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Essays on the Economics of Migration and Labor Market Entry

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Abstract

In this thesis I explore the labour market dynamics of specific categories, namely migrants and young individuals, relating to their entry into new labour markets and their performance in the subsequent years.

In the first chapter, joint with Joseph-Simon Görlach, we analyse the factors that shape the assimilation of migrants after their arrival in the host country. Using US data, we estimate the effect that two different channels of assimilation, skill growth and substitutability, play in determining the wage of migrants. We find that the equivalence of wage growth with productivity, a common assumption in the literature, is in fact inaccurate. Immigrants, as they accumulate skills, become much more productive than reflected in their wage growth, because at the same time they are also becoming more substitutable with natives and long-term migrants, with whom they compete for jobs. This increased competition dampens immigrants' wages. Therefore, substitutability must be taken into account alongside skill accumulation when evaluating assimilation and the impact of immigration on wages.

In the second chapter, I explore the patterns of migration within the European Union, differentiating by skill level and contrasting it with recent trends in the USA and other developed countries. Unlike the other examples, the European Union is experiencing a rising rate of inter-State migration, and I show that this is due in particular to the rising migration rate of high-skilled workers. After showing that

the common explanations advanced for migration rate patterns in other countries cannot explain this peculiarity of rising EU migration, I show that a significant cause instead is the streamlining of higher education, through the Bologna Process, and in particular the mutual recognition of qualifications from other EU countries.

Finally, the third chapter explores the impact of job instability on young people's entry into the labor market. In particular, I focus on the effect of the expansion of more precarious opportunities, such as internships, on the wage of more stable long-term contracts, such as apprenticeships. I develop a simple model whereby two effects can be simultaneously at play when new legislation increases the share of precarious contracts that can be activated: 1) a selection effect increases the average wage of permanent contracts since the most productive workers are offered a permanent contract; 2) a dumping effect decreases the average wage since entrant workers are willing to bargain lower wages in exchange for the stability of a permanent contract. Using a 2017 legislative change to internships in Italy I am able to exploit interregional variation and estimate the overall effect on apprenticeship wages, revealing that the dumping effect is prevalent.

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Introduction

Labor mobility and entry into new labor markets are fundamental factors in the equilibrating mechanisms of economies, especially following asymmetric shocks. Migrants entry new national labor markets seeking job opportunities, higher wages, better match for their skills. In doing so, they impact not only their personal economic condition, but also the macroeconomic conditions of the country they leave and the country they migrate to.

For this reason, labor mobility has received much attention in the economic literature, both regarding migration flows, as individuals move across borders, and regarding outcomes of migration, as individuals assimilate in the host country. In the first two chapters of this thesis, I explore these aspects of labor migration, and conclude with a third chapter on the impact of temporary work expansion in Italy for labor market entrants.

The first chapter, in co-authorship with Joseph-Simon Görlach, tackles the question of what happens to immigrants after their arrival in the host country. Immigrants' assimilation as they adapt to the host country is often measured by wage growth relative to native workers. This confounds two opposing forces: an adaptation of immigrants' skills that raises their productivity, and an increasing substitutability with earlier immigrants and natives, which puts downward pressure on wages.

Based on a labor demand framework, we decompose the wage growth of immigrants into these components, and estimate the model's structural parameters using U.S. Census data. Since cohort sizes may be endogenous to wage growth through immigration and emigration choices, we instrument immigrant numbers with variation in economic conditions in the country of origin and relative price levels. Results show that both skill growth and substitutability progress as migrants assimilate to the host country, that skill growth strongly exceeds wage growth, and that a small positive short-run effect of immigration on the average wage of natives dissipates as immigrants become more substitutable to natives.

The second chapter shifts the focus on the migration flow and the differential patterns of skilled versus unskilled migration. While full labor mobility is a frequent assumption in currency-union models, it is far from an accurate depiction of the reality. Structural factors can affect moving costs for workers, and for this reason policies aiming to ease labor relocation have been at the forefront of the European Union's agenda.

This opens an interesting window for comparison with the case of the United States, which has been often considered the benchmark for a low-barrier labor market. Previous research has highlighted a declining pattern in interstate mobility in the US, which seems to be the case for many other countries. I show that instead, in the EU labor migration has actually been increasing rather than decreasing. I then test an important reason proposed to explain the US decline in interstate migration on the European case, to check whether these factors are not relevant or whether simply there are stronger forces driving the increased mobility.

I show the different pattern for skilled and unskilled workers, finding that the driving force for migration is the skilled labor force. As a result, I check the strength of the main institutional channel likely to account for the trend differential in skilled workers mobility, which is assimilation of educational standards. As EU countries streamline their educational systems – in particular tertiary education – to become more standardized, it is easier for each country to recognize the degrees and qualifications of EU migrants. Therefore, it will be easier for skilled migrants to move and they will be able to do so at a lower cost.

I take advantage of the Bologna process, a set of reforms made in several countries with the goal of creating uniform university degrees, and in particular the common recognition frameworks, to show the impact of standardized degree recognition processes. The higher migration rate for high-skilled individuals following the reform indeed shows that an easier recognition of one's qualification is an important factor driving intra-EU migration, and suggests that the EU should strive towards easier degree recognition in order to reduce the barriers to labor flows.

Finally, entry into a new labor market occurs not only for those who migrate to another country, but also for those who transition from education into the world of work. The outcomes of this transition are, however, strongly impacted by the labor market conditions upon entry. The literature has long shown that negative business cycle conditions at graduation have a long-term impact on a series of labor-market outcomes: unemployment, skill mismatch, and lower earnings. A related but much less explored question is the effect of unstable entry into the labor market regardless of business cycles. A significant trend in several countries is the increasing use of more precarious contracts as entry-level in firms.

An important example in Italy is the post-graduate internship (*tirocinio extracurriculare*), intended as a formative opportunity and in theory not qualified as a job contract. In the past decade, however, it has grown significantly and become almost a standard hiring format for many firms, keeping young workers suck among short-term positions for several years before receiving a stable role. Due to its low costs to the firm, the internship also tends to replace the formally dual entry-level contract combining education and work, i.e. the apprenticeship, which is a more stable form of employment as it is configured as an open-ended contract.

Institutional changes to these entry-level instruments change the bargaining relationship between employer and employee, thus leading to important effects not only on the share of contracts offered of one type versus another, but also on the wage structure of these contracts. While some research has been conducted on the effects of more flexible contracts on the wages of temporary workers, little research exists on the effects of these temporary contracts on the wages of permanent workers.

As workers and employers negotiate the contractual position of new employees, the increase of temporary work could have two opposite effects on the wages: on the one hand, purely by a selection effect, permanent contracts are offered to the better skilled workers, thus resulting in a higher average permanent-contract wage; on the other hand, employers have gained bargaining power as workers are willing to compromise on wages in order to achieve more contractual stability. This latter effect can be thought of as a form of wage dumping.

It is therefore an interesting question to determine which of these two effects prevails. Taking advantage of a legislative liberalization in that occurred in Italy in 2017, which increased the possibility for firms to offer internships, I estimate the magnitude of these changes in order to determine the direction of the overall effect of more job instability on more stable jobs. Using a difference-in-differences and a triple difference setup, I find that the dumping effect is prevalent and causes an overall negative impact on apprenticeship earnings.

Chapter 1

Skills and Substitutability: A New View on Immigrant Assimilation

Joint with Joseph-Simon Görlach

1.1 Introduction

The estimation of immigrants' career profiles has been a long-standing branch of economic migration research, not least because it indicates the economic contribution of immigrants to a host country's economy. In a partial equilibrium perspective, immigrants' wage growth is interpreted as an improvement in language and other skills valued in the host country labor market.

In this paper we emphasize a second important determinant of wages that changes as immigrants assimilate to the host country: their increasing substitutability with earlier immigrants and native workers. Immigrants becoming ever closer substitutes to native workers escalates competition and puts downward pressure on migrant wages, offsetting part of the gains from skill accumulation. Quantifying these two dimensions of assimilation is important for understanding and measuring the impact of immigration on many outcomes, including on the wage of natives. We specify a nested constant elasticity of substitution (CES) production function where nests reflect the contribution of immigrants who have been in the U.S. for different numbers of years, or of U.S.-born workers.[1] As they spend more time in the host country, immigrants move through these nests, experiencing a change both in skill efficiency and in the substitutability with natives and other migrants. We employ a nonlinear method of moments estimator to match the model to U.S. Census and American Community Survey (ACS) data over two decades, and to estimate series of productivity and substitutability parameters. Immigration and return migration choices imply that the number of migrants of any given cohort may be endogenous to wage growth. We thus leverage variation in GDP growth and relative price levels in migrants' countries of origin to instrument cohort sizes observed in U.S. data.

A vast body of economic research examining wage progress of immigrants has emerged following Chiswick (1978)'s exploration of the "Americanization" of foreign-born workers, not least because of an interest in understanding immigrants' contribution to the host economy, for instance through taxes (Auerbach and Oreopoulos, 1999; Storesletten, 2000) or social security contributions (Kırdar, 2012). Immigrants' wages may increase with time spent in the host country for many reasons, ranging from the acquisition of language skills and integration to better job matches. While many studies are silent about the exact source of wage growth, some highlight specific mechanisms. In a recent paper, Lessem and Sanders (2020) specify a job ladder model, in which wages increase through occupation transitions. Surovtseva (2022) provides evidence for trade-related occupational upgrading. Fouka et al. (2020)

[1]The production function also distinguishes workers by education level, and in a robustness check by ethnic origin. We show that the model can be extended to other individual characteristics too.

highlight perceptions of outgroup distance by the majority population. A survey of earlier papers is provided by Cadena et al. (2015).

In accounting for changes in labor market competition as immigrants assimilate, and equilibrium effects on observed wages, our analysis dissects two thus-far convoluted aspects: that Americanization not only entails the acquisition of host country-specific skills, but it also leads to an increased substitutability with U.S. workers. Host country-specific skills accumulate for instance through experience in particular sectors in which immigrants are employed disproportionately (such as elderly care or agricultural harvesting). These often differ from the main sectors of employment in countries of origin, so that immigrants may arrive with a relatively low level of productive skills. The second channel, i.e. immigrants becoming more substitutable with natives, may involve a change in the tasks that immigrants are willing to perform at a given wage (such as physically tiring, dangerous or late-night work, see e.g. Orrenius and Zavodny, 2009). Other characteristics, such as language knowledge, entail an increase in both skill and substitutability.

Natives typically enjoy an advantage in communication intensive occupations. Over time, however, immigrants move into these occupations, in rising competition to earlier migrants and natives. Figure 1.1a shows that during the first years in the U.S., the share of immigrants working in service and professional jobs increases at the expense of manual-operative and production-repair jobs. We directly observe this pattern in raw data from the American Community Survey (ACS), by plotting occupation shares across the number of years immigrants have spent in the country. Panel (b) of the Figure, instead, uses a methodology akin to Peri and Sparber (2009). Specifically, it merges ACS occupation data with the U.S. Department of Labor's O*NET database,[2] and tracks the importance of different skills in a cohort's

[2]The O*NET database assigns to each occupation scores detailing the importance of each of

occupation composition over time. The graph reveals that immigrants indeed move to less manual and more communicative jobs, approaching the shares of native workers. Immigrant assimilation hence not only occurs through the acquisition of skill efficiency, but also involves competition for more similar jobs to natives, suggesting that substitutability between immigrant and native workers changes too. Our main analysis will account for compositional changes and selective out-migration that may explain part of this trend.

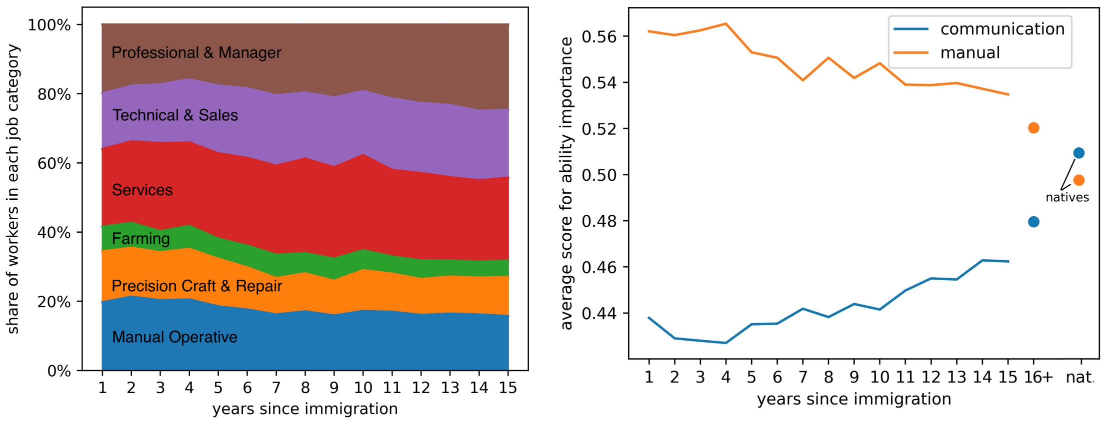
An important channel through which immigrants become more substitutable with native workers, and with long-term migrants, is by learning the language of their host country. Lewis (2013) highlights the importance of language ability for imperfect substitutability, evaluating the different impact of immigration on those who speak English well and those who do not, particularly for Spanish speakers. Appendix Figure 15 supports the notion that immigrants move into communication intensive jobs due to accumulation of an essential skill, language, over time spent in the host country.

We examine wage growth within a factor demand framework, drawing methodologically on evaluations of the wage impact of immigration. Whereas some studies rely on regional variation in immigration to estimate the effect on wages, Borjas et al. (1997) and Borjas (2003) advocate for a more aggregate factor-proportions approach to facilitate an analysis at the national level and thus avoid contamination of the estimates due to native relocation. Manacorda et al. (2012), Ottaviano and Peri (2012), and more recently Allen et al. (2019), Burstein et al. (2020), and Amior and Manning (2022) use similar setups, in which estimates of the impact of immigration crucially hinge on whether immigrants complement or substitute native labor in pro-

52 distinct abilities. These abilities are divided into those focusing on manual-physical tasks and those pertaining to the communication-language realm.

Figure 1.1: Immigrants' occupational assimilation over time.

(a) Occupational categories (b) Manual vs. communication jobs



Notes: Panel (a) plots for each year since arrival the share of immigrants in each occupational category as detailed by the OCC1990 variable (IPUMS data). Panel (b) plots the fractions of immigrants' jobs which respectively use manual and communication abilities, constructed from the IPUMS and O*NET databases following the procedure of Peri and Sparber (2009). In both cases, the cohort of immigrants tracked is the 2003 cohort, the last one for which the full 15 years can be tracked in our data.

duction. Building on the insights from this literature when analyzing immigrants' wage growth, we expand the labor demand framework to account for immigrant assimilation and time spent in the host country. Time since arrival differentiates immigrants and is a distinctive characteristic along which the applicability of immigrants' skills to the host country changes, as well as their substitutability with other workers. This time dimension introduces a dynamic element to the above literature, where in our nested CES production function immigrants successively move through different nests, each of which is characterized by a unique parameter pair determining skill efficiency and the elasticity of substitution with other inputs. This builds on D'Amuri et al. (2010), who, for an evaluation of the wage impact of immigration to Germany, distinguish two broad groups of immigrants according to the time they have spent in the country. We refine the analysis and distinguish immigrants year by year during the first decade they spend in the U.S., when assimilation is strongest. Within this fine differentiation of the early years in the country, we then decompose wage growth for immigrants into skill accumulation and changes in substitutability, an aspect not examined thus far.

We find that both skills and substitutability increase with the number of years spent in the host country, at a similar pace for immigrants of different education levels. In particular, our estimates show a sizable assimilation to natives through closer substitutability, with a change in the elasticity of substitution from about 3 at arrival to 18 after ten years in the country for college-educated and to 22 for less educated immigrants (conditional on other observables). The series of elasticities of substitution between immigrants and natives which we estimate rationalize differences in the estimates by Manacorda et al. (2012) and Ottaviano and Peri (2012): low elasticities of substitution for new arrivals (on whom Manacorda et al., 2012, focus), and higher elasticities for migrants who have been in the country for longer (akin to the inflow over a 10 year time period considered by Ottaviano and Peri, 2012). This increasing substitutability offsets part of the wage gains arising from skill accumulation, resulting in observed wages growing at a lower rate than skill efficiency taken in isolation. Specifically, we find that in the first 10 years since arrival, for less educated workers 67 percent of the skill gap between first-year immigrants and natives is closed, while for college educated workers 57 percent of the skill gap between first-year immigrants and natives is closed. On the other hand, in the same time span only 8 percent of the wage gap is closed for less educated workers, and 16 percent for college educated workers. The difference between education groups arises from low educated immigrants becoming closer substitutes to low educated natives faster than the rate at which substitutability progresses for college educated immigrants.

We furthermore use the model to evaluate dynamic wage effects. Whereas much of the literature estimating the wage impact of immigration evaluates substitutability upon entry and estimates short-run effects in static settings,[3] our model can be used to predict effects over time as immigrants assimilate. We find a negligible positive short-run impact of immigration on the average wage of natives, which quickly dissipates as immigrants become more substitutable to natives.

Besides wage growth for immigrants of a given cohort, our estimation exploits the fact that equilibrium wages of different cohorts are determined in different years, and are thus subject to variation in the workforce composition. Specifically, the size and distribution of the migrant population in terms of host-country tenure varies over time and determines equilibrium wages. Since the number of migrants resident in the U.S. at any point in time may be endogenous to immigrants' wage growth,[4] we further use changes in economic conditions in immigrants' countries of origin as an exogenous source of variation that affects emigration and return migration. The importance of home-country incomes has been highlighted both for the initial emigration decision and for return migration (see e.g. Kırdar, 2012, 2013; Djajic et al., 2016; Lessem, 2018). Other studies emphasize relative price levels between home and host country as an important factor, in particular for temporary migrants remitting and repatriating foreign savings (Yang, 2006; Albert and Monras, 2019; Akay et al., ming). Following this evidence, we exploit variation in per capita GDP growth in countries of origin and relative price levels to instrument the sizes of immigrant cohorts residing in the U.S. over time.

[3]Exceptions include Braun and Weber, 2016, and Monras, 2020, who examine labor market adjustments over time following the post-World War II refugee migration shock in Germany and the Mexican peso crisis, respectively.

[4]Several studies have emphasized the correlation between wage levels and out-migration (Hu, 2000; Duleep and Dowhan, 2002; Lubotsky, 2007; Abramitzky et al., 2014; Rho and Sanders, 2021), and this concern plausibly carries over to wage growth (Dustmann and Görlach, 2015), on which we focus in this paper.

Our baseline specification distinguishes different education levels. However, we also estimate a version of the model in which immigrants are not allocated to specific age or education cells. While this forces us to execute the analysis on a more aggregate level, it avoids a misspecification if immigrants downgrade relative to their formal qualification (see Dustmann et al., 2013). Our focus on an aggregate national production function also avoids issues of sorting by immigrants into regional markets, as documented by Altonji and Card (1991), and more recently by Cadena and Kovak (2016) and Albert and Monras (2019). It further alleviates concerns of internal migration by natives in response to immigration, as studied by Borjas (2006), Amior (2021), and Piyapromdee (2021). Our estimates are robust to different specifications of the instruments, for states with different immigration and industry structures, as well as to controlling for demographic variables. When differentiating immigrants' and natives' ethnic origin, we find that Hispanic immigrants become close substitutes to natives of Hispanic origin faster than other immigrants to non-Hispanic natives.

In related work, Albert et al. (2020) investigate the effect of cohort sizes on immigrants' wage progress. Our paper differs along several dimensions: First, they specify a production function of two inputs, general skills and host country-specific skills, where immigrants differ from natives in the latter. Instead, we treat immigrants who have spent different numbers of years in the host country explicitly as different inputs in production, with a separate elasticity of substitution and skill efficiency parameter for each year that a migrant has spent in the host country, separately for immigrants with and without college education. This allows us to directly estimate easily interpretable structural parameters and flexible assimilation profiles (separately in terms of skill efficiency and of substitutability) during the first years after arrival, when the rate of assimilation to natives is highest. Second, we account for the possibility that cohort sizes may be endogenous to wage growth in the host country by instrumenting cohort size using variation in economic conditions in immigrants' countries of origin. Third, we treat the U.S. as one labor market. This comes at the expense of losing cross-state variation in the data, but wards off concerns about sorting issues of immigrants within the U.S. and native relocation in response to immigrant inflows (Borjas, 2006; Amior, 2021; Piyapromdee, 2021). A further difference is that instead of estimating the equilibrium model on men only, with the implied assumptions, we include women within the stocks of all types of immigrants as well as natives, and take out wage differences by sex and other observables prior to the main estimation of the model's structural parameters.[5] Despite these important differences, our results feature a number of remarkable similarities with their findings and complement their analysis well. In particular, both estimations find that once the increased competition with natives and with earlier immigrants is accounted for, observed wage growth implies considerably stronger skill growth than if competition effects are ignored.

The paper is organized as follows: Section 1.2 details the theoretical model and the identification strategy. Section 1.3 describes the data and the estimation, whereas Section 1.4 shows the results alongside the model's counterfactual predictions. Section 1.5 presents robustness checks, and Section 1.6 concludes.

1.2 Model and Identification

In this section, we present a factor demand model which accounts for immigrants' assimilation to the host society, and allows for a distinction of skill accumulation and changes in the substitutability with natives and earlier migrants.

[5] We also re-estimate our model with data on men only.

1.2.1 Model Setup

Production is characterized by a nested CES production function for the national economy in year t, in which the outer nest uses S types of (aggregate) labor inputs $L_{s,t}$:

$$Y_t \quad = \quad A_t \left(\sum_{s=1}^{S} \beta_s L_{s,t}^{\sigma} \right)^{\frac{1}{\sigma}}, \tag{1.1}$$

with $\sum_{s=1}^{S} \beta_s = 1$ and $s \in 1, 2, ..., S$. Furthermore, let $N_{s,t}$ denote the native labor force of type s in year t, and $M_{\tau,s,t}$ the stock of type s immigrants observed in year t who have spent τ years in the country, with $\tau \in \{1, ..., T\}$. These immigrants thus arrived in the U.S. in year $t - \tau$. Each aggregate $L_{s,t}$ consists of the respective stocks of natives and immigrants of different tenure:[6]

$$\begin{aligned} L_{s,t} \quad = \quad & \Bigg((\alpha_{1,s} M_{1,s,t})^{\rho_{1,s}} + \Bigg((\alpha_{2,s} M_{2,s,t})^{\rho_{2,s}} + \Bigg(\quad \cdots \quad \Bigg((\alpha_{T-1,s} M_{T-1,s,t})^{\rho_{T-1,s}} \\ & + \left((\alpha_{T,s} M_{T,s,t})^{\rho_{T,s}} + (\alpha_{N,s} N_{t,s})^{\rho_{T,s}} \right)^{\frac{\rho_{T-1,s}}{\rho_{T,s}}} \Bigg)^{\frac{\rho_{T-2,s}}{\rho_{T-1,s}}} \quad \cdots \quad \Bigg)^{\frac{\rho_{2,s}}{\rho_{3,s}}} \Bigg)^{\frac{\rho_{1,s}}{\rho_{2,s}}} \Bigg)^{\frac{1}{\rho_{1,s}}}. \end{aligned} \tag{1.2}$$

where the sequence of parameters $\{\alpha_{\tau,s}\}_{\tau=1}^{T}$ captures the increase in immigrants' skill efficiency over time, whereas the sequence $\{\rho_{\tau,s}\}_{\tau=1}^{T}$ governs the substitutability with respect to native born workers and earlier immigrants of the same type s.[7] In our empirical application, we let s indicate workers' education level, which

[6]The model can be extended to include additional symmetric CES layers within $L_{s,t}$ in equation (1.1), as long as their innermost nest takes the form of equation (1.2).

[7]We also have estimated an alternative model in which parameters $\alpha_{\tau,s}$ are not exponentiated by ρ_{τ}. Results are qualitatively similar and available upon request. For other contexts, the properties of asymmetric nested CES functions have been analyzed for instance by Krusell et al. (2000) and Baqaee and Farhi (2019).

likely is the most defining characteristic determining earned wages. After $\tau = T$ years immigrants do not assimilate further, but may still differ from native workers. Substitutability parameters $\rho_{\tau,s}$ take values on $(-\infty, 1]$, where lower values indicate increasing complementarity, whereas $\rho_{\tau,s} = 1$ implies perfect substitutability across input factors and an elasticity of substitution going to infinity. Assuming constant returns to scale, as documented for a large set of industries by Basu and Fernald (1997), the values of $\alpha_{\tau,s}$, considered with their respective group substitutability, are normalized so as to sum to 1; in other words, the skill efficiency of natives is given by

$$\alpha_{N,s} = \left(1 - \sum_{k=1}^{T} \alpha_{k,s}^{\rho_{k,s}}\right)^{1/\rho_{T,s}}.$$

In Appendix A, we show that an extension of this model which includes capital as a complementary factor would leave wage *growth*, on which this paper focuses and which we derive below, unaffected. This distinguishes our analysis from studies that evaluate effects on wage *levels*, which will depend on the rate at which the stock of capital adjusts to immigration. Market frictions like in the elaborate study by Battisti et al. (2018) have been considered in economic analyses of immigration only in frameworks that do not account for immigrant assimilation. While the latter is the focus of our paper, our model in turn abstracts from frictions. In a competitive market equilibrium, workers are paid their marginal product, so that with inelastic labor supply we obtain a log equilibrium wage for the most recent immigrants of

$$\ln w_{1,s} = \sigma \ln A + (1-\sigma)\ln Y + \ln \beta_s + (\sigma - \rho_{1,s})\ln L_s + \rho_{1,s}\ln \alpha_{1,s} - (1-\rho_{1,s})\ln M_{1,s},$$

where we omit calendar time t for ease of notation. To write the wages of immigrants who have been in the country for more than one year, it is useful to define a term

$F_{\tau,s}$ as

$$F_{\tau,s} \equiv \left((\alpha_{\tau,s} M_{\tau,s})^{\rho_{\tau,s}} + \left((\alpha_{\tau+1,s} M_{\tau+1,s})^{\rho_{\tau+1,s}} + \quad \dots \left((\alpha_{T-1,s} M_{T-1,s})^{\rho_{T-1,s}} \right.\right.\right.$$
$$\left.\left.\left. + \left((\alpha_{T,s} M_{T,s})^{\rho_{T,s}} + \left(1 - \sum_{k=1}^{T} \alpha_{k,s}^{\rho_{k,s}}\right) N_s^{\rho_{T,s}} \right)^{\frac{\rho_{T-1,s}}{\rho_{T,s}}} \right)^{\frac{\rho_{T-2,s}}{\rho_{T-1,s}}} \dots \right)^{\frac{\rho_{\tau,s}}{\rho_{\tau+1,s}}} \right)^{\frac{1}{\rho_{\tau,s}}},$$

so that for $\tau \geq 2$, log wages for earlier immigrants can be written as

$$\begin{aligned} \ln w_{\tau,s} &= \sigma \ln A + (1-\sigma) \ln Y + \ln \beta_s + (\sigma - \rho_{1,s}) \ln L_s + \rho_{\tau,s} \ln \alpha_{\tau,s} \\ &\quad -(1-\rho_{\tau,s}) \ln M_{\tau,s} + \sum_{k=2}^{\tau} (\rho_{k-1,s} - \rho_{k,s}) \ln F_{k,s}. \end{aligned}$$

Finally, log wages for native-born workers are

$$\begin{aligned} \ln w_{N,s} &= \sigma \ln A + (1-\sigma) \ln Y + \ln \beta_s + (\sigma - \rho_{1,s}) \ln L_s + \ln(1 - \sum_{k=1}^{T} \alpha_{k,s}^{\rho_{k,s}}) \\ &\quad -(1-\rho_{T,s}) \ln N_s + \sum_{k=2}^{T} (\rho_{k-1,s} - \rho_{k,s}) \ln F_{k,s}. \end{aligned}$$

Combining these expressions, and holding everything else constant, we can write log wage growth for immigrants who have been in the country for τ years as

$$\begin{aligned} \ln \frac{w_{\tau+1,s}}{w_{\tau,s}} &= \rho_{\tau+1,s} \ln \alpha_{\tau+1,s} - \rho_{\tau,s} \ln \alpha_{\tau,s} + (\rho_{\tau,s} - \rho_{\tau+1,s}) \ln F_{\tau+1,s} \\ &\quad +(1-\rho_{\tau,s}) \ln M_{\tau,s} - (1-\rho_{\tau+1,s}) \ln M_{\tau+1,s}, \end{aligned} \tag{1.3}$$

whereas the log relative wage for long-term immigrants as compared to natives is

$$\ln \frac{w_{T,s}}{w_{N,s}} = \rho_{T,s} \ln \left(\alpha_{T,s}\right) - \ln \left(1 - \sum_{k=1}^{T} \alpha_{k,s}^{\rho_{k,s}}\right) - (1-\rho_{T,s}) \ln \frac{M_{T,s}}{N_s}. \tag{1.4}$$

Crucially, in these expressions for wage growth by years since arrival, upper-nest parameters such as σ and β_s cancel out, thus leaving wage growth within each workers group as a function of that group's parameters only when everything else is kept constant. The expressions for wage growth across multiple years and arrival cohorts in equations (1.3) and (1.4), therefore, constitute a system of equations that identifies the model's structural parameters governing the two dimensions of assimilation, and which will be estimated in the following sections.

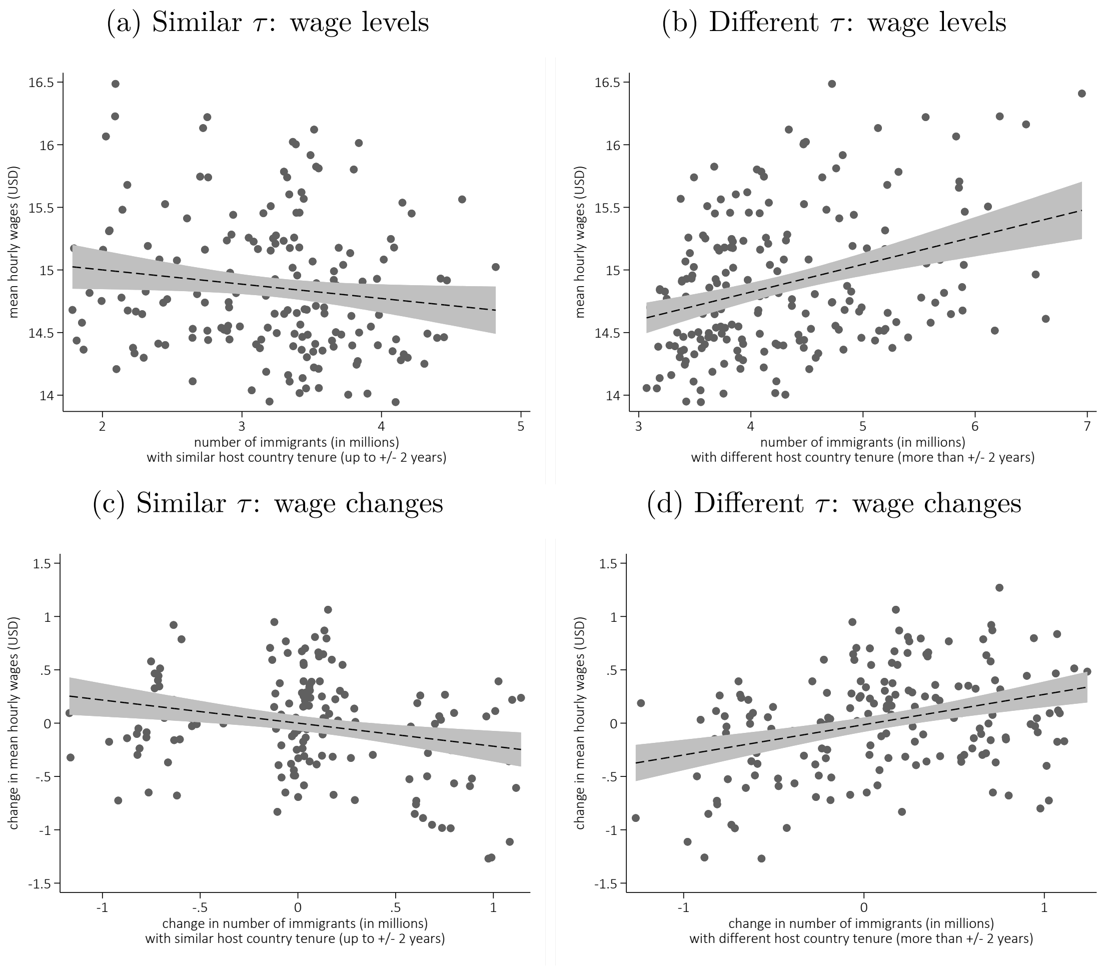
1.2.2 Identification

Estimating a national production function, and thus considering the aggregate of regional economies, eliminates concerns about immigrants' regional sorting into local labor markets, and about internal migration by native workers. Below, we furthermore explain how a potentially endogenous size of immigrant populations as well as variation in cohort quality will be accounted for in the estimation.

The non-linear equations (1.3) and (1.4) provide a set of moment conditions based on the observed relation between immigrant populations of different tenure in the host country and wage growth over time. Different from models that have been used in the literature to estimate the wage impact of immigration, our specification of the production function requires that most of its parameters (all parameters $\{\alpha_{\tau,s}\}_{\tau=1}^{T}$ and $\{\rho_{\tau,s}\}_{\tau=1}^{T}$ of the inner nests) be estimated jointly.

