Immigrant Labor, Child-Care Services, and the Work-Fertility Trade-Off in the United States*

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Abstract

The negative correlation between female employment and fertility in industrialized nations has weakened since the 1960s, particularly in the United States. We suggest that the continuing influx of low-skilled immigrants has led to a substantial reduction in the trade-off between work and childrearing facing American women. The evidence we present indicates that low-skilled immigration has driven down wages in the US child-care sector. More affordable child-care has, in turn increased the fertility of college graduate native females. Although childbearing is generally associated with temporary exit from the labor force, immigrant-led declines in the price of child-care has reduced the extent of role incompatibility between fertility and work.

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1 Introduction

A negative relationship between female market work and childbearing has been widely documented at the individual level. However, considering aggregate data on female labor force participation and fertility rates, the correlation between the two seems to be deteriorating, and this change is most pronounced in the United States (Engelhardt, Kögel and Prskawetz 2004). Relative to other industrialized countries with comparable total fertility rates, female labor force participation is substantially higher in the United States. Similarly, the US fertility rate is much higher than that in developed nations with comparable labor force participation rates (cf. Brewster and Rindfuss 2000). This is especially puzzling, given that government family policies in the United States are far less generous than in comparison countries. We argue that a partial explanation for this phenomenon is the continuous influx of low-skilled immigration to the United States, which has led to an increase in the affordability of child-care services.

There is a large literature that concentrates on the extent to which immigration constrains the opportunities for natives via a crowding-out effect. Much of the existing debate focuses on natives' wages and employment rates.[1] However, an increasing amount of attention is being paid to the impact of immigration in other domains such as schooling, for which existing evidence suggests a negative impact (Borjas 2000, Borjas 2004, Gould, Levy, and Paserman 2005), and crime rates, for which both positive and negative effects have been found (Sampson et al. 2005, Borjas et al. 2006). In this paper we consider a potential complementarity between low-skilled immigration and high-skilled native women in terms of their fertility and work decisions, focusing on magnitude of the trade-off, or "role incompatibility," between the two.

Our analysis makes use of city-level variation in the concentration of immigrants. Given that immigrant location decisions are likely to be related to local economic conditions that also affect native work-employment decisions, we adopt an instrumental variables strategy to identify a causal effect of immigration on child-care costs. Specifically, we rely on the propensity of new entrants to locate in areas with high concentrations of existing immigrants

[1] David Card (e.g. Card 1990, Altonji and Card 1991, Butcher and Card 1991) and George Borjas (2003) are particuarly relevant in this regard. See chapter 7 in Bean and Stevens (2003) for a more comprehensive discussion of the labor-market effects of immigration.

from the same country (e.g. Bartel 1989, Card 2001). Thus, the predicted inflow of immigrants based on their historical distribution across metropolitan areas provides a source of variation in the current distribution that is unrelated to current labor market conditions.

Using a similar approach, Cortes (2006) establishes a baseline association between low-skilled immigration and the price of locally traded goods and services in US cities. We narrow the focus to the costs of a market-provided service that is a particularly important substitute for its home-produced equivalent with respect to fertility decisions, namely child-care. We find that low-skilled immigration has led to substantial reductions in the wages of workers in the child-care sector as a whole.

We then analyze the effects of the lowered cost of child services on the fertility and labor force participation decisions of college-graduate U.S.-born women. Our main finding is that lower child-care costs due to immigration significantly increased the joint likelihood of child-bearing and labor force participation between 1980 and 2000, indicating a reduction in role incompatibility between the two. This result complements the analysis of Cortes and Tessada (2007), who find that low-skilled immigration to the U.S. has led to increases in the hours worked among employed female native college graduates, an effect that is most evident among those with young children.

The remainder of the paper proceeds as follows. In Section 2, we place our paper within the context of the related literatures on fertility, labor supply, and immigration. Section 3 follows with a description of the data and the methods used in the analysis. Results are discussed in Section 4, while Section 5 provides additional discussion and concluding remarks.

2 Background

Economic models of household decision-making focus on the allocation of time across market work, production of household goods, and rearing children (Becker 1965, Willis 1973). The highly time-intensive nature of child-rearing implies a trade-off between labor supply and fertility, particularly for females because their traditional role has been to perform household work.

Differences in wages, human capital, tastes, and endowments result in a negative relationship between fertility and female employment and can manifest itself in terms of the number of hours worked among the employed (the intensive margin) or the decision to work at all (the extensive margin). In the sociology literature, this phenomenon is often referred to in terms of an incompatibility between the roles of mother and worker. Although there is a greater focus in sociology on the institutional constraints that affect decisions at the extensive margin (e.g. rigidities in employment schedules), it is the underlying time constraint that drives role incompatibility.[2] We will use "role incompatibility" as shorthand for the "trade-off between female employment and fertility."

2.1 Empirical Relationships Between Fertility and Work

Numerous studies have documented a negative association between fertility and female labor supply at the individual level. As emphasized by Lehrer and Nerlove (1986) and Browning (1992), given the common link of time-allocation, fertility and work are simultaneous and intertwined outcomes of a joint decision-making process. Consequently, much of the empirical work on fertility and labor supply has focused on identifying a causal effect of childbearing on employment, using a variety of mechanisms to provide variation in fertility that is exogenous to the other determinants of work decisions. Examples include twin-births (Rosenzweig and Wolpin 1980), the sex composition of existing children (Angrist and Evans 1998) and access to contraception (Bailey 2006).

A parallel literature is devoted to describing and understanding the sources of variation in the extent of role incompatibility. Although cross-country studies (e.g., Ahn and Mira 2002) have shown a reversal in the correlation between the total fertility rate (TFR) and female labor force participation (LFP), substantial differences among countries in family policies, child-care availability, unemployment rates, stability of labor contracts, and gender norms may explain this phenomenon (Brewster and Rindfuss 2000, Adsera 2004, de Laat and Sevilla Sanz 2007). Within-country data indicate that the negative relationship between fertility and female employment remains negative but has weakened substantially since the 1960s (Engelhardt et al.

[2] See Stycos and Weller (1967) and Rindfuss and Brewster (1996).

2004).

