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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 work and fertility decisions of high-skilled women born in the United States. The evidence we present indicates that low-skilled immigration to large metropolitan areas between 1980 and 2000 lowered the cost of market-based household services. Using a novel estimation technique to analyze joint decision making, we find that college-educated native females responded, on average, by increasing fertility and reducing short-run labor force participation. These changes were accompanied by a weakening of the negative correlation between work and fertility, as well as an increase in the proportion of women who both bore children and participated in the labor force. Taken in combination, our estimates imply that the continuing influx of low-skilled immigrants substantially reduced the work-fertility tradeoff facing educated urban American women.

Journal of Economic Literature Classification: D10, F22, J13, J22, R23

Keywords: Child care, fertility, household services, labor supply, immigration

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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 (Clark and King 2008). 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 debate focuses on the potentially negative impact of immigration on the wages and employment rates of natives (Altonji and Card 1991; Bean and Stevens 2003; Borjas 2003; Butcher and Card 1991; Card 1990, 2001). There has also been an increasing amount of research that considers the impact of immigration on other domains such as schooling (Borjas 2000; Borjas 2004; Gould, Lavy, and Paserman 2005) and crime rates (Sampson, Morenoff and Raudenbush 2005; Borjas, Grogger and Hanson 2006). Perhaps driven by the observation that immigrants in the United States have become disproportionately low-skilled, most research has largely concentrated on the extent to which immigration constrains opportunities for low-skilled natives via a crowding-out effect.[1] Substantially 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 cost and availability of market-provided household services in large urban areas and how high-skilled U.S.-born females respond in terms of their childbearing and work behavior. We pay particular attention to the magnitude of the tradeoff, or "role incompatibility," between fertility and labor force participation.

Our analysis makes use of inter-city differences as a source of variation in the concentration of immigrants. Because immigrant location decisions are likely to be related to local economic conditions, standard regression-based analysis does not adequately capture the causal impact of low-skilled immigrants on local labor market outcomes. As a result, we use an instrumental variables approach that relies on the propensity of new entrants to locate in areas with high concentrations of existing immigrants from the same country (Bartel 1989; Card 2001).

[1] 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.

Thus, the predicted flow 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 (2008) establishes a baseline impact of low-skilled immigration on the general price of locally traded 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, namely child care, food preparation, and housekeeping. We show that low-skilled immigration led to substantial reductions in wages and increases in employment in these household service occupations between 1980 and 2000. The changes in the price and 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 employment choices is unclear due to the joint nature of the decision (Blau and Robins 1989).

In order to shed light on this question we develop a novel estimation technique that allows us to examine the impact of low-skilled immigration on the simultaneous fertility and labor force participation outcomes of U.S.-born women. This analysis again relies on the instrumental variables strategy discussed above to establish a causal relationship. We focus in particular on non-Hispanic college graduates in order to isolate the effects of low-skilled immigration that result from changes in the household services markets.[2] 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 tradeoff between work and fertility was not one-for-one. These results complement the analysis of Cortes and Tessada (2008), who find that low-skilled immigration to the U.S. led to increases in the number of hours worked conditional on being employed, but a decrease in the rate of labor force participation among native female college graduates. In our analysis of role incompatibility, we find that the negative correlation between work and fertility among urban high-skilled women

[2] 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.

was also dampened by the inflows of immigrants to their metropolitan area. Finally, we combine these three estimates to analyze a more concrete measure of role incompatibility, finding that low-skilled immigration led to substantial increases in the joint likelihood of childbearing and labor force participation.

The paper proceeds as follows. In Section 2 we assess the potential importance of low-skilled immigration as a source of the substantial "unexplained" increases in labor supply to the childcare sector (Blau 1991) that are associated with the sharp declines in role incompatibility witnessed in the United States. We also place our paper within the context of the literature concerning fertility, labor supply, and childrearing costs. Section 3 describes the data and methods that we use in our analysis, laying out the instrumental variables estimation strategy used to isolate the causal effect of low-skilled immigration. We describe the novel 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. Our main results regarding the effects of low-skilled immigration on household services markets and on female fertility and work outcomes are presented in Section 4. This analysis includes specification checks concerning the validity of the instrumental variables method and the extent to which geographic and educational selection among natives might affect our results. Section 5 provides additional discussion and concluding remarks, and the two appendices describe our empirical framework in greater detail.

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 childrearing implies a tradeoff between labor supply and fertility, particularly for females because their traditional role has been to perform household work. In the sociology literature, this phenomenon is often referred to in terms of an incompatibility between the roles of mother and worker. Although sociologists tend to emphasize the institutional constraints that affect work decisions, it is the underlying time constraint that drives role incompatibility (Rindfuss and Brewster 1996; Stycos and Weller 1967). 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.

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 and childbearing 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 variation in role incompatibility and understanding its sources. Cross-country studies focus on the substantial differences in family policies, childcare availability, unemployment rates, stability of labor contracts, and gender norms that may result in international differences in childbearing and work patterns (Adsera 2004; Brewster and Rindfuss 2000; de Laat and Sevilla Sanz 2007). Within-country data analyses indicate a correlation between fertility and female labor force participation that is consistently negative. 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 paper we consider the extent to which low-skilled immigration to the United States has reduced the cost of childbearing, altered work and fertility patterns, and ultimately led to a 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 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). Figure 1 provides evidence of this phenomenon based on labor force participation and fertility patterns observed among adult fecund women in the United States between 1970 and 2000.[3] As seen in the figure, labor force participation among women with a child younger than one rose from just over 20% in 1970 to around 60% in 2000. Among college graduates, there was 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%, a threefold increase.

Labor force participation rates of recent mothers may not, however, 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. We find it preferable to focus on the changes in the joint rate of childbearing and labor force participation since it accounts for changes in both outcomes simultaneously. An increase in the joint likelihood reflects increases in fertility among working women and/or increases in labor force participation among mothers, both of which may be thought of as measuring a reduction in role incompatibility. The joint rate of childbearing and labor force participation is small in absolute terms, which reflects the relative infrequency of childbirth. However, as seen in Figure 2, the joint likelihood of fertility and work among adult fecund women almost doubled between 1970 and 2000. Among college graduates the joint likelihood more than doubled, increasing from approximately 2.3% to 4.7%. 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.

[3] The figures in this section draw on data from the March Current Population Surveys (CPS), 1969-2001 (King et al. 2004). The samples are comprised of women ages 18-39, who we refer to as "adult fecund women." Educational attainment is based on the temporally consistent classification system developed by the IPUMS group (Ruggles et al. 2004). 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 the CPS data to smooth out year-to-year fluctuations.

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. Additional calculations using CPS microdata indicate that the negative correlation between work and childbearing among adult fecund college graduate women fell by 50% between 1970 and 2000. However, using the correlation coefficient as a metric for role incompatibility is problematic in that it does not provide quantifiable information about changes in behavior. The more structured statistical model that we employ in our empirical analysis allows us to translate the weakening of the correlation between work and fertility into changes in the joint likelihood of childbearing and labor force participation, thus obtaining more interpretable evidence on the decline in role incompatibility.

2.2 Childrearing costs and role incompatibility

The costs of childrearing play a central role in the theoretical discussion of role incompatibility (Rindfuss and Brewster 1996). Hence, 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 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 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, an unambiguous effect is a reduction in role incompatibility: 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 likelihood of fertility among women in the workforce. More succinctly, there should unambiguously be an increase in the joint likelihood of work and fertility.

A remaining question, then, is what may have driven down the costs of childrearing in the United States. One possibility is the Family Medical Leave Act (FMLA) of 1993. Consistent with the FMLA mandate, Berger and Waldfogel (2004) demonstrated that qualifying mothers were more likely to utilize parental leave and then return to employment after childbirth. Since women on maternity leave are considered by the Census to be in the labor force, part of the increase in the joint likelihood of fertility and work after 1993 may be attributable to the FMLA. However, the findings in Blau (2001) also point to changes in the market for child care as another explanation. Americans have come to rely heavily on privately-funded child care purchased outside the home (Gornick, Meyers and Ross 1997; Rindfuss and Brewster 1996). Yet, despite large increases in the demand for child care, 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 sector. We suggest that immigration may partially explain this phenomenon.

2.3 Low-Skilled Immigration and Household Services

Cortes (2008) analyzes the effect of low-skilled immigration on the relative prices of locally-traded goods and services in major U.S. cities using an instrumental variables strategy similar to ours. She concludes that immigration improves the purchasing power of high-skilled natives. Whereas Cortes uses an agglomerate of non-traded goods, we focus on three service markets that potentially provide substitutes for time-intensive childrearing tasks undertaken by parents: child care, food services, and housekeeping.[4]

The large numbers of low-skilled immigrants who arrived in the United States after 1965 are likely to have made a substantial impact on household services markets. As can be seen in Table 1, by the end of the 20th century, immigrants were overrepresented in child care,

[4] Occupation definitions are based on the consistent classification (1990 basis) system available from the IPUMS project (Ruggles et al. 2004). 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.

food services and housekeeping, relative to other occupations.[5] These sectors also generally employed individuals at the lower end of the educational distribution. In the market for child care, for example, 53% of workers in 2000 were "low-skilled," defined as having completed years of schooling equivalent to a high school degree or less, while less than 10% had a college degree. By comparison, the corresponding numbers in all non-household service occupations were roughly 41 and 27%, respectively. Moreover, the immigrants employed in child care were excessively low-skilled both by comparison to immigrants in non-household service occupations and natives employed in child care. Although immigrants as a whole were only slightly overrepresented in child care, low-skilled immigrants constituted a substantially larger share of the workforce in child care relative to non-household service occupations. In child care, 8.7% of workers were low-skilled immigrants, as compared to 6.5% of the workforce in the non-household service sectors. This pattern was even more striking in the food service and housekeeping sectors, in which 14.3% and 28.0% of workers were low-skilled immigrants, respectively.

