Comparing the Domestic Labor Market Impacts of a South-North and South-South Migration: The Cases of Costa Rica and the United States

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Abstract

In this paper, we calibrate a structural model of the native wage distribution to the economies of two countries: Costa Rica and the United States. We then use these empirical models to simulate the likely effects of labor market competition with immigrants on native wages. For Costa Rica we find very little evidence of imperfect substitutability between immigrants and natives of similar observable human capital levels. In contrast, there is fairly strong evidence of imperfect substitutability between immigrants and natives in the U.S. A plausible interpretation of this finding is that the ability to speak the receiving nation's language is an important determinant of the degree of substitutability between otherwise similar immigrants and natives. The wage simulation results yield little evidence of an impact of immigration on native wages in Costa Rica, and suggest modest effects on the least skilled natives in the U.S. The paper is organized as follows: The first section discusses the motive to undertake the comparative analysis presented in the paper and specifically discusses issues associated with degrees of substitutability between migrants and the labor force in the receiving country. Section 2 discusses similarities and differences between the two migratory processes discussed in the paper. Section 3 lays out the aggregate labor market model and derives the factor price elasticities needed to determine impacts of migrant flows on the native labor force. Sections 4 and 5 discuss general statistics derived from the data used in the analysis and estimates country specific labor elasticities of substitution. Section 6 uses the parameters obtained previously and simulates the impact of migrant flows on native wages. Section 7 presents conclusions and policy implications of findings.

Key words: Wage differentials, wage structure, immigrant workers.

JEL codes: J31, J61

1. Introduction

The impact of labor market competition with immigrants on the wages and employment of the native born is one of the most contentious issues surrounding immigration policy debates. In many national contexts, immigrants tend to be relatively low skilled and supply their labor to occupations often filled by the lowest native earners in the host country. To be sure, and at least in the South-North context, the greater the differences are between the skill distributions of natives and immigrants, the greater are the net aggregate welfare gains accruing to natives in the host country (Smith and Edmonston, 1997) – i.e., the gains from trade are largest when the differences in relative endowments (skills) are the largest. However, the aggregate gain likely masks heterogeneity in the distribution of effects.

The distribution of labor market effects of immigration depend on a number of factors pertaining to the nature of the immigrant flow and the structure of economic production in the receiving economy. To begin, the resulting impact in the labor market on native wages will be largest when the aptitudes of immigrants are generally unlike those of native-born workers. With differences in the skill distributions, one would predict that those native workers who most resemble immigrant in terms of standard measures of human capital (e.g., educational attainment, work experience level) would suffer the largest adverse labor market effects. Those native workers that are least like immigrants would experience the greatest gains. One would expect larger impacts the larger the size of the immigrant population, holding all else equal.

Another important determinant of the impact of immigration on native wages concerns the ability of employers to substitute between workers of different skill levels. The degree of substitutability between skill groups mediates the impact of immigration on native wages in a manner that one cannot sign a priori. For example, low substitutability between immigrants and natives with similar levels of observed human capital results in a larger negative effect of new immigration on the wages of previous immigrants and smaller effects on similarly skilled natives (Ottaviano and Peri, 2007). Such imperfect substitution may occur when immigrants and natives speak different languages or when immigrants originate in nations with different social norms that are differentially valued in the host labor market. Alternatively, low substitutability between workers of different levels of educational attainment often means that immigrant flows that are imbalanced along this dimension will have large concentrated impacts on natives that look most like immigrants (Card, 2009). The assumed low-substitutability prevents the diffusion of the supply shock to other education groups, thus concentrating adverse effects on those natives in education groups disproportionately exposed to the migrant flow.

In this paper, we calibrate a structural model of the native wage distribution to the economies of two countries: Costa Rica and the United States. We then use these empirical models to simulate the likely effects of labor market competition with immigrants on native wages. The two chosen nations provide an interesting contrast along a number of dimensions. First, both nations have relatively large foreign born populations, constituting roughly twelve percent of the U.S. resident population and eight percent of the Costa Rican resident population in 2005 (Marquette, 2007). Second, in both countries immigrant labor is concentrated in the agricultural and service sectors. Finally, the immigrant populations of the two nations are disproportionately comprised of immigrants from countries with much lower per-capita GDP than of the host country.

The experiences of these countries also differ along a number of dimensions. First and foremost, the U.S. per capita income is over four times that of Costa Rica. Hence, our analysis compares the impacts of a large South-North migration between developing countries and the world's largest developed economy with that of a South-South migration between developing countries[1]. While the South-North immigration phenomena has been extensively analyzed, the South-South dimension has yet to be fully explored (Hujo and Piper, 2007). Second, while the majority of immigrants to the U.S. come from countries where English is not the primary language, the overwhelming majority of immigrants to Costa Rica comes from Spanish-speaking countries. Finally, we observe a much greater degree of balance in the native and immigrant skill distributions in Costa Rica relative to the U.S.

Our empirical analysis yields several interesting findings. First, the Costa Rica data yield very little evidence of imperfect substitutability between immigrants and natives of similar observable human capital levels. In contrast, there is fairly strong evidence of imperfect substitutability between immigrants and natives in the U.S. A plausible interpretation of this finding is that the ability to speak the receiving nation's language is an important determinant of the degree of substitutability between otherwise similar immigrants and natives.

The simulation results yield little evidence of an impact of immigration on native wages in Costa Rica, and suggest modest effects on the least skilled natives in the U.S. The lack of an impact on Costa Rican natives is driven largely by the balance in the educational attainment skill distributions of immigrants and natives in this country. For the U.S. the negative effects observed for high school dropouts depend critically on the assumed substation elasticity between education groups. Employing the parameter implied by U.S. census data yields simulated adverse wage effects of immigrant competition that are negligibly small. Our results

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[1] Specifically, the World Bank classifies Nicaragua as a middle income country while Costa Rica is classified as a upper middle country.

demonstrate how these results change when we assume less substitutability across education groupings.

2. Immigration to Costa Rica and the United States[2]

The recent past has observed increases in the proportions of the resident populations that are foreign born in both countries under analysis. Figures 1 and 2 document growth in the proportion foreign born overall and among prime age working residents. Prior to 1997, the main Costa Rican household surveys did not ask about the nativity of the respondent. Thus, Figure 1 is limited to the period beginning in 1997. For the United States, we graph the proportion foreign born for each decade from 1970 and for 2005.

In both nations, the proportion foreign born rises steadily over the time period depicted. In Costa Rica, the foreign born account for eight percent of the resident population in 2006, compared with six percent in 1997. In the U.S. the immigrant population increased from roughly 4.8 percent in 1970 to 12.4 percent in 2005. In both countries, the proportion foreign born is higher among prime age working residents. This likely reflects economically-motivated migration and is supported by previous research (Marquette, 2007; Vargas, 2005).

Table 1 compares the distribution of the immigrant populations in the two countries by country of origin. Over 85 percent of immigrant Costa Rican residents come from Latin American countries where Spanish is the official language. Nearly 70 percent of the immigrant population is from neighboring Nicaragua, the nation with which Costa Rica shares a large geographic border to the north. By contrast, most immigrants to the U.S. do not come from English speaking countries. The data in Table 1 suggest that at least 85 percent of the foreign

[2] This section uses data obtained from the US. Public Use Micro Data Sample of the US Census of Population and the Encuesta De Hogares de Propósitos Múltiples collected by the National Institute of Census and Statistics, Costa Rica.

born in the U.S. come from non English-speaking countries. The largest single source country is Mexico, with nearly one-third of the U.S. foreign born of Mexican origin. However, the rest of Latin America (22 percent of the foreign born) and Asia (25 percent of the foreign born) are also important source areas of immigrants.

As our introductory discussion indicated, the impact of immigration on native wages will depend on the degree to which the skill distributions of immigrants and natives overlap. The two most common gauges of human capital are educational attainment and potential work experience (usually gauged as age minus years of schooling minus six). Table 2 compares the educational attainment distributions of the native and foreign born for both countries among those who are working and who have 40 or fewer years of potential work experience. Given the large differences in the distribution of educational attainment between the two countries and differences in the survey questions pertaining to education, we group the data differently for each nation. For Costa Rica, we define the four educational attainment groupings less than primary, primary school, secondary school, and university. Individuals fall into these categories if they have some education at this level – e.g., those with one year of university education are placed in the university category. For the U.S., we define the four groups as less than high school, high school graduate, some college, and college graduate. This categorization is the standard used in the immigration literature (Borjas, 2003; Ottaviano and Peri, 2007; Raphael and Ronconi, 2009). Our categorization for Costa Rica is largely driven by the structure of the question in the Costa Rican household survey.

There are several patterns in Table 2 that merit discussion. First, ignoring those with less than a primary education (less than 2 percent of working Costa Rican natives), the distribution of the immigrant and natives in Costa Rica is quite balanced. The foreign born are only slightly over-represented among those with only primary education and moderately under-represented among those with secondary and university education. Immigrants constitute a large proportion of the least educated (0.35). However, very few working Costa Rican natives fall in this category as do less than six percent of the foreign born.

There is greater imbalance between the educational attainment distributions of foreign and native-born residents of the U.S. For example, while less than ten percent of natives have less than a high school degree, fully 29 percent of the foreign born fall into this category (with the foreign born constituting 28 percent of those with less than a high school degree). The foreign born are moderately under-represented among those with a high school degree and very under-represented among those with some college. The proportions with a college degree or more are quite similar for immigrants and natives.

