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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 question: what was the effect of surging immigration on average and individual wages of U.S.-born workers during the period 1990-2004? We emphasize the need for a general equilibrium approach to analyze this problem. The impact of immigrants on wages of U.S.-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-experience-gender group (because they choose different occupations and have different skills). Moreover, accounting for a reasonable speed of adjustment of physical capital we show that most of the wage effects of immigration accrue to native workers within a decade. These two facts 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 run and in the long run. This positive effect results from averaging a positive effect on wages of U.S.-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 each year. 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. Two facts have contributed to feeding the debate. On the one hand, the proportion of uneducated workers (without a high school degree) has become increasingly large among recent immigrants and it is estimated that an increasing share of these workers is comprised of undocumented immigrants, mainly from Mexico and Central America[2]. On the other hand, the real wage of uneducated U.S.-born workers has performed very poorly: it even declined in real terms during recent decades (see, for example, Autor, Katz and Kearney, 2005). However, when scrutinizing the poor wage performance of uneducated U.S. workers in the context of labor market competition from immigrants the recent empirical literature has provided a mixed set of results.

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 larger negative effects (Borjas, Friedman and Katz, 1997)[3]. Recently, the latter view of a significant 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 (2007) and Borjas (2006) who use a similar empirical method, argues 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.

Our paper builds on the model presented in Borjas (2003), and takes a fresh look at some critical issues which imply significant 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 workers who differ by education, experience, nativity and gender should build on a production function that describes how these different types of workers interact with each other and with

[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] See for instance Passel (2006).

[3] 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.

physical capital to produce output. Then, one can derive the demand for each type of labor, which depends on the 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 these can be used to estimate empirically the elasticities of substitution (relative wage elasticities) between workers. 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 with respect 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 on wages.

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-gender group. Whether it is because immigrants tend to choose a different set of occupations, because they are a selected 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-gender 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 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. Since 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), and the empirical evidence as well, implies that capital adjusts in order to maintain a constant real return (and capital output ratio). 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 years 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 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 aforementioned effects, we significantly 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) the average wage of U.S. native workers increased moderately (+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 workers at the expense of a decrease in wages of foreign-born workers (namely, previous immigrants). Second, the group of least educated U.S.-born workers *suffers a 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 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 is the reason for this attenuation. Third, *all other groups of U.S.-born workers* (with at least a high school degree), accounting 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, 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 receive the largest non-economic benefits from the immigration of spouses, relatives or friends, making them more 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 revisit, 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 to 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[4]. 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 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? The consensus emerging from the literature is that the first (macro) effect on average U.S. 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 consequently 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 complementarities 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, adverse impact on 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),

[4]This review does not intend to be exhaustive. For a recent and articulate overview of the estimates of the effect of immigration on wages see Longhi, Nijkamp and Poot (2005).

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

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; Welch, 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, gender 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. We also allow males and females to be imperfectly substitutable in order to check whether within education-experience cells the gender composition of immigrants affected the gender-specific wage of natives. Within this framework we estimate three sets of elasticities: (i) between U.S.- and foreign-born workers within education-experience groups, separately for males and females as well as together; (ii) between experience levels within education groups; and (iii) between education groups. There is scant literature estimating the first set of elasticity parameters. Often, however, imperfect substitution between natives and immigrants has been cited as the reason for finding small wage effects of immigration on natives and larger negative effects on wages of previous immigrants (see Longhi, Nijkamp and Poot, 2005, page 468-469 for a discussion of this issue). We are only aware of two studies that explicitly estimate the elasticity of substitution between natives and immigrants, namely 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[6], and Cortes (2006) who considers low-skilled workers and uses metropolitan area data to find a rather low elasticity of substitution between U.S.- and foreign-born workers[7]. 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

[6] Jaeger (1996) is also the only previous work we know of that considers male and female workers separately when analyzing the substitutability between U.S.- and foreign-born.

[7] A recent paper by Manacorda, et al. (2006) applies a similar framework as this paper to British data and finds values of the elasticity of substitution between natives and immigrants similar to those found in the present paper.

on capital accumulation and different types of shocks, the speed of adjustment of capital to deviations from its long run 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 that occurred in the period 1990-2004 on average wages as of 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 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, 2007).

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 Kearney, 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 recent work (Ottaviano and Peri, 2005, 2006, 2007) and recent work by David Card (Card 2007) points out a *positive and significant* effect of immigration on the average wage of U.S. natives across U.S. states and metropolitan areas. This positive and significant effect survives 2SLS estimation, using instruments that should be exogenous to city-specific unobservable productivity shocks. The complementarities in production illustrated in this paper are able to reconcile the negative partial effects of immigrants on wages estimated in previous studies (such as Borjas 2003) with the positive average effect of immigration on native wages at the local level[8].

[8]The city model is developed in greater detail in Ottaviano and Peri (2007).

3 Theoretical Framework

In order to evaluate the effects of immigrants on the wages of natives and other foreign-born workers when each group differs by education, experience and gender 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. 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 first 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 and, within each of those groups, men and women are allowed a further degree of imperfect substitutability within the same education and experience cell. While the nested CES function imposes restrictions on the elasticities of substitution across skill groups it has the advantage of being parsimonious in parameters and widely used. Moreover, it 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, α, is reasonably 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; θ_{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 so that $k = 1$ denotes high school dropouts (HSD) , $k = 2$ denotes high school graduates (HSG), $k = 3$ college dropouts (COD) and $k = 4$ college graduates (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 θ_{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). Since 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 $\eta > \delta$. Distinct from most of the existing literature, we define L_{kjt} as a CES aggregate of U.S.-born and foreign-born workers. Denoting the number of workers with education k and experience j who are, respectively, U.S.-born and foreign-born, by H_{kjt} and F_{kjt}, and the elasticity of substitution between them by $\sigma > 0$, our assumption is that:

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

The terms θ_{Hkjt} and θ_{Fkjt} measure the specific productivity levels of foreign- and U.S.-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 standardized so that $(\theta_{Hkjt} + \theta_{Fkjt}) = 1$. 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 are differentiated enough to be imperfect substitutes for U.S.-born workers, even within the same education and experience group. Different productive characteristics between men and women could also imply imperfect substitutability between genders within the groups of native and foreign-born workers. While the gender gap literature does not explicitly explore the possibility that men and women are imperfect substitutes in production, recent studies emphasize the use of different skills and different task performances between men and women (Bacolod and Blum, 2006; Black and Spitz-Oener, 2007). Hence in Section 5 below, we test the degree of substitutability between U.S.-born men and women and foreign-born men and women allowing any degree of substitutability between men and women within education-experience-nationality cells. In practice, we allow each of the terms H_{kjt} and F_{kjt} to be CES aggregates of gender-specific labor units that can be represented as:

$$H_{kjt} = (\theta_{HMkjt} H_{Mkjt}^{\frac{\lambda-1}{\lambda}} + \theta_{HWkjt} H_{Wkjt}^{\frac{\lambda-1}{\lambda}})^{\frac{\lambda}{\lambda-1}} \ , F_{kjt} = (\theta_{FMkjt} F_{Mkjt}^{\frac{\lambda-1}{\lambda}} + \theta_{FWkjt} F_{Wkjt}^{\frac{\lambda-1}{\lambda}})^{\frac{\lambda}{\lambda-1}} \tag{5}$$

where H_{Mkjt} and F_{Mkjt} measure male employment and H_{Wkjt} and F_{Wkjt} measure female employment among natives and immigrants, respectively, and the terms θ_{HMkjt}, θ_{HWkjt} θ_{FMkjt} θ_{FWkjt} capture their respective productivity. In the empirical part we test whether the restriction $1/\lambda = 0$, which would imply perfect substitution between men and women, is rejected by the data or not. 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 or not and whether their substitutability depends on their gender or only on their foreign origin.[9]

A simplification adopted here is that the elasticity of substitution between U.S.- and foreign-born, σ, is taken to be equal across education and experience groups. Such a simplification seems empirically sound– in fact in Ottaviano and Peri (2006) we allowed different elasticities between different education groups (namely σ_k, for $k = HSD, HSG, COD, COG$) and we could never reject the restriction of identical values across groups ($\sigma_k = \sigma$). At the same time such a restriction allows us to estimate much more precisely the parameter σ since we can pool observations from all education and experience groups.

