Female Work and Fertility in the United States: Effects of Low-Skilled Immigrant Labor*

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Abstract
This paper examines the effects of low-skilled immigration on the fertility and work decisions of high-skilled women born in the United States. We find that low-skilled immigration to urban areas between 1980 and 2000 lowered the market cost of household services. College-educated native females responded by increasing fertility and reducing labor force participation. However, low-skilled immigration also weakened the negative correlation between work and fertility. Together, these changes resulted in an increase in the joint likelihood of childbearing and labor force participation. Our results imply that the continuing influx of low-skilled immigrants has substantially reduced fertility-work tradeoffs facing educated women.

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1 Introduction

The foreign-born population of the United States has quadrupled since the passage of the Immigration and Nationality Act in 1965, and the immigrant share of the U.S. population is at its highest point since the beginning of the 20th century. Among politicians and academics, this has led to substantial interest in the socioeconomic consequences of the recent waves of immigration to the United States. Much of the existing research focuses on the potentially negative impact of immigration on the wages and employment rates of natives (Borjas 2003; Card 2001). Less attention has been paid to the potential benefits accruing to natives from immigration. In this paper, we consider the impact of low-skilled immigration on the market cost of household services in large urban areas. We then consider the childbearing and labor supply responses of high-skilled U.S.-born females. We pay particular attention to the magnitude of the tradeoff, or "role incompatibility," between fertility and labor force participation.

Our analysis makes use of geographic variation in the concentration of immigrants. Because immigrant location decisions are likely to be related to current local economic conditions, we use an instrumental variables approach that relies on the propensity of new entrants to locate in areas with high historical concentrations of immigrants from the same country (Bartel 1989; Card 2001). Using a similar approach, Cortes (2008) estimates the broad impact of low-skilled immigration on an agglomerated bundle of local goods and services in U.S. cities. We present a more detailed examination of market-provided services that offer close substitutes for time-intensive childrearing tasks: child care, food preparation, and housekeeping. Our estimates indicate that low-skilled immigration between 1980 and 2000 reduced wages at a range of quantiles of the distribution, which suggests reductions in the price of household services across the quality spectrum. The wage decreases were accompanied by increases in occupation-specific employment, providing further evidence that we are identifying the effects of labor supply shocks.

Decreases in the price and increases in the availability of household services brought about by low-skilled immigration should imply considerable reductions in the cost of childrearing. However, the theoretical impact of lower childrearing costs on childbearing and labor force participation (LFP) choices is unclear due to the joint nature of the decision (Blau and Robins 1989). We consider these relationships empirically using an estimation technique that allows us to examine the simultaneous impact of low-skilled immigration on the fertility and LFP outcomes of U.S.-born women. This analysis again relies on the enclave-based instrumental variables strategy to establish a causal relationship. We focus on non-Hispanic college graduates in order to identify effects of immigration that result from changes in low-skilled service markets.[1] Our estimates indicate that inflows of low-skilled immigrants resulted in a higher rate of childbearing in this population of high-skilled women living in large urban areas. Increases in fertility were accompanied by slightly lower reductions in labor force participation rates, suggesting that the baseline tradeoff between work and fertility was not one-for-one.

These results complement the analysis of Cortes and Tessada (2009), who find that low-skilled immigration to large U.S. metropolitan areas led to increases in the number of hours worked conditional on being employed. Our findings are also consistent with a previous analysis demonstrating that inflows of immigrants to a metropolitan area attenuate the negative correlation between work and fertility among urban high-skilled women (Furtado and Hock 2010). As pointed out there, the joint likelihood of childbearing and labor force participation represents a more tangible measure of role incompatibility. The structured statistical model

[1] As discussed below, the sample restriction limits the degree of direct labor-market competition and non-market interactions that might arise from inflows of low-skilled immigrants that are predominantly Latin American in origin.

adopted in this paper allows us to combine our estimates to obtain a total effect of immigration on the joint likelihood, as well as to decompose the total effect into a component arising from changes in the rates of fertility and LFP and a component attributable to changes the correlation between the two. We find that low-skilled immigration increased the joint likelihood of childbearing and LFP and that between 25 and 40 percent of this change was attributable to a weakening of the negative fertility-work correlation.

The paper proceeds as follows. In Section 2, we place our paper within the context of the literature on fertility, labor supply, and childrearing costs, describing how low-skilled immigration might have reduced role incompatibility in the United States. A brief description of the data used in our analysis follows in Section 3. Our analysis of household services markets is described in Section 4, which also lays out the instrumental variables estimation strategy used to isolate the causal effect of low-skilled immigration. In Section 5, we describe the empirical model of simultaneous decision making that underpins our investigation of childbearing and labor force participation patterns and discuss how the estimated parameters may be interpreted. After presenting the main results, we conduct specification checks concerning the validity of the estimation method and the extent to which geographic and educational selection among natives might affect our results. Finally, Section 6 provides additional discussion and concluding remarks.

2 Background

The highly time-intensive nature of childrearing implies a tradeoff between fertility and labor supply, particularly for females because their traditional role has been to perform household work (Becker 1985; Willis 1973). In the sociology literature, this phenomenon is often referred to in terms of an incompatibility between the roles of mother and worker (Stycos and Weller

1967).[2] Given the common link of time-allocation, fertility and work are simultaneous outcomes of a joint decision-making process. 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. We are not aware of any research that attempts to quantitatively assess the determinants of this secular trend. In this section, we describe how low-skilled immigration to the United States may have, by reducing the cost of childrearing, altered work and fertility patterns, and ultimately contributed to the observed decline in role incompatibility.

2.1 Role Incompatibility in the United States

The most commonly cited evidence on the decline in role incompatibility is the substantial increase over the latter part of the 20th century in the propensity of mothers to work, especially in the years soon after giving birth (cf. Hotz, Klerman, and Willis 1997). However, labor force participation rates of recent mothers may not necessarily be informative as to changes in the tradeoff between work and fertility. The main issue is that observed changes in the conditional rate of employment may be affected by selection into (and out of) motherhood, rather than changes in the likelihood of work among otherwise-comparable mothers.

By contrast, concurrent changes in fertility and LFP are reflected in the joint likelihood, which, as a result, may better summarize role incompatibility than the conditional likelihood of work. 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

[2] We will use "role incompatibility" as shorthand for the "tradeoff between female employment and fertility." Throughout the paper we will also use the term "work" to denote labor force participation, rather than employment *per se*, since the former indicates an intention to be employed.

among adult fecund women almost doubled between 1970 and 2000.[3] Among college graduates it more than doubled, increasing from approximately 2.5% to 4.9%. These trends suggest that the work-fertility tradeoff weakened substantially over this time frame, particularly so for college-educated women facing a high opportunity cost of time spent out of the labor force.

Other evidence on the decline in role incompatibility relies on the correlation between fertility and labor force participation, which has been becoming steadily less negative in the United States (Englehardt et al. 2004, Furtado and Hock 2010). However, using the correlation coefficient as a metric for role incompatibility is problematic because it does not directly quantify changes in observable outcomes. The structured statistical model in our empirical analysis allows us to translate the weakening of the correlation, in combination with changes in the marginal likelihoods of childbearing and LFP, into changes in the joint likelihood. As a result, we are able to tie changes in the fertility-LFP correlation into the more readily interpretable evidence on the decline in role incompatibility.

2.2 Childrearing Costs, Role Incompatibility, and U.S. Immigration

It seems plausible that the reductions in role incompatibility witnessed in the United States were driven by reductions in childrearing costs. The relationships between the cost of childrearing, fertility, and work decisions derived from even a simple economic model of simultaneous decision-making are fairly complicated (Blau and Robins 1989). A decrease in childrearing costs may increase desired fertility due to a standard price effect and increase desired labor supply by reducing the relative value of time spent at home. However, the baseline time costs associated

[3] 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, who we refer to as "adult fecund women." Throughout the paper, educational attainment is based on the temporally consistent classification system developed by the IPUMS group (Ruggles et al. 2010). We also 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.