Figures 3 and 4 compare the age distributions among working residents by nativity for each country. As age is the key determinant of potential experience, these figures are meant to convey the degree of overlap (or lack thereof) along this human capital dimension. There is a remarkable degree of similarity across the two countries. In both figures the immigrant age distribution is more heavily concentrated among those between their early twenties and early forties (an age range with relatively high degree of attachment to the labor force). Natives on the other hand are more heavily concentrated among the relatively young and the relatively old.[3]

This empirical portrait yields several findings that are of direct importance in our more formal analysis to follow. First, in the U.S. immigrants are quite likely to come from a country where the official language is not English while in Costa Rica, most immigrants are Spanish speaking. This fact alone suggests that otherwise similar natives and immigrants are likely to be

[3] We also produced comparable density estimates for the entire age distributions irrespective of labor market behavior. The relative concentration of immigrants is similar in both countries to what is depicted in Figures 3 and 4.

more substitutable in Costa Rica relative to the U.S. Second, while immigrants in Costa Rica are slightly less educated than natives, the educational attainment distributions for these two groups are otherwise quite similar to one another. Combined with endogenous capital accumulation, such balanced shocks that do not appreciably alter a nation's factor proportions are unlikely to yield large impacts on the national wage structure. Finally, both countries exhibit a high concentration of immigrants among those of prime working age. The impact of this concentration will depend on the degree of substitutability between workers of different experience levels. To gauge these substitution possibilities, we now turn to our model of the aggregate economy and the empirical methods we use to calibrate the model to each of the countries analyzed here.[4]

3. The Model

In this section, we lay out the aggregate model of the national labor market. We begin by discussing the aggregate production function and the relationship between factor supplies and wages. We then derive the theoretical expressions for the factor price elasticities needed to simulate the impact of the empirically observed immigrant labor supply shock. Finally, we layout the details of the estimation of Equations used to estimate the substitution elasticities across different dimensions of skill.

[4] There is a large literature on the impact of immigration on native wages in the United States (Raphael and Ronconi, 2007). This literature tends to fall into two methodological strands: research designs that exploit regionally concentrated labor supply shocks (Card, 1990; 2005; Friedberg, 2001; Hunt, 1992) and, and research designs that rely on national level models to simulate wage effects (Borjas, 2003; 2005; Ottaviano and Peri, 2007) The regional shock research generally finds quite modest effects of immigration on wages. National level simulation studies are quite sensitive to the parameter values chosen in the model calibration. We discuss this issue in greater detail below. Research on the impact of immigration on the wage structure of developing countries has been more limited. In the particular case of Costa Rica, analysis of the impact of immigration has been centered on issues associated with health, poverty levels, and education among others (Marquette, 2007; Vargas, 2005). We are aware of only one other analysis of the impact of immigration on wages of native Costar Ricans (Gindling, 2009).

Complimenting the analysis done elsewhere we assume that overall production in each economy is described by the multi-layer constant elasticity of substitution (CES) production function (Borjas, 2003; Card and Lemieux, 2001; Ottaviano and Peri, 2007). In particular, production in each economy is modeled by

(1) $$Q_t = [a_{0t} K_t^{\nu} + a_{1t} L_t^{\nu}]^{\frac{1}{\nu}}, \quad where \quad \nu = 1 - \frac{1}{\sigma_{KL}}$$

(2) $$L_t = \left(\sum_{k=1}^{4} e_{tk} L_{tk}^{\eta} \right)^{\frac{1}{\eta}}, \quad where \quad \eta = 1 - \frac{1}{\sigma_{educ}}$$

(3) $$L_{tk} = \left(\sum_{j=1}^{8} x_{tkj} L_{tkj}^{\delta} \right)^{\frac{1}{\delta}}, \quad where \quad \delta = 1 - \frac{1}{\sigma_{\exp}}$$

(4) $$L_{tkj} = \left(\sum_{i=1}^{2} m_{tkji} L_{tkji}^{\varepsilon} \right)^{\frac{1}{\varepsilon}}, \quad where \quad \varepsilon = 1 - \frac{1}{\sigma_{immig}}$$

where t indexes time, k indexes four labor groups defined by educational attainment (using the categorizations described in Table 2), j indexes eight potential years of experience groups (0 to 4, 5 to 9, 10 to 14, 15 to 19, 20 to 24, 25 to 29, 30 to 34, 35 to 40) that are defined equivalently for each country, and i indexes nativity $(1 = native, 2 = immigrant)$. As we apply this model to each country separately, we omit country-specific subscripts to simplify notation.

Equation (1) combines capital and total labor in year t to produce national output Q_t, where a_{ot} and a_{1t} are productivity coefficients for capital and labor, respectively, and σ_{KL} is the elasticity of substitution between capital and labor. In turn, the total labor supply aggregate, L_t, is a CES aggregation of sub-categories of labor defined by the four educational groups, L_{tk}, given by Equation (2) where the e_{tk} provide the corresponding productivity coefficients and σ_{educ} is the elasticity of substitution between education groups. The labor supply of each educational group, L_{tk}, is further assumed in Equation (3) to be a CES aggregation of labor supply for each of the eight experience groups, L_{tkj}, with corresponding productivity coefficients x_{tkj} and an elasticity of substitution between experience groups within an education branch given by σ_{exp}. Finally, labor supplied within a given education-experience cell is assumed to be a CES aggregation of native labor, L_{tkj1}, and immigrant labor, L_{tkj2}, with a corresponding elasticity of substitution between immigrants and natives given by σ_{immig} and productivity coefficients given by m_{tkji}. In the event that immigrants and natives are perfect substitutes, the production function collapses to that described by the first three Equations only.

The wages of workers in group *tkji* are determined by their marginal product, which in turn will depend on the supply of capital, the overall supply of labor, the supply of labor in education group *tk,* the supply of labor in education-experience group *tkj*, and their own-supply of labor L_{tkji}. Assuming a product price of one, the wage is determined by the Equation

(5) $$w_{tkji} = Q_t^{1-\nu} a_{1t} L_t^{\nu-\eta} e_{tk} L_{tk}^{\eta-\delta} x_{tkj} L_{tkj}^{\delta-\varepsilon} m_{tkji} L_{tkji}^{\varepsilon-1},$$

where the right hand side of Equation (5) is derived by taking the derivative of the production function with respect to L_{tkji}. Given the relationship between the parameters in the exponents of the terms on the right-hand-side of (5), we can express these Equations in terms of the four elasticities of substitution. Taking logs and substituting in the substitution elasticities gives

(6) $$\ln w_{tkji} = \ln Q_t^{1-\nu} a_{1t} + \left[\frac{1}{\sigma_{educ}} - \frac{1}{\sigma_{KL}}\right]\ln L_t + \ln e_{tk} + \left[\frac{1}{\sigma_{\exp}} - \frac{1}{\sigma_{educ}}\right]\ln L_{tk} + \ln x_{tkj} + \left[\frac{1}{\sigma_{immig}} - \frac{1}{\sigma_{\exp}}\right]\ln L_{tkj} + \ln m_{tkji} - \frac{1}{\sigma_{immig}}\ln L_{tkji}.$$

Assuming that the supply of labor is perfectly inelastic within the year-education-experience-nativity cell of Equation (6), various specifications of Equation (6) can be used to estimate the substitution elasticities underlying the wage determination process. The Equation highlights how an increase in own factor supply suppresses wages, all else constant. The Equation can also be used to derive the effect on wages of a factor supply change for other nativity groups with similar skills (operating through L_{tkj}), of other experience groups within one's education group (operating through L_{tk}), and of other education groups (operating through the effect on L_t and capital accumulation). These cross-group wage elasticities are particularly important since in simulating the impact of immigration on the wage of any group, we would want to account for both the own-groups supply shock as well as the impact of supply shocks occurring elsewhere in skill distribution.

How immigration impacts the wages of specific native skill groups

Immigration over a given time period impacts the wages of a native in a specific skill group through four avenues: (1) its impact on the supply of workers within her year-education-experience cell, (2) its impact on the supply of workers in her year-education cell, (3) its effect on the overall aggregate supply of labor, and (4) its impact on capital accumulation. The impact of a skill-specific supply shock on a specific skill group of natives, however, will vary by education and experience group. An increase in immigrants within one's own year-education-experience group impacts wages through all four channels. Immigration within one's education group but outside one's experience group affects one's wages only through avenues (2) through (4), since the aggregate supply within one's year-education-experience cell is not impacted. Finally, immigration shocks to skill groups outside one's educational group impacts own wages only through the impact on aggregate labor supply and capital accumulation. Since the capital accumulation effect on wages is always positive (increasing capital increases the capital-labor ratios which unambiguously increases everyone's wages ceteris paribus), immigration shocks to skill cells that are closest to that of a specific native will have the largest negative effects while shocks to skill groups that are most unlike the specific native in question (outside one's education group) will have very small negative or even positive wage effects.

Following previous work, we assume that the economy is on its long-run balanced growth path (Borjas, 2005; Ottaviano and Peri, 2007), implying that capital accumulates at the rate needed to ensure a constant return to capital. Under this assumption regarding the growth path of the capital stock, Equation (6) becomes

(7) $$\ln w_{tkji} = \ln B_t + \frac{1}{\sigma_{educ}} \ln L_t + \ln e_{tk} + \left[\frac{1}{\sigma_{\exp}} - \frac{1}{\sigma_{educ}}\right] \ln L_{tk} + \ln x_{tkj} + \left[\frac{1}{\sigma_{immig}} - \frac{1}{\sigma_{\exp}}\right] \ln L_{tkj} + \ln m_{tkji} - \frac{1}{\sigma_{immig}} \ln L_{tkji}.$$

where the term B_t will be a function of the constant return to capital but varies over time due to technological progress.[5] Thus an increase in the supply of any labor skill group will induce a positive effect on wages through capital accumulation, partially offsetting any decrease in wages due to the greater labor supply.

To derive the full effects of a specific immigration-induced supply shock on the wages of a given native skill group, we must differentiate the log wage for native group *tkj1* (as given in

[5] Specifically, assume that the first level of the CES production function is defined by the Cobb-Douglas production function $Q_t = A_t K_t^{\alpha} L_t^{1-\alpha}$. If capital accumulates to maintain a constant return to capital of r, the production function reduces to $Q_t = B_t L_t$ where $B_t = \left[\frac{r}{A_t \alpha}\right]^{\frac{\alpha}{\alpha-1}}$.