2.2 Declining Role Incompatibility in the United States

For the remainder of the paper we focus our attention on the United States, where previous research suggests a particularly stark the decline in the trade-off between work and fertility. We provide summary evidence of this phenomenon based on the LFP and fertility patterns observed among women aged 20 to 40. Our samples are drawn from the March Current Population Survey (CPS; King et al. 2004).[3]

The most commonly cited evidence on the decline in role incompatibility is the large increase over the latter part of the 20th century in the propensity of mothers to work, especially among those with young children (cf. Hotz, Klerman, and Willis 1997). As seen in Figure 1, the increase in the labor force participation among women aged 20 to 40 with a child younger than one rose from just over 20 percent in 1970 to around 60 percent in 2000. Among college graduates, there has been an even sharper increase. In 1970 the LFP rate among college graduate mothers of young children was approximately the same as the overall rate. By 2000 it had risen to approximately 70 percent, a threefold increase.

A potential concern is that looking at the conditional likelihood of work in this manner is not necessarily informative as to changes in the trade-off between work and fertility. Treating the sample frequencies as estimates of the underlying probabilities of employment and childbearing, we can decompose the conditional likelihood of work, L, given fertility, C, as

$$P\left(L \mid C\right) = \frac{P\left(L \text{ and } C\right)}{P\left(C\right)} \; .$$

The observed increase in the conditional likelihood could be driven by increases in the joint probability of working and having a child, reductions in the unconditional likelihood of childbearing, reductions in childbearing among the unemployed, or any combination of the three.

[3]In the graphs presented below, we apply a 3-year moving average to the CPS data so as to smooth out year-to-year fluctuations.

There has, in fact, been a secular increase in the joint likelihood of work and fertility. While it is smaller in magnitude, it is still large in relative terms. Figure 2 shows that the joint likelihood of working and bearing children in the full population of fecund women almost doubled between 1970 and 2000.[4] Among college graduates the joint likelihood more than doubled, increasing from approximately 2.2 percent to over 5 percent.

Other evidence on the decline in role incompatibility relies on the correlation between fertility and LFP, which has been becoming steadily less negative in the United States. Our calculations using CPS microdata indicate that the negative correlation between work and childbearing among fecund college graduate women has fallen by 50 percent between 1970 and 2000. However, as with the rate of employment among mothers, changes in the correlation between fertility and LFP can arise simply due to changes in the rate of childbearing.[5] The more structured statistical model that we employ in our empirical analysis allows us make inferences on changes in the joint likelihood, thus obtaining, in some sense, "cleaner" evidence on the decline in role incompatibility.

2.3 Work, Fertility and Child-care Costs

A common thread in the literature on role incompatibility is the role that government and labor market institutions play in determining the costs of child-rearing. Government family policy is far less generous in the United States than in other industrialized nations in terms of child allowances, paid parental leave mandates, and state-subsidized child-care (Rindfuss and Brewster 1996, Henneck 2003). As a result, American families tend to rely heavily on child-care services provided by the market.

Most of the analysis of the cost of market-provided child-care in the US focuses on effect of childcare costs on the conditional likelihood of work, and generally finds an inverse relationship between the two (see Blau and Robins 1988 and Connelly 1992).[6] However, as

[4] Throughout the paper, our variable measuring fertility is the presence of a child less than or equal to one year old.

[5] For example, a change in childbearing concentrated on the unemployed would alter the correlation between work and fertility but have no effect extent of the trade-off facing employed women.

[6] Similarly, Gelbach (2002) finds that a child's enrollment in public school, which may be thought of as inexpensive child supervision, also increases maternal labor supply.

described above, changes in the conditional likelihood of work may result from changes in the propensity to bear children as well as changes in the joint likelihood of LFP and fertility, the latter of which is our measure of interest.

Whereas the effects of child-care costs on childbearing and work decisions seem straightforward, Blau and Robins (1989) point out that the implications derived from even a simple model of simultaneous decision-making are actually quite complicated. A decrease in the cost of child-care cost is expected to lead to an increase in desired fertility due to a standard price effect. Similarly, cheaper child-care services would increasing desired labor supply due to a lower opportunity cost of market work. However, the baseline time costs associated with childbearing might offset the increase in desired labor supply, effectively reducing labor force participation. It is also possible that the increase in desired labor supply is sufficient to induce a lower likelihood of childbearing. Thus, the net effects on fertility and LFP are ambiguous. However, a common effect is a reduction in role incompatibility: the decrease in the cost of child-care should reduce the likelihood and duration of labor force exit among women who bear children and increase the likelihood of fertility among women in the workforce. More succinctly: there should unambiguously be an increase in the joint likelihood of work and fertility.

Blau and Robins (1989) provide essentially the only other work that looks at the effects of child-care costs on simultaneous employment and fertility decisions. Their analysis is conditional on initial employment status and uses geographic variation in average per-child weekly expenditures as the main measure of child-care costs. Blau and Robins find that higher local child-care expenditures are associated with lower rates of employment among all women and with decreases in childbearing among the non-employed. However, their measure of child-care costs is slightly problematic as higher local expenditures might be the product of a greater local demand for child-care due to preferences or unobserved labor market conditions. In our analysis, we adopt an instrumental variables approach to isolate exogenous variation in the local cost of care.

2.3.1 Immigrants and Child-Care

Using an instrumental variables approach similar to ours, Cortes (2006) considers the effect of low-skilled immigration on the relative prices of non-traded goods. She concludes that immigration improves the purchasing power of high-skilled natives when considering an agglomerate of bundle of locally-puchased goods and services. We limit our analysis to the wages of child-care workers since they are particularly relevant to the work-fertility trade-off.

Child-care is a relatively low-wage occupation (Blau 1993, Helburn and Howes 1996) which may reflect the limited qualifications required even at child-care centers. As can be seen in Table 1, 60 percent of child-care workers in 2000 were low-skilled, defined as having less than or equal to a high school degree, while only about 10 percent had a college degree. By comparison, the corresponding numbers in all non-child-care occupations were approximately 46 and 26 percent, respectively. Immigrants were also overrepresented in the child-care sector comprising 17 percent of employment versus 14.5 percent in all other occupations. Moreover, the immigrants employed in child-care were excessively low-skilled both by comparison to immigrants in other occupations and natives employed in child-care.