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. In fact, Lehrer and Kawasaki (1985) suggest that when adequate childcare is not affordable, women may devote all of their energy into their domestic roles, thus increasing fertility. Hence, the net effects of changes in childrearing costs on fertility and LFP are ambiguous. Nevertheless, a decrease in the cost of childrearing should reduce the likelihood and duration of labor force exit among women who bear children and should increase the fertility rate of women in the workforce. This should unambiguously lead to an increase the joint likelihood of work and fertility.

Despite large increases in the demand for child care in the United States, there has been only a slow rise in its price, which Blau (2001) attributes to a large "unexplained" increase in the supply of labor to the childcare market. We suggest that the large numbers of low-skilled immigrants who arrived in the United States after 1965 are likely to have contributed to this phenomenon.[4] Cortes (2008) showed that low-skilled immigration led to reductions in a pooled index of non-traded goods and services in major U.S. cities. We provide a more detailed examination of three service markets that potentially provide strong substitutes for time-intensive childrearing tasks undertaken by parents: child care, food preparation, and housekeeping.[5]

As can be seen in Table 1, by the end of the 20th century immigrants were over-represented in these three household services occupations, relative to other occupations, whereas natives were underrepresented. Immigrants were also more likely to be "low-skilled," defined as

[4] Our calculations using data from the U.S. Census indicate that in 1970 roughly one quarter of both working-age immigrants and natives had advanced beyond high school. By 2006 over 60% of working-age natives had completed some post-secondary education, while the majority of working-age immigrants had a high school degree or less.

[5] Occupation definitions are based on the consistent classification (1990 basis) system available from the IPUMS project (Ruggles et al. 2010). What we call "housekeeping services" is generally referred to as "private household services," an occupation that includes maids, butlers, and lodging quarters cleaners. We use the alternative name to avoid confusion with "household services," which we will use as a catch-all term for the three sectors of interest.

never having attended college. Low-skilled immigrants constituted 9.3% of child care workers in 2000, as compared to 6.2% of the workforce in the non-household service occupations. This pattern was even more striking in food preparation and housekeeping, in which 14.1% and 28.4% of workers were low-skilled immigrants, respectively. Hence it seems likely that inflows of low-skilled immigration increased the labor supplied to all three markets, which should have reduced the final cost and increased the availability of market-based household services that substitute for a potential mother's time in home production.

2.3 Empirical Research on Fertility, Work, and Childcare Costs

A number of studies have investigated the relationships between childbearing, labor supply and the cost of child care, as measured by both price and availability. Some studies take work decisions as given and look at the relationship between childcare costs and fertility. For example, Mörk, Sjögren, and Svaleryd (2009) examined Swedish childcare subsidy reform, finding that lower childcare costs led to higher fertility. Another line of research investigates the relationship between childcare costs and the likelihood that mothers work (Blau and Robins 1988; Connelly 1992). Several recent studies have also considered the availability of pre-school or kindergarten, which might be thought of as inexpensive child supervision, finding a positive impact on maternal labor supply (Baker, Gruber and Mulligan 2008; Cascio 2008). Nonetheless, as described above, changes in conditional likelihoods may be driven by differential selection into motherhood.

Only a handful of papers have considered the effects of childcare costs on both employment and fertility outcomes. Mason and Kuhlthau (1992) examined mothers' perceptions as to whether the availability of child care constrained their employment and fertility decisions. Blau and Robins (1989) analyzed transitions among employment and fertility states as related to geographic variation in weekly childcare expenditures. Taking a different approach, Stolzenberg and Waite (1984) examine how variation in the individual-level association between fertility and LFP is explained by conditions in the local childcare market. All of these studies provide results suggesting that lower childcare costs reduce role incompatibility, but rely on potentially endogenous cost measures.

In our analysis we use an instrumental variables approach to isolate exogenous variation in the relative size of the local immigrant population, which we show to be causally associated with both lower childrearing costs and greater availability of services. Based on an identification strategy similar to ours, Cortes and Tessada (2009) provide evidence that low-skilled immigration to the United States led to an increase in the hours worked among highly-skilled females, conditional on participating in the labor force. Farré, Gonzalez, and Ortega (2009) reach similar conclusions for native females with high earnings potential in Spain. However, both Khananusapkul (2004) and Cortes and Tessada (2009) find the opposite effect when considering the likelihood of labor force participation. Adopting a simultaneous decision-making framework, we consider whether the negative relationship between immigration and female labor force participation may be explained by women exiting the workforce to bear children. We are also able to explicitly examine the degree to which low-skilled immigration reduced role incompatibility, as measured by the joint likelihood of childbearing and labor force participation. Broadly, our analysis proceeds in two steps. First, we consider the extent to which low-skilled immigration to the United States has, as a result of expansions in labor supply, reduced the cost of market-provided services that are close substitutes for time-intensive childrearing activities undertaken by parents. Second, we determine whether and how immigration has altered childbearing and labor force participation outcomes of American females.

3 Data

Throughout, our empirical estimates rely on geographic and temporal differences among large metropolitan areas as a source of variation in the concentration of low-skilled immigrants. Our main sample was drawn from the U.S. Census Bureau's 1980, 1990, and 2000 public-use microdata sample (PUMS) files, while the 1970 census provided additional data used to construct the instrumental variable. All data were obtained from the Integrated Public Use Microdata Series (IPUMS, Ruggles et al. 2010).

3.1 Nativity and Skill Groups

Because our goal is to estimate the impact of immigration on work and fertility attributable to changes in markets for household services, our analysis concentrates on low-skilled immigrants and high-skilled non-Hispanic native females. Sharply differentiating immigrants and natives by skill minimizes the possibility of competition for jobs, which might directly affect female employment prospects. Analyzing non-Hispanic native females avoids non-market channels of influence, such as social norms and peer effects, which might arise from inflows of low-skilled immigrants to the United States, the bulk of whom are from Latin America\. We define skill classes based on education. "Low skilled" implies having, at most, a high school degree and never having attended college, while "high skilled" implies having completed a bachelor's degree. Throughout the paper, we classify individuals born in Puerto Rico, the U.S Virgin Islands, and other outlying possessions as natives since they are citizens by birth.[6]

[6] Re-classifying these individuals as immigrants leads to estimated effects of low-skilled immigration that are very similar, although slightly less conservative, than the results presented below.

3.2 The Metropolitan Sample

The underlying geographic sampling units defined by the Census Bureau have changed over time. The resulting inconsistencies in the degree to which the population of a metropolitan area (MA) is covered in the microdata files makes it difficult to construct metro-level variables that are comparable across years. Shifting boundaries introduces noise into the aggregate measures. Of greater concern is that the boundaries of a substantial number of MAs were severely truncated in the 2000 Census microdata files. This may result in systematic bias given that individuals in far suburbs or exurban areas are likely to exhibit different behavior than residents of the core metropolitan area. To reduce the potential influence of these inconsistencies on our estimates, we rely on geographic definitions of MAs that provide the most comparable population coverage between 1980 and 2000.

To construct a temporally consistent sample, we established a "benchmark" geography based on 1990 Census definitions.[7] Relying on GIS map files provided by the IPUMS project, we then defined MAs in each year based on the set of sampling units that most consistently covered the population living in the benchmark territory. If two MAs were agglomerated by the Census into a single metropolitan area in any year between 1970 and 2000, we agglomerated them in all years. We tolerated a deviation from benchmark population coverage of up to 5 percent in any year. If all configurations of sampling units for a given year deviated by more than five percent from the benchmark population, we attempted to expand the definition of the MA incrementally until we were able to find a configuration such that the year-to-year deviations from the expanded benchmark population were tolerable. Since this process of expansion could potentially occur indefinitely (rendering the concept of a metropolitan area meaningless) we

[7] We relied on Census definitions from 1990 because that year represents the mid-point of our sample frame. The 1990 Census also contained a "metro" file with relatively more accurate boundaries for many MAs than the "state" file. No comparable metro file was available in 2000.

stopped expanding the MA definition once the resulting population was more than 10 percent larger than the original benchmark population. Additional details on this procedure, as well as Stata programs, will be available in an online appendix.

