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# RETHINKING THE EFFECTS OF IMMIGRATION ON WAGES

Gianmarco I.P. Ottaviano
Giovanni Peri

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**ABSTRACT**

This paper asks the following important question: what was the effect of surging immigration on average and individual wages of U.S.-born workers during the period 1990-2004? Building on section VII of Borjas (2003) we emphasize the need for a general equilibrium approach to analyze this problem. The impact of immigrants on wages of US born workers can be evaluated only by accounting carefully for labor market and capital market interactions in production. Using such a general equilibrium approach we estimate that immigrants are imperfect substitutes for U.S.-born workers within the same education and experience group (because they choose different occupations and have different skills). Moreover, accounting for reasonable speed of adjustment of physical capital we show that most of the wage effects of immigration accrue to native workers already within a decade. These two facts, overlooked by the previous literature, imply a positive and significant effect of the 1990-2004 immigration on the average wage of U.S.-born workers overall, both in the short and in the long run. This positive average effect resulted from a positive effect on wages of all US-born workers with at least a high school degree and a small negative effect on wages of U.S. born workers with no high school degree.

Gianmarco I.P. Ottaviano
University of Bologna
Dip Scienze Economiche
Strada Maggiore 45, 40125 Bologna
ITALY
ottavian@economia.unibo.it

Giovanni Peri
Department of Economics
University of California, Davis
One Shields Avenue
Davis, CA 95616
and NBER
gperi@ucdavis.edu

# 1 Introduction

During the last three and a half decades the United States has experienced a remarkable surge in immigration. The share of foreign-born workers in the labor force has steadily grown from 5.3% in 1970 to 14.7% in 2005[1], progressively accelerating; in the period between 1990 and 2005, almost one million immigrants entered the country every year[2]. In parallel to this surge, the debate about the economic effects of immigrants on U.S. natives, and particularly on their wages, has gained momentum both inside academia and in the policy and media arenas. The debate has been particularly lively in the wake of a bill passed by the U.S. House of representatives calling, among other provisions, for criminalization of illegal aliens. A second bill, passed by the U.S. Senate, calls for a road to legalization of the same group[3]. From the academic perspective, two facts have contributed to feeding the debate. First, the recent empirical literature about the effect of immigrants on the wages of natives has provided a mixed set of results. Second, the group of uneducated workers (without a high school degree) has become increasingly large among recent immigrants, while at the same time the real wage of uneducated U.S.-born workers has performed very poorly: it even declined in real terms during the recent decades (see, for example, Autor, Katz and Kearny, 2005). It is certainly tempting to attribute the poor wage performance of uneducated U.S. workers to the competition of immigrants, as such connection would provide an easy solution to the problem of wage decline: halt immigration.

Ten years ago an influential survey by Friedberg and Hunt (1995) summarized the literature concluding that, "the effect of immigration on the labor market outcomes of natives is small." Since then, a number of studies have re-examined the issue, refining the estimates by accounting for important problems related to the endogeneity of immigrant inflow and the internal migration of U.S. workers. Even with more accurate and sophisticated estimates at hand, a consensus has yet to be reached: some economists identified only small effects of immigration on wages (Card, 2001) while others found large negative effects (Borjas, Friedman and Katz, 1997)[4]. Recently, however, the latter view of a large negative impact of immigration on wages, particularly of uneducated workers, seems to have gained momentum. An influential article by George Borjas, (2003), followed by Borjas and Katz (2005) and Borjas (2006) that use a similar empirical method, argue, using national data from five decennial U.S. censuses (1960-2000) that U.S. workers lost, on average, about 3% of the real value of their wages due to immigration over the period 1980-2000 and that this loss reached almost 9% for native workers without a high school degree (Borjas, 2003, Table IX, page 1369) at least in the short run.

[1] Authors' calculations using 1 percent Integrated Public Use Microdata Series (IPUMS) data for the year 1970 and Current Population Survey March Supplement, Ruggles et al (2006), for the year 2005.

[2] While remarkable, such rapid increases are not unprecedented for the U.S. Large inflows from Europe during the period 1880-1910 brought the percentage of foreign-born very close to 15% in the year 1910, and previous episodes of very intense immigration (e.g. 1.5 million Irish immigrants between 1845 and 1854, in the wake of a great famine) caused similar surges.

[3] The bills were, respectively, the Border Protection, Anti-Terrorism, and Illegal Immigration Control Act (H.R. 4437) passed in December 2005 by the U.S. House of Representatives and the Comprehensive Immigration Reform Act ( S.2611) passed in May 2006 by the U.S. Senate.

[4] We are aware of only one previous paper, Friedberg (2001), that finds a positive *partial* effect of immigration on native wages. In most cases, however, that effect is not significant.

Our paper builds on section VII of the article by George Borjas (2003), but takes a fresh look at some critical issues with substantial revisions of several results. The key idea is that the effects of immigration on wages can only be measured within a *general equilibrium* framework. More specifically, a study on the effects of immigration on wages of different types of workers by education, experience and nativity should build on a production function that describes how these different types of workers interact with each other and with physical capital to produce output. Then, one can derive the demand for each type of labor, which depends on productivity and employment of the other labor types as well as on physical capital. Finally, market clearing conditions can be used to obtain wage equations from the labor demands and supplies, and use them to estimate the elasticities of substitution (relative wage elasticities) between workers empirically. Going back to the production function, these estimates can then be used to assess the effect of immigration (a change in the supply of different types of workers) on wages (the marginal productivity of different types of workers). In contrast, several existing empirical studies directly estimate a reduced-form wage equation for native workers with certain characteristics (such as educational or occupational groups) obtaining the elasticity of wages to new immigrants in the same group. Such an approach only provides the "partial" effect of immigration on wages (as it omits all cross-interactions with other types of workers and with capital) and as such is uninformative on the overall effect of immigrants.

The *general equilibrium* approach is accompanied by two novel features of our analysis. First we remove the usual assumption that foreign- and U.S.-born workers are perfect substitutes within the same education-experience group. Be it because immigrants tend to choose a different set of occupations (as we document below), because they are a selected, motivated and generally talented group, or because they have some culture-specific skills it seems reasonable to allow them to be imperfect substitutes for natives even within an education-experience group and to let the data estimate the corresponding elasticities of substitution. While acknowledging that in principle "[im]migrants may complement some native factors in production... and overall welfare may rise" (Friedberg and Hunt, 1995, page 23), most studies thus far have only focused on the partial effects of immigrants on the wages of those native workers who are their closest substitutes (i.e. within the same occupation, education-experience or skill groups). By modeling labor as a differentiated input in general equilibrium, we enlarge the picture to better capture the effects of immigration within and between different groups.

The second novel feature of our analysis is a more careful consideration of the response of physical capital to immigration. As physical capital complements labor it is important to account for its adjustment in the short and in the long run. In particular when evaluating the "short-run" response of wages to immigration it seems rather artificial to maintain a fixed stock of capital, while accumulating immigration flows occurring over ten or twenty years, as is currently done in the literature. Immigration happens gradually over time (not at the beginning of the decade) and investors respond continuously, although with sluggishness, to increased marginal productivity of capital caused by immigration. As for the long-run response of capital, any model of growth (Solow, 1956; Ramsey, 1928) as well as empirical evidence imply that capital adjusts to maintain its real return (and capital output ratio) constant. For the short-run, we use estimates of the speed of adjustment of capital taken from the growth and the real business cycle literature to evaluate the average wage impact of immigration. We are also able to assess how long it takes for full adjustment to take place. Rather than reporting the effects of fourteen year of immigration for fixed capital and for fully adjusted capital we are able to estimate the effect of immigration during the 1990-2004 period as of year 2004 and then we show that within the following 5 years the largest part of "the long run effect" has set in.

Once we account for the afore mentioned effects, we deeply revise several commonly estimated effects of immigrants on the wages of U.S. natives. First, in the long-run the average wage of U.S.-born workers experienced a *significant increase* (+1.8%) as a consequence of immigration during the 1990-2004 period. Even in the short run (as of 2004) average wage of US native workers had a moderate increase (+0.7%) because of immigration. This result stems from the imperfect substitutability between U.S. and Foreign born workers so that immigration increases the wages of U.S.-born at the expenses of a decrease in wages of foreign-born workers (namely, previous immigrants). Second, the group of least educated U.S.-born workers *suffers a significantly smaller wage loss than previously calculated*. In the long run native workers only lost 1.1% of their real wage due to the 1990-2004 immigration. Even in the short run (as of 2004) the negative impact was a moderate 2.2% real wage loss. The methodology used in the previous literature would estimate much larger losses, around -8% in the short run and -4.2% in the long run. The fact that uneducated foreign-born do not fully and directly substitute for (i.e. compete with) uneducated natives, but partly complement their skills, is the reason for this attenuation. Third, *all other groups of U.S.-born workers* (with at least an high school degree) who accounted for 90% of the U.S.-born labor force in 2004, *gained from immigration.* Their real wage gains in the long run range between 0.7% and 3.4% while even in the short run they either gain (high school graduates) or have essentially no wage change (college graduates). Finally, even considering only the *"relative" effect* of immigration on real wages of natives, namely its contribution to the widening of the college graduates-high school dropouts wage gap and of the college graduates-high school graduates wage gap, we find only a small contribution of immigration to the first and a negative contribution (i.e. reduction of the gap) to the second for the 1990-2004 period. The group whose wage was *most negatively affected by immigration* is, in our analysis, *the group of previous immigrants;* however, it is they who probably have the largest non-economic benefits from the immigration of spouses, relatives or friends making them willing to sustain those losses.

The remainder of the paper is organized as follows. Section 2 summarizes the relevant literature. Section 3 introduces the aggregate production function, derives the demand for each type of labor and identifies the key parameters for calculating the elasticity of wages to the inflow of immigrants. Section 5 presents the data, illustrates some preliminary evidence of differences between native and foreign born workers in the labor market and produces the key estimates of the relevant elasticities. Using those estimates, Section 6.1 evaluates the effect of immigration on the wages of U.S. natives for the period 1990-2004. We re-visit, in Section 6.2, the distinction between short and long run analysis and consider the short-run effects (as of year 2004) and the long-run effects (during the following five years and with full capital adjustment) and we compare our results to previous findings on the effects of immigration. Finally, in Section 6.4, we analyze by how much immigration contributed the increased wage dispersion during the period 1990-2004. Section 7 concludes the paper.

## 2 Review of the Literature

There is a long list of contributions in the literature dealing with the impact of immigrants on the wages of natives. Some of these studies explicitly consider the contribution of immigration to increased wage dispersion and to the poor performance of real wages of the least educated since 1980. Two questions are typically analyzed by the existing literature. The first is imbued with a "macro" flavor: Does the inflow of foreign born workers have a positive or negative net effect on the average productivity and wages of U.S.-born workers? This question requires that we aggregate the wages of quite heterogeneous workers. The second question is more "micro" in focus: How are the gains and losses from immigration distributed across U.S.-born workers with different levels of education (and experience)? The consensus emerging from the literature is that the first (macro) effect on average US wages is negligible in the long run, as capital accumulates to restore the pre-migration capital-labor ratio, however, for fixed capital in the short run there can be a large depressing effect of immigration on wages. Most of the literature represents immigration as an increase in labor supply for a given capital stock (Borjas, 1995, 2003), and readily finds a negative impact of immigration on average wages (in the short run) and a positive impact of immigration on the return to capital due to complementarity between the two factors. The recent debate, however, has focused on the effects of immigration on the *relative* wages of more and less educated U.S.-born workers. Some economists argue for a large relative impact adverse to less educated workers (Borjas, 1994, 1999, 2003, 2006; Borjas, Freeman and Katz, 1997), while others favor a smaller, possibly insignificant, effect (Butcher and Card, 1991; Card, 1990; Card, 2001; Friedberg 2001; Lewis, 2005; National Research Council, 1997). The size and significance of the estimated relative wage effects from immigration remain controversial, and possibly depend at least in part on the use of local versus national data.

The present article uses a framework from which both the "macro" (average) and the distributional (relative) effects of immigration can be derived. We argue that only within such a framework, based on the aggregate production function and general equilibrium outcomes, can one measure and discuss either of these effects. Our approach builds on the model employed in section VII of Borjas (2003) and uses national data in performing estimations. This approach avoids the problems arising from internal migration of natives and from endogenous location choice and attenuation bias when using metropolitan or state data[5].

The modern analysis of the effects of immigrant inflows on the wages of natives began with studies that treated foreign-born as a single homogeneous group of workers (Grossman, 1982; Altonji and Card, 1991), imperfectly substitutable with U.S.-born workers. A number of studies on the relative supply of skills and relative wages of U.S.-born workers made clear, however, that workers with different levels of schooling and experience are better considered as imperfectly substitutable factors (Katz and Murphy, 1992; Welsh, 1979; Card and Lemieux, 2001). As a consequence, more recent analysis has been carried out partitioning workers among imperfectly substitutable groups (by education and experience) while assuming perfect substitution of native- and foreign-born workers within each group (Borjas, 2003). The present article combines the two approaches in the sense that both can be seen as special cases nested in our more general framework. Specifically, we assume the existence of an aggregate production function that combines workers and physical capital, while using education, experience and place of origin (U.S. versus elsewhere) to categorize imperfectly substitutable groups. Following Borjas (2003), we choose a constant elasticity of substitution (CES) technology but, differently from that article, we treat the two groups of U.S.- and foreign-born workers as not perfectly substitutable and we partition them across eight experience levels and four educational attainment classes. This allows for the imperfect substitutability of individuals between different country origins and different education-experience levels; imperfect substitutability may arise from the different abilities, occupational choices or unobserved characteristics of workers. Within this framework we estimate three sets of elasticities: (i) between U.S.- and foreign-born within education-experience groups; (ii) between experience levels within education groups; and (iii) between education groups. There is scant literature estimating the first set of elasticity parameters. The few works we are aware of include Jaeger (1996) which only used 1980-1990 metropolitan data and whose estimates may be susceptible to attenuation bias and endogeneity problems related to the use of local data, and Cortes (2005) who only considers low-skilled workers and uses metropolitan areas data to find a very low elasticity of substitution between U.S.- and foreign-born workers. The other two sets of elasticities (between experience and between education groups) have been estimated in several studies (Card and Lemieux, 2001; Katz and Murphy, 1992; Angrist, 1995; Ciccone and Peri, 2005) and are found to be around 2 (across education groups) and around 4 (across experience groups).

As for physical capital, we explicitly consider its contribution to production and treat its accumulation as driven by market forces that equalize its real returns in the long run. In particular, we revise the usual approach that considers capital as fixed in short-run simulations. The growth literature (Islam, 1995 Caselli et al. 1996) and real business cycle literature (e.g. Romer, 2006, Chapter 4) has estimated, using yearly data on capital accumulation and different types of shocks, the speed of adjustment of capital to deviations from its long-run

[5] See Borjas (2006) and Borjas, Freeman and Katz (1997) for a discussion of these issues.

growth path. Adopting 10% per year as a reasonable estimate of the speed of adjustment of physical capital in the U.S. (confirmed by our own estimates for the 1960-2004 period) we analyze the impact of yearly immigration on average wages as capital adjusts. We can evaluate the effect of immigration occurred in the period 1990-2004 on average wage as of year 2004 and we can evaluate its effects after five or ten more years. This is an important departure from the literature, which has not paid much attention to the actual response of physical capital to immigration. When evaluating the wage effects of immigration, the prevalent assumption has been that of a fixed capital stock in the short run (Borjas, 1995; Borjas, 2003, Borjas, Freeman and Katz, 1997; Borjas and Katz, 2005).

Some studies on the effects of immigration on wages have specifically focussed on immigration (along with trade) as a proposed explanation for the worsening of income distribution in the U.S. during the years following 1980. In particular, Borjas, Friedman and Katz (1997) found that immigration contributed to the widening of the wage gap between high school dropouts and high school graduates during the 1980-1995 period but did not contribute to the widening of the college graduate-high school graduate wage gap. In light of new studies (notably Autor, Katz and Kearny, 2005, 2006) that further document the evolution of college graduate-high school graduate and high school graduate-high school dropout wage gaps during the 1990-2005 period, and in light of our new results that reduce the adverse impact of immigration on wage distribution, we revisit this arm of the literature by calculating the contribution of immigration to wage dispersion for the 1990-2004 period.

Finally, as mentioned earlier, several studies on the *relative* wage effects of immigrants have analyzed local data (e.g. for metropolitan areas) accounting for the internal migration response of U.S. natives (Card, 2001; Card and Di Nardo, 2000; Lewis, 2005) and correcting for the endogeneity of immigrant location choice (both factors would cause an attenuation bias in the estimates). These studies find a small negative partial effect of immigrants on wages. On the other hand, our previous work (Ottaviano and Peri, 2005a, 2005b, 2006) has pointed out a positive effect of immigration on the average wage of U.S. natives across U.S. metropolitan areas. This positive and significant effect survived 2SLS estimation, using instruments that should be exogenous to city-specific unobservable productivity shocks[6]. The complementarity in production illustrated in this paper could also be at work at the city level. Accordingly, the model proposed in this paper is able to reconcile the negative partial effects observed in previous studies by other authors with the positive average effect of immigration at the local level in our previous findings[7].

[6] We build the instrumental variables by using the initial share of foreign-born workers in a city, grouped by country of origin, and attributing to each group the average immigration rate for that nationality during each decade in the period (1970-2000). First introduced by Card (2001), this instrument is correlated with actual immigration in the metropolitan area if new immigrants tend to settle prevalently where fellow countrymen already live.

