states for those aged 20–25 range from below 0.8 to more than 1. There is also considerable heterogeneity across states in the proportion of households with a migrant abroad (with a range from less than 1 % of households to more than 15 % of households).[4]

Figures 4 and 5 also reveal a strong inverse relationship between each sex ratio and the proportion of households with a migrant abroad. In Fig. 4, the slope coefficient on the line fit through the data cloud equals −0.901 and is significant at the 1 % level of confidence. The strength of this relationship is reflected in the relatively high R^2 from this bivariate scatterplot (0.439). The slope coefficient for the regression line in Fig. 5 is −0.921 and is also highly

[4] The cross-state variation in emigration rates is a fascinating phenomenon in and of itself. A recent thorough empirical analysis of cross-state and cross-cohort migration rates finds that both economic push factors as well as network effects play important roles in determining the cross-state variation in migration rates depicted in Figs. 4 and 5. Hanson and McIntosh (2010) showed that interdecade migration rates from Mexican states between 1960 and 2000 depend positively on the size of particular birth cohorts (consistent with a labor market supply push argument) and negatively on initial per capita gross domestic product (GDP) and growth in per capita GDP. The authors also found substantial heterogeneity in these effects for states that are traditional sending states relative to states with lower historical migration rates. Specifically, they showed that these economic push and pull factors matter more in states with relatively high emigration rates in 1924 and states that are in close proximity to the main railroad passing through the country to the U.S. border (a transportation hub for recruiters of Mexican workers during the pre–World War II period (Cardoso 1980)). Independent of these factors, proximity to the U.S. border is not a particularly strong predictor of emigration rates.

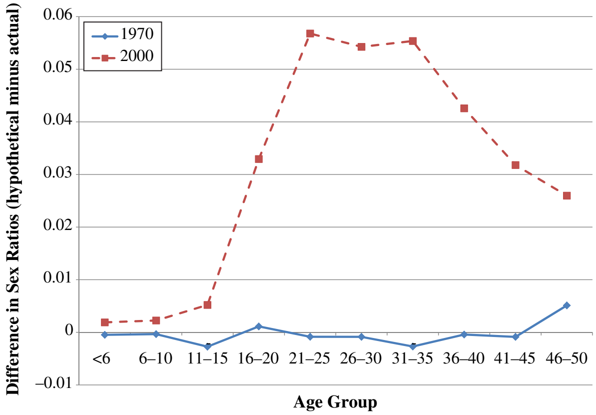

Fig. 3 Difference between age-specific sex ratios (males to females) for all Mexican nationals (residing in the United States and Mexico combined) and for Mexican nationals residing in Mexico, 1970 and 2000

significant, with the migration rate variable explaining slightly more than one-half the cross-state variation in sex ratios. Although these slope coefficients fall considerably short of −2, they are certainly statistically distinguishable from zero and strongly indicate that gender-biased international migration is altering Mexican sex ratios in a geographically concentrated manner. Hence, changes in Mexican sex ratios over time, the decline in sex ratios among the resident population of Mexico relative to those for the overall population of Mexican nationals (United States and Mexico combined), and the strong inverse cross-sectional correlation between sex ratios and migration rates all indicate that northern migration has lowered the relative availability of men in Mexico.[5]

Estimation Strategy

I test for an effect of the changing relative availability of Mexican men on several socioeconomic outcomes for women. Specifically, I test for an effect on the proportion who have never married, the proportion without children, the proportion who have never married but have borne children, average educational attainment, the proportion enrolled in school, and the proportion working. I also test for measures of within-marriage mismatch for formed unions. In particular, I analyze whether variation in the relative supply of men affects average male–female age differences within marriage as well as the proportion of women marrying younger men. I also analyze similar outcomes regarding within-marriage educational disparities.

[5] It is also the case that the low sex ratios beyond age 15 observed in Mexico are unique relative to other Latin American countries. In an earlier working paper version of this study, I presented a comparison of Mexican sex ratios by age to those for four other Spanish-speaking Latin American countries. In all comparisons, Mexican sex ratios among those over age 15 are relatively lower, with particularly large differences relative to Panama, Costa Rica, and Venezuela, and more modest differences relative to Colombia. This is not particularly surprising given the relatively low rate of migration from these alternative countries to the United States.