Cortes and Tessada (2007) provide evidence that low-skilled immigration during the 1990s increased the labor supply of highly skilled US-born females at the intensive margin. Specifically, they find an increase the number of hours they worked among women with college degrees, conditional on being in the labor force.[7] While this increase is strongest among working women with young children, Cortes and Tessada find the opposite effect when considering the extensive margin of employment. Specifically, low-skill immigration reduced the likelihood of women's participation in the labor force. By adopting a simultaneous decision-making framework, we are able to ascertain whether this effect arises due to women temporarily exiting the labor force to bear children. We are also able to explicitly consider the degree to which immigrant-led reductions in the cost of child-care attenuated the trade-off between work and fertility.

[7] For women with professional degrees, this shift in time allocation appears to have been coupled with a reduction in the time devoted to household work.

3 Data and Methods

Broadly, our analysis proceeds in two steps. First we consider the extent to which immigration has resulted in lower wages in the child-care sector via expansions in the supply of low-skilled labor. Second, we determine whether and how reduced cost of child-care has altered employment and fertility decisions of college-graduate females born in the United States. Throughout, our estimation relies on geographic and temporal differences in immigrant flows as a source of variation. The fundamental unit of analysis is the primary metropolitan statistical area (PMSA), and our data are drawn from the 1970 through 2000 US Census microdata samples via the Integrated Public Use Microdata Series (IPUMS; Ruggles et al. 2004).

The geographic sampling units defined by the Census Bureau have changed over the years, resulting in substantial inconsistencies in the population coverage of a number of PMSAs. This is particularly problematic in 2000; the Census did not release a "metro" sample in that year and only the core of a number of PMSAs is identifiable. Shifting boundaries introduces noise into the estimation of PMSA-level variables, as well as systematic bias for 2000 estimates, given the severe geographic truncation in that year.

We have attempted to create a set of geographically consistent PMSA sampling units based on the maps of the underlying sampling units. We exclude from our sample those PMSAs for which creating consistency requires a substantial contraction or expansion of the geographic definition, but retain those that have experienced natural growth in their boundaries.[8] In a number of instances we had to use the one percent micro-data samples; to reduce the noise in the PMSA-level variables, we restricted the sample to metropolitan areas with over 500,000 residents. Our final sample uses data from the 59 PMSAs listed in the Appendix.

3.1 Immigration and Child-Care Costs

We use the wages of child-care workers in our metro sample as a measure of the costs of market-provided child-care services based on data from 1980, 1990 and 2000.[9] The effects of

[8] The details of our coding procedure are available upon request.

[9] Occupation definitions are based on the consistent classification (1990 basis) system developed by the IPUMS group (Ruggles et al. 2004).

immigration on wages are analyzed in a labor supply-demand framework. Given the relative concentration of immigrants in the child-care sector (cf. Table 1), variation in immigrants as a share of the metropolitan workforce might be treated as a supply-shifter in the market for child-care. A basic regrssion approach to estimating the effects of immigration on wages in metro area m in year t takes the following form:

$$w_{mt} = \beta s_{mt} + \mu_m + \mu_t + \phi f_{mt} + \eta d_{mt} + \gamma I_{mt} + \varepsilon_{mt} \tag{1}$$

where w is the log of the median wage of child-care workers in and s is the share of the working age population (20 to 64) born abroad.[10] The fixed effects μ_m account for persistent differences between PMSAs in the market for child-care, while μ_t represents time fixed effects that capture secular changes in the nation as a whole.

The remainder of the year- and PMSA-specific explanatory variables attempt to isolate the supply-side effect of immigrants by controlling for factors that are likely to affect the demand for child-care. The variable f denotes the proportion of fecund women in their peak childbearing years, which we define as ages 20-30. The proportion of fecund women who have completed college is measured by d, while I controls for log of income per male college-graduate worker.[11] The final element of (1) is a randomly distributed error term. To account for heteroskedasticity due to sampling, we use the number of observations used to construct the wage measure as sampling weights. Additionally employing a cluster-robust variance estimator allows for arbitrary patterns of within-PMSA correlation.

If immigration represents a supply shift, we expect the estimate of β to be negative. However, the endogeneity of the location decisions of immigrants poses a problem for interpreting β in (1) as a causal effect of immigration. For example, an increase the demand for child-care services that is not captured by the controls might attract a greater number of immigrant child-care workers. Alternatively, local economic expansions might simultaneously increase

[10] Given the relatively small sample sizes, we use the median, rather than the average, for wage variables to reduce the impact of outliers.

[11] We use income data only for males because female income is expected to be endogenous with respect to any changes in childbearing and labor force participation decisions brought about by changes in the child-care sector.

the demand for child-care and result in a larger inflow of immigrants. Both scenarios would lead to an attenuated (less negative) estimate of the causal effect of immigration when using ordinary least squares (OLS).

3.1.1 Identification

Given the above discussion, we utilize a straightforward extension of the instrumental variables (IV) strategy employed by Card (2001) that relies on historical immigrant enclaves to disentangle the effects of new immigration on wages from the "pull" factors described above. Specifically, we base our identification strategy on the propensity of new immigrants to locate in areas with a relatively large existing concentration of co-ethnics (e.g. Bartel 1989, Massey et al. 1993, Munshi 2003).

Our instrument for the immigrant proportion of working-age adults, s_{mt}, is

$$\sum_b \phi_m^b \times \Delta_{(t,1970)} N^b \ , \tag{2}$$

where ϕ_m^b is the proportion of immigrants from country-of-birth b living in metro area m in 1970 and $\Delta_{(t,1970)}N^b$ is the overall inflow to the United States from country b between time 1970 and time t. Two conditions need to hold for this to be a valid procedure. First, due to the econometric problems with weak instruments (Stock, Wright, and Yogo 2002), the instrument must have strong predictive power. Second, the instrument must meet the exclusion restriction. If both hold, then using (2) as an instrumental variable for s_{mt} will result in an estimate of β that is not biased due to the problems of endogeneity described above.[12] We address both of these conditions in turn.

While immigrants are generally over-represented in the child-care sector (see Table 1), there are substantial between-group differences. For example, South Asians were 40 percent less likely to be employed in this sector than the average worker in 2000. At the opposite end of the spectrum, the share of Dominicans employed in child-care was almost twice the national

[12]IV will also reduce attenuation of the estimate of β due to measurement error in the percentage foreign born. Of particular concern is the Census' potential undercount of the large number of undocumented immigrants.

average. Consequently, we focus our attention on immigrant groups with a concentration in the child-care sector that was above the national average in 2000. We also limit our selection to immigrant groups in which (a) there were at least 25,000 members present in 1970, (b) there was a positive inflow to the US between 1970 and 2000, and (c) over 55 percent working age adults of the group had no more than a high school education.[13] The first two restrictions help to maximize the strength of the instrument, while the third is intended to limit the degree of potential competition for employment among the immigrant groups of interest and the college-educated women analyzed below. Descriptive statistics on the immigrant groups meeting these criteria are given in Table 2, and the strength of the instrument will be verified empirically.