[7] The city model is developed in greater detail in Ottaviano and Peri (2005b).

## 3 Theoretical Framework

To evaluate the effects of immigrants on the wages of natives and other foreign-born workers when each group differs by education, experience and other characteristics, we need a model of how the marginal productivity of a given type of worker changes in response to changes in the supply of other types. At the same time, it is important to account for the response of physical capital to immigration. In the macro and growth literature, a simple and popular way of doing this is to assume an aggregate production function in which aggregate output (the final good) is produced using a combination of physical capital and different types of labor.

### 3.1 Production Function

Following Borjas (2003) who builds on Card and Lemieux (2001), we choose a nested CES production function, in which physical capital and different types of labor are combined to produce output. Labor types are grouped according to education and experience characteristics; experience groups are nested within educational groups, that are in turn nested into a labor composite. U.S.-born and foreign-born workers are allowed a further degree of imperfect substitutability even when they have the same education and experience. While the nested CES function imposes some restrictions on the elasticities of substitution across skill groups it has the advantage of being parsimonious in parameters, widely used and yields results easily comparable with a large body of articles in the labor and macro literature. The aggregate production function we use is given by the following expression:

$$Y_t = A_t L_t^{\alpha} K_t^{1-\alpha} \tag{1}$$

where $Y_t$ is aggregate output , $A_t$ is total factor productivity (TFP), $K_t$ is physical capital, $L_t$ is a CES aggregate of different types of labor (described below), and $\alpha \in (0,1)$ is the income share of labor. All variables, as indicated by the subscripts, are relative to year $t$. The production function is a constant returns to scale (CRS) Cobb-Douglas combination of capital $K_t$ and labor $L_t$. Such a functional form has been widely used in the macro-growth literature (recently, for instance, by Jones, 2005 and Caselli and Coleman, 2006) and is supported by the empirical observation that the share of income going to labor, $\alpha$, is constant in the long run and across countries (Kaldor, 1961; Gollin, 2002). The labor aggregate $L_t$ is defined as:

$$L_t = \left[ \sum_{k=1}^{4} \theta_{kt} L_{kt}^{\frac{\delta-1}{\delta}} \right]^{\frac{\delta}{\delta-1}} \tag{2}$$

where $L_{kt}$ is an aggregate measure of workers with educational level $k$ in year $t$; $\theta_{kt}$ are education-specific productivity levels (standardized so that $\sum_k \theta_{kt} = 1$ and any common multiplying factor can be absorbed in the TFP term $A_t$). As is standard in the labor literature, we group educational achievements into four categories: high school dropouts (denoted as $HSD$) , high school graduates ($HSG$), college dropouts ($COD$) and college graduates ($COG$), so that $k = \{HSD, HSG, COD, COG\}$. The parameter $\delta > 0$ measures the elasticity of substitution between workers with different educational achievements. Within each educational group we assume that workers with different experience levels are also imperfect substitutes. In particular, following the specification used in Card and Lemieux (2001), we write:

$$L_{kt} = \left[ \sum_{j=1}^{8} \theta_{kj} L_{kjt}^{\frac{\eta-1}{\eta}} \right]^{\frac{\eta}{\eta-1}} \tag{3}$$

where $j$ is an index spanning experience intervals of five years between 0 and 40, so that $j = 1$ captures workers with $0-4$ years of experience, $j = 2$ those with $5-9$ years, and so on. The parameter $\eta > 0$ measures the elasticity of substitution between workers in the same education group but with different experience levels and $\theta_{kj}$ are experience-education specific productivity levels (standardized so that $\sum_j \theta_{kj} = 1$ for each $k$ and assumed invariant over time, as in Borjas, 2003) . As we expect workers within an education group to be closer substitutes than workers across different education groups, our prior (consistent with the findings of the literature) is that $\eta > \delta$. Finally, distinct from most of the existing literature, we define $L_{kjt}$ as a CES aggregate of home-born and foreign-born workers. Denoting the number of workers with education $k$ and experience $j$ who are, respectively, home-born and foreign-born, by $H_{kjt}$ and $F_{kjt}$, and the elasticity of substitution between them by $\sigma_k > 0$, our assumption is that:

$$L_{kjt} = \left[ \theta_{Hkjt} H_{kjt}^{\frac{\sigma_k-1}{\sigma_k}} + \theta_{Fkjt} F_{kjt}^{\frac{\sigma_k-1}{\sigma_k}} \right]^{\frac{\sigma_k}{\sigma_k-1}} \tag{4}$$

Foreign-born workers are likely to have different abilities pertaining to language, quantitative skills, relational skills and so on. These characteristics, in turn, are likely to affect their choices of occupation and their abilities in the labor force, therefore foreign-born workers should be differentiated enough to be treated as imperfect substitutes for U.S.-born workers, even within the same education and experience group. As we expect workers within the same education and experience group to be closer substitutes than workers across different education and experience groups, our working hypothesis is that $\sigma_k > \eta$. We analyze this issue in detail in Section 5 below. Ultimately, we allow the empirical analysis to reveal whether U.S.-born workers and foreign-born workers within the same education and experience group are perfect substitutes ($\sigma_k = \infty$) or not.[8] We also allow, as indicated by the subscript $k$, that the elasticity of substitution between U.S.- and foreign-born workers differs across education groups (more on this below). Finally, the terms $\theta_{Hkjt}$ and $\theta_{Fkjt}$ measure the specific

[8] The standard assumption in the literature has been, so far, to impose that $L_{kjt} = H_{kjt} + F_{kjt}$, i.e. that once we control for education and experience, foreign-born and natives are workers of identical type.

productivity levels of foreign- and home-born workers and they may vary across groups and years (in the empirical identification we impose a systematic structure on their time variations) . They are also standardized so that $(\theta_{Hkjt} + \theta_{Fkjt}) = 1$.

### 3.2 Physical Capital Adjustment

Physical capital adjustment to immigration may not be immediate. However, investors respond continuously to inflows of labor and to the consequent increase in the marginal productivity of capital; how fast they respond is an empirical question. Further, immigration is not an unexpected and instantaneous shock. It seems odd, therefore, to treat the short-run effect as the impact of immigration for fixed capital stock, which prompts the question: for how long is capital fixed and why?. Immigration is an ongoing phenomenon, distributed over years, predictable and rather slow. Despite the acceleration in legal and illegal immigration after 1990, the inflow of immigrants measured less than 0.6% of the labor force each year between 1960 and 2004. It is reasonable, therefore, to think of this issue more dynamically with investments continuously responding to the flow of immigrant workers. In a dynamic context the relevant parameter in order to evaluate the impact of immigration on average wages is the speed of adjustment of capital. In the long run, on the balanced growth path, such as in the Ramsey (1928) or the Solow (1956) models, the variable $\ln(K_t/L_t)$ follows a constant positive trend growth determined only by total factor productivity ($\ln A_t$) and not affected by the size of $L_t$. Therefore the average wage in the economy, that depends on $K_t/L_t$ does not depend on immigration in the long run. Shocks to $L_t$, such as immigration, however, may affect temporarily the value of $K_t/L_t$ causing it to be below its long run trend. How much and for how long will $\ln(K_t/L_t)$ be below trend (and the average wage be reduced) as a consequence of immigration depends on the yearly inflow of immigrants and on the yearly rate of adjustment of physical capital. The theoretical and empirical literature on the speed of convergence of a country's capital per worker to its own balanced growth path (Islam, 1995 Caselli et al. 1996), as well the business cycle literature on capital adjustment (Romer, 2006) provide estimates for such speed of adjustment that we can use together with data on total yearly immigration to obtain the effect of immigration 1990-2004 on average wages in 2004 and in the following 5 to 10 years as capital continues to adjust. We devote the next section, 3.2.1, to show in detail the connection between average wages and capital-labor ratio. In our empirical analysis we first focus on the long run effects of immigration (Section 6.1), allowing for full capital adjustment, as natural reference. Then in section 6.2 we use the estimated speed of capital adjustment to show the effect of fourteen years of immigration (1990-2004) on wages as of year 2004, and we compare those results with the traditional way of computing "short run' effects on wages.

### 3.2.1 Partial Adjustment, Total Adjustment and Wages

Given the production function in (1) the effect of physical capital $K_t$ on the wages of individual workers operates through the effect on the marginal productivity of the aggregate $L_t$. Let us call $w_t^L$ the compensation to the composite factor $L_t$,which is equal to the average wage in the economy[9]. In a competitive market it equals the marginal productivity of $L_t$, hence:

$$w_t^L = \frac{\partial Y_t}{\partial L_t} = \alpha A_t \left(\frac{K_t}{L_t}\right)^{1-\alpha} \tag{5}$$

Assuming either international capital mobility or capital accumulation, along the balanced growth path of the Ramsey (1928) or Solow (1956) models, the real interest rate $r$ and the aggregate capital-output ratio $K_t/Y_t$ are both constant in the long run and the capital labor ratio $K_t/L_t$ grows at constant rate equal to $\frac{1}{1-\alpha}$ times the growth rate of technology $A_t$. This assertion is also supported in the data, as the real return to capital and the capital-output ratio in the U.S. did not exhibit any trend in the long run while the capital-labor ratio grew at constant rate (Kaldor, 1961). In particular, this is true for our period of consideration 1960-2004 as depicted in Figure 1 and Figure 2, where the capital-output ratio ($K_t/Y_t$) and the de-trended log capital-labor ratio, $\ln(K_t/L_t)$, show cyclical movements but remarkable tendency to mean reversion in the long-run[10]. In order to show the effect of different patterns of capital adjustment on the average wage ($w_t^L$) it is useful to write the capital stock as $K_t = \kappa_t L_t$ where $\kappa_t$ is the capital-labor ratio. Hence $w_t^L$ (from equation 5) can be expressed in the following form:

$$w_t^L = \alpha A_t \left(\kappa_t\right)^{1-\alpha} \tag{6}$$

Calculating the marginal productivity of capital and equating it to the interest rate, $r$ augmented by capital depreciation $\delta$, we obtain the expression for the balanced growth path capital labor ratio, $\kappa_t^* = \left(\frac{1-\alpha}{r+\delta}\right)^{\frac{1}{\alpha}} A_t^{\frac{1}{\alpha}}$. Substituting it into equation (6) implies that the average wage in balanced growth path, $\left(w_t^L\right)^* = \alpha \left(\frac{1-\alpha}{r+\delta}\right)^{\frac{1-\alpha}{\alpha}} A_t^{\frac{1}{\alpha}}$ does not depend on the total supply of workers $L_t$. Hence, in the short run, the change in labor supply due to immigration affects average wages only if (and by the amount that) it affects the capital-labor ratio. Assuming that the technological progress ($\Delta A_t/A_t$) is exogenous to the immigration process, the percentage change of average wages due to immigration can be expressed as a function of the percentage response of $\kappa_t$ to immigration. Taking partial log changes of (6) relative to immigration we have:

[9] The "average wage" $w_t^L$ is obtained by averaging the wages of each group (by education, skill and nativity) weighting them by the share of the group in total employment.

[10] We describe the capital data and their dynamic behavior empirically in section 6.2.

$$\frac{\Delta w_t^L}{w_t^L} = (1-\alpha)\left(\frac{\Delta\kappa_t}{\kappa_t}\right)_{immigration} \tag{7}$$

where $(\Delta\kappa_t/\kappa_t)_{immigration}$ is the percentage deviation of the capital-labor ratio from $k_t^*$ due to immigration. With full capital adjustment and the economy in balanced growth path, $(\Delta\kappa_t/\kappa_t)_{immigration}$ equals 0. At the same time, if one assumes fixed total capital, $K_t = \overline{K}$, then $(\Delta\kappa_t/\kappa_t)_{immigration}$ equals the negative percentage change of employment due to immigration: $-\frac{\Delta F_t}{L_t}$, where $\Delta F_t$ is the increase in foreign-born workers in the period considered and $L_t$ is the labor aggregate at the beginning of the period. In the extreme case in which we keep capital unchanged over fourteen years of immigration, 1990-2004, the inflow of immigrants, equal to roughly 11% of the initial labor force, combined with a capital share $(1-\alpha)$ equal to 0.34, implies a negative effect on average wages of 3.5 percentage points.

Accounting for the sluggish yearly response of capital and for yearly immigration flows, however, we can estimate the *actual* response of the capital-labor ratio to immigration flows in the 1990-2004 period without the extreme assumption that capital be fixed for 14 years. We do this in section 6.2 when we revisit the short-run /long-run analysis.

### 3.3 Effects of Immigration on Wages

We now use the production function (1) to calculate the demand functions and wages for each type of labor at a given point in time. Choosing output as the numeraire good, in a competitive equilibrium the (natural logarithm of) the marginal productivity of U.S.-born workers $(H)$ in group $k, j$, equals (the natural logarithm of) their wage:

$$\ln w_{Hkjt} = \ln\left(\alpha A_t \kappa_t^{1-\alpha}\right) + \frac{1}{\delta}\ln(L_t) + \ln\theta_{kt} - \left(\frac{1}{\delta} - \frac{1}{\eta}\right)\ln(L_{kt}) + \ln\theta_{kjt} - \left(\frac{1}{\eta} - \frac{1}{\sigma_k}\right)\ln(L_{kjt}) + \ln\theta_{Hkjt} - \frac{1}{\sigma_k}\ln(H_{kjt}) \tag{8}$$

We assume that the relative efficiency parameters $(\theta_{kt}, \theta_{kjt}, \theta_{kjHt})$ as well as total factor productivity $A_t$ depend on technological factors and are therefore independent from the supply of foreign-born.

Let us define the change in the supply of foreign-born due to immigration between two censuses in group $k$, $j$ as $\Delta F_{kjt} = F_{kjt+10} - F_{kjt}$. We can use the demand function (8) to derive the effect of immigration on native wages. The overall impact of immigration on natives with education $k$ and experience $j$ can be decomposed into three effects that operate through $L_{kjt}$, $L_{kt}$ and $L_t$. First, a change in the supply of foreign-born workers with education $k$ and experience $j$ affects the wage of natives with identical education and experience by changing each one of the terms $L_{kjt}$, $L_{kt}$ and $L_t$ in expression (8). Second, a change in the supply of foreign-born workers with education $k$ and experience $i \neq j$ affects the wage of natives with identical education $k$ but different experience $j$ by changing the terms $L_{kt}$ and $L_t$. Third, a change in the supply of foreign-born workers with education $m \neq k$ affects native workers with different education $k$ only through a change in $L_t$. Aggregating the changes in wage resulting from immigration in each skill group as well as the response of capital-labor ratio $\kappa_t$ yields the wage change due to immigration for each home-born worker. The exact expression of each of the effects described above is provided in Appendix A.

Before showing the formula for the *total* effect of immigration on the wage of a home born worker of education $k$ and experience $j$, let us show the formula for a *partial* effect of the type emphasized in large part of the previous literature. If we only consider the impact of immigrants with education $k$ and experience $j$ on the wages of natives with identical education and experience, keeping the aggregates $L_{kt}$, $L_t$ and $\kappa_t$ constant, we obtain what a large part of the previous literature calls the "effect of immigrants on wages". This, in fact, measures a *partial* effect, keeping supply in all other skill groups constant and keeping constant the aggregates $L_{kt}$ and $L_t$ . Such effects have been estimated in the existing literature by regressing the wage of natives $\ln(w_{Hkjt})$ on the total number of immigrants in the same group $k, j$ in a panel across groups over census years, controlling for year-specific effects (absorbing the variation of $L_t$) and education-by-year specific effects (absorbing the variation of $L_{kt}$) (e.g. Borjas 2003). The resulting partial elasticity expressed as the percentage variation of native wages $(\Delta w_{Hkjt}/w_{Hkjt})$ in response to the percentage variation of foreign employment in the group $(\Delta F_{kjt}/F_{kjt})$, is given by the following expression:

$$\varepsilon_{kjt}^{partial} = \left.\frac{\Delta w_{Hkjt}/w_{Hkjt}}{\Delta F_{kjt}/F_{kjt}}\right|_{L_{kt}, L_t \text{ constant}} = \left[\left(\frac{1}{\sigma_k} - \frac{1}{\eta}\right)\left(\frac{s_{Fkjt}}{s_{kjt}}\right)\right] \tag{9}$$

The variable $s_{Fkjt}$ is the share of overall wages paid in year $t$ to foreign workers in group $k, j$, namely $s_{Fkjt} = \frac{w_{Fkjt}F_{kjt}}{\sum_m \sum_i (w_{Fmit}F_{mit} + w_{Hmit}H_{mit})}$. Analogously, $s_{kjt} = \frac{w_{Fkjt}F_{kjt} + w_{Hkjt}H_{kjt}}{\sum_m \sum_i (w_{Fmit}F_{mit} + w_{Hmit}H_{mit})}$ is the share of the total wage bill in year $t$ accounted for by all workers in group $k, j$.

By construction, the elasticity $\varepsilon_{kjt}^{partial}$ captures only the effect of immigration on native wages operating through the term $\left(\frac{1}{\eta} - \frac{1}{\sigma k}\right)\ln(L_{kjt})$ in (8). According to the standard assumption of the existing literature, U.S.- and foreign-born workers are perfect substitutes within group $k, j$ $(\sigma_k = \infty)$ and share the same efficiency $(\theta_{kjHt} = \theta_{kjFt})$ which implies $s_{Fkjt}/s_{kjt} = \varkappa_{Fkjt}/\varkappa_{kjt}$, where $\varkappa_{Nkjt}$ denotes the share of total employment represented by workers of nativity $N$ $(= H, F)$, education $k$, experience $j$ in year $t$, namely $\varkappa_{Fkjt} = \frac{F_{kjt}}{\sum_m \sum_i (F_{mit} + H_{mit})}$. Given these assumptions, expression (9) simplifies to $\varepsilon_{kjt}^{partial} = -\frac{1}{\eta}$: the harder it is to substitute between workers with different levels of experience (i.e. the lower $\eta$), the stronger is the negative impact that immigrants have on the wages of natives with similar educational and experience attainment. In the general case that we consider $(0 < \sigma_k < \infty)$, $\varepsilon_{kjt}^{partial}$ is still negative but smaller in absolute value than $\frac{1}{\eta}$, the reason being that the negative wage effect of immigrants on natives is partly attenuated by their imperfect substitutability.