A potential concern for our later analysis is that labor supply shifts attributable to low-skilled immigration might not affect the cost of the household services actually purchased by college-educated women. This may be particularly an issue in the market for child care. Although workers at formal childcare centers generally receive relatively low wages (Blau 1993; Helburn and Howes 1996), there is substantial variation across centers in the quality of care provided (Phillipsen, Cryer, and Howes 1995). Previous research indicates a positive relationship between educational attainment of childcare providers and observable measures of childcare quality (Blau 2000; Whitebook, Howes, and Phillips 1989). Moreover, educated and high-income women may demand a better quality of care (Blau and Hagy 1998; Hotz and Kilburn 1991). Thus, although low-skilled immigrants may increase the pool of childcare workers available, this could be irrelevant to the segment of the quality distribution from which college-educated women are obtaining child care.

[5] Throughought 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. 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.

To definitively establish whether low-skilled immigration affects the childcare costs of high-skilled natives would require data on the characteristics of childcare workers of the final consumer. To our knowledge, such data 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. The wage bill accounts for between 60% and 80% of the operating expenses at formal and home-based childcare centers (Blau and Mocan 2002; Helburn and Howes 1996), and likely represents an even higher share of the final costs of informal childcare providers. Consequently, we draw inference on the quality of care provided by childcare workers based on their position in the salary distribution. Using data from the 2000 census public-use microdata files, we have re-calculated the characteristics of childcare workers listed in Table 1 by wage quartile. The disaggregated tabulations, available on request, indicate that the foreign-born share of childcare workers increases with the wage, while the fraction that is low-skilled decreases. The net result is that the representation of low-skilled immigrants among childcare workers decreases only slightly to 8.0% at the top quartile, as compared with 8.9% in the third quartile of the wage distribution. Hence it seems likely that inflows of low-skilled immigration increased the labor supplied to all segments of the childcare market resulting in reductions in the wages of both high- and low-quality providers. We test this hypothesis more formally below, but if this is the case, then low-skilled immigration should have, indeed, reduced the cost of child care services utilized by college graduate women.

2.4 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. Some studies take work decisions as given, looking at the association between childcare availability, which affects the opportunity cost of care due to time and search, and fertility (Hank and Kreyenfeld 2003; Kravdal 1996; Lehrer and Kawasaki 1985). This research generally finds a negative, but statistically weak, relationship between childcare availability and childbearing. It is, however, difficult to interpret these findings as causal because childcare availability is not likely to be exogenous. As discussed in Hank (2002), observed regional variation in fertility may be explained by the spatial distribution of individual characteristics as opposed to the availability of child care.

The rather scarce literature on the effect of childcare availability on fertility contrasts with a much more extensive literature on the relationship between childcare costs and the likelihood that mothers work. This strand of research generally indicates a negative relationship between the price of child care and the conditional likelihood of work (Blau and Robins 1988; Connelly 1992; Han and Waldfogel 2001). Most of these analyses suffer from endogeneity concerns similar to those discussed above if the price of care responds to unobserved changes in local preferences for children and work and/or unmeasured local labor market conditions. Several more recent studies rely on variation in policies concerning government-subsidized child care and the availability of 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; Gellbach 2002). Nonetheless, as described above, changes conditional likelihood may be driven by differential selection into motherhood. For example, Mason and Kuhlthau (1992) find that mothers with a stronger labor force attachment were more likely to report that the availability of child care affected birth timing and the number of births.

There are a handful of papers that examine the effects of childcare costs on both employment and fertility outcomes. Mason and Kuhlthau (1992) examine mothers' responses to survey questions on whether the availability of child care constrained their employment and fertility decisions, finding a larger perceived effect on employment. Blau and Robins (1989) analyze transitions among employment and fertility states basing their measure of childcare costs on geographic variation in average per-child weekly expenditures. They report that higher local childcare expenditures are associated with lower rates of employment among all women and with decreases in childbearing among the non-employed. Taking a different approach, Stolzenberg and Waite (1984) conclude that labor force participation is less affected by fertility among married women living in areas with less expensive child care (as measured by the wages of childcare workers) and a greater number of child care workers. These findings suggest that lower childcare costs reduce role incompatibility, but the results must be interpreted with caution due to the endogeneity of the cost measures used. Local childcare expenditures almost certainly reflect underlying local preferences and/or economic conditions. Wages of childcare workers, which in turn affect the supply of labor to the childcare sector, are also endogenous due to price pass-through effects. Although we do not consider childcare costs explicitly in our analysis of fertility and labor supply outcomes, we use an instrumental variables approach to isolate exogenous variation in relative size of the local immigrant population, which we show to be causally associated with both lower costs and greater availability of care.

Based on an instrumental variables approach similar to ours, Cortes and Tessada (2008) provide evidence that low-skilled immigration to the United States has led to an increase in the hours worked among highly-skilled females, conditional on participating in the labor force. However, both Khananusapkul (2004) and Cortes and Tessada (2008) find the opposite effect when considering the unconditional likelihood of labor force participation. Neither of these studies considers the role of fertility in explaining the relationship between immigration and labor supply. 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 reduced role incompatibility, as measured by the joint likelihood childbearing and labor force participation.

3 Data and Methods

Broadly, our analysis proceeds in two steps. First, we consider the extent to which immigration 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 employment and childbearing decisions of American females. 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 non-Hispanic college-graduate native females. Sharply differentiating immigrants and natives by skill minimizes the possibility of competition for jobs, which might directly affect female wages and 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 came from Latin America. In both steps of the analysis we use an instrumental variables approach that allows us to estimate a causal effect of low-skilled immigration.

Throughout, our empirical models 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 is drawn from the U.S. Census Bureau's 1980, 1990, and 2000 public-use microdata samples, 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. 2004). The underlying geographic sampling units defined by the Census Bureau have changed over the years, resulting in a number of inconsistencies in the degree to which the population of a metropolitan area is covered in the microdata files. As discussed in Appendix A, this is likely to result in noisy and biased estimates of metro-level variables. In order to mitigate these effects, we attempted to establish reasonable consistency within our sample between 1980 and 2000 in terms of population coverage. We started with the 100 most populous metropolitan statistical areas (MSAs), as measured in 1990.[6] After recoding for consistency, our final sample includes 70 metropolitan areas, which accounted for approximately 80% of the 1990 working-age immigrant population in the United States. Appendix A describes the criteria used in our recoding procedure, lists the included and excluded MSAs, and compares the population characteristics of the two sets of metropolitan areas.

3.1 Immigration and Childrearing Costs

The first step of our analysis considers the effects of immigration on the cost of market-provided child care, food services, and housekeeping. Based on data from our metro sample, we use wages of workers occupied in these sectors as a measure of the price of the associated

[6] The practical rationale for the population-based restriction is laid out in Appendix A, but focusing on large metropolitan statistical areas in this manner also makes sense, given the extent to which immigration is an urban phenomenon. We use "metropolitan statistical area" as a general term for Census-defined MSAs and for primary metropolitan statistical areas (PMSAs). Additionally, as described in Appendix A, in a few instances PMSAs had to be agglomerated to obtain geographic consistency over time.

services.[7] 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 here are equally applicable to the markets for food services and housekeeping. Our empirical analysis covers all three markets.

Low-skilled immigrants are overrepresented in the childcare sector, relative to most other occupations (cf. Table 1). As a result, an increase in the low-skilled immigrant share of the metropolitan workforce should increase the supply of labor, driving the wages of childcare workers down. As discussed above, although low-skilled immigrants may drive down the average wage, this could be irrelevant to high-skilled women who demand a higher-than-average quality of child care. Given the relationship between quality and wages discussed earlier, we address this issue by considering the effects of immigration on wages at various percentiles of the local salary distribution. A potential pitfall of this strategy is that an influx of low-skilled immigrants might reduce an upper-percentile wage simply because their arrival in between census years results in more mass at the bottom of the wage distribution. To circumvent this 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:

$$w_{mt} = \beta LSI_{mt} + \boldsymbol{\mu}_m + \boldsymbol{\mu}_{kt} + \lambda IncControl_{mt} + \varepsilon_{mt} \; . \tag{1}$$

The dependent variable w_{mt} 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 25th, 50th, or 75th percentile.[8] The low-skilled immigrants share of the overall working-age population

[7] As already noted, payments to labor account for 60-80% of the final cost of childcare. This means that immigrant-induced declines in wages should translate into substantial reductions in the final cost of care. 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. Finally, wages are a smaller, but not inconsequential, component of production costs for restaurants (Lee, Schluter, and O'Roark 2000).

[8] 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.

immigration is denoted by LSI. We designate as "low skilled" those individuals who have completed education equivalent to a high school degree or less and define "working" ages as 20 to 64. Educational attainment is based on the consistent educational recode developed by the IPUMS group (Ruggles et al. 2004). MSA-specific intercepts are indicated by $\boldsymbol{\mu}_m$, while $\boldsymbol{\mu}_{kt}$ represents time fixed effects specific to the kth Census region. The variable $IncControl_{mt}$ denotes the log of income per capita among working-age males who have completed college, assigning imputed values to individuals whose income has been top-coded using a region-specific Pareto extrapolation. 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. For each services market, we estimate the three log-quantile equations jointly. To account for heteroskedasticity in the error term of (1), we weight by the number of observations used to construct the dependent variable. Additionally clustering standard errors by MSA allows for arbitrary patterns of correlation of the error term between percentiles and over time.

If immigration represents a supply shift, an increase in LSI should result in lower wages and so we expect the estimate of β to be negative. However, if the location decisions of immigrants are endogenous, it is difficult to interpret β as a causal estimate of low-skilled immigration. Metropolitan areas with booming economies might simultaneously exhibit an increased willingness to pay for child care and be more attractive to new migrants. A higher local demand for household services arising for other reasons could also draw in workers occupied in those sectors. In both scenarios, we would see more low-skilled immigrants arriving in areas with high pre-existing wages in the childcare sector. Thus, ordinary least squares will tend to result in an attenuated (less negative) estimate of the causal effect of immigration.