Equation (7)) with respect to a change in immigrant supply within the same education-experience cell, with respect to immigrant supply within one's education group but outside one's education-experience cell, and with respect to immigrant supply outside one's education group. These three wage elasticities are given by the expressions

(8) $$\varepsilon_{own} = \frac{\partial \ln w_{tkj1}}{\partial \ln L_{tkj2}} = \frac{1}{\sigma_{educ}} \frac{s_{tkj2}}{s_t} + \left[\frac{1}{\sigma_{\exp}} - \frac{1}{\sigma_{educ}} \right] \frac{s_{tkj2}}{s_{tk}} + \left[\frac{1}{\sigma_{immig}} - \frac{1}{\sigma_{\exp}} \right] \frac{s_{tkj2}}{s_{tkj}}$$

(9) $$\varepsilon_{cross-\exp} = \frac{\partial \ln w_{tkj1}}{\partial \ln L_{tkj'2}} = \frac{1}{\sigma_{educ}} \frac{s_{tkj'2}}{s_t} + \left[\frac{1}{\sigma_{\exp}} - \frac{1}{\sigma_{educ}} \right] \frac{s_{tkj'2}}{s_{tk}}, \quad where \quad j \neq j'$$

(10) $$\varepsilon_{cross-educ} = \frac{\partial \ln w_{tkj1}}{\partial \ln L_{tk'j2}} = \frac{1}{\sigma_{educ}} \frac{s_{tk'j2}}{s_t}, \quad where \quad k \neq k'$$

where s_t is labor's share of income in year *t*, s_{tk} is the share of income accruing to labor in education group *k* in year *t*, s_{tkj} is the share of income accruing to labor in group *tkj*, and s_{tkj2} is the share of income accruing to immigrant labor in group *tkj*.

The factor-price elasticities given by Equations (8) through (10) are the key inputs to the wage simulations that we perform below. Define the variable M_{kj} as the percentage increase in immigrant supply between a given starting and ending period and the column vector M as the complete set of shocks for the 32 education-experience groups. Using the elasticites in Equations (8) through (10), we can construct a square elasticity matrix Π where the rows are defined by the education-experience group of natives for whom we wish to analyze wage effects, the columns are defined by the education-experience group experiencing an immigrant labor supply shock. Elements of the matrix where $k_{row} = k_{column}$ and $j_{row} = j_{column}$ are given by the own-elasticity in Equation (8), elements of the matrix where $k_{row} = k_{column}$ and $k_{row} \neq k_{column}$ are given by the cross experience group elasticity in Equation (9), while elements of the matrix where $k_{row} \neq k_{column}$ are given by the cross education group elasticity in Equation (10). With this matrix and the supply shock vector, the vector of simulated effects caused by the vector of empirically observed immigration shocks between the start and end dates are given by the Equation

(11) $$Wage\ Effect = \Pi M,$$

where the individual elements of this vector are given by the expression

(12) $$Wage\ Effect_{kj} = \frac{1}{\sigma_{educ}} \sum_{k=1}^{4} \sum_{j=1}^{8} \frac{s_{tkj2}}{s_t} M_{kj} + \left[\frac{1}{\sigma_{\exp}} - \frac{1}{\sigma_{educ}} \right] \sum_{j=1}^{8} \frac{s_{tkj2}}{s_{tk}} M_{kj} + \left[\frac{1}{\sigma_{immig}} - \frac{1}{\sigma_{\exp}} \right] \frac{s_{tkj2}}{s_{tkj}} M_{kj}$$

As is evident in Equation (12), immigrant supply shocks to one's one education-experience group impacts ones wages through all three terms, shocks within one's education group but outside one's experience group only affect the first two terms, while shocks outside of one's education-experience cell only impact wages through the first term.

Calibrating the model

To implement the wage simulations described by Equations (11) and (12), we need three pieces of information. First, we need to estimate the proportion of labor income accruing to various aggregations of labor input (i.e., we need estimates of the ratios s_{tkj2}/s_{tkj}, s_{tkj2}/s_{tk}, and s_{tkj2}/s_t). We estimate these relative income shares directly from household surveys for the two analysis countries. For the U.S. the shares are estimated as of 2005 while for Costa Rica we estimate relative shares for 2006.

Second, we need to characterize the magnitude of the immigration shock for each education-experience cell. For the U.S., we define the immigrant shocks M_{kj} as the difference between the immigrant supply level (the measurement of which is discussed below) in 2005 less the immigrant supply level in 1970, all divided by the immigrant supply level in 2005. Thus, we are simulating the effect of reducing immigration to 1970 levels relative to 2005 labor supplies. For Costa Rica, we simulate the impact of reducing the 2006 immigrant population to (1) 1997 levels and (2) to zero.

Finally, performing these wage simulations requires that we choose values for the three substitution elasticities σ_{immig}, σ_{exp}, and σ_{educ}. We turn now to a discussion of how we estimate these elasticities.

Equation (7) above gives an expression for log wages for those in group *tkji* and demonstrates the inverse relationship between own factor supply and wages. Estimating the substitution elasticity between immigrants and natives can be done as follows. For immigrants and natives in the same year-education-experience groups, the first six terms in Equation (7) are the same. Thus subtracting the log of immigrant wages from log native wages within the same skill group eliminates these common terms and yields the expression

(13)
$$\ln \frac{w_{tkj1}}{w_{tlk2}} = \ln \frac{m_{tkj1}}{m_{tkj2}} - \frac{1}{\sigma_{immig}} \ln \frac{L_{tkj1}}{L_{tkj2}}$$

indicating that the relative log wages of natives and immigrants should vary inversely with relative log factor supplies. Equation (13) highlights an important intuition. Specifically, if immigrants and natives within a specific education-experience cell are imperfectly substitutable, an increase in the relative supply of natives should lead to a decrease in the relative wages of natives. The decrease will be larger the smaller the elasticity of substitution between immigrants and natives (σ_{immig}). On the other hand, if immigrants and natives are perfect substitutes within skill group (i.e., $\sigma_{immig} = \infty$) an increase in relative native supply will suppress immigrant and native wages equally, yielding no relationship between relative supplies and relative wages. An equivalent statement of this relationship is that the impact of an immigration-induced supply shock on the wages of comparably skilled natives will be larger the larger the substitution elasticity between immigrants and natives. With low substitutability, the wage impacts of the supply shock will be concentrated on the own wages of immigrants.

Thus, the inverse of the coefficient on relative supplies in Equation (13) provides an estimate of the elasticity of substitution between immigrants and natives. Below we estimate various specifications of Equation (13) where we use different combinations of year, education, and experience fixed effects to proxy for the ratio of productivity coefficients, $ln(m_{tkj1}/m_{tkj2})$. We assume that after adjusting for these fixed effects, the remaining variation in relative supplies is exogenous.

Estimating the elasticity of substitution between experience groups requires aggregating immigrant and native labor into the higher aggregate labor supply units. The average wage paid to immigrants and natives will equal the marginal effect of an increase in this labor aggregate on total output (calculated by differentiating Q with respect to L_{tkj}). Taking this derivative[6] and taking logs yields the wage expression

$$(14) \qquad \ln w_{tkj} = \ln B_t + \frac{1}{\sigma_{educ}} \ln L_t + \ln e_{tk} + \left[\frac{1}{\sigma_{\exp}} - \frac{1}{\sigma_{educ}} \right] \ln L_{tk} + \ln x_{tkj} - \frac{1}{\sigma_{\exp}} \ln L_{tkj}.$$

Thus, a regression of log wages for group *tkj* on a series of appropriate fixed effects and own-factor supply yields an estimate of $-1/\sigma_{\exp}$. To identify the elasticity of substitution across education groups, note that (14) can be rewritten as

$$\ln w_{tkj} = \ln B_t + \frac{1}{\sigma_{educ}} \ln L_t + \ln e_{tk} - \frac{1}{\sigma_{educ}} \ln L_{tk} + \ln x_{tkj} - \frac{1}{\sigma_{\exp}} [\ln L_{tkj} - \ln L_{tk}].$$

[6] The derivative assumes that capital accumulates endogenously to hold the return to capital constant.

(15)

where the coefficient on aggregate labor supplied in each education group provides an estimate of $-1/\sigma_{educ.}$

Estimating Equation (14) requires the inclusion of several sets of fixed effects to account for the first five terms on the right hand side of the Equation. The first two terms vary with time only and thus can be captured by a series of time effects. The third and fourth terms vary with time and education group and thus can be captured by time-education fixed effects. The term $\ln x_{tkj}$ varies across all observations, and thus an identifying restriction is needed. We assume, in line with previous research, that these effects vary by education and experience groups but do not vary over time (Borjas, 2003; Ottaviano and Peri, 2007). Thus, we estimate the key coefficient in Equation (14) with the regression model

(16) $$\ln w_{tkj} = \beta_t + \pi_{tk} + \theta_{kj} - \frac{1}{\sigma_{\exp}} \ln L_{tkj}$$

We estimate Equation (16) using instrumental variables where the log immigrant supply is used as an instrument for $\ln L_{kjt}$

To estimate the cross-education group elasticity using Equation (15) we must again impose some identifying restrictions on the first few terms of the Equation. First, we need to calculate the aggregate supply values, L_{tk}. With an estimate of the elasticity of substitution between experience groups, one could construct this aggregate from the third level of the CES production function given by Equation (3) above. However, previous research has found that estimates that simply sum up the supply measures from the next level of dis-aggregation tend to yield nearly identical results. Here we measure L_{tk} by simply summing across experience groups within year-education cells.

Since the key variable in this model varies by year and education group only, we cannot include a full set of year-education fixed effects. Instead we parameterize the relationship between education groups, time and wages by proxying the first three terms in Equation (15) with a set of year fixed effects and education group specific linear time trends. Again we impose the restriction that $\ln x_{tk}$ varies across education and experience groups but not by time. Thus, we estimate the model.

