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ABSTRACT

Can Immigrants Help Women "Have it All"? Immigrant Labor and Women's Joint Fertility and Labor Supply Decisions*

This paper explores how inflows of low-skilled immigrants impact the tradeoffs women face when making joint fertility and labor supply decisions. I find increases in fertility and decreases in labor force participation rates among high skilled US-born women in cities that have experienced larger immigrant inflows. Most interestingly, these changes have been accompanied by decreases in the strength of the negative correlation between childbearing and labor force participation, an often-used measure of the difficulty with which women combine motherhood and labor market work. Using a structured statistical model, I show that the immigrant-induced attenuation of this negative correlation can explain about 24 percent of the immigrant-induced increases in the joint likelihood of childbearing and labor force participation in the U.S. between the years 1980 and 2000.

JEL Classification: D10, F22, J13, J22, R23

Keywords: child care, fertility, labor force participation, immigration, tetrachoric correlation

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1 Introduction

The highly time-intensive nature of childrearing implies a tradeoff between fertility and labor supply, particularly for women given their traditional role of performing household work (Becker 1985; Willis 1973). Within-country empirical analyses indicate a consistently negative association between fertility and female labor force participation. However, Engelhardt, Kögel, and Prskawetz (2004) find that this relationship has weakened substantially since the 1960s, particularly in the United States, suggesting that women are finding it easier to combine their roles as workers and mothers.

Cross-country studies reveal a positive correlation between fertility and labor force participation since the 1980s (Adsera 2004; Brewster and Rindfuss 2000). There is evidence suggesting that high unemployment and unstable employment contracts in Southern Europe combined with the generous maternity benefits in Scandinavia can at least partially explain why countries with the highest fertility rates also have the highest female labor force participation rates (Adsera 2004). Cross-country differences in the cost, availability, and quality of child care, driven mostly by child care subsidies, might also explain why women in certain countries can more easily combine motherhood with labor force participation (Brewster and Rindfuss 1996). Formal child care is typically provided by the private market in the US, and any childcare subsidies are targeted to single mothers at the margin of receiving welfare payments (Blau and Tekin 2007). However, in light of the evidence that immigrant inflows are associated with decreases in the price (Furtado 2014) and likely improvements in the convenience of child care (Cortes and Pan 2013), this paper examines whether and how the increased immigration to the US between 1980 and 2000 has contributed to the decreasing tradeoffs women face when making joint work and fertility decisions.

Consistent with the notion that combining mother and worker roles has become easier over time, fertility has been increasing for college-educated women (Shang and Weinberg 2012), the very women with the highest labor force participation rates. Although home appliances, such as microwave ovens, and time-saving products, such as frozen foods, have been found to play a role in explaining increased female labor force participation rates (Greenwood et al. 2005, Coen-Pirani et al. 2010), most of the diffusion of these technologies seems to have occurred before 1980 (Greenwood et al. 2005). Hazan and Zoabi (2014) present evidence of a U-shaped relationship between education and fertility among the newest cohorts of U.S. women. They explain this with a model showing that high wage women substitute housekeeping and babysitting services for their own time in household production thereby allowing them to increase fertility without sacrificing their careers. The authors' empirical results suggest that this new pattern is attributable to the change in the cost of child care relative to high skilled women's wages.

Another line of research investigates the relationship between childcare costs and the likelihood that mothers work (Blau and Robins 1988; Connelly 1992, Kimmel 1998, and Kornstad and Thoreson 2007). In a survey article, Blau and Currie (2006) highlight the wide-ranging estimated elasticities of employment with respect to child care costs and attribute cross-study discrepancies largely to variation in biases from different estimation strategies. Newer studies estimate mothers' labor supply decisions by exploiting cross time and regional variation in the price or availability of care resulting from differences in the implementation of childcare-related public policies (Baker, Gruber and Mulligan 2008; Cascio 2009, Lefebvre and Merrigan 2008, Lundin et al. 2008, Havnes and Mogstad 2011). While these studies generally find that decreased childcare costs increase maternal labor supply, the estimates differ quite substantially. In fact, some studies find little to no causal impact of subsidized child care on mother's employment (Cascio 2009; Havnes and Mogstad 2011). It may be that subsidized child care simply crowds out informal childcare arrangements (Havnes and Mogstad 2011), but it is also possible that women respond to less expensive child care by having an additional child which then depresses labor supply, perhaps temporarily. Exploiting variation generated from a Swedish childcare subsidy reform, Mörk, Sjögren, and Svaleryd (2013) find that lower childcare costs lead to more childbearing.

Another potential source of variation in the price, convenience, and quality of child care comes from cross-city differences in the availability of low-skilled immigrant labor. My analysis complements a growing literature, pioneered by Cortes and Tessada (2011), examining the relationship between immigration and labor supply decisions of highly skilled native women. Cortes and Tessada find that low-skilled immigration to large U.S. metropolitan areas led to increases in the number of hours worked by women at the top of the wage distribution. Similar conclusions have been drawn for Spain (Farre, Gonzalez, and Ortega 2011), Italy (Barone and Mocetti 2011), and Hong Kong (Cortes and Pan 2013).

Fertility rates of high skilled US born women have also been shown to increase in response to immigrant inflows (Furtado 2014). Consistent with the labor supply impacts of immigration literature, Furtado (2014) finds increased likelihoods of women working 50 hours a week or more in response to immigration, but perhaps paradoxically, she also finds decreases in labor force participation rates, a result potentially attributable to women temporarily leaving the labor force upon giving birth but then working long hours upon returning.[1]

Regardless of whether women respond to immigrant-induced decreases in childrearing costs by increasing labor supply and decreasing fertility, increasing fertility and decreasing labor supply, or even increasing both, lower childcare costs should make it less important to leave the labor force after giving birth. Thus, we should expect that in cities with larger immigrant inflows, the negative correlation between childbearing and labor force participation should become less strong. This is precisely the result in Furtado and Hock (2010). However, the problem with using correlations to measure the difficulty with which women combine their roles as worker and mother is that correlations cannot directly quantify changes in observable outcomes of individuals.

In this paper, I start by considering the direct impact of immigration on the likelihood that a woman chooses to both give birth *and* participate in the labor market shortly thereafter, certainly a more tangible measure of mother-worker role compatibility. Next, I estimate the impact of immigrant inflows on the likelihood of childbearing, the likelihood of participating in the labor force, and the correlation between the two. The structure of the model allows me to then combine these estimates to construct an indirect measure of the impact of immigration on the joint likelihood of childbearing and labor force

[1] Cortes and Tessada's (2011) estimates of the impact of immigration on labor force participation of high skilled women in the U.S. are also negative, but statistically insignificant. Barone and Mocetti (2011) estimate positive but insignificant effects in their study of Italy, while Farre et al. (2011) and Cortes and Pan (2013) do find a positive and statistically significant relationships for high skilled females in Spain and Hong Kong, respectively. My results may differ from all of these given that my sample consists only of women of childbearing age.

participation. I then decompose this effect into three components: the effect on childbearing, the effect on labor force participation, and the effect on the correlation between the two. In so doing, I can then determine how much of the increase in the joint likelihood of childbearing and working is attributable to a decrease in the tradeoffs women face when making fertility and labor supply decisions.

