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Fertility Responses of High-Skilled Native Women to Immigrant Inflows

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Abstract Despite debate regarding the magnitude of the impact, immigrant inflows are generally understood to depress wages and increase employment in immigrant-intensive sectors. In light of the overrepresentation of the foreign-born in the childcare industry, this article examines whether college-educated native women respond to immigrant-induced lower cost and potentially more convenient childcare options with increased fertility. An analysis of U.S. Census data between 1980 and 2000 suggests that immigrant inflows are indeed associated with native women's increased likelihoods of having a baby, and responses are strongest among women who are most likely to consider childcare costs when making fertility decisions—namely, married women and women with a graduate degree. Given that native women also respond to immigrant inflows by working long hours, this article concludes with an analysis of the types of women who have stronger fertility responses versus labor supply responses to immigration.

Keywords Fertility · Childcare · Immigration · Labor supply

Introduction

The foreign-born population of the United States has quadrupled since the passage of the Immigration and Nationality Act of 1965 (Hart-Celler Act). Among politicians and academics, this increase has led to substantial interest in the socioeconomic consequences of the recent waves of immigration to the United States. Much of the existing research has focused on the potentially negative impact of immigration on the wages and employment rates of natives (Borjas 2003; Card 2001; Ottaviano and Peri 2012). Less attention has been paid to the potential benefits accruing to natives from

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immigration. This article considers the impact of low-skilled immigrant inflows on native fertility decisions and provides evidence that childcare markets are driving responses.

Decreases in the price and increases in the availability of childcare brought on by low-skilled immigration should imply reductions in the cost of child rearing. However, the theoretical impact of lower child-rearing costs on childbearing is unclear given that women may respond to these lower costs by increasing labor supply (Blau and Robins 1989) instead of increasing fertility. Cortés and Tessada (2011) found that low-skilled immigration to large U.S. metropolitan areas results in increases in the number of hours worked by women at the top of the wage distribution. If these labor supply responses are sufficiently large, then immigrant-induced decreases in childcare costs may decrease the likelihood of having a second or third child. Thus, the relationship between immigrant inflows and childbearing is essentially an empirical question.

Any analysis making use of geographic variation in immigrant concentration to study immigrant impacts must address the fact that immigrant location decisions are not exogenous. Even estimates from fertility models that control for observed and unobserved characteristics of cities that stay constant over time are biased upward if low-skilled immigrants have become increasingly likely to settle in cities where high-skilled native-born women are developing stronger preferences for large families. On the other hand, if cities with booming economies are attracting more immigrants while providing better labor market opportunities for high-skilled women, then standard estimates of the effect of immigrant concentration will be biased downward. To address these potential concerns, I take an instrumental variables approach, common in the immigration literature, which relies on the propensity of new entrants to locate in areas with high historical concentrations of immigrants from the same country of origin (Bartel 1989; Card 2001).

Using 1980–2000 U.S. Census data on U.S.-born college-educated women of childbearing age, I find that in models controlling for city fixed effects, increases in the share of low-skilled immigrants in a city are associated with an increased probability that women in that city have recently given birth. Instrumental variables models suggest an even stronger impact, implying that immigrants are less attracted to high-fertility cities. The increases in the probability of giving birth are likely to translate into increases in completed fertility given that older women, who cannot easily adjust future fertility, are the most influenced by immigrant inflows.

For evidence that immigrants are impacting native fertility decisions via childcare markets, I start by showing that metropolitan areas receiving more immigrants have larger decreases in the median wages of childcare workers. Suggestive of immigrant-induced labor supply shocks, these cities also tend to have a greater share of the labor force working in childcare, although this latter effect is small.

Next, I examine whether it is indeed the women most likely to use formal childcare options—as opposed to caring for their own children full-time or using friends and family for childcare—who are most sensitive to immigrant inflows. Results suggest that women with graduate degrees are more responsive than women with only college degrees, pointing to a role played by childcare markets given that higher-skilled women are less likely to live near family members (Molloy et al. 2011) and have higher opportunity costs of leaving the labor force. Interestingly, unmarried women have

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barely perceptible, albeit statistically significant, responses—a finding that makes sense if unmarried women are less likely to have carefully planned pregnancies.

For further evidence that immigrants affect fertility outcomes through childcare markets, I exploit variation in the country-of-origin composition of immigrants in different cities in different years. Immigrants from certain countries, such as Paraguay and Cameroon, are especially likely to work in the childcare industry, although virtually no immigrants from Albania and Bulgaria work in this industry. I find that native-born women have strong fertility responses to immigrant inflows from "high-childcare" countries and no statistically significant responses to inflows from "low-childcare" countries.

As noted earlier, this analysis complements a growing literature showing that women tend to work more hours in response to reduced childcare costs. Although women may respond to lower childcare costs by both increasing hours at work and having an additional child, some women may respond to lower childcare costs by working longer hours, whereas others may respond by having an additional child. Because some policy-makers may be focused on increasing fertility rates, but others are more interested in eliminating gender wage gaps, it may be useful to know how different types of women respond to lower childcare costs so that policies can be targeted appropriately.

To examine this issue, I start by reproducing the result in the literature that immigrant inflows tend to increase labor supply of high-skilled women, especially at the top of the hours worked distribution (Cortés and Tessada 2011). I then show that immigration is associated with an increased likelihood that women both work long hours and have recently given birth, a result suggestive of a childcare channel. Finally, I examine the characteristics that are associated with strong fertility relative to labor supply responses to immigrant inflows. Results suggest that women with graduate degrees are relatively more likely than women with just college degrees to respond to immigrant inflows by having an additional child.

Background

The relationships between childcare costs and fertility derived from even simple models are fairly complicated (Blau and Robins 1989). A decrease in child-rearing costs may increase desired fertility because of a standard price effect and increase desired labor supply by increasing the opportunity cost of time spent at home. Hazan and Zoabi (2015) presented a model showing how high-wage women might substitute housekeeping and babysitting services for their own time in household production, thereby allowing them to increase fertility without sacrificing their careers. On the other hand, the time costs associated with childbearing, such as time spent on late-night feedings, might offset the increase in desired labor supply, at least temporarily, for women who choose to have an additional child. It is also possible that the increase in desired labor supply is sufficient to induce a lower likelihood of childbearing if, for example, additional hours lead to promotions, which make women rethink original plans to have a third or fourth child. Lehrer and Kawasaki (1985) suggested that when adequate childcare is not affordable, women devote all their energy to their domestic roles, thus

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increasing fertility. Hence, the net effect of changes in childrearing costs on fertility is an empirical question.

A number of studies have considered the relationship between childcare subsidies and fertility. Examining a Swedish childcare subsidy reform, Mörk et al. (2013) found that lower childcare costs led to higher fertility. Gonzalez (2011) and Cohen et al. (2013) also uncovered fertility responses to changes in child benefits in Spain and Israel, respectively. It is difficult to determine whether these results extend to a U.S. context, where childcare subsidies are relatively small, at least for families in the middle and upper ends of the earnings distribution.

