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Revisiting the Labor Demand Curve

The Wage Effects of Immigration and Women's Entry into the US Labor Force, 1960–2010

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INTERNATIONAL FOOD POLICY RESEARCH INSTITUTE

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Contents

Tables

Figures

ABSTRACT

The debate over the wage effects of immigration for native workers is an old one. One side of the debate claims that immigration has little if any negative impact on wages among natives, whereas others suggest that immigration has large, negative effects on native wages. On the latter side of the debate, many point to the work of Borjas (2003), who takes a national view of the US economy and estimates a wage elasticity of -0.4 with respect to immigration. In this paper, we replicate and update Borjas with the 2010 US census data, and use the method to study an even larger, concurrent labor supply shock, namely the entry of women into the labor force. We both find a much lower wage elasticity than Borjas to immigration (-0.2) and estimate a positive, statistically significant relationship between men's wages and women's entry into education-experience cells when wages are annualized. We take this evidence to suggest that the Borjas model is misspecified as it inadequately specifies substitution between immigrants and natives, and inadequately controls for structural change in the US economy.

Keywords: immigration, labor force, women, United States

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1. INTRODUCTION

The debate over the wage effects of immigration for native workers in the United States is an old one. The 1997 report *The New Americans* by the National Research Council (Smith and Edmonston 1997), which gave a comprehensive analysis of the economic, fiscal, and demographic consequences of immigration, concluded that immigration had a small negative effect on the wages of competing native workers, reducing them by 1 or 2 percent for every 10 percent increase in population of immigrant workers (Murray, Batalova, and Fix 2006, 3). However, while the actual effect measured by different researchers since that time has varied widely from slightly positive to strongly negative, the methods used to analyze this effect have also evolved and diversified. In recent debate about immigration reform, opponents have argued, among other hypothetical costs of immigration, that additional immigration would depress wages among natives (Rector and Richwine 2013).

Empirical analysis of the impact of immigration on native wages in the United States is both challenging and revealing given that various, complex factors affect this calculation. Borjas in particular has studied the economic impacts of immigration in the United States since the early 1990s and has argued that treating immigration as an increase in national labor supply is the most accurate method of analysis (Borjas, Freeman, and Katz 1996). Borjas (2003) analyzed the national labor force as a whole from 1960 to 2000 by aggregating into groups (or "cells") all male workers with the same general combination of education and experience, and concluded that a 10 percent increase in the number of immigrants in a skill group was associated with a 3 to 4 percent decline in wages. This estimate was substantially larger than previous estimates, especially considering that this elasticity applies to the entire workforce, in aggregate, and not just the lowest skilled.

A major critique of the Borjas model is that within education-experience cells, it implicitly assumes perfect substitutability between native and immigrant workers. Ottaviano and Peri (2012) relax this assumption and estimate the elasticity of substitution between natives and immigrants, and using a nested constant elasticity of substitution framework they estimate the elasticity of immigration on wages. Even among high school dropouts, they find the elasticity is much lower, attributing the low estimate to immigrants' tendency to enter job sectors that are different than those in which natives work. Studying immigration to the United Kingdom, Dustmann, Frattini, and Preston (2005) relax the assumption that immigrants should be preassigned to specific skill categories, and study the impacts of immigration on the distribution of wages. They find that immigration slightly depresses wages among the bottom 20th percentile of the distribution but increases wages at the upper end of the distribution.

In this paper, we advance an alternative critique of the Borjas (2003) model. We argue that the model is inappropriate to study labor demand because it fails to account fully for simultaneity bias. The bias develops because the model cannot account for unobserved factors changing labor supply. We demonstrate simultaneity bias through a replication of Borjas (2003). If Borjas's model accurately represents the impact of supply shocks on men's wages, the model should generate an unbiased estimate of the elasticity of men's wages to any labor force supply shock. After providing a direct replication, we update the dataset to include the 2010 census data, and re-estimate the model.

However, the supply shock of immigration into the US labor force was not the largest supply shock occurring between 1960 and 2010. The entry of women into the labor force was larger in magnitude than that of immigrants, and in fact Edwards and Lange (2013) name women's entry into the labor force as one of the three most important changes in the US labor force over that period, the others being increasing returns to education and educational upgrading. Acemoglu, Autor, and Lyle (2004, 56) call female labor force participation "one of the most profound labor market transformations of the past century". If Borjas's model is appropriate for estimating the labor demand curve for immigrants, it should also be appropriate for estimating the effect of women entering the US labor force since 1960 on men's wages. Moreover, it could be that women act as an omitted variable in the Borjas model, and as a result the effect of immigration on men's wages is overstated.

We therefore explore whether the concurrent supply shock of women entering the labor force affects the elasticity of men's wages to immigration. A concern with including women in labor models is that women are far more likely than men to refrain from labor force participation or to move in and out of the labor force. To account for the irregular pattern of women's labor force participation, we adjust the work experience of women using a calculation for the average difference between potential and actual work experience originally presented by Regan and Oaxaca (2009) and average fertility levels from the Centers for Disease Control and Prevention (CDC) to estimate actual average experience among women at specific ages.

We find that in fact the two supply shocks are not correlated with one another, once we control for differences in education, experience, and year fixed effects. However, rather than demonstrating that the entry of women into education-experience cells also depresses men's wages, we estimate insignificant coefficients on weekly wages among men and positive, statistically significant coefficients on annual wages. Moreover, considering immigration, when we add 2010 data to Borjas's model, we estimate lower coefficients than Borjas and elasticities that are more in line with the rest of the literature, despite an acceleration of migration between 2000 and 2010. Together, these results suggest that Borjas's method is affected by omitted variable bias, and does not generate unbiased estimates of the wage elasticities of supply shocks.

The paper proceeds as follows. First, we describe the literature on the wage effects of immigration in more detail. Second, we describe the data used in the analysis, and third, we describe the conceptual and empirical model. The fourth section provides descriptive results and the fifth section regression results. The final sections discuss explanations for the results in more detail and suggest potential avenues for further research.

2. LITERATURE REVIEW

A large literature exists concerning the effects of immigration on native wages in the United States.1 Serious study on this topic began in the early 1980s, increased throughout the 1990s and 2000s, and only recently declined somewhat in volume. An early study by Altonji and Card (1991) used 1970 and 1980 US census data for 120 major US cities to measure the effects of immigration on weekly wages of less-skilled natives. Their results vary by native subgroup, but show that overall a 10 percent increase in the metro population due to immigration reduced weekly earnings of low-skilled natives by about 1.2 percent. In a more influential study around the same time, Card (1990) famously examined the labor force outcomes of the Mariel boatlift episode in which Fidel Castro's temporary suspension of emigration restrictions permitted more than 125,000 mostly low-skilled Cubans to leave the country for the United States, where the vast majority settled in Miami. Using the resulting shock to the Miami labor market as an opportunity for a natural experiment, Card found no effect on wages or unemployment rates of either Cubans or non-Cubans. It appeared that Miami's labor market had absorbed the sudden influx of labor through industries that adjusted their production methods toward more labor-intensive technologies (Okkerse 2008).

Another early but influential publication was the National Research Council's 1997 report titled *The New Americans* (see Smith and Edmonston 1997) but known simply as the "National Academy Report" (Murray, Batalova, and Fix 2006, 3). For the first time it offered a comprehensive review of research on the economic, demographic, and fiscal impacts of immigration in the United States, both for immigrants and natives. On the topic of wages, it concluded that a 10 percent increase in the size of the labor force reduces wages of competing workers by 3 percent. However, it emphasized that there are overall wage gains for natives because most are in skill levels that complement immigrant labor. These conclusions have generally persisted as the consensus on the wage effects of immigration, though various studies have since produced disparate findings.

As Okkerse (2008) explains, the different methods of estimating the wage effect of immigration over the past three decades have generally fallen into one of two broad categories: simulation-based analyses that use existing theoretical models, or empirical studies that use survey data to estimate effects.[2] Simulation-based analysis can be further divided into partial equilibrium and computable general equilibrium approaches. The partial equilibrium (or "factor proportions") approach can be summarized as comparing "a nation's actual supplies of workers in particular skills groups to those it would have had in the absence of immigration and then using outside information on the elasticities of substitution among skill groups to compute the relative wage consequences of the supply shock" (Borjas 1999, 1753, quoted in Okkerse 2008). The general equilibrium approach is similar to this, but accounts for endogenously determined relative prices and quantities. It has the benefit of allowing capital stock to be flexible in the form of output mixes or technology adjustments (Okkerse 2008, 5).

Early empirical analyses, meanwhile, initially focused on the city or state as the unit of analysis. As with Altonji and Card (1991), these studies used fixed effects or similar models to compare immigration and wage levels across time in different cities in order to determine a causal effect. However, these studies suffer from the endogeneity (or, in this case, reverse causation) problem that occurs when immigrants choose a city specifically for its labor market conditions; for example, immigrants may choose a city for its relatively high wages. Researchers have typically used an instrumental variable approach to deal with this problem (see Altonji and Card 1991; Pischke and Velling 1997). However,

[1] Indeed, many literature reviews have been published on this topic. This review will rely heavily on Okkerse (2008), both for content and structure, since that review was unique in highlighting both the results of each study and the methods used. This review will also be restricted to studies of the United States. This should not be limiting, however, as most studies of the impact of immigration on labor market outcomes in Europe focus on unemployment effects, rather than wages, since wage variability is much lower in Europe than the United States (Okkerse 2008).

[2] Note that the studies we list under each method of analysis may not be mutually exclusive vis-à-vis their use of one method or another. Borjas (2003), for example, first uses a national-level empirical model and then uses elasticities found therein to estimate a partial equilibrium analysis.

instruments are almost always susceptible to criticism, and no perfect instrument has been found for immigration, particularly in these contexts (Okkerse 2008). A second concern with these models is internal outmigration by natives from the affected area, which can offset the negative wage effects of immigration, a possibility that is difficult to control for. Some researchers have argued that this phenomenon is not very prevalent (see Card 2001 and 2004; Butcher 1998; Kritz and Gurak 2001), while others have argued the opposite (see Filer 1992; White and Liang 1998; Frey 1995). Regardless, it remains a major weakness of the area analysis approach (Okkerse 2008).

