Missing Men and Female Labor Market Outcomes: Evidence from large-scale Mexican Migration*

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Abstract

There is little evidence of the effects of changes in sex ratios on labor market participation and outcomes in developing countries. We estimate the effects of a lower male-female sex ratio on female labor market participation and outcomes in Mexico. To identify this effect, we construct a Card (2001) style demand pull instrument. This instrument and the variation in migration rates across states, cohorts and over time, allow us to isolate the exogenous male demand shock in Mexican labor markets. Consistent with existing literature, we find that a reduction in the male-female sex ratio results in increased schooling for women and fewer children. Additionally, our results show an increase in female labor force participation, notably in white collar or "brain" jobs, and an increase in the fraction of women who are top earners. Our results are robust to different measures of migration. We find that despite the increase in higher skilled jobs for women, occupations are becoming more segregated, suggesting that the increase is from white collar and brain jobs previously dominated by women.

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Keywords: Sex ratio, Mexico, Female labor force participation, Female labor market outcomes, Migration.

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1. Introduction

The natural ratio of men to women is approximately one to one. However, sex-selective abortion, infanticide, diseases, famines, violence, wars, incarceration and migration can alter this ratio and lead to missing men or missing women. The relative scarcity of men or women has important consequences on: fertility, marriage and labor markets, and in investments in human capital.

The relative scarcity of men, due to wars, violence and incarceration has effects on female marriage prospects and labor force participation. Women are less likely to marry and have children, while at the same time out-of-wedlock child bearing increases (Abramitzky et al. 2011; Charles and Luoh 2010). The evidence on female labor force participation and labor market outcomes is less clear: Acemoglu et al. (2004) find increases in labor force participation of women in the US after WWII, while Goldin (2001) finds a lower effect on female employment. A lower ratio of men to women, and the subsequent changes in female labor market outcomes such as labor force participation and types of employment need to be documented further, in particular in the context of developing countries.[1] Our contribution to the existing literature lies in addressing a question that has not been explored in the context of a developing country and in adapting a methodology known in the migration literature, previously used to isolate supply shocks in the United States.

In this paper, we document the effects of a lower male-female sex ratio on the female labor market using large-scale migration of Mexican men to the US as a shock to the sex ratio,.[2] Using Mexican census data from 1960 to 2010 we document the dramatic decline in the sex ratios, particularly for young cohorts. We also document the variation in sex ratios across states and

[1] Related literature on sex in the labor market in Mexico by Juhn et al. (2013, 2014) document changes in occupations for Mexican females due to trade liberalization, where females increased employment and wage shares in blue-collar jobs but not in white-collar jobs.

[2] For direct effects of migration on the Mexican labor market see Chiquiar and Hanson 2005 and Mishra 2007.

cohorts and the extent to which international migration explains this variation. We find evidence that the percentage of households with an international migrant is negatively and significantly related to the sex ratio for a given age cohort and state. This provides evidence that changes in the sex ratio are related to large-scale migration of Mexican men.

We next, estimate the effects of the decline in the sex ratio on female labor market outcomes. In particular, we explore whether women were able to break into higher skilled, traditionally male dominated occupations as fewer men are available for these jobs. To identify these effects we use variation in migration rates across states, cohorts and over time. We also control for the potential endogeneity of the sex ratio by instrumenting for it using a Card (2001) style demand pull variable. The instrument, which uses predicted migration based on U.S. demand for labor, exploits the fact that the vast majority of Mexican migrants who go abroad go to the U.S., and the specific destinations of these migrants are driven by historic patterns. Using three different data sources on the sending states in Mexico and the receiving states in the U.S. we document this persistence, and show that U.S. demand is a good predictor of sex ratios. We provide numerous tests of the instrument, showing the relationship is not driven by particular sending or receiving states, and is only based on U.S. locations where potential networks exist.

We find that declines in the relative number of men have significant impacts on the labor force outcomes of women. The declines in the sex ratio lead to large increases in women's labor force participation (between 3-6%) %, as well as the percentage of women in white collar jobs (between 3 and 6%%) and in brain-intensive (between 3 and 6%) as opposed to brawn-intensive jobs. Interestingly, however, we find no associated decrease in occupational segregation, which suggests a portion of the skill increase is happening in female dominated professions. We also find that the female share in top 25 percent of earners increased, providing further evidence that women move into higher skilled occupations in the absence of men. Our results are robust to different measures for migration shocks driven by U.S. demand.

The rest of the paper is structured as follows: in the next section we discuss related literature; in Section 2 we present a framework for sex ratios and migration; Section 3 describes the main data source and shows descriptive statistics; Section 4 discusses our empirical strategy with the OLS and IV regressions; Section 5 presents the results and tests of the instrument; Section 6 explores two channels through which changing sex ratios may impact women's labor market outcomes: the marriage market and human capital accumulation; and Section 7 concludes.

2. Literature Review

Female labor market outcomes have received considerable attention in the developed and developing country literature. In particular, increases of female labor force participation and incomes earned outside the home may have a beneficial effect on the status, education and health of women in society, which in turn increases development and growth of a country. Reversely, economic development itself leads to changes in the labor market and better opportunities for women (Duflo 2012).For instance, Qian (2008) finds that higher female earnings lead to changes in the sex ratio as the value of female skills increases and therefore the value of female children. While Qian (2008) looks at the effect of female incomes on the sex ratio, our paper looks at the effects of a large shock to the sex ratio on the labor market in a developing country context.

The effects of a change in the sex ratio on marriage markets and labor markets have been documented in various contexts. The most well documented changes in the labor market for women in the developed country context are those in the U.S. labor market from 1930 to the 1950s due to World War II (Acemoglu et al. 2004; Goldin 1991, 2006). World War II, a shock to labor supply and labor demand, changed the number of men available to work in the U.S. labor market in the short-term but also changed the nature of work for women in the short-term to long-term. Acemoglu et al. 2004 find an increase in female labor force participation due to the mobilization and this in turn lowered both male and female wages and increased the earnings inequality between men of different education levels. Goldin (1991) does not find a continued permanent shift of labor force participation of women because many of the women that were employed during World War II exited by 1950. However, Goldin (2006) documents this period as the start of a continued increase in married female labor force participation, with decreases in the constraints facing married women's employment.

In their analysis of the change in the supply of men that drive female labor and marriage outcomes, Abramitzky et al. (2011) examine how World War I impacts the availability of men in France on the marriage market. The effect of a changed sex ratio in this context, where men were scarce or missing, was a change in marriage rates. Amongst decreases in divorce rates, increases in out-of-wedlock births and decreases in age gap the propensity for the men to marry increased while for women it decreased. Charles and Luoh (2010) in the context of high male imprisonment in the U.S. for a specific ethnic group found that the female marriage rates and the quality of male spouses decreased alongside increases in female education and labor supply.

In the context of migration, Angrist (2002) studies an increase in the sex ratio and its positive effect on the female marriage rate and negative effect on female labor force participation. Most of the literature is set in a different context: we contribute to this literature with our study employing large-scale Mexican migration, where disproportionally more men leave, to understand the effect of a change in the sex ratio on subsequent outcomes in the labor and marriage market for women in a developing country.

The most closely related paper to our paper is by Raphael (2013), who employs the Mexican census and the event of Mexican male migration to change the sex ratio. He finds that changes in the sex ratios, with less males in Mexico, increases: the number of never married women, female educational attainment and the proportion of women employed. He does not find evidence for females marrying younger or less educated men. The focus of his paper is on the marriage market and educational attainment of women and to a lesser extent on the labor market outcomes. Labor force participation of women and in particular occupation are not analyzed in Raphael's paper. Pagan and Sanchez (2000) have study the changes of female labor force participation and occupational choice, such as self-employment or employment in specific sectors, in Mexico from the beginning of the early 1990s. Recent evidence from Juhn et al. (2013, 2014) on the effect of trade liberalization on female labor market outcomes finds that female wages and employment in blue collar occupations improve but not in white collar occupations.

Mexican migrants are largely drawn from the middle of the Mexican wage distribution and are not necessarily negatively selected into migrating to the U.S. Chiquiar and Hansen (2005) find that migrants tend to be more educated than Mexicans that do not migrate. Further, the migrants tend to be young, male, from rural areas, and from the middle of the education distribution (Kaestner and Malamud 2014). Evidence on migration patterns of males and females finds that single adult men with low education levels migrate to the U.S. while internal Mexican migration is dominated by married women and men with higher levels of education (Aguayo-Téllez, Ernesto and Martínez-Navarro 2013). In terms of effects of migration on the labor market in Mexico, Mishra (2007) finds positive effects of Mexican migration on the wages of a specific skill group in Mexico. Migration might be due to labor demand conditions in the Mexican states, but it also depends on labor demand conditions in the U.S. states where the migrants go to.