Card (2001) and Cortes and Tessada (2007) present strong arguments in favor of the exclusion restriction. In our case, the rationale is as follows. Given that we include time and PMSA fixed effects in our empirical specification, the exclusion restriction will hold so long as the initial distribution of immigrants across metro areas is uncorrelated with changes in the relative demand for child-care services across cities 15 to 25 years later. A secondary concern is that immigrants do not have a direct effect on the child-care sector per se and that wages are decreased through a general increase in the pool of labor. Even if this were true, it would still represent an indirect causal influence of immigration. Moreover, it is empirically testable and our evidence indicates that wage effects are driven by expansions in employment in the child-care sector.

3.2 Employment and Fertility Choices

The second step of the analysis attempts to ascertain whether the lowered cost of child-care due to immigration has altered female work and childbearing patterns. We limit our attention to non-Hispanic native women with college degrees, aged 23-39 and not living in group quarters. This represents a pool of potential mothers who would be both most likely to utilize child-care services and least likely to have their own wages directly affected by the low-skilled immigrants included in our instrumental variable.

[13]Note also that we include Puerto Ricans among our immigrant groups. While they are US citizens, they have maintained a semi-autonomous identity and have followed a similar, although perhaps accelerated, process of assimilation as international immigrants.

Our analysis is based on a simultaneous latent variables framework. Female employment and fertility decisions are specified according to

$$
\begin{aligned}
C^*_{imt} &= \alpha' x_{imt} + \varepsilon^C_{imt} &\quad (3)\\
L^*_{imt} &= \delta' x_{imt} + \varepsilon^L_{imt} \ , &\quad (4)
\end{aligned}
$$

where C^*_{imt} and L^*_{imt} are latent variables describing desired childbearing and labor force participation (LFP) of woman i living in metropolitan area m in year t. The associated binary outcomes are L_{imt} and Y_{imt}, where $L_{imt} = 1$ is observed if $L^*_{imt} > 0$ and likewise for fertility.[14] Given that there is no generally applicable exclusion restriction for identifying the effects of childbearing on LFP or vice versa, both equations are will generate reduced-from estimates of the coefficients.[15] That is, the estimates will capture the *net* effect of the explanatory variables on work and fertility decision. The vector of explanatory variables, x_{imt}, contains fixed effects for PMSA and time, controls for age, race, income, and the number of young children born prior to the year of observation.

Also incorporated in x is our key explanatory variable, the wage of child-care workers in the local labor market – w_{mt} from above. Given that the wage bill represents between 60 and 80 percent of the operating cost of even formal child-care centers (Helburn and Howes 1996), changes in wages should translate into substantial changes in the cost of child-care facing women.

Based on the discussion in Section 2.3, the net effects of w on the unconditional likelihood of childbearing and LFP are ambiguous, but a lower cost of child-care should imply a decrease in the role incompatibility between the two. We model this effect in terms of the correlation of the error terms in (3) and (4) among groups of individuals within a PMSA in a given year:

$$
corr\left(\varepsilon^C_{imt}, \varepsilon^L_{imt}\right) = \theta' x_{gmt} \equiv \rho_{gmt} \ , \quad (5)
$$

[14]Labor force participation is based on the self-reported measure. Childbearing is inferred from the presence of an own-child younger than one year old.

[15]The approach of Angrist and Evans (1996) is somewhat limiting in that it requires that the sample be restricted to women with at least two prior births.

for individuals $i \in g$. Assuming that the error terms are bivariate normally distributed, ρ_{gmt} is the tetrachoric correlation between fertility and labor force participation among members of group g in metro area m in year t. Given the time constraints that define role incompatibility, it is expected that ρ_{gmt} will be generally negative, but that reductions in the cost of care will result in a less negative tetrachoric correlation.

In estimating the model, we cannot treat wages in the child-care sector as exogenous for reasons similar to those described above. For example, if desired childbearing and/or labor force participation increases due to unobserved changes in economic conditions or social norms, this will cause a demand-led increase in the wages of child-care workers. Control functions have been developed (e.g. Rivers and Vuong 1988, Blundell and Powell 2004) and widely applied to address endogenous regressors in binary choice models. However, our interest in explicitly parameterizing the tetrachoric correlation does not allow this approach to be extended to the simultaneous choice model described above. Consequently we rely on a slight generalization of Amemiya's (1974) bivariate probit specification for grouped data, which allows a straightforward application of instrumental variables.

3.2.1 Grouped Bivariate Probit

Aggregating observations according to characteristics, metro area and time, analyzing sample proportions allows us to recover the model coefficients, but requires that there be no groups in which either of the binary outcomes is the same for all its members. To avoid uniform outcomes within groups we use divide the sample of college graduate women into two broad groups (g): ages 23-30 and ages 31-39.[16]

Given a bivariate normal distribution, the marginal distribution for the expected rate of childbearing is normal:

$$\pi_{gmt}^{C} \equiv E\left[\frac{1}{n_{gmt}} \sum_{i \in g} C_{igmt}\right] = \Pr\left(C_{igmt} = 1\right) = \Phi\left(\alpha' x_{gmt}\right) \tag{6}$$

[16]Disaggregating further by race is not possible even when using the Census microdata samples. Consequently, we include in x_{gmt} measures of the average characteristics of the group.

where x is a vector of aggregated group characteristics, described below, and $\Phi(\cdot)$ is the standard normal cumulative density function. Similarly,

$$\pi_{gmt}^{L} = \Pr(L_{igmt} = 1) = \Phi(\delta' x_{gmt}) \ , \tag{7}$$

while the expected joint rate of childbearing and LFP is

$$\pi_{gmt}^{CL} = \Pr(C_{igmt} = 1, L_{igmt} = 1) = F\left(\alpha' x_{gmt}, \delta' x_{gmt}, \rho_{gmt}\right) , \tag{8}$$

where $F(\cdot)$ is a standard bivariate normal distribution function.