Citing both the problems with area-specific empirical analysis mentioned above, Borjas et al. (1996) argued in favor of factor analysis, which considers immigration as an increase in the national labor supply. Along these lines, subsequent empirical studies have used national-level industries (for example, De New and Zimmerman 1994; Mühleisen and Zimmerman 1994), occupations (for example, Camarota 1998; Card 2001; Orrenius and Zavodny 2007), and skill groups (for example, Borjas 2003) as the unit of analysis. As later studies began to acknowledge that it was more realistic to distinguish different labor inputs along a skill dimension, the latter unit of analysis has received the majority of scholarly attention. Dustmann and Fabbri (2005) used educational attainment, while Card (2001) used occupation to define skill (Dustmann, Glitz, and Frattini 2008, 479). Borjas (2003) was the first to advocate for a combination of experience and education, followed by Aydemir and Borjas (2011). Some studies have gone a step further, relaxing the assumption of perfect substitutability of immigrants and natives (within any skill category) using nested production technologies (for example, Manacorda, Manning, and Wadsworth 2012; Ottaviano and Peri 2012).

Among these studies, Borjas (2003) represents the most significant departure from the typical findings on the native wage effects of immigration. The study shows an immigration elasticity of native wages of -0.38, meaning a 10 percent increase in immigration leads to a 3 to 4 percent decline in native wages.[3] Thus, between 1980 and 2000, immigrant influxes caused average wages among natives to decline by around 3.2 percent.

[3] This estimate uses the coefficient on immigrant share calculated by Borjas (2003) using the 2000 census data in place of pooled 1999–2001 Current Population Survey (CPS) data, which is reported in the paper's footnotes (page 1349). It thus differs slightly from the elasticity of -0.40 stated in the body of the paper.

3. DATA AND VARIABLES

The data used in this paper are from the US Census Bureau's decennial census for 1960, 1970, 1980, 1990, and 2000 and from the 2010 American Community Survey (ACS). All of the samples are nationally representative. Surveys in 1960, 1970, and 2010 were 1 percent samples of the US population, while the 1980, 1990, and 2000 surveys were 5 percent samples.[4] The US census counts every resident of the country using a household survey. Most surveys are completed by mail with remaining households interviewed in person by a census worker. The ACS data are collected in a similar fashion as the decennial census, collecting data from household units and group quarters. In the 2010 ACS, there was a 97.5 percent response rate for housing units and a 97.6 percent response rate for group quarters.

The analysis is restricted to individuals aged 18 to 64 who participate in the civilian labor force. Borjas (2003) argues in favor of measuring the labor market effects of immigration in the United States for native men by dividing the national labor force into groups based on skill, which he defines as a mix of the individual's education and work experience. Welch (1979) and Card and Lemieux (2001) establish that similarly educated workers with different levels of experience are not perfect substitutes. Borjas therefore defines four levels of education (high school dropouts; high school completion or General Educational Development [GED]); some college; and bachelor's degree or above) and eight experience cells by years of experience, using five-year increments up to 40 years.[5]

Experience among men is measured as an approximation to a Mincer variable for years of experience among men, estimated as a function of age and years of education.[6] Women traditionally take the primary role in rearing children within families, and if they choose to work, they tend to take more time out of the labor force than do men. To attempt to measure female experience accurately, we use an estimate of average work experience among women from Regan and Oaxaca (2009). They used the 1979 National Longitudinal Survey of Youth, which contains data on actual work experience, to show that in 1990 white female heads of household aged 18 to 55 had an average discrepancy of 5.4 years between potential and actual work experience, as opposed to an average discrepancy of just 1.1 years for white male heads of household in the same age range. We assume that this gap is largely due to fertility, and we take the fertility rates for each single-year age group in 1960, 1970, 1980, 1990, and 2000, as reported by the CDC (2010), and use these fertility rates as cumulative weights to build up the 5.4-year gap between the ages of 18 and 49, since fertility is largely completed by 49.[7] The result is an estimate of the difference between potential and actual experience for each single-year age group of women in each year; those are then used to adjust the potential work experience for women in the labor force before assigning them to education-experience cells.[8] Although measurement error may still be large for individual women's actual work experience, the fact that we aggregate to five-year cells and measure average wages means that this procedure should significantly reduce measurement errors in the aggregated data.

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[4] All datasets were accessed through the Integrated Public Use Microdata Series (IPUMS-USA). This service allows users to access multiple samples from different surveys under a single data file.

[5] Borjas creates an index of congruence in occupation distributions within education groups to demonstrate that immigrants with similar education and experience are a better substitute for natives in the same cell than other natives with similar education but different experience.

[6] As in Borjas (2003), work experience for men is defined as the worker's age in the given year minus the assumed age of entry in the labor market. It is assumed that high school dropouts enter the labor market at age 17, high school graduates at age 19, persons with some college at age 21, and college graduates at age 23. In this paper, female work experience is calculated the same as that of men, but is then adjusted downward to account for irregularities in women's labor force participation relative to men.

[7] Since fertility rates for 2010 were not available, 2010 data are weighted using the 2000 fertility rate.

[8] One potential shortcoming of this method is that women with different levels of education likely have different fertility rates, with fertility rates of more educated women lagging those of less educated women (for example, Skirbekk 2008).

Variable Construction

The dependent variables used in the paper are the logarithm of average weekly and annual wage levels among native men aged 18 to 64 within each education-experience cell. The census and ACS data contain a variable for annual pretax wage and salary income, pertaining to the previous calendar year for each census and the previous 12 months for the ACS. In the data, amounts are expressed in contemporary dollars, so we deflate these to 1980 dollars by using the CPI-U (Consumer Price Index for All Urban Consumers) series. The data also contain a variable for the number of weeks a respondent worked for profit, for pay, or as an unpaid family worker in the past year (with a year again defined as the prior 12 months). The variable for log weekly earnings is the log of the annual wage and salary income divided by weeks worked.[1]

We define the labor shares of immigrants and women as follows. We measure the share of immigrants in a specific education-experience cell, indexed by time. The variable can be written as $p_{ijt} = \frac{M_{ijt}}{M_{ijt}+N_{ijt}}$, where M_{ijt} is the number of immigrants and N_{ijt} is the number of natives with education level *i* and experience level *j* in year *t*. Counts of both immigrants and natives are limited to men who do not reside in group quarters and participate in the civilian labor force.[2] Immigrants are classified as all individuals born abroad, whether they are noncitizens or naturalized citizens; all others are classified as natives. The share of women in each education-experience cell, again indexed by year, follows a parallel definition. It is defined as $w_{ijt} = \frac{F_{ijt}}{F_{ijt}+G_{ijt}}$, where *F* is the number of women in each specific education-experience-time cell, and *G* is the number of men (native and immigrant) in an education-experience-time cell.

We illustrate changes in the share of immigrants within each education category of the male US labor force between 1960 and 2010 (Figure 3.1).[3] While immigrants generally made up an increasing share of the labor force for all four education categories, the increase is most pronounced among the least educated. In 1960, immigrants represented just 6.4 percent of male workers with less than a high school education in the United States.[4] By 2000, immigrants represented 35.7 percent of this group, and 44.7 percent by 2010. In comparison to this nearly sevenfold increase from 1960 to 2010, the next-most dramatic immigrant supply shock to an education category occurred among the college educated, which slightly doubled from 7.2 percent in 1960 to 15.2 percent in 2010.

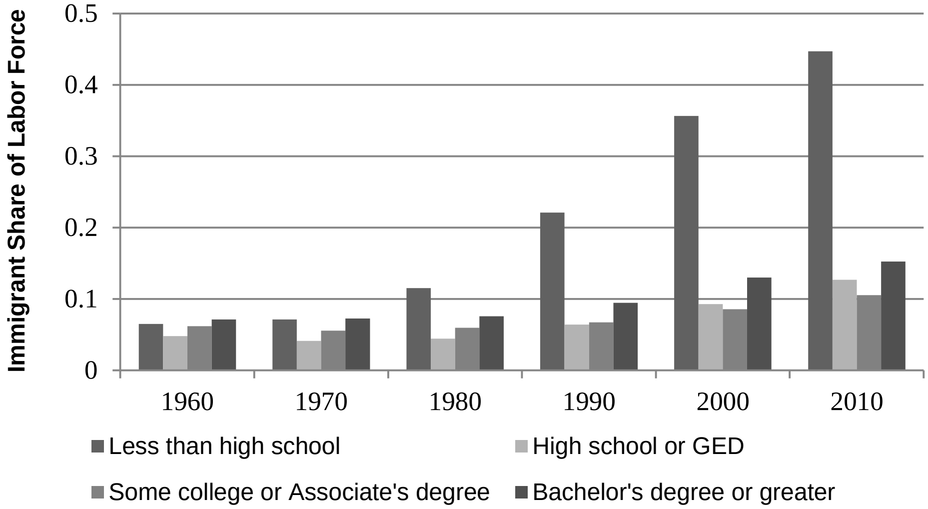
[1] To actually compute the earnings, we make adjustments as in Borjas to replicate his work as closely as possible. Quoting from Borjas (2003, 1391–1392), adjustments are as follows: "This variable is calculated for native men who do not live in group quarters, are employed in the civilian labor force, are not enrolled in school, report positive annual earnings, weeks worked, and weekly hours, and are not self-employed. In the 1960, 1970 and 1980 Census, the top-coded annual salary is multiplied by 1.5. In the 1960 and 1970 Census, weeks worked in the calendar year prior to the survey are reported as a categorical variable. I impute weeks worked for each worker as follows: 6.5 weeks for 13 weeks or less, 20 for 14–26 weeks, 33 for 27–39 weeks, 43.5 for 40–47 weeks, 48.5 for 48–49 weeks, and 51 for 50–52 weeks. The average log annual earnings or average log weekly earnings for a particular education-experience cell is defined as the mean of log annual earnings or log weekly earnings over all workers in the relevant population."