(17)
$$\ln w_{tkj} = \beta_t + \pi_k t + \theta_{kj} - \frac{1}{\sigma_{educ}} \ln L_{tk} - \frac{1}{\sigma_{\exp}} [\ln L_{tkj} - \ln L_{tk}].$$

where t is a time trend and π_k is a education group-specific trend coefficient. Note, the coefficient on the deviation in the third term provides an alternative estimate of $-1/\sigma_{exp}$. We estimate Equation (17) using instrumental variables where the two supply variables are instrumented with the corresponding values for immigrants.

4. Description of the Data

Estimating Equations (13), (16) and (17) requires year and skill-groups specific data on wages and factor supplies. Our simulations also require estimates of the size and distribution of immigration induced supply shocks. In this section, we describe the data used for each country and the specifics of how we measure wages and labor supply. While certain particulars of the data force us to stray from estimating the exact same models for both countries, we make a series of specification choices that maximize the comparability of the results.

The U.S. Data

Our U.S. data set covers the period from 1970 to 2005. We use data from the one percent 1970 Public Use Microdata Sample (PUMS) of the U.S. Census of Population and Housing, the five percent PUMS census data for 1980, 1990 and 2000, and data from the 2005 American Community Survey (ACS). Each sample is nationally representative and includes micro-level information on wages, labor supply, and nativity. Our chosen time frame (1970 through 2005) corresponds to a large increase in immigration to the United States following the passage of the 1965 Immigration Act.

In each year, we restrict the analysis to individuals ages 17 to 65 who do not reside in institutional group quarters, who have positive weeks worked, who work positive hours during the interview week (for 1970 through 1990) or who indicate that they usually work a positive number of hours per week (2000 and 2005), and who have positive values for annual wage and salary income. Our definition of education groups is as described in Table 2 while our definition of experience groups are the eight groups discussed in the previous section. Experience is defined as the individual's age less years of schooling, less six. The analysis is restricted to those with 40 years of experience or less. We use the microdata to estimate group specific wages and factor supplies for each year in our analysis period.

To measure wages, we calculate average weekly wages for males who usually work 35 hours or more (using either hours prior to the survey week or usual hours depending on the survey year). Following Borjas, Grogger and Hanson (2008), We follow previous research in using average weekly wages for full-time males to ensure that variation in the wage measure does not reflect variation in hours worked since the model relates factor supplies to prices (Borjas et al., 2008).

To measure labor supply, we aggregate total hours supplied to the market within various labor sub-aggregates using the entire sample of workers (i.e., not restricted to full time males). To calculate total hours, we calculate annual hours worked for each person, multiply by the survey weight, and then sum within year-education experience categories.

The Costa Rica Data

The data for the Costa Rican labor market comes from the Encuesta de Hogares de Propósitos Múltiples (EHPM) for each year from 1997 to 2006. The EHPM is a nationally representative household survey that includes information on labor market activity, wages, measures of human capital, and nativity. In addition, for most years a special panel on particular topics is included in the survey. These topics range from internet use to child labor and elderly heath issues among others. By 1997, there was already a numerically substantial immigrant population in Costa Rica. However, between 1997 and 2006 this population increased considerably in large part due to increasing migration from neighboring Nicaragua. While we have micro data extending back to 1989, the EHPM did not collect information on nativity prior to 1997. Thus, the analysis sample we use to estimate Equations (13), (16), and (17) focuses on the period from 1997 to 2006.

Given the lower average level of education attainment in Costa Rica and the younger ages at which many leave school, we restrict the analysis sample to those who are 14 years are over. We also restrict the analysis sample to those with 40 years of potential experience or less.

To measure year-education-experience specific wages, we first restrict the sample to males who work at least 40 hours per week. We use the higher hours threshold for Costa Rica due to the fact that full time (and the modal hours worked) is defined by 48 hours per week in contrast with the 40 hours per week threshold in the U.S. The EHPM provides earnings at the monthly level, where monthly hours are tabulated for those reporting weekly earnings by multiplying by 4.33. This imputed income at the monthly level is the only earnings field presented in the public use files. Thus, in contrast to the U.S. data, we use monthly earnings for full time males as the key dependent variable.

To measure factor supply, we multiply usual weekly hours by 4.33 to convert the data to usual monthly hours. We then multiply monthly hours by the survey weight to gauge the aggregate labor supply of workers presented by the observation included in the sample. Summing this variable within year-education-experience-nativity cells provides our gauge of factor supply. Similar to our treatment of the U.S. data, we use all workers with positive earnings and hours supplied (i.e., not restricted to full time men) to measure labor supply.

5. Country-Specific Estimate of the Labor Elasticities of Substitution

In this section, we present estimates of the key substitution elasticities that are used to calculate the responsiveness of native wages to immigration induced changes in supply. We begin with an analysis of the substitutability of immigrants and natives within the same education-experience group. Figures 5 and 6 present simple bivariate scatter plots of the log wage differentials between immigrants and natives against the comparable log supply differentials. Each data point represents one year-education-experience cell.

For Costa Rica, there is no evidence of a relationship between native-immigrant relative wages and native-immigrant relative factor supplies. While the plotted bivariate regression line has a slight negative slope, the coefficient from this regression is very small and statistically insignificant. In contrast, the U.S. data exhibits a clear negative relationship. The slope of the regression line in Figure 6 of -0.03 is statistically significant at the one percent level and implies an elasticity of substitution between similarly skilled immigrants and natives in the U.S. of 33.33.

Note the right-hand side of Equation (13) includes a term that gives the log of the ratio of the productivity coefficients for immigrants and natives. The scatter plots in Figures 5 and 6 implicitly assume that this productivity coefficient ratio is constant over time and across skill groups. To relax this assumption, Table 3 presents a series of model estimates where the dependent variable is the native-immigrant log wage differential, the key explanatory variable is the native-immigrant log supply differential, and the productivity coefficients ratio is proxied by different combinations of year, education, and experience fixed effects. The first three specifications consecutively add year, education, and experience fixed effects. The fourth specification drops the experience effects but adds experience-year effects. The final three specifications consecutively add experience effects, experience-year effects, and experience-education fixed effects. The standard errors in all models allow for clustering of the error variance-covariance matrix within education-experience cells and are thus robust to heteroscedasticity and serial correlation within skill groups.

Beginning with the results for Costa Rica in Panel A, all of the estimates are statistically insignificant in each model. While the standard errors are quite large in the most liberal specifications, the general impression drawn from the scatter plot and these more extensive model specifications is that there is little evidence of imperfect substitutability between similarly skilled natives and immigrants. In contrast, the first three U.S. models in Panel B yield statistically significant coefficients ranging from -0.03 to -0.055. While most of the more extensive specifications yield statistically insignificant results, the standard errors are quite large and cannot rule out fairly low values of the substitution elasticity between immigrants and natives.

Thus, we find no evidence of imperfect substitutability among immigrant and natives in Costa Rica and fairly strong evidence of a modest degree of imperfect substitutability in the U.S. One plausible interpretation of this difference is that the common primary language of immigrants and natives in Costa Rica and the disparity in English language ability between immigrants and natives in U.S. is driving this pattern. In recent research by Peri and Sparber *forthcoming*, the authors show that in the U.S. context similarly skilled immigrants and natives tend to specialize in occupations with different skill sets, with natives specializing in occupations requiring cognitive and verbal skills and immigrants specializing in occupations requiring manual skills[7]. Moreover, this research finds that the degree of specialization for each group is greatest the greater the relative supply of immigrants. This research suggests strongly that otherwise similar immigrants and natives have different relative skill endowments with host language ability being particularly important. The fact that immigrants and natives appear to be perfect substitutes when host language ability is held constant (the Costa Rican case) certainly lends further support to this line of reasoning.[8]

In Table 4, we turn to estimations of the elasticity of substitution across experience groups with similar levels of educational attainment. Recall, this substitution elasticity comes from a regression of the log of average wages at the year-education-experience level on the log

[7] Other researchers who have analyzed the US have found that when occupation is used as a proxy for skill an increase in the fraction of foreign born workers tends to lower the wages of natives in some blue collar occupations but that the effect is not statistically significant among natives in skilled occupations (Orrenious and Zavodny, 2007).

[8] Previous research for the U.S. found estimates of the immigrant-native elasticity of substitution on the order of 6 (Ottaviano and Peri, 2007) with similar results for Great Britain (Manacorda and Wadsworth, 2006) Other studies find no relationship using full-time full-year workers not enrolled in school to estimate the relative wage ratio in (13) (Borjas, Grogger and Hanson, 2008). Finally, estimates of $-1/\sigma_{immig}$ on the order -0.03 and using a cross section of metropolitan areas from the U.S. census was reported (Card, 2008).

supply measures at the comparable level. Our full specification of this Equation in (16) includes an extensive set of fixed effects to proxy for the other terms in the structural Equation given in (14). In Table 4, we provide three specifications of increasing complexity. The first includes year, education, and experience fixed effects. The second adds year-education interactions while the final specification includes year-education and education-experience interaction terms. Note, the final specification corresponds to that in Equation (16). Following previous research we use immigrant labor supply in the cell as an instrument for overall labor supply (Borjas, 2003; Ottaviano and Peri, 2007). Standard errors are clustered on the education-experience cells.

The results for Costa Rica suggest a high degree of substitutability between workers of different experience levels. The first two specifications suggest a value for $\sigma_{\exp}$ of approximately 14 (1/.07) while the final specification suggests perfect substitutability between workers of difference experience levels. The coefficient in the final specification, however, has a fairly large standard error and thus cannot rule out the elasticity implied by the first two specifications. The results for the U.S. are inconsistent across specifications. The most parsimonious model in the first column yields a coefficient of the wrong sign. Adding year-education interaction terms yields a significant negative coefficient of -0.048 implying a substitution elasticity of approximately 21. The final complete specification gives a fairly precisely measured coefficient of -0.109. In our simulations below, we therefore assume a value for this elasticity for the U.S. of 10.