A linearized representation of the model described in equations (6)-(8) can be derived as in Amemiya (1974) using first-order Taylor expansions of functions of the observed sample proportions $p_{gmt} = \left(p_{gmt}^{C}, p_{gmt}^{L}, p_{gmt}^{CL}\right)$ around their expected values, $\pi_{gmt} = \left(\pi_{gmt}^{C}, \pi_{gmt}^{L}, \pi_{gmt}^{CL}\right)$. Based on our characterization of the tetrachoric correlation in equation (5), the system of equations to be estimated is

$$\begin{aligned} c_{gmt} &= \alpha' x_{gmt} + u_{gmt} && (9) \\ \ell_{gmt} &= \delta' x_{gmt} + v_{gmt} && (10) \\ r_{gmt} &= \theta' x_{gmt} + w_{gmt} \ , && (11) \end{aligned}$$

where $c_{gmt} \equiv \Phi^{-1}\left(p_{gmt}^{C}\right)$, $\ell_{gmt} \equiv \Phi^{-1}\left(p_{gmt}^{L}\right)$, and r_{gmt} is the implicit solution to

$$p_{gmt}^{CL} = F(c_{gmt}, \ell_{gmt}, r_{gmt}) \equiv G\left(p_{gmt}^{C}, p_{gmt}^{L}, r_{gmt}\right) \ . \tag{12}$$

Given that $F(\cdot)$ is monotonically increasing in the third argument (Tihansky 1972), conditional on any observed proportions of women bearing children and participating in the labor force, an increase in the joint likelihood of doing both will translate into a higher estimate of r. That is, *ceteris paribus*, a less negative value of the estimated tetrachoric correlation implies a higher joint likelihood of fertility and employment.

3.2.2 Implementation

Based on Census micro-samples, we compute the sample proportions p_{gmt} as the weighted average of the binary outcomes in the corresponding group.[17] The corresponding values of c_{gmt} and ℓ_{gmt} can be quickly calculated using any statistical package. Given that $F(\cdot)$ is monotonic in the correlation coefficient, we apply a recursive binary chop to search for the r_{gmt} that solves

$$\left|p_{gmt}^{LY} - F\left(c_{gmt}, \ell_{gmt}, r_{gmt}\right)\right| < k$$

for a pre-defined level of precision, k, which we set to 2^{-50}.

Since all of the estimating equations (9)-(11) include the same explanatory variables, there is no efficiency gain from using seemingly unrelated regression techniques.[18] As a result, we estimate a series of independent equations of the form

$$y_{gmt} = \beta w_{mt} + \mu_m + \mu_{gt} + \gamma I_{mt} + \omega \bar{n}_{gmt} + \lambda^b \bar{p}_{gmt}^b + \lambda^o \bar{p}_{gmt}^o + \varepsilon_{gmt} \ , \qquad (13)$$

where y is one of the dependent variables (c, ℓ, r), w denotes log-wages of child-care workers, μ_m is a PMSA fixed effect, μ_{gt} is a time- and age-group-specific fixed effect, I_{mt} is the log of income per worker among college-graduate males in PMSA m in year t.[19] The average number of own-children ages 1-5 living with the women in each group is denoted by $\bar{n}$. Finally, the variables $\bar{p}^b$ and $\bar{p}^o$ denote the proportions of the group that are black and other non-white and are based on IPUMS single-race coding system that bridges the differing 1990 and 2000 Census classification schemes.

While there is no clear theoretical prediction regarding the sign of coefficient on w in the fertility and LFP equations, we expect a higher cost of child-care to reduce the negative correlation between the two. That is, β should be less than zero when the dependent variable is

[17] Census-provided sampling weights are used to ensure representativeness.

[18] We do, however, weight by the number of observations in each (year) x (PMSA) x (age-group) cell to account for heteroskedasticity and standard errors remain clustered at the PMSA level.

[19] Using actual spousal income is problematic if cheaper childbearing costs leads to reduced selection into marriage. Given the prevalence of assortative mating based on education (Mare 1991), the income of male college graduates provides a reasonable proxy for potential spousal income and allows us to retain unmarried women in the sample.

r. As already noted, estimating (13) using OLS may yield biased results due to the endogeneity of the wages of child-care workers with respect to local economic conditions, as well as any innate preferences over childbearing and employment among local women. In areas where mothers find it more desirable to work, perhaps due to favorable labor market opportunities, they may bid up the price of child-care. In this scenario, as well as when the wages of child-care workers are measured with noise, OLS will yield an estimate of β that is smaller in absolute value than the true coefficient.

To account for endogeneity and measurement error, we use the predicted flow of the low-skilled immigrants listed in Table 2 as an instrument for wages in the child-care sector. The criterion that the immigrant groups included in the instrument have an excessive share of working-age adults with no more than a high school education (58-82 percent) fulfills an additional exclusion restriction. Specifically, there is unlikely to be competition for jobs between the selected immigrants and the college-graduate native women in our sample, and so no direct labor-market-driven effect on the desired childbearing and LFP among the native women. Thus, instrumental variables estimation using (2) as an instrument for the wages of child-care workers should yield coefficient estimates that are closer to causal effects of immigrant-led changes in the cost of care.

4 Results

Wages in the Child-Care Sector

Our estimates of the effects of immigration on wages in the child-care sector are presented in Table 3. The first column gives the estimate based on OLS applied to our panel of PMSA-level data including only PMSA and year fixed effects.[20] The coefficient on the share of immigrants is negative and highly significant ($p < 0.01$), which suggests that immigration represents a supply, rather than demand, shock in the market for child-care services. Adding the demand-side controls lends credence to this hypothesis. As seen in the second column of Table 3, the demand controls are individually predictive of wages in the child-care sector. However, their

[20]Due to the large number of fixed effects and time-interacted controls, we only present the statistics that are pertinent to our variables of interest.

inclusion results in a minimal change in the coefficient on the immigrant share and a negligible increase in the overall explanatory power of the model.