[2] We relax this restriction in the empirical work and test reclassifying the natives variable as including women.

[3] Sampling weights are used in all calculations involving the 1990 census. IPUMS-USA includes in the downloaded dataset variables for weights of household and person-level variables.

[4] After immigration quotas were restricted through the Immigration Act of 1924, the proportion of immigrants in the US labor force began to decline. In the census data, the decline lasted from 1930 to 1970 before beginning to increase again (Borjas and Katz 2007).

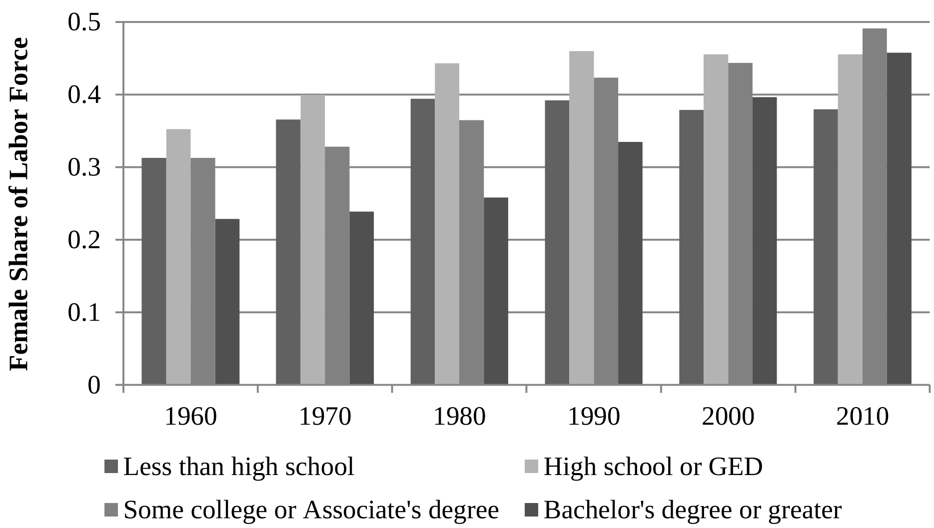
Figure 3.1 Immigrant share of labor force by education category, 1960–2010



Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Note: The immigrant share is defined among men in the civilian labor force who do not live in group quarters. Education categories are determined by the highest grade the individual has completed.

Although the immigrant share of the labor force rose rapidly between 1960 and 2010, women entered the labor force even more rapidly (Figure 3.2). The overall trend shows women as an increasing share of the labor force, except among the least educated workers. In contrast to immigrants, the change is most pronounced among the highest educated. Women represented just 23 percent of workers with a bachelor's degree or better in 1960, but made up 45 percent of this category by 2010. Similarly, women were just 31 percent of those with some college or an associate's degree in 1960, but were 49 percent of this category by 2010. Though largely taking place at opposite ends of the skill spectrum, the influx of women into the US labor market from 1960 to 2010 was perhaps even more dramatic than that of immigrants over the same period.

Figure 3.2 Female share of labor force by education category, 1960–2010



Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Notes: The female share is defined among all workers in the civilian labor force who do not live in group quarters. Education categories are determined by the highest grade the individual has completed.

4. CONCEPTUAL AND EMPIRICAL MODELS

The partial equilibrium framework in Borjas (2003) effectively considers how the productive component of the economy demands labor. Within a class of perfectly substitutable workers, if the supply of those laborers increases, we should observe a wage response on the labor demand side. He generally specifies the logarithm of average wages within a specific education-experience cell to be a function of the share of immigrants within the cell. If the share of immigrants increases within a cell, holding other factors constant, average wages offered to workers within that cell should decrease due to the increased competition for jobs within that cell. The unit of analysis is the education-experience cell as opposed to the individual.

Borjas (2003) uses the following fixed-effects regression model to determine the effects of immigration on wages in the United States from 1960 to 2000:

$$y_{ijt} = \theta p_{ijt} + s_i + x_j + \pi_t + (s_i * x_j) + (s_i * \pi_t) + (x_j * \pi_t) + \varphi_{ijt}, \qquad (1)$$

where y_{ijt} is the mean of log annual earnings or the mean of log weekly earnings for *native* men who have education level i ($i = 1, \ldots, 4$), experience level j ($j = 1, \ldots, 8$), and are observed at time t (t = 1960, 1970, 1980, 1990, 2000, 2010). As defined above, p_{ijt} is the share of immigrants in each education-experience cell at time t; θ is the effect of immigrant share on wages; s_i is a vector of fixed effects indicating the group's educational attainment; x_j is a vector of fixed effects indicating the group's work experience; and π_t is a vector of fixed effects indicating the time period. The fixed effects control for all observable and unobservable factors that might influence the wage levels by education level, by experience level, or for each year included. Terms that interact each pair of fixed effects control for unobservable differences in trends for specific education and experience levels, and the third interaction controls for any unique characteristics about specific education-experience cells that might influence wage levels.

To augment equation 1, we add the share of women w_{ijt} in an education-experience cell as a regressor:

$$y_{ijt} = \theta p_{ijt} + \beta w_{ijt} + s_i + x_j + \pi_t + (s_i * x_j) + (s_i * \pi_t) + (x_j * \pi_t) + \varphi_{ijt}. \qquad (2)$$

If p_{ijt} and w_{ijt} are correlated after we control for all of the fixed effects, then the estimate of θ generated by equation 1 would suffer from omitted variable bias. If they are not correlated, then equation 2 can be estimated while setting either θ or β to zero without biasing the estimated coefficient for the other parameter.

We restrict attention to the partial equilibrium model in Borjas (2003), and therefore we expect to estimate negative values for both θ and β. If we observe either that omitted variable bias exists or that the results from adding women to the model are questionable, we might question whether the estimation framework remains susceptible to simultaneity bias due to omitted variables. Before estimating equation 2, we consider some descriptive evidence on both the immigrant and female share variables.

5. DESCRIPTIVE RESULTS

To begin our analysis of the effects of immigrant and female penetration of labor force skill cells on native male wages, it is worth initially examining the supply shocks in more detail. First, we examine changes in the share of immigrants by education and experience level over time (Figure 5.1). The supply shock in immigrants generally rises over time in each education-experience cell. The rise is most pronounced among high school dropouts; at an extreme, among high school dropouts with 15 to 30 years of experience, immigrants make up less than 10 percent of the education-experience cell in both 1960 and 1970, but by 2010 immigrants make up 50 percent or more of the cell. Among college graduates, the immigrant share is also rising, but the increase is much less pronounced, with the largest increase over time taking place among those with 15 to 20 years of experience (from 5 percent in 1960 to 20 percent in 2010).

Figure 5.1 The immigrant supply shock, 1960–2010

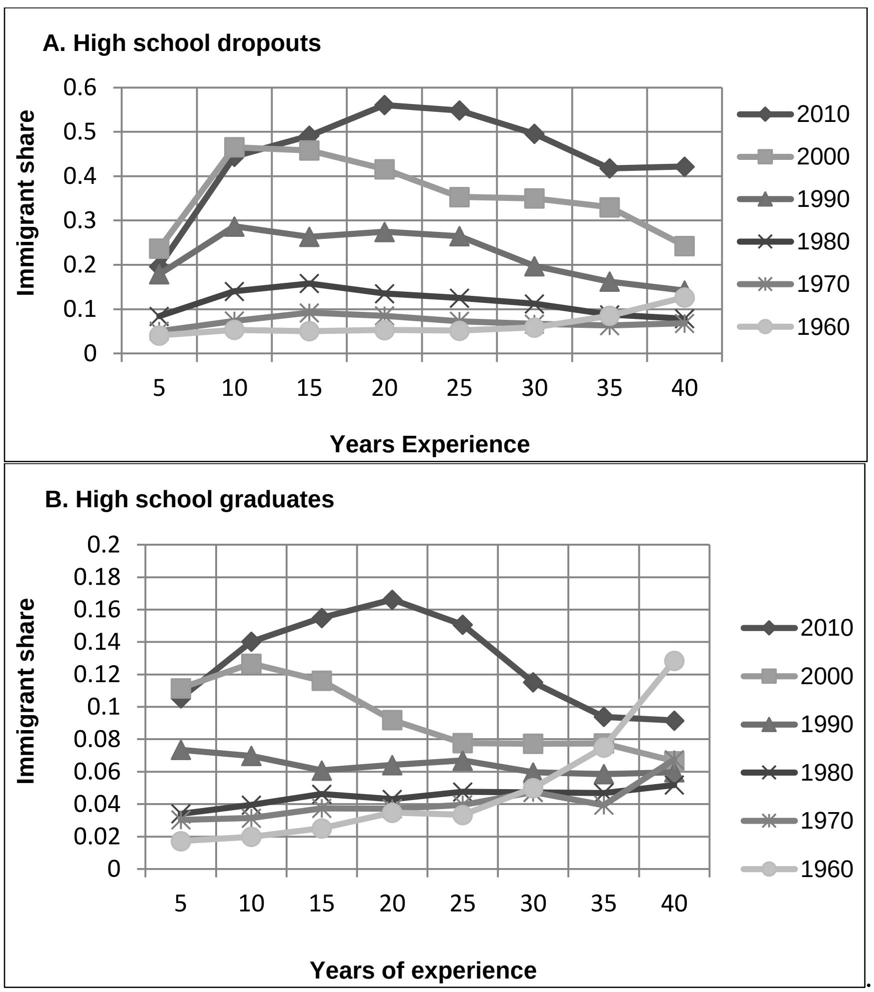
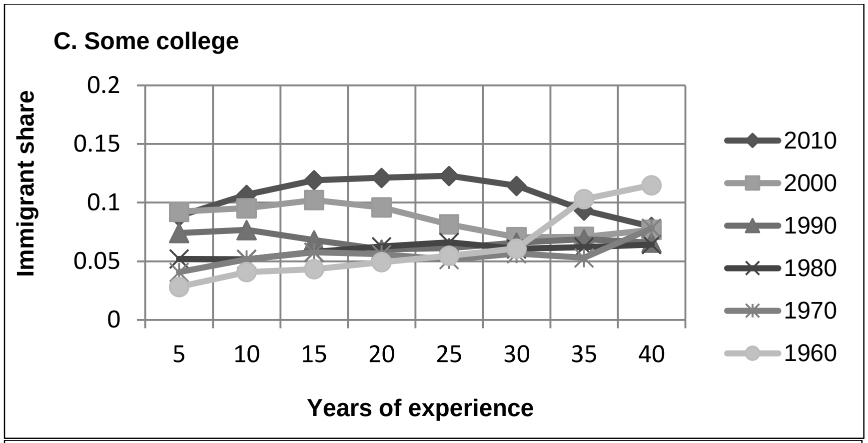
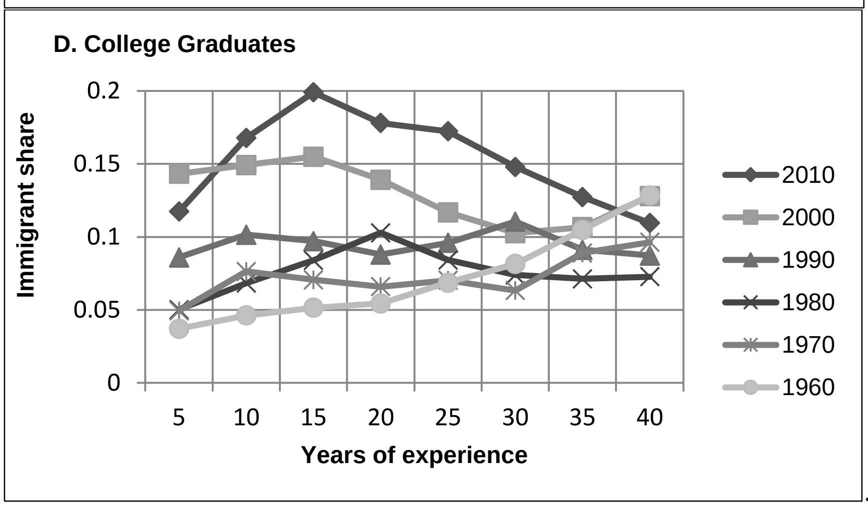