Starting with an initial universe of the 100 most populous MAs (as measured in 1990), after recoding for consistency, our final sample includes 77 metropolitan areas. Although they covered 55% of the working-age U.S. population in 1990, these MAs covered just over 80% of the working-age immigrant population, reflecting the fact that immigration is largely an urban phenomenon.[8] A similar proportion of low-skilled immigrants lived in the included MAs. By contrast, the 23 MAs excluded from the analysis only covered 7.5% of the working-age U.S. population and approximately 3% of the low-skilled immigrant population. Hence, estimates based on our sample of large metropolitan areas are likely to capture a substantial majority of the impact of low-skilled immigration.

4 Immigration and Childrearing Costs

The first step of our analysis considers the effects of immigration on the cost of market-provided child care, food preparation, and housekeeping. Due to its central role in the existing literature on childrearing costs, we focus our discussion largely on the market for child care. However, the methodological issues and estimation strategy that we outline are equally applicable to the markets for food preparation and housekeeping, which are also covered by our empirical analysis. The wage bill accounts for between 60% and 70% of the operating expenses at formal and home-based childcare centers (Blau and Mocan 2002; Helburn and Howes 1996), and likely

[8] We use the term "working age" to refer to individuals between the ages of 20 and 65.

represents an even higher share of the final costs of informal childcare providers.[9] Thus, we use wages of workers as a measure of the price of the associated services. A potential concern for our later analysis is that low-skilled immigration might not affect the cost of the childcare services actually purchased by college-educated women. In particular, educated and high-income women may demand a better quality of care (Blau and Hagy 1998; Hotz and Kilburn 1991).

To our knowledge, data linking the characteristics of childcare workers to characteristics of the final consumer do not exist. However, Blau and Mocan (2002) provide evidence that the cost of child care is a positive function of the underlying objectively-assessed quality. Thus, we draw inference on immigration's impact on the cost of services of various levels of quality by examining the effects of immigration on various quantiles of the wage distribution.[10] A potential pitfall of this strategy is that an influx of low-skilled immigrants might mechanically reduce an upper-percentile wage because their arrival results in more mass at the bottom of the wage distribution. Hence, the effect we estimate might not be relevant to high-skilled natives if they tend to hire established, higher-quality caregivers, whose equilibrium wages are only minimally affected by the newcomers. To remove the possibility that we are simply identifying a mechanical effect, we calculate the wage quantiles based on the sample of native and immigrant workers living in the United States for at least 10 years.

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

$$w_{qmt} = \beta_q LSI_{mt} + \boldsymbol{\mu}_{qm} + \boldsymbol{\mu}_{qkt} + \lambda_q IncControl_{mt} + \varepsilon_{qmt}. \qquad (1)$$

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[9] Although there are no formal estimates available, it is likely that payments to labor represent a similar proportion of the price of private household services. Wages are a smaller, but not inconsequential, component of production costs for restaurants (Lee, Schluter, and O'Roark 2000).

[10] A reanalysis of the data in Table 1 (available on request) indicates that low-skilled immigrants are well-represented at all quartiles of the wage distribution, decreasing slightly to 7.5% of the workforce at the top quartile, as compared to 9.5% in the bottom quartile of the wage distribution. Hence, it seems likely that low-skilled immigration would increase labor supply to all segments of the childcare market.

The dependent variable w_{mt}^{q} is the log of the hourly wage of native and non-recent immigrant household service workers in metropolitan area *m* in year *t,* as measured at the *q*th percentile. We specifically examine the 25th, 50th, and 75th percentiles in our analysis.[11] The low-skilled immigrant share of the overall working-age population immigration is denoted LSI_{mt}. MA-specific intercepts for each quantile are indicated by μ_{qm}, while μ_{qkt} represents time fixed effects specific to the *k*th Census region. The variable $IncControl_{mt}$ denotes the log of income per capita among working-age male college graduates.[12] Regressions are estimated using the estimated size of the workforce in each occupation as weights. We estimate cluster-robust standard errors, clustering by MA to allow for arbitrary patterns of correlation of the error term between percentiles and over time.

4.1 Identification

If immigration represents a supply shift, an increase in *LSI* should result in lower wages and so we expect the estimated βs to be negative. In our empirical analysis, we consider the degree to which immigrant-induced changes in wages were accompanied by expansions in employment in child care, food preparation, and housekeeping services. Our evidence suggests a substantial increase in labor supply to these occupations. However, immigrants may be drawn to areas with a higher willingness to pay for childcare, which might arise in a booming local economy or due to other time-varying unobserved factors. As a result, we would expect that ordinary least

[11] Note that we are applying least squares to analyze empirical quantiles obtained from grouped data. Knight (2001) points out that this approach will result in consistent estimates of the individual-level quantile regression coefficients when the explanatory variables only vary at the group level, as is the case in our model.

[12] College graduates are likely to be high demanders of household services and, for the most part, will have incomes that are not directly tied to wages in low-skill services markets. Females are not included in the income measure since their labor supply and earnings are potentially endogenous with respect to changes in the cost of household services. To account for top-coding, which was only an issue in 1980, we impute values for individuals whose income had been top-coded using a region-specific Pareto extrapolation

squares (OLS) estimates will yield attenuated (less negative) estimates of the causal effect of immigration. To address this concern, we rely on an instrumental variables (IV) approach to identify the causal impact of low-skilled immigration on wages in the household services occupations. Instrumental variables should also address attenuation bias in the estimated βs due to measurement error in the percentage foreign born, arising, for example, from an undercount of undocumented immigrants.

Our instrument is based on the propensity of new immigrants to locate in areas with a relatively large existing concentration of co-ethnics (e.g. Bartel 1989). Following a similar line of reasoning as Card (2001), Cortes (2008), and Cortes and Tessada (2009), our instrument uses historical enclaves to predict the flow of subsequent migrants across MAs. More specifically, the instrument for LSI_{mt} is

$$INST_{mt} = \sum_{b} \frac{N^{b}_{m,1970}}{N^{b}_{1970}} \times \left[NLS^{b}_{t} - NLS^{b}_{1970} \right]. \tag{2}$$

For each country of birth, b, the first term in equation (2) represents the fraction of all immigrants from country b living in MA m in 1970. The second term represents the net change in the number of low-skilled working age adults from country b between 1970 and time t.

To maximize the predictive power of our instrument, we focus on immigrant groups in which (a) there were at least 20,000 members present in 1970, (b) there was a net increase in the number of low-skilled working-age adults in the U.S. between 1970 and 2000, and (c) the majority of the overall increase in working age-adults consisted of low-skilled individuals.[13] Five immigrant groups meet these criteria: Dominicans, Ecuadorians, Haitians, Mexicans, and Portuguese. As indicated in Table 2, the share of working-age immigrants from these countries

[13] The third condition additionally limits the degree of potential competition for employment between members of the immigrant group and the college-educated women analyzed below.

that was low-skilled in 2000 ranged from 59 to 86%. By contrast, low-skilled individuals comprised 42% of the native population in 2000. Moreover, our calculations indicate that immigrants from countries listed in Table 2 were twice as likely as natives to work in the household service occupations we analyze.

The necessary criteria for the exclusion restriction to be met are very similar to those outlined by Cortes (2008). Given the MA and region-by-decade fixed effects included in the model, these criteria are as follows: (a) the initial distribution of immigrants must be uncorrelated with differential changes in relative economic conditions affecting the demand for household services across MAs within a region 10 to 30 years later, and (b) differential economic changes among MAs within a region should not affect the overall inflow of low-skilled immigrants to the United States. Although we cannot test them directly, we conduct a series of robustness checks that support the assumptions underlying the exclusion restrictions.