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 with a 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

[9]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 of either gender are workers of identical type.

positive trend growth determined only by total factor productivity ($\ln A_t$) and is not affected by the size of L_t. Therefore the average wage in the economy, which depends on K_t/L_t, does not depend on immigration in the long run. Shocks to L_t, such as immigration, however, may temporarily affect the value of K_t/L_t, causing it to be below its long run trend. How much and for how long $\ln(K_t/L_t)$ remains below trend 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 as 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 over 1990-2004 on average wages in 2004 and in the subsequent 5 to 10 years as capital continues to adjust. We devote the next section, 3.2.1, to showing in detail the connection between average wages and the 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 a 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 the 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[10]. In a competitive market it equals the marginal productivity of L_t, hence:

$$w_t^L = \frac{\partial Y_t}{\partial L_t} = \alpha A_t \left(\frac{K_t}{L_t}\right)^{1-\alpha} \tag{6}$$

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 a 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 a constant rate (Kaldor, 1961). In particular, this is true for our period of consideration 1960-2004. As depicted in Figure 1 the capital-output ratio (K_t/Y_t) shows small deviations around a constant mean over the 44 years considered. Moreover, the de-trended log capital-labor ratio, $\ln(K_t/L_t)$, shown in Figure 2 exhibits

[10] 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.

remarkably fast mean reversion, as evidenced by the fact that the path of the variable crosses the mean (equal to 0) eleven times in the sample. This suggests that deviations from the trend do not take more than 4 years, on average, to be eliminated[11]. 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 κ_t is the capital-labor ratio. Hence w_t^L (from equation 6) can be expressed in the following form:

$$w_t^L = \alpha A_t \left(\kappa_t\right)^{1-\alpha} \tag{7}$$

Calculating the marginal productivity of capital and equating it to the interest rate, r augmented by capital depreciation δ, 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 (7) 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 in average wages due to immigration can be expressed as a function of the percentage response of κ_t to immigration. Taking partial log changes of (7) relative to immigration we have:

$$\frac{\Delta w_t^L}{w_t^L} = (1-\alpha)\left(\frac{\Delta \kappa_t}{\kappa_t}\right)_{immigration} \tag{8}$$

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 ΔF_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.

[11] We analyze the capital data and their dynamic behavior empirically in section 6.2.

3.3 Effects of Immigration on Wages

We 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:

$$\ln w_{Hkjt} = \ln\left(\alpha A_t \kappa_t^{1-\alpha}\right) + \frac{1}{\delta}\ln(L_t) + \ln\theta_{kt} - \left(\frac{1}{\delta} - \frac{1}{\eta}\right)\ln(L_{kt}) + \ln\theta_{kjt} - \left(\frac{1}{\eta} - \frac{1}{\sigma_k}\right)\ln(L_{kjt}) + \ln\theta_{Hkjt} - \frac{1}{\sigma_k}\ln(H_{kjt}) \quad (9)$$

To keep notation less cumbersome and because men-women differences are not the focus of this paper, in the current section we maintain the aggregate male-female notation within education-experience-nativity groups. Hence H_{kjt} (F_{kjt}) represents the total input of male and female U.S.-born (foreign-born) workers of education k and experience j and w_{Hkjt} (w_{Fkjt}) represents the average wage of the group. In general H_{kjt} and F_{kjt} are the aggregates described by (5). However if men and women of the same education, experience and place of birth are perfectly substitutable the aggregates H_{kjt} and F_{kjt} are simply the sum of male and female U.S.-born (foreign-born) workers in the group[12]. We assume that the relative efficiency parameters, as well as total factor productivity A_t, depend on technological factors and are therefore independent from the supply of foreign-born.

We denote 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 (9) 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 (9). 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 wages resulting from immigration in each skill group, as well as the response of the capital-labor ratio κ_t, yields the wage change due to immigration for each U.S.-born worker[13].

Before expressing the formula for the *total* effect of immigration on the wage of a U.S.-born worker of education k and experience j, let us show the formula for a *partial* effect of the type emphasized in the 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 κ_t constant, we obtain what much of the previous literature calls the "effect of immigrants on wages". This, in fact, measures

[12] This, in fact, seems to be the empirically relevant case as we show in section 5.1.
[13] The exact expression for each of the effects described above is provided in Appendix 1 of Ottaviano and Peri (2006).

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:

$$\varepsilon_{kjt}^{partial} = \left.\frac{\Delta w_{Hkjt}/w_{Hkjt}}{\Delta F_{kjt}/F_{kjt}}\right|_{L_{kt},L_t \text{ constant}} = \left[\left(\frac{1}{\sigma} - \frac{1}{\eta}\right)\left(\frac{s_{Fkjt}}{s_{kjt}}\right)\right] \tag{10}$$

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 $\left(\frac{1}{\eta} - \frac{1}{\sigma}\right)\ln(L_{kjt})$ in (9). According to the standard assumption in the existing literature, U.S.- and foreign-born workers are perfect substitutes within group k, j $(\sigma = \infty)$ 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 (10) 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 η), the stronger is the negative impact that immigrants have on the wages of natives with similar education and experience. In the general case that we consider $(0 < \sigma < \infty)$, $\varepsilon_{kjt}^{partial}$ 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 σ and η, as well as data on wages and employment, the *partial* elasticity $\varepsilon_{kjt}^{partial}$ can be easily calculated (see Section 5.3 below). The problem is that this elasticity does not provide *any* indication of the total effect of immigration on the wages of natives in group k, j. The reason is that in order to calculate the total effect we also need to account for the changes in L_{kt} and L_t produced by immigration, as well as for the fact that immigration alters the supply of foreign-born workers in all other education and experience groups and, finally, for the response of κ_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:

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

It is easy to provide intuition for each term in expression (11) by referring to the labor demand equation (9). The term $\frac{1}{\delta}\sum_{m}\sum_{i}\left(s_{Fmit}\frac{\Delta F_{mit}}{F_{mit}}\right)$ is the 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 (9) 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. Since those immigrants are closer substitutes for natives in group k, j than workers in other education groups, 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 (9) which is negative if $\eta > \delta$. The term $\left(\frac{1}{\sigma}-\frac{1}{\eta}\right)\left(\frac{1}{s_{kjt}}\right)\left(s_{Fkjt}\frac{\Delta F_{kjt}}{F_{kjt}}\right)$ is the additional negative effect due to the supply of immigrants with the same education and experience as natives in group k, j. This effect operates through $\left(\frac{1}{\eta}-\frac{1}{\sigma}\right)\ln(L_{kjt})$ in (9) 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 (9). 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 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:

$$\left(\frac{\Delta w_{Fkjt}}{w_{Fkjt}}\right)^{Total} = \frac{1}{\delta}\sum_{m}\sum_{i}\left(s_{Fmit}\frac{\Delta F_{mit}}{F_{mit}}\right)+\left(\frac{1}{\eta}-\frac{1}{\delta}\right)\left(\frac{1}{s_{kt}}\right)\sum_{i}\left(s_{Fkit}\frac{\Delta F_{kit}}{F_{kit}}\right)+ \\ +\left(\frac{1}{\sigma}-\frac{1}{\eta}\right)\left(\frac{1}{s_{kjt}}\right)\left(s_{Fkjt}\frac{\Delta F_{kjt}}{F_{kjt}}\right)+(1-\alpha)\left(\frac{\Delta\kappa_t}{\kappa_t}\right)_{immigration}-\frac{1}{\sigma}\frac{\Delta F_{kjt}}{F_{kjt}} \tag{12}$$

The first four terms of expression (12) are identical to those in (11). Immigration in all other skill groups (and capital adjustment) has the same effect on the wages of U.S.- and foreign-born workers in group k, j.

However, immigrants in the k, j group itself have an extra negative impact on the wages of foreign-born in the same group, represented by the final term $-\frac{1}{\sigma}\frac{\Delta F_{kjt}}{F_{kjt}}$. This term is negative for $\sigma > 0$. Immigrants compete in occupations, sectors and jobs with previous immigrants more than natives and this causes an additional negative effect on the wage of foreign-born workers. For $\sigma = \infty$, the effects of immigration on the wages of workers in group k, j would be identical, independent of their nativity.

Using the percentage change in wages 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 can only 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 change in the average wage of native workers as a consequence of changes in each group's wage due to immigration is given by the following expressions:

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

where $\frac{\Delta w_{Hkjt}}{w_{Hkjt}}$ represents the percentage change in the wage of U.S.-born in group k, j due to immigration, and its expression is given in (11). Similarly, the percentage change in the average wage of foreign-born workers is:

$$\frac{\Delta\overline{w}_{Ft}}{\overline{w}_{Ft}} = \frac{\sum_k \sum_j (\frac{\Delta w_{Fkjt}}{w_{Fkjt}} \frac{w_{Fkjt}}{\overline{w}_{Ft}} \varkappa_{Fkjt})}{\sum_k \sum_j \varkappa_{Fkjt}} = \frac{\sum_k \sum_j (\frac{\Delta w_{Fkjt}}{w_{Fkjt}}) s_{Fkjt}}{\sum_k \sum_j s_{Fkjt}} \tag{14}$$

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

$$\frac{\Delta\overline{w}_t}{\overline{w}_t} = \sum_k \sum_j \left(\frac{\Delta w_{Fkjt}}{w_{Fkjt}} \frac{w_{Fkjt}}{\overline{w}_{Ft}} \varkappa_{Fkjt} + \frac{\Delta w_{Hkjt}}{w_{Hkjt}} \frac{w_{Hkjt}}{\overline{w}_{Ht}} \varkappa_{Hkjt} \right) = \sum_k \sum_j \left(\frac{\Delta w_{Fkjt}}{w_{Fkjt}} s_{Fkjt} + \frac{\Delta w_{Hkjt}}{w_{Hkjt}} s_{Hkjt} \right) \tag{15}$$

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 (15) is approximately 0 in the long run. However, if U.S.- and foreign-born workers are not perfectly substitutable, the overall effect on the wage of U.S.-born workers (13) need not be 0 but will be positive instead and the effect on the average wage of foreign-born workers (14) 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. Since 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 Censuses 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 effectively years of potential experience. They are calculated using the variable "age" and assuming that people without a 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 by dividing that value by the number of weeks worked[14]. 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")[15]. Gender is indicated as M (male) and W (female). The average wage for workers in a cell, (the variable w_{NGkjt} for $G = \{M, W\}$, $N = \{H, F\}$, $k = \{HSD, HSG, COD, COG\}$ 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_{Mkjt} F_{Mkjt}, H_{Wkjt} and F_{Wkjt}) is calculated as the weighted sum of workers belonging to that cell. These data also allow us to construct the variables $\varkappa_{Nkjt}$ and s_{Nkjt}, the

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

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

share of each group in the total wage bill and in total employment for each represented year t. These data are used to estimate the parameters δ, η, σ and λ needed to calculate the effects of immigration on the wage of each type of worker. When estimating the structural parameters δ, η, σ and λ we always use the whole panel of data, 1960-2004. When we simulate the effects 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 education and experience.