Figure 5.1. Continued





Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Note: Each education group is divided into eight five-year experience intervals; the values along the horizontal axis represent the upper bound of the corresponding cell.

We next examine changes in the share of women in each education-experience cell (Figure 5.2). As with immigrants, we observe large changes in the share of women in some of the education-experience cells. However, we note that the changes are relatively muted in the three lower education categories. For example, among high school dropouts, 2010 does not even have the largest shares of women working in any of the education-experience cells considering changes over time. Meanwhile, the shock is quite pronounced among college graduates, particularly those with less experience; by 2010, women make up the majority of all cells with less than 20 years of experience among college graduates, up from between 20 and 30 percent in 1960.

Figure 5.2 The female labor supply shock, 1960–2010

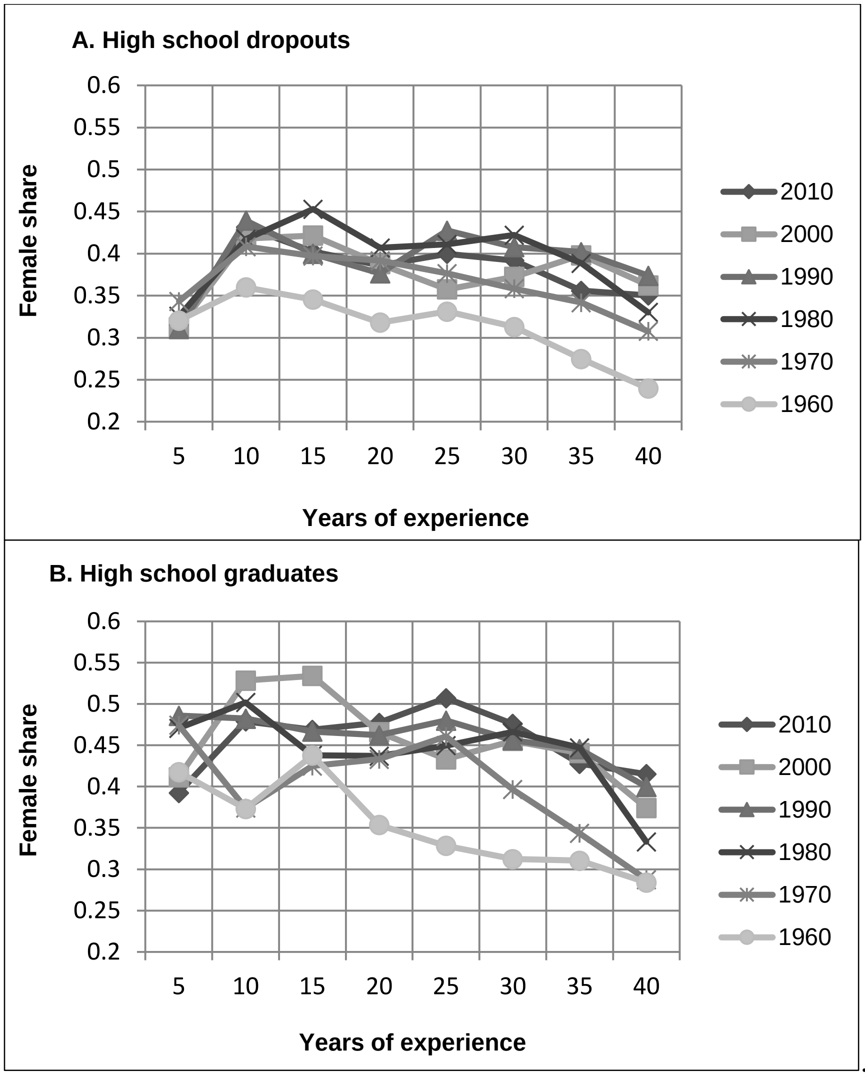
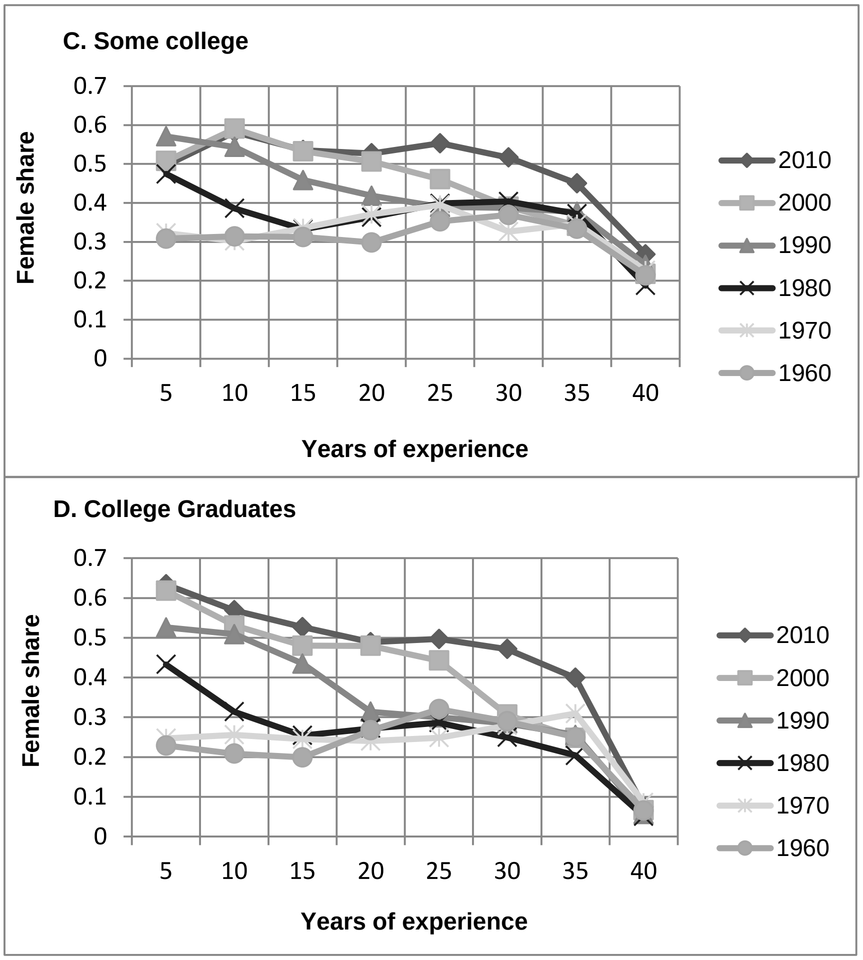


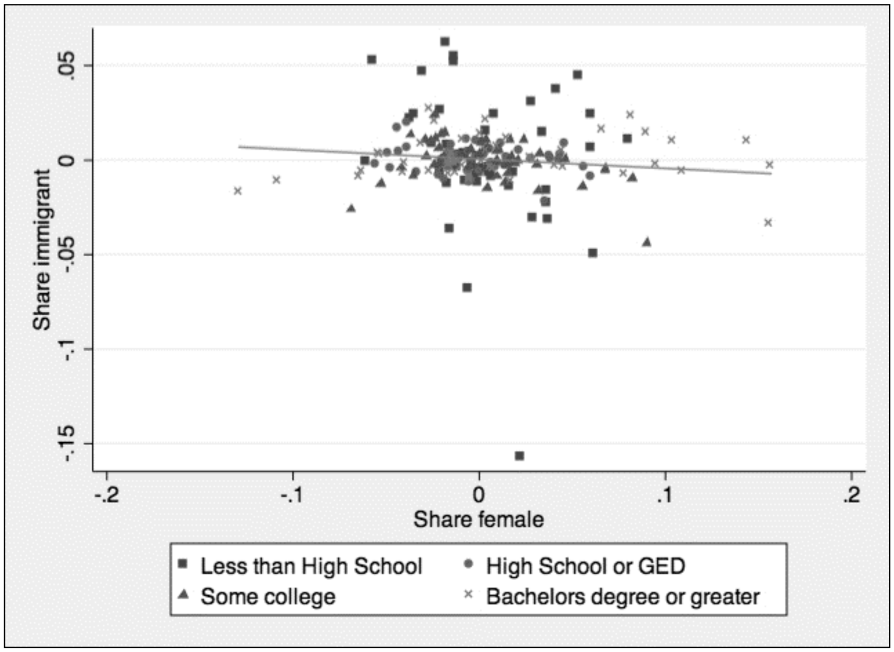
Figure 5.2 Continued



Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Note: Each education group is divided into eight five-year experience intervals; the values along the horizontal axis represent the upper bound of the corresponding cell.