3. Sex Ratios and Migration

3.1. Data and Descriptive Statistics

To examine the evolution of sex ratios we use data from the 1960, 1970, 1990, 2000 and 2010 Mexican Census, accessed through the IPUMS International webpage (maintained by the University of Minnesota). The 1980 Census data are not available due to a fire which destroyed many of the records. For each census year we calculate 8 year age cohorts by state, focusing on the years in which women are most active in the labor force (18-57). We have five cohorts in total—two younger cohorts, 18-25, 26-33—and three older cohorts, 34-41, 42-49, 50-57. Together this yields a total of 775 cohort-state-year combinations.[3]

Figure 1 shows the average sex ratios in Mexico for each of the census years. The figure shows that similar to most countries, the sex ratio for younger cohorts (ages 0 to 9) is well over one, reflecting the higher natural number of male births. The sex ratio declines sharply starting in the 18 to 25 cohort, where the ratio is well below one for all years, and remain below one until the 34 to 41 age cohort. In 1960 and 1970 the sex ratios then recover, moving back above one for the 42 to 65 year age group. For 1990, 2000 and 2010, however, the sex ratio remains below one for all remaining age cohorts. This highlights another feature of the graph, which is that the declines in the sex ratio are more acute in the later years (1990- 2010) than in the earlier years (1960 and 1970). Not only do the ratios fail to recover at the same rate as previous decade, the ratios are lower for all age cohorts after 1990. This shows that the problem of "missing men" became more, rather than less, acute over time.

[3] In the robustness section we redefine the age groups using 5 and 10 years and find that the results are XXX.

To show that the ratios in Figure 1 are distinctive we compare the sex ratios for Mexico in the year 2000 to those from the U.S. and Brazil, a country with similar income levels to Mexico, in the same year using Census data from IPUMS. The results, presented in Figure 2, confirm that Mexico's trajectory is unique. In the U.S. in the year 2000 the sex ratio for the 18 to 25 age cohort was 1.04, while that for Brazil was 0.999. For Mexico the ratio was 0.9, which means there was one fewer man for every ten women in that age cohort than in the U.S. or Brazil. This large gap between the sex ratios in Mexico and the other countries remains through the 34 to 41 age cohorts, highlighting the distinctive problem of missing men in Mexico.

3.2. International Migration

In the case of Mexico the changing sex ratios can be explained by international migration, particularly to the U.S. We begin by showing this through a simple theoretical framework, which follows that detailed by Raphael (2013).

Let M_{cs}^{t} = the total number of men in state s and cohort c in time t

M_{cs}^{t-1} = the total number of men in state s and cohort c in the previous period, t-1

Assuming no mortality, the difference between these two values is the number of men who have migrated out and the number of men who have migrated in. We assume only international migration drives these flows. This is based on data showing large differences in international migration rates but negligible differences in national migration rates by sex. For example, according to the 2009 Encuesta Nacional de la Dinámica Demográfica (ENADID, or the National Survey of Demographic Dynamics), a nationally representative survey that captures national and international migration over the past five years, 76% of international migrants to the U.S. over the 2004-2009 period were men, while only 50% of national migrants were men[4]. It can be shown that if the out and in migration rates are similar for men and women, the sex ratios do not change over time. Given that the national migration rates are roughly balanced across sexes this is less likely to drive changes in sex ratios.

Let M_{csm}^{t} equal the number of men in a cohort and state that migrate internationally (*m*) between t-1 and t, and M_{csr}^{t} equal the number of men in a cohort and state that return from abroad (*r*) between t-1 and t. The number of men in a given cohort and state in year t can be expressed as a function of the number of men in the previous period minus international migrants plus returning migrants.

$$M_{cs}^{t} = M_{cs}^{t-1} - M_{csm}^{t} + M_{csr}^{t}$$

The cohort of women can be expressed in the same manner:

$$W_{cs}^{t} = W_{cs}^{t-1} - W_{csm}^{t} + W_{csr}^{t}$$

To simplify the sex ratios further, we assume there is no in or out migration for women, which means $W_{sc}^{t} = W_{sc}^{t-1}$. The sex ratio in any given year t thus can be written as a function of the ratios in the previous year and the in (*r*) and out (*m*) migration rates for men:

$$\frac{M_{cs}^{t}}{W_{cs}^{t}} = \frac{M_{cs}^{t-1} - M_{csm}^{t} + M_{csr}^{t}}{W_{cs}^{t-1}} = \frac{M_{cs}^{t-1}}{W_{cs}^{t-1}} - \frac{M_{csm}^{t}}{W_{cs}^{t-1}} + \frac{M_{csr}^{t}}{W_{cs}^{t-1}}$$

This can be re-written to show that the sex ratio in any period is a function of the earlier sex ratio, and the out and return migration rates for men.

$$= \frac{M_{cs}^{t-1}}{W_{cs}^{t-1}} - \left[\frac{M_{cs}^{t-1}}{W_{cs}^{t-1}} + 1\right] outmigration_{cst} + \left[\frac{M_{cs}^{t-1}}{W_{cs}^{t-1}} + 1\right] returnmigration_{cst}$$

[4] These are based on the author's calculations from the 2009 ENADID, available on INEGI's website. We define national migrants as those who lived in a different state in Mexico 5 years ago.

In our empirical strategy we will exploit how international migration influences the sex ratio, to identify the effect of changes in the sex composition on female labor market participation and outcomes.

We use census data to establish if there is an empirical link between international migration and sex ratios across states. Starting in the year 2000 and continuing in 2010 the census includes a migration module that asks households if any member had moved to or returned from abroad in the past five years. We measure migration incidence as the percentage of households in a state that reports having a member of their household migrate abroad in the past five years.[5] We then regress the sex ratio for each cohort in a given state on the migration incidence for that state by census year. The results are shown in panel A of Table 1. They clearly show a negative and significant relationship between the incidence of international migration and the ratio of men to women in all age cohorts. For example, the coefficient on international migration incidence in the first column implies that, in the year 2000, a one standard deviation increase in the percentage of households with international migrants (3.6 percentage points) is associated with a decline in the ratio of men to women age 18-25 of 0.05 (5.5% of the mean). These numbers are large, they imply that states with the highest migration incidence (over 10% higher than states with the lowest migration incidence) are predicted to have sex ratios that are approximately 0.1 lower than states with the lowest migration incidence. This constitutes the difference between a fully balanced sex ratio of one to a very unbalanced ratio of 0.9. Furthermore, international migration incidence can explain a large percentage of the variation in sex ratios across states. For example, in the year 2000, 53% of the variation in the sex ratios of 18-25 year olds and 67% of the variation in the sex ratios of 26-33 year olds can be explained by international migration incidence.

[5] We also consider percentage that receive remittances and an index of this plus the percentage that receive return migrants. This measure of migration is the most comprehensive one available given the size and scale of the samples.

To further test if international migration can explain sex ratios we examine the relationship between another potential cause of missing men—homicides-- and sex ratios in the year 2010. From 2007 to 2010 Mexico experienced a dramatic increase in homicides due to the Drug War, which began in late 2006 after the government, for the first time, launched an attack on drug trafficking organizations. The homicides have been concentrated among young men and in a small number of states (authors' calculations from INEGI data). Given this variation across age cohorts and states, homicides over this period also might explain changing sex ratios. As shown in panel B of Table 2, however, this is not the case. Using two measures of homicides—the total per 100,000 inhabitants in the year 2010 and the totals for 2007 to 2010 per 100,000 inhabitants—we find no negative relationship between either measure and the sex ratio for any cohort. All of the coefficients are close to zero, none are significant and the R-squared values are quite low, showing that homicides explain very little of the variation in sex ratios across states.[6]

Finally, we calculate sex ratios for Mexican born individuals in the U.S., the main destination country for Mexican immigrants. We use data from the U.S. Census and American Community Survey, accessed through IPUMS. The results, presented in Figure 3, show a clear sex imbalance for all years except 1970 for the 18-49 year old age range. The sex ratios are well above one, and in some years and cohorts are as high at 1.5, which means there are 15 men for every 10 women in a particular age group. This provides further evidence that large-scale international migration led to declining ratios of men to women in Mexico.