An increase in $\{\alpha_{\tau,s}\}_{\tau=1}^{T}$, with τ being time spent in the U.S., leads to an increasing wage profile for immigrants of type s, whereas a rise in $\{\rho_{\tau,s}\}_{\tau=1}^{T}$ works in the opposite direction. If values of $\{\rho_{\tau,s}\}_{\tau=1}^{T}$ were constant, immigrants' wage growth with τ would indicate a proportional rise in $\alpha_{\tau,s}$. Intuitively, a separate identification of changes in $\rho_{\tau,s}$ under the model can be achieved by considering variation in the exposure to other immigrant cohorts (and native workers) over time. A larger number of immigrants who have spent a similar number of years in the U.S. puts downward pressure on wages earned by any given cohort if these immigrants are relatively close substitutes. The presence of more immigrants with a rather different τ, instead, would be more positively associated with a given cohort's wage if the two groups are less substitutable or even complementary in production.

Figure 1.2: Wages and cohort sizes.



(a) Similar τ: wage levels

(b) Different τ: wage levels

(c) Similar τ: wage changes

(d) Different τ: wage changes

Notes: Panel (a) plots wages $w_{\tau,t}$ against the number of immigrants $\sum_{k,|k-\tau|\leq 2} M_{k,t}$ with the same time since arrival +/- 2 years, excluding long-term migrants who have been in the U.S. for more than 10 years. Panel (b) plots wages $w_{\tau,t}$ against the number of immigrants $\sum_{k,|k-\tau|>2} M_{k,t}$ with a more than 2 year difference in the time since arrival. Panels (c) and (d) plot the same outcomes in first time differences. Shaded areas around the fitted lines indicate 95% confidence intervals.

Figure 1.2 supports this intuition. Its left-hand panels show the correlation between immigrants' hourly wage and the number of immigrants who have been in the U.S. for a similar amount of time, both in levels (Panel a) and in differences (Panel c). Each dot represents a t, τ combination, and the stock of *similar* immigrants includes the five cohorts with the same τ plus or minus two years.[8] The downward sloping fitted line indicates a negative association between the number of immigrants who are likely close substitutes on the one hand and immigrants' wages on the other. Note that the number of immigrants may be endogenous to wages, for which we account with the instrumentation described below. The horizontal axis in Figure 1.2b, instead, measures the number of immigrants with a *different* (by more than two years) number of years since arrival, and displays a positive correlation with wages, indicating some degree of complementarity. Panel (d) of the figure confirms this also for first time differences in both wages and the number of differently tenured migrants.

While Figure 1.2 illustrates a source of variation that can identify the substitutability of immigrants on top of their skill accumulation, two main challenges common in the literature on immigrant assimilation need to be accounted for: first, changes in immigrant composition have been widely documented (Borjas, 1985; Jasso et al., 2000), which may affect the wage level immigrants realize in the host country; second, the size of immigrant populations changes over time through out-migration, in a way that may be affected by wage growth. We address these issues in turn.

Cohort effects. In any year, immigrants who have spent different amounts of time in the host country, must have arrived at different points in time. Arrival cohorts may differ in immigrants' earnings potential, conditional on observable characteris-

[8]We exclude natives and long-term immigrants with more than 10 years since arrival from this graph.

tics. Isolating wage growth with time spent in the U.S. requires an elimination of these cohorts effects. Hence, rather than on observed raw wages, the estimation will be based on predicted wages $\tilde{w}_{\tau,s,t}$ after taking out cohort effects $t-\tau$. When doing so, we also eliminate observed individual characteristics, including an individual's age, sex and country of origin.

Out-migration. With several years of data, identifying information not only derives from immigrants' wage growth with time τ spent in the country, and their relative wage to natives, but also from how wage growth at a given τ varies between years t, across which the distribution of the migrant population in terms of migrant tenure τ changes. We use immigration data over two decades, and show in Figure 1.3b below that indeed the distribution of τ changes over time. Since many migrants do not stay permanently, both variation in new arrivals and out-migration contribute to changes in this distribution.

The decision to emigrate to and to stay in the host country, however, may be linked to wage growth expected or experienced there. In this case, reverse causality renders the stocks of migrants $M_{\tau,s,t}, M_{\tau+1,s,t}, ..., M_{T,s,t}$ staying for at least τ years (the right hand side of equation (1.3)) endogenous to the growth $\frac{w_{\tau}+1,s,t}{w_{\tau,s,t}}$ of wages (the left hand side). We solve this problem by using origin-specific information which determines migration flows, but is otherwise plausibly uncorrelated with wage growth in the destination country. The importance of economic conditions in sending countries has been documented for both the initial emigration decision and the decision to return (Kırdar, 2012, 2013; Djajic et al., 2016; Lessem, 2018; Görlach, ming). A number of related papers emphasize the role of relative price levels between the home and the host country, which determines the purchasing power of assets that can be accumulated abroad and repatriated for consumption and investment in the country of origin (Yang, 2006; Albert and Monras, 2019; Adda et al., ming; Akay et al., ming). Following that evidence, we use information on GDP per capita growth in the main countries of origin and price levels relative to the U.S. as exogenous variables affecting the incentive to emigrate or return. These variables have no direct impact on immigrant wage growth in the U.S., but will be related to wages through the flows of migrants. We use weighted means across the five main migrant sending countries to construct time series for both instruments, as described more in detail in the next Section.

In terms of our notation, realizations of both instruments in year $t-\tau$, when immigrants $M_{\tau,s,t}$ have decided to move to the United States, will affect the initial size $M_{0,s,t-\tau}$ of this cohort. However, the realizations in subsequent years $t-\tau+1, t-\tau+2, ...$ also affect later decision to stay or to leave, which equally contribute to the size of the observed stock $M_{\tau,s,t}$ in year t. As our baseline instruments we therefore use the realizations of origin-country GDP per capita growth and price level ratios averaged over all years from the time $t-\tau$ of arrival to time t when the stock $M_{\tau,s,t}$ is measured. That is, we define the instrument realization for each combination of τ and t as the average $\bar{\boldsymbol{z}}_{\tau,t} = \frac{1}{\tau}\sum_{i=t-\tau}^{t} \boldsymbol{z}_i$ across the τ years immigrants observed in year t have spent in the United States, where $\boldsymbol{z}_i$ contains the realizations of origin country GDP per capita growth and price level ratios in any given year.

We then interact these instruments with the group s-specific equations (1.3) and (1.4).[9] For each type s of workers, define residuals

$$
\begin{aligned}
u_{\tau,s,t} \quad \equiv \quad & \ln \frac{\tilde{w}_{\tau+1,s,t}}{\tilde{w}_{\tau,s,t}} - \Big(\rho_{\tau+1,s} \ln \alpha_{\tau+1,s} - \rho_{\tau,s} \ln \alpha_{\tau,s} + (\rho_{\tau,s} - \rho_{\tau+1,s}) \ln F_{\tau+1,s,t} \\
& + (1-\rho_{\tau,s}) \ln M_{\tau,s,t} - (1-\rho_{\tau+1,s}) \ln M_{\tau+1,s,t}\Big) \quad \text{for } \tau < T
\end{aligned}
$$

[9]Data sources, descriptives and first stage estimations for these instruments are presented in Section 1.3.3 and Appendix C.1. In Section 1.5.1 we also present the results using the realizations of these variables in the year $t-\tau$ of arrival only.

and

$$u_{T,s,t} \equiv \ln \frac{\tilde{w}_{T,s,t}}{\tilde{w}_{N,s,t}} - \left(\rho_{T,s} \ln \left(\alpha_{T,s} \right) - \ln \left(1 - \sum_{k=1}^{T} \alpha_{k,s}^{\rho_{k,s}} \right) - (1 - \rho_{T,s}) \ln \frac{M_{T,s,t}}{N_{s,t}} \right).$$

For each τ, s and t these residuals are functions $u_{\tau,s,t}(\boldsymbol{\theta})$ of the structural parameter vector $\boldsymbol{\theta}$ that contains the full sequences $\{\alpha_{\tau,s}\}_{\tau=1}^{T}$ and $\{\rho_{\tau,s}\}_{\tau=1}^{T}$. Using the vector $\bar{\boldsymbol{z}}_{\tau,t}$ collecting K instruments, we obtain — for each τ and s — the K non-linear moment conditions

$$\sum_{t} \bar{\boldsymbol{z}}_{\tau,t} u_{\tau,s,t}(\boldsymbol{\theta}) = 0. \tag{1.5}$$

The full system of KTS such conditions then identifies $\boldsymbol{\theta}$ (cf Hansen, 1982). Identification can be demonstrated by partial differentiation of the left hand side of equation (1.5) with respect to each parameter. The resulting gradient matrix of moments with respect to parameters must have full rank. In practice, this depends on the choice of instruments $\bar{\boldsymbol{z}}_{\tau,t}$. We thus compute the gradient matrix for the actual realizations of our instruments (see Section 1.3.3). We visualize this matrix in Appendix C.2. Figure 19 reveals that there is no collinearity and that the matrix hence has full rank.

1.3 Data and Estimation

In this section, we first describe the data used, as well as the procedure used to eliminate cohort effects and variation in observed characteristics. We then show the relevance of our instruments, and finally describe the procedure by which we estimate the model's structural parameters.

1.3.1 Data Sources

To estimate the model we use United States Census and American Community Survey (ACS) data for each year from 2000 to 2018. This time frame is determined by the availability of consistent data from migrants' countries of origin only in more recent times. In the estimated version of the model, we specify characteristic s to indicate whether individuals have any college education or not. Education likely is the most defining individual characteristic in terms of wages earned. Yet, we show that the assimilation paths for parameters $\{\alpha_{\tau,s}\}_{\tau=1}^{T}$ and $\{\rho_{\tau,s}\}_{\tau=1}^{T}$, on which our analysis focuses, barely differ across education groups. Besides origin country information, estimation of the model thus requires for each calendar year t and education level s the stock of natives and the stock of immigrants, where the latter is differentiated by years τ spent in the U.S., alongside wages earned by each group and individual characteristics. Time spent in the U.S. is determined as the difference between the survey year and the reported year of immigration. Hourly wages are computed by dividing annual earnings by the average number of hours worked per week and the number of weeks worked. All wages are deflated to 1999 prices.

For the instruments, we use data from the World Bank's World Development Indicators and the OECD's statistical database. Specifically, we use 1990-2018 series of real and nominal exchange rates and GDP per capita growth in the countries of origin of the largest immigrant groups: Mexico, China, India, the Philippines and Vietnam. Relative prices in countries of origins are calculated as the ratio of nominal to real exchange rates using purchasing-power-parity conversion factors.

We construct a vector of instruments containing weighted averages of per capita GDP growth and of relative prices in the main countries of origin. To prevent a potential endogeneity of the instruments, we use time-constant weights based on the percentage of immigrants from each country of origin in the pooled Census years 2000 and 2010. Since economic growth in a migrant sending country and the purchasing power of U.S.-dollars influence both the initial emigration and later decisions to stay in the U.S., we instrument each migrant stock $M_{\tau,s,t}$ with the mean of these variables during the years between arrival in $t-\tau$ and the time of the survey t.

1.3.2 Descriptives

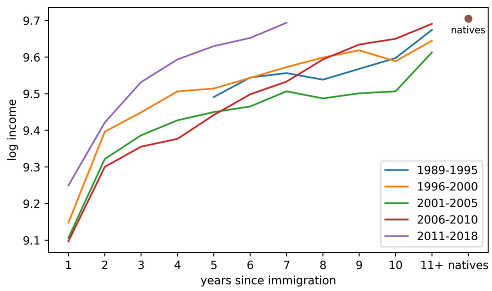
For the empirical implementation we set $T = 11$. That is, we focus on the early years after arrival when assimilation progresses most strongly, and track immigrants during their first decade in the U.S. Immigrants who have been in the country for more than 10 years are pooled in the category $M_{T,s,t}$. Table 2.1 shows the means of the variables used by the number of years since immigration. The last column shows the respective outcomes for natives.

Figure 1.3a displays annual income profiles for broad arrival cohorts, with the common pattern that earnings grow with time spent in the country. The right end of the graph shows that conditional on staying in the country long-term, migrants approach natives' incomes (averaged across all years and indicated as a disconnected dot). The figure shows that indeed cohort effects are not negligible. The data suggest that after a decline in incomes across different cohorts until the 2000s, there has been a noticeable increase in observed income for more recent cohorts.

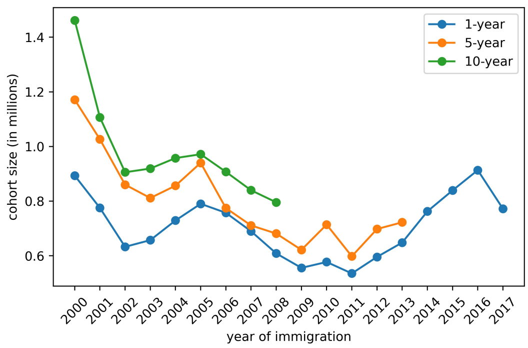
As our model relies on equilibrium conditions that equate the marginal product of labor with the wage, we focus on workers who report earning a wage income, and use that variable throughout the analysis. Since the ACS only reports yearly earnings, we compute an estimate of the hourly wage by dividing annual earnings by the number of weeks worked per year and the average number of hours worked

Figure 1.3: Variation in incomes and cohort sizes.

(a) Income paths by cohort



(b) Cohort sizes



Notes: The figure shows (a) raw annual income profiles (in logs) for immigrants of different arrival cohorts, and (b) variation in the size of immigrant cohorts over time (weighted according to the ACS individual weight). Source: ACS 2001-2018.

Table 1.1: Summary statistics

	Years since immigration											U.S.-
	1	2	3	4	5	6	7	8	9	10	>10	born
Percent of Total Population	0.34	0.35	0.37	0.37	0.39	0.38	0.38	0.39	0.38	0.45	11.12	85.09
Annual Incomes	20,829	21,981	22,291	22,736	22,903	23,477	24,122	24,304	24,951	24,672	30,303	30,014
Median Yearly Weeks	35.00	50.00	50.00	50.00	50.00	52.00	52.00	51.00	51.00	52.00	50.00	50.00
Median Weekly Hours	40.00	40.00	40.00	40.00	40.00	40.00	40.00	40.00	40.00	40.00	40.00	39.00
Percent Male	61.88	61.52	60.61	59.68	58.66	57.68	57.58	56.96	55.93	56.90	52.36	49.60
Age	33.22	33.06	33.61	33.95	34.55	35.12	35.71	36.26	36.78	37.34	48.95	46.79
Years of Schooling	12.95	12.69	12.61	12.59	12.49	12.49	12.52	12.44	12.51	12.30	12.46	13.19
Share with College	49.62	46.34	44.78	44.71	43.13	42.97	43.47	42.56	43.58	40.61	44.70	50.70
Immigrants' Origin (in percent):												
North America	3.04	2.16	2.00	1.89	1.87	1.78	1.68	1.65	1.67	1.51	2.83	
Central America & Mexico	37.05	40.91	42.40	42.49	44.85	45.29	45.04	46.04	44.80	48.73	42.90	
South America	7.99	8.05	8.19	7.88	7.90	7.51	7.75	7.74	7.75	7.60	6.29	
Europe	12.75	10.76	10.64	10.68	10.12	10.78	10.94	11.06	11.52	10.99	17.45	
Asia	32.03	30.75	29.65	29.61	28.28	27.77	27.67	26.94	27.88	25.31	27.00	
Africa	6.06	6.50	6.35	6.63	6.21	6.23	6.26	5.94	5.82	5.30	3.02	
Oceania	1.01	0.82	0.71	0.75	0.70	0.58	0.62	0.58	0.53	0.50	0.47	
N. of observations (in millions)	0.149	0.155	0.162	0.161	0.170	0.166	0.166	0.171	0.165	0.199	4.885	37.380

Source: U.S. Census 2000, ACS for 2001-2018. The table lists means (where indicated medians) among the population reporting a labor income.

per week. Taking out cohort effects removes the differences apparent in Figure 1.3a, and generates more similar wage profiles across cohorts. Since we further take out compositional changes in terms of individual characteristics, and in particular age, which contributes to wage growth over time, the wage profiles also become more concave and flatter in later years. This motivates our focus on immigrant assimilation during the very early years after arrival in the host country.

We interpret wages as equilibrium outcomes, which depend on the scale and composition of the immigrant population. This variation across years is an essential margin for identification of the structural parameters of the model. Figure 1.3b shows that indeed cohort sizes vary considerably across years—both the number of new arrivals per year and the number of immigrants who have been in the U.S. for five and ten years, respectively. Appendix Figure 16 further shows the distribution of immigrants by years spent in the country.

1.3.3 Instrumentation

We instrument the size of immigrant cohorts using price level ratios and per capita income growth in the main origin countries of migrants. Both determine the attractiveness of emigration and return migration in many economic models (see Dustmann and Görlach, 2016, for a survey). We construct time series for both determinants as weighted averages across the five main countries of origins for immigrants in the U.S. during our sample period.[10] Before calculating the moment conditions (1.5), we take out a quadratic trend from both instruments and endogenous variables so that identification derives from year-to-year fluctuations in these variables rather than long-term growth. For GDP per capita growth, we further control for the average of GDP growth across OECD countries (excluding the U.S.), to clean

[10] These are Mexico, China, India, the Philippines and Vietnam.

the instrument from global cycles. The instrument we ultimately use hence is the deviation of the origin country GDP per capita growth rate from the OECD average, net of quadratic trends. Appendix Figure 17 shows the strong co-movement of immigration flows and relative price levels, with higher purchasing power of U.S.-dollars abroad making the U.S. a more attractive destination for (temporary) migration. Growth of GDP per capita in origin countries, instead, exhibits a strong negative relation to arrivals: as countries of origin experience worse economic conditions, more workers choose to emigrate and stay abroad.

Table 1.2 lists the coefficients of a regression akin to a first stage in linear two stage least squares estimations. Each detrended instrument has strong predictive power for migrant stocks of different host country tenure. Columns (1)-(3) use our baseline instruments $\bar{\boldsymbol{z}}_{\tau,t}$, for which relative price levels and origin country GDP per capita growth are averaged across the τ years immigrants of a given cohort $t-\tau$ have been in the U.S. Columns (4)-(6) use each instrument's realization $\boldsymbol{z}_{t-\tau}$ in the year $t-\tau$ of arrival. Columns (3) and (6) include both instruments simultaneously, showing that each contains distinct information.

With these instruments at hand, and after taking out cohort effects and compositional changes from wages, we construct the vector of moment conditions (1.5). To do so, the residual matrix with elements $u_{\tau,s,t}(\boldsymbol{\theta})$ is multiplied, for each s, with the instrument matrix with elements $\bar{\boldsymbol{z}}_{\tau,t}$. Besides data, moment conditions (1.5) depend only on the structural parameter vector $\boldsymbol{\theta}$. We estimate these parameters by minimizing the (sum of squared) deviations of the moment conditions from their theoretical value. In Section 1.5.1, we conduct this estimation with GDP per capita growth and relative price levels in the year of arrival as instruments for the corresponding stocks of migrants. Additional details are provided in Appendix C.2, where in particular Figure 19 visualizes the gradient matrix of moments with respect to

Table 1.2: Instrument relevance

	Dependent Variable: Number of Immigrants (in Thousands)					
	(1)	(2)	(3)	(4)	(5)	(6)
Price Ratio (baseline)	653.688		405.566			
	(71.057)		(112.540)			
Price Ratio (arr. yr.)				293.933		262.304
				(48.172)		(48.201)
GDP/cap growth (baseline)		-115.239	-75.550			
		(21.855)	(25.259)			
GDP/cap growth (arr. yr.)					-17.369	-13.216
					(3.357)	(3.529)
Quadratic trend	Yes	Yes	Yes	Yes	Yes	Yes
OECD (excl US) GDP/cap growth	No	Yes	Yes	No	Yes	Yes
R-squared	0.076	0.079	0.099	0.066	0.035	0.085
F-statistic	84.631	27.804	39.567	37.231	26.762	26.722
Observations	380	380	380	380	380	380

Note: Dependent variable are the stocks of immigrants, denoted in thousands. Units of observation are t-τ-combinations. *Price ratio (baseline)* is computed as the ratio between nominal and PPP exchange rates, averaged over all years $t-\tau$ to t an immigrant cohort has been in the U.S. (denoted as price level in the U.S. over price level in the basket of main origin countries, see main text). *Price ratio (arr. yr.)* is the relative price level in the year of arrival $t-\tau$. GDP/cap growth (baseline) and GDP/cap growth (arr. yr.) are mean GDP growth per capita in origin countries during the years $t-\tau$ to t and in year t, respectively. All regressions include an intercept and a quadratic year trend, while with GDP per capita growth as an instrument, we also control for average GDP per capita growth in OECD countries excluding the U.S.

parameters to demonstrate local identification.

1.4 Results

The upper-level nest aggregates the most relevant characteristic, for which workers of different types exhibit a low degree of substitutability. In line with previous research on nested CES production functions (Ottaviano and Peri, 2012), in our main estimation we distinguish workers without and with college education as the most defining characteristic. The model can be extended to include additional CES layers, such as for age or experience. While this will affect the implied *level* of wages and estimates of, for instance, the effect of immigration on native wages, equations (1.3) and (1.4) for wage *growth* and the relative wage to natives are unaffected, as long as the innermost nest takes the form of equation (1.2). With our focus on education, the production function becomes

$$Y_t = A_t \left(\beta L_{l,t}^{\sigma} + (1-\beta) L_{h,t}^{\sigma} \right)^{\frac{1}{\sigma}} ,$$

with labor inputs $L_{l,t}$ and $L_{h,t}$ by workers of different education types. Each of these aggregates the number of natives and immigrants of different host country tenure within education group based on skill efficiency parameters $\{\alpha_{\tau,s}\}_{\tau=1}^{T}$ and substitutability parameters $\{\rho_{\tau,s}\}_{\tau=1}^{T}$, for $s \in \{l, h\}$, as specified in equation (1.2).

Our aim is to identify these parameter series in order to disentangle two central dimensions of assimilation. We first present our estimates for these parameters, before focusing on their implications for the interpretation of immigrants' wage growth, and testing the robustness of our results.

1.4.1 Parameter Estimates

Using a simplex minimization algorithm to solve the moment conditions (1.5), we obtain the values for skill efficiency $\boldsymbol{\alpha} \equiv (\alpha_{1,l}, \alpha_{2,l}, ..., \alpha_{T,l}, \alpha_{1,h}, \alpha_{2,h}, ..., \alpha_{T,h})'$ and substitutability $\boldsymbol{\rho} \equiv (\rho_{1,l}, \rho_{2,l}, ..., \rho_{T,l}, \rho_{1,h}, \rho_{2,h}, ..., \rho_{T,h})'$ visualized in Figure 1.4. Shaded areas indicate 99 percent confidence intervals,[11] and we list the exact point estimates in Appendix Table 3.

Both skill efficiency and substitutability increase over time. We display the parameter estimates for α_T and ρ_T for the inner-most nest as disconnected points in the figure. These parameters capture skill efficiency and substitutability of long-term immigrants who have been in the U.S. for more than 10 years. Recall that the skill efficiency parameters across all inputs are normalized to sum to one. Our point estimates for long term migrants' skill efficiencies are $\alpha_{T,l} = 0.100$ and $\alpha_{T,h} = 0.117$, and thus less than the values $\alpha_{N,l} = (1 - \sum_{k=1}^{T} \hat{\alpha}_{k,l}^{\rho_{k,l}})^{1/\rho_{T,l}} = 0.129$ and $\alpha_{N,h} = (1 - \sum_{k=1}^{T} \hat{\alpha}_{k,h}^{\rho_{k,h}})^{1/\rho_{T,h}} = 0.153$ attributed to natives' skill efficiency. Importantly, the growth paths of parameters are very similar for different education groups, with error bands being almost entirely overlapping.

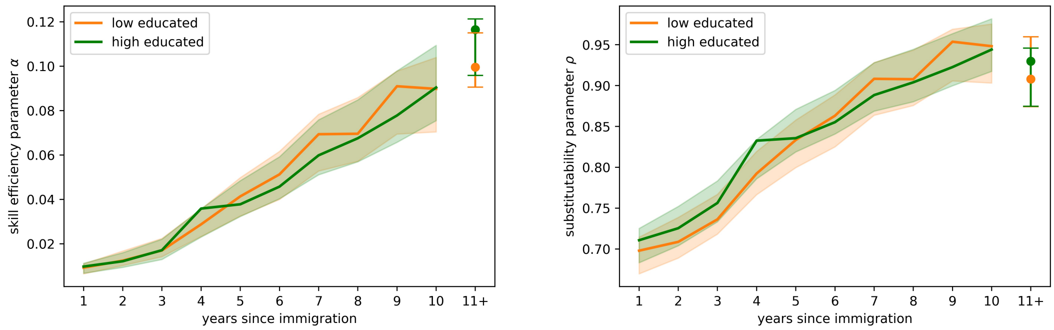
Identification of these profiles exploits both wage growth within immigrant cohorts, and variation in immigrant and native population stocks over time. Whereas wage growth has been the focus of the literature on immigrant wage assimilation, variation in competing or complementing input factors matters only in frameworks that analyze wages as an equilibrium outcome. Appendix Figure 18 illustrates the contribution of this second source of identification by showing the impact of changes in population stocks on wage growth as implied by the model.

Appendix Figure D shows the same estimation results performed keeping only

[11] We draw with replacement 1,000 bootstrap samples, for each of which we estimate the parameter vector. From this, we select the 0.5 and 99.5 percentiles as bounds of the interval.

Figure 1.4: Main parameter estimates.

(a) Skill efficiency $\boldsymbol{\alpha}$ (b) Substitutability $\boldsymbol{\rho}$



Estimated assimilation profiles in terms of (a) skill efficiency parameter series $\boldsymbol{\alpha}$, and (b) substitutability parameter series $\boldsymbol{\rho}$. Shaded areas represent 99% confidence intervals (bootstrapped with 1,000 replications). The orange line and shaded area represent the estimates for low educated workers; the green line and shaded area represent the estimates for high educated workers.

men in the sample. The resulting estimates exhibit little difference to the ones including women in Figure 1.4.[12] While the use of of male-only samples may be innocuous in some settings, wages in our framework are an equilibrium outcome. To avoid an assumption of fully gender-segregated labor markets, we keep women in our sample throughout the rest of the paper.

1.4.2 Elasticities of Substitution

The core of our analysis is the evolution of substitutability over time: as immigrants assimilate to the host society, their labor input becomes more substitutable to that of earlier immigrants and natives. A first result to note from Figure 1.4b is that all estimated values of $\boldsymbol{\rho}$ are greater than zero. Hence, even the most recent immigrants are (if only imperfect) substitutes to natives.

[12] While much of the literature restricts attention to men, Long (1980), Schoeni (1998), Butcher and DiNardo (2002), and Blau et al. (2011) present analyses of the earnings of female immigrants.

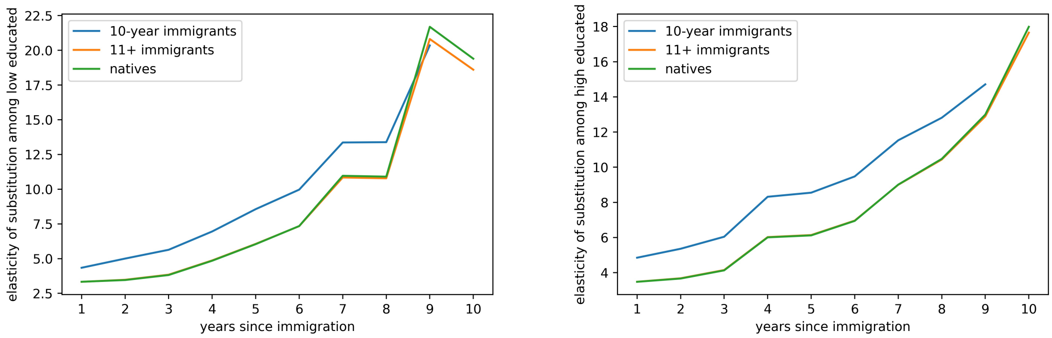
To obtain the elasticities of substitution with respect to native workers, we use our estimated values of $\boldsymbol{\rho}$ and $\boldsymbol{\alpha}$, and compute the predicted output, based on stocks of native workers and immigrants for each number τ of years immigrants have been in the U.S.[13] We then numerically compute the marginal products $\partial Y/\partial M_\tau$ and $\partial Y/\partial N$ to derive the elasticity of substitution with respect to natives. These elasticities are displayed as the green line in Figure 1.5. Recently arrived immigrants are characterized by a rather low level of substitutability with natives of the same education level in the first year after immigration, with an elasticity of substitution of around 3 in either education group. This plausibly derives from new immigrants entering jobs in sectors in which relatively few natives work, and our estimated elasticities are very much in line with results by Manacorda et al. (2012). For immigrants who arrived to the UK over a five year period, they find an elasticity of substitution with respect to natives of about 5. With more time spent in the U.S., the elasticity of substitution increases, reaching a level of about 20 for more substitutable longer-term migrants without college education, and an elasticity of substitution of 18 for college educated workers. This is similar to estimates by Ottaviano and Peri (2012), who consider immigrants who arrived to the U.S. within ten years prior to when data are collected, as well as to Albert et al. (2020), for whom these immigrants constitute the most recent group of arrivals.

This increasing substitutability during the years after arrival exerts—within skill groups—downward pressure on immigrants' wages. Wage growth is accordingly muted and appears weaker than the underlying rise in skill efficiency. In other words, immigrants' skill efficiency, reflected in the sequence of parameters $\boldsymbol{\alpha}$ in our model, increases more strongly than wage growth alone, neglecting equilibrium effects, would suggest.

[13] For the calculations in this section, we average these stocks across all years t in our data.

Figure 1.5: Elasticities of substitution.

(a) Low educated (b) High educated



Note: The figure shows the elasticity of substitution of 10-year migrants (blue line), all previous migrants (yellow line), and natives (green line) with respect to all migrants with host country tenure between 1 and 10 years. Panel (a) shows the elasticities of substitution within workers without college education, panel (b) within those with college education.

New immigrants not only compete with and assimilate to native workers, but also to earlier immigrants, who have been in the country for longer. The elasticity of substitution with respect to earlier immigrants is a parameter that has rarely been considered in the literature. Different papers have shown that earlier immigrants may be the ones experiencing the sharpest drop in wages as a result of continued immigration (see e.g. D'Amuri et al., 2010; Manacorda et al., 2012), an effect that depends on the elasticity of substitution between immigrants of different host country tenure. We are not aware of estimates for the change in this parameter during the very first years after arrival in the host country. Although Albert et al. (2020) do not directly estimate a substitution parameter, their model allows a simulation of the elasticity of substitution between broader groups of immigrant with 0-9, 10-19, etc years since arrival. Our estimation, instead, zooms into the very first of these intervals and follows immigrants yearly during this first decade in the U.S., when assimilation progresses most rapidly. D'Amuri et al. (2010) estimate a single value measuring the elasticity of substitution between immigrants who have been in Germany for up to five years and those who have arrived more than five years prior to the survey.

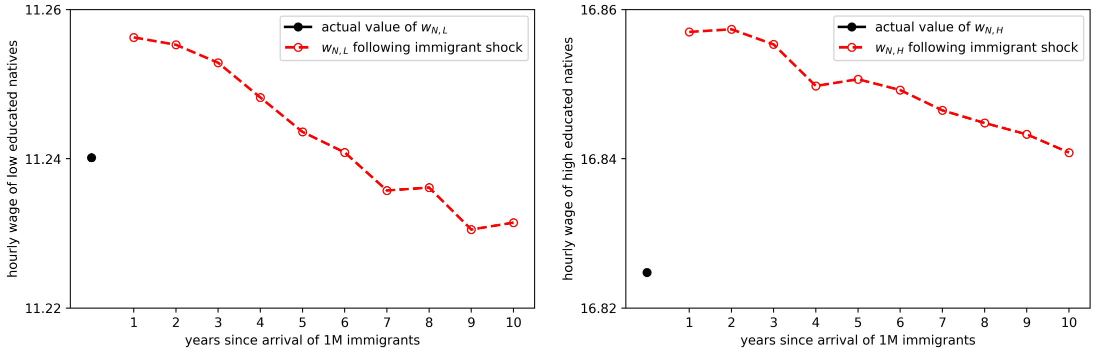
We can calculate the elasticity of substitution for each group of migrants with respect to other—more recent or less recent—immigrants. The longer immigrants have been in the country, the higher their substitutability with previous immigrants (as well as with natives). Figure 1.5 shows this increased substitutability of new immigrants with respect to earlier immigrants who have been in the country for 10 (blue line), as well as with respect to longer-term immigrants who have been in the U.S. for more than 10 years. The elasticity of substitution between 10-year immigrants and more recent ones is increasing, but consistently higher than the elasticity of substitution between those recent immigrants and immigrants who have been in the country for more than 10 years. Furthermore, we find that the elasticity of substitution between long-term immigrants (those with $\tau > 10$ years) and all earlier immigrants is only slightly higher than that of natives.

1.4.3 Impact of Immigration on Natives Wages

A strand of economic migration research that receives attention beyond academic circles examines the labor market impact of immigration on non-migrants (see Dustmann et al., 2016, for a survey). Estimated effects crucially hinge on the substitutability between native and immigrant workers. Whereas the literature has focused on the short-term impact of immigration on wages, our analysis of time-varying substitutability has implications for longer-term effects. We examine this by simulating the arrival of one additional million immigrants in the U.S. and following their impact on native workers throughout the first ten years after immigration. We perform this exercise holding all other variables constant, and assume that the immigrant inflow is a one-time shock. While our evaluation of dynamic wage effects is

Figure 1.6: Dynamic effects of immigration on natives' wages.