Using estimates of the parameters $\sigma_k$ and $\eta$ as well as data on wages and employment, the *partial* elasticity $\varepsilon_{kjt}^{partial}$ can be easily calculated (see section 5.4 below). The problem is that this elasticity does not provide *any* indication on the total effect of immigration on the wages of natives in group $k, j$. The reason is that, to calculate the total effect, we also need to account for the changes in $L_{kt}$ and $L_t$ produced by immigration, for the fact that immigration alters the supply of foreign-born workers in all other education and experience groups and for the response of $\kappa_t$ to immigration. Once we do so, and aggregate all the effects, the total effect of immigration on the wages of native workers in group $k, j$ is given by the following expression:

$$\begin{aligned}\left(\frac{\Delta w_{Hkjt}}{w_{Hkjt}}\right)^{Total} &= \frac{1}{\delta}\sum_m\sum_i\left(s_{Fmit}\frac{\Delta F_{mit}}{F_{mit}}\right)+\left(\frac{1}{\eta}-\frac{1}{\delta}\right)\left(\frac{1}{s_{kt}}\right)\sum_i\left(s_{Fkit}\frac{\Delta F_{kit}}{F_{kit}}\right)+ \\ &+\left(\frac{1}{\sigma_k}-\frac{1}{\eta}\right)\left(\frac{1}{s_{kjt}}\right)\left(s_{Fkjt}\frac{\Delta F_{kjt}}{F_{kjt}}\right)+(1-\alpha)\left(\frac{\Delta\kappa_t}{\kappa_t}\right)_{immigration}\end{aligned} \tag{10}$$

It is easy to provide intuition for each term in expression (10) by referring to the labor demand equation (8). The term $\frac{1}{\delta}\sum_m\sum_i\left(s_{Fmit}\frac{\Delta F_{mit}}{F_{mit}}\right)$ is a positive total effect on the productivity of workers in group $k, j$ due to the increase in supply of all types of labor that is, home labor benefits from the increase in aggregate labor caused by imperfectly substitutable workers. This effect operates through $\frac{1}{\delta}\ln(L_t)$ in (8) which is positive for $\delta > 0$. The term $\left(\frac{1}{\eta}-\frac{1}{\delta}\right)\left(\frac{1}{s_{kt}}\right)\sum_i\left(s_{Fkit}\frac{\Delta F_{kit}}{F_{kit}}\right)$ is the additional negative effect on productivity generated by the supply of immigrants within the same education group. As those immigrants are closer substitutes for natives in group $k, j$ due to similar education, they have an additional depressing effect on their wage. This effect operates through the term $\left(\frac{1}{\delta}-\frac{1}{\eta}\right)\ln(L_{kt})$ in (8) which is negative if $\eta > \delta$. The term $\left(\frac{1}{\sigma_k}-\frac{1}{\eta}\right)\left(\frac{1}{s_{kjt}}\right)\left(s_{Fkjt}\frac{\Delta F_{kjt}}{F_{kjt}}\right)$ is the additional negative effect due to the supply of immigrants with the same education and experience as natives in group $k, j$. This last effect operates through $\left(\frac{1}{\eta}-\frac{1}{\sigma k}\right)\ln(L_{kjt})$ in (8) and it is exactly the partial effect $\varepsilon_{kjt}^{partial}$ multiplied by the percentage change $\frac{\Delta F_{kjt}}{F_{kjt}}$. Finally, the term $(1-\alpha)\left(\Delta\kappa_t/\kappa_t\right)_{immigration}$ is the wage change due to imperfect capital adjustment and operates through $\ln\left(\alpha A_t\kappa_t^{1-\alpha}\right)$ in (8). Clearly, since the *total* effect aggregates the *partial* effect plus 40 other cross-effects (32 in the double summation and 8 in the single summation) and a capital-adjustment term, it will typically be very different from $\varepsilon_{kjt}^{partial} * \frac{\Delta F_{kjt}}{F_{kjt}}$. In fact, when immigration is large in groups with education and experience different from $k$ and $j$, the effect $\left(\frac{\Delta w_{Hkjt}}{w_{Hkjt}}\right)^{Total}$ tends to be positive while when immigration is large in the group with the same education and experience as $k$ and $j$, the effect $\left(\frac{\Delta w_{Hkjt}}{w_{Hkjt}}\right)^{Total}$ tends to be negative. In contrast, the effect $\varepsilon_{kjt}^{partial} * \frac{\Delta F_{kjt}}{F_{kjt}}$ would *always* be negative for reasonable parameters values.

As they are not perfect substitutes for U.S.-born workers, the impact of immigrants on wages of foreign-born workers in the U.S. would be somewhat different. The percentage change in wages of foreign-born of education $k$ and experience $j$ as a consequence of total immigration is:

$$\left(\frac{\Delta w_{Fkjt}}{w_{Fkjt}}\right)^{Total} = \frac{1}{\delta}\sum_m\sum_i\left(s_{Fmit}\frac{\Delta F_{mit}}{F_{mit}}\right)+\left(\frac{1}{\eta}-\frac{1}{\delta}\right)\left(\frac{1}{s_{kt}}\right)\sum_i\left(s_{Fkit}\frac{\Delta F_{kit}}{F_{kit}}\right)+ \tag{11}$$
$$+\left(\frac{1}{\sigma_k}-\frac{1}{\eta}\right)\left(\frac{1}{s_{kjt}}\right)\left(s_{Fkjt}\frac{\Delta F_{kjt}}{F_{kjt}}\right)+(1-\alpha)\left(\frac{\Delta\kappa_t}{\kappa_t}\right)_{immigration}-\frac{1}{\sigma_k}\frac{\Delta F_{kjt}}{F_{kjt}}$$

The first four terms of expression (11) are identical to those in (10). Immigration in all other skill groups (and capital adjustment) has the same effect on the wages of home- and foreign-born in group $k, j$ with the exception of immigrants in the $k$, $j$ group itself, represented by the final term $-\frac{1}{\sigma_k}\frac{\Delta F_{kjt}}{F_{kjt}}$. This term is negative for $\sigma_k > 0$ . The term captures an extra negative impact of immigration on the wages of foreign-born due to their perfect substitutability with immigrants in the same group. Immigrants compete in occupations, sectors and jobs with previous immigrants more than natives and this causes the additional negative effect on the wage of foreign born workers. For $\sigma_k = \infty$, the effects of immigration on the wages of workers in group $k, j$ would be identical, independent of their nativity.

Using the percentage changes in wage for each skill group, we can then aggregate and find the effect of immigration on several representative wages. The average wage for the whole economy in year $t$, inclusive of U.S.- and Foreign-born workers is given by the following expression: $\overline{w}_t = \sum_k\sum_j(w_{Fkjt}\varkappa_{Fkjt} + w_{Hkjt}\varkappa_{Hkjt})$. Similarly, the average wage of U.S.-born and foreign-born workers only can be expressed as weighted averages of individual group wages: $\overline{w}_{Ht} = \sum_k\sum_j(w_{Hkjt}\varkappa_{Hkjt})/\sum_k\sum_j\varkappa_{Hkjt}$ and $\overline{w}_{Ft} = \sum_k\sum_j(w_{Fkjt}\varkappa_{Fkjt})/\sum_k\sum_j\varkappa_{Fkjt}$, respectively (recall that the variables $\varkappa_{Nkjt}$ represent shares in total employment). The percentage changes in the average wage of native workers as a consequence of changes of each group's wage due to immigration is given by the following expressions :

$$\frac{\Delta\overline{w}_{Ht}}{\overline{w}_{Ht}} = \frac{\sum_k\sum_j(\frac{\Delta w_{Hkjt}}{w_{Hkjt}}\frac{w_{Hkjt}}{\overline{w}_{Ht}}\varkappa_{Hkjt})}{\sum_k\sum_j\varkappa_{Hkjt}} = \frac{\sum_k\sum_j(\frac{\Delta w_{Hkjt}}{w_{Hkjt}})s_{Hkjt}}{\sum_k\sum_j s_{Hkjt}} \tag{12}$$

where $\frac{\Delta w_{Hkjt}}{w_{Hkjt}}$ represents the percentage change in the wage of home-born in group $k, j$ due to immigration and its expression is given in (10). Similarly, the percentage change in the average wage of foreign-born workers is:

$$\frac{\Delta\overline{w}_{Ft}}{\overline{w}_{Ft}} = \frac{\sum_k\sum_j(\frac{\Delta w_{Fkjt}}{w_{Fkjt}}\frac{w_{Fkjt}}{\overline{w}_{Ft}}\varkappa_{Fkjt})}{\sum_k\sum_j\varkappa_{Fkjt}} = \frac{\sum_k\sum_j(\frac{\Delta w_{Fkjt}}{w_{Fkjt}})s_{Fkjt}}{\sum_k\sum_j s_{Fkjt}} \tag{13}$$

Finally, by aggregating the total effect of immigration on the wages of all groups, natives and foreign, we can obtain the effect of immigration on average wages:

$$\frac{\Delta \overline{w}_t}{\overline{w}_t} = \sum_k \sum_j \left( \frac{\Delta w_{Fkjt}}{w_{Fkjt}} \frac{w_{Fkjt}}{\overline{w}_{Ft}} \varkappa_{Fkjt} + \frac{\Delta w_{Hkjt}}{w_{Hkjt}} \frac{w_{Hkjt}}{\overline{w}_{Ht}} \varkappa_{Hkjt} \right) = \sum_k \sum_j \left( \frac{\Delta w_{Fkjt}}{w_{Fkjt}} s_{Fkjt} + \frac{\Delta w_{Hkjt}}{w_{Hkjt}} s_{Hkjt} \right) \tag{14}$$

Recall that the variables $s_{Nkjt}$ represent the share in total wages and notice that the correct weighting to obtain the percentage change on *average wages* is the share in the wage bill and not the share in employment. Due to constant returns to scale of the aggregate production function (1), while some of the wage changes are positive and others negative, when weighted by their wage shares the summation of these changes equals 0 once capital has adjusted fully (i.e. in the long run); hence, the change in the overall average wage in (14) is approximately 0 in the long run. However, if home- and foreign-born are not perfectly substitutable, the overall effect on the wage of home-born, (12) need not be 0 but will be positive instead and the effect on average wage of foreign-born (13) will be negative. We also adopt the same averaging procedure (weighting percentage changes by wage shares) in calculating the effect of immigration on specific groups of U.S.-born and foreign born workers. For instance, the changes in average wages of college Educated U.S. born workers is calculated as $\sum_j \left( \frac{\Delta w_{HCOGjt}}{w_{HCOGjt}} s_{HCOGjt} \right) / \sum_j s_{HCOGjt}$ and the change in average wages of foreign-born High-School dropouts is calculated as:$\sum_j \left( \frac{\Delta w_{FHSDjt}}{w_{FHSDjt}} s_{FHSDjt} \right) / \sum_j s_{FHSDjt}$, and so on.

# 4 Data Description and Preliminary Evidence

The data we use are from the integrated public use microdata samples (IPUMS) of the U.S. decennial Census and of the American Community survey (Ruggles et al, 2005). In particular we use the general (1%) sample for Census 1960, the 1% State Sample, Form 1, for Census 1970, the 5% State sample for the Censuses 1980 and 1990, the 5% Census Sample for year 2000 and the 1/239 American Community Survey (ACS) Sample for the year 2004. As those are all weighted samples we use the variable "personal weight" to construct all the average and aggregate statistics below. We consider people aged 17 to 65 not living in group quarters, who worked at least one week in the previous year and earned a positive amount in salary income. We convert the current wages to constant wages (in 2000 U.S. \$) using the C.P.I.-based deflator across years. We define the four schooling groups using the variable that identifies the highest grade attended (called "HIGRADEG" in IPUMS) for census 1960 to 1980 while we use the categorical variable (called "edu99" in IPUMS) for censuses 1990 and 2000 and ACS 2004. Years of experience are calculated using the variable "age" and assuming that people without an high school degree enter the labor force at age 17, people with high school degree enter at 19, people with some college enter at 21 and people with a college degree enter at 23. Finally, yearly wages are based on the variable salary and income wage (called "INCWAGE" in IPUMS). Weekly wages are obtained dividing that value by the number of weeks worked[11]. The status of "foreign-born" is given to those workers whose place of birth (variable "BPL") is not within the USA (or its territories overseas) and did not have U.S. citizenship at birth (variable "CITIZEN")[12]. The average wage for workers in a cell, (the variable $w_{Nkjt}$ for $N = \{H, F\}$, $k = \{HSD, HSG, COD, COG\}$ and $j = \{1, 2..., 8\}$) is calculated as the weighted average of individual wages in the cell using the personal weight ("PERWT") assigned by the U.S. census. The total number of workers in each cell ($H_{kjt}$and $F_{kjt}$) is calculated as the weighted sum of workers belonging to that cell. These data allow us to construct the variables $\varkappa_{Nkjt}$ and $s_{Nkjt}$, the share of each group in the total wage bill and in total employment for each represented year $t$. These data are also used to estimate the parameters $\delta$, $\eta$, $\sigma_k$ needed to calculate the effects of immigration on the wage of each type of worker. When estimating the structural parameters $\delta$, $\eta$, $\sigma_k$ we always use the whole panel of data 1960-2004. When we calculate and simulate the effect of immigration on real wages we focus on the most recent period 1990-2004. Before proceeding with the econometric analysis, let us present some salient features of the immigration and wage data as well as some simple statistics suggesting the plausibility of the hypothesis of imperfect substitutability between U.S. and foreign-born workers with similar observable skills.

Tables 1 and 2 report, respectively, the share of foreign-born workers in each education-experience group for each of the years considered and the real weekly wages (in 2000 U.S. $) for U.S. native workers in each education and experience group in each of the years considered. The wages for each group are calculated as described above. In the rows marked as "All Experiences" we report the total (average) value for the variables relative to the whole educational group. Two facts emerge even from a cursory look at the tables. First, Table 1 confirms that the distribution of foreign-born across educational groups has been uneven (increasingly so) over the period considered. In the year 2004, almost 35% of the workers with no degree were foreign-born, with some experience sub-groups counting more than 40% foreign-born. Following this group the one of college graduates shows the second highest concentration of foreign-born: almost 15% in the overall group, reaching 18% in some experience sub-groups. In contrast, the group of college dropouts contained only 9.5% of foreign-born workers and, in some experience sub-groups those were less than 8% of the total. As is well known, the relative distribution of foreign-born workers across educational groups in the U.S. has been U-shaped over the education spectrum and increasingly so over time. Foreign-born were, and increasingly are, over-represented among the groups of workers with the lowest and highest education and are under-represented among the two intermediate groups. The second fact emerging from Table 2 is the poor performance of real weekly wages of U.S.-born workers without a high school degree, especially during the last two and a half decades. In contrast, the performance of real wages of college educated has been spectacular, particularly during the last two and a half decades, with

[11] For the Census 1960 and 1970 only a categorical variables that measures weeks worked exists and is called "WKSWRK2". Individuals are assigned the middle value of the variable in the interval.

[12] The variable CITIZEN is not available in census 1960. For that year we consider all people born outside the U.S. as foreign-born.

the two intermediate schooling groups performing in between.

In order to provide a synthetic and effective representation of the two facts described above, focussing on the most recent 14 years, we present two figures. Figure 3 illustrates the percentage growth of immigrant employment for each of the four educational groups. The lightly shaded 3-D columns represent immigrants for the 1990-2000 period as percentage of 1990 employment in each education group. The darkly shaded 3-D columns represent immigration flows during the 1990-2004 period as a percentage of initial employment. The graph confirms the U-shaped distribution of immigrants along the educational spectrum, with a more pronounced U-shape when we consider the longer period of immigration (1990-2004). Figure 4, on the other hand, shows the growth rate of real wages of native workers by education group. The lightly shaded 3-D columns represent the real percentage change of yearly wages in the 1990-2000 period, the darkly shaded 3-D columns represent the 1990-2004 change. One sees very clearly the negative performance of real wages for the least educated (almost one percentage point loss in real wage each year) and the exceptional performance of real wages of college graduates (more than one percentage point gain each year). The natural questions stemming from these facts are: (i) How much of the negative performance in the wage of native dropouts is due to the large immigration flow in that group, (ii) How much of the college-high school dropout wage gap widening is due to immigration? and (iii) Given that the average wages of U.S. natives grew by around 12.5% in the 1990-2004 period, and overall immigration increased employment by almost 12%, would overall wage growth have been larger without the increase in labor supply due to immigration? We will address these questions in Section (6).