A handful of research has considered the effects of childcare costs on both employment and fertility outcomes using U.S. data. Mason and Kuhlthau (1992) examined mothers' perceptions of whether the availability of childcare constrained their employment and fertility choices. Blau and Robins (1989) analyzed how transitions among employment and fertility states respond to geographic variation in weekly childcare expenditures. Modeling female labor supply and fertility jointly within a dynamic model, Moffitt (1984) found that higher wages are associated with shifts in lifetime profiles of fertility and employment. Taking a different approach, Stolzenberg and Waite (1984) examined how variation in the individual-level association between fertility and labor force participation is explained by conditions in the local childcare market. All these studies provide results suggesting that lower childcare costs increase fertility but rely on potentially endogenous cost measures.

My analysis contributes to the childcare cost literature, but the main focus is on the effect of low-skilled immigration on fertility decisions of high-skilled native women. Despite large increases in the demand for childcare in the United States over the years, its price has risen slowly, which Blau (2001) attributed to a large "unexplained" increase in the supply of labor to the childcare market. Blau and Currie (2006) suggested that the large numbers of low-skilled immigrants may have contributed to this phenomenon.

Cortés (2008) showed that low-skilled immigration leads to reductions in prices of nontraded goods and services in major U.S. cities. Cortés and Tessada (2011) provided evidence that low-skilled immigration to the United States led to an increase in the hours worked among women at the top of the wage distribution. Similar conclusions have been reached for high-skilled native females in Spain (Farré et al. 2011), Italy (Barone and Mocetti 2011), and Hong Kong (Cortés and Pan 2013). Using harmonized data for Australia, Germany, Switzerland, the United Kingdom, and the United States, Forlani et al. (2015) showed that the positive effect of immigrant concentration on the number of hours worked by high-skilled women is stronger in countries with less-supportive family policies. Consistent with a role played by immigrant-induced changes in childcare prices, Amuedo-Dorantes and Sevilla (2014) found that low-skilled immigrant inflows result in changes in how mothers allocate their time with their children. In areas with larger immigrant concentrations, mothers spend less time on basic childcare tasks, such as bathing and feeding, but no less time on stimulating educational and recreational activities. Furtado and Hock (2010) showed that the correlation between fertility and labor force participation has become less negative in cities experiencing larger increases in their foreign-born populations. Furtado (2015) found that this change in the correlation explains about one-quarter of the 1980–2000 immigrant-induced increase in college-educated women's joint likelihood of having a

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child and working. To my knowledge, no other study has directly examined the effect of immigrant inflows on fertility rates of native-born women.

Data and Empirical Specification

Data

The main sample was drawn from the U.S. Census Bureau 1980, 1990, and 2000 public-use microdata sample (PUMS) files, and the 1970 sample was used to construct the instrumental variable. All data were obtained from the Integrated Public Use Microdata Series (IPUMS) (Ruggles et al. 2010).[1]

The analysis focuses on low-skilled immigrants of working age (age 20–64) and high-skilled non-Hispanic native females of childbearing age (22–42). Sharply differentiating immigrants and natives by skill minimizes the possibility of competition for jobs, which might directly affect female employment prospects. Analyzing non-Hispanic native females avoids nonmarket channels of influence, such as social norms and peer effects, which might arise from inflows of low-skilled immigrants to the United States, the bulk of whom are from Latin America and tend to have higher fertility rates. Skill classes are based on education. "Low-skilled" implies having, at most, a high school diploma and never having attended college; "high-skilled" refers to having completed a bachelor's degree or more. The native-born women who are still in school are dropped from the sample.

The underlying geographic sampling units defined by the Census Bureau have changed over time. The resulting inconsistencies in the degree to which the population of a metropolitan area is covered in the microdata files make it difficult to construct metro-level variables that are comparable across years. To reduce the potential influence of these inconsistencies on the estimates, I include in the analysis only those 117 metropolitan statistical areas (MSAs) that have consistent codes in the IPUMS between 1970 and 2000. Even MSAs with the same codes can consist of different counties and parts of counties in different years, but because counties that are in an MSA one year but not in another typically have small populations, remaining inconsistencies are unlikely to severely bias estimates. In practice, the consistent code restriction has very little impact on results because almost all the MSAs with a code in 1970 had consistent codes in 1980–2000.

[1] I ran models adding recent 2007–2011 American Community Survey (ACS) data to my sample, but standard errors were too large to draw any conclusions. One potential explanation for this is that the Great Recession induced more noise into fertility decisions. To examine this possibility, I reran the analysis using ACS data from before the recession, 2005–2007; this did not have any meaningful impact on results. I also considered the possibility that the instrumental variable, which is constructed from 1970 immigrant distributions, is simply not very predictive of immigrant concentrations 40 years later. This does not seem to be the case, either, because first-stage estimates were quite strong. Moreover, immigrant share coefficient estimates were statistically insignificant even in OLS models. These analyses lead to me to conclude that the explanation for the noisy estimates is related to how the ACS data are collected. Although census data are collected within a period of a few months in a particular year, ACS data are collected continuously over the course of several years. Thus, the constructed share low-skilled immigrant variable may be a very poor measure of immigrant concentration when women are making fertility decisions, especially those women sampled in 2007.

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Empirical Specification

Consider a basic fixed-effects model of the impact of low-skilled immigration using pooled data from multiple census years:

$$Y_{imt} = \beta_1 LSI_{mt} + \mathbf{X}_{imt}\beta_2 + \gamma_m + \gamma_{rt} + e_{imt},$$

where Y_{imt} is equal to 1 if woman i living in MSA m in year t has a child younger than 1 year in the household, and 0 otherwise.[2] The share of the working-age population that is low-skilled immigrant is denoted LSI.[3] MSA and region-year fixed effects are denoted γ_m and γ_{rt}, respectively, and e is an error term. The vector of controls, $\mathbf{X}$, includes a marriage dummy variable, a control for graduate degree, race dummy variables, and a full set of age dummy variables. To measure labor market opportunities for the high-skilled women in the sample, I also include the log of average yearly income among college-educated males living in the same MSA in the same year. Because many of these males will be married to the college-educated women in the sample, this variable will also partly account for within-family income effects. To measure norms and preferences regarding family life, I include the proportion of the woman's age group living in her MSA in the same year who are married, the proportion black, and the proportion nonblack and non-Hispanic. To minimize sampling error in constructing these variables, I use only two age groups (22–31 and 32–42), but results are robust to constructing three age groups and not separating into age groups at all. Standard errors are clustered on the MSA.

Immigrant location decisions cannot be taken as exogenous even conditional on the controls used in the analysis. Immigrants may be drawn to areas with a booming labor market for low-skilled workers and shrinking market for high-skilled female workers. The lower opportunity costs of time for high-skilled women may make childbearing more attractive. Also, immigrants may be attracted to cities with a high demand for childcare workers: that is, cities with high birth rates. For both reasons, ordinary least squares (OLS) estimates may yield upwardly biased estimates of the causal effect of immigration. Alternatively, immigrants may be attracted to cities with booming economies for both high-skilled and low-skilled workers. If high-skilled female workers are less likely to bear children when they have better labor market opportunities, then the least squares estimate of the effect of immigration will be biased downward. To address all these potential concerns, I rely on an instrumental variables (IV) approach to identify the causal impact of low-skilled immigration. Instrumental variables will also address attenuation bias in the estimated values of β due to measurement error in the share foreign-born variable.

[2] A mother who has given birth in the previous year but whose baby does not reside with her will not be counted in this fertility measure. Adoptive mothers and stepmothers, however, are treated as if they have given birth. Despite this, I use "having given birth" and "having had a baby in the past year" interchangeably with "having a young child in the household" throughout the article. I also examined the impact of immigration on the number of children under age 5 in the household as well as the likelihood that women have a child younger than ages 3 and 5. Results (available upon request) were robust across these different measures of fertility.