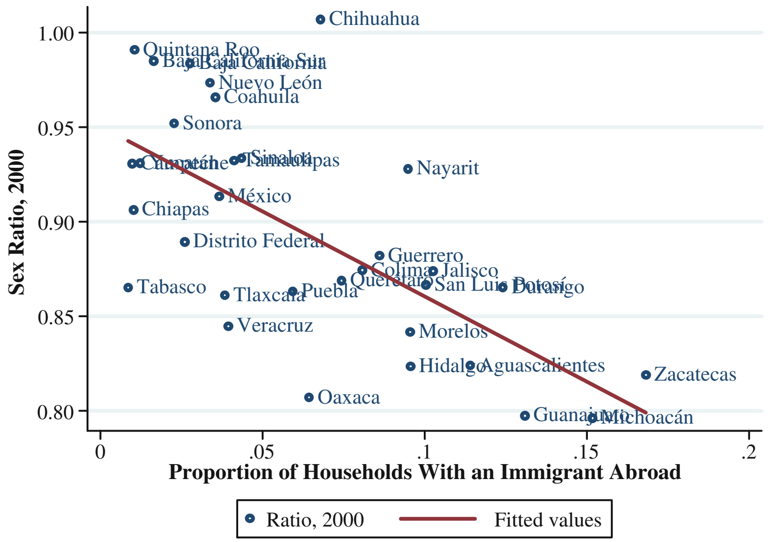

Fig. 4 Scatterplot of the 2000 ratio of males to females by state among those aged 20–25 against the 2000 proportion of households with a migrant residing abroad

The methodological strategy I employ in this article involves exploiting the cross-state variation in changes in the relative supply of men. Specifically, let c index three age groups (16–19, 20–25, and 26–30), s index the 32 Mexican states, and t index time. Define the variable $rsupply_{cst}$ as the relative supply of men to women in age group c in state s in year t, and $\Delta rsupply_{cst}$ as the

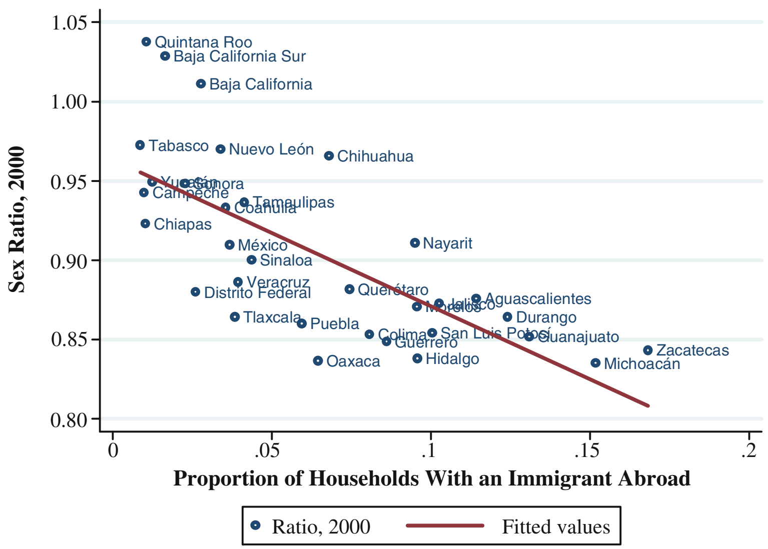

Fig. 5 Scatterplot of the 2000 ratio of males to females by state among those aged 31–35 against the 2000 proportion of households with a migrant residing abroad

change in the relative supply of men between periods $t-1$ and t. My estimates of the impact of the relative supply of men on outcomes for Mexican women derive from estimation of the following equation:

$$\Delta Outcome_{cst} = \alpha_c + \psi_t + \delta_s + \beta \Delta rsupply_{cst} + \varepsilon_{cst}, \tag{3}$$

where $\Delta Outcome_{cst}$ measures the change in an outcome variable between period $t-1$ and t measured at the state/age group level, α_c is an age-specific intercept permitting linear time trends that vary by age group, ψ_t are time period fixed effects, δ_s is a state-specific intercept, β provides the key estimate of the marginal effect of the change in the supply of men on the outcome variable, and ε_{cst} is a mean-zero disturbance term.