Before presenting the estimates of the structural parameters of our model, especially $\sigma_k$ , the elasticity of substitution between U.S.- and foreign-born, let us put forward three observations that seem to suggest imperfect substitutability between U.S.- and foreign-born workers in production (that is, $\sigma_k < \infty$). Even considering workers who have identical measurable human capital (education and experience), foreign- and U.S.-born differ in several respects that are relevant to the labor market. First, immigrants are a selected group of their original populations and have skills, motivations and tastes that may set them apart from natives. Second, in manual and intellectual works they have culture-specific skills (e.g. cooking, crafting, opera singing, soccer playing) as well as limits (knowledge of English language or American culture) creating comparative advantages in some and comparative disadvantages in other jobs. Third, and most important, due to the portability of skills, information via networks or historical accidents, foreign-born tend to choose different occupations than natives even for a given education and experience. As services of different occupations are imperfectly substitutable, this would imply imperfect substitutability between natives and foreign-born in the same education-experience group (i.e. an effect on relative wages when relative supply changes).

Differences in the occupational choice between natives and foreign-born with the same education are illustrated in Table 3. Following Welch (1990) and Borjas (2003) we calculate the "index of congruence" in the choice of 180 occupations (from the variable "occupation 1950" homogenized across censuses definitions) between the group of native workers and the group of foreign-born workers with the same education . The index of congruence is calculated by constructing a vector of shares in each occupation for each group and computing the centered correlation coefficient between these vectors for the two groups. A value of the index equal to 1 implies an identical distribution of workers among occupations for the two groups, a value equal to -1 implies an exactly "complementary" distribution. The first column in the table reports the U.S.-foreign born occupational congruence for each education group. By way of comparison, the remaining columns report the indices of congruence between natives in different education groups. The index of congruence between U.S.- and foreign-born with identical education is between 0.6 and 0.7 a value comparable to the congruence between native high school dropouts and native high school graduates (value 0.68 reported in the second column of Table 3). Also (see Welch, 1979), these index values are comparable to the congruence between U.S. born workers with different experience levels. Hence, given that an extensive literature shows imperfect substitutability between U.S. workers with different education and experience (Welch, 1979; Card and Lemieux, 2001), if part of imperfect substitutability is due to occupational choice we would also expect it to hold for natives and foreign-born with similar education.

A second interesting fact relative to the occupational choice of immigrants is that new immigrant tend to work disproportionately in those occupations where foreign-born are already over-represented. Possibly because of networking, information diffusion or the type of skills required, new immigrants are attracted to jobs and tasks already disproportionately occupied by previous immigrants. This implies a stronger wage competition (substitution) for those jobs compared to other jobs primarily held by natives and sometimes complementary in production. This tendency can be captured by regressing the increase in the share of foreign-born employment in an occupation on the initial share of foreign-born workers in that occupation, as follows:

$$\Delta(sh_{t,t+n}^{For})_i = \alpha + \beta \left(sh_t^{For}\right)_i + \varepsilon_i \tag{15}$$

where $\Delta(sh_{t,t+n}^{For})_i$ is the change in the share of foreign-born working in occupation $i$ between year $t$ and year $t+n$, $\left(sh_t^{For}\right)_i$ is the share of foreign-born in occupation $i$ in year $t$ and $\varepsilon_i$ is a zero-mean random error. A positive and significant estimate of $\beta$ means that new immigrants tend to disproportionately staff jobs with an already high density of immigrants (and mostly affect those wages). Table 4 reports the estimates of $\beta$ for the period 1980-1990 (first column) and 1990-2004 (second column). The first row reports the results pooling data for workers in each of the four schooling groups while the other rows report regression results considering each education group separately. For each period and each group, the coefficient $\beta$ is estimated to be very significant and positive. For instance, an occupation such as "Farm Laborers" (where foreign-born represented 46 percent of employment in 1990) experienced an increase in that share that was 20% greater than the increased foreign share of "Farm Managers" (only 6% foreign-born in 1990) during the 1990-2004 period. These two occupations have clearly significant complementarity (both types are needed to run a farm) and new immigrants competed mostly with existing ones for Laborer positions (hurting their wages) while complementing U.S.-born who work as Managers (possibly increasing demand for those). Similar stories can be told for several other occupations. This implies that new immigrants disproportionately affected occupations where old immigrants were already over-represented. Further, this implies that the wages mostly affected by immigration within each education group were likely those of previous immigrants.

# 5 Parameter Estimates

## 5.1 Estimates of $\sigma_k$

The model developed in Section (3.1) provides us with the framework to estimate the parameter $\sigma_k$ for each education group. Calculating the natural logarithm of the ratio of the wages of U.S.-born and foreign-born workers within the same group $k, j$ we obtain the following relation:

$$\ln(w_{Hkjt}/w_{Fkjt}) = -\frac{1}{\sigma_k}\ln(H_{kjt}/F_{kjt}) + \ln(\theta_{Hkjt}/\theta_{Fkjt}) \tag{16}$$

which defines the relative labor demand for foreign- and U.S.-born workers in group $k, j$. Equation (16) can be used to estimate the coefficient $\frac{1}{\sigma_k}$ (i.e. the elasticity of relative demand) as long as we identify a source of variation in relative supply $\ln(H_{kjt}/F_{kjt})$ that is independent of the variation of relative efficiency $\ln(\theta_{Hkjt}/\theta_{Fkjt})$.

Our estimation strategy works as follows. Due to technological reasons such as skill-biased technical change, sector biased technical change, increased international competition and others over the period 1960-2004, the profiles of the returns to education and to experience have changed differently across occupations. Accordingly, we allow the relative efficiency of U.S. and foreign-born, $\ln(\theta_{Hkjt}/\theta_{Fkjt})$, to have a systematic component that may vary by education and experience over time and we control for education by year fixed effects ($D_{kt}$) as well as experience by year fixed effects ($D_{jt}$). At the same time, different education-experience groups may include U.S.- and foreign-born workers of systematically heterogeneous quality; hence, we control for experience by education fixed effects ($D_{kj}$). Conditional on these controls, we assume that the residual decennial changes in relative employment within each education-experience cell over time, $\ln(H_{kjt}/F_{kjt})$, are due to random supply shocks such as demographic factors in the U.S. and in the immigration countries. Thus, using the IPUMS data from 1960 through 2004 we estimate the following regression:

$$\ln(w_{Hkjt}/w_{Fkjt}) = D_{kj} + D_{kt} + D_{jt} - \frac{1}{\sigma_k}\ln(H_{kjt}/F_{kjt}) + u_{kjt} \tag{17}$$

where $u_{kjt}$ is a residual random, zero-mean disturbance. In total we use 192 observations (8 experience by 4 education groups over 6 years: 1960, 1970, 1980, 1990, 2000, 2004) and we include education by experience $D_{kj}$ fixed effects, education by time, $D_{kt}$, fixed effects and experience by time, $D_{jt}$, fixed effects. The variables $w_{Hkjt}$, $w_{Fkjt}$, $H_{kjt}$, $F_{kjt}$ are constructed as described in section 4 above. Table 5 reports the estimates of the parameter $\frac{1}{\sigma_k}$. The first row of Table 5 reports the estimates when we impose the same elasticity of substitution between U.S. and foreign-born workers within each of the four education groups ($\sigma_k = \overline{\sigma}$), while the following four rows report the estimates when we allow $\sigma_k$ to differ by education group. Specifications 1 and 2 use all workers to calculate the supply and wages for each group, specification 3 and 4 use male workers only to calculate wages, while still all workers in calculating the supply. All four specifications are estimated using weighted OLS, weighting each observation by the total employment in the cell. Specifications 5 and 6 use unweighted least squares as the estimation method, while specifications 7 and 8 omit year 2004 (not a census year) and the initial year 1960 as immigration was extremely low before 1960 and foreign-born a very small group in 1960. We also tried other sub-samples beginning with year 1980 or ending in year 2000 (not reported) obtaining very similar results. The estimates of $\frac{1}{\sigma}$ are significant different from zero, quite stable across specifications and always between 0.1 and 0.2. The specifications using yearly wages tend to produce somewhat larger estimates (between 0.15 and 0.2), while those using weekly wages give estimates between 0.10 and 0.13. The standard errors are generally around 0.04. The estimates as a whole strongly support the idea of imperfect substitutability between U.S. and foreign-born workers, confirming our intuition. Moreover, the estimates imply an elasticity of substitution between the two groups within an education-experience cell between 5 and 10. Hence, we observe imperfect substitutability but, reasonably, not to the extent observed between educational groups (usually credited with a $1.5 - 2.5$ elasticity of substitution) and only slightly above the one observed between experience groups for U.S. natives (estimated between 3 and 5)[13].

The remaining rows of Table 5 show that allowing $\sigma_k$ to vary across education groups, the estimates become somewhat more imprecise. However, they are still significant at the 5% level in each specification and most of them are still between 0.1 and 0.2. We do not observe any clear pattern in differences of degree of substitution between groups. Interestingly, the group of high school dropouts exhibits estimates of $\frac{1}{\sigma_k}$ as large as (if not larger than) the other groups, confirming that imperfect substitutability between Foreign and U.S.-born workers is certainly a feature present at low levels of education and is possibly enhanced in that group by skill differences and labor market segmentation. The F-test of the hypothesis of equal elasticities across groups ($\sigma_{HSD} = \sigma_{HSG} = \sigma_{COD} = \sigma_{COG} = \overline{\sigma}$) is reported in the last row of Table 5 together with the p-value (in parenthesis) for rejecting the null. In no case, except for specification 7, can one reject the null of identical elasticities at usual confidence levels. As these results are confirmed by grouping foreign-workers by effective skills (see next

[13] See section 5.3 below for those estimates in our paper and in the literature.

section) for the remainder of the paper we will operate under the assumption of a common elasticity $\overline{\sigma}$.

### 5.2 Effective Skills

By grouping U.S.- and foreign-born individuals according to their years of working experience, one could be misclassifying the effective skills of foreign-born, assigning them to a group that is not their most natural term of comparison. Employers may value differently work experience accrued in the U.S. market from the one accrued abroad, hence an immigrant with some experience abroad should be reclassified using the "effective" value of that experience in terms of U.S.-accrued experience. The most natural comparison (and competitor) for an immigrant with 10 years of experience abroad may be a U.S. worker with only 4 years of experience, if experience abroad generates only 40% as much accumulation of valuable skills for the U.S. market. This point is noted and developed in Borjas (2003), section VI, and is relevant to the present study as well. If a reclassification by effective experience groups affects those differences between U.S. and foreign-born workers that are responsible for imperfect substitutability, the estimated elasticity of substitution $\sigma$ may change when using effective skill groups. We deal with this issue in the same way as Borjas (2003) did. First, we split the years of working experience of immigrants between experience in the U.S. and experience abroad. We do this by using the variable "year of immigration"[14] which defines the year of entry into the U.S. As the variable is categorical, the exact year of entry of an individual is chosen to be the year in the middle of the category interval. In so doing, we divide foreign born worker into two types: those who always worked in the U.S. (migrated before beginning their working period) and those who worked abroad for some years. We then use the "conversion" factors between foreign and U.S. experience calculated in Borjas (2003). Those factors are based on a wage regression that calculates (pooling 1980-1990 data) the wage growth associated with one year of working experience abroad, relative to the growth of wage associated with one year of working experience in the U.S. Specifically, for foreign-born who always worked in the U.S. no conversion is needed and one year of their experience equals one year of experience of a U.S.-born worker. For immigrants who worked abroad, the years of experience abroad are multiplied by a factor of 0.4 while the years of experience in the U.S. are multiplied by a factor of 1.6 (Borjas, 2003, page 1356). This implies an "under-accumulation" of useful skills per year when working abroad and an over-accumulation (catching up) during the years of U.S. working experience. Once we have calculated the effective experience we group foreign workers in the usual 8 groups (0 to 40 years by 5-years cells) using this new variable. This reclassification does not change the relative wages and relative supplies entering equation (17) much. The partial correlation of the new variables with the original variables, even after controlling for all the dummies, is between 0.96 and 0.98. Table 6 shows the elasticities, $\frac{1}{\sigma_k}$, calculated using the new groupings for specifications exactly corresponding to those in Table 5. Each estimate is very close to the

[14] The variable is called "YRIMMIG" in the IPUMS notation.

corresponding one in Table 5. The constrained estimates of the first row still range between 0.1 and 0.2 as well as, for the most part, the unconstrained ones. In this case as well, one cannot reject (except for specification 1) the hypothesis of equal elasticity across education groups at usual significance levels. For robustness purposes, we still produce estimates of the other parameters ($\eta$ and $\delta$) in the next section using both simple and effective experience groupings. We do not find, however, any relevant differences in our analysis between adopting the definition of "effective" experience and the definition of "simple" experience .

### 5.3 Estimates of $\eta$ and $\delta$

While equation (17) allowed us to estimate the parameters $\sigma_k$, we can also use it to infer the systematic component of the efficiency terms $\theta_{Hkjt}$ and $\theta_{Fkjt}$. In particular, those terms can be obtained from the estimates of the fixed effects $\widehat{D}_{kj}$, $\widehat{D}_{kt}$ and $\widehat{D}_{jt}$ as:

$$\widehat{\theta}_{Hkjt} = \frac{\exp(\widehat{D}_{kj})\exp(\widehat{D}_{kt})\exp(\widehat{D}_{jt})}{1+\exp(\widehat{D}_{kj})\exp(\widehat{D}_{kt})\exp(\widehat{D}_{jt})}, \widehat{\theta}_{Fkjt} = \frac{1}{1+\exp(\widehat{D}_{kj})\exp(\widehat{D}_{kt})\exp(\widehat{D}_{jt})} \tag{18}$$

where we have imposed the standardization that the two efficiency terms add up to one. Using the values of $\widehat{\theta}_{Hkjt}$ and $\widehat{\theta}_{Fkjt}$ from above and the estimate $\widehat{\sigma}$ we can construct the aggregate labor input, following 4, as $\widehat{L}_{kjt} = \left[\widehat{\theta}_{Hkjt} H_{kjt}^{\frac{\widehat{\sigma}-1}{\widehat{\sigma}}} + \widehat{\theta}_{Fkjt} F_{kjt}^{\frac{\widehat{\sigma}-1}{\widehat{\sigma}}}\right]^{\frac{\widehat{\sigma}}{\widehat{\sigma}-1}}$. Indeed, the production function (1) and marginal pricing imply the following relationship between the compensation going to the composite labor input $L_{kjt}$ and its supply:

$$\ln(\overline{W}_{kjt}) = \ln\left(\alpha A_t^{\frac{1}{\alpha}} \kappa_t^{\frac{1-\alpha}{\alpha}}\right) + \frac{1}{\delta}\ln(L_t) + \ln\theta_{kt} - \left(\frac{1}{\delta} - \frac{1}{\eta}\right)\ln(L_{kt}) + \ln\theta_{kj} - \frac{1}{\eta}\ln(L_{kjt}) \tag{19}$$

where $\overline{W}_{kjt} = w_{Fkjt}(F_{kjt}/L_{kjt}) + w_{Hkjt}(H_{kjt}/L_{kjt})$ is the average wage paid to workers in the education-experience group $k$, $j$ and can be considered as the compensation to one unit of the composite input $L_{kjt}$[15]. Equation (19) provides the basis to estimate the parameter $\frac{1}{\eta}$ that measures the elasticity of relative demand for workers with identical education and different experience. Empirical implementation is achieved by rewriting it as:

$$\ln(\overline{W}_{kjt}) = D_t + D_{kt} + D_{kj} - \frac{1}{\eta}\ln(\widehat{L}_{kjt}) + e_{kjt} \tag{20}$$

where five period fixed effects $D_t$ control for the variation of $\ln\left(\alpha A_t^{\frac{1}{\alpha}} \kappa_t^{\frac{1-\alpha}{\alpha}}\right) + \frac{1}{\delta}\ln(L_t)$, time by education fixed effects $D_{kt}$ control for the variation of $\ln\theta_{kt} - \left(\frac{1}{\delta} - \frac{1}{\eta}\right)\ln(L_{kt})$ and education by experience fixed effects $D_{kj}$ capture the terms $\ln\theta_{kj}$ that we assume is constant over time. Once we control for these systematic shifts in demand our identifying assumption, closely tracking Borjas (2003), is that the remaining variation

[15] The wage $\overline{W}_{kjt}$ is an average of the wages paid to U.S.- and foreign-born workers in group $k, j$. The averaging weights are equal to the employment of each group relative to the composite $L_{kjt}$ which are very close to their share of the employment of group $k, j$.

in employment of foreign-born is due to supply shifts. Under this assumption, we consistently estimate the coefficient $-\frac{1}{\eta}$ in regression (20) by 2SLS, using $\ln(F_{kjt})$, the supply of foreign-born workers in each group, as an instrument for the variable $\ln(\widehat{L}_{kjt})$. Table 7 reports the estimated values of $\frac{1}{\eta}$. The first row of Table 7 reports the estimates based on yearly wages, while the second row uses weekly wages. Specification 1 of Table 7 uses the parameters estimated in specification 1 and 2 of Table 5 to construct $\overline{W}_{kjt}$ and $\widehat{L}_{kjt}$.Specification 2 of Table 7 uses the parameters estimated in specification 1 and 2 of Table 6 (effective skill grouping) to construct $\overline{W}_{kjt}$ and $\widehat{L}_{kjt}$. The estimated values of $\frac{1}{\eta}$ are between 0.2 and 0.3 with standard errors around 0.10. This implies a value of $\eta$ between 3.3 and 5. These value are very similar to those previously estimated in the literature; the parameter $\eta$ was first estimated in Card and Lemieux (2001). Their preferred estimates of $1/\eta$ for the United States over the period 1970-1995 (as reported in their Table III, columns 1 and 2) are between 0.2 and 0.26, thus implying a value of $\eta$ between 4 and 5. Borjas (2003) also produces an estimate of $1/\eta$. He uses immigration as a supply shifter but assumes perfect substitutability between U.S.- and foreign-born workers. His estimate is equal to 0.288 (with standard error 0.11), implying a value of $\eta$ equal to 3.5. In order to check how sensitive the estimate of $\eta$ is to imperfect substitutability between U.S.- and foreign-born workers, we also re-estimate the parameter $\eta$ assuming $\sigma = \infty$ in the construction of $\widehat{L}_{kjt}$. This is done in specification 3 and 4. The point estimates of $1/\eta$ decrease very slightly (by 0.02-0.03) and their difference with specification 1 and 2 is not significant.