[3] I would have liked to use the share of low-skilled female immigrants, but this variable is almost perfectly correlated with the share of low-skilled male immigrants. Given that I cannot include both separately in the regressions, I chose to use the total share of low-skilled immigrants in order to be consistent with prior research in this literature.

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The instrument is based on the propensity of new immigrants to locate in areas with a relatively large existing concentration of coethnics (e.g., Bartel 1989). Following a similar line of reasoning as Card (2001), Cortés (2008), and Cortés and Tessada (2011), the instrument uses historical enclaves to predict the flow of subsequent migrants across MSAs. More specifically, the instrument for *LSI* is

$$INST_{mt} = \sum_b \frac{N^b_{m,1970}}{N^b_{1970}} \times \left[NLS^b_t - NLS^b_{t-10}\right].$$

For each country of birth, b, the first term in this equation represents the fraction of all immigrants from country b living in MSA m in 1970. The second term represents the net change in the total number of low-skilled working-age adults from country b between year t and the previous decade. Immigrants from countries listed as "unspecified" are not used in the construction of the IV. Also, I merge several countries in order to provide consistency over the different decades in the sample. Details are available upon request.

The necessary criteria for the instrument to be valid are very similar to those outlined by Cortés (2008). These are as follows: (1) the 1970 distribution of immigrants must be uncorrelated with differential changes in relative economic conditions affecting fertility across MSAs within the same region 10 to 30 years later; and (2) differential economic changes among MSAs should not affect the overall inflow of low-skilled immigrants to the United States. Although testing these assumptions directly is impossible, other studies have provided evidence pointing to their plausibility (e.g., Cortés and Tessada 2011).

Descriptive Statistics and Baseline Results

Table 1 provides descriptive statistics of the variables in the analysis, both in total and separated by whether immigrant share in a person's MSA is above or below the mean in the sample. Recall that the sample consists of non-Hispanic native-born women between the ages of 22 and 42 with at least a college degree. Interestingly, the women in cities with a high percentage of immigrants are slightly *less* likely to have given birth in the previous year, which might be explained by differences in the proportion of women who are married in these two types of cities. There are also more women defined as "other race"—the bulk of whom are Asian—in cities with a high percentage of immigrants, but the means of the other variables are very similar to each other in cities with high and low percentages of immigrants.

Table 2 presents baseline empirical results. To provide a sense for the basic cross-sectional relationship between the number of immigrants in a city and fertility, column 1 provides estimates from an OLS model with the full set of controls but without including MSA fixed effects. Estimated coefficients on the control variables imply that married women as well as women with a graduate degree are more likely to have an infant in the household. Black women are more likely than white women to have recently given birth, but women in the "other race" category are less likely. Women living in areas where more women of their age group are married also have higher fertility rates, even when their own marital status is held constant.

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Table 1 Descriptive statistics

	Total		Low Percentage Immigrant		High Percentage Immigrant	
	Mean	SD	Mean	SD	Mean	SD
Child	0.072	0.259	0.075	0.264	0.068	0.251
Share Working-Age Low-Skilled Immigrant	0.083	0.075	0.036	0.021	0.161	0.068
Age	32.322	5.666	32.228	5.663	32.478	5.667
Graduate Degree	0.284	0.451	0.275	0.446	0.299	0.458
Married	0.601	0.490	0.630	0.483	0.553	0.497
Black	0.090	0.287	0.089	0.285	0.092	0.288
Other Race	0.019	0.138	0.009	0.093	0.037	0.190
Log Mean Income of Males With College in MSA, Year	10.771	0.481	10.692	0.462	10.902	0.484
Proportion Married in Age Group, MSA, Year	0.601	0.134	0.630	0.116	0.553	0.147
Proportion Black in Age Group, MSA, Year	0.090	0.061	0.089	0.069	0.092	0.047
Proportion Other Race in Age Group, MSA, Year	0.019	0.048	0.009	0.009	0.037	0.073

Notes: The variable "Child" takes the value of 1 when a woman has a child younger than 1 year of age residing in the household. The variable "Other Race" is equal to 1 if the person is nonwhite, nonblack, and non-Hispanic. The low percentage immigrant sample includes people residing in MSAs where the fraction foreign-born is below the mean for the entire sample. The high percentage immigrant sample includes people residing in MSAs where the fraction foreign-born is at or above the mean for the entire sample. The "Share Working-Age Low-Skilled Immigrant" and "Log Mean Income of Males With College" variables are constructed by MSA and year. The "Proportion Married," "Proportion Black," and "Proportion Other Race" are constructed by MSA, year, and age group. The two age groups are 22–31 and 32–42. Means are weighted by census-provided person weights.

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Table 2 Baseline regressions: Dependent variable is child

	OLS 1	OLS 2	IV 3	IV 4
Share Working-Age Low-Skilled Immigrant	0.026**	0.065*	0.293**	0.119
	(0.009)	(0.033)	(0.112)	(0.090)
Share Working-Age Low-Skilled Immigrant × Age 29 to 35				–0.001
				(0.022)
Share Working-Age Low-Skilled Immigrant × Age 36+				0.149**
				(0.021)
Graduate Degree	0.005**	0.005**	0.005**	0.005**
	(0.001)	(0.001)	(0.001)	(0.001)
Married	0.115**	0.115**	0.115**	0.115**
	(0.001)	(0.001)	(0.001)	(0.001)
Black	0.004**	0.004**	0.004**	0.004**
	(0.001)	(0.001)	(0.001)	(0.001)
Other Race	–0.005*	–0.005*	–0.005*	–0.005*
	(0.002)	(0.002)	(0.002)	(0.002)
Log Mean Income of Males With College Degree	0.011	0.016	–0.011	–0.006
	(0.007)	(0.023)	(0.023)	(0.019)
Proportion Married in Age Group, MSA, Year	0.085**	0.157**	0.159**	0.148**
	(0.012)	(0.021)	(0.021)	(0.022)
Proportion Black in Age Group, MSA, Year	0.025	0.120*	0.066	0.048
	(0.015)	(0.053)	(0.060)	(0.051)
Proportion Other Race in Age Group, MSA, Year	–0.013	–0.423**	–0.447**	–0.259*
	(0.011)	(0.115)	(0.131)	(0.105)
Age Fixed Effects	Yes	Yes	Yes	Yes
Region-Year Fixed Effects	Yes	Yes	Yes	Yes
MSA Fixed Effects	No	Yes	Yes	Yes
First-Stage F (excluded instrument)			87.37	
Mean of Dependent Variable	0.072	0.072	0.072	0.072
N	607,790	607,790	607,790	607,790

Note: Standard errors, shown in parentheses, are clustered by MSA.
$^*p < .05$; $^{**}p < .01$

The simple OLS without MSA fixed-effects estimate of the immigration coefficient suggests that a 1 percentage point increase in the share of low-skilled immigrants in an MSA—with a mean percentage low-skilled immigrant in the sample of 8.3—is associated with only a 0.026 percentage point increase in the probability that a high-skilled native-born woman has an infant in the household. Not much credence should be placed on this figure given that there may be several unobserved city-level characteristics that are both attractive to immigrants and make high-skilled women prefer larger (or smaller) families. To address these city-specific time-invariant unobservable characteristics, MSA fixed effects are added in column 2. The estimated immigration

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coefficient is larger in this fixed-effects model, suggesting that in the cross-section, cities that tend to have large immigrant populations also tend to have lower fertility rates, even conditional on observable characteristics. The fixed-effects model suggests that a 1 percentage point increase in the share of low-skilled immigrants in an MSA is associated with a 0.065 percentage point increase in the probability that high-skilled women in that MSA give birth.