The fixed effects included in the model allow us to partially address these concerns. Persistent differences across metropolitan areas in economic conditions, household services markets, and immigration will be absorbed by the MSA fixed effects, $\boldsymbol{\mu}_m$. The region-year fixed effects, $\boldsymbol{\mu}_{kt}$, account for both decade-to-decade changes in the nation as a whole, for example the passage of the Family Medical Leave Act between 1990 and 2000, and differential changes between regions over time, for example, the relatively slow economic growth in the Midwest after 1980. The decade-by-region fixed effects also act as price deflators, given the log specification of the dependent variable. The fixed effects included in equation (1) do not, however, capture differential changes among MSAs within a region over time.

Although the income control accounts for certain aspects of differential economic change, equation (1) omits many time-varying factors that may affect the markets for child care and other household services. One issue is that quite a few of these factors, for example the employment rate, the age structure, the educational distribution, and occupational structure, may be a consequence of, as well as a cause of, low-skilled immigrant inflows. Hence, including them in the model would artificially absorb part of the structural effect of low-skilled immigration. Even the income control that we do include potentially suffers from this limitation if males adjust their labor supply in response to cheaper household services. The second, and more fundamental, problem is that variables such as preferences and norms, are fundamentally non-measureable and cannot be directly controlled for in the empirical model. Consequently, we ultimately rely on an instrumental variables approach to identify the causal relationship between low-skilled immigration and wages in the household services sectors. Instrumental variables will also address attenuation bias in the estimate of β due to measurement error in the percentage foreign born, arising, for example, from an undercount of undocumented immigrants.

3.1.1 Instrumental Variables Estimation

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; Massey et al. 1993; Munshi 2003). Following a similar line of reasoning as Card (2001), Cortes (2008), and Cortes and Tessada (2008), our instrument uses historical enclaves to predict the flow of subsequent migrants across MSAs. More specifically, the instrument for the low-skilled immigrant share of working-age adults in the local labor market (LSI_{mt}) can be written as

$$INST_{mt} = \sum_b \frac{N^b_{m,1970}}{N^b_{1970}} \times \left[NLS^b_t - NLS^b_{1970} \right] \ . \tag{2}$$

For each country of birth, b, the first term in equation (2) represents the fraction of all immigrants from country b living in MSA 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. Two conditions need to hold for this to be a valid instrument. 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: it should have no direct impact on the outcome variables. If both of these conditions hold, then two-stage least squares can be used to estimate the causal impact of low-skilled immigration. We discuss each of the two conditions in turn.

To maximize the predictive power of our instrument, we focus our attention 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.[9] Five immigrant groups meet these criteria: Dominicans, Ecuadorians, Haitians, Mexicans, and Portuguese. Descriptive statistics on these groups are given in Table 2. As indicated in the table, the share of working-age immigrants from these countries that was low-skilled in 2000 ranged from 59 to 86%. By contrast, low-skilled individuals comprise 42% of the native population. In our results we verify the strength of the instrument empirically using the method indicated in Stock et al. (2002).

In order for the exclusion restriction to be met, the instrument should be related to outcomes only through the allocation of low-skilled immigrants across metropolitan areas. The necessary criteria for this to be true are very similar to those outlined by Cortes (2008). Given the MSA 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 across MSAs within a region 15 to 25 years later, and (b) differential economic changes among MSAs within a region should not affect the overall inflow of low-skilled immigrants to the United States. If the exclusion restriction holds, then the fitted

[9] 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.

values of LSI based on

$$LSI_{mt} = \delta INST_{mt} + \boldsymbol{\mu}_m + \boldsymbol{\mu}_{kt} + \lambda IncControl_{mt} + \xi_{mt}$$

will not be correlated with unobserved local economic shocks. Substituting these fitted values into equation (1) will yield an estimate of β that does not suffer from bias arising from the concerns about endogeneity and measurement error described above. A secondary concern is that low-skilled immigrants do not have a direct effect on the supply of labor to the household services sectors *per se* and that wages are decreased by a different mechanism. Even if this were true, it would still represent an indirect causal impact of immigration. Moreover, it is empirically testable. Our calculations indicate that, as of 2000, immigrants from the selected countries listed in Table 2 were twice as likely as natives to be occupied in the household service sectors we analyze. We also present instrumental variables results below suggesting that immigrant-induced changes in wages were accompanied by expansions in employment in child care, food services, and housekeeping services. Hence, we are able to conclude that low-skilled immigration affects not just the cost, but also the availability of market-provided household services.

3.2 Fertility and Work Outcomes

The second step of the analysis estimates the effects of low-skilled immigration on female childbearing and labor force participation decisions. We attempt to isolate the channel of influence arising from immigrants' effects on low-skilled household service markets by limiting our sample of females to non-Hispanic natives with college degrees, aged 23-39, and not living in institutional group quarters. This represents a pool of potential mothers who would be most likely to purchase market-provided household services and least likely to have their employment prospects or preferences directly affected by the low-skilled immigrants included in the instrumental variable described above. 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 later present additional results that justify the enrollment-based restriction

We use two techniques to analyze the effects of low-skilled immigration on the joint likelihood of work and fertility, which, for the reasons discussed in the background section, is our primary measure of role incompatibility. The first relies on a bivariate model of fertility and labor supply decisions, while the second directly estimates the impact of immigration on the joint likelihood in a univariate framework. Although it is less straightforward to implement, the bivariate approach yields deeper insight into the chilbearing and work behaviors whereby low-skilled immigration exerts its effects. Directly estimating the effect of low-skilled immigration on the joint likelihood serves as a robustness check on this indirect method. Since immigrant location decisions may be affected by the local demand for childcare and other household services, low-skilled immigration to a metropolitan area cannot be treated as exogenous with respect to childbearing and work decisions. As a result, both techniques will ultimately rely on instrumental variables estimation.

3.2.1 The Bivariate Model

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

$$\begin{aligned} C^*_{imt} &= \beta_1 LSI_{mt} + \boldsymbol{\omega}'_1 \mathbf{v}_{imt} + \varepsilon^C_{imt} && (3) \\ L^*_{imt} &= \beta_2 LSI_{mt} + \boldsymbol{\omega}'_2 \mathbf{v}_{imt} + \varepsilon^L_{imt} \ , && (4) \end{aligned}$$

where C^*_{imt} and L^*_{imt} describe the desirability of childbearing and labor force participation (LFP) of woman i living in metropolitan area m in year t. The associated binary outcomes are C_{imt} and L_{imt}, where $C_{imt} = 1$ is observed if $C^*_{imt} > 0$ 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 $\mathbf{v}_{imt}$, 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.

Low-skilled immigration may affect both the pecuniary cost and the time cost (given changes in availability) of obtaining multiple household services. It is not, however, possible to include specific measures of these effects in equations (3) and (4) because instrumental variables estimation requires having at least as many instruments as endogenous variables. Given that we only have one instrument, 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 will, 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.

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, $corr\left(\varepsilon_{imt}^{C}, \varepsilon_{imt}^{L}\right)$ is, by definition, the *tetrachoric correlation*. There is substantial variation in the tetrachoric correlation between fertility and labor force participation over time and across space.[10] As discussed in Appendix B, the tetrachoric correlation can be understood as the degree to which changes in childbearing not explained directly by the model translate into changes in labor force participation, and vice-versa. This might reflect, for example, 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. As such, we expect the tetrachoric correlation to be negative, which is almost universally the case in our sample. We explore how the correlation is affected by low-skilled immigration based on the parameterization

$$corr\left(\varepsilon_{imt}^{C}, \varepsilon_{imt}^{L}\right) \equiv \rho_{gmt} = \beta_3 LSI_{mt} + \boldsymbol{\omega}_3' \mathbf{v}_{gmt} \; . \tag{5}$$

[10] Our calculations using CPS microdata discussed in Section 2.1 indicate a gradual weakening of the tetrachoric correlation in the United States over time. There is also evidence of substantial cross-sectional differences in the correlation across MSAs in our sample. In the international context, Ahn and Mira (2002) provide evidence that the cross-country correlation between fertility and employment changed dramatically between 1970 and 1995.

The index g denotes a group of women who share common characteristics $\mathbf{v}_{gmt}$; when we estimate the model we will use age groups, but for now we define a "group" in general terms. If an increase in LSI results in cheaper and more available market-based household services, we expect β_3 to be negative. That is, low-skilled immigration should dampen the negative relationship between childbearing and labor supply.

Endogenous regressors in univariate binary choice models may be addressed using control functions and related strategies (e.g. Rivers and Vuong 1988; Blundell and Powell 2004). Our interest in explicitly parameterizing ρ makes this approach somewhat difficult to extend to the simultaneous choice setting. 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. Aggregating observations according to group characteristics, metro area and time, we can recover the model coefficients by analyzing sample proportions.

Let p^C_{gmt}, p^L_{gmt}, and p^{CL}_{gmt} denote the observed proportions of the n_{gmt} women in group g in metro area m in year t that bear children, participate in the labor force, and do both, respectively. As in Amemiya (1974), a first-order Taylor expansion of the sample proportions around their expected values results in the system of linear equations:

$$
\begin{aligned}
c_{gmt} &= \beta_1 LSI_{mt} + \boldsymbol{\omega}_1' \mathbf{v}_{gmt} + u^c_{gmt}\ , \quad (6)\\
\ell_{gmt} &= \beta_2 LSI_{mt} + \boldsymbol{\omega}_2' \mathbf{v}_{gmt} + u^\ell_{gmt}\ , \quad (7)\\
\text{and } r_{gmt} &= \beta_3 LSI_{mt} + \boldsymbol{\omega}_3' \mathbf{v}_{gmt} + u^r_{gmt}, \quad (8)
\end{aligned}
$$

where $c_{gmt} \equiv \Phi^{-1}\left(p^C_{gmt}\right)$ and $\ell_{gmt} \equiv \Phi^{-1}\left(p^L_{gmt}\right)$ denote the inverse standard normal cumulative distribution (or "normit") function applied to the observed rates of childbearing and LFP. The variable r_{gmt} denotes the empirical tetrachoric correlation between outcomes among women in group g, metro area m, and year t, and the specification in equation (8) uses the parameterization given by equation (5). The algorithm used to compute r_{gmt} based on the observed sample proportions, p^C, p^L and p^{CL} is described in Appendix B. The appendix also provides the formulas used to compute the average marginal effects of the observed change in low-skilled immigration in our sample on the likelihoods of childbearing, of participating in the labor force, and of doing both.