Estimation of Eq. (3) relies on cross-state variation in the *change* in the relative supply of men. That is to say, I am assessing how within-state changes in outcome variables correlate with within-state changes in the key explanatory variable. By specifying Eq. (3) in first differences, average state-level variation in outcome levels are differenced out of the equation. Moreover, the inclusion of state-specific fixed effects in the change specification controls for linear time trends in the outcome levels that vary across states. The inclusion of age-specific fixed effects permits linear time trends in each outcome variable that may vary across age groups. Finally, the inclusion of year fixed effects permits the average changes to differ across census-year pairings. This is particularly important given that for one of my comparisons, the temporal change spans a 20-year period (in particular, the change between the 1970 and 1990 censuses).

The identifying assumption behind Eq. (3) is that the cross-state variation in the between-census-year changes in the relative supply of men is exogenous after the portion of variance in this explanatory variable that is attributable to variation across age groups, states, and census-year pairings is purged. Perhaps the key threat to the internal validity of the regression model in Eq. (3) comes from the possibility that women may internally migrate within Mexico in response to changes in marriage market conditions within their home state as well as in response to cross-state differences in marriage market conditions.

For example, suppose that women vary with respect to their desire to be married and/or to have children. In response to a decline in the relative supply of men, those women who are set on marrying may migrate to states with marriage market conditions more favorable to women. Such endogenous selection of women across states results in women with strong preferences for marriage locating in states with favorable marriage market conditions, and women less determined to marry in states with relatively poor conditions. Unobservable heterogeneity along this dimension of preferences may induce spurious correlation between variation in the relative supply of men and the outcome variable of interest. That is, average preference toward marriage should exert a negative impact on the proportion never married and be positively correlated with the relative supply of men. Thus, the omission of geographic variation in such preferences would exert a negative bias on the estimate of the relative supply of men on the proportion of women who have never married.

To address such concerns, I exploit the fact that in each census year, the Mexican census collects information on one's state of residence as well as one's state of birth. Variation in the relative supply of men in one's age group based on one's state of birth (and the state of birth of remaining nonmigrant Mexican men) will not be influenced by the internal migration of men and women. Shortly, I estimate the specification in Eq. (3) using (1) ordinary least squares (OLS) and the relative supply of men in one's current state of residence, and (2) instrumental variables in which the relative supply of men based on one's state of birth is used as an instrument for the relative supply of men in one's state of residence. Table 1 presents a series of regressions of the relative supply measure based on state of residence on the relative supply measure based on state of birth. In all three model specifications, the variable based on state of birth is a highly significant predictor of the relative supply measure based on state of residence.

To measure the relative supply of men to women of a specific age, I make use of the empirically observed propensity of married men of specific age groups to marry women of specific age groups. To be specific, I define $p_{c|i}$ as the conditional probability that a married man of age i is married to a woman of age c, where $\sum_c p_{c|i} = 1$. Let men_{ist} be the total population of men of age i in state s in year t. If the distribution of married men across the spousal age distribution is indicative of men's preferences, then the supply of men of age i to women of age c in a given state and year is given by $supply_{icst} = p_{c|i} \times men_{ist}$. The total supply of men to women in a specific state, age, and year group is found by summing over i, or $supply_{cst} = \sum_i supply_{icst}$. Finally, dividing this total supply measure by the relevant population count for women provides a gauge of the relative supply of men ($rsupply_{cst} = supply_{cst} / women_{cst}$).

I estimate the conditional probabilities $p_{c|i}$ using data on all married men in the 1990 census who can be matched to their spouses. A visual inspection of the 1990 census data reveals that although men are certainly more likely to marry women who are relatively close to them in age (although both distributions reveal a propensity to marry younger women), there is a fair degree of pairing outside narrow five-year age bands. Hence, the relative supply measure employed here should more precisely capture how the supply of men of a given age group is distributed across women of different ages. I calculate the relevant male and female populations at the state/age group level for all census years included in the analysis and then tabulate relative supply according to the formula previously cited.