3.3. Measures of labor force outcomes

[6] For example, in the year 2010 international migration explained 56% of the variation across states in the sex ratios of 18-25 year olds, while homicides explain only 3% of the variation. Thus this large event which affected young men in particular states does little to explain sex ratios.

We use the Mexican Census to define various labor force outcomes for women. We are particularly interested in examining if women move into more skilled occupations over time as a result of changing sex ratios. We start with labor force participation, defined as the percentage of women in a cohort and in a state that reports having an occupation. We use this in lieu of a measure based on an individual's employment status, as this question was not added until the 1990 census. However, in the years where both are available the overlap is quite high. In 1990 98.7% of those who are employed report an occupation, while in 2000 and 2010, 98% and 99.1%, respectively, do so.

We next consider the percentage of women in the labor force who are employed in jobs defined as white collar or "brain" jobs. White collar jobs are defined using two digit occupational codes and include categories such as professionals, technicians, managers, administrators, directive officials, and office workers. All other categories are defined as blue collar.

"Brain" jobs are defined using one digit occupational categories (ISCO codes), and include legislators, senior officials, managers, professionals, technicians, associate professionals, and clerks (ISCO=1,2,3 or 4). All other categories, which include service workers, agricultural and fishery workers, crafts and related trade workers and elementary occupations are defined as "brawn".[7]

We next consider the distribution of men and women across occupations by constructing an index of occupational segregation. We use the index defined by Beller (1985):

$$S_t = 1/2 \sum_i | m_{it} - f_{it} |$$

Where i denotes an occupation (2 digit code), m_{it} denotes the percentage of the total male labor force that is in the given occupation and f_{it} denotes the percentage of the female labor force that is

[7] Our categorization system is available upon request. This categorization following the current literature on brawn-intensive and brain-intensive occupations, which a number of authors have employed in the Mexican context (Vogl 2014; Rendall 2013; Bhalotra et al. (forthcoming)).

in the occupation. Thus $\sum_{i=1}^{I} m_{it} = \sum_{i=1}^{I} f_{it} = 1$. The measure therefore captures the sum of the absolute value of the difference between the percentage of the male labor force that is in the occupation and the percentage of the total female labor force in the same occupation. Without gender segregation, male and female workers should be equally distributed across occupations and the percentage of the labor force working in each occupation should be the same, so the value of the index would be zero. Larger values therefore would indicate a higher level of deviation from an equal distribution and more gender segregation in that industry.[8]

Finally we consider wages, measured by the log of average monthly earned income, and hours worked.[9] We also consider the percentage of women who fall within the top 25% of earners for a given state and year. The wage and hours worked data are only available for the 1990, 2000 and 2010 census. Thus the sample size for these outcomes is smaller. All earned income values are converted to year 2000 pesos using CPI data from INEGI.

Summary statistics for all of the labor force measures are shown in Table 2. The table shows clearly the increase in labor force participation for women over time, as the percentages increase significantly for all cohorts from 1960 to 2010. The table also shows an increase in women's participation in higher skilled jobs, as the percentage of women in white collar increases for all cohorts over time, while the percentage of women in "brain" jobs in increases for all but the youngest cohort. We see these increases reflected in average monthly income, which increases

[8] For example, consider two highly gendered occupations; drivers and conductors and household servants. In the 1960 Mexican Census 4.26% of the male labor force worked as drivers or conductors, as opposed to 0.25% of the female labor force. The absolute difference is 4.01% for this industry. For household servants, in the 1960 Census 22.9% of the female labor force worked in this profession, as compared to only 0.74% of the male labor force. The absolute difference is 22.16%. Compare these values to those from a less segregated occupation, like chemists and pharmacists. In 1960, the percentage of the male labor force working in this occupation was 0.29%, while the percentage of the female labor force working in this profession was 0.19%. The absolute difference is therefore only 0.1%. Therefore if the distribution of male and female workers is more unequal, more occupations will have higher absolute differences and the value of the index increases.

[9] The averages by state, cohort and year are constructed including individuals who report zero income.

significantly from 1990 to 2010, and even more strongly in the percentage of women who are in the highest earner category. This suggests the benefits from higher skilled jobs are more highly concentrated in the upper end of the distribution of female workers. Interestingly, the occupational segregation index shows that over time there has been a larger decrease in segregation for younger cohorts, but not much of a change for the older cohorts. And although we see a general decline over time in occupational segregation, the declines from 1960 to 2010 are not large, ranging from 3.7 to 10.7%. Indeed, despite the large increase in women participating in white collar jobs, the segregation within more specific job types does not appear to have declined much.

4. Empirical Strategy

Our basic model is as follows:

$$LFOutcome_{cst} = \beta_0 + \beta_1 SexRatio_{cst} + \gamma' X_{st} + \delta_t + \delta_c + \varepsilon_{cst}$$

The dependent variable is the particular labor force outcome for women in the 8 year age cohort *c*, in state *s* and year *t*. This is modeled as a linear function of the sex ratio for cohort *c* in state *s* and year *t*, state and year controls (*Xst*) and year and cohort fixed effects. If a decline in men relative to women led to increases in women entering high skilled occupations we expect β_1 to be negative for labor force participation, white collar, and brain jobs. In other words, a decline in the sex ratio is associated with an increase in women in high skill jobs. If this has subsequent effects on wages, we would expect β_1 to be negative as well. The impact on hours worked, however, is ambiguous, since depending on whether the income or substitution effect dominates, workers may reduce or increase hours worked with higher paying jobs. Meanwhile if a decline in men relative to women leads to a decrease in occupational segregation, then we expect β_1 to be positive. In other words, a decline in the sex ratio would be associated with less occupational segregation.

There are two concerns over the endogeneity of the sex ratio. First, there are concerns that the sex ratio could be influenced by women's re-location decisions, if for instance, women with characteristics that impact their labor market outcomes move to states with more (or less) favorable sex ratios. As a result, the sex ratios may reflect the outmigration of men, but also the in-migration of women with certain characteristics. Thus to ensure that the sex ratios do not purely reflect the changing location decisions of women, we limit the sample to non-movers, defined as those who reside in their state of birth. This is the closest we can come to eliminating movers, as the census does not include location histories.

The larger concern about endogeneity, however, stems from labor demand shocks. For example, if an industry that employs a large section of the population shrinks, men may migrate in search of work while women who stay face reduced employment opportunities. To control for this we include the percentage of the state's labor force by census year that is employed in each industry, as defined by 2 digit codes. This captures some changes in the employment landscape for men and women over time. These measures, however, likely do not fully capture labor demand shocks, and therefore to further address this concern, we instrument for the sex ratios using a measure that relies on shocks to the demand for Mexican migrant labor from the U.S. The instrument exploits the fact that international migration is the key source of changes in sex ratios over time and the vast majority of Mexican migrants go to the United States.

Migration flows are a function of the supply of labor from Mexico and the demand for Mexican labor in U.S. markets. The supply of labor likely is endogenous to the labor force outcomes of women who stay behind, but under the assumption that this demand is driven by factors that are idiosyncratic to local U.S. markets, the demand for labor from abroad is not. If demand from the U.S. is de-linked from factors in Mexico that impact women's labor force outcomes, we can use a measure of this demand as an instrument for sex ratios. Key to the feasibility of this measure is the fact that Mexican states send migrants to different locations in the U.S. and labor demand, in turn, varies across these locations. We detail this below. The differences in where Mexican states send migrants to the U.S. and in the demand for Mexican labor in the U.S. generate the variation needed to predict sex ratios across Mexican states and time.

To generate a measure of the demand for Mexican migrants in the U.S. we create an instrument that is very similar to that developed by Card (2001):[10]

Let M_g = number of male migrants from Mexico in U.S. state *g* as of a given year. This stock is a function of the demand conditions in U.S. state g and supply conditions in the sending Mexican state.

Let τ_{gc} = fraction of Mexican workers in U.S. state *g* as of a given year that are in cohort c. This captures differences in migration patterns by age group.

The combination of sending Mexican state and U.S. receiving state is determined by λ_{gs} = fraction of Mexican born in U.S. state *g* that are from Mexican state *s*. These weights are time invariant and rely on the persistence of migration patterns over time.

The instrumental variable therefore is the predicted male migration to the U.S., defined as $M_g \tau_{gc} \lambda_{gs}$. This is the number of men in each age cohort predicted to be missing in each Mexican state due to demand conditions in the U.S.