To ascertain the impact of LSI on the joint likelihood, our method uses the estimates of β_1, β_2, and β_3 to calculate not only the overall effect of immigration, but also to separately calculate the effect arising from differential changes in the marginal likelihoods and the effect arising due to changes in the tetrachoric correlation between fertility and work. This indirect estimate of the overall impact of immigration on the joint likelihood is based 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. 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

$$d_{gmt} = \beta_4 LSI_{mt} + \boldsymbol{\omega}_4' \mathbf{v}_{gmt} + u^d_{gmt} \ , \tag{9}$$

where $d_{gmt} \equiv \Phi^{-1}\left(p^{CL}_{gmt}\right)$. 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. Given that low-skilled immigration's effects on household services markets should weaken role incompatibility, we expect LSI to have a positive effect on the joint likelihood of work and fertility ($\beta_4 > 0$).

3.2.2 Implementation and Instrumental Variables Estimation

If there are any groups in which any of the binary outcomes is uniform across its members, it is not possible to estimate the empirical tetrachoric correlation. We divide the sample of college-graduate women into two broad age groups (g): women ages 23-30 and women ages 31-39. Due to the sample sizes, disaggregating further by race and marital status is not possible. Consequently, we include measures of the average characteristics of the group as explanatory variables. Based on equations (6) through (9) we estimate a system of four regressions of the form

$$y_{gmt} = \beta LSI_{mt} + \boldsymbol{\mu}_m + \boldsymbol{\mu}_{kt} + \boldsymbol{\mu}_g + \boldsymbol{\lambda}_{mt}' IncControl_{mt} + \boldsymbol{\theta}' \mathbf{x}_{gmt} + u_{amt} \ ,$$

where y is one of the four dependent variables (c, ℓ, r, and d). Low-skilled immigration (LSI), the MSA 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. A measure of potential spousal income is already included in the metro controls, namely income per capita among male college graduates in the metropolitan area. Each group-metro-year observation is weighted by the number of underlying women to address heteroskedasticity and clustering standard errors by MSA accounts for an arbitrary pattern of correlation in the error terms across equations, groups, and time.[11]

As already noted, standard least squares techniques are likely to be biased due to geographic selection among immigrants. In areas where high-skilled mothers find it more desirable to work due to changes in unobserved factors, they may bid up the wages in the household services sectors, making the MSA more attractive to low-skilled immigrants. If this is the case, as well as when LSI is measured with noise, ordinary least squares will yield estimates of the βs that are smaller in absolute value than the true coefficients. Two-stage least squares based on the instrument in equation (2) and the overwhelmingly low-skilled immigrant groups listed in Table 2 should yield causal estimates on the parameters of interest for our sample of college graduate females.

[11] A two-step feasible generalized least squares (FGLS) estimator could be derived to address the within group correlation of error terms across the four equations. FGLS would not, however, address error terms that are correlated between age groups and over time within an MSA unless the form of the correlation is explicitly known.

4 Results

4.1 Wages and Employment in Household Services

Focusing for the moment on the median wage among native and non-recent immigrant child-care workers, Table 3 compares the impact of low-skilled immigration using different specifications applied to our panel of MSA-level data. The first three columns present estimates based on ordinary least squares (OLS). Without MSA 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 MSA fixed effects are included. This reversal in sign suggests that low-skilled immigrants are drawn to metropolitan areas where wages in the childcare market are persistently high, emphasizing the difficulty in interpreting the OLS coefficient causally. Controlling for income per capita among college graduates yields a slightly more negative estimate, which lends credence to the hypothesis of selective migration among immigrants based on changing local economic conditions within MSAs over time.

Shifting to the instrumental variables (IV) strategy described above, the first-stage F statistic of 28.7 substantially exceeds the weak-instruments critical value given in Stock et al. (2002). Hence, the enclave-based instrument does a very good job in predicting the low-skilled immigrant share of the working-age population in our sample of MSAs. The second-stage IV estimate of the effect of low-skilled immigration reported in the fourth column of Table 3 is roughly 2.5 times as large as the OLS estimate reported in the third column. This may be taken as confirmatory evidence that immigrant location decisions based on labor market fluctuations substantially biases the coefficient estimates obtained from OLS upward, although we cannot rule out attenuation bias arising from noisy measurement of the low-skilled immigrant population.[12] For the remainder of our analysis, we focus exclusively on instrumental variables estimates.

[12] As an additional specification check, we have re-estimated the models in Table 3 using the wages of college graduates 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 ($p = 0.90$). These results suggest that the immigrant flows allocated by the enclave-based instrument are fairly unresponsive to local economic conditions, as compared to the overall inflow of low-skilled immigrants.

Given the semi-log specification of the wage regression, the estimated coefficient represents the percentage change in the wages of childcare workers caused by a one percentage point increase in LSI. Between 1980 and 2000, the low-skill immigrant share of the working-age population rose from 6.5% to 10.6% in the average working-age native's MSA.[13] Based on the IV point estimate, this increase would result a 15.1% reduction in the median wage. As previously discussed, changes in the median wage might not be a relevant measure of 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 expect low-skilled immigrants to have the strongest impact. However, even at the 75th percentile our estimates imply that the average increase in LSI between 1980 and 2000 would cause an 8.8% reduction in the wages of childcare workers, which should translate into a lower cost of high-quality care.

The relative effects of low-skilled immigration on the wages of food services workers and housekeepers are slightly weaker at most quantiles, as seen in Panels B and C of Table 4. Even so, the estimated effects 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 1990; Card 200). 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 services sectors.[14] We may combine these numbers with the estimates of the effect of LSI on wages to compute a

[13] As described further in Appendix B, we compute weighted averages, rather than simple averages, to determine the change for the "representative" native's MSA working-age population. For ease of exposition, we will use "representative," "typical," and "average" synonymously. We discuss below the implications of treating the changes in outcomes associated with the average change in low-skilled immigration as counterfactual estimates.

[14] Workforce shares are essentially aggregated measures of underlying binary outcomes. For consistency with our other specifications we use a normit transformation to construct the dependent variable. Log or logit transformations result in similar relative effects.

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 0.38 to 0.90. The comparable ranges in food and housekeeping services are 2.13-2.94 and 2.75-3.14, 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 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.

4.2 Fertility and Work Decisions of High-Skilled Natives

We next consider the extent to which low-skilled immigration has affected the childbearing and labor supply decisions of high-skilled non-Hispanic native females. The first three columns of Table 5 display coefficient estimates based on the grouped bivariate probit model described above. We report only results from instrumental variables estimation, which theoretically isolates exogenous changes in the local population of low-skilled immigrants. (Consistent with the biases previously discussed, the unreported coefficient estimates obtained from OLS are all smaller in magnitude than the IV coefficients given in Table 5.) Our IV estimates 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 MSA rose from 6.6% to 10.5%. Based on the IV coefficient estimates and the formulas provided in Appendix B, a 3.9 percentage point increase in LSI implies a likelihood of childbearing that is 0.89 percentage points higher. This corresponds to roughly 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 -0.67 percentage points.

Taken together, our results so far suggest that high-skilled women in our sample of MSAs responded to immigrant-induced reductions in childrearing costs by exiting from the labor force to bear children. This pattern of behavior, along with a 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 4, low-skilled immigration also attenuated the negative correlation between childbearing and labor force participation, thereby reducing the extent of the tradeoffs women had to make.

In order to make the implications for role incompatibility more concrete, we estimate the effects of low-skilled immigration on the joint likelihood of work and fertility. Between 1980 and 2000, the joint likelihood in our sample of urban non-Hispanic college graduate women rose from 3.4% to 5.7%, an increase similar to that seen in Figure 2 above. Although there were a host of social and economic changes over that time frame that may have reduced role incompatibility, we may use our instrumental variables estimates in the first three columns of Table 5 to assess the contribution of immigration. Based on the formulas in Appendix B, the 3.9 percentage point increase in LSI observed in the average high-skilled woman's MSA would result in a 0.65 percentage point increase in the likelihood of bearing children while remaining in the labor force. Moreover, 53.5% 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.

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 also 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 potentially 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 MSA during our sample period led to significant and substantial increases in the joint likelihood of childbearing and labor force participation.

4.2.1 Specification Checks

The bivariate probit model is useful because it allows us to examine a variety of channels whereby low-skilled immigration might exert its effects on fertility and work decisions. However, the effect of LSI on the joint likelihood is calculated indirectly by combining three sets of coefficient estimates. As previously discussed, this indirect method might be unreliable because it relies so heavily on the statistical structure of the model. In the fourth column of Table 5, we use a simpler univariate model to estimate the effect of low-skilled immigration on the likelihood of bearing children and participating in the labor force. The resulting direct estimate of the change in LSI experienced by the average high-skilled woman on the joint likelihood is virtually identical to the indirect estimate, differing by only 0.001 percentage points. This suggests that the latent bivariate structure from which our model is derived does not substantially skew our estimates of the effects of low-skilled immigration.

Regardless of the estimation structure employed, a pervasive concern regarding the interpretation of our estimates arises due to 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 outcomes to fix ideas, but our discussion of the biases that might arise due to selection are also applicable to labor supply, the tetrachoric, and the joint likelihood of childbearing and labor force participation. From an economic or policy perspective, it is not problematic if our estimated effects of LSI are reflecting *changes* in fertility among either selected or unselected women. Even an increase in fertility among selected women represents an increase in childbearing that would not have taken place in the absence of low-skilled immigration, and so still represents a causal effect. However, if the baseline fertility rate among selected women differs from that among unselected women, then our estimated effects of LSI may simply be capturing changes in the composition of the sample.