The difficulty with any empirical analysis of immigration that exploits cross-city variation in foreign-born concentrations is that immigrants do not choose where to live randomly. They may be more likely to reside in areas with an industry mix more (or less) conducive to combining work and family for highly skilled women. It is also plausible that immigrants tend to move to areas with more demand for childcare services precisely because women in these areas want to both have larger families and to work. Following other studies in this literature, I take an enclave-based instrumental variables approach pioneered by Card (2001) which exploits historical settlement patterns of immigrants from different countries.

Using U.S. Census data from 1980 to 2000 in conjunction with 1970 data to construct the enclave-based instrument, I find that inflows of low-skilled immigrants between 1980 and 2000 resulted in a higher joint likelihood of childbearing and labor force participation among high-skilled women. These increases were a result of increases in fertility that were accompanied by reductions in labor force participation (LFP) rates. I find that that about 24 percent of the impact of immigration on the joint likelihood of childbearing and working was attributable to a weakening of the negative fertility-work correlation. This suggests that indeed low skilled immigration has made it easier for women to combine work and family responsibilities. Additional analyses suggest that these results are quite robust and are unlikely to be driven by omitted variable or selection bias.

The remainder of the paper is organized as follows. In Section 2, I describe the model of decision making that underpins my investigation of childbearing and labor force participation patterns. A brief description of the data used in the analysis follows in Section 2. After presenting the main results in Section 3, I discuss how the estimated parameters may be interpreted. I also conduct specification checks concerning the validity of the estimation method and the extent to which geographic selection might affect the results. Finally, Section 4 provides additional discussion and concluding remarks.

2 Fertility and Labor Force Participation Decisions

2.1 The Model

Female employment and fertility decisions can be described using the standard bivariate probit model framework:

$$C^*_{igmt} = \beta_1 LSI_{mt} + \boldsymbol{\omega}'_1 \mathbf{v}_{igmt} + \varepsilon^C_{igmt} \quad (1)$$
$$L^*_{igmt} = \beta_2 LSI_{mt} + \boldsymbol{\omega}'_2 \mathbf{v}_{igmt} + \varepsilon^L_{igmt} \quad (2)$$

where C^*_{igmt} and L^*_{igmt} describe the desirability of childbearing and labor force participation (LFP) for woman *i* who is a member of age group *g* and living in metropolitan area *m* in year *t*. As will be discussed in more detail below, we divide the sample into groups predominantly because correlations can only be calculated at some level of aggregation. Although we use age to define groups in our analysis, this model could be applied to groups based on other exogenous determinants of childbearing and labor force participation. The associated binary outcomes are C_{igmt} and L_{igmt}, where $C_{igmt} = 1$ is observed if $C^*_{igmt} > 0$ and likewise for labor force participation. The presence of a child at or below the age of 1 in the household is used to measure childbearing.[2] There is no generally applicable exclusion restriction to identify the effect of childbearing on employment or vice-versa. Consequently, both equations have the same right-hand-side variables and yield estimates of the net effects of these variables on work and fertility outcomes. The vector of controls, $\mathbf{v}_{igmt}$, includes city, region-year, and age group fixed effects as well as other demographic variables. The two error terms, ε^C_{igmt} and ε^L_{igmt} are distributed according to a joint normal distribution with a mean of zero.

[2] Using older children to measure childbearing would be problematic because parameters of interest will be identified off of within-MSA differences in the size of the immigrant population across decades. Decisions to have older children are not likely to be affected by relatively recent changes in the immigrant population.

Even under the assumption that low-skilled immigrant inflows only affect childbearing and labor force participation decisions of high skilled women by decreasing childcare costs, theory does not provide clear predictions for β_1 and β_2. As described in detail in Blau and Robins (1989), women may respond to lower childcare costs by having an additional child given the reduction in the price of childrearing. They may also work more hours given that their net take-home pay increases when childcare costs decrease. However, because of the highly time-intensive nature of caring for young children, labor force participation may, at least initially, decrease if women respond to lower childcare costs by having an additional child. Similarly, if upon entering the labor force (as a result of decreased childcare costs) women start valuing their roles as breadwinners, they may choose not to have a second or third child. In the end, the net impact of childcare costs, and hence low-skilled immigration, on fertility and labor supply decisions is an empirical question.

In addition to examining the propensities to work and to bear children, the standard bivariate probit model enables an analysis of the correlation between the two. If the error terms in equations (1) and (2) follow a bivariate normal distribution, $\rho_{gmt} = corr(\varepsilon^{C}_{igmt}, \varepsilon^{L}_{igmt})$ is, by definition, the *tetrachoric correlation*. The tetrachoric correlation can be understood as the degree to which changes in childbearing--that are not a result of immigration and the other variables in the model--translate into changes in labor force participation (alternatively, the degree to which exogenous changes in labor force participation translate into changes in childbearing). Hence the tetrachoric correlation would, for example, determine the effect of an unintended pregnancy on desired labor supply or the effect of an increase in the local demand for high-skilled labor on the desirability of childbearing. Thus, ρ should be negative as is almost universally the case in the sample. Although the use of bivariate models to study behaviors is certainly not new, an innovation of this paper is to explore how the correlation is affected by low-skilled immigration. I use the parameterization

$$\rho_{gmt} = \beta_3 LSI_{mt} + \boldsymbol{\omega}'_3 \mathbf{v}_{gmt} + e_{gmt}, \tag{3}$$

where $\mathbf{v}_{gmt}$ is a vector of characteristics of women in group *g* in metro *m* in year *t*, and e_{gmt} represents the un-modeled determinants of ρ. If an increase in low-skilled immigration results in less expensive, more convenient, or better quality market-based childcare services, β_3 should be positive. That is, low-skilled immigration should dampen the negative latent correlation between childbearing and labor supply.

2.2 Grouped Estimation with Instrumental Variables

Although I start with the individual-level data provided by IPUMS, the analyses are all conducted with metropolitan area-year-age group cells as the unit of observation. The reasons for creating these groups are twofold. First, the group-level model, which is based on a slight generalization of Amemiya's (1974) grouped bivariate probit specification, allows for a straightforward application of instrumental variables. More importantly, grouping is necessary in order to calculate the tetrachoric correlation, which, like any other correlation, is not defined at the individual level. Linking the effect of immigration on the correlation to its effect on the joint probability of work and childbearing, an innovation of our work, would not be possible without the grouped model.