It is useful to think about timing in these specifications. All variables in the models are measured in the same year, and it is impossible that the foreign-born population in a given year has a causal impact on the probability that a woman gave birth the year before. However, the fixed-effects specification exploits within-MSA between-decade changes in the size of the foreign-born population. Although this measure changes discretely from decade to decade, the actual foreign-born population changes continuously between decades. Therefore, for example, the 2000 foreign-born population is likely a fine measure of the foreign-born population around 1997, when women were making pregnancy decisions about children born in the year 1999. Surely, the 2000 measure is better than the 1990 measure.[4]

Estimates from the MSA fixed-effects models will also be biased if time-varying determinants of fertility are correlated with the number of immigrants in a city. If, for example, low-skill industries are replacing high-skill industries in a city, we may observe increases in fertility rates among high-skilled women alongside large immigrant inflows not because immigrants are providing inexpensive childcare but because women face lower opportunity costs of leaving the labor force to raise children. Alternatively, if immigrants tend to move to cities with booming economies for both the low-skilled and high-skilled labor force, the MSA fixed-effects models will yield underestimates of the true causal impact of immigrant inflows.

The IV results shown in column 3 of Table 2 suggest that the second scenario is more likely. Note that the F statistic of 87.4 reveals a strong first-stage relationship. As can be seen in Table 8 in Appendix 1, the estimated first-stage coefficient on the instrument is positive and has a p value of less than .001. The second-stage estimate suggests that a 1 percentage point increase (which amounts to about 0.13 standard deviations) in the share of low-skilled immigrants in a city yields a 0.29 percentage point increase in the likelihood that a high-skilled woman has a child younger than 1 year in the household.

In my sample, the share of low-skilled immigrants increased from 0.066 in 1980 to 0.098 in 2000. Meanwhile, the proportion of high-skilled women who gave birth within the previous year increased from 0.0712 in 1980 to 0.0796 in 2000. The IV estimates imply that the 3.2 percentage point increase in low-skilled immigration can on its own more than explain the increase in fertility ($0.29 \times 3.2 + 7.12 = 8.05 > 7.96$). However, between 1980 and 2000, many factors—including increased female labor force participation rates and a decreasing gender wage gap—would have decreased fertility (Hoffman and Averett 2009). My results imply that if the share of low-skilled immigrants had stayed at its 1980 level, the likelihood of recently giving birth in the year

[4] Yearly estimates of the foreign-born population can be derived from the Current Population Survey (CPS) for years following 1994 but not before that. Another problem with CPS data is that sample sizes are significantly smaller than those in the census.

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2000 would have been 0.09 percentage points *lower* than in 1980 rather than 0.84 points higher.

The measure of fertility used in this article tells us whether immigrant inflows are associated with the probability of having a child in a particular year, but it is possible that large immigrant inflows change the timing of births without changing completed fertility. Addressing this issue is not as simple as using children ever born as the dependent variable because my identification strategy relies on cross-decade changes, and women's births typically do not fall neatly toward the end of any particular decade. Another problem is that 1990 is the last year in which the census asked for information on children ever born. To start, I will note that I am not overly concerned by this issue in light of new research showing that the long-term fertility responses to changes in unemployment rates are even larger than the short-term responses (Currie and Schwandt 2014). I would not expect the pattern to be significantly different when considering the dynamic responses of fertility to immigrant inflows.

Nevertheless, to assess whether immigrant inflows are likely to impact completed fertility, I examine how immigration affects women of different ages. The last column of Table 2 shows that it is women older than 35 whose fertility rates are most sensitive to immigrant inflows. The estimates suggest that a 1 percentage point increase in immigrant share increases birth likelihoods of 22- to 28-year-olds by (an imprecisely estimated) 0.12 percentage points, without much of a difference for 29- to 35-year-olds. On the other hand, the same 1 percentage point increase in immigrant share leads to a 0.27 percentage point increase in the likelihood that women older than 35 give birth (0.12 for the base sample of 22- to 28-year-old women plus an additional 0.15 points for those older than 35). Given that the oldest women in the sample cannot decrease future fertility to compensate for increases in current fertility, it seems likely that when women face immigrant-induced lower childcare costs, they do increase completed fertility.[5] Moreover, the smaller and statistically insignificant estimates of coefficients for younger women do not necessarily imply that younger women are not sensitive to immigrant inflows. In response to a decrease in childcare costs believed to be long lasting, all women may increase their desired number of children. However, older women must respond right away by giving a birth, whereas younger women have the option of waiting. This would suggest that the small estimates for younger women hide future increases in completed fertility.

Mechanisms

An Analysis of Childcare Labor Markets

The baseline estimates show that high-skilled women respond to immigrant inflows by increasing fertility. However, even if the estimates can be interpreted as causal, they do

[5] If women can anticipate future immigrant inflows (specifically from countries with large representations in their MSAs in 1970, given the IV identification), then it is possible that the 36- to 42-year-olds are merely compensating for previous decreases in fertility. To examine this, I considered the impact of future immigrant inflows on current period fertility. If women responded to future immigrant inflows by decreasing current period fertility, I would have expected a negative coefficient on these future immigrant inflows. Instead, I estimated a positive but statistically insignificant coefficient on the future immigrant share variable. These results are available upon request.

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not guarantee that immigrants affect fertility outcomes through childcare markets. As a first step toward showing that immigrants are in fact affecting fertility through childcare costs, I examine whether immigrant inflows lead to decreases in childcare costs as measured by wages of childcare workers. The wage bill accounts for between 60 % and 70 % of the operating expenses at formal and home-based childcare centers (Blau and Mocan 2002; Helburn and Howes 1996), and likely represents an even higher share of the final costs of informal childcare providers. Thus, it seems reasonable to use wages of childcare workers as a measure of the price of childcare.

Consider a basic fixed-effects model of the impact of low-skilled immigration again using pooled data from multiple census years:

$$w_{mt} = \alpha_1 LSI_{mt} + \alpha_2 Income_{mt} + \mu_m + \mu_{rt} + \varepsilon_{mt}.$$

The dependent variable w_{mt} is the log of the median hourly wage childcare workers in metropolitan area m in year t. An influx of low-skilled immigrants might mechanically reduce the median wage because their arrival results in more mass at the bottom of the wage distribution. To remove this possibility, I calculate median wages using a sample of native and immigrant workers living in the United States for at least 10 years. The variable $Income_{mt}$ denotes the log of income per capita among working-age male college graduates.[6] The other variables are defined as in Eq. (1). Regressions are estimated using the population of high-skilled women in the MSA-year as weights. Again, I keep only MSAs that are coded in the same way by the IPUMS between 1970 and 2000.