4.2 Results

Focusing for the moment on the median wage among native and non-recent immigrant childcare workers, Table 3 compares the impact of low-skilled immigration using different specifications applied to our panel of MA-level data. The first three columns present estimates based on OLS. Without MA fixed effects, there is a positive and statistically significant relationship between *LSI* and the median wage in child care, but this relationship is negative when MA fixed effects are included. This suggests that low-skilled immigrants are drawn to metropolitan areas where wages in the childcare market are persistently high. Controlling for income per capita among college graduates yields a more negative estimate, which lends credence to the hypothesis of selective migration among immigrants based on changing local economic conditions within MAs over time.

Shifting to the instrumental variables (IV) strategy described above, the first-stage F statistic of 32 substantially exceeds the weak-instruments critical value given in Stock and Yogo (2002). The second-stage IV estimate of the effect of *LSI* reported in the fourth column of Table 3 is roughly 2.6 times as large as the OLS estimate reported in the third column. This may be taken as confirmatory evidence that immigrant location decisions substantially bias the coefficient estimates obtained from OLS upward, although we cannot rule out attenuation bias arising from noisy measurement of the low-skilled immigrant population.[14] For the remainder of our analysis, we focus exclusively on instrumental variables estimates.

The estimated coefficient of −3.2 represents the percentage change in the median wage of childcare workers caused by a one percentage point increase in the low-skilled immigrant share of the working-age population. In order to make this number more meaningful, we scale it by the percentage point change in *LSI* experienced by the average native working-age individual during the sample period. Between 1980 and 2000, the low-skill immigrant share of the working-age population rose from 6.1% to 10.1% in the representative working-age native's MA. Based on the IV point estimate, this 4 percentage point increase would result in a 12.9% reduction in the median wage.

As previously discussed, changes in the middle of the distribution might not be relevant to the cost of care facing college-educated women. In the first three columns of Panel A of Table 4, we analyze the potentially disparate impacts of low-skilled immigration on wages at the 25th, 50th, and 75th percentiles. The relative effect is greater at the lower percentiles where one might

[14] As a specification check, we have re-estimated the models in Table 3 using the wages of college graduate workers as the dependent variable, dropping the income control from the explanatory variables for obvious reasons. We find that the OLS fixed-effects relationship between *LSI* and college graduate wages is positive and significant ($p < 0.01$), but that the IV estimate of the coefficient on LSI is small, negative, and insignificant ($p = 0.72$). These results suggest that the predicted immigrant flows, as allocated by the enclave-based instrument, are uncorrelated with local economic conditions.

expect low-skilled immigrants to have the strongest impact. However, our estimates imply that the increase in *LSI* between 1980 and 2000 experienced by the representative MA would cause an 8.1% reduction in the wages of childcare workers even at the 75th percentile. Hence, low-skilled immigration should have led to a decline in the cost of high- as well as low-quality care.

The relative effects of low-skilled immigration on the wages of food preparation workers and housekeepers are slightly weaker at most quantiles, as seen in Panels B and C of Table 4. 0Even so, the estimates in all three household services markets are considerably larger than most existing estimates of the wage effects low-skilled immigration (Friedberg and Hunt 1995; Card 2001). However, much of this research is based on examining broad skill classes, rather than specific occupations. The particular household service markets we examine are relatively labor intensive, as compared with the larger low-skilled labor market, providing little room for capital adjustments.

The fourth column of Table 4 indicates that low-skilled immigration resulted in expansions in the share of the local workforce concentrated in the household service occupation.[15] We may combine these numbers with the estimates of the effect of *LSI* on wages to compute a quasi-elasticity of demand. Specifically, we divide the percentage change in the labor force concentration attributable to a one percentage point increase in *LSI* by the associated percentage change in wages to obtain a quasi-elasticity ranging in absolute value from 0.58 to 1.38 in child care. The comparable ranges for food preparation and housekeeping are 1.38-2.06 and 1.56-2.24, respectively. These numbers imply that low-skilled immigration between 1980 and 2000 had a larger impact on the final price of child care than on its availability, while the opposite was true

[15] Workforce shares are essentially aggregated measures of underlying binary outcomes. For consistency with our later specification of female outcomes we use a normit transformation to construct the dependent variable, and the effect of the average MA's change in low-skilled in immigration is calculated as described below. Log or logit transformations result in similar effects.

for food and housekeeping services. Nonetheless, the reductions in wages and increases in availability caused by low-skilled immigration are sizeable in all of the household services markets. Both effects should have resulted in substantial reductions in the effective cost of childrearing.

5 Fertility and Labor Force Participation

The second step of the analysis examines the effect of low-skilled immigration on female childbearing and labor force participation decisions. We use a structured bivariate statistical model to gain insight into a variety of the mechanisms whereby low-skilled immigration may exert its effects. Specifically, the model allows us to separately estimate the impact of low-skilled immigration on: (i) the likelihood of bearing children, (ii) the likelihood of LFP, and (iii) the correlation between the two. Given that the influence of immigrants is largely being channeled through low-skilled household services markets, we are essentially examining the consequences of lower childrearing costs, which have a theoretically ambiguous effect on the marginal likelihoods, (i) and (ii). The bivariate model additionally provides a behavioral interpretation of the observed correlation between fertility and labor force participation, (iii). Finally, we can combine the three sets of estimates to predict the effect of low-skilled immigration on the joint likelihood, decomposing the total effect into changes in the marginal likelihoods and changes in the correlation between fertility and work.

5.1 The Bivariate Model

Female employment and fertility decisions can be described using the simultaneous latent variables framework:

$$C^*_{igmt} = \beta_1 LSI_{mt} + \boldsymbol{\omega}_1' \mathbf{v}_{igmt} + \varepsilon^C_{igmt} \tag{3}$$

$$L_{igmt}^{*} = \beta_2 LSI_{mt} + \boldsymbol{\omega}_2' \mathbf{v}_{igmt} + \varepsilon_{igmt}^{L} \quad (4)$$

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*.[16] 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. 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 the work and fertility outcomes. The vector of controls, $\mathbf{v}_{igmt}$, described in detail below, expands on the fixed effects and controls from the analysis of wages above by additionally incorporating demographic information on the college-graduate women in the sample.

As shown above, low-skilled immigration affects both the price and the availability of multiple household services. Moreover, the impacts of immigration may differ by the quality of services being provided. Order conditions do not permit us to identify all of the possible influences of low-skilled instrumental variables in an IV set-up, and it is not clear how to reduce the multiple dimensions of cost into a single index. As a result, we include *LSI* directly in the model as a proxy for the various impacts of low-skilled immigration on childrearing costs via market-based household services. Based on this interpretation of *LSI*, theory does not clearly predict whether β_1 and β_2 will be positive or negative. Our estimates of these parameters can, therefore, provide an empirical answer to the question of how the fertility and labor supply of high-skilled U.S.-born women respond to changes in childrearing costs.

[16] As described below, we divide the sample into groups because the tetrachoric correlation 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 attributes.

In addition to examining the propensities to work and to bear children, our model allows us to analyze the correlation between the two. If the error terms in equations (3) and (4) follow a bivariate normal distribution, $\rho_{gmt} = corr(\varepsilon_{igmt}^{C}, \varepsilon_{igmt}^{L})$ is, by definition, the *tetrachoric* correlation. Based on the conditional expectation $E\left[L_{igmt}^{*} \mid \varepsilon_{igmt}^{C}\right] = E\left[L_{igmt}^{*}\right] + \rho_{gmt}\varepsilon_{igmt}^{C}$ the tetrachoric correlation can be understood as the degree to which changes in childbearing not explained by common group-level covariates translate into changes in LFP; a parallel relationship can be written for the fertility response to within-group labor supply shocks. Hence the tetrachoric correlation would, for example, determine the extent of 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. Consequently, we expect that $\rho < 0$, which is almost universally the case in our sample. There is substantial variation in the tetrachoric correlation between fertility and labor force participation over time and across space.[17] We explore how the correlation is affected by low-skilled immigration based on the parameterization

$$\rho_{gmt} = \beta_3 LSI_{mt} + \boldsymbol{\omega}_3' \mathbf{v}_{gmt} + e_{gmt}, \tag{5}$$

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

[17] Furtado and Hock (2010) demonstrate a gradual weakening of the tetrachoric correlation in the United States from 1970 to 2000. There are also substantial cross-sectional differences in the correlation across MAs in our sample.