The coefficients in the individual-level model presented in the previous section can be estimated by analyzing sample proportions and using group-level explanatory variables ($\mathbf{v}_{gmt}$). Given the bivariate normal distribution of the error terms, the expected rates of childbearing and LFP follow univariate normal distributions:

$$\pi^C_{gmt} = \Phi\left(\beta_1 LSI_{mt} + \boldsymbol{\omega}_1' \mathbf{v}_{gmt}\right) \text{ and } \pi^L_{gmt} = \Phi\left(\beta_2 LSI_{mt} + \boldsymbol{\omega}_2' \mathbf{v}_{gmt}\right) \tag{4}$$

Let p^C_{gmt}, p^L_{gmt}, and p^{CL}_{gmt} denote the observed proportions of the women in group *g* in metropolitan area *m* in year *t* that bear children, participate in the labor force, and do both, respectively. A first-order Taylor expansion around the expected values of the sample proportions results in the linear equations:

$$c_{gmt} = \beta_1 LSI_{mt} + \boldsymbol{\omega}_1' \mathbf{v}_{gmt} + u^1_{gmt}, \tag{5}$$

$$\ell_{gmt} = \beta_2 LSI_{mt} + \boldsymbol{\omega}_1' \mathbf{v}_{gmt} + u^2_{gmt}, \tag{6}$$

where $c_{gmt} = \Phi^{-1}(p^{C}_{gmt})$ and $l_{gmt} = \Phi^{-1}(p^{L}_{gmt})$ denote the inverse standard normal cumulative distribution (normit) function applied to the observed rates of childbearing and LFP. Thus, we can estimate these equations with linear univariate models despite the bivariate structure of the overall model. Moreover, based on equation (3), the expression for the empirical analogue of the population tetrachoric correlation obtained from the data can be expressed as

$$r_{gmt} = \beta_3 LSI_{mt} + \boldsymbol{\omega}_3' \mathbf{v}_{gmt} + u^{3}_{gmt}. \qquad (7)$$

Equations (5)-(7) correspond to Amemiya's (1974) equations (4.11)-(4.13), with the expression in (7) additionally relying on the parameterization of the tetrachoric correlation described above.

The empirical analog of the tetrachoric correlation is significantly less straightforward to compute than the analogues of the expected rates of childbearing and labor force participation. It is calculated based on the population relationship:

$$\pi^{CL}_{gmt} = F(\Phi^{-1}(\pi^{C}_{gmt}), \Phi^{-1}(\pi^{L}_{gmt}), \rho_{gmt}) \equiv G(\pi^{C}_{gmt}, \pi^{L}_{gmt}, \rho_{gmt}) \qquad (8)$$

where π^{CL}_{gmt} represents the expected share of women who simultaneously bear children and participate in the labor force and $F(\cdot)$ denotes the standard bivariate normal distribution function. Using the observed proportions (p^{C}, p^{L}, and p^{CL}) as analogues of the expected values in equation (8) allows me to calculate the empirical tetrachoric correlation, r_{gmt} based on the sample of outcomes. Although there is no closed-form solution for r_{gmt}, since $F(\cdot)$ is monotonic in the third argument (Tihansky 1972), we can apply a recursive binary chop algorithm to search for the value of r_{gmt} that solves

$$\left| p^{CL}_{gmt} - G(p^{C}_{gmt}, p^{L}_{gmt}, r_{gmt}) \right| < \xi ,$$

where ξ represents a pre-defined level of precision, which we set to 2^{-50}.[3] Note that monotonicity of $F(\cdot)$ also implies that a higher value of ρ will, *ceteris paribus*, translate into a higher joint likelihood of childbearing and labor force participation.

If there are any groups in which any of the dependent variables take on the same value across its members, the data become uninformative and it is not possible to estimate the empirical tetrachoric correlation. Consequently, I divide the sample of college-graduate women into only two broad age groups, women ages 22-32 and women ages 33-43, and include measures of the average characteristics of the group as explanatory variables. Even with these broadly specified categories, there were a few MSA-year-age group cells with all ones or all zeros. In these cases, I replace zeros with 0.001 and ones with 0.999 when calculating the normits and estimating ρ.

The model described by equations (5) through (7) is estimated using three separate linear equations of the form

$$y_{gmt} = \beta LSI_{mt} + \boldsymbol{\mu}_m + \boldsymbol{\mu}_{kt} + \boldsymbol{\mu}_g + \lambda_{mt} IncControl_{mt} + \boldsymbol{\theta}'\mathbf{x}_{gmt} + u_{gmt}, \qquad (9)$$

where y is one of the three dependent variables (c, ℓ, r). The low-skilled immigrant share of the overall working-age population immigration is denoted LSI. Metropolitan area-specific intercepts are indicated by $\boldsymbol{\mu}_m$, while $\boldsymbol{\mu}_{kt}$ represents time fixed effects specific to the kth Census region. The variable *IncControl* denotes the log of income per capita among working-age male college graduates. Added to these variables are age-group fixed effects ($\boldsymbol{\mu}_g$) and a vector of demographic controls ($\mathbf{x}_{gmt}$). The vector $\mathbf{x}_{gmt}$ includes the share of women of women in each MSA-year-age group cell who are married and the corresponding proportions of the group that self-identify as being black and that self-identify as being a member of another non-white, non-Hispanic race. Each group-MSA-year cell is weighted by the estimated population of women represented by the cell. Robust standard errors are clustered by MSA.

In the following sections, I describe the data as well as the instrumental variables strategy and some baseline results relating immigrant inflows to the likelihood that women have both recently given birth

[3] Stata code for this calculation will be available in an online appendix to this paper.

and are participating in the labor market.[4] I then come back to the model just presented to gain insight into a variety of the mechanisms whereby low-skilled immigration may exert its effects. Specifically, the model will allow me to separately estimate the impact of low-skilled immigration on: (i) the likelihood of bearing children, (ii) the likelihood of participating in the labor market, and (iii) the correlation between the two. Most importantly, however, the model will be used to indirectly calculate the effect of low-skilled immigration on the joint likelihood of working and giving birth, decomposing the total effect into changes in the marginal likelihoods and changes in the correlation between fertility and work.

3 Data

In order to provide readers with a basic sense of what has been happening to women's work and fertility decisions over time, I start by plotting the proportion of women of child-bearing age that both work and have given birth in the previous year (a 3-year moving average is plotted to smooth out year-to-year fluctuations).[5] As shown in Figure 1, the joint rate of childbearing and labor force participation is small in absolute terms, which reflects the relative infrequency of childbirth. However, the joint rate among women ages 18-39 almost doubled between 1970 and 2000. Among college graduates it more than doubled, increasing from approximately 2.5% to 4.9%.

My main analysis relies on data from the 1980-2000 Integrated Public Use Microdata Series (IPUMS) (Ruggles 2010) of the US Census.[6] Additionally, the 1970 Census is used to construct the instrument for immigrant share. The sample consists of US-born non-Hispanic college-educated women

[4] My measure of having given birth in the previous year is the presence of a child under the age of one in the household. Some of these children may have been adopted. Moreover, women who have given birth but are not residing with the child will not be counted in my measure of childbearing.

[5] The figure draws on yearly data from the March Current Population Surveys (CPS), 1969-2001 (King et al. 2010). These data cannot be used for the main analysis because sample sizes are smaller and questions about immigration were only asked starting in 1994.