Table A1 in the Appendix reports the share of foreign-born workers in each education-experience group for each of the years considered, pooling men and women together. In the rows marked as "All Experience Levels" we report the share of foreign-born for the whole educational group. One fact emerges even from a cursory look at the table. The distribution of foreign-born across educational groups has been uneven (and increasingly so) over the period considered. In 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, college graduates represent 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, they were less than 8% of the total. Foreign-born were, and increasingly are, over-represented among the groups of workers with the lowest and highest education levels and are under-represented among the two intermediate groups.

Table A2 in the Appendix reports 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. The fact that emerges 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 workers has been very strong, particularly during the last two and a half decades, with the two intermediate schooling groups performing in between.

A synthetic and effective representation of the two facts described above, focussing on the most recent 14 years, is presented in two figures. Figure 3 illustrates the percentage growth of immigrant employment for each of the four educational groups. The lightly shaded columns represent immigrants for the 1990-2000 period as a percentage of 1990 employment in each education group. The darkly shaded 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 1990-2004. Figure 4, on the other hand, shows the growth rate of real wages of native workers by education group. The lightly shaded columns represent the real percentage change of yearly wages in the 1990-2000 period and the darkly shaded 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? (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, let us put forward some observations and facts that seem to suggest imperfect substitutability between U.S.- and foreign-born workers in production. Even considering workers who have identical measurable human capital (education and experience) and gender, foreign- and U.S.-born workers differ in several respects that are relevant to the labor market. First, immigrants are a selected group from their original populations and have skills, motivations and tastes that may set them apart from natives. Second, in manual and intellectual work 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 jobs and comparative disadvantages in other jobs[16]. Third, due to comparative advantages, migration networks or historical accidents, foreign-born tend to choose different occupations than natives, even for given education and experience levels. In particular, new immigrants tend to work disproportionately in those occupations where foreign-born are already over-represented. This should imply stronger wage competition (substitution) in those jobs compared to other jobs primarily held by natives. Since services of different occupations are imperfectly substitutable, this would imply imperfect substitutability between natives and foreign-born even in the same education-experience group[17].

Differences in the occupational choices of natives and foreign-born with the same education are illustrated in Table 1. 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 Census definitions), between the group of native workers and the group of foreign-born workers that share the same education group. 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 within each education group. By way of comparison, the remaining columns report the indices of

[16] Peri and Sparber (2007) develops and tests this story of different comparative advantages in production tasks between U.S.- and foreign-born workers.

[17] Evidence in support of this fact is presented in Ottaviano and Peri (2006). There we find a positive and very significant correlation between the initial share of immigrants in an occupation and the inflow of new immigrants in that occupation over the subsequent decade.

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 (a value of 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.

5 Parameter Estimates

5.1 Estimates of λ and σ

The model developed in Section (3.1) provides us with the framework to estimate the parameters λ and σ. Calculating the natural logarithm of the ratio of the wages of male and female workers within the same nativity ($N = H, F$) and skill (k, j) group we obtain the following relation:

$$\ln(w_{NMkjt}/w_{NWkjt}) = -\frac{1}{\lambda}\ln(N_{Mkjt}/N_{Wkjt}) + \ln(\theta_{NMkjt}/\theta_{NWkjt}) \quad \text{for} \quad N = H, F \tag{16}$$

Similarly, the natural logarithm of the ratio of wages of U.S. to foreign-born workers within the same skill (k, j) group for men (M) is:

$$\ln(w_{HMkjt}/w_{FMkjt}) = -\frac{1}{\sigma}\ln(H_{kjt}/F_{kjt}) - \frac{1}{\lambda}\ln(H_{Mkjt}/F_{Mkjt}) + \ln(\theta_{Hkjt}/\theta_{Fkjt}) \tag{17}$$

while for women (W) it is:

$$\ln(w_{HWkjt}/w_{FWkjt}) = -\frac{1}{\sigma}\ln(H_{kjt}/F_{kjt}) - \frac{1}{\lambda}\ln(H_{Wkjt}/F_{Wkjt}) + \ln(\theta_{Hkjt}/\theta_{Fkjt}) \tag{18}$$

Equation (16) defines the relative male-female labor demand within each N, k, j group, while equations (17) and (18) define the relative labor demand for male and female (respectively) foreign- and U.S.-born workers in group k, j. The first equation can be used to estimate the coefficient $\frac{1}{\lambda}$ which can then be substituted into (17) and (18), and those can then be used to estimate $\frac{1}{\sigma}$. In each case the estimates are consistent as long as we identify a source of variation in the relative supply of $\ln(N_{Mkjt}/N_{Wkjt})$ and of $\ln(H_{kjt}/F_{kjt})$ that is independent of the variation of relative male-female productivity $\ln(\theta_{NMkjt}/\theta_{NWkjt})$ and of the U.S.-foreign-born productivity levels $\ln(\theta_{Hkjt}/\theta_{Fkjt})$.

Our estimation strategy proceeds as follows. Due to technological reasons such as skill-biased technical change, sector-biased technical change, increased international competition, and other reasons 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 male-female and U.S.-foreign-born efficiencies to have a systematic component that may vary by education and experience over time. We control for education-year fixed effects (D_{kt}), and experience-year fixed effects (D_{jt}). At the same time, different education-experience groups may include male-female and U.S.- and foreign-born workers of systematically heterogeneous quality; hence, we control for experience-education fixed effects (D_{kj}). Conditional on these controls, we assume that the residual decennial changes in relative male-female and U.S.-foreign-born employment within each education-experience cell over time are due to random supply shocks such as demographic factors in the U.S. and in the immigrants' home countries. Using the IPUMS data from 1960 through 2004 we first estimate the following regression:

$$\ln(w_{NMkjt}/w_{NWkjt}) = D_{kj} + D_{kt} + D_{jt} - \frac{1}{\lambda}\ln(N_{Mkjt}/N_{Wkjt}) + u_{kjt} \tag{19}$$

where u_{kjt} is a residual, random, zero-mean disturbance. We estimate the regression separately for U.S.-born ($N = H$) and foreign-born ($N = F$) workers using 192 observations in each case (8 experience by 4 education groups over 6 years: 1960, 1970, 1980, 1990, 2000, 2004) and then pooling U.S.- and foreign-born for a total of 384 observations. The variables w_{HMkjt}, w_{FMkjt}, w_{HWkjt}, w_{FWkjt}, H_{Mkjt}, F_{Mkjt}, H_{Wkjt} and F_{Wkjt} are constructed as described in Section 4 above. Table 2 reports the estimates of the parameter $\frac{1}{\lambda}$. The first row is estimated by pooling cells of U.S.- and foreign-born workers, the second is estimated on cells of U.S.-born workers only and the third on cells of foreign-born workers only. All specifications use weighted least squares as the estimation method (weighting each observation by the total employment in the cell). Specifications in Column 2 omit year 2004 (not a census year) and specifications in Column 3 omit year 1960, since immigration was extremely low before 1960. Robust standard errors are clustered by education-experience groups. Two results emerge clearly from all specifications. First, the estimates of $\frac{1}{\lambda}$ are always negative and small in absolute value. Second, they are never statistically significant. A test of $\frac{1}{\lambda} = 0$ against $\frac{1}{\lambda} > 0$ never rejects the null at any confidence level. While the estimates using foreign-born only are quite imprecise, those on U.S. natives and on all individuals are rather precise. Considering the pooled sample, not only are we unable to rule out $\frac{1}{\lambda} = 0$, but we can reject at a standard confidence level the hypothesis that $\frac{1}{\lambda} = 0.04$ against the alternative of $\frac{1}{\lambda} < 0.04$. These results imply that even rather small degrees of imperfect substitutability between men and women are ruled out by the data. Essentially, men and women in the same education-experience-nativity cell are close-to-perfect substitutes. Hence, using $\frac{1}{\lambda} = 0$, equations (17) and (18) can be simplified into a single estimating equation:

$$\ln(w_{Hkjt}/w_{Fkjt}) = D_{kj} + D_{kt} + D_{jt} - \frac{1}{\sigma}\ln(H_{kjt}/F_{kjt}) + u_{kjt} \quad (20)$$

where $H_{kjt} = H_{Mkjt} + H_{Wkjt}$ and $F_{kjt} = F_{Mkjt} + F_{Wkjt}$ and the fixed effects and the error terms are defined as in (19). Table 3 reports the estimates of the parameter $\frac{1}{\sigma}$ obtained using weighted least squares on (20). To confirm the results of Table 2 and check that gender composition does not affect the elasticity of substitution between U.S.- and foreign-born workers, σ, we implement regression (20) alternatively on all workers (first row of Table 3) or, separately, on men workers only (second row) and on women workers only (third row). We also regress relative U.S.-foreign male wages on relative U.S.-foreign female employment (fourth row) or relative U.S.-foreign female wages on relative U.S.-foreign male employment (fifth row).