(a) Low educated natives (b) High educated natives



Notes: The figure shows the evolution of natives' wage (in USD) following the arrival of one additional Million immigrants, as they progress in time through years spent in the U.S., experiencing changes in both substitutability and skill efficiency. The first full dot is the actual average native wage; the dashed line shows the wage response following the arrival of immigrants with the same education composition as the observed immigrant population. Panel (a) shows the effect on the wage of low educated natives; Panel (b) on the wage of high educated natives.

novel, it abstracts from other channels through which immigration affects economic outcomes for natives, including through innovation (Hunt and Gauthier-Loiselle, 2010; Arkolakis et al., 2020) and fiscal contributions (Auerbach and Oreopoulos, 1999; Storesletten, 2000; Kırdar, 2012).

We simulate the effects of the arrival of an immigration wave with an education composition equal to that observed for the existing immigrant population (averaged across all years). To calculate the effect, we add these new immigrants to the average stock observed for each number of years since migration and compute the impact on natives' wages using equation 1.3. Different from the previous analysis, this requires a quantification of A, σ and β. In order to do this, we set $\sigma = 0.5$ following previous estimations in the literature (such as Ottaviano and Peri, 2012). We consequently estimate A and β as the values that make equation 1.3 for both education types match the empirical log wage for natives in each group, averaged across all years of observation.

The dynamic wage effect for natives is displayed in Figure 1.6. In line with much of the literature, we find little changes in native wages. Qualitatively, the imperfect substitutability at arrival with natives is reflected in the small positive short-run effect that the immigrant shock has on natives of each skill group, similar to the results by Ottaviano and Peri (2012) and Manacorda et al. (2012). The effect of immigrant assimilation on natives' wages in our framework is theoretically ambiguous: on the one hand, a higher skill efficiency of immigrants with a high complementarity to natives has a positive impact on the marginal product of native labor; on the other hand, the increased substitutability between immigrants and natives puts downward pressure on natives' wage. Empirically, we find that the latter dominates, so that the positive short-run effect dissipates as immigrants assimilate. Yet, for high educated natives, about half of the initial positive effect persists even in the long run. Appendix Figure 23 explicitly distinguishes immigration of workers with different education levels, and shows that the initial positive effect on the wage of natives of the respectively other skill group is amplified further through immigrant assimilation. Low educated natives suffer a decrease in wages following low educated immigration at all time horizons.

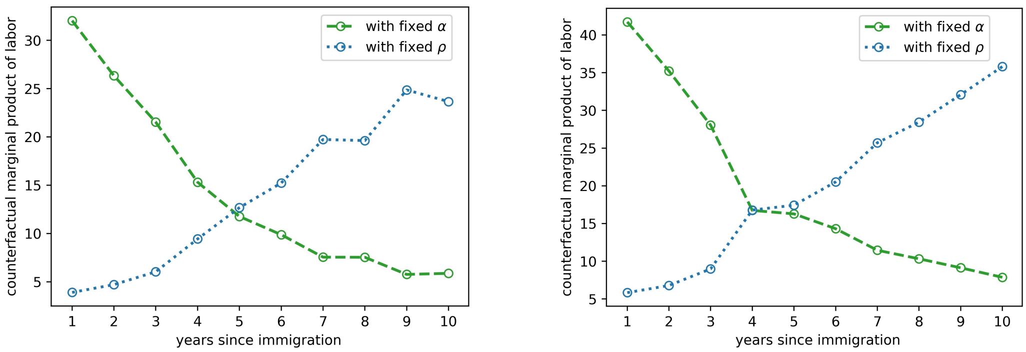
1.4.4 Interpreting Immigrant Wage Profiles

Our estimates indicate that immigrants' assimilation not only entails a growth in skill productivity, but also an increase in the elasticity of substitution with native workers and earlier immigrants. To demonstrate the magnitude of both channels, we simulate the counterfactual wage profile of immigrants that would result if one of the two channels was shut down. In a first exercise, we fix substitutability ρ_τ at its median value of 0.85 for all τ in both skill groups. In other words, we isolate the assimilation of immigrants only in terms of skill efficiency, without a change in substitutability. We then use the model to predict the marginal products of labor across τ.[14]

Figure 1.7: Counterfactual marginal products.

(a) Counterfactual $\{\partial Y/\partial M_{\tau,l}\}_{\tau=1}^{T}$ for $s = l$ (less than college)

(b) Counterfactual $\{\partial Y/\partial M_{\tau,h}\}_{\tau=1}^{T}$ for $s = h$ (with college)



Note: The figures show counterfactual wage profiles that would be obtained under alternative models when fixing (i) substitutability at its median value of 0.85 while letting skill efficiency evolve (blue dotted line), or (ii) skill efficiency at its median value of 0.046 while letting substitutability evolve (green dashed line). Panel (a) shows this for immigrants with less than college education; panel (b) for college educated immigrants.

The blue dotted line in Figure 1.7 shows the predicted marginal product for this case. The profile shows a steep increase in the marginal product arising from the growth in α_τ over time, strongly exceeding the growth in wages when equilibrium effects are accounted for. For comparison, the skill gap between first-year immigrants and natives closes by 67 percent within the first 10 years since arrival for low educated workers and by 57 percent for high educated workers, whereas only 8 percent of the wage gap is closed for low educated workers and 16 percent for high educated workers. Figure 1.7 also shows the opposite case, that is, allowing substitutability ρ_τ to grow as the estimation suggests, but fixing skill efficiency α_τ at its median value

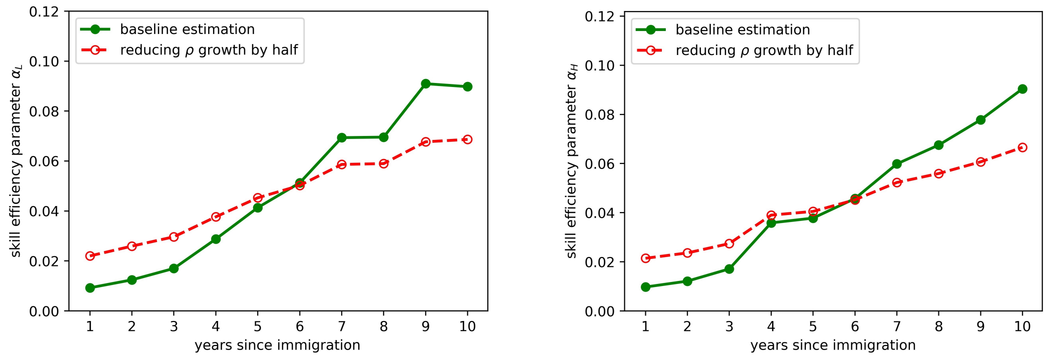
[14] As before, we set $\sigma = 0.5$, and estimate A and β to have equation 1.3 match the empirical average log wage of natives.

of 0.046 for all τ for both skill groups. The downward sloping dashed green profile of the predicted marginal products of labor for this case illustrates the competition effect as immigrants become increasingly substitutable to native workers.

Figure 1.8: Estimation bias.

(a) Counterfactual $\{\alpha_\tau\}_{\tau=1}^{T}$ for $s = l$ (less than college)

(b) Counterfactual $\{\alpha_\tau\}_{\tau=1}^{T}$ for $s = h$ (with college)



Note: The figure shows the actual skill efficiency profile $\boldsymbol{\alpha}$ estimated under the full model (solid line) versus a counterfactual profile (dashed line) that would be obtained if substitutability growth was half as steep, rotated around the median value of ρ_τ. The exercise is performed for low educated workers (left panel) and high educated workers (right panel).

Existing studies of immigrant assimilation interpret observed wage progress directly as immigrants' skill accumulation. We show that increasing substitutability disrupts this mapping, so that a partial equilibrium view under-predicts the skill efficiency growth for immigrants. To illustrate this, we re-estimate the sequence of skill efficiency parameters $\boldsymbol{\alpha}$ while fixing $\boldsymbol{\rho}$ at values that grow at half the rate of our actual estimates.[15] Figure 1.8 shows that under the (wrong) assumption of slower assimilation in terms substitutability, skill progress is strongly underestimated. The graph shows that the estimated skill efficiency profile obtained under the restricted

[15]We keep as fixed the median value of ρ_τ, i.e. at $\tau = 6$, and reduce the steepness of the substitutability parameter profile by reducing by half the wedge between each value of $\rho_{\tau\neq 6}$ and ρ_6. With this slower substitutability growth, we re-estimate the sequence of skill efficiency parameters $\boldsymbol{\alpha}$ which best fit the empirical wage data.

model (dashed line) is flatter than in the full model where the substitutability dimension of assimilation is fully accounted for (solid line), illustrating the shortcomings of partial analyses that do not account of this dimension of assimilation.

1.5 Robustness

In this section, we demonstrate the robustness of our estimated structural parameter series $\boldsymbol{\alpha}$ and $\boldsymbol{\rho}$ to various modifications in the definitions and preparations of the data used. Since Figure 1.4 shows little difference in the estimated parameter sequences for low- and high-educated workers, in what follows we pool education groups. In this way, we offer a clear and simple benchmark for the robustness checks that will be performed.

Without the upper nest distinguishing workers of different education levels, the underlying model in this section thus assumes a production function

$$\begin{aligned} Y_t \;=\; & A_t\Bigg((\alpha_1 M_{1,t})^{\rho_1} + \Bigg((\alpha_2 M_{2,t})^{\rho_2} + \Bigg(\quad \ldots \quad \Bigg((\alpha_{T-1} M_{T-1,t})^{\rho_{T-1}} \\ & + \Big((\alpha_T M_{T,t})^{\rho_T} + (\alpha_N N_t)^{\rho_T}\Big)^{\frac{\rho_{T-1}}{\rho_T}}\Bigg)^{\frac{\rho_{T-2}}{\rho_{T-1}}} \quad \ldots \quad \Bigg)^{\frac{\rho_2}{\rho_3}}\Bigg)^{\frac{\rho_1}{\rho_2}}\Bigg)^{\frac{1}{\rho_1}}, \end{aligned} \tag{1.6}$$

Following the same steps as in Section 1.2, log wage growth for immigrants, and the relative log wage between long-term immigrants and natives now become

$$\begin{aligned} \ln\frac{w_{\tau+1,t}}{w_{\tau,t}} \;&=\; \rho_{\tau+1}\ln\alpha_{\tau+1} - \rho_\tau \ln\alpha_\tau + (\rho_\tau - \rho_{\tau+1})\ln F_{\tau+1,t} \\ &\quad +(1-\rho_\tau)\ln M_{\tau,t} - (1-\rho_{\tau+1})\ln M_{\tau+1,t} \\ \ln\frac{w_{T,t}}{w_{N,t}} \;&=\; \rho_T\ln(\alpha_T) - \ln\left(1 - \sum_{k=1}^{T}\alpha_k^{\rho_k}\right) - (1-\rho_T)\ln\frac{M_{T,t}}{N_t}. \end{aligned}$$

This yields a reduced system of equations, which can be used to estimate pooled parameter sequences $\{\alpha_\tau\}_{\tau=1}^{T}$ and $\{\rho_\tau\}_{\tau=1}^{T}$. These estimates are listed in Table 4 in the Appendix. In Appendix G we also show the implied elasticities of substitution, the dynamic effect on natives' wages, as well as the counterparts to Figures 1.7 and 1.8 in the previous section. In what follows, we scrutinize the robustness of these parameter estimates.

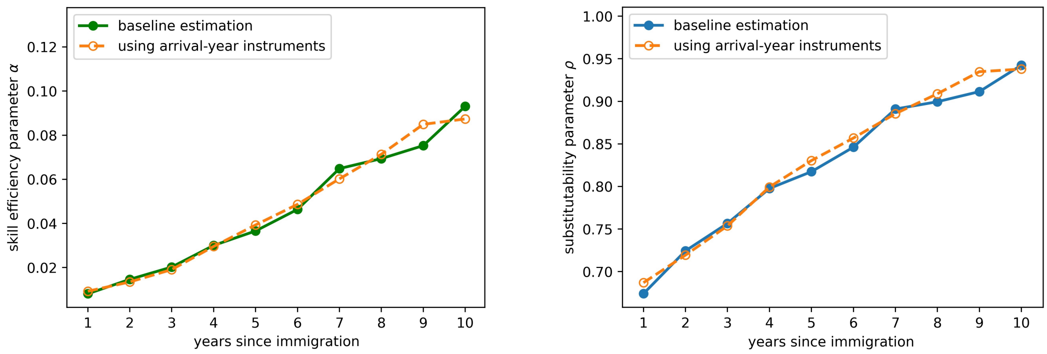
1.5.1 Timing of Instrumental Variables

The instruments underlying our results vary both across the year immigrants are surveyed and the year they arrived in the U.S. Identification in fact requires only the latter source of variation.[16] In this case, economic conditions in the country of origin and price-level ratios in the year of migration predict the scale and composition of a given immigrant cohort in terms of migrants' expected length of stay in the U.S. This composition will later determine how many migrants of the initial arrival cohort stay longer term, and how many will leave. As such, instruments $\boldsymbol{z}_{t-\tau}$, which only contain information for the year of migration $t-\tau$, can still predict later stocks of migrants who arrived in that year and are still in the U.S. Columns (4)-(6) of Table 1.2 shows the predictive power of this variant of the instruments for the stocks of migrants residing in the U.S. Figure 1.9 directly contrasts the estimated parameter values resulting under the two different sets of instruments, and illustrates their robustness. We find that the specified timing of the instruments matters somewhat for long-term immigrants, but has virtually no effect on the estimated assimilation profiles during the first years in the U.S., on which we focus.

[16] We find that variation along the other dimension—survey years—alone yields a weak first stage, and thus do not explore this further.

Figure 1.9: Different timing of instruments.

(a) Skill efficiency $\boldsymbol{\alpha}$ (b) Substitutability $\boldsymbol{\rho}$



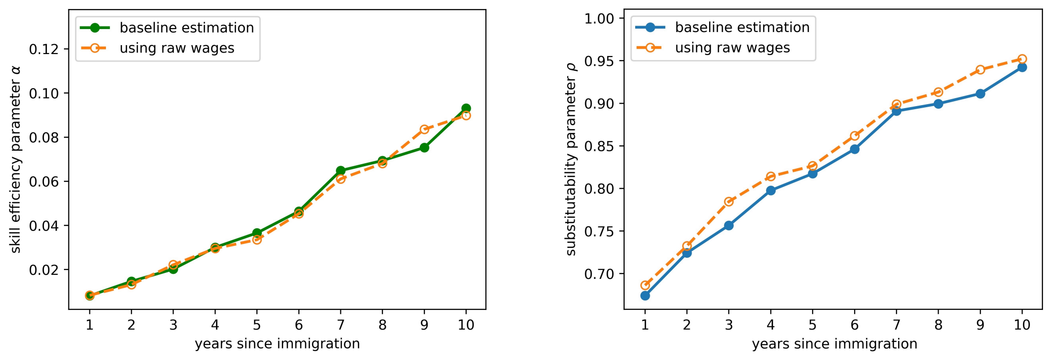
Note: Panel (a) shows skill efficiency parameter series $\boldsymbol{\alpha}$, panel (b) substitutability parameter series $\boldsymbol{\rho}$. The figure shows the parameter estimates obtained under two sets of instruments defined with different timing. Solid lines indicate our main estimates, dashed lines the parameter values obtained if immigrant stocks are instrumented using origin country GDP/capita growth and price-level ratios in the year of arrival only.

1.5.2 No Cleaning of Demographic and Cohort Effects

Our model considers a national production function to which all workers in the economy contribute. We thus do not assume that workers of different individual characteristics compete on segmented labor markets. To account for compositional changes in the labor force, and in particular for variation in the productivity or substitutability of different immigrant cohorts, we first regress wages on observed individual characteristics, including the year of arrival (see Section 1.2.2). We then predict wages for each category of immigrants or natives net of these characteristics and cohort components, which we use in the estimation of the structural model. To show that our results do not hinge on this procedure, we re-estimate the series of productivity and substitutability parameters on the raw data, maintaining the compositional variation in terms of individual characteristics. Figure 1.10 shows the result of this estimation. Again, the general path is very similar to the baseline estimates. The main difference is a higher estimated substitutability for $\tau \leq 10$. A potential explanation for this is that wages by less educated men, who are over-represented among short-term migrants (see Table 2.1), respond more strongly to labor supply changes. On the other hand, we estimate a lower substitutability parameter ρ_T for the rather heterogeneous group of long-term immigrants when observables are not controlled for. Neither of these differences, however, is highly significant.

Figure 1.10: No demographic controls or cohort effects.

(a) Skill efficiency $\boldsymbol{\alpha}$ (b) Substitutability $\boldsymbol{\rho}$



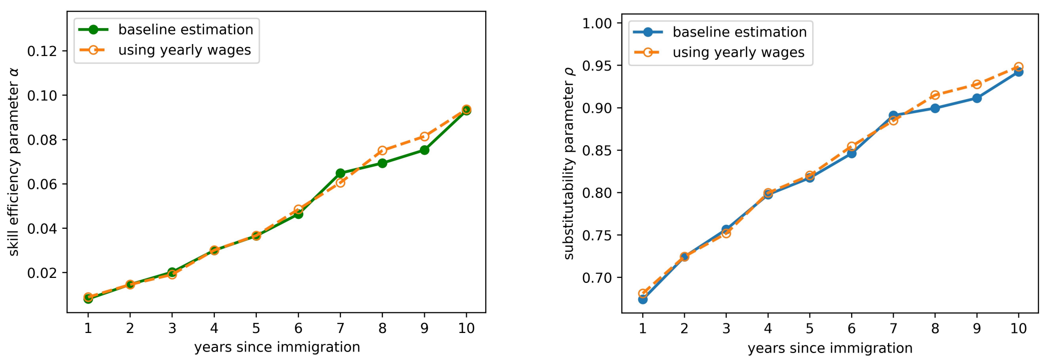
Note: Panel (a) shows skill efficiency parameter series $\boldsymbol{\alpha}$, Panel (b) substitutability parameter series $\boldsymbol{\rho}$. Solid lines indicate estimates in the baseline estimation after controls are taken out, dashed lines the parameter estimates obtained using unadjusted wages (only deflated).

1.5.3 Using Yearly Wages

As previously explained, we use hourly wages, which we compute from annual earnings observed in the data, divided by the number of weeks worked per year and the number of hours worked per week. While hourly wages are our preferred outcome measure, determined by the marginal product of labor derived from the model, dividing by weeks and hours may raise measurement error. In order to check whether this potential source of bias plays a role, for instance because the precision with which working time is reported varies with time spent in the U.S., we re-estimate the model with annual earnings exactly as reported (only deflated to the base year 1999). Figure 1.11 shows again that the general path is very similar to the baseline result.

Figure 1.11: Using annual rather than hourly wages.

(a) Skill efficiency $\boldsymbol{\alpha}$ (b) Substitutability $\boldsymbol{\rho}$



Note: Panel (a) shows skill efficiency parameter series $\boldsymbol{\alpha}$, panel (b) substitutability parameter series $\boldsymbol{\rho}$. Solid lines indicate estimates in the baseline estimation with hourly wages, dashed lines show the parameter estimates obtained using the data for yearly wages as reported by respondents.

1.5.4 Alternative Nesting: Hispanic

Our estimates in Figure 1.4 suggest little differences in terms of channels contributing to wage assimilation across education groups, likely the most important dimension across which workers differ in terms of substitutability. The flexibility of our model as defined as Section 1.2 allows us to consider alternative characteristics for which the evolution of skills and substitutability differs with years since immigration.

An important feature of the U.S. labor market is the large share of natives of Hispanic origin, since immigrants from Latin American countries jointly form a major part of U.S. immigration. We therefore redefine the model distinguishing Hispanic vs. non-Hispanic workers as the outer nest. That is, the production function (1.1) becomes

$$Y_t = A_t \left(\beta L^{\sigma}_{Hisp,t} + (1-\beta) L^{\sigma}_{nonHisp,t}\right)^{\frac{1}{\sigma}},$$

where $L_{Hisp,t}$ and $L_{nonHisp,t}$ aggregate natives and immigrants with different lengths of stay as in equation (1.2), however distinguishing both natives and immigrants with and without Hispanic origin. The estimates resulting from this estimation are depicted in Figure 1.12. Relative to their level of skill at arrival, immigrants of Hispanic origin accumulate skills ($\boldsymbol{\alpha}$) at a higher rate. similarly, although less substitutable with natives of Hispanic origin upon arrival, the substitutability of immigrants ($\boldsymbol{\rho}$) grows at a higher rate than for non-Hispanic workers. Note that this comparison is *within* the groups of Hispanic and non-Hispanic workers, respectively. That is, the higher substitutability of Hispanic immigrants after five years in the United States is with respect to natives of Hispanic origin, which Panel (b) of Figure 1.12 compares to the substitutability of non-Hispanic immigrants to non-Hispanic natives.

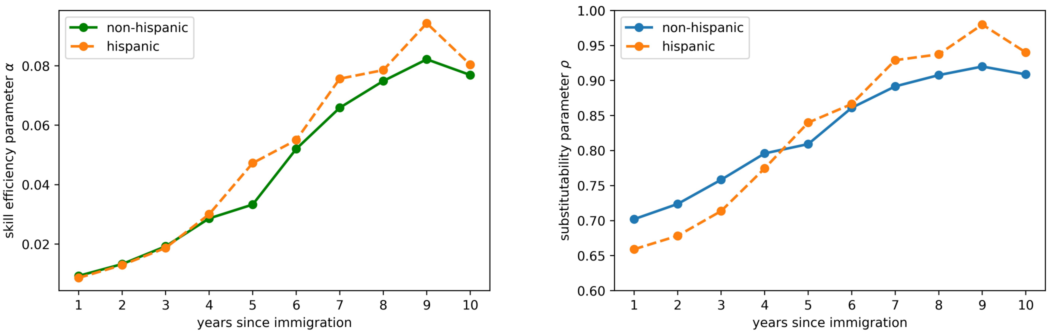
Yet, these differences are statistically insignificant. This exercise, however, does show both the flexibility of the model, and the broad robustness of the findings across an ethnic lines.

1.5.5 Heterogeneity

Our paper aims to provide a novel understanding of the assimilation of immigrants in the United States at the national level. This national aggregation avoids issues of endogenous sorting into regional labor markets. In this subsection, however, we explore the effects of heterogeneity across parts of the country. We test this

Figure 1.12: Alternative model distinguishing workers with and without Hispanic origin in outer nest.

(a) Skill efficiency $\boldsymbol{\alpha}$ (b) Substitutability $\boldsymbol{\rho}$



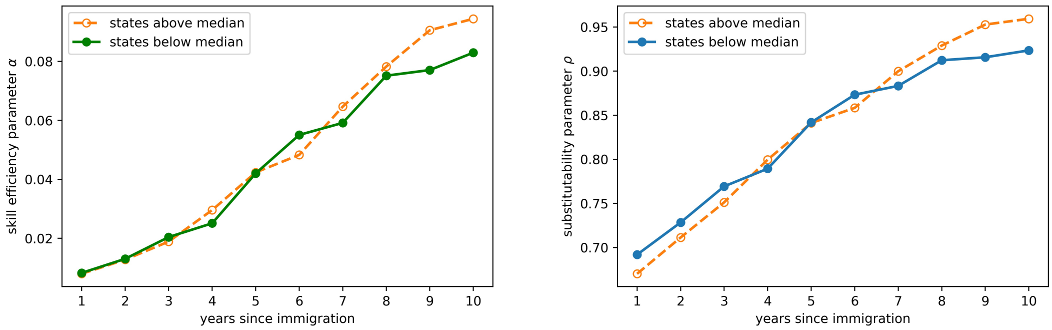
Note: Panel (a) shows skill efficiency parameter series $\boldsymbol{\alpha}$, panel (b) shows substitutability parameter series $\boldsymbol{\rho}$. The figure shows the parameter estimates resulting from the model using Hispanic and non-Hispanic as the first nest.

heterogeneity first by focusing again on the Hispanic origin. However, now we consider the concentration of Hispanic natives in the population. We re-estimate the model separately for states with above and below median shares of natives of Hispanic origin in the population. Figure 1.13b indicates a steeper assimilation in terms of substitutability in states with higher shares of the population being of Hispanic origin. Note that among very recent immigrants in our sample, the share of immigrants from Asia is relatively high (Table 2.1). In line with this, we find a slightly lower substitutability for the most recent immigrants (with $\tau \leq 3$) in states with a higher concentration of Hispanics. Other groupings of states yielded similar results.

As we show in the introduction, immigrants upon arrival tend to be concentrated in production and manual task intensive jobs. Assimilation profiles thus may differ in states where the economy is more geared towards manufacturing as opposed to those more dependent on the tertiary sector. Specifically, a strong manufacturing sector implies more jobs that new immigrants can enter early on. On the other

Figure 1.13: Heterogeneity: By share of Hispanic population.

(a) Skill efficiency $\boldsymbol{\alpha}$ (b) Substitutability $\boldsymbol{\rho}$



Note: Panel (a) shows skill efficiency parameter series $\boldsymbol{\alpha}$, panel (b) shows substitutability parameter series $\boldsymbol{\rho}$. The figure shows the parameter estimates obtained by dividing states into those with above median Hispanic shares in the population, and those with lower shares.

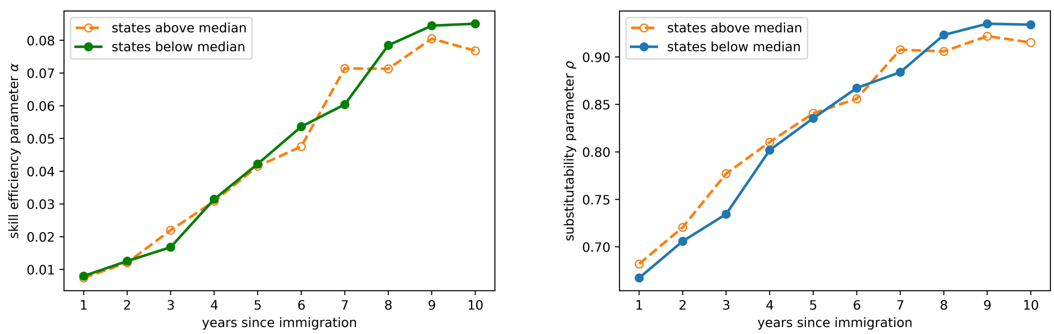
hand, a transition into more communication intensive jobs that are otherwise held by natives, and the ensuing rise in immigrants' substitutability, may progress more rapidly in states with a larger service sector. We therefore distinguish states with above- and below-median shares of manufacturing jobs. Figure 1.14 shows that substitutability for the most recent immigrants is slightly higher in states with a relatively high share of manufacturing jobs. On the other hand, assimilation in terms of substitutability is also flatter in these states. Differences are, however, small and statistically insignificant. Yet, taken at face value, these point estimates tell a plausible story about the integration of immigrants into the host economy.

1.6 Conclusion

In this paper, we disentangle two core elements of immigrants' assimilation into the host country's labor market: the increase in immigrants' skills and the substitutability of immigrants with natives and earlier immigrants. In accounting for

Figure 1.14: Heterogneity: By share of manufacturing jobs.

(a) Skill efficiency $\boldsymbol{\alpha}$ (b) Substitutability $\boldsymbol{\rho}$



Note: Panel (a) shows skill efficiency parameter series $\boldsymbol{\alpha}$, panel(b) substitutability parameter series $\boldsymbol{\rho}$. The figure shows the parameter estimates obtained by dividing states between those with above median shares of manufacturing jobs, and those with lower shares.

substitutability, we emphasize an aspect of assimilation that the literature examining immigrant wage assimilation has mostly neglected. Methodologically, our treatment of immigrants' wages as an equilibrium outcome draws on frameworks used for estimating labor market effects of immigration. Studies in this field have focused on impact upon arrival, and we extend this in several directions. First, we allow for a changing substitutability over time as immigrants assimilate. Second, we provide novel estimates on the substitutability not only between immigrants and natives, but also among immigrants of different host country tenure. Third, our model allows us to trace out dynamic wage effects as immigrants assimilate into the host country.

We account for endogenous cohort sizes, which are determined through immigration and out-migration, by leveraging variation in economic conditions in countries of origin. We estimate intuitive patterns of assimilation in both productivity and substitutability. Immigrants who have just arrived are imperfect substitutes, with an elasticity of substitution with respect to natives of about 3, but reach elasticities of about 20 within ten years in the U.S. that is slightly higher for less educated workers. The steep increase in substitutability during the first few years after arrival and the intensified competition it entails imply that wage growth strongly under-predicts the true gains in skill formation by immigrants.

These results highlight the importance of estimating productivity and substitutability jointly when examining immigrant wages and their assimilation in the host economy. While our empirical application focuses on the U.S., it has profound implications for different institutional contexts. For instance, where international migration takes place under the roof of a supranational body like the European Union, a coordination by national governments to ease the accreditation of foreign qualifications will directly affect the elasticity of substitution between immigrants and other workers and thus have important effects on immigrant assimilation.

A A Model with Capital

Suppose production was characterized by the production function laid out in Section 1.2, however augmented with capital. Specifically, assume output for the national economy is given by:

$$Y = AK^{\kappa}\left(\sum_{s=1}^{S}\beta_s L_s^{\sigma}\right)^{\frac{1-\kappa}{\sigma}},$$

where we omit calendar time t for ease of notation. Maintaining the same notation as in Section 1.2, the log wage for the most recent immigrants is given by

$$\begin{aligned}\ln w_{1,s} &= \frac{\sigma}{1-\kappa}\ln(AK^{\kappa}) + \left(1-\frac{\sigma}{1-\kappa}\right)\ln Y + \ln(1-\kappa)\\ &\quad + \ln\beta_s + (\sigma-\rho_{1,s})\ln L_s + \rho_{1,s}\ln\alpha_{1,s} - (1-\rho_{1,s})\ln M_{1,s}.\end{aligned}$$

Log wages for earlier immigrants are

$$\begin{aligned}\ln w_{\tau,s} &= \frac{\sigma}{1-\kappa}\ln(AK^{\kappa}) + \left(1-\frac{\sigma}{1-\kappa}\right)\ln Y + \ln(1-\kappa)\\ &\quad + \ln\beta_s + (\sigma-\rho_{1,s})\ln L_s + \rho_{\tau,s}\ln\alpha_{\tau,s}\\ &\quad -(1-\rho_{\tau,s})\ln M_{\tau,s} + \sum_{k=2}^{\tau}(\rho_{k-1,s}-\rho_{k,s})\ln F_{k,s}.\end{aligned}$$

Finally, log wages for native-born workers are

$$\begin{aligned}\ln w_{N,s} &= \frac{\sigma}{1-\kappa}\ln(AK^{\kappa}) + \left(1-\frac{\sigma}{1-\kappa}\right)\ln Y + \ln(1-\kappa)\\ &\quad + \ln\beta_s + (\sigma-\rho_{1,s})\ln L_s + \ln(1-\sum_{k=1}^{T}\alpha_{k,s}^{\rho_{k,s}})\\ &\quad -(1-\rho_{T,s})\ln N_s + \sum_{k=2}^{T}(\rho_{k-1,s}-\rho_{k,s})\ln F_{k,s}.\end{aligned}$$

Intuitively, the *level* of wages increases with capital for all groups of workers. Yet, holding everything else constant, we obtain the same expressions for log wage *growth*, on which the estimation is based, as in the model without capital in Section 1.2. Specifically, the log wage growth for immigrants who have been in the country for τ years is given by

$$\begin{aligned}\ln \frac{w_{\tau+1,s}}{w_{\tau,s}} &= \rho_{\tau+1,s} \ln \alpha_{\tau+1,s} - \rho_{\tau,s} \ln \alpha_{\tau,s} + (\rho_{\tau,s} - \rho_{\tau+1,s}) \ln F_{\tau+1,s} \\ &\quad +(1-\rho_{\tau,s}) \ln M_{\tau,s} - (1-\rho_{\tau+1,s}) \ln M_{\tau+1,s},\end{aligned}$$

whereas the log relative wage for long-term immigrants as compared to natives is

$$\ln \frac{w_{T,s}}{w_{N,s}} = \rho_{T,s} \ln \left(\alpha_{T,s}\right) - \ln \left(1 - \sum_{k=1}^{T} \alpha_{k,s}^{\rho_{k,s}}\right) - (1-\rho_{T,s}) \ln \frac{M_{T,s}}{N_s}.$$

B Data Appendix

We combine U.S. Census data (for the census year 2000) and American Communities Survey (for the years 2001-2018). We use the IPUMS database where variables have been already harmonized in order to overcome different variable classifications, breaks in the series, and other possible issues arising from the use of different survey types (Ruggles et al., 2020).

From these datasets, we use the following variables: census year, person's sampling weight, annual earnings, average number of hours worked per week, number of weeks worked in the previous year, a consumer price index deflating earnings to 1999-prices, the "ADJUST" variable recommended by the ACS given the year-round nature of the survey, and for immigrants their year of arrival. Additionally, we collect the following demographic variables to be used in robustness checks that eliminate compositional changes in terms of observed individual characteristics: for the entirety of the population (both immigrants and natives) sex, age, and education level in years of schooling completed; and for immigrants additionally the country of origin. A separate variable indicating whether the individual is Hispanic or not is used in the heterogeneity exercise of the robustness section. Time spent in the U.S. is determined as the difference between the survey year and the year of immigration. Finally, the data used in the introduction for the evolution of occupations was taken from using the "OCCSOC" and "OCC1990" variables of the ACS and the O*NET database classifications as described by Peri and Sparber (2009).

The notion that immigrants move into communication intensive jobs due to accumulation of an essential skill over time spent in the host country is supported by looking as the language knowledge of immigrants. Figure 15 shows, for each year of host country tenure, the fraction of immigrants who speak English very well, well, not well, or not at all. As their English skills improve, immigrants transition across these categories, with an increasing share of immigrants speaking English very well and a decreasing share of immigrants not speaking English at all. The intermediate categories (“well" and “not well") experience large shares of both inflows from lower and outflows to higher categories of language knowledge.