[6] I do not use the more recent American Community Survey (ACS) data for several reasons. First, the ACS is conducted every year, but it is only a one percent sample of the population. Many researchers merge several years of ACS data to increase sample size, but this is problematic for the purposes of this paper because of my measures of fertility and foreign-born share. For example, if I were to use the 2007-2011 five year sample, I would at least partially be using the foreign born population in 2011 to predict pregnancy decisions in 2006. Second, the Great Recession may have induced a fair amount of noise into women's fertility and labor force participation decisions making it difficult to estimate parameters precisely despite the increase in sample size. Finally, on a more practical note, first stage estimates in the IV analysis were weak when using the more recent data.

between the ages of 22 and 42, inclusively. I focus on college-educated women both because they are more likely to use market-provided child care and because their labor market opportunities are less likely to be directly affected by low-skilled immigrant inflows. I do not include Hispanic women in the sample given their greater likelihood of being impacted by the mostly Hispanic low-skilled immigrant inflows for reasons unrelated to childcare markets.

A potential difficulty in exploiting within Metropolitan Statistical Area (MSA), cross-decade variation in immigrant share is that MSA boundaries change over time. Some of these changes reflect seemingly arbitrary decisions of the U.S. Office of Management and Budget (OMB), but others are due to natural expansions or contractions of economic activity in the outskirts of an MSA. If immigrants are more likely to settle in expanding cities and women living in outer suburbs make systematically different fertility and labor supply decisions, the changing MSA boundaries may result in biased estimated immigrant share coefficients. To at least partially address this issue, I follow Cortes and Tessada (2011) and use only MSAs with the same codes in the IPUMS between 1970 and 2000. This does not perfectly resolve the problem in that counties represented in MSAs listed with the same code may still have changed over the years. However, because these counties tend to have small populations, they are unlikely to severely bias results.

Table 1 presents descriptive statistics of the main variables used in the study. Given the nature of the analysis, all descriptive statistics in the table are constructed from the already grouped variables. The number of high-skilled women in the population represented by each of the MSA-year-age group cells are used as weights when calculating the statistics in the table. In constructing the immigrant share variable, I divide the number immigrants aged 18-64 with no more than a high school degree by the total population of the MSA in the same age range in the same year. The statistics in the low immigrant share column were constructed using cells in MSAs where the share immigrant was at or below the mean in the overall sample while statistics in the high immigrant share columns were constructed using the remainder of the overall sample.

Interestingly, the share of women in MSA-year-age group cells who have recently given birth and participate in the labor market is higher in cities with smaller immigrant populations. The correlation between fertility and labor force participation is also less negative in these cities. The table suggests that these relationships are driven by differences in the likelihood of childbearing since labor force participation rates are nearly identical in low and high immigration cities. The negative association between share foreign born and work-life balance cannot be taken too seriously given that the cities that attract more immigrants differ in ways that might affect women's fertility and labor supply decisions. For example, college-educated men in low immigration cities tend to have smaller yearly incomes. If high-skilled women in these cities also have lower wages, the opportunity cost of childbearing and specifically, leaving the labor force after giving birth, will be lower.

Turning now to the characteristics of the college-educated women in the sample, Table 1 shows that marriage rates are higher among women in cities with fewer immigrants suggesting that family norms, as opposed to immigrant shares, may explain the higher fertility rates in low-immigrant cities. Also of note is that while the proportion of college-educated women in the sample that is black is practically the same across city-types, women in the sample are significantly more likely to be non-white, non-Hispanic, and non-black in high immigration cities.

4 Results

4.1 Direct Estimation

While the structure of the model will eventually allow me to decompose immigration's impact into several different components, I start by directly estimating the effect of immigrant inflows on the joint likelihood that high skilled women have recently born children and participate in the labor force. This portion of the analysis could have been conducted using individual-level data, but because I want to later compare the direct and indirect methods of estimating immigration impacts, I estimate the following equation using the grouped data probit model:

$$d_{gmt} = \beta LSI_{mt} + \boldsymbol{\mu}_m + \boldsymbol{\mu}_{kt} + \boldsymbol{\mu}_g + \lambda_{mt} IncControl_{mt} + \boldsymbol{\theta}'\mathbf{x}_{gmt} + e_{gmt},$$

The right hand side variables are defined as before and d_{gmt} is equal to $\Phi^{-1}(p_{gmt}^{CL})$. Column 1 of Table 2 displays results from a simple univariate probit model with grouped data which includes only age group fixed effects, region-year fixed effects and controls for share married, share black, and share non-Hispanic, non-white, and non-black of the high-skilled women in the sample. Note that because there are no Hispanic women in the sample, the proportion white is the omitted category. In contrast to what theory would predict but consistent with the discussion of the descriptive statistics, high skilled women living in cites with larger low-skilled immigrant shares are *less* likely to have given birth in the previous year and at the same time participate in the labor market even when controlling for year-region fixed effects and age group fixed effects.

A natural concern with this model is that cities with large immigrant populations may be systematically different from cities with small immigrant populations in ways that affect women's family and work decisions for reasons unrelated to childcare markets. To partially address this issue, MSA fixed effects are added in Column 2. When identification comes from within MSA variation in the foreign born share over time, the sign of the estimate of the share immigrant coefficient reverses suggesting that the largest increases in the proportion of high skilled women who have had recently given birth and work occur in cities with the largest increases in low-skilled immigrant populations.

A problem with this specification arises if unobserved factors cause changes within a city in both the size of the immigrant population and women's work-fertility decisions. For example, changing gender norms may lead to more working mothers who demand childcare services and the resulting increase in childcare wages attracts immigrants to a city. To address these types of issues, I take a commonly used instrumental variables approach which exploits the tendency of immigrants to locate in established communities of coethnics (Card 2001). Using 1970 data, I start by constructing for each MSA the proportion of immigrants from each origin country that reside in the MSA. I then multiply this value by the overall flow of immigrants from that country to the entire US within the previous decade. Finally, I sum this figure over all origin countries in the MSA. Mathematically,

$$Inst_{mt} = \sum_b \frac{N^b_{m,t_0}}{N^b_{t_0}} (N^b_t - N^b_{t-10})$$

where *b* refers to country of origin and N refers to number of immigrants. Thus, N^b_{m,t_0} is the number of immigrants from country *b* living in MSA *m* in the base year t_0 , which I take to be 1970, and $N^b_{t_0}$ is the total number of immigrants from country *b* living in the US in the base year. The inflow of immigrants from country *b* to the US within the previous decade is $N^b_t - N^b_{t-10}$. The resulting variable is correlated with the size of the foreign born population in an MSA but is not likely to be directly related to differences across MSAs, within the same year, in labor market opportunities or fertility preferences of high skilled females.

The first stage estimate of the effect of the instrument on the share low skilled immigrant is positive, as expected, and statistically significant ($p<.001$). The associated F statistic is 33.75 pointing to the instrument's strong predictive power. Second stage results are shown in the third column of Table 2. Notice that the IV estimate of the effect of immigrant share is positive and larger in magnitude than the fixed effects estimate shown in the previous column. This suggests that immigrants tend to be drawn to cities with either lower fertility rates, lower labor force participation, or both.