Aggregating one level further, we can construct the CES composite $\widehat{L}_{kt}$ .We obtain the estimates $\widehat{\theta}_{kj}$ from the experience by education fixed effects in regression (20), as follows: $\widehat{\theta}_{kj} = \exp(\widehat{D}_{kj})/\sum_j \exp(\widehat{D}_{kj})$. Then we use the estimated values of $\eta$ to construct, according to the formula (3), $\widehat{L}_{kt} = \left[\sum_{j=1}^{8} \widehat{\theta}_{kj} L_{kjt}^{\frac{\widehat{\eta}-1}{\widehat{\eta}}}\right]^{\frac{\widehat{\eta}}{\widehat{\eta}-1}}$. The production function chosen, together with marginal cost pricing, implies that the compensation going to the labor input $L_{kt}$ and its supply satisfy the following expression:

$$\ln(\overline{W}_{kt}) = \ln\left(\alpha A_t^{\frac{1}{\alpha}} \kappa_t^{\frac{1-\alpha}{\alpha}}\right) + \frac{1}{\delta}\ln(L_t) + \ln\theta_{kt} - \frac{1}{\delta}\ln(L_{kt}) \tag{21}$$

where $\overline{W}_{kt} = \sum_j \left(\frac{L_{kjt}}{L_{kt}}\right) \overline{W}_{kjt}$ is the average wage in education group $k$[16]. Following the same strategy as we did before, we use the above expression as the basis for the estimation of $\frac{1}{\delta}$. In so doing, we rewrite (21) as follows:

$$\ln(\overline{W}_{kt}) = D_t + (Time\ Trend)_k - \frac{1}{\delta}\ln(\widehat{L}_{kt}) + e_{kt} \tag{22}$$

where the time dummies $D_t$ absorb the variation of the terms $\ln\left(\alpha A_t^{\frac{1}{\alpha}} \kappa_t^{\frac{1-\alpha}{\alpha}}\right) + \frac{1}{\delta}\ln(L_t)$ and the terms (*Time*

[16] The weight for the wage of each group equals the size of the composite input for that education-experience cell, $L_{kjt}$, relative to the size of the composite input for the whole education group $L_{kt}$. This is very similar to the share of group $k, j$ in the employment of educational group $k$.

$Trend)_k$ are education-specific time trends. These trends control for the systematic component of the efficiency terms $\ln\theta_{kt}$ that are assumed to follow a time trend specific to each educational group. Conditional on these controls, our identifying assumption is that any other change in employment of foreign-born within a group is a supply shift. Hence, we can estimate the equation (22) by 2SLS using $\ln(F_{kt})$ ( where $F_{kt} = \sum_j F_{kjt}$) as instrument for $\ln(L_{kt})$. Table 8 reports the estimates of $\frac{1}{\delta}$ . The first row uses yearly wages in the calculations, while the second uses weekly wages. Specifications 1 and 2 of Table 8 use the estimated values of $\widehat{\eta}$ and of $\widehat{\theta}_{kj}$ from specifications 1 and 2 of Table 7 to construct $\widehat{L}_{kt}$ and $\overline{W}_{kt}$. Specifications 3 and 4 use $\eta = \infty$ and symmetric weights $\theta_{kj}$ to construct $\widehat{L}_{kt}$ and $\overline{W}_{kt}$ .The estimated values are mostly between 0.4 and 0.5 (with standard errors around 0.15), consistent with an elasticity of substitution across education groups around 2. The parameter $\delta$ is certainly the most analyzed in the literature. Its key role in identifying the impact of increased educational attainment (as well as of skill-biased technological change) on wages made it the object of analysis in Katz and Murphy (1992), through Angrist (1995), Murphy et al. (1998), Krusell et al. (2000) and Ciccone and Peri (2005). The estimates for that parameter range between 1.4 and 2.5. Our estimates of $1/\delta$ fall between 0.4 and 0.5, and imply a $\delta$ in the vicinity of 2, which is consistent with previous estimates.

## 5.4 Partial Effects of Immigration on Wages

Before using the estimated values of the parameters $\delta, \eta$ and $\sigma$ and the formulas derived in section 3.3 to calculate the effects of immigration on wages let us use those estimates to point out an important corollary to our analysis. Most existing empirical studies on the effect of immigration on wages (including Borjas, Freeman and Katz, 1997, Card 2001, Friedberg, 2001, Card, 2001, section IV of Borjas, 2003 -but not section VII- and Borjas, 2006) carefully estimate the partial elasticity of native wages to immigration within the same skill group (expressed in our equation 9) and treat it as "the effect of immigration on wages". As we illustrated in section 3.3, this is simply a partial effect uninformative of the actual overall effect of immigration on wages unless we consider the whole distribution of immigrant skills, the cross effects among groups and the effect of capital adjustment. More importantly, the partial elasticity (9) is likely to be negative in any reasonable model as long as immigrants are closer substitutes to natives in the same group (education-experience) than they are to natives in other skill groups. Using estimates from Tables 5 and 7, the term $\left(\frac{1}{\sigma_k} - \frac{1}{\eta}\right)$ is calculated to be negative and between -0.10 and -0.20. This implies, for instance, that the percentage change in the wage of native workers in group $k, j$, $\Delta w_{Hkjt}/w_{Hkjt}$, would be between -1.1% and 2.2% in response to an inflow of immigrants equal to 11% of the initial employment in the group[17]. We use 11% as this equals the inflow of immigrants over the 1990-2004

[17] This value is calculated using formula (9) and multiplying the two sides by $\Delta F_{kjt}/F_{kjt}$ so that we obtain: $\Delta w_{Hkjt}/w_{Hkjt} = \left[\left(\frac{1}{\sigma_k} - \frac{1}{\eta}\right)\left(\frac{s_{Fkjt}}{s_{kjt}}\right)\frac{\Delta F_{kjt}}{F_{kjt}}\right]$. Then, notice that the term $\left(\frac{s_{Fkjt}}{s_{kjt}}\right)\frac{\Delta F_{kjt}}{F_{kjt}}$ is approximately equal to $\frac{\Delta F_{kjt}}{H_{kjt}+F_{kjt}}$ if the share of wage of foreign-born in group $k, j$ is similar to its share of employment in that group. Using $\frac{\Delta F_{kjt}}{H_{kjt}+F_{kjt}} = 11\%$ and $-0.2 < \left(\frac{1}{\sigma_k} - \frac{1}{\eta}\right) < -0.1$ we obtain a real wage change of $-2.2\% < \Delta w_{Hkjt}/w_{Hkjt} < -1.1\%$.

period as a percentage of total initial employment. If one fails to notice the *partial* nature of the elasticity used in the calculations above, one could be tempted to generalize these findings saying that immigration caused a negative 1.1 to 2.2 percent change (1990-2004) on average wages of natives and that groups such as high school dropout, for which the inflow of immigrants was as high as 20% of initial employment, lost as much as 4.4% of their wage. No such generalization is possible, however, as expression (9) only accounts for the effect on wages of immigrants in the same skill group, for fixed $L_{kt}, L_{kjt}$ and omits all the cross-group effects from immigrants in other skill groups. In fact (as we see in section 6.1, below) while sharing the same negative partial elasticity $\left(\frac{1}{\sigma} - \frac{1}{\eta}\right)$, the wages of natives across groups have very different responses to immigration, some being positive and other being negative, due to the relative sizes of skill groups and the relative strength of cross-groups effects. Limiting our attention to the elasticity $\varepsilon_{kjt}^{partial}$ or even emphasizing this effect too much would be misleading in evaluating the effect of immigration on wages.

# 6 Immigration and Wages: 1990-2004

We are now ready for the third and final step in producing our estimates (simulations) of the effects of immigration on wages of U.S. and foreign-born workers. The first step of the procedure (Section 3) required specifying a production function and deriving labor demand curves and the elasticity of wages to immigration of workers with different skills. The second step (Section 5) required estimation of the relevant structural parameters (elasticities of substitutions). The third step (this section) uses these estimates and the actual flow if immigrants by group during the 1990-2004 period, in the formulas previously derived, in order to calculate the effects of immigrants on wages of U.S. and foreign-born in individual groups as well as overall.

## 6.1 The long-run Effects of immigration on Wages

Table 9 contains the relevant simulation results, relative to the impact of immigration for the 1990-2004 period on wages of U.S. and foreign born workers, in the long run. We focus on the 1990-2004 period as it is the most recent covered by available Census and ACS data and it is the period of largest immigration in recent U.S. history. To obtain the simulated effects we proceeded as follows: First, using the formulas (10) and (11), the estimated parameters $\delta, \eta$ ,$\sigma$ and the percentage change in foreign born workers by skill group ($\Delta F_{kj,1990-2004}/F_{kj,1990}$), we calculate the percentage change of real wage for U.S.-born and foreign-born workers in each skill group $(k, j)$. Then we obtain the average wage change in each education group for foreign- and U.S.-born by weighting the percentage change of each experience sub-group by its wage share in the education group. This provides the entries in rows one to four and six to nine in Table 9. Then we average the changes across education groups for U.S.- and foreign-born separately, again weighting them by their wage shares as described in the formulas (12) and (13). Those values are reported in rows five and ten (those in bold fonts). Finally, we average the changes for the two groups (U.S.- and foreign-born workers), still using wage-share weights (as described in formula (14)) to obtain the overall wage change, reported in the last row (also in bold fonts). The upper part of Table 9 can be compared to the results obtained in the previous literature that mostly focus on the effect of immigration on wages of U.S.-born workers. The lower part of Table 9 reports the effects of immigration on wages of foreign-born, rarely considered in the previous literature. The table reports the "long-run" effects, namely the wage effects once capital has fully adjusted, $(\Delta\kappa_t/\kappa_t)_{immigration} = 0$. In section 6.2 below we focus on the effects as of year 2004 and on how long will take for full adjustment to set in. The four specifications (columns) of Table 9 are reported to better understand the differences with the traditional estimates, implied by our new findings of imperfect substitutability between U.S.- and foreign-born workers. Specification 4 calculates the effects under the traditional assumptions of perfect substitutability between U.S.- and foreign-born workers in each group $k, j$.Proceeding leftward, columns 1 to 3 introduce imperfect substitutability between U.S.- and foreign-born workers, where column 3 uses the highest estimate of $\sigma$ (equal to 10), Column 2 uses the median estimate $\sigma = 6.6$, and column 1 uses the lowest estimate $\sigma = 5$.

Let us begin focussing on the effect of immigration on the wages of natives (upper part of Table 9). The introduction of our novel feature (imperfect substitutability) has two important effects: first it modifies the effect of immigration on average wages of natives from null (0.1%) to positive (between 1.2% and 2.3%), second it reduces the adverse distributional effect of immigrants on wages of U.S.-born workers. Both effects are stronger the lower the substitutability between U.S. and foreign-born workers. Considering the median estimate of $\sigma = 6.6$, our estimates imply a positive long-run effect of immigration on wages of workers with at least an high school degree. In particular college graduates benefit from immigration (+0.7% in wages) while under perfect substitutability they were hurt by it (as shown by the -1.5% in column 4 ) and high school graduates benefit up to 3.5% point in their real wage growth. Considering native workers with no high school degree, their long-run real wage loss due to immigration was evaluated by Borjas and Katz (2005), table 11 at -4.8%[18]. Column 4 of Table 9 reproduces that negative result obtaining a 4.2% loss in real wage of high school dropouts when we impose perfect substitutability between U.S.- and foreign born workers. Our preferred estimates, however, shown in column 2 of Table 9, report only a small negative effect (-1.1%) on wages of native dropouts. Overall our results show a significant positive effect of immigration on average U.S. wages, and on each group of worker with at least an high school degree and only a very small negative effect on wages of workers without high school degree in the long run.

In our preferred specification 2 of Table 9 the group whose wages are mostly hurt from immigration are foreign born workers, i.e. previous immigrants (see lower part table 9). On average they lost 19% of their real

[18] The Borjas and Katz (2005) estimates refer to the effects of immigration between 1980 and 2000.

wages while some groups (i.e. college graduates) lost up to 24% of their wage. Recall that, due to the assumption of constant returns to scale in the aggregate production function, once capital fully adjusts to immigration the average overall wage (last row) does not change. Hence, our hypothesis of imperfect substitutability simply shifts the distributional effects of immigration by increasing the wage competition effect of immigrants on other foreign-born and decreasing it for U.S.-born workers. If the negative effect on wages of foreign-born may seem large this is due to the massive inflow of immigrants 1990-2004 relative to the initial size of the foreign-born employment. Immigrants in the labor force have more than doubled in the period 1990-2004; in particular, foreign-born workers have increased by 140% ($\Delta F_{1990-2004}/F_{1990} = 1.4$) during that period. Hence even with a wage elasticity of that group relative to the U.S.-born group equal to 0.10 (in column 3, $\sigma = 10$, hence the relative wage elasticity $\frac{1}{\sigma} = 0.1$) one obtains a relative wage change of around 14%, split, as we see in column 3, into an increase of native wages by 1.2% and a decrease in wages of foreign-born by 13.3%. Notice that if $\sigma=\infty$ the effects of immigration on wages are identical for U.S.- and foreign-born in the same education-experience group. The small differences reported in column 4 between the effects on U.S.- and foreign-born wages are due to the different composition in employment distribution by experience and education between the two groups.

Are the effects on wages of foreign-born workers reasonable? First of all, simply considering the relative U.S./foreign-born wages there are some skill groups that experienced large immigrant inflows and a substantial deterioration of their wage relative to natives. For instance among the high school dropout between 20 and 35 years of experience, until 1980 wages of U.S. and foreign born where almost identical, while in 2004 U.S.-born were earning 15-20% more than foreign born workers. A similar relative decline can also be observed for high school Graduates in the 20-30 year experience range. The worsening of wages of foreign-born relative to U.S.-born (see for instance Borjas 1999, page 27), which is usually attributed to worsening in the quality of immigrants, is interpretable in the light of our results as an effect of increased wage competition between foreign-born in those occupations that overwhelmingly employ immigrants. Moreover, the reason that we do not observe larger native-foreign wage differentials in all skill groups is probably that immigrants choose sectors/occupations/jobs with booming demand so that the systematic components of $\theta_{Fkjt}$ by year and skill (that we controlled for in equation 17) partly offset the negative effect of increased supply. Finally, another reason why the efficiency term $\theta_{Fkjt}$ may vary, in its systematic part, to offset the increase in supply of foreign-born, $\Delta F_{kj,1990-2004}$, has been proposed by Lewis (2005) and Card and Lewis (2005): Those sectors/job where immigrants skills (in terms of education and experience) are more abundant induce technological choices "biased" towards those skills and use them more effectively, which increases the relative efficiency $\theta_{Fkjt}$. The large negative effects on wages of other foreign-born are, therefore, in part offset by systematic improvements in relative efficiency.

Finally, let us provide an explanation of an apparent puzzle raised by our results. In light of our analysis, previous immigrants are the group whose wage suffers most from new immigrants. Why, then, are they consistently among the strong supporters of more open immigration policies (see e.g. Hatton and Williamson, 2005 and Mayda, 2006)? Obviously, while they may forego as much as 1% wage growth per year due to new immigrants they are also the group that gains most from a non-economic point of view. As immigration (legal and illegal) in the U.S. works mostly through family re-union, network connections and personal ties, new immigrants are likely to be spouses, siblings, friends and acquaintances of foreign-born residents in the U.S. and hence probably have a huge personal, affective and amenity value to them, well above the negative wage effect that we identified.