Shifting to the instrumental variables (IV) strategy described above, the first-stage test statistics fall well above conventional cut-off points for weak instruments (cf. Stock et al. 2002), which indicates that our instrument does a very good job in predicting the immigrant share of the PMSA working-age population.[21] The second-stage estimate of the effect of immigration is reported in the third column of Table 3. The IV estimate remains highly significant (p $<$ 0.01) and rises in absolute value as compared to the OLS estimate reported in the second column. This can be interpreted as evidence that endogeneity of immigrant location decisions tends to attenuate the OLS estimate of the effect of immigration on wages. Alternatively, OLS estimates might be biased toward zero due to measurement error arising from undocumented immigrants.

The overall pattern of negative and significant coefficients might seem surprising given that the majority of previous research indicates small or nonexistent wage effects of immigration on natives (Friedberg and Hunt 1995, Card 1990, Card 2001 – but see Borjas 2003 for a notable exception). However, as discussed in Bean and Stevens (2003) much of this research is based on examining broad skill classes, rather than specific occupations. Child-care might also represent a relatively unique case, as it is a very labor-intensive occupation, which provides little room for capital adjustments to accommodate growth in the labor force.

Given the semi-log specification, the estimated coefficient represents the percentage change in the wages of child-care workers caused by a one percentage point increase in the immigrant share of the working-age population. Between 1980 and 2000, the average share of immigrants in our PMSA sample rose from 9.3 percent to 18.7 percent. Based on our IV point estimate, an increase of the same magnitude in the immigrant share of the working-age population would lead to a decline in wages in the child-care sector of approximately 15.7 percent.[22]

[21] Stock et al. 2002 suggest a cut-off point for the Cragg-Donald statistic of 16, which our instrument far surpasses. One limitation of the Cragg-Donald statistic, however, is that it is not robust to heteroskedasticity and clustering. However, the cluster-robust F-statistic associated with the instrument is also well above the recommended threshold of ten.

[22] Note that the growth in the average share of immigrants over time in our sample is substantially less than the between-PMSA standard deviation within any given year.

The fourth column of Table 3 indicates that the wage effects of immigration arise from expansions of labor supply in the child-care sector. The IV estimate indicates that an increase the local share of low-skilled immigrants from 9.3 to 18.7 percent increases the local employment concentration in child-care by over half a standard deviation. Thus, due to increases in the supply of child-care workers, immigration has led to sizeable reductions in the wages of child-care workers.

Fertility and Work Decisions

Table 4 considers the extent to which these cost reductions have altered women's childbearing and work patterns using the grouped bivariate probit model described above. Based on OLS, wages in the child-care sector have a negative relationship with fertility, but no statistically significant relationship with the other dependent variables. However, as already noted, OLS estimates are likely to be biased toward zero. We focus on the instrumental variables estimates, which theoretically isolate exogenous changes in the wages of child-care workers arising from the supply push of immigrants.

The IV estimates presented in panel B of Table 4 indicate that lower wages in the child-care sector are associated with both higher fertility and lower labor force participation rates. Over the sample frame, the marginal effect of the immigrant-induced 15.7 percent reduction in wages in the child-care sector on the likelihood of childbearing was approximately 1 percentage point.[23] The comparable estimate for the effect on the likelihood of labor force participation was 0.95 percentage points. Taken together, these results suggest that a lower cost of care induces temporary exit from the labor force to bear children. However, the reductions in labor supply are slightly smaller than the associated increases in fertility. This is consistent with the attenuation of the negative correlation between work and fertility in response to decreases in the cost of care as indicated in the final column of Table 4.

Role Incompatibility

Between 1980 and 2000, the joint likelihood of labor force participation and childbearing in

[23] Marginal effects are evaluated at the mean.

our metro sample of non-Hispanic college graduate women rose from 3.5 percent to 5.8 percent, an increase similar to that seen in Figure 2 above. While there were a host of social and economic changes over that time frame that reduced role incompatibility, we can utilize our instrumental variables estimates to assess the contribution of low-skilled immigration. Specifically, consider the following counterfactual scenario: suppose that the low-skilled immigrant share of the workforce remained constant at its 1980 level. Our estimates above suggest that wages in the child-care sector in 2000 would have been 15.7 percentage points higher than they were in actuality.

The first column of Table 5 lists the average 2000 values of the rate of childbearing (p_0^C), the rate of labor force participation (p_0^L), and the tetrachoric correlation between them (r_0) for our metro sample. The second column presents counterfactual values for these variables assuming a 15.7 percentage point increase in wages in the child-care sector in 2000. For fertility and labor force participation the counterfactual values ($\tilde{p}^C$ and $\tilde{p}^L$, respectively) are based on the marginal effects given above. The counterfactual tetrachoric correlation $\tilde{r}$ is calculated based on the IV regression coefficient reported in Table 4.

In the "no-immigration" counterfactual scenario, the bivariate probit structure of the model (see equation (12)) suggests that the share of women giving birth and remaining in the labor force in 2000 would have been

$$\tilde{p}^{CL} = G\left(\tilde{p}^C, \tilde{p}^L, \tilde{r}\right) = 4.8\%$$

in the absence of immigration. As already noted, the actual joint likelihood rose from 3.5 percent in 1980 to 5.8 percent in 2000. Thus, reductions in the cost of child-care due to low-skilled immigration potentially explains up to one-third of the secular increase in the joint likelihood of work and fertility.

5 Conclusion

In this paper, we have provided a partial explanation for why the role incompatibility between market work and childbearing is less severe for American women, relative to their counterparts in other highly-developed nations. The United States does not have generous family leave policies, nor does it provide large cash benefits for childbearing. However, the U.S. does receive more immigrants than any other nation in the world.

This research builds on a growing body of work highlighting the potentially beneficial effects that immigration has on the purchasing power of natives (Cortes 2006). In order to isolate a causal impact of immigration, we relied on a commonly used instrumental variables approach to account for the simultaneity of the location decisions of new migrants with respect to local labor market conditions. Using predicted settlement patterns based on historical enclaves, we found that low-skilled immigration to the United States between 1980 and 2000 led to substantial reductions in wages in the child-care sector.

Our results indicate that a lower cost of child-care due to immigration has, in turn, significantly altered the employment and fertility decisions of college graduate female natives. By adopting a simultaneous choice framework we are able to explain Cortes and Tessada's (2007) finding that low-skill immigration reduces the likelihood of employment among high-skilled females. Namely, immigrant-driven reductions in the cost of child-care have increased the fertility native college graduates, resulting in temporary exit from the labor force. However, lower child-care costs have also reduced the role incompatibility facing new mothers and thus attenuated the negative relationship between fertility and market work.