A potential problem with even the IV estimate arises if the distribution of different ethnic groups across cities in 1970 has a direct impact on women's work and fertility decisions in different cities ten to 30 years later. This may occur, for example, if the composition of a city's foreign born population in 1970 leads to a specific industry mix which might make work more or less attractive. To examine whether this is likely to be problematic, in column 4 I add to the model a measure of women's potential wage income:

the log of average yearly income for college-educated males in the MSA and year.[7] Estimates barely change suggesting that this is unlikely to be a problem.

The estimates of the coefficients tell us the impact of the variables on the normit of the joint likelihood of childbearing and labor force participation. To interpret these coefficients in a more meaningful way, we compute marginal effects in the standard way except that when calculating the average, each observation is weighted by the proportion of high-skilled women represented by the MSA-year-age group cell. The low-skilled immigrant share of the labor force in the average high-skilled woman's MSA rose from about 6.6% in 1980 to 9.8% in 2000. As can be seen in Table 2, our final and preferred specification implies that such a change in the share of low-skilled immigrants in a woman's MSA leads to a .58 percentage point increase in the probability that she has given birth in the previous year and participates in the labor market.

I conclude therefore that immigrant inflows do increase the proportion of childbearing-age women that have recently given birth and participate in the labor market. This is certainly consistent with the notion that immigrant-induced better childcare options make combining work and family less difficult. In the next section, I examine whether the immigration effects operate mostly through direct impacts on fertility and labor participation decisions or more indirectly by changing the tradeoffs women must make when deciding between the two.

4.2 Indirect Estimation

This part of the analysis, I consider the impact of immigration on fertility and labor force participation rates in addition to the correlation between the two. Coefficients in models depicted by equation 9 are estimated using the IV presented in the previous section. As can be seen in Table 3, the IV coefficients indicate that low-skilled immigration leads to significantly higher fertility rates and lower labor force participation rates. Using the IV point estimates to compute the average partial effects (APEs), a 3.2 percentage point increase in LSI, the change between 1980 and 2000, implies a likelihood of childbearing

[7] I use male income instead of female income because it is less likely to be affected by gender norms and childcare costs.

that is 0.97 percentage points higher.[8] This corresponds to about 14 percent of the observed fertility rate in the 2000 sample. The estimated effect of the average increase in low-skilled immigration on the likelihood of labor force participation is –1.32 percentage points.

Taken together, the changes in the marginal likelihoods results suggest that high-skilled women in our sample of MSAs respond to immigrant-induced reductions in childrearing costs by exiting from the labor force to bear children. This pattern of behavior, along with a generally negative tetrachoric correlation, indicates that high-skilled women face tradeoffs between work and fertility. As seen in the third column of Table 5, low-skilled immigration also attenuates the negative correlation between childbearing and labor force participation.

I return now to the bivariate probit structure of the model presented in Section 2 in order to indirectly calculate the marginal effect of immigration using estimated βs from the three univariate models. Specifically, equations (3) and (4) imply that the expected joint likelihood can be written as

$$\pi_{gmt}^{CL} = F\left(\beta_1 LSI_{mt} + \boldsymbol{\omega}_1' \mathbf{v}_{gmt}, \beta_2 LSI_{mt} + \boldsymbol{\omega}_2' \mathbf{v}_{gmt}, \beta_3 LSI_{mt} + \boldsymbol{\omega}_3' \mathbf{v}_{gmt}\right) \tag{10}$$

where again F denotes the standard bivariate normal distribution function. My interest is not in predicting the joint likelihoods directly. Instead I use this equation to calculate the average partial effect of immigration in a way that allows me to decompose the total impact of immigration on the joint likelihood into the three different components. Taking the derivative of π_{gmt}^{CL} with respect to LSI, we can compute the average partial effect of low-skilled immigration as $\hat{A}_{LSI}^{CL} = \sum_{gmt} h_{gmt}(d\hat{F}_{gmt} / dLSI)$, with h_{gmt} denoting the proportion of high-skilled skilled women represented by each age-MSA-year cell and $\hat{F}_{gmt}$ denoting the standard bivariate normal distribution function evaluated using the estimated coefficients

[8] Just as in the previous section, I compute the APE for the marginal likelihoods of childbearing and labor force participation as the weighted average of the partial derivatives of (4) across the sample, with the share of high-skilled women represented by each age-MSA-year cell used as weights. Scaling the APE by the change in LSI experienced by the representative woman in our sample between 1980 and 2000 yields the reported effects.

obtained from the univariate models in place of the true parameters.[9] Thus, the average partial effect (APE) can be writeen as

$$\hat{A}_{LSI}^{CL} = \hat{D}_1\hat{\beta}_1 + \hat{D}_2\hat{\beta}_2 + \hat{D}_3\hat{\beta}_3 \tag{11}$$

where $\hat{D}_j = \sum_{gmt} h_{gmt}\hat{F}_{gmt}^{j}$ and $\hat{F}^{j}$ denotes the *j*th partial derivative of $\hat{F}$. The first two terms in equation (11) represent the average change in the joint likelihood arising from the differential impacts of low-skilled immigration on the propensity to bear children and the propensity to work, respectively. The third term denotes the change in the joint likelihood attributable to changes in the tetrachoric correlation induced by low-skilled immigration. This can be interpreted as the effect of immigration on the joint likelihood arising from a weakened link between childbearing and labor force participation.

Again, to translate the APE (and its components) into more meaningful terms, I scale by the percentage point increase in low-skilled immigration experienced by the representative high-skilled woman between 1980 and 2000. Based on the point estimates from Table 5, the 3.2 percentage point increase in LSI would result in a 0.67 percentage point increase in the likelihood of bearing children while remaining in the labor force. Note that this calculation of the impact of immigration on the joint likelihood is itself indirect. It is obtained not by directly regressing the joint likelihood on share foreign born, but instead by putting together estimates of the low-skilled immigration coefficients in the fertility, labor force participation, and correlation models using the structure implied by the bivariate probit model. Thus, its validity rests on the statistical structure of the bivariate probit model as well as the implicit assumption that the women who change their fertility in response to immigration are the same women who change their labor supply.

In order to verify that the indirect estimate is not being affected by these assumptions, I compare the indirect estimate of the impact of immigration on the joint likelihood, shown in the last column of Table 3 to the direct estimate, first shown in Table 2 but reproduced in Table 3 for convenience. As can be seen in the Table 3, the marginal impact of immigration implied by the coefficient is similar to what is

[9] For convenience, I have included this derivative in Appendix 1.

obtained using the indirect method. This should assuage most concerns regarding the indirect methodology.

The indirect approach enables a decomposition of the total effect of immigration into a portion attributable to increasing fertility, a portion attributable to changing labor force participation, and a portion to decreasing the tradeoffs must make. As seen in Table 3, about 24% of the total effect of low skilled immigration on the joint likelihood is attributable to the weakened latent correlation between fertility and work, with the remainder arising from differential changes in childbearing and labor force participation rates. This implies that low-skilled immigrants not only directly affect women's fertility and labor supply decisions, but a substantial portion of immigrants' influence operates by making it less important for women to decrease labor force participation in response to additional births which are unrelated immigration.