To calculate both M_g and τ_{gc} we use data from 1% micro samples from the U.S. Census and the American Community Survey, as accessed through IPUMS. For each census year we calculate the total number of Mexican born men in each 8 year cohort living in each U.S. state.

[10] The use of an exogenous instrument for local migration demand is also similar to a Bartik-style instrument (see Theoharides 2013).

We next examine if migration patterns differ by age cohort. Table 3 presents τ_{gc} values for the year 2000 for the 12 largest receiving U.S. states.[11] The table clearly shows that variation exists in the distribution of Mexican men across age cohorts. For these states the percentage of the Mexican male population that is made up of 18-33 year olds ranges from 62% (North Carolina) to 32% (New Mexico). Meanwhile the percentage of this population that is made up of 34-57 year olds ranges from 38% (Texas) to 22% (North Carolina). Thus the number of younger and older cohorts that leave different Mexican states should vary depending upon their receiving state in the U.S.

To calculate λ_{gs} we ideally would have historical migration patterns from a year prior to any of our census periods. However, information that is representative at both the state level in Mexico and the U.S. is both scarce and recent. We therefore rely on data from recent years to assign Mexican men in the U.S. to specific states in Mexico. The data we use come from three different sources, with the years varying slightly across each source. We thus construct three sets of weights and compare the results to test the assumption that the λs are driven by historic trends rather than current supply shocks. We start by detailing each data source.

The first data source we use is the EMIF Norte Survey (Encuestas sobre Migracion en las Fronteras Norte y Sur de Mexico, or Surveys on Migration to the Northern and Southern borders of Mexico)[12] [13]. Using a probabilistic sampling methodology for mobile populations, this dataset collects information on migrants, 15 years old and older, coming from Mexico and in transit to the

[11] The table with results for all 50 states is available upon request. In the interest of visual clarity, we limit this table to 12 states. This provides sufficient variation in receiving locations in the U.S. while still being readable.

[12] Source: El Colegio de la Frontera Norte, Secretaría del Trabajo y Previsión Social, Consejo Nacional de Población, Instituto Nacional de Migración, Secretaría de Relaciones Exteriores, Encuesta sobre Migración en la Frontera Norte de México,www.colef.mx/emif. According to EMIF methodology document, 94 percent of migrants travel through one of 8 locations.
http://www.colef.mx/emif/metodologia/docsmetodologicos/Metodologia%20Emif%20Norte%20y%20Sur.pdf
[13] We thank Anne Le Brun for suggesting these data source in a Card style instrumental variable.

US. Most migrants crossing to the US travel through one of 23 locations. Migrants are interviewed in locations in the northern Mexican border and at airports during 12 months of the year. Since the majority of migrants are male and we are interested in changes in the sex ratio, we use migration patterns for men. In the survey, migrants are asked their state of birth in Mexico, and the US state to which they plan to travel. We use this information to calculate a matrix of migration flows from a particular Mexican state to a particular US state. The data are first collected in 1995 and to abstract from the idiosyncrasies of any given year we use an aggregate collected in several rounds starting in 1995 through 2011.[14] There are 2,415,855 observations in the aggregated dataset.

The second data source is a module on international migration included in the 2002 National Survey of Employment (ENE), conducted by the National Institute of Statistics and Geography (INEGI). This survey is representative at the state level in Mexico and asks households about members who have migrated abroad in the past five years. Members are asked the state in Mexico where the migrant lived before leaving and the specific state in the U.S. where they arrived. The survey therefore captures migration flows over the 1997 to 2002 period.

The third data source comes from recently released information from Mexican consulates in the United States on the issuance of identification cards, known as *matrículas consulares* (consular registration card). *Matrículas consulares* are issued to individuals living abroad as proof of Mexican citizenship and are accepted in many places in the U.S. as official identification. The registration documents have been issued originally on paper but in response to the number of undocumented migrants the card issued follows the same security standards as the Mexican passport and can serve as identification also in the United States (Riosmena and Massey 2010).[15] Both legal and illegal immigrants can apply for them, and it is estimated that 40% of all Mexicans living in the U.S. have

[14] Specifically, we use data from surveys in 1995, 2001, 2002, 2008, and 2011.
[15] http://www.ime.gob.mx/es/estadisticas-de-mexicanos-en-estados-unidos

one.[16] The data are compiled by the Institute for Mexicans Living Abroad, a government agency that is part of Mexico's State Department. The data are available on an annual basis from 2008 to the present, and to abstract from the immediate impact of the financial crisis, we use data from 2010. These data therefore capture the number of immigrants from particular states in Mexico that applied for a *matrícula consular* from a particular state in the U.S. in 2010.

The exclusion restriction partially rests on the argument that the weights are a result of historic migration patterns and therefore are independent of shocks within states in Mexico that compel people to migrate. Previous research has documented that new migrants tend to go to areas where earlier cohorts from the same origin have gone (Bartel 1989). To the extent however, that new migrants follow pre-established migration patterns, then we can use more recent information on migration patterns. While weights prior to 1995 do not exist, we do have multiple years of EMIF data and can examine the correlation of these weights over time. We test the assumption that migration patterns do not change much over time using the EMIF data set that records migration flows for the longest period available, from 1995 to 2011. Using 1995 and 2011 data the correlation of the lambdas are .83, corroborating the idea that migration patterns are stable over time. Correlations using data for 1995 and earlier years are between .7 and .85. We provide more checks of this assumption in Section 5.

To further explain the instrument and compare the weighting matrices we calculate the average predicted values across states for the year 2000, a year in which sex ratios were very low. These means are shown in Table 4, with standard deviations in parentheses. The means are the total number of predicted male migrants by cohort, such that the value of 1,589, found in the first row in column one, means that one thousand, five hundred and eighty nine men between the ages of 18 and 25 from a given state, on average, are expected to reside in the U.S. rather than that Mexican

[16] Correspondence from Dirección IME Global on 17th July 2014.

state. These predicted values vary across states depending on the size of the population and migration incidence. The table clearly shows that predicted migration is highest for the 26 to 33 year age cohort, and slowly declines in subsequent age groups. Unsurprisingly, the lowest level of migration is predicted for the 50 to 57 year old age cohort, with migration predicted to be close to one third of that for the youngest three cohorts. The table also shows a remarkable level of consistency across all three weighting regimes. For every cohort the average predicted values and their standard deviations are quite similar. This further supports the assertion that migration patterns are largely driven by historic factors, and likely do not change dramatically over time. Three different data sets from various years yield surprisingly similar predictions about migration totals in the year 2000.

5. Results

5.1. First Stage

We begin by showing the results from the first stage regression, where predicted male migration is used as an exogenous determinant of the ratio of men to women in any given cohort, state and year.

$$\frac{Men}{Women}_{cst} = \alpha_0 + \alpha_1 \,\mathrm{Pr}\,edictedMig\,rationU.S._{cst} + X_{st}'\gamma + \delta_c + \delta_t + e_{cst}$$

It is important to note that *Xst* contains the log of the cohort size, which controls for the fact that larger and smaller states will have different numbers of predicted male migrants. We also scale the predicted migration number by 100 for ease of interpretation. The instrument captures stocks rather than flows of migrants, and thus predicts the number of Mexican men in a given cohort living in the U.S. as of any given year, rather than the number who have left over a certain time period. The measure captures both the outmigration and return migration rates of these men, as higher rates of the former and lower rates of the latter will lead to a larger number of men living in the U.S. As shown in the theoretical framework, the out and return migration rates of men can predict changes in the sex ratio over time, assuming there is sex imbalance in both. Our measure therefore should be a good predictor of sex ratios across cohorts, states and time.

The first stage results are shown in Table 5. Each cell contains the estimated coefficient on the sex ratio for cohort *c*, in state *s* and year *t*. We show estimates for the full 1960-2010 sample and the reduced 1990-2010 sample. This is necessary because the income and hours worked measures do not include the years 1960 and 1970. All standard errors are clustered at the state- cohort level. For ease of interpretation we rescale the predicted demand flows to be on a per 100 individual basis.