5.2 Grouped Estimation with Instrumental Variables

Endogenous regressors in univariate binary choice models may be addressed using control functions and related strategies (e.g., Blundell and Powell 2004). Our interest in explicitly parameterizing ρ makes this approach difficult to extend to the simultaneous choice setting. Consequently we rely on a slight generalization of Amemiya's (1974) grouped bivariate probit specification, which allows a straightforward application of instrumental variables. Grouping is also necessary to calculate the tetrachoric correlation, which is not defined at the individual level. The model coefficients can be recovered 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:

$$\pi_{gmt}^{C} = \Phi\left(\beta_1 LSI_{mt} + \boldsymbol{\omega}_1' \mathbf{v}_{gmt}\right) \text{ and } \pi_{gmt}^{L} = \Phi\left(\beta_2 LSI_{mt} + \boldsymbol{\omega}_2' \mathbf{v}_{gmt}\right) \tag{6}$$

Let p_{gmt}^{C}, p_{gmt}^{L}, and p_{gmt}^{CL} denote the observed proportions of the women in group g in metro 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:

$$c_{gmt} = \beta_1 LSI_{mt} + \boldsymbol{\omega}_1' \mathbf{v}_{gmt} + u_{gmt}^{1}, \tag{7}$$

$$\ell_{gmt} = \beta_2 LSI_{mt} + \boldsymbol{\omega}_1' \mathbf{v}_{gmt} + u_{gmt}^{2}, \tag{8}$$

where $c_{gmt} = \Phi^{-1}(p_{gmt}^{C})$ and $l_{gmt} = \Phi^{-1}(p_{gmt}^{L})$ denote the inverse standard normal cumulative distribution (normit) function applied to the observed rates of childbearing and LFP. Moreover, based on equation (5), the expression for the empirical analogue of the population tetrachoric correlation obtained from the data can be expressed as

$$r_{gmt} = \beta_3 LSI_{mt} + \boldsymbol{\omega}_3' \mathbf{v}_{gmt} + u_{gmt}^3. \tag{9}$$

Equations (7)-(9) correspond to Amemiya's (1974) equations (4.11)-(4.13), with the expression in (9) additionally relying on the parameterization of the tetrachoric described above.

The empirical tetrachoric correlation is calculated based on the population relationship:

$$\pi_{gmt}^{CL} = F(\Phi^{-1}(\pi_{gmt}^{C}), \Phi^{-1}(\pi_{gmt}^{L}), \rho_{gmt}) \equiv G(\pi_{gmt}^{C}, \pi_{gmt}^{L}, \rho_{gmt}) \tag{10}$$

where π_{gmt}^{CL} 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 (10) allows us 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

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

where ξ represents a pre-defined level of precision, which we set to 2^{-50}. Note that monotonicity of $F(\cdot)$ also implies that a higher value of ρ will, *ceteris paribus*, translate into a higher joint likelihood. Below, we show how we can combine the changes in the tetrachoric implied by low-skilled immigration with changes in the marginal likelihoods to derive marginal effects for the joint likelihood.

If there are any groups in which any of the binary outcomes is uniform across its members, the data become uninformative and it is not possible to estimate the empirical tetrachoric correlation. Consequently, we divide the sample of college-graduate women into two broad age groups, women ages 22-30 and women ages 31-39, and include measures of the average characteristics of the group as explanatory variables. Even with these broadly specified categories, there were a few MA-year-age group cells with all ones or all zeros. In these cases, we replace zeros with 0.001 and ones with 0.999 when calculating the normits and estimating ρ.

We estimate the bivariate probit model described by equations (7) through (9) based on a stacked system of three equations of the form

$$y_{gmt} = \beta LSI_{mt} + \boldsymbol{\mu}_m + \boldsymbol{\mu}_{kt} + \boldsymbol{\mu}_g + \lambda_{mt} IncControl_{mt} + \boldsymbol{\theta}'\mathbf{x}_{gmt} + u_{gmt}, \qquad (11)$$

where y is one of the three dependent variables (c, ℓ, r). The variable *LSI*, the MA and region-year fixed effects ($\boldsymbol{\mu}_m$ and $\boldsymbol{\mu}_{kt}$), and $IncControl_{mt}$ are defined as in equation (1). Added to these variables are age-group fixed effects ($\boldsymbol{\mu}_g$) and a vector of demographic controls ($\mathbf{x}_{gmt}$). For each age group in a given metro area and year, $\mathbf{x}_{gmt}$ includes the share of women who are married and the proportions of the group that self-identify as being black and that self-identify as being a member of another non-white race—we use the IPUMS single-race coding system that bridges the 1990 and 2000 Census classification schemes. Because of concerns about endogeneity similar to those described above, we estimate equation (11) using two-stage least squares, again making use of the enclave-based instrument. Our analysis is limited to non-Hispanic natives with college degrees, aged 22-39, and not living in institutional group quarters. In our main analysis we additionally restrict the sample to women not currently enrolled in school since students' work and fertility decisions are expected to be substantially less sensitive to childrearing costs. We weight each group-MA-year cell by the estimated population of women represented by the cell. Robust standard errors are clustered by MA to allow for an arbitrary pattern of correlation in the error terms across equations, groups, and time.

5.3 Results

Coefficient estimates based on the grouped bivariate probit model are presented in Table 5. The IV coefficients indicate that *LSI* led to significantly higher fertility rates and lower labor force participation rates. The low-skilled immigrant share of the labor force in the average high-skilled woman's MA rose from 6.2% in 1980 to 10.1% in 2000. Using the IV point estimates to compute the average partial effects (APEs), a 3.9 percentage point increase in LSI implies a likelihood of childbearing that is 0.85 percentage points higher.[18] This corresponds to almost one tenth of the observed fertility rate in 2000. The estimated effect of the average increase in LSI on the likelihood of labor force participation is −0.72 percentage points.

Taken together, the changes in the marginal likelihoods results suggest that high-skilled women in our sample of MAs responded 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 faced tradeoffs between work and fertility. However, the reductions in labor force participation rates associated with low-skilled immigration were slightly smaller than the associated increases in childbearing rates, which implies that the tradeoff between work and fertility was not one-for-one.

As seen in the third column of Table 5, low-skilled immigration also attenuated the negative correlation between childbearing and labor force participation. In order to make the implications for role incompatibility concrete, we calculate the effect of *LSI* on the joint likelihood of fertility and LFP based on the point estimates in Table 5. Specifically, equations (5) and (6) imply that the expected joint likelihood can be written as

[18] For the marginal likelihoods of childbearing and labor force participation, we compute the APE as the weighted average of the partial derivatives of (6) across the sample, with the share of high-skilled women represented by each age-MA-year cell is 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 effects reported in the text.

$$\pi_{gmt}^{CL} = F\left(\beta_1 LSI_{mt} + \boldsymbol{\omega}_1'\mathbf{v}_{gmt}, \beta_2 LSI_{mt} + \boldsymbol{\omega}_2'\mathbf{v}_{gmt}, \beta_3 LSI_{mt} + \boldsymbol{\omega}_3'\mathbf{v}_{gmt}\right). \tag{12}$$

The estimated average partial effect is $\hat{A}_{LSI}^{CL} = \sum_{gmt} h_{gmt}(d\hat{F}_{gmt} / dLSI)$, with h_{gmt} denoting the proportion of high-skilled skilled women represented by each age-MA-year cell and $\hat{F}_{gmt}$ denoting the standard bivariate normal distribution function evaluated using the estimated coefficients in place of the true parameters. Based on equation (12), the APE can be decomposed as

$$\hat{A}_{LSI}^{CL} = \left[\hat{D}_1\hat{\beta}_1 + \hat{D}_2\hat{\beta}_2\right] + \hat{D}_3\hat{\beta}_3 \tag{13}$$

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 two terms inside the brackets in equation (13) represent the average change in the joint likelihood arising from the differential impacts of *LSI* 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 *LSI* on the joint likelihood arising from a weakened link between childbearing and LFP.