Table 3 presents results of this two-stage least squares (2SLS) analysis. All estimates are constructed by using the Card (2001) instrumental variables (IV) strategy described earlier. The estimated coefficient of –4.28 represents the percentage change in the median wage of childcare workers caused by a 1 percentage point increase in the size of the low-skilled immigrant population. This estimate is considerably larger than most existing estimates of the wage effects of low-skilled immigration (Card 2001; Friedberg and Hunt 1995). However, much of this research is based on examining broad skill classes rather than specific occupations. Childcare in particular is very labor-intensive compared with the larger low-skilled labor market, providing little room for capital adjustments. Childcare is also impossible to outsource to other countries.

As shown in the second column of Table 3, low-skilled immigration is associated with expansions in the share of the local workforce concentrated in the childcare occupation—a result suggestive of a labor supply (as opposed to demand) shock. Although statistically significant, the point estimate on immigration in the share of the labor force model is quite small, implying that low-cost immigrants tend to displace higher-cost natives in markets with larger immigrant inflows, but the change in the supply of childcare workers is mostly compositional.

As a way to check whether the estimated fertility effects described in the previous section are reasonable, I consider what these childcare wage impacts might imply for a

[6] College graduates are likely to be high demanders of household services and, for the most part, will have incomes that are not directly tied to wages in low-skill services markets. Females are not included in the income measure because their labor supply and earnings might be directly affected by wages of childcare workers. To account for top-coding, which was an issue only in 1980, I impute values for individuals whose income had been top-coded using a region-specific Pareto extrapolation.

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Table 3 Two-stage least squares (2SLS) regressions on household services markets

	Childcare		Private Households		Food Services	
	Median Wage	Proportion of Labor Force in Occupation	Median Wage	Proportion of Labor Force in Occupation	Median Wage	Proportion of Labor Force in Occupation
	1	2	3	4	5	6
Share Working-Age Low-Skilled Immigrant	–4.281**	0.044**	–0.811	0.025	–0.951**	0.088**
	(0.464)	(0.014)	(0.707)	(0.015)	(0.333)	(0.032)
Log Mean Income of Males With College in MSA, Year	0.797**	–0.010**	0.431**	–0.004†	0.503**	–0.022**
	(0.190)	(0.002)	(0.135)	(0.002)	(0.112)	(0.008)
Region-Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
MSA Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
First-Stage F (excluded instrument)	54.70	54.70	54.70	54.70	54.70	54.71
N	354	354	354	354	354	354

Notes: Standard errors, shown in parentheses, are clustered by MSA. Median wages are constructed using a sample of workers, including natives, who report working more than 0 hours in a typical week or worked more than 0 hours in the reference week and who had positive yearly wages in the previous year. The proportion of the labor force in occupation measures the number of workers who report having the occupation divided by the total number of workers (in the MSA and year).

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

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high-skilled woman employing a full-time childcare worker. According to the U.S. Bureau of Labor Statistics (BLS), the median pay of childcare workers in 2012 was \$19,510 per year (Bureau of Labor Statistics 2014). Thus, the 4.28 % decrease in childcare wages attributable to a 1 percentage point increase in immigrant share implies \$16.06 in monthly savings in childcare costs.

Recall that the same immigrant share increase results in a 0.29 percentage point increase in the probability of giving birth for the average woman in our sample (column 3 of Table 2). However, because only about 38.1 % of children under age 5 are cared for by a nonrelative on a consistent basis (Laughlin 2013), the 0.29 estimate might actually imply that for women who use purchased childcare, a 1 percentage point increase in immigrant share results in a 0.76 percentage point increase in the probability of having a child.

To determine whether these results make sense, I consider them in the context of those in a recently published study examining the fertility impacts of a change in a monthly child benefit in Israel. Cohen et al. (2013) found that a \$34 reduction in the monthly benefit for a marginal child reduces the probability of an incremental child by 0.99 percentage points, implying that a \$1 decrease in the cost of a child increases the probability of having one by 0.02 percentage points. This suggests that \$16.06 in monthly savings in childcare would result in a 0.47 percentage point increase in the probability of giving birth. If I make the very strong assumption that women in the United States respond to a change in the out-of-pocket costs of care by the same amount as women in Israel, then my 0.76 percentage point estimate of the fertility impact of immigration implies that almost two-thirds of immigration's impact operates via decreased out-of-pocket costs. The rest may be due to differences in quality or convenience of care provided by immigrant childcare workers or even decreases in prices of other substitutes for maternal care.

The remaining columns of Table 3 show the effects of low-skilled immigration on the wages and the share of the labor force working as housekeepers and food preparation workers. Immigrant inflows are associated with decreases in wages and increases in share of the labor force working in private households and in food preparation occupations, although estimates are not statistically significant for private household workers. Of course, my back-of-the-envelope calculation requires strong assumptions; thus, the estimate of how much of the immigrant effect operates through childcare wages must be interpreted with a great deal of caution. However, it is comforting that when I combine results from two separate regressions, they are consistent with existing literature.

One remaining potential concern with this analysis is that educated and high-income women demand a higher quality of care (Blau and Hagy 1998). If low-skilled immigrants generally provide low-quality care, then low-skilled immigrant inflows might not affect the cost of childcare services actually purchased by college-educated women. To my knowledge, data linking characteristics of childcare workers to characteristics of mothers do not exist. However, Blau and Mocan (2002) provided evidence that the cost of childcare is a positive function of its underlying objectively assessed quality. Thus, we can draw inferences on immigration's impact on the cost of services of various levels of quality by examining the effects of immigration on various quantiles of the

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childcare wage distribution. If immigrants generally provide low-quality care, then we would expect them to have the strongest impacts at the lower ends of the childcare wage distribution. Conversely, if they generally provide high-quality care, then we would expect the largest impacts to be at the top of the wage distribution.

Table 9 in Appendix 1 shows that low-skilled immigrant inflows impact wages at the bottom of the distribution more than wages at the top of the childcare wage distribution. However, a 1 percentage point increase in the immigrant share results in a 3.04 % decrease in wages even at the seventy-fifth percentile of the wage distribution, suggesting that low-skilled immigrant inflows impact childcare markets across the quality distribution. Interestingly, low-skilled immigrant inflows impact wages of housekeepers at the bottom end of the wage distribution but not the middle or top. Although the effects are smaller than for childcare workers, immigrant inflows have impacts on wages of food preparation workers across the wage distribution.

Heterogeneous Impacts of Immigrant Inflows on Fertility

For further evidence that immigrant inflows are influencing native fertility rates through childcare markets, I consider whether the types of women who are likely to be more sensitive to changes in childcare costs are in fact more responsive to immigrant inflows when making fertility decisions. The first two columns of Table 4 allow us to compare immigration impacts on women with a graduate degree with those on women with a college degree only. Although the more highly educated have higher household incomes and thus may be less sensitive to changes in childcare costs (see Cohen et al. 2013), they are also less likely to live near family members (Molloy et al. 2011) and more likely to work long hours, making them more dependent on nannies and other nonfamily full-time childcare providers. If these high-skilled women are more likely to work in jobs that often require unplanned late nights at the office (and have spouses with similar types of jobs), they may be especially likely to use nannies, who are often foreign-born, as opposed to formal childcare centers. Results shown in the first two columns of Table 4 suggest that fertility rates of women with graduate degrees are indeed more responsive to immigrant inflows than fertility rates of women with just college degrees.[7] A 1 percentage point increase in immigrant share results in a 0.46 percentage point increase in the likelihood that women with graduate degrees have a child but only a 0.21 percentage point increase for women with a college degree only.