Given the large difference between immigrants and women with respect to the education-experience cells in which we observe rapid growth, one might question whether the two are in fact correlated. We next test the null hypothesis of a zero correlation between the two as follows. First, we take both variables and regress them on a full set of dummy variables for education, experience, years, and the interactions between them. We save the residuals from both regressions and plot them against one another (Figure 5.3). The data suggest no correlation between the immigrant share and female share variables once fixed effects are removed; the slope of the regression line is not significantly different than zero. Since the immigrant and female share variables are uncorrelated, we do not expect that Borjas's result was affected by omitting the female share variable. The data generally show a bit more dispersion among the residuals of the female share variable than the immigrant share variable, particularly among the most educated workers, whereas the dispersion is largest among immigrants for the least educated workers. The illustration further supports the idea that immigrants and women were entering very different skill cells. It could be, of course, that other omitted variables related to labor supply could affect coefficient estimates for θ and β.

Figure 5.3 Scatter diagram relating immigrant and female entrance into the US labor market, 1960–2010 (fixed effects removed)



Source: IPUMS, 1960-2010 (Ruggles et al. 2010).

Notes: Each point in the scatter diagram represents the immigrant share among male workers and the female share among all workers in an education-experience-time cell. The data have been adjusted to remove any fixed effects of the skill cells across time. The slope of the regression line is -0.102, with a standard error of 0.159.

6. REGRESSION RESULTS

We initially directly replicate the partial equilibrium results in Borjas (2003) using equation 1 (Table 6.1, column 1). When we regress log weekly earnings among native men on the immigrant share in education-experience-time cells between 1960 and 2000, we estimate a coefficient on the immigrant share of -0.453, whereas using the same sample Borjas estimated a coefficient of -0.514.[13] As we replicated Borjas's variables starting from the raw data, the difference is likely due to small discrepancies in variable construction.[14]

Table 6.1 The impact of the share of immigrants and women on the logarithm of weekly earnings among native males

Time period	**1960–2000**	**1960–2010**	**1960–2000**	**1960–2010**	**1960–2000**	**1960–2010**
Explanatory variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Share immigrant	-0.453***	-0.348***			-0.451***	-0.343***
	(0.140)	(0.075)			(0.139)	(0.072)
Share female			-0.029	0.097	0.026	0.072
			(0.092)	(0.106)	(0.078)	(0.097)
Adjusted R^2	0.997	0.997	0.997	0.997	0.997	0.997
Observations	160	192	160	192	160	192

Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Notes: Standard errors are robust in all specifications, and all regressions are weighted by the relative size of the education-experience cell within the given year. Education, experience, and year fixed effects are included in all models, as well as interactions between education and experience fixed effects, education and year fixed effects, and experience and period fixed effects. ***$p < .01$, **$p < .05$, *$p < .10$.

Since our point estimate is 10 percent different in magnitude from that of Borjas, we next convert it into an elasticity. We define $m_{ijt} = {M_{ijt}}/{N_{ijt}}$ as the ratio of immigrants to natives in the labor supply of group (i,j,t) attributable to migration. The wage elasticity we are interested in is $\partial \log w_{ijt}/\partial m_{ijt} = \theta/(1+m_{ijt})^2$ (Borjas 2003, 1348–1349). If we evaluate the elasticity at the mean value of the increase in the male labor force, then we compute an average elasticity over time. Between 1960 and 2000, immigration increased the number of men in the labor force by 16.8 percent, so the wage elasticity is -0.33 according to our point estimate for θ. As our coefficient estimate is slightly lower than Borjas's using the same data, it is not surprising that the elasticity estimate is also somewhat lower in magnitude than his estimate (-0.38). Regardless, this estimate is still on the high end of the spectrum in the literature.

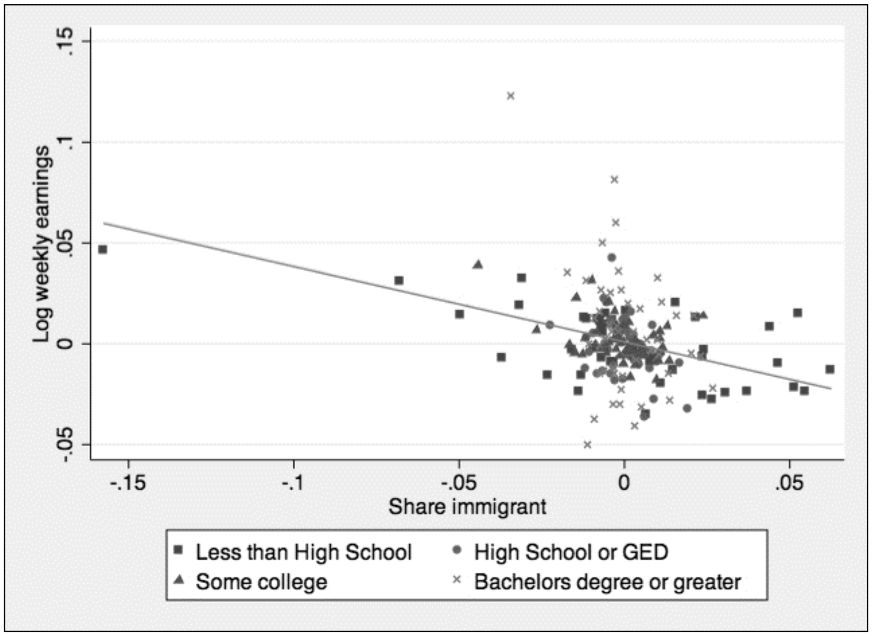
We next add the 2010 data and re-estimate equation 1 (Table 6.1, column 2). We estimate a coefficient of -0.348 on the share of immigrants, suggesting that when we average in the 2010 data, the relationship between wages and immigration weakens using the Borjas partial equilibrium model. We also graph the relationship in Figure 6.1, after removing fixed effects from both variables. Between 1960 and 2010, we compute that immigration increased the number of men in the labor force by 26.1 percent;

[13] This coefficient estimate is reported by Borjas (2003) in the footnotes of page 1349, and 2000 census data are used in place of the pooled 1999–2001 CPS data. It thus differs somewhat from the coefficient reported in the results table shown in the body of the paper, but it is directly comparable to the data used for our paper.

[14] It is worth noting that an unforeseen irregularity in the data construction process was the need to use the 1970 Form 2 Metro Sample. This was necessary given that the variable for current school attendance (used by Borjas to define the universe for the log weekly and annual earnings of native men) does not exist in the 1970 Form 1 Metro Sample. We assume that Borjas (2003) used that sample. While our calculations of mean log weekly earnings of men are not exactly the same as those of Borjas (2003), the correlation between the two sets of calculations is 0.990. In general, this shows that there is consistency in the re-creation of variables herein.

so the resulting wage elasticity is -0.22. This estimate is much more in line with the rest of the literature on the wage effects of migration; it is striking that adding 2010 to the analysis causes the point estimate for the elasticity to drop by over a full percentage point.

Figure 6.1 Scatter diagram relating native male weekly earnings and immigration, 1960–2010



Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Note: Each point in the scatter diagram represents the log weekly earnings of native male workers and the immigrant share among all male workers in an education-experience-time cell. The data have been adjusted to remove any fixed effects of the skill cells across time.

The latter result must begin to bring the estimating framework into question. Between 2000 and 2010, the immigrant supply shock actually accelerated. If the model estimated between 1960 and 2000 was correct, we would have expected, holding everything else constant, that as immigrants increased dramatically as a share of the overall labor force between 2000 and 2010, wages would decrease significantly on average. Combined with the 2008 financial crisis, one might expect that wages would decline significantly in most cells between 2000 and 2010, and we would expect coefficient estimates to remain consistent or perhaps even to increase as the supply shock accelerated. Instead, we estimate a smaller coefficient, suggesting an attenuation of wage impacts even as more immigrants joined the labor force. Although we observe some significant declines in wages particularly among the least experienced workers (Appendix Table A.1), these declines do not appear to affect all classes of laborers, with wages among the college educated in particular not declining much if at all (for some cells) between 2000 and 2010. Nevertheless, the change in coefficients may reflect the decline in jobs in the middle of the wage spectrum over the past two decades, a trend that was accelerated during the recession (for example, Autor 2010 and Holzer 2010). Since the data would have suggested a larger-magnitude elasticity if the model is correct, one must begin to question the estimating framework.

In the next two columns (Table 6.1, columns 3 and 4), we effectively estimate equation 2 while restricting $\theta = 0$. Doing so should isolate the impacts of female entry into the labor force on male wages, on a weekly basis. Using the 1960-to-2000 sample, we find a negative point estimate, but it is not statistically different from zero. When we add the 2010 data and re-estimate (column 4), we find a positive point estimate, though it is again not significantly different from zero. The scatter plot of wages on the share of women in each cell, after fixed effects have been removed, is either somewhat suggestive of nonlinearities in the relationship or suggestive of much more dispersion in the relationship among the most educated (Figure 6.1). Nonetheless, the results again call into question the estimating framework, as we would expect that such a large supply shock should have a negative correlation with wages.