As above, to translate the APE (and its components) into more meaningful terms, we scale by the percentage point increase in *LSI* experienced by the representative high-skilled woman. Based on the point estimates from Table 5, a 3.9 percentage point increase in LSI would result in a 0.65 percentage point increase in the likelihood of bearing children while remaining in the labor force. Moreover, almost 30% of the total effect of LSI 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.

Between 1980 and 2000, the joint likelihood in our sample of urban, non-Hispanic college graduate women rose from 3.3% to 5.7%, an increase similar to that seen in Figure 2.

The total effect of low-skilled immigration on the joint likelihood of fertility and work represents just over one fourth of the observed increase in the sample between 1980 and 1990. The estimate suggests that the joint likelihood would have been up to 11% lower in 2000 if the relative size of the low-skilled immigrant population were to revert to its 1980 level. These "counterfactual" numbers likely represent upper bounds on the long-term impact of low-skilled immigration, since 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, our instrumental variables estimates indicate that inflows of low-skilled immigrants to an MA during our sample period led to significant and substantial short-run increases in the joint likelihood of childbearing and labor force participation.

5.4 Specification Checks

5.4.1 Direct Estimation, Alternate Grouping, and Outliers

Our estimate of the impact of immigration on the joint likelihood is indirect. It relies on the implicit assumption that the women who change their fertility in response to decreases in the cost of household services are the same women who change their labor supply. Moreover, it relies heavily on the statistical structure of the model. In order to verify that the indirect estimate is not being affected by these assumptions, we directly estimate the effect of low-skilled immigration on the joint likelihood based on

$$d_{gmt} = \beta_4 LSI_{mt} + \boldsymbol{\omega}_4' \mathbf{v}_{gmt} + u_{gmt}^{d} \qquad (14)$$

where $d_{gmt} = \Phi^{-1}(p_{gmt}^{CL})$. Although the direct estimate does not rely on the structural assumptions used to calculate the indirect estimate, we compute the former using grouped data for comparability. Table 6 displays the estimated effect of low-skilled immigration on the likelihood of bearing children and participating in the labor force based on a grouped univariate probit model. The marginal impact of immigration implied by the coefficient is very similar to what was obtained using the indirect method, differing by less than one percent.

We have also estimated the grouped model using a variety of thresholds for the age categories, finding the results to be relatively insensitive to the cut-points. Increasing the number of age categories to three also yields similar estimates—we include results for the joint likelihood in the second column of Table 6. Taken together, these results suggest that the structured model is not skewing our estimates of the effects of low-skilled immigration.

To determine the influence of outliers, we re-estimated the model with subsets of the metropolitan areas considered in the main sample. We focus on the joint likelihood of childbearing and LFP, since this is our main behavioral measure of role incompatibility. We first drop the three MAs that experienced the largest influxes of low-skilled immigration between 1980 and 2000—Dallas, Los Angeles, and New York City—and then exclude California. As seen in the third and fourth columns of Table 6, the estimated effect of *LSI* on the joint rate of fertility and LFP rises when removing these potential outliers from the sample, although there is some fluctuation in the proportion of the overall effect explained by the latent correlation.

5.4.2 Validity of the Instrument

The validity of our IV results rests on the assumption that the distribution of immigrants across metropolitan areas in 1970 was not affected by factors that might, in the absence of subsequent immigration, affect later changes in the outcomes of high skilled U.S.-born women. Two potential violations of this assumption come to mind. First, immigrants may have been historically more highly represented in metropolitan areas with more persistent traditional family values. Second, differences in the historical concentration of low-skilled and immigrant workers may have led to cross-MA differences in the degree of the “polarization” of skills demand described by Autor, Katz, and Kearney (2008). In both cases our estimation strategy would falsely attribute changes in fertility and work outcomes to the rising share of low-skilled immigrants in the labor market predicted by the 1970 distribution of immigrants across MAs.

In the spirit of Cortes and Tessada (2009), our first approach to addressing these concerns relies on controlling for a series of MA-level characteristics in 1970. As proxies for cultural progressiveness, we use the share of working age women that were married, the share that had complete a BA, and the share who participated in the labor market. We additionally control for the proportion of the 1970 U.S.-born labor force that was white. To address the potential for differential skill polarization, we include the 1970 share of the labor force employed in high-skilled service industries (finance, professional services, and publishing) and the share employed in low-skilled service industries (entertainment services, personal services, and retail trade). We consider service industries specifically because of the particularly large growth in this sector in the last few decades. All of these variables are interacted with decade dummies to flexibly account for potential change over time.[19] As is evident from the final column of Table 6, the estimated effects of low-skilled immigration on the joint likelihood of childbearing and LFP are virtually unchanged when the 1970 controls are added. However, including the extra control variables to the regression results in more of the effect of *LSI* being channeled through changes the latent correlation, rather than differential changes in the likelihoods of fertility and work.

We have also examined the relationship between marriage patterns and low-skilled immigration as an additional check to determine if our IV estimates might be picking up cross-MA differences in the evolution of cultural norms, rather than changes in the cost of household

[19] Our list of 1970 variables is almost identical to the preferred specification in Cortes and Tessada (2009). We did not include the average 1970 wage of college graduates because doing so caused the first-stage F statistic to fall below conventionally accepted levels. However, unlike Cortes and Tessada, we control for the current income of male college graduates directly in all specifications.

services. In addition to the marriage rate, we considered the effect of *LSI* on the (negative) correlation between marriage and LFP, which might measure the prevalence of traditional homemaker norms.[20] The IV estimate for both outcomes was insignificant ($p > 0.4$) and trivial.

5.4.3 Sample Selection

The main sample of non-Hispanic native females is restricted by educational attainment and school enrollment. The sample is also implicitly restricted by geography in that we analyze the outcomes of women living in large metropolitan areas. Although it is likely that many women will not alter their educational and residential choices in response to immigrant-induced reductions in the cost of household services, some women may certainly do so. We will refer to the former group as being "unselected" with respect low-skilled immigration and the latter group as being "selected." We focus on fertility to fix ideas, but our discussion of potential biases from selection are also applicable to the other outcomes we consider.

Define *S(LSI)* as the number of women who entered the sample in response to low-skilled immigration, *U* as the number women who would have been in the sample regardless of the extent of immigration, and $N(LSI) = S(LSI) + U$. Then, the APE estimated by our model is

$$\hat{A}^C = \left[s \times \hat{A}_S^C + (1-s) \times \hat{A}_U^C \right] + \left[\frac{ds}{dLSI} \times (\pi_S^C - \pi_U^C) \right], \tag{15}$$

where $s(LSI) \equiv S(LSI) / N(LSI)$. The first bracketed term in equation (15) measures the (weighted) average of the marginal effects of low-skilled immigration on outcomes across sub-groups of women. From a causal perspective it is not problematic if our estimated effects of LSI are reflecting *changes* in fertility among either selected or unselected women. However, if the baseline fertility rate among selected women differs from that among unselected women, then

[20] We continue to use the tetrachoric correlation due to the limitations of other measures of binary association outlined in Furtado and Hock (2010).

our estimated effects of LSI may simply be capturing changes in the composition of the sample. This is reflected in the second bracketed term of equation (15).