Columns 3 and 4 of Table 4 present estimates of immigrant impacts on samples of married women and unmarried women. Married women respond to a 1 percentage point increase in immigrant share with a 0.44 percentage point increase in the likelihood of giving birth, but unmarried women respond with a practically 0 but statistically significant 0.04 percentage point increase. If a much smaller proportion of unmarried

[7] I include only women with at least a college degree in the main sample because of concerns that immigrant inflows directly impact the wages and types of jobs available to low-skilled native-born women. Given that college-educated women are not easily substituted with low-skilled immigrant labor, I am more comfortable arguing that the main effect of immigration on these women operates through childcare markets. Nevertheless, in Table 10 in Appendix 1, I also compare impacts for women with less than a college degree. Notice that immigrant inflows have smaller impacts on women without a college degree and no impact on the fertility decisions of women with less than a high school diploma.

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Table 4 Heterogeneous responses to immigrant inflows: Dependent variable is child

	Education		Marital Status		Race			Full Sample
	College	Graduate Degree	Unmarried	Married	White	Black	Other	
	1	2	3	4	5	6	7	8
Share Working-Age Low-Skilled Immigrant	0.205†	0.456**	0.035†	0.450*	0.402**	−0.350**	−0.129	
	(0.119)	(0.115)	(0.020)	(0.189)	(0.124)	(0.126)	(0.134)	
Share Working-Age Low-Skilled Immigrant From Low-Childcare Countries								−0.406 (0.345)
Share Working-Age Low-Skilled Immigrant From High-Childcare Countries								1.632* (0.784)
Mean of Dependent Variable	0.072	0.074	0.117	0.006	0.075	0.054	0.058	0.072
N	432,136	175,654	371,269	236,521	544,776	51,059	11,955	607,790

Notes: The married sample, shown in column 4, consists of married women with a spouse present. The unmarried sample, shown in column 3, consists of all others, including cohabiting, divorced, widowed, and never-married women. All regressions are run using 2SLS and include the full set of controls shown in Table 2, including MSA, region-year, and age fixed effects.

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

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women are likely to plan their pregnancies, it makes sense that they would be less sensitive to changes in childcare costs. Columns 5–7 of Table 4 present estimates of immigrant impacts constructed from samples separated by race. Results suggest that it is only white women's fertility patterns that increase in response to immigrant inflows.[8]

For further evidence that this analysis is measuring the impact of immigrant inflows via childcare markets, I exploit the fact that immigrants from certain countries are substantially more likely to work as childcare providers than immigrants from other countries. As a first step, I use 1990 census data to construct the proportion of immigrants from each origin country that list childcare as their occupation in the census. I define as "high-childcare" those origin countries in the top quartile of this distribution. All other origin countries are defined as "low-childcare."[9] After determining which immigrants are from low- and high-childcare countries, I then construct two new variables that are used in place of the share of working-age low-skilled immigrant variable: "share working-age low-skilled immigrants from high-childcare countries" and the corresponding share from low-childcare countries. Because there are relatively few immigrants from many of the countries with a large share of workers in childcare, low-skilled immigrants from high-childcare countries represent only about 14 % of the low-skilled immigrant workers in my sample. To instrument for the two immigrant share variables, I use the original instrument structure but construct one IV using only the high-childcare countries and the other one using only the low-childcare countries.

Regression results are shown in column 8 of Table 4. Holding constant the share of working-age low-skilled immigrants from low-childcare countries, a 1 percentage point increase in the share of working-age low-skilled immigrants from high-childcare countries increases the likelihood that a woman gives birth by 1.63 percentage points. On the other hand, holding constant the share of immigrants from high-childcare countries, an increase in the share from low-childcare countries has no statistically significant impact on birth rates.

Next, I examine whether immigrant inflows differentially impact women's decisions to have a first, second, third, or higher-order birth. Women who already have children are likely to be more knowledgeable about childcare markets when making new fertility decisions, and so one might expect immigrant inflows to have weaker effects on the decision to have a first child than a higher-order child. To test this hypothesis, I estimate the impact of immigrant inflows on the likelihood of giving birth (in the previous year) for women with different numbers of children in the household (before the previous year).

Results, shown in Table 5, suggest that immigration is not associated with any statistically meaningful change in the likelihood of having a first child (column 1) but has a strong impact on the likelihood of having a second child (column 2). The estimated immigrant share coefficient is much smaller when estimated from a sample

[8] In fact, these estimates imply that black women decrease fertility in response to immigrant inflows. Additional analyses, however, suggest that this result is not robust. In models with year fixed effects instead of region-year fixed effects, the estimated immigration coefficient for blacks is significantly smaller in magnitude and not statistically significant. Results from models with interactions between the race variables and immigrant inflows suggest that just like white women, black women increase fertility in response to immigration—just not as much.

[9] Proportions were constructed using data on immigrants in the labor force. A list of countries in each of the categories is provided in Appendix 2.

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of women who already have two children (column 3) than when estimated using women with only one child (column 2), but relative to the mean of the dependent variables, the effects are almost identical. For women with three or more children in the household, there is no evidence that new fertility decisions respond to childcare costs (column 4), but this may be an artifact of the small sample. Taken together, these findings are broadly consistent with the hypothesis that women with children are more responsive to immigrant inflows when deciding to have an additional child. However, an alternative explanation is that childless women may not be open to having children, regardless of their costs. Thus, caution must be used when interpreting the results in Table 5.

Labor Supply Responses to Immigrant Inflows

Although this article presents evidence suggesting that high-skilled native-born women respond to immigrant inflows by increasing fertility, a growing literature shows that women respond to immigrant inflows by increasing labor supply (Barone and Mocetti 2011; Cortés and Pan 2013; Cortés and Tessada 2011; Farré et al. 2011; Forlani et al. 2015). It is possible that lower childcare costs allow women both to have more children and to work long hours. Furtado and Hock (2010) showed that immigrant inflows to an MSA result in a less negative correlation between fertility and labor force participation in that MSA. However, it is also possible that some women respond to lower childcare costs by working more hours and not changing or even decreasing their desired number of children, while other women respond with increases in fertility even if it comes at the expense of working long hours in the labor market, at least temporarily.

Table 5 2SLS regressions by parity: Dependent variable is child

	Number of Children			
	0 vs. 1 Child	1 vs. 2 Children	2 vs. 3 Children	3+ Children
	1	2	3	4
Share Working-Age Low-Skilled Immigrant	0.112	0.842**	0.247**	–0.026
	(0.090)	(0.226)	(0.091)	(0.192)
Mean of the Dependent Variable	0.0561	0.165	0.0563	0.0438
N	341,243	99,766	115,492	51,289

Notes: The first column shows results of a regression conducted on a sample of women with 0 children before the previous year. The second column uses a sample with one child before the previous year. The third column uses a sample with two children before the previous year, and the last column uses a sample with three or more children before the previous year. Thus, the first column examines the impact of immigrant inflows on the decision to have a first child. The second, third, and fourth columns examine the impact on the decision to have a second, third, and fourth child, respectively. Standard errors, clustered by MSA, are shown in parentheses. All regressions include the full set of controls shown in Table 2, including MSA, region-year, and age fixed effects.