Finally, Table 5 presents estimates of the parameter needed to calculate the elasticity of substitution between education groups. For each country we present three specifications: (1) a model including education-group specific time trends, (2) a model including year fixed effects for a base education group, and education-specific linear time trends, and (3) a model including year fixed effect for a base education groups and linear and quadratic education-specific time trends. Recall, with the inclusion of year fixed effects and education-specific time trends, the specification allows the time path of wages for a particular education group to depart from the base year effects according to the trend terms. The key explanatory variables here are the labor supply measured at the year-education level and the difference between the labor supply measured at the year-education-experience level and at the year-education level. Recall from Equation (17) that the coefficient on the first provides an estimate of $-1/\sigma_{educ}$ while the coefficient on the difference term provides an alternative estimate of $-1/\sigma_{\exp}$. Again, we instrument the two labor supply terms with the comparable labor supply measures for immigrants only and allow for clustering of the error variance-covariance matrix by education-experience cells in the calculation of the coefficient standard errors.

The results for Costa Rica generally indicate low substitutability among workers of different education levels. As we move across the three specifications, the coefficient on labor supply measured at the year-education level increases, indicating decreasing estimates of the education elasticity of substitution. The three point estimates indicate a range for this parameter from 1.4 to 2.9. Although the coefficient are poorly measured (only the parameter value for the second specification is marginally significant), the patterns do suggest low substitutability in the Costa Rican labor market along this dimension. The alternative estimates of $\sigma_{\exp}$ generally conform with the estimates from the most extensive specification presented in Table 4.

The U.S. results are much less clear. The first specification yields an elasticity of substitution of approximately eight. The second specification yields an elasticity estimate of roughly 14, while the coefficient in the final specification is of the wrong sign. Like results from previous research there is difficulty in measuring this elasticity in U.S. (Card, 2009). The higher elasticity values for the U.S. may indeed reflect a greater degree of substitutability relative to Costa Rica. This might be the case if beyond some minimum level of educational attainment most workers are highly substitutable for one another. Given the much higher level of educational attainment in the U.S. (even for the least educated members of the workforce), this would provide a potential explanation for this disparity in parameter values.

Alternatively, wage trends by level of educational attainment in the U.S. over the time period studies may be sufficiently complex that our proxies for time trends may not be adequately capturing the myriad of factors that greatly increased U.S. wage inequality over this time period. In our simulations below, we present results with an elasticity of substitution between education groups implied by the parameter estimates in Table 5 along with alternative simulation results that employ lower degrees of substitution in order to highlight the sensitivity of the analysis to this particular parameter.

6. The Simulated Impacts of Immigration on Native Wages

The parameter estimates in the previous section provide the key inputs needed to compute the own and cross-input factor price elasticity given by Equations (8), (9), and (10). In conjunction with a vector of group-specific immigration shocks, these factor price elasticities permit us to simulate the impact of various counter-factual immigration scenarios. In this section, we first discuss our chosen parameters values based on the empirical work. We then discuss the immigration counterfactuals for each nation. Finally, we present the simulation results highlighting the impact of immigration on native wages.

Parameter Choices

The estimates of the elasticity of substitution between immigrants and natives differed by country. There is no evidence of imperfect substitution for Costa Rica, and consistent evidence of a finite elasticity for the United States. Thus, we assume an infinite substitution elasticity for the former and a value of 33 for the latter. Note, the value for the United States conforms with the empirical estimates presented in Table 3.

Given that we are assuming perfect substitutability between similarly skilled immigrants and natives in Costa Rica, the last labor dis-aggregation vanishes. In turn, the key factor price elasticities are now computed based on the derivative of log wages with respect to changes in labor supply at the year-education-experience level rather than the year-education-experience-nativity level. These alternative factor price elasticities are given by the Equations

(18)
$$\varepsilon_{own} = \frac{1}{\sigma_{educ}}\frac{s_{tkj}}{s_t} + \left[\frac{1}{\sigma_{\exp}} - \frac{1}{\sigma_{educ}}\right]\frac{s_{tkj}}{s_{tk}} - \frac{1}{\sigma_{\exp}}$$

(19)
$$\varepsilon_{cross-\exp} = \frac{1}{\sigma_{educ}}\frac{s_{tkj'}}{s_t} + \left[\frac{1}{\sigma_{\exp}} - \frac{1}{\sigma_{educ}}\right]\frac{s_{tkj'}}{s_{tk}}$$

(20)
$$\varepsilon_{cross-educ} = \frac{1}{\sigma_{educ}}\frac{s_{tk'j'}}{s_t}$$

where all terms are as defined above. In addition to this different set of formula for the factor-price elasticities, we must also redefine the vector of migration shocks for the purposes of simulating wage effects. In particular, rather than defining the immigration shock over a specific time period as the percentage increase over a base immigration supply level, now we must define the supply shock as the percentage increase in supply at the year-education-experience level caused by the immigrant inflow. Thus, for Costa Rica we use these alternative factor price elasticities and the alternative definition of the immigration shock vector to perform our wage simulations. For the U.S., we use the elasticities derived in Equations (8), (9), and (10).

Regarding the elasticity of substitution between experience groups, the Costa Rican regression results imply a value of at least 14, while the most extensively-specified model for the U.S. yields a value of 10. We experimented with several other values of this elasticity and found that the simulation results are generally not sensitive to this parameter choice. Hence, in all simulations we use these stated values.

Of particular importance is the value choice for the elasticity of substitution between education groups. The Costa Rican models suggest parameter values between 1.5 and 3. In all simulations below, we present two sets for Costa Rica using these two endpoint values of the range. The U.S. regression results suggest a relatively high degree of substitution between workers with different levels of educational attainment (on the order of 10). However, we also perform a simulations where we set this elasticity to 2.5 to illustrate the sensitivity of the results to this parameter choice. The value of 2.5 is consistent with values reported in previous research results (Card and Lemieux, 2001).

The Immigration Counterfactual

The factor-price elasticities provide us with estimates of the responsiveness of the wages of a given native skill group to changes in immigrant supply. To simulate the impact of immigration on wages, we must specify a counterfactual immigration scenario and the labor supply shock defined by the difference between reality and the counterfactual. For the U.S. we define the counterfactual in the following manner. Define M_{70kj} as the proportion of labor supply in education group k and experience group j from immigrants in 1970, define L_{05kj1} as the labor supply of natives in group kj in 2005, and L_{05kj2} as the labor supply of immigrants in group kj in

2005. Suppose that immigration policy between 1970 and 2005 was sufficiently stringent to hold immigration levels (and the distribution of immigrant across skill groups) to the 1970 value. Under this counterfactual, immigrant supply to group *kj* in 2005 would be given by the Equation

(21)
$$L^*_{05kj2} = L_{05kj1}\frac{M_{70kj}}{1-M_{70kj}}.$$

The proportionate effect on immigrant labor supply of reducing immigrant supply to this counterfactual level is then given by $(L_{05kj2} - L^*_{05kj2})/L_{05kj2}$. We calculate this hypothetical shock for each skill group. Based on these supply shocks, the wage simulations for the U.S. show how the increase in immigrant labor supply above and beyond what would have occurred had immigrant proportions been held to 1970 levels would have impacted native wages.

For Costa Rica, we explore the impact two counterfactual scenarios. We calculate the first vector of shocks by tabulating the proportion effect on the labor supply of a specific skill group of reducing the immigrant population to 1997 levels (the beginning period of our data). In a second set of simulations, we tabulated the proportionate effect on labor supply of reducing the immigrant population to zero.

Wage Simulations Results

Table 6 present several sets of simulations. For Costa Rica, we simulate the effects of reducing the immigrant population to 1997 levels as well as the effects of reducing immigrant labor supply to zero. For each counterfactual we report two sets of results, one employing an elasticity of substitution between education groups of three and the other setting this parameter value to 1.5 For the U.S., we simulate the effects of holding immigrant labor supply proportionally to 1970 levels. Again, there are two sets of results, one employing an elasticity of substitution across education groups of 10 and the other using the value of 2.5.

The results for Costa Rica are surprisingly consistent across simulations. For the overwhelming majority of native-born Costa Ricans (the approximately 98 percent in the top three education categories), immigration has essentially no effect on their wages. This is driven largely by the relative balance between the distribution of immigrants and natives across these educational groupings. For the least educated Costa Ricans (roughly 1.5 percent of the working population), immigration is predicted to have negative impacts on wages ranging from 2 percent to as high as 17 percent. The large impacts for this very small sub-segment of the native skill distribution is driven by the fact that immigrants account for a large share of labor supply in this group (roughly 35 percent) and thus the own-wage effect of the supply shock is quite large.

For the U.S., there are predicted negative effects for high school dropouts in both simulations and slight negative effects (on the order of 1 percent) for some high school graduate and college graduate experience groups in the simulation assuming a low degree of substitutability across education groups. The two sets of results illustrate the sensitivity of these simulations to the parameter value choices. High substitutability across education groups allows the shock to any one group to diffuse to other education groups, thus dulling the own-wage effect. When this degree of substitutability is low, however, the effect of an immigration-induced labor supply shock sticks to the portions of the skill distribution receiving the most immigrants.

In both nations, the least educated are those most likely to experience wage declines as a result of immigration. In Costa Rica, however, this group is extremely small while in the U.S. it is somewhat larger (around 9 percent).

7. Conclusions

The analysis presented in this paper uses a calibrated structural model to analyze the impact of immigration on two economies: Costa Rica and United States. Unlike previous analysis this paper compares the impact of immigration on wages within the South-North and the South-South contexts. In the South-South context this paper improves on the scarce and previous analysis of the impact of immigration on the wages of natives by using specific and detailed attributes of human capital such as age and education and linking this with the magnitude of immigration flows.