Composition bias may occur if, for example, college-graduate women with a preference for large families move to metropolitan areas with a large low-skilled immigrant population in order to take advantage of cheaper market-provided household services. Alternatively, young women within an MSA may recognize that cheaper child care will better allow them to realize the labor market returns to education due to an increased ability to work after bearing children. At the margin, it is likely to be the women with the strongest fertility preferences who respond to low-skilled immigration, thereby entering our sample. Hence compositional change due to selective migration and college completion may bias the estimated effect of LSI upward, relative to the causal changes in childbearing rates that we hope to measure. 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 are able to test for composition bias due to selective enrollment directly by estimating the likelihood of school attendance among college graduates living in our sample of MSAs using an empirical model similar to equation (9). The estimated instrumental variables coefficient on LSI from the normit regression is insignificant ($p = 0.83$) and trivial in magnitude. Hence, this form of composition bias has a negligible impact on our results. 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 of college-educated women into an MSA and changes that occur due to selective college completion within the MSA.[15] However, we are able to test for both forms of selection jointly based on the following model:

$$\ln N_{gmt} = \gamma LSI_{mt} + \boldsymbol{\mu}_m + \boldsymbol{\mu}_{kt} + \boldsymbol{\mu}_g + \lambda IncControl_{mt} + \chi \ln T_{gmt} + \varepsilon_{gmt} \ , \qquad (10)$$

[15] 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 MSAs, making it impossible to determine the migration status of many women.

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-MSA-year cell. The model in equation (10) essentially tests for extranormal growth in the population of high-skilled females, relative to the overall population of same-age females. As described in Appendix B, the degree to which composition bias affects our selected-sample estimates is roughly proportional to the coefficient on LSI. Instrumental variables regression indicates that the *p* value associated with the IV estimate of γ is 0.54, which allows us to reject the null hypothesis of composition bias due to selective migration and college completion.[16]

Substituting the number of non-Hispanic natives who had *not* completed college in place of N_{gmt} in equation (10), however, yields a negative and highly significant ($p \approx 0.01$) 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 (Card 2001; 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.

5 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 2008). 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 sub-

[16] Even 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 4.27 more children than unselected women between the ages of 23 and 39 in 2000. This differential seems implausibly large given that the average rate of fertility among all college graduates was 1.54 children per woman.

stantial reductions in the cost of market-provided child care, food, and housekeeping services in large metropolitan areas. Based on a novel bivariate estimation method, 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 also 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.

The results in this paper may shed light some light on a puzzle in the literature analyzing cross-country differences in the relationship between female market work and childbearing. Although a negative relationship has been widely documented at the individual level, at the aggregate level the correlation between the two seems to be deteriorating, and this change is most pronounced in the United States (Ahn and Mira 2002; Engelhardt et al. 2004). Relative to other industrialized countries with comparable total fertility rates, female labor force participation is substantially higher in the United States. Similarly, the U.S. fertility rate is much higher than that in developed nations with comparable labor force participation rates (cf. Brewster and Rindfuss 2000). These patterns are especially intriguing given that government family policies are far less generous in the United States than in comparison countries (Gornick et al. 1997; Henneck 2003). Our finding that low-skilled immigration attenuated the negative relationship between fertility and market points to a partial explanation for this American "exceptionalism" (Preston and Hartnett 2008: 26) among non-Hispanic college educated women. Future research could potentially explore this conjecture using cross-country data.

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 more frequent information on the size of the foreign-born population. 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.

Appendix A - Description of Metro Sample

The general concept of a metropolitan area used by the U.S. Census Bureau has been fairly consistent over time: an urban center along with adjacent communities that have strong social and economic ties to the center. However, in practice the Census Bureau has introduced a number of changes in geographic definitions resulting in temporal inconsistencies in the population covered by metropolitan areas in the public use microdata sample (PUMS) files. First, to account for population growth, the Census Bureau has expanded geographic boundaries so that many metropolitan areas encompass more territory in later years. Re-definitions have also resulted in metropolitan areas being merged and geographic components of a given area being dropped or moved to another metropolitan area. Finally, the Census attempts to protect the identities of individuals in geographic areas with small populations. As the identifiable sampling units used to construct the PUMS files have changed, so has the effective spatial definition of a number of metropolitan areas.[17]

The temporal inconsistencies in Census geography make it difficult to construct MSA-level variables that are comparable across years. Shifting boundaries introduces noise into the aggregate measures. Moreover, temporally expanding boundaries will result in a systematic bias given that individuals living in far suburbs or exurban areas are likely to have different characteristics compared to those residing closer to city centers. In order to reduce the impact of these temporal inconsistencies on our estimates, we developed a set of geographic definitions that provide the most comparable coverage of metropolitan populations between 1980 and 2000. Because the more recent metropolitan area boundary definitions generally encompass the most territory, we use Census 2000 geography as a benchmark.

Metropolitan areas consist of free-standing metropolitan statistical areas (FMSAs) and of primary metropolitan statistical areas (PMSAs) that are part of larger consolidated metropolitan statistical areas (CMSAs). We attempted to treat PMSAs as distinct metropolitan areas where possible, but in some cases we needed to combine benchmark-2000 MSAs together in

[17] The smallest sampling unit publicly identifiable in 1980 census PUMS files is the county group. Public use microdata areas (PUMAs) were developed for the 1990 census to replace county groups as the sampling unit. The boundaries of 1990 PUMAs did not follow county lines and also differed from the boundaries of the PUMAs defined for the 2000 census.

order to account for historically different definitions or overlapping sampling units. To simplify exposition, we refer to FMSAs, PMSAs and combined PMSAs, collectively, as "metropolitan statistical areas" (MSAs). The bivariate estimator used to analyze childbearing and labor outcomes is not defined when outcomes are uniform within a group. Smaller populations increase the chances of observing groups with all zeros or all ones for binary outcomes when using the microdata samples. Consequently, we limited our initial universe to the 100 most populous MSAs, as measured in 1990. We use the midpoint of our sample rather than the year 2000 so that the set of MSAs is not skewed toward metropolitan areas that are either growing or shrinking.

Our procedure to construct temporally comparable MSAs relied on GIS map files provided by the IPUMS project. We define MSAs based on the set of sampling units that was most consistent with the benchmark geography. Thus, individuals living in communities that would eventually become part of the metropolitan area are counted as part of the MSA population in earlier years. If feasible, we used the larger 5% PUMS files so that the aggregate MSA and group measures would have lower sampling errors. However, for some MSAs the 1% "metro" PUMS files, available in 1980 and 1990, resulted in a better match to the benchmark geography.[18] In many cases it was not possible to adjust sampling units so that MSAs were completely consistent between 1980 and 2000. We include in our sample only those metropolitan areas in which the population covered by the underlying sampling units was always within 5% of the population defined by the benchmark-2000 geography. Out of the initial universe of 100 MSAs, 29 failed to meet this criterion. One additional MSA was excluded from the sample due to uniform fertility and/or work outcomes among college-graduate females.

Table A1 lists the metropolitan areas that were included and excluded from the final sample. As can be seen in Table A2, the 70 included metropolitan areas are substantially more populous than the 30 excluded MSAs; the former group had an average 1990 population of 1.8 million, while in the latter group this number was approximately 825,000. Almost 51% of the

[18] There is no comparable 1% metro PUMS file associated with the 2000 census data. For a number of MSAs, the PUMAs used for the 5% file do not correspond very closely to the boundaries actually defined by the Census, resulting in either truncation of the MSA to the core or substantial overcoverage from the inclusion of neighboring territories.

U.S. population lived in the included MSAs in 1990. However, these MSAs covered 79.3% of the working-age immigrant population reflecting the fact that immigration is largely an urban phenomenon. Low-skilled immigrants appear to have followed a similar pattern of location, given that 79.7% of immigrants with a high school degree or less resided in included MSAs. By contrast, the smaller MSAs excluded from the analysis only covered 3.8% of the low-skilled immigrant population. Moreover, of non-Hispanic college-graduate native females aged 23 to 39 living in the United State in 1990, 59% resided in the included MSAs. Hence, estimates based on our sample of 70 large metropolitan areas are likely to capture a substantial majority of the impact of low-skilled immigration on the target population of high-skilled native females.

Appendix B - Technical Notes

B.1 The Tetrachoric Correlation

A variety of statistics have been proposed to measure association among categorical data. Given the focus of this paper, we limit this technical discussion to the special case of association between two dichotomous variables, setting aside measures of association among multiple dichotomous variables (e.g. polychoric correlation) and between ordinal and continuous data (e.g. biserial and polyserial correlation). When the traditional Pearson product-moment correlation approach is applied to two binary variables the result is referred to as the "phi coefficient." The product-moment coefficient is but one of many association statistics that can be computed from a two-by-two table of observed outcomes. Warrens (2008) provides a review of many of these measures including the tetrachoric correlation, which assumes that the outcomes are dichotomous manifestations of latent normally-distributed continuous variables.

Following the notation of the bivariate model described in the main text, let C^*_{imt} and L^*_{imt} denote the latent desirability of childbearing and labor supply, respectively. Ignoring covariates, the bivariate model can be written as:

$$
\begin{aligned}
C_{imt} &= \mathbb{I}\left(C_{imt}^{*} > 0\right) \text{ with } C_{imt}^{*} = \mu^{C} + \varepsilon_{imt}^{C}, \\
L_{imt} &= \mathbb{I}\left(L_{imt}^{*} > 0\right) \text{ with } L_{imt}^{*} = \mu^{L} + \varepsilon_{imt}^{L},
\end{aligned}
$$

where, in order to make notation more compact, we use $I(\cdot)$ to denote the indicator function. If the error terms ε_{imt}^{C} and ε_{imt}^{L} follow standard normal distributions, then $\rho \equiv corr(\varepsilon_{imt}^{C}, \varepsilon_{imt}^{L})$ is the unconditional tetrachoric correlation between fertility and work. As noted in Greene (2007:820), ρ is "the correlation that would be measured between the underlying continuous variables if they could be observed." These equations can be easily adapted to account for groups of individuals, g, with common characteristics. Using the notation in the main text, the grouped model is:

$$
\begin{aligned}
C_{igmt} &= \mathbb{I}\left(C_{igmt}^{*} > 0\right) \text{ with } C_{imt}^{*} = \beta_1 LSI_{mt} + \boldsymbol{\omega}_1' \mathbf{v}_{gmt} + \varepsilon_{imt}^{C}, && (11) \\
L_{igmt} &= \mathbb{I}\left(L_{igmt}^{*} > 0\right) \text{ with } L_{imt}^{*} = \beta_2 LSI_{mt} + \boldsymbol{\omega}_2' \mathbf{v}_{gmt} + \varepsilon_{imt}^{L}. && (12)
\end{aligned}
$$

In this case the correlation between the error terms, $\rho_{gmt} \equiv corr\left(\varepsilon_{igmt}^{C}, \varepsilon_{igmt}^{L}\right)$ is conditional on group membership.