Between 1980 and 2000, the joint likelihood in our sample of urban, non-Hispanic college graduate women rose from 2.8% to 4.5%. Depending on whether we use the direct or indirect methodologies, the total effect of low-skilled immigration on the joint likelihood of fertility and work represents 34 to 39 percent of the observed increase in the sample between 1980 and 2000. Of course, there are other margins along which household services markets and female decision making would have adjusted if no immigration had actually taken place after 1980. Nonetheless, the estimates in this paper indicate that inflows of low-skilled immigrants to a metropolitan area during the sample period led to significant and substantial short-run increases in the joint likelihood of childbearing and labor force participation.

Next, I present a series of specification and robustness checks to address some of the concerns readers may have regarding the baseline results. I start by considering the influence of outliers. Most of the MSA-years with the largest immigrant concentrations are in the state of California and so I re-estimated the model excluding California. As seen in Panel A of Table 4, the estimated effect of immigration on the joint rate of fertility and labor force participation rises when removing this high immigration state. The basic story of increasing fertility rates, decreasing labor force participation, and attenuation of the negative correlation remains the same.

I also consider the robustness of results to a slightly different way of grouping the data. As discussed above, grouped data is necessary to conduct the analysis of correlations and it allows for a straightforward application of instrumental variables techniques within a bivariate probit model. The construction of groups using two age categories, however, was rather arbitrary. To check for the sensitivity of results to this decision, I increased the number of age categories to three. As can be seen in Panel B of Table 4, this yields similar estimates suggesting that the structured model is not skewing the estimates of the effects of low-skilled immigration.

Perhaps of more concern than outliers or the grouped structure of the data is whether the instrument is correlated with the error terms in the estimating equations. More specifically, the validity of the IV results rests on the assumption that the distribution of immigrants across metropolitan areas in 1970 was not affected by or a factor determining other city-level characteristics that might, in the absence of subsequent immigration, affect later changes in the fertility and labor force participation outcomes of high-skilled U.S.-born women. A potential violation of this assumption is the following: Immigrants might have been historically more highly represented in metropolitan areas with more persistent traditional family values. This seems unlikely given that high-skilled women have higher fertility rates in cities with fewer immigrants (Table 1), but if it were true, my estimation strategy would falsely attribute changes in fertility and work outcomes to the rising share of low-skilled immigrants in the labor market.

To explore whether this is likely to be problematic, I examine the relationship between marriage patterns and low-skilled immigration. If there happens to be stronger family norms in cities receiving more immigrants as a result of historical ethnic enclaves, then my analysis should generate a positive relationship between immigration and marriage rates along with an increase in the joint probability of marriage and labor force participation. This cannot be considered a pure placebo test given the possibility that committed couples respond to immigrant-induced decreases in childcare costs by having a baby but get married before childbirth. It turns out, however, that my estimation strategy does not yield any statistically or economically significant relationships between immigrant inflows and marriage. As can be seen in Panel C of Table 4, the IV estimate on marriage rates is statistically insignificant ($p > 0.83$) and trivial. The estimated effect on the joint likelihood of marriage and labor force participation is not only statistically insignificant ($p > 0.27$) but has a negative sign.

Finally, I examine whether high-skilled native-born women move to cities with large foreign-born populations, perhaps to purchase less expensive household services. I test for this type of selection by estimating the following equation:

$$\ln N_{gmt} = \gamma LSI_{mt} + \boldsymbol{\mu}_m + \boldsymbol{\mu}_{kt} + \boldsymbol{\mu}_g + \lambda IncControl_{mt} + \chi \ln T_{gmt} + \varepsilon_{gmt}, \tag{12}$$

where N_{gmt} denotes the number of non-Hispanic female college graduates in age group g in metro area m in year t. All of the right hand variables have been previously defined, with the exception of T_{gmt}, which denotes the total number of non-Hispanic native females in the age-MSA-year cell. Equation (12) essentially tests for growth in the population of high-skilled females, relative to the overall population of same-age females. Although these results are not shown in tables, instrumental variables regression analysis indicates an IV estimate of γ of -2.36 with a corresponding p value of 0.26. This suggests that composition bias due to selective migration is not likely to be a problem for this analysis.

5 Conclusion

The results in this paper suggest that immigration to the US between 1980 and 2000 increased the fertility of U.S.-born college graduates of childbearing age. This rise in childbearing was accompanied by an increase in exits from the labor force. Most interestingly, low-skilled immigration resulted in a weakening of the negative correlation between fertility and work and a sizeable increase in the joint likelihood of childbearing and labor force participation. The structure of the model allows me to determine that about a quarter of the increase in the joint likelihood can be attributed to a weakening of the negative correlation between childbearing and labor force participation. Taken together, these findings indicate that low-skilled immigration substantially reduces the work-fertility tradeoff faced by educated urban American women.

Kremer (2006) argues that migration of foreign domestic household workers may raise wages of low-skilled native workers relative to high skilled workers and can create large fiscal gains for receiving countries. His model implies that if high-skill women respond to immigrant-induced better childcare options by entering the labor force or working more hours, the ratio of skilled to unskilled workers decreases thereby driving down skilled wages relative to unskilled wages. Given that home production is not taxed and high skill women are often married to high wage men, these women will typically pay a high marginal tax rate when entering the labor market thus generating large fiscal gains. Several analyses have concluded that high skill women do tend to work longer hours in response to immigrant inflows (Cortes and Tessada 2011, Barone and Mocetti 2011, Farre et al. 2011, Cortes and Pan 2013, Furtado 2014). This paper contributes to this literature by additionally showing that some women respond to better childcare options with increases in fertility and decreases in labor force participation. While the (most likely temporary) decreases in labor force participation imply smaller decreases in inequality and smaller fiscal gains than what is implied by Kremer's model, the increases in fertility might help sustain pension programs especially in countries with below-replacement fertility rates. Regardless of whether women choose to respond to immigrant inflows by increasing fertility or labor supply, this paper's finding that tradeoffs decrease in response to immigrant inflows implies welfare gains for the women making work-life decisions.