Overall the first stage results are strong. For all three sources of the weighting matrix the predicted migration value is negative and significant, showing that higher numbers of men predicted to be in the U.S. are indeed associated with lower male to female ratios in the sending states in Mexico. For example, the coefficients in columns one through three imply that an increase in the predicted number of men who migrate of 1000 leads to a decline in the sex ratio of 0.02. From the mean sex ratio across all cohorts, states and years, this constitutes a decline of 2.07%. Furthermore, it is possible to reject the null hypotheses that the first stage is weak or under-identified, as the F-statistics are above 10 in all cases. Thus predicted migration to the U.S., stemming from demand shocks, appear to be good predictors of sex ratios across cohorts, states and years.

5.2. Instrument Checks

Before proceeding we provide several checks of the validity of the instrument. The exclusion restriction holds if the predicted number of male migrations from different cohorts and Mexican states is independent of factors that determine both women's labor force outcomes and the incentives for men to move. The exogeneity of the instrument rests on the argument that the predicted migration numbers are generated by historic trends rather than contemporaneous supply shocks. If this assumption holds, an increase in employment opportunities in U.S. states where men from a given Mexican state historically have not gone should have no impact on the gender ratio in that Mexican state. For example, if men from Oaxaca don't migrate to Arizona, an increase in job opportunities in Arizona should do little to predict gender ratios in Oaxaca.

To check this assumption we create two, alternative instruments that rely on different weighting mechanisms. First, we create randomly assigned weights and allocate men in different cohorts across Mexican states using these random numbers. Second, we reverse the actual weights, such that Mexican states that have a high weight receive a low weight and Mexican states that have a low weight receive a high weight. For example, if Oaxaca has a low weight for Arizona (it makes up a small portion of the Mexican population in Arizona), we assign it a high weight. Meanwhile, if another state has a high weight, making up a large portion of the Mexican population in Arizona, we assign it a low weight. We do this for all three sources for our weights (the EMIF, ENE and Matricula Consular) and, in each case, allocate men in different cohorts across different Mexican states using these reversed weights. We then re-estimate the first stage equation using our alternative instruments. The results are provided in Panel A of Table 6. For comparison we include results from the first stage that use the actual instruments. For ease of interpretation we rescale the predicted demand flows to be on a per 1000 individual basis.

The results provide evidence in support of the exclusion restriction. While the actual instruments are associated with significant declines in the ratio of men to women, the alternative instruments are not. In the case of the randomly assigned weights, the coefficient is less than half the size of those from the actual instruments and is insignificant. In the case of the reverse weights, the coefficients are positive, extremely small, and insignificant. Thus predicted male outflows that either are independent of or inversely related to migration patterns do very little to predict sex ratios across states in Mexico. This suggests that male out-migration responds to conditions in U.S. states where there is an established network, but does not respond to conditions in U.S. states where these networks are negligible.

We also check if the first stage results are driven by particular receiving and and sending states. For receiving states the concern is Mexican immigrants from all Mexican states travel to a handful of locations in the U.S., in which case the instrument might not capture historic migration patterns and a response to U.S. demand conditions. For example, in the face of supply shocks all Mexican states may send men to California, the largest receiving state in the U.S., regardless of pre-existing networks or demand conditions in California. To test this we remove the largest receiving states in the U.S. Using the EMIF average weights, we define the largest U.S. receiving states as those in which either all Mexican states or all but one Mexican state has a presence. For example, according to the EMIF there is someone from every Mexican state in California. This exercise removes four states: California, Texas, Arizona and Florida. We then re-estimate the predicted migration flows using U.S. states except these four, and estimate the relationship between these modified predicted demand and the sex ratios. The results of this are shown in columns one through three in Panel B of Table 6. They show that the first stage is not driven by these main receiving U.S. states, as the coefficients on the modified demand are negative and significant. If anything, these states do less to predict sex ratios than those that receive migrants from fewer Mexican states, as the coefficients on the modified demand are two to three times larger than those for the original ones. This provides more evidence that the instrument indeed captures U.S. demand shocks.

For sending states, the concern is that the correlation between predicted migration flows and sex ratios is due to a small number of states with high rates of international migration. If this is the case, the instrument would fail to predict sex ratios for the vast majority of states. To test this we remove states with the highest level of international migration in the year 2000 (the first census where this information was included). We define high migration states as those where more than 10% of households said they had an international migrant. This removes two states: Michoacán and Zacatecas. The results, provided in columns five through seven in Panel B of Table 6 show the first stage results are not driven by these two states. The coefficient on predicted demand remains similar in size, negative and significant for all three weighting mechanisms, providing further evidence that the instrument captures demand rather than supply shocks.

5.3. Second Stage Results

The second stage results are shown in Table 7. Each cell contains the estimated coefficient on the sex ratio for cohort *c*, in state *s* and year *t*. Standard errors, clustered at the age cohort-state level, are presented in parentheses. For each outcome we present second stage IV coefficients for all three weighting matrices and OLS coefficients for comparison. It is important to note that the sex ratios have not been scaled, such that the mean across all cohorts and years is 0.964 and the standard deviation is 0.126.

The results show that decreases in the number of men to women have strong impacts on the labor force outcomes of women. Starting with labor force participation, all of the coefficients are negative and significant, with the IV coefficients ranging from -0.3 to -0.58. This means a decrease in the male to female ratio of 0.1 (or a decrease of one man per every ten women) leads to an estimated increase in women's labor force participation between 3% and 5.8%. Given that women's labor force participation averages 24.2%, this constitutes an increase of 12.4%-- a non-trivial amount. Thus a reduction in the relative number of men leads to large changes in labor force participation of women.

We also find that among women in the labor force, they are more likely to be in white collar or brain jobs if the sex ratio is low. Specifically, according to the coefficients in columns two through four, a decrease in the male to female ratio of 0.1 leads to an increase in the percentage of women in white collar jobs that ranges from 3.0% to 6.1%. Meanwhile, the percentage of women in brain jobs is predicted to increase between 3.7% and 6.1%. Given the mean values for both white collar and brain jobs, the highest values constitute increases of approximately 16.3 and 18.6%, respectively. These coefficients are significant for two out of the three weighting matrices. We therefore find strong evidence that "missing men" increase women's participation in high skilled jobs.

We would expect, then, that occupational segregation would decline in areas and among cohorts where sex ratios fall more dramatically, because women would be moving into white collar and brain jobs that were previously dominated by men. Interestingly, however, we do not find this. In fact, we find that a decline in the number of men to women is associated with an increase in occupational segregation. Thus among cohort-state pairs where the sex ratios are declining more over time, occupations are becoming more segregated by gender than before. This suggests that some of the increase is from white collar and brain jobs previously dominated by women.

Is the increase in labor force participation and the prevalence of women in higher skilled jobs reflected in earnings? On average, it is not, as none of the IV coefficients for log monthly income are significant, and in one case (EMIF as weighting matrix) the coefficient is positive. We do see, however, a significant increase in the percentage of women in the top 25% of earners. On average, a decline in the ratio of men to women of 0.1 is associated with an increase in the number of women in the top earner group that ranges from 3.7% to 5.1%. This suggests the gains from higher skilled jobs are concentrated at the upper end of the earnings distribution.

Finally, for hours worked we find no significant relationship between the decrease in the relative number of men and the average number of hours women work in a week. While the coefficients are positive, the values vary greatly across the models and none are statistically significant. To some degree the lack of significance is unsurprising, as we did not expect a relationship between increased skill and hours worked.

In sum, we find that declines in the number of men relative to women leads to significant increases in women's labor force participation and their number in high skilled professions. We do not find, however, that occupational segregation declines, suggesting that women moving into male dominated jobs does not completely explain rising levels of white collar and brain jobs. Meanwhile, we find that declines in the relative number of men increases in the number of women who are top earners, although it does not significantly increase average wages. This suggests the wage gains from entering higher skilled occupations are concentrated among the higher end of the income distribution.

6. Channels

There are two main channels through which the absence of men can impact women's labor market outcomes. The first is through marriage markets. Given the absence of potential mates, women are less likely to marry and have children and more likely to stay in school longer and enter the labor market. Raphael (2013) directly examines the marriage market channel within the context of Mexico and finds that a decrease in the relative number of potential male spouses increases the percentage of women who never marry and do not have children.