## 6.2 Reconsidering The Short Run Effects with yearly Capital Adjustment.

How long does it take for physical capital to adjust and restore its long-run returns? And in presence of sluggish adjustment of capital what are the effects of immigration on wages in the short run? As illustrated in section (3.3) accounting for capital adjustment simply adds a non-zero term, $(1-\alpha)\left(\frac{\Delta\kappa_t}{\kappa_t}\right)_{immigration}$ ,to the change in wage of each group. Hence the short-run wage response for each group and for the averages, will differ from the long run response by a common constant, due to the chosen Cobb-Douglas structure in which $\kappa_t$ only affects marginal productivity of workers through the overall average wage. A popular way, used in the growth and business cycle literature, to analyze the deviation of $\ln(\kappa_t)$ from its balanced growth path trend, is to represent its time-dynamics in the following way:

$$\ln(\kappa_t) = \beta_0 + \beta_1 \ln(\kappa_{t-1}) + \beta_2(trend) + \gamma\frac{\Delta F_t}{L_t} + \varepsilon_t \tag{23}$$

where the term $\beta_2(trend)$ captures the balanced growth path trajectory of $\ln(\kappa_t)$ equal to $\frac{1}{\alpha}\ln\left(\frac{1-\alpha}{r+\delta}A_t\right)$ and the term $\beta_1 \ln(\kappa_{t-1})$ captures the sluggishness of yearly adjustment to shocks. The parameter $(1-\beta_1)$ is commonly called "speed of adjustment" as it is the share of the deviation from the balanced growth path (trend) eliminated each year. Finally $\frac{\Delta F_t}{L_t}$ are the yearly immigration shocks and $\varepsilon_t$ are other shocks. Assuming that immigration shocks cause a proportional decrease in $\kappa_t$ for the same year ($\gamma = -1$) in order to calculate the effect of immigration on $\kappa_t$ over, say, the 1990-2004 period, one needs an estimate of the parameter $\beta_1$. Once we know $\beta_1$ and the sequence of yearly immigration flows, $\frac{\Delta F_t}{L_t}$, on can use (23) to obtain an impulse response of $\ln(\kappa_t)$ and its deviation from trend as of 2004 (short run) as well as for later years (long run). The previous migration literature has essentially assumed $\beta_1 = 0$ in the short run calculations cumulating the $\frac{\Delta F_t}{L_t}$ over one or two decades for fixed capital (implying a very large deviation from the trend). On the other hand it has assumed $\beta_1 = 1$, (full adjustment) in the long-run calculations. The recent empirical growth literature (Islam 1995, Caselli et al. 1996) and the recent business cycle literature (Romer, 2006, chapter 4), to the contrary, provide model-based and empirical estimates of $\beta_1$ . The recent growth literature estimates usually a 10% speed of convergence of capital to the own balanced growth path for advanced (OECD) economies (Islam, 1995, Caselli et al. 1996), implying $\beta_1 = 0.9$. Similarly the business cycle literature calculates speed of convergence of capital around 10% (Romer, 2006, Chapter 4) for closed economies and faster for open economies. Hence $\beta_1 = 0.9$ seems a reasonable estimate. We estimated ourselves a simple AR(1) process with trend for $\ln(\kappa_t)$.We constructed the variable $\kappa_t = (K_t/L_t)$ dividing the stock of U.S. capital at constant prices (Net Stock of Private and Government Fixed Assets from the Bureau of Economic Analysis, 2006) by the total non-farm employment from the Bureau of Labor Statistics (2006) for each year during the period 1960-2004. We estimated several specifications including changes in total employment as shock$\left(\frac{\Delta L_t}{L_t}\right)$, or immigrants only $\left(\frac{\Delta F_t}{L_t}\right)$[19] as shock and instrumented those with changes in population (to correct for endogeneity of employment). All estimates of $\beta_1$ ranged between 0.8 and 0.9 (speed of adjustment of 10 to 20% a year) with standard errors ranging between 0.02 and 0.08. We could never reject $\beta_1 = 0.9$, and we could always reject $\beta_1 = 1$ (no adjustment). Hence we consider 10% a reliable and, if anything, conservative estimate of the yearly speed of capital adjustment. Using the series of immigration rates 1990-2004 and the estimated parameters of capital adjustment $\beta_1 = 0.9, \gamma = -0.9$ (assuming that capital adjustment begins the same year as immigrants are received) the recursive equation (23) allows us to calculate $(\Delta\kappa_{1990-2004}/\kappa_{1990})_{immigration}$ as of year 2004 and what share of the deviation from trend remains in 2009. Using the formula (7) we can calculate the effect of $\Delta\kappa$ on the average wage and on each group's wage. Recall that assuming no adjustment of capital in the short run ( $\beta_1 = 1, \gamma = -1$) as the cumulated inflow of immigrants during the 1990-2004 period was 11% of the employment in 1990, implied an effect of immigration on average real wages equal to $(0.33) * (-11\%) = -3.6\%$, as of 2004. Using the actual 10% speed of adjustment of capital each year, however, we obtain only a $-3.4\%$ effect of immigration on capital-labor ratio corresponding to a mere -1.1% ($= 0.33 * 3.4\%$) effect on real wages as of 2004, and in five more years (2009) the negative effect on wages is reduced to 0.6%. Table 10 uses these adjustments of capital-labor ratio and shows the effects of immigration on wages as of year 2004 (column 1) and as of year 2009 (column 2). Those columns use the same parameter values as column 2 of Table 9, i.e. the median and preferred estimates of $\sigma$. We also report in column 3 the long-run effects for full capital adjustment (identical to column 2 of Table 9) and, for comparison, the "short-run" effects calculated assuming fixed capital (as in the previous literature) in column 4. Finally, the short-run effects with fixed capital and perfect substitutability between U.S.- and foreign-born workers are shown in column 5. Hence column 1 reports the newly calculated "short-run" effects of immigration while column 5 reports those calculated using the methods prevailing in the previous literature. The differences are remarkable. Average wage of U.S. born workers increased by 0.7% in our estimates as of 2004 rather than experiencing a decrease by 3.5%. US workers with no degree experience a loss of 2.2% of their real wage rather than a loss of almost 8%. College educated US born workers have essentially no change in their wage (rather

[19] We constructed $\Delta F_t$, for each year 1960-2004, using the following procedure. From the U.S. Department of Justice, Immigration and Naturalization Service, (2004) we obtain the number of (legal) immigrants for each fiscal year 1960-2004. We then distribute the net change of foreign-born workers in each decade (measured from census data and from the American Community Survey, which includes illegal immigrants as well as legal ones) over each year in proportion to the gross yearly flows of legal immigrants.

than a 5% loss) and the groups of high school graduates and college dropouts experience already in the short run significant gains rather than significant losses in their real wages. The benefits of immigration are realized for most workers already in the short run and certainly almost of all those benefits are enjoyed by 2009, with an average wage gain of more than 1% distributed as gains for the three groups with at least an high school degree and a small loss for high school dropouts. The wage losses, in the short run as well as in the long run, are concentrated among previous immigrants who experience most of the competition from new immigrants and forego sizeable wage increases as a consequence of that.

### 6.3 Robustness Checks

Table 11 shows the changes in the calculated long-run effects when we use different values for the parameters $\delta$ and $\eta$ in the simulations. While the values used in Table 10, equal to 2 and 4 respectively, seem to be right in the middle of the estimated range for these parameters (both in our estimates and in previous ones) some articles have estimated values of $\delta$ as low as 1.5 and as high as 2.5, while the range for $\eta$ is between 3 and 5. We reproduce simulations from Column 1-3 of Table 10 using, respectively, the low estimates of $\delta$ and $\eta$ (columns 1-3 in Table 11) and the high estimates of $\delta$ and $\eta$ (columns 4-6 in Table 11). While the average effects on wages of native and foreign-born workers are not sensitive to changes in those parameters, the distributional effects between education groups become stronger when we use lower estimates of $\delta$ and $\eta$. Considering columns 2 and 5 as the references, as they use the median estimate of $\sigma$, we see that the wage loss of U.S.-born high school dropout can be as large as -2.5% when $\delta$ =1.5. Still, this number is much smaller than the previous estimates. On the other hand, if we use the higher elasticity of substitution estimates ($\delta = 2.5$ and $\eta = 5$), unskilled natives barely suffer a loss of wage (-0.3%) from immigration. A similar widening of the distributional effects takes place, using the lower estimates, for wages of foreign-born workers across education group. Similar consequences of widening the distributional effects would be observed if we lowered $\delta$ and $\eta$ in the simulation with $\sigma = \infty$.

### 6.4 Contribution of Immigration to Average Wage and Wage Dispersion of U.S.-born workers

The differences in the real wage effects of immigration on natives shown in Table 10 between specification 1 (our preferred one) and specification 5 (representative of previously estimated short-run effects) are substantial. In order to put them in perspective, it is instructive to compare them with the actual changes in average wages of U.S.-born workers during the 1990-2004 period and with changes in the measures of their wage dispersion during the same period. Specification 1 in Table 10 implies an effect on average U.S. real wages 4.2% points larger than the usually estimated short-run effects reported in column 5 of Table 10 (+0.7% vs. -3.5%). This is a very large difference even when compared to the average growth rate of wages of U.S.-born workers in the period, which equals 12.5%, and is even more notable if compared to the typical changes of real wages over the business cycle (amounting, on average, to less than 0.5%). Roughly 60% of the difference between specifications is due to the hypothesis of yearly capital adjustment while about 40% is due to the imperfect substitutability between U.S. and foreign-born workers.

Even more interestingly, as immigration has been connected to increased wage dispersion (e.g. Freeman, Borjas and Katz 1997 and several others), we can compare what fraction of that increase could be due to immigration. There are several ways of measuring wage dispersion across educational groups, depending on which group we focus on. Columns 1 and 2 of Table 12 provide some standard measures of increased wage dispersion across educational groups during the period 1990-2004. In particular, Column 2 reports, in the first four rows, the percentage variation of the real wage for each of the four groups relative to the average real increase of wages[20] and, in the last 2 rows, the table shows the real increase in the college/high school dropouts wage premium and in the college/high school wage premium. All numbers are calculated for U.S.-born workers only. Column 1 reports the actual percentage changes for each real wage group (not net of the average) showing that high school dropouts experienced actually real wage loss in the period. Notice first of all that wage dispersion increased between any two groups since lower wage groups (lower education groups) had lower growth rates of wages. Particularly bad has been the performance of U.S.-born high school dropouts whose wages dropped by 24.4% relative to the average during the period. Also sub-average (but much less so) were the performances of wages of high school graduates (6.1% lower than average) and college dropouts (4.1% lower than average). On the other hand, wages of college graduates substantially out-performed the average (8.9% better). As a consequence, the wage premium (in ratio) between college graduates and high school dropouts increased by 33% during the period and the college/high school wage premium increased by 15%. These statistics are calculated using Census and American Community Survey IPUMS data on wage of all U.S.-born workers as defined in section 4. Column 3 shows the percentage changes in real wages attributed to immigration by our model, (Specification 2 of Table 9) and column 4 shows what share they represent of the actual 1990-2004 change. Looking at the first four rows immigration actually decreased wage dispersion for three groups (HSG, COD and COG), in that it helped the first two groups, that performed worse than average and hurt the last one that performed better. This is noted in Table 12 by the caption "attenuate dispersion" under the corresponding figures. As for native high school dropouts, immigration contributed to wage dispersion but it explains less than one eighth (0.12) of the difference in the performance of this group's wage with the average wage. Moreover, immigration does not contribute at all to explain the increased college-high school wage premium; if anything,

[20] The average increase is calculated by weighting the percentage wage increases of each group by the average wage share of that group in the 1990-2004 period. It is different from the change in the average wage which also includes the effect of changes in educational shares.

immigration caused a reduction in that premium as the last row of column 4 shows, and immigration only explains one twentieth (0.05) of the increase in the college-high school dropout premium (second to the last row of column 4). These numbers seem to show that immigration cannot be considered as an important candidate in explaining increased wage dispersion. Even giving immigration the best shot at causing wage dispersion by adopting the old assumption of $\sigma = \infty$[21] (column 5 and 6) one still obtains that immigration *attenuated* wage dispersion for three groups (helping those underperforming the mean and hurting those outperforming it) while it only contributed to the under-performance of high school dropout wages. However, even in this scenario, only one sixth (0.17) of their growth differential with the average wage and less than one tenth (0.087) of the increase in college-high school dropout premium can be attributed to immigration.

# 7 Conclusions

The main message of this paper is that only within a model that specifies the interactions between workers of different skills and between labor and physical capital (in a production function) can we derive marginal productivity, labor demands and analyze the effects of immigration on the wages of different types of workers. The existing literature on immigration has paid much attention to the estimates of the partial effect of immigrants on wages of U.S.-born workers with similar skills. Those estimates are *partial* in that they assume a constant supply of all other groups and of physical capital and therefore are not informative of the actual overall effects of immigration on wages. In taking the general equilibrium approach, instead, one realizes that the substitutability between U.S. and foreign-born workers with similar schooling and experience as well as the investment response to changes in the supply of skills are important parameters to evaluate the short and long run effects of immigration on wages. We therefore carefully tackle the tasks of estimating the elasticity of substitution between U.S.- and foreign-born workers and we account for physical capital adjustment in the short and long run. We find robust evidence that U.S.- and foreign-born workers are not perfect substitutes within an education experience group. This fact and the yearly adjustment of capital to immigration imply that average wages of natives benefit from immigration, even in the short-run. These average gains are, in the short and long run, distributed as a small loss to the group of high school dropouts and wage gains for all the other groups of U.S. natives. The group suffering the biggest loss in wage, rather than natives, is the contingent of previous immigrants, who compete for much more similar jobs and occupations with the new immigrants. Finally, our model implies that it is very hard to claim that immigration has been a significant determinant in the deterioration of wage distribution during the 1990's and 2000's.

[21] Assumptions on capital adjustment do not have any impact on relative wages but only on the average wage. Hence the relative changes in specifications 3 and 5 of Table 12 could be either for fixed or for fully adjusted capital.

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# A Appendix: Partial Effects of Immigration on Wages

The total effect of immigrants on the wages of natives in group $k, j$, as calculated in (10), is the combination of three types of effects. The first is the impact of foreign workers in the same education and experience group $k, j$ on the wages of natives in the same group. This effect is obtained by differentiating (8) with respect to $\ln(F_{kj})$ and expressing the results in terms of the percentage changes in the wage of group $k, j$ ($\Delta \ln w_{Hkj} = \Delta w_{Hkj}/w_{Hkj}$) that results from a percentage change of foreign-born in the same group ($\Delta F_{kj}/F_{kj} = \Delta \ln(F_{kj})$):

$$\frac{\Delta w_{Hkjt}}{w_{Hkjt}} = \left[\frac{1}{\delta} + \left(\frac{1}{\eta} - \frac{1}{\delta}\right)\left(\frac{1}{s_{kt}}\right) + \left(\frac{1}{\sigma_k} - \frac{1}{\eta}\right)\left(\frac{1}{s_{kjt}}\right)\right] s_{Fkjt} \frac{\Delta F_{kjt}}{F_{kjt}} \tag{24}$$

The second type of effect is the impact of foreign-born workers in a different experience group $i \neq j$ and the same education group $k$. Differentiating (8) with respect to $\ln(F_{ki})$ we obtain:

$$\frac{\Delta w_{Hkjt}}{w_{Hkjt}} = \left[\frac{1}{\delta} + \left(\frac{1}{\eta} - \frac{1}{\delta}\right)\left(\frac{1}{s_{kt}}\right)\right] s_{Fkjt} \frac{\Delta F_{kit}}{F_{kit}} \tag{25}$$

The third and last effects is the impact of foreign-born workers in a different education group $m \neq k$, whatever their experience group. Differentiating (8) with respect to $\ln(F_{mit})$ we obtain:

$$\frac{\Delta w_{Hkjt}}{w_{Hkjt}} = \frac{1}{\delta} s_{Fkjt} \frac{\Delta F_{mit}}{F_{mit}} \tag{26}$$

One can then easily combine the above effects to obtain the expression (10) reported in the text.

# Tables and Figures

**Table 1:**
**Share of Foreign Born Workers by Education and Experience**

| Group | | Year | | | | | |
|---|---|---|---|---|---|---|---|
| **Education** | **Experience** | **1960** | **1970** | **1980** | **1990** | **2000** | **2004** |
| **High School Dropouts** | 0 to 4 | 0.039 | 0.036 | 0.058 | 0.101 | 0.119 | 0.116 |
| | 5 to 9 | 0.061 | 0.060 | 0.138 | 0.264 | 0.375 | 0.354 |
| | 10 to 14 | 0.058 | 0.066 | 0.166 | 0.252 | 0.426 | 0.472 |
| | 15 to 19 | 0.056 | 0.072 | 0.143 | 0.262 | 0.416 | 0.493 |
| | 20 to 24 | 0.058 | 0.070 | 0.132 | 0.270 | 0.364 | 0.442 |
| | 25 to 29 | 0.055 | 0.061 | 0.124 | 0.222 | 0.363 | 0.377 |
| | 30 to 34 | 0.083 | 0.061 | 0.101 | 0.179 | 0.358 | 0.382 |
| | 34 to 40 | 0.122 | 0.058 | 0.086 | 0.161 | 0.281 | 0.345 |
| | **All Experiences** | **0.07** | **0.060** | **0.109** | **0.205** | **0.306** | **0.341** |
| **High School Graduates** | 0 to 4 | 0.019 | 0.025 | 0.032 | 0.057 | 0.095 | 0.107 |
| | 5 to 9 | 0.025 | 0.028 | 0.038 | 0.062 | 0.125 | 0.148 |
| | 10 to 14 | 0.028 | 0.033 | 0.046 | 0.057 | 0.118 | 0.167 |
| | 15 to 19 | 0.036 | 0.035 | 0.047 | 0.057 | 0.100 | 0.157 |
| | 20 to 24 | 0.035 | 0.037 | 0.051 | 0.062 | 0.085 | 0.119 |
| | 25 to 29 | 0.046 | 0.041 | 0.050 | 0.059 | 0.082 | 0.105 |
| | 30 to 34 | 0.065 | 0.040 | 0.051 | 0.060 | 0.081 | 0.097 |
| | 34 to 40 | 0.108 | 0.049 | 0.054 | 0.055 | 0.072 | 0.097 |
| | **All Experiences** | **0.038** | **0.034** | **0.044** | **0.059** | **0.095** | **0.124** |
| **College Dropouts** | 0 to 4 | 0.030 | 0.031 | 0.046 | 0.062 | 0.084 | 0.081 |
| | 5 to 9 | 0.042 | 0.047 | 0.051 | 0.071 | 0.097 | 0.104 |
| | 10 to 14 | 0.048 | 0.054 | 0.058 | 0.066 | 0.103 | 0.110 |
| | 15 to 19 | 0.055 | 0.058 | 0.065 | 0.063 | 0.095 | 0.117 |
| | 20 to 24 | 0.048 | 0.054 | 0.070 | 0.065 | 0.084 | 0.101 |
| | 25 to 29 | 0.052 | 0.058 | 0.065 | 0.070 | 0.076 | 0.087 |
| | 30 to 34 | 0.076 | 0.046 | 0.067 | 0.074 | 0.074 | 0.077 |
| | 34 to 40 | 0.099 | 0.057 | 0.067 | 0.072 | 0.077 | 0.076 |
| | **All Experiences** | **0.052** | **0.048** | **0.057** | **0.067** | **0.088** | **0.095** |
| **College Graduates** | 0 to 4 | 0.035 | 0.035 | 0.042 | 0.070 | 0.121 | 0.114 |
| | 5 to 9 | 0.045 | 0.064 | 0.062 | 0.090 | 0.143 | 0.173 |
| | 10 to 14 | 0.053 | 0.069 | 0.080 | 0.094 | 0.153 | 0.178 |
| | 15 to 19 | 0.056 | 0.060 | 0.097 | 0.087 | 0.138 | 0.160 |
| | 20 to 24 | 0.052 | 0.053 | 0.088 | 0.093 | 0.120 | 0.149 |
| | 25 to 29 | 0.064 | 0.058 | 0.073 | 0.107 | 0.105 | 0.126 |
| | 30 to 34 | 0.071 | 0.056 | 0.072 | 0.095 | 0.105 | 0.104 |
| | 34 to 40 | 0.088 | 0.070 | 0.072 | 0.088 | 0.125 | 0.122 |
| | **All Experiences** | **0.054** | **0.056** | **0.070** | **0.089** | **0.128** | **0.146** |

**Note:** Individuals included in calculations are those between 17 and 65 years, not living in group quarters who received non-zero income and worked at least one week during the previous year. Foreign-born are workers born outside of the US and not citizen at birth. Sources: Authors' Calculations on individual data from Census IPUMS and ACS from Ruggles et al (2006).