For the instruments, we use data from the World Bank’s World Development Indicators (WDI) and the OECD’s statistical database. Specifically, we use 1990-2018 series of real and nominal exchange rates and GDP per capita growth in the countries of origin of the largest immigrant groups: Mexico, China, India, the Philippines and Vietnam. To control for global economic trends, we further control for average GDP per capita growth in OECD countries.

The distribution of immigrants by years spent in the United States is shown in Figure 16, separately for data collected before and after 2005. Immigrants having

(a) Evolution of language ability

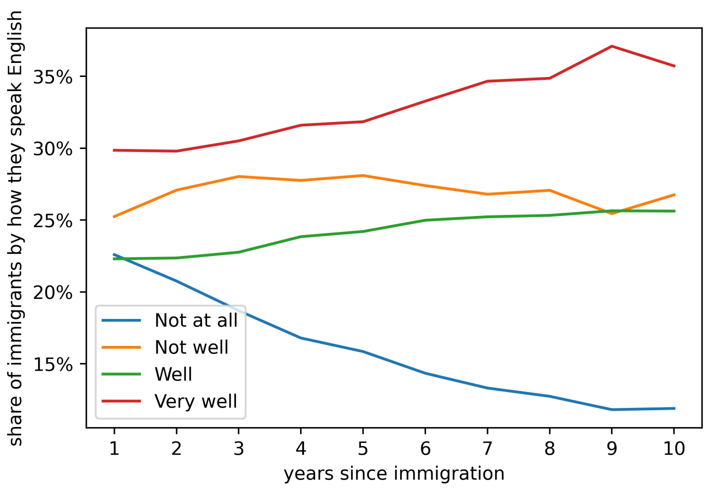


Figure 15: Immigrants' language assimilation over time. The figure plots for each year since arrival the percentage of immigrants by their English language ability. The plot is constructed from the IPUMS databases using the variable "SPEAKENG".

been in the U.S. for at least 40 years are grouped in the last bar. The gradual decline by years of tenure reflects a combination of out-migration and changes in the size of inflows over time. The latter is indicated also by the difference between more recent surveys (in blue) and the older data (in orange).

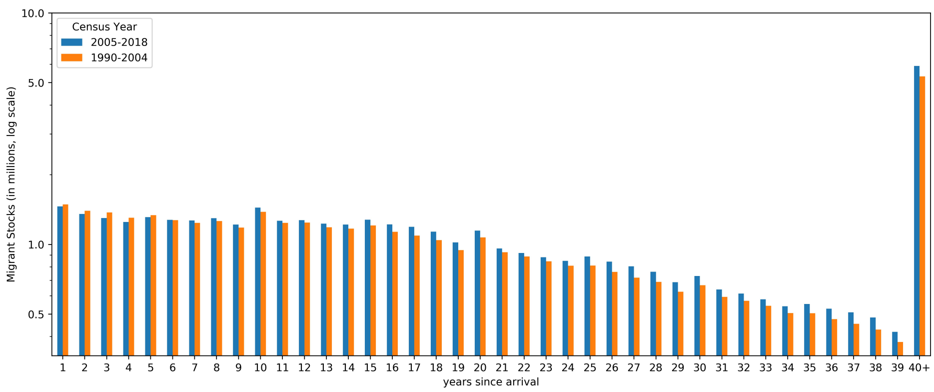


Figure 16: Distribution of immigrants by time spent in the U.S., separately for survey records until 2004 (in orange), and for surveys from 2005 onwards (in blue).

To account for compositional changes of the workforce in terms of observed characteristics, and in particular for cohort effects across waves of immigration, we eliminate these observables by regressing wages on full sets of cohort and years since immigration effects, as well as on an individual's age, sex, education and country of origin. That is, rather than using raw wages $w_{\tau,t}$ in the estimation of the model, we use wages $\tilde{w}_{\tau,t}$ net of cohort and other compositional effects in the construction of moment conditions (1.5).

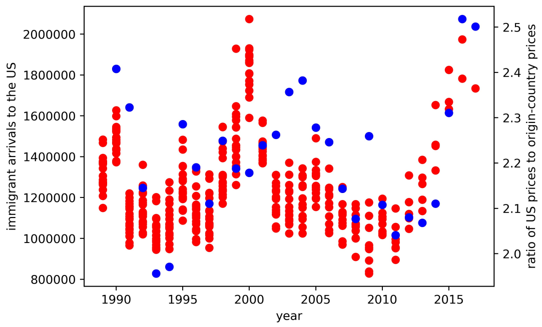
For this goal we use a two-step approach. In the first step, we consider both natives and immigrants, and run a weighted least-squares regression of wages on indicators for age, sex and education, using the weight variable provided in the census. By adding the residuals to the average wage we obtain wages net of compositional differences between immigrants and natives. In a second step, we focus only on immigrants, and do the same procedure controlling for country of immigration, year of immigration, and years spent in the United States. This time we add both the residuals and the effect of each indicator for years spent in the U.S. This way we preserve the variation due to assimilation.

C Estimation Details

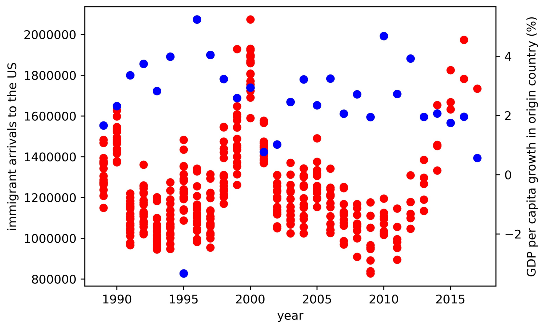
C.1 Instrumentation

Panels (a) and (b) of Figure 17 shows the strong co-movement of immigration flows with the instruments. The plots show the number of immigrants observed in our data by year of arrival (in red, with scale on the left axis), together with the realization of the instruments (in blue, scaled on the right axis). Figure 17a shows a strong positive correlation between relative price levels and immigrant numbers: a higher purchasing power of U.S.-dollars abroad makes the U.S. a more attractive destination for (temporary) migration. Growth of GDP per capita in origin countries, instead, exhibits a negative relation to arrivals: as countries of origin experience worse economic conditions, more workers will choose to emigrate and stay abroad (Figure 17b). Panels (c) and (d) visualize the first stage prediction relative to the observed actual stocks of immigrants.

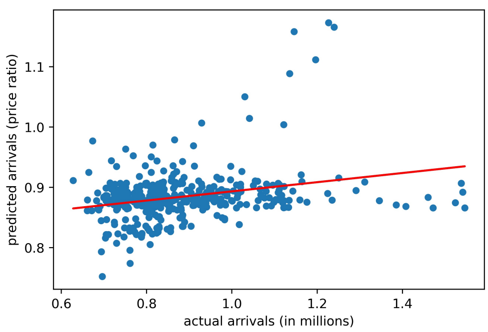
(a) U.S.-origin price ratio: Raw data



(b) GDP/capita growth at origin: Raw data



(c) U.S.-origin price ratio: Predicting migrant stocks



(d) GDP/capita growth at origin: Predicting migrant stocks

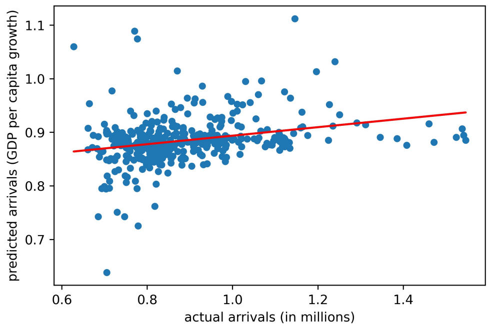


Figure 17: Instruments for cohort sizes over time. The top two panels show the stocks of immigrants who in different survey waves report having arrived in a given year (red dots and left axes) together with (a) the price level ratio of the U.S. relative to the weighted basket of origin countries, and (b) weighted GDP per capita growth in the five major countries of origin. Panels (c) and (d) show the predictive power of two instruments for the stocks of immigrants (each dot indicating a year of arrival-survey year combination). Sources: ACS 2001-2018, U.S. Census 2000, OECD Stats and World Bank WDI.

C.2 Structural Estimation

To verify the continuity of the objective function at each parameter estimate, we plot for each parameter the sum of squared residuals of the moment conditions as a function of a grid of possible parameter values. Figures 20 and 21 plot the objective function against each parameter. To further ensure that our moment conditions indeed identify the vector of structural parameters, Figure 19 visualizes the gradient matrix of each theoretical moment with respect to the vector of parameters, with darker shades indicating steeper gradients. It shows that all parameters are well identified. In particular, there is no white (zero gradient) row, and columns are not collinear.

To illustrate the contribution of variation in the workforce composition over time to identification, we compute the impact of changes in population stocks on wage growth. Specifically, for each $\tilde{\tau}$, we raise the stock of all migrants M_τ with $\tau \neq \tilde{\tau}$ as well as of natives by a given percentage, and compute the predicted wage growth $\ln \frac{w_{\tau+1}}{w_\tau}$ and $\ln \frac{w_N}{w_T}$, as predicted by equations 1.3 and 1.4 for the estimated parameters displayed in Figure 1.4. Figure 18 shows how wage growth changes when stocks are raised by respectively by 10% (orange dotted line), by 20% (blue solid line), and 50% (green dashed line). This responsiveness in wage growth for different levels of immigrant and native populations confirms their contribution to identification.

To further support identification of the model's structural assimilation parameters by the moments contained in the estimation criterion in equation (1.5), Figure 19 visualizes the gradient matrix of each moment with respect to each parameter. It shows that moments implied by the model indeed respond to changes parameters, and that there is no collinearity.

(a) Low educated (b) High educated

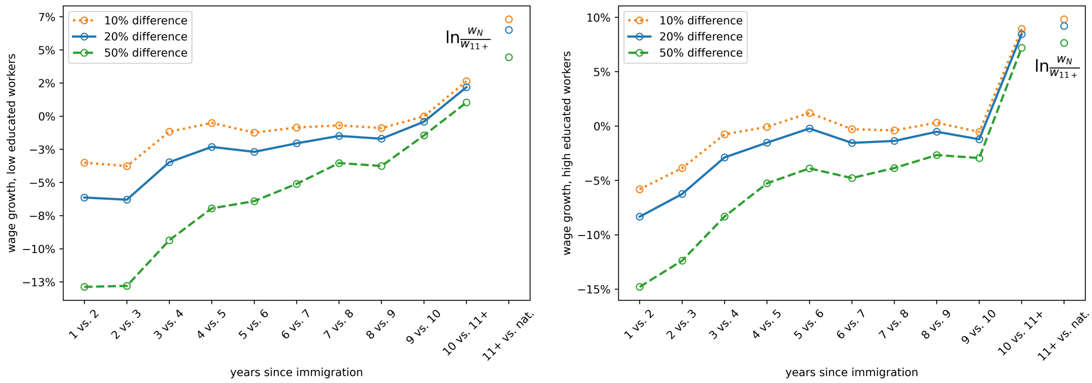


Figure 18: Impact of population stocks on predicted wage growth. The figure displays differences in wage growth predicted for different stocks of immigrants and natives. For each wage growth step on the horizontal axis, the figure shows the effect of raising all other stocks by 10% (orange dotted line), 20% (blue solid line), and 50% (green dashed line), respectively.

D Estimation using only men

Much of the literature, including the seminal papers by Chiswick (1978), Borjas (1985), Hu (2000) and Lubotsky (2007), restrict their analysis samples to men when estimating immigrant assimilation.[17] Since wages in our framework wages are an equilibrium outcome, all labor force changes, including through the immigration of women, must be accounted for. To check whether the inclusion of women in our analysis drives the results, we re-estimate the model using men only. Figure 22 shows the parameter estimates for this case. We do not find any relevant differences with respect to the baseline estimation.

[17]See, for instance, Long (1980), Schoeni (1998), Butcher and DiNardo (2002), and Blau et al. (2011) for analyses of the earnings of female immigrants.

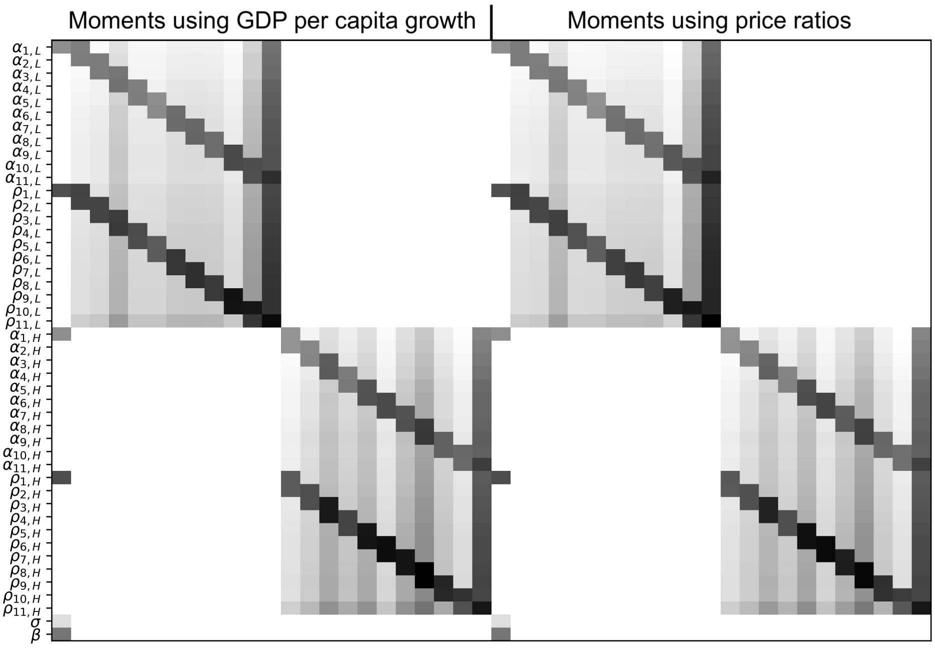


Figure 19: Sensitivity of moments to the model parameters. The figure illustrates the gradient matrix of moments used for identification with respect to parameters $\boldsymbol{\alpha}$ and $\boldsymbol{\rho}$. Darker shades indicate a stronger sensitivity of moments to changes in parameters.

E Parameter Estimates

In Table 3, we list the point estimates for the parameter values obtained under different instruments, and which were shown in the paper graphically.

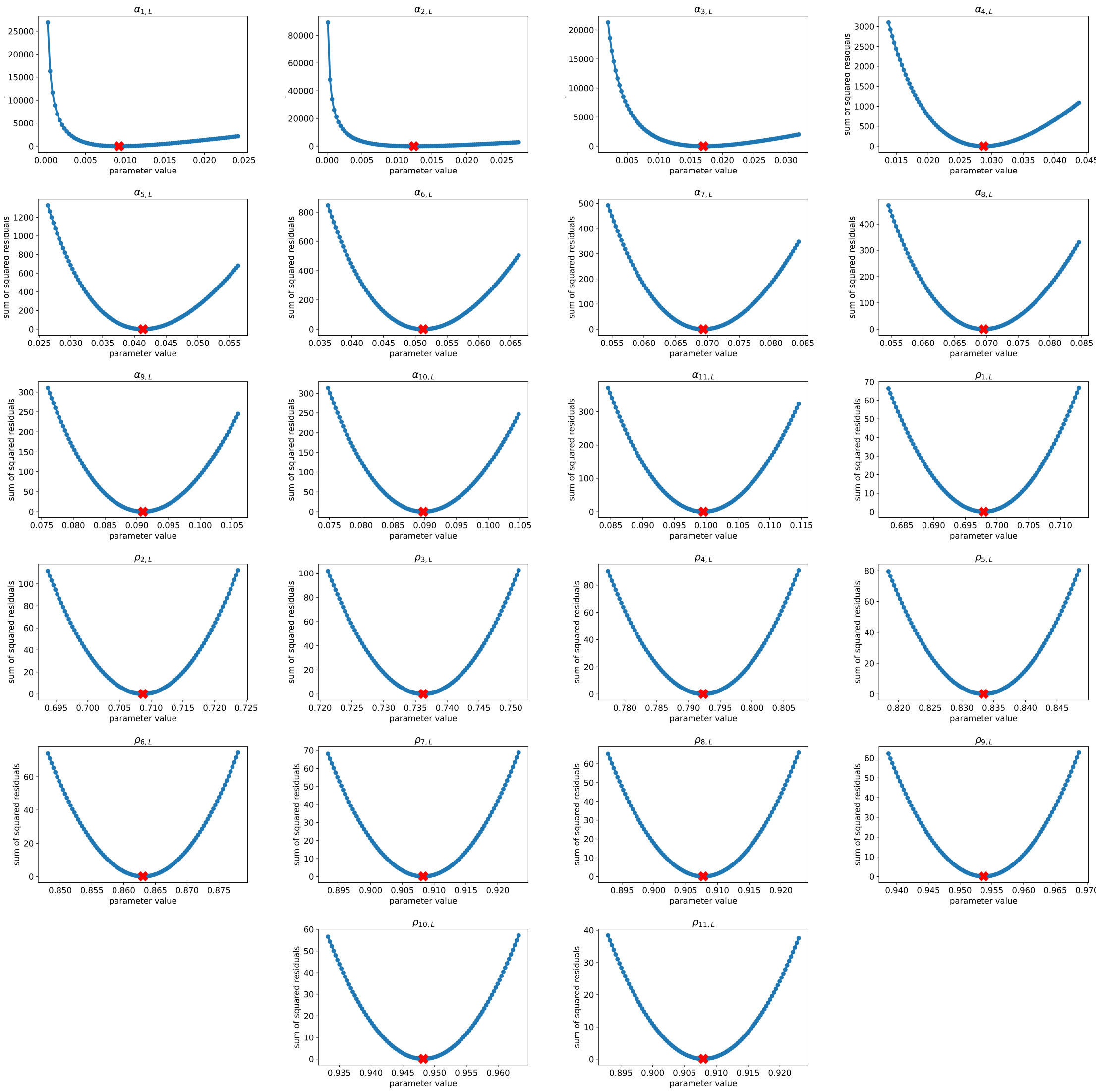


Figure 20: Parameter values pertaining to low educated workers, plotted against the corresponding estimation criterion. The minimizing value is shown in red.

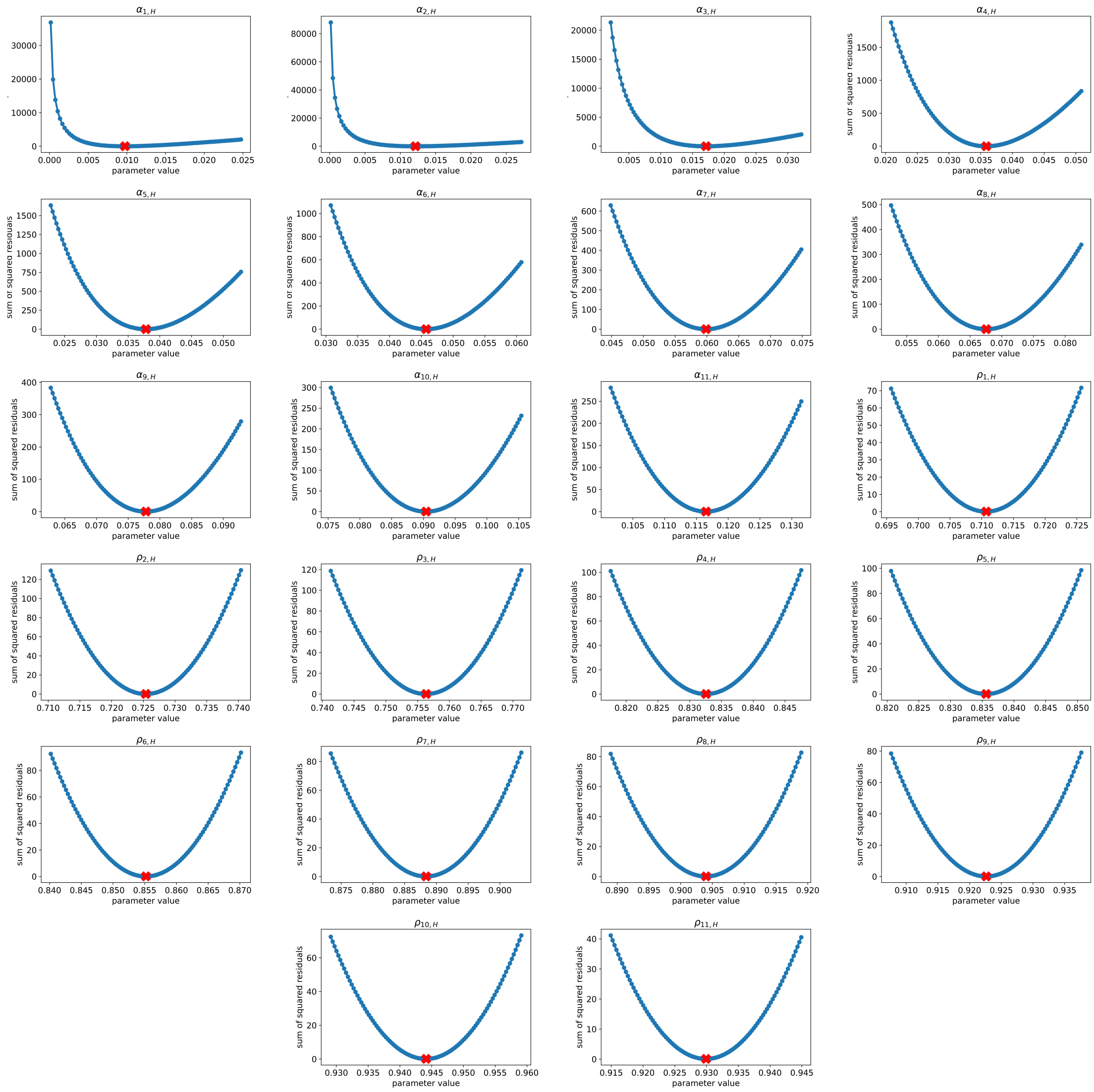


Figure 21: Parameter values pertaining to high educated workers, plotted against the corresponding estimation criterion. The minimizing value is shown in red.

(a) Skill efficiency $\boldsymbol{\alpha}$ (b) Substitutability $\boldsymbol{\rho}$

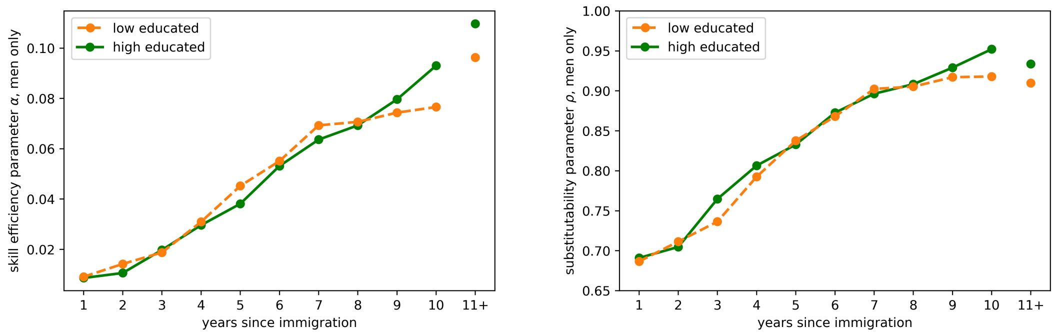


Figure 22: Panel (a) shows skill efficiency parameter series $\boldsymbol{\alpha}$, panel (b) shows substitutability parameter series $\boldsymbol{\rho}$. The figure shows the parameter estimates obtained using only men in the estimation.

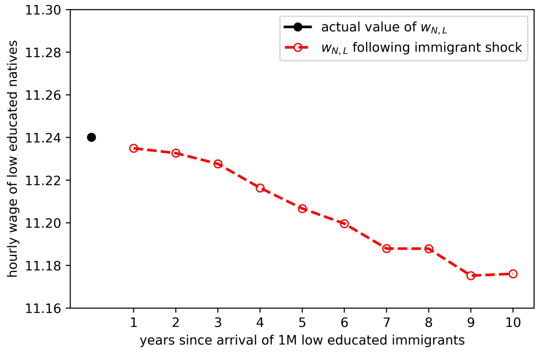
Table 3: Point estimates for parameters $\boldsymbol{\alpha}$ and $\boldsymbol{\rho}$.

τ	low educated		high educated	
	$\hat{\alpha}$	$\hat{\rho}$	$\hat{\alpha}$	$\hat{\rho}$
1	0.009	0.698	0.010	0.711
2	0.012	0.709	0.012	0.725
3	0.017	0.736	0.017	0.756
4	0.029	0.792	0.036	0.833
5	0.041	0.833	0.038	0.836
6	0.051	0.863	0.046	0.855
7	0.069	0.908	0.060	0.888
8	0.070	0.908	0.068	0.904
9	0.091	0.954	0.078	0.923
10	0.090	0.948	0.090	0.944
11+	0.100	0.908	0.117	0.930

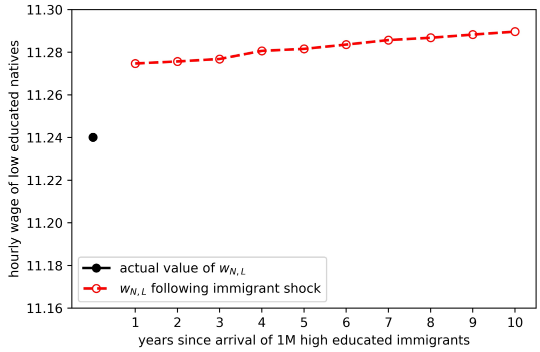
The table lists point estimates shown in Figure 1.4 of the paper.

F Dynamic Wage Impact by Immigrants' Skill

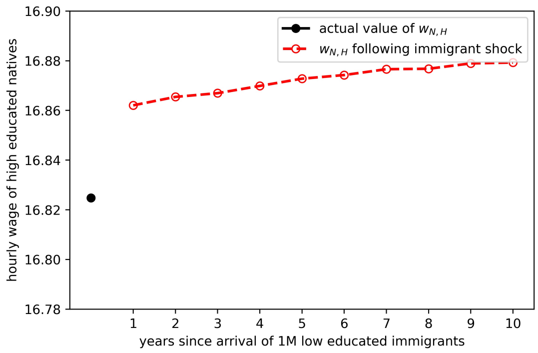
(a) Low educated natives wage after low educated immigrant shock



(b) Low educated natives wage after high educated immigrant shock



(c) High educated natives wage after low educated immigrant shock



(d) High educated natives wage after high educated immigrant shock

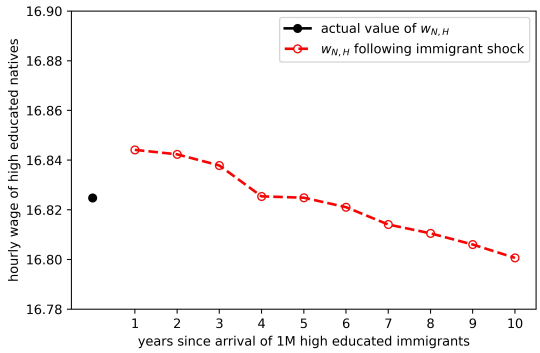


Figure 23: Effect on natives' wages (in USD) of one additional Million of immigrants, as they progress in time through years spent in the U.S., experiencing changes in both substitutability and skill efficiency. The first full dot is the actual average native wage; the dashed line shows the wage response to a one-million immigrant shock by years since arrival. The top two panels show the effect of one million low- (left) and high educated (right) immigrants on low educated natives. The bottom two panels show the effect of one million low- and high educated immigrants on high educated natives.

G Pooled Estimates

In Table 4, we list the point estimates for the parameter values obtained in the pooled skill-group estimation (see Section 1.5) and under different instruments, and which were shown throughout the robustness section graphically.

Table 4: Point estimates for parameters $\boldsymbol{\alpha}$ and $\boldsymbol{\rho}$ with pooled estimates.

τ	baseline		alternative instruments		raw wages		yearly wages	
	$\hat{\alpha}$	$\hat{\rho}$	$\hat{\alpha}$	$\hat{\rho}$	$\hat{\alpha}$	$\hat{\rho}$	$\hat{\alpha}$	$\hat{\rho}$
1	0.008	0.674	0.009	0.687	0.008	0.686	0.009	0.681
2	0.015	0.724	0.013	0.719	0.013	0.733	0.015	0.725
3	0.020	0.757	0.019	0.754	0.022	0.785	0.019	0.752
4	0.030	0.798	0.030	0.799	0.030	0.814	0.030	0.800
5	0.037	0.817	0.039	0.830	0.034	0.826	0.037	0.820
6	0.046	0.846	0.049	0.857	0.045	0.862	0.049	0.855
7	0.065	0.891	0.060	0.886	0.061	0.899	0.061	0.885
8	0.069	0.900	0.071	0.909	0.068	0.913	0.075	0.915
9	0.075	0.911	0.085	0.935	0.084	0.940	0.081	0.928
10	0.093	0.942	0.087	0.938	0.090	0.952	0.094	0.949
11+	0.132	0.995	0.110	0.934	0.122	0.924	0.115	0.958

The first two columns (baseline) list point estimates shown in Figure 1.4 of the paper. The remaining pairs of columns show estimates corresponding to Figures 1.9-1.11.

For comparison, we show the same plots developed in Section 1.4 but using pooled stocks of migrants, as explained in Section 1.5.

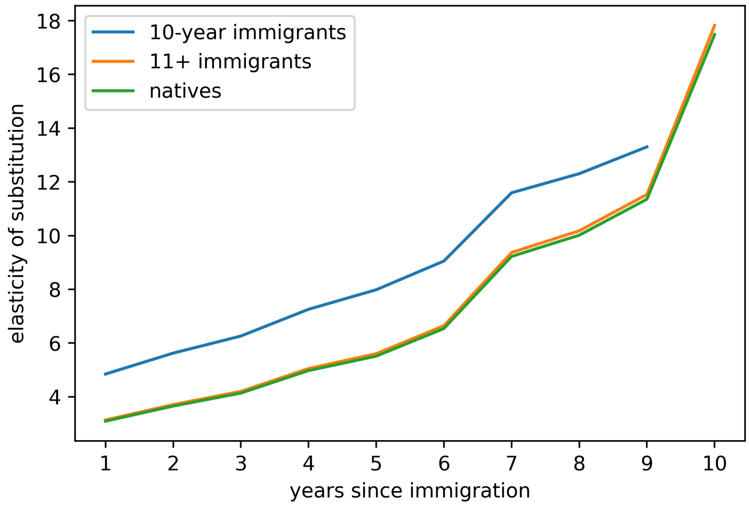


Figure 24: Elasticity of substitution of 10-year migrants (blue line), all previous migrants (yellow line), and natives (green line) with respect to all migrants between 1 and 10 years in the U.S.

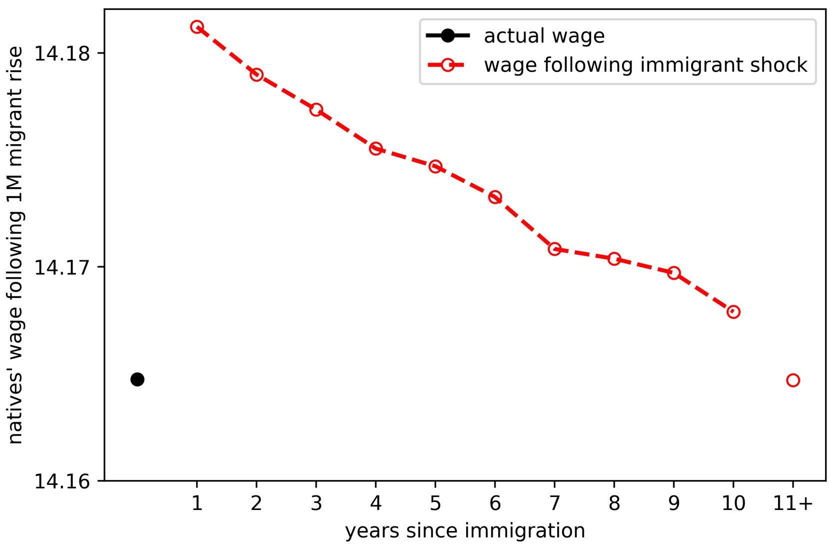


Figure 25: Effect on natives' wages (in USD) of one additional Million of immigrants, as they progress in time through years spent in the U.S., experiencing changes in both substitutability and skill efficiency. The first full dot is the actual average native wage; the dashed line shows the wage response to a one-million immigrant shock by years since arrival.

(a) Counterfactual $\{\partial Y/\partial M_\tau\}_{\tau=1}^{T}$ (b) Counterfactual $\{\alpha_\tau\}_{\tau=1}^{T}$

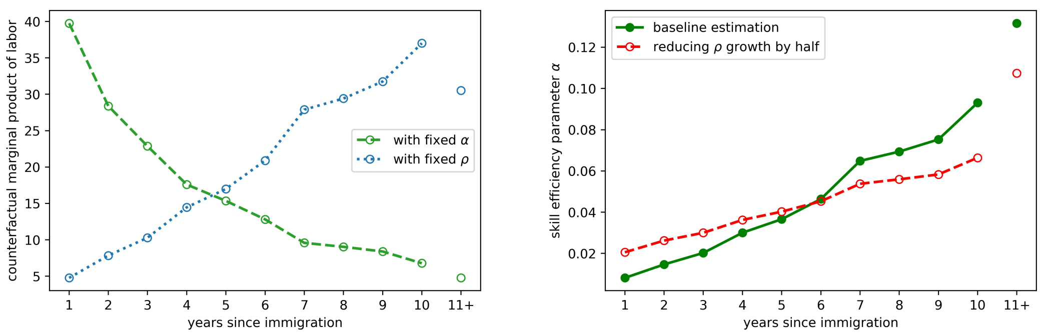


Figure 26: Counterfactuals marginal products and parameter estimates. Panel (a) shows counterfactual wage profiles that would be obtained under alternative models when fixing i. substitutability at its median value of 0.85 while letting skill efficiency evolve (blue dotted line), or ii. skill efficiency at its median value of 0.046 while letting substitutability evolve (green dashed line). Panel (b) shows the actual skill efficiency profile $\boldsymbol{\alpha}$ estimated under the full model (solid line) versus a counterfactual profile (dashed line) that would be obtained if substitutability growth was half as steep, rotated around the median value of ρ_τ.