We use the tetrachoric correlation as the main measure of association in our paper for two reasons. First, there is a well known problem with the Pearson product-moment based phi coefficient. Namely, although it equals zero under statistical independence, the range of possible values for the phi coefficient is affected by the marginal frequencies of the underlying binary variables and does not necessarily span the interval $(-1, +1)$. Considering a broad array of binary association measures, Warrens (2008) finds five for which, in addition to taking a value of zero under statistical independence, the theoretical range is always $(-1, +1)$ and is independent of of the marginal distributions: the tetrachoric correlation, Yule's Q, Yule's Y, Digby's H, and a measure of ecological association, Cole's C_7. Moreover he notes that "[t]he latter four coefficients have been studied as approximations to the tetrachoric correlation" (Warrens 2008:787).

The second reason that we use the tetrachoric correlation is that, unlike the other measures discussed by Warren, it has a clear interpretation in the context of our statistical model of decision making. The bivariate normal distribution of the latent error terms implies that:

$$\begin{aligned} E\left[C^{*}_{igmt}\middle|\,\varepsilon^{L}_{igmt}\right] &= E\left[C^{*}_{igmt}\right]+\rho_{gmt}\varepsilon^{L}_{igmt}\ , \\ \text{and}\ \ E\left[L^{*}_{igmt}\middle|\,\varepsilon^{C}_{igmt}\right] &= E\left[L^{*}_{igmt}\right]+\rho_{gmt}\varepsilon^{C}_{igmt}\ . \end{aligned}$$

Based on these two equations the conditional tetrachoric correlation measures the extent to which shocks to desired fertility translate into changes in desired labor force participation, and vice-versa – we give specific examples in the main text.

The structure of the model implies the following relationship between the true tetrachoric correlation and expected rates of childbearing (π^{C}), labor force participation (π^{L}), and jointly bearing children and working (π^{CL}) among the women in group g in metro area m in year t:

$$\pi^{CL}_{gmt}=F\left(\Phi^{-1}\left(\pi^{C}_{gmt}\right),\Phi^{-1}\left(\pi^{L}_{gmt}\right),\rho_{gmt}\right)\equiv G\left(\pi^{C}_{gmt},\pi^{L}_{gmt},\rho_{gmt}\right) \tag{13}$$

where $F\left(\cdot\right)$ denotes the standard bivariate normal distribution function. Given that $F\left(\cdot\right)$ is monotonically increasing in the tetrachoric correlation (Tihansky 1972), a higher value of ρ will, *ceteris paribus*, translate into a higher joint likelihood. Using the observed proportions (p^{C}, p^{L}, and p^{CL}) as analogues of the expected values in equation (13) allows us to calculate the empirical tetrachoric correlation, r_{gmt}, based on the sample of outcomes. There is no closed-form solution for r_{gmt}. However, since $F\left(\cdot\right)$ is monotonic in the third argument, we can apply a recursive binary chop algorithm to search for the value of r_{gmt} that solves

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

The parameter ξ represents a pre-defined level of precision, which we set to 2^{-50}.

B.2 Estimated Effects of Low-Skilled Immigration

In our analysis of household services markets, the coefficients on LSI in the semi-log wage regressions are approximately equal to the average percentage change in the wage associated with a one percentage point change in the low-skilled immigrant share of the working age population. In order to make the coefficient estimates more meaningful, we simply scaled them by $\bar{\Delta}^n_{LSI} \equiv \sum_m (n_{m,2000} LSI_{m,2000} - n_{m,1980} LSI_{m,1980}) \times 100$, where $n_{m,t}$ denotes the share of all U.S.-born adults ages 20-64 represented (based on Census person weights) in metropolitan area m at time t. The scaling factor $\bar{\Delta}^n_{LSI}$ represents the percentage point change in LSI experienced by the average native working-age individual between 1980 and 2000.

It is slightly less straightforward to calculate the effect of low-skilled immigration when analyzing grouped proportions data. Considering fertility first, the expected rate of childbearing is univariate probit based on equation (11). That is, $\pi^C_{gmt} = \Phi(\beta_1 LSI_{mt} + \boldsymbol{\omega}_1' \mathbf{v}_{gmt})$, where $\Phi(\cdot)$ is the standard normal cumulative distribution function. Consequently, the estimated average marginal effect of low-skilled immigration on the likelihood of childbearing among women in our sample is

$$\overline{ME}^C_{LSI} = \sum_{gmt} h_{gmt} \frac{d\hat{\pi}^C_{gmt}}{dLSI} = \sum_{gmt} h_{gmt} \phi\left(\hat{\beta}_1 LSI_{mt} + \hat{\boldsymbol{\omega}}_1' \mathbf{v}_{gmt}\right) \hat{\beta}_1 \,. \tag{14}$$

In equation (14) h_{gmt} denotes the share of high-skilled women represented by group g in metropolitan area m in year t, $\phi(\cdot)$ is the standard normal probability density function, and the hats on the coefficients indicate that we are using the parameter estimates. The marginal effect represents the impact of a one unit change in LSI on the probability of child-bearing. An equivalent expression can be written for the effect of low-skilled immigration on the likelihood of labor force participation. We scale equation (14) by $\bar{\Delta}^h_{LSI} = \sum_{gm,t} (h_{gm,2000} LSI_{m,2000} - h_{gm,1980} LSI_{m,1980}) \times 100$ to determine the percentage point effect of low-skilled immigration between 1980 and 2000 in the MSA of the representative high-skilled woman in our sample. The effects of LSI on the share of the representative workforce occupied in child care, housekeeping, and food services can be calculated similarly with the estimated share of high-skilled natives, h, replaced by the estimated share of all working-age natives, n. In Table 3, we additionally normalize the marginal effect by the observed average occupational concentration in 2000 so that it may be expressed as a percentage.

The bivariate probit model also allows us to determine the effect of LSI on the joint likelihood of fertility and labor force participation in an analogous manner. Specifically, our parameterization of the tetrachoric correlation, $\rho_{gmt} = \beta_3 LSI_{mt} + \omega_3' v_{gmt}$, in conjunction with equations (11)-(13) imply that the expected likelihood of both bearing children and working can be written as

$$\pi_{gmt}^{CL} = F\left(\beta_1 LSI_{mt} + \boldsymbol{\omega}_1' \mathbf{v}_{gmt}, \beta_2 LSI_{mt} + \boldsymbol{\omega}_2' \mathbf{v}_{gmt}, \beta_3 LSI_{mt} + \boldsymbol{\omega}_3' \mathbf{v}_{gmt}\right) \ . \tag{15}$$

The average marginal effect is $\overline{ME}_{LSI}^{CL} = \sum_{gmt} h_{gmt}(d\hat{\pi}_{gmt}/dLSI)$ with $\hat{\pi}_{gmt}^{CL}$ denoting the expected joint likelihood evaluated using the estimated coefficients in place of the true parameters in equation (15). The average marginal effect of immigration can be decomposed as

$$\overline{ME}_{LSI}^{CL} = \left[A_1\hat{\beta}_1 + A_2\hat{\beta}_2\right] + A_3\hat{\beta}_3 \ , \tag{16}$$

where $A_j = \sum_{gmt} h_{gmt} F_j\left(\hat{\beta}_1 LSI_{mt} + \hat{\boldsymbol{\omega}}_1' \mathbf{v}_{gmt}, \hat{\beta}_2 LSI_{mt} + \hat{\boldsymbol{\omega}}_2' \mathbf{v}_{gmt}, \hat{\beta}_3 LSI_{mt} + \hat{\boldsymbol{\omega}}_3' \mathbf{v}_{gmt}\right)$, and $F_j(\cdot)$ denotes the jth partial derivative of the standard bivariate normal distribution function. The two terms inside the brackets in equation (16) represent the average change in the joint likelihood arising due to 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. Based on the exposition above, this can be interpreted as the effect of LSI on the joint likelihood arising from a weakened link between childbearing and labor supply. As above, we may scale the marginal effect and its components by $\bar{\Delta}_{LSI}^{h}$ in order to determine the percentage point impact of low-skilled immigration on the joint likelihood of childbearing and LFP for the representative high-skilled native woman.

The decomposition in equation (16) provides insight into the mechanisms whereby low-skilled immigration affects the joint likelihood. However, for the reasons described in the text, this may not be accurate due to the indirect method employed. Hence we also provide a direct estimate of the effect that is obtained when the joint likelihood itself is treated as a binary dependent variable (equation (9) in the main text). The formula used to calculate the "direct" average marginal effect of LSI on the joint likelihood is equivalent to equation (14) with $\hat{\beta}_4$ and $\hat{\omega}_4$ substituted in place of $\hat{\beta}_1$ and $\hat{\omega}_1$.

B.3 Sample Selection

As noted in the main text, the effects of low-skilled immigration that we estimate may be biased by sample selection. In particular, if female education and migration decisions are affected by low-skilled immigration, then our results may be attributable to changes in the composition of the estimation sample. We derive below formulas that show how selection affects our results and derive an equation that allows us to test whether there is statistical evidence of sample selection.