In the final two columns of Table 6.1, we estimate equation 2, which includes both supply shocks on the right-hand side, for 1960–2000 and 1960–2010, respectively (columns 5 and 6). Given the lack of correlation between the share of immigrants and the share of women in each cell, it is not surprising that point estimates on the coefficients do not materially change. The wage elasticity of immigration would appear to remain about -0.22, using the full dataset.

We next use an alternative definition of wages as the dependent variable, which is the annual wages reported in the census. This variable is constructed as the weekly wage multiplied by the number of weeks worked, so it effectively includes a choice about the intensity of work. Here, our results differ substantially from Borjas, when considering the immigration shock (Table 6.2). Whereas Borjas estimated a coefficient of -0.68 significant at the 5 percent level, our replication shows a smaller coefficient and it is not significantly different from zero (column 1).[15] The difference likely has to do with small differences in the way that we replicated Borjas's variables; nonetheless, the fact that we cannot replicate the statistical significance suggests that this result is somewhat fragile. Moreover, when we add the 2010 data (column 2), the point estimate drops to -0.26, and it remains statistically no different from zero.

Table 6.2 Impact of immigrant share and female share on log annual earnings of native males (weighted regression)

Time period	**1960–2000**	**1960–2010**	**1960–2000**	**1960–2010**	**1960–2000**	**1960–2010**
Explanatory variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Share immigrant	-0.479	-0.260			-0.465*	-0.242
	(0.333)	(0.196)			(0.273)	(0.153)
Share female			0.218**	0.302**	0.214**	0.299***
			(0.102)	(0.112)	(0.092)	(0.107)
Adj. R-squared	0.997	0.996	0.997	0.996	0.997	0.996
Observations	160	192	160	192	160	192

Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Notes: Robust standard errors. All regressions are weighted by the relative size of the education-experience cell within the given year. Included but not reported for all models are year, education, experience fixed effects, as well as interactions between education and experience fixed effects, education and year fixed effects, and experience and period fixed effects. ***$p < .01$, **$p < .05$, *$p < .10$.

Results from the regression of average log annual wages among men on the share of women in each education-experience-year cell are even more interesting (Table 6.2, columns 3 and 4). Regardless of whether we include the 2010 data, we estimate a positive and significant coefficient on the share of women in the cell. In other words, when women enter a specific education-experience-year cell at a faster rate, men's wages increase rather than decrease. The coefficient estimates remain statistically significant when we add the immigrant supply shock variable to regressions as well (columns 5 and 6). If we took these results seriously, one might suggest that the labor supply curve has been upward, rather than downward, sloping, at least in response to the supply shock of women entering the labor force.

[15] Again, this estimate uses the coefficient Borjas reports in footnote 8 after using the 2000 census data in place of pooled CPS data.

In combination, these results suggest that Borjas's model is inadequate to describe the impact of a labor supply shock on wages, since the supply shock theoretically cannot induce an increase in wages. We next test the robustness of our main results to specification changes, and then we attempt to isolate the education groups in which the supply shocks are occurring, before returning to discuss some explanations for our findings.

Robustness Checks

First, recall that in Tables 6.1 and 6.2 we follow Borjas (2003) by weighting our main estimates by the size of each education-experience-year cell. We re-estimate equation 2 without weights using the two supply shocks sequentially, as well as both dependent variables (Table 6.3). Using the logarithm of average weekly wages as a dependent variable, the general results hold. We find the magnitude of the impact of immigrant penetration is a bit larger when we use only 1960–2000; it is a bit weaker when we include 2010 as well (Panel A, columns 1 and 2).[16] Using the log annual wages as a dependent variable (Panel B), we find the magnitudes of all coefficients increase, and the coefficient on the immigrant share variable becomes significant at the 5 percent level when included in the same model as the share of female workers, but only for the 1960–2000 data (column 5). The coefficients on the female share variables also become more precisely estimated, suggesting they remain positive, which they do for both the 1960–2000 subset and the 1960–2010 data.

Table 6.3 Impact of immigrant share and female share on log earnings of native males, without weights

Time period	**1960–2000**	**1960–2010**	**1960–2000**	**1960–2010**	**1960–2000**	**1960–2010**
Explanatory variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Panel A: Log weekly wages as dependent variable						
Share immigrant	-0.502***	-0.320***			-0.491***	-0.305***
	(0.136)	(0.068)			(0.123)	(0.063)
Share female			0.056	0.122	0.040	0.112
			(0.080)	(0.094)	(0.060)	(0.085)
Overall R-squared	0.016	0.027	0.015	0.080	0.010	0.029
Observations	160	192	160	192	160	192
Panel B: Log annual wages as dependent variable						
Share immigrant	-0.640*	-0.247	---	---	-0.569**	-0.199
	(0.353)	(0.188)			(0.251)	(0.130)
Share female	---	---	0.281***	0.362***	0.261***	0.356***
			(0.094)	(0.104)	(0.077)	(.098)
Overall R-squared	0.006	0.028	0.039	0.048	0.000	0.029
Observations	160	192	160	192	160	192

Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Notes: Robust standard errors. Included but not reported for all models are year, education, experience fixed effects, as well as interactions between education and experience fixed effects, education and year fixed effects, and experience and period fixed effects. ****p* < .01, ***p* < .05, **p* < .10.

[16] Borjas performs a similar exercise as in Table 6.3, yet his coefficient estimates decline rather than increase in absolute value.

Second, recall that we originally computed the size of the labor force in each education-experience-year cell using only the count of native men in the denominator. We next relax this assumption and include women in the labor force counts, particularly when calculating the immigrant share in each cell (Table 6.4). Using the logarithm of the weekly wages, we find the point estimate of the coefficient on immigrant share increases in magnitude (-0.380 from -0.348; Panel A, column 2) and remains statistically significant at the 1 percent level. Coefficients on the immigrant share using annual wages as the dependent variable also increase in magnitude, but remain insignificant when using the full 1960–2010 sample (Panel B). Not surprisingly, coefficient estimates on the share of women in each cell remain similar to the base estimates, as we have not changed those variables.

Table 6.4 Impact of immigrant share and female share on log weekly earnings of native males (variable for share immigrant includes women)

Time period	1960–2000	1960–2010	1960–2000	1960–2010	1960–2000	1960–2010
Explanatory variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Panel A: Weekly wages						
Share immigrant	-0.543***	-0.380***	---	---	-0.541***	-0.359***
	(0.136)	(0.081)			(0.124)	(0.096)
Share female	---	---	0.031	0.097	0.002	0.078
			(0.093)	(0.106)	(0.076)	(0.096)
Adjusted R-squared	0.997	0.997	0.997	0.997	0.997	0.997
Observations	160	192	160	192	160	192
Panel B. Annual wages						
Share immigrant	-0.662*	-0.319	---	---	-0.570*	-0.240
	(0.378)	(0.223)			(0.301)	(0.155)
Share female	---	---	0.218**	0.302**	0.188**	0.289**
			(0.102)	(0.112)	(0.090)	(0.106)
Observations	160	192	160	192	160	192

Source: IPUMS, 1960-2010 (Ruggles et al. 2010).

Notes: Robust standard errors. All regressions are weighted by the relative size of the education-experience cell within the given year. Included but not reported for all models are year, education, experience fixed effects, as well as interactions between education and experience fixed effects, education and year fixed effects, and experience and period fixed effects. ****p* < .01, ***p* < .05, **p* < .10.

We next follow Borjas and add a control for the logarithm of the number of native men in each education-experience-year cell (Table 6.5). The addition of this variable has the largest effect on the coefficient estimates for the regressions of annual wages on the share of immigrants in each cell (Panel B). Those coefficient estimates increase in magnitude becoming much larger and statistically significant at the 5 percent level (columns 1 and 2). However, the general pattern holds—adding the 2010 data decreases the magnitude of the coefficient estimates, and coefficients on the share of women in each cell remain positive and statistically significant at the 5 percent level. Interestingly, in this case the combination of the two shares changes coefficient estimates somewhat, particularly for the share of women. The coefficient on this variable declines slightly in magnitude and loses statistical significance when the two variables are both included in the model (columns 5 and 6). However, it remains positive and not statistically significant at the 5 percent level when the 2010 data are included in the regression (column 6).

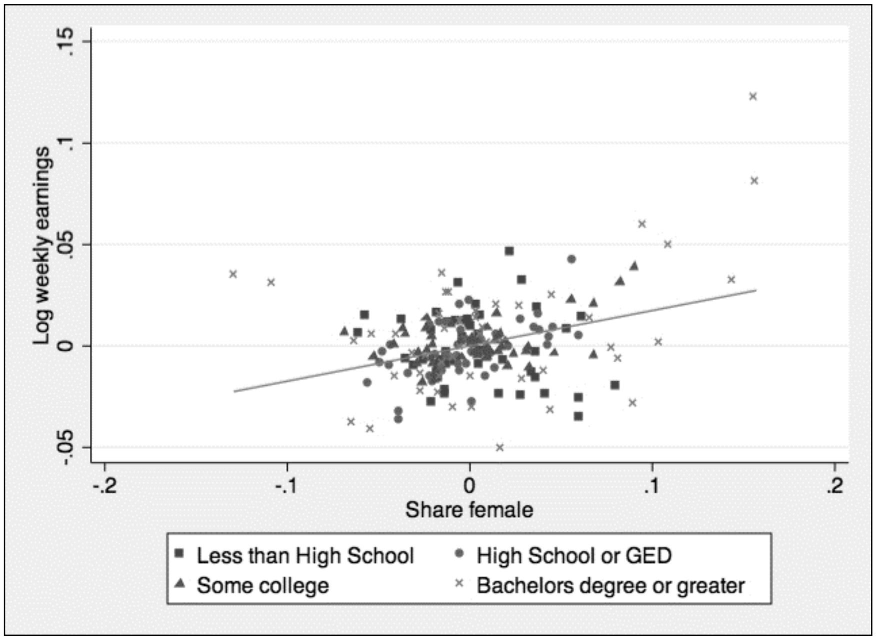
Table 6.5 Impact of immigrant share and female share on log weekly earnings of native males (all regressions include a variable for log of number of native males)

Time period	1960–2000	1960–2010	1960–2000	1960–2010	1960–2000	1960–2010
Explanatory variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Panel A: Log weekly wages						
Share immigrant	-0.434**	-0.389***	---	---	-0.376**	-0.282***
	(0.196)	(0.136)			(0.173)	(0.101)
Share female	---	---	0.115	0.171	0.064	0.121
			(0.090)	(0.102)	(0.083)	(0.102)
Adjusted R-squared	0.997	0.997	0.997	0.997	0.997	0.997
Observations	160	192	160	192	160	192
Panel B: Log annual wages						
Share immigrant	-0.725**	-0.586**	---	---	-0.581**	-0.372**
	(0.282)	(0.215)			(0.248)	(0.169)
Share female	---	---	0.238**	0.308**	0.159	0.241**
			(0.115)	(0.119)	(0.100)	(0.113)
Adjusted R-squared	0.997	0.996	0.997	0.996	0.997	0.996
Observations	160	192	160	192	160	192

Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Notes: Robust standard errors. All regressions are weighted by the relative size of the education-experience cell within the given year. Included but not reported for all models are year, education, experience fixed effects, as well as interactions between education and experience fixed effects, education and year fixed effects, and experience and period fixed effects. ****p* < .01, ***p* < .05, **p* < .10.