We therefore estimate the IV and OLS models taking never marrying and no children as our outcome variables. The results are shown in the last two rows of Table 8. Unlike Raphael, we find no significant effects of a decline in the relative number of men on the likelihood that women marry. Three of the four coefficients are positive, which suggest that a decrease in the number of men reduces the chance of never marrying (increases the change of marriage), but none are significant. For children, however, we find results that are similar to those of Raphael. The IV estimates show that a decrease in the relative number of men increases the probability that a woman does not have children. The differences in the results can be explained by the approach, as Raphael focuses on marriage markets and therefore constructs a sex ratio that considers the potential pool of male mates for women in an existing age cohort. Our sex ratio is simpler, matching women to men in the same age cohort. Since women marry men who are older, the absence of men in a woman's exact age cohort may have less of an impact on her marriage outcomes than a lack of men in the age cohorts she is most likely to marry.

The second channel through which sex ratios may impact labor force outcomes is human capital accumulation. As a result of there being fewer men, women may acquire more education. This may be due to fewer marriage market opportunities, an increase in the availability of education resources (due to less competitive from men), or an increase in the availability of high skill jobs, due to an absence of male candidates. We therefore investigate two human capital accumulation outcomes: total years of schooling and whether or not a woman went to college. The results, shown in the first two rows of Table 8, show a clear relationship and much stronger relationship between education and sex ratios. For years of education, the IV results show that a decline in the sex ratio of 0.1 leads to an increase in education of 1.1 to 1.7 years. Given that average years of education for women are 5.7 this constitutes an increase of 20% to 30%. The results also show that women in areas with a lower relative number of men are more likely to accumulate a higher level of human capital by going to college. According to the IV results, a decline in the sex ratio of 0.1 leads to an increase in the percentage of women in a cohort who go to college between 0.8 and 2.3%. Average college rates are 4.5% in the sample, highlighting this as a significant change.

These results indicate that the human capital channel is stronger than the marriage market channel, and that a key factor that pushes women into higher skill jobs in the absence of men is increased education. Higher college attendance rates, in particular, can explain why more women move into white collar and brain jobs.

7. Conclusion

In this paper we explore the impact of "missing men" on the labor market outcomes of working age women in Mexico. Mexico provides a unique case because it is a developing country that experienced dramatic changes in the relative ratio of men to women due to the international migration of men. The absence of potential employees and of potential mates may have changed the labor market opportunities of women. Using a model which instruments for the sex ratio and predicted migration stemming from U.S. demand we find that declines in the sex ratio have large effects on the labor market outcomes of women. We employ different data sources for our instrumental variable strategy and find the results to be consistent across them. Not only are these women more likely to enter the labor force, they are more likely to have higher skilled jobs. We also find that increased educational attainment may partially explain this result, as women who face lower relative number of men have more education and are more likely to attend college.

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ENADID Encuesta Nacional de la Dinámica Demográfica, INEGI

ENE Encuesta Nacional de Empleo, INEGI

EMIF. El Colegio de la Frontera Norte, Secretaría del Trabajo y Previsión Social, Consejo Nacional de Población, Instituto Nacional de Migración, Secretaría de Relaciones Exteriores, Encuesta sobre Migración en la Frontera Norte de México, www.colef.mx/emif.

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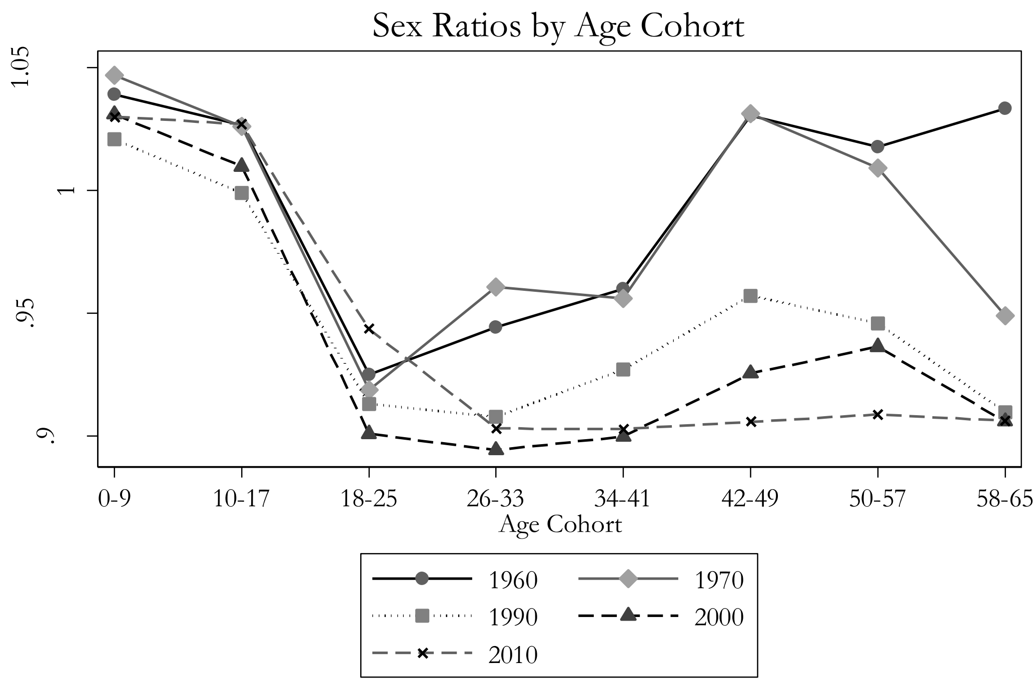
Figures and Tables

Figure 1: Sex Ratios by Census Year and Cohort, Mexico



Source: Mexican Census, accessed through IPUMS

Figure 2: Comparison of Sex Ratios

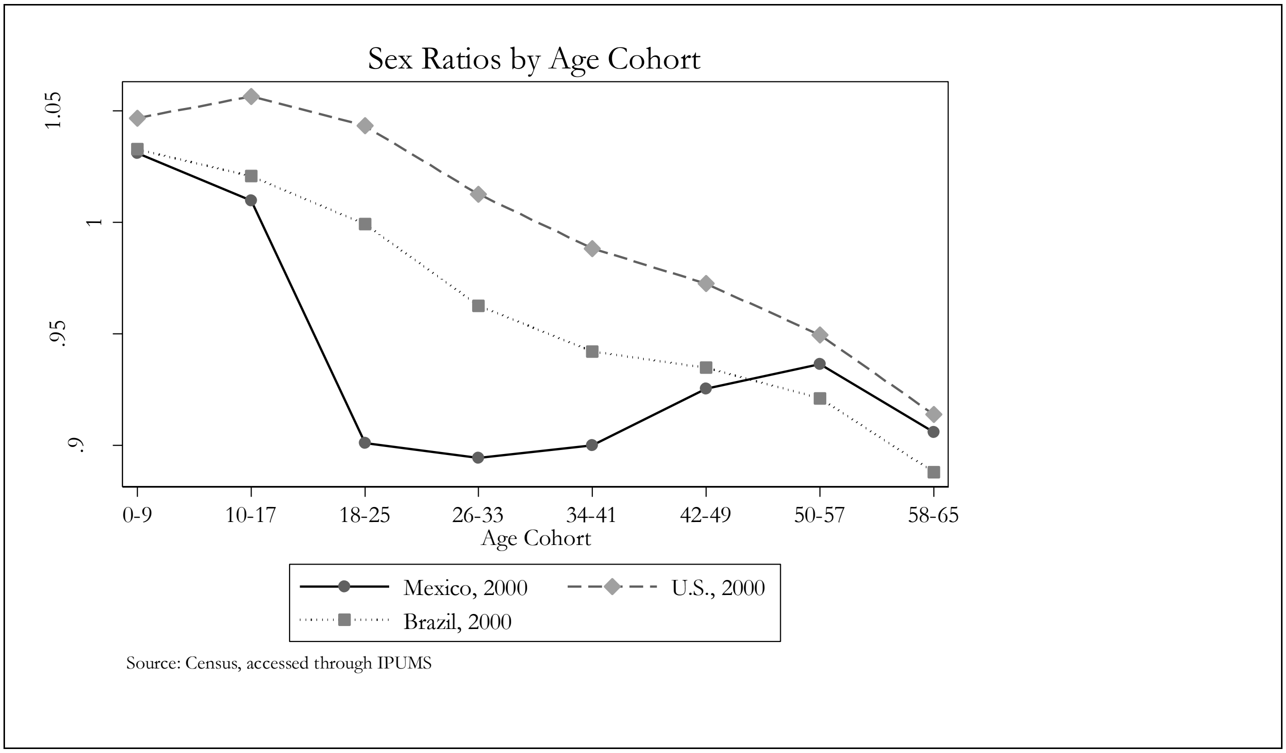


Figure 3: Sex Ratios for Mexican Born residents of the U.S.