It is straightforward to show that $ds / dLSI = \gamma \times (1-s)$, where $\gamma \equiv d \ln N(LSI) / dLSI$. When $\gamma > 0$, it is likely to be the women with the strongest fertility preferences that enter the sample in response to immigration. Since we expect such women to exhibit a higher rate of childbearing than the unselected women $(\pi_S^C > \pi_U^C)$, the estimated effect of low-skilled immigration may be biased upward, relative to the average causal effect across both types. Because of limitations associated with the migration data available in the Census PUMS samples, we are not able to distinguish between composition changes that arise due to selective migration and changes that occur due to selective college completion within an MA.[21] However, we are able to test for both forms of selection jointly based on:

$$\ln N_{gmt} = \gamma LSI_{mt} + \boldsymbol{\mu}_m + \boldsymbol{\mu}_{kt} + \boldsymbol{\mu}_g + \lambda IncControl_{mt} + \chi \ln T_{gmt} + \varepsilon_{gmt}, \qquad (16)$$

where N_{gmt} denotes the number of non-Hispanic females 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-MA-year cell. Equation (16) essentially tests for extranormal growth in the population of high-skilled females, relative to the overall population of same-age females, and the coefficient γ can be interpreted based on the discussion above.

Instrumental variables regression analysis indicates that the p value associated with the IV estimate of γ is 0.44, which suggests that composition bias due to selective migration and

[21] The Census collects information on residence five years prior to the date of the survey. However, in 1980 responses were coded for only half of the cases in the PUMS data. Moreover, in the 1990 and 2000 data, the geographic areas of previous residence do not always match with the geographic sampling units used to create consistent-boundary MAs, making it impossible to determine the migration status of many women.

college completion is not likely to be a problem for our analysis.[22] Substituting the number of non-Hispanic native females who had not completed college in place of N_{gmt} in equation (16), however, yields a negative and highly significant (p < 0.01) estimate of γ. In the absence of selective college completion, such an effect must arise from less-skilled women moving out of metropolitan areas experiencing large influxes of low-skilled immigration. This result is consistent with previous research (Cortes 2008) suggesting that low-skilled immigrants may displace low-skilled natives. Moreover, it implies that selective out-migration among less-skilled natives might have dampened the depressing effects of low-skilled immigration on the cost of household services.

Conditional on college completion and choice of location, a third form of composition bias may occur because we limit the sample to women not enrolled in school. Based on similar logic as above, college graduates with strong fertility and work preferences might exit the sample to enroll in an advanced degree program resulting in a downward-biased estimate of the causal effect of low-skilled immigration. We test for composition bias due to selective enrollment directly by estimating the likelihood of school attendance among college graduates living in the sample of MAs using an empirical model similar to equation (14). The estimated instrumental variables coefficient on LSI from the normit regression is insignificant (p = 0.82) and trivial in magnitude. Hence, this form of composition bias also has a negligible impact on our results.

[22] Taking the point estimate at its face, in order for composition bias to abolish our main results, selected women would, on average, need to have given birth to 4.1 *more* children than unselected women between the ages of 22 and 39. This differential seems implausibly large given that the average rate of fertility among all college graduates in our sample was 1.5 children per woman.

6 Conclusion

Our analysis builds on a growing body of work highlighting the potentially beneficial effects that immigration has on natives (Cortes 2008; Cortes and Tessada 2009). In order to isolate a causal impact of immigration, we relied on a common instrumental variables approach to account for the simultaneity of the location decisions of new migrants with respect to local labor market conditions. Using settlement patterns predicted from historical enclaves as instruments, we found that low-skilled immigration to the United States between 1980 and 2000 led to substantial reductions in the cost of market-provided child care, food, and housekeeping services in large metropolitan areas. Adapting a bivariate probit model to allow for endogenous explanatory variables, we then analyzed how high-skilled native-born women responded in terms of their fertility and work decisions. The results suggest that the immigrant-driven reductions in the cost of household services increased the fertility of non-Hispanic U.S.-born college graduates. The rise in childbearing was accompanied by a smaller increase in exits from the labor force. Moreover, 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. Taken together, our findings indicate that low-skilled immigration substantially reduced the work-fertility tradeoff facing educated urban American women.

A topic that warrants further exploration is how the effects of childrearing costs on work and fertility decisions vary over the life course. Although we do not expect the effects to be constant across ages and parities, we studied average behaviors in this paper due to data limitations. The model developed in this paper could be extended to analyze richer lifecycle dynamics using longitudinal data paired with higher-frequency data on immigrant inflows. Another 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 U.S.-born Hispanic college graduates. One might expect a particularly strong complementarity between high-skilled native Hispanics and low-skilled 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 remains a question 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. By contrast, Goldin’s (2004) assessment of detailed cohort data showed that, relative to older cohorts, women graduating from college in the 1980s have been significantly better able to combine both career and family. Our work suggests that women in large metropolitan areas are in fact facing smaller tradeoffs when making fertility and labor supply decisions, and that this has, in part, been driven by the continuing flow of low-skilled immigrant workers into the United States.

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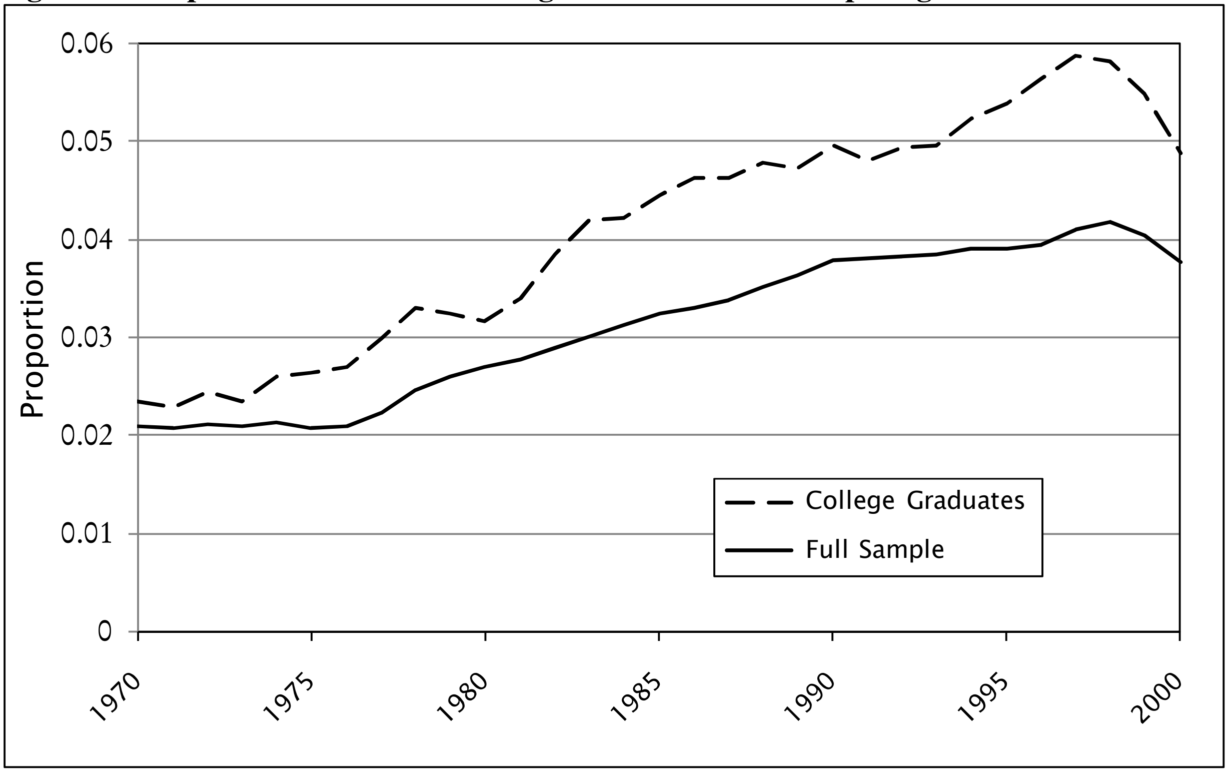
Figure 1. Proportion of Women Bearing Children and Participating in the Labor Force



Notes: See text for definition of the sample and variable.