The basic specifications in Column 1 of Table 3 use yearly wages, are estimated using weighted least squares (weighting each observation by the total employment in the cell) and include observations from years 1960, 1970, 1980, 2000 and 2004. Specifications in Column 2 use weekly wages. In Column 3 we do not weight the observations and simply use OLS, in column 4 we omit the observations from 2004 (not a census year) and in Column 5 we omit the observations from 1960, a year with very few foreign-born workers. The estimates of $\frac{1}{\sigma}$ are significantly different from zero in each estimate. They are quite stable across specifications and, except for two cases, they are always between 0.09 and 0.21. The specifications using weekly wages tend to produce somewhat smaller estimates (between 0.05 and 0.14), while those using yearly wages give estimates between 0.13 and 0.21. Interestingly, this implies that a portion of U.S.-born workers' adjustment in response to the complementarities of foreign-born workers is reflected in an increase in their weeks worked, which is consistent with higher relative productivity of natives in cells with higher relative supply of immigrants. The standard errors are generally around 0.04. There are no systematic differences in the estimates across rows so that the gender dimension does not seem to play a relevant role in estimating the elasticity of substitution between natives and foreign-born workers. This is consistent with the previous literature that did not, in general, find differences in the impact of immigrants across genders, once skill level is controlled for (see the discussion on page 468 of the meta-study by Longhi, Nijkamp and Poot, 2005). Let us also emphasize that there is an extremely high correlation across education-experience groups between relative U.S.-foreign- born employment calculated for both genders, H_{kjt}/F_{kjt}, and the relative U.S.-foreign-born employment for males only (H_{Mkjt}/F_{Mkjt}) or females only (H_{Wkjt}/F_{Wkjt})[18]. Hence the relative U.S.-foreign-born supply faced by natives in each education-experience group is similar whether they are men or women.

Table 4 is devoted to performing some further robustness checks in estimating $\frac{1}{\sigma}$. First, by grouping U.S.- and foreign-born individuals according to their years of working experience, one could be classifying incorrectly the

[18] The partial R^2 of the regression of $\ln(H_{Mkjt}/F_{Mkjt})$ on $\ln(H_{kjt}/F_{kjt})$ after controlling for all the dummies is 0.95 and for $\ln(H_{Wkjt}/F_{Wkjt})$ on $\ln(H_{kjt}/F_{kjt})$ it is 0.93.

effective skills of foreign-born, assigning them to a group that is not their most natural category for comparison. Employers may value differently work experience accrued in the U.S. market from that accrued abroad. Hence we re-classify foreign-born workers by splitting the years of working experience between experience in the U.S. and experience abroad and then we use the "conversion" factors between foreign and U.S. experience calculated in Borjas (2003)[19]. 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. The estimates of $\frac{1}{\sigma}$ using foreign-born workers grouped by effective experience are reported in the first row of Table 4. Specification 1 through 5 mirror the corresponding specifications in Table 3. The estimates range between 0.12 and 0.18 with standard errors around 0.05. Specifications in the second row of Table 4 consider total weeks worked in each cell (rather than total employment) as the measure of labor supply. This measure accounts for possible changes in the individual labor supply decisions. The estimates are still in the $0.14 - 0.20$ range and very significant. Finally, in the last row of Table 4 we estimate $\frac{1}{\sigma}$, restricting the data to only the groups of workers with no high school diploma or with just a high school diploma. These are the groups with the least education and are certainly those with the largest number of undocumented immigrants. One might be concerned that mismeasurement due to low coverage of undocumented immigrants could bias our estimates. Alternatively, complementarities between U.S. and foreign-born workers in these groups could be weaker than within other groups. These issues would result in significantly different estimates of $\frac{1}{\sigma}$ when we restrict our sample to these groups. The estimates in the third row of Table 4 are never significantly different from those obtained using all workers, and the point-estimates are slightly above those. This seems to indicate that the Census data do not suffer from significant mismeasurement of undocumented workers and that the degree of U.S.-foreign-born complementarity is not very different within different education groups.

The estimates of $\frac{1}{\sigma}$ as a whole strongly support the idea of imperfect substitutability between U.S.- and foreign-born workers. Moreover, in general, the estimates imply a value of σ 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 that observed between experience groups for U.S. natives (estimated between 3 and 5)[20].

5.2 Estimates of η and δ

We can use equation (20) to infer the systematic component of the efficiency terms θ_{Hkjt} and θ_{Fkjt}. In particular, those terms can be obtained using the estimates of the fixed effects $\widehat{D}_{kj}$, $\widehat{D}_{kt}$ and $\widehat{D}_{jt}$ from equation (20) as

[19] Those factors are based on a wage regression that calculates (pooling 1980-1990 data) the wage growth associated with one year of work experience abroad, relative to the growth of wages associated with one year of work experience in the U.S. Specifically, 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. work experience.

[20] See section 5.2 below for those estimates in our paper and in the literature.

follows:

$$\widehat{\theta}_{Hkjt} = \frac{\exp(\widehat{D}_{kj})\exp(\widehat{D}_{kt})\exp(\widehat{D}_{jt})}{1+\exp(\widehat{D}_{kj})\exp(\widehat{D}_{kt})\exp(\widehat{D}_{jt})}, \widehat{\theta}_{Fkjt} = \frac{1}{1+\exp(\widehat{D}_{kj})\exp(\widehat{D}_{kt})\exp(\widehat{D}_{jt})} \tag{21}$$

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:

$$\ln(\overline{W}_{kjt}) = \ln\left(\alpha A_t^{\frac{1}{\alpha}}\kappa_t^{\frac{1-\alpha}{\alpha}}\right) + \frac{1}{\delta}\ln(L_t) + \ln\theta_{kt} - \left(\frac{1}{\delta}-\frac{1}{\eta}\right)\ln(L_{kt}) + \ln\theta_{kj} - \frac{1}{\eta}\ln(L_{kjt}) \tag{22}$$

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}[21]. Equation (22) provides the basis for estimating the parameter $\frac{1}{\eta}$, which measures the elasticity of relative demand for workers with identical education and different experience levels. Empirical implementation is achieved by rewriting it as:

$$\ln(\overline{W}_{kjt}) = D_t + D_{kt} + D_{kj} - \frac{1}{\eta}\ln(\widehat{L}_{kjt}) + e_{kjt} \tag{23}$$

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 are constant over time. Once we control for these systematic shifts in demand our identifying assumption, closely tracking Borjas (2003), is that the remaining variation in employment of foreign-born workers is due to supply shifts. Under this assumption, we consistently estimate the coefficient $-\frac{1}{\eta}$ in regression (23) 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 5 reports the estimated values of $\frac{1}{\eta}$. The first row of Table 5 reports the estimates based on yearly wages, while the second row uses weekly wages. Specification 1 and 2 use the value $\sigma = 1/(0.16) = 6.25$ and $\widehat{\theta}_{Hkjt}, \widehat{\theta}_{Fkjt}$ estimated from the basic specification in the first row of column 1 of Table 3 to construct $\overline{W}_{kjt}$ and $\widehat{L}_{kjt}$. To check whether the gender of workers interacts with the elasticity of substitution across experience groups, specification 1 of Table 5 includes all workers in the calculation of $\overline{W}_{kjt}$ and $\widehat{L}_{kjt}$ while specification 2 uses male workers only. Finally, in specifications 3 and 4, in order to check how sensitive the estimate of η is to imperfect substitutability between U.S.- and foreign-born workers, we also re-estimate the parameter η assuming $\sigma = \infty$ in the construction of $\widehat{L}_{kjt}$. All the estimated values of $\frac{1}{\eta}$ are significantly different from zero and between 0.25 and 0.3, with standard errors below 0.10. This implies a point estimate of η between 3.3 and 4. These values are very similar to those previously estimated in the literature;

[21]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.

the parameter η 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 η between 4 and 5. Borjas (2003) also produces an estimate of $1/\eta$ using immigration as a supply shifter and assuming 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 η equal to 3.5.