Dropping subscripts for notational convenience, define $S\left(LSI\right)$ as the number of women who entered (or left) the sample in response to low-skilled immigration ("selected women") and U as the number women who would have been in the sample regardless of the extent of immigration ("unselected women"). Then the total number of high-skilled women in the estimation sample is $N\left(LSI\right) = S\left(LSI\right) + U$. Using these definitions, the likelihood of bearing children can be written as

$$\pi_N^C\left(LSI\right) = q\left(LSI\right) \times \pi_S^C\left(LSI\right) + \left[1 - q\left(LSI\right)\right] \times \pi_U^C\left(LSI\right), \tag{17}$$

where $q\left(LSI\right) = S\left(LSI\right)/N\left(LSI\right)$. Similar expressions can be written for the likelihood of labor force participation and joint likelihood. Based on equation (17), the marginal effect of LSI on the likelihood of childbearing in the estimation sample is

$$ME_N^C = \left[q \times ME_S^C + (1-q) \times ME_U^C\right] + \left[\frac{dq}{dLSI} \times \left(\pi_S^C - \pi_U^C\right)\right]. \tag{18}$$

The first bracketed term in equation (18) measures the (weighted) average of the marginal effects of low-skilled immigration on outcomes across sub-groups of women. Particularly interesting is ME_U^C, which represents the marginal effect of LSI on the unselected women. However, the average marginal effect across sub-groups also represents a causal impact of immigration. Even if the effect were entirely concentrated on selected women, it would still represent an increase in the likelihood of fertility among women, relative to a counterfactual scenario where there was no increase in low-skilled immigration at all.

The main problem for interpreting ME_N^C as a causal effect arises from the second bracketed term in equation (18), which captures the direct effect of the change in sample composition due to low-skilled immigration. It is straightforward to show that

$$\frac{dq}{dLSI} = \gamma \times (1 - q),$$

where $\gamma \equiv d \ln N / dLSI$. Hence the direct composition effect of selection is proportional to LSI's effects on the natural logarithm of the total number of women that meet the sample inclusion criteria. Based on the discussion in the main text, it is plausible that when there is immigration-based selection into the sample ($\gamma > 0$) the selected women exhibit a higher rate of childbearing than unselected women ($\pi_S^C > \pi_U^C$). If this is the case, then the estimated effect of low-skilled immigration will be biased upward, relative to the average causal effect across both types. We can calculate the extent of composition bias based on the following data: (a) an estimate of γ, which can be additionally used to generate an estimate of q, and (b) an assumed difference in the fertility rates of selected and unselected women ($\pi_S^C - \pi_U^C$).

Composition bias is not present, however, when $\partial \ln N / \partial LSI = 0$, a hypothesis that we may statistically test based on equation (10) in the main text. Moreover, the absence of composition bias implies that $S(LSI) = 0$ and, consequently, that that $q(LSI) = 0$. When this is the case, the weighted average of the marginal effects in the first bracketed term in equation (18) collapses to the marginal effect of LSI on a group of unselected women.

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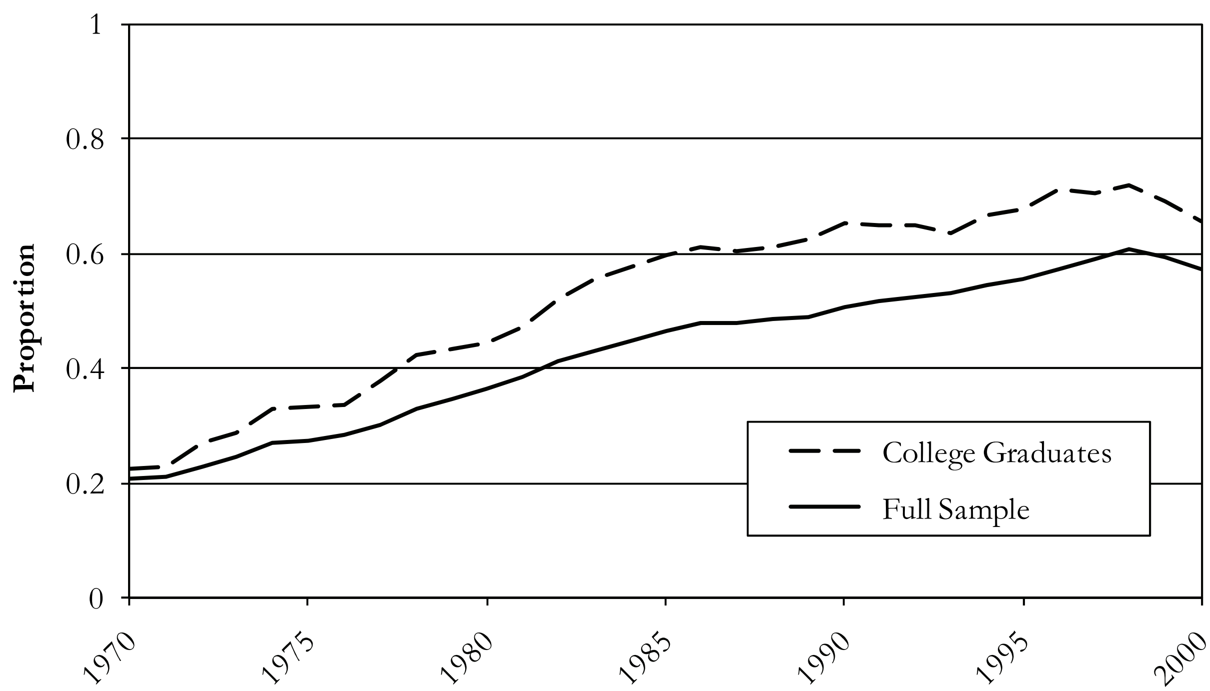
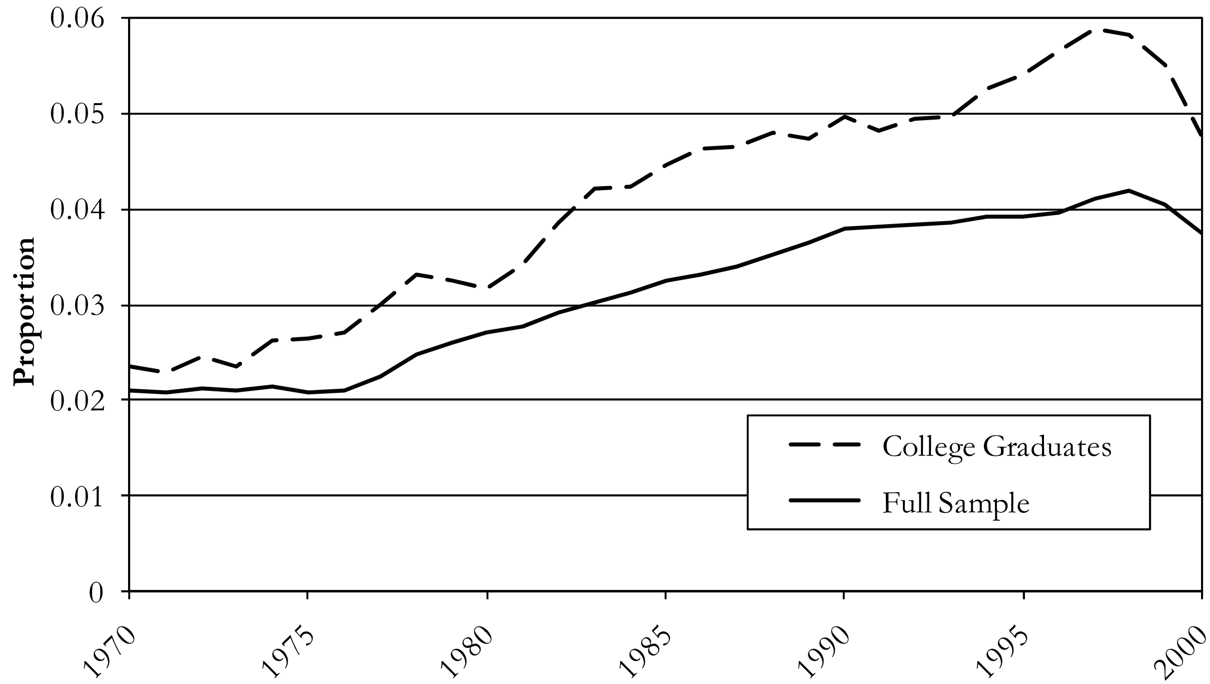
Figure 1. Proportion of Recent Mothers Participating in the Labor Force



Note: See text for sample and variable definitions.

Figure 2. Proportion of Women Bearing Children and Participating in the Labor Force



Note: See text for sample and variable definitions.

Table 1. Immigrant Concentration and Educational Distribution in Household Service Sectors, 2000

Characteristic	Child Care	Food Services	Housekeeping	All Other Occupations
Foreign-born	13.4%	18.4%	32.5%	13.0%
Native	86.6%	81.6%	67.5%	87.0%
Completed years of schooling, all workers:				
No post-secondary	53.0%	69.8%	84.7%	40.9%
Some post-secondary	37.3%	25.8%	12.7%	32.2%
Bachelors equivalent or higher	9.7%	4.4%	2.6%	26.9%
Completed years of schooling, immigrants:				
No post-secondary	64.5%	77.7%	86.2%	49.6%
Some post-secondary	24.6%	16.3%	9.7%	22.4%
Bachelors equivalent or higher	10.9%	6.0%	4.1%	28.0%

Note: Occupations are based on the consistent 1990-basis classification developed by the IPUMS group (Ruggles et al. 2004). Educational attainment is based on the consistent recode developed by the IPUMS group to bridge the 1980 and 1990 censuses.

Source: 1970 and 2000 Census public-use micro-data files (Ruggles et al. 2004).

Table 2. Characteristics of the Working-Age Population for 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	343,562	69.8%
Ecuador	26,300	213,645	128,509	62.0%
Haiti	21,125	317,459	186,415	59.3%
Mexico	549,125	6,689,664	5,691,821	85.7%
Portugal	70,400	97,869	60,729	75.6%

Note: The “working-age” and “low-skilled” segments of the population are as defined in the text.