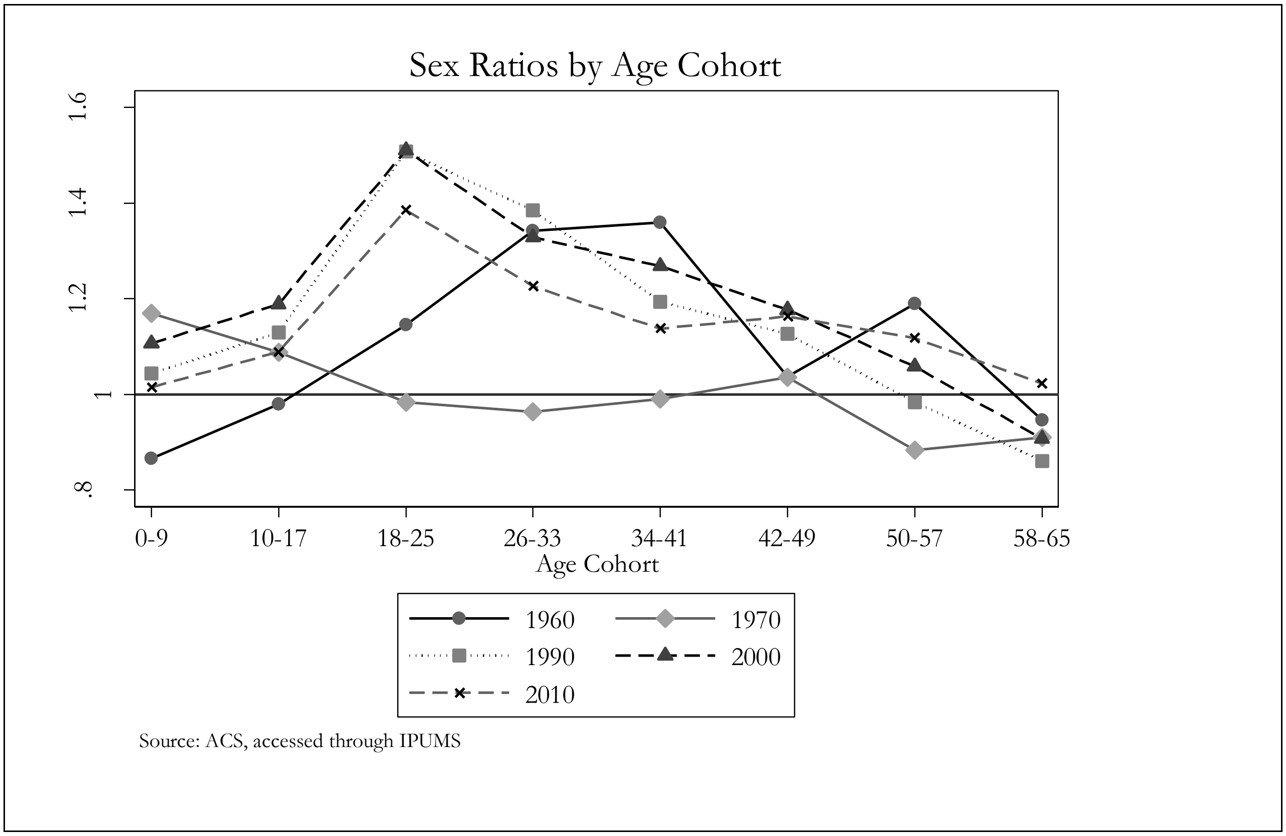


Table 1: Sex Ratios and International Migration Incidence by State

PANEL A: Sex Ratios and International Migration	18-25 year olds		26 to 33 year olds		34 to 41 year olds		42 to 49 year olds		50 to 57 year olds	
VARIABLES	2000	2010	2000	2010	2000	2010	2000	2010	2000	2010
% HHs with international migrant in past 5 years	-0.0135***	-0.0286***	-0.0141***	-0.0322***	-0.0112***	-0.0192***	-0.00522*	-0.0164***	-0.00602*	-0.0209***
	(0.00232)	(0.00466)	(0.00180)	(0.00605)	(0.00173)	(0.00494)	(0.00280)	(0.00511)	(0.00313)	(0.00664)
Constant	0.975***	1.022***	0.969***	0.982***	0.961***	0.951***	0.962***	0.943***	0.979***	0.963***
	(0.0124)	(0.0113)	(0.00962)	(0.0146)	(0.00930)	(0.0119)	(0.0150)	(0.0123)	(0.0168)	(0.0160)
Observations	32	32	32	32	32	32	32	32	32	32
R-squared	0.529	0.556	0.671	0.485	0.580	0.336	0.104	0.256	0.110	0.248
PANEL B: Sex Ratios and Homicides										
Census Year 2010 Only	18-25 year olds		26 to 33 year olds		34 to 41 year olds		42 to 49 year olds		50 to 57 year olds	
Homicides per 100,000 inhabitants year 2010 only	0.000430		0.000534		-8.70e-06		8.37e-05		-3.01e-05	
	(0.000325)		(0.000391)		(0.000289)		(0.000282)		(0.000365)	
Homicides per 100,000 inhabitants, years 2007-2010		0.000120		0.000168		-4.71e-05		1.80e-05		-4.25e-05
		(0.000124)		(0.000149)		(0.000109)		(0.000107)		(0.000138)
Constant	0.948***	0.949***	0.897***	0.896***	0.912***	0.918***	0.906***	0.907***	0.922***	0.926***
	(0.0151)	(0.0179)	(0.0182)	(0.0215)	(0.0135)	(0.0157)	(0.0131)	(0.0154)	(0.0170)	(0.0199)
Observations	32	32	32	32	32	32	32	32	32	32
R-squared	0.055	0.030	0.059	0.041	0.000	0.006	0.003	0.001	0.000	0.003

Standard errors in parentheses *** p<0.01, ** p<0.05, * p<0.1

Data source for sex ratios and percentage of HHs with international migrants is the Mexican census, as accessed through IPUMS.

Data source for homicides are municipal death records, while are compiled and accessible through INEGI's website.

Table 2: Summary Statistics on Labor Force Outcomes

Cohort	18-25	26-33	34-41	42-49	50-57
Women's Labor Force Participation					
1960	11.2%	8.0%	9.6%	9.7%	11.0%
1970	19.5%	12.7%	12.1%	13.4%	12.0%
1990	25.5%	25.4%	21.8%	17.2%	11.3%
2000	34.5%	38.4%	40.9%	35.8%	26.8%
2010	33.0%	45.1%	47.5%	46.0%	36.0%
Occupational Segregation					
1960	68.7	69.3	68.4	69.0	69.1
1970	71.9	75.1	75.8	76.0	71.4
1990	69.8	67.1	69.0	71.6	73.9
2000	61.4	62.5	64.1	66.7	69.4
2010	61.5	61.9	64.7	66.4	69.1
% Women in White Collar Jobs					
1960	37.5%	33.1%	24.5%	26.9%	14.7%
1970	36.5%	32.3%	21.6%	17.7%	16.5%
1990	48.0%	59.1%	46.4%	38.2%	26.8%
2000	32.5%	41.3%	39.6%	31.3%	22.3%
2010	58.6%	61.8%	56.1%	55.1%	49.5%
% Women in Brain Jobs					
1960	38.8%	35.2%	26.3%	28.7%	15.8%
1970	36.7%	33.2%	22.2%	17.9%	16.6%
1990	48.2%	58.9%	45.6%	36.7%	24.6%
2000	33.6%	42.7%	40.7%	32.1%	22.7%
2010	33.1%	40.6%	34.9%	34.1%	25.8%
Log(Average Earned Monthly Income (year 2000 pesos)), for women					
1990	6.40	6.71	6.56	6.31	5.83
2000	6.44	6.89	7.09	6.86	6.32
2010	7.71	8.06	8.07	8.15	8.08
% Women in Top 25% of Income					
1990	14.6%	18.0%	14.7%	11.0%	6.1%
2000	18.3%	24.4%	25.5%	19.6%	11.6%
2010	19.7%	36.4%	35.4%	38.0%	31.4%
Weekly Hours Worked, for women					
1990	42.13	38.90	39.64	40.38	40.59
2000	42.71	40.25	38.94	39.33	38.57
2010	41.08	40.01	39.66	39.13	38.70

Income and hours worked only available for 1990, 2000 and 2010.
Income scaled to year 2000 peso values using CPI information provide
Size of 1960-2010 sample is 775. Size of 1990-2010 sample is 480.