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 (23), as follows: $\widehat{\theta}_{kj} = \exp(\widehat{D}_{kj})/\sum_j \exp(\widehat{D}_{kj})$. Then we use the estimated values of η to construct, according to 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 satisfies the following expression:

$$\ln(\overline{W}_{kt}) = \ln\left(\alpha A_t^{\frac{1}{\alpha}} \kappa_t^{\frac{1-\alpha}{\alpha}}\right) + \frac{1}{\delta}\ln(L_t) + \ln\theta_{kt} - \frac{1}{\delta}\ln(L_{kt}) \tag{24}$$

where $\overline{W}_{kt} = \sum_j \left(\frac{L_{kjt}}{L_{kt}}\right) \overline{W}_{kjt}$ is the average wage in education group k[22]. Following the same strategy that we used before, we use the above expression as the basis for the estimation of $\frac{1}{\delta}$. In so doing, we rewrite (24) as follows:

$$\ln(\overline{W}_{kt}) = D_t + (Time\ Trend)_k - \frac{1}{\delta}\ln(\widehat{L}_{kt}) + e_{kt} \tag{25}$$

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 $(Time\ Trend)_k$ 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 equation (25) by 2SLS using $\ln(F_{kt})$ (where $F_{kt} = \sum_j F_{kjt}$ is used as an instrument for $\ln(L_{kt})$). Table 6 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 6 use the estimated values of $\widehat{\eta}$ and of $\widehat{\theta}_{kj}$ from specifications 1 and 2 of Table 5 to construct $\widehat{L}_{kt}$ and $\overline{W}_{kt}$. Specifications 3 and 4 use $\eta = \infty$ and symmetric weights θ_{kj} to construct $\widehat{L}_{kt}$ and $\overline{W}_{kt}$. Independently of specification and workers' gender, the estimated values are between 0.4 and 0.52 (with standard errors generally below 0.15), consistent with an elasticity of substitution across education groups around 2. The parameter δ 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),

[22] 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.

through Angrist (1995), Murphy et al. (1998), Krusell et al. (2000) and Ciccone and Peri (2005). The estimates for the parameter range between 1.4 and 2.5. Our estimates of $1/\delta$ fall between 0.4 and 0.5 and imply a δ in the vicinity of 2, which is consistent with previous estimates.

5.3 Partial Effects of Immigration on Wages

Before using the estimated values of the parameters λ, δ, η and σ 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; Section IV–but not Section VII–of Borjas, 2003; and Borjas, 2006) carefully estimate the partial elasticity of native wages to immigration within the same skill group (expressed in our equation (10)) and treat it as "the effect of immigration on wages"[23]. 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 (10) 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. Considering men and women within groups as perfectly substitutable and using estimates from Tables 3 and 5, the term $\left(\frac{1}{\sigma} - \frac{1}{\eta}\right)$ is calculated to be negative and between -0.05 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 -0.5% and -2.2% in response to an inflow of immigrants equal to 11% of the initial employment in the group[24]. We use 11% since this equals the inflow of immigrants over the 1990-2004 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, interpreting them as saying that immigration caused a negative 0.5 to 2.2 percent change (1990-2004) in average wages of natives and that groups such as high school dropouts, 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 (10) only accounts for the effect on wages of immigrants in the same skill group 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 positive and others negative, due to the relative sizes of skill groups and the relative strength of cross-group effects. Limiting our attention

[23]Even the recent meta-study by Longhi, Nijcamp and Poot (2005) considers this partial effect as the relevant estimate across studies. They find a representative value of -0.11 that is in our range of -0.05 to -0.20.

[24]This value is calculated using formula (10) 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 wages 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 $-2.2\% < \Delta w_{Hkjt}/w_{Hkjt} < -1.1\%$.

to the elasticity $\varepsilon_{kjt}^{partial}$, or 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 calculating 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 substitution). The third step (this section) uses these estimates and the actual flow of immigrants by group during the 1990-2004 period in the formulas previously derived to calculate the effects of immigrants on wages of U.S.- and foreign-born workers in individual groups as well as overall.

6.1 The Long Run Effects of Immigration on Wages

Table 7 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 proceed in five steps. First, in light of the result of perfect substitutability between men and women within cells ($\frac{1}{\lambda} = 0$), we aggregate across genders as follows: $H_{kjt} = H_{Mkjt} + H_{Wkjt}$ and $F_{kjt} = F_{Mkjt} + F_{Wkjt}$. Second, using formulas (11) and (12), the estimated parameters δ, η and σ as well as the percentage change in foreign-born workers by skill group ($\Delta F_{kj,1990-2004}/\ F_{kj,1990}$), we calculate the percentage change in real wages for U.S.-born and foreign-born workers in each skill group (k, j). Third, 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 1 to 4 and 6 to 9 in Table 7. Fourth, we average the changes across education groups for U.S.- and foreign-born separately, again weighting them by their wage shares as described in formulas (13) and (14). Those values are reported in rows 5 and 10 (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 (15)), to obtain the overall wage change reported in the last row, number 11, also in bold fonts. Rows 1 to 5 of Table 7 can be compared to the results obtained in the previous literature that mostly focuses on the effect of immigration on wages of U.S.-born workers. The lower part of Table 7 reports the effects of immigration on the 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 it will take for full adjustment to set in. The four columns of Table 7 are reported in order 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 3, 2 and 1 introduce imperfect substitutability between U.S.- and foreign-born workers, where column 3 uses the highest estimate of $\sigma = 10$, 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 7). 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%), and 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 a 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 (2007), Table 11 at -4.8%[25]. Column 4 of Table 7 reproduces that negative result obtaining a 4.2% loss in real wages for 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 7, report only a small negative effect (-1.1%) on wages of native dropouts. Overall our results show in the long run a significant positive effect of immigration on average U.S. wages, and on each group of workers with at least a high school degree, and only a small negative effect on wages of workers without a high school degree.

In our preferred specification 2 of Table 7, the group whose wages are hurt most by immigration are foreign-born workers, i.e., previous immigrants. On average they lost 19% of their real 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 workers and decreasing it for U.S.-born workers. If the negative effect on wages of foreign-born workers seems large, this is due to the massive inflow of immigrants over 1990-2004 relative to the initial size of foreign-born employment. Immigrants in the labor force have more than doubled in the period 1990-2004; in particular, foreign-born

[25]The Borjas and Katz (2007) estimates refer to the effects of immigration between 1980 and 2000.

workers have increased by 140% ($\Delta F_{1990-2004}/F_{1990} = 1.4$) during that period. Hence, even with a wage elasticity for 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 of 1.2% and a decrease in wages of foreign-born of 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 of foreign-born workers on wages reasonable? First of all, simply considering the relative U.S.- to foreign-born wages, reported in table A3, there are some skill groups that experienced large immigrant inflows and a substantial deterioration of their wage relative to natives. For instance, among high school dropouts between 20 and 24 years of experience, until 1970 wages of U.S.- and foreign-born workers were almost identical, while in 2004 U.S.-born were earning 12-16% more than foreign-born workers. At the same time, among workers with no high school degree, those with 0 to 4 years of experience did not experience a large increase in the share of foreign-born and the relative U.S.- to foreign-born wages in this group did not deteriorate. 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 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 θ_{Fkjt} by year and skill (which we controlled for in equation (20)) partly offset the negative effect of increased supply. Another reason why the efficiency term θ_{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 (2007): sectors/jobs 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 θ_{Fkjt}. The negative effects on wages of other foreign-born are, therefore, in part offset by systematic improvements in relative efficiency. Finally, the simulation in Table 7 is done for *a given level of employment* of native workers. The actual relative wages observed in Table A3 result from changes in employment of foreign-born as well as changes in employment of natives. Due to the estimated complementarities, this second change mitigates the negative wage impact on immigrants.

Finally, let us provide an explanation for an apparent puzzle raised by our results. In light of our analysis, previous immigrants are the group whose wages suffer most due to the arrival of new immigrants. Why, then, are they consistently among the strongest 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. Since immigration (legal and illegal) in the U.S. works mostly through family reunions, 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 are likely to have 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 the 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 the 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 κ_t only affects marginal productivity of workers through the overall average wage. A popular way to analyze the deviation of $\ln(\kappa_t)$ from its balanced growth path trend, used in the growth and business cycle literature, is to represent its time-dynamics in the following way:

$$\ln(\kappa_t) = \beta_0 + \beta_1 \ln(\kappa_{t-1}) + \beta_2(trend) + \gamma\frac{\Delta F_t}{L_t} + \varepsilon_t \tag{26}$$

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 " since 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 ε_t are other shocks. Assuming that immigration shocks cause a proportional decrease in κ_t for the same year ($\gamma = -1$), in order to calculate the effect of immigration on κ_t over, say, the 1990-2004 period, one needs an estimate of the parameter β_1. Once we know β_1 and the sequence of yearly immigration flows, $\frac{\Delta F_t}{L_t}$, one can use (26) 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 β_1 . The recent growth literature usually estimates 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 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 $\left(\frac{\Delta F_t}{L_t}\right)$ as shock and instrumented those with changes in population (to correct for endogeneity of employment)[26]. All estimates of β_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 over 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 (26) allows us to calculate $(\Delta\kappa_{1990-2004}/\kappa_{1990})_{immigration}$ as of year 2004 and the share of the deviation from trend that remains in 2009. Using formula (8) 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 ($\beta_1 = 1, \gamma = -1$), since the cumulated inflow of immigrants during the 1990-2004 period was 11% of the employment in 1990, this implies 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 the 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 8 uses these adjustments of the 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 7, i.e., the median and preferred estimates of σ. We also report in column 3 the long run effects for full capital adjustment (identical to column 2 of Table 7) 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. The average wage of U.S.-born workers increased by 0.7% in our estimates as of 2004, rather than experiencing a decrease of 3.5%. U.S. workers with no degree experience a loss of 2.2% of their real wage rather than a loss of almost 8%. College educated, U.S.-born workers have essentially no change in their wage (rather than a 5% loss) and the groups of high school graduates and college dropouts experience, even in the short run, significant gains rather than significant losses in their real wages.