Table 1. Distribution of Nativity and Educational Attainment in Household Service Occupations, 2000

Characteristic	Child Care	Food Services	Housekeeping	All Other Occupations
Nativity:				
Immigrant	14.0%	18.3%	33.5%	12.7%
Native	86.0%	81.7%	66.5%	87.3%
Educational attainment, immigrants				
No post-secondary	66.6%	76.9%	84.7%	48.7%
Some post-secondary	22.8%	16.7%	10.4%	22.3%
BA equivalent or higher	10.5%	6.4%	4.9%	28.9%
Educational attainment, natives				
Native, no post-secondary	52.7%	66.7%	82.1%	39.2%
Native, some post-secondary	38.0%	28.6%	15.9%	33.2%
Native, BA equivalent or higher	9.3%	4.7%	2.1%	27.6%

Notes: Estimates are based on 2000 Census data from IPUMS (Ruggles et al. 2010). Occupations are defined according to the consistent 1990-basis classification developed by the IPUMS group. Educational attainment is defined using the IPUMS consistent recode developed to bridge the 1980 and 1990 censuses.

Table 2. Characteristics of the Working-Age Population of Immigrant Groups Included in the Instrument

Country of Origin	Number in 1970	Total Change, 1970-2000	Low-Skilled Change, 1970-2000	Share Low-Skilled in 2000
Dominican Republic	41,275	505,759	343,962	69.8%
Ecuador	26,300	213,645	129,134	62.0%
Haiti	21,125	317,459	186,790	59.3%
Mexico	549,125	6,689,664	5,696,296	85.7%
Portugal	70,400	97,869	61,329	75.6%

Notes: Estimates are based on 1970 and 2000 Census data from IPUMS (Ruggles et al. 2010). The "working-age" and "low-skilled" segments of the population are as defined in the text.

Table 3. Analysis of Log Median Wage in Child Care Among Natives and Non-Recent Immigrants

Variable	Ordinary Least Squares (1)	Ordinary Least Squares (2)	Ordinary Least Squares (3)	Instrumental Variables
LSI	0.755***	-1.103***	-1.465***	-3.238***
	(0.184)	(0.369)	(0.393)	(0.384)
IncControl			0.474**	0.740***
			(0.194)	(0.202)
MSA fixed effects	No	Yes	Yes	Yes
Region-Year fixed effects	Yes	Yes	Yes	Yes
R^2	0.921	0.969	0.970	
First-stage F statistic				31.94

Notes: Occupation is based on the consistent 1990-basis classification developed by the IPUMS group (Ruggles et al. 2010). Each column represents a different specification applied to 1980, 1990, and 2000 data from the 77 metropolitan areas (MAs) described in the text (N = 231). Median wages are calculated using only natives and immigrants living in the United States for at least 10 years. *LSI* denotes the fraction of the local labor market consisting of immigrants without any post-secondary schooling, and *IncControl* is the log of income per capita among male college graduates. The fourth column relies on the enclave-based instrumental variable defined in equation (2). All specifications include time-varying region fixed effects. Each MA-year observation is weighted by the estimated number of childcare workers, and the robust standard errors in parentheses are clustered by MA.

*p < .10; **p < .05; ***p < .01

Table 4. Effects of Low-Skilled Immigration on Household Services Markets

Estimates by sector	Log Wage, Natives and Non-Recent Immigrants: 25th Percentile	50th Percentile	75th Percentile	Normit(Share of Labor Force)
(A) Child Care:				
Coefficient on *LSI*	-4.826***	-3.238***	-2.024***	1.348***
	(0.582)	(0.384)	(0.328)	(0.273)
Effect of average change in LSI, 1980-2000[a]	-19.20%	-12.90%	-8.07%	11.10%
(B) Food Preparation:				
Coefficient on *LSI*	-1.167***	-1.125***	-0.786***	0.786***
	(0.330)	(0.277)	(0.235)	(0.166)
Effect of average change in LSI, 1980-2000[a]	-4.65%	-4.48%	-3.13%	6.44%
(C) Housekeeping Services:				
Coefficient on *LSI*	-2.261***	-1.319***	-1.579***	1.252***
	(0.505)	(0.490)	(0.571)	(0.383)
Effect of average change in LSI, 1980-2000[a]	-9.01%	-5.26%	-6.29%	14.10%

Notes: Occupations are based on the consistent 1990-basis classification developed by the IPUMS group (Ruggles et al. 2010). Each coefficient estimate comes from a separate instrumental variables equation and is based on 1980, 1990, and 2000 data from the 77 metropolitan areas (MAs) described in the text. "Non-recent immigrants" refers to non-natives living in the United States for at least 10 years at the time of observation. The variable *LSI* and its instrument are described in the notes to Table 3. Each equation includes time-varying region fixed effects, MA fixed effects, and income per male working-age college graduate. Within each occupation, the three wage equations that span a row are estimated jointly, as are the three normit equations used to predict occupational shares in the final column. Each MA-year cell (N = 231) is weighted by the estimated population of individuals used to calculate the associated dependent variable, and the robust standard errors in parentheses are clustered by MA.

[a] Reported effects are the change in the underlying dependent variable that would be caused by the change in *LSI* experienced by the average working-age native in the sample.

*p < .10; **p < .05; ***p < .01

Table 5. Effects of Low-Skilled Immigration on Fertility and Work Outcomes of College-Educated Non-Hispanic Natives

	Normit, Fertility Rate	Normit, LFP Rate	Tetrachoric, Fertility and LFP
Coefficient on *LSI*	1.484**	-0.795**	0.790*
	(0.592)	(0.361)	(0.438)
Mean of underlying dependent variable in 2000	0.0898	0.8448	-0.3981
Effect of average change in *LSI*, 1980-2000[a]	0.0085	-0.0072	0.0306

Notes: Instrumental variables estimates are based on 1980, 1990, and 2000 data from the 77 metropolitan areas (MAs) described in the text. The data consist of MA-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 time-varying region fixed effects, MA 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. The variable *LSI* and its instrument are described in the notes to Table 3. Each of the MA-year-age group cells (*N* = 462) is weighted by the population of women represented by the cell, and the robust standard errors in parentheses are clustered by MA.

[a] Reported effects are the change in the underlying dependent variable that would be caused by the change in *LSI* experienced by the average member of the sample.

*p < .10; **p < .05; ***p < .01

Table 6. Effects of Low-Skilled Immigration on the Joint Likelihood: Main Specification and Robustness Checks

	Main Specification	Three Age Groups	Excluding Dallas, LA, and NYC	Excluding California	Adding 1970 Controls
Coefficient on *LSI*	1.737**	1.667**	2.425***	3.694**	1.840***
	(0.804)	(0.670)	(0.896)	(1.840)	(0.618)
Mean of joint rate of fertility and LFP in 2000	0.0568	0.0568	0.0586	0.0577	0.0568
Number of metropolitan areas	77	77	74	65	77
Effect of average change in *LSI*, 1980-2000:[a]					
Direct Estimate	0.0065	0.0062	0.0080	0.0115	0.0068
Indirect Estimate	0.0065	0.0068	0.0062	0.0095	0.0060
Proportion explained by weakened correlation	28.1%	34.6%	36.6%	24.5%	45.2%

Notes: Instrumental variables estimates are based on 1980, 1990, and 2000 data from the metropolitan areas (MAs) described in the text. The data consist of MA-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 5. The variable *LSI* and its instrument are described in the notes to Table 3, and the additional explanatory variables included in the final column are described in the text. Each of the MA-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 MA.

[a] Reported effects are the change in the underlying dependent variable that would be cause by the change in *LSI* experienced by the average member of the sample. Direct estimates rely on the coefficient on *LSI* reported in the table. The indirect estimate and the proportion explained by a weaker fertility-LFP correlation are based on auxiliary IV regressions and equation (13).

*p < .10; **p < .05; ***p < .01