Source: 1970 and 2000 Census public-use micro-data files (Ruggles et al. 2004).

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

	Ordinary Least Squares			Instrumental
Variable	(1)	(2)	(3)	Variables
LSI	0.684***	-1.037**	-1.282**	-3.297***
	(0.185)	(0.375)	(0.377)	(0.385)
IncControl			0.393*	0.593**
			(0.180)	(0.203)
MSA fixed effects	No	Yes	Yes	Yes
Region-Year fixed effects	Yes	Yes	Yes	Yes
R^2	0.93	0.97	0.97	
First-stage F statistic				28.70

Notes: Occupation is based on the consistent 1990-basis classification developed by the IPUMS group (Ruggles et al. 2004). Each column represents a different specification applied to 1980, 1990, and 2000 data from the 70 MSAs listed in Appendix A ($N = 210$). Median wages are calculated using only natives and immigrants living in the United States for at least 10 years. The explanatory variables and the instrumental variable are defined in the text. All specifications include time-varying region fixed effects. Each MSA-year observation is weighted by the number of childcare workers used to construct the wage measure to account for heteroskedasticity, and the robust standard errors in parentheses are clustered by MSA.

Source: 1970, 1980, 1990, and 2000 Census public-use micro-data files (Ruggles et al. 2004) and authors' calculations.

*p < .05; **p < .01; ***p < .001

Table 4. Instrumental Variables Estimates of the Effects of Low-Skilled Immigration on Household Services Markets

	Log Wage, Natives and Non-Recent Immigrants			Normit(Share of Labor Force)
Estimates by sector	25th Percentile	50th Percentile	75th Percentile	
(A) Child Care				
Coefficient on *LSI*	-5.030***	-3.297***	-2.140***	0.868***
	(0.653)	(0.386)	(0.358)	(0.322)
Effect of average change in *LSI*, 1980-2000, on the underlying dependent variable[a]	-20.62%	-13.52%	-8.77%	7.86%
(B) Food Services				
Coefficient on *LSI*	-1.400***	-1.301**	-1.013**	1.493***
	(0.388)	(0.381)	(0.302)	(0.234)
Effect of average change in *LSI*, 1980-2000, on the underlying dependent variable[a]	-5.74%	-5.34%	-4.15%	12.21%
(C) Housekeeping Services				
Coefficient on *LSI*	-2.743***	-2.399**	-2.681***	3.113***
	(0.794)	(0.712)	(0.727)	(0.634)
Effect of average change in *LSI*, 1980-2000, on the underlying dependent variable[a]	-11.25%	-9.84%	-10.99%	30.91%

Notes: Occupations are based on the consistent 1990-basis classification developed by the IPUMS group (Ruggles et al. 2004). Each coefficient estimate comes from a separate instrumental variables equation and is based on 1980, 1990, and 2000 data from the 70 MSAs listed in Appendix A. “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 defined in the text. Each equation includes time-varying region fixed effects, MSA fixed effects, and income per working-age male college graduate. Within each sector, 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 MSA-year cell ($N = 210$) is weighted by the number of individuals used to calculate the associated dependent variable, and the robust standard errors in parentheses are clustered by MSA.

Source: 1970, 1980, 1990, and 2000 Census public-use micro-data files (Ruggles et al. 2004) and authors’ calculations.

[a] Reported effects are based on the change in *LSI* experienced by the average working-age native in the sample and are calculated using the formulas in Appendix B.

**p < .01; ***p < .001

Table 5. Instrumental Variables Estimates of the Effects of Low-Skilled Immigration on the Fertility and Labor Force Participation Outcomes of Non-Hispanic College-Graduate Natives

	Normit, Fertility Rate	Normit, LFP Rate	Tetrachoric, Fertility and LFP	Normit, Joint Rate of Fertility and LFP
(A) Coefficient Estimates				
LSI	1.467**	-0.732*	0.845**	1.630*
	(0.509)	(0.324)	(0.307)	(0.661)
IncControl	0.223	-0.196**	-0.299***	0.038
	(0.187)	(0.065)	(0.079)	(0.180)
Share of women: black	1.079*	0.344†	-0.590†	0.927†
	(0.477)	(0.195)	(0.309)	(0.487)
Share of women: other non-white	-4.155***	-2.067***	1.580**	-4.188***
	(0.954)	(0.510)	(0.520)	(1.107)
Share of women: married	2.227***	-1.254***	0.456***	2.089***
	(0.178)	(0.111)	(0.114)	(0.191)
(B) Underlying Dependent Variable[a]				
Mean, 2000	0.0906	0.8425	-0.3993	0.0570
Effect of average change in *LSI* , 1980-2000	0.0089	-0.0067	0.0329	0.0065

Notes: Each column presents estimates from a separate instrumental variables equation and is based on 1980, 1990, and 2000 data from the 70 MSAs listed in Appendix A. The data consists of group-level 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, MSA fixed effects, and age-group fixed effects. Each of the MSA-year-group cells ($N = 420$) is weighted by the number of individuals used to calculate the grouped dependent variable, and the robust standard errors in parentheses are clustered by MSA.

Source: 1970, 1980, 1990, and 2000 Census public-use micro-data files (Ruggles et al. 2004) and authors' calculations.

[a] Characteristics of the underlying dependent variable and the average change in *LSI* are provided for the average college-graduate non-Hispanic native female in the sample. Reported effects of the average change in *LSI* are calculated using the formulas in Appendix B.

†p < .10; *p < .05; **p < .01; ***p < .001

Appendix Table A1. Metropolitan Areas Considered for Inclusion

(A) Metropolitan Areas Included in the Final Sample

Akron, OH; Austin, TX; Bakersfield, CA; Baltimore, MD; Baton Rouge, LA; Bergen-Passaic, NJ; Birmingham, AL; Boston-Lawrence-Lowell-Brockton, MA[a]; Bridgeport-Danbury-Stamford-Norwalk, CT[a]; Buffalo-Niagara Falls, NY; Charlotte-Gastonia-Rock Hill, NC/SC; Chicago, IL; Cleveland, OH; Columbia, SC; Columbus, OH; Dallas-Fort Worth-Arlington, TX[a]; Denver-Boulder-Longmont, CO[a]; Detroit, MI; El Paso, TX; Fort Lauderdale-Hollywood-Pompano Beach, FL; Fresno, CA; Gary-Hammond, IN; Grand Rapids, MI; Greensboro-Winston Salem-High Point, NC; Harrisburg-Lebanon-Carlisle, PA; Houston-Brazoria, TX[a]; Indianapolis, IN; Jersey City, NJ; Lancaster, PA; Lansing-East Lansing, MI; Las Vegas, NV; Little Rock-North Little Rock, AR; Los Angeles-Long Beach, CA; Miami-Hialeah, FL; Middlesex-Somerset-Hunterdon, NJ; Milwaukee-Waukesha, WI; Minneapolis-St. Paul, MN; Monmouth-Ocean, NJ; Nashville, TN; Nassau-Suffolk, NY; New Orleans, LA; New York, NY ; Newark, NJ; Norfolk-Virginia Beach-Newport News, VA; Orange County, CA; Orlando, FL; Philadelphia, PA/NJ; Phoenix-Mesa, AZ; Pittsburgh, PA; Portland-Vancouver, OR/WA; Raleigh-Durham, NC; Richmond-Petersburg, VA; Riverside-San Bernardino, CA; Sacramento-Yolo, CA[a]; Salt Lake City-Ogden, UT; San Antonio, TX; San Diego, CA; San Francisco-Oakland, CA[a]; San Jose, CA; Seattle-Everett, WA; St. Louis, MO/IL; Stockton, CA; Tacoma, WA; Tampa-St. Petersburg-Clearwater, FL; Tucson, AZ; Vallejo-Fairfield-Napa, CA; Ventura, CA; Washington, DC/MD/VA; West Palm Beach-Boca Raton, FL; Youngstown-Warren, OH.

(B) Metropolitan Areas Excluded from the Final Sample

Albany-Schenectady-Troy, NY; Albuquerque, NM; Allentown-Bethlehem-Easton, PA; Atlanta, GA; Charleston, SC; Cincinnati, OH-KY-IN; Dayton-Springfield, OH; Fort Wayne, IN; Greenville-Spartanburg, SC; Hartford, CT; Jacksonville, FL; Johnson City-Kingsport-Bristol, TN/VA; Kansas City, MO/KS; Knoxville, TN; Louisville, KY/IN; Memphis, TN/AR/MS; Mobile, AL[b]; New Haven-Meriden-Waterbury, CT[a]; Oklahoma City, OK; Omaha, NE/IA; Providence-Fall River-Pawtucket, RI/MA; Rochester, NY; Scranton-Wilkes-Barre-Hazleton, PA; Springfield, MA; Syracuse, NY; Toledo, OH; Tulsa, OK; Wichita, KS; Wilmington, DE/MD/NJ; Worcester, MA

[a] Indicates that the metropolitan area was formed by combining two or more Census-defined PMSAs or MSAs.

[b] Mobile, AL was excluded because no non-Hispanic native college graduates worked and participated in the labor force in 1990. The remaining metropolitan areas were excluded based on geographic inconsistencies.

Appendix Table A2. Population Characteristics in Included and Excluded MSAs, 1990

Variable	Included MSAs (N=70)	Excluded MSAs (N=30)	United States
Total population	126,007,523	24,743,766	248,107,630
Working-age population	76,063,254	14,724,786	145,978,410
Number of immigrants	12,127,065	691,189	15,290,363
Number of low-skilled immigrants	6,848,107	325,113	8,593,832
Number of potential sample females[a]	4,229,831	829,708	7,151,190

Source: 1970, 1980, 1990, and 2000 Census public-use micro-data files (Ruggles et al. 2004) and authors' calculations.

[a] "Potential sample females" refers to non-Hispanic female college graduates between the ages of 23 and 39 who were natives of the United States.