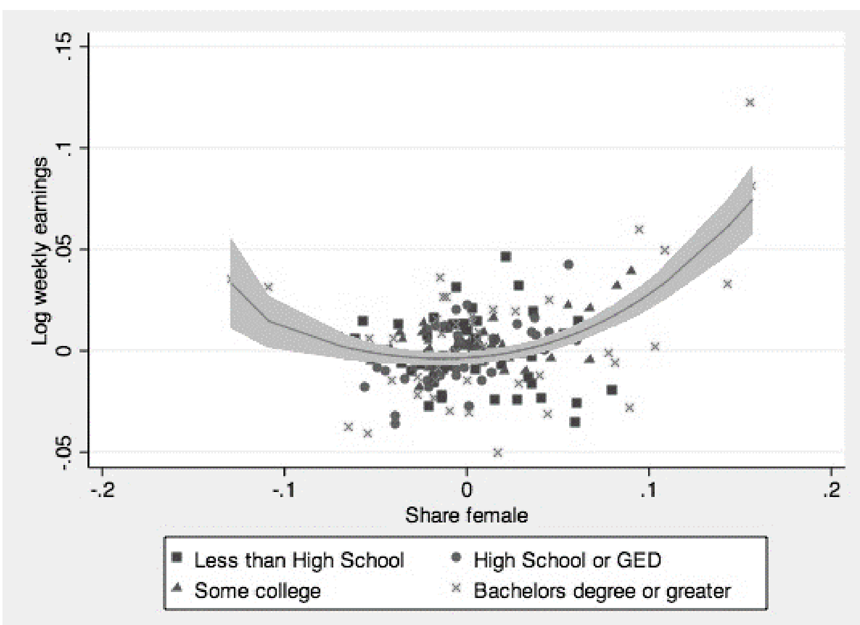
Finally, we examine the possibility that the effects of immigration and female entry on men's wages are nonlinear. Specifically, Figure 6.2 suggested a nonlinear relationship between the share of women entering a specific education-experience cell and the logarithm of men's weekly wages, whereas the relationship appears negative and linear for immigrants in Figure 6.1. We initially examine this hypothesis by using a local polynomial regression (Figure 6.3), and indeed we find a curve that appears quadratic, with an inflection point near zero. The local polynomial suggests that the relationship between men's wages and female entry might be negative, so long as the shocks are smaller than trends (removed through fixed effects) would suggest. However, when shocks are larger than expected through fixed effects, we would expect a positive and accelerating shock to wages. We also regressed log weekly wages on the female share and the female share squared, and find no significant coefficients. Nonetheless, these combined graphical and quantitative results are not consistent with a declining labor demand curve, and deserve additional explanation.

Figure 6.2 Scatter diagram relating native male weekly earnings and female entrance into the US labor market, 1960–2010



Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Note: Each point in the scatter diagram represents the log weekly earnings of native male workers and the female share among all workers in an education-experience-time cell. The data have been adjusted to remove any fixed effects of the skill cells across time.

Figure 6.3 Scatter diagram relating native male weekly earnings and female entrance into the US labor market, 1960–2010, with local polynomial fit



Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Note: Each point in the scatter diagram represents the log weekly earnings of native male workers and the female share among all workers in an education-experience-time cell. The data have been adjusted to remove any fixed effects of the skill cells across time.

Isolating the Effects of Education Groups

Recall that the descriptive results examined above suggested that immigrants and women have been entering the US labor force in very different skill cells in the past 50 years due to disparate education levels, and that changes in wages have not been equal across education levels during this time. Furthermore, scatterplots relating log native male weekly earnings to the share of immigrants and women in education-experience-year cells suggested that these correlations were strongly influenced by particular education groups (Figures 5.3 and 6.1). Thus, we next test whether excluding specific education groups changes the basic results. Since the figures suggest that immigrants are most likely to enter the labor force as high school dropouts and the fastest entry among women has been among the college educated, we try dropping both of these groups in estimating versions of equation 2.

First, we exclude high school dropouts from estimation (Table 6.6). Using log weekly wages as the dependent variable (Panel A), we find no significant coefficients on either the share of immigrants or women in each cell. This finding suggests that whatever correlation is found using this methodology, it is no longer present without high school dropouts; the negative coefficient between wages and immigration is fully driven by the least educated. Meanwhile, we continue to find positive, significant coefficients on the female labor force shock when we use annual wages as the dependent variable (Panel B).

Table 6.6 Impact of immigrant share and female share on log weekly earnings of native males, at least high school graduates (weighted regression)

Time period	1960–2000	1960–2010	1960–2000	1960–2010	1960–2000	1960–2010
Explanatory variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Panel A: Log Weekly Wages						
Share immigrant	-0.085	-0.078	---	---	-0.149	-0.175
	(0.565)	(0.328)			(0.563)	(0.347)
Share female	---	---	0.085	0.155	0.088	0.157
			(0.090)	(0.106)	(0.094)	(0.109)
Adj. R-squared	0.996	0.997	0.996	0.996	0.996	0.996
Observations	120	144	120	144	120	144
Panel B: Log Annual wages						
Share immigrant	-0.505	-0.324	---	---	-0.655	-0.514
	(0.699)	(0.427)			(0.630)	(0.413)
Share female	---	---	0.197*	0.299**	0.206*	0.307**
			(0.099)	(0.117)	(0.103)	(0.118)
Adj. R-squared	0.995	0.994	0.996	0.995	0.996	0.995
Observations	120	144	120	144	120	144

Source: IPUMS, 1960-2010 (Ruggles et al. 2010).

Notes: Robust standard errors. All regressions are weighted by the relative size of the education-experience cell within the given year. Included but not reported for all models are year, education, experience fixed effects, as well as interactions between education and experience fixed effects, education and year fixed effects, and experience and period fixed effects. ****p* < .01, ***p* < .05, **p* < .10.

We next drop college graduates from the full sample and re-estimate (Table 6.7). Not surprisingly, the negative coefficient estimate on the immigrant supply shock returns, when using weekly wages as the dependent variable (Panel A). The estimates are quite comparable to those using the full sample. We do not find major changes to the results pertaining to the female supply shock variable. The positive effect of women entering the labor force on annual earnings of native males also does not lose its statistical significance, though the magnitude of the coefficient does decrease somewhat.

Table 6.7 Impact of immigrant share and female share on log weekly earnings of native males, less than college graduates (weighted regression)

Time period	1960–2000	1960–2010	1960–2000	1960–2010	1960–2000	1960–2010
Explanatory variable	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
Panel A: Log Weekly Wages						
Share immigrant	-0.430***	-0.379***	---	---	-0.433***	-0.369***
	(0.127)	(0.091)			(0.121)	(0.082)
Share female	---	---	-0.010	0.078	0.013	0.034
			(0.079)	(0.098)	(0.063)	(0.083)
Observations	120	144	120	144	120	144
Panel B: Log Annual wages						
Share immigrant	-0.400	-0.298	---	---	-0.356	-0.228
	(0.282)	(0.205)			(0.255)	(0.180)
Share female	---	---	0.203**	0.274**	0.184*	0.246**
			(0.101)	(0.124)	(0.099)	(.120)
Observations	120	144	120	144	120	144

Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Notes: Robust standard errors. All regressions are weighted by the relative size of the education-experience cell within the given year. Included but not reported for all models are year, education, experience fixed effects, as well as interactions between education and experience fixed effects, education and year fixed effects, and experience and period fixed effects. ****p* < .01, ***p* < .05, **p* < .10.

In sum, these results suggest that the negative effect of immigration on weekly earnings of native men is being driven by those skill cells containing high school dropouts. Without these cells, the effect disappears entirely. And while women entering the labor force continues to have no statistically significant effect on weekly earnings of native males, the positive effect on annual earnings persists regardless of the exclusion of the highest (or lowest) educated from the analysis.

7. DISCUSSION

Our findings are strongly suggestive that empirical results from the Borjas partial equilibrium model of the national labor force should not be interpreted as causal. Still, they can be considered partial correlations, and such correlations deserve an explanation. Two particularly important features of the US economy over this period are that per capita gross domestic product (GDP) rose by 278 percent between 1960 and 2010 in real terms, suggesting substantial structural change in the economy, and that the 2000s were marked by a large sag in employment (Moffitt 2012), in which male labor force participation dropped back to 1991 levels from 2007, before the start of the Great Recession.