[26] We constructed ΔF_t, for each year from 1960 to 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.

The benefits of immigration are already realized for most workers in the short run and certainly most benefits are enjoyed by 2009, with an average wage gain of more than 1% distributed as gains for the three groups with at least a 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 undergo sizeable wage losses as a consequence.

6.3 Robustness Checks

Table 9 shows the changes in the calculated long run effects when we use different values for the parameters δ and η in the simulations. While the values used in Table 8, 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 scholars report values of δ as low as 1.5 and as high as 2.5, while the range for η is between 3 and 5. We reproduce simulations from columns 1-3 of Table 8 using, respectively, the low estimates of δ and η (columns 1-3 in Table 9) and the high estimates of δ and η (columns 4-6 in Table 9). 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 δ and η . Considering columns 2 and 5 as references, since they use the median estimate of σ, we see that the wage loss of U.S.-born high school dropouts 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 in wages (-0.3%) from immigration. A similar widening of the distributional effects of wages of foreign-born workers across education groups takes place using the lower estimates. The widening distributional effects would also be observed if we lowered δ and η in the simulation with $\sigma = \infty$.

6.4 Contribution of Immigration to the Average Wage and Wage Dispersion of U.S.-Born Workers

The differences in the real wage effects of immigration on natives shown in Table 8 between specification 1 (our preferred one) and specification 5 (representative of previously estimated short run effects) are important. 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 8 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 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, since immigration has been connected to increased wage dispersion (e.g., Freeman, Borjas and Katz 1997 and several others), we can inquire as to the fraction of that increase that 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 10 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 in the real wage for each of the four groups relative to the average real increase in wages[27] and, in the last 2 rows, the table shows the real increase in the college/high school dropout 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 actually experienced a 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. The performance of U.S.-born high school dropouts has been particularly bad, with wages dropping 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 (as a 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 wages 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 7) and column 4 shows the 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 which performed worse than average, and hurt the last one that performed better. This is noted in Table 10 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 respect to the average wage. Moreover, immigration does not contribute at all to explaining the increased college/high school wage premium; if anything, 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-last row of column 4). These numbers seem to show that immigration cannot be considered

[27] 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.

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 $\sigma = \infty$[28] (column 5 and 6) one still obtains the result that immigration *attenuated* wage dispersion for three groups (helping those which are under-performing, and hurting those outperforming, the mean) while it only contributed to the under-performance of high school dropout wages. However, even in this scenario, only one sixth (0.17) of the growth differential with respect to the average wage, and less than one tenth (0.087) of the increase in the 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 in evaluating 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 within education-experience and gender cells 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-gender 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 wage 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 wages is the contingent of previous immigrants, who compete with new immigrants for similar jobs and occupations. Finally, our model implies that it is hard to claim that immigration has been a significant determinant in the deterioration of the wage distribution of U.S.-born workers during the period 1990-2004.

[28]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 10 could be either for fixed or for fully adjusted capital.

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Tables and Figures

Table 1
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 using 180 different occupations and data from the 2004 American Community Survey in Ruggles et al. (2006). The exact definition of the index is provided in the main text. The index ranges between -1 and +1.

Table 2
Relative Male-Female Wage Elasticity within Education-Experience Cells

Table entries: 1/λ (male-female relative wage elasticity)	**1** **Basic**	**2** **Omit 2004**	**3** **Omit 1960**
All Workers	-0.04	-0.05	-0.04
	(0.04)	(0.04)	(0.05)
US-born only	-0.03	-0.05	-0.02
	(0.04)	(0.04)	(0.04)
Foreign-Born Only	-0.10	-0.09	-0.12
	(0.08)	(0.08)	(0.07)

Note: All Regressions include education-by-experience fixed effects, education-by-year fixed effects and experience-by-year fixed effects. The reported standard errors are heteroskedasticity robust and clustered by education-experience. The dependent variable is the natural logarithm of relative male-female weekly wage of workers in the same education-experience –nativity group, the explanatory variable is the relative labor supply of female to male workers in the same education-experience-nativity group. The first row includes in the regression education-experience cells for U.S.- and foreign-born workers. This implies a total of 384 observations for column 1 and 320 for columns 2 and 3. The second row only includes cells with US-born workers and the third row only includes cells with foreign-born workers. This implies 192 observations for column 1 and 160 for columns 2 and 3. Observations are weighted by the total employment in the cell, in all specifications.

Table 3
Relative U.S.-Foreign-Born Wage Elasticity within Education-Experience Cells, Overall and by Gender

Dependent Variable and Explanatory variable:	**1 Basic**	**2 Weekly Wages**	**3 Non weighted**	**4 Omit 2004**	**5 Omit 1960**
1/σ (US-Foreign born relative wage elasticity) Overall estimates and by gender groups					
Relative Wage and Employment U.S./Foreign Born, All Workers	0.16** (0.04)	0.14** (0.03)	0.14** (0.05)	0.14** (0.04)	0.18** (0.04)
Relative Wage and Employment U.S./Foreign Born , Male Workers	0.16** (0.05)	0.09** (0.03)	0.15** (0.07)	0.15** (0.05)	0.17** (0.06)
Relative Wage and Employment U.S./Foreign Born , Female Workers	0.18** (0.03)	0.05** (0.01)	0.13** (0.04)	0.20** (0.05)	0.20** (0.03)
1/σ (US-Foreign born relative wage elasticity) Estimates across gender groups					
Relative Wage of U.S./Foreign Born , Male Workers and Relative Employment of U.S./Foreign Born , Female	0.15** (0.04)	0.09* (0.03)	0.12** (0.06)	0.13** (0.04)	0.17** (0.04)
Relative Wage of U.S./Foreign Born Female Workers and Relative Employment of U.S./Foreign Born Male workers	0.17** (0.04)	0.05** (0.02)	0.13** (0.06)	0.21** (0.06)	0.20** (0.05)
Observations	192	192	192	128	128

Note: All Regressions include seducation-by-experience fixed effects, education-by-year fixed effects and experience-by-year fixed effects. Errors are heteroskedasticity robust and clustered by education-experience. The 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. Results in the first row are relative to regressions that use all workers to calculate employment and wages, results in the second row use male workers only, and results in the third row use female workers only. The fourth row uses relative wages of male workers and relative employment of female workers. The fifth row uses relative wages of female workers and relative employment of male workers. Observations are weighted by total employment in the cell, in all specifications except for 3. Specifications 1 to 3 are estimated using observations for 1960, 1970, 1980, 1990, 2000 and 2004.

Table 4
Relative U.S.-Foreign-Born Wage Elasticity within Education-Experience Cells: Robustness Checks

Table entries: 1/σ (US-Foreign born relative wage elasticity)	**1** **Basic**	**2** **Weekly Wages**	**3** **Non weighted**	**4** **Omit 2004**	**5** **Omit 1960**
Using Effective Experience to construct Education-experience cells	0.17**	0.12**	0.16**	0.16**	0.18**
	(0.05)	(0.04)	(0.05)	(0.06)	(0.05)
Using weeks-Person as measure of labor supply	0.14**	0.20**	0.12**	0.14**	0.14
	(0.04)	(0.03)	(0.06)	(0.05)	(0.04)
Restricted to the groups of workers with at most a high school degree	0.20**	0.18**	0.19**	0.19**	0.22**
	(0.03)	(0.03)	(0.04)	(0.02)	(0.03)

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. The dependent variable is the natural logarithm of relative wage of US and foreign born workers in the same education and experience group, the explanatory variable is the relative labor supply of US and foreign-born workers in the same education experience group. In the first row the effective experience for foreign-born workers is calculated by converting years of experience abroad into years of US experience. The period considered is 1970-2004, since in 1960 there is no information on the year when immigrants entered the country. In the second row we use week-person as measure of labor supply of each group and 1960, 1970, 1980, 1990, 2000 and 2004 as years. In the third row we select only the education-experience groups within the two educational groups HSD (high school dropouts) and HSG (high school graduates) for a total of 96 observations.

Table 5
Estimates of $1/\eta$: Relative Wage Elasticity across Experience Cells

Entries are estimates of $1/\eta$	**CES Foreign-U.S.- Born Using Estimated σ**		**Simple Sum Foreign- U.S.-Born (imposing σ=∞)**	
Specification	**1**	**2**	**3**	**4**
Sample	**All workers**	**Male only**	**All Workers**	**Male Only**
Yearly Wages	0.30** (0.09)	0.30** (0.10)	0.26** (0.06)	0.25** (0.07)
Weekly Wages	0.29** (0.09)	0.29** (0.10)	0.25** (0.07)	0.25** (0.08)
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 male workers only to calculate the wages and employment. In parenthesis we report the Heteroskedasticity Robust Standard error clustered by education group.

Table 6
Estimates of $1/\delta$: Relative Wage Elasticity Across Education Cells

Entries are estimates of $1/\delta$	**CES across Experience Groups, estimated η**		**Simple Sum Across Experience Groups (imposing $\eta=\infty$)**	
Specification	**1**	**2**	**3**	**4**
Sample	**All workers**	**Male Only**	**All workers**	**Male Only**
Yearly wages	0.45** (0.12)	0.46** (0.14)	0.51** (0.14)	0.52** (0.16)
Weekly Wages	0.40** (0.09)	0.40** (0.10)	0.44** (0.11)	0.44** (0.12)
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 include male individuals only in calculating wages and employment. Specifications 1 and 3 use all individuals in each education and experience group. In parenthesis we report the Heteroskedasticity Robust Standard error clustered by education group.