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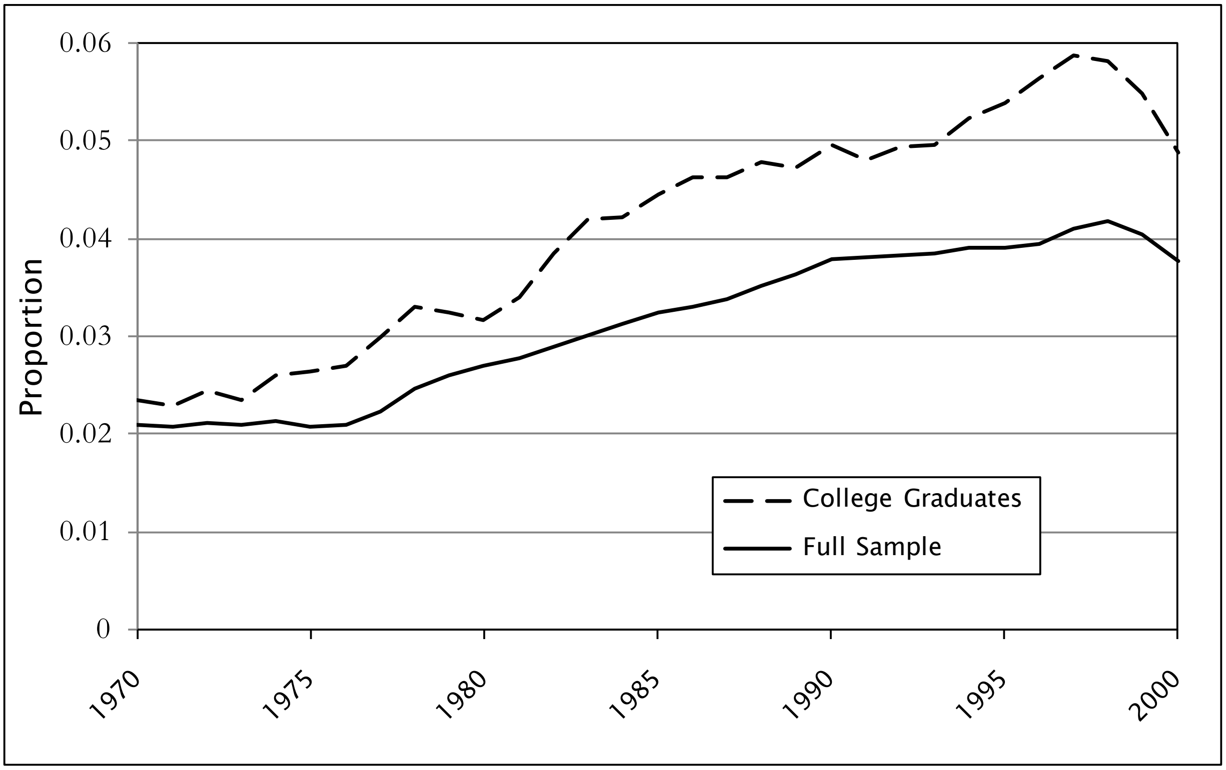
Figure 1. Proportion of Women Bearing Children and Participating in the Labor Force



Notes: The figure draws on data from the March Current Population Surveys (CPS), 1969-2001 (King et al. 2010). The sample is comprised of women ages 18-39. We define "childbirth" and "recent motherhood" based on the presence of an own-child less than or equal to one year old in the household. Each series of data has been plotted after applying a 3-year moving average to smooth out year-to-year fluctuations.

Table 1. Descriptive Statistics by Size of the Immigrant Population

	Total		Low Percent Immigrant		High Percent Immigrant	
	Mean	SD	Mean	SD	Mean	SD
Share Recently Given Birth and Participating in Labor Market	0.038	0.015	0.041	0.017	0.035	0.012
Fertility Rate	0.064	0.021	0.068	0.023	0.059	0.018
Labor Force Participation Rate	0.832	0.051	0.83	0.051	0.833	0.05
Tetrachoric Correlation: Fertility and Labor Force Participation	-0.386	0.14	-0.379	0.146	-0.393	0.134
Share Working Age Low-Skilled Immigrant	0.084	0.076	0.028	0.014	0.14	0.071
Log Mean Income of Males with College	10.774	0.478	10.678	0.464	10.872	0.474
Proportion Married	0.579	0.134	0.616	0.111	0.542	0.146
Proportion Black	0.095	0.064	0.094	0.069	0.096	0.058
Proportion Other Race	0.021	0.048	0.008	0.009	0.034	0.066
Number of Observations	708		478		230	

Notes: The share recently given birth and participating in the labor market variable is the proportion of women in the cell that have given birth within the previous year and participate in the labor market. The fertility rate is the share of women in the cell who have given birth within the previous year. The tetrachoric correlation is a measure of association between the two binary variables, having recently given and birth and participating in the labor market. See the text for further details. Working age refers to ages 18-64 and low-skilled is defined as having at most a high school degree. Given that the sample contains no Hispanics and whites are the omitted category, the share other race variable refers to the share that is non-Hispanic, non-black, and non-white. This group consists mostly of Asians and American Indians. The share working age immigrant and the log of the mean income of male college graduates variables are calculated within MSA-year cells. All other variables are averages within MSA-year-age group cells. The two age groups in the sample are 22-31 and 32-42. There are 708 MSA-year-age group cells in the sample (118 MSAs, 3 years, and 2 age groups). When computing the statistics in this table, cells are weighted by the share of the population represented by the cell. Of these, 478 are in MSA-years with a low-skilled immigrant share at or below the mean in the sample. The remaining 230 are in MSA-years with immigrant shares above the mean.

Table 2. Effects of Low-Skilled Immigration on the Joint Likelihood Using Grouped Probit Model

DEPENDENT VARIABLE: NORMIT OF JOINT LIKELIHOOD OF BIRTH AND LABOR FORCE PARTICIPATION	1	2	IV 3	IV 4
Share Working Age Low-Skilled Immigrant	-0.00458	0.703***	2.145***	2.274***
	(0.0848)	(0.225)	(0.715)	(0.830)
Proportion Married	1.040***	1.723***	1.719***	1.747***
	(0.0802)	(0.188)	(0.172)	(0.182)
Proportion Black	0.211*	0.908*	0.555	0.502
	(0.114)	(0.507)	(0.457)	(0.475)
Proportion Other Race	-0.0112	-4.450***	-4.776***	-4.625***
	(0.0726)	(0.810)	(0.873)	(0.858)
Log Mean Income of Males with College				-0.196
				(0.170)
Age Group Fixed Effects	Yes	Yes	Yes	Yes
Region-Year Fixed Effects	Yes	Yes	Yes	Yes
MSA Fixed Effects	No	Yes	Yes	Yes
Effect of Average Change in LSI, 1980-2000	-0.000011	0.0018	0.00545	0.00578
First Stage F (excluded instrument)			33.75	71.48
Number of Observations	708	708	708	708

Notes: The data consist of MSA-year-age group means of the dependent and explanatory variables for non-Hispanic U.S.-born college graduate women not enrolled in school. All equations include the explanatory variables described in the notes to Table 1. Each of the MSA-year-age group cells is weighted by the population of women represented by the cell, and the robust standard errors in parentheses are clustered by MSA. The marginal effects are the change in the underlying dependent variable that would be causes by the change in the share working age low-skilled immigrant experienced by the average member of the sample. *p < .10; **p < .05; ***p < .01

Table 3. IV Estimates of the Effects of Low-Skilled Immigration on Fertility and Work Outcomes