**$p < .01$

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To examine this, I start by reproducing the general labor supply results from the literature using my data and basic empirical specification. The first column of Table 6 shows the impact of immigrant inflows on the probability of working more than 0 hours in a typical week. Column 2 shows the impact on the probability of working 35 hours or more, column 3 the impact on 40 or more hours, and column 4 the impact on 50 or more hours. Consistent with the findings of Cortés and Tessada (2011), the largest effects are on labor supply at the high end of the hours-of-work distribution. Interestingly, immigrant inflows tend to decrease the probability of working more than 0 hours in a typical week.[10] Although this pattern may be surprising, it is consistent with a story of mothers with very young children temporarily exiting the labor force to care for children but working very long hours upon returning to the work force.

Regardless of whether women respond to immigrant inflows by working more hours or having additional children, we should expect increases in the joint likelihood of working in the labor market *and* recently having given birth if immigration operates via childcare markets. To test this empirically, I start by estimating the effect of immigrant inflows on the likelihood that a high-skilled woman both works more than 0 hours in a typical week and has an infant. As can be seen in column 5 of Table 6, a 1 percentage point increase in immigrant share results in a 0.19 percentage point increase in the joint likelihood, about 3 % of its mean in the sample. Given that immigrant inflows are associated with decreases in labor force participation rates, the fact that they increase the joint likelihood of participating and giving birth suggests that the decreases in the probability of working, shown in column 1 of Table 6, can be explained by the increases in the probability of giving birth.

In the next column, immigrant inflows are used to estimate the joint likelihood of working 50 or more hours in a typical week and having given birth in the past year. Although the estimated coefficient on immigrant share appears rather small, the mean of the dependent variable in this sample is only slightly more than 0.5 %. In fact, when compared with the mean, the impact of immigrants on the likelihood of working long hours and having a baby is about double the impact of immigrants on working at all and having a baby.

All these results are suggestive of immigration making it easier for high-skilled women to combine their roles as mothers and workers. However, it is unclear whether the increases in the joint likelihoods are driven mostly by increases in fertility, increases in working hours, or some combination of the two. In fact, the relative increases in fertility and labor supply could differ for different groups of women.

To explore this issue, I focus my study of labor supply on the decision to work more than 50 hours in a typical week and compare women who have graduate degrees with women who have a college degree only. The first and third columns of Table 7 simply reproduce results from Table 5 showing that more highly skilled women are more likely to give birth in response to

[10] Cortés and Tessada (2011) estimated negative but statistically insignificant effects of immigrant-induced increases in the low-skilled labor force on labor force participation.

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Table 6 Labor supply responses to immigrant share

	Usual Hours per Week					
	More Than 0	35+	40+	50+	0+ and Recent Birth	50+ and Recent Birth
	1	2	3	4	5	6
Share Working-Age Low-Skilled Immigrant	–0.398**	–0.150	0.220	0.846*	0.192*	0.039†
	(0.123)	(0.125)	(0.216)	(0.329)	(0.085)	(0.023)
Graduate Degree	0.064**	0.073**	0.065**	0.059**	0.011**	0.004**
	(0.002)	(0.004)	(0.004)	(0.002)	(0.001)	(0.000)
Married	–0.123**	–0.229**	–0.212**	–0.076**	0.091**	0.009**
	(0.002)	(0.004)	(0.004)	(0.002)	(0.001)	(0.000)
Black	0.046**	0.093**	0.059**	–0.056**	0.011**	–0.001†
	(0.004)	(0.005)	(0.012)	(0.005)	(0.001)	(0.000)
Other Race	0.003	0.036**	0.037**	–0.004	–0.003	0.000
	(0.007)	(0.009)	(0.008)	(0.008)	(0.002)	(0.001)
Log Mean Income of Males With College in MSA, Year	–0.018 (0.025)	0.044 (0.037)	0.057 (0.050)	0.066 (0.050)	–0.019 (0.020)	0.005 (0.004)
Proportion Married in Age Group, MSA, Year	–0.003 (0.033)	–0.237** (0.058)	–0.246** (0.055)	–0.040† (0.022)	0.129** (0.019)	0.013** (0.005)
Proportion Black in Age Group, MSA, Year	0.212** (0.057)	0.200* (0.089)	0.068 (0.114)	–0.206** (0.078)	0.037 (0.049)	–0.003 (0.010)
Proportion Other Race in Age Group, MSA, Year	–0.421** (0.132)	–0.821** (0.266)	–0.498* (0.254)	0.500** (0.126)	–0.434** (0.099)	–0.059* (0.025)
Age Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Region-Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes

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Table 6 (continued)

	Usual Hours per Week					
	More Than 0	35+	40+	50+	0+ and Recent Birth	50+ and Recent Birth
	1	2	3	4	5	6
MSA Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Mean of Dependent Variable	0.879	0.701	0.624	0.130	0.057	0.006
N	607,790	607,790	607,790	607,790	607,790	607,790

Notes: Standard errors are shown in parentheses. The dependent variable in column 1 is a dummy variable equal to 1 if the woman works more than 0 hours in a typical week; the dependent variable in column 2 equals 1 if the woman works at least 35 hours in a typical week; in column 3, at least 40 hours; and in column 4, at least 50 hours. The dependent variable in column 5 is a dummy variable equal to 1 if the woman works more than 0 hours and has given birth in the previous year. In column 6, the dependent variable equals 1 if the woman works 50 or more hours and has given birth in the previous year. All regressions are run using 2SLS.

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

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Table 7 2SLS fertility and labor supply regressions by education

	College Only		Graduate Degree	
	Child	Usually Work 50+ Hours	Child	Usually Work 50+ Hours
	1	2	3	4
Share Working-Age Low-Skilled Immigrant	0.205†	0.907**	0.456**	0.655*
	(0.119)	(0.331)	(0.115)	(0.333)
Mean of Dependent Variable	0.072	0.115	0.074	0.169
N	432,136	432,136	175,654	175,654
Ratio of Effect on Fertility to Effect on Work	0.23		0.70	
Ratio of Effect on Fertility to Effect on Work (adjusted by means of dependent variable)	0.36		0.52	

Notes: Standard errors are shown in parentheses. Regression results shown in the first two columns are constructed using a sample of women with no more than a college degree, and the last two columns are constructed using a sample of women with a graduate degree. The ratio of the effect on fertility to the effect on work divides the estimated coefficient in the first row of column 1 (or column 3) by the estimated coefficient in column 2 (or 4). When adjusted by the means of the dependent variable, the estimated fertility coefficient is first divided by average fertility of the given sample, and the estimated labor supply coefficient is first divided by the mean labor supply of the given sample. For example, for the college-only sample, the adjusted ratio of 0.36 is equal to (0.205 / 0.072) / (0.907 / 0.115). All regressions are run using 2SLS and include the full set of controls shown in Table 2, including MSA, region-year, and age fixed effects.

†$p < .10$; *$p < .05$; **$p < .01$

immigrant inflows than less-skilled women. The second and fourth columns show labor supply responses of these two groups. Interestingly, women with graduate degrees are less sensitive to immigrant inflows than women with just a college degree when it comes to the probability of working long hours. Although the increase in the probability of working long hours is larger than the increase in the probability of giving birth for both groups of women, if we consider the ratio of the estimated immigrant inflow coefficients and compare this ratio across the two groups, the relative fertility response is stronger for women with graduate degrees (.70 > .23). However, although both groups of women are about equally likely to have given birth in the previous year, Table 7 shows that women with graduate degrees are more likely to work long hours. If before taking the ratio of coefficients, we divide the fertility coefficient by the proportion of women with a small child in the home and divide the labor supply coefficient by the proportion of women who work more than 50 hours, the relative fertility response to immigrant inflows remains stronger for women with a graduate degree than women with just a college degree (.52 > 36), but the difference is smaller.[11]

[11] In a similar analysis, I compared married with unmarried women. Not surprisingly, married women were significantly more likely to respond to immigrant inflows by increasing fertility relative to labor supply. However, the relationship was not robust to weighting the estimates by the means of the dependent variables.