I restrict the analysis to three age groups of women—ages 16–19, 20–25, and 26–30—because most first marriages, first births, and human capital investment choices take place prior to age 30. Thus, if the relative supply of men is affecting outcomes, one would expect to see an effect for these three age groups. Relative supply is tabulated for each single-year age group and then averaged within the broader age bands that form the age dimension of variation in my data set.

Finally, all models presented here are weighted by the average number of observations used to compute the intercensus change in the dependent variable.

Table 1 First-stage regressions of the relative supply of males based on state of residence on the relative supply of males based on state of birth

	Dependent Variable = Relative Supply of Males Based on State of Residence		
Relative Supply of Males	0.82	0.82	0.77
Based on Sate of Birth	(0.03)	(0.03)	(0.03)
Fixed-Effect Specification			
Year effects	Yes	Yes	Yes
Age effects	No	Yes	Yes
State effects	No	No	Yes

Notes: Standard errors, shown in parentheses, are computed allowing for clustering in each included age-state cell. Observations in the data vary across three age groups (16–19, 20–25, and 26–30), 32 Mexican states, and three intercensus changes (1960–1970, 1970–1990, and 1990–2000), giving a total of 288 observations. For the proportion employed and the proportion enrolled, comparisons do not include the 1960–1970 changes, yielding a total of 192 observations. Each model also controls for the average age within each cell. All models are weighted by the average of the number of observations used to compute the intercensus change.

In addition, in all models, I tabulate robust standard errors that are clustered by state and age groups.

Estimation Results

Table 2 presents results from estimation of Eq. (3) for each outcome variable. For each outcome, the model presents results of an OLS model in which the key explanatory variable is the relative supply of men based on state of residence as well as results of an instrumental variables (IV) model in which the relative supply of men based on state of birth is used as an instrument for the relative supply measure based on state of residence.[6] Within each set, I present estimates for three specifications: a model including year fixed effects; a model with year and age effects; and a model with year, age, and state fixed effects. In all regressions, the average age of women within each group is also included in the specification.

The empirical relationship between the proportion of women who have never been married and the relative supply of men is fairly robust across specifications. In both the OLS and IV models controlling for year-specific and age-specific fixed effects, the relative supply of males exerts a significant (at the 1 % level) and negative effect. In all estimates, the OLS and IV results are quite similar, and formal Hausman tests of the relative supply of men fail to reject the null hypothesis of exogeneity in all comparisons. In the final specification allowing for state linear time trends, the coefficients are attenuated somewhat yet are still statistically significant at the 10 % level of confidence. For the proportion of women who have never had a child, relative supply

[6] The relevant first-stage models are presented in Table 1.

Table 2 Regression model estimates of the impact of changes in the relative supply of men on changes in outcome variable measures for all women

Dependent Variables	(1)	(2)	(3)
Δ Proportion Never-Married			
OLS	−0.17 (0.05)**	−0.16 (0.05)**	−0.11 (0.06)†
IV	−0.17 (0.06)**	−0.16 (0.05)**	−0.12 (0.07)†
Δ Proportion With No Children			
OLS	−0.20 (0.05)**	−0.19 (0.05)**	−0.15 (0.06)**
IV	−0.20 (0.05)**	−0.19 (0.05)**	−0.19 (0.07)**
Δ Proportion Never-Married and With Children			
OLS	0.01 (0.01)	0.01 (0.01)	0.01 (0.01)
IV	0.00 (0.02)	0.01 (0.01)	0.00 (0.02)
Δ Proportion Enrolled in School			
OLS	0.23 (0.09)**	0.18 (0.06)**	0.17 (0.09)*
IV	0.23 (0.09)**	0.17 (0.05)**	0.14 (0.07)*
Δ Years of Educational Attainment			
OLS	−2.71 (1.28)*	−2.67 (1.29)*	−1.98 (2.07)
IV	−3.47 (1.20)**	−3.44 (1.20)**	−2.82 (1.91)
Δ Proportion Employed			
OLS	−0.25 (0.11)**	−0.21 (0.08)**	−0.26 (0.10)*
IV	−0.31 (0.11)**	−0.26 (0.09)**	−0.29 (0.10)**
Fixed-Effect Specification			
Year effects	Yes	Yes	Yes
Age effects	No	Yes	Yes
State effects	No	No	Yes