Table 7
Calculated Percentage Changes in Real Wages Due to Immigrant Inflows: 1990-2004.
Long Run Effects ($\Delta\kappa/\kappa$=0)

Specification **Estimates of σ**	**1** **Low** σ=5	**2** **Median** σ=6.6	**3** **High** σ=10	**4** **σ , imposed = ∞**
% Real Wage Change of Us Born Workers Due to Immigration, 1990-2004				
1) HS dropouts US-born	-0.2%	-1.1%	-2.1%	-4.2%
2) HS graduates, US-born	+2.9%	+2.4%	+2.0%	+1.0%
3) CO dropouts, US-born	+3.7%	+3.4	+3.1%	+2.4%
4) CO graduates, US-born	+1.4%	+0.7%	0.0%	-1.5%
5) Average, US-born	**+2.3%**	**+1.8%**	**+1.2%**	**+0.1%**
% Real Wage Change of Foreign Born Workers Due to Immigration, 1990-2004				
6) HS dropouts Foreign-born	-20.2%	-16.3%	-12.3%	-4.4%
7) HS graduates, Foreign-born	-31.7%	-23.5%	-15%	+1.0%
8) CO dropouts, Foreign-born	-17.4%	-12.3%	-7.3%	+2.4%
9) CO graduates, Foreign-born	-31.6%	-24.2%	-16%	-1.6%
10) Average Foreign-born	**-26.3%**	**-19.8%**	**-13.3%**	**-0.9%**
11) Overall Average: 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 a 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 the 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 8
Calculated Percentage Changes in Real Wages Due to Immigrant 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 σ, 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 columns 1, 2, 3 and 4: σ=6.6, δ=2, η=4, α=0.66. Column 5 assumes : $\sigma=\infty$, δ=2, η=4, α=0.66. The inflow of immigrants in the period 1990-2004 as a 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 the capital-labor ratio due to immigration as of 2004 and 2009 is calculated using yearly immigration flows and the recursive formula (26) in the text. The effect of immigration 1990-2004 on the 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 9
Calculated Percentage Changes in Real Wages Due to Immigrant 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 **Value of σ_k**	**1** **Low** σ=5	**2** **Median** σ=6.6	**3** **High** σ=10	**4** **Low** σ=5	**5** **Median** σ=6.6	**6** **High** 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 a 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 10
Effect of Immigrants on Real Wage Dispersion of US Natives, 1990-2004

	1 Actual Percentage Change 1990-2004	2 Percentage Change Relative to the Average	3 Percentage Change (Relative to the Average Change) Due to Immigration, Our Model σ=6.6	4 Share of (2) Explained by (3)	5 Percentage Change (Relative to the Average Change) Due to Immigration $\sigma=\infty$	6 Share 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. The CPI deflator is used to convert the wages to constant 2000 $. The average growth of real wages between 1990 and 2004 was 12.5%. It is calculated weighting the percentage increases in real wages of each education group by their average wage shares in the period 1990-2004.