Table 3: Table of Taus for Select U.S. States (τ)

Percent of Mexican born men in a U.S. state that are in a specific age cohort

Year 2000	Cohort					Summary	
State	18-25	26-33	34-41	42-49	50-57	% young	% older
New Jersey	31.8%	26.2%	17.3%	6.3%	2.5%	58.0%	26.1%
New York	29.3%	28.7%	16.0%	6.5%	3.1%	58.0%	25.5%
Illinois	20.0%	24.5%	18.0%	11.5%	6.4%	44.5%	36.0%
Georgia	32.5%	27.4%	14.1%	6.2%	2.8%	59.9%	23.1%
North Carolina	34.5%	27.6%	13.8%	6.0%	2.4%	62.1%	22.1%
Texas	18.1%	20.7%	18.2%	12.4%	7.3%	38.9%	37.9%
Arizona	20.4%	20.9%	16.9%	10.8%	6.4%	41.3%	34.1%
Colorado	24.7%	23.8%	15.8%	9.3%	4.9%	48.6%	30.0%
New Mexico	14.0%	18.0%	17.8%	14.1%	8.9%	32.0%	40.8%
California	17.5%	23.3%	19.5%	12.2%	6.8%	40.8%	38.4%
Oregon	24.6%	27.1%	17.8%	8.1%	3.8%	51.7%	29.7%
Washington	23.1%	25.6%	16.4%	8.9%	5.0%	48.7%	30.3%

Data from U.S. Census, as accessed by IPUMS

Table 4: Predicted Male Migration to the U.S in the year 2000, by cohort

Weighting Source Predicted Male Migration	EMIF (1)	ENE (2)	Matricula Consular (3)
18 to 25 Age Cohort	1589.54	1588.71	1590.13
	(1651.0)	(1616.1)	(1487.7)
26 to 33 Age Cohort	1820.88	1819.62	1821.38
	(1917.7)	(1914.1)	(1748.3)
34 to 41 Age Cohort	1406.41	1405.41	1406.97
	(1488.2)	(1504.5)	(1353.2)
42 to 49 Age Cohort	858.63	857.97	858.91
	(904.0)	(928.3)	(820.6)
50 to 57 Age Cohort	475.75	475.22	475.78
	(499.1)	(517.0)	(453.1)

Coefficients are mean values across states. Standard deviations are in parentheses.

Table 5: First Stage IV Results

Outcome variable= Sex ratio (men/wome	1960-2010 Sample			1990-2010 Sample		
	EMIF	ENE	MatCon	EMIF	ENE	MatCon
	(1)	(2)	(3)	(7)	(8)	(9)
Predicted Migration, EMIF	-0.002***			-0.002***		
	(0.000)			(0.000)		
Predicted Migration, ENE		-0.002***			-0.002***	
		(0.000)			(0.000)	
Predicted Migration, Matricula Consular			-0.002***			-0.002***
			(0.000)			(0.000)
Constant	0.324	0.186	0.204	0.211	-0.154	0.023
	(0.316)	(0.316)	(0.316)	(0.252)	(0.248)	(0.247)
Observations	775	775	775	480	480	480
R-squared	0.292	0.295	0.294	0.510	0.528	0.515
Angrist-Pischke F test	17.25	22.08	15.43	47.23	60.86	56.28
Kleinbergen-Paap rk LM ChiSquared	10.13	12.67	11.30	16.73	22.12	20.78

Robust standard errors, clustered at the state-cohort level, in parentheses

All regressions include year, and cohort fixed effects, the percentage of the state's workforce employed in different industries (2 digit code) by year and the log of the size of the cohort (both men and women) by state and year. Income and Hours Worked are only available in the 1990, 2000 and 201

Data Source: Mexican Census, as accessed through IPUMS

Table 6: Instrument Robustness Checks

PANEL A: Alternative Instruments	Weighting Mechanism						
Outcome variable=	Random	EMIF	EMIF	ENE	ENE	MatCon	MatCon
Sex Ratio (men/women)	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Predicted Migration, Random Weigh	-0.00759						
	(0.0103)						
Predicted Migration, Actual Weights		-0.0151***		-0.0168***		-0.0172***	
		(0.00363)		(0.00357)		(0.00437)	
Predicted Migration, Inverse Weight			0.000186		0.000217		0.000188
			(0.000410)		(0.000408)		(0.000409)
Observations	775	775	775	775	775	775	775
R-squared	0.283	0.292	0.282	0.295	0.282	0.294	0.282
PANEL B: Modified Sample	Remove Largest U.S. Receiving States				Remove Largest Mex. Sending States		
Outcome variable=	EMIF	ENE	MatCon		EMIF	ENE	MatCon
Sex Ratio (men/women)	(1)	(2)	(3)		(5)	(6)	(7)
Predicted Migration, EMIF	-0.0345***				-0.0136***		
	(0.0113)				(0.00465)		
Predicted Migration, ENE		-0.0395***				-0.0163***	
		(0.0120)				(0.00454)	
Predicted Migration, Matricula Cons			-0.0389***				-0.0159***
			(0.0128)				(0.00552)
Observations	775	775	775		725	725	725
R-squared	0.288	0.290	0.289		0.268	0.270	0.269

Robust standard errors, clustered at the state-cohort level, in parentheses *** p<0.01, ** p<0.05, * p<0.1

All regressions include year, and cohort fixed effects, the percentage of the state's workforce employed in different industries (2 digit c
and the log of the size of the cohort (both men and women) by state and year.

Table 7: Second Stage IV Results

Coefficient on sex ratio (men/women)		IV, Source of Weighting Matrix		
	OLS	EMIF	ENE	MatCon
Outcome Variable	(1)	(2)	(3)	(4)
Women's Labor Force Participation	-0.0552**	-0.302*	-0.359**	-0.580***
	(0.0267)	(0.165)	(0.156)	(0.217)
Occupational Segregation	-2.111	-34.704**	-27.352**	-20.092
	(5.043)	(17.122)	(13.652)	(16.821)
% Women in White Collar Jobs	-0.104*	-0.461*	-0.606**	-0.301
	(0.0539)	(0.280)	(0.253)	(0.280)
%Women in Brain Jobs	-0.106*	-0.531*	-0.615**	-0.368
	(0.0550)	(0.275)	(0.244)	(0.272)
Observations	775	775	775	775
Log(Monthly Earned Income) women	-1.204***	0.151	-0.515	-0.276
	(0.307)	(0.814)	(0.660)	(0.752)
% Women in top 25% earners	-0.286***	-0.371**	-0.509***	-0.499***
	(0.0636)	(0.174)	(0.150)	(0.161)
Weekly Hours Worked women	0.119	0.426	5.372	5.544
	(1.504)	(5.141)	(4.238)	(4.591)
Observations	480	480	480	480

Robust standard errors, clustered at the state-cohort level, in parentl *** p<0.01, ** p<0.05, * p<0.1
All regressions include year, and cohort fixed effects, the percentage of the state's workforce emplo
in different industries (2 digit code) by year, and the log of the size of the cohort (both men and w
by state and year. Income and Hours Worked are only available in the 1990, 2000 and 2010 census
Data source: Mexican Census accessed via IPUMS

Table 8: Education, Marriage and Children Outcomes, Second Stage IV Results

		IV, Source of Weighting Matrix		
Coefficient on sex ratio (men/women)	OLS	EMIF	ENE	MatCon
Outcome Variable	(1)	(2)	(3)	(4)
Years Education, women	1.024**	-11.592***	-13.480***	-16.966***
	(0.435)	(3.715)	(3.763)	(5.149)
% women who are college educated	0.00263	-0.086	-0.095**	-0.232***
	(0.00770)	(0.056)	(0.047)	(0.088)
% women who are never married	0.132***	0.235	0.092	-0.002
	(0.0449)	(0.230)	(0.210)	(0.234)
% women who have no children	0.0508*	-0.789***	-0.758***	-0.815***
	(0.0296)	(0.213)	(0.191)	(0.249)
Observations	775	775	775	775

Robust standard errors, clustered at the state-cohort level, in pa *** p<0.01, ** p<0.05, * p<0.1

All regressions include year, and cohort fixed effects, the percentage of the state's workforce employed in different industries (2 digit code) by year, and the log of the size of the cohort (both men and women) by state and year.

Data source: Mexican Census accessed via IPUMS