	Normit, Birth Rate	Normit, LFP Rate	Tetrachoric, Birth and LFP	Normit, Joint Likelihood of Birth and LFP	
				Direct	Indirect
	1	2	3	4	5
Share Working Age Low-Skilled Immigrant	2.483	-1.674	1.033	2.274	
(LSI)	(0.836)	(0.607)	(0.363)	(0.830)	
Mean of Underlying Dependent Variable, 2000	0.07011	0.83421	-0.3294	0.04515	
Effect of Average Change in LSI, 1980-2000	0.00971	-0.01317	0.03294	0.00578	0.00667
Proportion Explained by Weakened Correlation					0.237
Number of Observations	708	708	708	708	

Notes: Column 4 reproduces results shown in column 4 of Table 2 for convenience. The indirect measure of the effect of low-skilled immigration, shown in column 5, is obtained by combining the estimates shown in columns 1-3 in the way described in the text. All models include time-varying region fixed effects, MSA fixed effects, age-group fixed effects, the log of income per male college graduate, and the following group-level characteristics: the proportion black, the proportion who are of another non-white race, and the proportion married. Each of the observation-cells is weighted by the population of women represented by the cell, and the robust standard errors in parentheses are clustered by MSA. Reported effects are the change in the underlying dependent variable that would be caused by the change in the share working age low-skilled immigrant experienced by the average member of the sample. *p < .10; **p < .05; ***p < .01

Table 3. Robustness Checks of the IV Estimates of the Effects of Low-Skilled Immigration on Fertility and Work Outcomes

PANEL A: DROP CALIFORNIA	Normit, Birth Rate	Normit, LFP Rate	Tetrachoric, Birth and LFP	Normit, Joint Likelihood of Birth and LFP	
				Direct	Indirect
	1	2	3	4	5
Share Working Age Low-Skilled Immigrant	3.438**	-3.269***	1.767**	3.601**	
(LSI)	(1.613)	(1.099)	(0.683)	(1.493)	
Mean of Underlying Dependent Variable, 2000	0.0712	0.834	-0.331	0.0458	0.0458
Effect of Average Change in LSI, 1980-2000	0.0116	-0.0220	0.0482	0.00796	0.00775
Proportion Explained by Weakened Correlation					0.3
Number of Observations	648	648	648	648	
PANEL B: THREE AGE GROUPS	Normit, Birth Rate	Normit, LFP Rate	Tetrachoric, Birth and LFP	Normit, Joint Likelihood of Birth and LFP	
				Direct	Indirect
	6	7	8	9	10
Share Working Age Low-Skilled Immigrant	2.278**	-0.912*	0.930*	1.989**	
(LSI)	(0.902)	(0.470)	(0.516)	(0.953)	
Mean of Underlying Dependent Variable, 2000	0.0701	0.834	-0.312	0.0451	0.0451
Effect of Average Change in LSI, 1980-2000				0.00494	0.00640
Proportion Explained by Weakened Correlation					0.213
Number of Observations	1062	1062	1062	1062	
PANEL C: MARRIAGE AND LFP	Normit, Marriage Rate	Normit, LFP Rate	Tetrachoric, Marriage and LFP	Normit, Joint Likelihood of Marriage and LFP	
				Direct	Indirect
	11	12	13	14	15
Share Working Age Low-Skilled Immigrant	0.109	-1.756***	0.349	-0.732	
(LSI)	(0.536)	(0.578)	(0.282)	(0.673)	
Mean of Underlying Dependent Variable, 2000	0.578	0.834	-0.368	0.446	0.446
Effect of Average Change in LSI, 1980-2000	0.00129	-0.0138	0.0111	-0.00899	-0.00839
Proportion Explained by Weakened Correlation					-0.104
Number of Observations	708	708	708	708	

Notes: Estimates in Panel A were constructed without California. Panel B uses the full sample but instead of two age groups (22-32, 33-43),MSA-year-age groups are constructed using three age groups (22-28, 29-35, 36-43). Panel C reverts to the two age groups but replaces the incidence of childbearing with marriage rates. Results reported in the first three columns were estimated using one regression run on stacked data for each of the three outcomes. All models include time-varying region fixed effects, MSA fixed effects, age-group fixed effects, the log of income per male college graduate, and the following group-level characteristics: the proportion black, the proportion who are of another non-white race, and the proportion married. Each of the observation-cells is weighted by the population of women represented by the cell, and the robust standard errors in parentheses are clustered by MSA. Reported effects are the change in the underlying dependent variable that would be caused by the change in the share working age low-skilled immigrant experienced by the average member of the sample. *p < .10; **p < .05; ***p < .01

Appendix 1

$$\widehat{F}^1 = \phi(\hat{\beta}_1 LSI_{mt} + \hat{\boldsymbol{\omega}}_1'\mathbf{v}_{gmt})\Phi\left(\frac{(\hat{\beta}_2 LSI_{mt} + \hat{\boldsymbol{\omega}}_2'\mathbf{v}_{gmt}) - (\hat{\beta}_3 LSI_{mt} + \hat{\boldsymbol{\omega}}_3'\mathbf{v}_{gmt})(\hat{\beta}_1 LSI_{mt} + \hat{\boldsymbol{\omega}}_1'\mathbf{v}_{gmt})}{\sqrt{1-(\hat{\beta}_3 LSI_{mt} + \hat{\boldsymbol{\omega}}_3'\mathbf{v}_{gmt})^2}}\right)$$

$$\widehat{F}^2 = \phi(\hat{\beta}_2 LSI_{mt} + \hat{\boldsymbol{\omega}}_2'\mathbf{v}_{gmt})\Phi\left(\frac{(\hat{\beta}_1 LSI_{mt} + \hat{\boldsymbol{\omega}}_1'\mathbf{v}_{gmt}) - (\hat{\beta}_3 LSI_{mt} + \hat{\boldsymbol{\omega}}_3'\mathbf{v}_{gmt})(\hat{\beta}_2 LSI_{mt} + \hat{\boldsymbol{\omega}}_2'\mathbf{v}_{gmt})}{\sqrt{1-(\hat{\beta}_3 LSI_{mt} + \hat{\boldsymbol{\omega}}_3'\mathbf{v}_{gmt})^2}}\right)$$

$$\widehat{F}^3 = \frac{1}{2\pi\sqrt{1-(\hat{\beta}_3 LSI_{mt} + \hat{\boldsymbol{\omega}}_3'\mathbf{v}_{gmt})^2}}\exp\left(\frac{-.5[(\hat{\beta}_1 LSI_{mt} + \hat{\boldsymbol{\omega}}_1'\mathbf{v}_{gmt})^2 + (\hat{\beta}_2 LSI_{mt} + \hat{\boldsymbol{\omega}}_2'\mathbf{v}_{gmt})^2 - 2(\beta_1 LSI_{mt} + \boldsymbol{\omega}_1'\mathbf{v}_{gmt})(\hat{\beta}_2 LSI_{mt} + \hat{\boldsymbol{\omega}}_2'\mathbf{v}_{gmt})(\hat{\beta}_3 LSI_{mt} + \hat{\boldsymbol{\omega}}_3'\mathbf{v}_{gmt})}{1-(\beta_3 LSI_{mt} + \boldsymbol{\omega}_3'\mathbf{v}_{gmt})^2}\right)$$