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Chapter 2

Skill Heterogeneity in Labor Mobility: the Key Role of Education for EU Assimilation

2.1 Introduction and literature review

The ever closer integration of today's globalized world, and in particular of the European Union member states, has made it easier than ever for labor to flow across countries. In the past, large migration waves were made up of predominantly lower-skilled workers, who emigrated to more productive countries in search of a better job, where they could quickly pick up many kinds of lower skill activities that the burgeoning industrial development offered in abundance. Those who had invested time and resources into getting a higher education for a more qualified job were much less willing to give up their domestic chances to seek fortune abroad. In other words, the costs of moving internationally were increasing in skill and ability.

However, nowadays, with a world more interconnected than ever before, it becomes a worthy question to wonder whether this skill-based trend has remained the same or whether it has changed. There are several studies that investigate the particular reasons and peculiarities of labor mobility (Vertovec, 2002), but there is surprisingly little previous work describing the labor mobility responsiveness with heterogeneous skills, distinguishing in particular between high-skilled and low-skilled workers.

Some interesting intuition in this direction can already be gained in the work of Salt (1997), who find that high-skilled workers have been displaying in recent years a quite high degree of volatility over the short-term. This suggests that as workers can move more easily and temporary jobs become more frequent (for example, short-term projects at foreign branches of multinational firms), labor flows are more volatile. This is more so the case for high-skilled labor, which is primarily employed in the tertiary and service sector of the economy, requiring less machinery, land and other physical capital that would make relocation a more long-term investment (Koser and Salt, 1997).

2.1.1 The importance of labor mobility for macroeconomic stability

The traditional macroeconomic literature of currency unions dates back to the pioneering work of Mundell and Fleming in the 1960s (Mundell, 1961; Fleming, 1962). Famously, these analyses compared the effectiveness of monetary policy with fixed and floating exchange rates, finding that fixed rates are optimal when applied to areas with high factor mobility – of both capital and labor. The reasoning is that the value added of a currency area is worth the cost of exchange rate rigidity if labor can move freely within the area to act as a stabilization mechanism against asymmetric shocks, since the exchange rate can no longer serve that purpose. The model was then expanded by McKinnon (1963) and Kenen (1969), becoming a workhorse of macroeconomics.

More recently, as different economic theories have been merged into a more comprehensive approach, important research has focused on analyzing optimal currency areas from a New Keynesian perspective. Starting with seminal work such as that of Obstfeld and Rogoff (1995), previous merely aggregate assumptions in Keynesian models of open economies began to be founded in microeconomic theory, making this approach a strong foundation to develop almost all new macroeconomic models. Gali and Monacelli (2008), importantly, was among the first papers to conceptualize within this framework the theory of optimal currency areas, distinguishing between a union-wide monetary policy and individual-state fiscal policies.

In many of these models, however, labor is treated as being perfectly mobile, in order to serve the same equilibrating purpose that it does in the Mundell-Fleming models. Similarly, the working paper by Farhi and Werning (2014) conceptualizes that reasoning by exploring in detail, and within a New Kaynesian framework, the effects of perfect labor mobility in case of asymmetric shocks in the union. The two authors find that labor mobility from a depressed region into a more productive one is neutral for non-movers in case of internal imbalances, but can be welfare enhancing when exogenous shocks hit. This perhaps surprising neutrality results from particular assumptions in their model, which make it so that the changes in labor supply as movers relocate are offset exactly by the opposite-direction changes in product demand caused by the relocating workers leaving/arriving. Regardless of whether this assumption is to be trusted or not, the paper assumes again perfect labor mobility and treats all workers as homogeneous, both features which are not good approximations of the current state of the European Union.

Finally, although not directly related to the issue of labor mobility, Galí and Monacelli (2016) make an interesting analysis of wage adjustment in currency unions. This is complementary to the present analysis because adjustment in wages plays alongside labor mobility a fundamental role in the equilibrating mechanism of a currency union that cannot alter exchange rates. Many countries have been advocating and pursuing labor market reforms, which would make wages less sticky and more responsive to business cycles to facilitate adjustment and reduce unemployment. The authors interestingly find that, because in a New Keynesian model employment is determined by aggregate demand rather than wage considerations, asymmetric shocks leave the monetary authority constrained and unresponsive, minimizing the impact of shocks on interest rates: thus the substitution effects on consumption are neutralized and in turn aggregate demand will not respond. Therefore, the authors conclude that making wages more responsive to business cycle fluctuations will not help in stimulating employment, and this reduces yet another potential stabilization tool, making labor mobility even more crucial in the analysis of equilibrium restoration.

Up until very recently little attention had been given to the joint effect of labor market frictions and the currency-union monetary constraints. This void has been filled by Kekre (2019), who characterizes optimal monetary policy using a standard New Keynesian framework integrated with search-and-matching features. He finds that labor market frictions amplify output shocks and that as a result it is best for the central bank to accommodate the more sclerotic member of the currency union, i.e. the state with more frictions. Again, however, he does not include labor mobility in the model, and clarifies that these results hold under the assumption of workers bound to their home markets.

Thus, despite all the work on these topics and the recognition that the issue is of importance, little consideration has been paid to imperfections in labor mobility. This is not to say that the topic has been entirely neglected. Dennis and İşcan (2006) develop a model about costly employment relocation, however they do so across productive sectors within the same country and without any concerns for currency union or monetary consequences. Similarly, an analysis focused on high-skilled workers performed by Grossmann and Stadelmann (2011) uses an overlapping generations model to study the effects of this migration on inequality.

Focusing more specifically on skill heterogeneity, Grogger and Hanson (2011) find empirically that countries with higher rewards to skill – in terms of after-tax wages – attract more educated migrants. In a more specific analysis of heterogeneous skill mobility, Franceschin and Görlach (2020) develop a dynamic equilibrium model with workers who differ by skill and tastes, and find that the benefits of migration are reaped mainly by high skilled workers, while low skilled workers might even be worse off. Finally, Basso et al. (2018) estimates a population elasticity to labor demand shocks as a measure of labor mobility. They do this distinguishing between natives and foreign-born and find that foreign-born individuals are much more mobile in response to local asymmetric shocks.

2.1.2 The peculiar case of the EU

It was still very much based on the early models of currency unions that thirty years later, in the 1990s, the European countries drafting the Maastricht treaty debated whether or not it was beneficial to develop a fixed-exchange monetary system and ultimately a common currency. Already at the time, critics warned that labor markets were far from being as mobile as Mundell's theory required and this flaw could have proved dangerous for the equilibrating mechanisms of the Union. Others replied that the increased benefits of risk-sharing would justify this leap forward, and that Europe was on a hopeful path to social and political integration, which would make it incrementally easier for workers to relocate cross-nationally.

After the struggle of the 2008-2013 crises, the issue of understanding how labor behaves in a currency union like the EU has become more and more urgent. More recent work also focus on the peculiarities of the Euro area, but through different lenses. Campolmi and Faia (2011) find that diverse labor market institutions across EU states can account for a significant degree of variety in inflation volatility. In a later work they also explore what the consequences of these labor market asymmetries might be for optimal monetary policy, yet they ignore labor mobility and explicitly admit it would be of interest to analyze it (Campolmi and Faia, 2015).

Search-and-matching analysis leads instead Ljungqvist and Sargent (1998) to suggest that the European states' failure to adapt to rapidly changing times is to blame for the peculiar EU unemployment behavior: unlike labor dynamics in the United States, the "old continent" displays more sclerotic patterns with high unemployment levels that struggle to revert to a natural rate, displaying a unit-root behavior. Many argued from the start that EU institutional features play a fundamental role in determining this behavior, from Nickell and Layard (1999) to Blanchard and Wolfers (2000), yet even today it is far from being resolved. As Galí (2015) shows, the more likely explanations for these patterns, such as hysteresis, crucially hinge on institutional aspects of employment. Clearly, the way European Union regulates its labor market can be fundamental in determining the macroeconomic outcomes of the entire continent.

The EU is particularly interesting because here the difference in behavior between high- and low-skilled workers might be particularly stark. There are three main reasons why I would suppose this. First, the language barrier across the different states in the Union is something rather unique, as opposed to other unions like the United States; foreign languages are mainly learned as a second language in schools, and therefore more educated workers have a greater chance of being familiar with them and have an easier time moving to a foreign country. Second, the great efforts undertaken by the EU's political institutions to pursue economic and social integration are making it much easier for firms in one country to offer their services elsewhere in the EU, requiring those workers to temporarily move where their skills are most needed, thus increasing short-term temporary high-skilled jobs. Third, the same political efforts are pursuing also educational integration, streamlining degree programs, standards, and certifications; this makes it possible for a skilled worker to move more and more easily and have his or her qualification recognized anywhere in the Union.

2.1.3 Paper Structure

These considerations open up interesting research questions into the mobility of workers with heterogeneous skill, and to explore some of the particular reasons for the differences. In this paper, I focus on the evolution of EU labor mobility comparing the patterns of high and low skilled workers over time. I show that, unlike many other unions internationally, in the European Union overall worker mobility has been increasing and that the bulk of this growth has been driven by increases in the migration of high skilled workers. Furthermore, in understanding the reasons behind this unique trend, I show that the factors behind the US decline in interstate mobility cannot be simply inverted to account for the greater EU mobility, and I focus in particular on the skill heterogeneity of migrants within the European Union.

Specifically, I explore on one key additional factor: the recognition of foreign degrees and qualifications, which accounts for an important incentive for high-skilled workers to migrate to other EU countries where their qualifications retain their value. I use survey data to show that as countries implemented the EU's common qualifications frameworks, making it easier to recognize other EU countries' professional degrees, the number of high skilled workers migrating increases remarkably. I use a differences in differences approach to estimate the size of this increase, using low-skilled workers migration as a control group and employing additionally an event study approach to ensure parallel trends and identifying yearly trends.

As a significant strand of recent literature has noted (Goodman-Bacon, 2021; Callaway and Sant'Anna, 2021; Sun and Abraham, 2021), the fact that EU countries implemented the common qualifications framework at different times results in a staggered difference-in-differences setup. This might bias the resulting estimates, and for this reason as a robustness check I use alternative estimations including a dependent variable detrended from year effects in order to center all countries around the same relative treatment year and average treatment effects using cohort-weights. All these exercises confirm a strong positive effect on the migration of high-skilled workers following the implementation of the common framework, reinforcing the importance of certifications portability to labour migration in the common market.

The rest of the paper is organized as follows. Section 2.2 presents the main trends in EU labor mobility, comparing them to the US and differentiating by skill level. Section 2.3 focuses more precisely on the impact of education and the Bologna process of degrees and qualifications recognition. Section 2.4 presents robustness checks for the econometric estimation and alternative specifications, and Section 2.5 concludes.

2.2 Data and empirical facts on EU labor mobility

2.2.1 Comparison with the US benchmark

When discussing about labor mobility, the benchmark often considered for a low-barrier labor market is the United States (e.g. Bryan and Morten (2019)). Previous research, however, in particular that by Kaplan and Schulhofer-Wohl (2017), has highlighted a declining pattern in interstate mobility in the US, which seems to be common to several countries. The authors show that the two main factors driving this trend are: 1) the decreased geographic specificity of occupation returns and 2) increased ease of learning about the destination prior to migration. Figure 2.1 shows the pattern of US mobility, as documented by different data surveys.

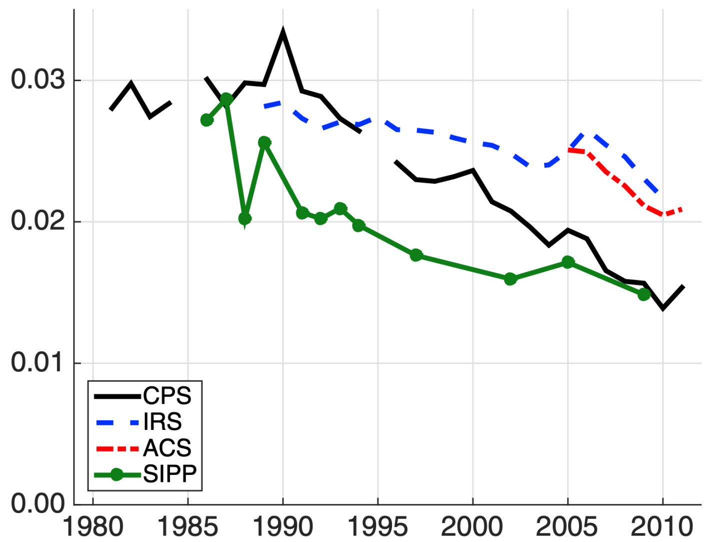


Figure 2.1: Migration rate in the US over time. Source: Kaplan and Schulhofer-Wohl (2017)

Using the yearly data files for the European Labor Force Survery (EU-LFS), I can replicate the exercise for the EU countries. Table 2.1 summarizes some key information about the dataset used. In order to have a cleaner sample size and avoid issues due to new member states joining the EU over the years, I restrict the sample to the countries of the EU15, i.e. those who were in the European Union prior to the large accession of 2004. These countries are: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden, United Kingdom. I therefore restrict the analysis to only these countries and only to within-EU15 migration, i.e. to only migrants who in the past 12 months have moved to a EU15 country coming from another EU15 country. To do so, I take advantage of the variable called "COUNTR1Y" which indicates the country of residence one year prior to the survey, in order to isolate those who have migrated recently.

Figure 2.2 shows the resulting pattern. As one can see, unlike in the case of the US, the migration rate has been increasing, especially in the more recent years. Despite some short-run fluctuations due to the scarcity of the data, the long-run trend is rather clearly positive.

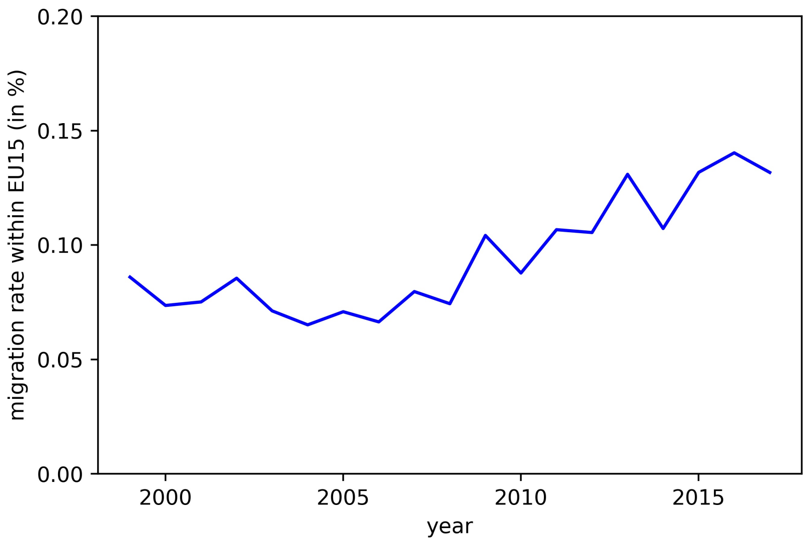


Figure 2.2: Evolution of the migration rate within the EU15 countries. The rate is calculated by dividing the share of individuals who report being in a different EU15 country the previous year by the total number of respondents. The vertical axis is expressed in percentage points.

Table 2.1: Summary statistics

	Full EU-15 sample	Only EU-15 Migrants	Only DE-FR-UK	Only DE-FR-UK migrants
N. of Observations	21663780	31653	6542916	13198
% Male	54.85	57.09	53.06	58.56
Mean Age	41.60	33.03	41.44	32.49
% of High-Skilled Workers	27.76	41.72	29.23	45.54
% of Medium-Skilled Workers	35.16	35.16	48.11	35.86
Mean Individual Sample Weight	157.1	265.1	283.1	435.1

Source: EU Labour Force Survey. The table lists means among the active labour force in the survey years between 2000 and 2015. The first two columns focus on all EU-15 countries: Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Luxembourg, Netherlands, Portugal, Spain, Sweden, United Kingdom. Columns 3 and 4 focus only on the three largest ones – Germany, France, and the United Kingdom – which are mainly used in the estimation. Columns 2 and 4, additionally, restrict the sample to recent migrants only.

As previously discussed, however, there is no reason to believe that these trends should be homogeneous across skill level groups. Therefore, in order to characterize better the composition of the EU migration rate, the same exercise is repeated distinguishing between high skilled, medium skilled and low skilled workers. As per the EU-LFS definitions, a low skilled worker is defined as an employed individual who has completed lower secondary education or below, and high skilled workers as employed individuals who have completed some level of tertiary education. Medium-skilled workers are defined as those workers who have completed upper secondary education but do not have tertiary education. Figure 2.3 shows the distinction of the migration rate across skill levels.

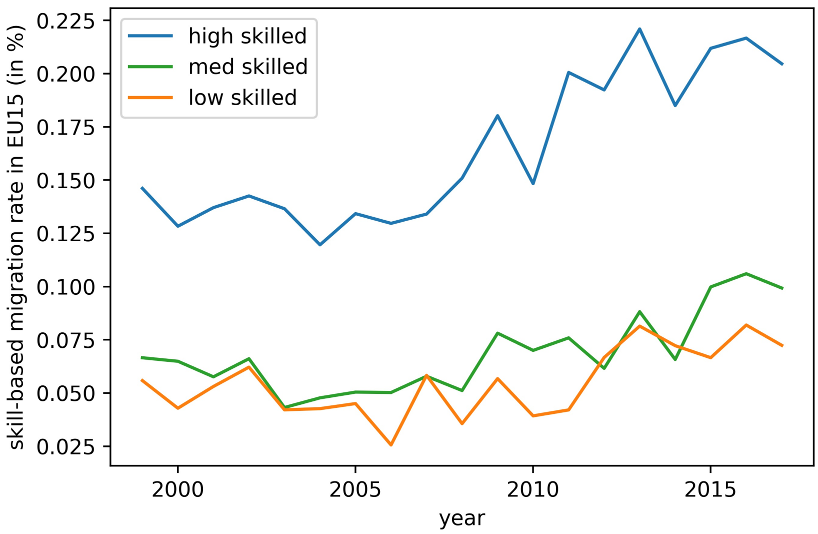


Figure 2.3: Migration rate by skill level. The migration rate is calculated as the number of respondents of a given skill level who report living in a different EU15 country the previous year, divided by the host country's population of that skill level. The blue solid line shows the migration rate for high-skilled workers, the green line for medium-skilled workers, and the yellow line for low-skilled workers. The vertical axis is expressed in percentage points. The data comes from the EU's LFS and the skill level follows the distinction reported by the LFS according to national educational systems.

As the Figure shows, low skilled workers have migrated at a rather constant rate over time, while medium-skilled migrants have a slightly higher growth rate than low-skilled ones but not as high as their high-skilled counterparts. High skilled migration clearly drives the bulk of the rise in the migration rate growth, and is mainly responsible for the spike in overall migration especially after 2005. A specific analysis aimed at explaining the European rise in migration rates, therefore, must take these differences across skill levels into consideration.

2.2.2 Ease of acquiring information

Kaplan and Schulhofer-Wohl (2017) identify two main factors driving the decline in US interstate migration. Namely, the reduction in the geographic specificity and income differential of occupations, thus providing less of a "pull factor" for individuals to move; and an increased ease with which individuals can acquire information about potential host states and their labor markets, thus being able to evaluate more accurately whether it is convenient for them to relocate there. This latter factor should reduce the number of "mistake" migrations, thus reducing the number also of return migrations following an unsuccessful or unsatisfying attempt to move and work abroad.

With the EU-LFS data, anonymized in order to preserve respondents' privacy, neither the income nor the specific regional information is available, thus making the testing of geographic specificity in the EU too imprecise to carry out. What can be tested, instead, is the second factor: the ease of acquiring information. The mechanism at play, following Kaplan and Schulhofer-Wohl (2017)'s reasoning, would be that if individuals have bad or scarce information about their destination, the number of "mistake" migrations would increase, driving up total migrations. To apply their procedure, I need a value for the return migration calculated as the share of arrivals who are natives of the host country, i.e. the share of individuals who return to their birth country from a foreign EU15 country.

I make use of the EU-LFS data to generate a variable for total yearly arrivals, calculated as the total number of individuals in each EU-15 country – both nationals and not – who were living in a different EU-15 country in the previous year. Of this set, I then also calculate the number of respondents who are also nationals of the country they moved to, i.e. those returning to their birth state. This is the same procedure used by Kaplan and Schulhofer-Wohl (2017) with the US census, albeit slightly more precise than theirs because the US census reports 5-year prior residence while the EU-LFS reports 1-year prior residence. Certainly an estimation of this kind is not perfect, as there is no way of knowing when the worker had emigrated abroad in the first place, so we cannot know the duration of the migration spell: a one- or two-year migration spell followed by a return home can more easily be interpreted as the result of unsuccessful migration, while a return after ten or twenty years less so. However, the purpose of the exercise is to replicate Kaplan and Schulhofer-Wohl (2017)'s mechanism to test whether it is able to explain some part of the EU migration rates.

Figure 2.4 shows the ratio of return immigration to total immigration, both at the aggregate level and differentiated by skill level and expressed in percentage points. The share of returning nationals among new arrivals is more or less stable around 0.4%. It certainly does not display any increasing trends, so return migration does not seem to be growing more rapidly than total migration. Therefore, the analysis of return migration, and any changes related to the quality of information as hypothesized by Kaplan and Schulhofer-Wohl (2017), cannot provide any significant explanation about the rising migration rates in the EU. We must thus look for alternative explanations.

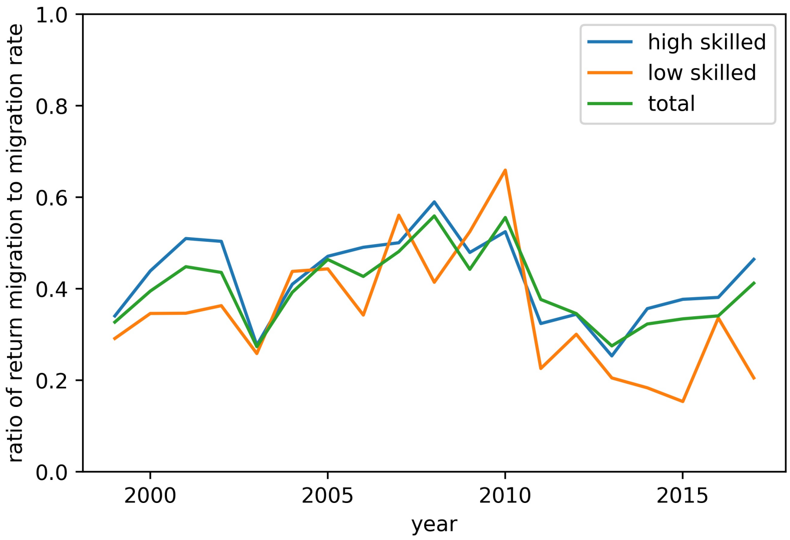


Figure 2.4: Ratio of return migration to total migration (expressed in percentage points), in the EU15 countries and differentiated by skill level. Total migration is calculated as the total number of workers who resided in a different EU-15 country in the previous year; return migration is calculated as the number of those recently-migrated workers who are nationals of the country they moved to from another EU-15 country. The blue line shows the ratio for high-skilled workers; the yellow line the one for low-skilled workers; the green line the aggregate ratio for all skill levels.

2.3 Estimating the role of degree recognition

The previous section highlights the uniqueness of the European Union's situation, and the need to find alternative mechanisms at play to explain the rise in migration rates, specifically for high-skilled workers. Certainly such rising trend cannot be explained by just one single factor, rather by a plurality of them. Here, I explore one of such factors, which has thus far not been thoroughly explored in the literature for migration: the ease with which high-skilled individuals can have their qualifications, acquired in the home country, recognized in other European countries.

2.3.1 The institutional setting and the Bologna Process

The EU has been involved in the development of a common institutional framework for educational and professional qualification, through the so-called Bologna Process. Launched in 1998-1999, the Bologna Process is an intergovernmental higher education reform process across the EU. Its main goal is to improve the quality and ease of recognition of European higher education systems, and to improve the conditions for exchange and collaboration within Europe, as well as internationally (EUA, 2020). It established goals for reform and developed shared instruments, in order to streamline education systems facilitate degree conversion and recognition among the participating countries.

As the name suggests, this is a long and gradual process, which makes an overall evaluation of its effect difficult to pinpoint causally. However, in this process there have been a few milestones, for example the reform of university degrees in the common 3+2 configuration, which allow for a specific individuation of a quasi-experimental approach. In particular, a particularly effective step in this process was the Bergen Conference of European Education Ministers, of May 2005, which adopted the common framework for qualifications in the European Higher Education Area (EHEA), which established the three main cycles of learning, descriptors for each cycle based on learning outcomes and competences, and credit ranges in the first and second cycles.

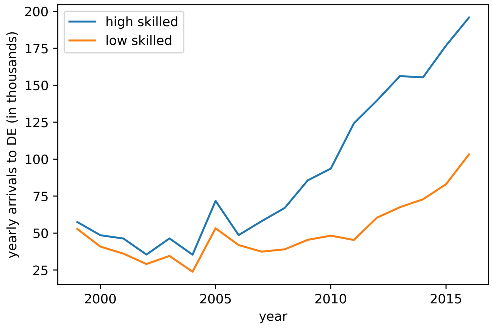
Perhaps more importantly, Ministers committed to elaborating national frameworks for qualifications compatible with the overarching framework for qualifications in the EHEA, thus adapting their national legislation to the common framework. Within this system, as soon as a country aligned its national qualification framework with the EHEA one, all degrees from other EU countries that had been also aligned with the EHEA common framework could much more easily be accepted and recognized. The first country to do so was Germany, who adopted the "Qualifications Framework for German Higher Education Degrees" already in 2005 (EHEA Germany, 2006), followed soon after by other European countries.

2.3.2 The effect of degree recognition

In order to estimate econometrically the impact of such changes, one must search specifically for the date in which each country aligned its national qualifications framework to the common European one, and then estimate the impact of this treatment on high-skilled workers in the subsequent years, using low-skilled workers as a control group. However, the limited size of the EU-LFS data allows me to perform this task with an acceptable degree of confidence only for the larger countries. This is because one must restrict the sample to workers who have migrated in the past year and who came from a EU15 country, and then further differentiate among high- and low-skilled workers. For most of the small and medium sized countries, there aren't enough respondents in the EU-LFS to distinguish a clear trend.

For larger countries such as Germany, France and the United Kingdom, instead, there are enough respondents to perform this exercise. Table 2.1 reports summary statistics on the sample restricted to these countries. The task then is to find the threshold year of implementation of the framework reforms, in order to define that as a treatment. Germany aligned its national qualification framework to the common EHEA standard in 2005 and had it operational since 2006; The UK began the process in 2007 with Scotland and had it operational nationwide in 2008 (EHEA UK, 2008). Figure 2.5 shows clearly how the alignment of national qualification frameworks to the common EHEA one has a striking impact on high-skilled migration in the following years.

(a) Migration to Germany



(b) Migration to the United Kingdom

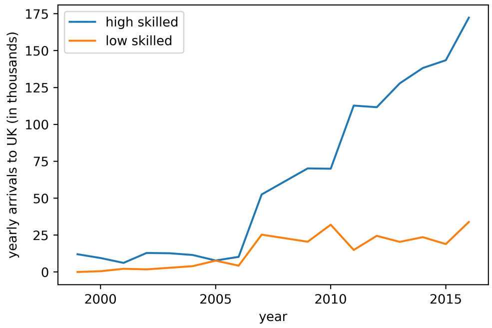


Figure 2.5: Migration to Germany and the United Kingdom over time, shown as number of arrivals by skill level. The blue solid line shows migration of high-skilled workers, the orange line the one of low-skilled workers. The vertical axis is expressed in thousands of workers who have arrived in the last year.

In order to estimate this effect more precisely, I run a difference in differences regression on the largest countries by pre-Bergen immigrant numbers in absolute value, which are Germany, the UK, and France. For each country, I define the "Recognition" variable as a binary indicator based on the year in which the country aligned its national qualification framework to the EHEA one, which I use as treatment. These years are 2006 for Germany and France, 2007 for the UK. Furthermore, for each country, I limit the regression to five years before and after this recognition threshold defined by the alignment year. High-skilled workers are therefore the treated group, low-skilled workers the control group. This way the estimation captures the effect of easier degree recognition for on high skilled migration compared to low skilled migration.

The model I estimate is thus the following:

$$M_{k,s,t} = \alpha + \gamma_k + \theta_{s,t} + \beta(\gamma_k \times \theta_{s,t}) + \phi_s + \delta_t + \epsilon_{k,s,t}$$

where $M_{k,s,t}$ is the migration flow of skill-group k in country s in year t (expressed in thousands), α is a constant intercept, γ_k controls for skill level permanent differences, ϕ_s and δ_t are respectively country and year fixed effects, $\theta_{s,t}$ is a dummy taking a value of 0 if the year t of country s is before treatment and 1 if it is after. The parameter of interest is the interaction term $\gamma_k \times \theta_{s,t}$, which captures the effect of the treatment on the high-skilled workers.

The results of this regression, and others including some or all fixed effects, are summarized in Table 2.2. Columns II and IV control for year fixed effects, while columns III and IV control for state fixed effects. These estimations show that indeed the effect of qualification recognition on high-skilled workers' migration is quite strong, and robust to fixed effects, as it is highly significant at the 5% level in all estimations, 1% when including state or state-by-year fixed effects. It suggests that an easier recognition of qualifications across the EU can rise the number of workers migrating to a large receiving country such as the UK or Germany by over 30 thousand, compared to low-skilled arrivals.

This estimation shows a first clear result on the importance of the common qualifications framework for high-skilled migration flows, which however presents two issues: first, migration flows are highly dependent on the size of the country considered, since a larger country can host more migrants than a smaller ones; second, Figure 2.5 shows a rising trend post-recognition rather than a flat effect, so it is worth exploring the effect on trends rather than simply levels of migration flows. In the next subsections I tackle these issues in turn.

2.3.3 Focusing on Migration Trends

Firstly, I focus on the trends of the post-recognition effect. It makes sense that once an opportunity becomes available, not all workers take advantage of it immediately. As soon as the foreign option becomes available, for example, the majority of unem-

Table 2.2: Effect of qualifications recognition on high-skilled migration

	I	II	III	IV
High-skilled	11.461	13.606	11.461	11.461*
	(9.190)	(9.257)	(6.930)	(5.580)
Recognition	3.518		8.080	
	(8.856)		(6.761)	
High*Recognition	33.340**	29.356**	33.340***	33.340***
	(12.524)	(12.325)	(9.444)	(7.605)
State FE	no	no	yes	no
Year FE	no	yes	no	no
State x Year FE	no	no	no	yes
intercept	21.146***	0.1061***		
	(6.499)	(0.0280)		
R-squared	0.453	0.525	0.702	0.899
R-squared Adj.	0.418	0.379	0.669	0.786
F-statistic	13.224	3.589	21.641	7.918

Difference-in-differences estimation of the impact of qualification framework adoption on the migration of high-skilled workers. The dependent variable is the number of arrivals of a given skill-level group, in thousands. "High-skilled" is a variable taking a value of 0 if the skill group is low-skilled workers, 1 if it is high-skilled workers. Treatment is defined as the variable named "Recognition", taking a value of 1 if the qualifications framework has been adopted, and 0 otherwise. Column I does not control for fixed effects; column II controls for only year effects; column III controls for state fixed effects only; column IV controls for state-by-year fixed effects. The data is taken from the EU-LFS, focusing on Germany, France and the United Kingdom as host countries and only on recent migrants. Years considered are five prior and five after the treatment for each country. Adoption of the framework, i.e. treatment years, are 2006 for Germany and France, 2007 for the United Kingdom.

ployed workers are already searching in the domestic labour market, so beginning to search abroad would become too costly if they believe they are about to receive an offer in the domestic market. In the following years, more and more individuals begin searching also abroad, at first maybe only those who know the host country better or already speak its language, then by word of mouth more searching workers begin exploring a wider range of opportunities. Also firms start getting to know foreign labour markets better and finding it increasingly easier to recruit there, and so on.

As a result, more and more workers take advantage of their degree's foreign recognition as years go on. This leads econometrically to a post-treatment effect that is not constant over time, but that increases year by year. Figure 2.5 in fact shows that this effect is increasing in the years after treatment in a somewhat linear fashion. For this reason, I perform the same estimation as before focusing, instead of on the number of new arrivals, on the growth rate of these arrivals. In other words, this means changing the dependent variable from migration flows to the change in migration flows from one year to the next. Table 2.3 shows the result of this exercise, comparing the growth rate relative to the previous year (in column II) it again with the baseline in absolute terms (in column I). As the table suggests, the growth rate of high-skilled migration is on average about 43% after the implementation of the common qualifications framework, clearly showing the impact of such a policy on high-skilled workers compared to lower educated ones.

2.3.4 Arrivals to Population Ratio

The previous paragraph has focused on growth trends mainly in order to account for a growing estimation effect over time, but using growth as a dependent variable has another important advantage: the possibility to generalize further the result.