**Table 2:**
**Weekly Wages of U.S. Natives in constant 2000 U.S. $ by Education and Experience**

| Group | | Year | | | | | |
|---|---|---|---|---|---|---|---|
| **Education** | **Experience** | **1960** | **1970** | **1980** | **1990** | **2000** | **2004** |
| **High School Dropouts** | 0 to 4 | 207 | 246 | 207 | 180 | 214 | 179 |
| | 5 to 9 | 324 | 384 | 370 | 358 | 406 | 357 |
| | 10 to 14 | 377 | 442 | 415 | 423 | 480 | 446 |
| | 15 to 19 | 403 | 452 | 444 | 455 | 507 | 489 |
| | 20 to 24 | 401 | 468 | 471 | 476 | 554 | 548 |
| | 25 to 29 | 403 | 486 | 482 | 500 | 579 | 549 |
| | 30 to 34 | 399 | 470 | 494 | 522 | 594 | 599 |
| | 34 to 40 | 402 | 463 | 500 | 521 | 608 | 585 |
| | **All Experiences** | **374** | **431** | **495** | **402** | **424** | **400** |
| **High School Graduates** | 0 to 4 | 307 | 355 | 334 | 313 | 350 | 325 |
| | 5 to 9 | 404 | 476 | 439 | 437 | 485 | 457 |
| | 10 to 14 | 454 | 526 | 487 | 504 | 553 | 567 |
| | 15 to 19 | 472 | 538 | 526 | 537 | 606 | 631 |
| | 20 to 24 | 484 | 546 | 544 | 559 | 642 | 645 |
| | 25 to 29 | 486 | 551 | 552 | 596 | 658 | 658 |
| | 30 to 34 | 485 | 565 | 560 | 606 | 665 | 681 |
| | 34 to 40 | 476 | 556 | 558 | 587 | 681 | 673 |
| | **All Experiences** | **463** | **501** | **478** | **507** | **579** | **576** |
| **College Dropouts** | 0 to 4 | 354 | 402 | 365 | 359 | 388 | 374 |
| | 5 to 9 | 473 | 565 | 495 | 522 | 571 | 570 |
| | 10 to 14 | 543 | 639 | 573 | 604 | 665 | 686 |
| | 15 to 19 | 574 | 672 | 631 | 656 | 731 | 755 |
| | 20 to 24 | 583 | 694 | 649 | 708 | 775 | 794 |
| | 25 to 29 | 573 | 706 | 660 | 749 | 805 | 846 |
| | 30 to 34 | 567 | 715 | 670 | 757 | 836 | 820 |
| | 34 to 40 | 572 | 669 | 666 | 731 | 855 | 832 |
| | **All Experiences** | **516** | **593** | **537** | **600** | **685** | **693** |
| **College Graduates** | 0 to 4 | 469 | 573 | 477 | 569 | 658 | 645 |
| | 5 to 9 | 611 | 763 | 639 | 786 | 904 | 976 |
| | 10 to 14 | 728 | 908 | 798 | 932 | 1155 | 1230 |
| | 15 to 19 | 779 | 983 | 906 | 1024 | 1287 | 1349 |
| | 20 to 24 | 776 | 1036 | 964 | 1147 | 1318 | 1357 |
| | 25 to 29 | 779 | 1038 | 992 | 1203 | 1340 | 1365 |
| | 30 to 34 | 789 | 996 | 997 | 1213 | 1430 | 1347 |
| | 34 to 40 | 782 | 950 | 953 | 1178 | 1413 | 1336 |
| | **All Experiences** | **693** | **863** | **761** | **950** | **1170** | **1201** |

**Note:** Individuals included in calculations are those between 17 and 65 years, not living in group quarters, which received non-zero income and worked at least one week during the previous year. Wages are in real US Dollars calculated using the CPI deflator with 2000 as base year. Natives are workers born within the US or U.S. citizen at birth. Sources: Authors' Calculations on individual data from Census IPUMS and ACS from Ruggles et al (2006).

## Table 3
## Index of Congruence in the choice of Occupations, Census 2004

| Index of Congruence | Foreign, Same Education | Natives HSD | Natives HSG | Natives COD | Natives COG |
|---|---|---|---|---|---|
| **Natives, HSD** | 0.65 | 1 | | | |
| **Natives, HSG** | 0.68 | 0.68 | 1 | | |
| **Natives, COD** | 0.61 | -0.25 | 0.22 | 1 | |
| **Natives, COG** | 0.70 | -0.73 | -0.91 | 0.35 | 1 |

**Note:** The Index of Congruence between the two groups (row and column headers) is calculated as the centered correlation coefficient (between -1 and +1) using 180 different occupations and data from the 2004 American Community Survey in Ruggles et al. (2006).

## Table 4
## Increasing specialization of foreign-born

| Dependent Variable: Growth in share of foreign-born in the occupation | | | |
|---|---|---|---|
| **Period:** | | 1980-1990 | 1990-2004 |
| **Explanatory variable: Initial share of foreign-born in the occupation** | **All education groups, pooled** | 0.69**<br>(0.09) | 0.64**<br>(0.05) |
| | **High School Dropouts** | 0.65**<br>(0.10) | 0.50**<br>(0.09) |
| | **High School Graduates** | 0.31**<br>(0.07) | 0.96**<br>(0.13) |
| | **College Dropouts** | 0.88**<br>(0.05) | 0.30**<br>(0.08) |
| | **College Graduates** | 0.88**<br>(0.09) | 0.35**<br>(0.06) |

**Note:** Each cell reports the coefficient from a separate regression. The growth in the share of foreign-born employment in each occupation is regressed on its initial value.

**Table 5**
**Relative U.S.-Foreign-Born Wage Elasticity within Education-Experience Cells**

| | All Workers, | | Male only | | Not weighted | | Omitting 1960, 2004 | |
|---|---|---|---|---|---|---|---|---|
| Specification | **1** | **2** | **3** | **4** | **5** | **6** | **7** | **8** |
| Dependent variable | **Yearly Wages** | **Weekly Wages** | **Yearly Wages** | **Weekly Wages** | **Yearly Wages** | **Weekly Wages** | **Yearly Wages** | **Weekly Wages** |
| $1/\overline{\sigma}$ | 0.17** | 0.11** | 0.18** | 0.10** | 0.15** | 0.13** | 0.20** | 0.11** |
| | (0.04) | (0.04) | (0.06) | (0.04) | (0.05) | (0.03) | (0.05) | (0.04) |
| $1/\sigma_{HSD}$ | 0.19** | 0.11** | 0.18** | 0.08** | 0.15** | 0.10** | 0.27** | 0.10** |
| | (0.06) | (0.04) | (0.07) | (0.04) | (0.07) | (0.04) | (0.06) | (0.04) |
| $1/\sigma_{HSG}$ | 0.17** | 0.14** | 0.17** | 0.09** | 0.14** | 0.12** | 0.15** | 0.08** |
| | (0.04) | (0.05) | (0.05) | (0.05) | (0.04) | (0.04) | (0.05) | (0.04) |
| $1/\sigma_{COD}$ | 0.19** | 0.12** | 0.23** | 0.16** | 0.19** | 0.15** | 0.18** | 0.13** |
| | (0.05) | (0.05) | (0.06) | (0.07) | (0.07) | (0.06) | (0.07) | (0.06) |
| $1/\sigma_{COG}$ | 0.10** | 0.08** | 0.12** | 0.09 | 0.19** | 0.13** | 0.10** | 0.08** |
| | (0.04) | (0.04) | (0.06) | (0.06) | (0.07) | (0.06) | (0.04) | (0.04) |
| Observations | 192 | 192 | 192 | 192 | 192 | 192 | 128 | 128 |
| Test F, All $\sigma$ are equal | 0.98 | 1.47 | 0.89 | 1.88 | 0.61 | 0.48 | 5.4 | 1.36 |
| (p-value) | (41%) | (24%) | (45%) | (15%) | (67%) | (69%) | (1%) | (27%) |

**Note:** All Regressions include education-by-experience fixed effects, education-by-year fixed effects and experience-by-year fixed effects. Errors are heteroskedasticity robust and clustered by education-experience. Dependent Variable is natural logarithm of relative wage of US and foreign born workers in the same education and experience group, the explanatory variable is the relative employment of US and foreign-born workers in the same education experience group. Observations are weighted by total employment in the cell, in all specifications except for 5 and 6.
In Specifications 1 and 2 the wages are calculated including all workers in the cells. In specifications 3 and 4 the wages are calculated including only Males in each cell. Specification 7 and 8 omit years 1960 and 2004. The F-statistic in the last row tests the hypothesis that all the elasticities of relative demand between US and foreign born worker are identical across educational group.

**Table 6**
**Relative U.S.-Foreign-Born Wage Elasticity within Education-Experience Cells**
**Foreign-born are grouped by Effective Experience**

| | **All Workers,** | | **Male only** | | **Not weighted** | | **Omitting 2004** | |
|---|---|---|---|---|---|---|---|---|
| Specification | **1** | **2** | **3** | **4** | **5** | **6** | **7** | **8** |
| Dependent variable | **Yearly Wages** | **Weekly Wages** | **Yearly Wages** | **Weekly Wages** | **Yearly Wages** | **Weekly Wages** | **Yearly Wages** | **Weekly Wages** |
| $1/\overline{\sigma}$ | 0.17** | 0.12** | 0.18** | 0.10** | 0.16** | 0.13** | 0.16** | 0.10* |
| | (0.05) | (0.04) | (0.06) | (0.05) | (0.05) | (0.05) | (0.06) | (0.04) |
| $1/\sigma_{HSD}$ | 0.21** | 0.14** | 0.24** | 0.11 | 0.19** | 0.16** | 0.17** | 0.10** |
| | (0.06) | (0.07) | (0.08) | (0.07) | (0.06) | (0.07) | (0.05) | (0.05) |
| $1/\sigma_{HSG}$ | 0.18** | 0.16** | 0.17** | 0.09 | 0.17** | 0.17** | 0.19** | 0.13** |
| | (0.05) | (0.07) | (0.07) | (0.05) | (0.05) | (0.05) | (0.07) | (0.06) |
| $1/\sigma_{COD}$ | 0.20** | 0.12** | 0.20** | 0.13** | 0.18** | 0.13** | 0.24** | 0.12** |
| | (0.08) | (0.06) | (0.08) | (0.05) | (0.09) | (0.06) | (0.09) | (0.05) |
| $1/\sigma_{COG}$ | 0.10** | 0.08** | 0.10** | 0.07* | 0.11** | 0.08** | 0.11** | 0.07* |
| | (0.05) | (0.04) | (0.05) | (0.04) | (0.05) | (0.04) | (0.05) | (0.04) |
| Observations | 160 | 160 | 160 | 160 | 160 | 160 | 128 | 128 |
| Test F, All σ are equal | 4.12 | 1.50 | 2.6 | 1.40 | 2.4 | 2.03 | 2.3 | 1.30 |
| (p-value) | (2%) | (22%) | (6%) | (26%) | (8%) | (13%) | (9%) | (32%) |

**Note:** Effective experience for foreign-born workers is calculated by weighting each year of their experience abroad by 0.4 and each year of experience in the US by 1.4. These "conversion factors" are estimated in Borjas (2003) using individual wages data. Period is 1970-2004 as in 1960 there is no information on the year when immigrants entered the country. All Regressions include education by experience fixed effects, education by year fixed effects and experience by year fixed effects. Errors are heteroskedasticity robust and clustered by education experience group. Dependent Variable is natural logarithm of relative wage of US and foreign born workers in the same education and experience group, the explanatory variable is the relative employment of US and foreign-born workers in the same education experience group. Observations are weighted by total employment in the cell, in all specifications except for 5 and 6.

In Specifications 1 and 2 the wages are calculated including all workers in the cells. In specifications 3 and 4 the wages are calculated including only Males in each cell. Specification 7 and 8 omit year 2004. The F-statistic in the last row tests the hypothesis that all the elasticities of relative demand between US and foreign born worker are identical across educational group.

**Table 7**
**Estimates of $1/\eta$ : Relative Wage Elasticity across Experience Cells**

| | CES Foreign-U.S.- born Using Estimated σ | | Simple Sum Foreign- U.S.-born (imposing σ=∞) | |
|---|---|---|---|---|
| **Specification** | **1** | **2** | **3** | **4** |
| **Skill groups** | **Education-Experience cells** | **Education- Effective Experience cells** | **Education-Experience cells** | **Education- Effective Experience cells** |
| Yearly Wages | 0.30**<br>(0.10) | 0.21**<br>(0.09) | 0.26**<br>(0.07) | 0.20**<br>(0.07) |
| Weekly Wages | 0.29**<br>(0.10) | 0.19**<br>(0.08) | 0.25**<br>(0.07) | 0.16**<br>(0.06) |
| Observations | 192 | 160 | 192 | 160 |

**Note**: Method of estimation is 2SLS using the log of foreign-born employed in the education experience group as instrument for the variable $\ln(L_{kjt})$ that is constructed as described in the text. All regressions include education by experience fixed effects and education by year fixed effects. Specifications 2 and 4 use effective experience groups, calculated as described in section 5.2 of the main text and include only years 1970, 1980, 1990, 2000 and 2004. Specification 1 and 3 use the regular experience groups and include data for 1960 as well. In Parenthesis we report the Heteroskedasticity Robust Standard error clustered by education group.

**Table 8**
**Estimates of $1/\delta$ : Relative Wage Elasticity across Education Cells**

| $1/\delta$ | **CES across experience groups, estimated η** | | **Simple Sum across experience groups ( imposing η=∞ )** | |
|---|---|---|---|---|
| **Specification** | **1** | **2** | **3** | **4** |
| **Skill Groups** | **By Education-Experience** | **By Education-Effective Experience** | **By Education-Experience** | **By Education-Effective Experience** |
| Yearly wages | 0.48**<br>(0.14) | 0.42**<br>(0.13) | 0.54**<br>(0.19) | 0.40**<br>(0.16) |
| Weekly Wages | 0.49**<br>(0.15) | 0.43**<br>(0.12) | 0.54**<br>(0.18) | 0.38**<br>(0.16) |
| Observations | 24 | 20 | 24 | 20 |

**Note:** Method of estimation is 2SLS using the log of foreign-born employed in the education group as instrument for the variable $\ln(L_{kt})$ that is constructed as described in the text. All regressions include 5 time fixed effects and 4 education-specific time trends. Specifications 2 and 4 use effective experience groups calculated as described in section 5.2 of the main text and include only years 1970, 1980, 1990, 2000 and 2004. Specification 1 and 3 use the regular experience groups and include data for 1960 as well. In parenthesis we report the Heteroskedasticity Robust Standard error clustered by education group.