We find that in a context where the immigrant and native population have similar native languages, immigrants and natives within easily observable human capital sub-groups appear to be perfect substitutes for one another. In the specific case of Costa Rica where most immigrants are from neighboring Nicaragua the data reveal infinite substitutability between otherwise similar native and immigrant workers. In the U.S, on the other hand, the immigrants come largely from non-English speaking populations and we observe that even within well-define skills groups the relative wages of immigrants depend on their relative supply. This is an interesting finding that merits further cross-national comparisons.

We also find very little evidence of an impact of immigration on native earnings in Costa Rica and modest evidence of a negative effect for low-skilled U.S. workers. The contrast in results follows primarily from the relative balance of immigrants and natives across skill groups in Costa Rica and the relative imbalance in the United States. However, this result could also be attributed to the fact that both immigration processes are the result of what can be termed as “economic induced” migration to both, Costa Rica and the United States.

The paper presents initial evidence of the magnitude and specific impact of migration flows on native wages when cultural differences between natives and immigrants are taken into account. Although a more detailed analysis of the impact of these differences is needed, the initial findings of this paper should provide guide to policy makers in both developed as well as developing countries as to what to expect when immigration becomes part of the social fabric and discussions about the impact of immigration on society are discussed.

These findings are important for policy makers in that they shed some light on the impact of immigration flows on particular segments of the workforce. An accurate prediction of the impact of migration on wages of particular segments of the native work force could enable policy makers and planners to more efficient use of public funds in sectors like education and infrastructure planning.

Finally, the methodology used in this paper provides a useful framework for future analysis of the South-South immigration process in other geographical contexts besides the Central American context. Given the importance of South-South migration worldwide policy makers will find the results presented here important for what to expect in terms of wages and occupational impacts in host countries so that more efficient planning can take place.

Table 1
Distribution of the Foreign Born Residents of Costa Rica and the United States by Country of Origin (%)*

	Costa Rica (2006)	United States (2005)
North America		
Canada	0.49	2.32
Mexico	1.17	30.74
United States	2.61	-
Central America		
Belize	0.05	0.12
Costa Rica	-	0.27
El Salvador	2.74	2.76
Guatemala	0.35	1.80
Honduras	0.40	1.08
Nicaragua	68.69	0.63
Panama	2.76	0.30
Caribbean	1.34	8.81
South America	8.96	6.87
Europe	1.89	10.91
Russia	0.18	3.03
Middle East	0.00	1.69
Asia	0.98	24.56
Africa	0.24	3.44
Other	-	0.47
Unknown	7.15	-

*Tabulated from the 2006 Encuesta de Hogares de Propositos Multiples for Costa Rica and the 2005 American Community Survey for the United States.

Table 2
Comparison of the Educational Attainment Distributions for Native-Born and Foreign-Born Working Residents of Costa Rica (2006) and the United States (2005)*

Panel A: Costa Rica

Educational Attainment Level	Distribution Across Educational Attainment Groupings		Proportion of Workers in Education groups that are Foreign Born
	Native-Born	Foreign-Born	
No Primary	0.015	0.057	0.350
Primary	0.427	0.430	0.121
Secondary	0.360	0.344	0.116
University	0.199	0.168	0.104

Panel B: United States

Educational Attainment Level	Distribution Across Educational Attainment Groupings		Proportion of Workers in Education groups that are Foreign Born
	Native-Born	Foreign-Born	
Less than high school	0.094	0.287	0.357
High school graduate	0.286	0.234	0.129
Some College	0.324	0.195	0.098
College graduate	0.297	0.284	0.148

* Tabulated from the 2006 Encuesta de Hogares de Propositos Multiples for Costa Rica and the 2005 American Community Survey for the United States.

Table 3
Estimated Results from Regressions of the Natural Log of the Native Immigrant Weekly Wage Ratio on the Native Immigrant Supply Ratio (Estimate of -1/ σ_{immig})

Panel A: Costa Rica, Dependent Variable = ln(Native Monthly Wage/Immigrant Monthly Wage)							
$\ln(L_{tkj1}/L_{tkj2})$	-0.017	0.003	-0.020	0.006	-0.018	0.001	-0.035
	(0.026)	(0.035)	(0.036)	(0.035)	(0.037)	(0.055)	(0.050)
Remaining Specification							
Year	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Educ	No	Yes	Yes	Yes	Yes	Yes	Yes
Exp	No	No	Yes	No	Yes	Yes	Yes
Year*Educ	No	No	No	Yes	Yes	Yes	Yes
Year*Exp	No	No	No	No	No	Yes	Yes
Educ*Exp	No	No	No	No	No	No	Yes
Panel B: United States, Dependent Variable = ln(Native Weekly Wage/Immigrant Weekly Wage)							
$\ln(L_{tkj1}/L_{tkj2})$	-0.030	-0.055	-0.034	-0.056	0.005	-0.040	-0.017
	(0.010)	(0.020)	(0.015)	(0.035)	(0.040)	(0.030)	(0.069)
Remaining Specification							
Year	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Educ	No	Yes	Yes	Yes	Yes	Yes	Yes
Exp	No	No	Yes	No	Yes	Yes	Yes
Year*Educ	No	No	No	Yes	Yes	Yes	Yes
Year*Exp	No	No	No	No	No	Yes	Yes
Educ*Exp	No	No	No	No	No	No	Yes

Standard errors are in parentheses and allow for clustering of the error variance-covariance matrix within education-experience cells. The monthly (weekly) wage differential is measured for full time male workers as described in the text. Ln supply differential is measures by the annual supply of hours for all workers in the education-experience-year-nativity cell for the U.S. and total weekly hours supplied in Costa Rica. The U.S. models are based on 1970 through 2005 and each have160 education-experience group observations. The Costa Rica models are based on the period from 1997 through 2006 and have 320 observation

Table 4
Estimated Results from IV Regressions of the Natural Log of Monthly (Weekly) Wages Among Full-Time-Male Workers on Log Weekly (Annual) Hours Supplied Using Log Hours Supplied by Immigrants as an Instrument at the Year-Education-Experience Level of Aggregation (Estimates of -1/$\sigma_{\exp}$)

Panel A: Costa Rica, Dependent Variable = ln Monthly Earnings			
Exp Var = ln(L_{tkj})	-0.073	-0.071	-0.024
	(0.026)	(0.031)	(0.067)
Remaining Specification			
Year	Yes	Yes	Yes
Educ	Yes	Yes	Yes
Exp	Yes	Yes	Yes
Year*Educ	No	Yes	Yes
Educ*Exp	No	No	Yes
Panel B: United States, Dependent Variable = ln Weekly Earnings			
Exp Var=ln(L_{tkj})	0.047	-0.048	-0.109
	(0.020)	(0.017)	(0.043)
Remaining Specification			
Year	Yes	Yes	Yes
Educ	Yes	Yes	Yes
Exp	Yes	Yes	Yes
Year*Educ	No	Yes	Yes
Educ*Exp	No	No	Yes

Standard errors are in parentheses and allow for clustering of the error variance-covariance matrix within education-experience cells. In each model, log supply is instrumented with the ln supply of immigrants in the particular year-education-experience cell. The U.S. models are based on 1970 through 2005 and each have160 education-experience group observations. The Costa Rica models are based on the period from 1997 through 2006 and have 320 observations

Table 5
IV Estimation Results from Regressions of ln Monthly (Weekly) Earnings on Supply Measured at the Year-Education-Experience Level (Providing an Estimate of $-1/\sigma_{\exp}$) and the Year-Education Level (Providing an Estimate of $-1/\sigma_{educ}$)

Explanatory Variables	Costa Rica: Dependent Variable = ln(monthly earnings)			United States: Dependent Variable = ln(weekly earnings)		
$\mathrm{Ln}(L_{tk})$	-0.344	-0.586	-0.736	-0.125	-0.073	0.113
	(0.254)	(0.307)	(0.496)	(0.057)	(0.032)	(0.059)
$\mathrm{Ln}(L_{tkj}) - \mathrm{Ln}L_{tk})$	-0.021	-0.013	-0.025	-0.095	-0.098	-0.105
	(0.072)	(0.064)	(0.070)	(0.046)	(0.039)	(0.041)
N	320	320	320	160	160	160
Remaining Specification						
Year effects	No	Yes	Yes	No	Yes	Yes
Educ-exp effects	Yes	Yes	Yes	Yes	Yes	Yes
Educ linear time trends	Yes	Yes	Yes	Yes	Yes	Yes
Educ quad. time trends	No	No	Yes	No	No	Yes

Standard errors are in parentheses and allow for clustering of the error variance-covariance matrix within education-experience cells.