Table 2.3: Estimation using migration growth as a dependent variable

	Migration levels (I)	Migration growth (II)
High-skilled	11.4607*	-0.2577*
	(5.5804)	(0.1439)
High*Recognition	33.3404***	0.4304**
	(7.6048)	(0.1961)
State x Year FE	yes	yes
R-squared	0.8991	0.9339
R-squared Adj.	0.7855	0.8596
F-statistic	7.918	12.562

Difference-in-differences estimation of the impact of qualification framework adoption on the migration flows of high-skilled workers. Column I shows the baseline using migration levels as a dependent variable, i.e. actual arrivals of workers in the host countries. Column II instead uses as dependent variable migration growth, i.e. the rate of change in yearly arrivals compared to the year before. "High-skilled" is a variable taking a value of 0 if the skill group is low-skilled workers, 1 if it is high-skilled workers. Treatment is defined as the variable named "Recognition", taking a value of 1 if the qualifications framework has been adopted, and 0 otherwise. The data is taken from the EU-LFS, focusing on Germany, France and the United Kingdom as host countries. Years considered are five prior and five after the treatment for each country. Adoption of the framework, i.e. treatment years, are 2006 for Germany and France, 2007 for the United Kingdom.

Having to focus only on the larger countries in the EU15, due to sample size reasons, the estimation in absolute numbers does not translate well to smaller countries, since the number of immigrants arriving to a host country is highly dependent on the size of that country. Focusing on growth rates of arrivals provides a relative measure of the qualifications framework's impact on high-skilled migration, which is more useful for reasoning about the EU as a whole rather than about a specific country's levels.

However, for certain analyses the growth rate of migration might a less informative parameter to consider than the magnitude actual migration flows. A researcher might want be more interested in understanding what is the impact of the Bologna process on the EU's labour market relative to the size of this labour market rather than just within the subset of migrants.

To do so, I create a different explanatory variable, which is the ratio of high- and low-skilled arrivals to the host country's population. This allows me to focus on actual yearly arrivals rather than their growth, yet it provides a relative measure – weighting arrivals by host-country population – that can be extended for a broader analysis of the EU's migration flows. As Figure 2.5 suggests, the share of high- and low-skilled workers with respect to the host country's population is quite small, hovering between 0.05 and 0.5 % of the total population. As a result, for the remainder of the paper I will multiply these arrivals-to-population ratios by 100, providing thus a value indicating the percentage share of high- or low-skilled arrivals in the population.

Table 2.4 provides a summary of the main estimations conducted thus far, both with and without fixed effects and including also trends, using the ratio of arrivals to host-country population. As the table's coefficients show, he implementation of the qualifications framework raises the share of high-skilled migrants in the total population by about 0.1%, which is pretty significant given pre-treatment population shares of about 0.1%. The treatment more than doubles the share of high-skilled arrivals in the host country's population.

This estimation, and this explanatory variable, forms a useful baseline to perform also further analysis, such as a more precise definition of yearly effects for every year post-treatment. An event study of this kind also allows to test for parallel trends, to which I now turn.

2.3.5 Event Study: Parallel Trends and Yearly Effects

The previous estimations, like all difference-in-differences, relies on the assumption of parallel trends pre-treatment. In the case of the EHEA, it means that had there not been the qualification framework adoption, the difference between high-skilled and low-skilled migration would have remained constant at its pre-adoption value. To test this assumption, one can interact the treatment variable with relative-time dummies for each year before and after the treatment. In other words, it requires creating a new set of time variables, which instead of representing calendar years, represent years before or after the adoption of the common European qualifications framework. This variable, which I call $D^j_{s,t}$ takes value of 1 if calendar year t is j years away from the year of implementation of the qualifications framework in country s; it takes value of 0 otherwise. By interacting it with the treatment-group variable γ_k, we ensure that the resulting coefficient β_j yields the effect of being j years away from the treatment for the high-skilled group.

Thus, if the value of β_j is not significantly different from 0 before the treatment, we can be confident that the parallel trends assumption is met and, therefore, the estimated effect is likely to be a valid one. Another advantage of this test is that at the same time it allows to identify the effect of the common qualifications framework

Table 2.4: Estimation using ratio of skill-group arrivals to total host-country population

	I	II	III	VI	V
High-skilled	0.0384	0.0469*	0.0384*	0.0384**	
	(0.0252)	(0.0242)	(0.0228)	(0.0185)	
Recognition	0.0093		0.0197		
	(0.0243)		(0.0222)		
High*Recognition	0.1073***	0.0916***	0.1073***	0.1073***	
	(0.0344)	(0.0322)	(0.0311)	(0.0252)	
Trend (high-skill)					
Trend (low-skill)					
State FE	no	no	yes	no	
Year FE	no	yes	no	no	
State x Year FE	no	no	no	yes	
Intercept	0.0637***	0.0770**			
	(0.0178)	(0.0327)			
R-squared	0.5322	0.6318	0.6342	0.8746	
R-squared Adj.	0.5030	0.5185	0.5944	0.7336	
F-statistic	18.204	5.576	15.948	6.201	

Difference-in-differences estimation of the impact of qualification framework adoption on the migration of high-skilled workers. The dependent variable is the ratio of arrivals of a given skill-level group to the host country total population, in percentage points. "High-skilled" is a variable taking a value of 0 if the skill group is low-skilled workers, 1 if it is high-skilled workers. Treatment is defined as the variable named "Recognition", taking a value of 1 if the qualifications framework has been adopted, and 0 otherwise. Column I does not control for fixed effects; column II controls for only year effects; column III controls for state fixed effects only; column IV controls for state-by-year fixed effects. The data is taken from the EU-LFS, focusing on Germany, France and the United Kingdom as host countries and only on recent migrants. Years considered are five prior and five after the treatment for each country. Adoption of the framework, i.e. treatment years, are 2006 for Germany and France, 2007 for the United Kingdom.

separately for each year since the treatment, unlike the main estimation which allows only to identify a global post-treatment effect or trend.

The estimation in this case becomes the following:

$$M_{k,s,t} = \alpha + \sum_{j \neq 0} \beta_j (D^j_{s,t} \times \gamma_k) + \gamma_k + \phi_s + \delta_t + \epsilon_{k,s,t}$$

The resulting estimates, with error bars around each year's estimates, are plotted in Figure 2.6. As the Figure shows, the years preceding the treatment, i.e. the alignment of qualifications to the European standard, the effect on high-skill migration is non-significant; however, after the alignment, the effect is growing and after the first year it becomes clearly positive and significantly different from zero. This exercise shows an important point, additionally: that the effect of qualifications alignment is growing over time, at least in the first five years since treatment. In the first year the effect seems to be positive but is still not very significant, the main impact is visible beginning with the second year after treatment, with an effect close to 0.14 percentage points. The effect grows past 0.18 percentage points with the fourth and fifth years after treatment. These values are higher than the 0.107 value found in the main estimation, as that represents a cumulative effect likely dampened by the low value of the first year after treatment.

Such a growing path makes sense as reforms might take a while to become actually implemented, especially for life-changing choices such as moving to another country. Therefore, the first year since the qualifications alignment the reform is still "new", and it can take several months if not years for its full effect to become visible. This increasing pattern makes sense also in the broader context of explaining the growing migration rates in the EU, especially for high-skilled workers, which aren't concentrated only around the implementation years, but are steadily growing over

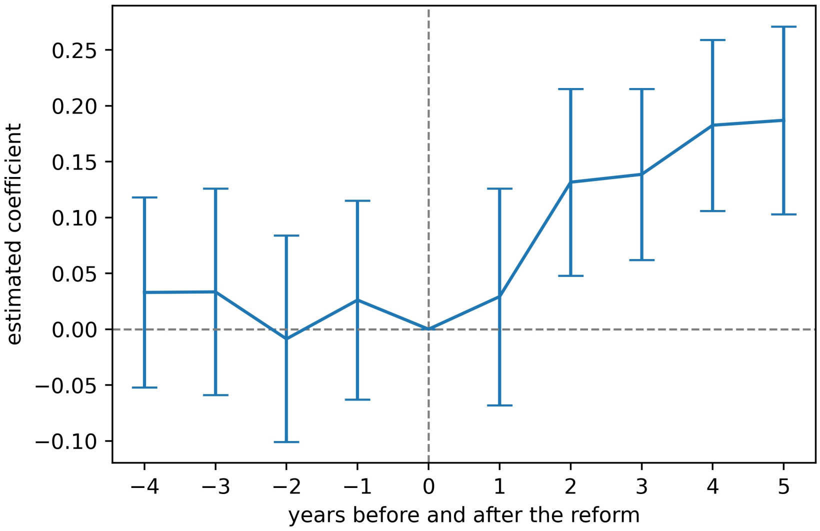


Figure 2.6: Yearly effects on high-skill migration from EU-15 countries relative to the year of implementation of the common qualifications framework in Germany, France and the United Kingdom. The dotted lines indicate the baseline for the year of implementation, data on the left of the vertical dotted line can be interpreted as a test of the parallel trends assumption. The vertical bars represent 95% confidence intervals around each estimate.

time.

That of parallel trends is not the only assumption which merits attention. The staggered implementation design of this quasi-experiment, having the various EU countries implemented the common qualifications framework in different years, might pose a threat to the validity of the estimation. Goodman-Bacon (2021) shows that regression designs where there is variation in the timing of treatment, i.e. when treatment is staggered, might present biased estimates, due to the difference-in-differences estimation actually representing the weighted average of all combinations of two-group/two-period estimates. I explore this critique and possible robustness checks to address it in Section 2.4.

2.3.6 Regulated Professions

Another important goal of the Bologna Process was the ability to recognize professional qualifications specifically regarding the so-called "regulated professions". These are specific occupations for which individuals have to be awarded a particular license by a board or a national entity, through an exam or some certified professional experience. Such categories include lawyers, architects, etc.

In order to facilitate the exercise of regulated professions across the EU, the Commission issued the Directive 2005/36/EC, with the goal of enabling the free movement of professionals such as doctors or architects within the EU. This legislation, sometimes referred to as the "Professional Qualifications Directive" provides rules and frameworks to recognize the qualifications of professionals from one EU country in order to practice in another one. The recognition can be either automatic, allowing all those with a minimum qualification or experience in the profession to practice in other EU countries, or more general setting out clear instructions and possible "compensatory measures" to apply for the license to practice abroad.

In order to exercise their occupation abroad, individuals in regulated professions must apply for the recognition of their qualification, and this process can culminate with a positive or negative decision by the national entity. The EU Regulated Professions database makes available the share of applications in each country that obtain positive decisions. Figure 2.7 plots the relationship between the share of positive decisions and high-skilled migration in the main EU15 countries. As is clearly visible, there is a strongly positive relationship between the two variables, meaning that the easier it is for a country to certify qualified foreigners to exercise their profession, the more attractive it will be for migration.

This further demonstrates the importance that the recognition of qualifications

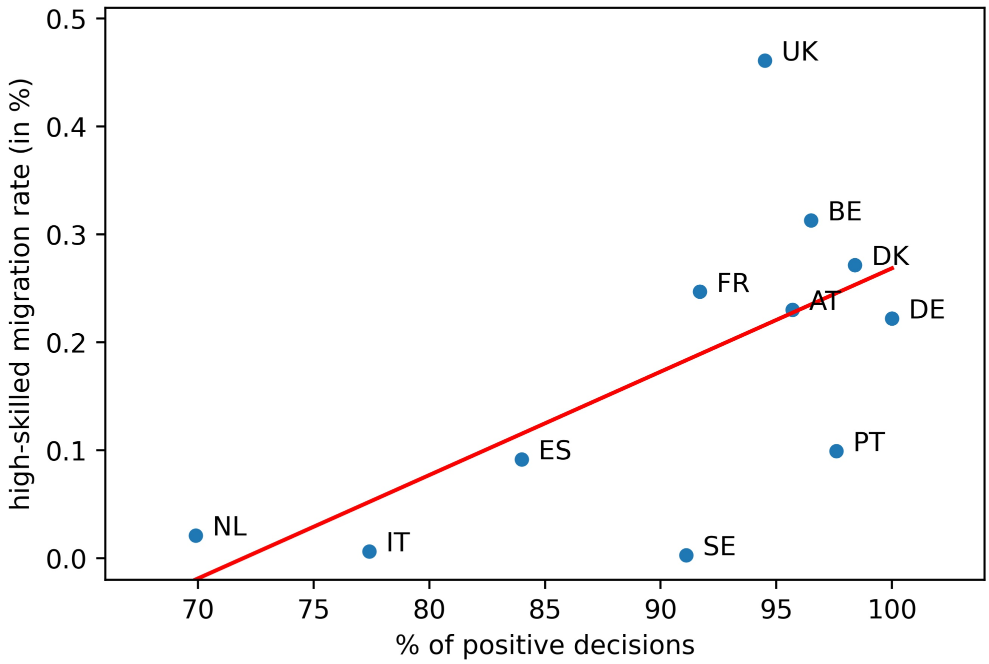


Figure 2.7: Scatter plot showing the relationship between the share of high skilled migrants in the population (in percentage points) and the recognition rate of each country, defined as the number of positively accepted requests for professional practice in the country from other EU15 countries. The red line shows the resulting fitted relationship between the data points. Data on migration comes from the EU-Labour Force Survey, while data on accepted applications comes from the EU Regulated Professions database.

has for the migration of high-skilled and highly qualified workers in the European Union and, given the traction that high-skilled migration has on the overall migration rate, for intra-EU migration in general. As EU countries become more intertwined and their economies more integrated, labour mobility as a re-equilibrating mechanism after asymmetric shocks requires an ever more efficient qualification recognition system.

2.4 Robustness Checks

In this section, I provide some robustness checks to test the validity of the results presented in the previous section. First, I focus on the inclusion of possible controls which have been used in the literature to explain migration – namely GDP per capita and GDP per capita growth, and the deterioration of national labour markets. Then, I focus on subgroups – men, women, youth – to check whether the phenomenon is robust across different demographics. Finally, I focus on the possible limitations of a staggered difference-in-differences design suggested by Goodman-Bacon (2021) and implement some of the methodologies recommended to recover the true effects, namely an alternative specification and event studies using cohort-based weights.

2.4.1 Including additional macro-level controls

A primary source of bias in the estimates might the presence of confounding variables, which could explain migration rate changes to a particular country. Most notably, macroeconomic conditions might attract or disincentivize migrants from choosing a given country as their destination. The primary variable to consider for this is GDP, as it is the most widely used variable for cross-country comparisons in terms of economic performance.

Using Eurostat data, I generate a variable indicating each country's GDP per capita both in absolute value, as reported by Eurostat, and in terms of changes, using year-on-year percentage growth. Each country's GDP variables are interacted with a country dummy ϕ_s taking a value of 1 for observations of that country, 0 otherwise. Table 2.5, columns II to V, shows the results of these estimations, both with and without state fixed-effects. While GDP per capita growth isn't particular significant, GDP per capita levels instead do have some impact on migration flows. In particular, once state-by-year fixed effects are accounted for, the GDP per capita of host countries is positively and significantly correlated with arrivals to that country. The inclusion of these GDP controls, compared to the baseline estimation in column I, reduces the error margin of the estimate but leaves the estimated effect of qualifications recognition largely unaffected.

Another important reason that has been frequently explored for the intra-EU migration rise is the deterioration of national labour markets, particularly in southern-European countries. As a result, more workers from these countries might wish to migrate to countries with stronger labour markets, of which Germany and the United Kingdom are indeed good examples. To check for this hypothesis, and verify whether they take out the explanatory power of the qualifications reform, I include two variables as control: the unemployment level in Southern-EU countries, and the migration of high-skilled workers from these countries to the United States.

The reasoning behind the first variable is pretty straightforward, since the deterioration of southern-European labour markets leads to an increase in the unemployment of those countries and therefore to higher emigration to other EU countries. The second variable tests, instead, whether there has been a rise specifically in high-skilled workers emigrating from those countries to the United States, thus perhaps reflecting an unusually high share of high-skilled emigration in the observed years to all international destinations, regardless of effect of the qualifications reform in EU destination countries. The US has long been a primary destination country for migrant workers from southern Europe, therefore controlling for high-skilled migration to the US might allow to isolate the effect of high-skilled migration changes due to labour market deterioration.

To do this, I use the American Communities Survey (ACS) data, which indicates for each migrant worker their country of origin and the year of arrival in the US, as well as their education level. I therefore isolate high-skilled workers coming from the three largest southern-European countries: Italy, Spain and Greece. I focus on workers coming from these three countries who have immigrated to the US in the previous year, and thus create a proxy variable of high-skilled immigration into the US from southern-European countries. For the unemployment rates, I focus on the same three sending countries, and exploiting Eurostat data I create a weighted average of their unemployment rates, using their average population size as weights.

The results of the estimation are shown in columns VI and VII of Table 2.5. Understandably, the unemployment rate of southern-European countries is positively related with migration flows, while high-skilled immigration to the US has a less significant impact. All controls, however, even the highly significant ones, do not take any effect away from the coefficient of interest – that of degree recognition for high-skilled workers – which remains mostly unchanged.

The last column of Table 2.5 shows a regression including both GDP per capita growth, US high-skilled immigration flows and southern-EU countries' unemployment as controls, as well as year-by-state fixed-effects. As throughout the various specifications including individual controls, the coefficient for the impact of qualifications recognition – what is called "High*Recognition" in the Table – remains robust to all controls in this cumulative regression as well. Thus these controls do have a significant impact on EU migration flows. As I already said previously, migration flows are explained by a plurality of factors and it would be foolish to imagine that only one factor could explain all the variation in migration rates. However, for the purpose of this paper it is key to note that these additional explanations do not invalidate or take significance away from the impact of the common qualifications framework's implementation.

Table 2.5: Effects on migration using controls

	I	II	III	IV	V	VI	VII	VIII
High-skilled	0.0384**	0.0384*	0.0384	0.0384**	0.0384**	0.0384**	0.0384**	0.0384**
	(0.0185)	(0.0199)	(0.0231)	(0.0185)	(0.0185)	(0.0185)	(0.0185)	(0.0185)
Recognition		-0.1033**	-0.0982*					
		(0.0479)	(0.0546)					
High*Recognition	0.1073***	0.1073***	0.1073***	0.1073***	0.1073***	0.1073***	0.1073***	0.1073***
	(0.0252)	(0.0271)	(0.0315)	(0.0252)	(0.0252)	(0.0252)	(0.0252)	(0.0252)
DE GDP		-0.0246*		0.0027***				
		(0.0140)		(0.0006)				
DE GDP growth			-0.1729		0.0045			0.0036
			(0.8350)		(0.0073)			(0.0067)
FR GDP		-0.0274*		0.0013				
		(0.0142)		(0.0008)				
FR GDP growth			-2.0892		-0.0011			-0.0012
			(1.7063)		(0.0019)			(0.0018)
UK GDP		-0.0269*		0.0027***				
		(0.0144)		(0.0008)				
UK GDP growth			-2.3145*		-0.0039			-0.0034
			(1.1585)		(0.0034)			(0.0033)
IT-ES-GR unemployment						0.0070***		0.0063***
						(0.0021)		(0.0021)
US immigration							0.0213*	0.0087
							(0.0122)	(0.0126)
Year-FE	No	Yes	Yes	No	No	No	No	
State x Year FE	Yes	No	No	Yes	Yes	Yes	Yes	
R-squared	0.8746	0.7885	0.7131	0.8746	0.8746	0.8746	0.8746	0.8746
R-squared Adj.	0.7336	0.6918	0.5819	0.7336	0.7336	0.7336	0.7336	0.7336
F-statistic	6.201	8.154	5.436	6.201	6.201	6.201	6.201	6.201

The table shows the coefficients of the estimation using controls. Column I shows the baseline estimation without controls. The first set of controls is GDP per capita, both in absolute terms and in growth; this set of estimation is displayed in columns II and IV for GDP per capita in absolute terms, in columns III and V in terms of growth. GDP per capita growth in each country is pre-interacted with a dummy variable ensuring each country's GDP only impacts migration to that country. Column VI includes only the immigration to the US from Southern European countries (Spain, Italy, Greece), and immigration is expressed in thousands of workers migrating. Column VI includes both GDP per capita growth and immigration to the US. Source: EU Labour Force Survey for most variables, Eurostat for GDP per capita, US American Communities Survey data for the US immigration variable.

2.4.2 Focusing on Subgroups

Another important aspect to check is the possible variation of the effect estimates across different population. The inclusion of different demographics in the data certainly reflects the aggregate migration trends, yet it is of interest to check whether the impact of qualifications recognition is more or less consistent across different population subgroups or whether it varies substantially.

To do so, I first run the estimation using only men, and report the results in Table 2.6. Then, I run the estimation focusing only on women, and report the results in Table 2.7. Finally, I restrict the sample to only young individuals under the age of 36, and report the results in Table 2.8.

As one can see from the Tables, there are noticeable and very interesting differences across these three sub-populations. Women display a slightly lower responsiveness to qualifications portability, perhaps due to lower career growth opportunities or to usually greater care and family duties, which keep them more likely in the origin country. Men, on the other hand, respond more strongly to the implementation, with an interaction coefficient of 0.118 whereas women's is 0.096. The most striking result concerns young individuals: they display a much larger high-skilled migration response to common qualifications, reaching 0.183, almost twice as much as the women-only subgroup. This is easily explainable by greater future returns in terms of career opportunities for young individuals, lower family ties and better knowledge of languages, among other factors.

Interestingly, each population's estimations are rather robust internally to the inclusion of fixed effects, yet important cross-group variations remain. However, all these sub-populations display very similar positive and significant estimates for the impact of common qualification frameworks on high-skilled migration.

Table 2.6: Effects on migration including only men

	I	II	III	IV
High-skilled	0.0427	0.0427	0.0427	0.0427*
	(0.0290)	(0.0273)	(0.0266)	(0.0240)
Recognition	0.0156	-0.0940	0.0280	-0.0941
	(0.0279)	(0.0609)	(0.0259)	(0.0574)
High*Recognition	0.1176***	0.1176***	0.1176***	0.1176***
	(0.0395)	(0.0373)	(0.0362)	(0.0327)
State FE	no	no	yes	yes
Year-FE	no	yes	no	yes
intercept	0.0668***	0.0895**	0.0845***	0.1279***
	(0.0205)	(0.0362)	(0.0218)	(0.0336)
R-squared	0.5206	0.6618	0.6126	0.7526
R-squared Adj.	0.4906	0.5461	0.5705	0.6495
F-statistic	17.375	5.719	14.546	7.301

Difference-in-differences estimation of the impact of qualification framework adoption on the migration of high-skilled workers, restricting the sample to men only. The dependent variable is the ratio of arrivals to host-country population (in percentage points) for a given skill-level group. Treatment is defined as the variable named "Recognition", taking a value of 1 if the qualifications framework has been adopted, and 0 otherwise. Column I does not control for fixed effects; column II controls for only year effects; column III controls for state fixed effects only; column IV controls for both year and state fixed-effects. The data is taken from the EU-LFS, focusing on Germany, France and the United Kingdom as host countries and only on recent migrants. Years considered are five prior and five after the treatment for each country. Adoption of the framework, i.e. treatment years, are 2006 for Germany and France, 2007 for the United Kingdom.

Table 2.7: Effects on migration including only women

	I	II	III	IV
High-skilled	0.0328	0.0328	0.0328	0.0328*
	(0.0224)	(0.0226)	(0.0199)	(0.0182)
Recognition	0.0022	-0.0733	0.0104	-0.0963**
	(0.0216)	(0.0504)	(0.0194)	(0.0435)
High*Recognition	0.0961***	0.0961***	0.0961***	0.0961***
	(0.0306)	(0.0309)	(0.0271)	(0.0248)
State FE	no	no	yes	yes
Year-FE	no	yes	no	yes
intercept	0.0601***	0.0712**	0.0854***	0.1054***
	(0.0159)	(0.0300)	(0.0163)	(0.0254)
R-squared	0.5174	0.6108	0.6375	0.7615
R-squared Adj.	0.4872	0.4776	0.5981	0.6621
F-statistic	17.154	4.587	16.179	7.661

Difference-in-differences estimation of the impact of qualification framework adoption on the migration of high-skilled workers, restricting the sample to women only. The dependent variable is the ratio of arrivals to host-country population (in percentage points). Treatment is defined as the variable named "Recognition", taking a value of 1 if the qualifications framework has been adopted, and 0 otherwise. Column I does not control for fixed effects; column II controls for only year effects; column III controls for state fixed effects only; column IV controls for both year and state fixed-effects. The data is taken from the EU-LFS, focusing on Germany, France and the United Kingdom as host countries and only on recent migrants. Years considered are five prior and five after the treatment for each country. Adoption of the framework, i.e. treatment years, are 2006 for Germany and France, 2007 for the United Kingdom.

Table 2.8: Effects on migration including only young individuals

	I	II	III	IV
High-skilled	0.0516	0.0516	0.0516	0.0516
	(0.0525)	(0.0497)	(0.0503)	(0.0496)
Recognition	0.0220	-0.2178*	0.0384	-0.1647
	(0.0506)	(0.1107)	(0.0490)	(0.1185)
High*Recognition	0.1828**	0.1828**	0.1828**	0.1828**
	(0.0716)	(0.0677)	(0.0685)	(0.0676)
State FE	no	no	yes	yes
Year-FE	no	yes	no	yes
intercept	0.1004***	0.1282*	0.0764*	0.1503**
	(0.0371)	(0.0657)	(0.0412)	(0.0692)
R-squared	0.4162	0.5868	0.4877	0.6100
R-squared Adj.	0.3797	0.4454	0.4320	0.4474
F-statistic	11.405	4.151	8.759	3.753

Difference-in-differences estimation of the impact of qualification framework adoption on the migration of high-skilled workers, restricting the sample to young individuals under 36 years of age. The dependent variable is the ratio of arrivals to host-country population (in percentage points). Treatment is defined as the variable named "Recognition", taking a value of 1 if the qualifications framework has been adopted, and 0 otherwise. Column I does not control for fixed effects; column II controls for only year effects; column III controls for state fixed effects only; column IV controls for both year and state fixed-effects. The data is taken from the EU-LFS, focusing on Germany, France and the United Kingdom as host countries and only on recent migrants. Years considered are five prior and five after the treatment for each country. Adoption of the framework, i.e. treatment years, are 2006 for Germany and France, 2007 for the United Kingdom.

2.4.3 Staggered DiD Setup

The estimations thus far have relied on a traditional difference-in-differences (DiD) setup. However, there are some key peculiarities about the EU countries' implementation of the Bologna Process which warrant further examination. Most notably, as already mentioned, not all countries implemented the process in the same year: Germany and France implement it in 2006 while the UK implements it in 2007. Therefore, at least in part, one can consider this an instance of staggered implementation, or staggered DiD.

In the case of a staggered implementation, some issues might arise as pointed out by Goodman-Bacon (2021). In fact, the DiD estimator in cases of staggered implementation is given by the weighted average of the coefficients of all 2x2 combinations, i.e. all pairs of two-time periods in which there is a treatment group whose treatment status changes and a control group whose treatment status does not change. The weights depend on the size of each group and on the distance of treatment from the window considered: groups receiving treatment in the middle of the time frame receive more weight, while groups treated at the beginning or end of the panel get less weight. Additionally, controls – if used – can bring about new variation, as they compare units with the same treatment variable but different predicted treatment based on the covariates.

In the case of the present analysis, these problems are not likely to affect the estimation dramatically: while it is true that not all treatment occurs in the same calendar year, there is only a one-year gap between the treatment of one country and the other two, out of a panel of 11 calendar years. Additionally, these treatments are located rather centrally in the panel, thus reducing the bias suffered by being located marginally in the panel. Finally, controls are not used if not in the robustness section, and thus cannot impact the main estimation.

In any case, and since the possibility of significant bias cannot be excluded a-priori, in the next two subsections I take advantage of the discussion on the topic and I explore two possible ways to deal with these problems, based on the suggestions of Goodman-Bacon (2021) and on related work, thus further testing the robustness of the results.

2.4.4 Alternative Specification

The first possibility I explore is to remove the time-asyncrony by expressing all time in relative terms. As the sample panel is long enough and the treatment instances close enough, an alternative specification of this kind should not alter the analysis significantly.

In particular, I first run a regression of each country's arrivals-to-population ratio on a time trend and GDP controls. I take out the effect of these variables and add back the residuals, thus creating a de-trended variable for each country. Then I create a new time variable which expresses time in terms of years before or after treatment, i.e. the adoption of the common framework. In a way, I am aligning all countries and years and calculating the group-time average treatment effects as defined by Callaway and Sant'Anna (2021). Another attractive feature of this their implementation is the possibility of conditional parallel trends, which however are not necessary here as unconditional parallel trends are already verified and there are no covariates.

This process yields a set of standard DiD panels all centered around the same relative year 0, which can be used in a simple estimation as previously done. The results of the estimation are described in Table 2.9, using both state and relative-year fixed effects. Interestingly, while the errors vary across specifications, the coefficient of interest – on the effect of common qualifications on high-skill migration – is the same and very robust. This suggests that the main estimation, with the staggered setup, does not significantly suffer from the weighting bias of the 2x2 DiD combinations.

Additionally, in this specification just like in the main one we can test for the parallel trends assumption and at the same time identify the yearly effect post-treatment. To do so, I run the same regression as in Section 2.3 to obtain the plot shown in Figure 2.8. As the figure shows, the error bars are larger yet the post-treatment estimates have a similar path; the growth of the effect is steeper in this specification, surpassing 0.2 in the fifth year. All yearly effects before treatment are not significantly different from zero, thus validating the parallel trends assumption.

2.4.5 Event Study

Another approach to solving the staggered adoption issue is the one of Sun and Abraham (2021). They carry out an event study by defining the unit-level treatment effect as the difference between the observed outcome relative to the never-treated outcome. To do so they define $CATT$, i.e. the cohort average treatment effects, as the cohort-specific difference in outcomes relative to never being treated, which can be estimated as a weighted average using weights that sum to one and are non-negative. In particular, the weights are the shares of cohorts that experience at least a certain number of years relative to treatment, normalized by the size of the group.

Based on this approach, I create a new variable reflecting the difference in skill-based migration for every year and country. I net the resulting difference from year and GDP trends, as done before, as well as from the pre-treatment difference between the two groups. Finally, I generate weights based on the size of each country-skill-year subsample and the LFS's weighting coefficient of its components.

Table 2.9: Alternative specification, using relative years

	I	II	III	IV
High-skilled	0.0384	0.0384	0.0384*	0.0384*
	(0.0239)	(0.0255)	(0.0211)	(0.0219)
Recognition	-0.0596**	-0.0482*	-0.0515**	-0.0483**
	(0.0230)	(0.0276)	(0.0206)	(0.0237)
High*Recognition	0.1073***	0.1073***	0.1073***	0.1073***
	(0.0325)	(0.0348)	(0.0288)	(0.0298)
State FE	no	no	yes	yes
Year-FE	no	yes	no	yes
intercept	0.0992***	0.0970***	0.1275***	0.1364***
	(0.0169)	(0.0285)	(0.0173)	(0.0266)
R-squared	0.4901	0.5145	0.6172	0.6609
R-squared Adj.	0.4583	0.3810	0.5756	0.5448
F-statistic	15.382	3.854	14.834	5.696

Difference-in-differences estimation of the impact of qualification framework adoption on the migration of high-skilled workers, using an alternative specification that focuses on relative years before/after treatment. The dependent variable is the ratio of arrivals to host-country population (in percentage points), detrended by taking out GDP and yearly trends. Treatment is defined as the variable named "Recognition", taking a value of 1 if the qualifications framework has been adopted, and 0 otherwise. Column I does not control for fixed effects; column II controls for only year effects; column III controls for state fixed effects only; column IV controls for both year and state fixed-effects. The data is taken from the EU-LFS, focusing on Germany, France and the United Kingdom as host countries and only on recent migrants. Years considered are five prior and five after the treatment for each country.

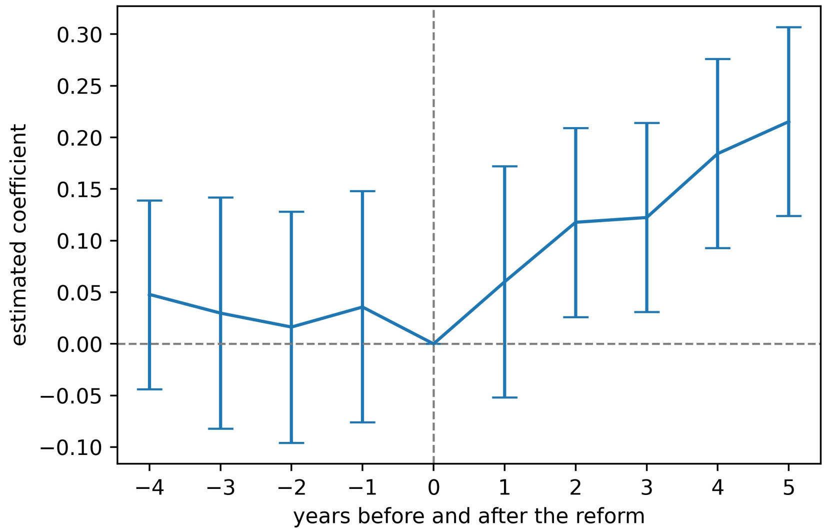


Figure 2.8: Yearly effects on high-skill migration (in percentage points) from EU-15 countries relative to the year of implementation of the common qualifications framework in Germany, France and the United Kingdom, using an alternative specification with detrended arrival-to-population ratios and relative years before/after treatment. The dotted lines indicate the baseline for the year of implementation, data on the left of the vertical dotted line can be interpreted as a test of the parallel trends assumption. The vertical bars represent 95% confidence intervals around each estimate.

This yields a series of weights and differences, which can be used calculate both overall effects (weighting all groups) and dynamic effects (each year by itself). The dynamic result can be contrasted to the parallel trends test conducted in Section 2.3. The result is shown in Figure 2.9, where the new path is indicated by the red line and the baseline estimation in blue. The result appears flatter, although in the fifth year after treatment it slightly surpasses the baseline value for the same year.

Overall, therefore, all robustness checks confirm the general trend of the estimation: after a country aligns its qualifications framework to the common European one, its high-skilled migration increase over time. In the first five years alone it can reach a value of around 0.2 percentage points increase.