One limitation of the current analysis is that, due to the exclusion restrictions required by the instrumental variables approach, we could not analyze the impact of low-skilled immigration on US-born Hispanic women. One might expect a particularly strong complementarity between high-skilled native Hispanics and low-skill migrants from Latin America. Similarly, we were not able to include in our sample any low-skilled natives, for whom time constraints might be more binding than the college graduate women we have analyzed. The extent to which these groups have benefited from the increased availability and affordability of child-care due to immigration should remain a topic for future research.

With respect to the highly-educated women that comprise our sample, the popular press has raised concerns about the so-called "Opt-out Revolution" (Belkin 2003, Wallis 2004), whereby women on the career track appeared to be increasingly likely to drop out of the labor force upon childbearing. These findings come in contrast to Goldin's (2004) assessment that, relative to older cohorts, women graduating from college in the 1980s have been significantly better able to combine both career and family. Our analysis indicates that women are in fact facing smaller trade-offs when making fertility and labor supply decisions. Although changes in overall labor market conditions, institutions and social norms have indubitably led to reductions in role incompatibility, our results indicate that this reduction is in part due to the continuing flow of immigrant child-care workers into the United States.

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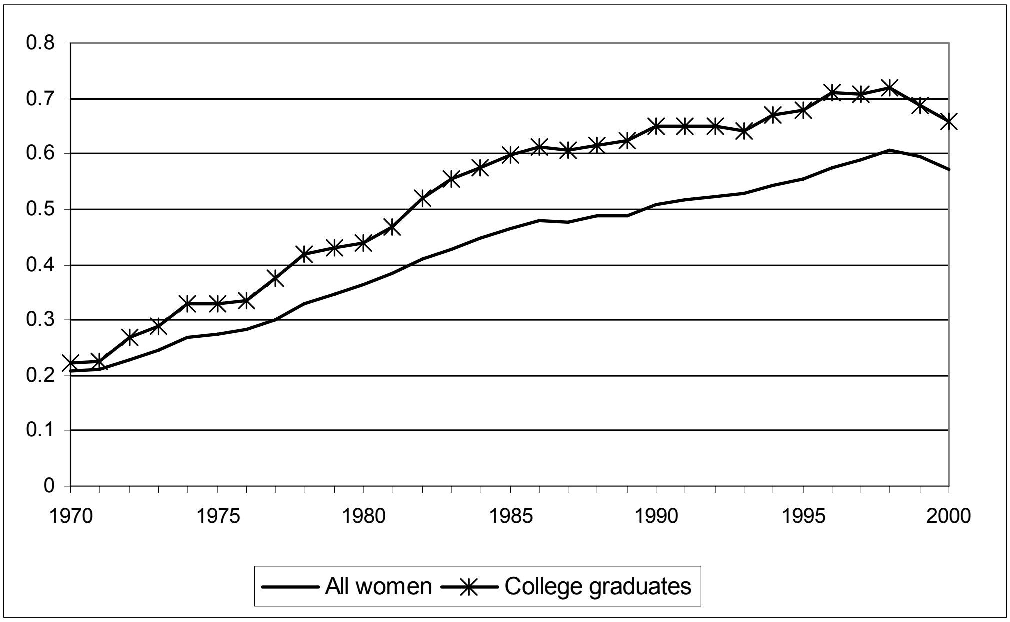
Figure 1: Employment Rates Among Mothers with Children Younger than One



Data Source: March Current Population Surveys, 1969-2001 (King et al. 2004)
Notes: The figure plots the rate of labor force participation among mothers aged 20-40 living an own-child children younger than one year old. College graduates are defined based on the consistent educational recode developed by the IPUMS group (Ruggles et al. 2004). The series are smoothed using a 3-year moving average centered on the given year.

Figure 2: Joint Rate of Child-Bearing and Employment

Data Source: March Current Population Surveys, 1969-2001 (King et al. 2004)
Notes: The figure plots the proportion of women aged 20-40 reporting labor force participation and living with children younger than one-year old. College graduates are defined based on the consistent educational recode developed by the IPUMS group (Ruggles et al. 2004). The series are smoothed using a 3-year moving average centered on the given year.

Table 1: Educational Distribution and Immigrant Concentration in Child-Care, 2000

Characteristic	Frequency by Occupation: Child Care	Frequency by Occupation: All Other
Less than high school	17.9%	12.6%
Less than or equal to high school	59.9%	45.8%
Bachelor's degree or higher	10.2%	25.7%
Percent Immigrant	17.0%	14.5%
Less than high school	38.2%	30.2%
Less than or equal to high school	72.0%	56.9%
Bachelor's degree or higher	9.5%	24.2%

Data Source: 2000 Census public-use micro-data file (Ruggles et al. 2004)
Notes: Child care is defined based on the consistent 1990-basis occupation classification developed by the IPUMS group (Ruggles et al. 2004).

Table 2: Characteristics of Immigrant Groups Used for Instrument

Country of Origin	Number in 1970	Change, 1970-2000	Low-Skilled in 2000
Dominican Republic	41,275	505,759	65.5%
Ecuador	26,300	213,645	58.1%
Mexico	549,125	6,689,664	82.3%
Puerto Rico	574,225	484,803	64.0%

Data Sources: 1970 and 2000 Census public-use micro-data files (Ruggles et al. 2004)
Notes: “Low-skilled” is defined as having no more than a high school degree based based on the consistent educational recode developed by the IPUMS group (Ruggles et al. 2004).