Over the study period, one way that low-skilled workers have likely improved their prospects is through increased human capital accumulation. There does not appear to be a large-scale movement among unskilled or potentially unskilled workers, native or otherwise, to attain more education; though the education level of workers has been rising somewhat in recent years, Heckman and LaFontaine (2010) show that in fact high school graduation rates peaked in 1969 and declined until 2000. The slight rise in high school graduation rates between 1968 and 2005 documented in education statistics among 18-to-24-year-olds (for example, Laird et al. 2007) is likely due to increases in GED completion among natives. Yet a larger and larger share of high school graduates continue their education (Edwards and Lange 2013). Consequently the labor force has fewer and fewer people with only a high school education over time, leaving high school dropouts even further behind as returns to higher education are increasing (Bowlus and Robinson 2012).

Still, in the context of including a period of declining labor demand, it might be surprising that we find a lower wage elasticity for immigration than Borjas. A potential explanation lies in the way that native high school dropouts have dealt with the employment decline of the 2000s, since less educated classes of workers have been more affected by the decline in employment. Many of the best jobs formerly held by unskilled workers have literally disappeared. Autor, Dorn, and Hanson (2013) show that the entry of China into specific industries, using trade statistics at the four-digit level, explains one-fourth of the aggregate decline in US manufacturing employment between 1990 and 2007. A similar analysis (Acemoglu et al. 2014) shows that a significant fraction of employment decline during the 2000s is attributable to manufacturing decline through both direct competition with China and through effects on upstream and downstream industries. As manufacturing employment declined, there has not been a concurrent rise in the share of GDP being generated through other low-skilled professions; the primary exception would be the accommodation and food industries, rising from 2.2 percent to only 2.9 percent of GDP.

Moreover, native workers losing manufacturing employment may not compete with immigrants in the same sectors. Even among high school dropouts, it is unclear that direct competition exists in many industries between native and immigrant high school dropouts. As argued by Ottaviano and Peri (2012), the two appear to largely enter different industries. As an example, Clemens (2013) finds that only seven natives completed the season working for the North Carolina Growers Association during the harvest, and only 268 of North Carolina's 489,000 unemployed even applied. The remainder of the 6,500 jobs were filled by migrants. Consequently, as found by Ottaviano and Peri (2012), immigrants and natives are not perfect substitutes; immigrants will take specific types of work that native workers will not take, and it is not obvious that in the fact of job losses high school dropouts stay in the labor force. To the extent that they live in nuclear families, the rise of female labor force participation actually may provide an alternative source of family income.

Indeed, the absolute number of high school dropouts in the potential labor force declined from 13.4 million to 11.9 million between 1992 and 2010, according to the Current Population Survey, and labor force participation rates among males are below 50 percent. A further possibility is that at least on the margin, the Social Security Disability Insurance (SSDI) Program has become akin to a reservation wage. The number of individuals receiving income from SSDI rose by 60 percent between 1984 and 2001 (Autor and Duggan 2003; Autor 2011) and continued to increase throughout the 2000s. Decreased demand for labor among high school dropouts, mixed with increasing ease of enrollment into the program, are estimated to have doubled the labor force exit propensity of displaced high school dropouts during this period (Autor and Duggan 2003). From a labor supply perspective, the possibility of receiving SSDI may appear to the decision maker as a reservation wage that takes him or her out of aggregate statistics on the supply of the labor force, particularly when faced with the choice of working at very difficult employment (for example, seasonal on-farm labor) versus receiving a check.

The positive correlation between men's wages on an annual basis and women's labor supply also deserves further explanation, and likely also lies with changes in the structure of the US economy over the 50-year period studied. As a share of GDP, some of the fastest-growing sectors of the economy require high-skilled education, such as finance and real estate, professional services, and education and health care. Labor demand has grown particularly rapidly in those sectors, and barriers to women's entry into those fields have concurrently fallen.[17] Hsieh et al. (2013) use an augmented Roy model, which includes labor market frictions, to examine changes in the allocation of talent in the US economy between 1960 and 2008 and estimate that improved allocation of talent, including women into high-skilled occupations, lifted aggregate GDP growth by 15 to 20 percent during that period. They find that using their model, which does not assume perfect substitutability, men's wages have fallen slightly over that period, suggesting the labor demand curve is indeed downward sloping.

To summarize, our results related to both immigrants and women can be explained by carefully considering how structural change may have led to different labor supply decisions among natives. Consequently, our estimates of equation 2 are not exactly estimates of labor demand, as they are likely affected by simultaneity bias related to labor supply that is not reflected in the fixed effects included in the equation. To truly identify the impacts of immigration on wages, more nuanced models allowing for imperfect substitution, such as the models used by Ottaviano and Peri (2012) or Hsieh et al. (2013), likely give more precise estimates. Even better would be a plausible source of exogenous variation for the impact of women's entry on wages, as used in Acemoglu, Autor, and Lyle (2004), though there is no obvious plausible source for immigration to the United States.

[17] In fact, the availability of immigrant labor may have helped reduce these barriers. Cortes and Tessada (2011) find that variation in the availability of immigrant labor across metropolitan areas helps explain increases in time spent at work among women in the upper quartile of the wage distribution.

8. CONCLUSION

In this paper, we replicate the Borjas (2003) model of wage effects on immigration, which has been used in policy debates to argue that immigration is generally harmful to wages among natives. Using the exact model and adding a round of the census (2010), we find that wages are not as elastic to immigration as Borjas found. Even if we believe the implicit assumptions behind the model, we estimate the wage elasticity to immigration is only -0.2, as opposed to -0.3 or -0.4. If one uses -0.2 as the wage elasticity of immigration with confidence, the rough calculation in Borjas (2014) using a wage elasticity of immigration of -0.3 substantially overstates the wage decline that would occur in the global North, implying that overall benefits to immigration are understated throughout the paper. In fact, with a wage elasticity of -0.2, overall welfare gains to immigration are positive ($3.1 trillion) even in Borjas's most pessimistic calculation. So it is not actually that difficult to find the "trillion dollars left on the sidewalk" (for example, Clemens 2011) by hindering migration.

That said, we use the same model to analyze the impacts of an even larger, concurrent shock to the US economy, namely the entry of women into the labor force. If we naively believe the model, we would believe that women's entry into the labor market actually has a positive effect on men's annual wages. This finding would suggest that perhaps the labor demand curve is not downward sloping, but upward sloping. Whereas one could tell a story consistent with this view—for example, that men and women are strong complements within specific education-experience cells, and the additional entry of women enhances such complementarities—it is more likely that this finding reflects a simultaneity bias between labor demand and supply in the model. Consequently, the measure of the share of immigrants in each education-experience cell is also likely reflective of labor supply decisions and the coefficient is likely also biased.

Therefore, our results call again into question the argument that immigration harms native wages in the United States. However, they also hint at interesting potential interactions between native workers and immigrants with low education levels. When college educated individuals are removed from the Borjas model, coefficient estimates increase, which could suggest harmful effects. Further research, either in terms of case studies such as Clemens (2013) or similar to the augmented Roy model used by Hsieh et al. (2013) across industries would help refine the impacts of immigration among low-skilled workers while taking into account imperfect substitutability.

APPENDIX: SUPPLEMENTARY TABLE

Table A.1 Log weekly earnings of male native workers, 1960–2010

Education	Years of experience	1960	1970	1980	1990	2000	2010
High school dropouts							
	1–5 years	5.015	5.179	5.125	4.958	5.023	4.723
	6–10 years	5.394	5.598	5.420	5.307	5.348	5.104
	11–15 years	5.587	5.753	5.563	5.474	5.503	5.308
	16–20 years	5.660	5.799	5.682	5.566	5.568	5.415
	21–25 years	5.677	5.846	5.744	5.637	5.623	5.493
	26–30 years	5.689	5.871	5.777	5.723	5.660	5.537
	31–35 years	5.663	5.843	5.815	5.762	5.696	5.577
	36–40 years	5.649	5.822	5.813	5.762	5.722	5.538
High school graduates or GED							
	1–5 years	5.390	5.535	5.472	5.306	5.325	5.020
	6–10 years	5.722	5.912	5.726	5.624	5.647	5.450
	11–15 years	5.862	6.003	5.884	5.781	5.777	5.617
	16–20 years	5.931	6.062	5.997	5.880	5.883	5.758
	21–25 years	5.963	6.113	6.034	5.965	5.935	5.819
	26–30 years	5.957	6.111	6.058	6.035	5.972	5.875
	31–35 years	5.932	6.112	6.056	6.030	6.000	5.893
	36–40 years	5.912	6.060	6.033	5.996	5.979	5.871
Some college or associate's degree							
	1–5 years	5.590	5.728	5.619	5.556	5.587	5.367
	6–10 years	5.880	6.059	5.848	5.850	5.867	5.741
	11–15 years	6.013	6.178	6.022	5.994	6.015	5.911
	16–20 years	6.090	6.230	6.137	6.079	6.099	6.019
	21–25 years	6.116	6.261	6.158	6.176	6.137	6.086
	26–30 years	6.102	6.260	6.178	6.228	6.151	6.109
	31–35 years	6.094	6.265	6.167	6.201	6.175	6.090
	36–40 years	6.047	6.195	6.129	6.119	6.120	5.995
College graduates							
	1–5 years	5.828	6.044	5.818	5.919	5.970	5.848
	6–10 years	6.105	6.340	6.092	6.233	6.264	6.216
	11–15 years	6.237	6.467	6.315	6.373	6.476	6.452
	16–20 years	6.328	6.532	6.438	6.464	6.561	6.568
	21–25 years	6.351	6.593	6.476	6.574	6.555	6.633
	26–30 years	6.353	6.580	6.483	6.597	6.558	6.634
	31–35 years	6.334	6.557	6.461	6.554	6.566	6.548
	36–40 years	6.295	6.479	6.378	6.432	6.430	6.409

Source: IPUMS, 1960-2010 (Ruggles et al. 2010).
Note: Limited to men aged 18 to 64, not living in group quarters, employed in the civilian labor force, not self-employed, not enrolled in school, with positive annual earnings, weeks worked, and weekly hours. GED = General Educational Development.

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