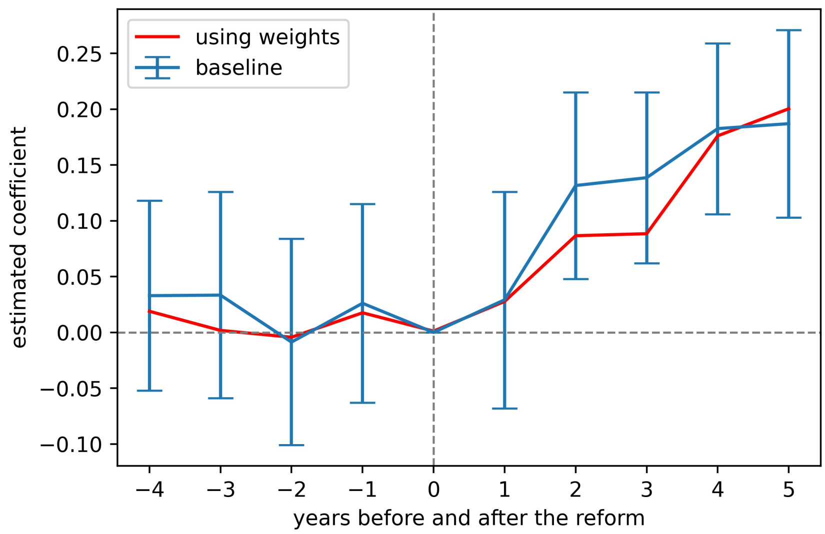


Figure 2.9: The red line depicts yearly effects on high-skill migration from EU-15 countries relative to the year of implementation of the common qualifications framework in Germany, France and the United Kingdom, using an event-study approach focusing on detrended differences and country-skill-year sample-size weights. The dotted lines indicate the baseline for the year of implementation. The blue line is the baseline yearly effects estimation, with vertical bars representing 95% confidence intervals around each estimate.

2.5 Conclusion

As universities and higher-learning institutions across the globe streamline their degrees and requirements, recognizing each other's qualifications, high-skilled labor has become much more mobile. Their education and training can be recognized easily in many countries, which makes it convenient also for more educated workers to find employment in higher-skill sectors in countries where their skill certifications are valid and their expertise is well employed.

In this paper, I show that in the European Union, unlike in the United States, the migration rate has been increasing in the past decades. This rise has been driven mainly by high-skilled workers, while the migration rate of low-skilled workers has been almost entirely flat. I show that this trend is driven by different factors to the ones that shape other countries' migration patterns, and that a crucial driver of this migration is the ease for individuals to have their degrees and professional qualifications recognized in the host country. In particular, in certain cases this can almost double the share of high-skilled migrants. These results are robust to various controls and different specifications, and are notably different across subgroups: women react slightly less than men, while young individuals are much more responsive than the aggregate population.

Given the crucial re-equilibrating role of migration in economic and currency unions, these results are fundamental for policymakers striving to achieve even stronger unity in the EU and better economic performance. Even more so in the aftermath of economic crises, the Bologna Process should be further encouraged and sustained, in order to increase the ease of relocating for workers, especially the high-skilled workers who are more responsive to changing economic conditions and can thus provide a stronger re-equilibrating force.

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Chapter 3

Entering Precariously: Wage Effects of Early-Career Insecurity

3.1 Introduction

The literature has long shown that aggregate insecurity and negative business cycle conditions at graduation have a long-term impact on a series of outcomes for young individuals entering the labour market: unemployment, skill mismatch, and lower earnings to cite a few (e.g. Gregg and Tominey (2005); (Liu et al., 2016); Rothstein (2021)). This makes young individuals one of the demographics most severely impacted by recessions and labour-market crises, such as the recent ones due to the Covid-19 pandemic or the conflict in Ukraine. Part of the reason for this higher cost paid by young labour market entrants is the often temporary nature of their contracts: as firms are hit by a negative shock or are more uncertain about future economic outcomes, they respond to losses by rescinding temporary contracts and reducing permanent hires, disproportionately affecting young workers.

The analysis of temporary contracts among the youth, therefore, becomes of particular relevance for policymakers wishing to address the economic and social effects of downturns, especially in the bigger picture of a declining population trend, which links young individuals' work conditions and salaries to an ever less sustainable pension system, challenged by an aging population, discontinuous careers and shrinking contributory payments. This paper takes a first step in this direction by exploring the dynamics of unstable entry into the labor market, regardless of business cycles. In particular, I focus on internships and apprenticeships, the main entry-level opportunities for young workers, as they are currently at the heart of a lively political debate in Italy.

In Italy the post-graduate internship (*tirocinio extracurriculare*), is intended as a training-work experience hybrid and in theory does not qualify as a job contract. However, its use as an active labour market policy has made it a very common hiring format for many entry-level jobs, increasing noticeably turnover among short-term positions. Recently, for the first time, in Italy internships have surpassed the number of apprenticeships, which are meant to be the main dual contract and represents a stable form of employment (ANPAL data). As Albanese et al. (2021) show, apprenticeships in Italy have been at the center of legislative incentives aimed at strengthening dual learning-working opportunities, which improve job turnover and enhance productivity.

Although they are different institutions, internships and apprenticeships are similarly distributed across sectors and education levels (as shown in the Appendix), so some degree of substitutability between the two contracts does exist. Institutional changes to these labor contracts, therefore, by changing the bargaining relationship between employer and employee, lead to important effects not only on the share of contracts offered of one type versus another, but also on the wage structure of these contracts. While some research has been conducted on the effects of more flexible contracts on the wages of temporary workers (Albanese and Gallo, 2020; Bosio, 2014), little research exists on the effects of institutional changes to the availability of these very short-term contracts on the wages of permanent workers.

Temporary contracts broadly have been the subject of a vast body of economic research (Dolado et al., 2002), following their dramatic rise in the past few decades in many European countries as a result of many institutional reforms. Consequently, many models have tried to describe key features of increasingly dual labour markets, with different approaches as to the coexistence of permanent and temporary contracts. For instance, Cahuc and Postel-Vinay (2002) used a standard search and matching model, simply adding an exogenous fraction of new matches as temporary contracts. Faccini (2014) postulates that all contracts begin as temporary and a share of them must be transformed into permanent contracts. Blanchard and Landier (2002) and Garibaldi and Violante (2005) endogenize the conversion into open-ended contract.

Instead, Garibaldi (2006) and Cao et al. (2010) assume that FTCs have a lower expected duration, but abstract from heterogeneity in workers' ability. Cahuc et al. (2016) show that permanent and temporary contracts coexist in a search market with random matching and wage bargaining, while Berton and Garibaldi (2012) show that temporary and permanent contracts can coexist as long as the job filling rate for permanent contracts is higher. Boeri and Garibaldi (2007) have shown that the attractiveness of temporary contracts takes over most new vacancies and can lead to bursts of sudden job growth. Other papers (Franceschin, 2021) allow for heterogeneous ability of workers and link the coexistence of both types of contract to on-the-job search.

There is consistent evidence showing that young individuals are one of the main demographics employed through temporary contracts Booth et al. (2002), which makes sense as temporary contracts are sometimes used as a "stepping stone" towards a permanent contract. Yet even within young individuals there are trends of differential sorting into temporary and permanent contracts: Portugal and Varejao (2022) show that high-skilled workers are more likely to be employed with a permanent contract, since that worker's ability can be more easily signaled and the firms on average invest more heavily on those workers.

As young individuals enter into the labour market, they often can be hired through an apprenticeship or through a fixed-term internship, as these are the contracts designed to accompany the school-to-work transition. While there is some literature that has focused on internships or on apprenticeship, to my knowledge these two opportunities for young people to enter into the labour market have not been linked in a joint analysis. This is where the present paper comes into effect to prove a broader look into the dynamics of internships and apprenticeships from a contractual bargaining perspective.

Since my goal is to focus on the wage dynamics for young workers, I take different elements from these papers to build a simple model that provides an intuitive theoretical framework to understand the choices that young individuals face. Crucially, in this model workers are heterogeneous in ability, and there is an exogenous probability of a negative shock that causes a loss to the job value for the firm. Due to the short and flexible nature of internships, the employer can terminate it very easily in case of adverse shock and thus only suffers the cost of the shock in the case of the permanent contract.

As workers and employers bargain on the wage and the contractual position of new employees, the increase of temporary work could have two opposite effects on the wages: on the one hand, purely by a selection effect, permanent contracts are offered to the better skilled workers, thus resulting in a higher average permanent-contract wage; on the other hand, workers are willing to compromise on wages in order to achieve more contractual stability. This latter effect can be thought of as a form of wage dumping.

It is an interesting question to determine which of these two effects prevails, especially in light of the growing efforts placed by the European Union on helping young NEETs (i.e. individuals below the age of 30 who are Not in Education, Employment or Training) enter the labour market and obtain high-quality jobs. In particular, internships were the subject of significant institutional changes that occurred in Italy in 2017, as they were expanded to a much wider set of workers. Taking advantage of a staggered regional implementation of the 2017 regulation, I use INAPP's PLUS survey data to estimate the effect that a higher proportion of internships has on apprenticeship wages, in order to determine the direction of the overall effect of increased job instability on stable jobs.

The Italian case is an interesting one to study. In Italy the share of temporary contracts has reached an all-time high (according to Istat monthly occupational data), and Scherer (2005) shows that Italy presents a longer spell of instability before young individuals can find a stable job, compared to other EU countries. Using a difference-in-differences setup as well as a triple difference estimation, I find that the increase in the share of internships that employers can offer has a negative effect on the wages of apprenticeships, despite a selection effect at play which I disentangle using the distribution of earnings for treated Regions. The effect is robust despite a rise in the number of apprenticeships across the board, which is however more pronounced in Regions where internships are liberalized. There is a slight wage premium on internships, yet the most significant effect is on permanent contracts.

These results are consistent with the existing wage-premium literature: Albanese and Gallo (2020) notably find a wage premium on temporary contracts for new hires, after controlling for observable characteristics; at the same time this confirms descriptively the inequality documented in the wage distribution literature (Bosio, 2014), supporting the notion that employers tend to offer permanent contracts to the most skilled workers, but at the same time suggesting that at the margin there is a stability-wage bargaining mechanism at play.

The framework used in this paper provides an analytical intuition of the main mechanisms at play, which could be extended to provide intuition about other short-term outcomes. For instance, job precariousness can have implications on social variables such as fertility and likelihood of leaving the parents' home, as well as other fatures of young individuals economic and social independence which could be explored empirically. Important precedents of this kind exist: Krahn et al. (2015) find that early employment instability was linked to lower income at age 32 and, among Canadian men, to lower occupational status and career satisfaction. Bertolini and Goglio (2019) find that these dynamics have a crucial impact on living independence from parents. Similarly, Virtanen et al. (2005) find higher likelihood of mental health issues among temporary workers.

A careful use of temporary contracts for labour-market entrants, therefore, is of crucial importance, and policymakers should make use of these insights to make sure that entry-level contracts represent a true opportunity for young individuals to achieve independence rather than a source of increased precariousness. The rest of the paper is structured as follows: Section 3.2 presents the model and its main implications, while Section 3.3 describes the Italian institutional setup and explores the effects of the 2017 reform of internships. Section 3.4 tests the robustness of the results and finally Section 3.5 concludes and suggests avenues to expand the current research.

3.2 Theoretical Model

In this section, I present a simple model which builds upon the traditional search and matching framework, allowing for the coexistence of both temporary and permanent contracts. I focus on the firms' preferences and the workers' preferences, finding the room in both agents' value functions for wage bargaining and for the response to an increase in the share of temporary contracts that can be offered. I present some key comparative statics analysis and make a set of predictions that result in two opposing effects on the wage of permanent contracts, which can be then tested in the estimation.

3.2.1 General Setup

I model a simple economy in a search and matching framework, characterized by workers who all participate in the labour force and identical firms that only use labour as an input. All parties are risk-neutral.

Workers are heterogeneous in skill, with an idiosyncratic productivity α that is distributed according to the cumulative distribution $F(\alpha)$. A worker's productivity is fully known ex-ante by all agents, and the wage reflects this productivity. Furthermore, the firm uses a constant-returns-to-scale technology to produce output y, which is also increasing in the productivity and always larger than the wage for a given level of ability:

$$\frac{\partial y}{\partial \alpha} > \frac{\partial w}{\partial \alpha} > 0 \tag{3.1}$$

This way the firm is always making an operational profit on a worker, and the profit is also an increasing function of productivity α. Two types of contracts exist in the economy: temporary contracts and permanent contracts. I allow for the possibility that, for a given level of productivity, the wage of a permanent contract $w_p(\alpha)$ and the wage of a temporary contract $w_t(\alpha)$ might differ, thus allowing for the possibility of a wage premium as described in the literature recalled above. Employed workers are subject to natural turnover and separate from their existing job with a Poisson process characterized by an arrival rate that differs across temporary and permanent contracts: the separation rate for temporary contracts s_t, due to their shorter duration, is larger than the separation rate for permanent contracts s_p.

The contracts differ additionally to the employer because with an instantaneous probability λ the firm undergoes an adverse shock which causes a loss κ to the employer. In the case of the temporary contract the employer can terminate without any cost the contract or, due to its short duration (i.e. a high separation rate), can simply wait for the contract to expire without replacing it with a new worker. Thus, in case of the shock λ, the cost κ is only paid for permanent contracts.

Firms create jobs by posting costly vacancies and can freely decide to offer either temporary or permanent jobs to a worker of observed productivity α. Keeping open a vacancy, regardless of the type of contract, involves a flow cost equal to c. For simplicity, I assume that the vacancy cost is identical for both contracts. The meeting of unemployed workers and vacant firms is described by a matching function m with constant returns to scale, which reflects matching frictions. Let $m(u, v)$ be the flow of new matches, where u denotes the measure of unemployed workers that are searching and can fill vacancies v posted by the firms. As standard, assume that m is concave and homogeneous of degree one in (u, v) with continuous derivatives. Consequently define $m(u, v)/v = q(\theta)$ as the arrival rate of workers for a vacancy and $h = m(u, v)/u = m(1, \theta) = h(\theta)$ as the transition rate from unemployment to employment, where $\theta = v/u$ is the labour market tightness.

To simplify the model, and square it with the high the levels of unemployment which often give employers high bargaining power, I assume that after matching the choice of contract is in the hands of the employer. Based on the worker's productivity α, the employer will offer a contract either permanent or temporary. Call $h_t(\theta, \alpha)$ (henceforth $h_t(\alpha)$ for simplicity) the probability that a worker of ability α receives a temporary job offer and $h_p(\theta, \alpha)$ (simplified as $h_p(\alpha)$) the probability that she receives the offer for a permanent one, such that $h_t(\alpha) + h_p(\alpha) = h(\alpha)$ for all values of α. Once the offer is received, the worker can either refuse it, accept it as it is or suggest a change in the kind of contract that is offered. At that point the employer will accept the counter-offer only if it as convenient as or more convenient than the initial proposal. The wage agreed upon in this process is fixed for the entire employment relationship without ex-post renegotiation. Any wage within the parties' bargaining set can be supported as an equilibrium.

The equilibrium of the model is characterized by free entry of firms, which drives the flow value of vacancies to zero. Additionally, if workers are out of work, they actively search for a job and enjoy a fixed exogenous benefit $b > 0$ as an outside option.

3.2.2 The Firm's Decision

The firm maximizes profits, by posting vacancies that have a cost c, assuming that the process is equally costly for a temporary and a permanent job. The Bellman equation describing the present discounted value of a vacancy to the firm is:

$$rV = -c + q(\theta)[\mathbb{E}(J_k(\alpha)) - V]$$

where $k \in \{t, p\}$ for, respectively, temporary and permanent contracts and

$\mathbb{E}(J_k(\alpha))$ is the expected continuation value from a new match given the distribution of α, which is known to the firm. Once a match is formed, the firm observes the value of α of the matched worker and chooses the type of contract accordingly to its preferences described below. The value of a job instead differs across the two types of contract. The value of a temporary job is the following:

$$rJ_t(\alpha) = y(\alpha) - w_t(\alpha) + s_t[V - J_t(\alpha)]$$

Instead the value of a permanent job is the following:

$$rJ_p(\alpha) = y(\alpha) - w_p(\alpha) + s_p[V - J_p(\alpha)] - \lambda\kappa$$

where κ is the cost of a negative shock to the value of a permanent job that occurs with probability $\lambda \in (0, 1)$. Therefore, to a worker of given α, a firm would strictly prefer to offer a temporary contract if $J_t(\alpha) > J_p(\alpha)$, which means:

$$\lambda\kappa > \underbrace{w_t(\alpha) - w_p(\alpha)}_{\text{wage premium}} + \underbrace{s_p[V - J_p(\alpha)] - s_t[V - J_t(\alpha)]}_{\text{relative separation cost}} \tag{3.2}$$

i.e. the firm will offer a temporary contract if the risk associated with the permanent contract is greater than the wage premium to be paid to the temporary worker and the relative cost of separation for a worker of productivity α in a temporary contract relative to a permanent one. This condition is important as, given the set of parameters β and the growing profits 3.1, it shows that there exists a value of productivity $\tilde{\alpha}$ above which the firm is willing to offer a permanent contract, as the cost of losing that worker due to the greater separation rate of a temporary contract (the relative separation cost) would be greater than the possible cost associated to the adverse shock in a permanent contract. Below this threshold productivity value, the firm prefers to offer a temporary contract. Additionally, it shows that there is some wiggle room for workers to negotiate on the wage premium.

Suppose, however, as it is the case in reality, that there is a legislative requirement setting a cap on temporary contracts τ, with τ being such that $F(\tilde{\alpha}) > \tau$, then the constraint is binding, and the firm will offer a temporary contracts to all workers with a productivity $\alpha \leq F^{-1}(\tau)$.

As a result, the probabilities of receiving a temporary and permanent job offer h, for a worker of a given productivity α, depend also on the legislative limit τ, and are influenced by it in opposite ways:

$$\frac{\partial h_t(\alpha, \tau)}{\partial \tau} > 0 \text{ and } \frac{\partial h_p(\alpha, \tau)}{\partial \tau} < 0 \tag{3.3}$$

Additionally, the separation rate for permanent contracts must be a function of τ as well. In fact, when companies are forced to offer permanent contracts to some workers whom they would like to only keep temporarily, they will be more likely to terminate those contracts early. In the case of the apprenticeship, it means that firms will be more likely to exercise the termination option at the end of the training period. Thus we also have that:

$$\frac{\partial s_p(\tau)}{\partial \tau} < 0 \tag{3.4}$$

The separation rate for temporary contracts s_t, instead, will not be affected by this change, since the maximum duration of temporary contracts is fixed legislatively, in Italy at a maximum value of 12 months, 24 in special circumstances. This limit is left untouched by a legislative change to the share of temporary contracts that a firm can activate.

As free entry drives the value of vacancies to zero ($V = 0$), we can plug in the equilibrium values of temporary and permanent jobs into 3.2 to get the value of the permanent contract wage that makes the firm indifferent between a temporary and a permanent contract:

$$w_p(\alpha)^f_{max} = \frac{r+s_p}{r+s_t}w_t(\alpha) + \frac{s_t-s_p}{r+s_t}y(\alpha) - \frac{1-s_p}{r}\lambda\kappa$$

This represents the maximum value of $w_p(\alpha)$, given an initial offer for a temporary contract with wage $w_t(\alpha)$, above which the firm prefers to maintain a temporary contract and below which it is willing to switch to a permanent contract. From this we can derive the value of the wage premium at which the firm is indifferent between temporary and permanent contract, which is:

$$w_t(\alpha) - w_p(\alpha) = \frac{1-s_p}{r}\lambda\kappa - \frac{s_t-s_p}{r+s_t}[y(\alpha) - w_t(\alpha)]$$

This shows that there exists a measure of workers with productivity $\alpha \in [\underline{\alpha'}, \overline{\alpha'}]$ for whom the firm would prefer to offer a temporary contract if $w_p(\alpha) \geq w_t(\alpha)$; but is willing to agree to a permanent contract if there is wage premium large enough between temporary and permanent contract wages. Such a wage premium must compensate for the difference between the adverse shock risk associated with permanent contracts and the loss of profit that the firms suffers because of the higher separation rate of a temporary contract. Since firm's profits are increasing in the worker's α due to 3.1, the above expression shows that the wage premium that a firm requires in order to agree on a permanent contract is decreasing in α. Therefore, above the threshold level of ability $\tilde{\alpha}$ the firm will require no wage premium and would spontaneously be willing to offer a permanent contract to all workers with $\alpha > \tilde{\alpha}$.

3.2.3 The Worker's Decision

On the worker's side, the Bellman equations for the expected value of a job offer, if accepted, is:

$$rE_k(\alpha) = w_k(\alpha) + s_k[U - E_k(\alpha)]$$

where $k \in \{t, s\}$ depending on whether the contract is temporary or permanent. The expected value of unemployment, instead, is:

$$rU = b + h_t(\alpha, \tau)[E_t(\alpha) - U] + h_p(\alpha, \tau)[E_p(\alpha) - U]$$

where due to equations 3.1 and 3.2, we can say that:

$$\frac{\partial h_t(\alpha, \tau)}{\partial \alpha} < 0 \text{ and } \frac{\partial h_p(\alpha, \tau)}{\partial \alpha} > 0$$

Given that $s_p > s_t$ and that market viability requires $E_k > U$, in the absence of a wage premium (i.e. if $w_t(\alpha) = w_p(\alpha)$) any worker prefers a permanent contract. Specifically, the worker would prefer the permanent contract as long as $E_p(\alpha) > E_t(\alpha)$, i.e. as long as:

$$w_p(\alpha) + s_p(\tau)[U - E_p(\alpha)] > w_t(\alpha) + s_t[U - E_t(\alpha)] \tag{3.5}$$

The value of the permanent contract wage that would make the worker indifferent between a permanent and a temporary contract is:

$$w_p(\alpha) = \frac{r + s_p(\tau)}{r + s_t}[w_t(\alpha) + s_p(\tau)U] - s_p(\tau)U \tag{3.6}$$

A reduction in the value of unemployment, therefore, caused by an increasing precariousness in the labour market with a lower share of valuable permanent contracts, will have a negative effect on the permanent contract's wage:

$$\frac{\partial w_p(\alpha)}{\partial U} = s_t \frac{r + s_p(\tau)}{r + s_t} - s_p(\tau) > 0$$

where the above condition and its positive value stems from the fact that $s_t > s_p$.

More precisely, by substituting the value for temporary and permanent contracts into the expression for the value of unemployment, the value of the permanent wage at which the worker is indifferent between a permanent and a temporary contract can be calculated as:

$$w_p(\alpha)^w_{min} = \frac{r + s_p(\tau) + h_t(\alpha, \tau) + h_p(\alpha, \tau)}{r + s_t + h_t(\alpha, \tau) + h_p(\alpha, \tau)} w_t(\alpha) + \frac{s_t - s_p(\tau)}{r + s_t + h_t(\alpha, \tau) + h_p(\alpha, \tau)} b \tag{3.7}$$

The previous condition shows how the worker's attitude towards a temporary versus a permanent contract substantially depends on the relationship between the wages of the two contracts, mediated by the separation rates and the outside option b. If the separation rates were equal, then the two wages would naturally coincide, i.e. there would be no difference between temporary and permanent contracts. If instead the separation rates are indeed much lower for permanent contracts, $s_p < s_t$, then the worker would be willing to accept a lower wage for a permanent contract, which is just a fraction of the temporary-contract wage plus a fraction of the outside option that is recuperated by temporary workers in the more likely case of unemployment spells.

So the maximum value of the wage premium for a worker will be:

$$w_t(\alpha) - w_p(\alpha) = \frac{s_t - s_p(\tau)}{r + s_t + h_t(\alpha, \tau) + h_p(\alpha, \tau)}[w_t(\alpha) - b]$$

In other words, the maximum wage premium for the worker is a function of the difference between the temporary contract wage and the outside option, which is the value that would be lost in the case of unemployment following a likely separation from a temporary worker. And in fact, this difference is mediated by the difference between the separation rates, i.e. the additional likelihood that indeed the worker would face unemployment by accepting a temporary contract instead of a permanent one.

3.2.4 Equilibrium and Comparative Statics

The matching process thus is the following: if a worker of a given productivity α is matched with a firm, and the value of α is such that 3.2 holds, then the worker will be offered a temporary contract with wage $w_t(\alpha)_{offer}$. If however, there exists a value (or set of values) $w_p(\alpha)_{counter}$ low enough to break equation 3.2 but at the same time high enough such that $w_p(\alpha)_{counter} > w_p(\alpha)^w_{min}$, then the worker will counter-offer a permanent contract with a lower wage $w_p(\alpha)_{counter}$ and the employer will accept. In other words, the switch from temporary to permanent occurs if $w_p(\alpha)^f_{max} > w_p(\alpha)^w_{min}$. This mechanism, therefore, gives rise to the wage premium of temporary-contract wages with respect to permanent-contract wages.

Having derived the job creation and worker preference conditions, we can define the equilibrium in this model as the set of conditions where:

1. the matching function $m(u, v)$ is verified;

2. free entry drives the value of vacancies to zero, so that:

$$\frac{c}{q(\theta)} = \mathbb{E}(J_k(\alpha))$$

3. firms and workers maximize their value functions;

4. wages and contracts are bargained to the wage premia conditions identified;

5. the share of temporary contracts in each firm is at most τ.

How is this process affected by a change in the legislatively imposed maximum share τ of temporary contracts that a firm can offer? As described above, due to conditions 3.3 and 3.4, an increase in the firm's share of temporary contracts is reflected in three variables: 1) an increase in the job offer rate for temporary contracts; 2) a decrease in the job offer rate for permanent contracts; and 3) a decrease in the separation rate of permanent contracts s_t.

The effect of these three changes on 3.7 is ambiguous, as it depends on the relationship between h_p and h_t. However, we do know that it will be certainly positive if $\frac{\partial^2 h_t}{\partial \tau^2} < \frac{\partial^2 h_p}{\partial \tau^2}$, which is not an unreasonable assumption. For workers at the margin, the arrival of temporary contract offers is likely already high, as that is the firm's preferred contract. Thus it is reasonable to think that a larger share of temporary contracts reduces the likelihood of receiving a permanent contract offer by more than it raises the likelihood of receiving a temporary contract offer.

On top of this reasoning, in fact, we can focus on the effect of the rise in s_t, which can be summarized with a simple comparative static of the form:

$$\frac{\partial w_p(\alpha)}{\partial s_p} = \frac{w_t(\alpha) - b}{r + s_t + h_t(\alpha) + h_p(\alpha)} > 0 \tag{3.8}$$

where the sign of the previous is positive for all workers who consider the offer, i.e. for all those who are offered a wage higher than the outside option. Thus, as the separation rate of permanent contracts decreases further away from that of temporary contracts, the worker is willing to reduce even further the wage for a permanent contract relative to the temporary one, in exchange for the increased stability of a permanent job.

The combination of these considerations leads to a negative overall effect on permanent-contract wages of the rise in the share of temporary contracts τ. This creates – or exacerbates of already present – the wage premium, representing a key prediction of the model which can be tested in the data.

3.2.5 Model Predictions

From the implications of the simple model outlined above, we can make a few key inferences. We can predict that, if a legislative liberalization allows employers to increase the share of workers that they can hire through temporary contracts (i.e. an increase in τ, reflected in a decrease in s_p), then we will observe that:

- the number of permanent contracts decreases;
- the number of temporary contracts increases;
- the types of contracts that shift from permanent to temporary are mainly the ones located in the lower end of the productivity (and thus wage) distribution;
- the minimum level of wage for a permanent contract that a worker is willing to accept decreases;
- the individual wage premium, conditional on observables, widens.

The phrasing of the last two predictions is of particular importance. First of all it is important to note that, by symmetry, the concept of wage premium could arise also due to firms wishing to bargain more for temporary contracts and offering higher wages for temporary contracts in exchange. However, I consider this less important in the context of the present paper for two reasons: firstly, the legislative limit on temporary contracts likely squeezes firms very close to the threshold and, especially if the constraint is binding, does not allow firms to offer additional temporary contracts. In Italy, a legislative limit of this kind was introduced to limit the ratio of fixed-term contracts over permanent-contracts to 20% for each employer.[1] Istat data suggests that in January of 2018, the total number of fixed-term workers over the number of permanent workers was about 19.8%, while in January 2022 it hovers at 20.3%.

Secondly, a rise in the share of allowed temporary contracts, in an environment where workers prefer permanent contracts, gives more bargaining power to the employers and less to the workers. Thus a widening in the wage gap in the legislative case considered in this paper is much more likely to be due to workers giving up some of their surplus in exchange for a permanent contract than to workers giving up their share of surplus to offer a high-wage temporary contract. While I don't exclude this latter option as a possibility, and test for it in the estimation section, I consider it secondary and focus on the effect on the wage of permanent contracts.

Additionally, the phrasing of the last two predictions is carefully tailored around individual effects. Thus far, the model has focused on the effects of a given worker of productivity α. When shifting the focus from the dynamics of a single contract's

[1]There are some exceptions and specific sector-wise implementation differences, which make in practice this limit not exactly 20% for all firms, but at the aggregate level very close to it. Additionally, there are some kinds of non-permanent contracts that do not fall under this limit, for instance seasonal work.

bargaining process towards an analysis at the aggregate level, as the estimation will do, interesting additional effects emerge. In particular, the third prediction suggests that jobs at the low end of the permanent-contract wage distribution are the main ones to make the switch from permanent to temporary contracts for new hires. This results in a selection effect that puts upward pressure on the average wage of new permanent contracts.

Therefore, following a rise in the share of temporary contracts, the overall effect on the average wage of permanent contracts is ex-ante ambiguous. It could be driven by one of two opposing effects:

1. a **selection effect**, which suggests that only the most productive workers are offered a permanent contract, thus increasing the average permanent-contract wage;

2. a **dumping effect**, which suggests that workers are willing to lower their wage in order to secure a permanent contract, thus reducing the average permanent-contract wage.

In the following estimation section, I take to the data to determine which of these two effects has prevailed in the Italian empirical experience.

3.3 Empirical Assessment: Internships vs. Apprenticeships

In this section I test the predictions of Section 3.2 on the Italian context, where the two main learning-working opportunities for young individuals entering the labour market are internships and apprenticeships. First I present some background on the Italian institutional setup, then I show the effect of the 2017 reform that liberalized internships, both on the average earnings from internships and from apprenticeships, as well as on the distribution of apprenticeship earnings, and present a counterfactual analysis to disentangle the main effects at play.

3.3.1 The Institutional Setup

In Italy, the standard labour contract is the open-ended contract (*contratto a tempo indeterminato*). There are then fixed-term contracts (*a tempo determinato*), seasonal contracts and on-call contracts to name a few temporary alternatives. However, there exist two specific contracts for young individuals designed to facilitate the matching between a young worker and an employer, including a component of training in the contract to reduce the skill mismatch.

The main one of these is the apprenticeship contract. Formally, apprenticeships are open-ended contracts characterized by an initial dual training-working period that can last anytime from six months to three years – up to five years for specific artisan crafts. At the end of the training period, if the worker has not successfully acquired the desired skill, the employer has the option to rescind the contract without any penalty. The apprenticeship is a long-standing tradition of the Italian labour market, which has therefore been revisited several times to better align it with the developments in the world of work. Currently, the apprenticeship is regulated by the Legislative Decree n. 81 of 2015, which deeply revised the Italian labour legislation and includes a comprehensive set of baseline requirements. It is then up to collective bargaining agreements to further and more specifically discipline apprenticeships, so that each sector and job has tailor-made apprenticeship features – including duration, amount of external vs. internal training, tasks, etc.

In addition, Legislative Decree 81/2015 includes substantial incentives – both in terms of salary and social security contributions – for employers who use apprenticeships, intended to increase the reliance on this kind of open-ended contract. Since 2015, therefore, the number of apprenticeships has been increasing in Italy.

At the same time, another instrument that has seen a dramatic increase is the extracurricular internship (in Italian *tirocinio extracurriculare*, sometimes also referred to as *stage*). [2] Designed to help young individuals get closer to the labour market and make useful experiences by learning a profession, the internship does not qualify as a work contract. It is instead a training contract where firm and worker commit to a learning on the job setup. This makes it much faster to activate and easy to manage, as it is free from all the requirements that are typical of a job contract – including collective bargaining and severance payments. Thus employers can easily terminate the internship or, given its short duration, simply not hire the worker at the end of the internship. This flexibility and the low labour costs associated with it make the internship a very attractive tool for employers. In many firms, therefore, entry-level positions have been increasingly filled by internships.

Since formally it is configured as a training opportunity rather than a job, the internship in Italy is constitutionally a matter of Regional regulation. The national Government and the 21 Regions and autonomous Provinces lay out periodically common guidelines that form the basis for internship regulation, which then have to be implemented by regional law in each Region or autonomous Province [3]. With a set of guidelines issued in 2013, Regions had agreed on several baseline features of the internship, including the limit to use the extracurricular internship mainly for

[2]Internships are formally divided into two kinds: curricular, if it takes place as part of the curriculum for a formal educational qualification; and extracurricular, or post-graduate, otherwise. In this paper, I only focus on the extracurricular kind, as it is the most common one and more frequently used as a *de facto* entry-level job contract.

[3]Italy is formally divided into 20 Regions, however the region Trentino-Alto Adige has a special status whereby it is divided into two autonomous Provinces. Thus in matters of governmental coordination, rather than 20 Regions there are 19 Regions and 2 autonomous Provinces. For the sake of brevity, henceforth I shall refer to them as simply "the Regions"

those who had just graduated or for disadvantaged groups.