**Table 9**
**Calculated Percentage Changes in Real Wages due to Immigrants Inflows: 1990-2004.**
**Long-run effects ($\Delta\kappa/\kappa$=0)**

| **Specification**<br>**Estimates of σ** | **1**<br>**Low**<br>σ=5 | **2**<br>**Median**<br>σ=6.6 | **3**<br>**High**<br>σ=10 | **4**<br>$\sigma$**, imposed =** ∞ |
|---|---|---|---|---|
| **% Real Wage Change of Us Born Workers due to immigration, 1990-2004** | | | | |
| HS dropouts US-born | -0.2% | -1.1% | -2.1% | -4.2% |
| HS graduates, US-born | +2.9% | +2.4% | +2.0% | +1.0% |
| CO dropouts, US-born | +3.7% | +3.4 | +3.1% | +2.4% |
| CO graduates, US-born | +1.4% | +0.7% | 0.0% | -1.5% |
| **Average, US-born** | **+2.3%** | **+1.8%** | **+1.2%** | **+0.1%** |
| **% Real Wage Change of Foreign Born Workers due to immigration, 1990-2004** | | | | |
| HS dropouts Foreign-born | -20.2% | -16.3% | -12.3% | -4.4% |
| HS graduates, Foreign-born | -31.7% | -23.5% | -15% | +1.0% |
| CO dropouts, Foreign-born | -17.4% | -12.3% | -7.3% | +2.4% |
| CO graduates, Foreign-born | -31.6% | -24.2% | -16% | -1.6% |
| **Average Foreign-born** | **-26.3%** | **-19.8%** | **-13.3%** | **-0.9%** |
| **Overall Average:**<br>**Native and US Born** | **0%** | **0%** | **0%** | **0%** |

**Note:** Values of the other parameters used in the estimation: δ=2, η=4, α=0.66. The inflow of immigrants in the period 1990-2004 as percentage of initial employment in the group were as follows: High School Dropouts: 20%, High School Graduates: 9.9%, College Dropouts: 6.5%, College Graduates: 14.1%, Overall 11.0%. The percentage change for the wage of each worker in group k, j is calculated using the formula (11) for US born and (12) for foreign-born. Then percentage wage changes are averaged across experience groups using the wage-share of the group in 1990 to obtain the Table entries. The averages for US and Foreign-born are obtained averaging the wage change of each education group weighted by its share in wage (as described in formulas 13 and 14). The overall average wage change adds the change of US and foreign-born weighted for the relative wage shares in 1990 (equal to 8.5% for foreign-born and 91.5% for US born).

**Table 10**
**Calculated Percentage Changes in Real Wages due to Immigrants Inflows: 1990-2004.**
**Short-run Effects, accounting for yearly capital adjustment.**

| **Specification Estimates of σ** | **1 As of 2004 (short run)** | **2 As of 2009** | **3 Long-run** | **4 Fixed K (traditional short-run)** | **5 Fixed K and $\sigma$ , imposed = ∞** |
|---|---|---|---|---|---|
| **% Real Wage Change of Us Born Workers due to immigration, 1990-2004** | | | | | |
| HS dropouts US-born | -2.2% | -1.7% | -1.1% | -4.8% | -7.9% |
| HS graduates, US-born | +1.3% | +1.8% | +2.4% | -1.2% | -2.6% |
| CO dropouts, US-born | +2.3% | +2.8% | +3.4 | -0.2% | -1.2% |
| CO graduates, US-born | -0.4% | +0.1% | +0.7% | -2.9% | -5.2% |
| **Average, US-born** | **+0.7%** | **+1.2%** | **+1.8%** | **-1.9%** | **-3.5%** |
| **% Real Wage Change of Foreign Born Workers due to immigration, 1990-2004** | | | | | |
| HS dropouts Foreign-born | -17.4% | -16.9% | -16.3% | -19.9% | -8.1% |
| HS graduates, Foreign-born | -24.6% | -24.1% | -23.5% | -27.1% | -2.6% |
| CO dropouts, Foreign-born | -13.4% | -12.9% | -12.3% | -15.9% | -1.2% |
| CO graduates, Foreign-born | -25.3% | -24.8% | -24.2% | -27.8% | -5.3% |
| **Average Foreign-born** | **-20.9%** | **-20.4%** | **-19.8%** | **-23.4%** | **-4.7%** |
| **Overall Average: Native and US Born** | **-1.1%** | **-0.6%** | **0%** | **-3.6%** | **-3.6%** |

**Note:** Values of the other parameters used in the estimation of column 1, 2, 3 and 4: σ=6.6, δ=2, η=4, α=0.66. Column 5, assumes : σ=∞, δ=2, η=4, α=0.66. The inflow of immigrants in the period 1990-2004 as percentage of initial employment in the group were as follows: High School Dropouts: 20%, High School Graduates: 9.9%, College Dropouts: 6.5%, College Graduates: 14.1%, Overall 11.0%. The formulas used to obtain single entries and averages are identical to those used in Table 9. The method used to construct the percentage changes in wages is identical to the one used in table 9. The change in capital-labor ratio due to immigration as of 2004 and 2009 is calculated using yearly immigration flows and the recursive formula (23) in the text. The effect of immigration 1990-2004 on capital-labor ratio as of 2004 (column 1) is -3.4% and it is -2.0% as of 2009 (column 2). To the contrary the effect assuming fixed capital (column 4 and 5) is -11%.

**Table 11**
**Calculated Percentage Changes in Real Wages due to Immigrants Inflows: 1990-2004.**
**Long-run effects. Robustness Checks, for different values of δ, η.**

| **Value of δ** | | **1.5** | | | **2.5** | |
|---|---|---|---|---|---|---|
| **Value of η** | | **3** | | | **5** | |
| **Specification**<br>**Value of $\sigma_k$** | **1**<br>**Low**<br>σ=5 | **2**<br>**Median**<br>σ=6.6 | **3**<br>**High**<br>σ=10 | **4**<br>**Low**<br>σ=5 | **5**<br>**Median**<br>σ=6.6 | **6**<br>**High**<br>σ=10 |
| **% Real Wage Change of Us Born Workers due to immigration, 1990-2004** | | | | | | |
| HS dropouts US-born | -1.6% | -2.5% | -3.5% | +0.6% | -0.3% | -1.3% |
| HS graduates, US-born | +3.3% | +2.8% | +2.3% | +2.7% | +2.2% | +1.8% |
| CO dropouts, US-born | +4.6% | +4.2% | +3.9% | +3.2% | +2.9% | +2.6% |
| CO graduates, US-born | 0.9% | +0.2% | -0.6% | +1.7% | +1.0% | -0.2% |
| **Average, US-born** | **+2.3%** | **+1.8%** | **+1.2%** | **+2.3%** | **+1.8%** | **+1.2%** |
| **% Real Wage Change of Foreign Born Workers due to immigration, 1990-2004** | | | | | | |
| HS dropouts Foreign-born | -21.0% | -17.8% | -14.0% | -19.3% | -15.3% | -11.4% |
| HS graduates, Foreign-born | -31.2% | -23.3% | -15.0% | -31.5% | -23.4% | -15.3% |
| CO dropouts, Foreign-born | -16.1% | -11.2% | -6.4% | -17.5% | -12.7% | -7.8% |
| CO graduates, Foreign-born | -32% | -24.8% | -17% | 31.1% | -23.8% | -16% |
| **Average Foreign-born** | **-26%** | **-19.6%** | **-13.6%** | **-26%** | **-19.6%** | **-13.6%** |
| **Overall Average: Native and Foreign-Born** | **0%** | **0%** | **0%** | **0%** | **0%** | **0%** |

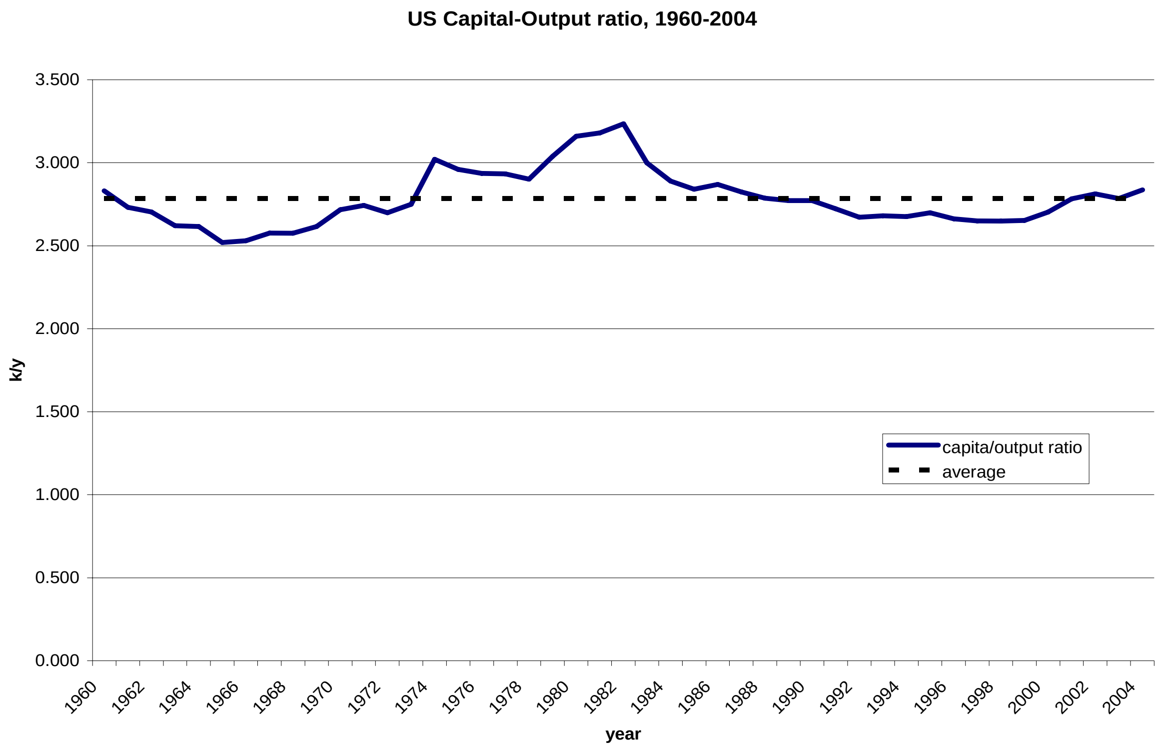
**Note:** Inflow of immigrants in the period 1990-2004 as percentage of initial employment in the group: High School Dropouts: 20%, High School Graduates: 9.9%, College Dropouts: 6.5%, College Graduates: 14.1%, Overall 11.0%. The formulas used to obtain single entries and averages are identical to those used in Table 9. The method used to construct the percentage changes in wages is identical to the one used in table 9.

**Table 12**
**Effect of Immigrants on Real Wage Dispersion of US natives, 1990-2004**

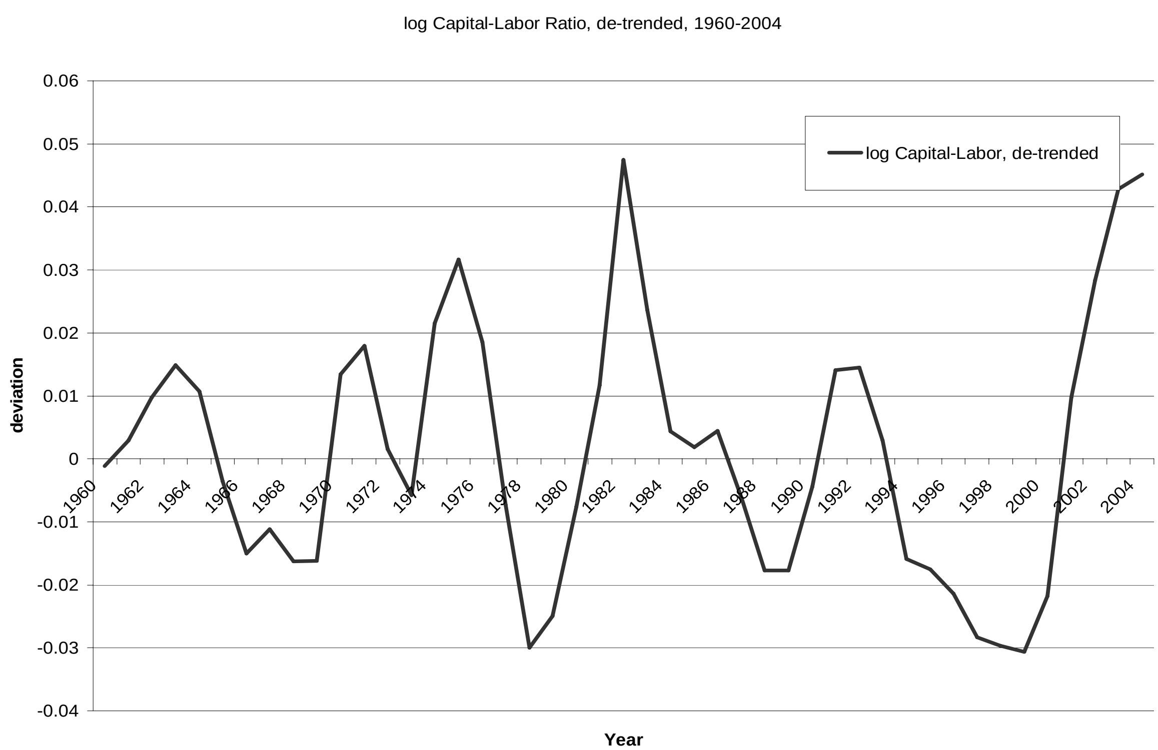
| | 1<br>Actual Percentage Change 1990-2004 | 2<br>Percentage Change relative to the average | 3<br>Percentage Change (relative to the average change) due to Immigration, Our model<br>σ=6.6 | 4<br>Share of (2) explained by (3) | 5<br>Percentage Change (relative to the average change) due to Immigration<br>σ=∞ | 6<br>Percentage of (2) explained by (5) |
|---|---|---|---|---|---|---|
| **Real percentage Changes in Wages of Education Groups 1990-2004** | | | | | | |
| **Real Wage of US-born, HS dropouts** | -11.9% | -24.4% | -2.9% | 0.12 | -4.3% | 0.17 |
| **Real Wage of US-born HS graduates** | 6.5% | -6.1% | +0.6% | -0.098 (Attenuate Dispersion) | +0.9% | -0.15 (Attenuate Dispersion) |
| **Real Wage of US-born CO dropouts,** | 8.5% | -4.1% | +1.4% | -0.34 (Attenuate Dispersion) | +2.3% | -0.56 (Attenuate Dispersion) |
| **Real Wage of US-born, CO graduates** | 21.5% | +8.9% | -1.1% | -0.12 (Attenuate Dispersion) | -1.4% | -0.015 (Attenuate Dispersion) |
| **Real percentage Changes in Wage Premia, 1990-2004** | | | | | | |
| **College/High School Dropout Wage Premium** | +33.3% | +33.3% | +1.8% | 0.05 | +2.9% | 0.087 |
| **College/High School Graduates Wage Premium** | +15% | +15% | -1.7% | -0.11 (Attenuate Dispersion) | -2.3% | -0.15 (Attenuate Dispersion) |

**Note:** The wages for each group are calculated considering all US-born workers between the ages of 17 and 65, from the IPUMS Census 1990 and the IPUMS American Community Survey 2004 as described in the main text. CPI deflator is used to convert the wages in constant 2000 $. The Average growth of real wages between 1990 and 2004 was 12.5%. It is calculated weighting the percentage increases in real wage of each education group by their average wage shares in the period 1990-2004.

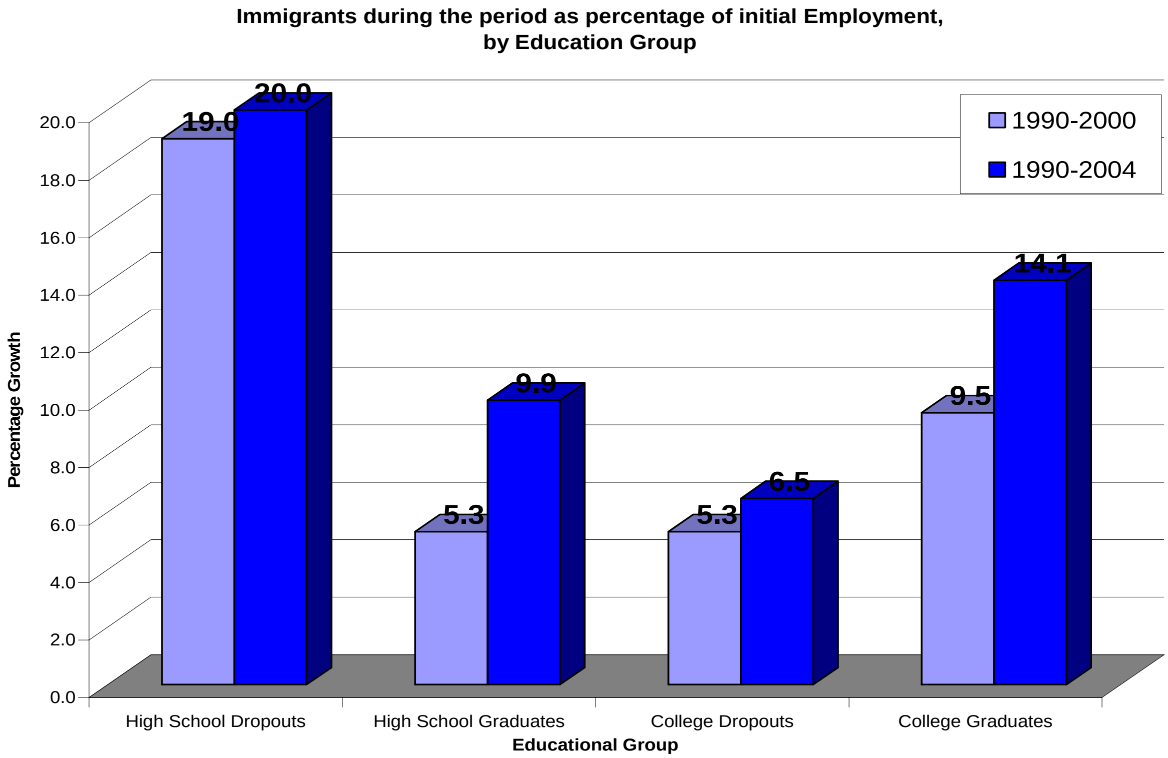
# Figure 1
# Capital-Output Ratio




# Figure 2
# De-trended log Capital-Labor Ratio



# Figure 3
## Immigration and Employment Growth, 1990-2004




# Figure 4
## Growth of Real Yearly Wages of US natives: 1990-2004.