Notes: Standard errors, shown in parentheses, are computed allowing for clustering in each included age-state cell. The figure in each cell is the coefficient on the between-census change on the relative supply of males. Observations in the data vary across three age groups (16–19, 20–25, and 26–30), 32 Mexican states, and three intercensus changes (1960–1970, 1970–1990, and 1990–2000), giving a total of 288 observations. For the proportion employed and the proportion enrolled, comparisons do not include the 1960–1970 changes, yielding a total of 192 observations. Each regression also controls for the average age within each cell. All models are weighted by the average of the number of observations used to compute the intercensus change.

$^{\dagger}p < .10$; $^{*}p < .05$; $^{**}p < .01$

exerts a negative effect that is statistically significant at the 1 % level of confidence in all specifications.

The results in Table 2 reveal a statistically significant positive relationship between the relative supply of men and school enrollment, suggesting that young women are more likely to be enrolled in school when men are relatively abundant. This estimate is fairly consistent across model specifications and significant at either the 1 % or 5 % level of confidence in all models. Completed years of schooling is negatively associated with changes in the relative supply of men, suggesting that women complete more schooling when men are relatively scarce. Finally, there is a strong robust negative effect of the relative supply of men on the proportion of women who are working. The coefficient estimate is significant in all specifications, with little variance between models.[7]

To put the results into context, Table 3 presents the results from the following thought experiment. Suppose we were to transplant a young woman aged 20–25 from the central state of Michoacán (where the relative supply of men to women of this age group in 2000 has the relatively low value of 0.89) to the southern state of Quintana Roo (where the relative supply of men to women of this age group in 2000 is the relatively high value of 1.11). By how much would each of the analysis outcomes change? Moreover, how large would these effects be relative to the base level of each outcome for all Mexican women?

The first column of figures in Table 3 presents the coefficient from the most complete specification reported in Table 2 (those in the final column of figures) using coefficient—either the OLS or IV coefficient—with the smaller value for each outcome. The second column presents the difference in the relative supply of men between Quintana Roo and Michoacán. The third column presents the product of the figures in the first two columns, providing an estimate of the impact that the hypothetical move would have on each outcome. The final column characterizes the implied effect relative to the base level. For the proportion never-married, never had a child, employed, and enrolled, I use the average for all Mexican women aged 20–25 for 2000. For years of schooling, I use the overall average level of schooling for all women aged 14–50 as the base.

The results suggest relatively modest effects of variation in the supply of men for the proportion of women who have never married and have never had a child. Moving from the state with one of the lowest relative supplies to the state with one of the highest would decrease the proportion never-married by roughly 6 % relative to the observed base level. The comparable figure for the proportion who have never had a child is 7 %. The relative impact on school enrollment among this group is appreciably larger, with the move yielding a 20 % increase in enrollment. In addition, the implied impact on employment is a reduction of 15 %. Finally, the relative impact on years of schooling is a modest 6 %.

Tables 2 and 3 explore the relationship between the relative supply of men and average outcomes measured for all women regardless of marital status. However, several of the studies reviewed earlier find that the relative supply of men may also affect characteristics of marriages that actually form. For example, several studies

[7] The results regarding work and school enrollment are certainly consistent with one another. Women are less likely to work when men are relatively abundant and more likely to be enrolled in school, suggesting that the need to work displaced formal education among the women analyzed. However, I also find completed schooling levels to be higher on average when men are relatively scarce. Although these findings regarding school enrollment and educational attainment may seem at odds with one another, it is important to keep in mind that the median woman aged 14–50 in Mexico has nine years of completed schooling, while the comparable median for early years in this analysis period is even lower. Hence, most women in our sample (the minimum age considered is 16) are several years beyond the termination of their formal education. It is therefore possible to see both a response of average years of school completion as well as change in enrollment akin to what is presented here. With poor marriage market prospects, women may increase their educational attainment from primary to middle school levels. At the same time, one can see a reduction in secondary and postsecondary enrollment for some because of the need to work in the formal economy.