In 2017, a new set of guidelines was agreed to by the Regions and the Government. These new guidelines mainly liberalized the use of the internship, most notably removing the limit restricting its use mainly for recent-graduates. Specifically, the 2017 guidelines stated that beneficiaries of internships could be, among other categories, all workers who were unemployed or even "at risk of unemployment". Since no further clarification was provided concerning what exactly qualifies as risk of unemployment, the effect was that anyone could effectively be offered an internship. This novelty increased considerably the share of potential workers – particularly young ones – whom a firm could "hire" through an internship. Additionally, *Garanzia Giovani*, the Italian Youth Guarantee scheme, i.e. a set of EU funding aimed at reducing youth unemployment and inactivity rates, greatly contributed to the rise of internship use; in fact, Regions would pay some of the remuneration of the internship alongside the firm, thus reducing further the effective cost of internships for employers.

The new guidelines, however, in order to become effective needed to be implemented by each Region through a regional law or regional government decree. This happened in a staggered fashion, thus creating the perfect setup for a quasi-experimental approach. 10 Regions implemented the new guidelines already in 2017, just a few months after the agreement: these are Piemonte, Val d'Aosta, Liguria, Trentino, Veneto, Marche, Lazio, Basilicata, Calabria, Sicilia. 5 Regions implemented the new guidelines throughout 2018: Lombardia, Friuli Venezia Giulia, Toscana, Abruzzo, Campania. 5 Regions implemented the new guidelines in 2019 or did not implement them at all (it is the case of Puglia, who decided to keep its legislation tied to the 2013 guidelines).

3.3.2 Data and Estimation

The staggered approach with which the Regions implemented the 2017 guidelines on internships allows to define two groups: a set of Regions that were treated in 2017, i.e. Regions in which in 2018 internships were available for any worker regardless of their academic background and timing; and a set of Regions in which in 2018 the reference legislation was still reflective of the 2013 guidelines and thus had stricter application rules. In the main specification, I consider the regions that implemented the new guidelines in 2018 as not treated, since throughout 2018 most of their legislation was still based on the 2013 guidelines, the new legislation came into effect towards the end of the year and only affected new activations, so that most sampled internships had been activated under the previous legislation. However, in Section 3.4 I check the robustness of the estimation under this assumption by excluding the Regions that implemented the guidelines in 2018, only leaving as control group the Regions that implemented them in 2019 or did not implement them at all.

This legislative set up yields a treated set of Regions and a control group made up of untreated Regions. This makes it a perfect setting to test for a pre/post reform treatment effect. A good dataset for this purpose is constructed by the *Istituto Nazionale di Analisi delle Politiche Pubbliche* (INAPP) and it is called the "Participation, Labour, Unemployment Survey" (PLUS for short). It is a recurring interview-based survey that focuses on labour-market outcomes for workers, and it includes key information for my intended analysis – such as age, earnings, type of contract, education level, region of residence, etc. What is particularly convenient of this dataset is that its two most recent waves were carried out precisely in 2016 and 2018, thus one year before and after the new guidelines on extracurricular internships.

Using these surveys, therefore, I can construct a dataset that only focuses on respondents doing an internship and respondents doing an apprenticeship (as well as those with an open-ended contract for the robustness checks in Section 3.4) in 2016 and 2018. The resulting dataset is made up of 1130 observations; it represents only a small fraction of the young individuals actually involved in apprenticeships and internships in those two years, yet it is useful to provide some key insights. Additionally, INAPP goes to great lengths to ensure the representativeness of its sample, and more information about the sampling procedure can be found on their website. Of course a larger sample, or better yet the use of administrative data, would greatly improve the reliability of the estimation – however, the training nature of internships (which are not subject to social security contributions) makes it difficult to gather the necessary data in one database and would require multiple sources of data to be accessed. This is a daunting task given current Italian privacy laws, and I leave it to future research to improve the level of detail about internship data. To my knowledge, the PLUS survey is the only reliable micro-level source of income data for internships in Italy.

A final necessary clarification is that the setting explored in this section is not a closed system: internships and apprenticeships are not the only two types of contract that can be offered to young individuals entering the labour market and thus do not cover the entirety of young workers, so flows from one contract do not necessarily go to the other. There are other contracts available: nothing prevents an employer from hiring a young worker, say, directly through a regular open-ended contract or through a seasonal contract. However, these are the only two contracts that explicitly include training and that are specifically geared at young individuals, and for this reason also significantly lower the cost of labour for employers compared to other contracts. Thus they are much closer substitutes to each other than to other contracts.

Given the national incentives on apprenticeships and the Garanzia Giovani funding on internships, both these contracts have grown in numbers in the 2016-2018 time frame (data from the Italian Ministry of Labour and Social Policies and ANPAL, the national Agency for active labour market policies). However, there are important differences across regions. Between 2016 and 2018, Regions that adopted the new internship guidelines in 2017 had an average growth rate of internships that is 65% greater than the late-adopters; the late adopters, on the other hand, in the same time period had an average apprenticeship growth rate 48% larger than 2017-adopters. There is therefore reason to believe that, while both contracts grew in the 2016-2018 time frame, the liberalization of internships caused a shift of a considerable share of new contracts from the apprenticeship towards the internship.

Thus the first prediction of the model, which is that temporary contracts should grow at the expense of permanent ones, finds substantiation. The next question, at the heart of the paper, concerns the effect on the wages of apprenticeships.

3.3.3 Effect on Apprenticeships

With the necessary clarifications of the previous paragraphs, I proceed to test the effect of the 2017 internship reform, which through the new guidelines liberalized the use of extracurricular internships. First and foremost, I explore the effect on the wages of apprenticeships. As shown in the model, the two effects at play are a selection effect and a dumping effect: the former should bring up the average wage for apprenticeships, the latter should bring it down.

To estimate which effect is prevalent, I first focus on the subset of the PLUS data of only those working through an apprenticeship and I run a difference-in-differences estimation of the form:

$$W_{i,r,t} = \delta_t + \beta(\mathbb{1}_{treat}[r] \times \delta_t) + \phi_r + \epsilon_{i,r,t}$$

where $W_{i,r,t}$ is the earnings of individual i in region r in year t, ζ is a constant intercept (which I drop when I include fixed-effects), $\mathbb{1}_{treat}[r]$ is a dummy variable taking a value of 1 if the region is in the treated group and 0 otherwise, and δ_t is an indicator taking value of 0 for 2016 observations and 1 for 2018 observations. In some estimations I also control for regional fixed effects, in which case I will include ϕ_r as a set of regional fixed-effects and not include a constant term. The coefficient of interest is β, which identifies the effect on the year-treatment interaction variable, thus estimating the average change in earnings in 2018 for regions where the internship reform was implemented.

Table 3.1 shows the results of this estimation. As is apparent from the results, from 2016 to 2018 there is an increase in the earnings of all apprenticeships of around 4500 euros per year, however this change is not distributed evenly across Regions: in Regions where the internship reform was implemented early (i.e. 2017-adopters), the average earnings are lower by about 2500 euros than the other Regions (the late-adopters). Column I of the table is conducted without Region fixed-effects and thus contains a "treated" variable, i.e. the variable γ_r in the equation above. Column II instead includes Region-specific fixed-effects and thus I drop γ_r from the estimation. In both columns, however, the effect has the same negative direction and a similar magnitude, in both cases above 2000 euros.

This gives reason to believe that indeed there might be a dumping mechanism at play, where, faced with a more likely prospect of a precarious contract such an internship, workers are willing to accept a permanent contract albeit at a lower wage. A selection effect might or might not be present, either way the dumping

Table 3.1: Effect on apprenticeship earnings

	Apprenticeship earnings I	Apprenticeship earnings II
Treated	1,572.24**	
	(770.68)	
Year2018	4,494.32***	4,532.18***
	(744.27)	(750.47)
Treated*Year2018	-2,318.87**	-2,506.06**
	(1,153.61)	(1175.38)
Intercept	14,837.07***	
	(507.87)	
Region-FE	No	Yes
R-squared	0.0528	0.0767
R-squared Adj.	0.0491	0.0523
F-statistic	14.365	3.142
N. of Observations	777	777

Difference-in-differences estimation of the impact of internship reform on the earnings of apprenticeships. Treatment is defined as belonging to the set of regions that implemented the reform, taking a value of 1 if the reform was implemented in 2017, and 0 otherwise. Column I does not control for regional fixed effects, while column II does. The data is taken from INAPP's PLUS survey, waves 2016 and 2018. Standard errors are in parentheses. Stars indicate p-values according to: * p<0.1, ** p<0.05, ***p<0.01

effect is prevalent. In the coming sections I will also attempt to disentangle these two effects for a more precise picture.

3.3.4 Effect on Internships

Next, although not the main purpose of this paper, it is reasonable to question whether the wage premium might arise also from a symmetrical reasoning on the temporary-contract side – i.e. if some firms might "compensate" workers by offering a higher wage in exchange for the flexibility of a temporary arrangement. To verify this, I run a similar procedure as the one described above, however with only respondents who reported being employed through an internship. The results are listed in Table 3.2.

As the table shows, the general rise in earnings between 2016 and 2018 is about the same for internships as it was for apprenticeships, however in this case the effect of the reform is much less clear. Although there seems to be a rise in internship earnings for young workers in Regions that implemented the reform in 2017, the effect is not significant and therefore the hypothesis of a wage premium driven by the rise in precarious contracts' wages does not find sufficient substantiation. A possible reason for this is that the internship reform mostly benefits firms by relaxing a constraint on their preferences, employers gain bargaining power, therefore there is not as much need for them to offer some of their surplus share in exchange for their preferred contract. A small selection effect, which should also put upward pressure on internship wages, might be at play, however again the lack of significance of the coefficient does not provide much clarity on this. Part of the reason for it might be the smaller sample of internship respondents in the PLUS survey compared to the apprenticeship, so perhaps the use of administrative data will be able to identify a more precise effect.

Table 3.2: Effect on internship earnings

	Internship earnings I	Internship earnings II
Treated	200.30	
	(974.28)	
Year2018	4,707.96***	4,717.77***
	(1,304.79)	(1,316.84)
Treated*Year2018	850.62	1,158.92
	(1,796.20)	(1,860.87)
Intercept	6,602.17***	
	(731.75)	
Region-FE	No	Yes
R-squared	0.1096	0.1469
R-squared Adj.	0.0998	0.0940
F-statistic	11.125	2.776
N. of Observations	275	275

Difference-in-differences estimation of the impact of internship reform on the earnings of internships. Treatment is defined as belonging to the set of regions that implemented the reform, taking a value of 1 if the reform was implemented in 2017, and 0 otherwise. Column I does not control for regional fixed effects, while column II does. The data is taken from INAPP's PLUS survey, waves 2016 and 2018. Standard errors are in parentheses. Stars indicate p-values according to: * $p<0.1$, ** $p<0.05$, *** $p<0.01$

3.3.5 Triple difference

Next, I focus more specifically on the rise in average earnings from 2016 to 2018. Such a high rise in wages might seem suspicious at first. Certainly those were years in which the economy was growing and inflation was rising, all coupled with a series of renewals of collective contracts, which often include wage rises for workers – including apprentices. However, it is worth questioning whether some other apprenticeship-specific mechanism might be driving the rise in average earnings from 2016 to 2018 and therefore bias the estimation.

To do this, I take advantage of PLUS's variety of respondents and detail of information. Thus far I have only used in the regressions respondents involved in internships and apprenticeships, who represent a small fraction of the sample. To test for an overall economy-wide change in earnings, I include in the estimation also workers hired through regular open-ended contracts, i.e. without the training component of the apprenticeship. This could raise the issue of substantial worker difference: internships and apprenticeships concern young workers, while open-ended contracts include all age groups, who might experience very different earnings trends. Therefore, I restrict the sample to young workers in open-ended contracts, below the age of 30.

Therefore, focusing on the subsample of open-ended contracts aged 16-30 and on apprenticeships, I can run a triple-difference estimation. I control not only for the year and the residence in a treated Region, but also for the type of contract used and the related interactions. The resulting estimation is the following:

$$W_{i,r,t} = \delta_t + \phi_r + \kappa_s + \beta_1(\mathbb{1}_{treat}[r] \times \delta_t) + \beta_2(\kappa_s \times \delta_t) + \beta_3(\mathbb{1}_{treat}[r] \times \kappa_s) \\ + \beta_4(\mathbb{1}_{treat}[r] \times \delta_t \times \kappa_s) + \epsilon_{i,r,t}$$

where κ_s represents a dummy for the type of contract employed, which takes a value of 1 if the contract is an apprenticeship and a value of 0 if it is a simple open-ended contract. The coefficient of the triple interaction β_4 is the focus of the estimation now as it captures the effect of the reform on apprenticeships in treated Regions.

Table 3.3 shows the coefficients resulting from the estimation. As before, column I does not contain Region fixed-effects and therefore includes a dummy variable for the treatment or control group; column II instead uses Region fixed-effects, thus the treatment variable is omitted due to collinearity with the time-unvarying fixed-effects.

The Table shows a few interesting results. First of all, the average earnings increase from 2016 to 2018 for all contracts, including open-ended contracts. Thus there is nothing unique to apprenticeships driving the coefficient on "year2018", which has about the same magnitude as in the previous estimations. There is also no significant difference in the way that earnings change for apprenticeships versus open-ended contracts, as the coefficient on the interaction term "apprenticeship*year2018" is not significant. There is a slightly larger rise in earnings, for all contracts, in treated Regions (2017-adopters) rather than untreated ones. Additionally, as it would be expected, apprenticeships pay a bit less on average than regular open-ended contracts; and this effect is not particularly different across treated and untreated Regions.

Table 3.3: Triple Difference Estimation

	Earnings I	Earnings II
Treated	-749.29	
	(548.04)	
Year2018	4,698.86***	4,548.92***
	(513.39)	(509.72)
Apprenticeship	-2,768.19***	-2,863.31***
	(772.26)	(768.32)
Treated*Year2018	1,692.31**	1,885.72**
	(809.07)	(805.61)
Treated*Apprenticeship	1,909.24*	1,762.70
	(1,152.01)	(1,145.16)
Apprenticeship*Year2018	-168.15	147.27
	(1,132.85)	(1,126.758)
Treated*Apprenticeship*Year2018	-3,988.457**	-4,290.36**
	(1,732.25)	(1,720.61)
Intercept	17,964.78***	
	(351.97)	
Region-FE	No	Yes
R-squared	0.0838	0.1088
R-squared Adj.	0.0815	0.1012
F-statistic	36.716	14.215
N. of Observations	2819	2819

Triple difference estimation of the impact of internship reform on the earnings of apprenticeships. The dependent variable is the earnings of all workers included in the estimation, who are respondents with apprenticeships and open-ended contracts (*tempo indeterminato*) aged 16-30. Treatment is defined as belonging to the set of regions that implemented the reform, taking a value of 1 if the reform was implemented in 2017, and 0 otherwise. Column I does not control for regional fixed effects, while column II does. The data is taken from INAPP's PLUS survey, waves 2016 and 2018. Standard errors are in parentheses. Stars indicate p-values according to: * $p<0.1$, ** $p<0.05$, *** $p<0.01$

Most importantly, the coefficient on the triple interaction term, describing the effect of treatment on the earnings of apprenticeships only in the treated Regions, is negative and significant at the 5% level. Additionally, it is very similar to the previous estimations which only included apprentices in the sample. Therefore, this triple difference exercise, allowing to compare the effect of the reform on apprenticeships against the larger backdrop of all open-ended contracts, confirms that there is a rather large overall earnings rise between 2016 and 2018. It is true for all contracts, but the impact of the internship reform affects apprenticeship earnings negatively for about 2800 euros.

3.3.6 Disentangling Selection and Dumping Effects

So far the estimation has identified a significantly negative overall effect: the reform that extends the use of internships reduces the wages of apprenticeships. What we can deduce from this, however, is simply that the dumping effect is larger than the selection effect, which instead would push the apprenticeship wage upwards. The magnitudes of these two effects are so far unknown, so I now attempt to disentangle the two and estimate their approximate magnitudes.

First of all it is worth questioning whether the selection effect is at all present. Figure 3.1 displays the frequency histograms showing the distribution of earnings for respondents in apprenticeships in 2016 and again in 2018, in the Regions that have adopted the new internship guidelines in 2017. As the comparison between the two pictures suggests, in the change from 2016 to 2018 the distribution has become less skewed and more symmetrically centered around the median. The contracts located at the lower end of the 2016 distribution seem to have indeed been the main ones to be removed in the shift to 2018. This gives reason to believe that indeed a selection effect might be at play.

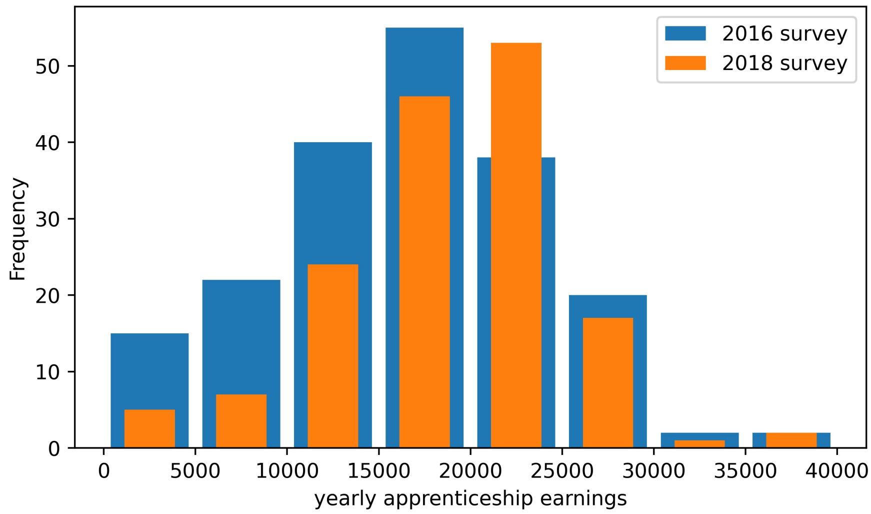

Figure 3.1: Distribution of yearly earnings in treated Regions for individuals working under an apprenticeship contract in 2016 and 2018. 2016 observations are represented in the blue wider columns, while 2018 observations are represented in the orange narrower columns. Data: INAPP's PLUS survey.

With this premise, substantiating the presence of a selection effect, I can design a counterfactual scenario in which only the selection effect were present. I simulate the distribution of earnings under a normal probability density function, with moments reflecting the actual 2016 distribution: a mean of 16474.91 euros and a standard deviation of 7467.40 euros.

Then, as mentioned at the beginning of this section, I focus on the differential growth in apprenticeship numbers among Regions that adopted the internship reform in 2017 and the late adopter Regions. If the 2017-adopters had witnessed the same growth rate of apprenticeships as late-adopters, they would have seen a further rise in apprenticeships of about 5.8% the 2016 value. Thus, for the sake of the counterfactual, I assume that 5.8% is the share of apprenticeships that were "lost" to internships due to the reform.

As the selection effect suggests and the previous figures confirm, I assume that the apprenticeships lost are indeed those located in the left tail of the distribution. Thus, from the probability density function constructed, I remove the bottom 5.8% and, focusing only on the remaining apprenticeship contracts, I calculate the resulting mean earnings. This would be the counterfactual value of the average apprenticeship earnings if only selection effects were present.

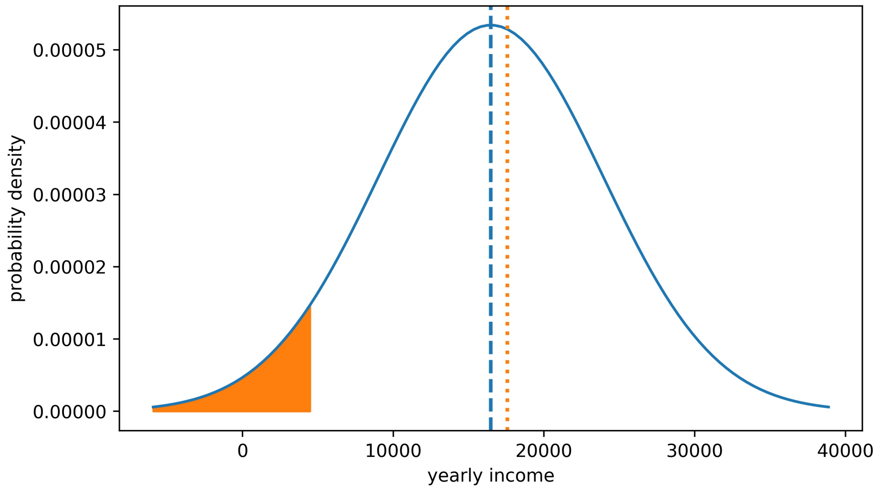

Figure 3.2: Probability density function showing the impact of the selection effect. The blue density function shows the normal distribution of earnings using the mean and standard deviation of the 2016 earnings distribution for apprenticeships. The blue dashed vertical line shows the actual mean earnings. The orange colored area shows the contracts that are lost if the bottom 5.8% of contracts are removed. The orange dotted vertical line shows the resulting mean of the remaining contracts after the orange area is removed.

Figure 3.2 shows this exercise graphically. The blue solid line shows the probability density function and the blue vertical dashed line shows the actual 2016 mean earnings of 16474.91 euros. The orange shaded area, instead, represents the bottom 5.8% of the distribution that is lost to internships and, focusing only on the area under the curve to the right of the shaded area, I calculate a counterfactual average earnings of 17574.95 euros, represented as the orange dotted vertical line. Thus, I can calculate that the selection effect has a magnitude of 1100.04, i.e. if no dumping were present then as a result of the reform the average apprenticeship wage should increase by approximately 1100 euros.

From this, I can indirectly calculate the magnitude of the dumping effect. A quick way to do it would be to rely on the estimation result, which suggests a total effect of -2506 euros. Thus the dumping effect would be the estimated treatment effect plus the 1100 euros of the selection effect: $-2506 - 1100 = -3506$ euros. This exercise suggests that the dumping effect has a magnitude about three times larger than the selection effect, and that therefore the prevailing force should indeed be the dumping one, as is the case from the estimations.

However, I also can use the calculated magnitude of the selection effect and the average earnings growth to calculate the magnitude of the dumping effect without using the estimated total effect. This way I can use the exercise as a test for the validity of the estimation results. First of all I need to add to the selection effect the rise in overall earnings between 2016 and 2018, which is 4532.18 euros. This gives a counterfactual average earnings of 22107.13 euros. Then, from this value I subtract the actual observed average earnings in the 2018 survey, which is of 18575.01 euros, and get a resulting dumping effect of magnitude of -3532.12 euros.

Now I can compare this result with the actual total effect estimated previously. By summing up the two effects calculated using the counterfactual (-3532.12 for the dumping effect and 1100.04 for the selection effect), I get a total effect of -2432.08, which is very close to the coefficient -2506.06 estimated in Table 3.1. This gives some reassurance to the soundness of these exercises, suggesting that the magnitudes of these effects are in the correct ballpark. Thus, following the Italian reform that expanded the use of internships, the dumping effect on apprenticeship earnings has been about three times larger than the selection effect, making the overall result a reduction in the average earnings for apprenticeships.

3.4 Robustness Checks

In this section, I test the robustness of the results obtained in the previous section, first by reducing the number of Regions in the control group, then by including regular open-ended contracts in the analysis to conduct a triple-difference estimation, and finally by using an alternative specification where the year trend is removed and only dumping and selection changes are at play.

3.4.1 Excluding Regions treated in 2018

As mentioned in the institutional setup, the implementation of the 2017 guidelines on internships by the various Regions of Italy took a staggered approach. The treated group for the estimation consists of the Regions that implemented the reform in 2017, while the control group consists of the Regions that implemented the reform in 2018 or in 2019, or did not implement it at all. The reason for the inclusion of 2018 in the control group is that 2018 legislation occurred throughout the year, mostly late in the year, often with a provision to enter into effect at a later date to give workers and firms a chance to adapt to the new regulations. Therefore, most sampled respondents would have been hired under the previous guidelines.

However, a reasonable objection might be raised against the validity of this approach: as long as there is the possibility that some workers in 2018-adopter Regions might have been sampled following the new guidelines, then the estimation might be flawed and suffer from this dampening bias.

As a result, I run the estimation again excluding the workers in the Regions that implemented the new legislation in 2018 from the sample. This way, the control group is made up of only 5 Regions, while the treated group remains made up of the same 10 Regions. The estimation procedure is the same of the previous section, and the results are listed in Table 3.4.

As the table shows, there is still a negative overall effect of the interaction variable, if anything more strongly negative than in the main specification. However, the overall average earnings change from 2016 to 2018 is also larger in magnitude, so probably the reason for this larger effect is simply a concentration in the remaining control Regions of a more substantial earnings increase. The excluded Regions, who implemented the legislation in 2018, had likely a dampening effect on the overall control group.

Taken at face value, therefore, these results seem to suggest indeed that the presence of the 2018-adopter Regions in the main estimation adds noise to the data, as it reduces the impact of the reform on apprenticeship earnings. Perhaps there are indeed some sampled workers in these regions for whom the effects of the reform were already in place. However, the overall effect is not dramatically different and in any case it confirms the finding of a negative impact of the reform on apprenticeship earnings.

3.4.2 Detrended Estimation

Finally, as a further test of the estimation and in particular trying to do away with the confounding trend of rising earnings, I perform an alternative estimation using a simpler specification. To do this, I take out the earnings growth trend using the data on open-ended contracts from the PLUS survey, subtracting the yearly effect from 2018 earnings. Then, I run the regression as in the main estimation, including the "year2018" variable just to make sure that the effect has been correctly removed.

The results of this exercise are shown in Table 3.5. As the Table shows, the yearly trend is insignificant and therefore has been removed correctly. The coefficient on

Table 3.4: Excluding 2018-adopter Regions

	Apprenticeship earnings I	Apprenticeship earnings II
Treated	2,138.82**	
	(978.45)	
Year2018	5,066.83***	5,164.90***
	(1,181.05)	(1,214.35)
Treated*Year2018	-2,891.38**	-3,138.78**
	(1,428.08)	(1,470.84)
Intercept	14,270.49***	
	(823.76)	
Region-FE	No	Yes
R-squared	0.0534	0.0673
R-squared Adj.	0.0473	0.0368
F-statistic	8.831	2.204
N. of Observations	474	474

Difference-in-differences estimation of the impact of internship reform on the earnings of apprenticeships. Treatment is defined as belonging to the set of regions that implemented the reform, taking a value of 1 if the reform was implemented in 2017, and 0 otherwise. Differently from the main estimation, this estimation excludes regions that implemented the reform in 2018 from the regression. Column I does not control for regional fixed effects, while column II does. The data is taken from INAPP's PLUS survey, waves 2016 and 2018. Standard errors are in parentheses. Stars indicate p-values according to: * $p<0.1$, ** $p<0.05$, *** $p<0.01$

Table 3.5: Detrended Estimation

	Detrended earnings I	Detrended earnings II
Treated	1,572.24**	
	(689.60)	
Year2018	33.23	62.38
	(665.98)	(670.83)
Treated*Year2018	-2,146.57**	-2,303.28**
	(1,032.25)	(1,050.64)
Intercept	14,837.07***	
	(454.44)	
Region-FE	No	Yes
R-squared	0.0112	0.0382
R-squared Adj.	0.0074	0.0127
F-statistic	2.918	1.500
N. of Observations	777	777

Estimation of the impact of internship reform on the earnings of apprenticeships, using an alternative specification that focuses on relative wage growth. The dependent variable is constructed as the apprenticeship earnings detrended by taking out the yearly trend of open-ended contracts wage growth. Treatment is defined as belonging to the set of regions that implemented the reform, taking a value of 1 if the reform was implemented in 2017, and 0 otherwise. The set of workers considered is only those respondents with apprenticeships, data on open-ended contracts is used only to determine average wage growth before detrending. Column I does not control for regional fixed effects, while column II does. The data is taken from INAPP's PLUS survey, waves 2016 and 2018. Standard errors are in parentheses. Stars indicate p-values according to: * $p<0.1$, ** $p<0.05$, *** $p<0.01$

treated Regions shows slightly higher earnings for apprentices in those Regions. Finally the interaction coefficient, the main one of interest, is similar to the main specifications in the estimation section.

This confirms the robustness of the empirical findings of this paper, and thus suggests that the 2017 reform that expanded the use of internships in Italy indeed did have a significantly negative effect on the wages of the more stable entry-level contract – apprenticeships.

3.5 Conclusion and Future Research

In the aftermath of the Covid-19 pandemic, the topic of youth employment and contract quality has gained relevance once again, as lockdowns hurt younger workers more heavily and companies face uncertainty in their economic prospects during the recovery, increasing their reliance on temporary contracts. In this paper, I have examined the effect on young individuals' wages and contract quality as firms and workers have different preferences for contracts.

I have developed a simple model of an economy with temporary and permanent contracts, where heterogeneous workers and identical firms bargain on contract type and wages in a search and matching environment. As the legislative constraint on the share of temporary contracts is relaxed, the model predicts two competing effects on the wages of permanent contracts: on the one hand, through a selection effect, firms move workers on the lower-end of the permanent-contract distribution towards newly available temporary contracts, thus mechanically increasing the average wage of the remaining permanent contracts; on the other hand, workers will be willing to lower the wage in exchange for a permanent contract, leading to a dumping effect which should lower the average wage of permanent contracts.

I test these predictions using Italian survey data before and after a 2017 reform that liberalized the use of internships, increasing the number of internships that firms can offer to potential new hires. Taking advantage of a staggered implementation schedule by Italian Regions, I am able to estimate an overall effect of the reform on apprenticeship earnings, which is significantly negative in treated Regions compared to untreated ones, suggesting that the dumping effect is prevalent. Using a counterfactual scenario, I estimate that the selection effect has a magnitude of about 1100 euros of yearly earnings, while the dumping effect is about three times larger and of the opposite sign. Thus, the dumping effect is prevalent in the overall impact on average wages.

Despite the simple model and the limitations in the data's size, this paper sheds some light on topical dynamics of labour market entry, for a demographic that is most severely impacted by precariousness. The discussion of whether and how to limit temporary contracts is ongoing in many European countries. Spain, for instance, just passed legislation drastically reducing the possible instances of temporary contract offer, and a few countries are considering tightening their legislation as well. In Italy, many parties leading up to the 2022 election cite labour market precariousness as a top concern of the campaign. Italy additionally has numerous legislative precedents of changes to temporary contract restrictions: from the liberalizations of the so-called "Legge Biagi" in the early 2000s to the so-called "Decreto Dignità" in the last few years.

In all these instances, ideally using administrative data, a similar analysis to the one conducted in this paper can be performed to estimate the effect of these legislative changes. At the same time, the model can be improved and expanded to make more and better predictions on numerous additional variables, including the wages of temporary workers, the effect on on-the-job search, etc.

The focus of research exploring the impact of precariousness on labour market entry can also be extended to other avenues of research, including long-term effects of early-career insecurity. The literature has analyzed the effects of job instability mainly with respect to its repercussions on productivity, on worker searching behavior, and on other short-term variables. Little attention has been paid to the long-term effect of entering in the labor-market with more precarious contracts, not only on labour market outcomes but also on social aspects such as fertility rates and home independence from parents.

Part of the reason for this absence is the lack of accurate long-term longitudinal data that keeps track of different types of contracts, and in particular of data that concerns internships, which makes it hard to follow individuals as they enter precariously into the labor market. Through an effort to collect and make more readily available for research more and better labour market data, the research of this paper could be extended to hopefully provide some new and useful insights for both researchers and policymakers about the longer-term effects of early career instability.

A Skills and sectors differences

In many countries, the kinds of jobs that are offered in apprenticeships or internships is not quite overlapping: the former is used more in manufacturing or in manual jobs, usually for lower-skilled workers; the latter is used more in the services sector, for higher-skilled workers.

In Italy, however, the situation is less different in the 2010s a large set of new legislation and government incentives was passed with the aim to extend the use of the apprenticeship beyond the more traditional sectors and to all qualification levels. Even for PhDs and other tertiary degrees a special apprenticeship was designed, called third-degree apprenticeship. Following these changes, all collective contracts were urged to include a specific discipline on apprenticeships and as a result most if not all sectors' contracts in Italy include a specific set of rules and incentives for hiring apprentices.

Table 6 shows the difference in the distribution of workers between the internship and the apprenticeship in Italy. The data for sectors is taken from administrative sources, from the Ministry of Labour and the ANPAL agency for active labour market policies. Their monitoring reports include detailed analyses of young individuals in apprenticeships and internships and are released periodically. As they are released separately, however, the aggregation of sectors is not consistent across reports. While all sectors are divided using the Italian classification of Ateco codes, the reports present an aggregation of different Ateco codes, which however are not consistent across reports. For a consistent comparison of the two, therefore, I have to aggregate the data into broader sector categories.

For the education levels of workers in these contracts, instead, the internship report is quite detailed, while no education-level data is published in the appren-

Table 6: Education levels and sectors of workers

	Apprenticeship	Internship
Sectors		
Agriculture	NA	1.3
Manufacturing	20.3	18.6
Construction	7.1	3.8
Commerce, service and professionals	61.4	57.8
Other services/Not classified	11.2	18.5
Education levels		
Lower secondary	4.9	1.7
Upper secondary	48.7	42.4
Tertiary	45.1	54.6

Difference is skill levels and sectors between young individuals in apprenticeships and in internships. Data for the skill levels is taken from the INAPP PLUS' sample used in the estimation of this paper. Data for the sectors is taken from ANPAL and Italy's Ministry of Labour monitoring reports of internships and apprenticeships. All values are expressed in percentage points.

ticeship report. As a result, in the Table I use the data from INAPP's PLUS survey, using the same samples used in the estimation. The data indicates the highest level of education attained. All values in the Table are expressed in percentage points. Since some observations did not report their education level, totals might not add up to 100 exactly.

Overall, the table shows that the education levels of workers and sectors of employment are mostly comparable across internships and apprenticeships. While they are not exactly the same, they are similar enough to suggest that there might indeed be some degree of substitutability between the two instruments, allowing for the dumping effect highlighted in the estimation section.

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