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Conclusion

This study builds on a growing body of work highlighting the potentially beneficial effects that immigration has on natives (Barone and Mocetti 2011; Cortés 2008; Cortés and Pan 2013; Cortés and Tessada 2011; Farré et al. 2011; Forlani et al. 2015). In order to isolate a causal impact of immigration, I relied on a common instrumental variables approach to account for the simultaneity of the location decisions of new migrants with respect to local labor market conditions. Using settlement patterns predicted from historical enclaves as instruments, I found that low-skilled immigration to the United States between 1980 and 2000 led to substantial reductions in the cost of market-provided childcare and that high-skilled native-born women responded with increases in fertility.

The popular press has raised concerns about the so-called Opt-out Revolution (Belkin 2003; Wallis 2004) and women still being unable to "have it all" (Slaughter 2012). These articles suggest that combining work and family responsibilities remains very difficult for women on the career track. By contrast, Goldin's (2004) assessment of detailed cohort data showed that relative to older cohorts, women graduating from college in the 1980s have been significantly better able to combine both career and family. This article suggests that women are in fact facing smaller tradeoffs when making fertility and labor supply decisions and that this may have been partly driven by the continuing flow of low-skilled immigrant workers into the United States.

The analysis provides a potential explanation for women's continued underrepresentation in top positions in business and academia despite the many new family-friendly policies over the years. Although policies that make it easier to combine work and family (such as subsidized childcare) may increase the amount of time women spend working in the labor market, those same policies also may increase the likelihood of having more children. In fact, the analysis in this article suggests that the very women who are most likely to break the glass ceiling—older women with graduate degrees—are the ones whose fertility decisions are most likely to respond to changes in childcare costs, at least the changes induced by immigrant inflows.

The findings in this article also have important implications for countries facing low fertility rates, such as southern European countries and Japan. Immigrants directly increase the size of the labor force; and given their high fertility rates, they tend to directly increase future population. My analysis suggesting that immigrant inflows also increase fertility rates of natives, particularly native women with graduate degrees, provides an additional avenue through which immigration policy can remedy below-replacement fertility rates.

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Appendix 1

Table 8 First-stage regression

Variables	Share Working-Age Low-Skilled Immigrant
IV / 100,000	0.008**
	(0.001)
Graduate Degree / 100,000	–4.377
	(7.477)
Married / 100,000	0.899
	(1.161)
Black / 100,000	–0.108
	(0.466)
Other Race / 100,000	–0.052
	(0.327)
Log Mean Income of Males With College Degree	0.105**
	(0.029)
Proportion Married in Age Group, MSA, Year	–0.012
	(0.012)
Proportion Black in Age Group, MSA, Year	0.186**
	(0.048)
Proportion Other Race in Age Group, MSA, Year	–0.139*
	(0.067)
Age Fixed Effects	Yes
Year Fixed Effects	Yes
MSA Fixed Effects	Yes
First-Stage F (excluded instrument)	87.37
N	607,790

Note: The IV is described in the text.

*$p < .05$; **$p < .01$

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Table 9 Impacts of immigration at various points of wage distribution

	Log of Wages at the:		
	25th Percentile	50th Percentile	75th Percentile
	1	2	3
Panel A. Childcare			
Share working-age low-skilled immigrant	–5.965**	–4.281**	–3.042**
	(0.688)	(0.464)	(0.479)
N	354	354	354
Panel B. Housekeeping			
Share working-age low-skilled immigrant	–1.448**	–0.811	–1.126
	(0.423)	(0.707)	(0.723)
N	354	354	354
Panel C. Food Services			
Share working-age low-skilled immigrant	–1.138**	–0.951**	–0.741*
	(0.270)	(0.333)	(0.309)
N	354	354	354

Notes: All estimates shown in this table are constructed from separate 2SLS regressions. All regressions include a control for (log) annual wage income among male college graduates as well as MSA and region-year fixed effects. Column 2 shows impacts of low-skilled immigrant inflows on median wages of the three household services industries. Columns 1 and 3 present estimates of the effect of immigrant inflows on wages at the 25th and 75th percentile, respectively, in the three industries.

$^{*}p < .05$; $^{**}p < .01$

Table 10 Heterogeneous responses to immigrant inflows by education: Dependent variable is child

	<High School Diploma	High School Diploma	Some College	College Degree	Graduate Degree
	1	2	3	4	5
Share Working-Age Low-Skilled Immigrant	0.089	0.175**	0.177*	0.204†	0.457**
	(0.159)	(0.066)	(0.074)	(0.118)	(0.114)
N	220,998	822,818	680,148	432,136	175,654
Mean of Dependent Variable	0.06	0.061	0.069	0.072	0.074

Notes: All estimates shown in this table are constructed from separate 2SLS regressions. Results in the first column are constructed using a sample of women with less than a high school diploma. The second column sample includes women with a high school diploma only. The third column includes women with some college completed but no degree. The fourth and fifth columns reproduce results shown in Table 5 for convenience. All regressions include the full set of controls shown in Table 2, including MSA, region-year, and age fixed effects.

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

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Appendix 2

High-childcare countries include (from lowest concentration of childcare workers to highest):

Indonesia, Brazil, Colombia, Spain, France, Argentina, Algeria, British West Indies, Ireland, Fiji, Wales, Norway, Uruguay, Peru, Chile, El Salvador, Belize/British Honduras, Sierra Leone, Liberia, Sri Lanka (Ceylon), Denmark, Honduras, Sudan, Bolivia, Guatemala, Bermuda, Cameroon, Greenland, and Paraguay.

Low-childcare countries include (from lowest concentration of childcare workers to highest):

Albania, Senegal, Tunisia, Uganda, Qatar, Yemen, PDR (South), Nepal, St. Helena and Ascension, Cyprus, United Arab Emirates, Lithuania, Zimbabwe, Latin America, ns, Saudi Arabia, Bulgaria, Yemen Arab Republic (North), Oman, Falkland Islands, Somalia, Morocco, Hungary, Vietnam, Laos, Ghana, Greece, Lebanon, Nigeria, Egypt/United Arab Rep., Yugoslavia, Turkey, Czechoslovakia, India, Syria, South Africa (Union of), China, Romania, Cuba, USSR/Russia, Western Samoa, Italy, Libya, Tanzania, Korea, Portugal, Philippines, New Zealand, Iraq, Kuwait, Jordan, Ethiopia, Thailand, Cambodia (Kampuchea), Haiti, Iran, Singapore, American Samoa, Pakistan, Israel/Palestine, Canada, Dominican Republic, Japan, Burma (Myanmar), Australia, Malaysia, Afghanistan, Latvia, Panama, Scotland, Mexico, Germany, Netherlands, Belgium, Poland, Tonga, Venezuela, Finland, Cape Verde, Switzerland, Sweden, Jamaica, Kenya, Austria, England, Ecuador, Costa Rica, and Nicaragua.

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