Table 6
Simulated Effects of Immigration on lnWages in Costa Rica and the United States

	Costa Rica				United States	
	Impact, 1997 and 2006		Overall impact		Impact, 1970 and 2005	
Education-Experience Group (U.S.)	$\sigma_{immig}=\infty$ $\sigma_{educ}=3$ $\sigma_{\exp}=14$	$\sigma_{immig}=\infty$ $\sigma_{educ}=1.5$ $\sigma_{\exp}=14$	$\sigma_{immig}=\infty$ $\sigma_{educ}=3$ $\sigma_{\exp}=14$	$\sigma_{immig}=\infty$ $\sigma_{educ}=1.5$ $\sigma_{\exp}=14$	$\sigma_{immig}=30$ $\sigma_{educ}=10$ $\sigma_{\exp}=10$	$\sigma_{immig}=30$ $\sigma_{educ}=2.5$ $\sigma_{\exp}=10$
< Primary (< high school)						
0 to 4	-0.02	-0.05	-0.09	-0.17	0.00	-0.01
5 to 9	-0.02	-0.05	-0.08	-0.15	-0.01	-0.08
10 to 14	-0.04	-0.07	-0.09	-0.17	-0.01	-0.11
15 to 19	-0.04	-0.07	-0.09	-0.16	-0.01	-011
20 to 24	-0.03	-0.06	-0.08	-0.15	-0.01	-0.09
25 to 29	-0.02	-0.05	-0.07	-0.15	-0.01	-0.05
30 to 34	-0.04	-0.07	-0.08	-0.15	0.00	-0.03
35 to 40	-0.03	-0.05	-0.10	-0.14	0.00	-0.01
Primary (High School)						
0 to 4	0.00	0.01	0.00	0.00	0.01	0.02
5 to 9	0.00	0.01	0.00	0.00	0.00	0.00
10 to 14	0.00	0.01	0.00	-0.01	0.00	-0.01
15 to 19	0.00	0.01	0.00	0.00	0.00	0.00
20 to 24	0.01	0.01	0.00	0.00	0.01	0.01
25 to 29	0.00	0.01	0.00	0.00	0.01	0.02
30 to 34	0.00	0.01	0.00	0.00	0.01	0.03
35 to 40	0.00	0.01	0.00	0.00	0.01	0.03
Secondary (Some College)						
0 to 4	0.00	0.00	0.00	0.00	0.01	0.02
5 to 9	0.00	0.00	0.00	0.00	0.01	0.02
10 to 14	0.00	0.00	0.00	0.00	0.01	0.01
15 to 19	0.00	0.00	0.00	0.00	0.01	0.01
20 to 24	0.00	0.00	0.00	0.00	0.01	0.01
25 to 29	0.00	0.00	0.00	0.00	0.01	0.02
30 to 34	0.00	0.00	0.00	0.00	0.01	0.03
35 to 40	0.00	0.00	0.00	0.00	0.01	0.03
Any College (College Grad)						
0 to 4	0.00	0.00	0.00	0.01	0.01	0.01
5 to 9	0.00	0.00	0.01	0.01	0.00	-0.01
10 to 14	0.00	0.00	0.00	0.00	0.00	-0.01
15 to 19	0.00	0.00	0.00	0.01	0.00	-0.01
20 to 24	0.00	0.00	0.00	0.01	0.00	0.00
25 to 29	0.00	0.00	0.01	0.01	0.00	0.01
30 to 34	0.00	0.00	0.00	0.01	0.01	0.02
35 to 40	0.00	0.00	0.00	0.01	0.01	0.01

Figure 1: Percent Foreign-Born Among All Costa Rican Residents and Among Working, Prime-Age Costa Rican Residents, 1997 through 2006

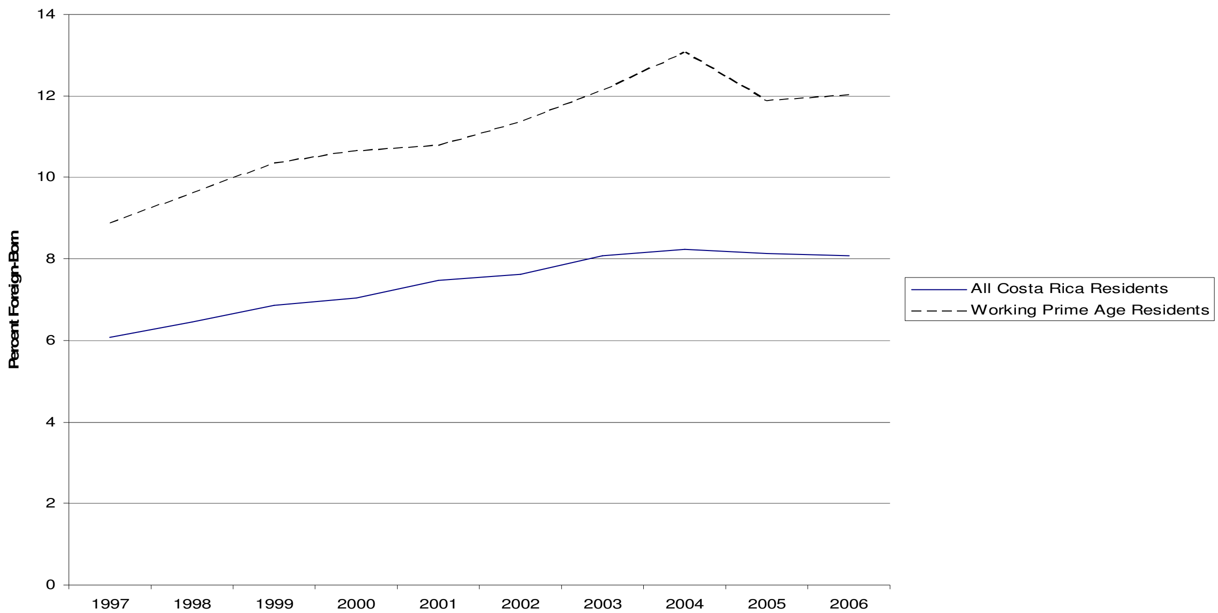


Figure 2: Percent Foreign Born Among All U.S. Residents and Among Working, Prime Age U.S. Residents, 1970 through 2005

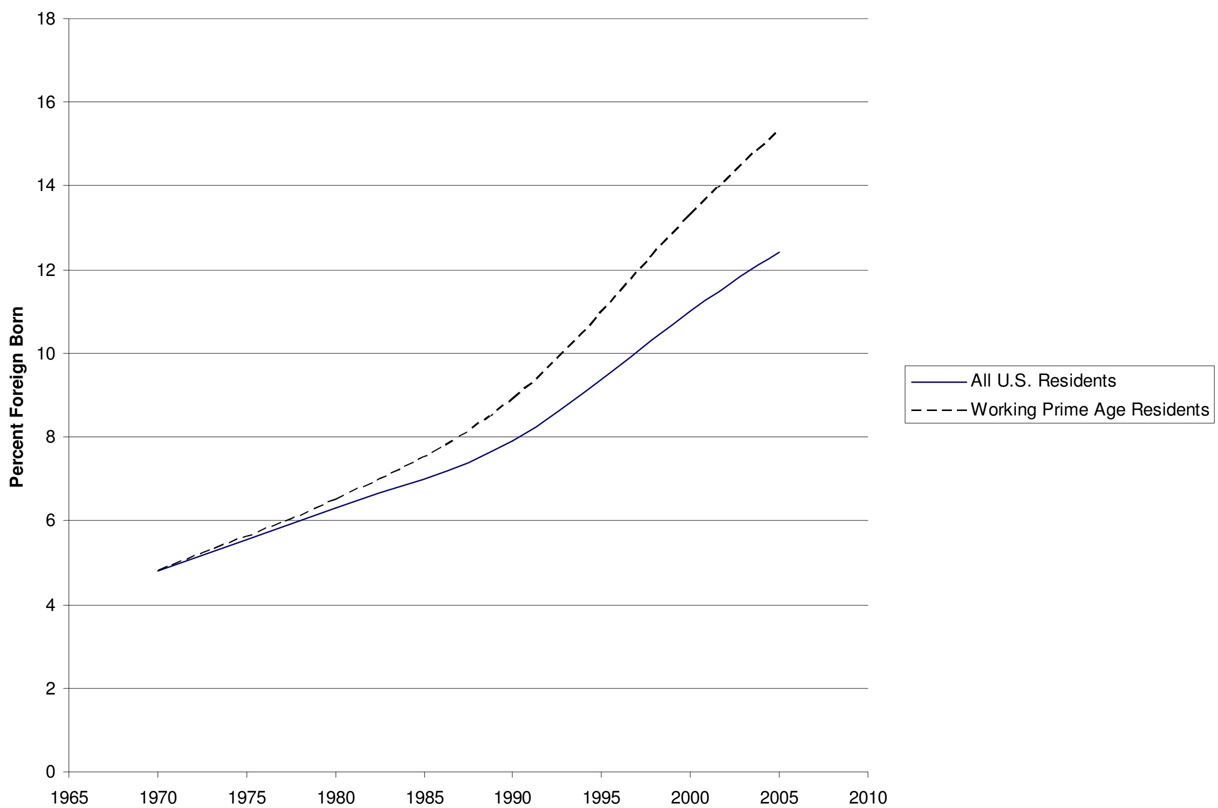


Figure 3: Comparison of the Age Distribution Among Working Native and Foreign-Born Costa Rican Residents 2006

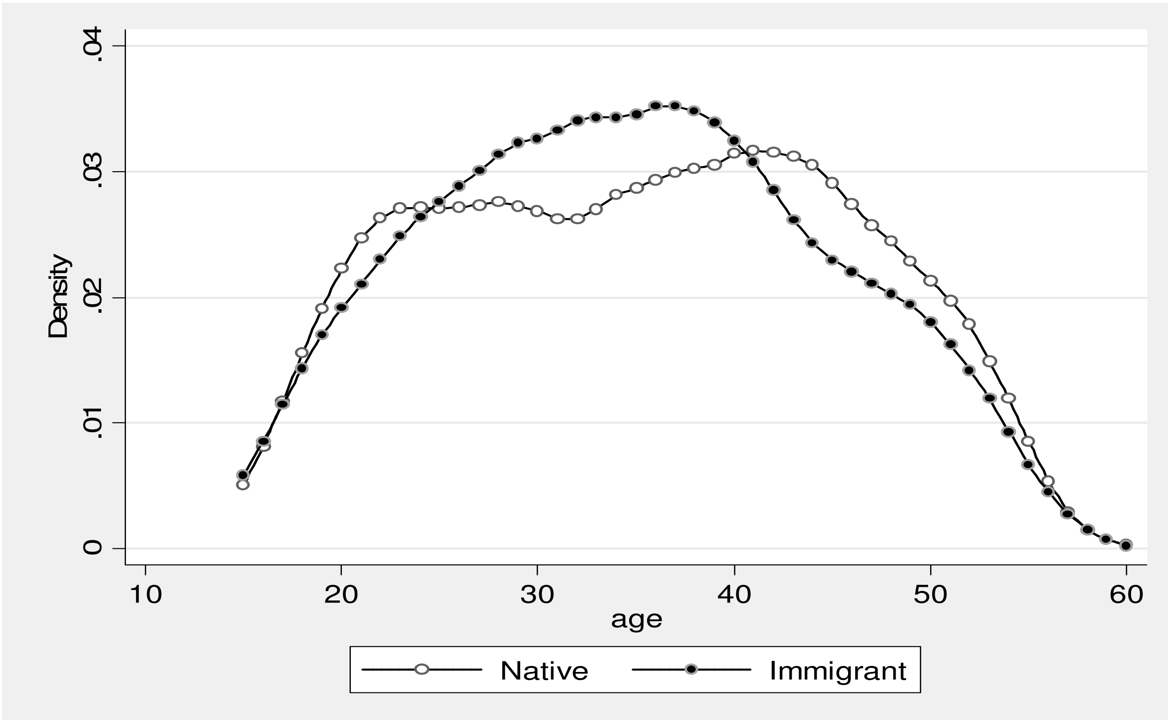


Figure 4: Comparison of the Age Distribution Among Working Native and Foreign-Born U.S. Residents, 2005