Figure 1
Capital-Output Ratio

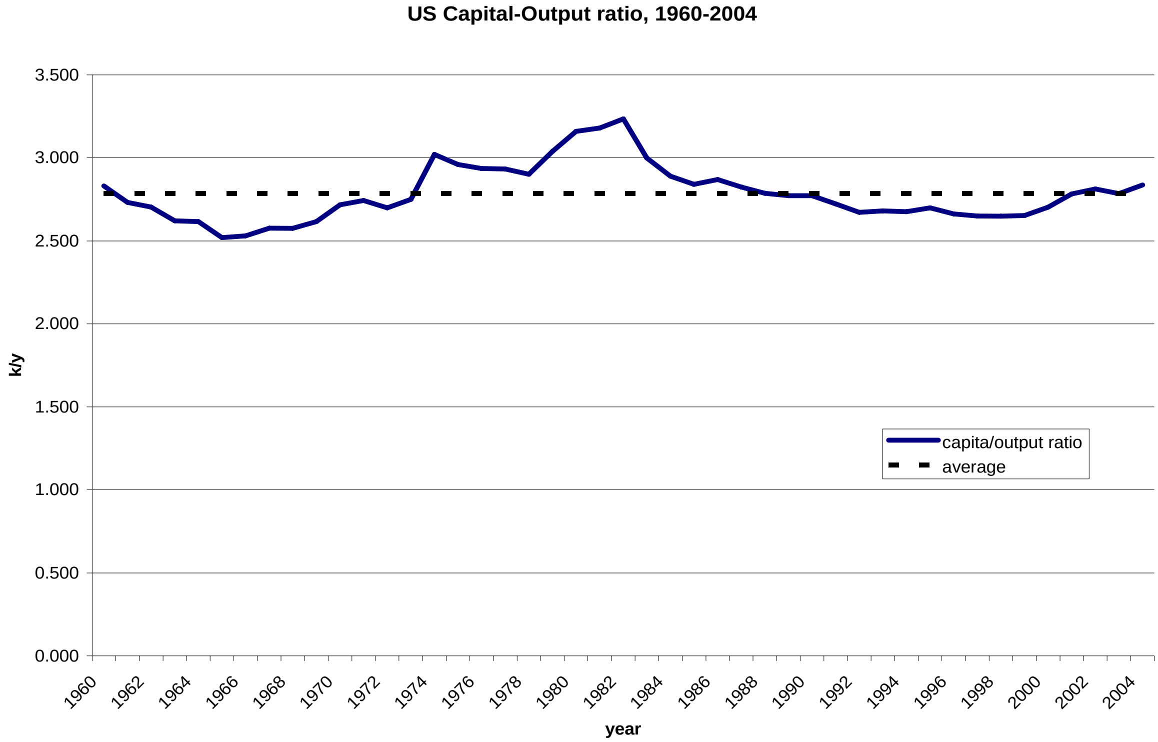


Figure 2
De-Trended Log Capital-Labor Ratio

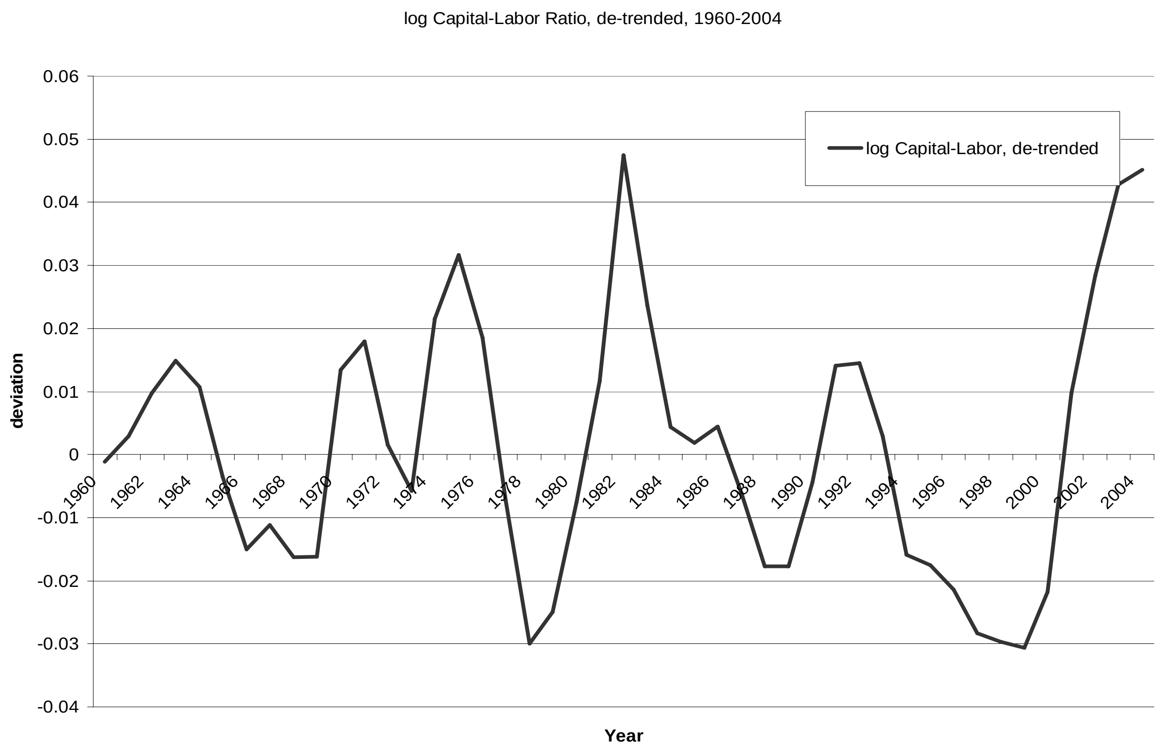


Figure 3
Immigration and Employment Growth, 1990-2004

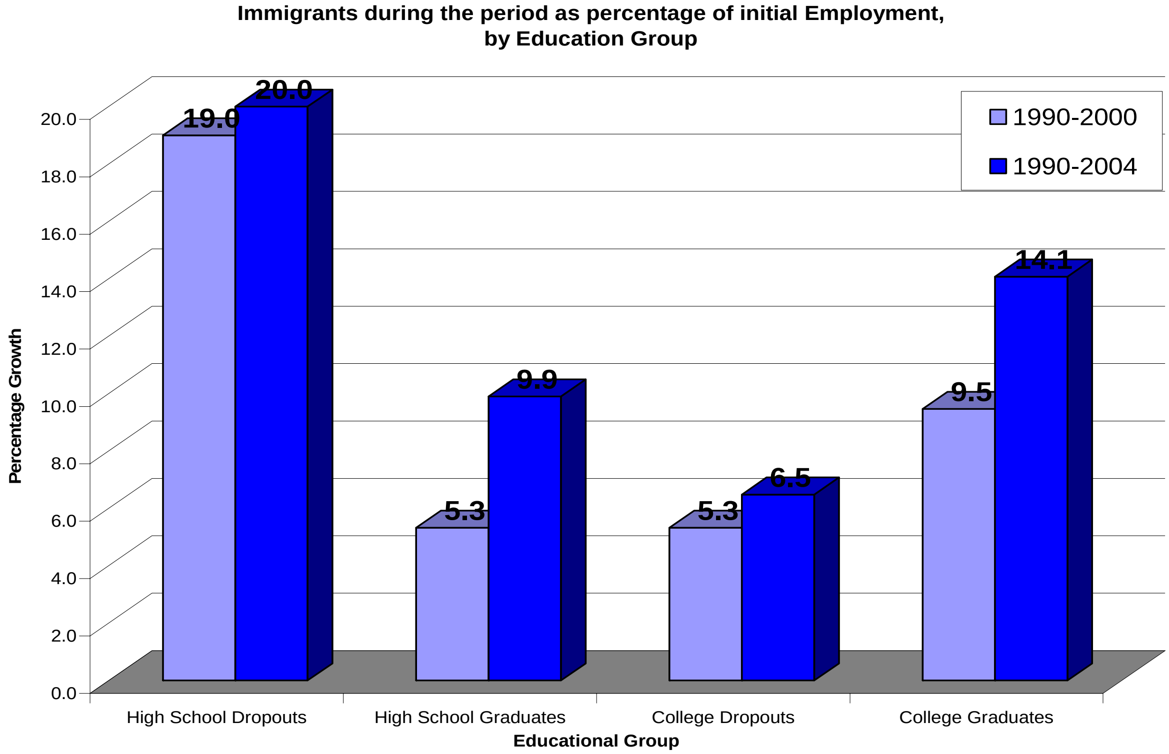
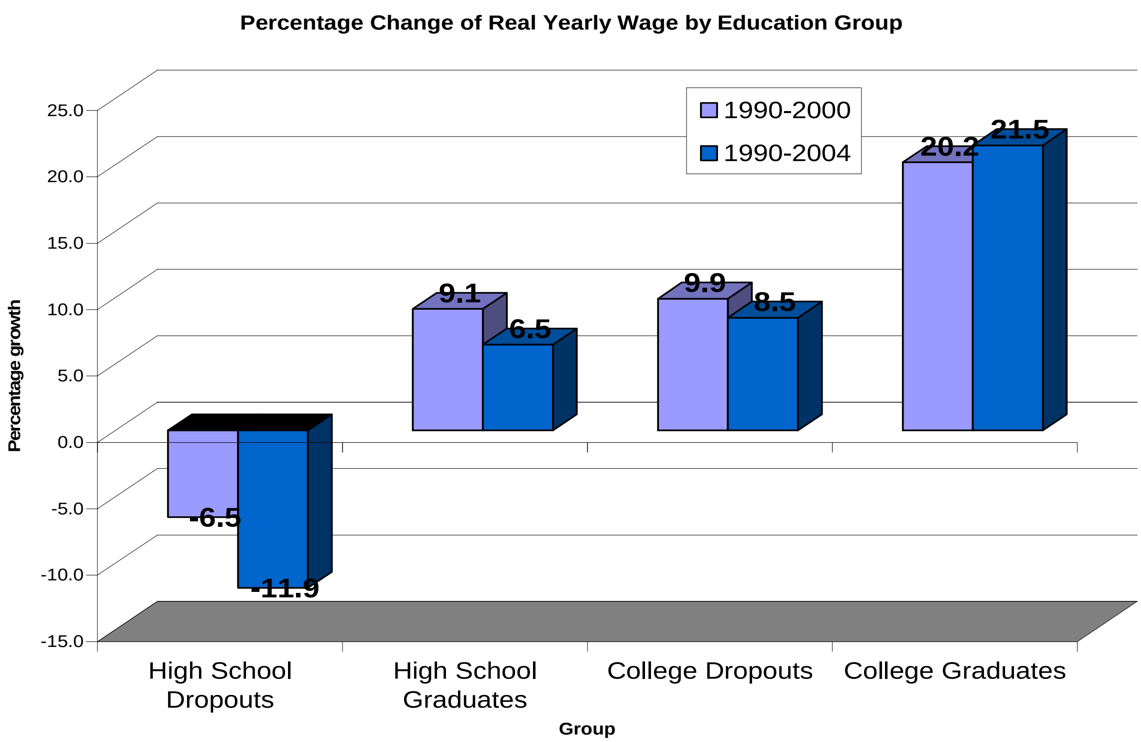


Figure 4
Growth of Real Yearly Wages of US Natives: 1990-2004.



Appendix

Table A1:
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 Experience Levels	**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 Experience Levels	**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 Experience Levels	**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 Experience Levels	**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 who are not citizens at birth. Sources: Authors' calculations on individual data from Census IPUMS and ACS from Ruggles, et al (2006).

Table A2:
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 Experience Levels	**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 Experience Levels	**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 Experience Levels	**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 Experience Levels	**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 who are US citizens at birth. Sources: Authors' calculations on individual data from Census IPUMS and ACS from Ruggles, et al (2006).

Table A3:
Relative Weekly Wages Foreign-Born/ US-Born Workers by Education and Experience

Group		Year					
Education	**Experience**	**1960**	**1970**	**1980**	**1990**	**2000**	**2004**
High School Dropouts	0 to 4	1.172	1.147	1.304	1.463	1.756	1.442
	5 to 9	0.969	1.003	0.956	0.984	1.082	1.624
	10 to 14	0.969	1.030	0.942	0.904	1.062	1.031
	15 to 19	0.987	1.024	0.954	0.937	0.936	0.980
	20 to 24	1.009	0.996	0.959	0.991	0.947	0.874
	25 to 29	1.008	0.966	0.925	0.868	1.038	0.859
	30 to 34	1.055	1.046	0.892	0.933	0.935	1.027
	34 to 40	1.075	1.007	0.882	0.909	0.859	0.921
	All Experience Levels	1.030	1.027	0.977	0.999	1.077	1.095
High School Graduates	0 to 4	0.936	1.046	1.008	1.000	1.070	1.056
	5 to 9	0.948	0.977	0.918	0.988	0.961	0.970
	10 to 14	0.885	0.963	0.961	1.030	1.004	0.909
	15 to 19	0.984	1.011	0.962	1.037	1.057	0.938
	20 to 24	1.032	0.971	0.947	1.027	0.949	0.986
	25 to 29	1.063	0.965	0.962	0.982	0.988	0.905
	30 to 34	1.043	1.058	0.902	0.930	0.936	0.953
	34 to 40	1.060	1.022	0.961	0.976	0.980	1.142
	All Experience Levels	0.994	1.002	0.953	0.996	0.993	0.982
College Dropouts	0 to 4	0.963	0.985	0.948	1.114	0.997	0.964
	5 to 9	0.920	0.966	0.955	1.013	1.053	1.142
	10 to 14	0.915	1.002	0.976	0.980	0.969	0.931
	15 to 19	0.976	0.928	0.962	1.017	0.986	0.982
	20 to 24	1.054	0.950	0.950	1.018	1.023	1.105
	25 to 29	1.011	0.919	0.954	0.952	0.967	0.915
	30 to 34	1.059	1.034	0.910	0.980	0.978	1.042
	34 to 40	1.005	1.053	0.953	1.026	0.959	0.967
	All Experience Levels	0.988	0.980	0.951	1.012	0.992	1.006
College Graduates	0 to 4	0.959	0.984	0.995	1.011	1.163	1.141
	5 to 9	0.929	0.917	0.959	0.934	1.120	1.002
	10 to 14	0.927	0.909	0.977	1.006	1.017	1.020
	15 to 19	0.946	0.966	1.005	1.047	0.945	0.967
	20 to 24	0.980	0.991	0.956	1.014	0.946	0.905
	25 to 29	0.994	0.941	0.989	1.083	1.049	0.951
	30 to 34	1.021	1.011	0.933	0.996	0.973	0.965
	34 to 40	0.976	1.008	0.899	1.017	0.918	0.905
	All Experience Levels	0.966	0.966	0.964	1.013	1.016	0.982

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. The entries are equal to the ratio of foreign-born to native weekly wages. Natives are workers born within the US or who are US citizens at birth. Foreign-born are workers born outside the US and who are not citizens at birth. Sources: Authors' calculations on individual data from Census IPUMS and ACS from Ruggles, et al (2006).