Table 3: Effects of Immigration on the Market for Child-Care Services

Dependent Variable	Log Median Wage in Child-Care			% of LF in Child-Care
Specification	OLS	OLS	IV	IV
% of Working-Age Adults Born Abroad	-0.849*** (0.293)	-0.885*** (0.230)	-1.816*** (0.284)	0.017*** (0.006)
% of Fecund Women, Ages 20-30		0.685 (0.651)	1.371** (0.647)	-0.011 (0.012)
% of Fecund Women, College Graduates		1.762*** (0.496)	0.826 (0.458)	0.032*** (0.012)
Log Income per Worker, Male College Graduates		0.311 (0.198)	0.727*** (0.183)	-0.014*** (0.004)
R-squared	0.9715	0.9795	--	--
Cragg-Donald Statistic	--	--	35.64	35.64
Cluster-Robust F, Instrument	--	--	67.24	67.24
Mean of dependent variable	1.680			0.010
S.D. of dependent variable	0.370			0.003
Number of observations	177			177

Data Sources: 1970, 1980, 1990 and 2000 Census public-use micro-data files (Ruggles et al. 2004)
Notes: Child care is defined based on the consistent 1990-basis occupation classification developed by the IPUMS group (Ruggles et al. 2004). Each column represents a different model specification applied to the panel 59 PMSAs described in the text between 1980 and 2000. All specifications include the PMSA and year fixed effects. The instrument used in the IV specifications is derived from the 1970 distribution of selected immigrant groups across PMSAs and the net flow to the U.S. after 1970 – see equation (2). PMSA-year observations are weighted by the number of child-care workers to account for heteroskedasticity, and the robust standard errors in parentheses are clustered at the PMSA level. Finally, *, **, and *** represent significance at the 10, 5 and 1 percent levels.

Table 4: Immigration, Wages in Child-Care, and Fertility-Work Decisions of Non-Hispanic Native College Graduates

Dependent Variable	c		ℓ		r	
Specification	OLS	IV	OLS	IV	OLS	IV
Metro-Level Variables:						
Log Wage in Child-Care	-0.108**	-0.386***	0.032	0.229*	-0.048	-0.184**
	(0.051)	(0.133)	(0.056)	(0.125)	(0.043)	(0.088)
Log Income per Worker Male College Graduates	0.608***	0.763***	-0.198*	-0.308*	-0.082	-0.006
	(0.193)	(0.244)	(0.119)	(0.168)	(0.120)	(0.137)
Within-Metro Group Variables:						
Proportion Black	0.185	0.177	0.120	0.126	0.003	-0.001
	(0.432)	(0.397)	(0.291)	(0.317)	(0.320)	(0.298)
Proportion Other Non-White	-4.31***	-4.75***	-1.875***	-1.568***	1.161***	0.951**
	(0.880)	(1.029)	(0.467)	(0.518)	(0.436)	(0.467)
Average Number of Own-Children Aged 1-5	0.978***	1.001***	-0.978***	-0.999***	0.380***	0.395***
	(0.230)	(0.235)	(0.126)	(0.129)	(0.095)	(0.095)
R-squared	0.6009	--	0.9558	--	0.7981	--
Cragg-Donald Statistic	--	42.21	--	42.21	--	42.21
Cluster-Robust F, Instrument	--	36.19	--	36.19	--	36.19
Mean of dependent variable	-1.402		1.023		-0.454	
S.D. of dependent variable	0.148		0.298		0.148	
Number of observations						

Data Sources: 1970, 1980, 1990 and 2000 Census public-use micro-data files (Ruggles et al. 2004)

Notes: Child care is defined based on the consistent 1990-basis occupation classification developed by the IPUMS group (Ruggles et al. 2004). All models are estimating using the grouped data for non-Hispanic U.S.-born college graduate women living in the 59 PMSAs described in the text between 1980 and 2000. The dependent variables c and ℓ are the normits of the unconditional likelihoods of child-bearing and labor force participation, while r is the tetrachoric correlation derived from equation (12). All specifications include PMSA and (year)×(age-group) fixed effects. The instrument used in the IV specifications is derived from the 1970 distribution of selected immigrant groups across PMSAs and the net flow to the U.S. after 1970 – see equation (2). PMSA-year observations are weighted by the number of observations used to calculate the grouped dependent variables and robust standard errors are in parentheses and are clustered at the PMSA level. Finally, *, **, and *** represent significance at the 10, 5 and 1 percent levels.

Table 5: Actual and Counterfactual Work and Fertility Measures, 2000

Variable	Actual Mean	Counterfactual Mean
Fertility Rate	0.0902	0.0799
Labor Force Partcipation (LFP) Rate	0.8432	0.8527
Tetrachoric Correlation, Fertility and LFP	-0.4001	-0.4316

Data Sources: 2000 Census public-use micro-data files (Ruggles et al. 2004) and authors' calculations.
Notes: Actual means are based on the sample of 59 PMSAs used in the main analysis. The counterfactual means assume a 15.7 percent increase in wages in the child-care sector.

Appendix Table 1: Consistent-Boundary Primary Metropolitan Statistical Areas Used for Estimation

Akron, OH	New Orleans, LA
Albany-Schenectady-Troy, NY	New York-Northeastern NJ
Allentown-Bethlehem-Easton, PA/NJ	Nassau Co, NY
Austin, TX	Jersey City, NJ
Baltimore, MD	Middlesex-Somerset-Hunterdon, NJ
Birmingham, AL	Newark, NJ
Boston, MA	Norfolk-VA Beach-Newport News, VA
Buffalo-Niagara Falls, NY	Orlando, FL
Charlotte-Gastonia-Rock Hill, SC	Philadelphia, PA/NJ
Chicago-Gary-Lake, IL	Phoenix, AZ
Gary-Hammond-East Chicago, IN	Pittsburgh-Beaver Valley, PA
Cleveland, OH	Portland-Vancouver, OR
Columbus, OH	Raleigh-Durham, NC
Dallas-Fort Worth, TX	Richmond-Petersburg, VA
Denver-Boulder-Longmont, CO	Riverside-San Bernadino, CA
Detroit, MI	Sacramento, CA
Fort Lauderdale-Hollywood-Pompano Beach, FL	St. Louis, MO-IL
Fresno, CA	Salt Lake City-Ogden, UT
Grand Rapids, MI	San Antonio, TX
Greensboro-Winston Salem-High Point, NC	San Diego, CA
Hartford-Bristol-Middleton-New Britain, CT	San Francisco-Oakland-Vallejo, CA
Houston-Brazoria, TX	San Jose, CA
Indianapolis, IN	Seattle-Everett, WA
Kansas City, MO-KS	Syracuse, NY
Los Angeles-Long Beach, CA	Tampa-St. Petersburg-Clearwater, FL
Anaheim-Santa Ana-Garden Grove, CA	Tucson, AZ
Miami-Hialeah, FL	Ventura-Oxnard-Simi Valley, CA
Milwaukee, WI	Washington, DC/MD/VA
Minneapolis-St. Paul, MN	West Palm Beach-Boca Raton-Delray Beach, FL
Nashville, TN	