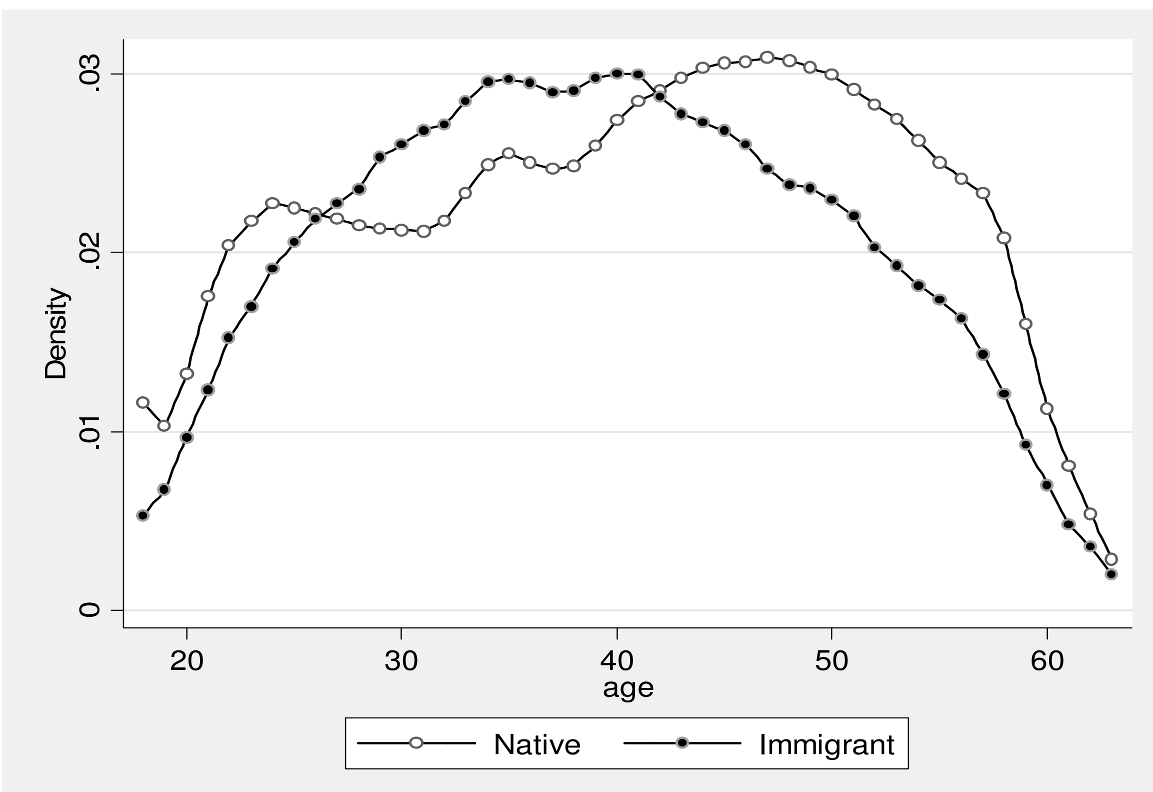


Figure 5: Scatter Plot of the Native-Immigrant ln Wage Differential Measured by Year (1997 through 2006), Education, and Experience Group Against Corresponding Log Supply Differential, Costa Rica

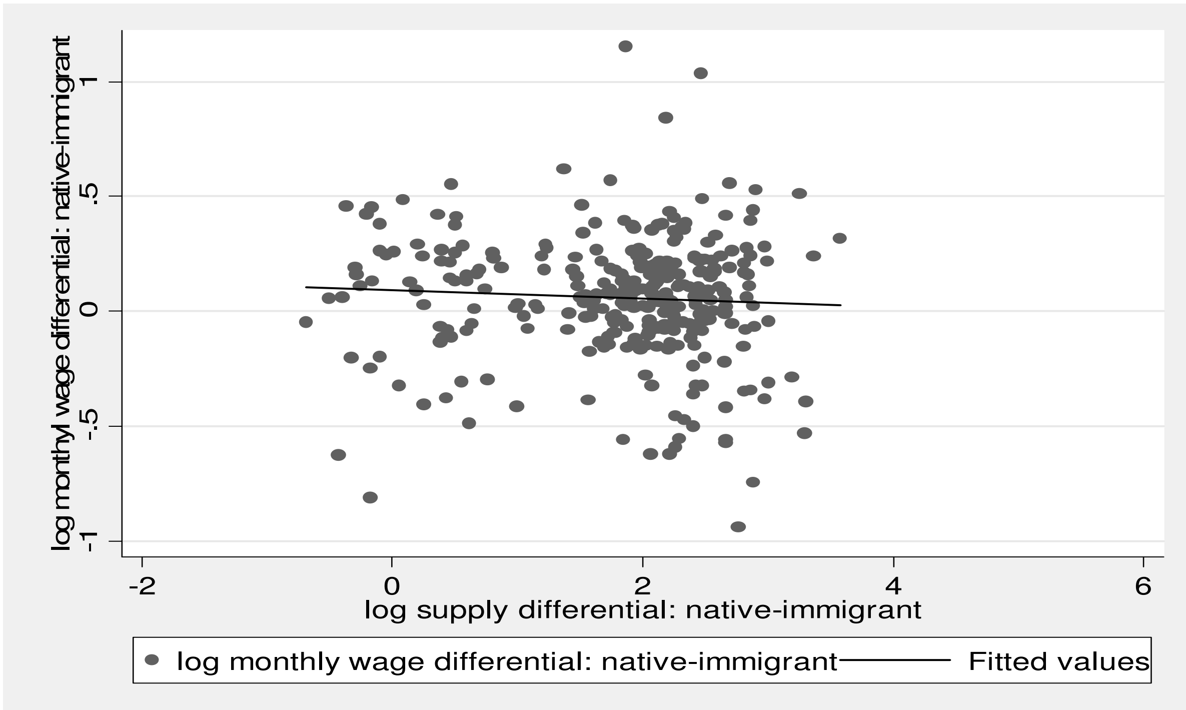
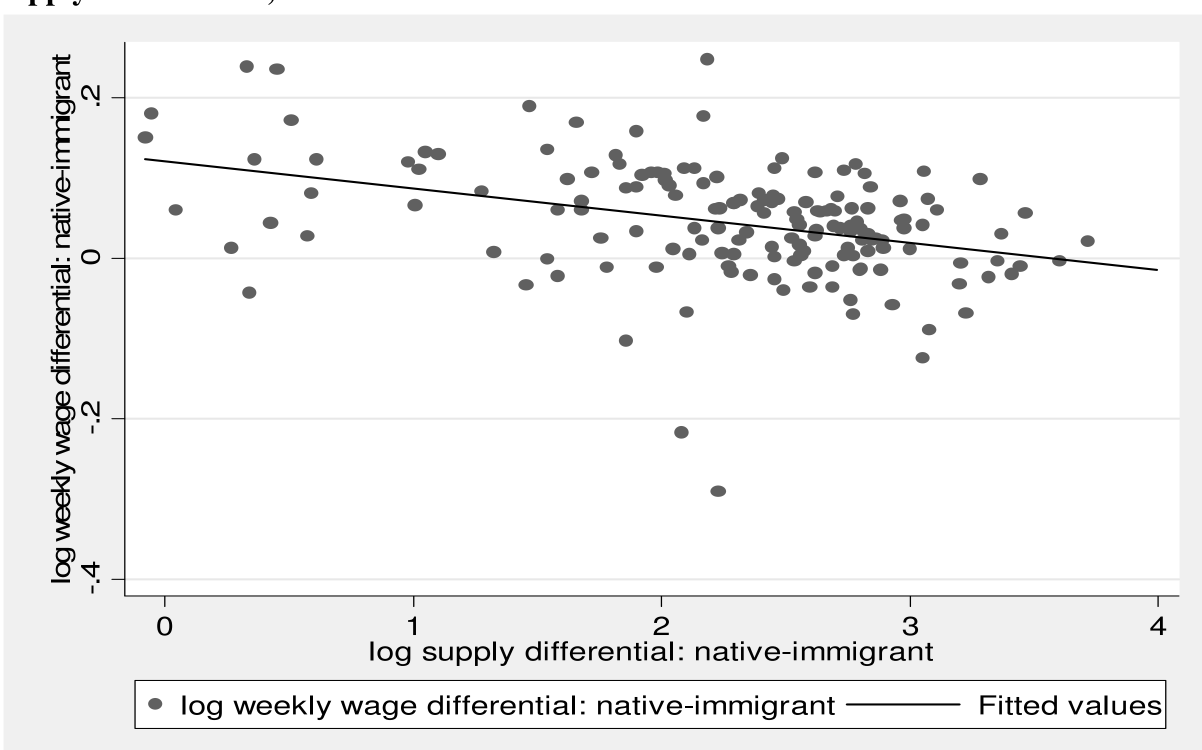


Figure 6: Scatter Plot of the Native-Immigrant Ln Wage Differential Measured by Year (1970 through 2005), Education, and Experience Groups Against the Corresponding Log Supply Differentials, United States



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