Table 3 Implied effect on select outcomes of moving a 20- to 25-year-old woman from Michoacán (where men are relatively scarce) to Quintana Roo (where men are abundant)

Outcome	Coefficient[a]	Difference in Relative Supply, Quintana Roo Minus Michoacán	Implied Impact	Impact Relative to Base Estimate[b]
Proportion Never-Married	−0.11	0.22	−0.02	−0.06
Proportion With No Children	−0.15	0.22	−0.03	−0.07
Proportion Enrolled	0.14	0.22	0.03	0.21
Years of Schooling	−1.98	0.22	−0.43	−0.05
Proportion Employed	−0.26	0.22	−0.06	−0.15

[a] Coefficients in this column are the coefficient estimates from the final column of Table 5. I use the smaller of either the IV or the OLS coefficients from this specification.

[b] For the proportion never-married, the proportion with no children, the proportion enrolled, and the proportion employed, the figures in this column present the ratio of the effect size in the previous column to the average value for all Mexican women aged 20–25 in 2000. For years of schooling, average education in 2000 for all women is used as the base value.

found evidence consistent with women dropping their standards when men are relatively scarce. Table 4 presents estimation results for the four outcomes meant to characterize several dimensions of the quality of the marriage from the perspective of the female. In particular, I explore the relationship between changes in relative male supply and four measures describing married women: (1) the median within-marriage age difference between husbands and wives, (2) the proportion of women married to a younger man, (3) the median within-marriage education difference between husbands and wives, and (4) the proportion of women married to a less-educated man.

Table 4 shows very little evidence that the quality of formed marriages is affected by the relative supply of males, at least along the dimension measured here. I find no evidence that the relative supply of men affects the median within-marriage difference in age between husbands and wives or the proportion of women married to younger men. Similarly, the educational difference between husbands and wives is unrelated to the relative supply of males in all models. In the most detailed specification for the change in the proportion of women married to less-educated men, the marginal effect of the relative supply of men becomes statistically significant and positive, suggesting that women in stronger marriage markets are more likely to marry men that are less educated than themselves—an effect in the opposite direction of what theory would imply. In general, however, there is little evidence of an impact of the relative supply of men on these outcomes.

Discussion

I documented quite large changes in the ratio of resident Mexican men to women since 1960 and the great deal of cross-state heterogeneity in these ratios. Sex ratios for prime-aged men and women are quite closely associated with the proportion of households in a state with a migrant abroad, with states with higher migration rates

Table 4 Regression model estimates of the impact of changes in the relative supply of men on changes in outcomes describing the spouses of married women

Dependent Variables	(1)	(2)	(3)
Δ Median Husband-Wife Age Difference			
OLS	0.54 (0.86)	0.43 (0.90)	0.21 (1.38)
IV	0.97 (1.24)	0.87 (1.25)	0.71 (2.13)
Δ Proportion Married to Younger Men			
OLS	−0.02 (0.02)	−0.01 (0.02)	0.01 (0.03)
IV	−0.03 (0.02)	−0.02 (0.02)	0.01 (0.04)
Δ Median Husband-Wife Education Difference			
OLS	−0.46 (0.42)	−0.48 (0.41)	−0.32 (0.50)
IV	−0.35 (0.36)	−0.36 (0.37)	0.04 (0.40)
Δ Proportion Married to Less-Educated Men			
OLS	0.05 (0.08)	0.08 (0.08)	0.27 (0.11)*
IV	0.06 (0.09)	0.09 (0.08)	0.28 (0.13)*
Fixed-Effect Specification			
Year effects	Yes	Yes	Yes
Age effects	No	Yes	Yes
State effects	No	No	Yes

Notes: Standard errors are shown in parentheses. Standard errors are computed allowing for clustering in each included age-state cell. The figure in each cell is the coefficient on the between-census change on the relative supply of males. Observation in the data vary across three age groups (16–19, 20–25, 26–30), 32 Mexican states, and three intercensus changes (1960–1970, 1970–1990, and 1990–2000), giving a total of 192 observations. Each regression also controls for the average age within each cell. All models are weighted by the average of the number of observations used to compute the intercensus change.

*$p < .05$

having relatively low sex ratios. Thus, it is certainly the case that emigration from Mexico has altered the internal demographic composition of the nation.

In addition, changes in the relative supply of men have geographically concentrated effects on several average socioeconomic outcomes for young Mexican women. In particular, states experiencing relatively large declines in the relative supply of men also experience relatively large increases in the proportion of young women who have never married, the proportion who have never had a child, female school enrollment rates, female educational attainment, and female employment rates. I find very little evidence that low sex ratios increase the proportion of married women who are paired with less-educated men or with men who are younger than themselves.

The findings reported here to some degree support the dynamic human capital accumulation model offered by Becker (1991). I did indeed find that a decline in marriage prospects as measured by lower male mate availability leads to greater labor force participation and an increase in formal educational attainment among young Mexican women. Moreover, the effect on the proportion never-married substantiates the hypothetical link between mate availability and the likelihood of marriage. The evidence presented is clearly consistent with young Mexican women incorporating the changing marriage market into their long-term decisions pertaining to the degree to which they engage in the formal labor market.

In light of the evidence of this rational behavioral response, it is surprising that there is little effect of mate availability on mate quality within formed marriages. This result stands in contrast to the results for African American women using similar estimation methods presented in Charles and Luoh (2010). This disparity in findings relative to research on the United States may be driven by several factors. First, the role of class in determining marriage pairings may exhibit greater influence in Mexico than it does among African Americans in the United States, among the French after World War I, or in other social and historical contexts where researchers have documented tangible evidence of a decline in standards in response to an adverse shift in the marriage terms of trade. Alternatively, finer outcomes of match quality, such as those employed in Harknett (2008), may be needed to detect changes in mate characteristics associated with relatively poor marriage odds for women. In the Mexican immigrant context, this may be particularly important given that Mexican immigrants to the United States generally have very low average levels of formal educational attainment and are characterized by an educational attainment distribution that is very low variance (Bohn et al. 2012).

Another contrast with the findings from previous research pertains to the proportion of never-married women who have borne children. Several researchers have found evidence of a negative relationship between mate availability and the likelihood that women have children out of wedlock. Clearly, the potential fertility responses to a paucity of marriageable men include either an increase in the likelihood of not having children, an increase in the likelihood of having a child out of wedlock, or increases in both probabilities. The U.S. research suggests responses along both margins, especially among African American women. The findings presented here indicate that for Mexican women, only the likelihood of never having a child increases. One might speculate that differences in cultural mores regarding marriage and fertility may explain this differential responsiveness. This is clearly a line of research worth pursuing in the future.

The findings here raise a number of questions that are certainly in need of further research. For example, as women's behavioral and social outcomes are influenced by the relative supply of men, it would be interesting to assess whether internal migration of men and women respond to variation in marriage market conditions. In other words, do women move from states where men are scarce to states where men are relatively abundant? Are internal migratory patterns for men opposite those of women? A further analysis of this question would certainly shed light on the process by which men and women seek out spouses and contribute to our understanding of the economics of household formation.

The results here also prompt the larger questions of how Mexican emigration has affected Mexican society. In a purely rational choice framework in which preferences are exogenously determined and individuals optimize subject to constraints, one would posit that the behavioral responses of women to changes in mate availability documented here would be reversed if emigration from Mexico to the United States were to dry up, as has occurred in recent years (Passsel et al. 2012). However, to the extent that women's preferences regarding work and marriage change, to the extent that female labor supply and human capital choices are determined in part by available older female role models, or to the extent that increase in female labor force participation and formal educational attainment are changing social mores, male-biased emigration flows may have permanently transformed gender roles in Mexico.

Related questions concern how the changes in female roles induced by gender-biased emigration are affecting women's relative position within social and economic hierarchies in Mexico. For example, does the increase in female labor force participation correspond to greater occupational mobility for women? Has the resultant economic independence altered the living arrangements of young Mexican women? For example, are they more likely to reside away from their parents?

More generally, greater attention should be paid to the impact of international migration on sending countries. The scale of north–south population movements is certainly sufficient to generate similar patterns in other large sending nations.

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